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The Rise in Inequality, the Decline in the Natural Interest Rate, and the Increase in Household Debt*

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I investigate the effect of rising income inequality on the natural rate of interest in an economy with "rich" households who have "capitalist spirit" type preferences over their wealth, "non-rich" households, and housing and credit markets. Simulating the increase in U.S. income inequality over the 1981–2016 period generates a downward trend in the natural rate in line with recent empirical estimates. The model also broadly captures the upward trend in the debt-to-income ratio and loan-to-value ratio of the bottom 90 percent of households, as well as the upward trend in the value of the housing stock during this period.

JEL Codes: E25, E52, E43, D14.

1. Introduction

The objective of this paper is to examine to what extent rising income inequality can explain the downward trend in (estimates of) the natural rate of interest, i.e., the real interest rate consistent with a closed output gap and stable inflation, in the United States since the 1980s. Figure 1 (first panel) suggests that the decline in the

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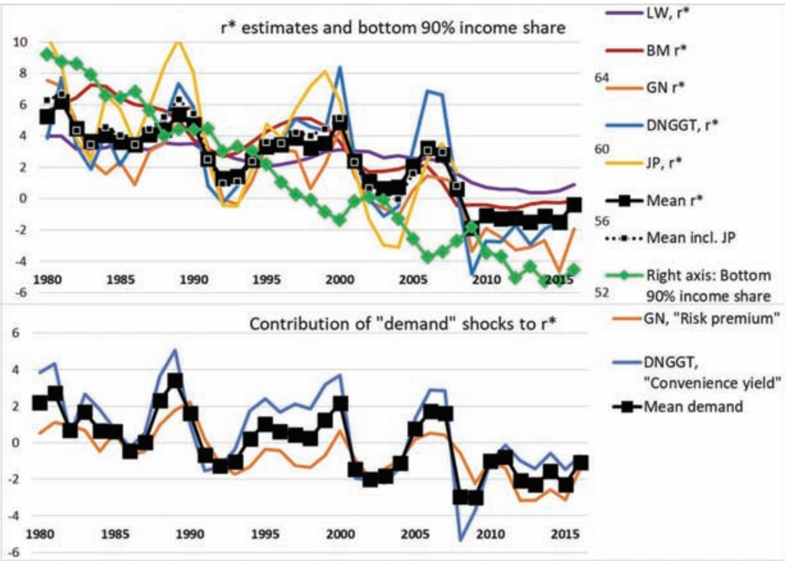
Figure 1. The Trend in Income Inequality and Estimates of the Natural Rate



Note: The r* values are annual averages of quarterly values. LW: Laubach and Williams (2016), updated estimates downloaded from the Federal Reserve Bank of New York webpage. BM: Brand and Mazelis (2019) modify the Laubach and Williams (2016) approach by, inter alia, closing the model with an interest feedback role and including the federal funds rate as an observable, using the unemployment rate to discipline the output gap estimate and the use of Bayesian estimation techniques. GN: Gerali and Neri (2019), DSGE model, r*: interest rate in the flexible-price economy, "Risk premium" shock: Increases the demand for government bonds at the expense of consumption and physical capital, first introduced by Smets and Wouters (2007). DNGGT: Del Negro et al. (2017), DSGE model, r*: interest rate in the flexible-price economy. "Convenience yield" shock: This shock has qualitatively the same impact on consumption and the demand for physical capital as the GN risk premium shock, but the estimation uses additional observables to identify the shock, and the estimated model differs in some respects from Gerali and Neri (2019). JP: Justiniano and Primiceri (2010), DSGE model, r*: interest rate in the flexible-price economy. The authors do not report a historical decomposition of r*. Mean r*: Unweighted average across all estimates except JP, as the latter is available only during 1980–2008. Bottom 90 percent income share: Share of the bottom 90 percent of households in net national income, World Inequality Database (WID); see Alvaredo et al. (2020).

natural rate has been large.[1] It also shows that the income share of the bottom 90 percent of U.S. households declined by about 12 percent from 1980 to 2016 (see Figure 1, first panel, the green diamond line; Alvaredo et al. 2020; and Piketty, Saez, and Zucman 2018). These two coinciding trends raise the possibility that the redistribution of income towards income-rich households might have increased the aggregate supply of savings and thus depressed the natural rate, as argued by Summers (2014) and Rachel and Smith (2017).

Another finding consistent with this conjecture is the major role of aggregate demand shocks in the downward trend of the estimates of Del Negro et al. (2017) and Gerali and Neri (2019) (see Figure 1, the bottom panel), which are based on estimated dynamic stochastic general equilibrium (DSGE) models.

I formalize the mechanism suggested by Summers (2014) and Rachel and Smith (2017) in a model with two distinct groups of households. One group represents the top 10 percent of the income distribution (referred to as "the rich"), and the other the remainder. Crucially, the rich derive utility from their wealth, i.e., they have "capitalist spirit type preferences" (CSP), a motive first suggested by Weber (1958), implying that they will increase their saving in response to a permanent income increase, in line with the evidence of Dynan, Skinner, and Zeldes (2004). In the model, income inequality may rise due to higher wage dispersion or an increase in the price markup of firms owned by rich households.

I first show that in an economy where the non-rich do not borrow, with CSP, the natural rate declines strongly in response to a decline in the bottom 90 percent income share. The non-rich lower their consumption by the amount of their income decline, while at the initial interest rate, the rich attempt to save part of the increase in their permanent income. Thus the interest rate needs to decline to equilibrate the government bond market. By contrast, without CSP, rich households do not attempt to save in response to the inequality increase and thus the interest rate remains unchanged.

[1]Moreover, Rachel and Summers (2019) argue that available estimates tend to mask an even more dramatic decline in the "private sector" natural rate of about 7 percent across advanced economies since the 1970s, which was partially offset by the expansionary effects of the simultaneous increase in government debt, as well as the obligations implied by the presence of pay-as-you-go pension systems and government-funded health care.

I then extend the model by allowing for homeownership, a credit market subject to frictions where the non-rich borrow from the rich via financial intermediaries using their home as collateral, and physical capital as an additional factor of production owned by the rich. In this setup, in the presence of CSP, an increase in inequality lowers the natural rate and also increases borrowing, as the non-rich use the decline in their cost of credit to postpone the decline in their consumption and housing demand. Furthermore, the lower interest rate increases the relative housing demand of both rich and non-rich households, thus increasing house prices. The increase in house prices becomes stronger if I assume that non-rich housing demand is subject to a "consumption cascade," meaning it depends positively on the total consumption of the rich, in line with the evidence of Bertrand and Morse (2016).

I then replicate the decline in the bottom 90 percent of wage earners' share in labor income and the decline in the labor share over the 1980–2016 period using the aforementioned wage dispersion and price markup shocks. The simulated increase in inequality generates a decline in the natural interest rate of between 3 and 4 percentage points, in line with what the aforementioned empirical estimates tend to attribute to "aggregate demand" type shocks. The simulation also broadly captures the upward trend in the debt-to-income ratio and loan-to-value ratio (LTV) of the bottom 90 percent of households observed over the 1983–2007 period. This scenario of an increase in bottom 90 percent indebtedness funded by saving on the part of the top 10 percent, which is in turn fueled by an increase in their income share, is consistent with the empirical findings of Mian, Straub, and Sufi (2020b) of a "saving glut of the rich." Furthermore, the simulation matches much of the rising trend in the value of the housing stock relative to gross domestic product (GDP) once non-rich housing demand is subject to a consumption cascade. Finally, the markup increase ensures that the simulation matches the empirically observed absence of a downward trend of measures in the real return on business capital noted by Caballero, Farhi, and Gourinchas (2017), Eggertsson, Robbins, and Wold (2018), and Farhi and Gourio (2018).

There is an evolving literature modeling a link between the increase in inequality and the decline in the natural rate, to which my analysis contributes as follows. I study a permanent increase in income inequality, while in the models used in most existing contributions, only transitory increases in inequality cause a decline in the natural rate. For instance, Eggertsson, Mehrotra, and Robbins (2019) show that an increase in inequality within the middle generation of a three-generation overlapping generations (OLG) model with credit constraints may reduce the natural rate.[2] Furthermore, in the heterogeneous agent models of Auclert and Rognlie (2018) and Rachel and Summers (2019), an increase in inequality driven by higher income uncertainty increases precautionary saving and thus lowers the interest rate. Contributions in this literature attribute between 0.8 and 1 percentage point of the decline in the natural rate to the increase in income inequality. However, Kopczuk, Saez, and Song (2010) and DeBacker et al. (2013) provide evidence that increases in permanent (not transitory) earnings variance drove the increase in inequality observed in recent decades in the United States. Furthermore, Kopczuk, Saez, and Song (2010) report that short- and long-term income mobility has been either stable or declining since the 1950s.

By contrast, Straub (2017) and Mian, Straub, and Sufi (2020a) consider permanent increases in income inequality. Straub (2017) analyzes a heterogeneous-agent 65-generation OLG model where all agents have non-homothetic preferences over bequests, which generates a positive relationship between permanent income and the saving rate. When replicating the increase in U.S. labor income inequality since the 1970s, he finds a 1 percentage point decline in the interest rate. However, his model abstracts from borrowing and assets other than physical capital. Perhaps for this reason, his model predicts an increase of the capital-output ratio of about 25 percentage points over the 1980–2016 period, while the ratio of the private non-residential capital stock as estimated by the Bureau of Economic Analysis (BEA) to GDP does not display such a trend. By contrast, my best-performing model version avoids this prediction. More recently, Mian, Straub, and Sufi (2020a) analyze the effect of

[2]Lancastre (2016) extends this approach by adding a bequest motive where agents care about the sum of the bequest and their children's middle-age income, and assumes that parents' and children's "middle-age" income is negatively correlated, which appears at odds with the evidence (Charles and Hurst 2003; Lee and Solon 2009; Bjoerklund and Jaentti 2011).

a permanent increase in inequality within a two-dynasty perpetual youth model with non-homothetic preferences over bequests, and find that it lowers the natural rate and increases the indebtedness of non-rich households. A key difference between my analysis and theirs is that their assumptions regarding the utility from wealth imply a downward-sloping long-run "saving supply curve," i.e., a negative steady-state relationship between the real interest rate and the wealth of the rich. Therefore, an increase in the supply of government bonds decreases the real interest rate, which is at odds with the empirical evidence of Gale and Orszag (2004), Engen and Hubbard (2005), Laubach (2009), and Krishnamurthy and Vissing-Jorgensen (2012). By contrast, in my model an increase in the supply of safe assets increases the real interest rate. Furthermore, Mian, Straub, and Sufi (2020a) do not focus on assessing the ability of the model to match long-run trends, and do not have a housing market.

A link between the increase in inequality and rising indebtedness of non-rich households has been argued by Rajan (2010) and modeled by Kumhof, Ranciere, and Winant (2015) in an endowment economy with CSP and credit to the non-rich as the *only* asset. My analysis goes beyond Kumhof, Ranciere, and Winant (2015)'s on various dimensions. Firstly, they do not analyze the effect of income inequality on the natural rate of interest.[3] Secondly, I find that Kumhof, Ranciere, and Winant (2015)'s prediction of an empirically relevant increase in non-rich debt is robust to allowing rich households to invest in assets *other than* non-rich debt. Finally, unlike their endowment economy, my model distinguishes between different sources of rising income inequality, namely wage dispersion and price markups.

Other contributions have investigated the influence of demographic shifts like the decline in population growth and the increase in life expectancy on the natural rate (e.g., Bielecki, Brzoza-Brzezina, and Kolasa 2018; Eggertsson, Mehrotra, and Robbins 2019; and Papetti 2019; and Brand, Bielecki, and Penalver 2018 for a survey). However, Mian, Straub, and Sufi (2021) cast doubt on the aspect of the demography hypothesis, emphasizing the role of

[3]Their model has no safe interest rate, and the risky interest rate paid by borrowers does not display a trend when they simulate the empirical inequality increase over the 1983–2008 period.

saving by the baby-boom generation entering working age by showing that saving rates vary little across the age distribution (while varying substantially across the income distribution). Furthermore, they document that the increase in life expectancy was associated with declining saving rates in the bottom 90 percent of the income distribution.

Zou (1994, 1995) shows that a long line of economists—including but not limited to John M. Keynes, Max Weber, Joseph Schumpeter, Karl Marx, and Adam Smith—have argued that entrepreneurs' wealth accumulation is driven by CSP. To my knowledge, the first formal macroeconomic application of CSP occurs in the context of growth models, in order to shed light on the large differences in growth rates across countries and the lack of convergence in per capita income levels. Following a technical note by Kurz (1968), Zou (1994) investigates the effect of CSP on growth in a deterministic context, while Bakshi and Chen (1996) and Smith (1999) consider stochastic environments. Cole, Mailath, and Postlewaite (1992) develop a microfoundation of CSP in an "AK" type model with status-dependent mating. Furthermore, a concept related to CSP, so-called preferences over safe assets, has been shown recently to considerably alleviate the so-called forward-guidance puzzle, i.e., the finding that in DSGE models, the effect of forward guidance is implausibly strong (e.g., Michaillat and Saez 2018; Rannenberg 2019, 2021).

Moreover, CSP has been shown to address certain counterfactual predictions of optimizing models at the micro level. Firstly, in a life-cycle context, CSP avoids the prediction that following retirement, old people decumulate assets and that their saving rates are lower (see Zou 1995 and Carroll 2000),[4] and allows to match the large wealth-to-income ratio (see Carroll 2000, Reiter 2004, and Francis 2008) and larger saving rates (Dynan, Skinner, and Zeldes 2004) of income-rich households. Secondly, models with CSP can match the empirically estimated higher marginal propensity to save out of permanent income of rich households (see Dynan, Skinner, and Zeldes 2004; Kumhof, Ranciere, and Winant 2015). Thirdly, Gechert and

[4]Relatedly, Zou (1995) and Carroll (2000) argue that a pure bequest motive (as opposed to utility from wealth at all ages) is not supported by the evidence because the childless old do not seem to dis-save faster than those with children.

Siebert (2022) provide evidence for utility from wealth in a multi-period experiment where half of the participants choose to maintain a stock of wealth even though spending all cash on hand immediately would maximize the participants' final payout.

Finally, my modeling approach forms part of a literature analyzing the macroeconomic consequences of household heterogeneity by dividing households into distinct groups which differ regarding important characteristics—for instance, their consumption smoothing opportunities, asset holdings, or impatience (see Iacoviello 2005; Gali, Lopez-Salido, and Valles 2007; Bilbiie 2008, 2020; Debortoli and Gali 2017; Sterk and Ravn 2017; and Broer et al. 2020).

The remainder of the paper is structured as follows. Section 2 develops and analyzes a stylized model. Section 3 develops the model with household borrowing and a housing market, which I refer to as the "full model." Section 4 discusses the effects of an increase in inequality in the full model, and the simulation of the empirically observed increase in inequality.

2. A Simple Model

The model features two distinct household groups, namely a mass of n_S rich and a mass of n_{CC} non-rich households, as well as monopolistically competitive firms owned by rich households and employing rich and non-rich household labor. The model thus precludes the possibility that the observed increase in income inequality might be the consequence of individual households' greater income mobility between different income groups. However, Kopczuk, Saez, and Song (2010) and DeBacker et al. (2013) provide evidence that increases in permanent (not transitory) earnings variance drove the increase in inequality observed in recent decades. Furthermore, Kopczuk, Saez, and Song (2010) report that short- and long-term income mobility has been either stable or declining since the 1950s. Throughout, $\tilde{X}_S$ (or $\tilde{x}_S$) denotes a rich household choice variable expressed in per capita terms, while the corresponding economy-wide variable is denoted as X_S (or x_S), and is computed as $X_S = \tilde{X}_S n_S$, *unless otherwise mentioned* (and analogously for non-rich households).

2.1 Households

Throughout, I index rich households with the subscript S. Rich households derive utility from (per capita) consumption $\tilde{C}_{S,t}$, and their stocks of safe real financial assets $\tilde{b}_{S,t}$ (consisting of government bonds). Their objective function is thus given by

$$E_t \left\{ \sum_{i=0}^{\infty} \beta_S^i \left[\frac{\tilde{C}_{S,t+i}^{1-\sigma_S}}{1-\sigma_S} + \frac{\tilde{\phi}_b}{1-\sigma_b} \left(\tilde{b}_{S,t+i} \right)^{1-\sigma_b} \right] \right\},$$

where β_S denotes their utility discount factor, and $\tilde{\phi}_b$, σ_S, and σ_b are non-negative constants. A rich household's budget constraint is given by

$$\tilde{b}_{S,t} = \frac{R_{t-1}}{\Pi_t} \tilde{b}_{S,t-1} + w_{S,t} \tilde{N}_{S,t} + \tilde{\Xi}_t - \tilde{T}_{S,t} - \tilde{C}_{S,t},$$

where R_t, $w_{S,t}$, $\tilde{\Xi}_t$, $\tilde{T}_{S,t}$, and Π_t denote the nominal interest rate on safe assets, the real wage, the real profits of intermediate goods firms, real lump-sum taxes, and the inflation rate, respectively. The assumption that only the rich own firms and government bonds is motivated by the extreme concentration of stocks, business ownership, and bonds (e.g., Kuhn and Rios-Rull 2016). From now on, I will refer to the model where the rich derive utility from their wealth (i.e., $\phi_b > 0$) on top of consumption as the model with "capitalist spirit" type preferences (CSP), while I refer to the $\tilde{\phi}_b = 0$ case as NOCSP. The first-order conditions (FOCs) with respect to consumption and government bonds are given by

$$\Lambda_{S,t} = \frac{1}{C_{S,t}^{\sigma_S}} \tag{1}$$

$$\Lambda_{S,t} = \beta_S E_t \left\{ \Lambda_{S,t+1} \frac{R_t}{\Pi_{t+1}} \right\} + \phi_b \left(b_{S,t} \right)^{-\sigma_b}, \tag{2}$$

where per capita variables have been eliminated using $\tilde{X}_S = \frac{X_S}{n_S}$, $\Lambda_{S,t} \equiv \frac{\tilde{\Lambda}_{S,t}}{n_S^{\sigma_S}}$ denotes the (scaled) marginal utility of consumption, and $\phi_b \equiv \frac{\tilde{\phi}_b n_S^{\sigma_b}}{n_S^{\sigma_S}}$.[5] If $\phi_b > 0$, $\phi_b \left(b_{S,t}\right)^{-\sigma_b}$ represents an extra marginal benefit from saving over and above the utility associated with the future consumption opportunity saving entails (represented by $\beta_S E_t \left\{\Lambda_{S,t+1} \frac{R_t}{\Pi_{t+1}}\right\}$). CSP weakens the effect of an increase in permanent income and thus a decline in $\Lambda_{S,t+1}$ on $\Lambda_{S,t}$, since the two become less than proportional. To gain some intuition, compare the bond market equilibrium in the CSP and NOCSP case, assuming that the economy is initially in the steady state in both period t and $t+1$. The presence of the extra benefit $\phi_b \left(b_{S,t}\right)^{-\sigma_b}$ with CSP implies that, for the bond market to clear, the present value $\beta_S \frac{R_t}{\Pi_{t+1}}$ which the household attaches to $\Lambda_{S,t+1}$—the net effect of the reward of waiting and the household's impatience—has to be smaller than in the NOCSP case, thus reducing the importance it attaches to a decline in $\Lambda_{S,t+1}$. Furthermore, this weakening of intertemporal consumption smoothing compounds the more distant in time the anticipated future consumption increase is located, as $\Lambda_{S,t+1}$ is no longer proportional to $\Lambda_{S,t+2}$ either, and so on and so forth. As a result, with CSP a 1 percent permanent increase in saver household income will ceteris paribus not cause a 1 percent increase in consumption, but instead an increase in both saving and consumption. The marginal propensity to save out of a permanent income increase will be larger the smaller the curvature parameter σ_b. Relatedly, the above implies that for $\phi_b > 0$,

$$R_t < \frac{1}{E_t \left\{\frac{\beta \Lambda_{S,t+1}}{\Lambda_{S,t}\Pi_{t+1}}\right\}} \equiv DIS_t, \tag{3}$$

i.e., the nominal interest rate may be smaller than the discount rate which the household applies to future income streams DIS_t.

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[5]The FOCs with regard to per capita consumption and bonds are given by $\tilde{\Lambda}_{S,t} = \frac{1}{\left(\tilde{C}_{S,t}\right)^{\sigma_S}}$ and $\tilde{\Lambda}_{S,t} = \beta_S E_t \left\{\tilde{\Lambda}_{S,t+1} \frac{R_t}{\Pi_{t+1}}\right\} + \tilde{\phi}_b \left(\tilde{b}_{S,t}\right)^{-\sigma_b}$, respectively. Substituting $\tilde{C}_{S,t} = \frac{C_{S,t}}{n_S}$ and $\tilde{b}_{S,t} = \frac{b_{S,t}}{n_S}$, and defining $\Lambda_{S,t} \equiv \frac{\tilde{\Lambda}_{S,t}}{n_S^{\sigma_S}}$ and $\phi_b \equiv \frac{\tilde{\phi}_b n_S^{\sigma_b}}{n_S^{\sigma_S}}$ yields Equations (1) and (2).

I assume that non-rich households, denoted as CC, simply consume their disposable income. Their behavior is thus described by

$$C_{CC,t} = w_{CC,t} N_{CC,t} - T_{CC,t}. \tag{4}$$

Households are endowed with a fixed amount of hours $\tilde{N}_S$ and $\tilde{N}_{CC}$ which they supply to firms, implying that

$$N_{CC,t} = n_{CC} \tilde{N}_{CC} \tag{5}$$

$$N_{S,t} = n_S \tilde{N}_S. \tag{6}$$

2.2 Firms

A continuum of perfectly competitive final goods firms produces final consumption and investment goods by aggregating a continuum of differentiated intermediate goods using a constant elasticity of substitution (CES) technology. The differentiated goods are supplied by monopolistically competitive intermediate goods firms. Intermediate goods firm j combines the labor supplied by the two household types using a Cobb-Douglas technology:

$$Y_t(j) = \left(\frac{N(j)_{S,t}}{\eta_S} \right)^{(1-\omega_{CC}-d_{CC,t})} \left(\frac{N(j)_{CC,t}}{\eta_{CC}} \right)^{(\omega_{CC}+d_{CC,t})}, \tag{7}$$

where $d_{CC,t}$ represents a shock to the production elasticity of rich and non-rich households which I will employ to increases in labor income inequality, and $\eta_S, \eta_{CC} > 0$ denote normalizing constants.[6] A negative value of $d_{CC,t}$ lowers the demand for non-rich household labor and thus their real wage, while increasing the demand for rich household labor. The shock can be viewed as a proxy for skill-biased technological change and the "race between education and technology" (Goldin and Katz 2007). Note that under my assumption of

[6] In particular, I assume $\eta_S = N_S$ and $\eta_{CC} = N_{CC}$. Hence since $N_{S,t} = N_S$ and $N_{CC,t} = N_{CC}$ (see Equations (6) and (5)), the employment levels cancel out in the aggregate. This assumption ensures that even for $N_S \neq N_{CC}$, $d_{CC,t}$ does not affect N_t output. It follows Straub (2017), who considers exactly the same type of shock (see his Appendix E, p. 68, first equation).

flexible prices (and thus an exogenous price markup) the shock will not change the overall labor income share. The firm's first-order conditions are given by

$$w_{S,t} = mc_t \left(1 - \omega_{CC} - d_{CC,t}\right) \frac{Y_t}{N_{S,t}} \tag{8}$$

$$w_{CC,t} = mc_t \left(\omega_{CC} + d_{CC,t}\right) \frac{Y_t}{N_{CC,t}} \tag{9}$$

$$\frac{1}{\mu_P + d_{\mu,t}} = mc_t, \tag{10}$$

where μ_P denotes the steady-state markup of prices over marginal costs and $d_{\mu,t}$ a shock to the markup, which I will use to generate increases in inequality which are accompanied by a decline in the labor share.

2.3 Government

There is a government consuming G_t units of output. It levies lump-sum taxes on households in order to keep the government debt-to-GDP ratio and the GDP share of government expenditure constant at $Target_{bgov2GDP}$ and $Target_{G2GDP}$, respectively. For simplicity, I assume that fiscal policy keeps total lump-sum taxes of non-rich households constant. The central bank successfully pursues a perfect inflation target, implying that the actual real interest rate equals the natural rate. Hence policy is described by

$$b_{gov,t} = \frac{R_{t-1}}{\Pi_t} b_{gov,t-1} + G_t - T_{S,t} - T_{CC,t} \tag{11}$$

$$Target_{bgov2GDP} = \frac{b_{gov,t}}{4Y_t} \tag{12}$$

$$Target_{G2GDP} = \frac{G_t}{Y_t} \tag{13}$$

$$T_{CC,t} = T_{CC} \tag{14}$$

$$\Pi_t = \Pi. \tag{15}$$

2.4 Equilibrium

Equilibrium in goods, capital, and labor markets implies

$$Y_t = C_{S,t} + C_{CC,t} + G_t \tag{16}$$

$$b_{S,t} = b_{gov,t}. \tag{17}$$

The only exogenous variables are the shocks to the production elasticity of households $d_{CC,t}$ and the price markup $d_{\mu,t}$.

2.5 Calibration

One time period in the model corresponds to one quarter. I assume a labor endowment $\tilde{N}_{CC} = \tilde{N}_S = \frac{1}{3}$ of $\frac{1}{3}$, as well as $\eta_S = N_S$ and $\eta_{CC} = N_{CC}$. I assume a price markup μ_p of 1.05 (see Table 1). I calibrate the remaining parameters such that the steady-state values of the model match the empirical targets reported in Table 2. In the model without CSP (i.e., where $\phi_b = 0$), there are in total five parameters calibrated in this fashion, marked with an asterisk (*), namely the rich consumption utility curvature σ_S, the rich household discount factor β_S, the non-rich share in labor income ω_{CC}, and the government expenditure and debt targets $Target_{bgov2GDP}$ and $Target_{G2GDP}$. The empirical targets are the intertemporal elasticity of substitution, which I set in line with the mean estimate reported in the meta-analysis of Havranek (2015), the real ex post federal funds rate, the GDP share of government expenditure, the government debt-to-GDP ratio, and the income share of the bottom 90 percent of households, which I assume to be the real-world counterparts of the non-rich in the model. I compute all targets as averages over the 1973–80 period, as the historical simulation of Section 4.2 starts in 1981 (the bottom 90 percent income share is essentially constant during 1973–80), with the exception of the government debt-to-GDP ratio and the government expenditure share. These targets I compute as averages over the 1981–2016 period since I hold these variables constant throughout the paper. Finally, I set the share of the non-rich in the total tax burden equal to their pre-tax income share.

In the model with CSP, I calibrate the two CSP-related parameters (ϕ_b, σ_b) by using two additional empirical targets. The first target is an estimate of the "discounting wedge" θ_t, defined as $\theta_t \equiv \frac{R_t}{DIS_t}$, where DIS_t denotes the nominal individual discount

Table 1. Parameters: Simple Model

Parameter	Parameter Name	Value NOCSP ($\theta = 1$)	Value CSP ($\theta = 0.97$)
β_S	Rich Household Utility Discount Factor	0.9951*	0.9652*
σ_S	Rich Utility Curvature Consumption	2*	2*
$\tilde{N}_{CC}$, $\tilde{N}_S$	Labor Endowments	$\frac{1}{3}$	$\frac{1}{3}$
ω_{CC}	Non-rich Share in Total Labor Income	0.7*	0.7*
μ_P	Price Markup	1.05	1.05
$Target_{bgov2GDP}$	Gov. Debt-to-GDP Ratio	0.44*	0.44*
$Target_{G2GDP}$	Government Expenditure-to-GDP Ratio	0.2*	0.2*
$\frac{T_{CC}}{T_{CC}+T_S}$	Share of the Non-rich in the Total Tax Burden	67%	67%
σ_b	Rich Utility Curvature of Real Financial Assets	—	0.40*
ϕ_b	Rich Utility Weight on Real Financial Assets	0	0.51*

rate which the household applies to future nominal income streams (defined by Equation (3)), with $\theta = \beta_S \frac{R}{\Pi}$. Note that $\theta < 1$ implies a smaller value of β_S than in the NOCSP case (which corresponds to $\theta = 1$), given the unchanged target for the real interest rate. Conditional on an assumption for σ_b, the steady-state relationship implied by the Euler equation (23) allows to back out ϕ_b as

$$\phi_b = (1-\theta)\,\Lambda_S\,(b_S)^{-\sigma_b}\,. \tag{18}$$

I set $\theta = 0.97$, close to the choice of Rannenberg (2019), who obtains evidence on θ by drawing on 34 empirical estimates of the (time-varying) nominal individual discount rate which the household applies to future nominal income streams, $DIS_t = \frac{1}{E_t\left\{\frac{\beta\Lambda_{t+1}}{\Lambda_t\Pi_{t+1}}\right\}}$ (all based on choices resulting in real, rather than hypothetical, money flows).[7] Given estimates of DIS_t, Rannenberg (2019) exploits the fact that for sufficiently small weights on safe assets in the utility function (i.e., θ smaller than but close to one), $\theta_t = \frac{R_t}{DIS_t}$ is approximately constant across time in the model. This property is

[7] See his Table 2, p. 16, and the associated discussion.

Table 2. Targets: Simple Model

Target	Value NOCSP	Value CSP	Source
IES $\frac{1}{\sigma_S}$	0.5	0.5	Havranek (2015)
Real Short-Term Interest Rate $\left(\frac{R}{\Pi}\right)^4 - 1$	2%	2%	Federal Funds Rate Minus Core PCE Inflation, APR, FRED
$\frac{G}{Y}$	20%	20%	Government Expenditure GDP Share, BEA
$\frac{b_{gov,S,}}{4Y}$	44%	44%	Federal Debt Held by the Public, Percentage of GDP, FRED
Non-rich National Income Share NIS_{CC} (see note)	66%	66%	Bottom 90% Net National Income Share, Pre-tax, WID
MPS Top 5%	0	51%	Target CSP Case: Dynan, Skinner, and Zeldes (2004)
Discounting Wedge θ	1.0	0.97	Target CSP Case: Literature Discount Rates; See Discussion in the Text

Note: FRED = Federal Reserve Economic Data (Federal Reserve Bank of St. Louis). BEA = Bureau of Economic Analysis. IES = intertemporal elasticity of substitution. WID = World Inequality Database; see Alvaredo et al. (2020) and Piketty, Saez, and Zucman (2018) for the U.S. data used here for details. The non-rich national income share NIS_{CC} corresponds to the concept used in the WID. In particular, when computing the pre-tax income of an adult, the WID allocates the property income of the government $-b_{gov}\left(\frac{R}{\Pi}-1\right)$ across adults according to the percentile's pre-tax factor income share, which in my model equals the national income share (due to the absence of pensions and transfers). The allocation of the government's property income to individuals is necessary in order to ensure that, across adults, pre-tax income adds up to national income. Hence the national income share of the non-rich households is given by $NIS_{CC} \equiv \frac{w_{CC}N_{CC}-b_{gov}\left(\frac{R}{\Pi}-1\right)NIS_{CC}}{Y}$. Solving for NIS_{CC} yields the expression in the table. $NIS_{CC} = \frac{w_{CC}N_{CC}}{Y\left(1+\frac{b_{gov}}{Y}\right)}$. The targets for $\frac{G}{Y}$ and $\frac{b_{gov}}{4Y}$ are calculated as averages over the 1981–2016 period. All other ratios for which annual data were available were calculated as averages over the 1973–80 period.

a consequence of intertemporal substitution by the household: An increase in R_t shifts consumption from t to $t+1$, thus reducing the marginal utility of future consumption and increasing DIS_t. Hence $\theta \approx \frac{R_t}{DIS_t}$, which given the assumed steady-state value of the real interest rate $\frac{R}{\Pi}$ then allows to pin down β. Among those studies used in Rannenberg (2019), the contributions of Pleeter and Warner (2001), Harrison, Lau, and Williams (2002), and Harrison et al. (2005) are of particular relevance in my context. Harrison, Lau, and Williams (2002) and Harrison et al. (2005) report estimates for (income-) rich households, while Pleeter and Warner (2001) elicit the discount rate of officers of the U.S. armed forces choosing between two severance packages during the 1992–95 military drawdown.[8] The values of θ obtained from these studies are between 0.95 and 0.97.

Finally, following Kumhof, Ranciere, and Winant (2015), the second target that I use to calibrate the CSP is an estimate of the rich households' marginal propensity to save (MPS) out of an increase in their permanent income. Here I draw on the evidence computed by Kumhof, Ranciere, and Winant (2015) based on the saving rate regressions of Dynan, Skinner, and Zeldes (2004) for households in the top 5 percent of the income distribution. Following Kumhof, Ranciere, and Winant (2015), I compute the rich household MPS in the model from microsimulation of a permanent income increase, described in Appendix A.

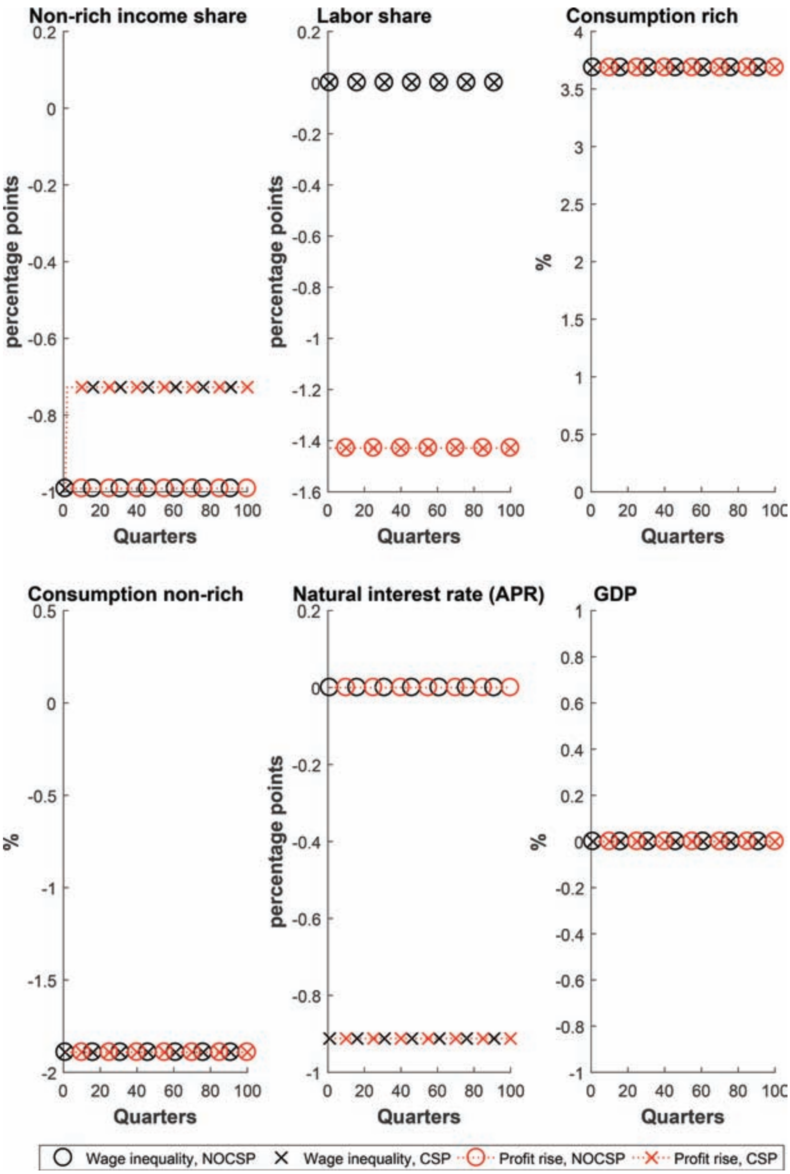
2.6 An Increase in Inequality in the Simple Model

All simulations in this paper are performed using the Newton-type solver for deterministic non-linear simulations as implemented in Dynare 4.6.3. (see Juillard 1996 and Adjemian et al. 2011).[9] Figure 2 displays separately the effect of a permanent increase in the markup (red lines) and the labor income share of rich households (black lines). Both shocks are calibrated such that the share of non-rich households in total household income declines by 1 percentage point on impact. Both inequality shocks also have effects of identical magnitudes on the consumption of rich households (which increases) and non-rich households (which decreases), while the labor share is affected only by the markup shock. Furthermore, CSP do not change the effect of the increase in inequality on any of the variables except for the interest rate and the non-rich income share. The interest rate declines by almost 1 percentage point. Hence, with CSP, the increase in rich households' permanent income does not in itself trigger an equal increase in rich household consumption. Since rich household wealth $b_{s,t}$ is constant as a result of the government's fiscal policy

[8] The authors report that virtually all of the officers in their sample have a college degree, while according to the Current Population Survey the same was true for only 24.5 percent of individuals in the same age group.

[9] This solver treats the deterministic simulation as a system of simultaneous equations in n endogenous variables in T periods.

Figure 2. Impact of a Permanent Increase in Inequality—Simple Model



(see Equation (12)), for bond and goods markets to clear, the interest rate has to decline (see Equation (2)). The interest rate decline partially compensates for the increase in the labor or profit income of the rich, implying that the non-rich income share recovers by about 0.2 percentage point in the second quarter.

3. The Full Model

In the full model rich households invest in financial intermediary deposits, physical capital, and housing (on top of government bonds) and derive utility from these assets. Non-rich households derive utility from consumption and housing and borrow from financial intermediaries. Throughout, $\tilde{X}_S$ (or $\tilde{x}_S$) denotes a rich household choice variable expressed in per capita terms, while $X_S \equiv \tilde{X}_S n_S$ denotes the corresponding economy-wide variable (and analogously for non-rich households).

3.1 Rich Households

Rich households derive utility from consumption $\tilde{C}_{S,t}$; their stocks of safe real financial assets $\tilde{b}_{S,t}$; the value of their physical capital $Q_t\tilde{K}_t$, where Q_t denotes the price of capital goods; and their housing stock $\tilde{H}_{S,t}$. Their objective function is thus given by

$$E_t\left\{\sum_{i=0}^{\infty}\beta_S^i\left[\frac{\tilde{C}_{S,t+i}^{1-\sigma_S}}{1-\sigma_S}+\frac{\tilde{\phi}_{H,S}}{1-\sigma_{H,S}}\tilde{H}_{S,t+i}^{1-\sigma_{H,S}}+\frac{\phi_{\tilde{b}}}{1-\sigma_b}\left(\tilde{b}_{S,t+i}\right)^{1-\sigma_b}\right.\right.$$
$$\left.\left.+\frac{\tilde{\phi}_K}{1-\sigma_K}\left(Q_{t+i}\tilde{K}_{t+i}\right)^{1-\sigma_K}\right]\right\}. \tag{19}$$

From now on, I will refer to the model where the rich derive utility from real safe assets and physical capital (i.e., $\tilde{\phi}_b, \tilde{\phi}_K > 0$) on top of housing and consumption as the CSP model, while I refer to the case where the rich do not derive utility from these two assets ($\tilde{\phi}_b = \tilde{\phi}_K = 0$) as the model without CSP (NOCSP). Note that, even in the presence of CSP, I assume that housing remains a durable consumption good and thus, unlike for the other two assets, utility depends on the physical housing stock $\tilde{H}_{S,t}$ rather than its market value $Q_{H,t}\tilde{H}_{S,t}$, in line with the literature (Iacoviello 2005, 2015; Iacoviello and Neri 2010; Clerc et al. 2015). However, results are robust to assuming that utility depends on the market value of the housing stock instead.

A rich household's budget constraint and capital accumulation equation are given by

$$\begin{aligned}\tilde{b}_{S,t} = & \frac{R_{t-1}}{\Pi_t}\tilde{b}_{S,t-1} + w_{S,t}\tilde{N}_{S,t} + r_{K,t}\tilde{K}_{t-1} + \Xi_t \\ & - Q_{H,t}\left(\tilde{H}_{S,t} - \tilde{H}_{S,t-1}\right) - \tilde{T}_{S,t} - \tilde{C}_{S,t} \\ & - \left(\tilde{I}_t + \tilde{K}_{t-1}\frac{\epsilon_I}{2}\left(\frac{\tilde{I}_t}{\tilde{K}_{t-1}} - \delta\right)^2\right) \qquad (20)\end{aligned}$$

$$\tilde{K}_t = (1-\delta)\,\tilde{K}_{t-1} + \tilde{I}_t, \qquad (21)$$

where $r_{K,t}$, $Q_{H,t}$, $\tilde{I}_t$, δ, and $\tilde{K}_{t-1}\frac{\epsilon_I}{2}\left(\frac{\tilde{I}_t}{\tilde{K}_{t-1}} - \delta\right)^2$ denote the real capital rental, the real house price, investment, the depreciation rate of physical capital, and quadratic capital stock adjustment costs, respectively.

I assume that safe assets $\tilde{b}_{S,t}$ comprise both government bonds and financial intermediary deposits. The simplifying assumption that government bonds and financial intermediary deposits are perfect substitutes in equilibrium, or assumptions to that effect, are common in the DSGE literature on models with banking (e.g., Gertler and Karadi 2011, 2013; Iacoviello 2015), and appear a reasonable approximation in the context of this paper, which focuses on long-run trends. The first-order conditions with respect to consumption, safe assets, capital, investment, and housing can be written in terms of economy-wide variables as

$$\Lambda_{S,t} = \frac{1}{C_{S,t}^{\sigma_S}} \qquad (22)$$

$$\Lambda_{S,t} = \beta_S E_t\left\{\Lambda_{S,t+1}\frac{R_t}{\Pi_{t+1}}\right\} + \phi_b b_{S,t}^{-\sigma_b} \qquad (23)$$

$$Q_t = E_t\Bigg\{\beta_S \frac{\Lambda_{S,t+1}}{\Lambda_{S,t}}\Bigg[r_{K,t+1} + \frac{I_{t+1}}{K_t}\epsilon_I\left(\frac{I_{t+1}}{K_t} - \delta\right) - \frac{\epsilon_I}{2}\left(\frac{I_{t+1}}{K_t} - \delta\right)^2 + (1-\delta)\,Q_{t+1}\Bigg] + Q_t\frac{\phi_K\,(Q_tK_t)^{-\sigma_K}}{\Lambda_{S,t}}\Bigg\} \tag{24}$$

$$Q_t = 1 + \epsilon_I\left(\frac{I_t}{K_{t-1}} - \delta\right) \tag{25}$$

$$Q_{H,t} = \frac{\phi_{H,S}}{H_{S,t}^{\sigma_{H,S}}\Lambda_{S,t}} + \beta_S E_t\left\{\frac{\Lambda_{S,t+1}}{\Lambda_{S,t}}Q_{H,t+1}\right\}, \tag{26}$$

where Q_t denotes the real value of an additional unit of capital to the household.[10] Finally, for future reference, I define total rich household consumption $C_{S,T,t}$ as

$$C_{S,T,t} = C_{S,t} + CH_{S,t} \tag{27}$$

$$CH_{S,t} = \frac{\phi_{H,S}}{\Lambda_{S,t}H_{S,t}^{\sigma_{H,S}}}H_{S,t}, \tag{28}$$

where $\frac{\phi_{H,S}}{\Lambda_{S,t}H_{S,t}^{\sigma_{H,S}}}$ denotes rich households' imputed or "shadow" rent, i.e., the value of an additional unit of housing to rich households expressed in consumption units. $CH_{S,t}$ denotes rich households' housing consumption as defined in the national accounts, i.e., the product of the imputed rent and $H_{S,t}$.

A key difference between my analysis and that of Mian, Straub, and Sufi (2020a) is that they assume that the wealth entering the utility function equals the present value of future income from *all* sources, including labor income if present. This assumption combined with $\sigma_S > \sigma_b$ implies that inducing rich households to hold more wealth requires a *decline* in the interest rate (see Appendix D for further details). Correspondingly, Mian, Straub, and Sufi (2020a) find that an increase in the supply of government bonds lowers

[10] The steps to replace per capita variables with economy-wide variables are analogous to the simple model (see footnote 5), with $\Lambda_{S,t} \equiv \frac{\tilde{\Lambda}_{S,t}}{n_S^{\sigma_S}}$, $\phi_K \equiv \frac{\tilde{\phi}_K n_S^{\sigma_K}}{n_S^{\sigma_S}}$ and $\phi_{H,S} \equiv \frac{\tilde{\phi}_{H,S} n_S^{\sigma_H}}{n_S^{\sigma_S}}$.

the real interest rate, which is at odds with the empirical evidence of Gale and Orszag (2004), Engen and Hubbard (2005), Laubach (2009), and Krishnamurthy and Vissing-Jorgensen (2012). By contrast, I assume that rich households derive utility from accumulated real and financial wealth only (like, e.g., Carroll 2000; Reiter 2004; Francis 2008; Piketty 2011; and Kumhof, Ranciere, and Winant 2015) and that rich households have substantial labor income, which allows my model to match this evidence. I will make use of this property when calibrating the safe asset curvature parameter σ_b in Section 3.6 below.

3.2 Borrower (Non-Rich) Households

Borrowing households are indexed with CC and derive utility from consumption and housing. The objective of a borrower household is given by

$$E_t \left\{ \sum_{i=0}^{\infty} \beta_{CC}^{i} \left[\frac{\tilde{C}_{CC,t+i}^{1-\sigma_{CC}}}{1-\sigma_{CC}} + \frac{\tilde{\phi}_{H,CC,t+i}}{1-\sigma_{H,CC}} \tilde{H}_{S,t+i}^{1-\sigma_{H,CC}} \right] \right\},$$

where I allow the utility weight on housing to be time varying (but exogenous to the individual borrower household). I assume that non-rich households are sufficiently impatient such that their borrowing is positive in equilibrium. Furthermore, I assume that borrowing is subject to a costly friction, possibly in the form of a default cost. The friction becomes more severe the larger a household's LTV ratio $\frac{b_{CC,t}}{H_{CC,t}Q_{H,t+1}}$, possibly because the likelihood of (strategic) default increases. The financial intermediary passes on these costs in full to borrower households, implying that the borrowers' expected interest rate $E_t \{R_{L,t+1}\}$ on its period t borrowing is determined by

$$\frac{E_t \{R_{L,t+1}\}}{R_t} = \left(1 + E_t \left\{ f\left(\frac{\tilde{b}_{CC,t}}{\tilde{H}_{CC,t}Q_{H,t+1}} \right) \right\} \right) \tag{29}$$

with $f'() > 0$. These assumptions capture in a simple fashion the empirical finding that non-rich households are more likely to be subject to borrowing constraints, but that their constraint is lessened by an increase in the value of their home, as argued by Mian and Sufi (2011, 2014). Appendix J shows that a positive relationship between the household's LTV and its cost of borrowing may be microfounded by assuming idiosyncratic shocks to the value of a borrower's house, costly strategic default, and that the borrower's house serves as collateral in a state-contingent debt contract, following Lozej, Onorante, and Rannenberg (2018). The simulation results which I discuss in Section 4 of the main text are broadly robust to adopting this microfoundation (see Appendix K for details).

The budget constraint of borrowing households is given by

$$\frac{R_{L,t}}{\Pi_t}\tilde{b}_{CC,t-1} + \tilde{C}_{CC,t} + Q_{H,t}\left(\tilde{H}_{CC,t} - \tilde{H}_{CC,t-1}\right) = \tilde{b}_{CC_t} + w_{CC,t}\tilde{N}_{CC,t} - \tilde{T}_{CC,t}. \tag{30}$$

The FOCs with respect to consumption $\tilde{C}_{CC,t}$, real loans $\tilde{b}_{CC,t}$, and housing $\tilde{H}_{CC,t}$, expressed in terms of economy-wide variables, are given by

$$\Lambda_{CC,t} = \frac{1}{C_{CC,t}^{\sigma_{CC}}} \tag{31}$$

$$\Lambda_{CC,t} = \beta_{CC} E_t \left\{ \Lambda_{CC,t+1} \left[\frac{R_{L,t+1}}{\Pi_{t+1}} + \frac{\frac{dR_{L,t+1}}{db_{CC,t}}\left(\frac{b_{CC,t}}{H_{CC,t}Q_{H,t+1}}\right)}{\Pi_{t+1}} b_{CC,t} \right] \right\} \tag{32}$$

$$Q_{H,t} = \frac{\phi_{H,CC,t}}{\Lambda_{CC,t} H_{CC,t}^{\sigma_{H,CC}}} + \beta_{CC} E_t \left\{ \frac{\Lambda_{CC,t+1}}{\Lambda_{CC,t}} \left(Q_{H,t+1} - \frac{\frac{dR_{L,t+1}}{dH_{CC,t}}\left(\frac{b_{CC,t}}{H_{CC,t}Q_{H,t+1}}\right)}{\Pi_{t+1}} b_{CC,t} \right) \right\}, \tag{33}$$

where $\frac{dR_{L,t+1}}{db_{CC}}\left(\frac{b_{CC,t}}{H_{CC,t}Q_{H,t+1}}\right)$ denotes the effect of an increase in borrowing $b_{CC,t}$ on the loan rate $R_{L,t+1}$ implied by the loan supply curve (29).[11] Hence when trading off today's and tomorrow's consumption,

[11] The steps to eliminate per capita variables are mostly analogous to the simple model (see footnote 5), with $\phi_{H,t} \equiv \frac{\phi_{\tilde{H},t} n_S^{\sigma_H}}{n_S^{\sigma_S}}$. For instance, the debt-FOC is

borrower households take into account both the expected interest rate on the additional unit of borrowing $\frac{R_{L,t+1}}{\Pi_{t+1}}$ and the expected increase in the interest rate burden on their existing stock of borrowing $\frac{dR_{L,t+1}}{db_{CC}}\left(\frac{b_{CC,t}}{H_{CC,t}Q_{H,t+1}}\right)$ resulting from the worsening of the borrowing friction. Correspondingly, $\frac{dR_{L,t+1}}{dH_{CC,t}}\left(\frac{b_{CC,t}}{H_{CC,t}Q_{H,t+1}}\right)$ denotes the implied (negative) effect of an increase in the housing stock on the loan rate (holding $b_{CC,t}$ constant). I assume that $f\left(\right)$ is described by a simple linear function

$$f\left(LTV_t\right) = \chi_{CC} LTV_t \tag{34}$$

with $LTV_t = \frac{b_{CC,t}}{H_{CC,t}Q_{H,t+1}}$. Finally, I assume that the utility weight on housing of the non-rich $\phi_{H,CC,t}$ may depend on lagged rich household total consumption (including housing consumption) $C_{T,S,t-1}$

$$\phi_{H,CC,t} = \phi_{H,CC}\left(\left(\frac{C_{T,S,t-1}}{C_{T,S}}\right)^{\nu_{cascade}}\right)^{\sigma_{H,CC}} \tag{35}$$

with $\phi_{H,CC} > 0$ and $\nu_{cascade} \geqq 0$. Hence, a 1 percent increase in lagged rich household total consumption $C_{T,S,t-1}$ increases the housing demand of the non-rich by $\nu_{cascade}$ percent. The motivation for this assumption is the so-called catching up with the richer Joneses (Drechsel-Grau and Schmid 2014) type of behavior. Specifically, there is microeconometric evidence that households care about their consumption relative to a reference group richer than themselves, and that an increase in the consumption of that richer

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given by $\tilde{\Lambda}_{CC,t} = \beta_{CC} E_t\left\{\tilde{\Lambda}_{CC,t+1}\left[\frac{R_{L,t+1}}{\Pi_{t+1}} + \frac{\frac{dR_{L,t+1}}{db_{CC,t}}\left(\frac{\tilde{b}_{CC,t}}{\tilde{H}_{CC,t}Q_{H,t+1}}\right)}{\Pi_{t+1}}\tilde{b}_{CC,t}\right]\right\}$. Substituting $\frac{E_t\left\{R_{L,t+1}\right\}}{d\tilde{b}_{CC,t}} = E_t\left\{f'\left(\frac{\tilde{b}_{CC,t}}{\tilde{H}_{CC,t}Q_{H,t+1}}\right)\right\}R_t\frac{1}{\tilde{H}_{CC,t}Q_{H,t+1}}$ yields $\tilde{\Lambda}_{CC,t} = \beta_{CC}E_t\left\{\tilde{\Lambda}_{CC,t+1}\left[\frac{R_{L,t+1}}{\Pi_{t+1}} + \frac{R_t}{\Pi_{t+1}}f'\left(\frac{\tilde{b}_{CC,t}}{\tilde{H}_{CC,t}Q_{H,t+1}}\right)\frac{\tilde{b}_{CC,t}}{\tilde{H}_{CC,t}Q_{H,t+1}}\right]\right\}$. Applying $\Lambda_{CC,t} \equiv \frac{\tilde{\Lambda}_{CC,t}}{n_{CC}^{\sigma_{CC}}}$ and $\tilde{X_S} = \frac{X_S}{n_S}$ to replace the remaining per capita variables, and then using $\frac{E_t\left\{R_{L,t+1}\right\}}{db_{CC,t}} = E_t\left\{f'\left(\frac{b_{CC,t}}{H_{CC,t}Q_{H,t+1}}\right)\right\}R_t\frac{1}{H_{CC,t}Q_{H,t+1}}$ yields Equation (32).

reference group boosts their own consumption (see Kuhn et al. 2011; Drechsel-Grau and Schmid 2014; Bertrand and Morse 2016), thus giving rise to so-called consumption cascades (Frank, Levine, and Dijk 2014; Belabed, Theobald, and van Treeck 2017). I limit the consumption cascade effect to the non-rich utility from housing due to the evidence in Bertrand and Morse (2016), who find that, in response to a 1 percent increase in the (total) consumption of the top 10 percent of households in a given state, the bottom 90 percent increase both the amount of housing services that they consume and the share of housing in their consumption basket.[12] They provide evidence that the disproportional effect on non-rich housing consumption may be related to the high visibility and thus status intensity of housing consumption. In the historical simulation of Section 4.2, this feature will help the model to match the upward trend in the value of the housing stock relative to GDP and the bottom 90 percent debt-to-income ratio by boosting the effect of rising inequality on housing demand, which in turn relaxes the borrowing constraint of non-rich households by lowering their LTV_t. Finally, analogously to rich households, housing consumption is defined as[13]

$$CH_{CC,t} = \left[\frac{\phi_{H,CC,t}}{\Lambda_{CC,t} H_{CC,t}^{\sigma_{H,CC}}} - \beta_{CC} E_t \left\{ \frac{\Lambda_{CC,t+1}}{\Lambda_{CC,t}} \frac{\frac{dR_{L,t+1}}{dH_{CC,t}} \left(\frac{b_{CC,t}}{H_{CC,t} Q_{H,t+1}} \right)}{\Pi_{t+1}} b_{CC,t} \right\} \right] H_{CC,t}. \quad (36)$$

[12]See Table 2, column 3, and Internet Appendix Table A9, rows 2 and 4.

[13]Due to the borrowing friction non-rich households face, on top of $\frac{\phi_{H,CC,t}}{\Lambda_{CC,t} H_{CC,t}^{\sigma_{H,CC}}}$, their marginal benefit of increasing $H_{CC,t}$ includes the present value of a decline in their cost of borrowing $-\beta_{CC} E_t \left\{ \frac{\Lambda_{CC,t+1}}{\Lambda_{CC,t}} \frac{\frac{dR_{L,t+1}}{dH_{CC,t}} \left(\frac{b_{CC,t}}{H_{CC,t} Q_{H,t+1}} \right)}{\Pi_{t+1}} b_{CC,t} \right\}$ generated by the decline in their LTV_t caused by the increase in $H_{CC,t}$, holding $b_{CC,t}$ constant.

3.3 Firms and Total Output

The firm's technology now features physical capital, which it combines with labor using a CES technology,

$$Y_{f,t}(j) = \left[(1-\alpha_K) \left(A_N \left(\frac{N(j)_{S,t}}{\eta_S} \right)^{1-\omega_{CC}-d_{CC,t}} \left(\frac{N(j)_{CC,t}}{\eta_{CC}} \right)^{\omega_{CC}+d_{CC,t}} \right)^{\frac{\epsilon-1}{\epsilon}} + \alpha_K \left(A_K K(j)_{t-1} \right)^{\frac{\epsilon-1}{\epsilon}} \right]^{\frac{\epsilon}{\epsilon-1}}, \tag{37}$$

implying the following FOCs:

$$r_{K,t} = mc_t \alpha_K \left(\frac{Y_{f,t}}{A_K K_{t-1}} \right)^{\frac{1}{\epsilon}} A_K \tag{38}$$

$$w_{S,t} = mc_t (1-\alpha_K)(1-\omega_{CC}-d_{CC,t}) \left(\frac{Y_{f,t}}{A_N N_t} \right)^{\frac{1}{\epsilon}} \frac{N_t A_N}{N_{S,t}} \tag{39}$$

$$w_{CC,t} = mc_t (1-\alpha_K)(\omega_{CC}+d_{CC,t}) \left(\frac{Y_{f,t}}{A_N N_t} \right)^{\frac{1}{\epsilon}} \frac{N_t A_N}{N_{CC,t}} \tag{40}$$

$$N_t = \left(\frac{N_{S,t}}{\eta_S} \right)^{1-\omega_{CC}-d_{CC,t}} \left(\frac{N_{CC,t}}{\eta_{CC}} \right)^{\omega_{CC}+d_{CC,t}} \tag{41}$$

$$\frac{1}{\mu_P + d_{\mu,t}} = mc_t. \tag{42}$$

GDP (or value-added) Y_t equals the sum of total firm output $Y_{f,t}$ net costs associated with the financial friction and adjusting the capital stock, and total housing services CH_t:

$$Y_t = Y_{f,t} - b_{CC,t-1} f(LTV_t) \frac{R_{t-1}}{\Pi_t} - \Phi \left(\frac{I_t}{K_{t-1}} \right) K_{t-1} + CH_t \tag{43}$$

$$CH_t = CH_{S,t} + CH_{CC,t}. \tag{44}$$

3.4 Government

I assume that the government sets the share of non-rich agents in total lump-sum taxes payable equal to their share in pre-tax national income NI_t:

$$Target_{T_{CC}2T_t} = \frac{T_{CC,t}}{T_{S,t} + T_{CC,t}} \tag{45}$$

$$Target_{T_{CC}2T_t} = NIS_{CC,t} \tag{46}$$

$$NI_t = Y_t - \delta K_{t-1} \tag{47}$$

$$NIS_{CC,t} = \frac{w_{CC,t}N_{CC,t} - b_{CC,t-1}\left(\frac{R_{L,t}}{\Pi_t} - 1\right)\frac{R_{t-1}}{\Pi_t} + \frac{H_{CC,t}}{H_t}CH_t}{\left(1 + \frac{b_{gov,t}}{NI_t}\left(\frac{R_{t-1}}{\Pi_t} - 1\right)\right)NI_t.} \tag{48}$$

These equations replace Equation (14), while the remaining equations describing the government sector are unchanged. In particular, as in the simple model, the central bank successfully pursues a perfect inflation target, implying that the actual real interest rate always equals the natural rate. The computation of the non-rich national income share $NIS_{CC,t}$ is consistent with the World Inequality Database (WID) (see Appendix B for further details).

3.5 Equilibrium

The equilibrium conditions of goods market, the market for safe assets, and the housing market are given by

$$Y_t = C_{S,t} + C_{CC,t} + CH_t + I_t + G_t \tag{49}$$

$$b_{S,t} = b_{CC,t} + b_{gov,t}$$

$$H = H_{S,t} + H_{CC,t}, \tag{50}$$

where I assume a constant economy-wide housing stock H. Thus the endogenous variables of the model are determined by (22)–(50) and (8)–(15). The only exogenous variables are the shocks to the production elasticity of households $d_{CC,t}$ and the price markup $d_{\mu,t}$.

3.6 Calibration

3.6.1 Baseline Calibration

I set the capital stock adjustment cost curvature $\epsilon_I = 7$, in line with the estimate of Cummins, Hassett, and Olinar (2006) (see Table 3) and the rate of depreciation $\delta = 0.025$. In line with the literature on housing in DSGE models, I assume a partial equilibrium income effect on housing demand of 1 percent (i.e., $\sigma_{H,S} = \sigma_{H,CC} = \sigma_{CC} = \sigma_S$; see Iacoviello 2005, 2015; Iacoviello and Neri 2010; Clerc et al. 2015). I assume an elasticity of substitution between capital and labor ϵ of 1. As in the simple model, I set the remaining parameters such that the steady-state of the model matches a range of empirical targets, reported in Table 4. In the full model without CSP, there are in total 11 parameters calibrated in this fashion (σ_S, β_S, β_{CC}, $\phi_{H,S}$, $\phi_{H,CC}$, μ_p, α_K, ω_{CC}, $Target_{G2GDP}$, $Target_{bgov2GDP}$, χ_{CC}). They are pinned down by the intertemporal elasticity of substitution, the real short-term interest rate, the borrower debt-to-annual-income ratio, the residential housing-stock-to-annual-GDP ratio, the share of borrowers in total residential real estate, the labor share, the share of private fixed investment in GDP, the national income share of the bottom 90 percent of households as reported by the WID, the GDP share of government expenditure on goods and services, the government debt-to-GDP ratio, and a measure of the spread of the mortgage rate over the risk-free rate. Appendix I shows how, given those targets, the parameter values supporting them can be computed for in (almost) closed form.

The CSP are now described by four parameters, σ_b, ϕ_b, σ_K, and ϕ_K, two more than in the simple model of Section 2. Therefore I add two additional empirical targets to calibrate the preference parameters, on top of the MPS of the rich and the discounting wedge θ already used for calibrating the simple model. The first is the spread between the return on capital $r_K - \delta$ and the real risk-free rate $\frac{R}{\Pi}$. I measure this spread as a simple average of an empirical estimate of the external finance premium and the equity risk premium (see Table 4 for details). The second is the evidence of Gale and Orszag (2004), Engen and Hubbard (2005), and Laubach (2009) on the effect of a 1 percentage point increase in the government debt-to-annual-GDP ratio on the real interest rate on U.S. government bonds, for

Table 3. Full Model, Parameter Values

Parameter	Parameter Name	Value NOCSP ($\theta = 1$)	Value CSP ($\theta = 0.97$)
β_S	Rich Utility Discount Factor	0.9951*	0.9652*
β_{CC}	Borrower Utility Discount Factor	0.9868*	0.9868*
σ_S,σ_{CC}	Utility Curvature Consumption	2*	2*
$\sigma_{S,H}$,$\sigma_{CC,H}$	Utility Curvature Housing	2	2
$\phi_{H,S}$	Rich Utility Weight on Housing	0.17*	1.30*
$\phi_{H,CC}$	Borrower Utility Weight on Housing	0.64*	0.73*
$\nu_{cascade}$	Consumption Cascade	0	0; 0.7
$\tilde{N}_{CC}$,$\tilde{N}_S$	Labor Endowments	$\frac{1}{3}$	$\frac{1}{3}$
μ_P	Price Markup	1.17*	1.06*
α_K	Output Elasticity w.r.t. Capital	0.19*	0.24*
ϵ	Elasticity of Substitution Capital/Labor	1	1; 0.3
ω_{CC}	Borrower Share in Labor Income	0.83*	0.79*
δ	Depreciation Rate Physical Capital	0.025	0.025
ϵ_I	Capital Adjustment Cost Curvature	7	7
χ_{CC}	Financial Intermediation Cost, Linear	0.015*	0.015*
$Target_{bgov2GDP}$	Government Debt Target	44%*	44%*
$Target_{G2GDP}$	Government Expenditure Share Target	20%*	20%*
σ_b	CSP: Utility Curvature, Real Financial Assets	—	0.69*; 1.16*
σ_K	CSP: Utility Curvature, Physical Capital	—	5*/1.16*
ϕ_b	CSP: Utility Weight on Real Financial Assets	0*	3.32*
ϕ_K	CSP: Utility Weight on Physical Capital	0*	222.8*

Note: Values marked with an asterisk (*) are set to match the targets reported in Table 4. Wherever a cell in the CSP column reports two values, the first refers to the baseline CSP calibration discussed in Section 3.6.1. Without loss of generality, I assume $A_K = \frac{Y}{K}$ and $A_N = \frac{Y}{N}$, where $\frac{Y}{K}$ and $\frac{Y}{N}$ refer to the initial steady state of the model. This normalization implies that in the initial steady state, Equations (38) to (40) are the same in the Cobb-Douglas and the CES case. See Appendix I.

Table 4. Targets

Target	Value Data	NOCSP	CSP	Source
Real Short-Term Interest Rate $\frac{R}{\Pi}$	2%	2%	2%	Federal Funds Rate Minus Core-PCE Inflation, APR, FRED
Intertemporal Elasticity of Substitution (IES) $\frac{1}{\sigma_S}$	0.5	0.5	0.5	Havranek (2015)
Labor Share $\frac{wN}{Y}$	66%	66%	66%	WID, % of Factor-Price Gross National Income
Non-res. Investment-to-GDP Ratio $\frac{I}{Y}$	13%	13%	13%	Share of Non-residential Fixed Investment in GDP, BEA
$\frac{K}{4Y}$ (Not Targeted)	126%	130%	130%	Private-Capital-Stock-to-Annual-GDP Ratio, BEA
$\frac{G}{Y}$	20%	20%	20%	Government Expenditure as Percentage of GDP, BEA
$\frac{b_{gov}}{4Y}$	44%	44%	44%	Federal Debt Held by the Public, Percentage of GDP, FRED
Borrower Share in Labor Income (Not Targeted) $\frac{w_{CC}N_{CC}}{w_{CC}N_{CC}+w_S N_S}$	72%	78%	79%	Bottom 90% of Wage Earners Labor Income Share, WID
Borrower National Income Share NIS_{CC}	66%	66%	66%	Bottom 90% National Income Share, Pre-tax, WID
Residential Housing Stock to GDP $\left(\frac{Q_H H}{4Y}\right)$	106%	106%	106%	Flow of Funds, FRB
Borrower Share in Residential Real Estate $\left(\frac{Q_H H_{CC}}{Q_H H}\right)$	70%	70%	70%	Bottom 90% Share in Residential Real Estate, SCF, FRB, 1983
Borrower Debt-to-Annual Income Ratio	38%	38%	38%	Debt Secured by Primary Residence, SCG, FRB, 1983
Mortgages Rate Minus Risk-Free Rate $\frac{R_L}{R} - 1$	1.68%	1.68%	1.68%	30-Year Mortgage Rate Minus 30-Year Treasury, FRED
MPS Top 5%	51%	0	56%	Kumhof, Ranciere, and Winant (2015), Appendix III, Based on Dynan, Skinner, and Zeldes (2004)'s Estimates
Discounting Wedge θ	0.97	1.0	0.97	Target CSP Case: See Section 2.5
Return on Capital Minus Risk-Free Rate $ep = \frac{r_K - \delta + 1}{\frac{R}{\Pi}} - 1$	4%	0	4%	Target CSP Case: See Note Below for Details
$\frac{d\text{Interest rate gov. bonds}}{d\text{Gov. Debt ratio}}$	0.03–0.06 p.p.	0	0.055 p.p.	Target CSP Case: See Note Below for Details

Note: The targets for $\frac{G}{Y}$, $\frac{b_{gov}}{4Y}$ and $\frac{R_L}{R} - 1$ are calculated as averages over the 1981–2016 period. All other ratios for which annual data were available were calculated as averages over the 1973–80 period. FRED: Federal Reserve Economic Data. BEA: Bureau of Economic Analysis. SCF: Survey of Consumer Finances. WID: World Inequality Database. FRB: Federal Reserve Board. Since the FRB publishes comprehensive summary statistics for the SCF only starting in 1989, the 1983 data required to compute the bottom 90% mortgage-debt-to-income ratio, their share in residential real estate, and their LTV were computed as follows:

- Bottom 90% mean income: Computed from Table 1 in Avery et al. (1984b), Table 1, which divides the household population into different household income intervals.
- Bottom 90% mean mortgage debt (secured by primary residence): Avery et al. (1984a), Table 1 (which reports mean mortgage debt and the incidence of owing mortgage debt for selected income groups), and using the aforementioned population shares as weights.
- Bottom 90% and top 10% mean home value: From Avery et al. (1984b), Tables 5 and 7, which contain home ownership incidence and net equity of homeowners for selected household income intervals.

The target for the spread between the return on capital and the risk-free rate $r_K - \delta - \left(\frac{R}{\Pi} - 1\right)$ and the risk-free rate is a simple average of

- The spread between the yield on Moody's seasoned BAA-rated corporate bonds and 10-year Treasury bonds, 1981–2016 average, which equals 2.4%.
- The average estimate of the equity risk premium reported in Duarte and Rosa (2015), Table 7, which equals 5.7%.

The range of empirical estimates of $\frac{d\text{Interest rate gov. bonds}}{d\text{Gov. Debt ratio}}$ are obtained from Gale and Orszag (2004), Engen and Hubbard (2005), and Laubach (2009). The private capital stock used to calculate $\frac{K}{4Y}$ equals private non-residential fixed assets from the BEA's NIPA, "Table 2.1. Current-Cost Net Stock of Private Fixed Assets, Equipment, Structures, and Intellectual Property Products by Type."

which these authors find a range of 0.03 to 0.06 percentage point. I assume that the corresponding model counterpart is the effect of a 1 percentage point permanent increase in the government debt-to-annual-GDP ratio on the steady-state real interest rate. In practice, the value of $\frac{d\text{Interest rate gov. bonds}}{d\text{Gov. Debt ratio}}$ implied by the model is closely linked to the value of safe asset curvature parameter σ_b.[14] Appendix C provides further validation of the consumption and saving behavior of the households in the model by drawing on additional microeconometric evidence which the calibration does not target.

3.6.2 CSP Model Variants

On top of the baseline calibration discussed above, I consider three variants of the CSP model. Firstly, I consider an “identical curvature” version, where I drop the target for $\frac{d\text{Interest rate gov. bonds}}{d\text{Gov. Debt ratio}}$, and instead assume $\sigma_b = \sigma_K$, while keeping the target for the rich marginal propensity to save out of increases in permanent income. The approach results in $\sigma_b = \sigma_K = 1.2$. For this calibration, a given percentage increase (decline) in consumption (the marginal utility of consumption) implies the same percentage increase in the demand for safe assets and capital in the long run. Secondly, I consider a “CES” version, which assumes an elasticity of substitution between capital and labor of $\epsilon = 0.3$, as estimated by the metaregression of analysis of Gechert et al. (2019). Finally, I consider a version with an active “consumption cascade,” with $\nu_{cascade} = 0.7$, in line with the evidence of Bertrand and Morse (2016), who find that a 1 percent increase in the total consumption of the top 10 percent of households increases the consumption of housing services by the bottom 90 percent of households living in the same state by about 0.7 percent.

4. Results in the Full Model

In this section, I first investigate the effect of one-off wage inequality and price markup shocks (Section 4.1), and then perform a historical

[14]Note that the value of $\frac{d\text{Interest rate gov. bonds}}{d\text{Gov. Debt ratio}}$ implied by the calibration is at the upper bound of the empirical range. Targeting a lower value would imply a smaller value of σ_b and a larger value of σ_K, and would strengthen the results discussed in Section 4.

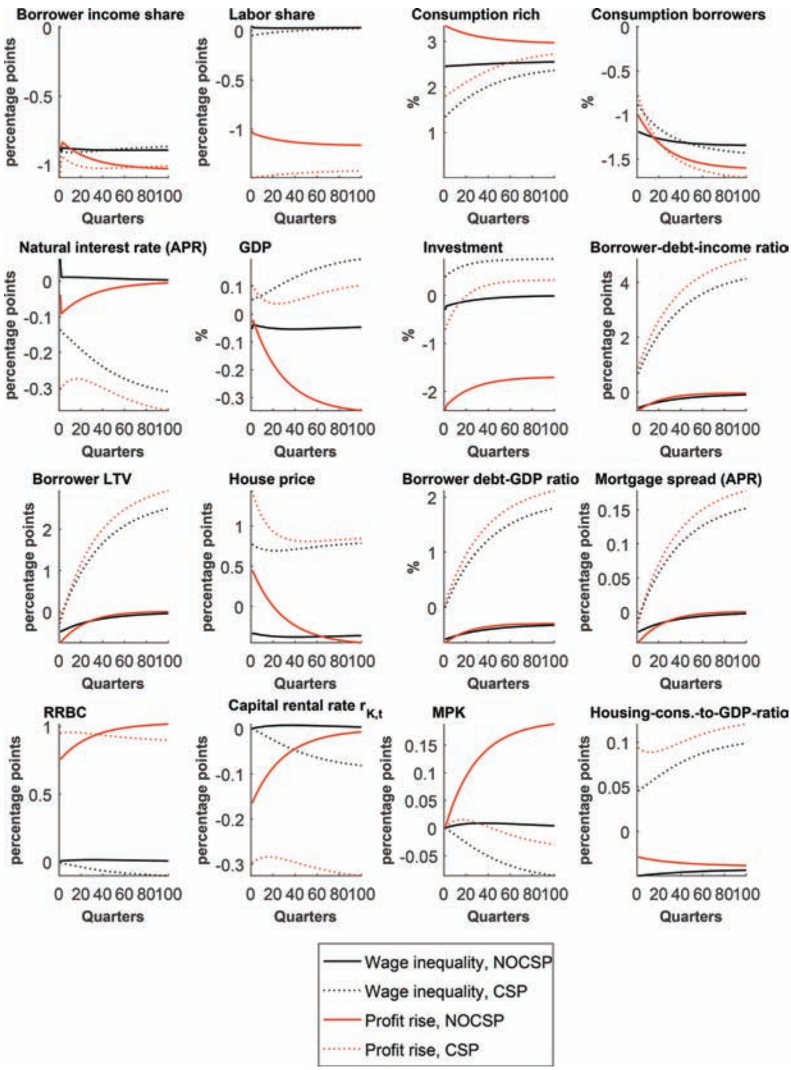
simulation which replicates the empirically observed increase in income inequality over the 1981–2016 period. The results reported below are broadly robust to assuming that rich households' capitalist spirit motive extends to their housing stock (see Appendix H) and to explicitly modeling the borrowing friction assumed above (see Appendix K).

4.1 A One-Off Permanent Increase in Inequality

I first consider the effect of a permanent decline in the share of borrower households in total labor income, caused by a permanent decline (increase) in the elasticity of output with respect to the labor supplied by borrower (rich) households, i.e., $d_{CC,t}$ becomes permanently negative (see Figure 3). I calibrate the value that $d_{CC,t}$ takes such that the on-impact decline in the borrower household income share equals approximately 1 percentage point. Without CSP (solid black line), rich households increase their consumption and housing demand on impact by approximately the magnitude of their permanent income change, similar to the simple model. Borrower households lower their consumption by approximately the decrease in their permanent income, and permanently reduce their housing demand and borrowing, implying that the household debt-to-GDP ratio declines permanently. As a result, without CSP, the effect on the natural interest rate is small, and actually positive on impact before quickly returning to zero.

By contrast, with CSP (dotted black line), rich households increase their consumption by only about half as much as without CSP, implying that in order to equilibrate asset and goods markets, the natural interest rate declines. The decline in the natural rate is initially passed on in full to borrowers, which motivates them to postpone the ultimate decline in their consumption of goods and housing services and thus to reduce their consumption on impact by only half as much as in the NOCSP case. As a result, borrower's housing demand declines more slowly than in the NOCSP case and their debt increases. At the same time, the lower interest rate increases the relative housing demand of both rich and non-rich households, implying that the value of the housing stock increases. The increase in house prices tends to relax the borrower households borrowing constraint, which tends to further strengthen their consumption and housing

Figure 3. Impact of a Permanent Increase in Inequality



Note: The black lines display the effect of a one-off permanent decline in the elasticity of output with respect to the labor supplied by non-rich (rich) households ($d_{CC,t}$ permanently declines; see Equation (37)). The red lines display the effect of a one-off increase in the price markup ($d_{\mu,t}$ permanently increases). CSP refers to the model with capitalist spirit type preferences. The safe interest rate R_t and the risk spread $E_t R_{L,t+1} - R_t$ are expressed as annualized percentage rates. The borrower debt-to-income ratio is based on annualized borrower income. The borrower debt-to-GDP ratio is based on annualized GDP. RRBC: real return on business capital, defined as $RRBC_t \equiv \frac{(1-mc_t)Y_{f,t}+(r_{K,t}-\delta)K_{t-1}}{Q_t K_t - 1}$. MPK: marginal product of capital.

demand. As a result, their debt-to-income ratio increases steeply, while their LTV actually decreases during the first year due to the higher house price before turning positive.

Apart from expanding their lending to borrower households via the financial intermediary, saver households also use their additional income to increase their investment, as the decline in the safe interest rate and the decline in the marginal utility of consumption relative to the marginal utility of physical capital renders physical capital relatively more attractive. As a result, the capital stock, GDP, and the capital-output ratio increase, and consequently the marginal product of capital, the capital rental $r_{K,t}$, and the real return on business capital all decline. Note that, in spite of the GDP increase, the effect of shock on non-rich households' welfare (as measured by their objective) is unambiguously negative due to the large decline in their consumption of goods and their housing stock. This result extends to all variants of the CSP model considered below, and extends to the price markup increase.

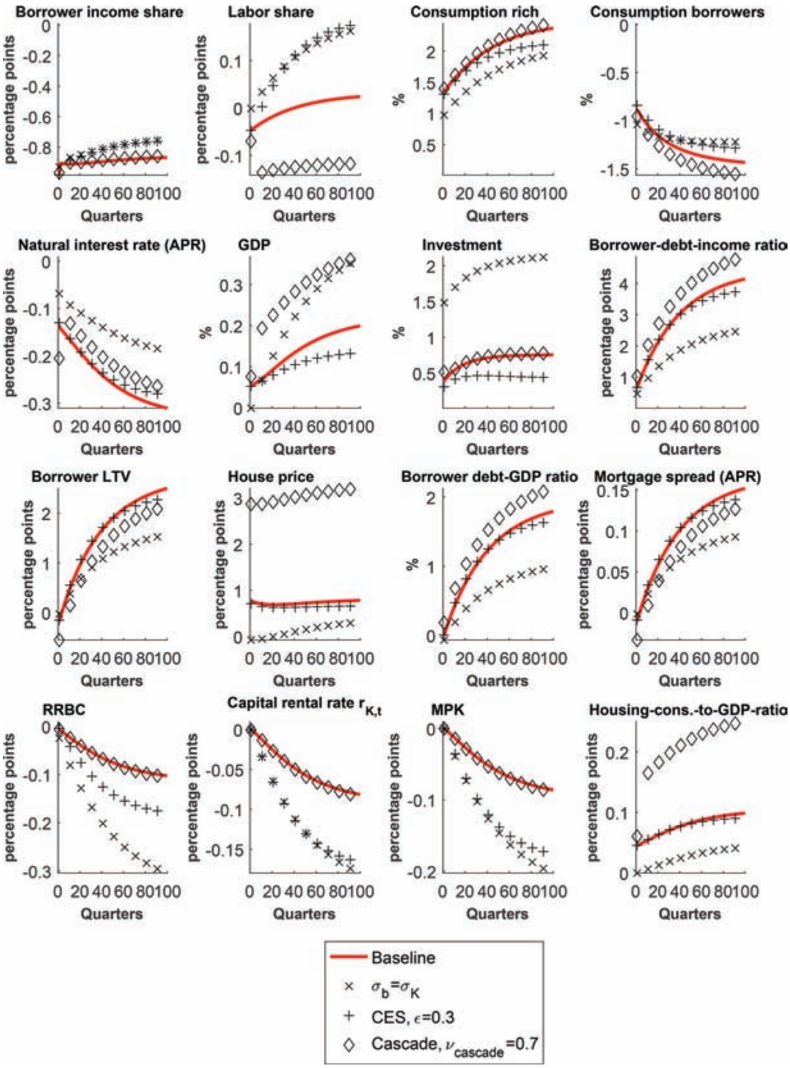
An increase in inequality driven by a permanent rise in the price markup (the red lines) differs from an increase in wage inequality in that it lowers the labor share and the demand for capital goods by the monopolistically competitive firms, and thus investment. However, GDP still increases slightly due to an increase in housing consumption CH_t.[15] Due to the decline in the demand for capital goods, the real interest rate and the capital rental $r_{K,t}$ decline more than for an increase in wage inequality (see Equations (10) and (38)), while the marginal product of capital increases subsequently due to the decline in the capital stock, though only temporarily. The real return on business capital increases due to the increased profits from monopolistic competition, which outweigh the decline in the net rental income from capital $(r_{K,t} - \delta)\, K_{t-1}$. Furthermore, the decline in their investment expenditure allows rich households to

[15]This increase due to a rise in rich households' housing consumption $CH_{S,t}$, which in turn is driven by the increase in their "imputed rent" due to the fact that their consumption increases by more than their housing stock, and thus their marginal utility of consumption declines relative to their marginal utility from housing (see Equation (28)). Compared with the wage inequality shock, the increase in rich household consumption and thus the increase in their imputed rent is enhanced by the decline in their investment spending.

increase their consumption even more than for an increase in wage inequality, implying that house prices now increase more due to a larger "wealth effect" on rich households housing demand, i.e., a sharper decline in their marginal utility of consumption relative to their marginal utility of housing. Apart from these differences, the impact of the markup shock and the wage inequality shock is very similar. The same is true for the effect of allowing for CSP, which again implies, inter alia, a decline in the natural rate, an increase in borrower household debt, an increase in house prices exceeding the increase observed in the absence of CSP, and a higher trajectory for investment than in the NOCSP case due to the lower real interest rate and the increase in the marginal utility of physical capital relative to the marginal utility of consumption.

Figure 4 compares the effect of a wage inequality increase in multiple variants of the CSP model. Assuming identical curvature ($\sigma_b = \sigma_K$) implies that an increase in safe assets reduces the marginal utility of capital (safe assets) by less (more) than in the baseline case. Hence, in the identical curvature case, a given percentage decline in the marginal utility of consumption increases the rich households' demand for capital and deposits by the same percentage. Thus, the demand for capital (bank deposits) increases more strongly (weakly) in the identical curvature case than in the baseline CSP case (see Equations (24) and (23), respectively). Correspondingly, both investment and GDP increase by substantially more than in the baseline CSP case (compare the black "X" line with the red solid line), while household borrowing increases by less and the decline in the natural rate is smaller. Furthermore, the medium-term decline in the capital rental $r_{K,t}$ now closely matches the decline in the natural interest rate, while it remains considerably smaller in the baseline case. By contrast, reducing the elasticity of substitution between capital and labor to the estimate of Gechert et al. (2019) of 0.3 (see the black "+" line) greatly reduces the increase in investment compared with the baseline case, because the optimal capital-labor ratio increases less in response to a given decline in the interest rate if capital and labor are less easily substitutable. In particular, unlike in the Cobb-Douglas case, the labor share increases, as the increase in the capital stock no longer fully compensates for the decline in the capital rental rate associated with the decline in the interest rate.

Figure 4. Impact of a Permanent Increase in Inequality: CSP Model Variants



(*continued*)

Figure 4. (Continued)

Note: The black lines display the effect of a one-off permanent decline in the elasticity of output with respect to the labor supplied by non-rich (rich) households ($d_{CC,t}$ permanently declines; see Equation (37)) for the baseline calibration of the CSP model and three variants (see Section 3.6.2 for details). "Baseline" refers to the calibration discussed in Section 3.6.1. "$\sigma_b = \sigma_K$" indicates the "identical curvature" variant. "CES" indicates the variant where the elasticity of substitution between capital and labor is set to $\epsilon = 0.3$. "Cascade" indicates that the borrowers marginal utility of housing depends on rich households' total consumption (see Section 3.2 and Equation (35) for details). The safe interest rate R_t and the risk spread $E_t R_{L,t+1} - R_t$ are expressed as annualized percentage rates. The borrower debt-to-income ratio is based on annualized borrower income. The borrower debt-to-GDP ratio is based on annualized GDP. RRBC: real return on business capital, defined as $RRBC_t = \frac{(1-mc_t)Y_{f,t}+(r_{K,t}-\delta)K_{t-1}}{Q_t K_t - 1}$. MPK: marginal product of capital.

Allowing for a spending cascade ($\nu_{cascade} = 0.7$) on top of CSP strongly raises the effect of an increase in wage inequality on household debt and the value of the housing stock (see line marked by "black diamonds"). With an active cascade, the increase in rich household consumption directly raises the housing demand of borrowers, which they fund via additional debt. The stronger rise in house prices implies that, in spite of borrowers' higher debt trajectory, their LTV is actually lower than in the absence of the spending cascade (compare the black dotted line and the black diamond line). By contrast, the observed decline in the natural interest rate is smaller. Furthermore, the increase in GDP is larger than in the baseline case, due to an increase in housing consumption CH_t compared with the baseline simulation, driven by the larger marginal utility of housing of non-rich households resulting from the cascade (see Equation (36)).

Finally, note that due to the existence of additional uses for the savings of rich households not present in the simple model of Section 2 (i.e., residential housing, physical capital, lending to the non-rich via financial intermediaries), the simulated decline in the safe real interest rate is smaller than in Figure 2. Nevertheless, as discussed in the following section, when the actual increase in inequality over the 1980–2016 period is replicated in the model, the resulting decline in the natural rate is substantial.

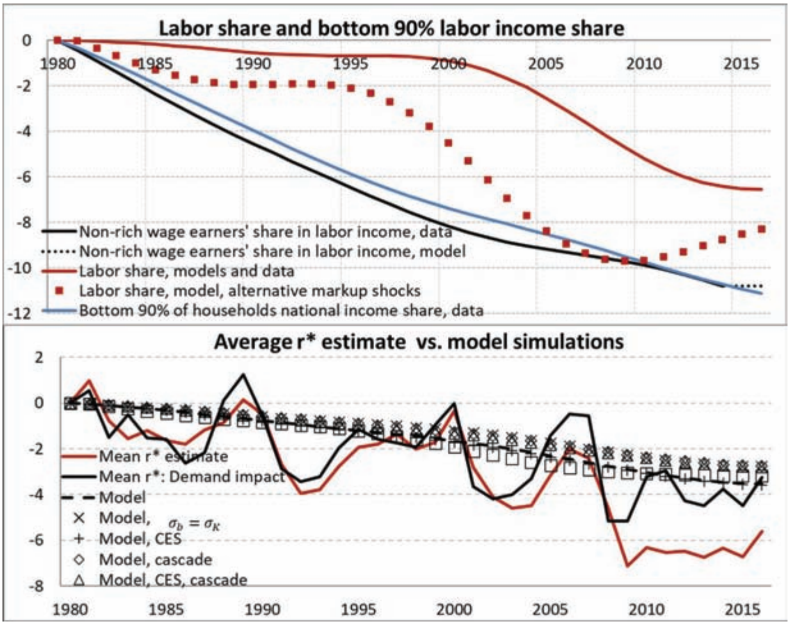
4.2 Simulation of the Empirically Observed Increase in Inequality

I now feed two stylized facts of the U.S. income distribution over the 1981–2016 period into the model with CSP. The first of these is the recent decline in the U.S. labor share, which starts at around the late 1990s (see Figure 5, first panel, the solid red line). In order to match the path of the labor share, I assume a sequence of unexpected permanent increases in the shock to the price markup $d_{\mu,t}$. This approach is similar to Caballero, Farhi, and Gourinchas (2017) and Farhi and Gourio (2018), who target the labor share to back out the path of the price markup in a simulation exercise studying the decline in the safe real interest rate and other trends. Hall (2018), Barkai (2020), and De Loecker, Eeckhout, and Unger (2020) provide evidence that pure profits (i.e., profits exceeding the opportunity cost of capital) and correspondingly price markups have indeed risen in the corporate sector. In order to focus the discussion on long-run trends, I remove fluctuations with an amplitude of 2 to 16 years from labor share series and all other annual data reported below using the asymmetric full-sample band-pass filter of Christiano and Fitzgerald (2003), assuming a unit root with drift.

The second stylized fact I reproduce in the model is the decline in the share of the bottom 90 percent of wages earners in total labor income reported in the WID over the 1980–2014 period (see the first panel of Figure 5, the solid black line), using the wage inequality shock $d_{CC,t}$.[16] Hence I implicitly assume that, as in the model, the bottom 90 percent of wage earners correspond to the bottom 90 percent of households. This assumption appears a reasonable approximation for two reasons. Firstly, the initial steady-state labor income share of non-rich households ω_{CC} implied by my calibration (and my target for the national income share of non-rich household in particular) is close to the empirical share of the bottom 90 percent of wage earners in total labor income (see Table 3), even though I do not explicitly target this value. Secondly, the share of the bottom 90 percent of wage earners in labor income and the national income share of the bottom 90 percent of households move almost

[16] Since the final data point is in 2014, for the final two years of the simulation I assume $d_{CC,2016} = d_{CC,2015} = d_{CC,2014}$.

Figure 5. Simulation 1981–2016: Income Distribution and Natural Interest Rate



Note: The label "Model" indicates the results of simulation using the model with CSP developed in Section 3. The simulation subjects the model to a sequence of unexpected but permanent shocks to the price markup $d_{\mu,t}$ and the labor income share of non-rich households $d_{CC,t}$. The trajectory of $d_{CC,t}$ is set to replicate the path of the bottom 90 percent of wage earners' share in total labor income in the data. The trajectory of $d_{\mu,t}$ is set such that the simulation replicates the path of the labor share in the data, unless the line is labeled "Model, Alternative Markup Shocks." In that case $d_{\mu,t}$ is set to replicate the path of an estimate of the pure profit share (see Appendix G for details). Changes in $d_{CC,t}$ and $d_{\mu,t}$ occur every four quarters, starting with the first simulation quarter. The corresponding 1980 value has been subtracted from all displayed series.
First Panel: The label "Labor Share, Models and Data" indicates the BLS labor share which all simulations other than the "Alternative Markup Shocks" simulation are forced to match.
Second Panel: The line labeled "Average r* Estimate" is the simple average across the r* estimates plotted in the upper panel of Figure 1 (the black squared line). The line labeled "Average Estimated Demand Impact on r*" is the simple average across the estimated impact of demand shocks on r* in Del Negro et al. (2017) and Gerali and Neri (2019) plotted in the lower panel of Figure 1 (the black squared line). The lines labeled "Model, $\sigma_b = \sigma_K$", "Model, CES", "rstar, Model, cascade" refer to the variants of the CSP model discussed in Section 3.6.2. "Model, CES, cascade" indicates an active spending cascade and an elasticity of substitution between capital and labor of $\epsilon = 0.3$.

(continued)

Figure 5. (Continued)

Data Sources and Treatment: Bottom 90 percent of wage earners' labor income share: WID. Labor share: Labor share non-farm business sector, Bureau of Labor Statistics. Bottom 90 percent of households' share in pre-tax national income: WID. For "Average r* Estimate" and "Average Estimated Demand Impact on r*" see the note below Figure 1. From all annual data except the rstar estimates, I remove fluctuations with an amplitude of 2 to 16 years using the asymmetric full sample band-pass filter of Christiano and Fitzgerald (2003), assuming a unit root with drift.

in parallel over the 1980–2014 period, thus suggesting substantial overlap between the two groups (compare the blue and black solid lines). Finally, the simulated path of the bottom 90 percent share in national income closely matches its empirical counterpart (see the graph in Appendix E).[17] Since the WID data are available at an annual frequency only and my focus is on the trends in any case, I assume that the changes in $d_{CC,t}$ and $d_{\mu,t}$ occur every four quarters, starting with the first simulation quarter.[18]

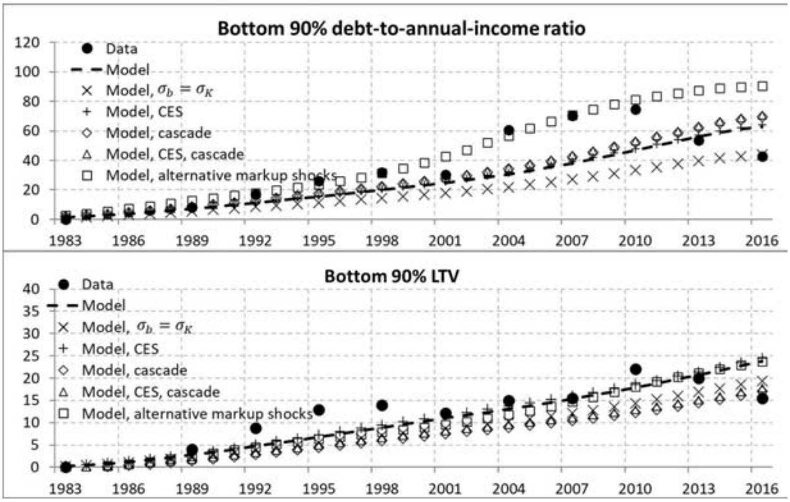
As can be seen from the second panel of Figure 5, the simulated increase in inequality generates a decline in the natural interest rate of between 3 and 4 percentage points over the 1981–2016 period, depending on the model variant. The simulation is thus able to broadly replicate the downward trend in the part of the natural rate which the aforementioned empirical estimates attribute to aggregate demand.

It should be remembered that the model's only drivers are the two aforementioned shocks to the income distribution, and that the model represents a hypothetical flexible price equilibrium, i.e., a situation where the output gap is closed. It thus abstracts from a multitude of potentially relevant influences, and would not be expected a

[17] While the model abstracts from the potential role of progressive income taxation in attenuating the effect of rising pre-tax inequality on household finances, note that the bottom 90 percent of households' post-tax disposable income share reported in the WID moves almost in parallel with the share of the bottom 90 percent of households in pre-tax national income over the 1980–2014 period.

[18] More specifically, I conduct a sequence of "perfect foresight" simulations, with one simulation corresponding to each occurrence of the shocks (i.e., 36 simulations in total). Of course, each simulation uses the fourth quarter of the preceding simulation as a starting point.

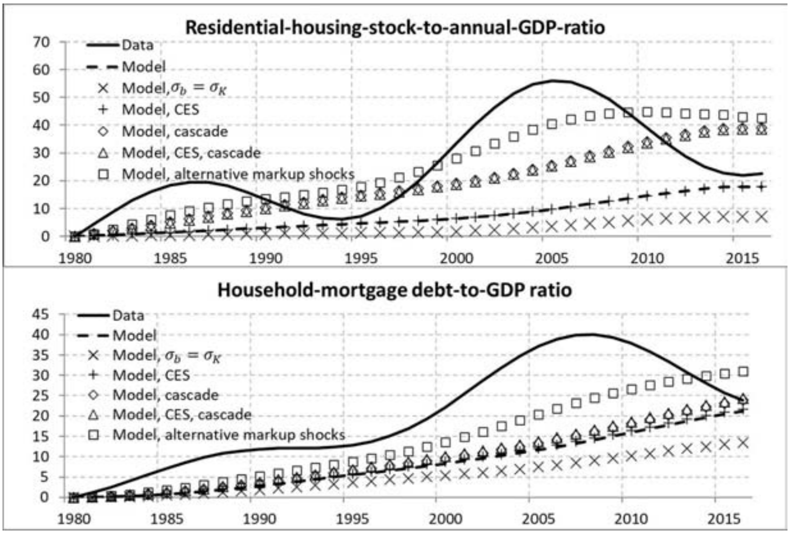
Figure 6. Simulation 1981–2016: Borrower Debt-to-Income Ratio and LTV



Note: This graph plots the simulated path of the non-rich debt-to-annualized-income ratio $\frac{b_{CC,t}}{4w_{CC,t}N_{CC,t}}100$ and LTV $\frac{b_{CC,t}}{Q_{H,t}H_{CC,t}}100$ and their empirical counterparts. See the note below Figure 5 for details on the model variants used in the simulation and the simulation setup. Note that the corresponding 1980 value has been subtracted from all displayed series.
Data Sources: The 1989–2016 values are computed from summary statistics provided by the Federal Reserve Board's Survey of Consumer Finances (SCF) 1989–2016. Bottom 90% debt-to-income ratio = (Bottom 90% mean mortgage debt)/(Bottom 90% mean income)*100. Bottom 90% mean income: Table 1. Bottom 90% mean mortgage debt (secured by primary residence): Table 13 and Table 13 means. Bottom 90% LTV = (Bottom 90% mean mortgage debt)/(Bottom 90% mean home value (primary residence)). Bottom 90% mean home value (primary residence): Table 13 and Table 13 means. Regarding the computation of the 1983 values, see the note below Table 4.

priori to match the data year by year. That being said, the simulation speaks to a number of important trends observed during the 1981–2016 period. All model variants closely track the upward trend in the bottom 90 percent of households' mortgage debt-to-annual-income ratio from the early 1980s until about 2001 (see Figure 6). The model variants other than the "$\sigma_b = \sigma_K$" model (see the line marked with "X") replicate between about two-thirds (for the baseline calibration) or more (CES +cascade; see the "black triangle" marked line)

Figure 7. Simulation 1981–2016: Housing Stock and Total Mortgage Debt-to-GDP Ratios

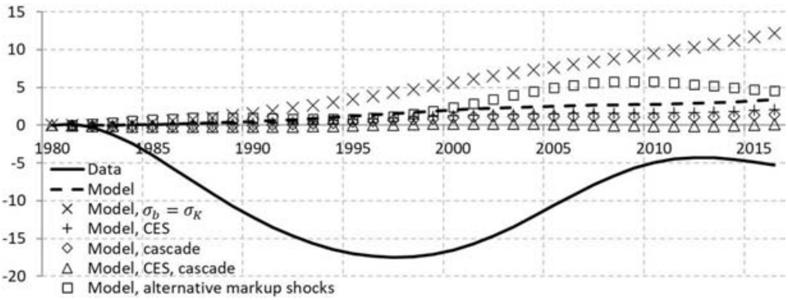


Note: This graph plots the simulated path of the housing-stock-to-annualized-GDP-ratio $\frac{Q_{H,t}H_t}{4_{Y,t}}100$ and the household-mortgage-debt-to-annualized-GDP ratio $\frac{b_{CC,t}}{4_{Y,t}}$ and their respective empirical counterparts. See the note below Figure 5 for details on the model variants used in the simulation, the simulation setup, and data treatment.
Data Sources: Federal Reserve Board/Flow of Funds, Bureau of Economic Analysis.

of the peak increase in the debt-to-income ratio compared with its 1980 value (attained in 2010). Furthermore, the simulation roughly tracks the rising trend in the bottom 90 percent LTV observed in the data.

The model also generates an empirically relevant rising trend in the ratio of nominal residential housing stock to GDP and the ratio of household mortgage debt to GDP (see Figure 7). Apart from a rising trend, in the data both variables display substantial volatility, especially post-2001, which my simple simulation exercise cannot capture. While the baseline model is able to replicate about one-fifth of the peak of the residential-housing-stock-to-GDP ratio, attained in 2005, the replicated share increases to almost one-half

Figure 8. Simulation 1981–2016: Private Non-residential Producible Capital-Stock-to-GDP Ratio



Note: This graph plots the simulated path of the private non-residential-capital-stock-to-annual-GDP ratio $\frac{K_t}{4_{Y,t}}100$ and its empirical counterpart. See the note below Figure 5 for details on the model variants used in the simulation, the simulation setup, and data treatment.
Data Sources: The private non-residential capital stock was computed from the BEA's NIPA, "Table 2.1. Current-Cost Net Stock of Private Fixed Assets, Equipment, Structures, and Intellectual Property Products by Type" as "Non-residential equipment" + "Nonresidential structures" + "Nonresidential intellectual property products," and divided by nominal GDP. Hence it includes only producible capital, i.e., it excludes the value of land.

in the model variants with a cascade. Furthermore, the simulations other than the $\sigma_b = \sigma_K$ case replicate between a third and 40 percent of the peak of the mortgage-debt-to-GDP ratio, respectively.

Furthermore, I also compare the model's predictions regarding the ratio of non-residential capital stock to GDP with the corresponding BEA estimate (see Figure 8). In the data, the ratio of non-residential capital to GDP does not exhibit a clear trend over the 1980–2016 period but remains below its 1980 value throughout.[19]

[19]That being said, an alternative "market value" estimate of the capital stock based on the market value of corporate equity (see Piketty and Zucman 2014, and Piketty, Saez, and Zucman 2018 for the most recent data) does yield a strong upward trend of the capital-output ratio over the 1980–2016 period, driven mainly by an increase in the difference between the market and book value of corporations, i.e., an increase in Tobin's Q. The "market-value" measure circumvents some practical challenges posed by the perpetual inventory method statistical agencies like the BEA rely on to estimate the value of produced fixed assets, like the need to estimate the replacement costs of fixed assets for which no centralized

The "identical curvature" model predicts an increase of 12 percentage points by 2016, close to its prediction regarding the mortgage-debt-to-GDP ratio for that model, as a result of an identical long-run "wealth effect" on the demand for capital and bank deposits by rich households in this variant of the CSP model. By contrast, the baseline model predicts an increase of less than 3 percentage points due its smaller (larger) wealth effect on rich households' demand for capital (deposits). With a CES technology and/or an active cascade, the predicted increase in capital intensity declines to 2 percentage points or less. Hence the model simulation combines an empirically relevant decline in the natural interest rate with the absence of a trend in the capital-output ratio. Correspondingly, with a CES technology, the ratio of non-residential investment to GDP increases by merely 0.1 to 0.3 percentage point by 2016 (see Table E.2 in Appendix E). By contrast, the model by Straub (2017) predicts an increase in the capital-output ratio of about 25 percentage points over a similar period when the increase in labor income inequality is fed into the model, while at the same time predicting a decline in the natural rate of merely 1 percent. His result may be partially due to the fact that his model abstracts from household borrowing, and producible physical capital is the only asset. Furthermore, my simulation features not merely an increase in wage inequality but also an increase in the markup in order to capture the decline in the labor share. As shown in Section 4.1, in response to a markup shock the CSP model predicts a decline in investment during the first five years, followed by only slight increase in the long run (see Figure 3), even with a Cobb-Douglas technology.

Relatedly, a number of authors have noted that, in spite of the decline in safe interest rates, measures of the return of capital like the one by Gomme, Ravikumar, and Rupert (2015) have remained stable, and perhaps even increased somewhat (e.g., Caballero, Farhi, and Gourinchas 2017; Eggertsson, Robbins, and Wold 2018; Farhi

market exists (see Piketty and Zucman 2014). Unfortunately, it may also fluctuate independently of any physical capital accumulation, as stressed more recently by a team including the same two authors (see Alvaredo et al. 2020). Relatedly, Farhi and Gourio (2018), Corhay, Kung, and Schmid (2020), and Greenwald, Lettau, and Ludvigson (2021) argue that a crucial driver of the aforementioned increase in Tobin's Q is an increase in firm profits at the expense of wages, attributed by Farhi and Gourio (2018) and Corhay, Kung, and Schmid (2020) to rising market power.

and Gourio 2018). Figure F.1 in Appendix F shows that the simulation broadly matches this feature of the data as well, due to the rising price markup.

The models predict a small increase in GDP Y_t as a result of the increase in inequality, driven by an increase in housing consumption CH_t. For instance, the baseline model predicts that by 2016 (i.e., after 36 years), GDP exceeds its level in the absence of the inequality increase by 1.8 percent, or by 1.5 percent with a CES technology (see Table E.1 in Appendix E). In the presence of the consumption cascade, the GDP increase rises to 3.4 percent and 3.0 percent, since the cascade triggers a stronger increase in CH_t, as discussed in Section 4.1. Correspondingly, the predicted effect on average annual GDP growth over the 1980–2016 period ranges between merely 0.04 and 0.09 percentage point. The predicted increase in the share of housing consumption in GDP is broadly in line with the data (see Figure E.2 in Appendix E).

Finally, as a robustness check, I repeat the simulation of the model with a CES production function and consumption cascade for an alternative markup shock series. Specifically, instead of targeting the path of the labor share in the data, I set $d_{\mu,t}$ such that the simulation matches an estimate of the share of pure profits, i.e., profits in excess of the opportunity cost of capital, in firm value-added PS_t, whose model counterpart is given by $PS_t = 1 - \frac{1}{\mu_P + d_{\mu,t}}$. I compute this measure from empirical evidence of De Loecker, Eeckhout, and Unger (2020) (see Appendix G for details). This approach yields a higher $d_{\mu,t}$ trajectory than before, implying that the simulated post-1996 labor share decline is larger than in the data (see Figure 5, the red square), though the gap narrows towards the end of the simulation period. However, the interest rate declines slightly more and the simulation better matches the increase in household indebtedness (at both the macro and the micro level) and the value of the housing stock (see Figures 5 (lower panel) through 7; compare the empty square and triangle)

5. Conclusion

This paper links four empirical trends observed during the post-1980 period: The upward trend in the income share of the top 10 percent of U.S. households, the downward trend in empirical estimates of the natural rate of interest, the simultaneous increase in indebtedness of non-rich households, and the increase in the value of the residential housing stock relative to GDP. For that purpose it develops a model with two household groups, the bottom 90 percent and the top 10 percent, where rich households have capitalist spirit type preferences (CSP) over their wealth. With CSP, an increase in income inequality increases the saving of rich households and lowers the natural rate of interest, while also increasing borrowing by the non-rich, as the non-rich use the decline in the interest rate to postpone the decline in their consumption of goods and housing services. House prices increase as well.

Replicating the increase in the share of the top 10 percent of wage earners in total labor income and the decline in the labor share observed over the 1980–2016 period within the model, I find a decline in the natural rate of a magnitude in line with available empirical estimates of the impact of demand shocks on the natural rate. The simulation also replicates the major part of the increase in the bottom 90 percent debt-to-income ratio and roughly tracks the increase in the bottom 90 percent LTV ratio. At the economy-wide level, it replicates a large part of the upward trend in the value of the housing stock and total household mortgages relative to GDP.

My analysis suggests that the natural rate may remain at its current low level for a long time, as the distribution of income tends to change only slowly, and thus the scope for "conventional" monetary policy may remain limited. Furthermore, to the extent that the tax and transfer system may change the distribution of income in an efficient manner, it implies a potentially important role for fiscal policy in determining the distance of the economy from the zero lower bound (ZLB) and the overall level of household debt in the economy.

Appendix A. Microsimulation Used to Compute Saver Households' MPS

I describe the microsimulation I use to calibrate the wealth curvature parameter for the case of the simple model of Section 2. The procedure in the full model is fully analogous. For the purpose of the microsimulation (or partial equilibrium simulation), I exogenize the real interest rate $\frac{R}{\Pi}$ and denote all income exogenous to the choices of the household as $Y_{S,t}$. $Y_{S,t}$ thus equals the sum of labor income and the profits of the monopolistically competitive firms. Household behavior is then described by

$$b_{S,t} = \frac{R}{\Pi} b_{S,t-1} + Y_{S,t} - C_{S,t} \tag{A.1}$$

$$\Lambda_{S,t} = \frac{1}{C_{S,t}^{\sigma_S}} \tag{A.2}$$

$$\Lambda_{S,t} = \beta_S E_t \left\{ \Lambda_{S,t+1} \frac{R}{\Pi} \right\}) + \phi_b \left(b_{S,t}\right)^{-\sigma_b}. \tag{A.3}$$

I then simulate a permanent increase in exogenous income $Y_{S,t}$ occurring in $t = 1$. I compute the marginal propensity to save (MPS) over a horizon of six years (24 quarters) as

$$MPS_{S,1-24} = \frac{b_{S,24} - b_0}{\sum_{t=1}^{24} \left(Y_{S,t} - Y_{S,0}\right) + \left(b_{S,t} - b_0\right)\left(\frac{R}{\Pi} - 1\right)}. \tag{A.4}$$

The rationale for the six-year horizon is that the saving rate regression of Dynan, Skinner, and Zeldes (2004) from which the value of the MPS I target is computed measures saving as the change in net worth from 1983 to 1989 (see Kumhof, Ranciere, and Winant 2015 for further details on how to compute the MPS in a way consistent with these empirical estimates). Given the calibration of the other parameters as described in Section 2.5, I use σ_b to set $MPS_{S,1-24}$ to the empirical target value.

Note that in the presence of housing and physical capital, the numerator of (A.4) features the change in the value of the holdings of these assets as well, while the denominator features the associated change in the rental income from capital caused by the shock.

Appendix B. The Non-rich National Income Share $NIS_{CC,t}$ in the Full Model

The non-rich national income share is defined as

$$NIS_{CC,t} \equiv \frac{\begin{array}{c} w_{CC,t}N_{CC,t} - b_{CC,t-1}\left(\frac{R_{L,t}}{\Pi_t} - 1\right)\frac{R_{t-1}}{\Pi_t} \\ +\frac{H_{CC,t}}{H_t}CH_t - NIS_{CC,t}b_{gov,t}\left(\frac{R_{t-1}}{\Pi_t} - 1\right) \end{array}}{NI_t},$$

and Equation (48) in the main text is derived by solving for $NIS_{CC,t}$. The numerator represents total pre-tax income of non-rich households and arises as follows. $w_{CC,t}N_{CC,t}$ represents the non-rich households' labor income, while $-b_{CC,t-1}\left(\frac{R_{L,t}}{\Pi_t} - 1\right)$ and $\frac{H_{CC,t}}{H_t}CH_t$ represent their (negative) mortgage income and their share in total imputed rental income (from providing housing services CH_t), respectively. Note that the computation of non-rich households' capital income $-b_{CC,t-1}\left(\frac{R_{L,t}}{\Pi_t} - 1\right)\frac{R_{t-1}}{\Pi_t} + \frac{H_{CC,t}}{H_t}CH_t$, and in particular the allocation total of imputed rental income according to non-rich households' share in total housing $\frac{H_{CC,t}}{H_t}$, is consistent with the way capital income is allocated to adults in the WID. In the WID, for the United States, the share of a group of adults in economy-wide income generated by any asset class corresponds exactly to the share of that group in the total stock of that asset, meaning that their share of imputed rental income across adults equals their share in the total housing stock (see Saez and Zucman 2016 and Piketty, Saez, and Zucman 2018). Finally, the $-NIS_{CC,t}b_{gov,t}\left(\frac{R_{t-1}}{\Pi_t} - 1\right)$ term arises because the WID distributes the property income of the government $-b_{gov,t}\left(\frac{R_{t-1}}{\Pi_t} - 1\right)$ across adults according to their pre-tax national income share. The allocation of the government's property income to individual adults is necessary in order to ensure that across adults, individual pre-tax income adds up to national income.

Appendix C. Additional Validation of the Consumption and Saving Behavior of Households in the Model

This section provides additional validation of the consumption and saving behavior of households in the full model using microevidence. The second line of Table C.1 displays the MPS of rich households

Table C.1. MPS Out of Permanent Income and MPC Out of Wealth in the Full Model and the Data

	Model	Empirical
MPS Out of Permanent Income: Rich	56%	51%
MPS Out of Permanent Income: Non-rich	11%	19%/11%/14%
MPC Out of Wealth: Rich	1.3%	1.2%
MPC Out of Wealth: Non-rich	3.3%	4.3%

Note:

MPS Out of Permanent Income Changes of Non-rich and Rich Households

The empirical counterpart of the rich households MPS out of permanent income changes equals Kumhof, Ranciere, and Winant (2015)'s estimate of the MPS of the top 5 percent of households (see their Appendix III). The nearest available empirical counterpart of the non-rich MPS is the MPS of the bottom 80 percent of households, which I can compute from results in Dynan, Skinner, and Zeldes (2004). The first value is based on Dynan, Skinner, and Zeldes (2004)'s "Median instrumental variable regressions of the saving rate on income using lagged income as an instrument" (Table 5, column 2), which is based on SCF data and measures saving as the change in net worth from 1983 to 1989. This is the saving regression from which Kumhof, Ranciere, and Winant (2015) compute their estimate of the MPS of the top 5 percent. I compute the MPS for each of the first four quintiles q by applying the following formula: $\widehat{MPS}_q = \frac{\hat{\alpha}_q Median(y_q) - \hat{\alpha}_q Median(y_{q-1})}{Median(y_q) - Median(y_{q-1})}$ from Kumhof, Ranciere, and Winant (2015), where $Median\ (y_q)$ and $\hat{\alpha}_q$ denote the 1989 median household income in quintile q (obtained from the SCF, Table 1: Before-tax family income, 1989–98 surveys) and the average saving rate estimated by Dynan, Skinner, and Zeldes (2004) for that quintile, respectively. For the first quintile ($q = 1$), I set $Median(y_{q-1}) = 0$. I then calculate the bottom 80 percent MPS as a weighted average of the $\widehat{MPS}_q$ values across the first four quintiles, using the respective share in total income of the bottom 80 percent of households as a weight. The second and third empirical estimate in the "Non-rich" line I compute from the direct estimates of the MPS reported by Dynan, Skinner, and Zeldes (2004) for each quintile, which are based on PSID data (see the note below their Figure 3, and their Table 9, column 2). I then calculate a weighted average across the first four quintiles, using the median income they report in their Figure 3 to calculate the weights. The model MPS of rich and non-rich households is computed from a partial equilibrium simulation of a permanent-income increase as described in Appendix A.

MPC Out of Wealth

I compute the empirical counterpart of the rich and non-rich household wealth MPC reported above as a simple average across estimates of the wealth MPC of the top 10 percent and bottom 90 percent of households in the wealth distribution from a number of European countries. To the best of my knowledge, there are no U.S. estimates of how the MPC out of wealth varies across the wealth distribution. The estimates and sources are listed in the table below. I computed the bottom 90 percent wealth MPC as a weighted average of the respective MPC estimates for the percentiles encompassed in the bottom 90 percent. I used the share of the respective percentile in total wealth as a weight; see Arrondel, Lamarche, and Savignac (2019), Table 2, and Di Maggio, Kermani, and Majlesi (2020), Table I, financial wealth. I thank Frédérique Savignac for kindly sharing the net wealth totals of the wealth percentiles for which Garbinti et al. (2020) estimate the wealth MPC.

(*continued*)

Table C.1. (Continued)

		MPC Out of Wealth	
Source	**Country**	**Top 10%**	**Bottom 90%**
Arrondel, Lamarche, and Savignac (2019), Table 4	France	0.5%	1.4%
Garbinti et al. (2020), Table 5	Belgium	1.2%	4.8%
	Cyprus	0.3%	2.1%
	Germany	0.6%	3.5%
	Spain	0.8%	4.5%
	Italy	2.3%	5.4%
Di Maggio, Kermani, and Majlesi (2020), Table III	Sweden	2.8%	8.1%

computed from a microsimulation of a permanent income increase (see Appendix A for details) and corresponding empirical evidence computed by Kumhof, Ranciere, and Winant (2015) from Dynan, Skinner, and Zeldes (2004). Recall that this evidence was one of the targets used in Section 3.6.1 to calibrate the wealth utility curvature parameters. The third line reports the same objects for non-rich households in the model and their empirical counterparts, which I computed from Dynan, Skinner, and Zeldes (2004) (see the note below the table for details). The model and empirical values are close, even though I did not use an estimate of the non-rich household MPS as a target for the calibration of the model. Hence my assumption that only rich households have CSP is consistent with the finding that the permanent-income MPS of rich household far exceeds the permanent-income MPS of the rest of the population.

An alternative statistic which one might consider is the rich households' marginal propensity to consume (MPC) out of wealth, which Mian, Straub, and Sufi (2020a) target to calibrate their model with preferences over wealth. Table C.1 reports the MPC computed from a microsimulation of a one-off exogenous increase in rich-household wealth and a one-off exogenous increase in non-rich household wealth. It also reports an average of recent empirical estimates from European countries of the wealth MPC of the top 10 percent and the bottom 90 percent of the wealth distribution, respectively.[20] In the microsimulation for rich households, I assume that the exogenous wealth increase is split between capital (whose adjustment is subject to costs) and other assets according to their respective initial shares in total rich household assets. More formally, the budget constraint and capital accumulation equation used in the microsimulation are given by

$$b_{S,t} = \frac{R}{\Pi} b_{S,t-1} + w_S N_S + r_K K_{t-1} + \Xi - Q_H \left(H_{S,t} - H_{S,t-1}\right) - T_S - C_{S,t} - \left(I_t + K_{t-1} \Phi \left(\frac{I_t}{K_{t-1}}\right)\right) + \left(1 - \frac{K}{H_S Q_H + b_S + K}\right) \epsilon_{wealth,t}$$

$$K_t = (1-\delta) K_{t-1} + I_t + \frac{K}{H_S Q_H + b_S + K} \epsilon_{wealth,t}$$

with $\epsilon_{wealth,1} > 0$ and $\epsilon_{wealth,t} = 1$ for $t > 1$. All variables exogenous to the individual household kept constant. I compute the MPC as the ratio of the first-year average increase in consumption to the first-year average increase in wealth:

$$MPC_{wealth,1Y} = \frac{\sum_{t=1}^{4} \left(C_{S,t} - C_{S,0}\right)}{\sum_{t=1}^{4} \left(b_{S,t} - b_{S,0}\right) + Q_H \left(H_{S,t} - H_{S,0}\right) + \left(K_t - K_0\right)}.$$

As can be seen from Table C.1, the model closely matches the empirical wealth MPC of both rich and non-rich households.[21]

Appendix D. The Long-Run Saving Supply Curve of Rich Households

This appendix derives the slope of the long-run "saving supply curve" of rich households and explains how it differs from Mian,

[20] To the best my knowledge, there are no U.S. estimates of how the MPC out of wealth varies across the wealth distribution.

[21] Garbinti et al. (2020) differ from the other studies in that they use the four-year change in consumption and wealth as dependent and independent variable, respectively. The model counterpart is given by $MPC_{wealth,4Y} = \frac{\sum_{t=13}^{16} dC_{S,t}}{\sum_{t=13}^{16} db_{S,t} + Q_H dH_{S,t} + dK_t}$. However, $MPC_{wealth,4Y}$ is with 1.46 percent very close to $MPC_{wealth,1Y}$.

Straub, and Sufi (2020a). Abstracting for simplicity and easier comparability with Mian, Straub, and Sufi (2020a) from physical capital and housing, the steady-state budget constraint and safe assets first-order condition of rich households are given by

$$C_S = \left(\frac{R}{\Pi} - 1\right) b_S + w_S N_S + \Xi - T_S$$

$$1 = \beta_S \frac{R}{\Pi} + \frac{\phi_b b_S^{-\sigma_b}}{C_S^{-\sigma_S}},$$

where $\phi_b b_S^{-\sigma_b}$ and $C_S^{-\sigma_S}$ represent the marginal utility of safe assets and consumption, respectively. Accordingly, after substituting the first equation into the second equation, the slope of the steady-state saving supply curve can be derived using implicit differentiation:

$$\frac{db_S}{d\frac{R}{\Pi}} = -\frac{\beta_S + \sigma_S \phi_b \left(\left(\frac{R}{\Pi} - 1\right) b_S + w_S N_S + \Xi - T_S\right)^{\sigma_S - 1} (b_S)^{1-\sigma_b}}{\frac{\phi_b\left(\left(\frac{R}{\Pi}-1\right)b_S + w_S N_S + \Xi - T_S\right)^{\sigma_S}}{b_S^{\sigma_b}} \left[\frac{\sigma_S\left(\frac{R}{\Pi}-1\right)}{\left(\frac{R}{\Pi}-1\right)b_S + w_S N_S + \Xi - T_S} - \frac{\sigma_b}{b_S}\right]}. \tag{D.1}$$

Hence $\frac{db_S}{d\frac{R}{\Pi}} < 0$ (as in Mian, Straub, and Sufi 2020a) iff $\sigma_S > \sigma_b \frac{\left(\left(\frac{R}{\Pi}-1\right)b_S + w_S N_S + \Xi - T_S\right)}{\left(\frac{R}{\Pi}-1\right)b_S}$. For this inequality to hold, the utility curvature of consumption has to be smaller than that of wealth (i.e., $\sigma_S > \sigma_b$). This condition is met both in my calibration and in Mian, Straub, and Sufi (2020a). However, on top of that, the income derived from the utility-generating asset (here $\left(\frac{R}{\Pi} - 1\right) b_S$) has to be sufficiently large relative to income from alternative sources (here $w_S N_S + \Xi - T_S$). In Mian, Straub, and Sufi (2020a), income from sources other than the asset is zero.[22] Therefore $C_S = \left(\frac{R}{\Pi} - 1\right) b_S$, implying that a 1 percent increase in b_S increases C_S by 1 percent as well. With $\sigma_b < \sigma_S$, the marginal utility of consumption $C_S^{-\sigma_S}$

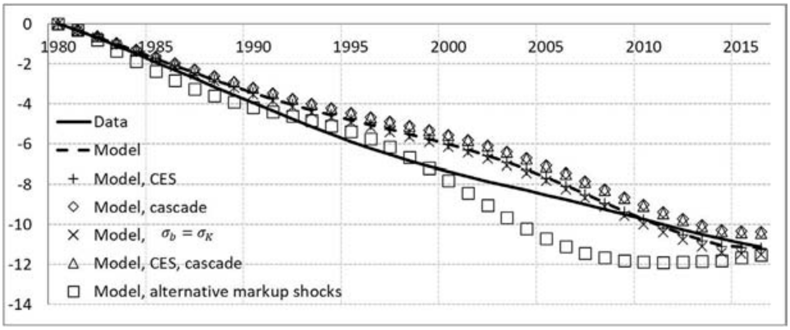
[22]In their baseline model rich household income is generated by lending to non-rich households and a "Lucas tree," with households deriving utility from the combined value of these two assets. In the variant of the model with labor and capital, the measure of wealth-generating utility equals the sum of lending to non-rich households, the value of the capital stock, and the present value of future labor income ("human capital"). Hence, the wealth concept in the utility function is always the capitalized income from all sources, including labor income.

declines by a larger percentage than the marginal utility from bonds $\phi_b b_S^{-\sigma_b}$ (the "income effect" dominates the "substitution effect"), implying that rich households would like to hold *even more* safe assets at the original interest rate. Hence, for the bond market to return to equilibrium after an increase in the supply of b_S, $\frac{R}{\Pi}$ has to decline to make bonds relatively less attractive, and to limit the consumption increase and hence the decline in marginal utility of consumption $C_S^{-\sigma_S}$. By contrast, in my model, the inequality does not hold due to the presence of substantial labor income $w_S N_S$, implying that the impact of a 1 percent increase in b_S increases C_S by much less than 1 percent, generating a much weaker "income effect." Therefore the saving-supply curve slopes upward.

This result extends to the case with physical capital. Specifically, for my calibration, an increase in any of the two rates of return $\frac{R}{\Pi}$ and r_K increases rich households steady-state demand for both safe assets and capital (i.e., $\frac{\partial b_S}{\partial \frac{R}{\Pi}}, \frac{\partial K}{\partial \frac{R}{\Pi}}, \frac{\partial b_S}{\partial r_K}, \frac{\partial K}{\partial r_K} > 0$). The positive long-run cross-effects (i.e., $\frac{\partial K}{\partial \frac{R}{\Pi}}, \frac{\partial b_S}{\partial r_K} > 0$) are the result of an "income effect." For instance, an increase in $\frac{R}{\Pi}$ increases the households interest income and thus eventually consumption, thereby increasing the marginal utility of capital relative to the marginal utility of consumption.

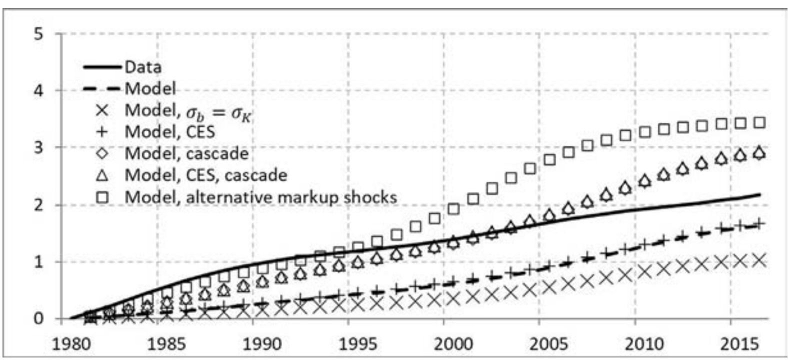
Appendix E. Further Results from the Historical Simulation of the Full Model

Figure E.1. Simulation 1981–2016: Bottom 90 Percent National Income Share



Note: This graph displays the share of the bottom 90 percent of households in national income $NIS_{CC,t}$ (see Equation (48)). See the note below Figure 5 for details on the model variants used in the simulation, the simulation setup, data treatment, and data sources.

Figure E.2. Simulation 1981–2016: Share of Housing Consumption in GDP



Note: This graph displays the share of housing consumption in GDP $\frac{CH_t}{Y,t} * 100$ and its empirical counterpart. See the note below Figure 5 for details on the model variants used in the simulation, the simulation setup, and data treatment. **Data Sources:** BEA, NIPA Table 2.5.5, Personal Consumption Expenditures by Function, line 19 "Housing."

Table E.1. Simulation 1981–2016: 2016 Values of Selected Variables

	Baseline	$\boldsymbol{\sigma_b = \sigma_K}$	**CES**	**Cascades**	**CES, Cascades**	**Alt. Markup Shocks**
Non-rich Share in Labor Income	–10.8	–10.8	–10.8	–10.8	–10.8	–10.8
Labor Share	–6.6	–6.6	–6.5	–6.6	–6.5	–8.3
Interest Rate	–3.5	–2.8	–3.6	–2.8	–2.8	–3.2
Borrower Debt-to-Income Ratio	62.8	44.0	64.0	69.5	70.7	90.7
Loan-to-Value Ratio	23.8	19.4	24.5	17.1	17.8	23.7
Housing Stock-to-GDP Ratio	17.8	7.0	17.8	39.0	38.7	42.6
Household Debt-to-GDP Ratio	21.1	13.5	21.6	23.9	24.3	30.9
Non-residential Capital-to-GDP Ratio	3.4	12.2	2.1	1.4	0.1	0.4
Non-residential-Investment-to-GDP Ratio	0.5	1.8	0.3	0.3	0.1	0.2
GDP	1.8	3.5	1.5	3.4	3.0	3.3
GDP Excluding Housing Consumption	0.0	2.3	–0.4	0.1	–0.3	–0.6

Note: This table displays the 2016 values of selected model variables implied by the simulation of Section 4.2. The column labels indicate the simulation variant. For details on the same, see the note below Figure 5. All variables are expressed as deviations from their 1980 values, with the exception of GDP, which is expressed as a percentage deviation. GDP or income appearing in the denominator of ratios is annualized. Note that GDP is defined by Equation (49), while "GDP Excluding Housing Consumption" is defined as $C_{S,t} + C_{CC,t} + I_t + G_t$.

Table E.2. Simulation 1981–2016: New Steady-State Values of Selected Variables

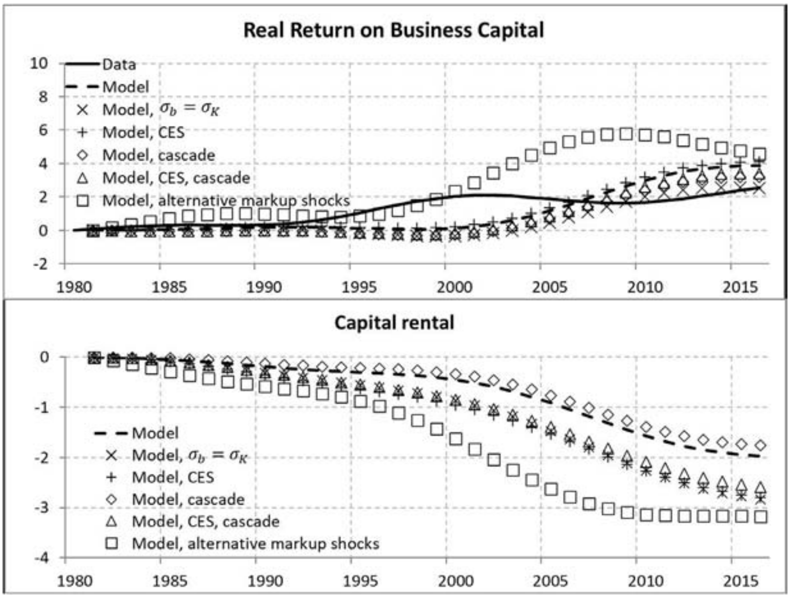
	Baseline	$\sigma_b = \sigma_K$	CES	Cascades	CES, Cascades	Alt. Markup Shocks
Non-rich Share in Labor Income	–10.8	–10.8	–10.8	–10.8	–10.8	–10.8
Labor Share	–6.5	–6.6	–6.1	–6.5	–6.1	–8.0
Interest Rate	–4.0	–3.1	–3.9	–3.2	–3.1	–3.5
Borrower Debt-to-Income Ratio	78.2	53.8	76.6	88.7	86.5	100.6
Loan-to-Value Ratio	33.2	26.1	32.8	26.5	26.1	28.8
Housing Stock-to-GDP Ratio	17.5	7.3	17.2	38.2	37.6	42.2
Household Debt-to-GDP Ratio	27.5	17.5	27.1	31.7	31.2	35.2
Non-residential Capital-to-GDP Ratio	5.3	19.0	3.3	3.5	1.4	1.3
Non-residential-Investment-to-GDP Ratio	0.5	1.9	0.3	0.3	0.1	0.1
GDP	2.2	5.0	1.6	3.8	3.2	3.5
GDP Excluding Housing Consumption	–0.1	3.6	–0.6	0.0	–0.5	–0.7

Note: This table displays the new steady-state values of selected model variables implied by the simulation of Section 4.2. The column labels indicate the simulation variant. For details on the same, see the note below Figure 5. All variables are expressed as deviations from their 1980 values, with the exception of GDP, which is expressed as a percentage deviation. GDP or income appearing in the denominator of ratios is annualized. Note that GDP is defined by Equation (49), while "GDP Excluding Housing Consumption" is defined as $C_{S,t} + C_{CC,t} + I_t + G_t$.

Appendix F. The Return on Business Capital in the Model and the Data

A number of authors have noted that, in spite of the decline in safe interest rates, measures of the return on capital have remained stable, and perhaps even increased somewhat (e.g., Caballero, Farhi, and Gourinchas 2017; Brand, Bielecki, and Penalver 2018; Eggertsson, Robbins, and Wold 2018; and Farhi and Gourio 2018). Figure F.1 confirms this finding by displaying the widely cited measure of Gomme, Ravikumar, and Rupert (2015) (see the black solid line). All model simulations succeed in avoiding a decline in the return on business capital, though the simulated increase is somewhat delayed relative to the data. By contrast, the capital rental rate $r_{K,t}$ declines. The reason for this divergence is the rise in pure profits caused by the increase in the markup (see also the discussion of the effect of a markup shock in Section 4.1). My analysis is similar to that of Caballero, Farhi, and Gourinchas (2017), Eggertsson, Robbins, and Wold (2018), and Farhi and Gourio (2018) in that they also rely on an increase in the price markup to match the decline in the labor share and to prop up the return on capital. On top of that, in my model, CSP allows the markup increase to contribute to the decline in the risk-free rate via an increase in saving. By contrast, Caballero, Farhi, and Gourinchas (2017), Eggertsson, Robbins, and Wold (2018), and Farhi and Gourio (2018) have to rely on separate exogenous forces to match this decline, namely an increase in the household discount factor and an increase in risk.

Figure F.1. Simulation 1981–2016: Real Return on Business Capital and Real Capital Rental



Note: This graph plots the simulated path of the annualized real return on business capital (RRBC) $\frac{(1-mc_t)Y_{f,t}+(r_{K,t}-\delta)K_{t-1}}{Q_tK_{t-1}}400$, an empirical counterpart proposed by Gomme, Ravikumar, and Rupert (2015), and the simulated path of the annualized capital rental $400r_{K,t}$. See the note below Figure 5 for details on the model variants used in the simulation, the simulation setup, and data treatment. RBBC: real return on business capital.
Data Sources: Pre-tax return on business capital of Gomme, Ravikumar, and Rupert (2015) (see their Figure 2). I downloaded the updated series from Paul Gomme's webpage.

Appendix G. The Pure Profit Share in the Historical Simulation and an Alternative Parameterization of $d_{\mu,t}$

This appendix compares the simulated share of pure profits in firm value-added PS_t with empirical estimates, and also provides details on how the alternative series of markup shocks $d_{\mu,t}$ used in the robustness experiment discussed at the end of Section 4.2 is computed. "Pure profits" refers to the profits resulting from monopolistic competition, the value obtained after subtracting not just labor costs but also the (opportunity) costs of capital $r_{K,t}K_{t-1}$ from firm value-added. Hence in the model, PS_t is defined as

$$PS_t \equiv \frac{Y_{f,t} - w_{S,t}N_{S,t} - w_{CC,t}N_{CC,t} - r_{K,t}K_{t-1}}{Y_{f,t}} = 1 - \frac{1}{\mu_P + d_{\mu,t}}, \tag{G.1}$$

where the final expression is derived using Equations (37)–(42). Figure G.1 plots the simulated path of PS_t arising from my baseline simulation setup, i.e., where I invert $d_{\mu,t}$ by imposing that the labor share in the model matches the path of the trend labor share in the data as plotted in Figure 5, and two empirical estimates of PS_t. Barkai (2020) estimates the share of pure profits of the U.S. non-financial corporate sector in its gross value-added over the period 1984 to 2014 by estimating its cost of capital from BEA capital stock data and a financial-market-based estimate of the required return on capital. For their sample of publicly traded firms, De Loecker, Eeckhout, and Unger (2020) report an estimate of the average (revenue-weighted) profit rate, i.e., pure profits as a fraction of sales over the 1981–2016 period, which I convert to a fraction of value-added by multiplying it with the economy-wide gross-output-to-value-added ratio, since gross output measures economy-wide revenue.[23]

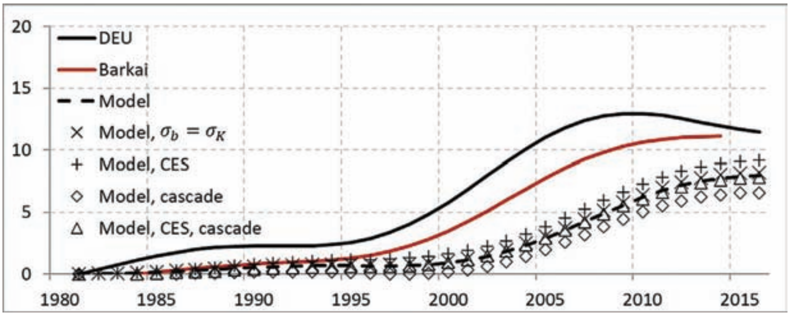
Both the estimates of Barkai (2020) and De Loecker, Eeckhout, and Unger (2020) display a gradual increase of about 1 and 2.5 percentage points, respectively, until about 1995, followed by a much

[23]Note that the validity of this conversion is *not* impaired by heterogeneity in the sales-to-value-added ratio of the firms in the De Loecker, Eeckhout, and Unger (2020) sample, since they report a *revenue-weighted* average of the profit rate. Therefore the value they report in fact equals the ratio of total profits to total sales in the sample. More formally, and using their notation, the profit rate of firm i $\pi_{i,t}$ is defined as $\pi_{it} \equiv \frac{\Pi_{i,t}}{S_{i,t}}$, where $\Pi_{i,t}$ and $S_{i,t}$ denote profits and sales, respectively. The revenue-weighted sample average of the profit rate π_t is then given by

$$\pi_t = \sum_i \left(\frac{S_{i,t}}{S_t}\right)\pi_{it} = \sum_i \frac{S_{i,t}}{S_t}\frac{\Pi_{i,t}}{S_{i,t}} = \frac{\sum_i \Pi_{i,t}}{S_t} = \frac{\Pi_t}{S_t},$$

where $\Pi_t = \sum_i \Pi_{i,t}$ and $S_t = \sum_i S_{i,t}$ denote total profits and sales, respectively. Hence I can compute the share of total profits in total value-added as $PS_t = \pi_t\left(\frac{S_t}{Y_{f,t}}\right)$. Since De Loecker, Eeckhout, and Unger (2020) do not report a $\frac{S_t}{Y_t}$ time series for their sample, I compute it as the ratio of All-industry Gross Output from the BEA's "GDP by industry" account to GDP. The ratio fluctuates between 1.7 and 1.9 over the 1980–2016 period.

Figure G.1. Simulation 1981–2016: Pure Profit Share



Note: The figure displays the path of PS_t (see Equation (G.1)) for the various model variants, and two empirical estimates. The series labeled "Barkai" is obtained from Barkai (2020) (his Figure 3B), who estimates the share of pure profits of the U.S. non-financial corporate sector in its gross value-added over the period 1984 to 2014 as the residual obtained by subtracting the cost of capital and labor from gross value-added. He estimates the cost of capital from BEA capital stock data and a financial-market-based estimate of the required return on capital. The series labeled DEU is computed from De Loecker, Eeckhout, and Unger (2020) (their Figure VIII). De Loecker, Eeckhout, and Unger (2020) report an estimate of the average (revenue-weighted) profit rate, i.e., pure profits as a fraction of sales, over the 1981–2016 period for their firm-level data set of listed companies. Their computation of profits subtracts variable costs, the opportunity costs of capital, and overhead costs from firm revenues (see their Equation (13) for details). I convert it to a fraction of value-added by multiplying it with the economy-wide gross-output-to-GDP ratio, which I computed from the BEA's "Industry Economic Account Data: GDP by Industry." I remove fluctuations with an amplitude of 2 to 16 years from all data series using the asymmetric full-sample band-pass filter of Christiano and Fitzgerald (2003), assuming a unit root with drift.

larger and steeper increase until the end of their respective samples. The PS_t trajectory in my simulation has a shape similar to the two empirical measures and tracks the estimate of Barkai (2020) closely until the second half of the 1990s. However, its inflection point is somewhat delayed compared with the two empirical measures, and the overall increase in PS_t is smaller.

Therefore, as a robustness check, I repeat the simulation of the empirical increase in inequality from Section 4.2, but now choose $d_{\mu,t}$ such that the increase in PS_t equals the path I computed from De Loecker, Eeckhout, and Unger (2020)'s results, implying that the price markup increases by more than when I target the labor share.

I limit the discussion here to the model variant with a consumption cascade and a CES technology; additional results are available upon request. The direct consequence of the stronger increase in $d_{\mu,t}$ is that the simulated post-1996 labor share decline is now larger than in the data (see Figure 5, top panel, the red square), though the difference becomes smaller towards the end of the simulation period. However, the simulated decline in the natural rate is slightly larger, and the simulation performs better at matching the increase in household indebtedness and the value of the housing stock (see Figures 5 (lower panel) through 7; compare the empty square and triangle).

Appendix H. Results with CSP Over Housing

This section examines the robustness of the results discussed in the main text to assuming that in the CSP model rich households derive utility from the market value of their housing stock $Q_{H,t}\tilde{H}_{S,t}$. Thus with CSP, the objective of rich households becomes

$$E_t\left\{ \sum_{i=0}^{\infty} \beta_S^i \left[\frac{\tilde{C}_{S,t+i}^{1-\sigma_S}}{1-\sigma_S} + \frac{\tilde{\phi}_{H,S}}{1-\sigma_{H,S}} \left(Q_{H,t+i}\tilde{H}_{S,t+i} \right)^{1-\sigma_{H,S}} \right.\right.$$
$$\left.\left. + \frac{\phi_{\tilde{b}}}{1-\sigma_b} \left(\tilde{b}_{S,t+i} \right)^{1-\sigma_b} + \frac{\tilde{\phi}_K}{1-\sigma_K} \left(Q_{t+i}\tilde{K}_{t+i} \right)^{1-\sigma_K} \right] \right\} \quad \text{(H.1)}$$

with $\tilde{\phi}_{H,S}, \tilde{\phi}_b, \tilde{\phi}_K > 0$. The economic interpretation of utility depending on $Q_{H,t}\tilde{H}_{S,t}$ is that the capitalist spirit motive extends to housing, and thus, analogously to the non-residential capital stock, it is the real *value* of the housing stock that matters for utility. By contrast, in Equation (19) in the main text, housing represents simply a durable consumption good, and thus utility depends on "the size and quality of the house," in line with the approach adopted by the literature on housing in DSGE models (Iacoviello 2005, 2015; Iacoviello and Neri 2010; Clerc et al. 2015), but not its price.

The only first-order condition affected by this change is the one for housing,

$$Q_{H,t} = \frac{Q_{H,t}^{1-\sigma_{H,S}} \phi_{H,S}}{H_{S,t}^{\sigma_{H,S}} \Lambda_{S,t}} + \beta_S E_t \left\{ \frac{\Lambda_{S,t+1}}{\Lambda_{S,t}} Q_{H,t+1} \right\}, \quad \text{(H.2)}$$

where the marginal utility from housing is now given by $\frac{Q_{H,t}^{1-\sigma_{H,S}}\phi_{H,S}}{H_{S,t}^{\sigma_{H,S}}}$ and thus with $\sigma_{H,S} > 1$ (which is what I assume) depends negatively on the house price, an effect absent from the model of the main text (see Equation (26)). Correspondingly, rich households' imputed rental income (or housing consumption) becomes

$$CH_{S,t} = \frac{Q_{H,t}^{1-\sigma_{H,S}}\phi_{H,S}}{\Lambda_{S,t}H_{S,t}^{\sigma_{H,S}}}H_{S,t}. \tag{H.3}$$

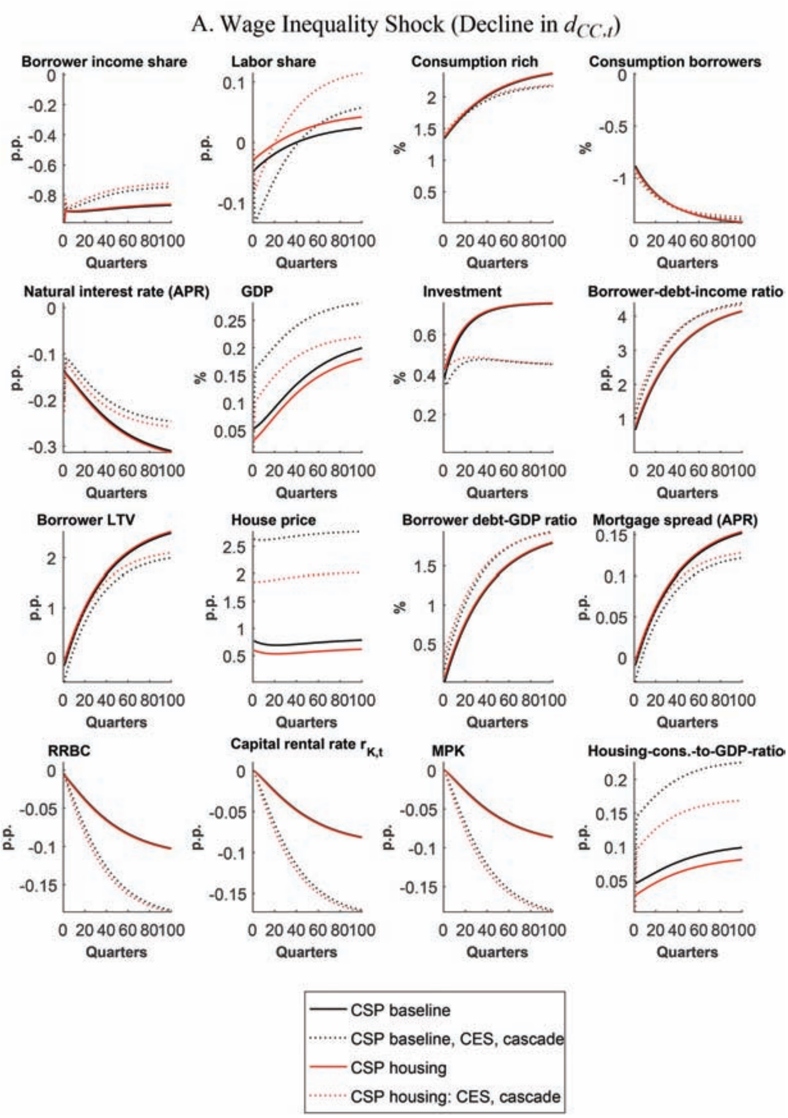
As can be obtained from Figure H.1, assuming that the capitalist spirit motive extends to housing (the lines labeled "CSP housing," in red) has a marginal impact on most variables. The exception is the house price, which increases by between one-tenth and one-quarter less than if the capitalist spirit motive does not extend to housing (the lines labeled "CSP baseline," in black), as I assume in the main text. The reason for the lower house price in the "CSP housing" model follows directly from the just-mentioned direct negative effect of the house price on the marginal utility from housing of rich households. This negative effect also dampens the increase housing consumption with respect to the baseline CSP case, implying a smaller increase in GDP Y_t.

Figures H.2–H.4 compare the results of the historical simulation in the model with baseline CSP and the model with housing CSP. Again the trajectories of the displayed variables are mostly very close across the two CSP specifications, except for the predicted increase in the value of the housing stock, which by 2016 is about one-fifth smaller with housing CSP than with baseline CSP (see Figure H.4).

In line with the discussion above, the increase in the share of housing consumption in GDP is smaller with housing CSP.

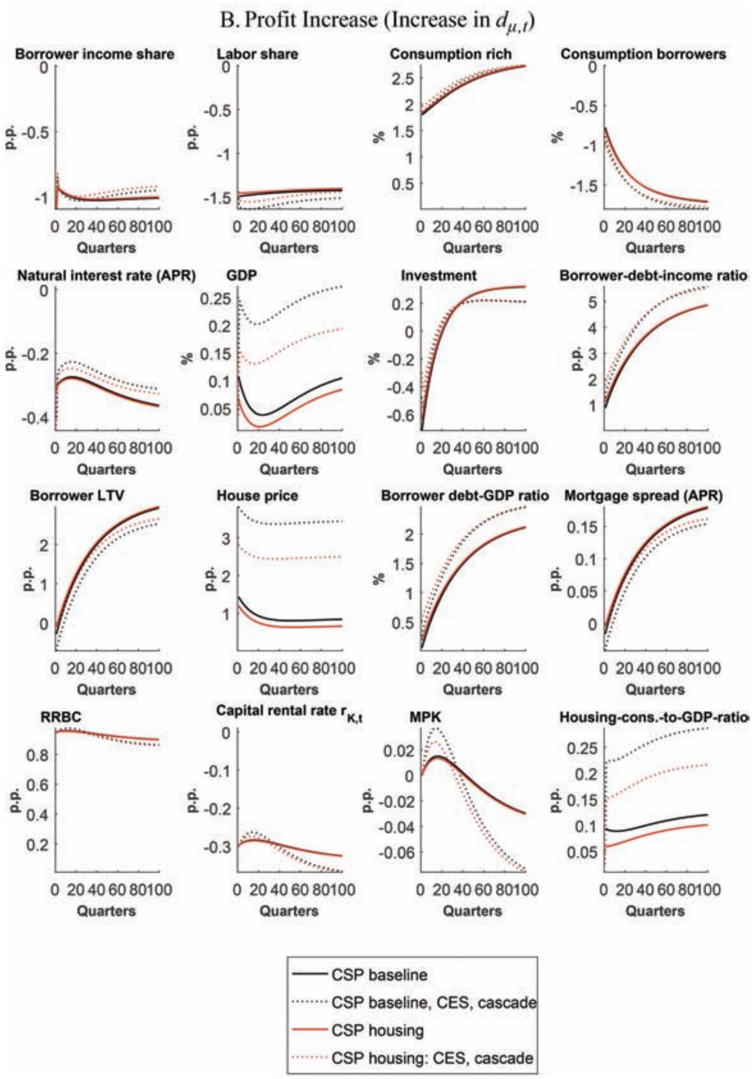
The assumption made with Equation (H.1) that, given the values of the other arguments of the utility function, rich households care *only* about the market value of their house, implying that they would be indifferent with respect to an arbitrarily large decline in their physical housing stock $\tilde{H}_{S,t+i}$ as long as its market value $Q_{H,t+i}\tilde{H}_{S,t+i}$ remains unchanged, is clearly extreme. A more plausible alternative would be to posit that utility depends on a Cobb-Douglas aggregate of the physical housing stock and its market value, i.e., $\left(Q_{H,t+i}\tilde{H}_{S,t+i}\right)$ in (H.1) would be replaced

Figure H.1. Impact of a Permanent Increase in Inequality: Role of Loan Supply Assumption



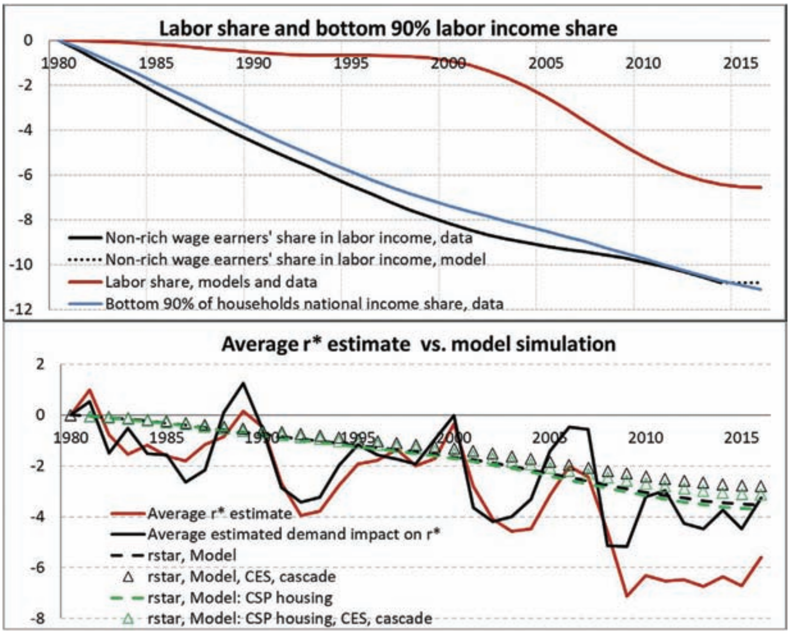
(*continued*)

Figure H.1. (Continued)



Note: The black lines ("CSP baseline") refer to results computed using the model of Section 3. The red lines ("CSP housing") refer to a modified version of the model where rich households derive utility from the market value of the housing stock (see Equation (19)). "cascade" indicates that the model allows for an effect of rich household total consumption on non-rich housing demand (i.e., $\nu_{cascade} > 0$; see Equation (35)). The safe interest rate R_t and the risk spread $E_t R_{L,t+1} - R_t$ are expressed as annualized percentage rates (APR). The borrower debt-to-income ratio is based on annualized borrower income. The borrower debt-to-GDP ratio is based on annualized GDP.

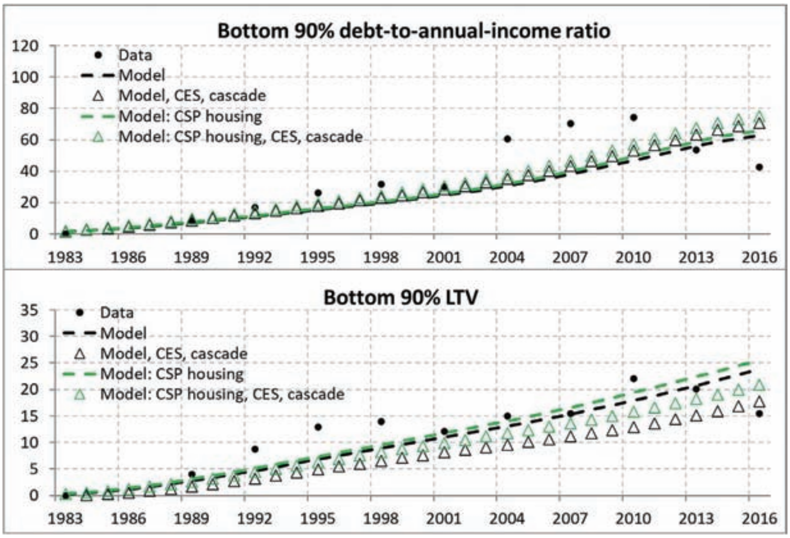
Figure H.2. Simulation 1981–2016: Income Distribution and Natural Interest Rate



Note: The label "Model: CSP housing" refers to a modified version of the model where rich households derive utility from the market value of the housing stock (see Equation (19)). For details on the meaning of the other labels, see the note below Figure 5.

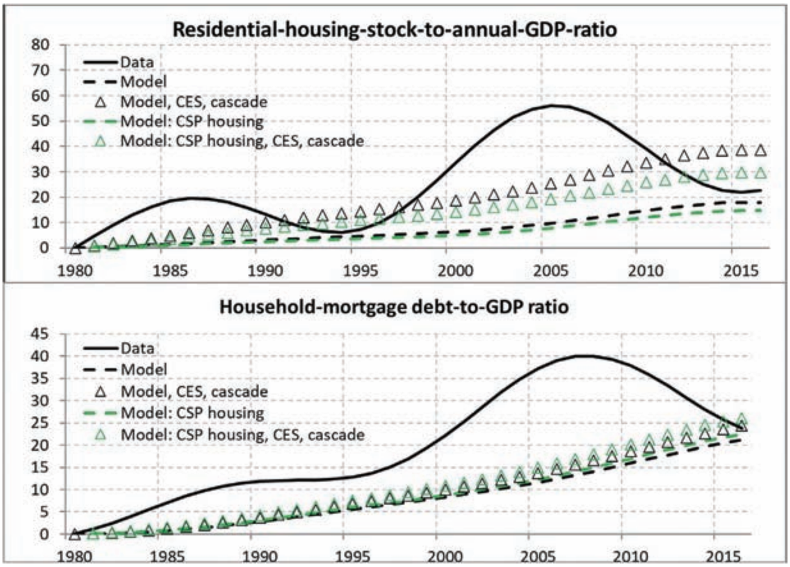
by $\left(Q_{H,t}\tilde{H}_{S,t}\right)^{\alpha_{MV}}\tilde{H}_{S,t}^{1-\alpha_{MV}} = Q_{H,t}^{\alpha_{MV}}\tilde{H}_{S,t}$, with $0 \leq \alpha_M \leq 1$. For $0 < \alpha_{MV} < 1$, simulation results (not shown) are in between those obtained for the baseline CSP and the housing CSP case.

Figure H.3. Simulation 1981–2016: Borrower Debt-to-Income Ratio and LTV



Note: See the note below Figure H.2 for details on the labels reading "Model . . .". See the note below Figure 6 for details on the data sources.

Figure H.4. Simulation 1981–2016: Housing Stock and Total Mortgage Debt-to-GDP Ratios

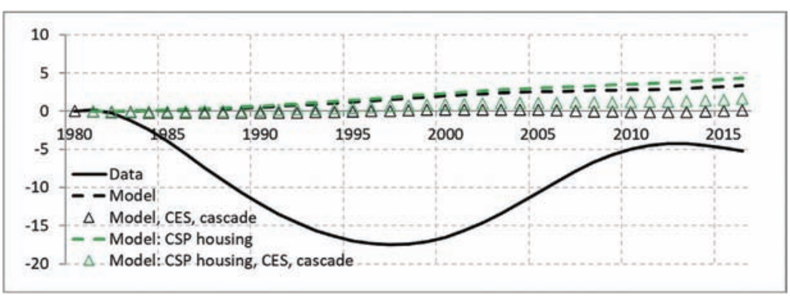


Note: See the note below Figure H.2 for details on the labels reading "Model . . .". See the note below Figure 7 for details on the data sources.

Appendix I. Steady-State Solution of the Full Model in (Almost) Closed Form

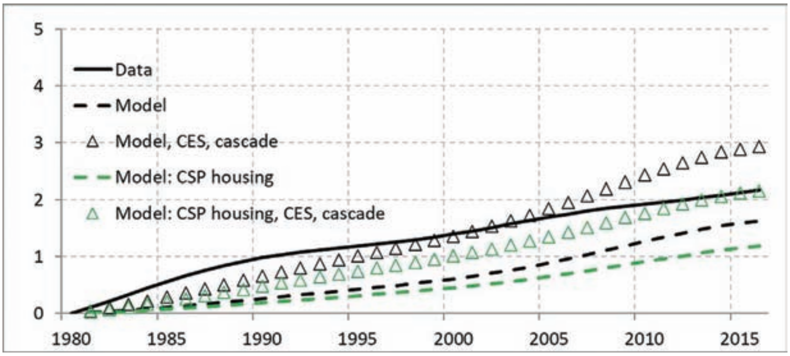
This appendix shows how to compute the steady state of the full model in almost closed form, given 13 of the 15 empirical targets listed in Table 4 (i.e., all targets except for the MPS of the rich and the effect of a permanent increase in the supply of government bonds on the safe interest rate $\frac{d\frac{R}{\Pi}}{d\text{Target}_{bgov2GDP}}$), and using 13 of the 15 parameters whose values are marked with an asterisk (*) in Table 3 (i.e., $\sigma_S/\sigma_{CC}, \beta_S, \beta_{CC}, \phi_{H,S}, \phi_{H,CC}, \phi_b, \phi_K, \chi_{CC}, \alpha_K, \mu_P, \omega_{CC}, Target_{G2GDP}, Target_{bgov2GDP}$) to support those targets. For that purpose, in the first iteration of the computation, I will set an (intermediate) target for the non-rich households' LTV $\left(\frac{b_{CC}}{Q_H H_{CC}}\right)$ instead of their debt-to-annual income ratio $\left(\frac{b_{CC}}{4w_{CC}N_{CC}}\right)$. Furthermore, in order to render the steady state identical regardless of whether I assume a CES or Cobb-Douglas production function, I assume $A_N = \frac{Y_f}{N}$ and $A_K = \frac{Y_f}{K}$.

Figure I.1. Simulation 1981–2016: Private Non-residential Producible Capital-Stock-to-GDP Ratio



Note: See the note below Figure E.2 for details on the labels reading "Model . . .". See the note below Figure 8 for details on the data sources.

Figure I.2. Simulation 1981–2016: Share of Housing Consumption in GDP



Note: See the note below Figure E.2 for details on the labels reading "Model . . .". See the note below Figure 7 for details on the data sources.

From the target for the IES, the definitions of θ and ep and the targets for θ, $\frac{R}{\Pi}$, mp, and ep follows:

$$\sigma_S = \sigma_{CC} = \frac{1}{IES} \tag{I.1}$$

$$\beta_S = \frac{\theta}{\left(\frac{R}{\Pi}\right)} \tag{I.2}$$

$$R_L = R\,(1 + ep) \tag{I.3}$$

$$r_K = (1 + ep)\,\frac{R}{\Pi} + \delta - 1. \tag{I.4}$$

From the targets for $\left(\frac{Q_H H}{4Y}\right)$ and $\left(\frac{H_{CC}}{H}\right)$,

$$\frac{Q_H H_{CC}}{Y} = \left(\frac{Q_H H}{4Y}\right) 4 \left(\frac{H_{CC}}{H}\right) \tag{I.5}$$

$$\frac{Q_H H_S}{Y} = \left(\frac{Q_H H}{4Y}\right) 4 \left(\frac{H_S}{H}\right). \tag{I.6}$$

Furthermore, multiplying the target for $\left(\frac{b_{CC}}{Q_H H_{CC}}\right)$ and $\frac{Q_H H_{CC}}{Y}$, we have

$$\frac{b_{CC}}{Y} = \left(\frac{b_{CC}}{Q_H H_{CC}}\right)\left(\frac{Q_H H_{CC}}{Y}\right). \tag{I.7}$$

Using Equations (29), (34), and (32) then allows to back out the slope of the loan supply curve χ_{CC}, $\left(\frac{dR_L}{db_{CC}}(..)\, b_{CC}\right)$, and β_{CC}, and to calculate share of the cost of the financial friction in output $\frac{b_{CC}}{Y} f(..)\, R$:

$$\chi_{CC} = \left(\frac{R_L}{R} - 1\right)\frac{1}{\left(\frac{b_{CC}}{Q_H H_{CC}}\right)} \tag{I.8}$$

$$\frac{dR_L}{db_{CC}}(..)\, b_{CC} = R\chi_{CC}\left(\frac{b_{CC}}{Q_H H_{CC}}\right) \tag{I.9}$$

$$\beta_{CC} = \frac{1}{\frac{R_L}{\Pi} + \frac{\left(\frac{dR_L}{db_{CC}}(..)b_{CC}\right)}{\Pi}} \tag{I.10}$$

$$\frac{b_{CC}}{Y} f(..)\,\frac{R}{\Pi} = \frac{\frac{dR_L}{db_{CC}}(..)\, b_{CC}}{\Pi}\left(\frac{b_{CC}}{Y}\right). \tag{I.11}$$

Furthermore, rearranging the first-order conditions with respect to housing (26) and (33) allows to find the share of housing consumption in GDP $\left(\frac{CH}{Y}\right)$ as follows:

$$Q_H H_S (1-\beta_S) = \frac{\phi_{H,S} H_S^{1-\sigma_{H,S}}}{\Lambda_S} = CH_S$$

$$H_{CC} Q_H (1-\beta_{CC}) = \frac{\phi_{H,CC}}{\Lambda_{CC}} H_{CC}^{1-\sigma_{H,CC}} - \beta_{CC}\frac{\frac{dR_L}{dH_{CC}}\left(\frac{b_{CC}}{H_{CC}Q_H}\right)}{\Pi} b_{CC} = CH_{CC}$$

or

$$\frac{CH_S}{Y} = \left(\frac{Q_H H_S}{Y}\right)(1-\beta_S) \tag{I.12}$$

$$\frac{CH_{CC}}{Y} = \left(\frac{H_{CC} Q_H}{Y}\right)(1-\beta_{CC}) \tag{I.13}$$

$$\frac{CH}{Y} = \left(\frac{CH_S}{Y}\right) + \left(\frac{CH_{CC}}{Y}\right). \tag{I.14}$$

From (43), we have the ratio of GDP to firm output as

$$\frac{Y}{Y_f} = \frac{1}{1 - \frac{CH}{Y} + \frac{b_{CC}}{Y} f\left(LTV\right) \frac{R}{\Pi}}. \tag{I.15}$$

I can now calculate the economy's supply side. From Equation (21) and the target for $\left(\frac{I}{Y}\right)$,

$$\frac{K}{Y_f} = \frac{\left(\frac{I}{Y}\right)}{\delta}\left(\frac{Y}{Y_f}\right). \tag{I.16}$$

From (39) and (40), and using $A_N = \frac{Y_f}{N}$ and $A_K = \frac{Y_f}{K}$,

$$\begin{aligned}\left(\frac{wN}{Y}\right)\left(\frac{Y}{Y_f}\right) &= \frac{w_S N_S + w_{CC} N_{CC}}{Y_f} \\ &= \frac{mc\left(1-\alpha_K\right)\left(\frac{Y_f}{A_N N}\right)^{\frac{1}{\epsilon}}\left[\left(1-\omega_{CC}\right)NA_N + \omega_{CC} NA_N\right]}{Y_f} \\ &= mc\left(1-\alpha_K\right).\end{aligned}$$

Combining this expression with (38) to eliminate mc and using (42) allows to back out α_K and μ_P, and finally mc:

$$\alpha_K = \frac{r_K \frac{K}{Y_f}}{\left(\frac{wN}{Y}\right)\left(\frac{Y}{Y_f}\right) + r_K \frac{K}{Y_f}} \tag{I.17}$$

$$\mu_P = \frac{1-\alpha_K}{\left(\frac{wN}{Y}\right)\left(\frac{Y}{Y_f}\right)} \tag{I.18}$$

$$mc = \frac{1}{\mu_P}. \tag{I.19}$$

Total firm output for the CES and Cobb-Douglas case are given by

$$Y_f = \left[(1-\alpha_K)\left(A_N N\right)^{\frac{\epsilon-1}{\epsilon}} + \alpha_K \left(A_K K\right)^{\frac{\epsilon-1}{\epsilon}}\right]^{\frac{\epsilon}{\epsilon-1}}$$

$$Y_f = \left(A_N N\right)^{1-\alpha_K}\left(A_K K\right)^{\alpha_K}.$$

Substituting my assumptions $A_N = \frac{Y_f}{N}$ and $A_K = \frac{Y_f}{K}$ into either equation yields $Y_f = Y_f$, implying that the steady-state level of firm output is indeterminate. Hence the $A_N = \frac{Y_f}{N}$, $A_K = \frac{Y_f}{K}$ assumption yields an additional degree of freedom, which, without loss of generality, I close by fixing Y_f at an arbitrary level. I can then directly compute a range of other variables, using the values computed or calibrated above,

$$K = \left(\frac{K}{Y_f}\right) Y_f \tag{I.20}$$

$$A_K = \frac{Y_f}{K} \tag{I.21}$$

$$Y = \left(\frac{Y}{Y_f}\right) Y_f \tag{I.22}$$

$$K = \left(\frac{K}{Y_f}\right) Y_f \tag{I.23}$$

$$I = \delta K \tag{I.24}$$

$$G = \left(\frac{G}{Y}\right) Y \tag{I.25}$$

$$Q_H = \left(\frac{Q_H H}{4Y}\right) \frac{4Y}{H} \tag{I.26}$$

$$H_{CC} = \left(\frac{H_{CC}}{H}\right) H \tag{I.27}$$

$$H_S = H - H_{CC} \tag{I.28}$$

$$CH = \left(\frac{CH}{Y}\right) Y \tag{I.29}$$

$$CH_{CC} = \left(\frac{CH_{CC}}{Y}\right) Y \tag{I.30}$$

$$b_{CC} = \left(\frac{b_{CC}}{Y}\right) Y \tag{I.31}$$

$$Target_{G2GDP} = \left(\frac{G}{Y}\right) \tag{I.32}$$

$$Target_{bgov2GDP} = \left(\frac{b_{gov}}{4Y}\right) \tag{I.33}$$

$$b_{gov} = \left(\frac{b_{gov}}{4Y}\right) 4Y \tag{I.34}$$

$$T = \left(\frac{R}{\Pi} - 1\right) b_{gov} + G \tag{I.35}$$

$$NI = Y - \delta K, \tag{I.36}$$

where (I.24), (I.36), and (I.35) follow from the capital accumulation equation (21), (47,) and the government budget constraint (11), respectively.

Substituting (40) into (48) allows to back out the share of non-rich households in total labor income ω_{CC} as

$$\omega_{CC} = \frac{NIS_{CC}\left(\left(1 + \frac{b_{gov}}{NI}\left(\frac{R}{\Pi} - 1\right)\right) NI\right) + b_{CC}\left(\frac{R_L}{\Pi} - 1\right)\frac{R}{\Pi} - \frac{H_{CC}}{H} CH}{mc\left(1 - \alpha_K\right) Y_f}. \tag{I.37}$$

Using my assumption $\eta_S = N_S$ and $\eta_{CC} = N_{CC}$, total employment (from Equation (41)) and A_N are given by

$$N = 1 \tag{I.38}$$

$$A_N = \frac{Y_f}{N}. \tag{I.39}$$

From (39), (40), $A_N = \frac{Y_f}{N}$, and $A_K = \frac{Y_f}{K}$ follows

$$w_S = mc_t\left(1 - \alpha_K\right)\left(1 - \omega_{CC}\right)\frac{Y_f}{N_S} \tag{I.40}$$

$$w_{CC} = mc\left(1 - \alpha_K\right)\omega_{CC}\frac{Y_f}{N_{CC}} \tag{I.41}$$

$$w = mc\left(1 - \alpha_K\right) A_N, \tag{I.42}$$

which then allows to compute, from Equations (46), (45), (30), and (49),

$$Target_{T_{CC}2T} = NIS_{CC} \tag{I.43}$$

$$T_{CC} = Target_{T_{CC}2T}T \tag{I.44}$$

$$T_S = T - T_{CC} \tag{I.45}$$

$$C_{CC} = w_{CC}N_{CC} - T_{CC} - \left(\frac{R_L}{\Pi} - 1\right) b_{CC} \tag{I.46}$$

$$C_S = Y - C_{CC} - G - I - CH_t. \tag{I.47}$$

From the consumption FOCs (22), (31), and the investment FOC (25),

$$\Lambda_S = \frac{1}{C_S^{\sigma_C}} \tag{I.48}$$

$$\Lambda_{CC} = \frac{1}{C_{CC}^{\sigma_{CC}}} \tag{I.49}$$

$$Q = 1. \tag{I.50}$$

Using the FOCs with respect to housing ((26) and (33)), safe assets (Equation (23)), and capital (Equation (24)), and setting σ_b and σ_K for now to some arbitrary value, allows to back out the utility weights for housing $\phi_{H,S}$ and $\phi_{H,CC}$, safe assets ϕ_b, and physical capital ϕ_K as

$$\phi_{H,S} = Q_H \left(1 - \beta_S\right) \Lambda_S H_S^{\sigma_{H,S}} \tag{I.51}$$

$$\phi_{H,CC} = Q_H(1 - \beta_{CC}) \Lambda_{CC} H_{CC}^{\sigma_{H,CC}} + \beta_{CC} \frac{\frac{dR_L}{dH_{CC}}(..)}{\Pi} b_{CC}\Lambda_{CC} H_{CC}^{\sigma_{H,CC}} \tag{I.52}$$

$$\phi_b = (1 - \theta) \Lambda_S b_S^{\sigma_b} \tag{I.53}$$

$$\phi_K = \Lambda_S \left(1 - \beta_S \left[r_K + (1 - \delta)\right]\right) (K)^{\sigma_K}. \tag{I.54}$$

Finally, I adjust the target for the $\left(\frac{b_{CC}}{Q_H H_{CC}}\right)$ in order to set the implied debt-to-annual-income ratio $\left(\frac{b_{CC}}{4 w_{CC} N_{CC}}\right)$ to its target value.

After pinning down the steady state as just described, I then set σ_b and σ_K to match the empirical targets for the effect of a permanent increase in the supply of government bonds on the safe interest rate $\frac{d\frac{R}{\Pi}}{d\text{Target}_{bgov2GDP}}$, and the MPS of the rich, respectively, as described in Section 3.6.1. Note that, thanks to the (almost) closed-form solution, all changes in σ_b and σ_K are absorbed by ϕ_K and ϕ_b, implying that the steady state remains unaffected.

Appendix J. A Simple Microfoundation of the Borrowing Friction

The increasing relationship between the loan rate $R_{L,t}$ and the borrower LTV assumed in the main text may be microfounded by assuming borrowing is subject to frictions similar to Lozej, Onorante, and Rannenberg (2018). Specifically, I assume that the household's housing wealth is subject to idiosyncratic uncertainty which resolves at the beginning of the period, and that a household j defaults if its housing wealth is less than its real debt $\frac{R_{Lj,,t}}{\Pi_t} b_{CC,j,t-1}$. More formally, default occurs if

$$\omega_{j,t} H_{CC,j,t-1} Q_{H,j,t} < \frac{R_{L,j,t}}{\Pi_t} b_{CC,j,t-1},$$

where $\omega_{j,t}$ denotes an i.i.d. random variable with mean one. Hence the default threshold of household j is given by

$$\overline{\omega_{j,t}} = \frac{\frac{R_{L,j,t}}{\Pi_t} b_{CC,j,t-1}}{H_{CC,j,t-1} Q_{H,j,t}}. \tag{J.1}$$

I assume that if $\omega_{j,t}$ falls below the default threshold and the household therefore defaults, the loss given default incurred by the financial intermediary is fixed at a fraction LGD of the loan, with $0 \leq LGD \leq 1$. Furthermore, in order to abstract from the effect of loan losses on the financial intermediary, I follow Bernanke, Gertler, and Gilchrist (1999) and assume that the debt contract is contingent on the realization of aggregate variables to ensure that, in every quarter, the financial intermediary earns an average nominal rate of return $R_{t-1} FIC$, where $FIC - 1$ represents non-bankruptcy related costs of financial intermediation, which I assume to be a fixed fraction of the total loan amount. Hence, the interest rate adjusts accordingly ex post and is given by

$$R_{L,j,t} = \frac{R_{t-1}FIC}{1 - LGDJ\left(\overline{\omega_{j,t}}\right)}, \tag{J.2}$$

where $J\left(\overline{\omega_t}\right)$ denotes the cumulative distribution function of $\omega_{j,t}$. This equation replaces the ad hoc loan supply relationship assumed in the main text (i.e., Equation (29)). Finally, defaulting households face a cost $LGD\frac{R_{L,j,t}}{\Pi_t}b_{CC,j,t-1}$, implying that otherwise identical defaulting and non-defaulting households face identical debt-related payments at the beginning of period t.[24] After $\omega_{j,t}$ has been revealed and some households default, resources are redistributed between borrower households such that their housing wealth is again identical before they make their consumption and saving decisions. With these assumptions, the borrowing household's budget constraint is identical regardless of default, and I therefore drop the j subscript from now on:

$$\begin{aligned} \frac{R_{L,t}}{\Pi_t}b_{CC,t-1} + C_{CC,t} + Q_{H,t}\left(H_{CC,t} - H_{CC,t-1}\right) \\ = b_{CC_t} + w_{CC,t}N_{CC,t} - T_{CC,t}. \end{aligned} \tag{J.3}$$

The FOCs with respect to consumption $C_{CC,t}$, real loans $b_{CC,t}$, housing $H_{CC,t}$, and the expected loan interest rate $R_{L,t+1}$ imply

$$\Lambda_{CC,t} = \frac{1}{C_{CC,t}^{\sigma_{CC}}} \tag{J.4}$$

$$\Lambda_{CC,t} = \beta_{CC}E_t\left\{\Lambda_{CC,t+1}\left[\frac{R_{L,t+1}}{\Pi_{t+1}} + \frac{\frac{dR_{L,t+1}}{db_{CC,t}}\left(\bar{\omega}_{t+1}, R_{L,t+1}, b_{CC,t}\right)b_{CC,t}}{\Pi_{t+1}}\right]\right\} \tag{J.5}$$

[24] Specifically, at the beginning of period t, a non-defaulting households repays $\frac{R_{L,,t}}{\Pi_t}b_{CC,t-1}$, while a defaulting household repays $(1 - LGD)\frac{R_{L,,t}}{\Pi_t}b_{CC,t-1}$ but faces default costs $LGD\frac{R_{L,,t}}{\Pi_t}b_{CC,t-1}$, which amount to the same debt-related payment as that of the non-defaulting households. This assumption is necessary to ensure that a change in the lending rate caused by an increase in the expected probability of default E_tJ_{t+1} has an effect on household behavior.

$$Q_{H,t} = \frac{\phi_{H,t,CC}}{\Lambda_{CC,t} H_{CC,t}^{\sigma_{H,CC}}} + \beta_{CC} E_t \Bigg\{ \frac{\Lambda_{CC,t+1}}{\Lambda_{CC,t}} \left(Q_{H,t+1} - \frac{\frac{dR_{L,t+1}}{dH_{CC,t}} \left(\bar{\omega}_{t+1}, R_{L,t+1}, H_{CC,t}\right)}{\Pi_{t+1}} b_{CC,t} \right) \Bigg\}, \tag{J.6}$$

where $\frac{dR_{L,t+1}}{db_{CC,t}} \left(\bar{\omega}_{t+1}, R_{L,t+1}, b_{CC,t}\right)$ and $\frac{dR_{L,t+1}}{dH_{CC,t}} \left(\bar{\omega}_{t+1}, R_{L,t+1}, H_{CC,t}\right)$ denote the implicit function derivatives of $R_{L,t+1}$ with respect to $b_{CC,t}$ and $H_{CC,t}$, respectively, given by

$$\frac{dR_{L,t+1}\left(\bar{\omega}_{t+1}, R_{L,t+1}\right)}{db_{CC,t}} = \frac{LGDJ'\left(\bar{\omega}_{t+1}\right) \frac{\bar{\omega}_{t+1}}{b_{CC,t}} R_{L,t+1}}{\left(1 - LGDJ\left(\bar{\omega}_{t+1}\right) - LGDJ'\left(\bar{\omega}_{t+1}\right) \bar{\omega}_{t+1}\right)} \tag{J.7}$$

$$\frac{dR_{L,t+1}}{dH_{CC,t}} \left(\bar{\omega}_{t+1}, R_{L,t+1}, H_{CC,t}\right) = \frac{-LGDJ'\left(\bar{\omega}_{t+1}\right) \frac{\bar{\omega}_{t+1}}{H_{CC,t}} R_{L,t+1}}{\left(\left(1 - LGDJ\left(\bar{\omega}_{t+1}\right)\right) - LGDJ'\left(\bar{\omega}_{t+1}\right) \bar{\omega}_{t+1}\right)}. \tag{J.8}$$

For $\frac{dR_{L,t+1}(\bar{\omega}_{t+1}, R_{L,t+1})}{db_{CC,t}} > 0$ and $\frac{dR_{L,t+1}}{dH_{CC,t}} \left(\bar{\omega}_{t+1}, R_{L,t+1}, H_{CC,t}\right) < 0$ it has to be true that $\frac{1}{LGD} > J\left(\bar{\omega}_{t+1}\right) + J'\left(\bar{\omega}_{t+1}\right) \bar{\omega}_{t+1}$. For this inequality to hold, it is sufficient if its right-hand side is less than one. That condition is met under the assumptions I adopt below. I assume a logistic form for $J\left(\overline{\omega_t}\right)$:

$$J\left(\overline{\omega_t}\right) = \frac{1}{1 + e^{-\frac{\bar{\omega}_t - 1}{\sigma_h}}}. \tag{J.9}$$

Hence, loan supply is now determined by the three parameters FIC, LGD, and σ_h, while the slope parameter of the ad hoc loan supply curve χ_{CC} drops out of the model. To identify the parameters of the model, I therefore need two additional targets on top of those listed in Table 4. These two additional targets are listed in Table J.1. I pin down FIC by adopting a target value for the ratio of non-bankruptcy-related costs to net interest income $\frac{(FIC-1)}{R_L - R}$, which I estimate based on the FDIC Quarterly Banking Profile. Specifically,

Table J.1. Additional Targets Microfounded Borrowing Friction

Target	Value Data	NOCSP	CSP	Source
Mortgage Default Rate (APR) $J()$	4.2%	4.2%	4.2%	Single-Family Residential Mortgages, FRED (1991–2016)
Non-default-Costs-to-Interest-Income Ratio $\frac{(FIC-1)}{R_L-R}$	59%	59%	59%	FDIC QBP, (1984–2016); See note below for details

Note: This table lists the two additional targets I use in the model with the microfounded borrowing friction. For the other targets, see Table 4. The empirical counterpart of the non-default-costs-to-interest-income ratio $\frac{(FIC-1)}{R_L-R}$ is computed from the FDIC Quarterly Banking Profile as $\frac{\left(\frac{\text{Total noninterest expense}}{\text{Total Assets}}\right)}{\left(\frac{\text{Net interest income}}{\text{Total loans and leases}}\right)}$.

I calculate $\frac{\left(\frac{\text{Total noninterest expense}}{\text{Total Assets}}\right)}{\left(\frac{\text{Net interest income}}{\text{Total loans and leases}}\right)} \approx 59\%$. I set the steady-state default rate $J(\bar{\omega})$ equal to the average "Delinquency Rate on Single-Family Residential Mortgages, Booked in Domestic Offices, All Commercial Banks," which equals an annualized 4.2 percent. Using the (implied) values of FIC, $J(\bar{\omega})$, the (unchanged) target value for $R_L - R$ and Equation (J.2) pin down the required LGD value. Table J.2 displays the complete list of targets and parameter values. Values which differ from the model with the ad hoc borrowing friction of the main text are in bold.

Figure K.1 in Appendix K compares the loan supply curve implied by the ad hoc borrowing friction to the microfounded loan supply curve (Equation (J.2)). The curves cross at an LTV of 0.27, which is the steady-state value implied by the empirical targets.

Table J.2. Full Model, Parameter Values

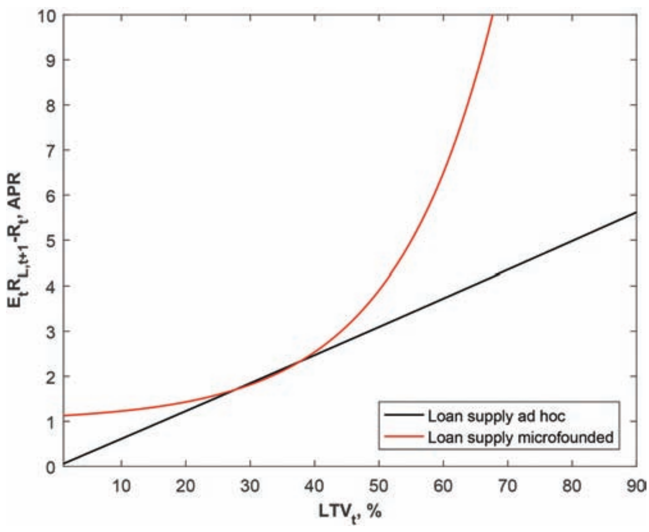
Parameter	Parameter Name	Value NOCSP ($\theta = 1$)	Value CSP ($\theta = 0.97$)
β_S	Rich Utility Discount Factor	0.9951*	0.9652*
β_{CC}	Borrower Utility Discount Factor	0.9881*	0.9881*
σ_S, σ_{CC}	Utility Curvature Consumption	2*	2*
$\sigma_{S,H}$, $\sigma_{CC,H}$	Utility Curvature Housing	2	2
$\phi_{H,S}$	Rich Utility Weight on Housing	0.16*	1.29*
$\phi_{H,CC}$	Borrower Utility Weight on Housing	0.58*	0.66*
$\nu_{cascade}$	Consumption Cascade	0	0/0.7
$\tilde{N}_{CC}$,$\tilde{N}_S$	Labor Endowments	$\frac{1}{3}$	$\frac{1}{3}$
μ_P	Price Markup	1.18*	1.06*
α_K	Output Elasticity w.r.t. Capital	0.19*	0.24*
ε	Elasticity of Substitution Capital/Labor	1	1; 0.3
ω_{CC}	Borrower Share in Labor Income	0.84*	0.8*
δ	Depreciation Rate Physical Capital	0.025	0.025
ε_{I}	Capital Adjustment Cost Curvature	7	7
FIC	**Non-default Intermediation Cost**	1.0025*	1.0025*
σ_h	**Volatility Housing Value Shock**	0.16*	0.16*
LGD	**Loss Given Default**	0.16*	0.16*
$Target_{bgov2GDP}$	Government Debt Target	44%	44%
$Target_{G2GDP}$	Government Expenditure Share Target	20%	20%
σ_b	CSP: Utility Curvature, Real Financial Assets	—	0.69*
σ_K	CSP: Utility Curvature, Physical Capital	—	5*
ϕ_b	CSP: Utility Weight on Real Financial Assets	0*	3.28*
ϕ_K	CSP: Utility Weight on Physical Capital	0*	220.43*

Note: Values marked with an asterisk (*) are set to match the targets reported in Table J.1.

Appendix K. Results with the Microfounded Loan Supply Curve

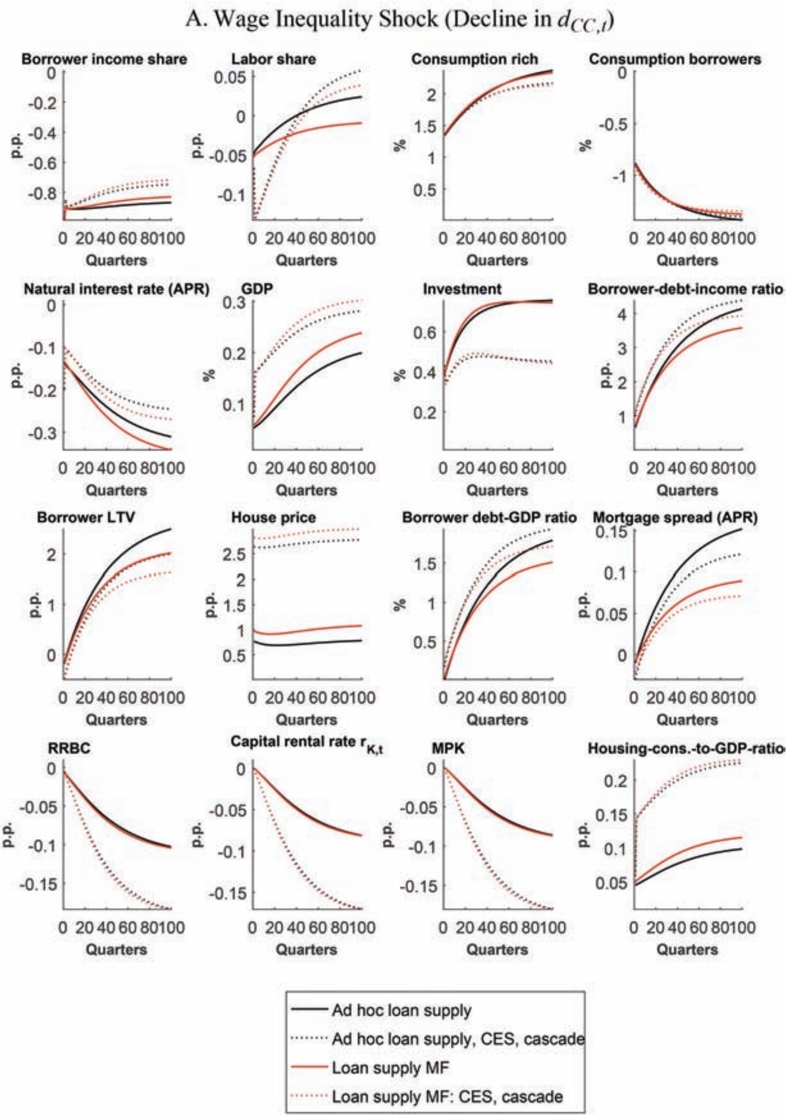
As can be seen from Figure K.2, exchanging the ad hoc loan supply curve assumed in the main text (black lines) for the more microfounded one from Appendix J (red lines) has virtually no impact on the simulated effects of a one-off permanent increase in wage inequality or the price markup. The reason is that around the initial steady state, the slope of the ad hoc and microfounded loan supply curves are in fact quite similar (see Figure K.1).

Figure K.1. Loan Supply Curve



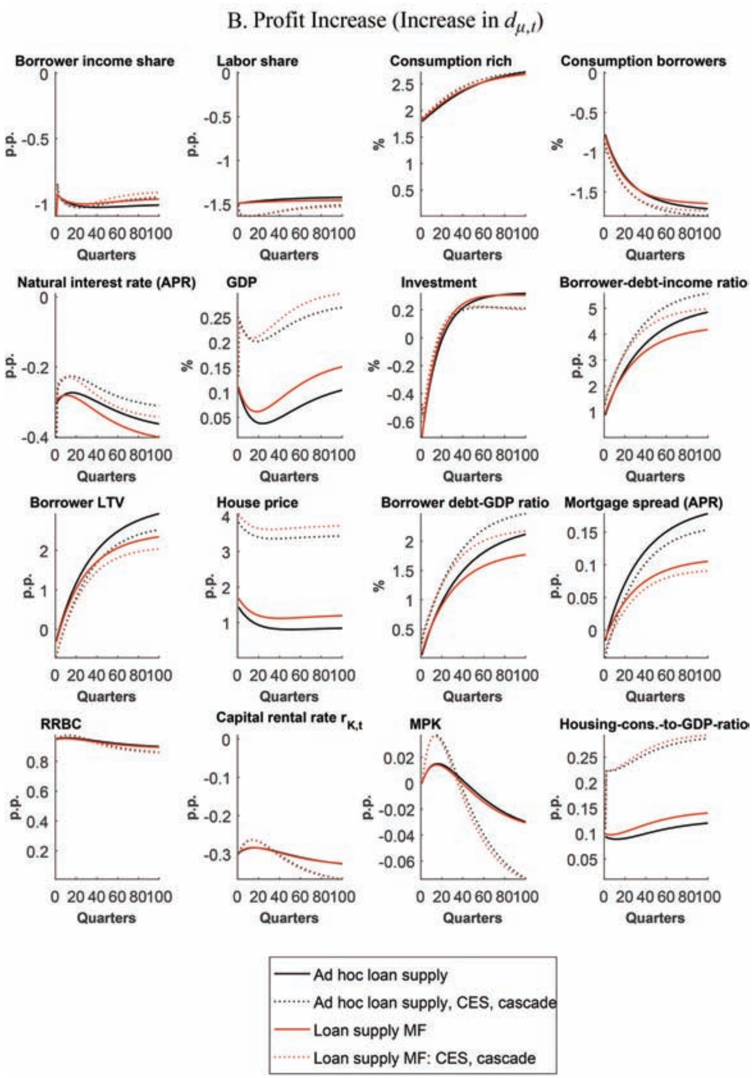
Note: The linear black line displays the relationship between the spread of the mortgage rate over the risk-free rate (i.e., $E_t R_{L,t+1} - R_t$) and the loan-to-value ratio $\left(LTV_t = \frac{b_{CC,t}}{H_{CC,t} Q_{H,t+1}}\right)$ implied by the borrowing friction assumed in the main text (see Equations (34) and (29)), with χ_{CC} as in Table 3. The non-linear red line displays the analogous relationship implied by the microfounded borrowing friction derived in Appendix J (see Equations (J.2), (J.1), and (J.9)) and the calibration discussed in Appendix K.

Figure K.2. Impact of a Permanent Increase in Inequality: Role of Loan Supply Assumption



(*continued*)

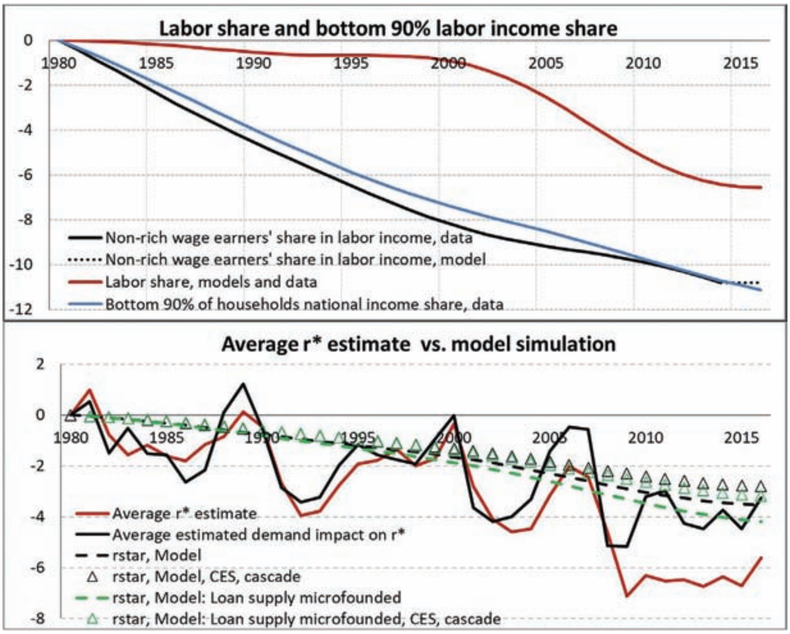
Figure K.2. (Continued)



Note: The black lines ("Ad hoc loan supply") refer to results computed using the model of Section 3. The red lines ("Loan supply MF") refer to a modified version of the model where the borrowing friction is microfounded as described in Appendix J, implying that Equations (J.1), (J.2), and (J.9) replace Equations (34) and (29). "cascade" indicates that the model allows for an effect of rich household total consumption on non-rich housing demand (i.e., $\nu_{cascade} > 0$; see Equation (35)). The safe interest rate R_t and the risk spread $E_t R_{L,t+1} - R_t$ are expressed as annualized percentage rates (APR). The borrower debt-to-income ratio is based on annualized borrower income. The borrower debt-to-GDP ratio is based on annualized GDP.

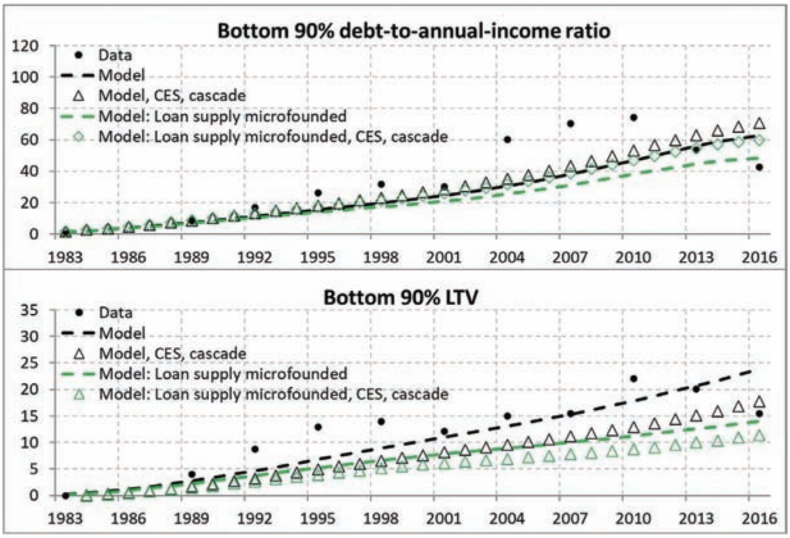
The historical simulation described in Section 4.2 yields a somewhat larger decline in the natural rate (see Figure K.3) and a somewhat smaller increase in household indebtedness than when assuming the ad hoc loan supply curve (see Figures K.4 and K.5). The reason is that, in this simulation, the model is hit by a sequence of shocks moving the economy further away from the initial steady state than in the simulation of a one-off increase in inequality. The LTV therefore moves into the region where the microfounded loan supply curve becomes substantially steeper than the ad hoc loan supply curve used in the main text (see 19), implying that borrowing does not expand as much in response to a decline in the risk-free rate. By contrast, the increase in the value of the housing stock is somewhat larger because, with the steeper loan supply curve, a given increase in house prices has a larger negative effect on $E_t R_{L,t+1}$ and thus the demand for houses via $\frac{\Lambda_{CC,t+1}}{\Lambda_{CC,t}}$. Differences between the ad hoc and microfounded loan supply cases are smaller in the model with a consumption cascade.

Figure K.3. Simulation 1981–2016: Income Distribution and Natural Interest Rate



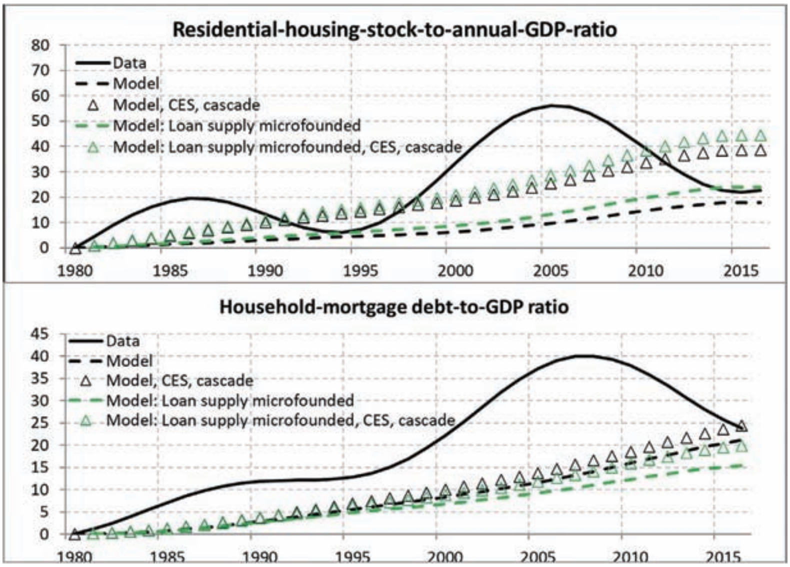
Note: The label "Model: Loan supply microfounded" refers to a modified version of the model where the borrowing friction is microfounded as described in Appendix J, implying that Equations (J.1), (J.2), and (J.9) replace Equations (34) and (29). For details on the meaning of the other labels, see the note below Figure 5.

Figure K.4. Simulation 1981–2016: Borrower Debt-to-Income Ratio and LTV



Note: See the note below Figure K.3 for details on the labels reading "Model . . .". See the note below Figure 6 for details on the data sources.

Figure K.5. Simulation 1981–2016: Housing Stock and Total Mortgage Debt-to-GDP Ratios



Note: See the note below Figure K.3 for details on the labels reading "Model . . .". See the note below Figure 7 for details on the data sources.

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Making Waves: Monetary Policy and Its Asymmetric Transmission in a Globalized World*

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This paper compares the international transmission of European Central Bank (ECB) and Federal Reserve System (Fed) monetary policy in a unified framework, identifying pure monetary policy shocks purged of bias from central bank information effects. The estimates reveal a stark asymmetry in the global spillovers from ECB and Fed monetary policy: Fed monetary policy shocks have a significant impact on euro-area financial conditions and real activity, while ECB monetary policy shocks do not have a similar effect on the United States. Fed monetary policy shocks also affect real and financial variables in the rest of the world more than ECB monetary policy shocks.

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1. Introduction

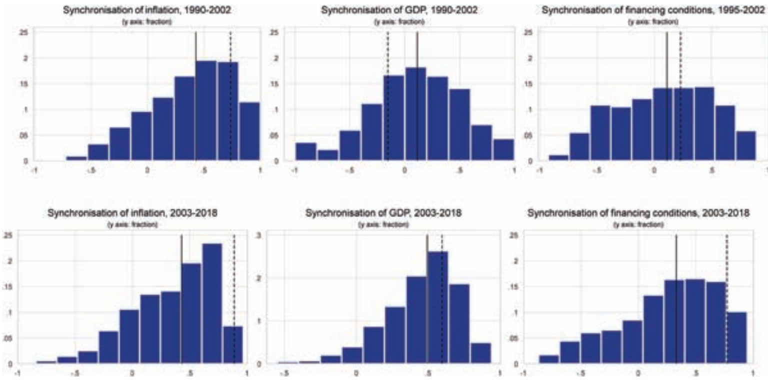
The global financial crisis and the recourse to unconventional monetary policy measures it entailed have created renewed interest in the international dimension of monetary policy. Events since then have triggered an intense debate about the potential of monetary policy in systemic economies to propagate to the rest of the world (Mantega 2010).[1] A related discussion has revolved around the question of whether a "global financial cycle"—also fueled by monetary policy in systemic economies—is undermining the ability of central banks in the rest of the world to affect domestic financial conditions and eventually control prices, real activity, and financial stability (Rajan 2013; Rey 2016).

Indeed, the increasing international co-movement of key macroeconomic and financial variables over the recent decades points to closer international linkages that may underpin growing monetary policy spillovers (Figure 1). Splitting the period between 1990 and 2018 into two halves, the distributions of bilateral cross-country correlations between inflation, real gross domestic product (GDP) growth, and financing conditions have become more skewed to the right. For the euro area–U.S. country pair, the correlations have increased substantially for all three indicators, i.e., inflation, real GDP growth, and financing conditions. Especially sizable is the rise in the correlation between euro-area and U.S. financing conditions across the two periods. Some explanation for this could be the larger and more frequent common shocks in the later period, but the greater spillovers from country-specific shocks and the associated systematic responses by central banks in systemic economies could play a non-trivial role as well.

While a large and growing literature has explored spillovers from monetary policy in the main systemic economies—namely by the Fed and ECB—there remain gaps in our understanding. One of these relates to assessing the differences across spillovers from Fed and ECB monetary policy. Existing work has typically explored Fed *or* ECB monetary policy spillovers individually, but not together in a consistent and unified methodological framework evaluating both their bilateral and global impact. As a result, a comparison

[1] See Draghi (2016).

Figure 1. Distribution of Pairwise Cross-Country Correlations of Inflation, GDP Growth, and Financing Conditions



Source: World Development Indicators; Haver, Refinitiv Datastream, and Bloomberg.
Note: The solid line indicates the median correlation and the dashed line the correlation between the United States and the euro area. The data cover 53 advanced and emerging economies at annual frequency for GDP and inflation and 34 countries at monthly frequency for financing conditions. Financing conditions are calculated as a weighted average of five financial variables.

of Fed and ECB monetary policy spillovers across existing studies is generally problematic, as findings are based on rather different methodological approaches and data. Specifically, it remains unclear whether differences in the existing evidence on Fed and ECB monetary policy spillovers reflect a feature of the world or simply sampling and model uncertainty.

The main contribution of this paper is to document and compare spillovers from Fed and ECB monetary policy using a consistent and unified methodological framework. We estimate spillovers from Fed and ECB monetary policy using identical vector autoregressive (VAR) models, identification approaches, and data samples. We estimate Bayesian VAR models with the same set of U.S. and euro-area endogenous variables, employ high-frequency interest rate surprises around Federal Open Market Committee and ECB Governing Council meetings to identify monetary policy shocks (Jarociński and Karadi 2020), and consider data for the same countries over the time period from 1999 to 2018. We first analyze the domestic effects and transatlantic spillovers between the United States and the euro area elicited by Fed and ECB monetary policy, respectively, and then spillovers to the rest of the world with a focus on emerging market economies (EMEs).

Three key findings improve our understanding of the domestic and international effects of Fed and ECB monetary policy. First, our results document that even in a highly globalized economy both Fed and ECB monetary policy have a sizable impact on domestic financial conditions, real activity, and inflation. An exogenous ECB and Fed monetary policy tightening raises domestic risk-free rates and corporate bond yields, depresses domestic equity markets, is followed by an appreciation of the domestic currency, slows real activity, and reduces inflation.

Second, we document a stark asymmetry in transatlantic spillovers, with the Fed having a much more encompassing impact on the euro-area economy than the ECB on the U.S. economy. The largest spillover from Fed monetary policy materializes in euro-area financial markets; the spillover to euro-area real activity is more subdued and on euro-area inflation very short lived. The impact of ECB monetary policy on U.S. economic variables is instead small in most dimensions, including on U.S. financial markets, real activity, and inflation. Hence, while deeper transatlantic goods and financial market integration has entailed a greater role of Fed monetary policy for the euro-area economy, the role of ECB monetary policy for the U.S. economy remains limited.

Third, we document that there is a "hierarchy" in Fed and ECB monetary policy spillovers to EMEs. Consistent with the dominant role of the U.S. dollar in the international monetary and trade system, Fed monetary policy elicits large spillovers to financial conditions and real activity in EMEs. By contrast, spillovers from ECB monetary policy are largely confined to trade (and, perhaps surprisingly, commodity prices).

Our paper is related to and contributes to existing literature. A very large and still-growing literature has explored the spillovers from Fed monetary policy (for a subset of recent work, see Passari and Rey 2015; Ammer et al. 2016; Georgiadis 2016; Dedola, Rivolta, and Stracca 2017; Gerko and Rey 2017; Dées and Galesi 2019; Iacoviello and Navarro 2019; Degasperi, Hong, and Ricco 2020). Another—much smaller—literature explores the spillovers from ECB monetary policy (Babecká-Kucharčuková, Claeys, and Vašíček 2016; Bluwstein and Canova 2016; Potjagailo 2017; Moder 2019; Feldkircher, Gruber, and Huber 2020; ter Ellen, Jansen, and Larsson Midthjell 2020). In general, the results in this literature suggest that spillovers from Fed monetary policy to the rest of the world—in particular, to EMEs—are large, while those from ECB monetary policy are confined to Europe and neighboring regions. Comparing the findings for Fed and ECB spillovers is, however, not generally feasible due to differences in identification assumptions, the choice of time and country samples, and model specifications. Moreover, none of the existing work zooms in on bilateral spillovers between the United States and the euro area.

A third—even smaller—literature has estimated spillovers from both Fed and ECB monetary policy, but with a more limited scope than in our study and different methodologies. Rogers, Scotti, and Wright (2014), Curcuru, De Pooter, and Eckerd (2018), Curcuru et al. (2018), and Kearns, Schrimpf, and Xia (2018) focus on short-term spillovers from monetary policy surprises to financial markets. Chen et al. (2017) examine spillovers from Fed and ECB unconventional monetary policy in a global VAR model with sign restrictions. Hajek and Horvath (2018) also consider a global VAR model, but only explore generalized impulse responses rather than identified monetary policy shocks. Walerych and Wesolowski (2020) consider Bayesian panel VAR models with Taylor-rule residuals as monetary policy shocks. In general, these papers find that Fed monetary policy elicits larger spillovers than ECB monetary policy at the global level but do not discuss bilateral spillovers between the United States and the euro area. Moreover, several of these papers do not examine the response of macroeconomic variables but focus on the impact on financial variables. And none of these papers purges interest rate shocks from central bank information effects, which has been shown to be empirically important (Jarociński and Karadi 2020).[2] Finally, Miranda-Agrippino, Nenova, and Rey (2020) estimate Bayesian VAR models to estimate the spillovers from U.S. and Chinese monetary policy. However, they use different identification approaches—namely high-frequency interest rate surprises for the

[2]A related paper by Jarociński (2020) focuses instead on the transatlantic spillovers of the information effects.

United States and recursiveness assumptions for China—and do not consider bilateral spillovers.

The rest of the paper proceeds as follows. Section 2 briefly reviews the main international transmission channels for monetary policy to set the stage on the importance of the financial channel in a highly globalized world. Section 3 introduces our empirical methodological framework and why it is helpful in isolating interest rate surprises. In Section 4 we present our empirical findings concerning the domestic and transatlantic effects of Fed and ECB monetary policy. The international effects of ECB and Fed monetary policy on EMEs are considered in Section 5. In Section 6 we conclude.

2. Transmission Channels of Monetary Policy Spillovers

To set the stage for our empirical analysis we briefly review how monetary policy spillovers propagate via the aggregate demand, the expenditure-switching, and a multi-faceted financial channel.

To the extent that a contractionary monetary policy action curbs home consumption and investment, it also reduces the demand for imported goods, and thus for exports of the economy's trading partners. As a result, spillovers through the *aggregate demand channel* reduce output in trading partners. The magnitude of the monetary policy spillover through the aggregate demand channel rises with the weight of the home economy in its trading partners' overall trade. Therefore, monetary policy of economies with a large weight in the global economy should have a commensurately large effect on aggregate demand worldwide.

Monetary policy affects the exchange rate, which due to price stickiness tends to alter in the short run the relative price of imported and domestically produced goods, which then gives rise to an *expenditure-switching channel.* How a nominal appreciation of the home currency affects the relative price between domestically produced goods and imports depends on the degree of exchange rate pass-through (ERPT) to import prices and the ensuing expenditure switching. A key determinant of ERPT over shorter horizons is the currency in which export and import prices are sticky. First, under producer-currency pricing (PCP), traded goods prices are sticky in the currency of the producer, ERPT is complete, and an appreciation of the home currency improves the terms of trade, inducing expenditure switching away from domestically produced goods towards goods produced in the rest of the world at home and abroad. Second, under local-currency pricing (LCP), some or even all export prices are sticky in the currency of the importer, and ERPT and expenditure switching are muted.[3] Third, under dominant-currency pricing (DCP), all export (and import) prices worldwide are sticky in just a few major currencies, and expenditure switching depends on the source of the shock and on the specific bilateral trade relationship in question.[4] The U.S. dollar is currently the dominant invoicing currency in global trade (Boz et al. 2020), so that DCP should be particularly relevant for the global transmission of Fed monetary policy. Especially EMEs invoice the bulk of their imports and exports in U.S. dollars regardless of the destination. For the euro area, in contrast, a substantial fraction of non-U.S. imports and exports are invoiced in euros rather than in dollars; the share of euro-area countries' total imports and exports invoiced in euros amounts to 71 percent and 74 percent, respectively (see Boz et al. 2020). For example, under DCP an appreciation of the U.S. dollar is inconsequential in terms of expenditure switching in the United States, as its import and export prices are sticky in U.S. dollars. In contrast, in all countries in the rest of the world a multilateral appreciation of the U.S. dollar entails a widespread rise in import prices, which induces expenditure switching away from imports towards domestically produced goods. Moreover, because imports in economies in the rest of the world decline regardless of the source, rest-of-the-world exports decline commensurately. In contrast, a multilateral appreciation of a non-U.S. dollar currency (like the euro) would have only limited expenditure-switching effects under DCP. In the non-U.S. economy expenditure switching in this case affects only imports but not exports. Trade in the rest of the world that does not involve this non-U.S. economy is entirely unaffected by the multilateral appreciation of its currency.[5]

[3] See Betts and Devereux (1996, 2000) as well as Devereux and Engel (2003).

[4] See Gopinath et al. (2020).

[5] An additional transmission channel for monetary policy spillovers from the dominant-currency-issuing economy to the rest of the world operates through the endogenous response of monetary policy: as all import prices in the rest of the

The role of the *financial channel* is particularly important in view of the strong international integration of financial markets. In a financially integrated world, monetary policy in a large currency area may affect financial conditions and thereby aggregate demand in the rest of the world. First, when a country supplies a global safe asset, its monetary policy can have a direct effect on aggregate demand abroad: a home monetary policy tightening increases the global demand for home assets and thus directly reduces global aggregate demand.

Second, exchange rate valuation effects in cross-border assets and liabilities change the value of foreign-currency denominated collateral, and thereby borrowing and leverage.[6] For example, when a firm borrows in foreign currency, home-currency depreciation tightens borrowing constraints and reduces the firm's borrowing capacity. Third, a monetary policy tightening depresses the value of domestic assets via a higher discount factor and lower expected cash flows. Some holders of these assets are leveraged investors, including financial intermediaries. The decline in asset values tightens their balance sheet constraints and raises their borrowing costs. This domestic balance sheet channel propagates across borders via asset price equalization and the synchronization of credit spreads and borrowing costs of leveraged cross-border investors (see, e.g., Devereux and Yetman 2010, and Dedola and Lombardo 2012). The U.S. dollar is the dominant currency in global financial markets, and hence U.S. monetary policy has a disproportionate impact on global financial conditions (Rey 2016; Gourinchas, Rey, and Sauzet 2019; Obstfeld 2020; Miranda-Agrippino and Rey 2022).

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world are sticky in the dominant currency—regardless of their source—a multilateral appreciation of the dominant currency raises local-currency import prices and thereby consumer price inflation. Depending on the degree of openness, this might induce local monetary policy to tighten, putting downward pressure on production (see Mukhin 2018; Corsetti et al. 2021; Zhang 2022;). Georgiadis and Schumann (2019) discuss export-import U.S. dollar invoicing share differentials under partial DCP as another conduit for output spillovers from U.S. monetary policy.

[6]Bruno and Shin (2015) describe the consequences of the co-movement of U.S. dollar exchange rates and the leverage of global banks. They refer to this relationship between domestic and global financial conditions as the "risk-taking channel of [local] currency appreciation." See Kearns and Patel (2016), Hofmann, Shim, and Shin (2017), and Avdjiev et al. (2019) for empirical evidence.

3. Empirical Framework

3.1 Identification of Monetary Policy Surprises

We construct exogenous interest rate surprises from asset price movements over narrow time windows around monetary policy announcements. The basic idea of this identification approach is that in a sufficiently narrow time window it is unlikely that factors other than monetary policy announcements move financial markets systematically. Therefore, any systematic movements in interest rates over the narrow time window represent the financial market effect of the monetary policy announcement. Moreover, such abrupt movements in interest rates represent a surprise: If financial markets had anticipated a change in the monetary policy stance, e.g., as a systematic reaction of monetary policy to the state of the economy, it would have already been priced in and interest rates would not have moved over the narrow time window.

However, interest rate surprises might not coincide with monetary policy surprises. In particular, the interest rate surprises might be contaminated by a central bank information effect. Central banks may move financial markets not only by surprises in their monetary policy stance for a given state of the economy, but also by affecting financial market beliefs about the state of the economy. For example, financial markets may interpret an unexpected interest rate cut as the central bank having a more pessimistic view about the state of the economy; in this case, financial markets could revise downwards their own beliefs about the state of the economy. This central bank information effect is different from a monetary policy shock, both conceptually as well as in terms of its macroeconomic effects (Melosi 2017; Cieslak and Schrimpf 2018; Nakamura and Steinsson 2018; Romer and Romer 2000; Miranda-Agrippino and Ricco 2021). Jarociński and Karadi (2020) document that central bank information effects can distort the estimation of the effects of monetary policy—in particular, for the persistence of the interest rate response and the magnitude of the price-level response.

We follow Jarociński and Karadi (2020) and purge interest rate surprises from central bank information effects using changes in stock prices in the same narrow window around the monetary policy announcement. Specifically, if stock prices move in the same direction as interest rates around the time of the announcement, we label the interest rate surprise a central bank information effect. If, by contrast, stock prices and interest rates move in opposite directions, we classify this as a monetary policy shock. This corresponds to the "poor man's" identification approach of Jarociński and Karadi (2020).

The "poor-man's" approach makes the simplifying assumption that the total interest rate surprise is either entirely a monetary policy shock or entirely a central bank information effect. We also consider the "rotational sign restrictions" approach of Jarociński and Karadi (2020), under which the total interest rate surprise that we observe is assumed to be a combination of both types of shocks in each month, i.e., in a typical month both shocks contribute to the overall interest rate surprise. It turns out that the monetary policy surprises based on the "poor man's" approach represent a better instrument for ECB monetary policy in our setup. Since our results are not very sensitive to this choice, for the sake of comparability we consider the "poor's man" approach to construct both ECB and Fed monetary policy surprises. Nevertheless, we discuss the results from "rotational sign restrictions" for the Fed whenever the results diverge in important ways.[7]

3.2 Data and Sample

Our data set consists of 168 Fed and 296 ECB monetary policy announcements between 1999 and 2018. The changes in interest rates and stock prices are measured in the time window starting 10 minutes before and ending 20 minutes after a central bank announcement. In the case of the Fed, the timing of the announcement typically coincides with that of the press release. In the case of the ECB, the time window is generally longer, starting 10 minutes before the press release and ending 20 minutes after the end of the press conference. In these windows we define the Fed interest rate surprise as the first principal component of the changes in federal funds futures and Eurodollar futures with remaining maturities from one month up to one year. Similarly, we define ECB interest rate surprises as the first principal component of the changes in EONIA (euro overnight

[7] Jarociński and Karadi (2020) discuss the merits of the two identification approaches.

index average) swaps with maturities from one month up to one year. By including maturities of up to one year, these surprises capture not just changes in current policy rates but also the expectations for interest rates up to one year in the future, reflecting forward guidance and other non-standard monetary policy measures.[8]

Our ECB and Fed monetary policy surprises are uncorrelated. The systematic components of ECB and Fed monetary policy of course both respond endogenously to the state of the economy and hence also to synchronized business cycles more generally (Belke and Gros 2005). As a result, the observed ECB and Fed monetary policy stance is correlated over time. But the exogenous, unsystematic surprise components of ECB and Fed monetary policy are uncorrelated. Therefore, any co-movement between euro-area and U.S. variables that we may estimate in response to an ECB or Fed monetary policy shock must represent monetary policy spillovers rather than the effects of common shocks.

3.3 A Bayesian VAR Model with Monetary Policy Surprises

We estimate the impact of ECB and Fed monetary policy based on a series of separate Bayesian VAR models. In each case, the VAR model includes the one-year government bond yield, stock prices, the corporate bond spread, industrial production, and the consumer price index (CPI/HICP). We add the monetary policy surprises to the VAR model as the first endogenous variable. We restrict the coefficients of the first equation to zero, reflecting the assumption that the monetary policy shock is independently and identically distributed over time. After estimating the VAR models with a standard Minnesota prior, we compute the impulse responses to a shock to the first equation, assuming a recursive structure. Note that this is less restrictive than in a VAR model in which a shock to a monetary

[8]The Fed surprises come from the updated data set of Gürkaynak, Sack, and Swanson (2005) and the ECB surprises from the data set of Jarociński and Karadi (2020). Similar monetary policy surprises are used in a large body of literature that includes, e.g., Kuttner (2001), Bernanke and Kuttner (2005), Gürkaynak, Sack, and Swanson (2005), and many others. For robustness checks we also consider longer-term rates. Extracting ECB monetary policy surprises from movements in three-year overnight index swaps during the effective lower bound period increases the magnitude, but not the time-series pattern, of monetary policy surprises, and therefore our results remain unchanged.

policy indicator such as the federal funds rate is used. In particular, in our specification the variable in the VAR that is assumed to not respond contemporaneously—and in fact not at all due to the zero constraints on all coefficients—is the high-frequency interest rate surprise on monetary policy announcement days cleansed from central bank information effects aggregated from daily to monthly frequency. The set of endogenous variables and the estimation of the VAR model in the baseline are the same as in Jarociński and Karadi (2020).[9]

We compute the transatlantic spillovers from Fed monetary policy shocks by entering the Fed shocks in the euro-area VAR model. Analogously, for the domestic effects we enter the Fed shocks in the U.S. VAR model. The effects of the ECB shocks are obtained analogously. The responses of other variables, which are not part of the baseline VAR model, are computed by adding them one by one as the last variable to the respective baseline VAR model.

Several variables we consider have a bilateral nature—for example, the exchange rate or the spread between U.S. and euro-area bond yields. In these cases, we use a bilateral VAR model to estimate their impulse responses. The bilateral VAR model includes the exchange rate, the spread between U.S. Treasuries and one-year bund yields, the corporate bond spread of the country experiencing the shock, and, separately for the United States and the euro area, industrial production and consumer price indices.[10] Monetary policy shocks are typically found to account only for a small fraction of the total variation of the observed monetary policy stance in the data. In the case of policy rates, the typical (exogenous) shock in an average month is only of about 2 or 3 basis points.

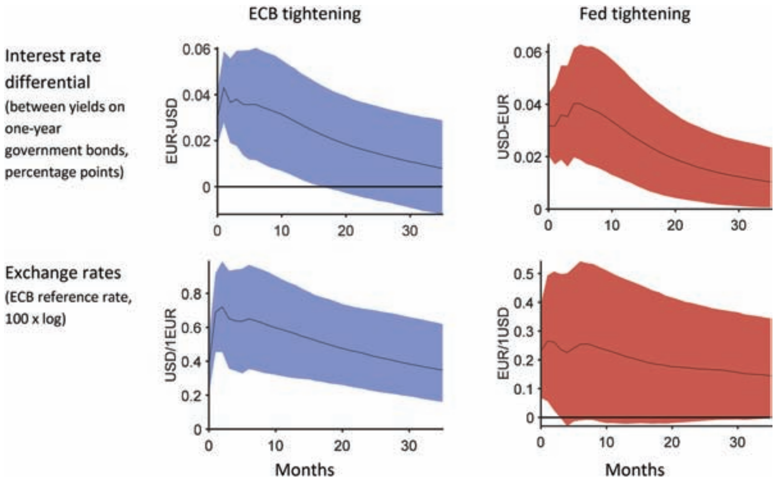
4. Domestic and Bilateral Effect of U.S. and Euro-Area Monetary Policy

In the following we report the responses to an exogenous monetary policy tightening of one standard deviation, which corresponds to

[9] Jarociński and Karadi (2020) also provide details on the rotational sign restrictions approach, which yields a similarly good instrument for Fed monetary policy shocks but a weaker one for ECB shocks.

[10] Appendix A.1 provides further details on the specification of the Bayesian VAR models. Appendix B summarizes the definitions of the response variables.

Figure 2. Responses of the Bilateral Interest Rate Differential and Exchange Rate to an Exogenous ECB and Fed Monetary Policy Tightening (bilateral model)



Note: The solid line plots the median impulse response, surrounded by the 68 percent confidence band. The left-hand column shows the responses to an ECB tightening and the right-hand column responses to a Fed tightening.

a contemporaneous increase in domestic one-year bond yields by almost 2.8 basis points for the ECB and close to 2.0 basis points for the Fed.[11] In Figure 2 and in the following figures we show how key domestic and foreign variables are estimated to respond in the 36 months following a tightening by the ECB (left column) and the Fed (right column). The point estimate of the impulse response is plotted as a solid line, surrounded by the 68 percent confidence band as shaded area (blue for euro-area variables and red for U.S. variables). The responses reflect the general equilibrium effects of the exogenous monetary policy tightening, and hence include the effects of the endogenous policy responses on the other side of the Atlantic.

[11]The ECB and Fed monetary policy shocks studied here have thus a similar, but not identical, impact on one-year bond yields. To compare monetary policy shocks with a (counterfactually) identical impact, the quantities in the following figures can be rescaled. Such rescaling reduces the relative magnitude of the effect of ECB policy versus Fed policy shocks but, by construction, does not affect significance. Because the model is linear, the impulse responses to a monetary policy loosening can be obtained by flipping the sign of the responses.

4.1 Effect on Interest Rate Differentials and Bilateral Exchange Rates

The evidence in the first row of Figure 2 highlights that both after an exogenous ECB and after an exogenous Fed tightening the interest rate differential between the two regions (defined as home minus foreign government bond rate) widens significantly. The magnitude and persistence of the responses of the interest rate differential for one-year bonds is very similar across the ECB and the Fed shocks.[12] The impulse responses in the second row of the figure document that, in line with uncovered interest rate parity, the domestic currency appreciates both in the case of ECB and in the case of Fed shocks, as domestic interest rates increase relative to foreign rates. After an ECB monetary policy tightening, the euro appreciates particularly sharply against the U.S. dollar. Likewise, after a Fed monetary policy tightening, the U.S. dollar appreciates as well, but the effect is smaller and less persistent.[13]

4.2 Effect on Real Activity and Prices at Home and Abroad

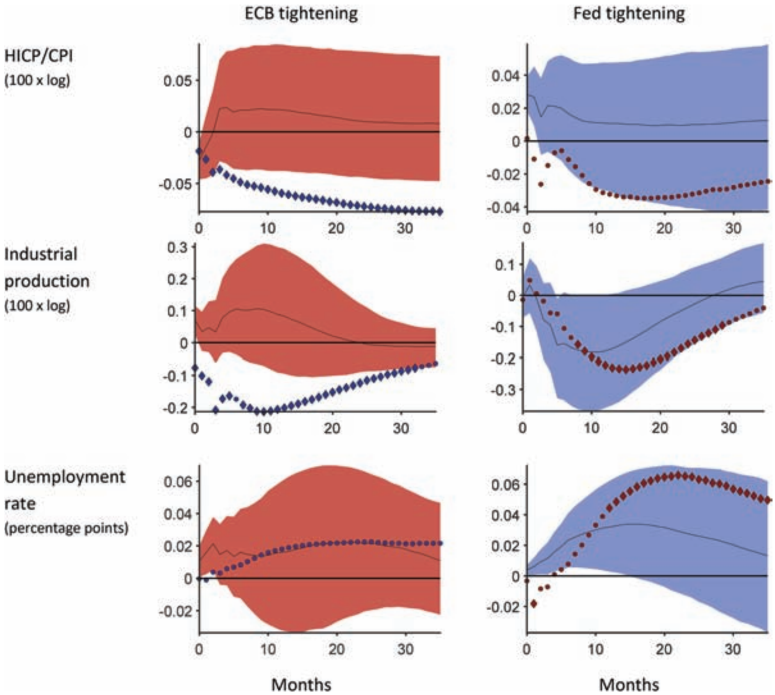
Figure 3 shows the domestic effect of an exogenous monetary policy tightening by the ECB (left column) and by the Fed (right column), together with the corresponding spillover. Domestic effects are reported as a dotted line; they are marked by diamonds instead of the dots whenever the one-standard-error (68 percent) band around the domestic response excludes zero. Spillovers are shown as a solid line surrounded by the one-standard-error (68 percent) confidence band as shaded area (blue for euro-area variables and red for U.S. variables, as before).

The two columns in Figure 3 confirm that both ECB and Fed monetary policy shocks have a substantial impact on domestic consumer prices and real activity, in line with the results in Jarociński and Karadi (2020). After an ECB tightening, euro-area consumer

[12]The effect on interbank lending rates on impact is considerably stronger after a Fed monetary policy shock. Over a one-year horizon, however, the effects from the two monetary policy sources are very similar, as shown in Appendix C.

[13]Under the "rotational sign restrictions" described in Jarociński and Karadi (2020), the U.S. dollar appreciation is persistently statistically significant, but even then, it remains smaller for at least one year than the euro appreciation after an ECB tightening.

Figure 3. Bilateral Spillovers from ECB and Fed Exogenous Monetary Policy Tightening to Real Activity and Consumer Prices



Note: The solid line plots the median impulse response surrounded by the 68 percent confidence band. In the left-hand column these are the responses to an ECB tightening, in the right-hand column responses to a Fed tightening. Quantities for the United States are plotted in red, quantities for the euro area in blue. The dotted lines plot the responses of the corresponding domestic variables, with diamonds symbolizing significance at the 68 percent level. In the left-hand column these are the responses of euro-area variables, in the right-hand column the responses of U.S. variables.

prices drop immediately, statistically significantly, and persistently. The effect of a Fed tightening on U.S. consumer prices is not statistically significant in our baseline. Under the more general “rotational sign restrictions” approach, however, the effect of Fed monetary policy on U.S. consumer prices is comparable in size to that of ECB monetary policy on euro-area consumer prices. The fall in real activity is statistically significant both in the case of an ECB and in the case of a Fed tightening. Domestic unemployment (in the bottom row) also rises in response to an ECB and a Fed tightening, even if the effect is statistically significant in the latter case only. The more limited impact of ECB monetary policy on euro-area unemployment is consistent with the more comprehensive employment protection compared with the United States.

Turning to spillovers, Figure 3 suggests that the cross-border effects of ECB and Fed monetary policy on consumer prices are small and short-lived. An ECB tightening is followed by a marginal decline in U.S. consumer prices on impact, which however turns insignificant very quickly. The spillover from an ECB tightening to U.S. consumer prices is thus very small compared with the sizable domestic effect of a Fed tightening. Against the background of the discussion in Section 2, this evidence is consistent with the absence of a quantitatively significant aggregate demand effect on the one hand (as shown in the middle and bottom rows) and, on the other hand, the large share of U.S. imports from/exports to the euro area being invoiced in U.S. dollars and hence immune to exchange rate variation.

A Fed tightening has a somewhat stronger impact on euro-area consumer prices in the short term. There is a small but statistically significant increase in euro-area consumer prices for about one quarter in response to the Fed tightening. The upward pressure on euro-area consumer prices may be due to an increase in import prices due to the appreciation of the dollar. Indeed, the euro-area GDP deflator, which is not directly exposed to exchange rate changes, responds negatively, though not significantly, to a Fed monetary policy shock.[14] The finding of a statistically significant rise in euro-area consumer prices in the absence of a corresponding increase in the GDP deflator is consistent with a non-trivial share of euro-area imports, not only from the United States, being invoiced in U.S. dollars as discussed in Section 2 under the DCP case; for example, while on average only about 4 percent of euro-area countries' total imports are sourced from the United States, about 24 percent are invoiced in dollars (see Boz et al. 2020).

[14] Please refer to Appendix C for this and more impulse responses.

Turning to real activity, we find that there are no spillovers to unemployment and industrial production from an ECB tightening. In contrast, real activity spillovers from a Fed tightening are large. The impact of an ECB shock on U.S. unemployment is not statistically significant beyond one quarter, while the impact on U.S. industrial production is even slightly expansionary. An expansionary spillover from an ECB tightening to U.S. industrial production may in theory occur as U.S. exporters temporarily gain competitiveness relative to their local competitors in the euro area. However, looking beyond the one-month horizon, spillovers to U.S. industrial production are short-lived and statistically insignificant. The impact of a Fed tightening on real activity in the euro area is sizable and much more long-lasting. Euro-area unemployment and industrial production respond by at least as much as their U.S. counterparts over the course of an entire year, and initially the spillover is even larger than the domestic effect in the United States.[15] After one year the increase in the unemployment rate in the euro area is still close to one-half of the domestic increase in the United States. In the next two subsections we delve into the transmission channels of U.S. spillovers, looking at the responses of trade and financial conditions.

4.3 Effect on Trade

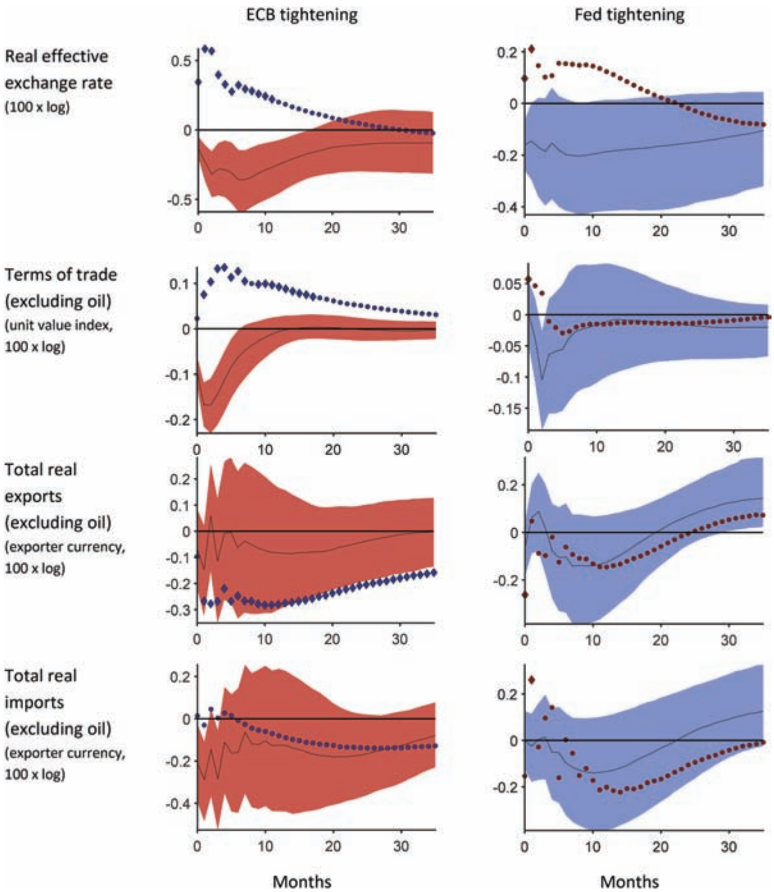
The main international transmission channel of monetary policy in textbook models operates via the effect of exchange rates on exports (e.g., Mundell 1963). In this subsection we explore the relative importance of the aggregate demand and expenditure-switching effects in transmitting spillovers from monetary policy tightening.

Figure 4 shows that a tightening by both the ECB and the Fed induces an appreciation (i.e., an increase in the dotted lines in the first row of Figure 4) of the respective real effective exchange rate (REER). The larger appreciation of the euro REER is in line with our estimates suggesting that the effect of an ECB tightening on the euro-dollar exchange rate is larger than that of a Fed tightening.[16]

[15] The estimated spillover from Fed shocks to the euro area based on "rotational sign restrictions," which—as noted—capture U.S. monetary policy shocks better, is larger.

[16] The response of exchange rates to Fed monetary policy shocks is also weaker under "rotational sign restrictions."

Figure 4. Response of Euro-Area and U.S. Trade to an Exogenous Monetary Policy Tightening



Note: The left-hand column shows the responses to an ECB tightening, and the right-hand column the responses to a Fed tightening. Quantities for the United States are plotted in red, quantities for the euro area in blue. The dotted lines are the responses of the domestic variables, with diamonds symbolizing significance at the 68 percent level. In the left-hand column these are the responses of euro-area variables, in the right-hand column the responses of U.S. variables. The solid line shows the median impulse response of the corresponding spillover with a 68 percent confidence band. The real effective exchange rate is the number of trade-weighted foreign currency units per home currency unit. Within any given graph, the dotted line and the solid line therefore differ in their denominator.

Because of the substantial trade links between the euro area and the United States, both central banks also affect each other's REER. A Fed tightening triggers a persistent and marginally statistically significant depreciation of the euro REER (a decrease in the solid line in the right column). An ECB tightening leads to a depreciation of the U.S. dollar REER for more than one year. Of course, these results are at least in part due to the weight each currency has in the effective exchange rate of the other.

The dotted line in the right column in the second row of Figure 4 shows that the U.S. terms of trade vis-à-vis the rest of the world barely respond to a Fed tightening. This limited response is consistent with large LCP and PCP components in U.S. imports and exports, respectively, under which U.S. export and import prices are both sticky in dollars.[17] The small improvement in the U.S. terms of trade on impact may be due to some share of bilateral imports from the euro area being invoiced in euros. The left column shows that the U.S. terms of trade respond more strongly to an ECB tightening, which is consistent with the considerably stronger response of the bilateral exchange rate already seen in Figure 2.

The drop of U.S. exports in response to a Fed tightening shown in the right column of the third row is consistent with expenditure switching and negative demand effects in the rest of the world due to contractionary spillovers, but it is not estimated very precisely. Similarly, the drop in U.S. imports in the last row in the right column is consistent with negative demand effects and a lack of expenditure switching in the United States given the stickiness of import prices in dollars, but it is again not estimated precisely. The left column shows that U.S. exports do not respond to an ECB tightening, while U.S. imports fall somewhat, consistently with the more pronounced response of the U.S. terms of trade in the second row.

The second row of the left column of Figure 4 shows that an ECB tightening improves (i.e., raises) the euro area's terms of trade statistically significantly and persistently. In fact, while a large share of euro-area countries' total imports and exports is invoiced in euros, namely about 71 percent and 74 percent, respectively, a non-trivial share is invoiced in other currencies (see Boz et al. 2020). Similarly,

[17] From the perspective of the United States, DCP is equivalent to PCP in exports and LCP in imports.

after a Fed tightening euro-area terms of trade deteriorate for several months, as shown in the right column in the second row of Figure 4. This matches the observation that—consistent with the DCP paradigm—the dollar plays an outsized role among non-euro invoicing currencies in euro-area imports, i.e., that the dollar is used widely for invoicing of trade that does not involve the United States.

An ECB tightening strongly affects euro-area exports, but not imports, as shown in the last two rows of the left column of Figure 4. Euro-area exports decline statistically significantly and persistently after an ECB tightening. Such a drop is consistent with their prices being sticky in euros to a large degree. The limited response of euro-area imports suggests that the expenditure-switching and aggregate demand channels largely offset each other; the possibility that euro-area imports barely move because their prices are largely sticky in euros is hard to reconcile with the above evidence that the euro-area terms of trade improve in response to an ECB tightening.

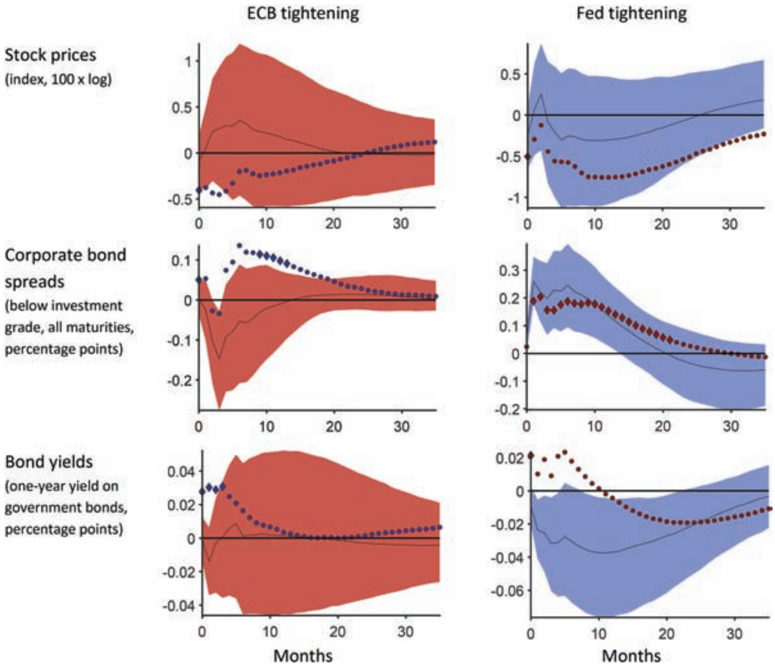
The third and fourth row in the right column show that while euro-area trade slows down in response to a Fed monetary policy tightening, this response is not estimated very precisely. This may be because the share of total euro-area exports absorbed by the United States (about 6 percent of total euro-area countries' exports) or whose prices are sticky in dollars (about 20 percent of total euro-area countries' exports; see Boz et al. 2020) is not too large. Similarly, only a small share of the euro area's imports is sourced from the United States, and only a limited share is invoiced in dollars and therefore subject to expenditure switching in the face of a dollar appreciation.

Overall, despite their effect on exchange rates and the terms of trade, there are only relatively limited spillovers from ECB monetary policy to U.S. trade and from Fed monetary policy to euro-area trade.

4.4 Effect on Financial Conditions

The comparable size of the real economies of the euro area and the United States contrasts sharply with the unequal global importance of their financial sectors and currencies. The global dominance of U.S. financial markets and the U.S. dollar renders the financial

Figure 5. Financial Spillovers of an Exogenous Monetary Policy Tightening



Note: The left-hand column shows the responses to an ECB tightening, and the right-hand column the responses to a Fed tightening. Quantities for the United States are plotted in red, quantities for the euro area in blue. The dotted lines are the responses of the domestic variables, with diamonds symbolizing significance at the 68 percent level. In the left-hand column these are the responses of euro-area variables, in the right column the responses of U.S. variables. The solid line shows the median impulse response of the corresponding spillover with a 68 percent confidence band. The stock index shown for the euro area is the Euro Stoxx 50, and the index for the United States the S&P 500. The corporate bond spread is the option-adjusted spread between a corporate bond with BBB or below investment-grade rating and a government bond. The government bonds used are bunds for the euro area and Treasuries for the United States.

channel between the United States and the euro area almost unidirectional, and accounts for the spillovers from U.S. monetary policy on euro-area real activity.

Our findings suggest that financial spillovers from ECB and Fed monetary policy are asymmetric. Figure 5 shows the spillovers to three financial variables: stock prices, spreads of speculative-grade corporate bonds, and sovereign bond yields. For stocks, as shown in the top row of the figure, even though both ECB and Fed tightening reduce domestic prices on impact, bilateral spillovers are not statistically significant. Note that this evidence is not inconsistent with a very short-lived positive correlation between euro-area and U.S. stock markets after a monetary policy shock over a couple of days that is no longer detectable at monthly frequency.

At the same time, we find that Fed monetary policy strongly affects interest rates and financing conditions in the euro area. The middle row of Figure 5 shows the spread between a basket of corporate bonds below investment grade (i.e., rated BBB or below, with interest rates adjusted for any included option value) and government bonds. After a Fed tightening, there is a statistically significant and persistent increase in the spread of euro-area speculative-grade corporate bonds.[18] In fact, this financial spillover is as large as the response of domestic (U.S.) corporate bond spreads, shown by the red dotted line in the right column. The bottom row shows that one-year bund yields decline after a Fed tightening,[19] pointing to some (systematic) offsetting response by the ECB to mitigate at least some of the effects of Fed monetary policy shocks on the euro-area economy.[20] Clearly, financial spillovers are much smaller in the case of an ECB tightening than in the case of a Fed tightening, as responses of U.S. variables in the left column of the figure are small and never statistically significant. A corollary of this finding is that monetary policy spillovers do not significantly contribute to the observed co-movement between euro-area and U.S. financial variables such as stock prices and sovereign rates.[21]

[18]These corporate bonds were not part of the ECB's asset purchase program. It is conceivable that in recent years they have been more isolated from Fed monetary policy shocks, but this period is too short to significantly dampen the estimated spillover from monetary policy shocks. The impulse responses show the average spillover over 20 years.

[19]Under "rotational sign restrictions," after a Fed tightening euro-area one-year interest rate swaps decline significantly as well.

[20]Recall that all euro-area impulse responses are net of the effect of any systematic ECB policy response.

[21]Because the effects shown are long-term averages, they are robust to isolated spillover episodes, e.g., in the context of unconventional monetary policy. The period of unconventional monetary policies in the euro area is too short to confirm or reject a possible change in spillovers from the ECB to the United

In contrast, an ECB monetary policy tightening does not have a comparable effect on U.S. financial conditions. U.S. corporate bond spreads—unlike euro-area corporate bond spreads—do not increase after an ECB tightening. And although one-year government bond yields in the euro area respond significantly to the ECB monetary policy shock, there are no discernible spillovers to U.S. bond yields.[22]

In summary, a Fed tightening tightens financial conditions in the euro area as captured by corporate bond spreads, while an ECB tightening does not affect financial conditions in the United States.

5. International Effects of ECB and Fed Monetary Policy

This section explores whether the asymmetry found in the previous section is specific to the euro-area–U.S. bilateral links or whether it can be extended to the global economy.

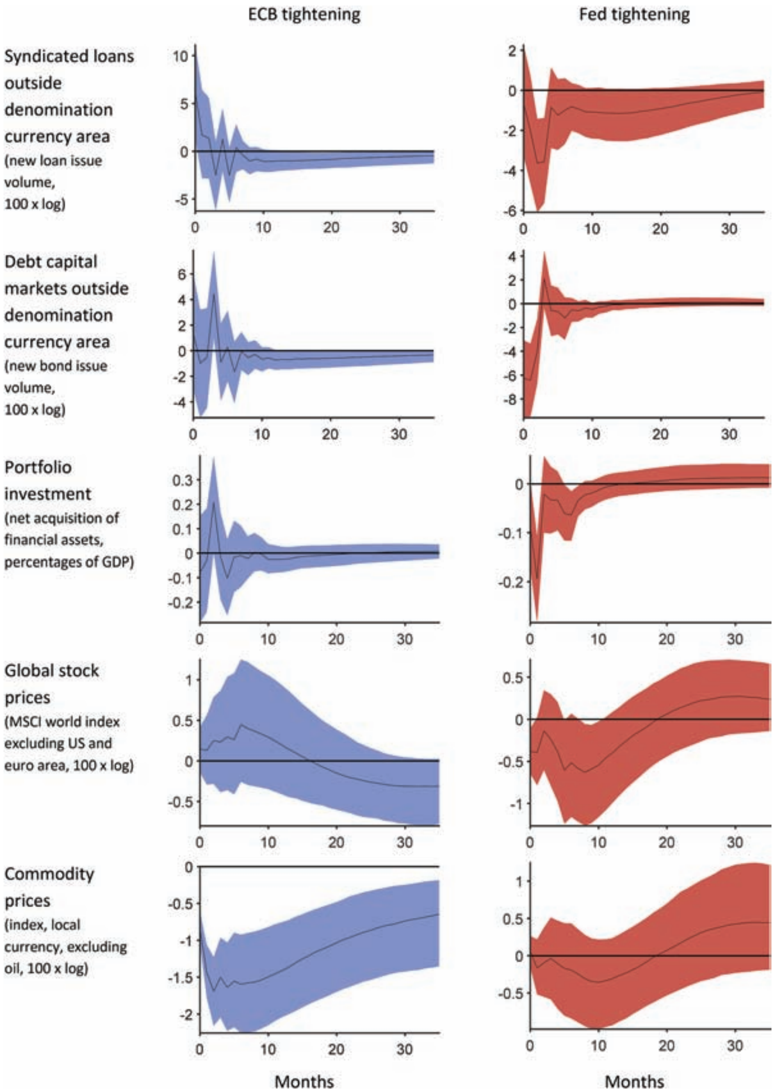
5.1 Effect on Global Financial Markets

Figure 6 presents the effects of ECB and Fed monetary policy on global financial markets. The right column documents that a Fed tightening contracts borrowing denominated in U.S. dollars worldwide. A prominent example is new debt issued by borrowers whose business activity is mostly located outside of the United States (and other countries which use the U.S. dollar as their official currency). After a Fed tightening, new issuance of dollar-denominated syndicated loans outside the United States drops by up to 4 percent, and new dollar-denominated debt capital, which includes all sectors (government, financial, and non-financial) drops by even more (first and second row of Figure 6, respectively). One might conjecture that this drop is mainly due to the financial sector, but this is not the

States in recent years. Using a different approach, Curcuru et al. (2018) find that between the euro area and the United States, spillovers of conventional monetary policy (as measured by changes in expected interest rates) on 10-year yields were not significantly different from spillovers of unconventional monetary policy (as measured by changes in term premia).

[22] After both ECB and Fed tightening, the respective domestic bond yields respond very similarly on impact. During the first half-year after the shock a considerable term spread opens up, as shown in Appendix C. Only thereafter do longer maturities follow short-term rates, as noted by Hanson and Stein (2015).

Figure 6. Effects of an Exogenous Monetary Policy Tightening on Global Financial Markets



Note: The solid line shows the median impulse response surrounded by the 68 percent confidence band. In the left-hand column these are the responses to an ECB tightening, in the right-hand column the responses to a Fed tightening. Quantities related to the euro area are plotted in blue, quantities related to the United States in red.

case. Non-financial corporations outside the United States reduce their issuance of U.S. dollar debt by just as much.[23] This reveals a direct link between U.S. monetary policy and investment activity in the rest of the world, i.e., a powerful financial spillover channel from Fed monetary policy to the rest of the world.

In contrast, the effect of an ECB tightening on borrowing in euros outside the euro area, shown in the left column of the figure, is not contractionary and is barely statistically significant. Euro-denominated borrowing outside the euro area is less common than U.S. dollar-denominated borrowing outside the United States.[24] But even in relative terms the effect of ECB monetary policy on foreign euro borrowing is much smaller and insignificant.

Likewise, U.S. international net portfolio investment drops significantly after a Fed tightening. As shown in the third row of Figure 6, after a Fed tightening, U.S. residents acquire a significantly smaller amount of foreign financial assets, i.e., claims against non-U.S. residents, in net terms. Since U.S. liabilities in net terms drop simultaneously, cross-border financial investment positions shrink, consistent with the hypothesis of Fed monetary policy greatly affecting the global financial cycle (Miranda-Agrippino and Rey 2022). Again, the analogous effects of ECB tightening have the opposite sign and are not statistically significant.

Next, a Fed tightening also depresses global stock markets, while an ECB tightening is inconsequential. Global stock prices, summarized by the MSCI index excluding both U.S. and euro-area stocks in the fourth row of Figure 6, fall on impact and a few months after a Fed tightening, but are unchanged after an ECB tightening. It is also worth noting that global stock prices respond more strongly to the Fed's monetary policy than the euro-area stock prices presented earlier in Section 4.4.

[23] Please refer to Appendix C for this and further impulse responses.

[24] Since 1999 on average about 31 percent of U.S. dollar-denominated syndicated loans and about 22 percent of debt capital have been issued outside the United States, whereas of the corresponding euro-denominated assets only about 4 percent and 7 percent, respectively, have been issued outside the euro area. The volume in these foreign euro-denominated markets over the same period was about one-sixth of that in foreign U.S.-dollar-denominated markets for syndicated loans, and one-third in the case of debt capital.

An ECB tightening has a statistically significant impact on non-oil commodity prices, however. Although the spillovers from ECB monetary policy via financial conditions are negligible, there seems to be an indirect impact via the euro-area business cycle. An ECB tightening leads to a drop in non-oil commodity prices quoted in U.S. dollars and, in combination with the euro appreciation, to an even larger drop in commodity prices quoted in euros. The effect of a Fed tightening is statistically significant only under the "rotational sign restrictions."

5.2 *Effect on EMEs*

The first row of Figure 7 shows that financial spillovers to EMEs from Fed monetary policy are much more consequential than those from ECB monetary policy. After a Fed tightening, the statistically significant drop in EME stock prices is larger than that in both euro-area and global stock prices shown in Figures 5 and 6; after an ECB tightening, EME stock prices barely move. This mirrors the strong spillovers from Fed tightening to financial markets of EMEs found by Hoek, Kamin, and Yoldas (2020).[25]

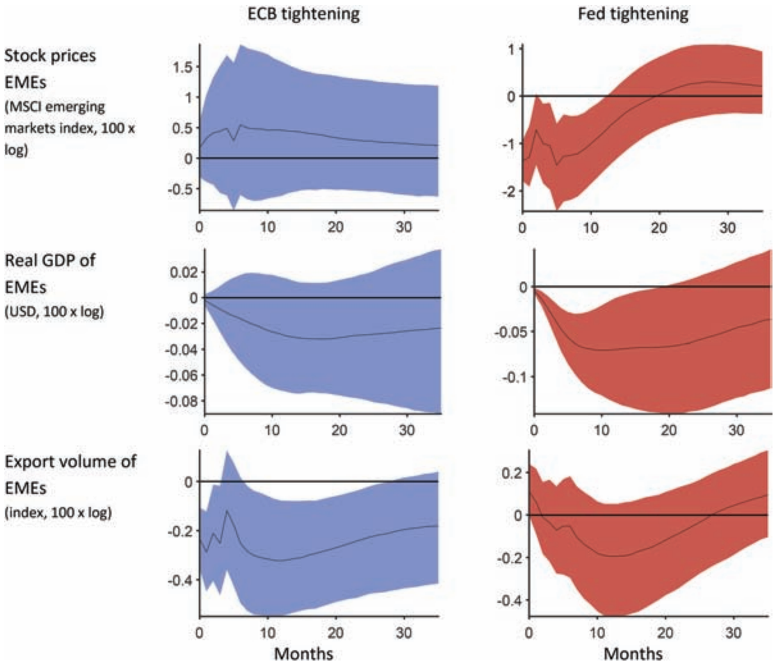
The evidence for asymmetric effects of Fed and ECB monetary policy on EMEs extends to real activity (Avdjiev et al. 2019). Monetary policy tightening by both the ECB and the Fed reduces real GDP in EMEs, but the effect is statistically significant only in the case of a Fed tightening.

Both the ECB and the Fed affect EMEs' trade. After an ECB tightening, exports (including energy) decline persistently, but with only a limited impact on overall GDP. Exports similarly fall after a Fed tightening, although the decline is statistically significant only under the "rotational sign restrictions" approach.[26] This suggests that Fed monetary policy affects EME real activity beyond trade.

[25] Hoek, Kamin, and Yoldas (2020) consider identification of the drivers of Fed tightening based on high-frequency moves in U.S. Treasury yields and stock prices around Federal Open Market Committee announcements and U.S. employment report releases. They interpret positive co-movements between stocks and interest rates around these events as growth shocks and—as we do in this paper—negative co-movements as monetary shocks.

[26] Likewise, the remaining impulse responses to a Fed shock in the figure are statistically significant (or more precisely estimated) under the "rotational sign restrictions" approach.

Figure 7. Effects of an Exogenous Monetary Policy Tightening on EMEs



Note: The solid line shows the median impulse response surrounded by the 68 percent confidence band. In the left-hand column these are the responses to an ECB tightening, in the right-hand column the responses to a Fed tightening.

To better understand the transmission channels of monetary policy, we focus on non-oil trade, which however requires narrowing the analysis to a smaller set of countries.[27] We focus on spillovers to Brazil, Russia, India, China, and South Africa, often referred to as "BRICS" countries. After an ECB and Fed tightening, non-oil exports from, respectively, the euro area and the United States to BRICS countries decline persistently. Imports from BRICS countries decline with a delay, but this is only statistically significant in case of the Fed tightening. Overall, the effects on trade with the BRICS countries are similar for both Fed and ECB tightening. The

[27]Please refer to Appendix C for these impulse responses.

demand effects that dampen imports from BRICS countries about half a year after the tightening are, however, more pronounced in case of the Fed, in line with the stronger impact of Fed monetary policy on, e.g., U.S. unemployment and the weaker appreciation of the U.S. dollar (see Figures 2 and 3).

5.3 A Hierarchy of Spillovers from Monetary Policy

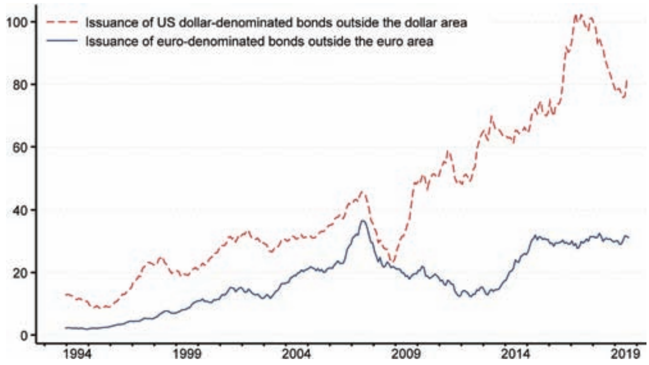
The comparison of the effects of ECB and Fed monetary policy in a unified and coherent framework suggests a pronounced asymmetry in their spillovers. Overall, a hierarchy of monetary policy spillovers emerges, which places the Fed ahead of the ECB in terms of the global impact of its monetary policy. Financial spillovers from Fed monetary policy spread to the euro area and other countries, affecting real activity. Trade spillovers from ECB monetary policy barely affect the United States but spread to other countries. Finally, significant spillovers from ECB and Fed monetary policy may imply policy trade-offs in EMEs.

The asymmetry of monetary policy spillovers is most pronounced in financial channels. The importance of this channel of Fed monetary policy has been illustrated during the COVID-19 crisis, when aggressive action by the Fed, especially in mid-March 2020, resulted in a normalization of financing conditions globally. At least three well-known factors contribute to the asymmetry of the international impact of ECB and Fed monetary policy: the central role of U.S. financial markets, the dominant role of the U.S. dollar, and the relatively low trade openness of the United States, with respect to the euro area.

The first factor is the central role of U.S. financial markets. U.S. financial markets represent a global financial hub, whose size and global interconnectedness can be seen, for example, in the importance of the U.S. dollar lending market. Figure 8 shows the issuance volume of bonds outside of their home-currency area. Since 2009 the U.S. dollar bond market has been about three times as large as the euro bond market.[28] In fact, it has been argued in this vein that

[28]This difference reemerged after the sovereign bond crisis. Since then, the euro's relevance in bond markets has been falling behind that of the U.S. dollar, enforcing the dominant role of the United States in global financial conditions. In

Figure 8. Issuance Volume of Bonds Outside the Home-Currency Area (EUR billions, 12-month moving average)



Source: Dealogic and ECB calculations.

Fed monetary policy is a major driver of the global financial cycle (Miranda-Agrippino and Rey 2020).[29]

A second factor is that the U.S. dollar remains the globally dominant currency.[30] A dominant trade-invoicing currency changes

the period up to the year 2007, both euro and dollar bond markets grew rapidly. The euro market grew disproportionally strongly, boosted by the strong euro appreciation, so that by 2007–08 it had largely caught up with the dollar bond market. But after the sovereign bond crisis, the euro bond market only recovered to pre-crisis levels, whereas the dollar bond market kept growing at a constant rate.

[29]For international capital flows, Fed monetary policy may be less important than financial shocks. Habib and Venditti (2019) find that changes in global risk caused by "pure" financial shocks have an even larger effect on capital flows than Fed monetary policy shocks.

[30]The dominant role of the U.S. dollar is particularly sizable in terms of foreign exchange rate holdings, issuance of international debt by international borrowers, and international loans in foreign currency and foreign exchange turnover. Two other pieces of evidence confirm the persistent centrality of the U.S. dollar in global financial markets during periods of high economic tensions. The first piece of evidence is the correlation between the effective exchange rate of the U.S. dollar and global (non-U.S.) stock markets during period of turbulence. The second piece of evidence is the positive correlation between the U.S. dollar and the VIX index. The centrality of the U.S. dollar is explained, besides historical reasons such as the importance of the dollar as an anchor currency, by the safety and liquidity of dollar-denominated assets in crisis periods (Maggiori, Brent, and Schreger 2019) and by EME firms' incentives to issue dollar-denominated bonds

transmission via the expenditure-switching channel. Our results suggest, however, that the central role of the U.S. dollar in trade invoicing may be only one aspect of dollar dominance (Boz, Gopinath, and Plagborg-Moller 2017; Gopinath et al. 2020). More important is the U.S. dollar's dominance in the pricing of financial assets, which amplifies the exchange rate effect of Fed monetary policy, affecting global financial conditions. As important financial assets are denominated in U.S. dollars, dollar exchange rate fluctuations affect balance sheet positions worldwide. Thereby, Fed monetary policy can influence not only borrowing and lending in any currency but also any capital-intensive economic activity, even global value chains (Bruno and Shin 2015, 2020).

The third factor which helps explain the asymmetry across the two regions is the stronger importance of trade for the euro area than for U.S. GDP. Indeed, trade openness measured in terms of imported goods and services relative to GDP is approximately 25 percent for the euro area and only 15 percent for the United States. Figures 4 and 7 highlight the pronounced responsiveness of euro-area trade to exchange rate movements caused by monetary policy shocks, especially in trade with EMEs.

The first two factors contribute to the potential special role of Fed monetary policy as a driver of the global financial cycle. The global effects of Fed monetary policy are strong, and are not limited to EMEs, but are seen also in the bilateral spillovers to the euro area (see, e.g., Figure 3). Further adding to this might be other countries mimicking Fed monetary policy (Mukhin 2018; Corsetti et al. 2021; Georgiadis and Zhu 2021).

The third factor helps explain why ECB spillovers tend to transmit mainly via trade channels. These spillovers via trade are stronger for other economies than bilaterally with the United States. For trade partners other than the United States, the euro area's import share of goods is about 90 percent, so that ECB monetary policy can spill over to smaller countries via, specifically, the demand and expenditure-switching channels.

(Bruno and Shin 2017) in periods of favorable U.S. dollar carry trades. The relative strength of the euro during the recent pandemic episode signals the relevance of the enhanced institutional architecture of the Economic and Monetary Union and of high-quality marketable euro-denominated assets in determining the future role of the euro (Ilzetzki, Reinhart, and Rogoff 2020).

6. Concluding Remarks

In this paper we estimate spillovers from Fed and ECB monetary policy in a unified and state-of-the-art econometric framework. We find that bilateral spillovers between the United States and the euro area are generally small. Our finding of a small impact of ECB monetary policy on the U.S. economy could be due to two reasons: First, it might reflect that the Fed has been able and determined to fully offset spillovers from ECB monetary policy; alternatively, it might reflect that ECB monetary policy spillovers to the United States have been small in the first place. We also find that the Fed monetary policy spillovers to the euro area are statistically significant—at least for unemployment—and there is evidence of a stronger responsiveness of euro-area financial conditions to Fed monetary policy conditions. This finding could imply that the ECB faces trade-offs between price stability and stabilization of output and/or financial conditions in the very short term. Notwithstanding this possibility, our estimates suggest that the ECB has been successful in preserving price stability in the face of Fed monetary policy spillovers at least over the medium term.

Looking at spillovers from Fed and ECB monetary policy to EMEs, our findings suggest a key role for the trade and financial channels. Moreover, our findings point to a pronounced asymmetry across ECB and Fed monetary policy spillovers. The large spillovers from Fed monetary policy via financial channels are consistent with the central role of U.S. financial markets and the U.S. dollar, and the spillovers from ECB monetary policy through trade are consistent with the euro area's strong integration in world trade. The policy trade-offs in EMEs implied by our finding of non-trivial spillovers could be rooted in more pronounced and widespread frictions in these economies.

Appendix A

A.1 High-Frequency Identification of Monetary Policy Shocks

As highlighted in Section 3, we use financial market reactions to monetary policy announcements to identify monetary policy shocks. The approach follows Jarociński and Karadi (2020).

The data set consists of 248 Federal Reserve monetary policy announcements since 1990 and 283 ECB monetary policy announcements since 1999. Financial market reactions are measured in the time window starting 10 minutes before and ending 20 minutes after a central bank announcement. In the case of the Federal Reserve, the announcement time is typically the release time of the press release. In the case of the ECB, the time window is longer and ends 20 minutes after the end of the press conference (when there is one). In these windows we record interest rate surprises and stock price surprises. The Federal Reserve interest rate surprises are defined as the mean of the changes in federal funds futures and Eurodollar futures with remaining maturities from one month up to one year. The ECB interest rate surprises are defined as the mean of the changes in EONIA swaps with maturities from one month up to one year. By including maturities up to one year, these surprises capture not just the changes of the current policy rates but also of the expectations for interest rates up to one year into the future, reflecting forward guidance and non-standard policies. The Federal Reserve stock price surprises are measured as the change in the S&P 500 stock index, and the ECB stock price surprises are measured as the change in the Euro Stoxx 50 stock index. We aggregate these surprises to monthly frequency.

In the next step we isolate the monetary policy shocks from among the interest rate surprises by purging the information effects from them. This is based on the sign restriction: interest rates and stock prices are assumed to co-move negatively after a monetary policy shock, as is implied by a wide range of models. Therefore, we treat as monetary policy shocks only those interest rate surprises which co-move negatively with stock prices in the respective month. A more sophisticated alternative is to decompose interest rate and stock price surprises into two orthogonal components and "rotate" them so that one is associated with a negative co-movement and the other with the positive co-movement of interest rate and stock price surprises. Jarociński and Karadi (2020) find that in the large sample for the United States the two approaches yield similar results, but in the euro-area sample the former, simpler approach, dubbed "poor man's sign restrictions," yields a stronger instrument for monetary policy (i.e., is associated with a stronger increase in the one-year bond yield) than the more sophisticated sign restrictions approach. Therefore, we use this approach for both the United States and the euro area for comparability. The Federal Reserve monetary policy surprise in April 2001 is larger than six standard deviations of monetary policy shocks. We exclude this stark outlier from the analysis.

A.2 Estimation of the Impulse Responses

We track the responses of the economy to the identified shocks using a VAR. The baseline VAR for each country includes the one-year government bond yield, stock prices, the corporate bond spread, industrial production, and the respective consumer price index (CPI/HICP). We add the identified shock to this VAR as the first variable. We restrict the coefficients of the first equation to zero, reflecting the fact that the shock is independently and identically distributed. After estimating the VAR with the standard Minnesota prior, we compute the impulse responses to the first shock identified recursively. The variables and the estimation of the baseline VAR are the same as in Jarociński and Karadi (2020). This paper also provides details on the rotational sign restrictions approach, which yields a stronger instrument for Federal Reserve monetary policy shocks, but not for ECB shocks.

We compute the effect of the Federal Reserve policies on the euro area (plotted in the graphs as solid line) by combining the Federal Reserve shocks with the euro-area variables in the VAR. Analogously, for the domestic effects of the Federal Reserve policies (plotted with dots) we combine Federal Reserve shocks with the U.S. variables in the VAR. To obtain the effect of the ECB policies, the setup is simply mirror-inverted: the domestic effect of ECB policies is based on the effect of ECB shocks on euro-area variables; the spillover effect to the United States is based on the effect of ECB shocks on U.S. variables. The responses of other variables, which are not part of the baseline VAR specification, are computed by adding them one by one as the last variable to the respective baseline VAR.

Several variables that we study are bilateral: the exchange rate, the spread between the United States and the euro-area bond, etc. In these cases, we use a bilateral VAR specification[31] to compute

[31] Georgiadis (2017) compares the performance of bilateral VARs with multilateral (global) VARs.

their impulse responses. The bilateral VAR includes the exchange rate, the spread between U.S. Treasuries and one-year bund yields, the corporate bond spread of the country experiencing the shock, and, separately for the United States and the euro area, industrial production and consumer price indices.

Appendix B. Data

The monetary and central bank information shocks are part of the online appendix to Jarociński and Karadi (2020).

All series cover the period January 1999–December 2018, unless otherwise noted.

Table B.1. Variable Descriptions

Variable	Euro Area	United States
Interest Rate Differential	Yield spread between the one-year bund and the benchmark one-year Treasury as calculated by Thomson Reuters, end-of-month. Source: Thomson Reuters.	Yield spread between the benchmark one-year Treasury and the one-year bund as calculated by Thomson Reuters, end-of-month. Source: Thomson Reuters.
Exchange Rates	ECB reference USD/EUR exchange rate, rebased to EUR/USD, monthly average. Source: ECB Statistical Data Warehouse (SDW).	ECB USD/EUR reference exchange rate, monthly average. Source: SDW.
HICP/CPI	Harmonised Index of Consumer Prices (HICP) for the euro area in changing composition, working-day and seasonally adjusted, monthly. Source: SDW.	Consumer Price Index (CPI) for all urban customers: all items, seasonally adjusted, monthly. Source: Federal Reserve Bank of St. Louis (FRED database).
Industrial Production	Industrial production index, excluding construction, monthly. Source: SDW.	Industrial production index, excluding construction, monthly. Source: FRED.
Unemployment Rate	Standardized unemployment rate for euro area in fixed composition (19 countries), seasonally but not working-day adjusted, monthly. Source: SDW.	Civilian unemployment rate, seasonally adjusted, monthly. Source: FRED.
Real Effective Exchange Rate	Real broad effective exchange rate of the euro area, trade weighted, deflated by relative consumer prices, as calculated by the Bank for International Settlements, monthly. Source: Haver Analytics.	Real broad effective exchange rate of the United States, trade-weighted, deflated by relative consumer prices, as calculated by Bank for International Settlements, monthly. Source: Haver Analytics.
Terms of Trade (Excluding Oil)	Export prices over import prices in euro, unit value index, monthly, sample: 2000–18. Source: Haver Analytics.	Export prices over import prices in U.S. dollars, unit value index, monthly, sample: 2000–18. Source: Haver Analytics.

(continued)

Table B.1. (Continued)

Variable	**Euro Area**	**United States**
Total Real Exports (Excluding Oil)	Euro-area trade data, deflated by price indicators, excluding oil trade, monthly, sample: 2000–18. Sources: Haver Analytics, ECB calculations.	U.S. trade data, deflated by price indicators, excluding oil trade, monthly, sample: 2000–18. Sources: Haver Analytics, ECB calculations.
Total Real Imports (Excluding Oil)	Euro-area trade data, deflated by price indicators, excluding oil trade, monthly, sample: 2000–18. Sources: Haver Analytics, ECB calculations.	U.S. trade data, deflated by price indicators, excluding oil trade, monthly, sample: 2000–18. Sources: Haver Analytics, ECB calculations.
Stock Prices	Dow Jones Euro Stoxx 50, historical close, end-of-month. Source: SDW.	S&P Dow Jones S&P 500 index, end-of-month. Source: SDW.
Corporate Bond Spreads	Average spread between euro-area corporate bonds and euro-area government bonds; ICE BofAML Euro High Yield Index option-adjusted spread, basket of corporate bonds below investment grade (i.e., rated BBB or below), spreads averaged across maturities, monthly average. Source: FRED.	Average spread between U.S. corporate bonds and U.S. government bonds; ICE BofAML U.S. High Yield Master II option-adjusted spread, basket of corporate bonds below investment grade (i.e., rated BBB or below), spreads averaged across maturities, monthly average. Source: FRED.
Bond Yields	Thomson Reuters benchmark one-year German government bond bid yield, end-of-month. Source: Thomson Reuters.	Thomson Reuters benchmark one-year Treasury bid yield, end-of-month. Source: Thomson Reuters.
Syndicated Loans Outside Denomination Currency Area	Volume of syndicated loans denominated in euro (or the former currency of one of the 11 initial euro-area countries) and newly issued outside of countries with the euro as official currency (pre-1999: outside of the 11 initial member countries), monthly, unit: euro. Sources: Dealogic, ECB calculations.	Volume of syndicated loans denominated in U.S. dollars and newly issued outside of countries with the U.S. dollar as official currency, unit: U.S. dollar, monthly. Sources: Dealogic, ECB calculations.

(*continued*)

Table B.1. (Continued)

Variable	Euro Area	United States
Debt Capital Markets Outside Denomination Currency Area	Volume of newly issued debt capital denominated in euro (or the former currency of one of the 11 initial member countries) outside of countries with the euro as official currency (pre-1999: outside of the 11 initial member countries), monthly, unit: euro. Sources: Dealogic, ECB calculations.	Volume of newly issued debt capital denominated in U.S. dollars outside of countries with the U.S. dollar as official currency, unit: U.S. dollar, monthly. Sources: Dealogic, ECB calculations.
Portfolio Investment—Financial Assets	Net acquisition of financial assets, euro area in fixed composition (19 countries) vis-à-vis rest of the world, not seasonally adjusted, monthly, at market value. Source: SDW (balance of payments and international investment position).	Net purchases of foreign securities by U.S. residents, not seasonally adjusted, monthly, at market value. Sources: Treasury International Capital (TIC) Data, Haver Analytics.
Global Stock Prices	Weighted average of MSCI country indices underlying the MSCI World and MSCI emerging markets indices excluding the United States and the euro area, local currency, all series rebased to 100 in January 2010, countries weighted by market capitalization in U.S. dollar, end-of-month. Source: Bloomberg, ECB calculations.	
Commodity Prices in U.S. Dollars	Market prices of raw materials, excluding energy, in U.S. dollars, monthly average. Source: OECD.	
Commodity Prices in Euro	Market prices of raw materials, excluding energy, in euro, monthly average. Source: OECD.	
Stock Prices in Emerging Economies	MSCI emerging markets index, U.S. dollar, end-of-month. Source: Bloomberg.	
Real GDP of Emerging Economies	GDP at prices and exchange rates in 2010, seasonally adjusted, in U.S. dollars, cubic spline interpolation from quarterly data, sum of the following countries: Bolivia, Botswana, Brazil, Chile, China, Costa Rica, Ecuador, El Salvador, Hong Kong, India, Indonesia, Israel, Jordan, Kazakhstan, South Korea, Malaysia, Mexico, Paraguay, Peru, Philippines, Poland, Russia, Singapore, South Africa, Taiwan, Thailand, Turkey, Uruguay. Source: Haver Analytics.	
Export Volume of Emerging Economies	Export volume index including energy, 43 countries, monthly. Source: Haver Analytics and CPB Netherlands Bureau for Economic Policy Analysis World Trade Monitor.	

Table B.2. Description of Additional Variables Used in the Online Appendix

Variable	Euro Area	United States
Interest Rate Differential	Spread between one-month EURIBOR and LIBOR, end-of-month. Source: SDW.	Spread between one-month LIBOR and EURIBOR, end-of-month. Source: FRED.
Core HICP/CPI	HICP for all items excluding energy and food, for euro area in changing composition, working-day and seasonally adjusted, monthly. Source: SDW.	CPI for all urban customers, all items less food and energy in U.S. city average, seasonally adjusted, monthly. Source: FRED.
GDP Deflator	GDP deflator interpolated from quarterly data, employing a similar strategy as for the United States. Source: SDW.	GDP deflator interpolated from quarterly data, following Stock and Watson (2010). Source: FRED.
Real GDP	Real GDP interpolated from quarterly data, employing a similar strategy as for the United States. Source: SDW.	Real GDP interpolated from quarterly data, following Stock and Watson (2010). Source: FRED.
Total Real Trade Balance (Excluding Oil)	Euro-area trade data, deflated by price indicators, excluding oil trade, monthly, sample: 2000–18. Source: Haver Analytics, ECB calculations.	U.S. trade data, deflated by price indicators, excluding oil trade, monthly, sample: 2000–18. Source: Haver Analytics, ECB calculations.
Interest Rate Swaps	One-year interest rate swap based on six-month EURIBOR, monthly average. Source: Bloomberg.	One-year interest rate swap based on three-month LIBOR, monthly average. Source: Bloomberg.
Term Spreads	Spread between 10- and 1-year yields based on the estimated German government debt yield curve, end-of-month. Source: Thomson Reuters.	Spread between 10- and 1-year Treasury constant maturity rates, end-of-month. Source: Thomson Reuters.

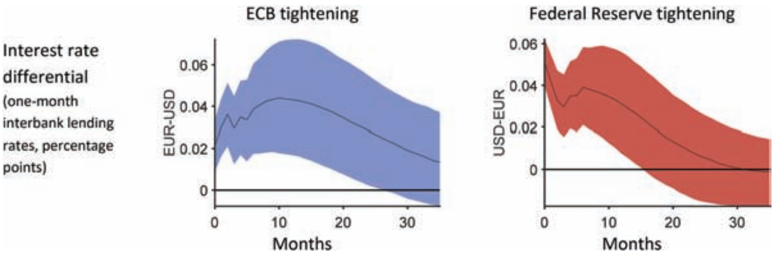
(*continued*)

Table B.2. (Continued)

Variable	Euro Area	United States
New Issuance of Bonds Outside Denomination Currency Area by Non-financial Corporations	Volume of debt capital denominated in euro (or the former currency of one of the 11 initial euro-area countries) and newly issued outside of countries with the euro as official currency (pre-1999: outside of the initial member countries), non-financial corporations only, monthly, unit: euro. Sources: Dealogic, ECB calculations.	Volume of debt capital denominated in U.S. dollars and newly issued outside of countries with the U.S. dollar as official currency, non-financial corporations only, unit: U.S. dollar. Sources: Dealogic, ECB calculations.
New Issuance of Bonds Outside Denomination Currency Area, Low or Junk Rating	Volume of debt capital denominated in euro (or the former currency of one of the 11 initial euro-area countries) and newly issued outside of countries with the euro as official currency (pre-1999: outside of the initial member countries), low or junk rating only, monthly, unit: euro. Sources: Dealogic, ECB calculations.	Volume of debt capital denominated in U.S. dollars and newly issued outside of countries with the U.S. dollar as official currency, low or junk rating only, monthly, unit: U.S. dollar. Sources: Dealogic, ECB calculations.
Portfolio Investment—Financial Liabilities	Net incurrence of financial liabilities, euro area in fixed composition (19 countries) vis-à-vis rest of the world, not seasonally adjusted, monthly, at market value. Source: SDW (balance of payments and international investment position).	Net foreign purchases of U.S. securities, not seasonally adjusted, monthly, at market value. Sources: TIC data, Haver Analytics.
Oil Price (in U.S. Dollars, in Euros)	Spot price of West Texas Intermediate. Source: FRED.	
Real Exports to BRICS (Excluding Oil)	Sum of real exports from euro area to all BRICS countries, deflated by price indicators, excluding oil, monthly, sample: 2000–18. Sources: Haver Analytics, ECB calculations.	Sum of real exports from the United States to all BRICS countries, deflated by price indicators, excluding oil, monthly, sample: 2000–18. Sources: Haver Analytics, ECB calculations.
Real Imports from BRICS (Excluding Oil)	Sum of real imports by the euro area from all BRICS countries; see real exports to BRICS (excluding oil).	Sum of real imports by the United States from all BRICS countries; see real exports to BRICS (excluding oil).

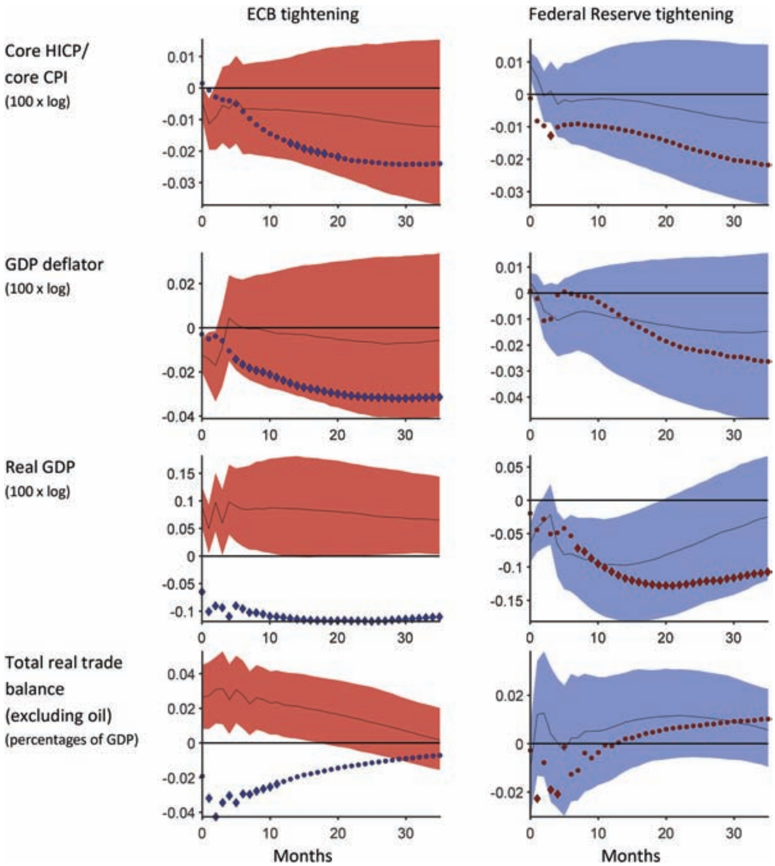
Appendix C. Additional Impulse Responses

Figure C.1. Response to an Exogenous Monetary Policy Tightening (bilateral model)



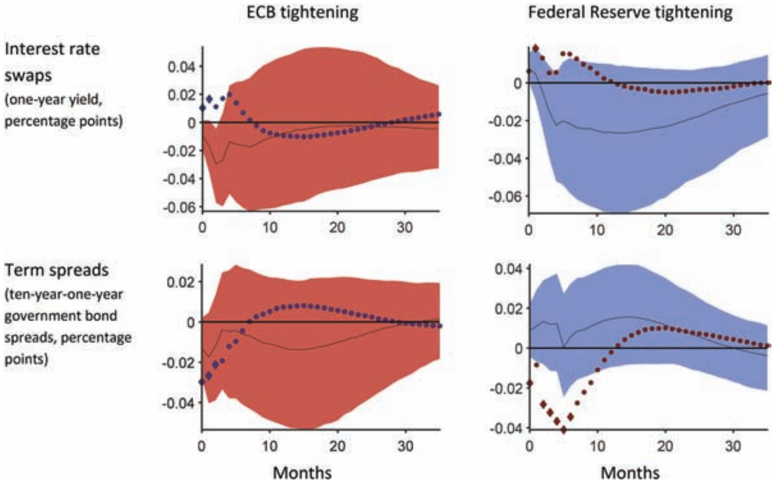
Note: The solid line shows the median impulse response surrounded by the 68 percent confidence band based on the bilateral model.

Figure C.2. Spillovers from an Exogenous Monetary Policy Tightening to Real Activity and Prices



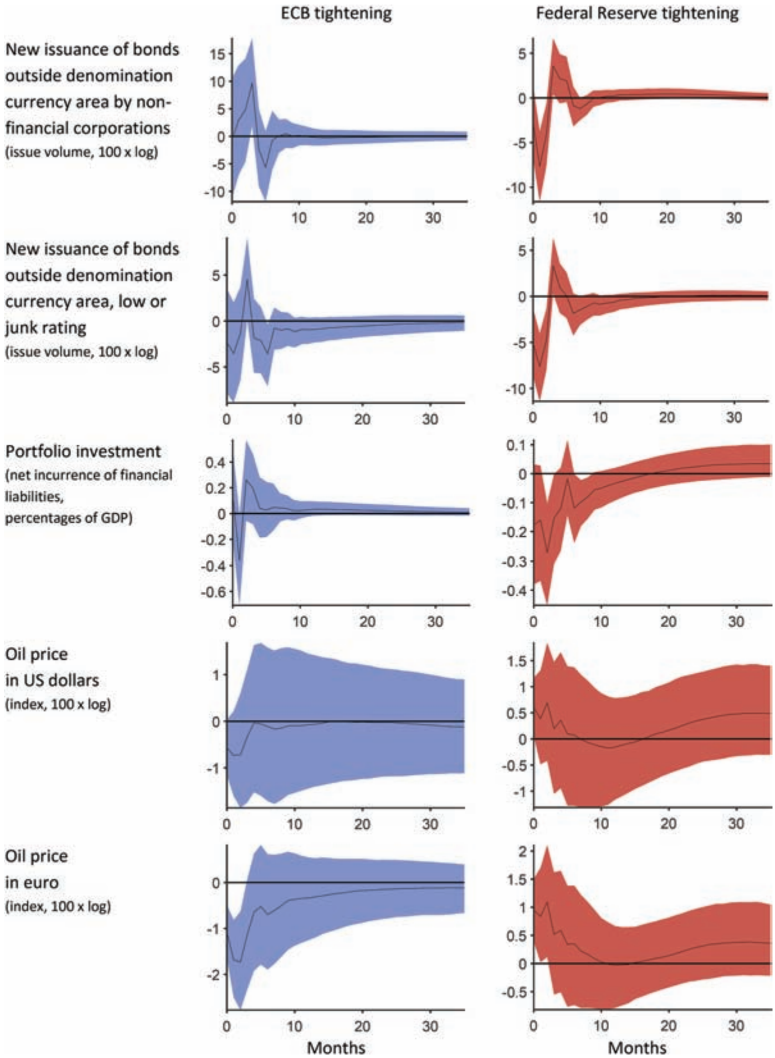
Note: The solid line plots the median spillover surrounded by the 68 percent confidence band. In the left-hand column these are the responses to an ECB tightening, in the right-hand column the responses to a Federal Reserve tightening. Quantities for the United States are plotted in red, quantities for the euro area in blue. The dotted lines plot the responses of the corresponding domestic variables, with diamonds symbolizing significance at the 68 percent level. In the left-hand column these are the responses of euro area variables, in the right-hand column the responses of U.S. variables.

Figure C.3. Spillovers from an Exogenous Monetary Policy Tightening to Financial Conditions



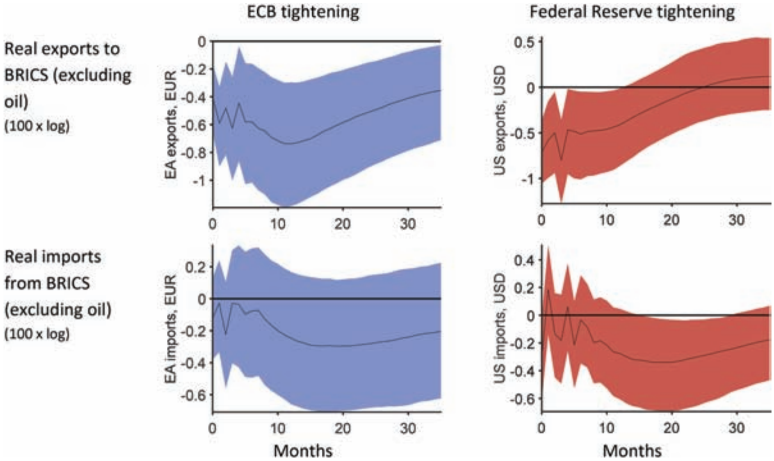
Note: The solid line plots the median spillover surrounded by the 68 percent confidence band. In the left-hand column these are the responses to an ECB tightening, in the right-hand column to a Federal Reserve tightening. Quantities for the United States are plotted in red, quantities for the euro area in blue. The dotted lines plot the responses of the corresponding domestic variables, with diamonds symbolizing significance at the 68 percent level. In the left-hand column these are the responses of euro area variables, in the right-hand column the responses of U.S. variables.The government bonds used are the bund for the euro area and the Treasury for the United States.

Figure C.4. Effects of an Exogenous Monetary Policy Tightening on Global Financial Markets



Note: The solid line shows the median impulse response surrounded by the 68 percent confidence band. In the left-hand column these are the responses to an ECB tightening, in the right-hand column to a Federal Reserve tightening. Quantities related to the euro area are plotted in blue, quantities related to the United States in red.

Figure C.5. Effects of an Exogenous Monetary Policy Tightening by the ECB or the Federal Reserve on Emerging Economies



Note: The solid line shows the median impulse response surrounded by the 68 percent confidence band. In the left-hand column these are the responses to an ECB tightening, in the right-hand column the responses to a Federal Reserve tightening. (EA = euro area)

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Central Bank Credibility and Monetary Policy*

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In this paper, a numerical measure of central bank credibility is proposed that can be incorporated into a New Keynesian model under bounded rationality. This measure arises due to the existence of the changes in private beliefs, which are different from those of the central bank. It is shown that central bank credibility matters for macroeconomic stability. Specifically, as credibility increases, macroeconomic variables vary less since private agents' expectations are more anchored. Through this channel, the model generates endogenous volatility changes. Finally, the credibility of the Federal Reserve and the European Central Bank are computed based on the proposed method.

JEL Codes: E3, E52, E58, D8.

1. Introduction

Over the past 20 years, many central banks in advanced countries have introduced inflation targeting as their way of conducting monetary policy. By introducing inflation targeting, the notion of central bank credibility became more important for policymakers. Despite the importance of credibility, there has been little research quantifying time-varying central bank credibility that is determined endogenously by the central bank's actions and economic outcomes in a

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New Keynesian framework, which has been a workhorse for central banking studies in recent decades. In stylized New Keynesian models, it is implicitly assumed that the central bank is fully credible, in the sense that private agents believe that the central bank follows predetermined rules and tries to achieve the policy target, even if the rules and/or target are unknown to the public. Therefore, the results derived from previous studies might be too optimistic and overestimate the efficiency of monetary policy.[1]

In this paper, a numerical measure of central bank credibility is proposed using a version of a New Keynesian model with imperfect knowledge. Then, implications of credibility on macroeconomic stability are analyzed. The proposed credibility measure is defined as the tendency of private agents to rely on the central bank's forecast announcements in forming expectations of future endogenous variables, and is determined by the relative performance of the central bank's forecasts compared with those of private agents. That is, we consider central bank communications, especially announcements of forecasts, as important sources for shaping credibility. As private agents cannot observe internal procedures of the central bank, such as the actual targets and policy rules, the central bank's announcements are the sole sources for measuring its credibility.

The formulation of credibility proposed here reflects the views of policymakers. For instance, Svensson interpreted a large gap between the Riksbank's announced repo rate path and market expectations as being low credibility.[2] Yellen (2006) also described credible monetary policy as a situation in which "market participants correctly anticipate the actions that the Fed will make in response to economic news and shocks." In addition, Blinder (2018) pointed out that ignoring the messages that were sent by the central bank is almost the same thing as not believing the central bank.

In this paper, we find that perpetual assessments of central bank credibility can generate endogenous changes in volatility of endogenous variables, such as inflation and the output gap. Specifically, the baseline model that is estimated based on the U.S. economy

[1]Goodfriend and King (2015) also pointed this out: "Forecasts, and policy, should not be based solely on forecasts from a model that assumes full credibility in the stated policy path." See Goodfriend and King (2015) for details.

[2]See Goodfriend and King (2015).

generates a 33 percent higher standard deviation of inflation compared with the associated rational expectations model. In addition, we also show that volatility of endogenous variables increases as credibility deteriorates. Specifically, comparing volatility of inflation between periods with the credibility level above and below the median, inflation varies 81 percent more when the credibility level is below the median. As the central bank builds a credible reputation, private expectations become more anchored around the zero steady state, which is also the target of the central bank, and this helps to stabilize the economy by preventing vicious feedback cycles produced by the self-referential effects of expectations formation.

This paper contributes to the monetary policy credibility literature. Various researchers have provided different definitions of the credibility of central banks. One of the widely used definitions is related to imperfect information and/or knowledge of private agents regarding monetary policy (Cukierman and Meltzer 1986; Ball 1994, 1995; Erceg and Levin 2003; Schaumburg and Tambalotti 2007). This stream of research describes imperfect credibility as a status wherein private agents cannot directly observe the objectives of the central bank, such as the inflation target, or as the possibility that the central bank may renege on a pre-announced policy path. Our research differs from those, as we allow endogenous changes in the level of central bank credibility evaluated by private agents, based on the central bank's announcements and economic outcomes within a single framework.

On the other hand, many previous studies in the literature have measured the credibility of a central bank and/or analyzed how credibility affects macroeconomic stability. These studies share the same concept of credibility, which is either a deviation of (long-term) inflation expectations from the central bank's target rate or a range between the realized inflation and the target rate (Johnson 1997, 1998; Bomfim and Rudebusch 2000; Demertzis, Marcellino, and Viegi 2010). Ours differs from those in some aspects. First, we explicitly specify the belief structure and build a microfounded model that is consistent with the underlying belief. Second, the credibility measure they suggest is not bounded, so it is difficult to interpret and compare. However, the credibility measure proposed in this paper is bounded to between 0 and 1, so it is easier to interpret and compare across countries and periods.

There have been studies that connect bounded rationality to credibility issues similar to ours. Gibbs and Kulish (2017) study disinflation using a structural model that is close to ours. The credibility measure therein represents the proportion of households that form expectations rationally. Our approach differs from it in some ways. First, the level of credibility is fixed in their model, making it impossible to study the interaction between credibility and macroeconomic outcomes. In addition, they do not specify explicitly the underlying belief structure that can result in the expectation formation used in their study. Hommes and Lustenhouwer (2019) examine the stability of a heterogeneous expectations model under a bounded rationality assumption. Although they incorporate an endogenous credibility measure that is similar to ours, the expectation formations in their model are too restrictive: some stick to policy targets and the others rely on past observations to forecast future variables. Hence, their credibility measure and model cannot account for the forward-looking behavior.

This paper is organized as follows. The microfoundations of the model are explained in Section 2. Belief structures and the credibility measure are proposed in Section 3. Section 4 estimates the model. Section 5 depicts historical credibility of the Federal Reserve and European Central Bank. Section 6 discusses implications of credibility on macroeconomic stability. Finally, Section 7 concludes.

2. Model

This section develops a variant of the canonical New Keynesian model that shares microfoundations with many other models. Details of the microfoundations can be found in Preston (2005). In this model, near rationality is assumed, and households hold subjective beliefs. The formation of expectations will be discussed in detail in the subsequent section.

A continuum of households, i, on the unit interval maximize their lifetime utility

$$\hat{E}_t^i \sum_{T=t}^{\infty} \psi_T \beta^{T-t} \left\{ \frac{c_T(i)^{1-\sigma}}{1-\sigma} - \chi n_T(i) \right\}, \tag{1}$$

where $\beta \in (0,1)$ is the discount factor, $\chi > 0$ measures disutility of labor, and σ is the relative risk aversion parameter, subject to the following budget constraint:

$$c_t(i) + b_t(i) \leq \frac{1 + i_{t-1}}{1 + \pi_t} b_{t-1}(i) + w_t n_t(i) + \Gamma_t(i) \tag{2}$$

and the no-Ponzi condition

$$\lim_{T\to\infty} \hat{E}_t^i \left(\Pi_{j=0}^{T-t} \frac{1 + i_{t+j}}{1 + \pi_{t+j+1}} \right)^{-1} B_t(i) \geq 0. \tag{3}$$

The variables $c_t(i)$, $n_t(i)$, $b_t(i)$, i_t, π_t, w_t, and $\Gamma_t(i)$ denote consumption, labor supply, real bond holdings, nominal interest rate, net inflation rate, real wage, and real dividends from firms. ψ_t is the exogenous preference shifter. The operator $\hat{E}_t^i$ denotes private agents' subjective expectations based on information up to time t.

A continuum of monopolistically competitive firms maximize profits,

$$\hat{E}_t^i \sum_{T=t}^{\infty} \alpha^{T-t} Q_{t,T} \left[p_t(i) y_T(i) - P_T w_T n_T(i) \right], \tag{4}$$

subject to the linear production technology, $y_T(i) = n_T(i)$, and the demand function derived from the households' problem

$$n_T(i) = y_T(i) = \left(\frac{p_t(i)}{P_T} \right)^{-\theta_T} Y_T, \tag{5}$$

where α is the Calvo (1983) parameter and denotes the probability of not being able to reset prices in subsequent periods, and $p_t(i)$ and P_t are the prices charged by firm i and the aggregate price level. In addition, $y_t(i)$ and Y_t present the output produced by firm i and the aggregate output, respectively. $\theta_t > 1$ depicts the elasticity of demand across differentiated goods and follows an exogenous process. The stochastic discount factor $Q_{t,T}$ is given as

$$Q_{t,T} = \beta^{T-t} \frac{P_t Y_t^{\sigma}}{P_T Y_T^{\sigma}}. \tag{6}$$

In a symmetric equilibrium, private agents share the same subjective beliefs; thus, aggregate subjective expectations are the same as individual expectations, although private agents are not aware of this. The log-linear approximation around the zero-inflation steady state gives the following decision rules for consumption and price streams:

$$\hat{c}_t(i) = \hat{E}_t^i \sum_{T=t}^{\infty} \beta^{T-t} \left[(1-\beta)\hat{w}_{T+1} - \frac{1}{\sigma}\left(\hat{\imath}_T - \hat{\pi}_{T+1} - \beta\left(\hat{\psi}_T - \hat{\psi}_{T+1} \right) \right) \right] \tag{7}$$

$$\hat{p}_t(i) = \hat{E}_t^i \sum_{T=t}^{\infty} (\alpha\beta)^{T-t} \left[(1-\alpha\beta)(\hat{w}_T + u_T^*) + \alpha\beta\hat{\pi}_{T+1} \right], \tag{8}$$

where the hatted variables are the log deviations from their steady states, except $\hat{p}_t(i) = \ln(p_t(i)/P_t)$, $\hat{\imath}_t = \ln[(1+i_t)/(1+\bar{\imath})]$, and $u_t^* = \ln(\theta_t/\bar{\theta})$. In the remainder, the hat notations that denote the log deviations from the steady states are dropped for simplicity, as there is no confusion that arises from this notational simplification.

We define the output gap as $x_t = y_t - y_t^n = \sigma^{-1} w_t$. That is, the output gap is the difference between the actual output and the natural output, which is the level of output in a flexible-price environment. Aggregating and imposing market clearing conditions to Equations (7) and (8) yields the following equations, which are counterparts of the dynamic IS and Phillips curve in a canonical New Keynesian model.[3]

$$x_t = \hat{E}_t \sum_{T=t}^{\infty} \beta^{T-t} \left[(1-\beta)x_{T+1} - \frac{1}{\sigma}(i_T - \pi_{T+1} - r_T^n) \right] \tag{9}$$

$$\pi_t = \hat{E}_t \sum_{T=t}^{\infty} (\alpha\beta)^{T-t} \left[\kappa x_T + (1-\alpha)\beta\pi_{T+1} + u_T \right] \tag{10}$$

[3]The decision rules are comparable with the other models with bounded rationality, for instance, Gabaix (2020), which extends a canonical New Keynesian model by incorporating the sparsity-based limited attention suggested in Gabaix (2014, 2016), as the current consumption gap is determined by the streams of future consumption gaps and real interest rates in both models. However, expectation formation differs as Gabaix (2020) assumes a specific term structure of attention allocations while we explicitly specify subjective beliefs held by private agents.

The slope of the Phillips curve is given by $\kappa = (1 - \alpha\beta)(1 - \alpha)/\alpha$. u_t is the autoregressive cost-push shock due to variations in a firm's markup reflecting fluctuations in elasticity of demand θ_t. An exogenous disturbance $r_t^n = \beta(\hat{\psi} - \hat{E}_t\hat{\psi}_{t+1})$ arises due to changes in household preferences, and it can be interpreted as fluctuations in the natural interest rate.[4]

The model is closed with a version of the Taylor rule that describes the behavior of the central bank.

$$i_t = \phi_\pi \pi_t + \phi_x x_t + m_t \tag{11}$$

The central bank reacts to inflation and the output gap. m_t is the monetary policy shock and follows AR(1) process.

3. Beliefs, Forecasts, and Credibility

In this research, we rely on a near-rationality assumption and a learning mechanism, as private agents have imperfect knowledge about how the central bank conducts its policy. In addition, as a byproduct, this assumption sidesteps the technical problem that arises in rational expectations models due to the presence of higher-order beliefs. However, this departure from rational expectations is minimized by assuming that credibility only matters in non-fundamental drift terms in the perceived law of motion (PLM) and that the other parts of the model share the same structure as the rational expectation model without the credibility measure, as shown in detail in the subsequent sections.

3.1 Belief Structures

Private agents have the prediction model expressed below:

$$\begin{aligned} z_t &= H\bar{a}_{t-1} + \Omega s_t + e_t^1 \\ \bar{a}_t &= F\bar{a}_{t-1} + \nu_t, \end{aligned} \tag{12}$$

[4]Equations (9) and (10) can be reduced to the ordinary Euler equation and dynamic Phillips curve if subjective expectations are formed rationally. This can be proven by leading the equation and applying the law of iterated expectations. See Preston (2005) for details.

where $z_t = [x_t \; \pi_t \; i_t]'$ is a vector containing endogenous variables that private agents need to predict and $s_t = [r_t^n \; u_t \; m_t]'$ denotes a vector that observable exogenous disturbances are stacked. e_t^1 is the prediction error of private agents' prediction model. $\bar{a}_t$ indicates a vector of n_a unobserved time-varying terms, possibly random-walk drifts. Following Eusepi, Giannoni, and Preston (2015), it is assumed that the number of the underlying driving forces of drift terms, n_a, is equal to two and denote a^π and a^x, respectively.[5] These terms are labeled as the nominal and real factor. The nominal factor reflects uncertainty about the inflation target of the central bank, and the real factor represents fundamental uncertainty about long-term technological advance.[6] This specification is supported by empirical analysis in Eusepi, Giannoni, and Preston (2015). It is also intuitively plausible based on the Fisher equation. Movements in long-term expectations of the interest rate cannot be decoupled from those of inflation provided that the real interest rate does not deviate substantially from its steady state. Under this assumption, the drift term attached to the interest rate prediction equation can be expressed as $\lambda_1 a^x + \lambda_2 a^\pi$. Following Eusepi, Giannoni, and Preston (2015), we assume $\lambda_2 = 1$ and interpret λ_1 as the relative contribution of the real factor to the nominal interest rate.

Exogenous disturbances follow the stationary AR(1) process

$$s_t = \Phi s_{t-1} + \varepsilon_t, \tag{13}$$

where ε_t are i.i.d. shocks. It is also assumed that parameters that govern reactions to fundamental disturbances, Ω, are known to private agents and coincide with their rational expectation counterparts.[7] This assumption simplifies analysis and helps place greater focus on the dynamics of long-term expectations, whereas it minimizes deviation from the rational expectations model. Hence, private agents only need to learn about the unobserved drift terms.

[5] a denotes the estimate of $\bar{a}$, which is unobservable to private agents.

[6] The real factor can be also interpreted as a shock to the higher-order belief, or to private agent sentiments, à la Angeletos and La'O (2013). For example, persistent positive waves of the real factor can be considered as strong optimism for the real activities in the economy, as perceived by private agents.

[7] This assumption does not change the qualitative results of this paper, since learning procedures for constants and coefficients attached to structural disturbances are totally separated.

Based on the estimated time-varying drift terms a_{t-1} up to time $t-1$ and the beliefs given above, private agents compute their own forecasts as

$$\hat{E}_t^P[z_T] = HF^{T-t}a_{t-1} + \Omega\Phi^{T-t}s_t, \tag{14}$$

where $\hat{E}_t^P$ denotes the private forecast based on the information up to time t. If time-varying drift terms $\bar{a}$ are more persistent than structural disturbances s_t, long-term forecasts are driven more by time-varying drift terms than by disturbances. For example, if $\bar{a}$ have unit roots, then infinite-horizon long-term forecasts are simply given by the current time-varying drift terms.

$$\lim_{T\to\infty} \hat{E}_t^P[z_T] = Ha_{t-1} \tag{15}$$

Following Eusepi, Giannoni, and Preston (2015), a_t is updated using the steady-state Kalman filter recursion

$$a_t = Fa_{t-1} + K(z_t - Ha_{t-1} - \Omega\Phi s_{t-1}), \tag{16}$$

where the time-invariant Kalman gain matrix is

$$K = PH'(HPH' + R)^{-1}. \tag{17}$$

P is given as $E[(\bar{a}_t - a_t)(\bar{a}_t - a_t)']$ and R denotes private agents' prior beliefs about the covariance matrix of the observation error terms e_t^1 in their prediction equation (12).[8]

The central bank is also near rational and has its own PLM, which is used to predict future endogenous variables. The PLM of the central bank is given as follows:

$$z_t = \Omega s_t + e_t^2, \tag{18}$$

which coincides with the rational expectations solution to the model. Here, e_t^2 is the prediction error of the central bank's prediction model. Compared with that of private agents, the PLM of the

[8]Since the Kalman gain is determined by the private agents' prior beliefs and is time invariant, the steady-state Kalman filter can be understood as a version of the constant-gain learning process, which is widely used in the learning literature—for instance, Eusepi and Preston (2011).

central bank clearly shows that the central bank's forecasts are well anchored, as there are no time-varying terms $\bar{a}$. This formulation seems reasonable, as many central banks predict their economy by using country-specific dynamic stochastic general equilibrium (DSGE) models with rational expectations. In addition, Mokhtarzadeh and Petersen (2021) show that central banks interested in maintaining inflation stability should communicate their predictions solely based on rational expectations in their laboratory experiment. On the other side, the actual law of motion nests that of the rational expectations under this assumption, as shown later. This characterization sets a natural comparison of the model, thereby facilitating interpretation of the results.

The central bank announces its own forecasts of endogenous variables after observing the realization of disturbances

$$\hat{E}_t^{CB}[z_T] = \Omega\Phi^{T-t}s_t, \tag{19}$$

where $\hat{E}_t^{CB}$ denotes the forecast of the central bank based on information up to time t.

The belief structure can be justified under the assumption that no commitment device exists. This is a reasonable assumption, as there is no credible commitment device. It is neither allowed nor possible to inform the public credibly of the central bank's internal decisionmaking procedure, including objectives and rules. A notion of credibility arises due to this point. Even if a central bank deliberately conducts monetary policy and follows its target, private expectations can diverge from the target because the central bank cannot commit to the target and communicate its policy credibly. Private agents continuously evaluate the central bank's resolutions and abilities to achieve the target based on the history of outcomes. The formal definition of credibility used in this paper is presented below.

3.2 Credibility Measure

The tendency of private agents to rely on the central bank's announced forecasts in forming expectations of future endogenous variables defines the credibility.

Definition 1 (Credibility). *Central bank credibility ξ_t is the time-varying relative weight attached to the announced forecasts* *made by the central bank, other than private agents' own forecasts, when private agents derive ensemble forecasts and form expectations. Specifically, private agents use the following subjective expectations:*

$$\hat{E}_t z_T = \xi_{t-1} \hat{E}_t^{CB}[z_T] + (1 - \xi_{t-1}) \hat{E}_t^{P}[z_T]. \tag{20}$$

As the proposed measure determines the relative weight among forecasts, it can be determined naturally based on the relative accuracy of two forecasts. To be precise, the credibility measure, ξ_t, is determined by the following dynamic predictor selection problem, modified from that of Brock and Hommes (1997) and a possibly sluggish law of motion.

$$\tilde{\xi}_t = 1 - \left(\frac{\exp\left(\delta_1 U_t^P\right)}{\exp\left(\delta_1 U_t^P\right) + \exp\left(\delta_1 U_t^{CB}\right)} \times \mathcal{D}_t \right), \tag{21}$$

where

$$U_t^k = -\sum_{j=0}^{\infty} \omega^j \left[\left(z_{t-j} - \hat{E}_{t-j-1}^k [z_{t-j}] \right)' W \left(z_{t-j} - \hat{E}_{t-j-1}^k [z_{t-j}] \right) \right] \tag{22}$$

$$\mathcal{D}_t = 1 - \exp\left(-\delta_2 \left(\hat{E}_{t-1}^{CB}[z_t] - \hat{E}_{t-1}^{P}[z_t] \right)' \left(\hat{E}_{t-1}^{CB}[z_t] - \hat{E}_{t-1}^{P}[z_t] \right) \right) \tag{23}$$

for $k \in \{CB, P\}$ and

$$\xi_t = \xi_{t-1} + \eta(\tilde{\xi}_t - \xi_{t-1}). \tag{24}$$

U^k is the fitness measure of central bank ($k = CB$) or private ($k = P$) prediction. The fitness measure is a discounted weighted sum of the negative of the past squared prediction errors. $\omega \in [0, 1]$ is the memory parameter that controls the degree of discounting of past prediction errors in measuring the fitness. The fitness measure becomes more persistent when ω approaches one. $\mathcal{D} \in [0, 1]$ measures the distance between two predictors and approaches zero as two predictors become similar. W is a weighting matrix and δ_1 and δ_2 are the intensity of choice parameters. $\eta \in [0, 1]$ controls inertia of the credibility measure. As it approaches one, inertia does not exist and the law of motion becomes $\xi_t = \tilde{\xi}_t$. This sluggish process allows slow adjustments, rather than drastic changes, in credibility.

The credibility measure is determined by two parts: the distance between private and central bank forecasts ($\mathcal{D}_t$), and the relative accuracy of these forecasts (U^{CB}, U^P). Firstly, if the central bank's forecasts are close to those of private agents ($\hat{E}^{CB}_{t-1}[z_t] - \hat{E}^{P}_{t-1}[z_t] \rightarrow 0, \mathcal{D}_t \rightarrow 0$), there is no reason for private agents to disregard the forecasts announced by the central bank, because they believe that the central bank shares their views on the economy. In that case, private agents will perceive the central bank to be credible and will make use of it ($\tilde{\xi}_t \rightarrow 1$).[9] At the same time, if two predictors differ ($\mathcal{D}_t \rightarrow 1$) and private forecasts are more accurate than those of the central bank ($U^P > U^{CB}$), private agents will doubt the credibility of the announced policy targets and forecasts of the central bank, and will place more reliance on their own forecasts ($\exp\left(\delta_1 U^P_t\right) / \exp\left(\delta_1 U^P_t\right) + \exp\left(\delta_1 U^{CB}_t\right) \rightarrow 1, \tilde{\xi}_t \rightarrow 0$). Therefore, poor central bank forecasts decrease a central bank's credibility.

Kocherlakota (2011) stated the same point:

> I've been emphasizing the importance of communication and communication matters greatly. But, ultimately, the public's beliefs about the FOMC's inflation objective will also depend on inflationary outcomes. If annual inflation averages less than 1.5 percent for more than three or four years, onlookers will begin to suspect that the FOMC's true objective for inflation is lower than its declared "two percent or a bit under." Correspondingly, if inflation is persistently higher than 2 percent, then the public will begin to believe that the FOMC's true objective for inflation is higher than 2 percent. In either case, inflation expectations could become unmoored, and the FOMC could lose control of inflation itself. Communication can only be effective if the FOMC also retains credibility.

[9] Some policymakers acknowledged this point in their speeches. For example, Bernanke (2004) argued that "The . . . way in which clear and open communication enhances the effectiveness of monetary policy. . . is by helping to align financial-market participants' expectations about the future course of monetary policy more closely with the policy committee's own plans and projections."

This is also in line with Plosser (2010)'s prescription of achieving credibility.

One advantage of this formulation is that the aggregate economic outcomes do not necessarily influence credibility directly. That is, higher inflation realizations than the announced target do not necessarily decrease credibility. If the economy is hit by a severe exogenous shock, aggregate variables can deviate from the pre-announced policy targets. Since private agents understand this possibility and make the same mistakes in predicting macro variables, they take a lenient stance in relation to these deviations. Put differently, this case translates into larger losses in both U^P and U^{CB}, and credibility ξ_t does not change much.[10]

Although this formulation seems similar to those of previous studies, such as Bomfim and Rudebusch (2000) and Hommes and Lustenhouwer (2019), it differs from them in several ways. First, our formulation does not employ past realizations as predictors to make expectations behave adaptively. In aforementioned research, the credibility measure is simply the degree of forward-lookingness. On the contrary, both predictors used in this paper are forward looking. Second, it is more plausible to account for the central bank's real-time communications. In Hommes and Lustenhouwer (2019), the credibility measure represents the proportion of households that use policy targets as their future forecasts. However, as the interim

[10] One might raise the question as to whether central bank credibility can be independent of the central bank's ability of keeping inflation stable if the central bank's staff become very good at predicting inflation and gross domestic product (GDP), to a point where the private sector's expectations totally rely on the central bank's forecasts. In this model, central bank credibility is not independent of the central bank's ability to keep inflation stable in the longer-run perspective. Definitely, credibility rises if the central bank's predictions are good enough. However, it is necessary to understand what makes the central bank's predictions better. In general, a central bank's predictions become more accurate when non-fundamental beliefs (a_t) that affect the private predictions are suppressed. These non-fundamental beliefs emerge when inflation and the output gap diverge from the pre-announced targets continuously beyond the private agents' tolerance. While temporary deviations due to large shocks are tolerated, continuing deviations may trigger divergent beliefs and drops in credibility. Therefore, deliberate policymaking for stabilizing inflation and the output gap is required to achieve high credibility and to make private agents rely more on the central bank's predictions. For this reason, the proposed measure is related to a central bank's ability to stabilize the economy in the longer-run horizon.

targets may not coincide with pre-announced policy targets, it is plausible that households would just use the central bank announcements in predictions, as in this paper, rather than rely on the long-run policy targets.

The proposed measure is also close to the one used by Gibbs and Kulish (2017). In their paper, two forecasts formed by rational expectations and adaptive learning are used to derive the ensemble forecast, and credibility is described as the proportion of households that have the rational expectations. Their measure is closely related to ours, as past realizations matter as well through the learning process but the weight placed on adaptive learning is not changing over time endogenously.

Compared with the credibility measures used in the previous research, the proposed measure possesses several advantages. First, our credibility measure is ex ante bounded in the unit interval, making it easier to interpret and compare. Second, what matters in determining credibility is the relative accuracy of forecasts, not the distance between either the actual or the predicted inflation and pre-announced inflation target. Hence, the deviation of inflation from the target, which arises due to large shocks, does not necessarily damage credibility because the public understands that it is inevitable and that their own predictors also perform badly.

3.3 Equilibrium

Combining expectations formation equations (14), (19), and (20) with policy rule equations (9), (10), and (11), the actual laws of motion (ALM) for this economy can be obtained.[11]

$$z_t = C_0 a_{t-1} + \Omega s_t \tag{25}$$

C_0 is the factor loading matrix, and this changes endogenously, as the credibility measure ξ_{t-1} is contained in this. The ALM and Equations (21), (22), and (24) governing the dynamics of the credibility measure fully characterize the model.

One advantageous feature of the model is that we can recover the ordinary rational expectations model if central bank credibility

[11]See Appendix A for details.

is fixed at the maximum level in any of the cases. That is, when the central bank credibility measure ξ_t is exogenously imposed to be one, the ALM become identical to those in the rational expectations model. Hence, the ordinary rational expectations model can be considered as a special case, in which the central bank is fully credible under any circumstances. The ALM, however, diverge away from that of the rational expectations model, and the effect of the subjective beliefs magnifies as the credibility measure moves away from one.

4. Historical Credibility of Central Banks

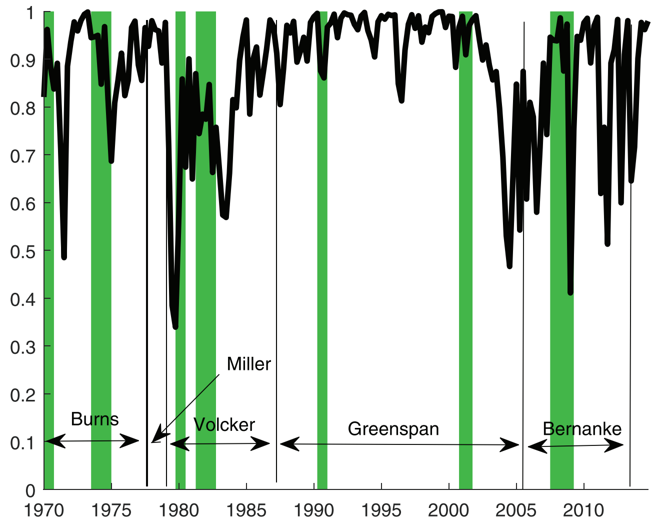
The credibility measure suggested above can be readily computed based on the observed data using Equations (20)–(24) if data exist about the economic forecasts of private agents and the central bank. To understand the credibility measure better, we construct the historical credibility series of the Federal Reserve (the Fed) and the European Central Bank (ECB). In this exercise, we only use the one-quarter-ahead inflation forecasts for data compatibility.[12] For the Fed, we compute credibility for 1968:Q4 to 2014:Q4 using Greenbook and SPF forecasts.[13] Similarly, credibility of the ECB is computed by the European Survey of Professional Forecasters and the ECB's own announcements on economic projections.

Figure 1 shows the historical series of the Federal Reserve's credibility computed by the SPF and Greenbook forecasts and the names of the Fed's chairperson. It shows that the Fed's credibility has fluctuated mostly between 0.3 and 1, but retains relatively high levels. While there have been substantial drops in credibility in 1972, from 1979 to 1980, in the mid-2000s, and during the global financial crisis, mean credibility is higher than 0.85 for the entire period. The shaded

[12]The choice of values of ω, δ_1, and δ_2 are discussed in Section 5. In the estimation, it is shown that considering only inflation forecast errors results in the similar outcomes compared with the case that takes all three variables—output gap, inflation, and interest rate—into account.

[13]There are multiple numbers of Greenbook forecasts in a given quarter. In this procedure, we choose the first Greenbook forecast in the quarter to comply with the assumption that forecasts are announced at the beginning of the period in the model. Lastly, the median forecast from the SPF is used for private forecasts.

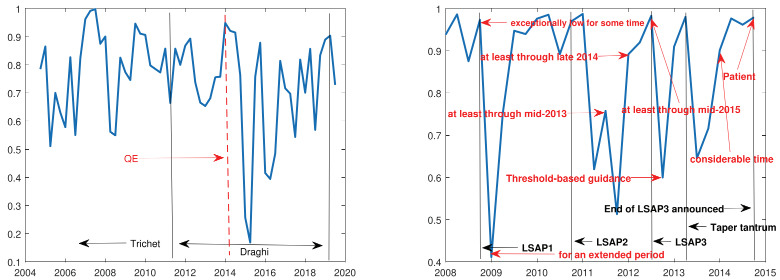
Figure 1. Credibility Measure of the Federal Reserve, 1970:Q1–2014:Q4



areas represent recessions identified by the National Bureau of Economic Research (NBER). This figure shows that shifts in credibility do not show any clear relationship with the business cycle.

Although it is hard to identify one specific factor that determines credibility, since that is determined collectively by various macroeconomic outcomes, the computed credibility of the Fed hints that major changes in the ways of conducting monetary policy are particularly important. In the first few years of the 1970s, frequent modifications in monetary policymaking took place as the techniques required for setting and pursuing money targets were developed. The figure also shows that the Fed's credibility was at a low level at the beginning of Volcker's tenure, but the Fed built credibility slowly over the 1980s. The reason behind this shift might be attributed to a dramatic change in policy course, which attempted to tame persistently high inflation. While credibility remained relatively high until the early 2000s, it shows a persistent drop in the mid-2000s. The Fed's credibility shifts rapidly during the global financial crisis period. Since the Fed introduced new policies sequentially during this period, that might drag credibility downward, as will be

Figure 2. Credibility of the European Central Bank, 2005:Q1–2019:Q3 (left) and Credibility of the Fed during the Global Financial Crisis (right)



presented in details below. Experiencing drops in credibility after introducing new policies seems reasonable. As private beliefs are persistent, realizing the expected effects takes time. Hence, credibility may decrease initially. As the expected effects emerge, the central bank starts to gain credibility.

Lastly, we provide two examples to which our measure of credibility may apply. To do so, we expand the analysis to credibility of the ECB and to the impacts of unconventional monetary policies conducted during the global financial crisis in the United States.

The left panel in Figure 2 depicts the credibility of the ECB. One notable feature is that the ECB's credibility has been persistently lower than that of the Fed. Specifically, the average credibility of the ECB is more than 15 percent lower than that of the Fed. It provides an example of the way in which the proposed measures can be used to compare credibility across economies, and calls for deeper future research on this matter.

The right panel provides the Fed's credibility and major policy events during the global financial crisis. The black and red arrows indicate policy changes regarding quantitative easing and forward guidance, respectively. It is difficult to identify the factors that drive credibility, but we may find some stylized facts regarding the influence of monetary policy on credibility. For instance, as is postulated above, major policy shifts precede drops in credibility. To go into more depth, the figure shows that the Fed's credibility tends to decrease following the announcements regarding quantitative easing.

Interestingly, the impacts of forward guidance seem unclear. This finding is also valid in the case of the ECB. The announcement that the ECB began purchasing government bonds (quantitative easing, or QE) preceded the largest drop in credibility.

If private beliefs are persistent and influential concerning economic outcomes, we can provide an explanation for the above findings. First, policies that only affect private beliefs or expectations, such as forward guidance, may have small impacts in credibility. If credibility is low, forward guidance may not be able to exert meaningful influence, hence it would not be able to enhance credibility. On the other hand, if credibility is high, private agents will change their beliefs immediately following the projections announced by the central bank. Therefore, credibility would not show drastic changes in this case either. Second, it will take a long time to show intended outcomes after major policy changes if economic outcomes are substantially affected by private beliefs, which evolve relatively slowly. This may temporarily damage credibility, as the policy effects come with a delay. We are going to show that these requirements seem to be satisfied in the estimated model.

5. Estimation

As this paper focuses on quantitatively measuring the influence of central bank credibility on the economy, it is necessary to estimate the model to discipline the parameters. As the model is highly nonlinear, we estimate the model by implementing a bootstrap particle filter Metropolis-Hastings algorithm to evaluate the likelihood functions and to derive posterior distributions of parameters.[14]

5.1 Data

We include six observables to estimate the model. For inflation and short-term interest rate measures, the GDP deflator and the federal funds rate (FFR) are used. For the output gap measure, the output gap published by the Congressional Budget Office is used. We also include a private forecasts series on the level of GDP, GDP deflator, and FFR from the Federal Reserve Bank of Philadelphia's Survey

[14]See Herbst and Schorfheide (2015) for theoretical discussion on this method.

of Professional Forecasters (SPF) to match the model-generated private forecasts. The sample period covers 1981:Q3 to 2007:Q4. Since the model lacks the zero lower bound constraint, we only use data from before the financial crisis.[15]

5.2 Calibrated Parameters

We calibrate some parameters before estimating the model. This set of parameters includes the subjective time discount factor, β; two intensity of choice parameters, δ_1 and δ_2; weights on past prediction errors, ω; and the weighting matrix, W. The subjective time discount factor β is set to 0.995. δ_1 and δ_2 are chosen to minimize the distance between the credibility measure estimated in the model and the one calculated directly from the private forecasts and the Fed's forecasts in Figure 1. As a result, $\delta_1 = 1.62$ and $\delta_2 = 23$ are chosen.[16] The parameter that governs weights on past prediction errors, ω, is set to zero for tractability. While there is no memory regarding the past loss due to prediction errors, the persistence of credibility can be still captured by the sluggish evolution of credibility measure which is controlled by η. The weighting matrix W is set as the identity matrix. This means that private agents put the same weight on the prediction errors on the output gap, inflation, and interest rate.[17]

5.3 Method

We estimate the model by constructing a Metropolis-Hastings particle filter (MHPF). The proposal parameters are drawn from a Markov chain repeatedly for 60,000 iterations. We discard the first 10,000 draws to remove any influence from the initial condition.

[15] Extending the sample period beyond 2007 using the shadow rate proposed by Wu and Xia (2016) does not change the estimates substantially.

[16] As long as δ_1 and δ_2 are sufficiently large, the predictor selection problem mimics the classical choice behavior. That is, the agents always choose the predictor that is more accurate, and the weight put on the other predictor becomes negligible. Hence, the quantitative results do not change considerably even if these parameters are changed. See Appendix D.

[17] The case where private agents only consider prediction errors on inflation results in similar outcomes. Specifically, the estimated credibility and private beliefs are almost identical to the one from the baseline case. See Appendix D.

Table 1. Estimated Parameters

Parameters	Distribution	Prior Mean	S.D.	Posterior Mode	[0.05, 0.95]
α	Beta	0.7	0.08	0.73	[0.72,0.73]
σ	Gamma	2	0.5	1.77	[1.76,1.78]
ϕ_x	Gamma	0.12	0.1	0.17	[0.17,0.17]
ϕ_π	Gamma	2	0.15	2.12	[2.12,2.13]
ρ_r	Beta	0.7	0.12	0.73	[0.72,0.73]
ρ_u	Beta	0.7	0.12	0.73	[0.73,0.74]
ρ_m	Beta	0.7	0.12	0.79	[0.78,0.79]
λ_1	Normal	0	0.5	–0.25	[–0.28,–0.25]
f_1	Uniform	0	1	0.74	[0.73,0.75]
f_2	Uniform	0	1	0.99	[0.99,0.99]
σ_{ε^r}	IGamma	0.1	2	0.58	[0.58,0.58]
σ_{ε^u}	IGamma	0.1	2	0.15	[0.14,0.16]
σ_{ε^m}	IGamma	0.1	2	0.50	[0.49,0.51]
K_{11}	Uniform	–1	1	0.34	[0.34,0.36]
K_{12}	Uniform	–1	1	0.12	[0.12,0.12]
K_{13}	Uniform	–1	1	–0.28	[–0.30,–0.28]
K_{21}	Uniform	–1	1	–0.09	[–0.09,–0.08]
K_{22}	Uniform	–1	1	0.30	[0.29,0.31]
K_{23}	Uniform	–1	1	0.14	[0.13,0.15]
η	Beta	0.7	0.12	0.66	[0.64,0.66]
Marginal Log-Likelihood	–1002.9				

Note: The priors and posteriors, Para(1) and Para(2), correspond to the mean and standard deviation of the Normal, Gamma, Inverse Gamma, and Beta distributions and to the lower and upper bounds for the Uniform distribution.

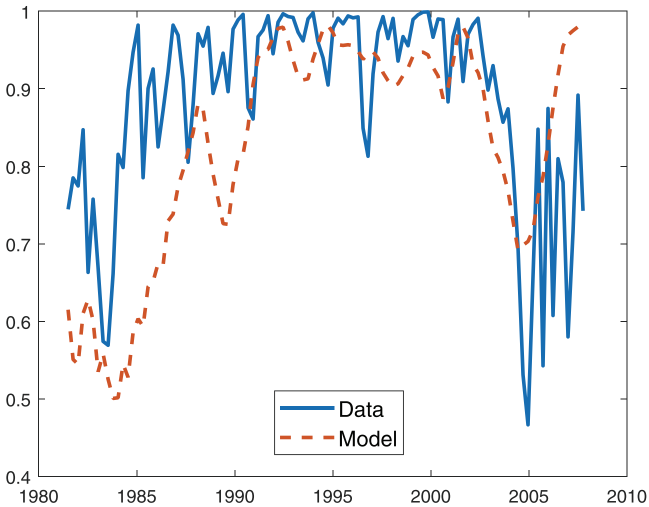
Next, we collect one draw in every five draws to thin out the chain to reduce the autocorrelation of the chain. Then we construct marginal posterior densities from the remaining 10,000 draws. The scaling parameter is set to vary along the iteration so that it guarantees achieving the acceptance rate in between 0.2 and 0.4. The resulting total acceptance rate is around 0.28. Lastly, the number of particles is set to 10,000.[18]

5.4 Estimation Outputs

Table 1 shows the prior and posterior distributions for the model parameters. We choose diffuse priors for f_1, f_2, K, η, and λ_1, as

[18]See Appendix B for details regarding the measurement and transition equations and iteration process.

Figure 3. Comparison between Estimated and Directly Computed Credibility



they are model specific, and so there is no consensus on the choice of these prior distributions. Other priors are quite general in the literature. In the next section, we use the modes of posteriors when simulating the model.

Some results are worth mentioning. The results for f_1 and f_2 show that the long-term inflation forecasts are highly persistent, and not anchored, but that the perception regarding the long-term business outlook is less persistent. In addition, the estimate of η shows that the credibility measure is moderately persistent.

Next, we compare the estimated credibility and the directly computed credibility derived in the previous section to validate the estimation. Figure 3 presents the credibility series directly computed by the forecast data (blue solid line) and the estimated credibility the non-linear filter (red dashed line). While the estimated credibility shows more sluggish movement, the two series move in a similar manner. Numerically, the correlation coefficient between the two series is 0.48. The correlation reaches 0.57 when the empirical credibility series is smoothed by a moving average with a five-quarter window.

Finally, before analyzing the model, it is necessary to check whether the model results in stable dynamics under the given parameters. Appendix C analyzes the model stability and confirms that the model is stable under the current calibration.

6. Quantitative Results

In this section, we examine the role of central bank credibility on the overall macroeconomic stability by simulating the model. To this end, the model is simulated 50,000 times, for 1,200 periods each, to gauge the effects of endogenously evolving central bank credibility on macroeconomic stability. The initial 1,000 periods are discarded to remove any effects from initial conditions. The model parameters are set to their posterior modes derived from the estimation procedure.

6.1 Endogenous Volatility Changes

First, we compare the macroeconomic volatility obtained in the baseline model to that computed in the rational expectations model to examine whether endogenous central bank credibility can create additional volatility in macroeconomic variables.[19] To this end, standard deviations in the output gap and inflation are computed and compared with those from the rational expectations model.

The results are summarized in Table 2. This table contains the mean standard deviations obtained from the simulations across models and subperiods. In general, introducing the credibility measure in a New Keynesian model raises the volatility of both the output gap and inflation. Specifically, the standard deviation of inflation is 33 percent higher in our model, whereas that of the output gap changes by only 7 percent. We derive two observations from the above results. First, introducing central bank credibility can endogenously generate additional volatility in the economy. Second, shifts in central bank credibility can substantially affect the volatility of inflation, but that of the output gap is affected much less.

Next, the standard deviations of the macro variables are calculated for different levels of credibility to analyze the relationship

[19]The rational expectations model is evaluated at the same parameter values with the exception of the belief structure.

Table 2. Volatility of Macro Variables

	RE	Baseline	$\frac{\text{Baseline}}{\text{RE}}$
$\sigma(x)$	1.37	1.47	1.07
$\sigma(x_h)$	1.36	1.37	1.01
$\sigma(x_l)$	1.37	1.54	1.12
$\sigma(x_l)/\sigma(x_h)$	1.01	1.12	
$\sigma(\pi)$	0.61	0.81	1.33
$\sigma(\pi_h)$	0.54	0.54	1.00
$\sigma(\pi_l)$	0.66	0.98	1.48
$\sigma(\pi_l)/\sigma(\pi_h)$	1.22	1.81	

Note: The average standard deviations are obtained from the simulations. Subscripts h indicate standard deviations computed from periods with credibility higher than the median and l with lower than the median.

between credibility and macroeconomic volatility. To this end, simulated series are divided into two groups of subperiods. One contains variables at periods when credibility is higher than the median, and the other when credibility is lower than the median.[20] For better contrast of the results, we also report standard deviations derived by simulating the same sets of shocks under the rational expectations model since there might be the possibility that higher volatility is caused by larger shocks rather than the endogenous credibility channel.[21]

Table 2 clearly shows that the variations of endogenous variables decrease in credibility. The standard deviations of the output gap and inflation are 12 percent and 81 percent higher in the low-credibility periods. It is appropriate to ask whether this difference is fully caused endogenously by introducing credibility concerns in the model, because some portions of the increments can be caused just by selecting low-volatility realization periods as high-credibility periods. This concern can be resolved by comparing the $\sigma(x_l)/\sigma(x_h)$ and $\sigma(\pi_l)/\sigma(\pi_h)$ ratios between the baseline and rational expectations

[20] The median is around 0.63.

[21] Since there is no change in credibility in the rational expectations (RE) model, we report $\sigma(x_h)$, $\sigma(\pi_h)$, $\sigma(x_h)$, and $\sigma(\pi_h)$ of the RE model based on the credibility level of the baseline model calculated from the same shocks.

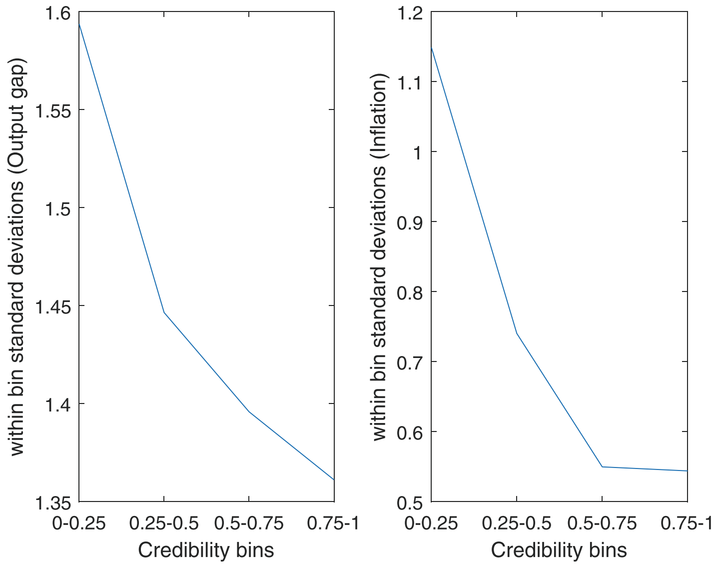
models. As shown in Table 2, accounting for credibility of the central bank significantly increases the ratio for inflation. The endogenous evolution of credibility generates an additional 59 percent increment in the standard deviation of inflation (81 percent minus 22 percent) while that of the output gap shows 11 percent rise in volatility compared with the rational expectations model. Thus, we can conclude that an increase in the standard deviation of inflation and the output gap is largely generated by introducing endogenous credibility into the model. One thing worth noticing is that volatility is almost the same across the models when credibility is in the upper half. This is reasonable, as our model converges on the rational expectations model as credibility enhances.[22]

Next, it is still not clear whether there is a monotonic negative relationship between the credibility and volatility of endogenous variables. To answer this question, we divided the simulated series into finer credibility bins, as in Figure 4. It clearly shows that the standard deviations of the output gap and inflation increase monotonically as credibility decreases. In particular, the variability of inflation increases exponentially as credibility decreases. To be precise, the increments in the volatility of inflation rise from 2 percent to 55 percent as credibility decreases, while the increments in the output gap increase only from 3 percent to 10 percent. This suggests that most of the benefits of higher credibility accrue in the low credibility region. This result is in line with the results discussed in Schaumburg and Tambalotti (2007).

The quantitative results can be connected to the recent experimental studies on central bank credibility (Ahrens, Lustenhouwer, and Tettamanzi 2017; Mokhtarzadeh and Petersen 2021). In these studies, central bank credibility is measured by the fraction of forecasters that have the same (or close) projected value as the one announced by the central bank. As predicted in this paper, it is

[22]It is well known that introducing a learning procedure instead of rational expectations assumption produces an additional layer of interaction between economic outcomes and monetary policy and results in more volatile macroeconomic dynamics compared with rational expectations models as shown above. For instance, Orphanides and Williams (2004) find that policies that fail to maintain control over inflation are vulnerable to episodes in which the public's expectations become decoupled from the policy objectives under imperfect knowledge environment using a perpetual learning model.

Figure 4. Standard Deviations of Output Gap and Inflation across Credibility Quartiles



shown that credibility decreases when the central bank makes larger prediction errors in these experiments. The experimental results can explain the theoretical results provided in this paper reasonably well. First, when credibility is higher, volatility is almost identical to that which arises in the rational expectations model. Second, when credibility is lower, the standard deviations (or similarly, mean square deviations from the targets) of the output and inflation increase. Specially, volatility of the output increases by quantitatively small amounts whereas that of inflation increases substantially. These results are surprisingly in line with the theoretical predictions provided in this model.

Combining the above results with the estimated credibility of the Fed, we can postulate that shifts in the Fed's credibility is one possible explanation for the Great Moderation, referring to the period of low macroeconomic variability between the mid-1980s and the global financial crisis. Since the Fed's credibility during the Great Moderation was higher than during other periods (Table 3), additional volatility injected because of the lower credibility had

Table 3. Average Credibility of the Fed across the Periods

Period	1970–1984	1985–2007	2008–2014
Mean Credibility	0.83	0.91	0.86

been suppressed. This might have contributed to relatively stable macroeconomic developments. For this reason, lower macroeconomic volatility during the Great Moderation might be partially attributed to an increase in the Fed's credibility.

6.2 Underlying Mechanism: Feedback Effect of Private Beliefs

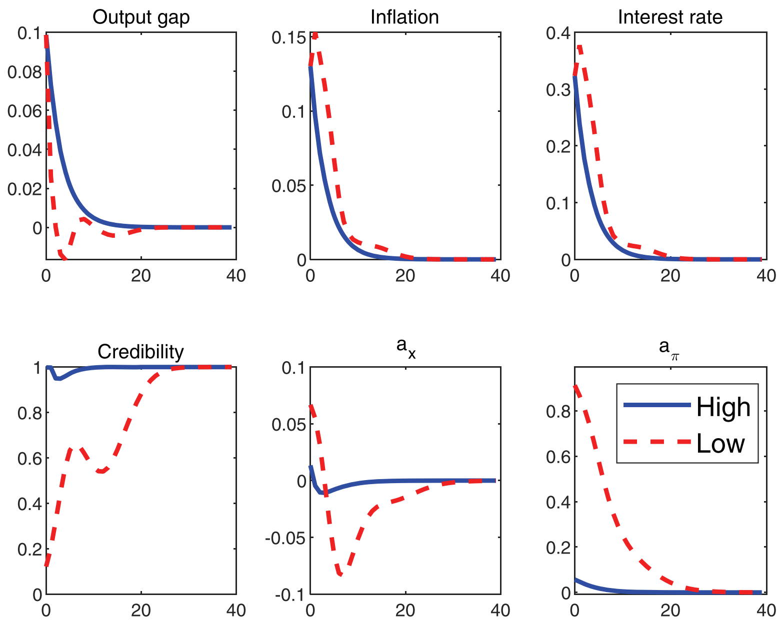
In this subsection, we analyze the underlying mechanism that creates additional volatility in the economy by examining the impulse response of the model. Before proceeding with the results, it is necessary to emphasize that the impulse response is not unique, and depends on a realization of history due to the non-linearity of the model. In particular, the level of credibility and the real and nominal factors that occur when a shock hits the economy are important determinants of the impulse response. For this reason, we derive impulse response functions with different initial conditions for two factors, a^x and a^π, and for credibility ξ.[23] Since two initial values for subjective factors are required to generate impulse response functions, we assume $a_{-1} = a_{-2}$ for simplicity. Finally, it is also noteworthy that impulse response functions obtained in case $a = 0$ and $\xi = 1$ is almost identical to those in the rational expectations model.

Impulse responses are calculated by differencing simulated series with and without a specific temporary shock. Finally, we present impulse responses to the preference shock in the main text, and those to the cost-push and monetary policy shock are delegated to Appendix E, as the underlying mechanism is the same.

In Figure 5, it is assumed that the private agents perceive that the long-term inflation is higher than the zero inflation

[23] Alternatively, we may derive generalized impulse responses as in much of the empirical literature. However, we do not follow this strategy because our method helps to understand better the transmission mechanisms of the model.

Figure 5. Impulse Responses to a One-Standard-Deviation Preference Shock When a_π Is Positive



Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^{\pi}_{-1} = 0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^{\pi}_{-1} = 0.9$.

target ($a^{\pi}_{-1} > 0$) and the exact values differ across initial credibility. To be exact, a^{π}_{-1} is assumed to be 0.01 and 0.9 for the initial credibility level 1 (high credibility), and 0.4 (low credibility) cases.[24]

When $\xi_{-1} = 1$, subjective factors a^x and a^π do not affect the output gap and inflation, even if these factors have non-zero values. Therefore, their effects on endogenous variables are limited in subsequent periods. For this reason, impulse responses are similar to those of the rational expectations model when initial credibility is equal to one: the output gap, inflation, and interest rate increase and return to the steady state monotonically.

[24]We choose these values for realistic model simulations based on the estimated nominal factor for each credibility level for the U.S. economy.

As credibility deteriorates, however, the model dynamics change considerably. When a^{π} is positive, a positive preference shock reduces credibility, as a positive a^{π} makes private forecasts relatively more accurate compared with the central bank's forecasts. Specifically, private forecasts become more accurate as credibility deteriorates since private agents anticipate higher inflation than the central bank at time zero due to the positive perception on the nominal factor. Thus, credibility drops more than 0.2 points immediately in the low credibility case and remains at the lower level for longer periods.

Similarly, the nominal factor is revised upward and returns to zero slowly in all cases, as realized inflation is higher than expected. This shift in the nominal factor again feeds back to even higher inflation and this feedback effect gets larger as credibility decreases since the influence of the nominal factor becomes more substantial. As credibility decreases, private expectations of future inflation increase because of a higher a^{π}. This increases inflation instantly and feeds into a higher nominal factor. This stimulates inflation again and the self-referential cycle continues. For this reason, inflation even overshoots in the low-credibility case.

Initially, the output gap increases sharply due to the positive preference shock. However, as credibility drops, private agents expect higher nominal interest rates in future periods, and the central bank actually increases the nominal interest rate to stabilize higher inflation. Therefore, the output gap shows a more contracted path compared with the case with a higher credibility.

While the initial real factor is zero, it is revised upward due to the positive reaction of the output gap. Then, as private agents update their perceptions on the real factor, it comes back to zero while the deviation is larger in the low-credibility case.

6.3 Improving Credibility

Finally, in this subsection, we discuss how a central bank can enhance its credibility. There are many things that can affect central bank credibility. For instance, even a single word spoken by policymakers could affect central bank credibility. Hence, it is not an easy task to discuss every possible option for improving credibility. For

Table 4. Monetary Policy Rules and Average Credibility

ϕ_π	**1.12 (–1)**	**1.62 (–0.5)**	**2.12 (Benchmark)**	**2.62 (+0.5)**	**3.12 (+1)**
Mean Credibility	89	95	100	101	102
ϕ_x	**0.07 (–0.1)**	**0.17 (Benchmark)**	**0.27 (+0.1)**	**0.37 (+0.2)**	
Mean Credibility	100	100	101	102	

Note: The average credibility in the benchmark model is normalized to 100. Therefore, the average credibility under different specifications can be interpreted as the percentage changes compared to the benchmark case.

this reason, we only focus on the role of the central bank's systematic reactions to economic developments in shaping credibility. To this end, we analyze how the average credibility changes as we vary the parameters governing the monetary policy rule specified in Equation (11).

Table 4 presents how the average credibility changes as the monetary policy reaction function varies. The upper panel shows the changes in credibility when the central bank's response to inflation changes, while the lower panel presents the changes in credibility as the reaction to the output gap changes. Compared with the benchmark case that represents the current policy practice, stronger reactions to both inflation and the output gap result in a higher mean credibility, though the increments are quite small. This is reasonable since stronger responses to inflation and the output gap make them easier to forecast by pushing them closer to their respective targets.[25]

[25] However, it is uncertain whether stronger responses to inflation and the output gap are welfare improving, since a stronger reaction to inflation results in a more volatile output gap, while a stronger response to the output gap leads to more volatile inflation. This suggests that the optimal monetary policy may depend on central bank credibility. Although we do not analyze the optimal monetary policy under the credibility restriction since that is out of the scope of this paper, we believe that it might be an interesting future research topic.

On the contrary, weaker responses to inflation sharply reduce credibility, while weaker reactions to the output gap do not change credibility substantially. This non-linear relationship between monetary policy reactions and credibility suggests that the Fed is efficiently conducting its policy so that it achieves high credibility without creating excessively volatile macroeconomic responses to monetary policy.

7. Conclusion

In this research, a numerical measure of central bank credibility is proposed and its effects on macroeconomic stability are examined. The main contributions of this paper are to show to what extent accounting for credibility affects macroeconomic stability. Specifically, it is shown that volatility in the output gap and inflation increase as credibility deteriorates due to the self-referential effect of private beliefs. This model can generate endogenous volatility changes based on the shifts in credibility without relying on exogenous volatility regime changes. This result theoretically confirms the idea that maintaining a credible reputation helps to anchor private expectations and to achieve macroeconomic stability. Despite their importance, these results have not yet been fully analyzed in a New Keynesian framework with endogenous central bank credibility concerns, and this paper provides a useful benchmark that can be easily analyzed.

The findings derived in this research have important implications for many issues in monetary policy. For instance, the definition of credibility used in this paper is related to the forward guidance, especially the Delphic effects of forward guidance, as studied in, among others, Campbell et al. (2012, 2017). A central bank's forecasts as described in this model are closely related to forward guidance, as they hint at a future course for the economy and communicate information held by the central bank. Most analyses of forward guidance, however, implicitly assume full credibility so that private agents believe what the central bank says and focus on information flow from the central bank to the public. If the private agents do not think that the central bank is fully credible, the Delphic effects that were described might disappear. Therefore, this research suggests that it is necessary to take the credibility issue into account when the effects of forward guidance are examined.[26]

Although central bank credibility has been considered an important feature that shapes the efficacy of monetary policy, it has not been sufficiently taken into account in quantitative monetary policy studies. Nonetheless, credibility matters in many cases, as documented in this paper, hence more serious research regarding this issue is warranted.

Appendix A. Actual Laws of Motion

First, obtain a minimum state variable (MSV) solution for the rational expectations model below by method of undetermined coefficients.

$$\begin{aligned} x_t &= E_t x_{t+1} - \sigma^{-1}(i_t - E_t \pi_{t+1} - r_t^n) \\ \pi_t &= \kappa x_t + \beta E_t \pi_{t+1} + u_t \\ i_t &= \phi_\pi \pi_t + \phi_x x_t + m_t, \end{aligned} \tag{A.1}$$

A unique and bounded solution exists if $\kappa(\phi_\pi - 1) + (1 - \beta)\phi_x > 0$ holds. Note that this condition is satisfied in a current calibration. The MSV solution is given as

$$\begin{aligned} x_t &= a_{11} r_t^n + a_{12} u_t + a_{13} m_t \\ \pi_t &= a_{21} r_t^n + a_{22} u_t + a_{23} m_t \\ i_t &= a_{31} r_t^n + a_{32} u_t + a_{33} m_t, \end{aligned}$$

where

$$\begin{aligned} a_{21} &= \frac{1}{\left(\frac{1-\beta\rho_r}{\kappa}\right)((1-\rho_r)\sigma + \phi_x) - \rho_r + \phi_\pi} \\ a_{11} &= a_{21}\frac{1-\beta\rho_r}{\kappa} \\ a_{31} &= \phi_\pi a_{21} + \phi_x a_{11} \end{aligned}$$

[26]In their experiment paper, Ahrens, Lustenhouwer, and Tettamanzi (2017) emphasize the importance of credibility in shaping the effectiveness of forward guidance.

$$
\begin{aligned}
a_{22} &= \frac{1-\rho_u+\frac{\phi_x}{\sigma}}{\kappa\left[\frac{\phi_\pi-\rho_u}{\sigma}+\frac{1-\beta\rho_u}{\kappa}\left(1-\rho_u+\frac{\phi_x}{\sigma}\right)\right]}\\
a_{12} &= a_{22}\frac{(1-\beta\rho_u)}{\kappa}-\frac{1}{\kappa}\\
a_{32} &= \phi_\pi a_{22}+\phi_x a_{12}\\
a_{23} &= \frac{-1}{\left[\frac{1-\beta\rho_m}{\kappa}(\sigma(1-\rho_m)+\phi_\pi)+\phi_\pi-\rho_m\right]}\\
a_{13} &= a_{23}\frac{1-\beta\rho_m}{\kappa}\\
a_{33} &= \phi_\pi a_{23}+\phi_x a_{13}+1.
\end{aligned}
$$

Combining Equations (14), (19), and (20), obtain future expectations on endogenous variables

$$
\begin{aligned}
\hat{E}_t x_T &= (1-\xi_{t-1})f_1^{T-t}a_{t-1}^x+a_{11}\rho_r^{T-t}r_t^n+a_{12}\rho_u^{T-t}u_t+a_{13}\rho_m^{T-t}m_t\\
\hat{E}_t \pi_T &= (1-\xi_{t-1})f_2^{T-t}a_{t-1}^\pi+a_{21}\rho_r^{T-t}r_t^n+a_{22}\rho_u^{T-t}u_t+a_{23}\rho_m^{T-t}m_t\\
\hat{E}_t i_T &= (1-\xi_{t-1})(\lambda_1 f_1^{T-t}a_{t-1}^x+\lambda_2 f_2^{T-t}a_{t-1}^\pi)+a_{31}\rho_r^{T-t}r_t^n\\
&\quad +a_{32}\rho_u^{T-t}u_t+a_{33}\rho_m^{T-t}m_t.
\end{aligned}
$$

Inserting these expectations into policy rules, Equations (9), (10), and (11), gives the following system of equations:

$$
\begin{aligned}
x_t+\frac{i_t}{\sigma} =& \frac{(1-\xi_{t-1})f_1}{1-\beta f_1}\left(1-\beta-\frac{\lambda_1\beta}{\sigma}\right)a_{t-1}^x\\
&+\frac{(1-\xi_{t-1})f_2}{\sigma(1-\beta f_2)}(1-\lambda_2\beta)a_{t-1}^\pi\\
&+\frac{\sigma(1-\beta)\rho_r a_{11}+\rho_r a_{21}-\beta\rho_r a_{31}+1}{\sigma(1-\beta\rho_r)}r_t^n\\
&+\frac{\sigma(1-\beta)\rho_u a_{12}+\rho_u a_{22}-\beta\rho_u a_{32}}{\sigma(1-\beta\rho_u)}u_t\\
&+\frac{\sigma(1-\beta)\rho_m a_{13}+\rho_m a_{23}-\beta\rho_m a_{33}}{\sigma(1-\beta\rho_m)}m_t
\end{aligned}
$$

$$\pi_t - \kappa x_t = \frac{\kappa(1-\xi_{t-1})\alpha\beta f_1}{1-\alpha\beta f_1} a^x_{t-1} + \frac{(1-\alpha)\beta(1-\xi_{t-1})f_2}{1-\alpha\beta f_2} a^\pi_{t-1}$$
$$+ \frac{\rho_r(\kappa\alpha\beta a_{11} + (1-\alpha)\beta a_{21})}{1-\alpha\beta\rho_r} r^n_t$$
$$+ \frac{\kappa\alpha\beta\rho_u a_{12} + (1-\alpha)\beta\rho_u a_{22} + 1}{1-\alpha\beta\rho_u} u_t$$
$$+ \frac{\rho_m(\kappa\alpha\beta a_{13} + (1-\alpha)\beta a_{23})}{1-\alpha\beta\rho_m} m_t$$

$$i_t - \phi_x x_t - \phi_\pi \pi_t = m_t. \tag{A.2}$$

In matrix form, this system can be expressed as follows:

$$\begin{bmatrix} 1 & 0 & \frac{1}{\sigma} \\ -\kappa & 1 & 0 \\ -\phi_x & -\phi_\pi & 1 \end{bmatrix} \begin{bmatrix} x_t \\ \pi_t \\ i_t \end{bmatrix} = \begin{bmatrix} d_{11} & d_{12} & d_{13} & d_{14} & d_{15} \\ d_{21} & d_{22} & d_{23} & d_{24} & d_{25} \\ 0 & 0 & 0 & 0 & 1 \end{bmatrix} \begin{bmatrix} a^x_{t-1} \\ a^\pi_{t-1} \\ r^n_t \\ u_t \\ m_t \end{bmatrix}. \tag{A.3}$$

By inverting the leading matrix in the left-hand side, we can obtain the actual laws of motion,

$$\begin{bmatrix} x_t \\ \pi_t \\ i_t \end{bmatrix} =$$
$$\begin{bmatrix} \frac{\sigma d_{11} - \phi_\pi d_{21}}{h} & \frac{\sigma d_{12} - \phi_\pi d_{22}}{h} & a_{11} & a_{12} & a_{13} \\ \frac{\kappa\sigma d_{11} + (\phi_x + \sigma) d_{21}}{h} & \frac{\kappa\sigma d_{12} + (\phi_x + \sigma) d_{22}}{h} & a_{21} & a_{22} & a_{23} \\ \frac{\sigma(\phi_x + \kappa\phi_\pi) d_{11} + \phi_\pi \sigma d_{21}}{h} & \frac{\sigma(\phi_x + \kappa\phi_\pi) d_{12} + \phi_\pi \sigma d_{22}}{h} & a_{31} & a_{32} & a_{33} \end{bmatrix} \begin{bmatrix} a^x_{t-1} \\ a^\pi_{t-1} \\ r^n_t \\ u_t \\ m_t \end{bmatrix}, \tag{A.4}$$

where $h = \phi_x + \kappa\phi_\pi + \sigma$. Note that the actual laws of motion (ALM) and future expectations are identical to those of the rational expectations model without credibility when ξ_{t-1} converges to one. To see this, note that $d_{ij} = 0$ when $\xi_{t-1} = 1$ for all i and j. Therefore, ALM reduces to an MSV solution of underlying rational expectations model as ξ_{t-1} converges to one.

Appendix B. Estimation Procedure

The measurement equation is given as

$$\begin{bmatrix} x_t \\ \pi_t \\ i_t \\ \hat{E}_t^P[x_{t+1}] \\ \hat{E}_t^P[\pi_{t+1}] \\ \hat{E}_t^P[i_{t+1}] \end{bmatrix} = \begin{bmatrix} \mathcal{I}_3 & \mathcal{O}_{3\times 8} \\ \mathcal{O}_{3\times 3} & \mathcal{M}_{3\times 8} \end{bmatrix} \begin{bmatrix} x_t \\ \pi_t \\ i_t \\ a_t^x \\ a_t^\pi \\ \xi_t \\ a_{t-1}^x \\ a_{t-1}^\pi \\ r_t^n \\ u_t \\ m_t \end{bmatrix} + o_t, \tag{B.1}$$

where

$$\mathcal{M}_{3\times 8} = \begin{bmatrix} 0 & 0 & 0 & 0 & 0 & f_1 & 0 & \rho_r a_{11} & \rho_u a_{12} & \rho_m a_{13} \\ 0 & 0 & 0 & 0 & 0 & 0 & f_2 & \rho_r a_{21} & \rho_u a_{22} & \rho_m a_{23} \\ 0 & 0 & 0 & 0 & 0 & \lambda_1 f_1 & f_2 & \rho_r a_{31} & \rho_u a_{32} & \rho_m a_{33} \end{bmatrix}. \tag{B.2}$$

The transition equations consist of laws of motion which are derived in the main text and deterministic identity equations.

The estimation procedure can be summarized by the following:

(i) Draw $\mathcal{V}$ from the Markov chain (proposal parameter).

(ii) Generate particles of exogenous disturbances z.

(iii) For $t = 1 : T$,
- Propagate state variables x_t given particles z_t and initial states x_{t-1}.
- Evaluate likelihood functions $p(y_t \mid x_t, z_t, \mathcal{V})$.
- Resample the particles weighted by their likelihoods.
- Approximate the time t likelihood $\hat{p}(y_t \mid \mathcal{V})$ weighted by likelihoods of each particle.

(iv) Approximate the likelihood function $\hat{p}(y_{1:T} \mid \mathcal{V})$.

(v) With probability $\alpha(\mathcal{V} \mid \theta^{i-1}) = \min\left\{1, \frac{\hat{p}(y_{1:T}|\mathcal{V})p(\mathcal{V})}{\hat{p}(y_{1:T}|\theta^{i-1})p(\theta^{i-1})}\right\}$, set $\theta^i = \mathcal{V}$ otherwise, $\theta^i = \theta^{i-1}$ where $p(\theta)$ is the prior distribution.

We set the standard deviations of observation errors to 20 percent of the standard deviations of corresponding actual data series following Herbst and Schorfheide (2015).

Appendix C. Model Stability

Combining Equation (16) and Equation (25) gives

$$a_t = (F + KC_0 - KH)a_{t-1} + K\Omega\varepsilon_t. \tag{C.1}$$

As explained in Eusepi, Giannoni, and Preston (2015), the self-referentiality of beliefs can lead to instability. This instability arises if any eigenvalue of the matrix $F+KC_0-KH$ lies outside the unit circle. This approach is, however, not possible in this case, because the model stability depends on the evolution of the credibility measure ξ. Specifically, C_0 contains ξ, so the evolution of credibility affects the dynamics of beliefs. For this reason, we analyze a Jacobian matrix of this non-linear system following Hommes and Lustenhouwer (2019) and Branch and McGough (2010), among many others.

$\tilde{\xi}_t$ is determined by the following:

$$\begin{aligned}\tilde{\xi}_t = 1 - &\frac{\exp\left(\delta_1 U_t^P\right)}{\exp\left(\delta_1 U_t^P\right) + \exp\left(\delta_1 U_t^{CB}\right)} \\ &\times \left(1 - \exp\left(-\delta_2 \left(\hat{E}_{t-1}^{CB}\pi_t - \hat{E}_{t-1}^{P}\pi_t\right)^2\right)\right).\end{aligned} \tag{C.2}$$

We introduce an auxiliary variable q_t which is defined as below.

$$q_t = \frac{1 - \exp(\delta_1 U_t^{CB} - \delta_1 U_t^P)}{1 + \exp(\delta_1 U_t^{CB} - \delta_1 U_t^P)} = \tanh\left(-\frac{\delta_1}{2}(U_t^{CB} - U_t^P)\right) \tag{C.3}$$

Using this auxiliary variable and evaluating the distance between two predictions, $\tilde{\xi}_t$ can be simplified as below.

$$\tilde{\xi}_t = 1 - \frac{q_t + 1}{2}\left(1 - \exp(-\delta_2 f_2^2 (a_{t-2}^{\pi})^2)\right) \tag{C.4}$$

The model can be summarized as a system of non-linear equations.

$$\begin{aligned} z_t &= M_1(a_{t-1}, \xi_{t-1}, s_t) \\ a_t &= M_2(a_{t-1}, z_t, s_{t-1}) \\ \xi_t &= M_3(q_t, a_{t-2}, \xi_{t-1}) \\ q_t &= M_4(a_{t-2}, s_{t-1}, z_t) \\ a_{t-1} &= I(a_{t-1}) \\ s_{t+1} &= M_5(s_t) \\ s_t &= I(s_t) \end{aligned} \tag{C.5}$$

Before analyzing the stability, we show that there is a steady state in this system.

Proposition 1. *A steady state of the model exists and this steady state satisfies* $x = 0$, $\pi = 0$, $i = 0$, $a = 0$, $s = 0$, $q = 0$, *and* $\xi = 1$.

Proof. It is easy to show that $x = 0$, $\pi = 0$, $i = 0$, $a = 0$, $s = 0$, $q = 0$, and $\xi = 1$ solve the system of equations.

As this steady state with $x = 0$, $\pi = 0$, and $i = 0$ coincides with that of a rational expectations model, we label this steady state as the fundamental steady state.

The next proposition provides the global stability result of the model. The main idea behind this result is that, as zero credibility is the most de-stabilizing condition, the system is globally stable if the eigenvalues of the Jacobian are inside the unit circle when the credibility measure ξ is fixed at zero for all t.

Proposition 2. *The fundamental steady state is globally stable under the baseline calibration.*

Proof. If full credibility, $\xi = 1$ for all t, is imposed, the economy has a globally stable steady state, which is the fundamental steady state, as the belief terms a do not affect the dynamics, and the steady state is exactly the same with that of the underlying rational expectations model. As ξ moves away from one to zero, the influence of the belief terms a on the system increases and the economic dynamics become more unstable. Therefore, if the economic system is stable under zero credibility, it is globally stable. Under the assumption that $\xi = 0$ for all t, the system can be written as follows:

$$\begin{aligned} a_t &= M_2(a_{t-1}, M_1(a_{t-1}, s_t), s_{t-1}) \\ s_{t+1} &= M_5(s_t) \\ s_t &= I(s_t). \end{aligned} \tag{C.6}$$

Then, we can obtain the following Jacobian matrix:

$$\begin{bmatrix} f_1 - kf_{11} + \Delta_{11} & -kf_{12} + \Delta_{12} & ka_{11} & ka_{12} & ka_{13} & -ka_{11}\rho_r & -ka_{12}\rho_u & -ka_{13}\rho_m \\ -kf_{21} + \Delta_{21} & f_2 - kf_{22} + \Delta_{22} & ka_{21} & ka_{22} & ka_{23} & -ka_{21}\rho_r & -ka_{22}\rho_u & -ka_{23}\rho_m \\ 0 & 0 & \rho_r & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & \rho_u & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & \rho_m & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 & 0 & 0 & 0 & 0 \end{bmatrix}, \tag{C.7}$$

where

$$\begin{aligned} \Delta_{ij} = K_{i1}\frac{\sigma dk_{1j} - \phi_\pi dk_{2j}}{\phi_x + \kappa\phi_\pi + \sigma} + Ki2\frac{\kappa\sigma dk_{1j} + (\phi_x + \sigma)dk_{2j}}{\phi_x + \kappa\phi_\pi + \sigma} \\ + K_{i3}\frac{\sigma(\phi_x + \kappa\phi_\pi)dk_{1j} + \phi_\pi\sigma dk_{2j}}{\phi_x + \kappa\phi_\pi + \sigma} \end{aligned}$$

$$dk_{11} = \frac{f_1}{1 - \beta f_1}\left(1 - \beta - \frac{\lambda_1\beta}{\sigma}\right), \quad dk_{12} = \frac{f_2}{1 - \beta f_2}\left(\frac{1 - \lambda_2\beta}{\sigma}\right)$$

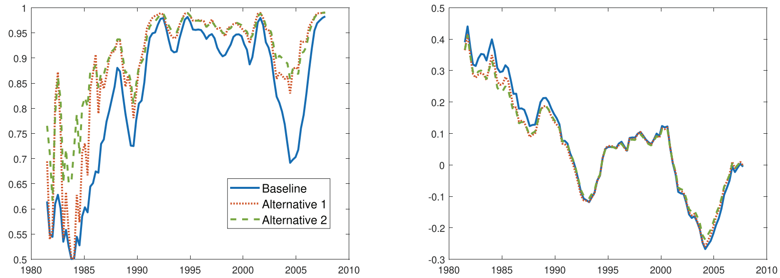
$$dk_{21} = \frac{\kappa\alpha\beta f_1}{1 - \alpha\beta f_1}, \quad dk_{22} = \frac{(1 - \alpha)\beta f_2}{1 - \alpha\beta f_2},$$

where $kf_{ij} = f_j K_{ij} + \lambda_j f_j K_{i3}$, $ka_{ij} = K_{i1}a_{1j} + K_{i2}a_{2j} + K_{i3}a_{3j}$, and K_{ij} denotes (i, j) element of the Kalman gain matrix. It has three zero and five non-zero eigenvalues, which are all inside the unit circle under baseline calibration. Therefore, the fundamental steady state is globally stable.

Appendix D. Alternative Calibration

In this appendix, we provide evidence that using alternative calibrations for δ_1, δ_2, and W does not change the quantitative result considerably. To be precise, we illustrate this robustness by showing

Figure D.1. Estimated Credibility with Different Calibrations (left) and Estimated a^{π} with Different Calibrations (right)



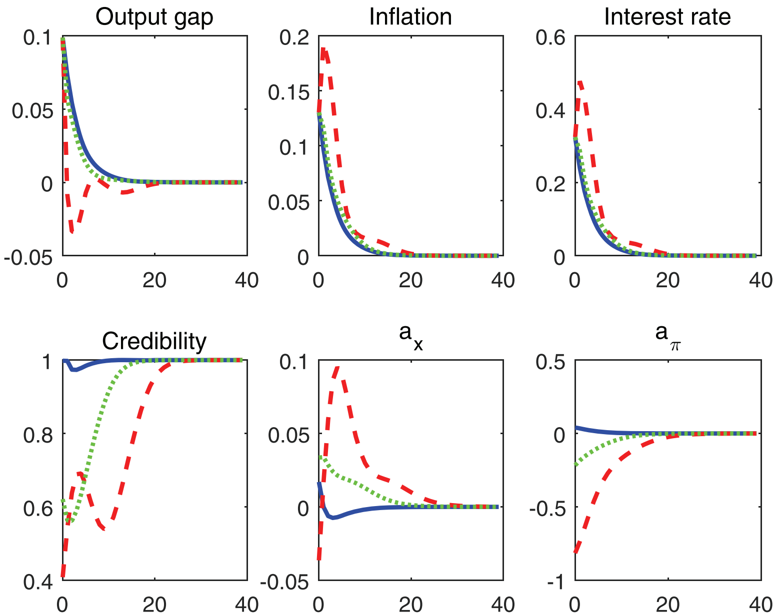
the estimated credibility and private beliefs across different calibrations. Figure D.1 presents the estimated results. The blue solid lines represent the estimated series obtained from the baseline calibration. The red dotted lines and the green dashed lines are estimated results with different calibrations. Specifically, “Alternative 1” shows the case that $\delta_1 = \delta_2 = 15$ while “Alternative 2” stands for the case with different W shown below.

$$W = \begin{bmatrix} 0 & 0 & 0 \\ 0 & 1 & 0 \\ 0 & 0 & 0 \end{bmatrix}$$

We choose $\delta_1 = \delta_2 = 15$ as an alternative calibration in this exercise, but choosing any values with sufficiently large δ_2, say greater than 2, produces very similar result. Under “Alternative 2” assumption, the private agents only care about the forecast errors associated with inflation. The estimated output shows that the results are quite similar to the baseline case.

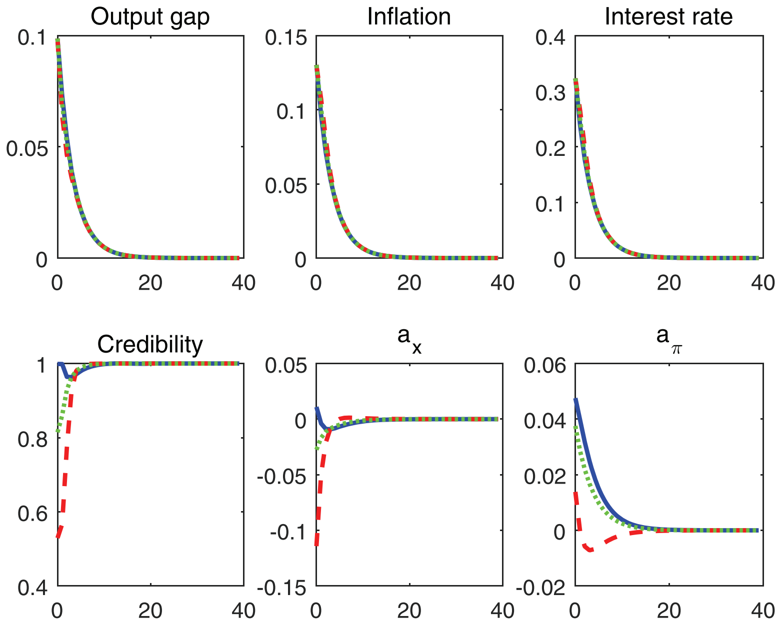
Appendix E. Additional Impulse Responses

Figure E.1. Impulse Responses to One-Standard-Deviation Preference Shock When a_π Is Negative



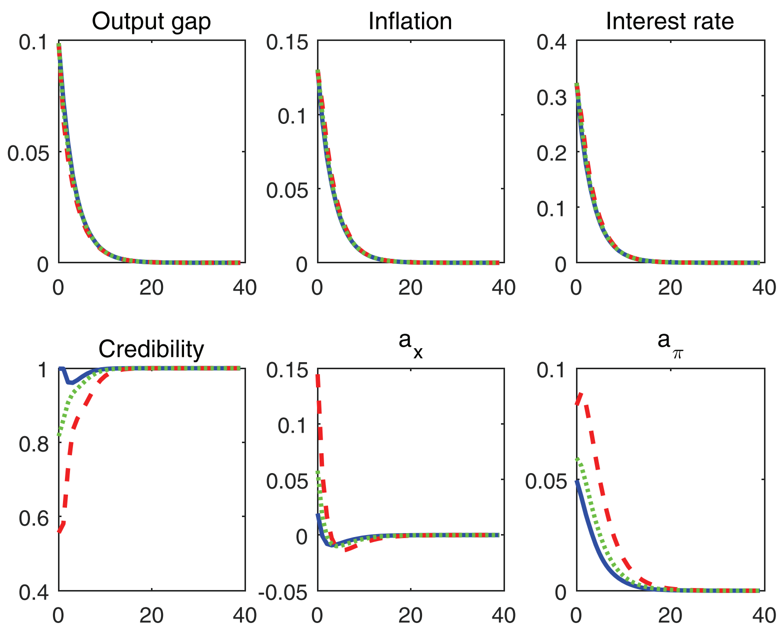
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^{\pi}_{-1} = -0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^{\pi}_{-1} = -0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^{\pi}_{-1} = -0.3$.

Figure E.2. Impulse Responses to One-Standard-Deviation Preference Shock When a_x Is Negative



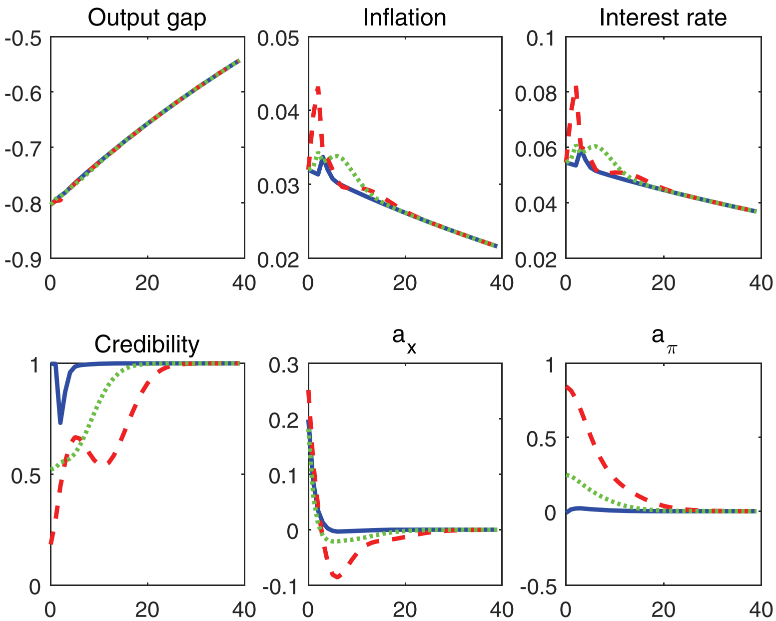
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^x_{-1} = -0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^x_{-1} = -0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^x_{-1} = -0.3$.

Figure E.3. Impulse Responses to One-Standard-Deviation Preference Shock When a_x Is Positive



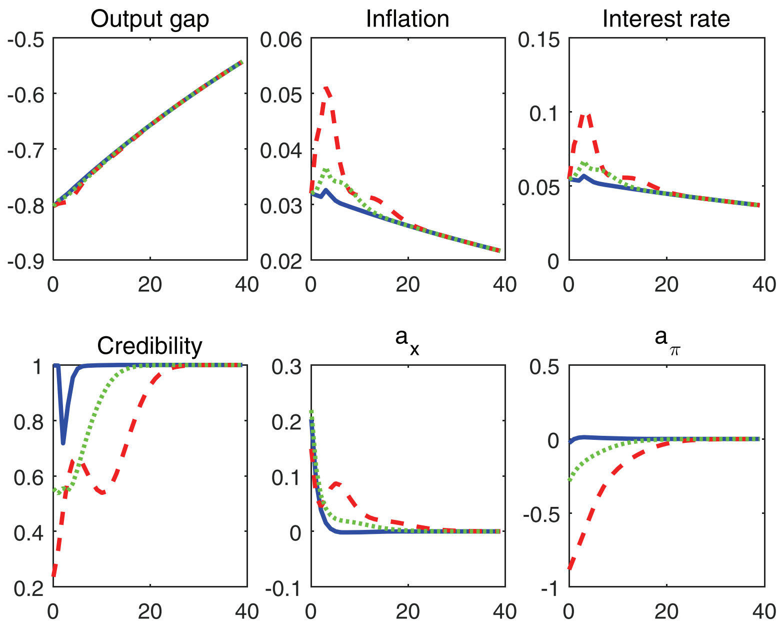
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^{x}_{-1} = 0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^{x}_{-1} = 0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^{x}_{-1} = 0.3$.

Figure E.4. Impulse Responses to One-Standard-Deviation Cost-Push Shock When a_π Is Positive



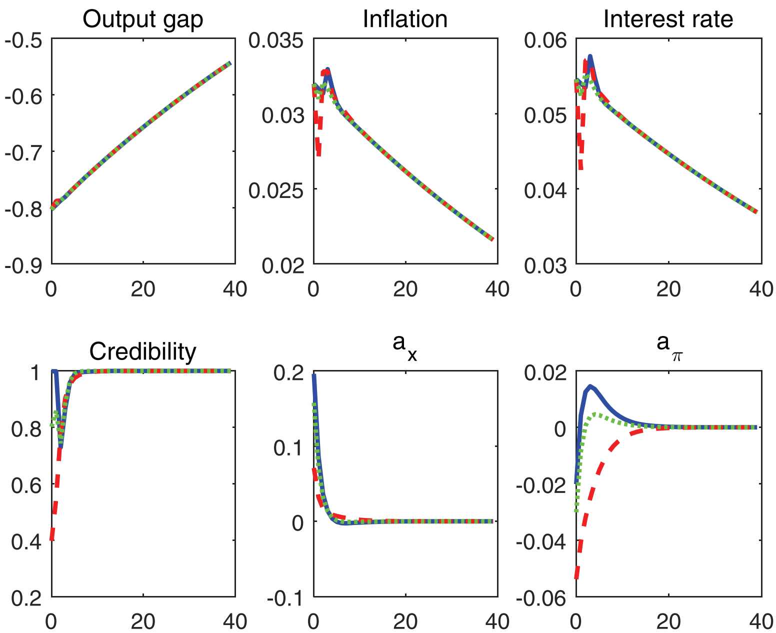
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^{\pi}_{-1} = 0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^{\pi}_{-1} = 0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^{\pi}_{-1} = 0.3$.

Figure E.5. Impulse Responses to One-Standard-Deviation Cost-Push Shock When a_π Is Negative



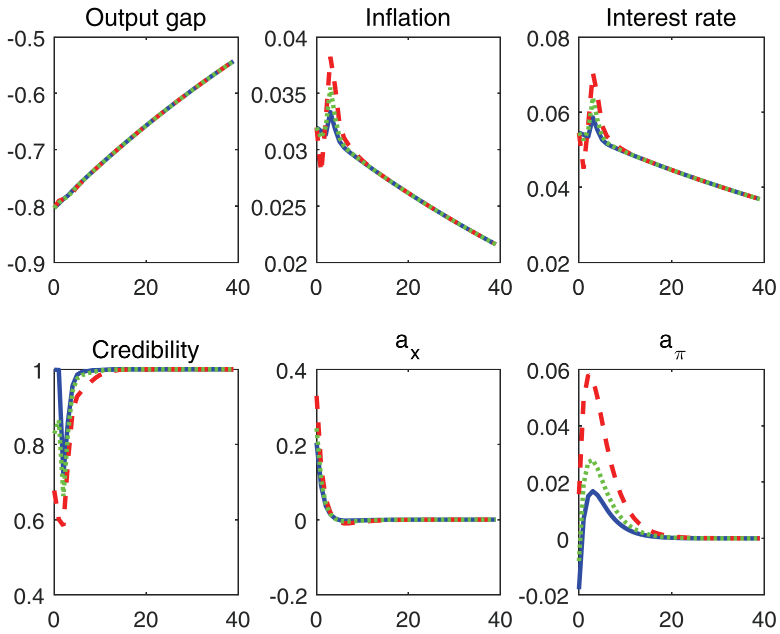
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^{\pi}_{-1} = -0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^{\pi}_{-1} = -0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^{\pi}_{-1} = -0.3$.

Figure E.6. Impulse Responses to One-Standard-Deviation Cost-Push Shock When a_x Is Negative



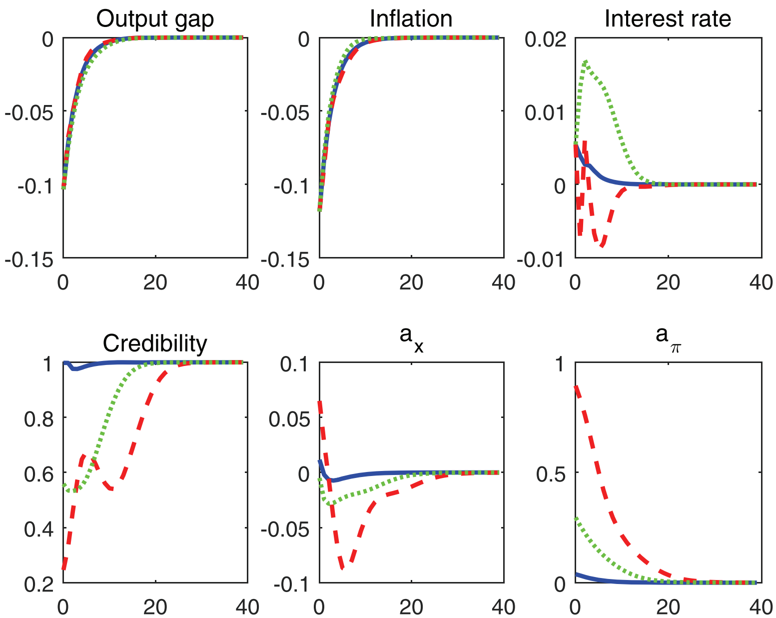
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^x_{-1} = -0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^x_{-1} = -0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^x_{-1} = -0.3$.

Figure E.7. Impulse Responses to One-Standard-Deviation Cost-Push Shock When a_x Is Positive



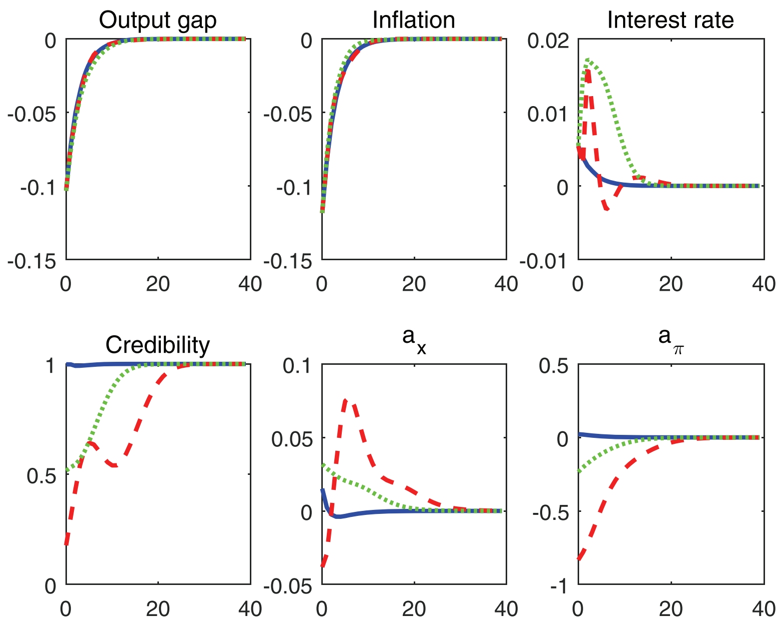
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^x_{-1} = 0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^x_{-1} = 0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^x_{-1} = 0.3$.

Figure E.8. Impulse Responses to One-Standard-Deviation Monetary Policy Shock When a_π Is Positive



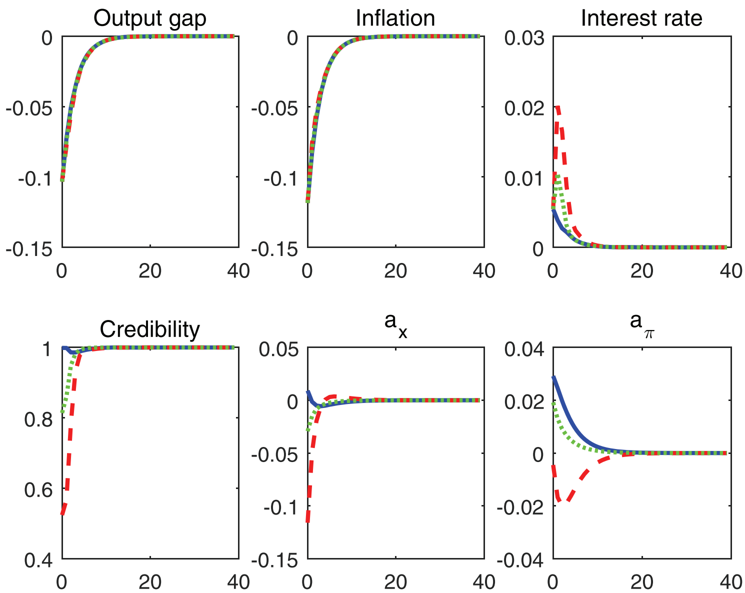
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^{\pi}_{-1} = 0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^{\pi}_{-1} = 0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^{\pi}_{-1} = 0.3$.

Figure E.9. Impulse Responses to One-Standard-Deviation Monetary Policy Shock When a_π Is Negative



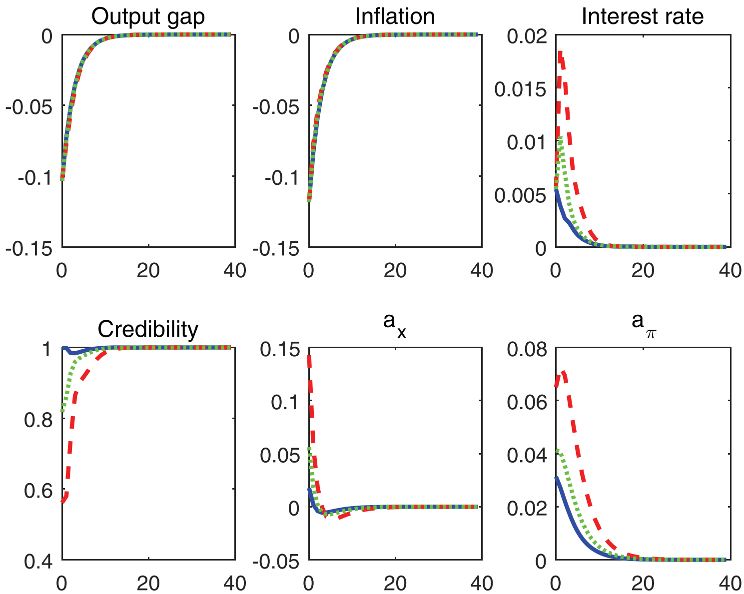
Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^{\pi}_{-1} = -0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^{\pi}_{-1} = -0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^{\pi}_{-1} = -0.3$.

Figure E.10. Impulse Responses to One-Standard-Deviation Monetary Policy Shock When a_x Is Negative



Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^x_{-1} = -0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^x_{-1} = -0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^x_{-1} = -0.3$.

Figure E.11. Impulse Responses to One-Standard-Deviation Monetary Policy Shock When a_x Is Positive

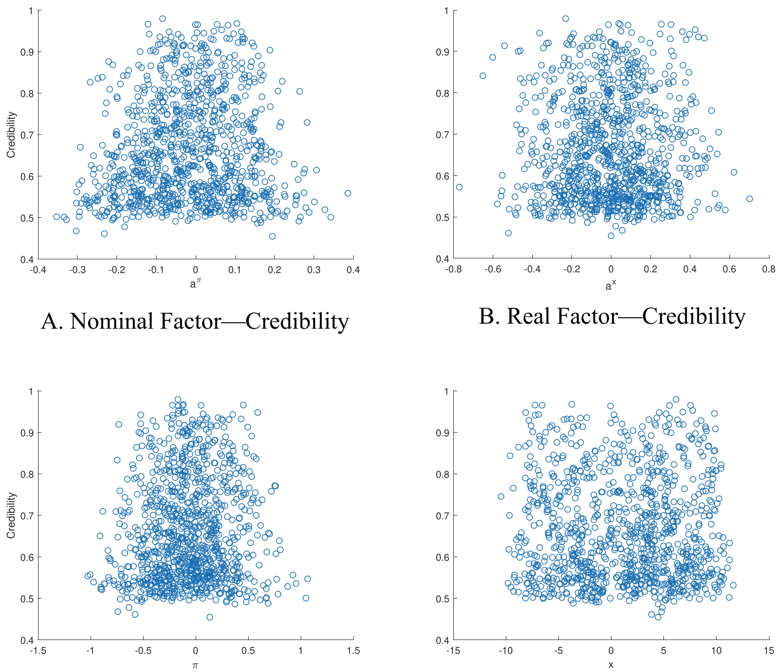


Note: The blue solid line represents impulse responses from $\xi_{-1} = 1$, $a^x_{-1} = 0.01$ and the red dashed line represents impulse responses from $\xi_{-1} = 0.4$, $a^x_{-1} = 0.9$. The green dotted line represents impulse responses from $\xi_{-1} = 0.6$, $a^x_{-1} = 0.3$.

Appendix F. Scatter Plots of Simulated Series

We provide additional evidence that a combination of lower credibility and shifting private beliefs undermines macroeconomic stability based on the simulated time series. Figure F.1 shows the relationship between credibility and economic outcomes. The upper panels present the relationship between private beliefs and credibility while the bottom panels illustrate the co-movements of macro variables and credibility. The nominal factor spreads out as credibility deteriorates. This leads to more volatile realizations of inflation that is caused by a self-referential effect examined above. The real factor shows the same pattern as in the nominal factor. It suggests that the private beliefs are closely related to central bank credibility as asserted. However, the output gap does not show any distinctive pattern across the different levels of credibility. This is in line with the above result that the volatility of the output gap is not significantly affected by credibility.

Figure F.1. Scatter Plots of Simulated Series



A. Nominal Factor—Credibility

B. Real Factor—Credibility

C. Inflation—Credibility

D. Output—Credibility

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After the Storm: Natural Disasters and Bank Solvency*

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This study examines how natural disasters affect the solvency of banks. It explores (i) whether and how natural disasters affect bank solvency, (ii) how accounting and regulatory measures of bank solvency reflect a bank's true affectedness, and (iii) whether the effects vary across different types of banks. Analyzing a comprehensive data set on natural catastrophes and detailed financial statements for 9,928 banks that operate in 149 countries, the main finding is that damages from disasters matter: they negatively affect capital ratios, and the severity of their impact depends on a bank's location, capitalization, and business model. Particularly, the results show that accounting measures of solvency are more sensitive to disasters than are regulatory measures. Evidence of a bank's sensitivity to natural disasters and the suitability of capital ratios to assess this sensitivity may both be helpful for financial institutions and regulatory authorities in designing appropriate risk mitigation strategies.

JEL Codes: G21, G28, G32.

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1. Introduction

The recent rise in the frequency and severity of natural disasters such as floods, droughts, wildfires, and extreme winds (hurricanes, typhoons, tornadoes, etc.) is often attributed to climate change, and climate change itself to our production and consumption behavior (Rummukainen 2012; Mechler and Bouwer 2015; World Economic Forum 2018). Because natural disasters primarily affect the real economy, research on their economic effects has mainly focused on their impact on production and growth (Hallegatte 2014; Arouri, Nguyen, and Youssef 2015; Lesk, Rowhani, and Ramankulty 2016). Only recently has research started to explore how such disasters affect financial institutions and the broader financial markets. The relevance of natural disasters for the risk and the risk management of individual institutions can be explained through their claims—e.g., via loans, bonds, and stocks—on the real economy. In addition, financial institutions, particularly banks and the banking network, may be exposed to operational risks if disasters hit the institutions' physical locations and computer systems. Finally, if banks have to cut their lending following a disaster, it reduces their income opportunities and imposes capital constraints on their customers (Brei, Mohan, and Strobl 2019).

There is a growing concern among financial regulators and central banks that damages may affect the financial system as a whole (Batten, Sowerbutts, and Tanaka 2016). The Network for Greening the Financial System (NGFS), which comprises many of the world's most influential central banks and supervisory authorities, has recently outlined the need to incorporate climate risks into financial policies and regulatory frameworks (NGFS 2018). In addition, the European Union's High-Level Expert Group on Sustainable Finance (HLEG) has repeatedly argued that the financial system is a crucial component in any intended moves to shift the overall economy towards a more sustainable system, i.e., a system that balances the needs of our economy, society, and ecology (HLEG 2018). Financial institutions, especially banks, are expected to provide the financial expertise, backing, and networking necessary for the transition towards sustainability (SFSG 2018).

Whether and how a given financial institution is affected by a natural disaster is difficult to assess. Its claims against exposed counterparties (e.g., mortgage loans, business loans, etc.) may be affected with varying levels of intensity. In addition, even if a given loan has to be written off because, e.g., a firm is forced out of business or a residential property is damaged beyond repair and the homeowners have to default on their loans, the disaster may create new demand for loans as restructuring and rebuilding activities commence (Cortés and Strahan 2017; Barth, Sun, and Zhang 2019).

Challenges may arise from disasters themselves (physical risks) as well as from changes in the legal framework (transition risks). A further complexity arises from the interconnectivity of different actors in the financial markets that makes them reciprocally vulnerable to risks. For example, interbank lending in the money markets or the participation of banks in insurance companies can indirectly transfer risks among institutions (Battiston et al. 2017). In a similar fashion, the impact from disasters depends on the risk management strategies of both banks and their customers, i.e., the instruments applied to hedge the damages from disasters (Benson and Clay 2004). The multiple factors affecting banks and the banking system may explain why evidence of the effects from natural disasters is mixed, and a more granular perspective is needed.

Against the backdrop of the rising frequency and severity of natural disasters in recent years and the complex effects external shocks have on bank stability, this study aims to explore whether and how damages from natural disasters translate into potential solvency problems for banks, whether the effect varies across different types of banks, and how different measures of solvency reflect this effect. Particularly, we address the following research questions:

- To what extent do natural disasters affect bank solvency?
- Do natural disasters affect accounting based measures of solvency as much as they affect regulation based measures?
- Are different types of banks affected differently by natural disasters?

Our study focuses on bank solvency because a bank's ability to withstand risks and remain solvent even under adverse conditions is existential for both its own stability as well as the soundness of the financial system as a whole (Flannery and Giacomini 2015). Recent research provides additional evidence that the capital ratio of banks has an impact on bank lending in the context of natural disasters. Rehbein and Ongena (2022) find that banks with low capital levels tend to lend less in the aftermath of a disaster. Moreover, banks that have lower solvency ratios and are affected by floods also reduce the lending to companies not directly affected by a disaster.

Banking regulations typically focus on ensuring that banks maintain sufficient capital. The ability of the banking system to manage risks is driven both by individual institutions' ability to absorb damages and by the diversification of risks within the system (Batten, Sowerbutts, and Tanaka 2016). Although banking regulations have undergone considerable refinements in recent years, particularly after the 2007/08 subprime mortgage crisis, they are only now starting to consider natural disasters as a potential risk factor (European Banking Authority 2019). Our study aims to shed light on the possibility that natural disasters may pose the next big threat for our economy and for our financial system. That way, banks and bank regulators can better prepare themselves for the predicted increase in the severity and frequency of such events.

The remainder of our paper is structured as follows: Section 2 provides an overview of the research related to bank solvency and natural disasters, and further outlines the contribution of this study in the context of the existing literature. Section 3 develops the underlying hypotheses about banks' sensitivity to natural disasters in the context of existing theories and discusses the measurement of this sensitivity. Section 4 explains the database and Section 5 the methodological approach of the study. Section 6 discusses our results and explores both the affectedness of specific types of banks from natural disasters as well as the suitability of the accounting capital ratio and the regulatory capital ratio to assess this effect. Robustness tests in Section 7 further support the results, and Section 8 concludes.

2. Evidence from the Literature

Thompson (1998) is one of the first authors to include environmental factors in the risk analysis of banks. He examines the composition of assets of six major banks headquartered in the United Kingdom and assigns risk weights depending on the inclusion of environmentally critical industries. His approach is conceptual, with simplified assumptions about the risk characteristics of industries and bank portfolios. In line with this, Klomp (2014) investigates the association between natural disasters and bank stability. His approach focuses on a country's banking system as a whole and on the system's aggregated z-score. Using data for 169 countries, he concludes that natural disasters increase the likelihood of bank default. Battiston et al. (2017) model the climate risk of the financial system as a whole. Their model is based on the assumption that climate risk affects the equity holdings of financial institutions in carbon risk-sensitive industries. They find that first-round effects manifest as losses in critical equity holdings, while second-round effects are driven by the connectivity of financial institutions that have been hurt in the first round.

Cortés and Strahan (2017) investigate the lending behavior of banks in the aftermath of a natural disaster. They ask how banks that operate in multiple local markets adjust their lending when credit demand in a particular local market increases after a natural disaster. Based on data for the mortgage lending of small banks in different counties of the United States, they find that these banks tend to cut loans in non-core connected markets and increase the securitization of mortgages. In a similar analysis of U.S. banks, Barth, Sun, and Zhang (2019) conclude that natural disasters incentivize institutions to attract more deposits in order to meet the higher loan demand, and that therefore they raise interest rates on both deposits and loans. Koetter, Noth, and Rehbein (2020) obtain comparable results when analyzing the lending adjustments of German banks with credit relationships to corporates affected by the 2013 flooding of the river Elbe. The authors find that after the flooding, banks lend more to disaster-hit firms (in the form of emergency lending) than to non-affected firms. In addition, banks source their lending primarily through local savings deposits rather than through wholesale funding.

After differentiating between affected and non-affected banks, Schüwer, Lambert, and Noth (2019) apply a similar approach to assess the adjustment strategies of U.S. banks following a catastrophic event. Using Hurricane Katrina as a case study, they examine how natural disasters affect a bank's lending, asset allocation, and capital ratios. The authors further distinguish between independent banks and banks affiliated within bank holding companies

(BHCs) and find evidence that suggests that independent banks increase their risk-based capital ratios. In another study in which they examine the impact of multiple disasters on banks in the United States, Noth and Schüwer (2018) focus on bank stability and bank performance. They analyze bank accounting ratios such as the return on assets, z-score, and equity to assets and find that disasters weaken both bank performance and stability.

Previous studies on the effects of natural disasters and bank solvency take different approaches. Brei, Mohan, and Strobl (2019) investigate the effects of hurricanes on the aggregate banking system of seven countries in the Caribbean, and Nguyen et al. (2020) focus on the affectedness of individual banks from natural disasters, particularly earthquakes and tsunamis, in seven East Asian countries. They find that the disasters hit bank liquidity via deposits, but do not observe negative effects on solvency. Nguyen et al. (2020) measure the default risk of banks using their z-score, and Brei, Mohan, and Strobl (2019) use in addition the Tier 1 capital ratio of the banking system.

Typically, changes in financial regulation are driven by past experiences and aim to address the vulnerabilities that these experiences have revealed in the financial and economic system.[1] However, it is questionable if this approach is sufficient to avoid future financial crises. Rather, a complete approach that also includes emerging risks is called for. The current solvency requirements should be extended to ensure that banks introduce factors in their capital reserve calculations that account for their susceptibility to the increasing likelihood and severity of natural disasters, particularly with respect to their lending, financing, and investment activities. Accordingly, risk-weighted assets should be adjusted while leaving the overall capital requirements at the same level (van Gelder and vander Stichele 2011). This approach is also propagated by Batten, Sowerbutts, and Tanaka (2016), who argue that weather-related natural disasters can trigger financial instability and may cause severe damages to the balance sheets of banks.

[1]For instance, the Basel III Accord was largely developed in response to the recent subprime mortgage crisis. In line with the accord, the European Union's Capital Requirements Directive (CRD) obliges banks to set aside a minimum percentage of their capital to cover any potential defaults on their loans and investments.

A recent report by the Cambridge Institute for Sustainability Leadership (CISL) recommends that the Basel Committee for Banking Supervision should explicitly acknowledge environmental risk and its increasing impact on the stability of the financial system (CISL 2016). The report encourages regulators and banking institutions to adopt new practices to address environmental issues and incorporate a forward-looking approach to ensure the sustainability of bank lending activities. From a more comprehensive perspective, Aiyar, Calomiris, and Wieladek (2015) argue that credit instruments that are not subject to capital regulation or that constitute no risk weights will cause undesirable negative effects for the credit supply of banks. Credit risks from natural disasters as a more recent phenomenon might not yet be considered adequately in risk-weighted assets or regulatory requirements.

In addition to studies related to banks and banking regulations, recent work has investigated the effect of disasters on other types of institutions as well as on the financial value of investments. These studies emphasize potential channels of disaster risk transmission and frequently call for novel methodological approaches. Building on the new climate-economy literature, Balvers, Du, and Zhao (2017) posit that temperature shocks restrict the growth of companies and impose a higher cost of equity. Based on the arbitrage pricing theory and a specification for expected temperature levels, they consider temperature shocks as a systematic risk factor and examine the loading of asset prices to the temperature risk factor. The loading is negative and equates to a higher cost of equity capital of approximately 0.22 percent. Another consequence of climate change is the rise of sea levels with further effects on the price of properties in coastal areas and their use as collateral. Bernstein, Gustafson, and Lewis (2018) categorize properties into buckets of similar size, elevation, and zip code, yet with a different exposure to sea-level rise. They find that properties exposed to sea-level rise trade at a discount of 6.6 percent compared with those that are not exposed.

This study contributes to the literature by assessing whether and by how much bank solvency is affected by natural disasters. We provide a cross-country analysis based on individual banks worldwide and thus complement existing studies examining fewer countries (Brei, Mohan, and Strobl 2019; Nguyen et al. 2020) or focusing on aggregate systemwide evidence (Klomp 2014). Specifically, we investigate how different characteristics, business models, and locations of banks affect their solvency following a natural disaster. Moreover, we assess the suitability of two alternative measures of bank solvency to reflect banks' sensitivity to natural disasters and their interaction with bank characteristics and locations.

3. Hypothesis Development

Prior studies in this area discuss different measures of solvency and note that solvency can be expressed from a balance sheet perspective as a form of accounting equity or from a supervisory point of view as a more refined risk-based measure of regulatory capital (Flannery and Giacomini 2015; Hogan 2015). We thus employ two different types of bank capital ratios in our analysis: (i) the equity ratio (accounting capital ratio) and (ii) the Tier 1 capital ratio (regulatory capital ratio). Accounting equity comprises all balance sheet components of a bank's proprietary capital including both common equity and preferred equity (Cohen and Scatigna 2016). It can be interpreted as an institution's risk-bearing capacity based on standard accounting principles. In contrast, the Tier 1 capital takes a regulatory and specific risk-based point of view, with Tier 1 capital generally defined as high-quality equity capital (Basel Committee on Banking Supervision 2011, 2017).

In order to obtain numbers that are comparable across banks and years, we standardize the different types of capital. We use the volume of total assets (TA) to standardize the volume-based accounting equity (equity ratio), and the risk-weighted assets (RWA) to standardize Tier 1 capital as risk-adjusted capital (Tier 1 capital ratio). Risk-weighted assets are based on the Basel II regulation that in essence have also been retained in the Basel III Accord (Dermine 2015). Risk-weighted assets do not comprise all balance sheet assets of a bank, and the weight of included assets may be below 100 percent or even zero.

Our first hypothesis is in line with the general assumption that natural disasters have a negative impact on customers and bank operations and may thus cause losses (Benson and Clay 2004; Brei, Mohan, and Strobl 2019; Nguyen et al. 2020). Negative effects relate to the assets and/or counterparties of banks and to the banks' infrastructure. As natural disasters are a class of emerging risks, it is likely that banks have not yet priced them or built reserves. Specifically, we postulate the following:

HYPOTHESIS 1. *Natural disasters negatively affect the solvency of banks, measured via either the equity ratio (Hypothesis H_{1A}) or the Tier 1 capital ratio (Hypothesis H_{1B}).*

Because the two capital ratios are standardized using a different denominator, we can test the behavior of the simple volume-weighted equity ratio with respect to disasters and compare it with the behavior of the risk-weighted Tier 1 capital ratio. On one hand, risk weights are calibrated depending on the type of assets and/or counterparty and considered more adequate for supervisory risk assessment; however, they may be more complex and less robust on the other hand (Dermine 2015; Hogan 2015). Moreover, because Tier 1 capital is generally understood to be a more refined measure of a bank's capitalization, we further propose the following:

HYPOTHESIS 2. *The regulatory capital ratio is more sensitive to disaster risk than the accounting capital ratio.*

There are also arguments in disfavor of this hypothesis. As regulators usually align the risk weights of assets based on experiences, they may not fully reflect the impact from emerging risks such as natural disasters and contribute to an "ill-defined concept of bank capital ratios" (Aiyar, Calomiris, and Wieladek 2015, p. 976). Bischof, Laux, and Leuz (2020) argue that the Tier 1 capital ratio is a license to operate, and banks manage it actively to keep it at a stable level. The authors further make the point that based on prudential filters regulators may add back some losses (e.g., unrealized fair value losses) in the calculation of regulatory capital. As risk weights of top-rated companies and countries are very low and often zero, this further obstructs the adaptability and sensitivity of the Tier 1 ratio.

We further assume that the magnitude of effects on solvency depends on the characteristics and locations of individual banks. Particularly, the business model of banks and their size may affect the damage they are exposed to from natural disasters. Our third hypothesis, which we also test with respect to the accounting and regulatory capital ratio, therefore reads as follows:

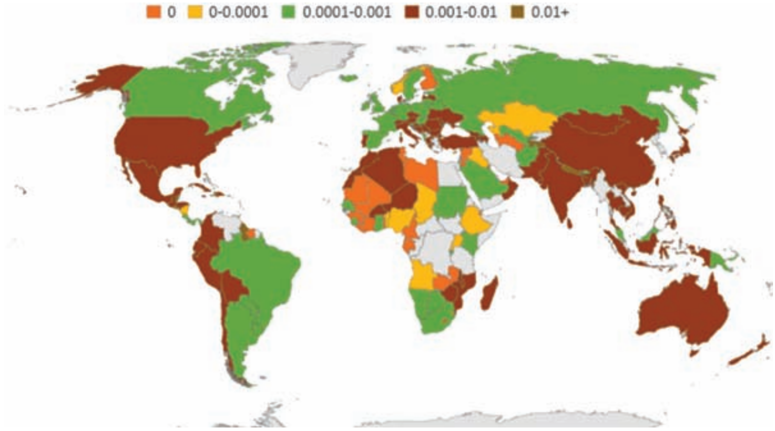
HYPOTHESIS 3. *Disasters affect banks differently depending on the individual banks' characteristics.*

4. Data and Data Preparation

This study uses data from the Emergency Events Database (EM-DAT) and a merged data set of banks' financial statements from Bankscope and Fitch. EM-DAT is provided by the Centre for Research on the Epidemiology of Disasters (CRED) at the University of Leuven, and contains detailed data on damages and other relevant information about various types of catastrophes around the globe. The data are collected from a variety of public and private sources, and since 2000, the center has enhanced the data by geocoding each disaster (CRED 2016). Natural (non-technological) disasters include critical meteorological (e.g., droughts, floods, storms) and geophysical events (e.g., earthquakes, volcanic eruptions). Figure 1 provides an overview of the average annual damages per country caused by recorded disasters during the period 2000–17. The different shades refer to the weighted damage ratio of each country, i.e., the ratio of the total annual damages in a given country to the country's gross domestic product (GDP), averaged across our sample period. Figure 2 provides an overview of the damages caused by different types of disasters during our sample period. The proportion of damages attributable to the three main categories of disasters (earthquakes/tsunamis, floods, and strong winds) varies considerably over time and often depends on one or two "mega-disasters" that caused most of the damages during a given year. For instance, in 2005, Hurricane Katrina was responsible for a large proportion of the natural-disaster-related damages during that year, while in 2011 the earthquake leading to the Fukushima nuclear catastrophe represented a mega-disaster.

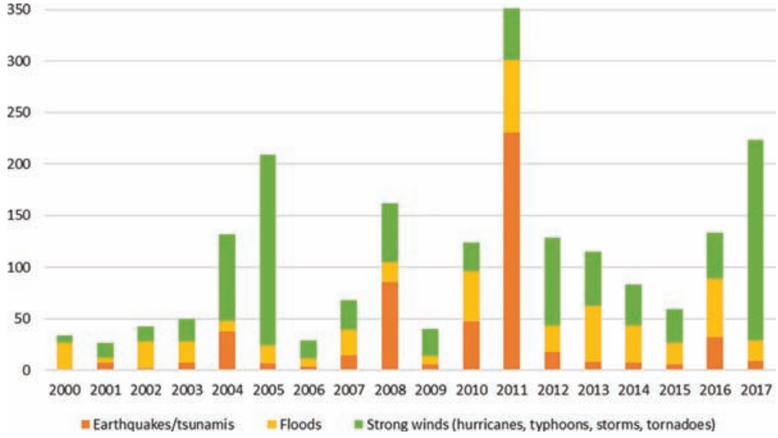
Bureau van Dijk (Bankscope) and Fitch Solutions (Fitch) provide detailed data on banks' accounting and financial statements. Bankscope includes extensive information for the years 2000 to 2014 yet limits the range of data offered thereafter. Therefore, we merge data from Bankscope through the year 2014 with information from Fitch for the years 2013 to 2017. When matching the two databases, we perform numerous checks to ensure the consistency of institutions and parameters included. A first issue is that the names of banks in Bankscope can differ from those in Fitch. In some cases, banks with similar names may be located in different countries, or banks can have several subsidiaries that are located in different cities in a given

Figure 1. Ratios of Disaster Damages/GDP: Country-Level Averages for the 18-Year Period (2000–17)



Note: Grey indicates that no data are available.

Figure 2. Distribution of Worldwide Disaster Damages across Different Types of Disasters, by Year (US$ billion)



country yet display the same name. Furthermore, for some variables, the way Bankscope records or calculates the data can be different from Fitch, and thus variables with the same name in Bankscope and Fitch are not always identical.[2]

In a next step, we employ the year 2013 data on total assets from both Bankscope and Fitch (i.e., the year in which the two databases overlap) and calculate the following variable which we then use to further compare the banks in each database:

$$ADTA = \left| \frac{BTA - FTA}{BTA} \right|, \tag{1}$$

where ADTA is the absolute difference in total assets between two banks, BTA is the value of total assets for a bank in Bankscope, and FTA is the value of total assets for a matched bank in Fitch.

The distribution of ADTA is shown in Appendix A. If the absolute difference in the value of total assets is smaller than 0.1 (10 percent), we consider the match to be authentic. In contrast, if the absolute difference exceeds 0.1, we drop the matched bank pair.

In addition to total assets, we check the consistency of other variables in Bankscope and Fitch. We again examine the year 2013 data for 2,895 banks in Fitch, and compare the variable values with those of their matched counterparts in Bankscope. We use two different methods for our comparison (see Appendix B). The first method is based on two correlation measures (the normal correlation and the correlation after trimming each variable at the 1 percent and 99 percent levels). The respective results are displayed in columns 1 and 2 of the table in Appendix B. The second method employs the absolute difference ratio, calculated in the same fashion as the absolute difference in total assets above. If the difference ratio is larger than 0.1 (i.e., a variable value in Fitch is 10 percent larger or smaller than in Bankscope), we assign a value of "1" (wrong matching); if not, we

[2]We use the Stata command "matchit" to fuzzy-match the bank names (Stata 2017). This command calculates similarity scores, ranging from 0 to 1, between every paired bank from Bankscope and Fitch. After matching the names, we ensure that the countries and cities provided as bank locations in the Bankscope database are exactly the same as the matched banks in Fitch. If the locations do not match, we delete the matched banks. Afterwards, we check the rest of the matched banks manually, to ensure that they are very likely to be the same banks.

assign a value of "0" (correct matching). In addition, if the value in Bankscope is 0 (making it impossible to be used as a denominator in our percentage difference calculation), then we assign a value of "0" if the value in Fitch is also 0 (correct matching); otherwise, we assign a value of "1" (wrong matching). The percentage of "1s" (i.e., the percentage of wrong matches) for each variable is shown in column 3 of the table in Appendix B. We mark the variables we use in this paper in bold and with grey shading. They exhibit good quality matching with a correlation higher than 0.99 and a percentage of difference ratio (at the 0.1 level) of less than 10 percent.

The Bankscope and Fitch databases include banks from around the world that file their financial statements in different currencies. In total, we have 9,928 banks in our sample with complete data on all variables. These banks are located across 149 different countries. Table 1 reports the geographical distribution.

Some authors suggest keeping data in the original currency and thus avoid translation effects (Cohen and Scatigna 2016). However, in order to achieve better comparability (e.g., in terms of size), we convert all non-US$ figures at the respective exchange rate at the end of the accounting period. For most of our variables, potential biases caused by exchange rate fluctuations are avoided, as we work with standardized data (e.g., capital in absolute terms divided by assets in absolute terms). Hence, any potential biases arising from currency fluctuations in the nominator and denominator should compensate each other.

Appendix C provides definitions for all variables used in our analysis, and Table 2 reports summary statistics for the variables. The number of observations of the Tier 1 capital ratio (124,997) is considerably smaller than that of the equity ratio (164,046). The discrepancy is due to the fact that banks have not always been obliged to publish regulatory capital ratios. It is worth noting that the Tier 1 capital ratio (Tier 1 capital divided by risk-weighted assets) has a median of 14.50 percent, much higher than the 6 percent required by Basel III.

5. Methodology

We assess a bank's sensitivity to risk based on a series of ordinary least squares (OLS) and quantile regression approaches. We employ alternative measures of disaster damages as our main independent

Table 1. Sample Distribution: The Number of Banks per Country

Seq. Country	N	Seq. Country	N	Seq. Country	N	Seq. Country	N
1 Afghanistan	5	39 Egypt	19	76 Liberia	1	113 Romania	17
2 Albania	12	40 El Salvador	9	77 Libya	2	114 Russian	54
3 Algeria	2	41 Estonia	8	78 Lithuania	6	115 Rwanda	6
4 Angola	4	42 Ethiopia	1	79 Luxembourg	27	116 Saint Kitts	1
5 Antigua	1	43 Finland	23	80 Macau	5	117 Saint Lucia	1
6 Argentina	1	44 France	33	81 Madagascar	1	118 Saudi Arabia	10
7 Armenia	11	45 Gabon	2	82 Malawi	6	119 Senegal	4
8 Australia	23	46 Gambia	2	83 Malaysia	43	120 Serbia	10
9 Austria	119	47 Georgia	11	84 Maldives	2	121 Seychelles	3
10 Azerbaijan	22	48 Germany	1,211	85 Mali	3	122 Sierra Leone	3
11 Bahamas	6	49 Ghana	20	86 Malta	4	123 Singapore	12
12 Bahrain	15	50 Greece	5	87 Mauritania	1	124 Slovakia	8
13 Bangladesh	38	51 Grenada	1	88 Mauritius	11	125 Slovenia	12
14 Barbados	1	52 Guatemala	2	89 Mexico	20	126 South Africa	12
15 Belarus	10	53 Guinea	1	90 Moldova	7	127 Spain	49
16 Belgium	13	54 Guyana	3	91 Mongolia	2	128 Sri Lanka	15
17 Benin	1	55 Haiti	1	92 Montenegro	6	129 Sudan	5
18 Bhutan	2	56 Honduras	1	93 Morocco	4	130 Suriname	3
19 Bolivia	6	57 Hong Kong	16	94 Mozambique	10	131 Sweden	66
20 Bosnia	16	58 Hungary	15	95 Namibia	7	132 Switzerland	79
21 Botswana	7	59 Iceland	2	96 Nepal	2	133 Tajikistan	2
22 Brazil	59	60 India	9	97 Netherlands	22	134 Thailand	21
23 Bulgaria	15	61 Indonesia	56	98 New Zealand	4	135 Togo	1
24 Burundi	1	62 Iraq	1	99 Nicaragua	4	136 Trinidad	4
25 Cambodia	15	63 Ireland	9	100 Niger	1	137 Tunisia	3
26 Canada	3	64 Israel	7	101 Nigeria	15	138 Turkey	18
27 Cape Verde	3	65 Italy	338	102 Norway	93	139 Uganda	13
28 Chile	18	66 Jamaica	5	103 Oman	9	140 Ukraine	6
29 China	104	67 Japan	231	104 Pakistan	21	141 Emirates	22
30 Congo	2	68 Jordan	14	105 Panama	18	142 United Kingdom	95
31 Costa Rica	11	69 Kazakhstan	24	106 New Guinea	2	143 United States	6,010
32 Croatia	24	70 Kenya	24	107 Paraguay	2	144 Uruguay	14
33 Cyprus	8	71 South Korea	7	108 Peru	3	145 Vanuatu	2
34 Denmark	53	72 Kuwait	3	109 Philippines	11	146 Venezuela	14
35 Djibouti	2	73 Latvia	14	110 Poland	23	147 Vietnam	8
36 Dominica	1	74 Lebanon	19	111 Portugal	77	148 Yemen	5
37 Dominican Rep.	24	75 Lesotho	2	112 Qatar	10	149 Zambia	13
38 Ecuador	15					Total	9,928

Note: This table reports the distribution of our sample of 9,928 banks across 149 countries. We list countries alphabetically and report the number of banks (N) for each country.

Table 2. Summary Statistics

Variable Name	No. of Observations	Mean	Std. Dev.	Median	5th Percentile	95th Percentile
Equity Ratio	164,046	0.1166	0.0945	0.0973	0.0448	0.2253
Tier 1 Capital Ratio	124,997	0.1812	0.1359	0.1450	0.0914	0.3688
Damage Ratio (60 Days)	194,186	0.0023	0.0078	0.0009	0.0000	0.0072
Damage Ratio (180 Days)	194,186	0.0022	0.0068	0.0010	0.0000	0.0097
Total Assets (in US$ billion)	164,056	3.2179	11.3531	0.2753	0.0295	14.9945
Net Loans Ratio	163,168	0.5978	0.1853	0.6258	0.2363	0.8522
Customer Deposits Ratio	161,893	0.7581	0.1744	0.8127	0.3893	0.9105
Net Income to Equity Ratio	163,865	0.0730	0.0892	0.0705	–0.0467	0.2075
Real GDP Growth Rate	194,157	0.0204	0.0675	0.0227	–0.1035	0.1260
Growth Rate of Credit to Private Sector	172,545	0.0128	0.0619	0.0097	–0.0886	0.0920
Real GDP per Capita (in US$ Thousands)	194,186	32.0650	14.0286	37.0949	0.7586	42.0992

Note: This table provides summary statistics for our sample. For each variable, we report the number of bank-year observations, together with the mean, standard deviation, median, 5th percentile, and 95th percentile. The number of bank-year observations varies due to missing data for some banks.

variables and different specifications of bank solvency as our dependent variables. A major challenge in our analysis is to relate the two types of variables in a meaningful way. For instance, the EM-DAT database we use to assess damages from natural disasters reports disasters over periods of different length, e.g., single-day tornados or blizzards versus longer periods for floods and droughts. In addition, the impact of disasters on banks may be immediate, e.g., if they expose banks to operational risks, or long-term if disasters first affect the banks' customers and then gradually transform into credit risks.

To address these issues, we follow Klomp (2014) and design our main explanatory variable of interest (*Damage Ratio*) as follows: We assume that all banks in one country experience the same repercussions from a given disaster, and further that the impact of the disaster fully materializes and affects banks within one single year or two consecutive years.[3] For example, we assume that the shortest period during which a given disaster occurs and affects a bank is two months (60 days). In addition, we assume that disaster j affects country i approximately m days before the end of year t, and that the total number of disasters that occur in year t for country i is n. The proportion of damages attributed to year t ($damage_{ijt}$) and year $t+1$ ($damage_{ij(t+1)}$) is thus calculated as follows:

$$\textit{If } m \geq 60\text{: } damage_{ijt} = total\ damage\ of\ disaster\ j\ in\ country\ i$$

$$Otherwise\,(m < 60): \begin{cases} damage_{ijt} = \left(\frac{m}{60}\right) * total\ damage\ of \\ \quad disaster\ j\ in\ country\ i \\ damage_{ij(t+1)} = \left(\frac{60-m}{60}\right) * total\ damage\ of \\ \quad disaster\ j\ in\ country\ i \end{cases} \tag{2}$$

$$DamageRatio_{it} = \left(\sum_{j=1}^{n} damage_{ijt}\right) / GDP_{it}.$$

[3] Klomp (2014) also allocates disaster damages to two different years. However, he only uses large-scale disasters and equally assigns 50 percent of the damage to the disaster year and the subsequent year. In contrast, we include all disasters listed in the EM-DAT database and divide the damages resulting from each disaster into two years based on the specific timing of the disaster during a given year.

To account for the different time patterns that characterize both disasters themselves and their effects, we consider periods of varying length during which damages may materialize. Specifically, in addition to the aforementioned 60 days, we also assume that damages manifest within 90 days and 180 days after the beginning of the disaster. Because the results for the different periods are very similar, we only report the results for an impact period of 60 and 180 days, and consider other periods as part of our robustness tests.

Following the prior literature on bank capitalization, we control for several characteristics of banks: Size, measured as the natural logarithm of total assets (Barrios and Blanco 2003; Brewer, Kaufman, and Wall 2008; Schepens 2016), the loan ratio, measured as net loans over total assets (Altunbas et al. 2007; Demirgüç-Kunt, Detragiache, and Merrouche 2013; Schepens 2016), profitability, measured as the ratio of net income over equity (Brewer, Kaufman, and Wall 2008; Schaeck and Cihák 2012), and the deposit level, measured as the ratio of total customer deposits over total assets (Barrios and Blanco 2003; Demirgüç-Kunt, Detragiache, and Merrouche 2013).

Furthermore, because country-specific variables can affect each nation's banking system, we include several country-levels controls that have been used in previous research in this area. These include the level of national development, measured as the natural logarithm of a country's annual real GDP per capita; economic growth, measured as the annual growth in the real GDP; and the credit activity of a country, measured as the growth of credit to the private sector. We also examine other country-specific control variables such as the world government index (Kaufmann, Kraay, and Mastruzzi 2011), a country's trade balance, and changes in each country's exchange rate. The resulting models either suffer from multicollinearity problems or are associated with large reductions in our sample size due to missing values. We thus decided not to report the respective regressions here. However, even with these variables included, the results remain similar.

Our resultant regression model can be written as follows:

$$\Delta ratio_{kit} = \mu * ratio_{kit-1} + \beta * \text{ DR}_{it} + \alpha_m * B^s_{kit} + \gamma_h * C^h_{it} + \theta_t + \varphi_i + \delta_{kit} + \omega_{kit} + \varepsilon_{kit} \quad (3)$$

and

$$\Delta ratio_{kit} = ratio_{kit} - ratio_{kit-1}, \tag{4}$$

where $ratio_{kit}$ represents the equity ratio or Tier 1 capital ratio for bank k in country i in year t, and $ratio_{kit-1}$ is the corresponding ratio in the preceding year. DR_{it} is our explanatory variable of interest (in this case the weighted damage ratio during the 60 days (or 180 days) following a disaster). B^s_{kit} is a vector of s bank-specific control variables, and C^h_{it} is a vector of h country-specific control variables. θ_t represents time fixed effects, and φ_i the country fixed effects. δ_{kit} are the accounting standard fixed effects, and ω_{kit} are the bank specialization fixed effects.

6. Results

Before commencing with our multivariate analysis, we first examine the Pearson correlation coefficients for all variable pairs in Table 3. All correlations—except for two—between the variables are well below 0.5. Exceptions include the correlation between the lagged Tier 1 capital ratio and the lagged equity ratio (0.8206), where a high correlation is expected. However, the two variables are never employed in the same model, thus mitigating any multicollinearity concerns in our multivariate analysis. Similarly, and again as expected, the damage ratio (60 days) and the damage ratio (180 days) exhibit a high correlation (0.9425). The two variables are used as alternative damage proxies and thus never coexist in one model, again mitigating any multicollinearity concerns.

We next commence our multivariate analysis by examining how banks' solvency ratios are affected by natural disasters (Hypotheses H_{1A} and H_{1B}). In addition, we explore whether the relationship is different when employing the Tier 1 capital ratio, instead of the equity ratio, as a dependent variable (Hypothesis H_2). Because the sensitivity to natural hazards is unlikely to be uniform across institutions, we differentiate between banks located in countries with different land masses as well as between different types of banks (based on their business model) as well as different ex ante capitalization levels of banks.

Table 3. Correlation Matrix

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
(1) Change in Equity Ratio	1											
(2) Lagged Equity Ratio	–0.3068*	1										
(3) Lagged Tier 1 Capital Ratio	–0.2712*	**0.8206***	1									
(4) Damage Ratio (60 Days)	–0.0036	0.0050*	0.0111*	1								
(5) Damage Ratio (180 Days)	–0.0046*	0.0057*	0.0094*	**0.9425***	1							
(6) Log (Total Assets)	0.0414*	–0.2452*	–0.2774*	–0.0328*	–0.0352*	1						
(7) Net Loans Ratio	–0.0045*	–0.2795*	–0.4838*	0.0062*	0.0056*	0.0304*	1					
(8) Customer Deposits Ratio	–0.0138*	–0.3933*	–0.3165*	0.0468*	0.0498*	–0.2616*	0.1929*	1				
(9) Lagged Net Income to Equity Ratio	0.0999*	–0.0505*	–0.1180*	0.0266*	0.0295*	0.0508*	0.0209*	0.0237*	1			
(10) Real GDP Growth Rate	–0.0072*	–0.0304*	0.0203*	0.0272*	0.0289*	0.0019	0.0153*	0.0264*	0.0464*	1		
(11) Growth Rate of Credit to Private Sector	–0.0340*	0.0851*	0.0315*	–0.0163*	0.0074*	–0.0339*	–0.0020	–0.0590*	0.1581*	0.0955*	1	
(12) Log (Real GDP per capita)	0.0264*	–0.1599*	–0.0262*	–0.0056*	–0.0111*	–0.1318*	0.1866*	0.2486*	–0.1112*	0.0737*	–0.2081*	1

Note: This table presents the Pearson correlation coefficients between the dependent variables (the annual change in the equity ratio and the annual change in the Tier 1 capital ratio) and the explanatory variables. Although the lagged Tier 1 capital ratio is highly correlated (**0.8206**) with the lagged equity ratio, they do not coexist in any model. Similarly, the damage ratio (60 days) and the damage ratio (180 days) exhibit a high correlation (**0.9425**), but are not jointly used in any model. * indicates significance at the 10 percent level.

6.1 The Sensitivity of Banks' Equity Capital to Natural Disasters

Table 4 provides our regression results for Hypothesis H_1. To ensure the robustness of our results, we perform separate regressions for our full (worldwide) sample of banks, banks in the United States (US only), and banks in other countries (non-US). Columns 1 to 4 of Table 4 show how the weighted 60-day damage ratio affects the equity ratio ($\Delta E/TA$) for the three geographical subsamples (with column 4 repeating the full-sample analysis of column 3, but employing a non-winsorized sample). The coefficients for the damage ratio are consistently negative and significant, suggesting that natural disasters indeed have a detrimental effect on banks' capital ratios. Columns 5 to 8 of Table 4 employ the same model specifications as those employed in columns 1 to 4, but use the weighted damage ratio measured over a period of 180 days as the main variable of interest. The results are qualitatively and quantitatively very similar to those in the first four columns.

There are likely several reasons why natural disasters affect a bank's capital ratio. One explanation is that while banks protect their lending activities by requiring assets as collateral, the occurrence of natural disasters may destroy or at least reduce the value of the assets in question, hence reducing the bank's capacity to recover the outstanding loan balance via its collateral. Accordingly, if a borrower defaults on his/her loan and the bank manager realizes that the bank cannot recover the borrowed money through the collateral, the bank has to write off the borrowed amount from its books and, by extension, the bank equity. Consequently, losing collateral as a result of a natural disaster is likely the main channel through which natural disasters affect a bank's equity. Furthermore, disasters may affect banks directly—for instance, by damaging a bank's offices or its technical infrastructure. In summary, there is a multitude of reasons why banks that lend in high-risk areas should prepare for and create reserves to protect themselves against natural disasters and prevent any associated deterioration in their capital ratios.[4]

[4]It is worth noting here that higher capital requirements (e.g., those mandated by Basel III) have been shown to increase banks' lending rates and, consequently, have been blamed for the comparatively slow economic recovery following the

Table 4. The Effect of Natural Disasters on Banks' Equity Ratios, Employing Damage Ratios Calculated Over 60 and 180 Days

	Damage Period: 60 Days				Damage Period: 180 Days			
	(1) US Only	(2) Non-US	(3) Full Sample	(4) Full Sample	(5) US Only	(6) Non-US	(7) Full Sample	(8) Full Sample
Dependent Variable: ΔE/TA	**ΔE/TA Winsorized**			**ΔE/TA Not Winsorized**	**ΔE/TA Winsorized**			**ΔE/TA Not Winsorized**
Damage Ratio	–0.359***	–0.012**	–0.016**	–0.018**	–0.522***	–0.013*	–0.019***	–0.022**
	(0.000)	(0.037)	(0.011)	(0.016)	(0.000)	(0.051)	(0.007)	(0.011)
Lagged Equity Ratio	–0.304***	–0.187***	–0.222***	–0.237***	–0.304***	–0.187***	–0.222***	–0.237***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Log (Total Assets)	–0.003***	–0.003***	–0.003***	–0.003***	–0.003***	–0.003***	–0.003***	–0.003***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.024***	–0.006***	–0.015***	–0.023***	–0.024***	–0.006***	–0.015***	–0.023***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.112***	–0.032***	–0.059***	–0.083***	–0.112***	–0.032***	–0.059***	–0.083***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.029***	0.025***	0.036***	0.001	0.029***	0.025***	0.036***	0.001
	(0.000)	(0.000)	(0.000)	(0.500)	(0.000)	(0.000)	(0.000)	(0.500)
Real GDP Growth Rate	0.068***	0.001	–0.004**	–0.002	0.063***	0.001	–0.004**	–0.002
	(0.000)	(0.714)	(0.047)	(0.390)	(0.000)	(0.751)	(0.042)	(0.395)
Growth Rate of Credit to Private Sector	–0.012***	–0.018***	–0.016***	0.000	–0.008***	–0.019***	–0.017***	0.000
	(0.000)	(0.000)	(0.000)	(0.453)	(0.000)	(0.000)	(0.000)	(0.452)
Log (Real GDP per capita)	0.048***	0.003**	0.003**	0.003**	0.050***	0.003**	0.003**	0.003**
	(0.000)	(0.016)	(0.022)	(0.040)	(0.000)	(0.017)	(0.024)	(0.041)
Country FE		Yes	Yes	Yes		Yes	Yes	Yes
Year FE		Yes	Yes	Yes		Yes	Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Accounting standard FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	88,918	53,145	142,063	142,063	88,918	53,145	142,063	142,063
Adjusted R^2	0.316	0.154	0.217	0.215	0.316	0.154	0.217	0.215

Note: This table presents the results for a series of models in which we regress the equity ratio of banks on the weighted damage ratio and other control variables over the 2000–17 period for the 142,063 firm-year observations in our sample for which data on the equity ratio are available. The dependent variable in columns 1 to 3 and 5 to 7 is the change in the equity ratio (winsorized). The dependent variable in columns 4 and 8 is the change in the equity ratio (not winsorized). In columns 1 to 4, the independent variable of interest is the damage ratio calculated over a period of 60 days; in columns 5 to 8, the independent variable of interest is the damage ratio calculated over a period of 180 days. Columns 1 and 5 report results for the U.S.-only subsample; columns 2 and 6 report results for the non-U.S. subsample; and columns 3, 4, 7, and 8 report results for the full sample. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

When examining the other explanatory variables, we observe that bank size (measured by the *natural log of total assets*) negatively correlates with the bank equity ratio, which is in line with prior research on bank solvency (Barrios and Blanco 2003; Altunbas et al. 2007; Schaeck and Cihák 2012; Schepens 2016). Similarly, and also in line with the extant literature, we observe that profitability (measured by the *lagged net income to equity ratio*) is positively related to the equity ratio (Brewer, Kaufman, and Wall 2008; Schaeck and Cihák 2012; Panier, Pérez-Gonzalez, and Villanueva 2013; Berger, Öztekin, and Roman 2018); and that the *net loan ratio* (net loans/total assets) is, generally, negatively correlated with the equity ratio (Altunbas et al. 2007; Schepens 2016).

The prior banking literature exhibits mixed evidence regarding the effect of disruptions on the equity ratio of banks. Studies on financial crises (De Jonghe and Öztekin 2015, and Gambacorta and Shin 2018) suggest that the equity ratio of banks is procyclical: when a financial crisis hits the market, the equity ratio of banks increases (likely due to capital injections). Similarly, Koetter, Noth, and Rehbein (2020) and Bos, Li, and Sanders (2018) argue that capital adequacy (as proxied by the equity ratio) and lending (in the form of total outstanding loans) increase after large-scale natural disasters. In contrast, Noth and Schüwer (2018) find evidence that suggests that U.S. banks that engage in mortgage lending experience a decline in bank capital following a natural disaster. Klomp (2014) shows that banks' default risk increases (and the equity ratio decreases) in the years following a large natural disaster. Brei, Mohan, and Strobl (2019) analyze a sample of seven countries in the Caribbean and find that banks experience changes in funding and lending after hurricanes, yet they do not detect any effects on risk and equity. Nguyen et al. (2020) confirm this result for banks operating in East Asia. Our results complement this research.

2008/09 financial crisis and a reduction in global GDP growth, estimated at approximately 0.3 percent per year. A well-measured response to climate change with appropriately defined natural disaster prone risk weightings for banks' assets is therefore called for.

6.2 The Sensitivity of Banks' Tier 1 Capital to Natural Disasters

In order to compare the sensitivity of our two solvency measures, we reestimate the same regressions we employed in Table 4 with the *Tier 1 capital ratio* as the independent variable. We thus address our hypothesis (H_2) that suggests that regulatory capital ratios more distinctly reflect changes in risk than accounting based measures of capital. Columns 1 to 3 (4 to 6) of Table 5 show how the weighted 60-day (180-day) damage ratio affects the Tier 1 capital ratio of banks in our three geographical subsamples (U.S. banks, non-U.S. banks, and the full sample). Except for the United States, the coefficients are not significant and not always negative, suggesting that natural disasters have a smaller effect on regulatory capital ratios than they have on the accounting-based equity ratios we examined in Table 4.

For the subsample of U.S. banks, the coefficients for the damage ratio in our accounting equity analyses (Table 4) are considerably larger than those in our regulatory capital regressions (Table 5). We also find that, in general, disasters have a larger impact on the equity ratio of U.S. banks than on the equity ratio of non-U.S. banks. This is likely driven by the fact that since about the 1980s, the damages caused by disasters in the United States increased considerably more than those in other countries. For instance, in 2017, the United States accounted for 83 percent of damages from storms worldwide (Munich Re 2018, p. 52; World Economic Forum 2018, p. 12).

Contrary to Hypothesis H_2, we note that disasters do not have a large impact on the Tier 1 capital ratio. If anything, our results show that, in comparison with the equity ratio, the Tier 1 capital ratio is less significantly and uniformly influenced by natural disasters. There are several possible reasons: first, regulations may force banks to keep the required amount of Tier 1 capital at a specific and constant level; second, in order to protect against failure, bank management will, by itself, have an incentive to keep the Tier 1 capital ratio at a safe level (Abou-El-Sood 2015; Bischof, Laux, and Leuz 2020); third, the denominator of the Tier 1 capital ratio (a bank's risk-weighted assets), does not sufficiently take natural disaster risk into account, causing regulatory weightings to remain largely unaffected by disasters. Our lack of support for Hypothesis H_2 is in line with prior research findings in this area. For instance,

Table 5. The Effect of Natural Disasters on Banks' Tier 1 Capital Ratios, Employing Damage Ratios Calculated Over 60 and 180 Days

	Damage Period: 60 Days			Damage Period: 180 Days		
Dependent Variable: ΔT1R/TA (Winsorized)	**(1) US Only**	**(2) Non-US**	**(3) Full Sample**	**(4) US Only**	**(5) Non-US**	**(6) Full Sample**
Damage Ratio	–0.292***	0.003	–0.019	–0.407***	–0.003	–0.024
	(0.000)	(0.935)	(0.625)	(0.000)	(0.954)	(0.582)
Lagged Tier 1 Capital Ratio	–0.217***	–0.183***	–0.192***	–0.218***	–0.183***	–0.192***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Log (Total Assets)	–0.004***	–0.003***	–0.004***	–0.004***	–0.003***	–0.004***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.089***	–0.044***	–0.074***	–0.089***	–0.044***	–0.074***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.121***	–0.024***	–0.086***	–0.121***	–0.024***	–0.086***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.041***	0.028***	0.054***	0.041***	0.028***	0.054***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Real GDP Growth Rate	0.083***	0.009	–0.005	0.078***	0.009	–0.005
	(0.000)	(0.181)	(0.287)	(0.000)	(0.180)	(0.285)
Growth Rate of Credit to Private Sector	–0.016***	–0.048***	–0.035***	–0.013***	–0.048***	–0.035***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Log (Real GDP per capita)	0.064***	–0.004	–0.009***	0.066***	–0.004	–0.009***
	(0.000)	(0.177)	(0.000)	(0.000)	(0.175)	(0.000)
Country FE		Yes	Yes		Yes	Yes
Year FE		Yes	Yes		Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes	Yes	Yes
Accounting standard FE	Yes	Yes	Yes	Yes	Yes	Yes
N	86,231	21,601	107,832	86,231	21,601	107,832
Adjusted R^2	0.238	0.265	0.177	0.238	0.265	0.177

Note: This table presents the results for a series of models in which we regress the Tier 1 capital ratio of banks on the weighted damage ratio and other control variables over the 2000–17 period for the 107,832 firm-year observations in our sample for which data on the Tier 1 capital ratio are available. The dependent variable is the change in the Tier 1 capital ratio (winsorized). In columns 1 to 3, the independent variable of interest is the damage ratio calculated over a period of 60 days; in columns 4 to 6, the independent variable of interest is the damage ratio calculated over a period of 180 days. Columns 1 and 4 report results for the U.S.-only subsample; columns 2 and 5 report results for the non-U.S. subsample; columns 3 and 6 report results for the full sample. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

Schüwer, Lambert, and Noth (2019) document that the regulatory capital ratio increases (rather than decreases) after a disaster.[5]

We conclude that we cannot find clear evidence of a higher risk sensitivity of the Tier 1 capital ratio. Rather, the equity ratio appears to be a more appropriate measure of natural disaster risk and should be considered for regulatory purposes. In addition, as noted above, a revised risk-weighting of assets that does not only take historical credit and liquidity into consideration (as per Basel III), but weighs assets based on their expected susceptibility to natural disasters, may lead to a better inclusion of natural disaster risks in banking regulations. Similar results should be achieved from a fairer risk-weighting of assets that takes the geographical lending habits (and thus the proneness to natural disasters) of a given bank into consideration.

With respect to our other explanatory variables, we observe that—in line with previous research in this area (e.g., Brewer, Kaufman, and Wall 2008)—bank size negatively correlates with the Tier 1 capital ratio, and that profitability (the *lagged net income to equity ratio*) positively relates to the equity ratio.

6.3 Bank Solvency Strategies to Prepare for Natural Disasters (ex ante Tests)

It is plausible that banks may anticipate natural disasters and respond in advance. To address this possibility, we perform a series of tests in which we include the forward damage ratio (damage ratio one year ahead) among our explanatory variables. Assuming that banks can correctly predict upcoming challenges from natural disasters, they should be able prepare themselves by increasing their equity and raising their risk premiums to build reserves. However, our full-sample results in Table 6 provide little evidence for this conjecture and suggest no significant change in banks' equity ratio (column 3), Tier 1 capital ratio (column 6), or net interest margin (column 9) in the year preceding a given disaster. However, banks in the United States (columns 1, 4, and 7) appear to be more forward-looking and show signs of strengthening their balance

[5]The authors show that higher risk-based capital ratios are the result of banks prioritizing lower risk-weighted assets such as government securities.

Table 6. Ex Ante Tests

	Dependent Variable: ΔE/TA (Winsorized)			Dependent Variable: ΔT1R/TA (Winsorized)			Dependent Variable: ΔNIM (Winsorized)		
	(1) US Only	(2) Non-US	(3) Full Sample	(4) US Only	(5) Non-US	(6) Full Sample	(7) US Only	(8) Non-US	(9) Full Sample
Lagged Equity Ratio	–0.305***	–0.195***	–0.228***						
	(0.000)	(0.000)	(0.000)						
Lagged Tier 1 Capital Ratio				–0.257***	–0.226***	–0.235***			
				(0.000)	(0.000)	(0.000)			
Lagged Net Interest Margin							–0.032	–0.012**	–0.016***
							(0.182)	(0.010)	(0.003)
Forward 1 Damage Ratio	0.038*	0.006	0.006	0.210***	0.069	0.043	0.090***	–0.010	–0.009
	(0.084)	(0.335)	(0.352)	(0.000)	(0.326)	(0.482)	(0.000)	(0.175)	(0.228)
Log (Total Assets)	–0.003***	–0.003***	–0.003***	–0.004***	–0.004***	–0.004***	–0.000***	–0.000	–0.000***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.280)	(0.000)
Net Loans Ratio	–0.024***	–0.006***	–0.016***	–0.091***	–0.047***	–0.078***	0.004***	0.002***	0.003***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.112***	–0.031***	–0.060***	–0.091***	–0.023***	–0.071***	–0.002***	0.000	–0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.559)	(0.148)
Lagged Net Income to Equity Ratio	0.029***	0.025***	0.036***	0.040***	0.030***	0.051***	–0.010***	–0.006***	–0.009***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Real GDP Growth Rate	0.045***	–0.000	–0.006***	0.055***	0.011	–0.005	–0.002	–0.003**	–0.002***
	(0.000)	(0.933)	(0.009)	(0.000)	(0.138)	(0.317)	(0.252)	(0.015)	(0.000)
Growth Rate of Credit to Private Sector	–0.008***	–0.017***	–0.015***	–0.013***	–0.048***	–0.033***	–0.001**	0.001	0.002**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.026)	(0.286)	(0.019)
Log (Real GDP per capita)	0.047***	0.004***	0.003***	0.061***	–0.003	–0.008***	–0.016***	–0.000	0.000
	(0.000)	(0.008)	(0.007)	(0.000)	(0.375)	(0.009)	(0.000)	(0.615)	(0.874)
Country FE		Yes	Yes		Yes	Yes		Yes	Yes
Year FE		Yes	Yes		Yes	Yes		Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Accounting standard FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	88,918	53,145	142,063	86,231	21,601	107,832	87,636	48,767	136,403
Adjusted R^2	0.317	0.162	0.224	0.254	0.192	0.240	0.065	0.038	0.075

Note: This table presents the results for a series of models in which we regress changes in the equity ratio, Tier 1 capital ratio, and net interest margin of banks on the forward damage ratio and other control variables over the 2000–17 period for different subsamples of our data set. The dependent variable in columns 1 to 3 is the change in the equity ratio (winsorized). The dependent variable in columns 4 to 6 is the change in the Tier 1 capital ratio (winsorized). The dependent variable in columns 7 to 9 is the change in the net interest margin (winsorized). Columns 1, 4, and 7 report results for the U.S.-only subsample; columns 2, 5, and 8 report results for the non-U.S. subsample; and columns 3, 6, and 9 report results for the full sample. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

sheet by increasing (reducing) equity (debt), injecting liquidity, and expanding their profit margin.

These results have several implications. First, the forward damage ratio significantly affects the asset structure of U.S. banks, which indicates the quality of their prediction and the actions taken to prepare for natural disasters. Second, as can be observed in the aftermath of natural disasters, some governments tend to adopt relatively loose credit policies to support post-disaster reconstruction, which may lead to an increase in banks' bad debt thereafter. However, U.S. banks can effectively mitigate the impact of natural disasters on loan quality and credit risk by raising the lending rate in advance and increasing cash reserves, thereby alleviating panic in the capital market and reducing market risks.

6.4 The Influence of Bank Characteristics

The results up to now provide evidence for our full sample of banks. However, banks around the globe operate under different conditions, pursue divergent business models, and are subject to differing types of disasters as well as variations in country-level factors characterizing each country's legal environment, economic development, and banking regulations. To address these issues, we perform a series of robustness tests in which we examine whether our results hold for different subsamples of our data based on the characteristics of both banks and/or the countries they operate.

6.4.1 Business Models

Banks vary considerably with respect to the way they conduct their business, and it is important to explore whether a bank's business model affects its susceptibility to adverse consequences from a natural disaster. We therefore investigate if the risk sensitivity of banks to catastrophic events depends on their respective business models, i.e., their strategy towards customers, products, and regions, and the associated diversification potential. The assumption is that more diversified institutions (whose lending and investment portfolio includes claims with low correlations) are better able to absorb and deal with large damages than undiversified banks. In this respect, damages from disasters may be considered a specific class of risk that allows for diversification effects.

Our analysis focuses on the three predominant business models, namely bank holding companies (BHCs), commercial banks, and savings banks. A bank holding company typically operates across multiple regions and product markets through the participation in different entities. As a result, the potential for geographical diversification is generally higher for BHCs than for commercial banks that also have a broad product portfolio yet display a smaller network of national and international branches. Savings banks operate under a third type of business model. Their lending portfolio is often regionally focused and they tend to be smaller, increasing their exposure to local disasters. In summary, we expect more diversified (and less concentrated) banks such as BHCs to be less affected by disasters than commercial banks and, in particular, savings banks.

Although our results are not fully as expected, our assumption that a banks' business model matters is confirmed. Table 7 shows that in our global sample, only BHCs exhibit a significant and negative coefficient. In the U.S. subsample, BHCs experience the most negative effect from natural disasters whereas commercial banks have a lower, albeit still significant, coefficient. U.S. savings banks exhibit a non-significant and economically small coefficient. Further investigation is needed to explore the causes. In particular, it is likely that a more refined geographical matching of disasters and bank lending activities will affect our results. For instance, there are several thousand savings banks that operate across the United States, and while a certain proportion of these banks is likely to be severely affected by a disaster, the remainder are likely to be unaffected because they are geographically removed from the disaster. On average, this makes savings banks appear unexposed, even though the individual exposures within this group may vary widely. Future research may also consider if business models are still to be conceived as a proxy for diversification as far as damages from natural disasters are concerned. Natural catastrophes represent to a certain extent a systemic risk, and traditional patterns of diversification may fail to sufficiently protect the institutions. In addition, it is worth exploring if the benefits from diversification are potentially overcompensated by higher idiosyncratic risks institutions assume with respect to natural disasters.

Table 7. The Effect of Natural Disasters on Banks' Equity Ratios—Commercial Banks vs. Bank Holding Companies and Savings Banks

	Commercial Banks			Bank Holding Companies			Savings Banks		
Dependent Variable: ΔE/TA (Winsorized)	**(1) US Only**	**(2) Non-US**	**(3) Full Sample**	**(4) US Only**	**(5) Non-US**	**(6) Full Sample**	**(7) US Only**	**(8) Non-US**	**(9) Full Sample**
Damage Ratio	–0.360***	–0.012**	–0.014**	–0.406***	0.062	0.120	–0.083	–0.157***	–0.022
	(0.000)	(0.047)	(0.021)	(0.000)	(0.478)	(0.274)	(0.359)	(0.008)	(0.721)
Lagged Equity Ratio	–0.346***	–0.235***	–0.267***	–0.142***	–0.175***	–0.146***	–0.189***	–0.052***	–0.120***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Log (Total Assets)	–0.003***	–0.006***	–0.003***	–0.001***	–0.006***	–0.002***	–0.003***	–0.000**	–0.002***
	(0.000)	(0.000)	(0.000)	(0.003)	(0.003)	(0.000)	(0.000)	(0.012)	(0.000)
Net Loans Ratio	–0.025***	–0.005**	–0.017***	–0.012***	0.012	–0.008***	–0.021***	–0.002**	–0.014***
	(0.000)	(0.012)	(0.000)	(0.000)	(0.202)	(0.000)	(0.000)	(0.014)	(0.000)
Customer Deposits Ratio	–0.132***	–0.041***	–0.079***	–0.033***	–0.058***	–0.041***	–0.077***	–0.003**	–0.037***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.002)	(0.000)	(0.000)	(0.038)	(0.000)
Lagged Net Income to Equity Ratio	0.029***	0.025***	0.036***	0.014***	0.037	0.022***	0.019**	0.009*	0.022***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.202)	(0.000)	(0.015)	(0.059)	(0.000)
Real GDP Growth Rate	0.071***	0.004	0.000	0.018	0.012	–0.009	0.078***	0.019**	–0.005
	(0.000)	(0.273)	(0.973)	(0.111)	(0.448)	(0.534)	(0.000)	(0.018)	(0.150)
Growth Rate of Credit to Private Sector	–0.014***	–0.021***	–0.022***	–0.011***	–0.031*	–0.042***	0.016**	–0.010	0.002
	(0.000)	(0.000)	(0.000)	(0.002)	(0.054)	(0.001)	(0.015)	(0.166)	(0.606)
Log (Real GDP per capita)	0.048***	0.003	–0.000	0.021***	0.017	0.006	0.056***	–0.008**	0.005**
	(0.000)	(0.118)	(0.900)	(0.000)	(0.140)	(0.476)	(0.000)	(0.033)	(0.011)
Country FE		Yes	Yes		Yes	Yes		Yes	Yes
Year FE		Yes	Yes		Yes	Yes		Yes	Yes
Accounting Standard FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	71,106	20,985	92,091	10,314	1,060	11,374	7,423	11,206	18,629
Adjusted R^2	0.370	0.182	0.264	0.097	0.108	0.114	0.168	0.122	0.119

Note: This table presents the results for a series of models in which we regress the equity ratio of banks on the weighted damage ratio and other control variables over the 2000–17 period, for different subsamples of our data set based on the business model of each bank. The dependent variable is the change in the equity ratio (winsorized). The independent variable of interest is the damage ratio calculated over a period of 60 days. Columns 1 to 3 report the result for the subsample of commercial banks. Columns 4 to 6 report the result for the subsample of bank holding companies, and columns 7 to 9 report the result for the subsample of savings banks. Columns 1, 4, and 7 report results for the U.S. sample; columns 2, 5, and 8 report results for the non-U.S. sample; and columns 3, 6, and 9 report results for the full sample. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

6.4.2 Ex ante Capitalization Levels

Another attribute that may affect the sensitivity of banks to disasters is the extent of their ex ante capitalization. We expect that banks with higher capital can better mitigate and control damages from disasters, as they are better equipped to offset losses. Particularly, single losses may affect large and well-capitalized banks to a lesser extent than smaller institutions with less capital. We consider a bank's total equity as a proxy for size and the equity ratio as a proxy for the bank's equity base. We then use a series of quantile regressions to examine whether our results differ across banks with different ex ante equity ratios.

Table 8 provides the results for the 0.25, 0.50, and 0.75 quantiles. When examining the coefficients for the damage ratio across all countries and for the United States only, we observe that banks with a lower equity ratio (i.e., the 0.25 quantile) exhibit a higher (negative) sensitivity to damages than banks at the 0.75 quantile. The sensitivity decreases continuously from the 0.25 quantile to the 0.75 quantile. In the subsample of non-U.S. banks, the coefficients of the damage ratio in the 0.25 and 0.50 quantile regressions are not significant but become significantly negative in the 0.75 quantile. Overall, these results suggest that higher ex ante equity ratios appear to reduce the impact of natural disasters on a bank's solvency. Another potential explanation is that banks with a higher capitalization have been hit less frequently by disasters in the past and therefore have been able to maintain higher levels of equity capital.

6.4.3 Bank Location

From a spatial perspective, the severity of disasters and the magnitude of the associated damages may vary among countries and affect some banks more than others based on their location. In addition, natural disasters of the same magnitude may hit smaller countries more extensively, while their impact on large countries may be comparatively small. For example, a single tsunami may destroy much of the infrastructure of any Caribbean state. In contrast, the 2008 earthquake in China's Sichuan province had a destructive effect on this province but had a relatively small impact on banks in surrounding provinces, because they are geographically far removed.

Table 8. The Effect of Natural Disasters on Banks' Equity Ratios—Quantile Regression Results for the 0.25, 0.50, and 0.75 Quantiles

	0.25 Quantile			0.5 Quantile			0.75 Quantile		
Dependent Variable: ΔE/TA (Winsorized)	**(1) US Only**	**(2) Non-US**	**(3) Full Sample**	**(4) US Only**	**(5) Non-US**	**(6) Full Sample**	**(7) US Only**	**(8) Non-US**	**(9) Full Sample**
Damage Ratio	–0.294***	–0.003	–0.022***	–0.243***	–0.006	–0.018***	–0.190***	–0.012**	–0.014***
	(0.000)	(0.688)	(0.002)	(0.000)	(0.244)	(0.000)	(0.000)	(0.046)	(0.002)
Lagged Equity Ratio	–0.139***	–0.135***	–0.132***	–0.051***	–0.038***	–0.031***	–0.056***	–0.017***	–0.032***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Log (Total Assets)	–0.001***	–0.000***	–0.000***	–0.000***	–0.000***	–0.000***	–0.000***	–0.001***	–0.001***
	(0.000)	(0.003)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.004***	0.005***	0.001*	–0.003***	0.001***	–0.001***	–0.004***	–0.003***	–0.002***
	(0.000)	(0.000)	(0.098)	(0.000)	(0.002)	(0.000)	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.019***	0.001	–0.004***	–0.023***	–0.002***	–0.009***	–0.025***	–0.007***	–0.012***
	(0.000)	(0.126)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.020***	0.015***	0.020***	0.008***	0.006***	0.010***	–0.001***	0.000	–0.000
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.006)	(0.790)	(0.654)
Real GDP Growth Rate	0.063***	0.002*	–0.006***	0.050***	0.002*	–0.005***	0.034***	0.000	–0.006***
	(0.000)	(0.076)	(0.000)	(0.000)	(0.073)	(0.000)	(0.000)	(0.712)	(0.000)
Growth Rate of Credit to Private Sector	–0.015***	–0.011***	–0.013***	–0.015***	–0.011***	–0.013***	–0.015***	–0.010***	–0.013***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Log (Real GDP per capita)	0.017***	0.000	0.002***	0.006***	–0.000	0.001***	0.004***	–0.001	0.002***
	(0.000)	(0.852)	(0.000)	(0.000)	(0.703)	(0.000)	(0.000)	(0.360)	(0.000)
Country FE		Yes	Yes		Yes	Yes		Yes	Yes
Year FE		Yes	Yes		Yes	Yes		Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Accounting Standard FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	88,918	53,145	142,063	88,918	53,145	142,063	88,918	53,145	142,063

Note: This table presents the results for a series of models in which we regress (using quantile regressions) the equity ratio of banks on the weighted damage ratio and other control variables over the 2000–17 period for the 142,063 firm-year observations in our sample for which data on the equity ratio are available. The dependent variable in all models is the change in the equity ratio (winsorized). The independent variable of interest is the damage ratio calculated over a period of 60 days. Columns 1 to 3 provide the results for the 0.25 quantile regression, columns 4 to 6 provide the results for the 0.50 quantile regression, and columns 7 to 9 provide the results for the 0.75 quantile regression. Columns 1, 4, and 7 report results of the U.S. sample, columns 2, 5, and 8 report results of the non-U.S. sample, and columns 3, 6, and 9 report results for the full sample. P-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

For smaller countries, the reduced land mass increases the likelihood that a disaster affects one or more of a bank's clients. Second, banks in smaller countries are less likely to be diversified. As a result, a country's land mass should be negatively related to a bank's post-disaster solvency.

Table 9 provides empirical evidence for several tests that address this issue. We divide our sample into two subsamples based on whether the land mass of their respective countries falls above or below the sample median. In column 1, we can see that disaster damages have a significantly negative impact on the equity ratio of banks in countries with a comparatively small land size. However, regardless of whether we include banks in the United States, the results are insignificant for larger countries (see columns 2 and 3). As before, disasters appear to have no impact on the Tier 1 capital ratio for either small or large countries.

We further estimate a weighted regression in which we assign a weight equal to one divided by the square root of a country's 2017 surface area based on data from the World Bank Indicators Database (https://data.worldbank.org/indicator) to all damage observations in that country. Our results remain robust with respect to this methodological variation.

Ceteris paribus, as the size of a country increases, it is more likely to be affected by a natural disaster. We thus adjust the natural disaster variable for a country's surface area to rule out any endogeneity bias, and the results remain robust. In the full sample, the banks' equity ratio is significantly reduced after they experience a natural disaster (columns 1 and 2 in Table 10), but the capital adequacy ratio is not significantly curtailed (column 3).

Next, we group countries into quintiles based on the 2017 GDP per capita in each country, again employing data provided by the World Bank Indicators Database. In rare instances where 2017 data were not available, we employ the most recent available GDP per capita for that country and then extrapolate it to the year 2017 (i.e., to the end of our sample period) using the GDP per capita growth rate during the previous five years as a growth factor. We then assign dummy variables to each of the five quintiles and estimate a "wealth fixed-effect regression" in which we include dummy variables for four of the five quintiles in our model (excluding the center quintile). The coefficients for the dummy variables denoting the lowest two

Table 9. The Effect of Natural Disasters on Banks' Equity and Tier 1 Capital Ratios—Small vs. Large Countries

	Dependent Variable: ΔE/TA (Winsorized)		Dependent Variable: ΔT1R/TA (Winsorized)	
	(1) Small Countries	(2) Large Countries (incl. US)	(3) Small Countries	(4) Large Countries (incl. US)
Damage Ratio	–0.018***	–0.011	–0.101	–0.013
	(0.010)	(0.499)	(0.522)	(0.725)
Lagged Equity Ratio	–0.213***	–0.228***		
	(0.000)	(0.000)		
Lagged Tier 1 Capital Ratio			–0.224***	–0.235***
			(0.000)	(0.000)
Log (Total Assets)	–0.005***	–0.003***	–0.005***	–0.004***
	(0.000)	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.007***	–0.016***	–0.032***	–0.080***
	(0.004)	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.041***	–0.063***	–0.035***	–0.073***
	(0.000)	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.023***	0.038***	0.025***	0.052***
	(0.000)	(0.000)	(0.008)	(0.000)
Real GDP Growth Rate	0.006	–0.005**	0.017	0.002
	(0.373)	(0.017)	(0.323)	(0.699)
Growth Rate of Credit to Private Sector	–0.019***	–0.015***	–0.061***	–0.023***
	(0.001)	(0.000)	(0.000)	(0.000)
Log (Real GDP per capita)	0.004	0.003**	–0.011	–0.011***
	(0.105)	(0.037)	(0.242)	(0.000)
Country FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes
Accounting Standard FE	Yes	Yes	Yes	Yes
N	13,471	128,592	4,681	103,151
Adjusted R^2	0.180	0.228	0.199	0.239

Note: This table presents the results for a series of models in which we regress the capital ratio (i.e., either the equity ratio or the Tier 1 capital ratio) of banks on the weighted damage ratio and other control variables over the 2000–17 period for different subsamples of our data set based on the size of the country in which each bank is headquartered. The dependent variable in columns 1 and 2 is the change in the equity ratio (winsorized). The dependent variable in columns 3 and 4 is the change in the Tier 1 capital ratio (winsorized). The independent variable of interest is the damage ratio calculated over a period of 60 days. Columns 1 and 3 report the result for the subsample of banks in countries whose land mass is smaller than the median of all countries. Columns 2 and 4 report the result for the subsample of banks in countries (including the United States) whose land mass is larger than the median of all countries. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

Table 10. The Effect (Area Adjusted) of Natural Disasters on Banks' Equity and Tier 1 Capital Ratios

	Dependent Variable: ΔE/TA (Winsorized)		Dependent Variable: ΔT1R/TA (Winsorized)	
	(1) Full Sample	**(2) Full Sample**	**(3) Full Sample**	**(4) Full Sample**
Lagged Equity Ratio	–0.191***	–0.223***		
	(0.000)	(0.000)		
Lagged Tier 1 Capital Ratio			–0.205***	–0.231***
			(0.000)	(0.000)
Damage Ratio (/Area)	–0.002***	–0.003**	–3.581	–7.203*
	(0.009)	(0.013)	(0.112)	(0.085)
Log (Total Assets)	–0.002***	–0.003***	–0.003***	–0.004***
	(0.000)	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.014***	–0.015***	–0.068***	–0.077***
	(0.000)	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.040***	–0.059***	–0.051***	–0.069***
	(0.000)	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.034***	0.036***	0.035***	0.050***
	(0.000)	(0.000)	(0.000)	(0.000)
Real GDP Growth Rate	–0.011***	–0.004**	–0.004	–0.004
	(0.000)	(0.042)	(0.202)	(0.379)
Growth Rate of Credit to Private Sector	–0.011***	–0.017***	–0.028***	–0.033***
	(0.000)	(0.000)	(0.000)	(0.000)
Log (Real GDP per capita)	0.000	0.003**	0.003***	–0.007***
	(0.134)	(0.011)	(0.000)	(0.003)
Country FE		Yes		Yes
Year FE		Yes		Yes
Specialization FE		Yes		Yes
Accounting Standard FE		Yes		Yes
N	142,063	142,063	107,832	107,832
Adjusted R^2	0.178	0.217	0.189	0.233

Note: This table presents the results for a series of models in which we regress changes in the equity ratio and Tier 1 capital ratio of banks on the damage ratio (adjusted for the land mass of the country in which a given bank is headquartered) and other control variables over the 2000–17 period for different subsamples of our data set. The dependent variable in columns 1 and 2 is the change in the equity ratio (winsorized). The dependent variable in columns 3 and 4 is the change in the tier 1 capital ratio (winsorized). Columns 2 and 4 report results with country, year, specialization, and accounting standard fixed effects. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

GDP per capita quintiles are significantly negative, suggesting that banks in poorer countries (i.e., countries with lower GDP per capita figures) experience larger declines in solvency following a natural disaster (see Table 11).

Considering that the occurrence of natural disasters closely relates to the geographical location of banks, we divide the sample into six groups based on the continent on which they are located. The results show that the equity ratio of banks in Africa exhibits the strongest negative effects (column 1 in Table 12), which may be attributed to an overall less resilient banking system in that region, followed by smaller adjustments in bank assets in Oceania and North America (columns 5 and 4). Even so, there is no significant change in the Tier 1 capital ratio on any continent, supporting our main hypothesis (columns 7–12). In other words, banks on all continents appear to meet the capital adequacy requirements for regulatory purposes, with little variation following a natural disaster. However, this finding also reveals that the Basel Accords fail to take into account the adequacy of capital requirements in the case of natural disasters.

6.5 Different Types of Disasters

Finally, we aim to analyze the consequences of different types of disasters on bank solvency. Among all disaster types, floods have the most significant impact on capital (column 2 in Table 13), followed by storms (column 1), while earthquakes have no significant impact (column 3). As Benson and Clay (2004) point out, geological disasters such as earthquakes are likely to cause Schumpeter's "creative destruction" and thus stimulate post-disaster economic development. Yet meteorological and hydrological disasters such as storms and floods that occur more frequently are likely to have a pronounced negative effect on the local economy, leading to a significant contraction in the equity ratio.

7. Additional Robustness Tests

To further ensure the robustness of our results, we perform some additional sensitivity checks. First, we are interested whether and

Table 11. The Effect of Natural Disasters on Banks' Equity and Tier 1 Capital Ratios—Wealthy vs. Poor Countries

	Dependent Variable: ΔE/TA (Winsorized)		Dependent Variable: ΔT1R/TA (Winsorized)	
	(1) GDP Low	(2) GDP High	(3) GDP Low	(4) GDP High
Lagged Equity Ratio	–0.197***	–0.259***		
	(0.000)	(0.000)		
Lagged Tier 1 Capital Ratio			–0.217***	–0.245***
			(0.000)	(0.000)
Damage Ratio	–0.113*	–0.539*	–0.126	–0.366
	(0.059)	(0.088)	(0.775)	(0.559)
Log (Total Assets)	–0.004***	–0.003***	–0.004***	–0.004***
	(0.000)	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.006***	–0.019***	–0.048***	–0.083***
	(0.000)	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.035***	–0.086***	–0.026***	–0.087***
	(0.000)	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.031***	0.035***	0.036***	0.051***
	(0.000)	(0.000)	(0.000)	(0.000)
Real GDP Growth Rate	0.001	–0.008***	0.009	–0.031***
	(0.799)	(0.005)	(0.265)	(0.000)
Growth Rate of Credit to Private Sector	–0.020***	–0.020***	–0.051***	–0.017*
	(0.000)	(0.000)	(0.000)	(0.056)
Log (Real GDP per capita)	0.004***	–0.004**	–0.004	–0.004
	(0.005)	(0.047)	(0.229)	(0.379)
Country FE	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes
Accounting Standard FE	Yes	Yes	Yes	Yes
N	41,194	100,869	17,409	90,423
Adjusted R^2	0.160	0.274	0.185	0.254

Note: This table presents the results for a series of models in which we regress changes in the equity ratio and Tier 1 capital ratio of banks on the damage ratio and other control variables over the 2000–17 period for different subsamples of our data set based on the GDP per capita of the country in which a bank is headquartered. The dependent variable in columns 1 and 2 is the change in the equity ratio (winsorized). The dependent variable in columns 3 and 4 is the change in the Tier 1 capital ratio (winsorized). Columns 1 and 3 report results for the lower GDP per capita subsample (with a GDP per capita below the full sample median); columns 2 and 4 report results for the higher GDP per capita subsample. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

Table 12. The Effect of Natural Disasters on Banks' Equity and Tier 1 Capital Ratios—Banks from Different Continents

	Dependent Variable: ΔE/TA (Winsorized)						Dependent Variable: ΔT1R/TA (Winsorized)					
	(1) Africa	(2) Asia	(3) Europe	(4) N. Amer.	(5) Oceania	(6) S. Amer.	(7) Africa	(8) Asia	(9) Europe	(10) N. Amer.	(11) Oceania	(12) S. Amer.
Lagged Equity Ratio	–0.320***	–0.242***	–0.158***	–0.298***	–0.070	–0.184***						
	(0.000)	(0.000)	(0.000)	(0.000)	(0.229)	(0.000)						
Lagged Tier 1 Capital Ratio							–0.483***	–0.218***	–0.160***	–0.255***	–0.202***	–0.385***
							(0.000)	(0.000)	(0.000)	(0.000)	(0.001)	(0.000)
Damage Ratio	–0.376***	0.044	–0.028	–0.010**	–0.030***	–0.021	0.001	0.028	–0.071	–0.406	0.176	0.018
	(0.005)	(0.365)	(0.702)	(0.021)	(0.007)	(0.174)	(0.998)	(0.629)	(0.615)	(0.125)	(0.413)	(0.769)
Log (Total Assets)	–0.006***	–0.004***	–0.002***	–0.003***	–0.001*	–0.008***	–0.012***	–0.006***	–0.002***	–0.004***	–0.002	–0.021***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.055)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.138)	(0.000)
Net Loans Ratio	0.021***	–0.013***	–0.005***	–0.020***	–0.005	–0.010	–0.133***	–0.063***	–0.031***	–0.087***	0.004	–0.120***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.469)	(0.142)	(0.000)	(0.000)	(0.000)	(0.000)	(0.849)	(0.000)
Customer Deposits Ratio	–0.063***	–0.050***	–0.025***	–0.111***	–0.002	–0.053***	–0.130***	–0.027***	–0.014***	–0.096***	–0.004	–0.092***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.661)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.821)	(0.000)
Lagged Net Income to Equity Ratio	0.042***	0.024***	0.020***	0.036***	0.021	0.030**	0.082***	0.048***	0.019**	0.052***	–0.017	0.040
	(0.000)	(0.000)	(0.000)	(0.000)	(0.197)	(0.013)	(0.005)	(0.000)	(0.012)	(0.000)	(0.528)	(0.225)
Real GDP Growth Rate	0.011	–0.004	0.013	0.008	–0.051**	–0.015	0.042	–0.010	0.012	–0.079	–0.030	0.054
	(0.152)	(0.484)	(0.139)	(0.515)	(0.027)	(0.225)	(0.196)	(0.323)	(0.479)	(0.318)	(0.517)	(0.121)
Growth Rate of Credit to Private Sector	–0.024***	–0.014*	–0.017***	–0.001	0.028	–0.037***	–0.025	–0.038***	–0.038***	–0.048	–0.073	0.014
	(0.001)	(0.088)	(0.005)	(0.919)	(0.234)	(0.009)	(0.390)	(0.001)	(0.001)	(0.352)	(0.373)	(0.823)
Log (Real GDP per capita)	0.004	0.001	–0.001	0.014**	0.017	0.011***	0.005	–0.005	0.001	0.005	0.044	0.017
	(0.375)	(0.668)	(0.787)	(0.040)	(0.513)	(0.004)	(0.821)	(0.252)	(0.834)	(0.851)	(0.626)	(0.412)
Country FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Year FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Accounting Standard FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
N	3,137	10,048	35,893	90,680	294	2,011	962	5,244	14,261	86,597	184	577
Adjusted R^2	0.301	0.205	0.122	0.319	0.193	0.144	0.355	0.195	0.133	0.265	0.187	0.274

Note: This table presents the results for a series of models in which we regress changes in the equity ratio and Tier 1 capital ratio of banks on the damage ratio and other control variables over the 2000–17 period for different subsamples of our data set based on the continent on which a given bank is headquartered. The dependent variable in columns 1 to 6 is the change in the equity ratio (winsorized). The dependent variable in columns 7 to 12 is the change in the tier 1 capital ratio (winsorized). Columns 1 and 7, 2 and 8, 3 and 9, 4 and 10, 5 and 11, 6 and 12 report results for banks headquartered in Africa, Asia, Europe, North America, Oceania, and South America, respectively. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

Table 13. The Effect of Natural Disasters on Banks' Equity and Tier 1 Capital Ratios—Different Types of Disasters

	Dependent Variable: ΔE/TA Winsorized)		
	(1) Full Sample	**(2) Full Sample**	**(3) Full Sample**
Lagged Equity Ratio	–0.222***	–0.222***	–0.222***
	(0.000)	(0.000)	(0.000)
Storm Damage Ratio	–0.011*		
	(0.087)		
Flood Damage Ratio		–0.050*	
		(0.060)	
Earthquake Damage Ratio			–0.009
			(0.418)
Log (Total Assets)	–0.003***	–0.003***	–0.003***
	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.015***	–0.015***	–0.015***
	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.059***	–0.059***	–0.059***
	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.036***	0.036***	0.036***
	(0.000)	(0.000)	(0.000)
Real GDP Growth Rate	–0.004**	–0.004**	–0.004**
	(0.038)	(0.040)	(0.040)
Growth Rate of Credit to Private Sector	–0.017***	–0.016***	–0.017***
	(0.000)	(0.000)	(0.000)
Log (Real GDP per capita)	0.003**	0.002**	0.003**
	(0.022)	(0.027)	(0.022)
Country FE	Yes	Yes	Yes
Year FE	Yes	Yes	Yes
Specialization FE	Yes	Yes	Yes
Accounting Standard FE	Yes	Yes	Yes
N	142,063	142,063	142,063
Adjusted R^2	0.217	0.217	0.217

Note: This table presents the results for a series of models in which we regress changes in the equity ratio of banks on the damage ratio associated with various types of natural disasters and other control variables over the 2000–17 period for the 142,063 firm-year observations in our sample for which data on the equity ratio are available. The dependent variable in all columns is the change in the equity ratio (winsorized). Columns 1, 2, and 3 report results for storm, flood, and earthquake damages, respectively. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

how much any potential damages from the previous year can influence the equity ratio of the current year. We therefore regress the change in the equity ratio of the current year against the damage ratio of the previous year (damage ratio_{t-1}). Our results, presented in columns 1 and 2 of Table 14, show that, regardless of whether we use the damage ratio over 60 or 180 days, its significance level decreases relative to our main regression results in Table 4, although the coefficient remains negative. This suggests that damages affect a bank's capitalization relatively quickly and that, as time passes, the impact decreases.

Next, we examine whether our results stay robust if we do not control for the equity ratio in the previous year. We estimate the respective regressions in column 3 and 4. The coefficients of both the 60- and 180-day damage ratio remain negative and highly significant. Moreover, we employ a system generalized method of moments (GMM) approach to estimate our regressions. The coefficients are still significantly negative. However, it is worth noting in this context that with fixed effect dummies (or any other dummy with many 0s or 1s), the results of a system GMM can be biased.[6]

Finally, our results remain robust after standardizing all variables. When doing so, we observe that for every standard deviation change in the damage ratio, the equity ratio decreases by 0.004 standard deviations (column 3 in Table 15). Although 0.004 is a small number, the damages from disasters can be surprisingly large: the highest damage ratio is 148.38 percent, which corresponds to approximately 192.70 standard deviations (0.77 percent).

8. Conclusion

This paper examines whether and how natural disasters affect bank solvency. Specifically, using a sample of 9,928 banks located in 149 countries and data on natural disasters that occurred around the globe during the period 2000–17, we examine how natural disaster damages affect banks' equity ratios and Tier 1 capital ratios.

Our major finding is that damages from natural disasters negatively affect bank solvency. The relationship varies across regions

[6]For additional details, please refer to Roodman (2009).

Table 14. The Effect of Natural Disasters on Banks' Equity Ratios—Additional Robustness Tests

	OLS Regressions				System GMM Regressions	
Dependent Variable: ΔE/TA (Winsorized)	**(1) Damage Period: 60 Days**	**(2) Damage Period: 180 Days**	**(3) Damage Period: 60 Days**	**(4) Damage Period: 180 Days**	**(5) Damage Period: 60 Days**	**(6) Damage Period: 180 Days**
Lagged Damage Ratio	–0.015*	–0.011				
	(0.050)	(0.229)				
Damage Ratio			–0.016***	–0.019***	–0.789**	–0.807**
			(0.010)	(0.008)	(0.016)	(0.037)
Lagged Equity Ratio	–0.222***	–0.222***			–0.725***	–0.689***
	(0.000)	(0.000)			(0.000)	(0.000)
Log (Total Assets)	–0.003***	–0.003***	0.000***	0.000***	–0.012***	–0.012***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.015***	–0.015***	–0.001	–0.001	–0.000	–0.002
	(0.000)	(0.000)	(0.235)	(0.235)	(0.977)	(0.833)
Customer Deposits Ratio	–0.059***	–0.059***	–0.005***	–0.005***	0.054***	0.046***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.036***	0.036***	0.048***	0.048***	–0.362***	–0.332***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Real GDP Growth Rate	–0.004**	–0.004**	–0.004	–0.004	0.056***	0.052***
	(0.045)	(0.042)	(0.106)	(0.108)	(0.001)	(0.000)
Growth Rate of Credit to Private Sector	–0.016***	–0.016***	–0.021***	–0.021***	–0.062**	–0.046**
	(0.000)	(0.000)	(0.000)	(0.000)	(0.040)	(0.032)
Log (Real GDP per capita)	0.003**	0.003**	–0.000	–0.000	–0.032	–0.053
	(0.023)	(0.023)	(0.872)	(0.867)	(0.456)	(0.167)
Year FE	Yes	Yes	Yes	Yes	Yes	Yes
Country FE	Yes	Yes	Yes	Yes	Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes	Yes	Yes
Accounting Standard FE	Yes	Yes	Yes	Yes	Yes	Yes
N	142,063	142,063	142,063	142,063	142,063	142,063
Adjusted R^2	0.217	0.217	0.034	0.034		
Hansen's p-value					0.360	0.148
Arelanno–Bond AR(1) p-value					0.000	0.000
Arelanno–Bond AR(2) p-value					0.557	0.086

Note: This table presents the results for a series of models in which we regress changes in the equity ratio of banks on the weighted damage ratio and other control variables over the 2000–17 period for the 142,063 firm-year observations in our sample for which data on the equity ratio are available. The dependent variable in all models is the change in the equity ratio (winsorized). In columns 1 and 2, the independent variable of interest is the lagged damage ratio. In columns 3 and 4, the independent variable of interest is still the damage ratio, but the lagged equity ratio is excluded from the control variables. In columns 5 and 6, the independent variable of interest is also the damage ratio, but we use system GMM regressions. In columns 1, 3, and 5, we use the damage ratio calculated over a period of 60 days, and in columns 2, 4, and 6, we use the damage ratio calculated over a period of 180 days. P-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

Table 15. Economic Significance

	Dependent Variable: ΔE/TA (Winsorized)			Dependent Variable: ΔT1R/TA (Winsorized)		
	(1) US Only	(2) Non-US	(3) Full Sample	(4) US Only	(5) Non-US	(6) Full Sample
Lagged Equity Ratio	–0.832***	–0.510***	–0.607***			
	(0.000)	(0.000)	(0.000)			
Lagged Tier 1 Capital Ratio				–0.668***	–0.552***	–0.605***
				(0.000)	(0.000)	(0.000)
Damage Ratio	–0.096***	–0.003**	–0.004**	–0.049***	–0.000	–0.003
	(0.000)	(0.037)	(0.011)	(0.000)	(0.935)	(0.625)
Log (Total Assets)	–0.169***	–0.217***	–0.186***	–0.157***	–0.163***	–0.167***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Net Loans Ratio	–0.151***	–0.035***	–0.098***	–0.402***	–0.198***	–0.339***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Customer Deposits Ratio	–0.660***	–0.190***	–0.347***	–0.374***	–0.099***	–0.285***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Lagged Net Income to Equity Ratio	0.110***	0.105***	0.128***	0.104***	0.072***	0.117***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Real GDP Growth Rate	0.091***	–0.001	–0.011***	0.072***	0.012	–0.010
	(0.000)	(0.714)	(0.047)	(0.000)	(0.181)	(0.287)
Growth Rate of Credit to Private Sector	–0.025***	–0.038***	–0.032***	–0.022***	–0.062***	–0.042***
	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)	(0.000)
Log (Real GDP per capita)	1.990***	0.152**	0.103**	1.831***	–0.110	–0.208***
	(0.000)	(0.016)	(0.022)	(0.000)	(0.177)	(0.000)
Country FE		Yes	Yes		Yes	Yes
Year FE		Yes	Yes		Yes	Yes
Specialization FE	Yes	Yes	Yes	Yes	Yes	Yes
Accounting Standard FE	Yes	Yes	Yes	Yes	Yes	Yes
N	88,918	53,145	142,063	86,231	21,601	107,832
Adjusted R^2	0.316	0.154	0.217	0.238	0.265	0.177

Note: This table presents the results for a series of models in which we regress standardized changes in the equity ratio and Tier 1 capital ratio of banks on the standardized damage ratio and other control variables over the 2000–17 period for different subsamples of our data set. The dependent variable in columns 1 to 3 is the change in the equity ratio (winsorized). The dependent variable in columns 4 to 6 is the change in the Tier 1 capital ratio (winsorized). Columns 1, 4, and 7 report results for the U.S.-only subsample; columns 2, 5, and 8 report results for the non-U.S. subsample; and columns 3, 6, and 9 report results for the full sample. All variables are standardized. Robust p-values are reported in parentheses. *, **, and *** indicate significance at the 10 percent, 5 percent, and 1 percent level, respectively.

and among different types of banks but provides compelling evidence that natural disasters represent a significant threat for the financial soundness of individual banks and, by extension, the stability of our financial system as a whole.

We hypothesized that the Tier 1 capital ratio—as a regulatory measure of bank solvency—would be more sensitive to natural disaster damages than the accounting-based equity ratio. However, natural disasters appear to affect the Tier 1 capital ratio to a lesser extent than the corresponding accounting ratio. Although this issue calls for further investigation, we conclude that the regulatory weights attributed to risky assets in the Tier 1 capital ratio specification are not adequate in capturing a bank's exposure to natural disasters. The regulatory risk weights stem from historical evidence and rely primarily on economic drivers of risk. However, the observable increase in the frequency and severity of natural disasters is a more recent phenomenon with roots that largely lie outside the financial system.

The results of our study have important implications for financial regulators and risk managers. In particular, financial regulators should consider modifying the assessment and weighting of solvency risks in light of the increasing damages caused by natural disasters. For instance, they may consider explicitly including disasters as a source of operational risk and increasing the risk weights for customers who are particularly exposed to these risks. Similarly, managers of institutions that lend in disaster-prone areas should include the expected damages from disasters in their calculations of the risk premium of loans. If the premium is priced correctly, i.e., when it accounts for higher damages from natural disasters, any losses in a bank's lending business should be largely compensated by the premium.

The negative effect of natural disasters on bank solvency varies depending on the specific profile of banks. Banks located in countries where damages from natural disasters have a relatively high impact (as compared with the GDP) show a higher degree of affectedness. This is also the case for banks with a low ex ante capitalization. In contrast, our study does not find significant and consistent results for banks with different business models. We conclude that natural disasters may exhibit a different propagation pattern and may affect regions, infrastructures, and institutions as a whole. Consequently, traditional diversification patterns appear to be irrelevant in this case.

A potential direction for further research on the link between bank solvency and natural disasters is to address the underlying transmission process of damages. This is challenging, as causes and effects may unfold in various forms. Natural disasters primarily affect a bank's customers but may, at the same time, jeopardize the infrastructure of banks themselves. Depending on the risk management strategies both banks and their customers employ, the effect of disasters on bank solvency may be different. In addition, the frequency and magnitude of disasters may change over time. Future research has to cope with this high degree of complexity and the dynamic nature of disasters.

Appendix A. Absolute Differences in Total Assets (ADTA) between Fitch and Bankscope

The table in this appendix examines differences in observations between the two databases (Fitch and Bankscope) used in our paper. We match banks by name and then employ a variable that measures the absolute difference in total assets (ADTA) to compare each match, where $ADTA = |BTA - FTA| / BTA$, BTA is the value of total assets of a given bank in Bankscope, and FTA is the value of total assets of the same bank in Fitch. Matches whose ADTA exceed 0.1 are excluded from our sample. Column 1 reports the number of banks whose ADTAs fall within each range bracket. Column 2 shows the percentage distribution of our sample across the different brackets. Due to missing data for several of our dependent and independent variables, the total number of observations reported here (11,881) decreases to 9,928 in Table 1.

Range of ADTA	(1) Frequency	(2) Percentage
0	9,803	82.51
0–0.0000001	723	6.09
0.0000001–0.000001	273	2.30
0.000001–0.00001	19	0.16
0.00001–0.0001	33	0.28
0.0001–0.001	76	0.64
0.001–0.01	168	1.41
0.01–0.1	468	3.94
0.1+	318	2.68
Total	11,881	100

Appendix B. Correlation between Matched Banking Variables from Fitch and Bankscope

The table in this appendix examines the correlation between various variables reported by Fitch and Bankscope for the year in which the two databases overlap (year 2013). Column 1 reports the Pearson correlation coefficients between the variable values in Fitch and the corresponding values in Bankscope. Column 2 reports the same correlations, except that variables are trimmed at the 1 percent and 99 percent level. Column 3 reports the mean percentage difference between the paired variables. Difference ratios are calculated as (the value in Bankscope – the value in Fitch)/the value in Bankscope. The variables we use in our paper (i.e., variables which exhibit a maximum difference of 10 percent) are bolded and highlighted in grey.

Variable Name	(1) Correlation	(2) Trimmed Correlation	(3) Percentage of Difference Ratio > 0.1
Total Liabilities & Equity	1	0.9999	0.00%
Total Assets	**1**	**0.9999**	**0.00%**
Net Interest Revenue	1	0.9999	4.21%
Number of Branches	1	0.9996	0.81%
Deposits & Short-Term Funding	0.9999	0.9997	1.29%
Fixed Assets	0.9999	0.9997	4.13%
Gross Loans	0.9999	0.9999	1.78%
Net Loans	**0.9999**	**0.9999**	**1.71%**
Number of Employees	0.9998	0.9999	1.30%
Total Customer Deposits	**0.9998**	**0.9997**	**0.81%**
Net Income	**0.9998**	**0.9981**	**5.62%**
Tier 1 Capital	**0.9997**	**0.9990**	**1.23%**
Intangibles	0.9997	0.9999	6.97%
Profit before Tax	0.9996	0.9990	5.52%
Derivatives	0.9996	0.9978	2.91%
Reserves for Impaired Loans/NPLs	0.9995	0.9967	6.54%
Loan Loss Reserves	0.9995	0.9965	8.90%
Total Earning Assets	0.9994	0.9994	2.04%
Impaired Loans	0.9991	0.9959	9.34%
Net Fees and Commissions	0.9990	0.9891	5.96%
Equity	**0.9987**	**0.9985**	**4.39%**
Long-Term Funding	0.9981	0.9894	14.27%
Loan Loss Reserves/Gross Loans	**0.9979**	**0.9957**	**7.94%**
Equity/Net Loans	0.9976	0.9915	4.95%
Equity/Total Assets	**0.9976**	**0.9954**	**4.18%**
Equity/Liabilities	0.9969	0.9920	5.48%
Tier 1 Capital Ratio	**0.9963**	**0.9952**	**1.38%**
Equity/Customer & Short-Term Funding	0.9959	0.9806	5.29%
Net Loans/Total Assets	0.9957	0.9951	0.80%
Trading Liabilities	0.9949	0.9853	2.96%
Tax	0.9948	0.9990	6.14%
Dividend Paid	0.9917	0.9803	13.46%
Subordinated Debts	0.9878	0.9574	4.36%
Impaired Loans/Gross Loans	0.9792	0.9865	7.94%
Net Loans/Deposits & Short-Term Funding	0.9765	0.9933	1.90%

(*continued*)

	(1)	(2)	(3)
Variable Name	**Correlation**	**Trimmed Correlation**	**Percentage of Difference Ratio > 0.1**
Net Interest Revenue/Average Assets	0.9762	0.9955	2.77%
Impaired Loans/Equity	0.9755	0.9750	11.90%
Net Charge-Offs	0.9683	0.9801	6.14%
Dividend Pay-Out	0.9571	0.9507	14.04%
Unreserved Impaired Loans/Equity	0.9487	0.9712	14.62%
Other Deposits and Short-Term Borrowings	0.9484	0.9554	8.78%
Non-Interest Expenses/Average Assets	0.9416	0.8846	51.47%
Net Interest Margin	0.8854	0.9637	34.14%
Other Operating Income/Average Assets	0.8801	0.5014	95.67%
Loan Loss Reserves/Impaired Loans	0.8612	0.9341	8.98%
Deposits from Banks	0.8459	0.8468	17.75%
Loans and Advances to Banks	0.8446	0.7786	17.12%
Liquid Assets	0.8369	0.7956	93.69%
Other Operating Income	0.6763	0.6308	95.42%
Return on Average Assets (ROAA)	0.6539	0.9895	8.30%
Return on Average Equity (ROAE)	0.6366	0.9812	10.27%
Other Securities	0.6242	0.4288	56.47%
Liquid Assets/Deposits & ST Funding	0.5550	0.4393	93.93%
Loan Loss Reserves	0.5016	0.4697	97.81%
Other Earning Assets	0.4416	0.3688	99.35%
NCO/Average Gross Loans	0.2672	0.9779	13.43%
Interbank Ratio	0.0548	0.1847	22.31%

Appendix C. Definitions and Descriptions of Variables

Name	Description	Sources
Equity Ratio	Equity/total assets, winsorized at the 1.5%–98.5% level	Bankscope & Fitch
Tier 1 Capital Ratio	Tier 1 capital/risk-weighted assets, winsorized at the 1.5%–98.5% level	Bankscope & Fitch
Damage Ratio	Total damages caused by natural disasters in a given country in year t, distributed across year t and year $t+1$ following Equation (2) and divided by the gross domestic product (GDP) of each country	EM-DAT International Disaster Database
Log (Total Assets)	Natural log of total assets, winsorized at the 1.5%–98.5% level	Bankscope & Fitch
Net Loans Ratio	Net loans/total assets, winsorized at the 1.5%–98.5% level	Bankscope & Fitch
Customer Deposits Ratio	Total customer deposits/total assets, winsorized at the 1.5%–98.5% level	Bankscope & Fitch
Net Income to Equity Ratio	Net income/equity, winsorized at the 1.5%–98.5% level	Bankscope & Fitch
Real GDP Growth Rate	Annual growth of the real GDP of a given country	World Bank
Growth Rate of Credit to Private Sector	Annual growth of domestic credit to the private sector (expressed as a percentage of GDP) in a given country, winsorized at the 1.5%–98.5% level	World Bank
Log (Real GDP per capita)	Natural log of the real GDP per capita of a given country	World Bank
Year FE	Binary variables that take on a value of 1 if a given observation falls within a year from 2000 to 2017, 0 otherwise	Bankscope & Fitch
Country FE	Binary variables that take on a value of 1 if a bank operates in one of 149 countries, 0 otherwise	Bankscope & Fitch
Specialization FE	Binary variables that take on a value of 1 if a bank operates under one of seven business models/specializations (bank holding companies, commercial banks, cooperative banks, investment banks, Islamic banks, real estate and mortgage banks, and savings banks), 0 otherwise	Bankscope & Fitch
Accounting Standard FE	Binary variables that take on a value of 1 if a bank employs one of five accounting standards (IAS, IFRS, Local GAAP, Regulatory, and U.S. GAAP) in a given year, 0 otherwise	Bankscope & Fitch

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On the Structural Determinants of Growth-at-Risk*

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We examine structural differences in growth vulnerabilities across countries associated with financial risk indicators. Considering trade openness, financial sector size, the public spending ratio, and government effectiveness, our findings suggest the existence of a structural gap and a risk sensitivity gap. Hence, structural country characteristics not only drive level differences in growth-at-risk (GaR) but also give rise to differences in the responsiveness of GaR to financial risks. Furthermore, we show that the impact of structural characteristics varies over the forecasting horizon. A proper understanding of structural country characteristics in the context of the GaR framework is important to facilitate the use of the concept in macroprudential policy.

JEL Codes: E27, E32, E44, F43, G01, G20, G28.

1. Introduction

The empirical growth-at-risk (GaR) concept introduced by Adrian, Boyarchenko, and Giannone (2019) suggests that deteriorating financial conditions are associated with increased downside risks to economic growth. While standard forecasts focus on the expected value of future gross domestic product (GDP) growth, the GaR approach places a particular emphasis on the probability and magnitude of potential adverse outcomes. Similar to the value-at-risk

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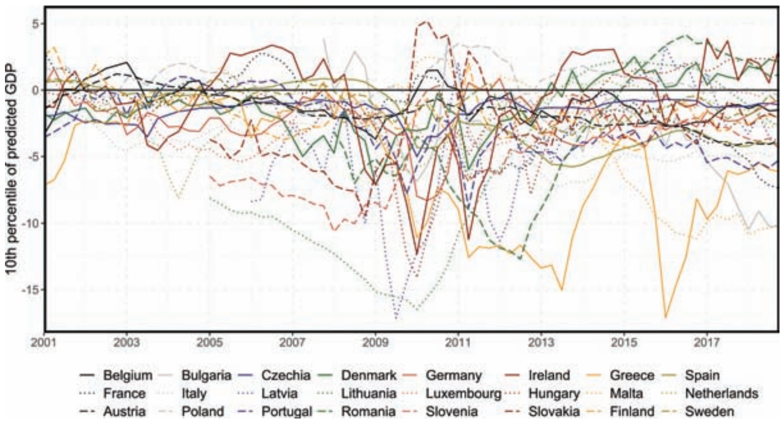
concept in finance, the GaR of an economy for a given time horizon is defined as a specific low quantile of the distribution of the projected GDP growth rate for the respective horizon (see, for instance, Suarez 2022). In this context, Adrian, Boyarchenko, and Giannone (2019) show that the left tail of the distribution of (projected) GDP growth is less stable and more affected by financial conditions than the upper quantiles of the distribution. Against this background, the GaR concept is a useful and intuitive policy tool to identify and quantify systemic risk and has therefore gained traction among policymakers in recent years.

In the last few years, the GaR concept has been extended in various directions. While financial conditions have turned out to be highly relevant for the conditional GDP growth distribution at relatively short time horizons (i.e., up to one year), risk indicators from the financial cycle literature also have been introduced into the GaR framework. In this context, recent empirical studies indicate that external imbalances, excessive credit growth, and house price booms are associated with increasing growth vulnerabilities in the medium term, typically defined as longer time horizons between six quarters and five years (Aikman et al. 2019; Arbatli-Saxegaard, Gerdrup, and Johansen 2020; Duprey and Ueberfeldt 2020). In a similar vein, characterizing the term structure of GaR, it has been highlighted that the sensitivity of downside risks to growth depends on the respective time horizon. These findings imply not only differing term structures depending on financial risk indicators but also a possible intertemporal trade-off, i.e., lower growth vulnerability at medium and long horizons may come at the cost of lower expected growth (or GaR) in the short term (Adrian et al. 2022).

By linking observed financial risk indicators as well as policy indicators to the distribution of projected growth outcomes, the GaR concept also is increasingly used as a measure of systemic risk at the individual country level (see, for instance, Adrian et al. 2019; European Systemic Risk Board (ESRB) 2019; Prasad et al. 2019). In this context, the application of the GaR concept enables policymakers to quantify the probability of adverse scenarios, thereby facilitating an appropriate and timely policy reaction. Previous studies suggest that downside risks can be mitigated to some extent by respective policy measures, e.g., by increasing the capitalization of the banking system (Aikman et al. 2019) or by applying other macroprudential or

Figure 1. Estimated 10th Percentile of Predicted GDP Growth



Note: The figure shows the time series of the predicted GaR one year ahead for each country in the sample. The predicted GaR is shifted forward to align the predictions with the realization.

monetary policy instruments (Duprey and Ueberfeldt 2020; Franta and Gambacorta 2020; Galán 2020). Furthermore, by examining the impact of policy variables on the vulnerability of GDP growth, the GaR concept also can be used as a potential measure to calibrate the current stance of macroprudential policy to safeguard financial stability (ESRB 2019; Suarez 2022).

To facilitate the use of the GaR concept as a measure of systemic risk or macroprudential policy stance, a proper understanding of cross-country differences is crucial. While some papers take into account selected country properties in their estimations (e.g., Arbatli-Saxegaard, Gerdrup, and Johansen 2020) or discuss this issue as an important area of future research (O'Brien and Wosser 2021; Suarez 2022), structural country characteristics have not been examined systematically so far in the respective strand of the literature. Such an analysis appears warranted, as the empirical GaR measure typically not only fluctuates substantially in the time dimension but also across countries (see Figure 1).[1] Our study contributes to

[1]Figure 1 shows one-year-ahead forecasts of the 10th percentile of predicted GDP growth, a frequently used measure of GaR, conditional on standard measures of financial risk estimated country by country. For each country, we estimate

the existing literature by putting a particular emphasis on structural country characteristics and their impact on empirical GaR estimates in a sample of European Union (EU) countries. The EU provides a particularly interesting setting. While countries are subject to a common regulatory framework, they vary substantially with respect to structural[2] country characteristics.

In Figure 1, two observations stand out. First, GaR estimates fell markedly in the run-up to the global financial crisis, in line with the narrative of a high financial risk episode. Second, a substantial degree of variation can be observed in the cross-section, giving rise to the potentially important role of cross-country heterogeneity. We focus on the latter aspect by examining the role of structural country characteristics in the context of GaR. Structural country characteristics can play an important role in at least three dimensions (see also Suarez 2022). First, countries can differ in their "standard" GaR values, i.e., the average GaR over time. For instance, economies with a more dispersed distribution of GDP growth are likely to exhibit lower average values of GaR. While this *structural gap* could be accounted for by including country fixed effects, it is nevertheless important to understand the drivers behind the cross-country differences in GaR, particularly from a policy perspective. Second, the reaction of GaR to changes in financial risk indicators may differ across countries. Such a *risk sensitivity gap* would become apparent when GaR estimates in individual countries show different reactions to changes in the financial risk indicator due to structural country characteristics.[3] For instance, economies with a strong financial

the following quantile function: $\hat{Q}_{y_{i,t+h}}(\tau \mid X_{i,t}) = X_{i,t}\hat{\beta}_{i,\tau}$, where $X_{i,t}$ is a vector containing the country-level index of financial stress (CLIFS), country-specific credit growth, current GDP growth, and a constant. Figure 1 shows the predicted 10th percentile of four quarters ahead, hence $\tau = 0.1$ and $h = 4$.

[2]While the expression "structural" often refers to the identification of causal effects in structural models, we use the term in a different context, i.e., structural country characteristics in the context of "non-cyclical."

[3]The gap vulnerability to risk, as defined by Suarez (2022), is a similar concept. While the risk sensitivity gap refers to differences in the coefficients of the risk parameters, the gap vulnerability to risk describes the resulting change in the "target gap," i.e., the deviation of GaR from mean (or median) growth. As this is the logical consequence of varying coefficients of the risk parameter, the two definitions are basically two sides of the same coin.

sector are likely to be less dependent on financial inflows and therefore may be less vulnerable to tightening financial conditions than countries with a less developed financial sector. Finally, structural differences across countries in principle also could result from a different effect of policy measures on the respective GaR, which can be referred to as the *policy sensitivity gap*. For instance, the magnitude of the effect of higher capital requirements in the banking sector may depend on the size of the financial sector, with limited effects on downside risks to growth in countries where the financial sector is small.

A better understanding of structural country characteristics driving differences in GaR across countries is a prerequisite to extend the use of the GaR concept in the context of policy design and assessment, as it should be taken into account when comparing GaR estimates (and the corresponding policy reactions) across countries. Our paper aims to fill this gap in the literature by focusing on the former two issues, i.e., the structural gap and the risk sensitivity gap.[4]

Specifically, we employ panel quantile regressions in which cross-country variation is modeled by including country-specific characteristics as well as interaction terms with financial risk indicators. Therefore, we examine potential drivers of the structural gap across countries by including various structural country characteristics in the panel quantile regression. As a result, we are able to shed light on the drivers of GaR across countries, which are usually disguised in country-by-country or panel fixed-effects regressions. Furthermore, we examine the interactions between structural characteristics and the respective financial risk indicator. Therefore, we investigate the impact of varying structural characteristics on the sensitivity of the GaR value with respect to the financial risk indicator, thus quantifying the respective risk sensitivity gap due to specific structural country characteristics.[5]

[4]The impact of the policy sensitivity gap is evidently also a relevant issue. However, given our sample, a reliable analysis is not feasible at the current juncture, in light of measurement errors due to challenges in quantifying macroprudential policy across countries and data availability.

[5]The panel framework is warranted not only to be able to evaluate structural country characteristics across countries but also to be able to take them into account in common regulatory frameworks. While forecasting performance

In light of a lack of previous empirical work on potential drivers of GaR across countries, we consult two adjacent strands of the literature to establish a conceptual framework. First, various country characteristics were identified to play a crucial role in explaining differences in terms of output drops in the global financial crisis and cross-country variation in business cycle volatility (see, among others, Blanchard et al. 2010; Crucini, Kose, and Otrok 2011; Lane and Milesi-Ferretti 2011; Rose and Spiegel 2011). From an ex post perspective, the financial crisis was associated with high downside risks. Characteristics that explain the realized output decline in the financial crisis may therefore also drive GaR estimates. Second, factors that explain heterogeneity in observed business cycle volatility also might help to understand GaR across countries, as countries with higher business cycle volatility show a more dispersed growth distribution than countries with limited growth volatility. As a result, the respective GaR values also are expected to be lower, as the distribution of the quantile projections reflects the distribution of GDP growth.

Four factors stand out as particularly relevant in shaping both the downturn during the global financial crisis and business cycle volatility: trade openness, public spending ratio, financial sector size, and government effectiveness. With respect to trade openness, previous studies suggest a positive link to GDP volatility (see, for instance, Loayza and Raddatz 2007; di Giovanni and Levchenko 2009; Kim, Lin, and Suen 2016), also because higher trade openness is associated with higher degrees of specialization in an economy. Thus, previous literature suggests a negative effect of higher trade openness on GaR. In contrast, the impact of public expenditures on output volatility is discussed more controversially in the literature. Carmignani, Colombo, and Tirelli (2011) report a positive link between government size and volatility, while earlier studies find a negative effect of public expenditures or government size on GDP growth volatility (Galí 1994; Fatás and Mihov 2001). Therefore, the direction of the effect also may depend on the type of taxes (Posch 2011) as well as on the type of the shock (Collard,

is somewhat attenuated in an integrated cross-country approach through pooling, the strength of the approach lies in the accurate evaluation of the sources of cross-country heterogeneity that would otherwise be difficult to capture appropriately.

Dellas, and Tavlas 2017). Thus, the impact of the public spending ratio on GaR remains mostly unclear. Regarding the size of the financial sector, empirical studies point to a dampening effect of more developed financial sectors on the volatility of GDP, consumption, and investment (Denizer, Iyigun, and Owen 2002; Beck, Lundberg, and Majnoni 2006; Manganelli and Popov 2015), although the effect seems to be less pronounced compared with trade openness, and the transmission channel may work via other structural country characteristics (Loayza and Raddatz 2007). At very high levels of financial depth, however, the effect weakens or even reverses, with high financial depth amplifying consumption, and investment volatility (Dabla-Norris and Srivisal 2013). In the context of GaR, we therefore expect a positive effect on GaR, although this effect could be reversed at higher levels of financial development. Finally, government effectiveness is generally found to be negatively linked to GDP volatility (see, for instance, Evrensel 2010) and is therefore likely to be positively linked to GaR. In summary, previous literature suggests increasing growth vulnerabilities (i.e., lower GaR) with increasing trade openness and decreasing levels of government effectiveness. The empirical effect of the ratio of public expenditures remains ambiguous, and the impact of the financial sector size may depend on the respective level of financial development, potentially resulting in a non-linear relationship between the two variables.

Our empirical analysis not only sheds light on whether stabilizing factors in the global financial crisis and with respect to growth volatility also mitigate growth risks,[6] i.e., whether these factors significantly contribute to the structural gap in GaR, but we also are able to examine the risk sensitivity gap associated with structural country characteristics. We find that structural country characteristics indeed play an important role in shaping cross-country variations in GaR. Both the structural gap and the risk sensitivity gap contribute significantly to structural differences in GaR across countries, whereby the magnitude of the effect differs by the respective financial risk indicator (i.e., financial stress versus credit growth) as well as by the respective time horizon. Higher trade openness and larger financial sectors lead to a structurally lower GaR value,

[6]Throughout the paper, lower GaR implies higher growth risks, and vice versa.

particularly at longer time horizons. Higher levels of government effectiveness mitigate growth risks across all time horizons, while the stabilizing role of a high public spending ratio is limited to the short run. The risk sensitivity gap seems to be most pronounced with respect to public spending ratio and trade openness but plays a less significant role in the context of financial sector size and government effectiveness.[7] Overall, our study highlights the importance of structural country characteristics when estimating GaR at the individual country level. We show that both the structural gap and the risk sensitivity gap play an important role, with the impact of structural characteristics varying with different time horizons, i.e., the term structure of GaR also may be driven by structural country characteristics. Finally, model evaluation exercises reveal that taking into account structural country exercises enhances the accuracy of projected growth risks compared with panel quantile regressions with fixed effects only.

The paper is structured as follows. Section 2 explains our empirical methodology and introduces a framework to examine both the structural gap and the risk sensitivity gap in the context of panel quantile regressions. Section 3 shows our empirical results, including our panel quantile estimations and the impact of the structural characteristics on the GaR term structure. Section 4 presents our model evaluation exercises, while Section 5 draws conclusions and discusses the policy implications of our empirical results.

2. Empirical Approach

2.1 Data

Our analysis is based on a cross-country unbalanced panel data set using time series from 24 European economies[8] over the period 1999:Q1–2019:Q4. The sample includes all European economies for

[7]We also find evidence for non-linearities in how financial sector size and government effectiveness affect GaR, although the effects are relatively small in magnitude compared with the overall effects of the respective structural characteristics (i.e., the structural gap).

[8]Austria, Belgium, Bulgaria, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Latvia, Lithuania, Luxembourg, Malta, Netherlands, Poland, Portugal, Romania, Slovakia, Slovenia, Spain, Sweden.

which a country-specific financial stress measure and the credit-to-GDP ratio are available. For these countries, we construct the annualized average GDP growth rates using the quarterly seasonally adjusted real GDP provided by Eurostat. The logarithm of these time series, $Y_{i,t}$, is then converted into the approximate annualized growth rates h periods ahead, $y_{i,t+h} = \frac{(Y_{i,t+h} - Y_{i,t})}{h/4}$.

In line with previous literature, we include a measure of financial stress as an explanatory variable. In our first model, we use the Composite Indicator of Systemic Stress (CISS) developed by Hollo, Kremer, and Lo Duca (2012) and published by the ECB as a measure of European-wide financial stress.[9] The CISS aggregates five market-specific subindices on the basis of weights reflecting their time-varying cross-correlation structure. Thus, the CISS takes into account both the level of individual subindices and the number of indicators suggesting high financial stress. As a result, the CISS reacts more strongly if more indicators show signs of financial stress simultaneously. In the second estimation, we follow a more traditional GaR framework and use country-specific financial stress measures, i.e., the Country-Level Index of Financial Stress (CLIFS), introduced by Duprey, Klaus, and Peltonen (2017). The construction of the index follows the approach of Hollo, Kremer, and Lo Duca (2012). Using both the CISS and the CLIFS allows us to check whether the impact of country characteristics on GaR is already implicitly captured by country-specific financial stress measures.

While financial stress measures are highly relevant for short-term GaR estimations, credit growth is frequently used as a signal for medium-term financial imbalances (see, e.g., Aikman et al. 2019; Galán 2020; Adrian et al. 2022). The Bank for International Settlements publishes credit-to-GDP ratios for a wide range of countries. Based on this data set, we use the two-year average of the log differences of the credit-to-GDP ratio as a measure of credit growth.

Finally, for each country, we collect time series of four different structural characteristics: trade openness, which we define as the ratio of exported goods to GDP; the size of the financial sector, defined as the ratio of gross value-added of the financial sector to

[9]While the CISS is a euro-area-wide indicator, we also replicate our analysis with the global financial stress index developed by Monin (2019). The results are very similar and can be seen in the appendix in Table A.1.

GDP; and the ratio of public expenditures to GDP[10] and government effectiveness, as measured by the Worldwide Governance Indicators (WGI) project (Kaufmann, Kraay, and Mastruzzi 2011).[11] To make the coefficients in our estimations comparable across all explanatory variables, all included factors are standardized with a mean of zero and a standard deviation of one.

2.2 Growth-at-Risk Methodology

Following Adrian, Boyarchenko, and Giannone (2019), we rely on quantile regressions, developed by Koenker and Bassett (1978), to estimate GaR. Because of the multicountry setup, we employ a panel quantile regression framework. A major concern when estimating panel quantile regression is the large number of fixed effects (α_i) for every cross-sectional unit, especially when N is large and T is relatively small (Koenker 2004). However, as T is much larger than N in our case, coefficients can be estimated consistently (Galvao and Montes-Rojas 2015; Adrian et al. 2022). We follow previous research and include fixed effects for each country, resulting in country-specific intercepts at each quantile (τ).[12]

Quantile regressions allow us to estimate the differential effects of the conditioning variables on the distribution of the dependent variable. In our study, we are interested in the effects on the lower part of the distribution of the dependent variable, i.e., the effects on GaR. In our model, the dependent variable, y_{t+h}, is the annualized average GDP growth 1 quarter to 16 quarters ahead ($h = 1, 2, 3, \ldots, 16$), and the vector of conditioning variables, X_t, includes a constant, current GDP growth, a measure of financial stress and credit growth, as well

[10]The data for the first three country characteristics are obtained from Eurostat.

[11]Government effectiveness captures "perceptions of the quality of public services, the quality of the civil service and the degree of its independence from political pressures, the quality of policy formulation and implementation, and the credibility of the government's commitment to such policies" (Kaufmann, Kraay, and Mastruzzi 2011). Thus, this indicator captures not only the intention of the regulations but also how they are implemented and whether they are credibly enforced.

[12]For inference, we use the block-bootstrap method, as shown in Kapetanios (2008). We use a block size of four quarters; however, changing the block size does not alter the interpretation of our results (see also Lahiri 2003).

as the structural characteristics we are mainly interested in. For each projection horizon h, we estimate the quantile function

$$\hat{Q}_{y_{i,t+h}}(\tau \mid X_{i,t}, \alpha_i) = \hat{\alpha}_{i,\tau} + X_{i,t}\hat{\beta}_\tau, \qquad (1)$$

where $\hat{\alpha}_{i,\tau}$ denotes the estimated country-specific fixed effects at quantile τ. To estimate $\hat{\alpha}_{i,\tau}$ and coefficients $\hat{\beta}_\tau$, the quantile weighted absolute value of errors is minimized:

$$(\hat{\beta}_\tau, \hat{\alpha}_{i,\tau}) = \arg\min_{\alpha_i,\beta_\tau} \sum_{i=1}^{n} \sum_{t=1}^{T-h} \rho_\tau(y_{i,t+h} - X_{i,t}\beta_\tau - \alpha_i), \qquad (2)$$

where ρ_τ is the standard asymmetric absolute loss function. As a measure for GaR, we use the 10th percentile of projected growth (in line with, e.g., Figueres and Jarociński 2020), hence, $\tau = 0.1$.

To assess the effect of structural characteristics on GaR, $X_{i,t}$ includes structural country characteristics, which we evaluate in the panel quantile regression. As explanatory variables, we consider various structural country characteristics, such as trade openness, the size of the financial sector, the public spending ratio, and government effectiveness.

In variants of the model, we consider interactions between the financial risk indicators and the included structural characteristics to take into account possible non-linearities. It is well documented that high financial stress leads to a widening of the lower tails of the distribution of projected growth (Adrian, Boyarchenko, and Giannone 2019). We thus introduce interactions to evaluate whether this form of non-linearity is further reinforced through structural country characteristics. By interacting the structural characteristics with the financial risk indicators, we allow the effects of the structural country characteristics to vary depending on current financial stress and observed credit growth.

Note that the coefficients of the structural country characteristics help to detect structural gaps, indicating whether these structural characteristics are associated with generally lower or higher GaR. The extent to which non-linearities in the impact of financial risk indicators are prevalent is indicative of the existence of risk sensitivity gaps highlighting particular sensitivities (i.e., varying responsiveness of GaR) in the face of high financial stress or credit growth.

Including interaction terms, we estimate the following panel quantile regression model:

$$\hat{Q}_{y_{i,t+h}}(\tau \mid X_{i,t}, \alpha_i) = \hat{\alpha}_{i,\tau} + X_{i,t}\hat{\beta}_\tau + Z_{i,t} \times FS_{i,t}\hat{\nu}_\tau + Z_{i,t} \times Credit_{i,t}\hat{\gamma}_\tau, \quad (3)$$

where $\alpha_{i,\tau}$ denotes the fixed effects, $Z_{i,t}$ is a subset of vector $X_{i,t}$ comprising structural country characteristics, and $FS_{i,t}$ and $Credit_{i,t}$ denote financial stress and credit growth, respectively, which also are elements of $X_{i,t}$.

Relating the concepts of the structural gap and the risk sensitivity gap to Equation (3), one may think of the partial effects of the structural characteristics, i.e., first derivatives. The structural gap can be thought of as the total effect, i.e., the coefficient on the respective measure plus the interaction terms for a given level of the financial risk measures. The risk sensitivity gap is captured by the interaction terms, as the responsiveness of GaR to the respective financial risk indicator depends on structural country characteristics.

3. Results

3.1 Main Results

First, we consider the CISS measure (Hollo, Kremer, and Lo Duca 2012) as a financial stress indicator, which is an aggregate measure that does not vary across countries. The fact that we use one and the same financial risk indicator across countries permits us a direct interpretation of how the propagation of financial stress to growth vulnerabilities is linked to country-specific structural characteristics. In contrast, credit growth is country specific. All measures in the regression are standardized to facilitate a direct comparison of the various factors in terms of magnitude.

Table 1 shows the coefficient estimates for the conditional 10th percentile for different specifications of the panel quantile regression model in which we evaluate the structural determinants of GaR. Columns 1–2 show results from a parsimonious, linear specification for forecasting horizons $h = 4$ and $h = 12$. These horizons are typically considered to assess short- and medium-term growth

Table 1. Main Results—CISS and Credit Growth

	Model 1		Model 2	
	***h* = 4** (1)	***h* = 12** (2)	***h* = 4** (3)	***h* = 12** (4)
CISS	–2.145***	–0.144**	–2.364***	–0.040
	(0.279)	(0.060)	(0.247)	(0.065)
Credit Growth	–0.225	–0.680***	–0.351	–0.986***
	(0.380)	(0.184)	(0.344)	(0.150)
Current GDP Growth	0.047	–0.573***	0.071	–0.548***
	(0.093)	(0.045)	(0.075)	(0.048)
Openness	–1.545**	–1.074***	–2.290***	–1.169***
	(0.645)	(0.237)	(0.507)	(0.219)
Financial Sector	0.036	–2.376***	0.958	–2.823***
	(1.141)	(0.704)	(1.303)	(0.735)
Public Expenditure	1.032***	0.065	1.051***	0.162
	(0.300)	(0.108)	(0.337)	(0.145)
Government Effectiveness	3.002***	2.250***	3.939***	2.609***
	(0.675)	(0.458)	(0.703)	(0.403)
Openness × CISS			0.214	0.106
			(0.238)	(0.099)
Financial Sector × CISS			0.674***	0.011
			(0.249)	(0.095)
Public Expenditure × CISS			1.271***	–0.099
			(0.278)	(0.092)
Government Effectiveness × CISS			0.184	0.124
			(0.251)	(0.098)
Openness × Credit Growth			–0.487*	–0.759***
			(0.291)	(0.173)
Financial Sector × Credit Growth			0.097	0.203***
			(0.223)	(0.064)
Public Expenditure × Credit Growth			–0.179	–0.046
			(0.176)	(0.143)
Government Effectiveness × Credit Growth			–0.181***	–0.121
			(0.367)	(0.164)
Observations	1,744	1,576	1,744	1,576

Note: The table shows the estimated coefficients of the conditional 10 percent quantile. Columns 1–2 show the results from the regression model in Equation (1) for the horizons (h) 4 and 12. Columns 3–4 show the results from the regression model in Equation (3) for the horizons (h) 4 and 12. The measure of financial stress is the CISS. Bounds are computed using 1,000 bootstrap samples. The significance level is denoted as follows: *p < 0.1; **p < 0.05; ***p < 0.01.

risks.[13] To gauge the role of non-linearities and to assess the prevalence of risk sensitivity gaps, we augment the model with interaction terms of structural characteristics and the two included financial risk indicators (columns 3–4).

Considering the effects of financial stress on the one hand, as measured by the CISS, and credit growth on the other, we observe that the impact of financial stress is particularly significant and pronounced over shorter horizons, while the role of credit growth becomes more important as h increases. This pattern is well documented in the literature (see, for instance, Adrian et al. 2022).

The coefficients on structural characteristics in the linear specification shown in columns 1–2 provide an intuition on the overall effects of openness, financial sector size, public expenditures, and government effectiveness. For $h = 4$, we document significant adverse effects of trade openness on the predicted 10th percentile of the conditional one-year-ahead forecast of GDP growth. By contrast, public expenditures and government effectiveness exert significant positive effects on short-term growth risks and tend to stabilize the economy. Government effectiveness appears to play a particularly important role: a one-standard-deviation surge in government effectiveness is associated with an increase in the 10th percentile of projected GDP growth in $h = 4$ by approximately 3 percentage points. The two effects broadly confirm the findings of previous literature focusing on the link between public spending ratio and output volatility (Galí 1994; Fatás and Mihov 2001) on the one hand and the effect of government effectiveness on the other (Evrensel 2010). The coefficient on financial sector size is not significant for $h = 4$.

As h increases, we observe significantly adverse effects of openness and financial sector size. While both characteristics are associated with higher growth risks (i.e., lower GaR), larger financial sectors are particularly detrimental. An increase in financial sector size by one standard deviation is associated with a decrease in the lower tail of projected GDP growth by more than 2 percentage points. While the effect of trade openness is well in line with the findings of previous literature, which suggests a positive link between GDP volatility and openness (see, for instance, di Giovanni

[13] Below, we elaborate on the term structure of GaR, discussing coefficients from $h = 1, \ldots, 16$.

and Levchenko 2009), the role of large financial sectors is somewhat more surprising, as most empirical studies indicate a dampening effect of more developed financial sectors on GDP volatility (e.g., Manganelli and Popov 2015). Previous literature also suggests, however, that this effect weakens or even reverses at high levels of financial depth, as large financial sectors may amplify consumption and investment volatility (Dabla-Norris and Srivisal 2013). Earlier studies thus point to a non-linear link between financial sector size and GDP volatility, which is also consistent with recent findings in the finance-growth nexus literature (see, for instance, Breitenlechner, Gachter, and Sinderman 2015). According to our findings, the negative effect of financial sector size seems to dominate in the GaR framework, although we will argue below that the link between the two variables is ambiguous depending on the forecasting horizon (see Section 3.2). With an increasing forecasting horizon, the effect of the public spending ratio becomes insignificant, suggesting that the stabilizing role of higher public expenditures works only in the short term. Furthermore, the effect of government effectiveness appears to diminish somewhat but nevertheless plays an important role for a forecasting horizon of three years ($h = 12$).

Next, we consider the interaction terms of the structural characteristics in the regression model to account for non-linearities in the effects of the explanatory variables on GaR (columns 3–4). For $h = 4$, we observe significant and positive coefficients on interactions with financial stress for financial sector size and public expenditures, indicating that these factors mitigate the adverse effects of financial stress on projected growth vulnerabilities to some extent as risks increase. We also observe significant and negative coefficients of openness and government effectiveness interacted with credit growth. This is to some extent surprising, as credit growth usually plays a secondary role in shaping short-term growth vulnerabilities. This finding suggests an overall negative effect of openness on short-term vulnerabilities, at least for countries with buoyant credit growth. A possible explanation for this effect is that more open economies typically also exhibit higher levels of financial openness, with high rates of credit growth possibly depending on cross-border wholesale funding. On the other hand, the negative interaction term for government effectiveness and credit growth indicates that the effect of higher levels of government effectiveness is less pronounced in the face of excessive credit growth.

Regarding the role of interactions in shaping medium-term projected growth risks, we observe significantly negative coefficients of the interactions between credit growth and openness as well as financial sector size. While the detrimental effects of openness on growth vulnerabilities become more pronounced with higher credit growth, probably for the same reasons explained above, the negative effect of financial sector size is somewhat mitigated with higher credit growth. In this context, a larger financial sector could be associated with lower dependencies on cross-border funding, thereby mitigating risks linked to higher credit growth. From this perspective, the stabilizing role of more developed financial sectors, as suggested in the literature (e.g., Beck, Lundberg, and Majnoni 2006), becomes more relevant in an environment of high credit growth. The coefficient on the interaction term is, however, relatively small in magnitude, suggesting a limited role of non-linearities associated with financial sector size.

In Table 2, we replicate the estimations from above using the CLIFS instead of the CISS as a measure of financial stress. In contrast to the CISS, which is an aggregate measure of financial stress, the CLIFS is country specific (Duprey, Klaus, and Peltonen 2017). We consider the CLIFS to take into account that structural characteristics may affect not only the transmission of financial stress but also its country-specific emergence.

Considering columns 1–2, it appears that the overall effects of structural characteristics on projected growth vulnerabilities are not sensitive to the financial stress measure used. Allowing for multiplicative terms in 3–4, however, we observe some differences in how financial stress and structural country characteristics interact in shaping short-term risks. While we have observed that the effects of financial sector size and public expenditure are mitigated in instances of high financial stress using the CISS, this effect becomes insignificant once we consider the CLIFS for $h = 4$. Considering the CLIFS, the interaction with openness becomes significant for $h = 12$, thus indicating that the adverse effects of trade openness diminish to some extent with increasing levels of financial stress. Estimates shown in Table 2 suggest that structural characteristics also affect the transmission of country-specific financial stress and

Table 2. Main Results—CLIFS and Credit Growth

	Model 1		Model 2	
	$h = 4$ (1)	$h = 12$ (2)	$h = 4$ (3)	$h = 12$ (4)
CLIFS	–1.082***	–0.036	–1.222***	–0.096
	(0.256)	(0.056)	(0.310)	(0.082)
Credit Growth	–0.559	–0.729***	–0.724**	–0.999***
	(0.365)	(0.188)	(0.368)	(0.140)
Current GDP Growth	0.046	–0.577***	0.062	–0.558***
	(0.074)	(0.046)	(0.084)	(0.043)
Openness	–1.588***	–0.885***	–2.028***	–1.208***
	(0.581)	(0.291)	(0.700)	(0.231)
Financial Sector	–1.220	–2.856***	0.093	–2.919***
	(1.293)	(0.787)	(1.596)	(0.714)
Public Expenditure	1.142***	0.083	1.201***	0.105
	(0.294)	(0.129)	(0.387)	(0.129)
Government Effectiveness	5.224***	2.254***	6.039***	2.632***
	(0.710)	(0.409)	(0.905)	(0.353)
Openness × CLIFS			0.340	0.214**
			(0.264)	(0.098)
Financial Sector × CLIFS			0.012	0.021
			(0.367)	(0.069)
Public Expenditure × CLIFS			0.346	0.052
			(0.252)	(0.076)
Government Effectiveness × CLIFS			0.341	0.084
			(0.267)	(0.071)
Openness × Credit Growth			–0.589*	–0.814***
			(0.310)	(0.141)
Financial Sector × Credit Growth			0.303	0.201***
			(0.214)	(0.067)
Public Expenditure × Credit Growth			–0.187	–0.125
			(0.206)	(0.140)
Government Effectiveness × Credit Growth			–1.221***	–0.097
			(0.431)	(0.146)
Observations	1,740	1,572	1,740	1,572

Note: The table shows the estimated coefficients of the conditional 10 percent quantile. Columns 1–2 show the results from the regression model in Equation (1) for the horizons (h) 4 and 12. Columns 3–4 show the results from the regression model in Equation (3) for the horizons (h) 4 and 12. The measure of financial stress is the CLIFS. Bounds are computed using 1,000 bootstrap samples. The significance level is denoted as follows: *$p < 0.1$; **$p < 0.05$; ***$p < 0.01$.

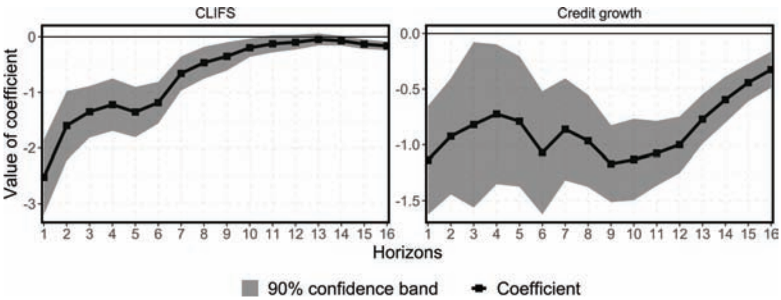
are generally robust to the variants of financial stress measures used in the analysis.

Overall, our findings clearly suggest that structural country characteristics play an important role in shaping variations in GaR. Over short-term horizons, we observe the stabilizing effects of public spending ratio and government effectiveness, with the latter being particularly pronounced. Considering interaction terms with financial stress, we observe that the stabilizing effect of public expenditures is particularly important when financial stress is high, whereas government effectiveness has a predominately linear effect on short-term GDP growth risks. Regarding medium-term growth risks, financial sector size and trade openness play an important and negative role in shaping growth vulnerabilities, while high levels of government effectiveness are still associated with higher GaR levels. Considering interactions with credit growth, we show that the adverse effects of larger financial sectors somewhat diminish with higher credit growth, while the negative effects of openness are further reinforced by increasing levels of credit growth.

The significant effects of structural characteristics, both with respect to GaR levels and the sensitivity of GaR to the underlying financial risk indicators, point to the prevalence of both structural and risk sensitivity gaps. In turn, our results have important macroprudential policy implications. Variations in the structural gap suggest that the appropriate macroprudential policy stance may, among other things, depend on structural characteristics, at least in an environment of homogeneous risk preferences across countries. Risk sensitivity gaps, as revealed by non-linearities in the effect of the included financial risk indicators depending on structural country characteristics, suggest that growth risks in some countries react more sensitively to increasing financial risks than in others. Thus, the appropriate reaction of macroprudential policy to variations in financial stress and credit growth also may depend on the respective (structural) country characteristics, as already suggested in theoretical considerations related to the GaR framework (Suarez 2022).

In the following sections, we focus on country-specific measures of financial stress (i.e., the CLIFS), primarily for two reasons. First, using country-specific financial stress measures is more common in previous literature (see, for instance, Aikman et al. 2019; Galán 2020; Adrian et al. 2022), thus facilitating a comparison of our empirical

Figure 2. Estimated Coefficients of the Financial Risk Indicators from 1 to 16 Quarters Ahead, Using the CLIFS as the Financial Stress Measure



Note: The figure shows the estimated coefficients of the financial risk indicators in the GaR estimation ($\tau = 0.1$) 1 to 16 quarters ahead. The black line represents the estimated coefficients, and the gray area shows the 90 percent confidence intervals; bounds are computed using 1,000 bootstrap samples.

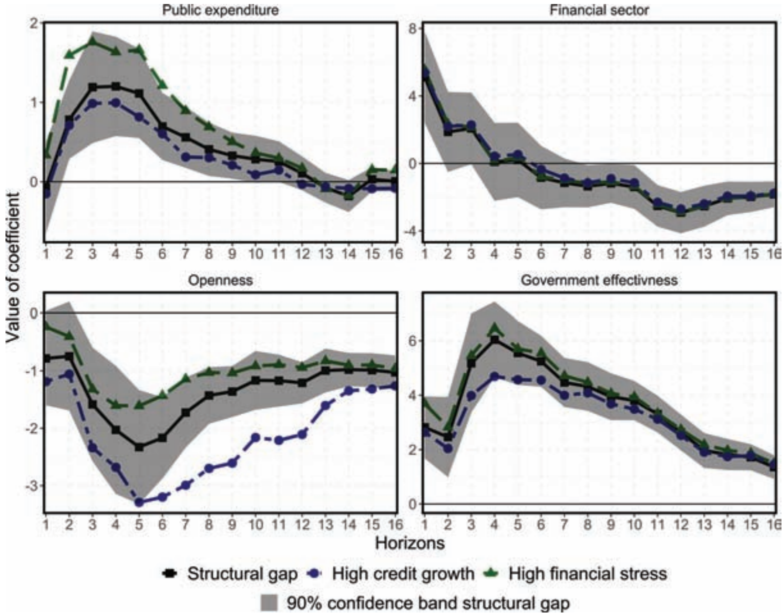
results to other studies. Second, using the CLIFS instead of the CISS is a more conservative approach to evaluate the effect of structural country characteristics on GaR, as those same factors may be associated with differences in financial stress across countries (i.e., more favorable structural country characteristics could be associated with lower contagion or higher resilience, thus resulting in more favorable financial stress at the individual country level).

3.2 Term Structure of GaR and Structural Characteristics

While the focus above is on the distribution of projected GDP growth one year ($h = 4$) and three years ahead ($h = 12$), we now extend our analysis to $h = 1, \ldots, 16$ quarters. Considering the effects of structural country characteristics on GaR for a series of forecasting horizons gives us an indication of how structural characteristics affect the term structure of GDP growth risks. Therefore, we extend the analysis by Adrian et al. (2022), who examine how financial conditions affect the term structure of GaR, to structural country characteristics.

We first evaluate how the two financial risk indicators affect the term structure of GaR. Figure 2 shows the evolution of the estimated coefficients of the CLIFS and credit growth $h = 1, \ldots, 16$ quarters

Figure 3. Estimated Coefficients for Structural Characteristics from 1 to 16 Quarters Ahead, Using the CLIFS as the Financial Stress Measure



Note: The figure shows the estimated coefficients of the structural characteristics in the GaR estimation ($\tau = 0.1$) 1 to 16 quarters ahead. The black line represents the estimated coefficients, and the gray area shows the 90 percent confidence intervals; bounds are computed using 1,000 bootstrap samples. The green and blue lines show the linear combination of the coefficients on the structural characteristics and the interaction terms with financial stress and credit growth (each evaluated at the 90th percentile).

ahead, based on the estimation of regression model (3). While we discuss these in more detail below, estimates using the CISS are shown in Figures A.1 and A.2 in the appendix. The gray area indicates the 90 percent confidence intervals. Consistent with previous literature (see, for instance, Aikman et al. 2019; Adrian et al. 2022), the CLIFS has the most adverse effects in the short term, while the negative impact of credit growth is economically and statistically significant for all time horizons.

In a similar vein, Figure 3 presents the evolution of the structural characteristics' coefficients for 1 to 16 quarters ahead. The black line shows the coefficients of the respective structural country characteristics. In addition, we show the coefficients of the structural characteristics plus the interaction term with financial stress (green line) and credit growth (blue line) evaluated at the 90th percentile of financial stress and credit growth, respectively. While the black (solid) line can be interpreted as a measure of the structural gap, the green and blue (dashed) lines point to the additional existence of risk sensitivity gaps in the case of strong deviations from the black line.

As already discussed above, higher public expenditures mitigate growth risks in the short run, as the respective coefficients are significantly positive from $h = 2$ to $h = 9$ (upper left panel in Figure 3). While we do not observe a risk sensitivity gap associated with increasing credit growth, higher financial stress can be mitigated to some extent by a high public spending ratio, once again pointing to a stabilizing role of larger public sectors in the short run.

Interestingly, the effect of the size of the financial sector strongly depends on the forecasting horizon, as evident in the upper right panel. In the very short run, larger financial sectors are associated with lower growth vulnerabilities but exercise strong detrimental effects on GaR in the medium run. Non-linearities therefore seem to play an important role not only in the finance-volatility nexus, as suggested by the literature (Dabla-Norris and Srivisal 2013), but also with respect to the GaR term structure. Interestingly, interactions with financial stress and credit growth do not play an important role in quantitative terms,[14] indicating that financial sector size is an important determinant of the structural gap but less so of the risk sensitivity gap.

The effects of openness are shown in the lower left panel of Figure 3. Non-linearities associated with increasing financial risk indicators are most pronounced with respect to openness. Notably, however, the impact of the two financial risk indicators, i.e., financial stress and credit growth, go in opposite directions. While openness mitigates growth risks in the face of high financial stress, growth risks

[14] As shown in Table 2, the interaction term is still statistically significant. Due to the large structural gap driven by financial sector size, however, the relatively small coefficient on the interaction term (i.e., the risk sensitivity gap) is hardly visible in this graphical illustration.

are amplified when credit growth is high. The figure clearly shows that trade openness is an important factor for both the structural and risk sensitivity gaps.

Finally, for government effectiveness, shown in the lower right panel, we see that this variable is a stabilizing factor for projected GDP growth, irrespective of the forecasting horizon. Interaction terms play an important role mostly with respect to high credit growth, and in particular for $h = 3, \ldots, 7$, giving rise to a risk sensitivity gap. Clearly, higher levels of government effectiveness are associated with a marked positive structural gap throughout the forecasting horizon.

Overall, considering a series of forecasting horizons, we document that structural country characteristics do strongly affect the term structure of GaR from the short to the medium run. However, the effects of structural characteristics across different forecasting horizons draw a rather heterogeneous picture. While public expenditures tend to affect projected growth risks in the short run, openness is more important at higher forecasting horizons. Government effectiveness has pronounced effects over forecasting horizons of at least three years, while a larger financial sector has mitigating effects over the short run but amplifies growth risks in the medium run.

3.3 Sensitivity Analyses

From the analysis above, it becomes evident that the structural characteristics of the public and the financial sector are important determinants of GaR. We now assess the sensitivity of our results by including different measures for public and financial sector characteristics.[15]

To capture financial system characteristics and potential systemic risk factors, we have focused so far on indicators of financial sector size. Evaluating a further aspect that has been identified as a potentially important aspect of the financial system (see also ESRB 2021) and complementing the analysis above, we now turn to banking sector concentration. Market concentration might be of particular interest from a policy perspective, as the related risks can be more readily addressed than in the case of a large financial sector.

[15]The corresponding Figures A.3, A.4, and A.5, are shown in the appendix.

Specifically, we study the effect of bank concentration measured by the Herfindahl-Hirschman Index (HHI). In the panel quantile regression, financial sector size is replaced by the HHI. As expected, we find that higher banking sector concentration is associated with higher growth risk. This is evident through negative coefficients on GaR (black line), at least over a forecasting horizon of one year, where we observe significant effects. In addition, the inclusion of interaction terms with both financial stress and credit growth points to the prevalence of a risk sensitivity gap, at least for some forecasting horizons. In line with previous findings documented by the ESRB (2021), countries with highly concentrated banking sectors exhibit higher growth risks and are especially vulnerable in the event of high credit growth.

In addition, we evaluate whether the effects of financial sector size are potentially driven by specific segments of the financial sector by considering subsectoral aggregates. Specifically, we replace total financial sector size with the ratio of other financial intermediaries' assets (excluding monetary institutions, pension funds, and insurance companies) to GDP. While the effects in the short run are consistent, medium-run effects turn insignificant when considering the relative size of the non-bank financial sector, indicating that mainly the relative size of the banking and insurance sector drives medium-run growth risks.

Instead of considering the public spending ratio, we also replicate our analysis with debt-to-GDP ratios. We find very similar effects, i.e., higher levels of debt are associated with lower growth risks, even though we do not observe a pronounced risk sensitivity gap associated with public debt. While this result may sound counterintuitive, one has to take into account that public expenditures and public debt are correlated (the correlation coefficient amounts to 0.44 in our sample). In other words, the (stabilizing) effect of flows seems to exceed the (potentially destabilizing) impact of stocks, i.e., high debt levels. This is not truly surprising, as GaR measures cyclical risks, and higher levels of public expenditures are likely associated with stronger automatic stabilizers on the one hand and also may enable the respective government to enact more effective countercyclical fiscal policies in a crisis, leading to lower cyclical downside risks. Against this background, it is conceivable that variations in public debt mainly reflect the variation in public expenditures rather than fiscal sustainability or fiscal space issues.

Overall, it appears that the relevance of country characteristics does not depend on the specific measure, confirming the importance of structural country characteristics in the GaR framework.

4. Model Evaluation

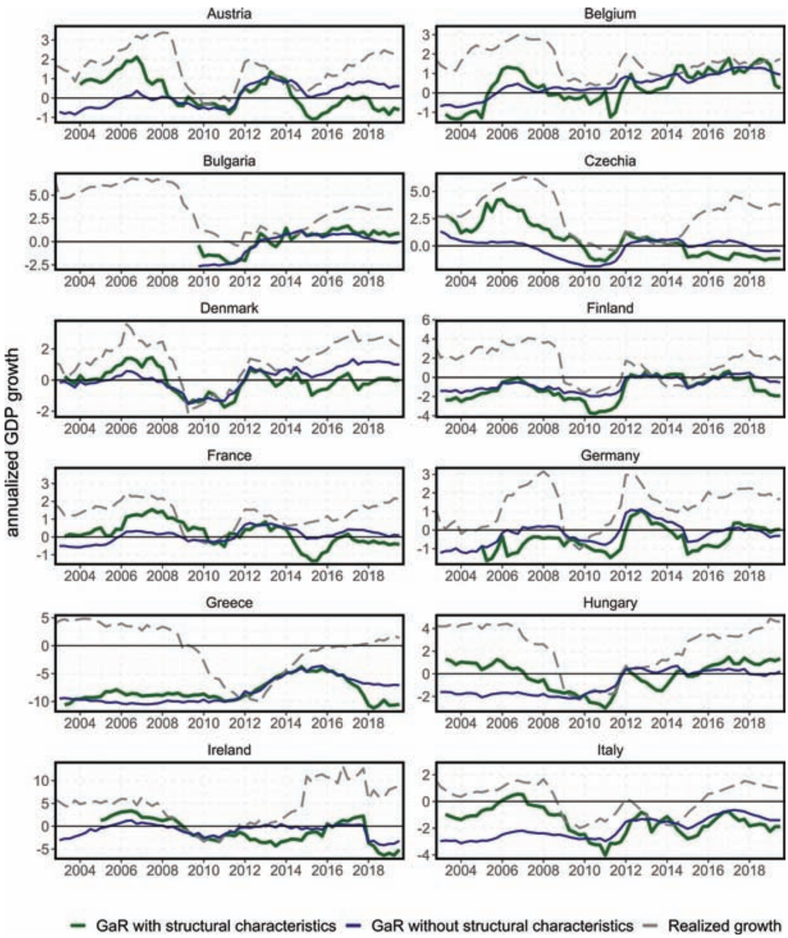
To evaluate the implications of the inclusion of structural characteristics for GaR, we next discuss the predicted values of the 10th percentile of projected GDP growth with and without structural characteristics. Moreover, building on predicted GaR, we examine potential forecasting gains by taking structural characteristics into account and consider model evaluation exercises.

Figures 4 and 5 show the predicted GaR three years ahead, estimated with and without country characteristics.[16] The gray (dashed) line is the realized annualized growth rate, the blue line represents the predicted GaR without taking into account structural characteristics, and the green line specifically considers structural characteristics in the form of Equation (3). Since we show the 10th percentile of projected GDP growth, realized GDP growth should be above the predicted GaR values approximately 90 percent of the time. To facilitate the interpretation of the figures, the predicted GaR is shifted forward to align the growth predictions with realizations for the respective quarters. Depending on data availability, the series of predicted GaR starts later for some countries.

While Figures 4 and 5 reveal the importance of structural country characteristics when estimating GaR, a detailed discussion of individual countries would clearly go beyond the scope of the paper. Generally, the effect of structural characteristics is both country and time specific. While, e.g., in Sweden, GaR values tend to be lower

[16]For the sake of brevity, we focus only on the GaR with a time horizon of three years. This perspective is probably more interesting for policymakers, as such a medium-term view may allow for a specific and appropriate policy reaction to increased systemic risks. We repeat the same analysis for GaR estimates one year ahead in the appendix, also confirming that structural country characteristics are important determinants of GaR, both with respect to the structural gap and the risk sensitivity gap.

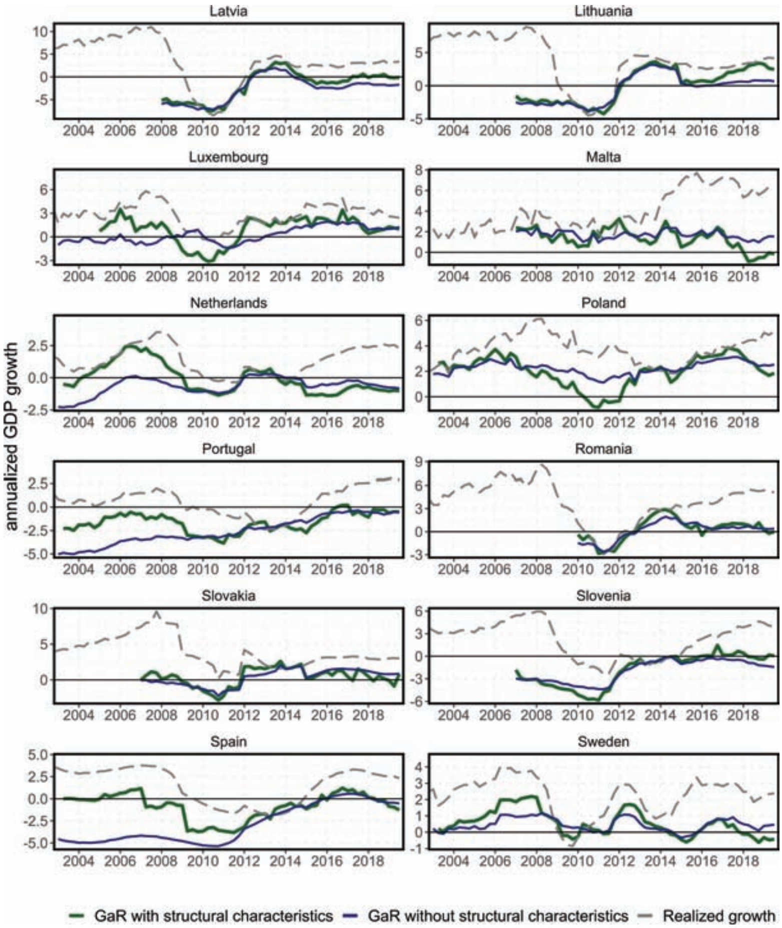
Figure 4. Predicted GaR Three Years Ahead with and without Structural Characteristics



Note: The figure shows the predicted GaR ($\tau = 0.1$) for a three-year forecasting horizon, estimated with and without the structural characteristics, together with realized GDP growth.

when structural characteristics are taken into account, the opposite holds true for, e.g., Malta. Moreover, it appears that in several countries, models incorporating structural characteristics predict higher values of GaR (e.g., Austria, Czech Republic, Hungary, Italy, Netherlands, Luxembourg, Portugal, Spain, Sweden) in the early

Figure 5. Predicted GaR Three Years Ahead with and without Structural Characteristics

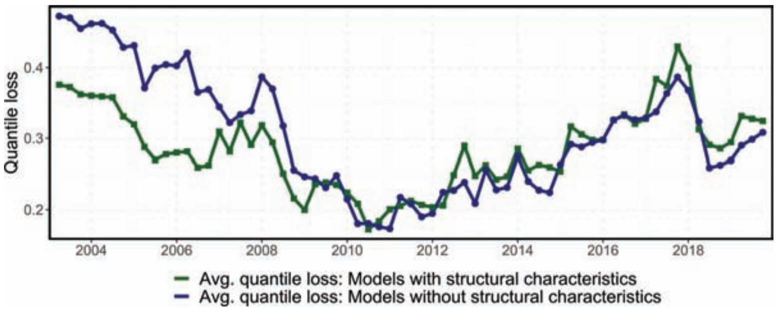


Note: The figure shows the predicted GaR ($\tau = 0.1$) for a three-year forecasting horizon, estimated with and without the structural characteristics, together with realized GDP growth.

2000s. However, after the global financial crisis, the wedge between predictions from models with and without country characteristics appears to shrink, indicating that both models perform similarly after the global financial crisis.

The fact that predicted GaR with and without structural characteristics deviates substantially, in particular in the run-up to the

Figure 6. Quantile Loss over Time



Note: This figure shows the average quantile loss of the panel model with and without structural characteristics for the forecasting horizon $h = 12$ over time.

financial crisis, begs the question of which model would have been more useful in identifying growth risks during that period. Considering in-sample average quantile loss across countries as an indication for the goodness of fit, it appears that the model with structural characteristics performs considerably better until the financial crisis.[17] Figure 6 shows the quantile loss over the sample period for $h = 12$. Especially in the run-up to the financial crisis, we observe substantially larger quantile loss for the model without structural characteristics, suggesting that a model incorporating structural characteristics is more accurate for this period.[18]

To evaluate the forecasting performance of GaR models that utilize country characteristics, we run several model evaluation exercises. We backtest the specification of the panel quantile regression model including country characteristics and interaction terms against two benchmarks using backtesting tools used in previous literature for forecasting horizons 1, 4, and 12 quarters (see, e.g., Brownlees and Souza 2021). As benchmarks, we consider a simple country-by-country quantile regression (BM 1), as shown in Figure 1,

[17]The overall quantile loss is evaluated below in further model evaluation exercises.

[18]Figure A.8 shows the average quantile loss across countries for $h = 4$. Even though quantile loss tends to be higher for the model without structural characteristics, differences in forecasting accuracy are overall less pronounced.

and a fixed-effects panel quantile regression without structural characteristics and interaction terms (BM 2).

First, we consider the unconditional coverage as the actual-over-expected ratio (AE ratio), i.e., incidences in which the actual GDP growth rate falls short of the respective GaR values (so-called hits) compared with the expected incidences implied by the quantile τ. A ratio below one means that the respective GaR model is too conservative and overestimates growth risks, while a ratio above one implies the opposite. Second, to further assess the in-sample goodness of fit, we use the dynamic quantile test (DQ test) of Engle and Manganelli (2004). The DQ test allows us to check for independence of hits in addition to the evaluation of the correct coverage.[19] Finally, we evaluate the average in-sample predictions based on the quantile loss function as an indication for forecasting accuracy, which also is frequently done in the context of value-at-risk evaluations (see, e.g., Gonzalez-Rivera, Lee, and Mishra 2004; Giacomini and Komunjer 2005; Brownlees and Souza 2021).

Table 3 reports the in-sample model evaluation, showing the three backtesting methods for forecasting horizons 1, 4, and 12 quarters for the two benchmark models and the panel specification with country characteristics (CC model). Generally, the CC model lies between the two benchmarks. Starting with the AE ratio, it becomes obvious that all models are too conservative and overestimate growth risks, except for the BM 1 model for $h = 1$, where it slightly underestimates growth risks. Over all predicted horizons, the AE ratio of the country-by-country model (BM 1) is closest to 1, whereas the fixed-effects panel quantile regression model overestimates GaR to the largest degree (BM 2). The specification with country characteristics also is too conservative but to a lesser extent than BM 2, indicating that taking country characteristics into account increases forecasting accuracy. Next, Table 3 shows the percentage of countries where the DQ test is not rejected at the 5 percent significance level. Generally, we see that the null hypothesis of an accurate model is rejected more often with increasing horizons. This means that the number of countries where GaR models are considered optimal by the DQ test shrinks with the forecasting horizon. While there is no difference

[19] Following Brownlees and Souza (2021), we use four lags of the hit sequence.

Table 3. In-Sample Model Evaluation

	AE Ratio			DQ Test			Average Quantile Loss		
	BM 1	**BM 2**	**CC Model**	**BM 1**	**BM 2**	**CC Model**	**BM 1**	**BM 2**	**CC Model**
$h = 1$	1.03	0.75	0.91	87.5%	87.5%	87.5%	0.68	0.82	0.76
$h = 4$	0.76	0.67	0.69	37.5%	29.2%	45.8%	0.57	0.63	0.61
$h - 12$	0.91	0.82	0.86	62.5%	20.8%	29.2%	0.22	0.30	0.28

Note: This table reports the in-sample model evaluation for the forecast horizons of 1, 2, and 12 quarters ahead and three models: the actual over expected ratio; the percentage of series for which the dynamic quantile test, based on the last four lags of the hit sequence, does not reject the null of model optimality at the 5 percent significance level; and the average quantile loss.

for $h = 1$, for $h = 4$, the specification with country characteristics performs best and is adequate for more countries than the two benchmark models. However, for the 12-quarters-ahead predictions, BM 1 is rejected to the lowest extent, while BM 2 is rejected most often. Finally, we compare the performance of each model according to the average quantile loss. We see that the BM 1 model performs 11 percent, 7 percent, and 22 percent better than the CC model, while the CC model improves the average quantile loss compared with BM 2 by 8 percent, 3 percent, and 7 percent, respectively. Hence, the CC model once again lies between the two benchmarks.

Our sample of 24 European countries is subject to considerable heterogeneity. As a result, it is not surprising that the panel regressions cannot keep pace with the country-by-country models in terms of forecasting accuracy, as the panel regression framework operates under the assumption that the countries load similarly on the financial risk measures employed. While this is important to document, we are mainly interested in the extent to which structural country characteristics affect growth risks. Against this background, the panel regression framework is warranted, especially in the context of a common regulatory framework across countries to implement an integrated modeling approach. Notably, the loss in accuracy can be reduced considerably when country characteristics are modeled explicitly.

5. Conclusion

The analysis in this paper aimed to understand the cross-country variation in growth vulnerabilities associated with financial stress and credit growth by putting a particular focus on the role of structural country characteristics. Our findings document that structural differences across countries play an important role in how financial risk indicators affect the projected distribution of future growth outcomes. By focusing on differences in trade openness, financial sector size, public spending ratio, and a measure of government effectiveness, we show that these structural characteristics not only lead to structural differences in GaR at the individual country level but also give rise to different reactions to varying levels of financial risk. Thus, our findings suggest the existence of both a structural gap in GaR due to structural country characteristics and a risk sensitivity gap, with structural differences across countries also leading to different degrees of responsiveness to varying financial risk indicators. Furthermore, our empirical results also show that structural country characteristics play a significant role in the context of the term structure of GaR, with the impact of the structural characteristics varying with the respective time horizons.

Our findings have important policy implications, particularly for macroprudential surveillance and the calibration of the respective policy tools. Since both the level of systemic risks, as measured by the structural gaps, and the responsiveness of systemic risk to changes in the financial risk indicators, as measured by the risk sensitivity gap, crucially depend on structural country characteristics, such cross-country differences should be explicitly considered both in the risk assessment and in the design of macroprudential policy. For example, a larger financial sector is, ceteris paribus, associated with generally higher growth risks. Thus, countries with large financial sectors may need a tighter macroprudential policy stance to mitigate possible downside risks to the same extent. However, as the negative effects of financial sector size diminish with higher financial stress (short term) and credit growth (medium term), growth risks in these countries will react less sensitively to surges in the respective financial risk indicators. Similar reflections can be made with regard to the remaining structural country characteristics that were examined in this paper. Furthermore, taking into account structural country characteristics in the transmission of financial risks, both in terms of GaR levels and sensitivity to the examined financial risk indicators, also may facilitate the use of the concept to assess the macroprudential policy stance at the individual country level. In fact, model evaluation exercises show that taking into account country characteristics helps to accurately identify and predict growth risks.

To make the GaR framework more readily applicable in a policy context, further research is necessary both in examining other potentially important structural determinants of GaR and concerning cross-country differences with regard to the sensitivity of GaR to policy measures, i.e., the policy sensitivity gap. The latter task is particularly challenging, as macroprudential policy is difficult to measure due to the multidimensional nature of the respective toolbox, and experience in applying many of those instruments is still limited.

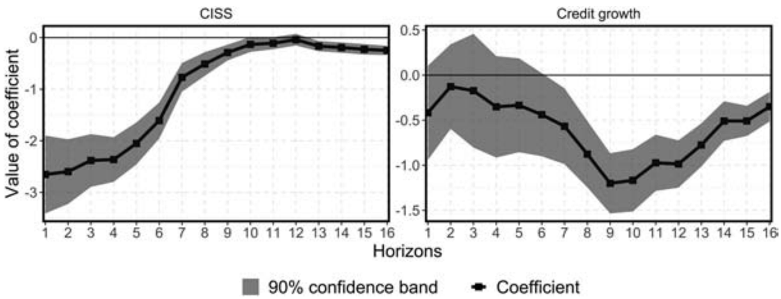
Appendix. Additional Table and Figures

Table A.1. OFR FSI and Credit Growth

	Model 1		Model 2	
	$h = 4$ (1)	$h = 12$ (2)	$h = 4$ (3)	$h = 12$ (4)
OFR FSI	–2.326***	0.081	–2.415***	0.051
	(0.199)	(0.058)	(0.165)	(0.061)
Credit Growth	0.041	–0.819***	0.163	–1.034***
	(0.351)	(0.185)	(0.348)	(0.166)
Current GDP Growth	0.070	–0.576***	0.083	–0.539***
	(0.052)	(0.054)	(0.072)	(0.051)
Openness	–2.974***	–0.738***	–2.885***	–1.149***
	(0.482)	(0.273)	(0.429)	(0.241)
Financial Sector	–1.162	–2.672***	–0.756	–2.714***
	(1.012)	(0.809)	(1.166)	(0.649)
Public Expenditure	0.645***	0.025	0.835***	0.109
	(0.215)	(0.142)	(0.307)	(0.127)
Government Effectiveness	3.862***	2.373***	4.600***	2.415***
	(0.615)	(0.414)	(0.633)	(0.390)
Openness × OFR FSI			–0.069	0.032
			(0.182)	(0.068)
Financial Sector × OFR FSI			0.191	0.022
			(0.223)	(0.084)
Public Expenditure × OFR FSI			0.871***	–0.012
			(0.236)	(0.082)
Government Effectiveness × OFR FSI			0.232	0.111**
			(0.159)	(0.055)
Openness × Credit Growth			–0.480*	–0.618***
			(0.267)	(0.159)
Financial Sector × Credit Growth			0.164	0.200***
			(0.175)	(0.060)
Public Expenditure × Credit Growth			0.007	–0.078
			(0.131)	(0.150)
Government Effectiveness × Credit Growth			–0.938**	–0.059
			(0.399)	(0.139)
Observations	1,744	1,576	1,744	1,576

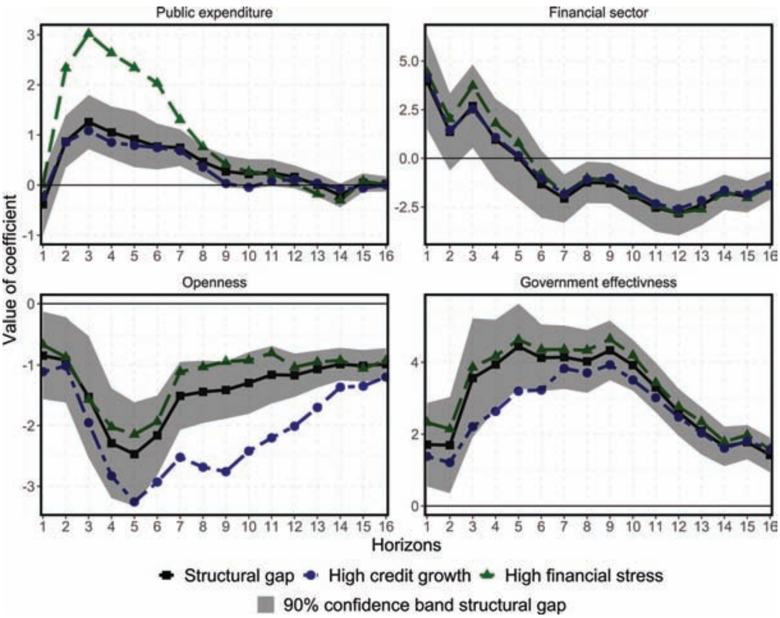
Note: The table shows the estimated coefficients of the conditional 10 percent quantile. Columns 1–2 show the results from the regression model in Equation (1) for the horizons (h) 4 and 12. Columns 3–4 show the results from the regression model in Equation (3) for the horizons (h) 4 and 12. The measure of financial stress is the Office of Financial Research Financial Stress Index (OFR FSI). Bounds are computed using 1,000 bootstrap samples. The significance level is denoted as follows: *p $<$ 0.1; **p $<$ 0.05; ***p $<$ 0.01.

Figure A.1. Estimated Coefficients of the Financial Risk Indicators from 1 to 16 Quarters Ahead, Using the CISS as a Financial Stress Measure



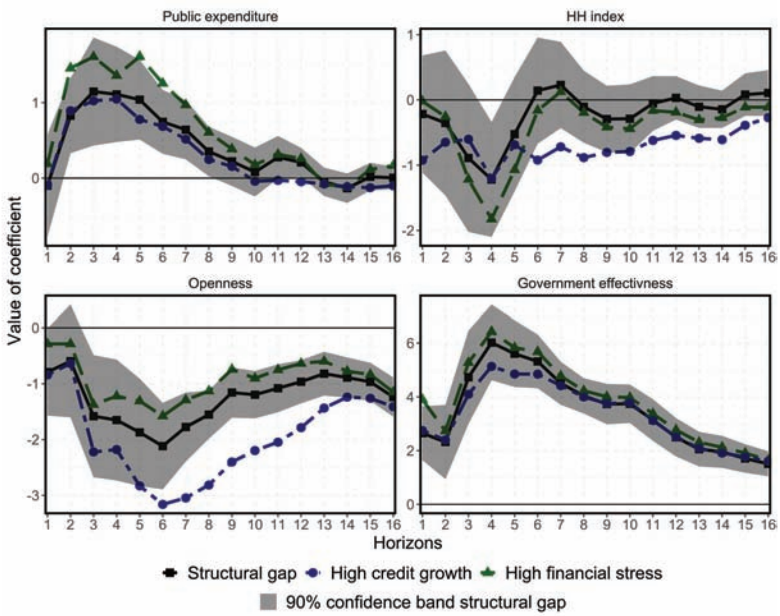
Note: The figure shows the estimated coefficients of the financial risk indicators in the GaR estimation ($\tau = 0.1$) 1 to 16 quarters ahead. The black line represents the estimated coefficients, and the gray area shows the 90 percent confidence intervals; bounds are computed using 1,000 bootstrap samples.

Figure A.2. Estimated Coefficients of the Structural Characteristics from 1 to 16 Quarters Ahead, Using the CISS as a Financial Stress Measure



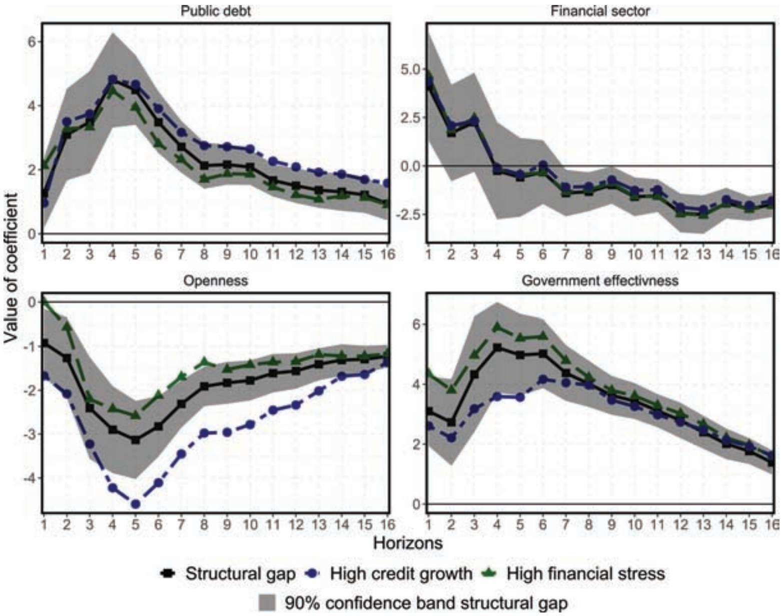
Note: The figure shows the estimated coefficients of the structural characteristics of the GaR ($\tau = 0.1$) 1 to 16 quarters ahead. The black line represents the estimated coefficients, and the gray area shows the 90 percent confidence intervals; bounds are computed using 1,000 bootstrap samples. The green and blue lines show the linear combination of the coefficients on the structural characteristics and the interaction terms with financial stress and credit growth (each evaluated at the 90th percentile).

Figure A.3. Estimated Coefficients of the Structural Characteristics from 1 to 16 Quarters Ahead, Using the CLIFS as the Financial Stress Measure



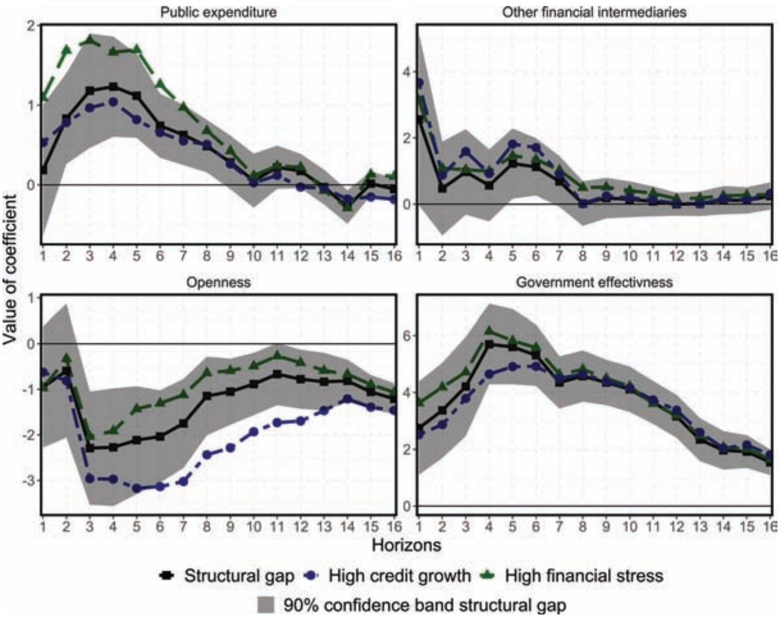
Note: The figure shows the estimated coefficients of the structural characteristics of the GaR ($\tau = 0.1$) 1 to 16 quarters ahead. Financial sector size is substituted with the HH index. The black line represents the estimated coefficients, and the gray area shows the 90 percent confidence intervals; bounds are computed using 1,000 bootstrap samples. The green and blue lines show the linear combination of the coefficients on the structural characteristics and the interaction terms with financial stress and credit growth (each evaluated at the 90th percentile).

Figure A.4. Estimated Coefficients of the Structural Characteristics from 1 to 16 Quarters Ahead, Using the CLIFS as the Financial Stress Measure



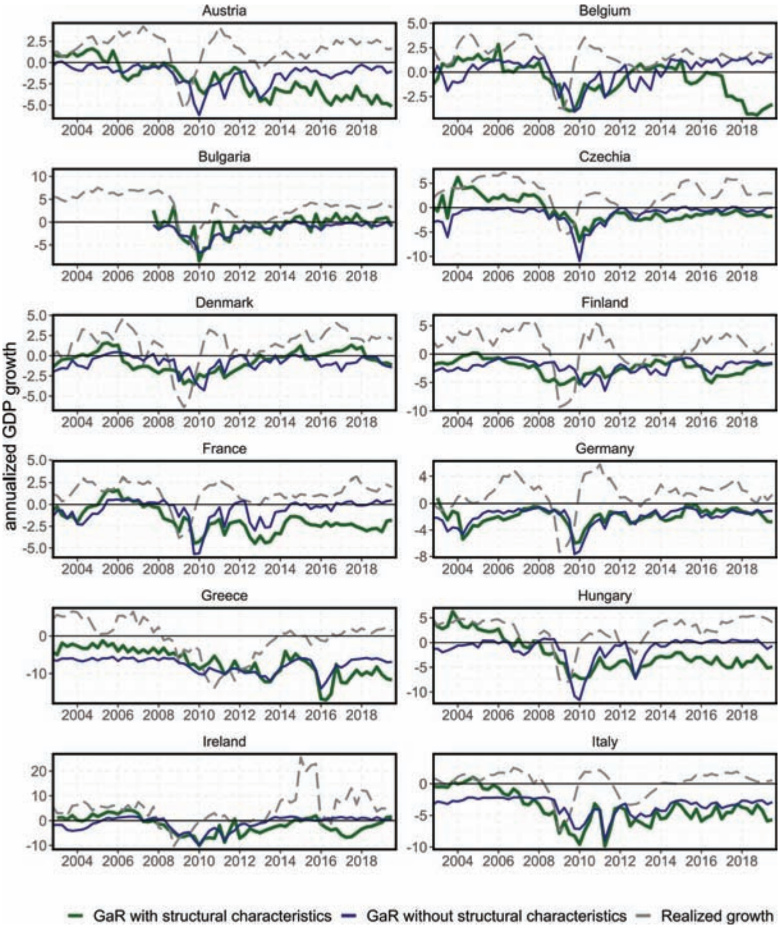
Note: The figure shows the estimated coefficients of the structural characteristics of the GaR ($\tau = 0.1$) 1 to 16 quarters ahead. Public expenditure is substituted with public debt. The black line represents the estimated coefficients, and the gray area shows the 90 percent confidence intervals; bounds are computed using 1,000 bootstrap samples. The green and blue lines show the linear combination of the coefficients on the structural characteristics and the interaction terms with financial stress and credit growth (each evaluated at the 90th percentile).

Figure A.5. Estimated Coefficients of the Structural Characteristics from 1 to 16 Quarters Ahead, Using the CLIFS as the Financial Stress Measure



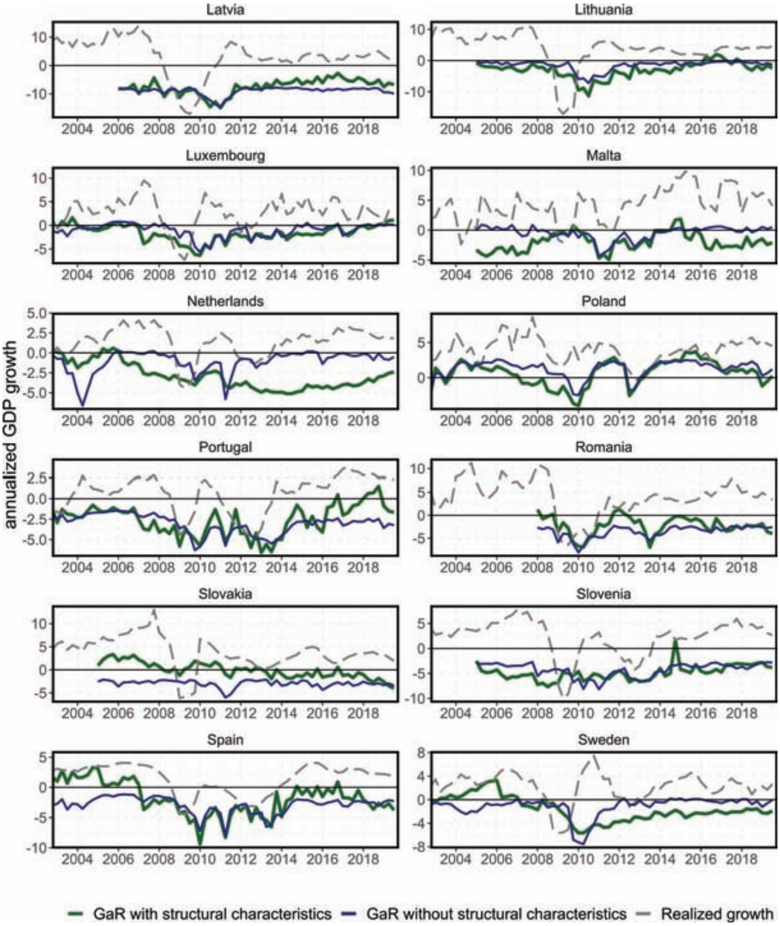
Note: The figure shows the estimated coefficients of the structural characteristics of the GaR ($\tau = 0.1$) 1 to 16 quarters ahead. Financial sector size is substituted with the ratio of assets of other financial intermediaries (excluding monetary institutions, pension funds, and insurance companies) to GDP. The black line represents the estimated coefficients, and the gray area shows the 90 percent confidence intervals; bounds are computed using 1,000 bootstrap samples. The green and blue lines show the linear combination of the coefficients on the structural characteristics and the interaction terms with financial stress and credit growth (each evaluated at the 90th percentile).

Figure A.6. Predicted GaR One Year Ahead with and without Structural Characteristics



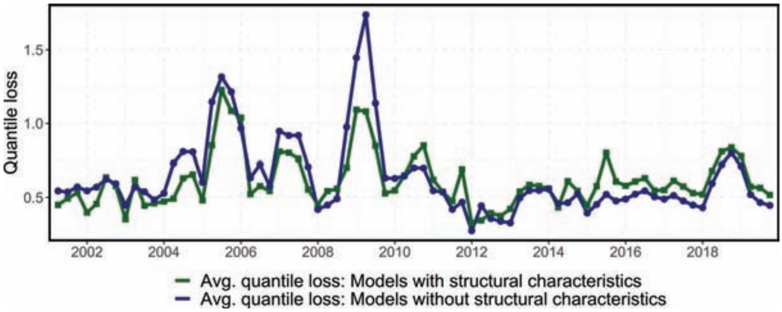
Note: The figure shows the predicted GaR ($\tau = 0.1$) for a one-year forecasting horizon, estimated with and without the structural factors, together with realized GDP growth.

Figure A.7. Predicted GaR One Year Ahead with and without Structural Characteristics



Note: The figure shows the predicted GaR ($\tau = 0.1$) for a one-year forecasting horizon, estimated with and without the structural factors, together with realized GDP growth.

Figure A.8. Quantile Loss over Time



Note: This figure shows the average quantile loss of the panel model with and without structural characteristics for the forecasting horizon $h = 4$ over time.

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Share Buybacks, Monetary Policy, and the Cost of Debt*

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Share buybacks have become common practice across U.S. corporations. This paper shows that firms finance these operations mostly through newly issued corporate bonds, and that the exogenous variation in the cost of debt—due to innovations in monetary policy—is key in explaining managers' incentives to repurchase their own shares. Under our identification strategy, we find that firms are more likely to repurchase in periods of accommodative monetary policy when the yield on their bonds adjusts in the same direction. This behavior has macroeconomic implications, as it diverts resources from investment and employment, thus reducing the transmission of accommodative monetary policy at firm level.

JEL Codes: E52, G11, G35, G32.

1. Introduction

Since 1985, U.S. corporations are allowed to buy back their own shares on the stock market. Very quickly buybacks have become common practice used to return cash to particular categories of investors,

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to send signals of confidence to markets, to concentrate firms' ownership, or also to adjust stock prices. However, these operations tend to divert resources from productive investments (Almeida, Fos, and Kronlund 2016) such that many raised concerns about the legitimacy of repurchase programs, particularly about the way managers use their financial resources and on the impact of buybacks on the real economy. These arguments became of interest to legislators and economists in the aftermath of the Great Recession, a period in which firms—despite having at their disposal substantial internal and external liquidity—devoted considerable resources to buybacks rather than to new investments and job openings.[1]

Much is already known about the negative effect of repurchases on real variables (Wang, Yin, and Yu 2021), on the market timing of repurchases (Stein 1996; Ma 2019; Baker and Wurgler 2002) and the reason why firms do buy back (Grullon and Michaely 2004; Hribar, Jenkins, and Johnson 2006). Yet, little is known about how firms finance this operation and to what extent the cost of financing affects managers' decision to buy back their own shares.

This paper aims to fill this gap in the literature and shows that buyback programs are mostly financed through new corporate debt issuance, and they are most likely and bigger in periods of accommodative monetary policy. In fact, for an exogenous fall in the federal funds rate, firms that benefit from a downward adjustment of their corporate yield tend to repurchase more by issuing more debt in the same quarter. Using low-cost debt to finance repurchases takes away resources from capital expenditures and new employment, thus reducing the effectiveness of accommodative monetary policy at firm level. The contribution of the paper is to properly quantify by how much the use of resources to repurchase programs is due to accommodative monetary policy, and to causally assess by how much the

[1]After the introduction of the stimulus package following the financial crisis of 2008, the Obama administration publicly opposed this managerial behavior since many bailed-out banks started using public money to buy back shares or compensate managers. Similar concerns were raised also in the aftermath of the current pandemic crisis. See, for example, the speech President Biden gave in Dunmore on July 9, 2020, or the most recent statement of Senator Elizabeth Warren of March 2, 2021. Both advocate for a reduction of repurchases, since they are primarily beneficial for managers and shareholders, they discourage new investments, and they hinder inequality reduction.

transmission of monetary policy on real variables is attenuated by the share buyback channel.

In light of this evidence, this paper not only unveils a new fact that informs on the use of share repurchases and the allocation of firms' financial resources, but it also highlights how these corporate decisions prevent a full transmission of an expansionary monetary policy on real variables. Hence, this work is also linked to a growing literature investigating how firm-level heterogeneity influences corporate dynamics and the transmission of macroeconomic shocks to the real economy (see, for example, Bloom, Bond, and Van Reenen 2007; Armenter and Hnatkovska 2011; Acharya, Almeida, and Campello 2013; Falato, Kadyrzhanova, and Sim 2013; and Bacchetta, Benhima, and Poilly 2014).

The first part of the paper shows some basic facts that motivated our investigation. First, we use corporate balance sheet data and provide evidence that, in the cross-section, firms use 75 cents of each dollar of newly issued debt to finance repurchase programs, whereas corporate cash plays a minor role. Second, we combine firm-level and macro data to show that repurchase programs are 3 percent more likely and 10 percent larger in periods in which a 1 percent negative monetary policy shock realizes. However, these estimates are biased due to the fact that there are many different channels through which monetary policy can operate and influence managerial decisions on share repurchases.

In the second part of the paper we deal with this problem. This is not a trivial task since the relationship between buybacks, monetary policy, and real variables is exposed to several sources of endogeneity: a firm can self-select into a repurchase program at any time and for reasons other than an exogenous change in the cost of debt. Similarly, there are factors—monetary policy included—that can simultaneously affect employment and investment such that the decision to repurchase, and the size of the buyback program, might be an endogenous outcome. To solve the endogeneity issue, quantify the correct effect of monetary policy on repurchase, and impute by how much the crowding out of buybacks on real variables is due to an accommodative monetary shock, we need a rigorous identification strategy. More specifically, we need an exogenous factor, orthogonal to firm characteristics and monetary policy itself, able to explain ex ante firms' repurchase behavior. This, in a first stage, would allow us to correctly evaluate how monetary policy influences buyback behavior by comparing the effect of monetary innovations between firms that are ex ante supposed to repurchase and those that are not. Thereafter, in a second stage, we can use this strategy to assess the causal crowding-out effect of repurchase on real variables, by how much an accommodative monetary policy exacerbates such effect and by how much share buybacks attenuate the overall transmission of an accommodative monetary policy shock.

To do so, we exploit a discontinuity in the likelihood of repurchasing that is driven by management earnings considerations. As shown in Hribar, Jenkins, and Johnson (2006), firms whose earning-per-share (EPS) ratio is below the analysts' forecast are more prone to launch an accretive buyback program in order to meet markets' expectations, build credibility, and avoid markets' future punishment.[2] This maneuver allows us to split the sample of firms into a "treatment" group, i.e., those who need to adjust the EPS to meet the target, and a "control" group, i.e., those who do not need to adjust the EPS. Both groups are very similar in terms of leverage, size, cost of debt, growth opportunities, and financial constraints, and exhibit also similar dynamics in investments and employment before the EPS forecast is announced. Moreover, monetary policy shocks and the implied changes of corporate debt cost are not correlated anyhow with the EPS forecast. Hence, all the identifying assumptions for a regression discontinuity design hold and the distance from the EPS forecast is a valid predictor of repurchase behavior.

Under this strategy, first we study how an exogenous fall in the cost of corporate debt—as explained by a monetary policy shock—affects both groups of firms around the discontinuity at the moment of the EPS forecast announcement and show that it has a significant positive impact only for the "treatment" group. In other words, if a manager needs to repurchase to satisfy EPS market expectations, (s)he is more likely and capable to do so if, at the same time, (s)he benefits from a fall in the cost of debt, i.e., if (s)he can raise money at a low cost to finance this operation. In particular, from this analysis,

[2] An accretive buyback program is one that raises the EPS by more than the opportunity cost of not saving resources.

we find that a 1 percent exogenous fall in the 10-year corporate bond yield leads to an increase of 0.5 percent of repurchase among firms in the "treatment" group. Thereafter, by using the distance from the EPS forecasts and monetary policy shocks as instruments, we study the causal effect of repurchases, the cost of debt, and their interaction on real variables.

From this analysis, the result is that repurchases causally lead to a considerable crowding-out effect on future investments and employment, and any accommodative monetary policy shock lowering the corporate cost of debt exacerbates such effect. In particular, we find that—through the repurchase channel—a 1 percent fall in the corporate cost of debt leads to an extra decrease of investments by 11,200 dollars and 0.10 employees for every million dollars of a firm's assets. Such diversion of resources from real variables calls into question the effectiveness of monetary policy and its transmission at firm level. By doing a simple back-of-the-envelope calculation, we find that indeed buybacks attenuate the transmission of expansionary monetary policy and, if the repurchase channel was muted, the transmission of a 1 percent accommodative shock on investments and employment would be, respectively, 11 percent and 15 percent stronger.

Related Literature. This paper is related to three strands of literature. The first is the vast literature on share buybacks. This tells us that repurchases are typically conducted when firms have the private information that their stock price is undervalued (Ikenberry, Lakonishok, and Vermaelen 1995; Stein 1996; Brockman and Chung 2001; Peyer and Vermaelen 2008), when they lack future growth opportunities (Grullon and Michaely 2004), to signal confidence to markets on strong future performance (Hribar, Jenkins, and Johnson 2006), to increase employees' effort (Babenko 2009), to mitigate the dilutive effect of stock option exercises (Kahle 2002; Bens et al. 2003), or to distribute excess capital (Dittmar 2000). Moreover, we know that repurchase programs follow market timing. For example, firms repurchase when the value of equity is relatively low with respect to other sources of financing (Ma 2019; Baker and Wurgler 2002). Finally, Almeida, Fos, and Kronlund (2016) tell us that share buybacks crowd out future capital investment, employment, and R&D (research and development) investment. Also Lazonick (2014) goes in this direction and cites repurchases as a possible explanation for why, in the post-recession era, firms have high corporate profitability but low growth in employment.

The second strand of literature this paper relates to is on earnings and EPS management. Our identification strategy is based on the fact that managers care about meeting market expectations on earnings, and it is well known that repurchases can help boost the EPS index (see, among the many, Burgstahler and Dichev 1997, Skinner and Sloan 2002, and Graham and Harvey 2005).

Third, this paper relates to the growing literature studying the role of firm heterogeneity for the transmission of macroeconomic shocks and for the comprehension of macroeconomic dynamics. For example, and consistently with the results of this paper, Bacchetta, Benhima, and Poilly (2014) show that firms exploit liquidity shocks to hoard cash for precautionary purposes at the detriment of employment. In the same vein, Armenter and Hnatkovska (2011), Falato, Kadyrzhanova, and Sim (2013), Acharya, Almeida, and Campello (2013), and Bloom, Bond, and Van Reenen (2007) show the effects of firms' precautionary behavior when productivity and uncertainty shocks materialize. Others, like Jeenas (2018) and Melcangi (2018), show that demand shocks and monetary shocks heterogeneously affect firms' employment choice depending on the capital structure of the firm, the degree of financial constraint, and the level of liquidity.

This paper develops as follows: Section 2 documents the financing and the timing of repurchase programs; Section 3 explains the identification strategy to study the causal crowding-out effect of repurchases on real variables and to impute correctly the attenuation of accommodative monetary policy due to buybacks; in Section 4 we do robustness checks; Section 5 reconciles the empirical evidence with a simple model showing the conditions under which a fall in the cost of debt allows for accretive repurchases; Section 6 concludes.

2. Repurchases, Debt, and Monetary Policy

In this introductory section, we describe the data and provide some basic evidence on how share buybacks, debt issuance, and monetary policy are all related. In particular, we show three facts. First, firms finance repurchase programs by issuing new debt and cutting their capital expenditure. Second, the timing and magnitude of buyback programs are correlated with unanticipated changes in monetary policy: they are more probable and larger in periods of accommodative monetary policy. Third, monetary shocks have a firm-specific effect on debt issuance through changes in the yield on corporate bonds.

2.1 Data and Sample Selection

We use two types of data: firm-level data and macroeconomic data on monetary policy shocks. Firm microdata come from different sources. We use Standard and Poor's Compustat to extract firms' fundamentals data at quarterly frequency between 1985 and 2016, with the exception of employment data, which are available at yearly frequency only. Following Almeida, Fos, and Kronlund (2016), we exclude regulated utility firms (Standard Industrial Classification (SIC) codes 4800–4829 and 4910–4949) and financial firms (SIC codes 6000–6999) as well as firms with missing or negative assets. Thereafter, we merge the Compustat sample with analysts' forecast data from the Institutional Brokers' Estimate System (IBES). Finally, we use data from the Trade Reporting and Compliance Engine (TRACE) to extract firm-level yields on newly and previously issued corporate bonds.[3] Regarding monetary policy shocks, we follow the literature on structural vector autoregression (SVAR) and recent developments as in Rossi and Zubairy (2011) and Ramey (2016) to extract innovations on the fund rate.[4] Table 1 shows summary statistics of the variables we use and describes their construction. In particular, as in Ma (2019), we define repurchases as the firm's net position on the equity market, i.e., difference between the value of the shares repurchased and the value of the newly issued shares normalized by total assets in the previous period. In this way, a negative value would stand for a net equity issuance, while a positive value would stand for a net equity repurchase. As the first panel of Table 1 reports, 24 percent of firms are net repurchasers across quarters. Among them, on average 3.1 percent of assets are repurchased every period with an average cash flow of 38 million dollars.

[3]Firm-level yields are calculated using equal weighted average on the different bonds issues of the same maturity.

[4]See Appendix A for details on the SVAR model we use to extract monetary policy shocks and its identifying assumptions.

Table 1. Descriptive Statistics

	Mean	SD	p1	p5	p25	p50	p75	p95	p99	N
Repurchase Statistics										
$\mathbb{I}(Repurchases > 0)$	.24	.43	0	0	0	0	0	1	1	831,649
For $\mathbb{I}(Repurchases > 0) = 1$										
Repurchases/Assets	.03	.06	.00	.00	.00	.01	.03	.12	.30	204,794
Repurchases ($M)	38.27	88.90	.00	.02	.36	3.141	25.56	229.51	474.184	204,794
EPS Distance Statistics										
Distance (%)	–.07	1.91	–7.33	–3.34	–.48	.04	.61	2.30	5.91	196,378
$\mathbb{I}(Distance \geq 0)$	.54	.49	0	0	0	1	1	1	1	196,378
$\mathbb{I}(Distance < 0)$	.46	.49	0	0	0	0	1	1	1	196,378
Firm Characteristics										
Market Cap. ($M)	2,630	14,901	.33	2.02	22.70	141.86	876.99	9,428	46,011	248,137
Market-to-Book	3.46	4.98	.15	.41	1.11	1.98	3.65	11.46	28.16	211,214
Assets ($M)	1,946	13,13	.02	.91	14.43	84.68	507.06	6,533	34,235	831,649
Money/Assets	.17	.19	.00	.00	.03	.09	.23	.63	.88	223,742
Profits/Assets	–.01	.19	–.79	–.42	–.03	.02	.06	.17	.33	586,650
Debt/Assets	.23	.20	.00	.00	.05	.18	.34	.65	.87	562,305
Investments/Assets	.04	.07	.00	.00	.00	.02	.05	.17	.40	723,171
Employment/Assets	16.98	22.01	.08	.41	2.44	5.72	12.12	36.44	110.49	653,749
10-Year Yield (%)	5.20	2.52	1.38	2.07	3.41	4.88	6.34	9.59	14.67	48,560
Q	2.43	2.61	.35	.65	1.09	1.55	2.58	7.51	14.97	234,911
*PE*10	21.74	352.78	.03	.15	1.10	3.50	10.65	52.94	212.20	95,314
$\mathbb{I}(Dividend > 0)$	.16	.37	0	0	0	0	0	1	1	831,649
Fin. Constraint	–2.58	0.72	–3.27	–3.25	–3.11	–2.79	–2.28	–1.27	–.15	351,375
Monetary Innovations										
Shock (%)	0.00	0.11								128

Note: All variables are built on quarterly data, with the exception of employment. *Repurchases* is the difference between stock purchases and stock issuances (in $M). *Distance* is the difference between the reported EPS and the median EPS forecast at the end of the quarter, normalized by the end-of-quarter stock price. *Market-to-Book* is the market value of common equity divided by the book value of common equity. *Money* is the total value of cash holdings (in $M). *Profits* is defined as net income plus depreciation (in $M). *Debt* is the value of total debt (in $M). *Investments* equates capital expenditure (in $M). *Employment* is the stock of employees (in Ks). This variable is available only at yearly frequency, but we use it at quarterly frequency when normalized by the (quarterly) value of assets. *Yield* is the firm's yield on a 10-year-maturity corporate bond. *Q* is the book value of liabilities plus the market value of common equity divided by the book value of assets. The measure of *Fin. Constraint* follows Hadlock and Pierce (2010). *PE*10 is the 10-quarter moving average of the price-earning ratio. *Shock* is the monetary shock obtained from an SVAR (see Appendix A).

The second panel reports statistics on firms' ex post EPS distance from the analysts' target and frequency for a (weakly) positive and negative distance from the target. Such distance is measured as the difference between the EPS forecast and the end-of-the-quarter EPS as reported by the firm. The (price-normalized) average distance is negative and 0.07 percent off the median analysts' consensus. Across quarters, 54 percent of the time firms are on target or above (i.e., they are reporting an end-of-the-quarter EPS at least as big as the forecast) while 46 percent of the time they are below the target. The third panel reports on other firm characteristics like market capitalization, the market-to-book value of the firm, assets, internal and external financial resources (cash holdings, profits, debt issuance), investments, employment, the cost of debt (measured as the yield on a 10-year corporate bond), the Q-value, the 10-quarter moving average of the price-earning ratio (PE10), an indicator on whether the firm has paid dividends in the previous four quarters, and a measure for financial constraint (built after Hadlock and Pierce 2010). The fourth panel reports the mean and standard deviation of monetary policy innovations as extracted from the SVAR. These shocks have mean zero and standard deviation equal to 11 basis points, very similarly to the (quarterly aggregated) high-frequency monetary shocks identified in Gertler and Karadi (2015) and Ottonello and Winberry (2018).

2.2 The Financing of Share Buybacks

How are share buybacks financed? We answer this question by considering the following "accounting equation" for firms conducting a (positive net) share repurchase:[5]

$$Repurchases_{i,t} = \beta_1 \Delta Debt_{i,(t,t-1)} + \beta_2 \Delta Cash_{i,(t,t-1)} + \beta_3 Investments_{i,t} + \beta_4 Dividends_{i,t} + \epsilon_{i,t}.$$

Under this specification, we want to understand how much of each dollar (of assets) that the firm spends on repurchases is financed through the change in debt (β_1), the change in cash holdings

[5]We consider only firms for which the difference of the value of the shares repurchased and the value of the newly issued shares is positive.

Table 2. Financing Buybacks

	***Repurchases* (1)**	***Repurchases* (2)**	***Repurchases* (3)**	***Repurchases* (4)**	***Repurchases* (5)**
$\Delta Debt$	0.75***	0.79***	0.40***	0.40***	0.40***
	(0.11)	(0.07)	(0.05)	(0.05)	(0.05)
$\Delta Cash$		–0.23	–0.00	–0.00	–0.00
		(0.31)	(0.03)	(0.03)	(0.03)
Investments			–0.62***	–0.62***	–0.62***
			(0.04)	(0.04)	(0.04)
Dividends				0.06	0.06
				(0.05)	(0.05)
Observations	180,436	163,278	144,858	144,858	144,858
Time FE	No	No	No	No	Yes
Industry FE	No	No	No	No	Yes
Controls	No	No	No	No	No

Note: Standard errors are in parentheses, clustered at the firm level. The unit of observation *Repurchases* is the difference between the value of stock purchases and stock issuances from the statement of cash flows, and we consider only firms for which such difference is strictly positive. $\Delta Debt$ is the change in the value of current total debt of the firm. $\Delta Cash$ is the change in firm money holding plus current net profit. *Investments* is equal to capital expenditure. *Dividends* is equal to the value of the dividends paid. All variables are normalized by the value of total assets in $t - 4$. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

including net profits from the current quarter (β_2), a reduction in capital expenditure (β_3), or in dividend distribution (β_4). All variables are normalized by the level of assets in $t - 4$, and errors are clustered at the firm level. As reported in the first column of Table 2, unconditionally on other sources of financing, a \$1 repurchase is explained by \$0.75 of new debt issuance. Controlling for the change in cash holding and quarterly profits (column 2) does not affect the role of debt by much. Moreover, the estimate for $\Delta Cash$ is insignificant, suggesting that firms do not use their liquidity from cash holdings or newly generated net profits to finance this operation. This is consistent with the trend across U.S. corporations of hoarding cash for precautionary savings (see Acharya, Almeida, and Campello 2013 and Falato, Kadyrzhanova, and Sim 2013). When controlling for all other variables in the accounting equation, as well as time and industry fixed effects—as in columns 3 to 5—the contribution of debt drops but remains quite substantial at 40 cents. On the other hand, now repurchase expenditure looks mostly financed by subtracting resources from new capital investments (around 62 cents).[6]

2.3 Share Buybacks and Monetary Policy

The fact that debt is an important source to finance buybacks suggests that these corporate operations might be sensitive to changes in the cost of money, i.e., changes in monetary policy. To check this fact in the data, we consider the following regressions:

$$\mathbb{I}(Repurchases_{i,t} > 0) = \alpha + \beta Shock_t + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t} \quad (1)$$

$$Repurchases_{i,t} = \alpha + \beta Shock_t + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t}, \quad (2)$$

where $\mathbb{I}(Repurchases_{i,t+1} > 0)$ takes value one when the firm is a net repurchaser in quarter t, $Shock_t$ is the exogenous innovation on the fund rate as predicted by our SVAR, and $X_{i,t-1}$ controls for firm-level characteristics such as Q-value of investment, return (profit) on assets, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, and a dummy indicating the quintile of asset to which the firms belong.[7] This set of variables will help to take into account other factors influencing the decision and capability to repurchase, such as market valuation, profitability, payoff policies, and size. θ is a year-quarter fixed effect.[8]

Table 3 reports results for models (1) and (2). Since our measure of monetary policy shock is estimated from an SVAR, errors are double-clustered at firm level and date.[9] As from the specification in column 1, we find that a 1 percent exogenous fall of the federal funds rate leads to an increase in the probability of repurchase by 3 percent. For the model in column 2, we find that a 1 percent exogenous fall in the federal funds rate leads to a 16 percent increase in the size of the repurchase program. Therefore, we conclude that monetary

[6]In Appendix B we repeat the same analysis with variables in levels, and we further disaggregate the potential sources of financing. Yet, we find that debt issuance finances at least 35 cents of each dollar spent on buybacks.

[7]The set of control variables X will remain the same throughout the paper, if not otherwise specified.

[8]The year-quarter fixed effect implies controlling for a year dummy and a quarter dummy separately.

[9]From now until the end of Section 3, errors are always clustered at this level.

Table 3. Net Repurchases and Monetary Policy Shocks

	$\mathbb{I}(\boldsymbol{Repurchases} > \mathbf{0})$ **(1)**	***Repurchases*** **(2)**
Shock	–0.031*** (0.006)	–0.156** (0.065)
Observations	213,761	213,761
Time FE	Yes	Yes
Industry FE	Yes	Yes
Controls (X)	Yes	Yes

Note: Standard errors are in parentheses, double-clustered at firm level and date. In column 1, the unit of observation is $\mathbb{I}(Repurchases > 0)$, an indicator variable taking value one if the firm is a net repurchaser, i.e., the difference between equity repurchased and new equity issuance is positive. In column 2, the unit of observation *Repurchases* is the difference between the value of stock purchases and stock issuances from the statement of cash flows, normalized by total asset in t – 4. *Shock* is an exogenous monetary innovation as from an SVAR (see Appendix A for details). The set of controls X includes Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

policy shocks have an effect on both the propensity and the level of repurchase.

2.4 Corporate Yield and Monetary Policy

Although debt is important for the financing of share buybacks, it is not plausible to assume that common monetary shocks affect firms' capital structure and new debt issuance in the same way. This will ultimately depend on the responsiveness of the firm's bond yield to the shock. The following regressions investigate the effect of the unanticipated monetary shock on firm-level bond yields and debt issuance:

$$\Delta Yield_{i,t} = \alpha_1 + \beta_1 Shock_t + X'_{i,t-1}\gamma_1 + \theta_t + \epsilon_{i,t} \tag{3}$$

$$\Delta Debt_{i,t} = \alpha_2 + \beta_2 \Delta Yield_{i,t} + X'_{i,t-1}\gamma_2 + \theta_t + \nu_{i,t}, \tag{4}$$

where the variable $Yield_i$ is firm i's yield on a 10-year-maturity corporate bond.

Table 4. Corporate Bond Yield, Debt Issuance, and Monetary Shocks

	Δ *Yield* **(1)**	**Δ *Debt*** **(2)**	**Δ *Debt*** **(3)**
Shock	0.606***		
	(0.127)		
Δ *Yield*		–0.001**	
		(0.000)	
$\Delta \widehat{Yield}$			–0.004**
			(0.002)
Observations	41,624	41,624	41,624
Time FE	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes
Controls (X)	Yes	Yes	Yes
Estimator	OLS	OLS	2SLS

Note: Standard errors are in parentheses, double-clustered at firm level and date. In column 1, the unit of observation Δ *Yield* is the change in the firm's yield on a 10-year-maturity corporate bond. In columns 2 and 3, the unit of observation Δ *Debt* is the change in the value of current total debt of the firm, normalized by total asset in $t - 4$. *Shock* is an exogenous monetary innovation as from an SVAR (see Appendix A for details). $\Delta \widehat{Yield}$ is the exogenous change in the 10-year corporate yield as predicted by monetary policy shocks, i.e., when *Shock* is used as instrument for the change in the cost of debt. Column 3 reports 2SLS estimates for Equation (4). Control X includes return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

Column 1 of Table 4 shows results for regression (3), for which we find that an exogenous innovation of 10 basis points over the fund rate leads to an increase by 6.1 basis points of the 10-year yield. The results of regression (4) are reported in column 2, where we show that a 1 percent fall in the yield is associated with an increase of debt issuance by 0.1 percent. Since the relationship between debt issuance and changes in the yield is endogenous, we instrument $\Delta Yield$ of Equation (4) with the exogenous monetary innovations, i.e., we use Equation (3) as first stage to predict the exogenous change in the yield $\Delta \widehat{Yield}$. Then, we use the latter to explain the causal effect of an exogenous change of the yield on debt issuance. As reported in column 3, the (two-stage least-squares, 2SLS) estimator is four times larger: if the yield falls by 1 percent, the firm will issue 0.4 percent more debt.

3. Identifying the Effect of Monetary Policy on Real Variables through the Repurchase Channel

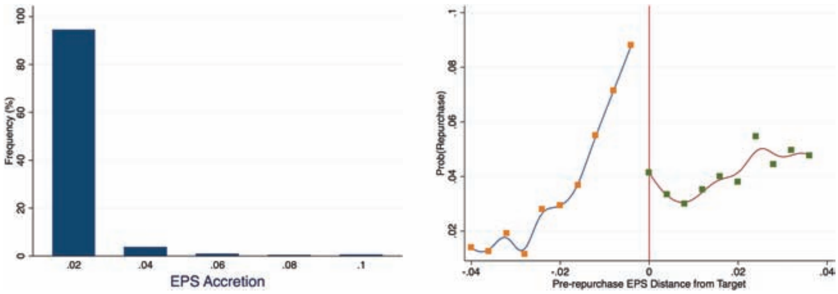
Despite the results of the previous section, it is important to stress that—when measuring the effect of monetary policy on the level of repurchase—our estimates are biased since monetary policy interacts with many firm's characteristics and time-varying variables (e.g., real variables) that can influence the size of the buyback program at the same time.

In fact, the option of buying back shares is always at managers' disposal and buybacks can happen for a long list of (endogenous) factors—such as poor growth prospects, lack of investment opportunities, or a need to adjust the balance sheet structure—that might correlate with monetary policy as well. Moreover, monetary policy might directly influence managers' choices over investment and employment such that repurchases might be a subsequent endogenous result. In other words, our identification is exposed to endogeneity problems mainly due to endogenous self-selection into a buyback program and reversed causality between repurchases and real variables (investment and employment). Therefore, monetary policy cannot explain alone repurchase behavior. In order to assess how much monetary policy encourages buybacks, first we must solve this issue. In particular, we need an exogenous factor, orthogonal to firm-level characteristics and monetary policy, capable to predict ex ante the repurchase behavior of the firm and to split the sample in two groups: repurchasers and non-repurchasers. This allows us to break the loop between monetary policy, repurchases, and real variables, and—more importantly—to assess correctly how exogenous changes in monetary policy affect both groups in their capability to repurchase. Finally, this strategy allows us to study the causal crowding-out effect of repurchase on real variables and evaluate by how much such crowding out is causally explained by accommodative monetary policy. By doing so, we can quantify the extent to which accommodative monetary policy transmits on real variables and by how much share buybacks reduce such transmission.

3.1 Identification Strategy

In order to overcome the endogeneity problem, we exploit a discontinuity in the level and probability of conducting an accretive repurchase. This discontinuity, first introduced by Hribar, Jenkins, and Johnson (2006) and more recently used in Almeida, Fos, and Kronlund (2016), exploits the misalignment between the firm EPS and the analysts' forecast. At the beginning of each quarter, analysts release their forecast for what the EPS of the firm will be at the end of that same quarter. Once the forecast is observed, managers decide whether to launch a buyback program to align their EPS at least with the level predicted by analysts. At the end of the quarter, firms announce their (adjusted) EPS along with information on the quantity and buying price of the repurchased shares. Hence, it is possible to reconstruct what the EPS would have been without repurchasing, i.e., the non-adjusted EPS (or the counterfactual) that would have prevailed without repurchasing. This information allows us to understand which firms were able to run an accretive buyback, by how much they were able to increase their EPS, and—for a given EPS forecast—which firms would have missed the EPS target without repurchasing. For example, say that analysts' EPS forecast for a certain firm is \$4 by the end of the quarter. For the same firm, we observe that the realized EPS is \$4.1 as announced at the end of the quarter. Thus, we check the number of shares held at the beginning of the period (say it was $N = 1{,}000$ million), the number of shares repurchased (say $n = 50$ million) and at what price (say $P = \$50$). Hence, we can build the forgone earnings due to buybacks as the opportunity cost of putting the amount $Pn = 2{,}500$ million into a deposit with a quarterly rate of $r^s = 5\%$ at the net of taxes (e.g., $\tau = 30\%$). In our example, the forgone net earnings are equal to $Pnr(1 - \tau) = 87.5$ million. Under this correction, the realized earnings (as reported at the end of the quarter) are equal to $4.1 * (1,000M - 50M) = 3,895$ million such that—if managers were not buying back their own shares—the EPS before adjustment would have been equal to $(3,895M + 87.5M)/1,000M = 3.98$ dollars per share. In this case, managers were able to beat the analysts'

Figure 1. EPS Accretion, Probability to Repurchase, and Distance from Target



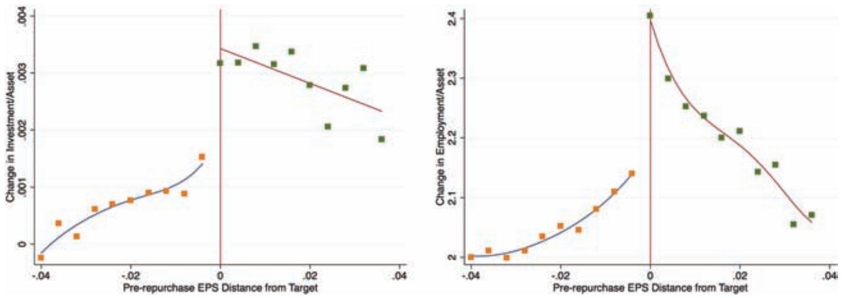
Note: The graph on the left-hand side plots the frequency of repurchases by EPS accretion bins. The accretion is the difference (normalized by the end-of-the-period stock price) between the adjusted EPS, i.e., the EPS as reported at the end of the quarter, and the EPS that would have prevailed if no buyback was conducted during the same quarter. The graph on the right-hand side plots the probability for a firm to buy back its own shares as a function of the distance of the pre-repurchase (non-adjusted) EPS and the analysts' EPS forecast (normalized by the end-of-the-period stock price).

forecast by 10 cents by increasing the EPS from \$3.98 to \$4.1. In this sense, the repurchase program was accretive because managers were able to boost the EPS above the level of inaction by 12 cents.

On the left-hand side of Figure 1, we plot the frequency of firms conducting a repurchase over bins of EPS accretion, i.e., the difference between announced EPS and pre-repurchase EPS (normalized by the stock price). More than 95 percent of firms conduct repurchases that allow to increase the EPS by 0 to 2 cents while only a few boost the EPS by more. As the numerical example suggests, this is because increasing the EPS by more than 2 cents through buybacks might be extremely expensive and too detrimental for earnings such that the operation would be overall ineffective.

Then, we exploit the distance from the EPS forecast and the pre-repurchase EPS to understand which firms are more likely to repurchase and by how much. The right-hand side of Figure 1 plots the share of repurchasing firms over the pre-repurchase distance from the forecast (normalized by the stock price). If firms that were already on target to exhibit an average probability of repurchasing around 4 percent, things are different for those on the left of the cut-off. In fact, those are the firms strategically more willing to buy back

Figure 2. Distance from Target and Changes in Investment and Employment



Note: The graph on the left-hand side plots the change in capital expenditure on the distance between the pre-repurchase (non-adjusted) EPS and the analysts' EPS forecast (normalized by the end-of-the-period stock price). The graph on the right-hand side plots the change in employment on the distance between the pre-repurchase (non-adjusted) EPS and the analysts' EPS forecast (normalized by the end-of-the-period stock price). The change in investment (employment) is defined as the difference between the average level of capital investment (employment) in the previous four quarters and the average level of capital investment (employment) in the following four quarters. Such difference is normalized by the level of assets in $t - 4$.

in order to correct the EPS and not disappoint capital markets.[10] This explains why, on the left-hand side of the cut-off, the probability to repurchase increases the closer a firm is to meeting the analysts' forecast. In fact, for firms ex ante closer to target, incentives to repurchase are high, since it is easier and does not take many resources to tilt the EPS to meet market expectations. Conversely, for firms far away from the cut-off, the probability is smaller, since any repurchase would not be large enough to put the EPS on target.

This heterogeneous propensity to repurchase has implications for the dynamics of real variables. In fact, since the EPS adjustment is costly, we expect firms that need to repurchase to invest less in new capital and new hires compared with firms that do not need to repurchase. Figure 2 shows that this is indeed the case. The left- and

[10] As documented in Bartov, Givoly, and Hayn (2002), Kasznik and McNichols (2002), Kinney, Bargstahler, and Martin (2002), and Hribar, Jenkins, and Johnson (2006), missing the EPS negatively affects the market value of the firm, stock market returns, and the credibility of the firm's management.

right-hand side plot, respectively, the change in capital investment and employment (normalized by the value of assets) as a function of the pre-repurchase distance from the EPS forecast (normalized by the stock price). Firms that learn that they need to adjust their EPS tend to invest and hire less with respect to firms that are sure of meeting the EPS target.

In light of this evidence, now we want to study if monetary policy can influence the repurchase behavior of firms, and with what implications for real variables. Our strategy proceeds as follows. First, we focus on the right-hand side of Figure 1 and follow Calonico, Cattaneo, and Titiunik (2014) to define $[-0.018\$, +0.018\$]$ as the optimal (symmetric) interval of firms around the discontinuity. By doing so, the discontinuity in the probability to buy back allows to separate firms into two comparable groups: "repurchasers" and "non-repurchasers," i.e., firms below and above the cut-off. Then, we assess across both groups how exogenous changes in monetary policy affect repurchase expenditure and real variables through variations in the corporate cost of debt. Finally, we exploit our identification strategy to investigate the causal effect of repurchases, changes in the corporate cost of debt, and their interaction on investment and employment.

3.2 Results

Repurchases, EPS Distance from Forecast, and the Cost of Debt. Here we study how variations in the cost of debt due to innovations in monetary policy differently affect repurchasing behavior of firms around the discontinuity. First, for each firm we define the pre-repurchase distance from the EPS target with the variable *Distance*, i.e., the stock price normalized difference between the pre-repurchase EPS and the EPS forecast. Under this definition, a firm i is off (on) target before the repurchase if $Distance < 0$ ($Distance \geq 0$). Second, following Calonico, Cattaneo, and Tittunik (2014), we keep observations only for firms with $Distance_{i,t}$ in the $[-0.018\$, +0.018\$]$ bracket. Third, we extract the exogenous change in the firm-specific yield, by using monetary policy innovations (*Shock*) as an instrument for the firm-specific cost of debt ($\Delta Yield$). Then, we study how being off target and receiving an exogenous change in the cost of debt affects this level of repurchase. In order to do so, consider the following:

$$
\begin{aligned}
Repurchases_{i,t} &= \alpha + \beta_1 \mathbb{I}(Distance_{i,t} < 0) + \beta_2 \widehat{\Delta Yield}_{i,t} \\
&+ \beta_3 \mathbb{I}(Distance_{i,t} < 0) * \widehat{\Delta Yield}_{i,t} + \beta_4 Distance_{i,t} \\
&+ \beta_5 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t} + \beta_6 Distance^2_{i,t} \\
&+ \beta_7 \mathbb{I}(Distance_{i,t} < 0) * Distance^2_{i,t} + \beta_8 Distance^3_{i,t} \\
&+ \beta_9 \mathbb{I}(Distance_{i,t} < 0) * Distance^3_{i,t} + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t}, \quad (5)
\end{aligned}
$$

where $Repurchases_{i,t}$ is the level of positive net repurchase normalized by the level of assets in $t-4$, $\mathbb{I}(Distance_{i,t} < 0)$ takes value one when the firm i is off target with respect to the analysts' EPS forecast, $Distance_{i,t}$ is the pre-repurchase distance from the forecast, the square and the cube of this measure and its interaction with $\mathbb{I}(Distance_{i,t} < 0)$ control for non-linear behavior both at the left- and right-hand side of the discontinuity cut-off, and $\widehat{\Delta Yield}$ is the change in the 10-year corporate yield as predicted (instrumented) by monetary policy shocks. We estimate this equation with 2SLS.[11] Table 5 reports results. As from column 1, firms that are off target buy back 0.8 percent more than those already on target, since they want to tilt the EPS to market expectations. In column 2, we control for exogenous variation in the cost of debt due to monetary policy, but the effect is not significant across firms on both sides of the discontinuity. In column 3, we control for the interaction between the dummy variable $\mathbb{I}(Distance_{i,t} < 0)$ and the change in the cost of debt $\widehat{\Delta Yield}_{i,t}$. As a result, the average level of repurchase is now 2 percent higher for those off target. More interesting is the effect of the cost of debt across groups: if the change in the yield does not matter for those already on target, it does matter for those off target. In particular, if the cost of debt falls by 1 percent, repurchases increase by 0.5 percent only for those that need to buy back in the same quarter. In other words, if a firm in the position to launch an accretive repurchase faces an exogenous increase of the cost of debt, then its

[11] The initial stage for the instrumentation of the firm-level cost of debt with monetary policy innovations is reported in Appendix D.

Table 5. Repurchases, Distance from the EPS Forecast, and the Cost of Debt

	Repurchases **(1)**	***Repurchases*** **(2)**	***Repurchases*** **(3)**	***Repurchases*** **(4)**
$\mathbb{I}(Distance < 0)$	0.008***	0.009***	0.021***	0.016***
	(0.002)	(0.002)	(0.005)	(0.004)
$\Delta \widehat{Yield}$		0.000	0.003	0.002
		(0.088)	(0.002)	(0.002)
$\mathbb{I}(Distance < 0) * \Delta \widehat{Yield}$			–0.005***	–0.005***
			(0.002)	(0.001)
Observations	44,419	30,738	30,738	30,738
Time FE	Yes	Yes	Yes	Yes
Industry FE	No	No	No	Yes
Controls (Polynomial, X)	No	No	No	Yes
Estimator	OLS	OLS	2SLS	2SLS

Note: This table reports 2SLS estimates of Equation (5). Standard errors are in parentheses, double-clustered at firm level and date. In columns 1 to 4, the unit of interest *Repurchases* is the difference between the value of stock purchases and stock issuances from the statement of cash flows. We consider only firms for which such difference is strictly positive, and we normalize it by total asset in $t - 4$. $\mathbb{I}(Distance < 0)$ is an indicator variable that takes value one if the firm is below the EPS forecast before EPS adjustment. $\Delta \widehat{Yield}$ is the exogenous change in the 10-year corporate yield as predicted by monetary policy shocks, i.e., when we instrument $\Delta Yield$ with the monetary innovations *Shock* as from an SVAR. In column 4, we control for a polynomial of the variable *Distance*, i.e., the difference between the EPS forecast and the pre-adjusted EPS of the firm, interacted with the indicator $\mathbb{I}(Distance < 0)$. X controls for return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

action will be limited and its capability to buy back a larger amount of shares will be reduced. Conversely, if the yield on debt falls for a firm about to launch an accretive repurchase, then the lower cost of debt expands the quantity repurchased. Therefore, we conclude that the cost of debt causally affects the size of a repurchase program, and it matters for those managers that need to buy back their own shares to satisfy market expectations. Column 4 shows results when controlling for a polynomial of the variable *Distance* interacted with the indicator variable $\mathbb{I}(Distance < 0)$, and the set of covariates X. Estimates do not differ much.

Real Variables, EPS Distance, and the Cost of Debt. Here we study how variations in the cost of debt due to innovations in monetary policy differently affect capital investment and employment of firms around the discontinuity. To do so, consider the following:

$$\begin{aligned}\bar{Y}_{i,(t+1,t+4)} - \bar{Y}_{i,(t-4,t-1)} &= \alpha + \beta_1 \mathbb{I}(Distance_{i,t} < 0) + \beta_2 \widehat{\Delta Yield}_{i,t} \\ &+ \beta_3 \mathbb{I}(Distance_{i,t} < 0) * \widehat{\Delta Yield}_{i,t} + \beta_4 Distance_{i,t} \\ &+ \beta_5 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t} + \beta_6 Distance^2_{i,t} \\ &+ \beta_7 \mathbb{I}(Distance_{i,t} < 0) * Distance^2_{i,t} + \beta_8 Distance^3_{i,t} \\ &+ \beta_9 \mathbb{I}(Distance_{i,t} < 0) * Distance^3_{i,t} + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t}, \qquad (6)\end{aligned}$$

where the dependent variable is the difference between the mean value of Y in the next and previous four quarters, with Y being either capital investment or employment. This difference is normalized by the level of assets in $t-4$.[12] All other variables are the same.

Table 6 reports results. As from column 1, firms that are off target and need to adjust the EPS reduce investments by 0.24 percent. When adding the instrumented change in the cost of debt $\widehat{\Delta Yield}$ and its interaction with the indicator $\mathbb{I}(Distance < 0)$ in column 2, we find that a 1 percent fall in the cost of debt leads to a 0.15 percent increase of investments for "non-repurchasing" firms, i.e., those above the EPS target. On the other hand, the same fall in the cost of debt leads to a smaller increase in investments for "repurchasing" firms, which is equal to $(0.15\% - 0.04\%) = 0.11\%$. As from column 3, firms off target cut employment by 27 percent. When considering the cost of debt in column 4, we find that a 1 percent fall in the cost of debt leads to a 6.6 percent increase of employees for "non-repurchasing" firms. On the other hand, the same change in the cost of debt leads to a smaller increase in employees for "repurchasing" firms, which is equal to $(6.66\% - 1.28\%) = 5.32\%$.

[12]Recall that employment data are available only at yearly frequency. As in Almeida, Fos, and Kronlund (2016), we replace the same value of employment for each quarter of the same year. Then, we proceed by calculating the four-quarter moving average of employment across time, and we normalize it by the value assets in $t-4$. Within a year, the resulting ratio varies due to quarterly movements in the level of assets. Across years, the ratio varies due to changes of both employment and value of assets.

Table 6. Real Variables, Distance from the EPS Forecast, and the Cost of Debt

	Δ*Inv.* (1)	**Δ*Inv.* (2)**	**Δ*Emp.* (3)**	**Δ*Emp.* (4)**
$\mathbb{I}(Distance < 0)$	–0.0024***	–0.0022***	–0.2719***	–0.3018***
	(0.0004)	(0.0006)	(0.0286)	(0.0416)
$\Delta\widehat{Yield}$		–0.0015***		–0.0651***
		(0.0004)		(0.0215)
$\mathbb{I}(Distance < 0) * \Delta\widehat{Yield}$		0.0004**		0.0128**
		(0.0002)		(0.0057)
Observations	38,427	25,985	35,161	23,587
Time FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Controls (Polynomial, X)	Yes	Yes	Yes	Yes
Estimator	OLS	2SLS	OLS	2SLS

Note: This table reports 2SLS estimates of Equation (6). Standard errors are in parentheses, double-clustered at firm level and date. In columns 1 and 2, the unit of interest Δ*Inv.* is the difference between the mean value of capital investments in the next four quarters and in the previous four quarters, normalized by total asset in $t - 4$. In columns 3 and 4, the unit of interest Δ*Emp.* is the difference between the mean level of employment in the next four quarters and in the previous four quarters, normalized by total asset in $t - 4$. $\mathbb{I}(Distance < 0)$ is an indicator variable that takes value one if the firm is below the EPS forecast before EPS adjustment. $\Delta\widehat{Yield}$ is the exogenous change in the 10-year corporate yield as predicted by monetary policy shocks, i.e., when we instrument $\Delta Yield$ with the monetary innovations *Shock* as from an SVAR. For both models, we control for a polynomial of the variable *Distance*, i.e., the difference between the EPS forecast and the pre-adjusted EPS of the firm, interacted with the indicator $\mathbb{I}(Distance < 0)$. X controls for return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

This evidence corroborates results from Table 5: the way firms on and off target manage a liquidity shock is very different. In fact, "repurchasing" firms exploit the lower cost of debt to buy back more and invest and hire less. On the contrary, "non-repurchasing" firms exploit the lower cost of debt to invest and hire more. This proves that any unanticipated monetary policy shock that leads to a downward adjustment in the corporate cost of debt transmits to real variables in different ways, depending on whether the firm is about to repurchase or not.

Real Variables, Repurchases, and the Cost of Debt. Here we investigate the causal impact of share buybacks, changes in the cost of debt, and their interaction on capital investments and employment. The following regression quantifies these three effects:

$$\begin{aligned}\bar{Y}_{i,(t+1,t+4)} - \bar{Y}_{i,(t-4,t-1)} &= \alpha + \beta_1 Repurchases_{i,t} + \beta_2 \Delta Yield_{i,t} \\ &+ \beta_3 Repurchases_{i,t} * \Delta Yield_{i,t} + \beta_4 Distance_{i,t} \\ &+ \beta_5 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t} + \beta_6 Distance_{i,t}^2 \\ &+ \beta_7 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t}^2 + \beta_8 Distance_{i,t}^3 \\ &+ \beta_9 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t}^3 + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t}. \quad (7)\end{aligned}$$

Under this specification, the parameter β_1 and β_3 will tell us, respectively, the local average treatment effect (LATE) of repurchases and changes in the cost of debt on real variables. The interaction of these two variables will explain whether variation in the cost of debt exacerbates the effect of repurchases on real variables ($\beta_1 + \beta_3$), and whether share repurchases reduce the effect of a change in the cost of debt on real variables ($\beta_2 + \beta_3$). We estimate Equation (7) using 2SLS, where the endogenous variables *Repurchases*, $\Delta Yield$ and their interaction are instrumented, respectively, with the indicator variable $\mathbb{I}(Distance < 0)$, the monetary policy innovation *Shock*, and the interaction of these two instruments. Table 7 shows results. As from column 1, we find that a 1 percent repurchase program leads to a 5.9 percent decline in investments. A 1 percent decrease in the corporate cost of debt works in the opposite direction and leads to an increase in investments by 0.25 percent. However, launching a 1 percent repurchase program contemporaneously to a 1 percent fall of the cost of debt exacerbates the crowding out of repurchase on investments ($\beta_1 + \beta_3 > \beta_1$). At the same time, the same shock attenuates the positive effect of a lower cost of debt on investments ($\beta_2 + \beta_3 < \beta_2$). When in column 2 we repeat our estimation under further controls, we find similar results. As from column 3, a 1 percent repurchase program leads to a decline in the employment stock by 1.2 units per million of assets. On the other hand, a 1 percent decrease in the corporate cost of debt causes an increase in employment by 0.05 units per million of assets. Also in this case, launching a 1 percent repurchase program contemporaneously to

Table 7. Real Variables, Repurchases, and the Cost of Debt

	Δ*Inv.* **(1)**	**Δ*Inv.*** **(2)**	**Δ*Emp.*** **(3)**	**Δ*Emp.*** **(4)**
Repurchases	–0.0596***	–0.0615***	–1.2591***	–1.4266***
	(0.0151)	(0.0155)	(0.4489)	(0.4614)
Δ *Yield*	–0.0025**	–0.0027**	–0.0497***	–0.0486***
	(0.0012)	(0.0012)	(0.0180)	(0.0179)
Repurchases * Δ *Yield*	0.0255***	0.0301***	–0.2632*	–0.2812**
	(0.0084)	(0.0091)	(0.1431)	(0.1216)
Observations	25,985	25,985	23,587	23,587
Time FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Controls (Polynomial, X)	No	Yes	No	Yes
Estimator	2SLS	2SLS	2SLS	2SLS

Note: This table reports 2SLS estimates of Equation (7). Standard errors are in parentheses, double-clustered at firm level and date. In column 1 and 2, the unit of interest Δ*Inv.* is the difference between the mean value of capital investments in the next four quarters and in the previous four quarters, normalized by total asset in $t-4$. In columns 3 and 4, the unit of interest Δ*Emp.* is the difference between the mean level of employment in the next four quarters and in the previous four quarters, normalized by total asset in $t-4$. *Repurchases* is the difference between the value of stock purchases and stock issuances from the statement of cash flows. We consider only firms for which such difference is positive, and we normalize it by total asset in $t-4$. The endogenous variables *Repurchases*, Δ*Yield*, and their interaction are instrumented, respectively, with the indicator variable $\mathbb{I}(Distance < 0)$, the monetary policy innovation *Shock*, and the interaction of the two. We control for a polynomial of the variable *Distance*, i.e., the difference between the EPS forecast and the pre-adjusted EPS of the firm, interacted with the indicator $\mathbb{I}(Distance < 0)$. Control X includes return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

a 1 percent fall of the cost of debt exacerbates the crowding out of repurchase on employment and—at the same time—it attenuates the positive effect of a lower cost of debt on this variable. Adding further controls in column 4 does not change significantly these results.[13]

[13]In Appendix E, we show how results change when using directly monetary policy shocks in Equation (7), instead of the (instrumented) corporate cost of debt. The difference between the results following this alternative identification

In light of this evidence, we conclude that a fall in the cost of debt does exacerbate the crowding-out effect of repurchase on real variables. Moreover, the repurchase channel attenuates the positive effect that a decline in the cost of debt has on investments and employment.

3.3 The Attenuation of Accommodative Monetary Policy on Real Variables Due to Repurchases

By how much do repurchases reduce the transmission of an accommodative monetary policy shock on real variables? For firms around the discontinuity, we compute a back-of-the-envelope calculation by using results from Sections 2.3 and 3.2. From Table 4, column 1, we know that a 1 percent exogenous innovation in the fund rate leads to a 0.61 percent increase of the 10-year yield. As from Table 5, column 4, we know that a 0.61 percent increase in the yield leads to a rise in repurchase by $0.5 \times 0.61 = 0.30\%$. Therefore, by using results from Table 7, we can write the expected change in real variables for a 1 percent accommodative monetary policy shock and its implied repurchase level equal to 0.30 percent as follows:

$$\mathbb{E}[\Delta Inv.|Shock_t = -1\%, Rep = 0.30\%] = -5.9 \times 0.3\% + 0.3 \times 0.61 - 2.5 \times 0.30\% \times 0.61 \approx 0.16$$

$$\mathbb{E}[\Delta Emp.|Shock_t = -1\%, Rep = 0.30\%] = -125.9 \times 0.3\% + 4.9 \times 0.61 - 26.3 \times 0.3\% \times 0.61 \approx 2.53.$$

In words, the overall effect of an expansionary monetary policy is positive and in line with what the basic macroeconomic theory predicts: investments grow by 160,000 dollars every million dollars of assets and employment grows by 2.53 employees every million dollars of assets. Yet, if we mute the repurchase channel, the transmission of

and those from Table 7 is explained by firm-level heterogeneity in the cost of debt which—overall—amplifies the effect on real variables of the repurchase channel. Therefore, we believe that—by considering the firm-level cost of debt as in Equation (7)—we do a better job in measuring the firm-level effect of an exogenous monetary policy innovation on the managerial incentive to buy back.

monetary policy is going to be stronger. In fact, the expected change in real variables for a 1 percent accommodative monetary policy shock and a repurchase level equal to zero is

$$\mathbb{E}[\Delta Inv.|Shock_t = -1\%, Rep = 0\%] = 0.3 \times 0.61 \approx 0.18$$
$$\mathbb{E}[\Delta Emp.|Shock_t = -1\%, Rep = 0\%] = 4.9 \times 0.61 \approx 2.98,$$

meaning that—if the repurchase channel is muted—a 1 percent accommodative monetary policy shock would increase investments by 180,000 dollars every million dollars of assets and employment by 2.98 units every million dollars of assets. Therefore, in light of this simple back-of-the-envelope calculation, we can say that the repurchase channel attenuates the transmission of a 1 percent accommodative shocks on investments and employment, respectively, by $[1 - (0.16/0.18)] = 11\%$ and $[1 - (2.53/2.98)] = 15\%$.

In light of these results, we conclude that—at least for firms around the discontinuity—share buybacks not only crowd out investments and employment but also represent a channel through which the transmission of an accommodative monetary policy shock is attenuated and the crowding-out effect on real variables exacerbates. This happens because firms that do repurchase exploit the lower cost of debt to finance these non-productive operations. As a consequence, they divert resources from the real economy.

4. Robustness Checks

4.1 Pre-existing Differences across Firms Above and Below the EPS Target

In order to validate our identification strategy, first we need to check whether firms around the discontinuity differ in major characteristics before the repurchase program is launched. This ensures that no other motive leads firms to repurchase their own share, but only the distance from the EPS forecast. Table 8 shows the difference in leverage, size, corporate yield, growth perspective (PE10), and financial constraint (measured following Hadlock and Pierce 2010) between firms below ($[-0.018\$, 0)$) and above the cut-off ($[0, 0.018\$]$). As

Table 8. Pre-repurchase Difference in Firm Characteristics

	Leverage **(1)**	***Size*** **(2)**	***Yield*** **(3)**	***PE*10** **(4)**	***Fin. Constraint*** **(5)**
Difference	–1.360 (2.920)	–0.009 (0.039)	–0.412 (0.513)	–0.866 (1.970)	–0.003 (0.004)

Note: The table reports the difference in characteristics between firms above and below the cut-off. Each difference is evaluated by regressing the firm characteristic on an indicator variable taking value one if the firm is below the cut-off. For each case, we control for time and firm's industry fixed effects. In column 1, the unit of interest *Leverage* is the ratio between the value of total corporate debt and equity. In column 2, the unit of interest *Size* is the logarithm of the total value of assets. In column 3, the unit of interest *Yield* is the yield for a 10-year-maturity corporate bond. In column 4, the unit of interest *PE*10 is the 10-quarter moving average of the price-earning ratio. In column 5, *Fin. Constraint* is a measure of the financial constraint of the firm built after Hadlock and Pierce (2010). Standard errors are in parentheses, clustered at firm level. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

evident, firms on and off target are very homogeneous in these dimensions.[14]

Second, we need to check that there are no discontinuous differences in employment and investment dynamics for firms around the EPS cut-off before the repurchase takes place. Hence, for firms with $Distance_{i,t}$ in the $[-0.018\$, +0.018\$]$ bracket, we carry out this exercise for the four j quarters preceding the repurchase by running the following regression:

$$\Delta\bar{Y}_{i,(t-1,t-1-j)} = \alpha + \beta\mathbb{I}(Distance_{i,t} < 0) + \theta_t + \epsilon_{i,t}, \qquad (8)$$

where $\Delta\bar{Y}_{i,(t-1,t-1-j)}$ is the change of the dependent variable between $t-1$ and $t-1-j$ (normalized by total asset in $t-4$), with $j = \{1, 2, 3, 4\}$. As from Table 9, there are no systematic pre-existing differences between firms at the left and at the right of the discontinuity in terms of outcome variables. The pre-repurchase

[14]In Appendix D, we show that the EPS forecast is not correlated with monetary policy or the firm-level exogenous variation in the cost of debt, neither in the quarter in which the EPS forecast is released nor in the previous one.

Table 9. Pre-repurchase Trend in Outcome Variables

	Δ*Investment* (1)	**Δ*Employment*** (2)
Changes (t – 2 to t – 1)	–0.000	0.000*
$\mathbb{I}$(*Distance* < 0)	(0.000)	(0.000)
Changes (t – 3 to t – 1)	–0.001*	0.000
$\mathbb{I}$(*Distance* < 0)	(0.000)	(0.000)
Changes (t – 4 to t – 1)	–0.001	0.000
$\mathbb{I}$(*Distance* < 0)	(0.001)	(0.000)
Changes (t – 5 to t – 1)	–0.000	0.000
$\mathbb{I}$(*Distance* < 0)	(0.003)	(0.000)
Time FE	Yes	Yes
Industry FE	Yes	Yes
Controls (X)	No	No

Note: Standard errors are in parentheses, clustered at firm level. In column 1, the unit of interest Δ*Investment* is the difference between the mean value of capital investments measured in the four quarters before period t – 1 (included) and in the four quarters before period t – 1 – j (included), with j = {1,2,3,4}. Each difference is normalized by total asset in t – 4. In column 2, the unit of interest Δ*Employment* is the difference between the mean value of employment measured in the four quarters before period t – 1 (included) and in the four quarters before period t – 1 – j (included), with j = {1,2,3,4}. Each difference is normalized by total asset in t – 4. $\mathbb{I}$(*Distance* < 0) is an indicator variable taking value one if the firm is currently below the EPS target. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

common trend assumption holds. This validates our strategy such that the results of Section 3.2 can be interpreted as causal.

4.2 The Financing of Share Buybacks around the EPS Target

In this section, we check if indeed firms off target use debt to repurchase compared with firms on target. In order to do so, we propose the same accounting equation as in Section 2.2. Columns 1 and 2 of Table 10 show the contribution of debt to repurchases for the sample of firms off target (whose distance from target is in the [–0.018\$,0\$) bracket) while columns 3 and 4 show results for firms on target (whose distance from target is in the [0\$,+0.018\$]). For firms off target, every dollar of assets repurchased is financed with

Table 10. Financing Buybacks

	***Repurchases* (1)**	***Repurchases* (2)**	***Repurchases* (3)**	***Repurchases* (4)**
$\Delta Debt$	0.204* (0.122)	0.185** (0.089)	0.001 (0.001)	0.008 (0.028)
Observations	21,261	15,733	20,102	17,175
Off-Target Sample	Yes	Yes	No	No
Time FE	No	No	No	No
Industry FE	No	No	No	No
Controls (Other Sources)	No	Yes	No	Yes

Note: Standard errors are in parentheses, clustered at firm level. In columns 1 to 4, the unit of interest *Repurchases* is the difference between the mean value of stock purchases and stock issuances from the statement of cash flows. We consider only firms for which such difference is positive, and we normalize it by total asset in $t - 4$. $\Delta Debt$ is the change in the value of current total debt of the firm, normalized by total asset in $t - 4$. When used, the control variables are all main other sources of the budget constraint of the firm: the change in firm money holding plus current net profit, capital expenditure, and the value of the dividends paid. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

20 cents coming from new debt issuance (column 1); when controlling for other sources of financing and expenditures (column 2), the result does not change. For firms already on target (columns 3 and 4), the contribution of debt is not significant. As we learned from Section 3.2, firms below and above target manage their resources in different ways: the former devote more resources to repurchases rather than investing in new capital and new hires, whereas the latter use their resources for productive purposes. This validates our analysis, confirming that debt is an important source for firms that need to launch an accretive repurchase to bring the EPS on target. Firms already on target do not use debt issuance to finance repurchases, but to fund new investment and employment.[15]

[15]In Appendix F, we use our identification strategy to study if repurchases, the corporate yield, and their interaction can explain changes in corporate debt. In other words, we consider the model of Equation (7), but with $\Delta Debt$ as dependent variable. In this case, we find that repurchases cause an increase in debt issuance when the cost of debt is low. This corroborates the idea that debt is

4.3 EPS Accretion and Monetary Policy

Here we study how much an exogenous monetary policy innovation affects the capability of managers to conduct an accretive repurchase, i.e., a repurchase that is able to increase the EPS. In order to test it, consider the following:

$$\mathbb{I}(Accretion_{i,t} > 0) = \alpha + \beta Shock_t + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t} \quad (9)$$

$$Accretion_{i,t} = \alpha + \beta Shock_t + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t}, \quad (10)$$

where $\mathbb{I}(Accretion > 0)$ is an indicator variable taking value one if the firm was able to increase the EPS over the quarter through a share repurchase; *Accretion* is the price-normalized difference between the reported EPS at the end of the quarter and the one that would have prevailed without launching a repurchase program. Table 11 shows results. As from column 1, a 1 percent fall in the federal funds rate leads to an increase in the (linear) probability of conducting an accretive repurchase of 21 percent. In column 2 we consider the level of accretion, and we find that a 1 percent fall in the federal funds rate leads to an increase in accretion by 0.2 cents.

5. A Simple Model of EPS Maximization

The identification strategy of Section 3 is entirely based on the evidence that firms that need to adjust their EPS tend to repurchase more, particularly when the cost of debt is low. This section shows that this empirical fact can be reconciled with a simple theoretical framework of EPS maximization.

EPS Adjustment and the Cost of Money. Consider the following definition for the earning-per-share ratio:

$$EPS = \frac{(1-\tau)[y - r^s nP]}{N - n},$$

where y is firm's profit at the net of production and financial costs, τ is the firm-specific taxation rate, P is the current stock price, n is the

important for the funding of repurchase programs—in particular, when the cost of debt is favorable.

Table 11. EPS Accretion and Monetary Policy Shocks

	$\mathbb{I}(Accretion > 0)$ (1)	*Accretion* (2)
Shock	–0.210*** (0.012)	–0.002*** (0.000)
Observations	44,419	44,419
Time FE	Yes	Yes
Industry FE	Yes	Yes
Controls (X)	Yes	Yes

Note: Standard errors are in parentheses, double-clustered at firm level and date. In column 1, the unit of observation is $\mathbb{I}(Accretion > 0)$, an indicator variable taking value one if the firm was able to boost the EPS through a repurchase program, i.e., if the difference between reported EPS and the EPS that would have prevailed without repurchasing is positive. In column 2, the unit of observation *Accretion* is the difference between reported EPS and the EPS that would have prevailed without repurchasing. *Shock* is an exogenous monetary innovation as from an SVAR (see Appendix A for details). Controls includes return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

number of own shares repurchased, r^s is the return on a three-month government bond, and N is the number of outstanding shares.

Correction upward of the EPS can occur through two channels: (i) through profit management (y),[16] or (ii) through share buybacks (n).[17] However, share repurchases are not always effective in increasing the EPS ratio, i.e., they are not always accretive. In fact, since n appears in both the numerator and denominator, a repurchase

[16] As shown in Burgstahler and Dichev (1997), it is very unlikely for listed firms to report losses. In fact, they would rather adjust their cash flow or reduce operating costs and working capital than report earnings below market expectations. In this regard, see also Degeorge, Patel, and Zeckhauser (1999) and Burgstahler and Eames (2006), who demonstrate that distance from the analysts' EPS or sales forecast triggers managerial strategic behavior on profits in order to immediately please shareholders.

[17] Share buybacks are used not only to tilt the EPS to market expectations (as discussed in Section 3.1) but also to build credibility and reputation on capital markets (see Vermaelen 1981, Grullon and Ikenberry 2000, and Bens et al. 2003).

program is accretive only if the change in the denominator dominates the change in the numerator. In light of this, it is not trivial to say that EPS accretion is more feasible or bigger for a change in the interest rate r^s. In fact, an exogenous change in the value of money changes also managers' incentive to issue new debt, to buy new capital, and to change their leverage position. This endogenous adjustment in the capital structure will ultimately affect production and profits (y). This should be taken into account when considering launching an accretive repurchase program.

For these reasons, we introduce a simple static model to show under which conditions a negative change in the cost of debt allows for accretive repurchases. Following the work of Stein (1996), we imagine a firm characterized by a leverage ratio d, choosing today the level of capital K, debt B, and the quantity n of shares to be repurchased.[18] The firm is a price taker on the equity, bond, and capital markets such that the stock price P, the cost on newly issued debt r^B, and the unitary cost of capital are all observed at the beginning of the period and taken as given. Also, we assume that the firm-specific cost of debt is proportional to the minimum return r^s on a saving account, i.e., $r^B = \kappa r^s$, with $\kappa > 1$. Once the factors of production and the capital structure are chosen, the firm starts production with a final output equal to $f(K) = zK^\alpha$, with $\alpha \in (0, 1)$ and z being the productivity of the firm.

Given this setup, managers who are willing to launch an accretive repurchase face the following problem:

$$\max_{K,B,n} \ \Omega = \frac{(1-\tau)[f(K) - r^B B - r^s nP]}{N - n} - \frac{\theta}{2}[B - dK]^2.$$

In words, they maximize the EPS of the firm (the first element of the objective function Ω), taking into account the quadratic cost that arises due to deviations from the original leverage ratio d (the second element of Ω).[19] Under this formulation, earnings are defined as the after-tax income generated from production once the firm pays the

[18] If $n < 0$, then the firm is a net equity issuer.

[19] Note that, under this formulation, d is the targeted leverage of the firm. Hence, for a level of capital K, the debt issued B should be equal to dk. Deviation from the targeted leverage leads to a quadratic cost with weight θ, a proxy for capital structure flexibility.

interests on debt and reports the forgone earnings if the amount of money spent in the repurchase was instead kept on a saving account. The maximization problem is subject to the firm's budget constraint $K = B - nP$, such that capital is financed through debt at the net of the amount of money allocated to repurchases. Substituting the budget constraint into the objective function reduces the problem by one dimension and gives us the following first-order conditions.

LEMMA 1. *Managers maximize the EPS under quadratic capital adjustment costs if*

(i) $\frac{\partial \Omega}{\partial B} = 0$, *i.e.,* $(1-\tau)[f' - r^B] = \theta(1-d)[B(1-d) + dnP](N-n)$

(ii) $\frac{\partial \Omega}{\partial n} = 0$, *i.e.,* $EPS = (1-\tau)P[f' + r^s] + \theta d[(1-d)B + dnP]P(N-n)$,

where condition (i) states that the net marginal income from an extra unit of debt must be equal to the marginal cost of changing the capital structure through higher bond issuance, while condition (ii) states that the level of repurchase is optimal if the adjusted EPS is equal to the sum of the marginal loss in net income from diversion of resources from production and savings on a safe asset, and the marginal cost of changing the capital structure due to higher buybacks. The solution of the system of equations (i) and (ii) leads to the equilibrium B^*, n^* *and therefore* $K^* = B^* - n^*P$.

In order to understand how changes in the cost of money affect the optimal level B^*, n^* and EPS^*, we perturbate conditions (i) and (ii) of Lemma 1 by a marginal change in the interest rate r^s. This leads to the following.

PROPOSITION 1. *For* θ *small and* N *large, a marginal decrease in the interest rate leads to higher debt issuance* ($\partial B^*/\partial r^s < 0$)*, higher repurchase* ($\partial n^*/\partial r^s < 0$)*, and higher EPS* ($\partial EPS^*/\partial r^s < 0$)*. In other words—for firms with high level of outstanding shares and high flexibility in capital structure—debt issuance, share buybacks, and EPS are all negatively correlated with changes in the cost of money.*

Proof. See Appendix G.1.

Under Proposition 1, we gain two insights. First, launching a repurchase program affects mechanically more the denominator than the numerator of the EPS ratio: for an extra share repurchased, the fall in net income is smaller than the fall in the number of outstanding shares. Second, for a fall in the interest rate, the capital structure of the firm changes in favor of debt despite the quadratic cost of over-leveraging, and managers buy back more.

The theoretical result of Lemma 1 mimics well the empirical evidence of Tables 5 and 11: if a firm has to boost (maximize) the EPS—i.e., if a firm is off target—a fall in the interest rate helps the EPS accretion through a larger repurchase program.

Implications for Capital Expenditure. Analytically, we still do not know what the results of Proposition 1 imply for the derivative $\frac{\partial K^*}{\partial r^s}$. In fact, once perturbing the budget constraint at the equilibrium, we obtain

$$\frac{\partial K^*}{\partial r^s} = \frac{\partial B^*}{\partial r^s} - \frac{\partial n^*}{\partial r^s} P, \tag{11}$$

the sign of which depends heavily on the parametrization of the model.

Under the assumption of θ small and N large and a baseline calibration,[20] we find $\partial K^*/\partial r^s > 0$. In words, for a fall in the interest rate, EPS-maximizing firms increase repurchases at the expenses of capital investment. This theoretical result is in line with what is shown in Table 6: for an exogenous fall in the cost of debt, firms that need to boost (maximize) their EPS in order to meet market expectations invest less in new capital. Conversely, if repurchase behavior was insensitive to changes in the interest rate (i.e., $\partial n^*/\partial r^s = 0$), then firms would exploit the lower interest rate to issue more debt, which will be entirely used to finance new capital[21] (i.e., $\partial K^*/\partial r^s = \partial B^*/\partial r^s < 0$).

[20] Given the mean values of debt-on-asset and 10-year yield (see Table 1), and given knowledge of 10-year average risk-free rate, we set $r^B = 0.052$, $d = 0.23$, $\kappa = 1.3$. Then, we assume $\alpha = 0.5$, $\theta = 1$, $z = 1$, and $N = 1$. Finally, we calibrate P in order to obtain an equilibrium level $n^*/K^* = 0.03$, as observed in the data (see Table 1). See Appendix G.1 and G.2 for details.

[21] We explore this case from a theoretical point of view in Appendix G.2, where we compare the effect of a change in the interest rate r^s for firms that are allowed to repurchase and firms that are not (i.e., firms that maximize the EPS by

6. Conclusion

This paper documents how debt and the cost of debt are key deciding factors for a manager when launching a repurchase program. In particular, we show that if a firm benefits from an exogenous fall in the corporate yield—caused by an accommodative monetary policy shock—and needs to buy back its shares, the amount of shares repurchased from the stock market is going to be larger. This proves that the cost of debt determines the size of repurchase programs and that firms mostly rely on new and low-cost debt to finance this market operation. Moreover, when conducting a repurchase of their shares, the same firms tend to reduce investment and employment since they devote their resources to these programs at the expense of new capital or employees. Thus, we conclude that share buybacks represent a channel through which the transmission of an accommodative monetary shock is attenuated.

The main contribution of the paper is that we are able to measure the causal impact of monetary policy on share buyback programs, to quantify the extent to which the crowding-out effect of repurchases on investment and employment is due to an accommodative monetary policy shock, and, finally, to assess by how much share buybacks reduce the effectiveness of an expansionary monetary policy. This is an empirical challenge that we solve by exploiting a discontinuity in the data triggered by managerial consideration over the EPS index. We use an information shock based on the distance of a firm's EPS from the analysts' forecast to split the sample in firms more prone to buy back their own shares, i.e., those below the analyst EPS forecast, and those that are less prone to buy back, i.e., those above the forecast. Then we show that a negative change in the cost of debt, as explained by an accommodative monetary policy shock, affects only managers below the target and allows them to launch a larger repurchase program and to easily adjust their EPS in order to meet market expectations. Through this instrumentation, we show that such repurchase behavior has a causal and negative impact on investments and employment and that an expansionary

choosing only the optimal level of debt). In the same appendix, we discuss in depth these theoretical results and we link them to the empirical evidence of Section 3.2.

monetary policy exacerbates this effect such that the overall transmission of an accommodative shock on real variables is attenuated. In particular we find that, if the repurchase channel was muted, the transmission of a 1 percent accommodative shock on investments and employment would be, respectively, 11 percent and 15 percent stronger.

Appendix A. An SVAR for Monetary Shocks

We follow Rossi and Zubairy (2011) and Ramey (2016) to extract monetary policy shocks from an SVAR. In particular, for a time-window spanning from 1985:Q1 to 2016:Q4, we consider this model:

$$Z_t = K + \Gamma_1(t) + \Gamma_2(t)d_t + A(L)Z_{t-1} + B(L)u_t^R + \epsilon_t,$$

where $Z_t = [G_t, Y_t, h_t, C_t, I_t, w_t, \pi_t, r_t]$, i.e., a vector containing series for government spending (G_t), real GDP (Y_t), hours worked in the non-farm business sector (h_t), non-durable and service consumption (C_t), gross private investments and durable consumption (I_t), wages in the non-farm business sector (w_t), GDP deflator inflation (π_t), and the three-month rate on government bonds. $\Gamma_1(t)$ and $\Gamma_2(t)$ are both a fourth-degree polynomial time trend, d_t is a dummy variable taking value equal to one for periods after the beginning of the Great Recession (2008:Q1–2016:Q4), and zero otherwise. In this way, not only do we control for a non-linear trend, but we also take into account the structural change that occurred to the economy with the Great Recession. Moreover, with u^R we include also a "narrative" measure of government spending shocks, based on defense news-shocks as from Ramey (2009). $A(L)$ and $B(L)$ are set to be lag polynomials of degree four, consistently with the existing literature on fiscal a monetary policy shocks. All variables, with the exception of the interest rate, are in logs. The monetary shocks are identified using a Cholesky decomposition.

Appendix B. The Financing of Share Buybacks

In Section 2.2, we show that newly issued debt is important for the financing of buyback programs. Here we corroborate this idea

Table B.1. Financing Buybacks

	***Repurchases* (1)**	***Repurchases* (2)**
$\Delta Debt$	0.35***	0.35***
	(0.08)	(0.08)
$\Delta Money$	–0.09***	–0.09***
	(0.03)	(0.03)
Profits	0.48***	0.48***
	(0.06)	(0.06)
Investments	–0.19***	–0.19***
	(0.06)	(0.06)
Dividends	0.00	0.00
	(0.03)	(0.03)
Observations	144,858	144,858
Time FE	No	Yes
Industry FE	No	Yes
Controls (X)	No	No

Note: Standard errors are in parentheses, clustered at firm level. The unit of observation *Repurchases* is the difference between the value of stock purchases and stock issuances from the statement of cash flows. $\Delta Debt$ is the change in the value of current total debt of the firm. $\Delta Cash$ is the change in firm money holding. *Profits* is the value of firm profit at the net of taxes. *Investments* is equal to capital expenditure. *Dividends* is equal to the value of the dividends paid. All variables are in levels, and expressed in U.S. dollars. The sample is composed of firms that are net repurchasers, i.e., firms for which *Repurchases* > 0. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

by considering the same specification of Section 2.2, but with variables expressed in levels and both cash holdings and profits on the right-hand side:

$$\begin{aligned} Repurchases_{i,t} = {} & \beta_1 \Delta Debt_{i,(t,t-1)} + \beta_2 \Delta Money_{i,t} + \beta_3 Profits_{i,t} \\ & + \beta_3 Investments_{i,t} + \beta_4 Dividends_{i,t} + \epsilon_{i,t}. \end{aligned}$$

Table B.1 reports results when the sample is composed of firms that conduct a positive net repurchase. As from column 1, each dollar spent in repurchases is financed with 35 cents of newly issued debt, 10 cents from a reduction of cash holdings, 48 cents from current profits, and 20 cents from a reduction of investments. The level of dividend payments is not significantly correlated with the level of repurchases. These magnitudes do not change when controlling for time and industry fixed effects (column 2).

Appendix C. Monetary Policy Shocks and EPS Surprise

Here, we check that the probability of being below the EPS forecast is not influenced anyhow by the (contemporaneous or lagged) monetary policy shock and the (contemporaneous or lagged) exogenous change in the firm' cost of debt (as explained by the monetary policy shock itself). In practice, we consider the following specifications for $j = \{0, 1\}$:

$$\mathbb{I}(Distance_{i,t} < 0) = \alpha + \beta Shock_{t-j} + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t} \tag{C.1}$$

$$\mathbb{I}(Distance_{i,t} < 0) = \alpha + \beta \Delta \widehat{Yield}_{i,(t-j,t-j-1)} + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t}, \tag{C.2}$$

where $\mathbb{I}(Distance < 0)$ is the indicator variable that takes value one if the firm is currently below the EPS forecast, *Shock* is the monetary policy innovation out of an SVAR (see Appendix A for details), $\Delta \widehat{Yield}$ is the exogenous change in the 10-year corporate yield as predicted by monetary policy shocks, i.e., we use the variable *Shock* as instrument for the cost of debt. X controls for firm-level characteristics such as return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to, θ controls for year fixed effects, and quarter fixed effects separately.

Table C.1 shows results (errors are doubled-clustered at firm level and date). From columns 1 and 2 we find that the monetary policy shock has no contemporaneous or lagged effect on the probability for the firm to be off target. From columns 3 and 4, we find the same when considering the exogenous change in the cost of debt (as explained by a monetary policy shock). These results validate our identification strategy: the two instruments used in the first-stage analysis of Section 3.2 are not significantly correlated.

Table C.1. Monetary Policy Shocks and EPS Surprise

	$\mathbb{I}(Distance < 0)$ (1)	$\mathbb{I}(Distance < 0)$ (2)	$\mathbb{I}(Distance < 0)$ (3)	$\mathbb{I}(Distance < 0)$ (4)
$Shock_t$	0.003 (0.005)			
$Shock_{t-1}$		–0.004 (0.004)		
$\Delta \widehat{Yield}_{(t,t-1)}$			0.483 (0.376)	
$\Delta \widehat{Yield}_{(t-1,t-2)}$				0.586 (0.394)
Observations	44,419	42,214	30,738	29,401
Time FE	Yes	Yes	Yes	Yes
Industry FE	Yes	Yes	Yes	Yes
Controls (X)	Yes	Yes	Yes	Yes
Estimator	OLS	OLS	2SLS	2SLS

Note: Standard errors are in parentheses, double-clustered at firm level and date. The unit of observation $\mathbb{I}(Distance < 0)$ is an indicator variable that takes value one if the firm is currently below the EPS forecast. *Shock* is an exogenous monetary innovation as from an SVAR (see Appendix A for details). $\Delta \widehat{Yield}$ is the exogenous change in the 10-year corporate yield as predicted by monetary policy shocks, i.e., when we use *Shock* to instrument the firm-level cost of debt. Control X includes return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

Appendix D. Share Buybacks, Monetary Policy, and the Cost of Debt: The First Stage

As explained in Section 3.2, we instrument the cost of debt $\Delta Yield$ with the variable $Shock$, i.e., the monetary policy innovations extracted from an SVAR. As in any two-stage least-squares regression, this implies that the endogenous variable will be regressed over all instruments, exogenous variables, and controls. In other words the initial stage is

$$\begin{aligned}\Delta Yield_{i,t} &= \zeta + \mu_1 \mathbb{I}(Distance_{i,t} < 0) + \mu_2 Shock_t \\ &+ \mu_3 \mathbb{I}(Distance_{i,t} < 0) * Shock_t + \mu_4 Distance_{i,t} \\ &+ \mu_5 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t} + \mu_6 Distance^2_{i,t} \\ &+ \mu_7 \mathbb{I}(Distance_{i,t} < 0) * Distance^2_{i,t} + \mu_8 Distance^3_{i,t} \\ &+ \mu_9 \mathbb{I}(Distance_{i,t} < 0) * Distance^3_{i,t} + X'_{i,t-1}\xi + \theta_t + \eta_{i,t},\end{aligned}$$

Table D.1. First Stage

	Δ *Yield* (1)	**Δ *Yield* (2)**	**Δ *Yield* (3)**	**Δ *Yield* (4)**
$\mathbb{I}(Distance < 0)$	–0.000	–0.000	–0.000	–0.001
	(0.001)	(0.001)	(0.001)	(0.001)
Shock		0.522**	0.528**	0.528**
		(0.254)	(0.250)	(0.252)
$\mathbb{I}(Distance < 0) * Shock$			–0.108	–0.098
			(0.103)	(0.115)
Observations	44,419	30,738	30,738	30,738
Time FE	Yes	Yes	Yes	Yes
Industry FE	No	No	No	Yes
Controls (Polynomial, X)	No	No	No	Yes

Note: Standard errors are in parentheses, double-clustered at firm level and date. In columns 1 to 4, the unit of interest is Δ*Yield*, the change in the firm-level 10-year corporate yield. $\mathbb{I}(Distance < 0)$ is an indicator variable that takes value one if the firm is below the EPS forecast before EPS adjustment. *Shock* is an exogenous monetary innovation as from an SVAR (see Appendix A). In column 4, we control for a polynomial of the variable *Distance*, i.e., the difference between the EPS forecast and the pre-adjusted EPS of the firm, interacted with the indicator $\mathbb{I}(Distance < 0)$. X controls for return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

where $\Delta Yield_{i,t}$ is the change in the 10-year corporate yield, $\mathbb{I}(Distance_{i,t} < 0)$ takes value one when the firm i is off target with respect to the analysts' EPS forecast, *Shock* is the monetary policy innovation as from an SVAR (see Appendix A), $Distance_{i,t}$ is the effective distance from the EPS forecast, the square and the cube of this measure and its interaction with $\mathbb{I}(Distance_{i,t} < 0)$ control for non-linear behavior both at the left- and right-hand side of the discontinuity cut-off. Table D.1 shows results (errors are doubled-clustered at firm level and date).

The fact that the coefficient on the indicator variable $\mathbb{I}(Distance < 0)$ is always insignificant proves that there is no difference in terms of cost of debt for firms around the EPS cut-off. Also the coefficient on the interaction term between the indicator variable and the monetary policy shock confirms that the information shock (the variable $\mathbb{I}(Distance < 0)$) does not have any impact on the cost of debt for both firms above and below the discontinuity. On the other hand, the monetary shock has a significant effect on the cost of debt across both groups. Notice finally that the effect of the monetary policy shock on the change in the cost of debt, estimated here for the sample in the $[-0.018, +0.018]$ bracket, is comparable in magnitude to what is found in Section 2.4.

Appendix E. Challenging our Identification Strategy

Here, we want to challenge what was done in Section 3 in order to understand to what extent our identification strategy is rigorous. In particular, we want to check whether using directly monetary policy shocks (*Shock*) instead of the changes in the cost of debt ($\Delta Yield$) in Equation (7) leads to similar results to those from Table 7 of Section 3.2.

First Stage: EPS Distance from Forecast and Monetary Policy Shocks. Here we study how monetary policy innovations affect repurchasing behavior of firms around the discontinuity. Similarly to what we do in Section 3.2, we consider only firms with $Distance_{i,t}$ in the $[-0.018\$, +0.018\$]$ bracket, and estimate the following first-stage regression:

$$\begin{aligned} Repurchases_{i,t} &= \alpha_0 + \beta_1 \mathbb{I}(Distance_{i,t} < 0) + \beta_2 Shock_t \\ &+ \beta_3 \mathbb{I}(Distance_{i,t} < 0) * Shock_t + \beta_4 Distance_{i,t} \\ &+ \beta_5 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t} + \beta_6 Distance^2_{i,t} \\ &+ \beta_7 \mathbb{I}(Distance_{i,t} < 0) * Distance^2_{i,t} + \beta_8 Distance^3_{i,t} \\ &+ \beta_9 \mathbb{I}(Distance_{i,t} < 0) * Distance^3_{i,t} + X'_{i,t}\gamma + \theta_t + \epsilon_{i,t}, \end{aligned}$$

where $Repurchases_{i,t}$ is the level of net repurchase normalized by the level of assets in $t-4$; $\mathbb{I}(Distance_{i,t} < 0)$ takes value one when the firm i is off target with respect to the analysts' EPS forecast; *Shock* is the monetary policy innovation as from an SVAR (see Appendix A); $Distance_{i,t}$ is the effective distance from the EPS forecast, and the square and the cube of this measure and its interaction with $\mathbb{I}(Distance_{i,t} < 0)$ control for non-linear behavior both at the left- and right-hand side of the discontinuity cut-off. $X_{i,t-1}$

Table E.1. EPS Surprise as Unique IV

	Repurchases (1)	*Repurchases* (2)	*Repurchases* (3)	*Repurchases* (4)	*Repurchases* (5)
$\mathbb{I}(Distance < 0)$	0.008***	0.009***	0.013***		
	(0.002)	(0.002)	(0.004)		
Shock		–0.001	–0.001		
		(0.002)	(0.001)		
$\mathbb{I}(Distance < 0)$ * *Shock*			–0.004***		
			(0.001)		
Repurchases				–0.047***	–1.276**
				(0.014)	(0.541)
Shock				–0.002***	–0.031**
				(0.000)	(0.015)
Repurchases * *Shock*				0.017***	0.158*
				(0.006)	(0.095)
Observations	44,419	44,419	39,778	38,427	35,161
Time FE	Yes	Yes	Yes	Yes	Yes
Industry FE	No	No	Yes	Yes	Yes
Controls (Polynomial, X)	No	No	Yes	Yes	Yes
First Stage	Yes	Yes	Yes	No	No
Second Stage	No	No	No	Yes	Yes

Note: Standard errors are in parentheses, double-clustered at firm level and date. In columns 1 to 3, the unit of observation *Repurchases* is the difference between the value of stock purchases and stock issuances from the statement of cash flows. We consider only firms for which such difference is positive, and we normalize it by total asset in $t-4$. $\mathbb{I}(Distance < 0)$ is an indicator variable that takes value one if the firm is currently below the EPS forecast. *Shock* is an exogenous monetary innovation as from an SVAR (see Appendix A for details). In column 4, the unit of observation $\Delta Inv.$ is the difference between the mean value of capital investments in the next four quarters and in the previous four quarters, normalized by total asset in $t-4$. In column 5, the unit of interest $\Delta Emp.$ is the difference between the mean level of employment in the next four quarters and in the previous four quarters, normalized by total asset in $t-4$. Column 3 is the first-stage regression where the endogenous variable *Repurchases* is instrumented with $\mathbb{I}(Distance < 0)$, *Shock*, and their interaction. Columns 4 and 5 report the second-stage regression when the dependent variable is, respectively, the change in investments and employment. We control for a polynomial of the variable *Distance*, i.e., the difference between the EPS forecast and the pre-adjusted EPS of the firm, interacted with the indicator $\mathbb{I}(Distance < 0)$. X controls for return on assets, Q-value of investment, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

controls for firm-level characteristics such as Q-value of investment, return (profit) on assets, a dummy indicating whether the firm has redistributed dividends in the previous four quarters, and a dummy indicating the quintile of asset to which the firms belong.

Table E.1 shows results (errors are doubled-clustered at firm level and time). Firms off target repurchase 1 percent more (column 1). When we control for the monetary policy shock, we find no significant impact on the size of the repurchase program (column 2). When we control for the interaction term $\mathbb{I}(Distance_{i,t} < 0) * Shock_t$ and further controls, we find a negative and significant impact of monetary policy among firms off target. In other words, if a firm needs to buy back to adjust the EPS in the same quarter in which a negative shock realizes, then the firm will be able to repurchase more. How much more? For a 1 percent negative innovation on the federal funds rate, firms off target buy back 0.4 percent more. Is this result comparable with the one from Table 5? There we found that 1 percent decrease in the corporate cost of debt among firms off target leads to an increase of repurchases by 0.5 percent. The difference is explained by heterogeneity in the firm-level cost of debt.

Second Stage: Repurchases, Monetary Shocks, and Real Variables. Here we complete our analysis by investigating the impact of share buybacks, monetary shocks, and their interaction on capital investments and employment. The following regression quantifies this effect:

$$
\begin{aligned}
\bar{Y}_{i,(t+1,t+4)} - \bar{Y}_{i,(t-4,t-1)} &= \alpha_1 + \xi_1 Repurchases_{i,t} + \xi_2 Shock \\
&+ \xi_3 Repurchases_{i,t} * Shock_t + \xi_4 Distance_{i,t} \\
&+ \xi_5 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t} + \xi_6 Distance^2_{i,t} \\
&+ \xi_7 \mathbb{I}(Distance_{i,t} < 0) * Distance^2_{i,t} + \xi_8 Distance^3_{i,t} \\
&+ \xi_9 \mathbb{I}(Distance_{i,t} < 0) * Distance^3_{i,t} + X'_{i,t-1}\omega + \theta_t + \epsilon_{i,t},
\end{aligned}
$$

where the dependent variable is the difference between the mean value of Y in the next four quarters and in the previous four quarters, with Y being either capital investments or employment. All variables are normalized by the level of assets in $t - 4$.

Columns 4 and 5 of Table E.1 report two-stage least-squares estimates when the variable *Repurchases* and $Repurchases * Shock$ are instrumented, respectively, with the variable $\mathbb{I}(Distance < 0)$ and $\mathbb{I}(Distance < 0) * Shock$. Errors are doubled-clustered at firm level and date. From this estimation we find that the crowding-out effect of 1 percent increase in repurchase on investment is 5 percent, while on employment it is 127 percent. When looking at the direct effect of the monetary policy shock, we find that a 1 percent accommodative shock leads to 0.2 percent (3.1 percent) increase in investments (employment). When looking at the interaction term, we find that conducting a 1 percent repurchase in coincidence of a 1 percent accommodative shock leads to 1.7 percent (15.8 percent) decrease in investments (employment) such that the crowding-out effect of repurchases on investments (employment) exacerbates. With respect to the results of Table 7 of Section 3.2, this different identifying equation gives slightly smaller coefficients. The difference is due to the fact that—here—we are not taking into account firm-level heterogeneity in the cost of debt. Since there is heterogeneous adjustment of the corporate yield across firms, not considering this dimension will under-estimate the effects of monetary policy and its interaction with repurchases.

Appendix F. Debt Issuance, Repurchases, and the Cost of Debt

From Section 4.2, we know that firms off target issue more debt to finance the required repurchases to tilt the EPS to market expectations. Here we move a step forward and study if repurchases do cause an increase in debt. For this, we consider again Equation (7), but with the change in debt ($\Delta Debt$) as dependent variable:

$$\begin{aligned}\Delta Debt_{i,t} &= \alpha + \beta_1 Repurchases_{i,t} + \beta_2 \Delta Yield_{i,t} \\ &+ \beta_3 Repurchases_{i,t} * \Delta Yield_{i,t} + \beta_4 Distance_{i,t} \\ &+ \beta_5 \mathbb{I}(Distance_{i,t} < 0) * Distance_{i,t} + \beta_6 Distance^2_{i,t} \\ &+ \beta_7 \mathbb{I}(Distance_{i,t} < 0) * Distance^2_{i,t} + \beta_8 Distance^3_{i,t} \\ &+ \beta_9 \mathbb{I}(Distance_{i,t} < 0) * Distance^3_{i,t} + X'_{i,t-1}\gamma + \theta_t + \epsilon_{i,t}.\end{aligned} \tag{F.1}$$

Table F.1 reports results of 2SLS estimates when the endogenous variables $Repurchases$, $\Delta Yield$ and $Repurchases * \Delta Yield$ are instrumented, respectively, with $\mathbb{I}(Distance < 0)$, $Shock$ and $\mathbb{I}(Distance < 0) * Shock$. Errors are doubled-clustered at firm level and date.

Despite a positive estimate, there is no significant evidence that repurchases cause per se a direct increase in debt. Yet, for an exogenous decline in the cost of debt, an increase of repurchases does cause

Table F.1. Debt, Repurchases, and the Cost of Debt

	Δ*Debt* (1)	**Δ*Debt* (2)**
Repurchases	0.0183	0.0183
	(0.0114)	(0.0114)
Δ *Yield*	–0.0061**	–0.0059**
	(0.0029)	(0.0028)
Repurchases * Δ *Yield*	–0.0066**	–0.0054*
	(0.0031)	(0.0029)
Observations	28,560	28,560
Time FE	Yes	Yes
Industry FE	Yes	Yes
Controls (Polynomial, X)	No	Yes

Note: This table reports 2SLS estimates of Equation (F.1). Standard errors are in parentheses, double-clustered at firm level and date. The unit of interest Δ*Debt* is the change in debt, normalized by total asset in $t-4$. *Repurchases* is the difference between the value of stock purchases and stock issuances from the statement of cash flows. We consider only firms for which each difference is positive, and we normalize it by total asset in $t-4$. The endogenous variables *Repurchases*, Δ*Yield*, and their interaction are instrumented, respectively, with the indicator variable $\mathbb{I}(Distance < 0)$, the monetary policy innovation *Shock*, and the interaction of the two. We control for a polynomial of the variable *Distance*, i.e., the difference between the EPS forecast and the pre-adjusted EPS of the firm, interacted with the indicator $\mathbb{I}(Distance < 0)$. Control X includes return on assets, Q-value of investment, dummy indicating whether the firm has redistributed dividends in the previous four quarters, a dummy indicating the quintile of asset the firms belong to. *, **, and *** indicate significance at the 90 percent, 95 percent, and 99 percent level, respectively.

a significant increase in debt issuance. This evidence corroborates the idea that the cost of debt is strategic for repurchase behavior: when the cost of debt is low, repurchases leads to an increase of debt issuance. Such result is line with Ma (2019), who shows that the relative change in the price of equity with respect to the cost debt is important to rationalize corporate decisions over capital structure and repurchase behavior. Almeida, Fos, and Kronlund (2016) and Wang, Yin, and Yu (2021) do not take into account how variations in the cost of debt interact and affect decisions over repurchases and debt issuance. For this reason, they put more emphasis on the role of internal resources for the financing of share buybacks.

Appendix G. Theoretical Model and Mapping with the Empirical Strategy

G.1 Proof of Proposition 1

Assume for simplicity that the tax rate is zero (i.e., $\tau = 0$). Then, consider the system of equation pinned down by condition (i) and (ii) of Proposition 1 and evaluate it at the equilibrium:

$$\begin{cases} [f'(B^* - n^*P) - r^B] = \theta(1-d)[B^*(1-d) + dn^*P](N-n^*) \\ EPS(B^*, n^*) = P[f' + r^s] + \theta d[(1-d)B^* + dn^*P]P(N-n^*). \end{cases}$$

Perturbate the latter for a small change in the interest rate r^s. Then we obtain the following:

$$\begin{cases} a\frac{\partial B^*}{\partial r^s} + b\frac{\partial n^*}{\partial r^s} = \kappa \\ c\frac{\partial B^*}{\partial r^s} + d\frac{\partial n^*}{\partial r^s} = \frac{\kappa B^* + NP}{N-n^*}, \end{cases} \tag{G.1}$$

where

$$\begin{aligned} a &= [f'' - \theta(1-d)^2(N-n^*)] \\ b &= \theta(1-d)[B^*(1-d) - dP(N-2n^*)] - Pf'' \\ c &= -\frac{Pf''}{N-n^*} - \theta dP(1-d)(N-n^*) \\ d &= f''P^2 - \theta dP[dP(N-2n^*) - (1-d)B^*]. \end{aligned}$$

Then, by using Cramer's rule, we can find the solution of system (G.1):

$$\begin{aligned} \frac{\partial B^*}{\partial r^s} &= \frac{\kappa d - b\frac{\kappa B^* + NP}{N-n^*}}{ad - cb} \\ \frac{\partial n^*}{\partial r^s} &= \frac{a\frac{\kappa B^* + NP}{N-n^*} - \kappa c}{ad - cb}. \end{aligned}$$

To understand the sign of $\frac{\partial B^*}{\partial r^s}$ and $\frac{\partial n^*}{\partial r^s}$, analyze first the sign of the denominator. For simplicity, we consider the case in which $\theta = 0$. Therefore we can write

$$ad - cb = \frac{(f'')^2P^2[N - n^* - 1]}{N - n^*}.$$

Assuming concavity of the production function $\alpha \in (0,1)$, $N-n^* > 1$ and θ small is sufficient for $ad - cb$ to be positive. Under these normative assumptions, which, respectively, imply decreasing returns to capital, an amount of outstanding shares bigger than $1+n^*$, and low cost in leverage change, we can write

$$\kappa d - b\frac{\kappa B^* + NP}{N - n^*} < 0$$

and

$$a\frac{\kappa B^* + NP}{N - n^*} - \kappa c < 0.$$

Therefore, under these normative assumptions, we conclude that

$$\frac{\partial B^*}{\partial r^s} < 0 \qquad \text{(G.2)}$$

$$\frac{\partial n^*}{\partial r^s} < 0. \qquad \text{(G.3)}$$

In light of these results, we can take a first-order Taylor expansion around the equilibrium EPS, and study how the latter changes for a small variation of the interest rate Δr^s:

$$\Delta EPS^* = [f' - \kappa^B B^*]\Delta B^* - P[f' + r^s]\Delta n^* \\ - [\kappa B^* + Pn^*]\Delta r^s + EPS^* \Delta n^*.$$

Under the simplifying assumption of $\theta = 0$, and by making use of condition (i) and (ii) of Proposition 1, we can write

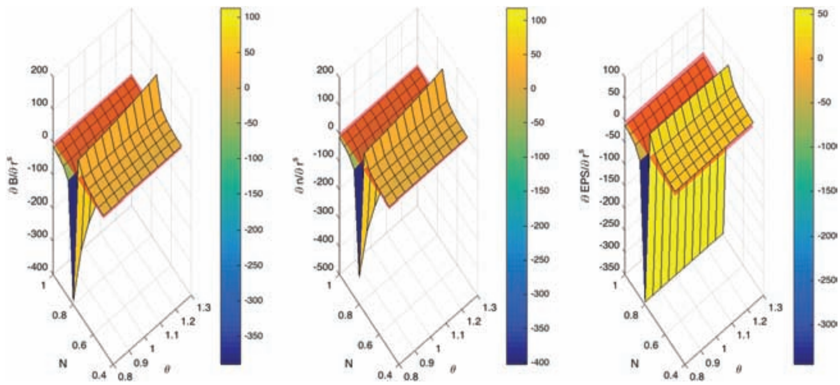
$$\Delta EPS^* = -[\kappa B^* + Pn^*]\Delta r^s,$$

which is negative for a positive increase in the interest rate r^s. Hence, we can approximate this result as

$$\frac{\partial EPS^*}{\partial r^s} = -[\kappa B^* + Pn^*] < 0. \qquad \text{(G.4)}$$

Yet, these results are not general since they are based on the assumption that $\theta = 0$, which requires $\alpha \in (0,1)$ and N large for the derivatives of Equations (G.2), (G.3), and (G.4) to be negative. Conversely, for high level of θ, α and a small level of N, the sign of

Figure G.1. The Role of θ and N for the Sign of $\frac{\partial B^*}{\partial r^s}$, $\frac{\partial n^*}{\partial r^s}$, and $\frac{\partial EPS^*}{\partial r^s}$

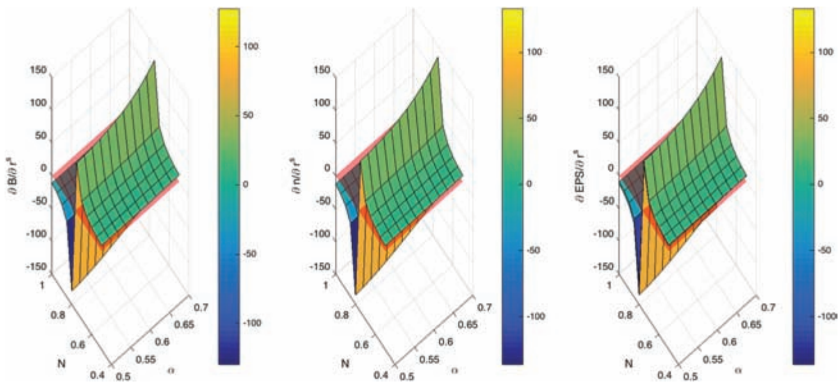


Note: This figure plots the derivative $\frac{\partial B^*}{\partial r^s}$, $\frac{\partial n^*}{\partial r^s}$, and $\frac{\partial EPS^*}{\partial r^s}$ on the (N, θ)-space.

the derivative could potentially change. Here we check this fact by studying how $\frac{\partial B^*}{\partial r^s}$, $\frac{\partial n^*}{\partial r^s}$, and $\frac{\partial EPS^*}{\partial r^s}$ behave in the parameter space of α, θ, and N. First, we numerically solve the model. The initial parameterization is $r^B = 0.052$, as the mean yield on a 10-year corporate bond (see Table 1), $r^s = 0.04$ as the mean 10-year Treasury bill observed in the 1985–2016 period (which implies $\kappa = 1.3$), $d = 0.23$ as the observed mean leverage, $\theta = 1$, $z = 1$, $N = 1$, and $\alpha = 0.50$. Then, we compute a grid search for P in order to obtain an equilibrium level n^*/K^* close to the ratio of repurchase on assets observed on average in the data (and equal to 0.03). The calibrated P is equal to 1.195, which implies a 10-year return on equity equal to 19 percent. Under this set of parameters the equilibrium levels of the endogenous variables are $EPS^* = 0.88$, $n^* = 0.023$, $B^* = 0.85$, $K^* = 0.82$. Hence, we evaluate Equations (G.2), (G.3) and (G.4) and study how each derivative behaves for values of θ, α, and N around their initial (pre-set) levels.

Figure G.1 plots the sign of each derivative in the (θ, N)-space. As shown, the sign of each derivative flips from negative to positive for low levels of N. In fact, if the number of outstanding share is above (below) a certain threshold, any increase in the interest rate will decrease (increase) also debt, the level of repurchase, and the

Figure G.2. The Role of α and N for the Sign of $\frac{\partial B^*}{\partial r^s}$, $\frac{\partial n^*}{\partial r^s}$, and $\frac{\partial EPS^*}{\partial r^s}$



Note: This figure plots the derivative $\frac{\partial B^*}{\partial r^s}$, $\frac{\partial n^*}{\partial r^s}$, and $\frac{\partial EPS^*}{\partial r^s}$ on the (N, α)-space.

EPS. Moreover, each derivative is an increasing function of θ. For example, if we consider the set of parameters for which the derivative is negative, we find that the higher is the cost of leverage change, the less negative is each derivative.[22] In other words, the optimal levels B^*, n^*, and EPS^* become less sensitive to variation in the interest rate r^s if the cost of changing the capital structure is too high.

In Figure G.2, we study the behavior of each derivative in the (α, N)-space and we reach similar conclusions. The derivative is a positive function of the parameter α,[23] and there is a portion of the parameter space for which the sign of each derivative flips from negative to positive. This happens for low levels of N.

In light of this argument, we conclude that the results of Proposition 1 hold true for firms that can adjust easily their capital structure, that have a large number of outstanding shares, and that exhibit decreasing returns to scale. Moreover, assuming a high N, a

[22]In the portion of parameter space where the derivative is negative, the limit of each derivative is zero for $\theta \to \infty$.

[23]For example, in the portion of parameter space where the derivative is negative, the limit of each derivative is zero for $\alpha \to \infty$.

relatively small θ and $\alpha \in (0, 1)$ is realistic, since this parameterization grants the normative negative relationship between the interest rate and the level of debt.

G.2 Implications for Capital Expenditure: Theory vs. Empirical Results

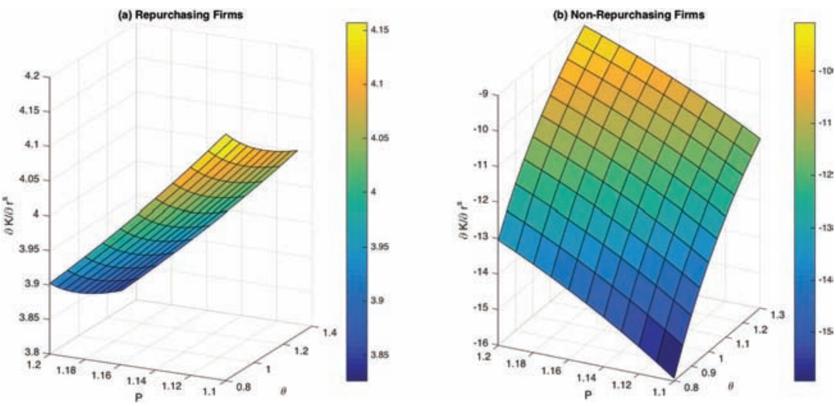
Since $K^* = B^* - n^*P$, we can immediately write

$$\frac{\partial K^*}{\partial r^s} = \frac{\partial B^*}{\partial r^s} - \frac{\partial n^*}{\partial r^s}P, \tag{G.5}$$

which can be either positive or negative depending on the value of $\frac{\partial B^*}{\partial r^s}$ and $\frac{\partial n^*}{\partial r^s}P$ at the equilibrium. For example—for a negative change in the interest rate r^s—if the extra expenditure in share buybacks $\frac{\partial n^*}{\partial r^s}P$ is higher (lower) than the the amount of money raised through debt issuance $\frac{\partial B^*}{\partial r^s}$, the level of capital K^* will decrease (increase). Overall, the sign of the derivative heavily depends on initial parameterization and equilibrium levels. Under the current parameterization and initial equilibrium, we find $\frac{\partial K^*}{\partial r^s} > 0$, i.e., for a decrease in the interest rate, debt alone cannot finance the optimal level of repurchase such that capital expenditure must be cut.

As we can understand from Equation (G.5), if N is large and $\alpha \in (0, 1)$, the value of $\frac{\partial K^*}{\partial r^s}$ evaluated at the equilibrium depends mostly on the stock price and the capability of the firm to adjust its capital structure. Therefore, here we study how the derivative behaves for changes in θ and P. As shown in Figure G.3(A), the derivative is bigger for low levels of P and θ. In words, for a common fall in the interest rate r^s, firms facing a relatively lower stock price and lower capital adjustment cost choose a higher optimal level of repurchase. By doing so, the same firms can increase the EPS by more. However, such increase in share repurchase is also more than proportional to the increase in debt. Therefore, in order for the firms' budget constraint to hold, the equilibrium level of capital must decrease relatively more for these firms. On the other hand, if the stock price and cost of changing capital structure are both too high, there is little the firm can do to increase the EPS, since it cannot easily change the level of debt and the cost of adjusting the EPS through repurchase is already too big. As a result, the effect

Figure G.3. The Sign of $\frac{\partial K^*}{\partial r^s}$: Repurchasing vs. Non-repurchasing Firms



Note: This figure plots the derivative $\frac{\partial K^*}{\partial r^s}$ on the (P, θ)-space for two types of firms. In Figure G.3(A), the firm can use repurchases to maximize the EPS. In Figure G.3(B), the firm is not allowed to use repurchases to maximize the EPS.

on the optimal level of capital will be smaller for the same change in r^s.

How does the derivative $\frac{\partial K^*}{\partial r^s}$ change if we mute the repurchase channel? Say that the economy is split into two group of firms. The first group is the one described so far: it is composed of firms that maximize the EPS by choosing the optimal level B^* and n^*. On the other hand, the second group of firms is not allowed to repurchase and therefore it maximizes the EPS by choosing B^* only. Under this assumption, the EPS maximization problem introduced in Section 5 becomes a simple profit-maximization problem for these (non-repurchasing) firms.

For comparability between the two groups, the optimal level n^* for the repurchasing firms is assumed as an exogenous parameter ($\bar{n} = n^*$) for the group of non-repurchasing firms. Hence, $\bar{n}P$ is a sunk cost for non-repurchasing firms. All other parameters are in common such that the equilibrium level EPS^* and B^* will be identical for both types of firms. In light of this, we can study how the derivative of K^* with respect to the interest rate r^s differs between the two groups and for different parameterization of θ and P.

Figure G.3(B) shows results for the non-repurchasing group. Differently from repurchasing firms, the derivative $\frac{\partial K^*}{\partial r^s}$ is strongly negative for this group. In other words, these firms exploit the lower interest rate to issue debt which is entirely used to finance capital investment. A higher level of capital allows the firm to increase profits and—consequently—the EPS. This effect is stronger for lower level of θ and P, when the non-repurchasing firm can easily adjust the capital structure and the sunk cost $P\bar{n}$ is small.

These results map well to what found in our empirical analysis. As shown in the right-hand side of Figure 1 of Section 3.1, only firms below and very close to the EPS forecast (left-hand side of the discontinuity) have the incentive to buy back since this maneuver allows to boost (maximize) the EPS and to put it on target. As the empirical evidence of Section 3.2 suggests—and consistently with the predictions of the model—the level of repurchase is very sensitive to variation in the cost debt only for firms off target: a lower interest rate incentivizes them to launch a bigger buyback program since this will allow them to boost the EPS by more and to out-beat market expectations. As also shown in Section 3.2—and consistently with the theoretical framework of this section—the same firms divert more resources towards the repurchase program at the expenses of capital investment, which declines.

Conversely, firms on target (right-hand side of the discontinuity) have no incentive to buy back since their EPS is already at the optimal level. Hence, it is plausible to assume—as we do in this theoretical section—that they are operating just under profit maximization. Consequently, they exploit the same lower interest for purposes different from share buybacks: they issue debt to finance capital expenditure.

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Macroeconomic Surprises and the Demand for Information about Monetary Policy*

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This paper studies the *demand* for information about monetary policy, while the literature on central bank transparency and communication typically studies the *supply* of information by the central bank or the reception of the information provided. We use a new data set on the number of views of the Federal Reserve's website and show that exogenous news about the state of the economy as reflected in U.S. macroeconomic news surprises raise the demand for information about monetary policy. Surprises trigger an increase in the number of views of the policy-relevant sections of the website, but not the other sections. Hence, we show that market participants not only revise their policy expectations after a surprise, but also actively acquire new information.

JEL Codes: E52, E58, E32.

1. Introduction

Central banks communicate with financial markets and the general public. The past two decades have seen a remarkable shift towards a higher degree of transparency of central banks about their intentions, the decisionmaking process, and their internal forecasts (Blinder et al. 2008). More recently, central banks such as the Federal Reserve (Fed) or the European Central Bank (ECB) reached the zero lower bound on nominal short-term interest rates and had

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to resort to unconventional tools to implement a further monetary easing. Among these new tools is forward guidance, i.e., enhanced communication with the public about monetary policy in the future. The adoption of forward guidance further increased the interest in central bank communication.

However, central bank communication is typically understood as the provision of information by the central bank to markets and the wider public. Hence, it is about the *supply* of information in order to reduce the information asymmetry between the central bank and the market. As of yet, the literature does not pay much attention to the *demand* for information. Papers study the response of asset prices to central bank announcements or the market reception of central bank communication (e.g., Gürkaynak, Sack, and Swanson 2005a, 2005b; Hansen and McMahon 2016; Cieslak and Schrimpf 2019; Hansen, McMahon, and Tong 2019; Neuhierl and Weber 2019; Swanson 2021). Much less is known about the demand side of central bank communication, that is, what type of information market participants strive for or at what time they want to update their information set about future monetary policy.

This paper studies the demand for information about the monetary policy of the Federal Reserve. We use a new data set that contains the daily number of views of the Federal Reserve Board's website between 2015 and 2019. The data, which have not yet been used to understand the interaction of market participants with the Fed, allow us to break down the demand for information into different aspects of central banking. This is possible because we know the views of each section of the Fed's website such as the "Monetary Policy" section, the "News & Events" section, or the "FOMC" subsection of the "Monetary Policy" section. We discuss the data set with all its limitations below.

As a matter of fact, the views of the Fed's website reflect a number of determinants. One of them is monetary policy itself. For example, views of the website explode around meetings of the Federal Open Market Committee (FOMC). Hence, website views are driven by the supply of and the demand for information. We believe that macroeconomic surprises are an exogenous source of variation of the demand for information and thus facilitate an identification of demand effects. It is not likely that the Fed adjusts its supply of information as a systematic response to particularly good or bad macroeconomic news. Therefore, we concentrate on the response of website views to exogenous macroeconomic surprises. Financial markets pay a lot of attention to U.S. data releases such as new nonfarm payroll employment figures on "Payroll Fridays." The literature shows that financial markets immediately adjust their assessment of future monetary policy based on the surprise component of news announcements (e.g., Fleming and Remolona 1999; Balduzzi, Elton, and Green 2001; Andersen et al. 2003, 2007; Gürkaynak, Sack, and Swanson 2005b; Hördahl, Remolona, and Valente 2020).[1] These news surprises are exogenous with respect to the number of website views on the release day and allow us to study exogenous movements in website views triggered by the news release.

The only two papers using central bank website data thus far are Haldane, Macauley, and McMahon (2019) and Jung and Kühl (2021). Haldane, Macauley, and McMahon (2019) use data on website traffic associated with the release of the Inflation Report of the Bank of England. Their data show an increase in website visits following the introduction of a "layered" form of presentation that addresses experts and the wider public differently. Jung and Kühl (2021) employ website traffic on the ECB's website using Google Analytics to show how ECB communication affects the public's demand for information about monetary policy. A higher search volume has an affect on inflation expectations. They use monetary policy shocks in a two-hour window around Governing Council meetings to show that the demand for information increases in the size of the policy shock. The adoption of forward guidance by the ECB reduced the demand for information, while complex unconventional programs such as the Asset Purchase Program raised the demand for information. The crucial difference with respect to these two papers is that we study the adjustment of the demand for information following exogenous macroeconomic news surprises rather than news issues by the central bank itself. We want to know whether market participants do not just reassess the expected path of monetary

[1] Beechey and Wright (2009) show that real interest rates are responsive to macroeconomic surprises and explain a large part of the responses of nominal interest rates on announcement days. Gilbert et al. (2017) contrast the market response to news surprises with their intrinsic value, i.e., their ability to forecast future real economic activity.

policy in light of macroeconomic surprises, but also actively acquire information about the Fed.

Based on our new data set, we study the response of website views to exogenous macroeconomic surprises. To the extent news surprises prompt markets to recalibrate their expectations about monetary policy, market participants should also actively acquire new information, e.g., by searching the Fed's website for new information, re-reading the last FOMC statement, or checking the last set of FOMC projections. In order to identify whether the change in website views is indeed motivated by the demand for information about monetary policy rather than other businesses of the Fed, we separately study the views of each section of the Fed's website. A release after which the number of views of the "Monetary Policy" section increases but interest in the remaining sections remains unchanged is interpreted as a shift in attention towards monetary policy.

We regress the number of website views on macroeconomic news surprises and interpret the estimated effect as a reflection of the *demand* for information. It is unlikely that changes in the *supply* of information systematically coincide with macroeconomic news and thereby render the demand-side interpretation invalid. Thus, we believe that macroeconomic surprises are shocks to information demand. In order to rule out that our results are driven by the supply of information rather than the demand, i.e., to help identify the effect, we proceed as follows. First, we control for events that clearly change the supply of information. These includes meetings of the FOMC, releases of FOMC minutes, and speeches of the Chair or the Vice-Chair of the Federal Reserve. Second, we also distinguish between news during the blackout period before meetings of the FOMC and news in normal times. During the blackout period, members of the FOMC and senior staffers of the Federal Reserve adhere to a strict embargo and abstain from any provision of news and views about the state of the economy or the future course of monetary policy.[2]

[2]One could try to draw an analogy between this paper and the literature on the estimation of supply and demand curves utilizing exogenous variation in supply or demand in a two-stage least-squares (2SLS) regression model. This

We find that macroeconomic surprises prompt a higher demand for information about monetary policy. The views of the policy-relevant sections of the website increase significantly on days with macroeconomic news. The views of the other sections of the website, in contrast, remain insensitive to news. Hence, market participants not only adjust their expectations of future monetary policy—they also actively acquire new information. The results remain unchanged if we control for measures of policy uncertainty. We find that during the blackout period, when the supply of new information from the Fed is negligible, the sensitivity of website views to labor market surprises is even larger than during normal times.

We also study whether positive or negative news surprises have asymmetric effects on information demand as well as the role of forecast disagreement before the news release for the subsequent demand for Fed information. The marginal effect of the size of the surprise is smaller for negative surprises. Forecast disagreement weakens the demand for information about monetary policy. This finding is in line with Pericoli and Veronese (2016), who show that the market response to news falls in the dispersion of beliefs reflected in the Bloomberg survey.

Our results can be interpreted through the lens of recent attempts to introduce inattention or rational information acquisition into macroeconomics and finance (see Sims 2003; Veldkamp 2012). Under a limited capacity to process information, attention is a precious resource. In contrast to the rational expectations paradigm, inattentive market participants (Reis 2006) do not continuously update their information set, but remain inattentive. Hoopes, Reck, and Slemrod (2015) find a similar pattern for taxpayers' online search for information. If the tax deadline approaches, online searches for capital-gains-taxation surges. After the deadline, agents remain inattentive. The optimally chosen level of attention implies that market participants sporadically update their information set. Market participants' expected benefit of acquiring information is dependent on the magnitude of the expected errors that would be made in the absence of updating. A large (absolute) macroeconomic

analogy, however, is not perfect, as we observe the quantity of information, i.e., the number of viewers, but there is no price of information.

surprise may be a signal that information based on previous periods is now less accurate for the future, and so would increase this expected benefit of acquiring new information.[3]

2. The Data Set

To fix ideas, consider a stylized reaction function of the Fed. The short-term interest rate, R_t, is set as a function of the time-varying state of the economy, θ_t, $R_t = f_t(\theta_t)$. The reaction function $f_t(.)$ changes over time. Macroeconomic news surprises provide information about θ_t to market participants. However, in order to be able to gauge the consequences of the signals about θ_t for interest rates, agents have to acquire information on $f_t(.)$ from the Fed's website. The macroeconomic news alone without the information about the form of the reaction function remains a noisy signal about monetary policy.

We measure the quantity of information consumed about monetary policy by the number of views of the Federal Reserve Board's website (https://www.federalreserve.gov/) and its main sections. This is made possible by a new data set to be introduced below. Apart from Haldane, Macauley, and McMahon (2019) and Jung and Kühl (2021), website views have not yet been used to study the flow of information between central banks and the public.

We believe website visits offer several advantages over alternative measures of attention to the Fed such as newspaper articles or the search volume on Google.com: First, in contrast to newspaper reports, clicking on the Fed's website reflects an active research for information. The appearance of an article in a leading newspaper, in contrast, is not informative about how often the article is actually read. Second, using visits to the Fed's website allows to distinguish between sections of the website. Hence, we can narrow the acquisition of new information down to, say, information about monetary policy as opposed to information about the payments system. Third, in contrast to the Google Trends search volume, the number of website views is informative about the absolute level of the demand for information. Data from Google Trends, in contrast, is normalized

[3] I am grateful to an anonymous reviewer for clarifying discussions on this point.

Figure 1. The Website of the Board of Governors of the Federal Reserve System

Note: Navigation bar on the home page of the Board of Governors of the Federal Reserve System: https://www.federalreserve.gov/, accessed on February 7, 2020.

such that the observation with the highest search volume is assigned a value of 100. Below, we also estimate our main regression model for daily visits to Wikipedia pages about the Federal Reserve and for the daily search volume of Fed-related terms on Google.

As a matter of fact, the Fed's website is just one out of many sources of information about monetary policy. In particular, financial professionals use news-wire services of data providers such as Bloomberg to search for information. However, even for professional market participants, the Fed's website remains an important source of information. Hayo and Neuenkirch (2015) conduct a survey among 195 market participants showing that respondents rely on self-monitoring their home central bank, while they retrieve information through media reports regarding foreign central banks' actions.

Figure 1 shows a screenshot of the Fed's website as of February 2020. We use the number of views of the home page as well as all the main sections of the site, that is, "About the Fed," "News & Events," "Monetary Policy," "Supervision & Regulation," "Payment Systems," "Economic Research," "Data," and "Consumers & Communities." While the "Monetary Policy" section is clearly related to monetary policy, the other segments are not. This distinction allows us to identify whether an increase in the number of views is indeed related to monetary policy. We also use data for the main subsection of the "Monetary Policy" section, i.e., the subsection "FOMC," that contains all the information about the Federal Open Market Committee including the meeting calender, minutes, transcripts, and projections. Figure 1 shows the navigation bar with the content of the "Monetary Policy" section unfolded in order to see the content of the "FOMC" subsection.[4]

We filed a Freedom of Information Act request to the Board of Governors of the Federal Reserve System and obtained data on the daily number of views of the aforementioned sections of the Fed's website for the time period October 2, 2015 to October 8, 2019.[5] The sample period mostly covers the period after the Fed returned to conventional monetary policy in December 2015. In October and November 2015, markets were already anticipating the "lift-off" from the zero lower bound in December 2015.

We obtained the number of views only. We could not obtain the number of distinctive users or the clicks per user. As we cannot distinguish between different types of viewers of the Fed website, we refer to the viewers as "market participants," knowing that this characterization is not free from assumptions. Since the Fed does not use persistent cookies or single-session cookies to track views, the number of unique views would be inaccurate.[6]

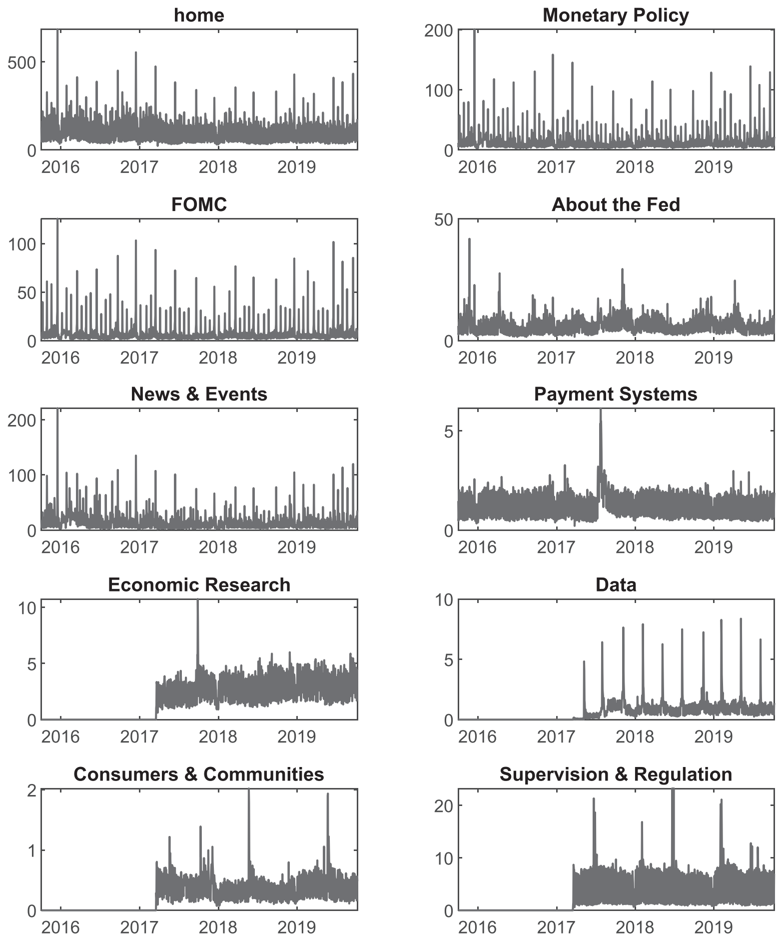
Figure 2 shows the daily number of views of each section of the Fed's website over the sample period. We see that the number of views as well as the volatility of views differ strongly across sections of the website. For the "Consumers & Communities" section, the number of views is the smallest and least volatile. The views of the "FOMC" and "Monetary Policy" sections, in contrast, are very volatile, with the number of clicks exploding on FOMC meeting

[4]The appendix uses snapshots of the Fed website from https://archive.org/web/ to document that the basic structure of the website did not change much over the sample period. The appendix also documents the provision of information on the website during the blackout period before meetings of the FOMC.

[5]For the "Economic Research," "Data," "Consumers & Communities," and the "Supervision & Regulation" subsections, the data start in 2017 due to a reorganization of these website sections.

[6]The Fed's website policies are explained here: https://www.federalreserve.gov/website-linking-policies.htm.

Figure 2. Number of Daily Views of the Federal Reserve Website



Note: The graphs show the number of daily views (in thousands) of each section of the Federal Reserve Board's website. The sample period covers October 2, 2015 to October 8, 2019.

days. On selected days, the number of views is 200 times higher than the average number of views of the "Consumers & Communities" section. The remarkable spikes in the traffic in the "Data" section occur on the days the Fed releases the Senior Loan Officer Opinion Survey on bank lending standards.

To account for the vast differences in average website views, we will base the empirical analysis below on the number of views in standard deviations from its section-specific mean. The regression model will also contain a number of dummy variables, i.e., for FOMC meetings, publication dates of FOMC minutes, and day-of-the-week dummies, which account for the strong daily patterns in the data.[7]

Importantly, the number of views of the policy-relevant sections fluctuates even in the inter-meeting period. We will study macroeconomic news announcements as one exogenous source of these fluctuations.

To understand the data series, we study the distribution of views of each website section and the correlation of views across website sections. Figure 3 shows histograms of views for each section. The histograms also document the skewness and the long tails of the distributions of views. Views of the policy-related sections of the website, i.e., the "Monetary Policy" section and the "FOMC" subsection, jump on days of important monetary policy decisions. This should not be a problem for our daily event study, as the days of macroeconomic news announcements do not coincide with FOMC meeting days.

Figure 4 shows the correlation of views across website sections as a heatmap. Daily views of the monetary policy section are strongly positively correlated with views of the FOMC subsection. We refer to these two sections as the policy-relevant sections of the website. The correlation between the policy-relevant sections and the "About the Fed" and the "News & Events" section is also positive, but smaller. The correlation coefficients in the figure suggest that views of the policy-related sections are only loosely correlated with views of the remaining sections.

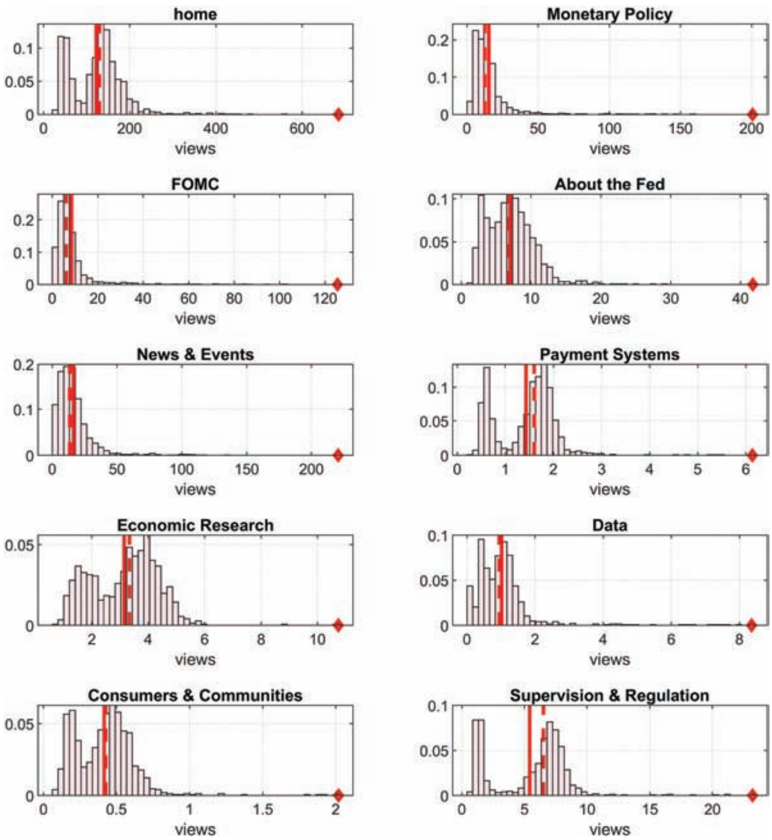
3. The Demand for Information After News Announcements

3.1 Macroeconomic News Surprises

The demand for information about the Federal Reserve and the number of clicks on the Federal Reserve's website are endogenous

[7]Similar daily patterns are also found in Jung and Kühl (2021) for the number of views of the ECB website.

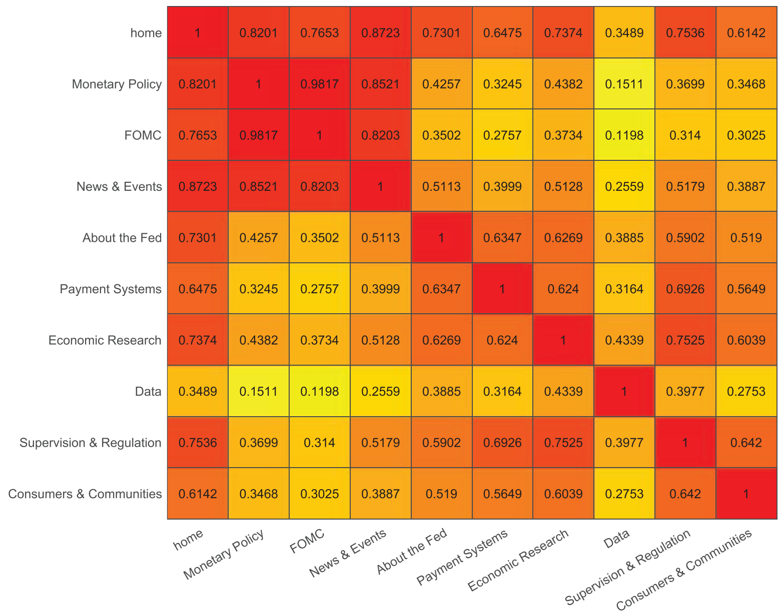
Figure 3. Distribution of Daily Views of Federal Reserve Website



Note: The graphs show histograms of the number of daily views (in thousands) of each section of the Federal Reserve Board's website. The solid (dashed) vertical line is the mean (median). The diamond marks the maximum number of views. The sample period covers October 2, 2015 to October 8, 2019.

variables. Hence, throughout the paper, we study only those changes in the demand for information about monetary policy which are triggered by exogenous news surprises about the U.S. economy. We focus on scheduled monthly releases of new macroeconomic figures, i.e., about the labor market or capacity utilization, which should prompt an adjustment of expectations about monetary policy. These news

Figure 4. Correlation of Views of Federal Reserve Website



Note: The graph shows the correlation of the daily number of views of different sections of the Federal Reserve's website. The darker the color, the higher is the correlation. The sample period covers March 20, 2017 to October 8, 2019, for which we have data on all website sections.

releases are the main source of public information about changes in the state of the U.S. business cycle.

Table 1 lists the indicators and their release schedule. Certainly, some of those indicators are more important than others. For example, news about non-farm payroll employment is by far the most closely watched monthly real economic indicator for the U.S. economy. The press devotes regular columns to the upcoming labor market report such as Bloomberg's "What to Expect From Tomorrow's Jobs Data" column or the *Wall Street Journal's* "5 Things to Watch in the [month] Jobs Report."

Many of our results will be based on non-farm payroll releases. However, we include a broad set of indicators in order to highlight the importance of non-farm payroll data. Besides non-farm payroll numbers, our indicators include the ISM (Institute for Supply

Table 1. Releases of U.S. Macroeconomic Indicators

Indicator	Release Schedule
(1) Non-farm Payroll Employment	On First Friday of the Month by the Bureau of Labor Statistics
(2) ISM Manufacturing Survey	On First Business Day of the Month by the Institute for Supply Management
(3) Industrial Production	Around the 15th of Each Month by the Federal Reserve Board Together with (4)
(4) Capacity Utilization	Around the 15th of Each Month by the Federal Reserve Board Together with (3)
(5) Personal Income	Last Business Day of the Month by the Bureau of Economic Analysis
(6) Retail Sales	Around the 15th of Each Month by the U.S. Census Bureau

Note: The table lists the set of macroeconomic indicators, whose surprise components we focus on in the empirical analysis. We maintain the releases of industrial production and capacity utilization in our data set, although these indicators are provided by the Fed itself, in order to investigate whether their release triggers interest in the policy-related sections of the Fed's website.

Management) manufacturing survey, industrial production, capacity utilization, retails sales, and personal income. In light of the short sample period, we do not use the first releases of gross domestic product (GDP) and other macroeconomic aggregates, which are available on a quarterly frequency only.

While most indicators are published by U.S. statistical agencies, some (industrial production and capacity utilization) are released by the Federal Reserve Board itself. We keep these indicators in the sample in order to assess whether the Fed's own data releases raise interest in the monetary policy sections of its website.

Financial markets should be driven by the unexpected part of the data release only. To isolate the surprise component of the news, we follow the large literature and contrast the release of indicator j at day t, I_t^j, where j is one of the indicators listed in

Table 1, with the median consensus forecast, F_t^j. Hence, the surprise component is

$$Surp_t^j = \frac{I_t^j - F_t^j}{\sigma_{F,j}}. \tag{1}$$

The difference between realization and consensus expectation should be the macroeconomic surprise. It is normalized by the standard deviation of the time series of forecasts. The constant used to normalize the difference should have no effect on the results. A positive $Surp_t^j$ reflects an unexpected improvement in macroeconomic conditions, while $Surp_t^j < 0$ is a surprise deterioration of economic activity.

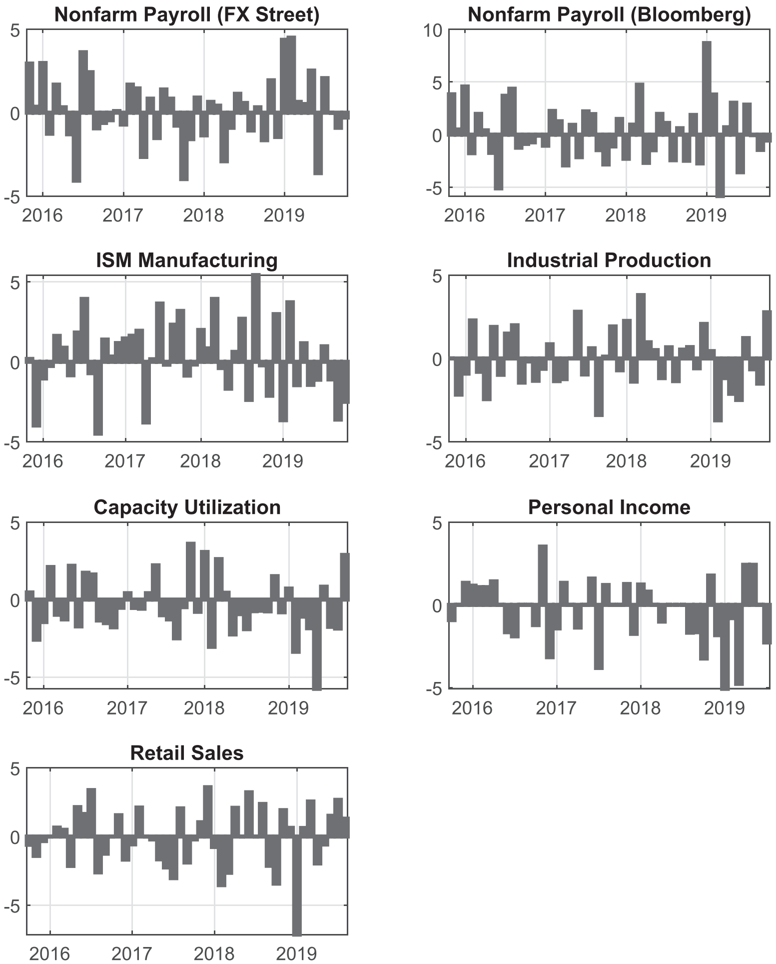
For each of the macroeconomic indicators, we use the surprise component available on Bloomberg, which is based on the Bloomberg survey of market participants. Since the non-farm payroll release is much more important than the other indicators, we use surprises based on the survey from FXStreet.com as an alternative data source. Below, we will study the difference between both surprise components for the non-farm payroll releases in detail and elaborate on the role of forecast dispersion for the response to news.

The survey forecasts, both the Bloomberg survey and the one conducted by FXStreet.com, collect forecasts from market participants in the days before the release. For each indicator, we have 48 news surprises throughout our sample period. The surprise components on announcement dates m are shown in Figure 5. In the online appendix (available at https://sites.google.com/view/peter-tillmann/startseite), we show the market response in a narrow window of 30 minutes after the release of non-farm payroll news, the most important macroeconomic indicator. The evidence clearly underlines the notion that markets adjust the expected interest path in light of labor market news.

3.2 The Regression Model

We now turn to one of our two main research questions: to what extent do exogenous news surprises change the demand for information about monetary policy? To address this question, we regress the number of website views, which we introduced before, on the macroeconomic surprises. On macroeconomic announcement days,

Figure 5. Macroeconomic News Surprises



Note: The graphs show the surprises elements of different macroeconomic news announcements on announcement days. The first surprise series is taken from FX Street; the remaining surprise series are drawn from Bloomberg. The surprises are defined as the difference between the news release and the median of the consensus forecast divided by the standard deviation of forecasts. The sample period covers October 2, 2015 to October 8, 2019.

a change in the attention of market participants with respect to monetary policy should be reflected in an increase in the number of views of policy-related sections of the website.

Let V_t^i be the logarithmic number (multiplied by 100) of views of the Fed's website section i on day t. We run a model separately for each website segment i, in which we regress the website views on the absolute news surprises of type j and a set of control variables, e.g.,

$$V_t^i = \beta_0 + \beta_1 |Surp_t^j| + \beta_2 T_t^{after} + \beta_3 V_{t-1}^i + \Gamma' X_t + \varepsilon_t, \qquad (2)$$

where β_0 is a constant and X_t is a vector of control variables to be explained below. In the baseline model, we include the three most important news surprise jointly. Nevertheless, we also report results for possible combinations of news surprises. We use the absolute news surprise, as positive and negative surprises would equally well trigger interest in monetary policy.

The variable T_t^{after} is the number of days that have elapsed since the last FOMC statement. Ehrmann and Sondermann (2012) show that the relative information content of news increases if the distance to important news of the past becomes larger. Therefore, we expect that the number of views increases the more time has elapsed since the previous policy meeting.[8]

We are primarily interested in the coefficient estimate for β_1. A positive coefficient would imply that a macroeconomic surprise raises the interest in Section i of the Fed's website. The vector X_t contains control variables that should reflect other determinants of market participants' interest in the Fed website. We include a dummy that is one on FOMC meeting days and zero otherwise. This variable captures the strong increase in the public's interest on meeting days. Since the interest in the Fed increases before the meeting and remains high even on the days after the announcement, we also include two leads and lags of the meeting dummy. The publication of the FOMC meeting's minutes three weeks after the FOMC statement should also raise the interest in the Fed. Hence, we construct a dummy that is one on those days that coincide with the publication of FOMC minutes and zero otherwise. These control variables, e.g.,

[8]A non-linear function of the number of days after the FOMC announcement is also possible. For example, interest could remain high after the announcement, decrease thereafter, and increase as the new FOMC meeting approaches. We experimented with different specifications, which, however, leave the main findings unchanged.

the dummies for FOMC meetings and the publication dates of minutes, also control for the supply of information in the policy-relevant sections of the website. Apart from small editorial changes, the Fed does not regularly supply other pieces of information in the "Monetary Policy" section. Hence, after controlling for these two events, we can be confident that fluctuations in website views are driven by the demand rather than supply of information.

Finally, we include a set of day-of-the-week dummies to control for the daily patterns of website views. This should capture the smaller number of views on Fridays and the weekend relative to other days of the week, weekly releases of data, and other factors.

4. Empirical Evidence

We discuss most of the results for a model that includes three separate series of the three most important news surprises, i.e., non-farm payroll news, news about the ISM manufacturing survey, and news about industrial production. We also report alternative specifications for non-farm payroll surprises, which are the most important news release. The appendix provides additional results.

4.1 Baseline Results

Table 2 presents the coefficient estimates for all sections of the website. As our baseline result, we find that labor market surprises raise the number of views of the policy-relevant sections of the Fed's website, i.e., the "Monetary Policy" section and the "FOMC" subsection. The effect is positive and highly significant. Put differently, either a negative or a positive news surprise raises the demand for information about monetary policy. A non-farm payroll surprise 1 percentage point in size raises the number of views of the "FOMC" section by more than 5 percent. News about industrial production also enter with highly significant coefficients, while news surprises about the ISM are significant for the "FOMC" subsection only.

Views of other sections of the website, which do not primarily contain information about monetary policy, remain unaffected by

Table 2. Response of Website Views to the Most Important News Surprises

Website Section	Absolute Surprise			Days after FOMC	Lagged Views, Dummies for Meeting, Minutes, and Day of the Week	# Obs.	R^2
	Non-farm	ISM	IP				
Home	2.026	2.608	4.525	0.065	Yes	1,464	0.91
	(0.643***)	(1.106**)	(1.186***)	(0.032**)			
Monetary Policy	4.204	1.882	4.367	0.236	Yes	1,464	0.91
	(1.001***)	(1.353)	(1.772**)	(0.038***)			
FOMC	5.136	2.413	5.340	0.344	Yes	1,464	0.03
	(1.158***)	(1.451*)	(1.813***)	(0.039***)			
About the Fed	2.180	0.012	0.501	0.020	Yes	1,464	0.86
	(1.158*)	(1.379)	(1.236)	(0.043)			
News & Events	3.017	1.413	2.181	0.014	Yes	1,464	0.83
	(1.504**)	(1.825)	(2.261)	(0.065)			
Payment Systems	1.079	2.026	0.210	0.084	Yes	1,464	0.87
	(0.910)	(1.138*)	(0.082)	(0.039**)			
Research Economic	1.197	2.056	–0.514	0.030	Yes	933	0.90
	(0.766)	(1.116*)	(1.105)	(0.041)			
Supervision & Regulation	1.806	4.202	0.213	0.060	Yes	932	0.92
	(1.011*)	(1.909**)	(0.957)	(0.057)			
Consumers & Communities	–0.088	1.904	–2.100	0.037	Yes	932	0.79
	(1.508)	(1.982)	(1.246*)	(0.061)			
Data	0.502	1.285	–2.120	–0.028	Yes	932	0.89
	(0.950)	(1.716)	(1.196*)	(0.109)			

Note: The dependent variable is the daily (log, multiplied by 100) number of views of the specific section of the Federal Reserve's website. The series of macroeconomic surprises are drawn from Bloomberg. "Days after FOMC" counts the number of days elapsed since the last FOMC statement. The regression also includes a constant. The sample period covers October 2, 2015 to October 8, 2019. Robust standard errors are in parentheses. A significance level of 1 percent, 5 percent, and 10 percent is denoted by ***, **, and *, respectively.

the surprise or even respond negatively. The only exception is the front page, for which the three alternative news surprises also raise the number of views. This is not surprising, as the front page serves as a starting point for navigating the Fed website.

We also find that the number of views of the policy-relevant sections increases in the number of days elapsed since the last FOMC statement. As the next FOMC meeting approaches, market participants' demand for information increases. With each day elapsing after the last meeting, the number of views of the "FOMC" site increases by 0.3 percent. All dummy variables enter the equation with highly significant coefficients. To save space, we do not report these coefficients here.

Macroeconomic surprises drive the demand for policy-relevant information. In the subsequent analysis, we therefore concentrate on the "Monetary Policy" and "FOMC" sections of the website and ignore the other sections in order to save space. In the appendix, we assess the properties of the most important news series and run a placebo experiment: we shift the non-farm payroll surprises one day or one week, respectively, into the past or the future and estimate the model again. We find a significant increase in views only for the original payroll surprises, not the placebo dates. This corroborates the notion that the non-farm payroll news systematically contain information that triggers the search for information by market participants.

4.2 The Blackout Period

As discussed in the introduction, website views reflect changes in both the supply of and the demand for information. However, macroeconomic news surprises should be a source of exogenous variation in the demand for information while leaving the supply unchanged. In this subsection, we take an additional identifying step and study the blackout period only. We can rule out changes to the supply side of information during the blackout period in advance of the FOMC meeting. Members of the FOMC and senior staff members of the Federal Reserve Board adhere to a strict blackout period starting on the Saturday 10 days before each FOMC meeting. During this period, they refrain from any public comment on monetary policy.[9] Hence, we can be certain that there is no change in the supply of monetary policy information during this period.[10] In the appendix to this paper, we show archived screen shots of the Fed website during the blackout period before the September 2018 FOMC meeting in order to document that there is indeed no important new information available on the website before the meeting.

We split the news surprises between surprises during and outside the blackout period. This provides us with a clear identification of the demand-driven increase in website traffic. An important caveat, however, remains: the news surprises are not equally distributed between blackout and non-blackout periods. Of the 48 non-farm payroll releases, 3 fall into the blackout period and 45 occurred before the blackout. For the ISM (industrial production) news, 10 (12) fall into the blackout period and 37 (36) fall in the non-blackout period.

Table 3 shows the results. The surprise components of labor market news and the ISM survey remain a highly significant driver of website views during normal times. During the blackout period, i.e., in the absence of changes to the supply of information, non-farm payroll news have an even larger impact on website views. The coefficient doubles during the blackout period for both the "Monetary Policy" and the "FOMC" section. A 1 percentage point surprise about the state of the labor market raises the number of views of the "Monetary Policy" section by 11 percent. During the blackout, labor market news also significantly affect views of the "News & Events" section. News about industrial production also become more important during the blackout. Note that this effect does not hinge on the fact that the FOMC meeting is approaching, as we still include the number of days elapsed after the previous meeting to account for that.

The fact that the effect of news on website visits becomes stronger is in line with the view that prices are more responsive to

[9]The rules on external communications for Federal Reserve staff during the blackout period are documented here: https://www.federalreserve.gov/monetarypolicy/files/FOMC_ExtCommunicationStaff.pdf.

[10]Ehrmann and Fratzscher (2009) and van Dijk, Lumsdaine, and van der Wel (2016) study the behavior of financial markets in the run-up to FOMC meetings and show that prices respond more strongly to central bank communication and macroeconomic announcements, respectively, during the Fed's blackout period.

Table 3. Response of Website Views to News Surprises: The Role of the Blackout Period

	Website Section	
	Monetary Policy	**FOMC**
Absolute Surprise Outside Blackout		
Non-Farm	4.050 (1.015***)	4.955 (1.166***)
ISM	3.559 (1.287***)	3.969 (1.446***)
IP	1.806 (2.286)	3.203 (2.448)
Absolute Surprise During Blackout		
Non-Farm	11.234 (3.942***)	13.906 (3.974***)
ISM	–8.832 (2.979***)	–7.525 (2.202***)
IP	8.570 (2.549***)	8.861 (2.145***)
Days after FOMC	Yes	Yes
Dummies	Yes	Yes
Lagged Views	Yes	Yes
# Obs.	1,464	1,464
R^2	0.91	0.93

Note: The dependent variable is the daily (log, multiplied by 100) number of views of the specific section of the Federal Reserve's website. The series of macroeconomic surprises are drawn from Bloomberg. We separate surprises during the blackout period before the FOMC meeting from surprises during normal times. The regression also includes a constant and all the control variables explained in the text. The sample period covers October 2, 2015 to October 8, 2019. Robust standard errors are in parentheses. A significance level of 1 percent, 5 percent, and 10 percent is denoted by ***, **, and *, respectively.

news during the blackout (Ehrmann and Fratzscher 2009).[11] If markets are more volatile, the incentives of market participants to pay attention are higher. Hence, the stronger effect of news during the

[11] I am grateful to an anonymous reviewer for this interpretation.

blackout period is consistent with the interpretation of the results in terms of the demand for information rather than the supply of information.

While these results corroborate our baseline finding, the effect of news about the ISM survey remains puzzling. During the blackout, views respond negatively to the surprise component of the survey.

4.3 Alternative Triggers of Attention

The regression model discussed before studies the response of website views to absolute news surprises. We now look at the responses to announcement dummies. We replace $|Surp_t^j|$ in the regression model by a dummy that is one on the announcement day and zero otherwise. Thus, the results are independent from the magnitude of the surprise. The first column in Table 4 reports the results. We find that views of the "Monetary Policy" section and the "FOMC" subsection strongly increase on days with non-farm payroll announcements and releases of industrial production.

Besides macroeconomic news surprises, the number of website views could also respond to other factors that trigger an increase in attention to the Fed. In the following, we study the impact of speeches of senior Fed policymakers, uncertainty about economic policy or monetary policy, respectively, tweets from President Trump, and dissent on the FOMC.

We construct a dummy variable that takes the value of one on a day when the Fed Chair or Vice-Chair gives a speech and is zero otherwise. This variable is included as an additional control variable in our regression equation. Table 4 reports the estimated coefficients. The key result remains unchanged: labor market news prompt an increase in the demand for information about monetary policy. Views of the "Monetary Policy" and "FOMC" sections strongly increase. Furthermore, views remain weakly sensitive to news about the ISM survey and industrial production. Speeches of the Chair and the Vice-Chair strongly raise the attention in the "Monetary Policy" and "FOMC" sections of the website. Hence, the number of website views is very responsive to changes in the supply of policy-relevant information.[12]

[12]Likewise, the Chair's testimony to Congress should raise the public's interest in the Fed and, as a result, the clicks on its website. Hence, we construct a

Fluctuations in uncertainty are another potential explanation for changes in website views. An increase in uncertainty, both about economic policy in general and monetary policy in particular, could trigger an increase in the demand for information. To rule out that our results are driven by uncertainty, we augment the regression equation with the level of two alternative uncertainty indices.[13] The first index is the Baker, Bloom, and Davis (2016) newspaper-based Economic Policy Uncertainty (EPU) index. While the EPU index reflects a general notion of policy uncertainty, our second index of monetary policy uncertainty (MPU) summarizes the uncertainty about the path of short-term interest rates over the coming 180 days. The uncertainty measure is calculated from Eurodollar futures following Swanson (2006). The results are also shown in Table 4. Monetary policy uncertainty drives views of the policy-related website sections, while general economic policy uncertainty remains insignificant for all parts of the website. Most importantly, the surprise component of non-farm payroll releases and industrial production remain significant drivers of the demand for information about Fed policy.

It could be argued that the attention to the Fed triggered by news releases is due to President (or then-candidate) Trump's public pressure on the Fed. The President mostly used his Twitter account to comment on the Fed's reluctance to ease monetary conditions and to attack Chair Powell personally.[14] A tweet from the President on announcement days rather than the announcement itself could be the true driver of attention to the Fed. Therefore, we search the Trump Twitter Archive for tweets on the state of the labor market or the Federal Reserve.[15] Specifically, we collect a variable that counts the daily number of tweets containing the word "Fed" and another variable with the daily number of tweets containing the word "jobs"

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dummy that is one for days that coincide with the Chair's testimony and zero otherwise. This dummy, however, remains always insignificant and is excluded from the regression.

[13]We use uncertainty in $t-1$ to avoid that reverse causality, i.e., a change in uncertainty due to new information about monetary policy in t.

[14]See Tillmann (2020) for evidence on the effect of Trump's pressure on interest rate expectations incorporated into long-term yields.

[15]See http://www.thetrumparchive.com/.

Table 4. Response of Website Views to News Surprises: Alternative Drivers of Information Demand

	Website Section											
	Monetary Policy						FOMC					
Announcement Dummies												
Non-farm	14.830						17.636					
	(3.220***)						(3.487***)					
ISM	2.895						4.913					
	(3.155)						(3.117)					
IP	6.864						8.584					
	(3.208**)						(3.184***)					
Absolute Surprise												
Non-farm		4.107	4.258	4.173	4.203	4.208		5.043	5.168	5.074	5.227	5.287
		(1.050***)	(1.004***)	(0.956***)	(1.020***)	(1.001***)		(1.200***)	(1.147***)	(1.090***)	(1.179***)	(1.152***)
ISM		1.861	2.019	1.804	1.795	1.857		2.400	2.511	2.231	2.253	2.367
		(1.367)	(1.351)	(1.310)	(1.334)	(1.351)		(1.457*)	(1.449*)	(1.369)	(1.411)	(1.448)
IP		4.544	4.368	4.399	4.448	4.431		5.503	5.344	5.389	5.499	5.443
		(1.758***)	(1.786**)	(1.759**)	(1.768**)	(1.741**)		(1.806***)	(1.820***)	(1.776***)	(1.815***)	(1.790***)
Fed Speeches		7.126						6.648				
		(2.179***)						(2.187***)				
EPU			–0.028						–0.017			
			(0.014**)						(0.015)			
MPU				13.050						23.659		
				(3.656***)						(3.987***)		

(*continued*)

Table 4. (Continued)

	Website Section											
	Monetary Policy						FOMC					
Trump												
#Tweets												
On Jobs					0.420						0.169	
					(0.902)						(0.979)	
On Fed					3.145						6.059	
					(0.916***)						(1.078***)	
FOMC Dissent						1.279						2.075
						(0.579***)						(0.626***)
Days after FOMC	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Lagged Views	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
# Obs.	1,464	1,464	1,459	1,459	1,464	1,464	1,464	1,464	1,459	1,459	1,464	1,464
R^2	0.91	0.91	0.91	0.91	0.91	0.91	0.93	0.93	0.93	0.93	0.93	0.93

Note: The dependent variable is the daily (log, multiplied by 100) number of views of the specific section of the Federal Reserve's website. The series of macroeconomic surprises are drawn from Bloomberg. We include the following variables: (i) the daily number of speeches given by the Fed Chair and the Vice-Chair, (ii) the Economic Policy Uncertainty (EPU) index of Baker, Bloom, and Davis (2016), (iii) an indicator of monetary policy uncertainty (MPU) constructed from Eurodollar futures following Swanson (2006), (iv) the number of tweets sent by President Trump that include reference to the labor market and the Fed, respectively, and (v) a measure of dissent in the last FOMC meeting. "Days after FOMC" counts the number of days elapsed since the last FOMC statement. The regression also includes a constant. The sample period covers October 2, 2015 to October 8, 2019. Robust standard errors are in parentheses. A significance level of 1 percent, 5 percent, and 10 percent is denoted by ***, **, and *, respectively.

or "labor market." We then run a specification of the regression equation in which we include these two additional variables.

Table 4 presents the results of the extended model. Tweets on the Fed are a significant driver of website views for the "FOMC" subsection only. Tweets on the labor market itself do not enter the regressions with a significant coefficient. Importantly, the main results remain unchanged: macroeconomic news surprises still shift attention to the policy section of the website, but not the remaining sections.

Decisions in the FOMC are frequently accompanied by dissenting votes. The literature shows that dissent contains information that is relevant for financial markets (Madeira and Madeira 2019). A possible source of dissent could be disagreement among voting FOMC members in the assessment of the macroeconomic situation. Thus, dissent could raise the market's interest in the upcoming release of macroeconomic indicators. Macroeconomic surprises, in turn, could motivate market participants to search online for the number and the name of dissenters in the previous meeting. Therefore, we aim at controlling for dissenting votes in the previous FOMC meeting. We draw on the data provided by Thornton and Wheelock (2014) and include the number of dissenting votes in our benchmark regression equation. All baseline findings remain unchanged. Dissent in the previous meeting strongly increases attention on release days.

4.4 The Response to Individual News Surprises

Thus far we have shown results for a selection of news surprises only. We now report the results for two alternative models in which we include all news surprises jointly. This allows us to compare the relevance of different surprises for market participants' information demand.

Table 5 shows the coefficients for two alternative specifications In the first, we include the Bloomberg non-farm payroll surprise, the ISM survey, capacity utilization, personal income, and sales surprises, while the second specification includes the FX Street non-farm payroll surprise, the ISM survey, industrial production, personal income, and sales surprises. Note that news about capacity utilization and industrial production are released on the same day.

Table 5. Response of Website Views to Individual News Surprises

	Website Section							
	Monetary Policy				FOMC			
Absolute Surprise								
Non-farm	4.083		0.944		4.984		1.716	
(Bloomberg)	(1.001***)		(1.043)		(1.154***)		(1.344)	
Non-farm		5.344		0.757		6.376		1.378
(FX Street)		(1.685***)		(1.776)		(1.879***)		(2.437)
ISM	1.956	1.841			2.566	2.431		
	(1.346)	(1.378)			(1.439*)	(1.482)		
IP		4.417		–2.120		5.163		–0.329
		(1.758**)		(2.505)		(1.842***)		(2.312)
Capacity	3.891		–3.724		4.645		–1.629	
Utilization	(2.010*)		(3.055)		(1.946**)		(2.696)	
Personal Income	–1.461	–1.462	–1.825	–1.834	–2.608	–2.625	–2.816	–2.821
	(1.411)	(1.410)	(1.221)	(1.219)	(1.285**)	(1.283**)	(1.137**)	(1.136**)
Sales	–0.659	–0.664	–0.000	0.120	0.077	0.079	0.928	1.062
	(1.158)	(1.169)	(1.509)	(1.527)	(0.908)	(0.922)	(1.305)	(1.302)

(*continued*)

Table 5. (Continued)

	Website Section							
	Monetary Policy				**FOMC**			
Absolute Release								
Non-farm			2.487	2.636			2.579	2.881
			(0.781***)	(0.748***)			(0.900***)	(0.922***)
IP				2.501				2.108
				(0.914)				(0.801***)
Capacity Utilization			3.131				2.573	
			(1.061***)				(0.931***)	
Personal Income			0.483	0.480			0.419	0.416
			(0.566)	(0.566)			(0.603)	(0.603)
Sales			–0.470	–0.456			–0.554	–0.555
			(0.381)	(0.400)			(0.429)	(0.435)
Days after FOMC	0.237	0.236	0.229	0.231	0.345	0.344	0.340	0.341
	(0.038***)	(0.038***)	(0.038***)	(0.038***)	(0.039***)	(0.039***)	(0.039***)	(0.039***)
Dummies	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Lagged Views	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
# Obs.	1,464	1,464	1,464	1,464	1,464	1,464	1,464	1,464
R^2	0.91	0.91	0.91	0.91	0.93	0.93	0.93	0.93

Note: The dependent variable is the daily (log, multiplied by 100) number of views of the specific section of the Federal Reserve's website. The series of macroeconomic surprises are drawn from Bloomberg and FX Street, respectively. The regression also includes a constant. The sample period covers October 2, 2015 to October 8, 2019. Robust standard errors are in parentheses. A significance level of 1 percent, 5 percent, and 10 percent is denoted by ***, **, and *, respectively.

News surprises on non-farm employment (for both the Bloomberg and the FX Street survey), industrial production, and capacity utilization shift attention to the monetary-policy-related sites of the Federal Reserve. This underlines the notion that macro surprises trigger an increase in the demand for information about monetary policy.

Table 5 also reports the estimated coefficients from a model that includes the absolute news release besides the absolute news surprise. While the former is, for example, the absolute growth rate of non-farm employment released, the latter is the unexpected component of employment growth. We find that the news surprises no longer enter significantly once the absolute releases are included.

The results suggest that market participants respond to the release, not the surprise component, when updating their information set. It seems that the factors that trigger an adjustment of market prices, i.e., news surprises, are not necessarily identical to the factors that prompt observers to revise their information. This remains an interesting issue for future research.

To summarize, this subsection shows that news surprises trigger market participants' demand for information about monetary policy. Fed observers actively start acquiring information about policy and do not just reassess their existing information set.

4.5 Alternative Indicators of Information Demand

The number of website views is an attractive indicator of the demand for information about the Federal Reserve. Nevertheless, we want to confirm the effect of macroeconomic surprises on information demand based on alternative indicators. The first alternative is the daily views of the Wikipedia entries on "FOMC" and "Federal funds rate." The access statistics can be obtained directly from the Wikipedia sites. The second indicator is the daily search volume on Google for "FOMC" or "Federal funds rate."

Table 6 shows that views of both Wikipedia sites respond significantly to news about the labor market. The public seems to collect information about the policymaking process when surprised by the performance of the labor market. News about industrial production raise interest in the "FOMC" Wikipedia page, while news about the

Table 6. Response of Alternative Attention Indicators to News Surprises

Indicator	Announcement Dummy			Days after FOMC	Dummies for Meeting, Minutes, and Day of the Week	# Obs.	R^2
	Non-farm	ISM	IP				
Wikipedia Views	6.823	–0.637	4.968	0.092	Yes	1,464	0.83
"FOMC"	(3.368**)	(2.527)	(2.836*)	(0.049*)			
Wikipedia Views	16.889	–2.334	2.691	0.094	Yes	1,464	0.89
"Fed Funds Rate"	(3.018***)	(2.562)	(2.914)	(0.033***)			
Google Searches	2.735	–11.225	–2.228	0.674	Yes	1,339	0.74
"FOMC"	(6.949)	(10.639)	(6.511)	(0.148***)			
Google Searches	1.126	–9.022	6.391	0.220	Yes	1,311	0.57
"Federal Funds Rate"	(8.442)	(7.132)	(8.371)	(0.121*)			

Note: The dependent variable is the daily (log, multiplied by 100) number of views of the Wikipedia entries or the search volume on Google. The regression includes a constant and all the variables and dummy variables introduced in the main text. "Days After FOMC" counts the number of days elapsed since the last FOMC statement. The sample period covers October 2, 2015 to October 8, 2019. Robust standard errors are in parentheses. A significance level of 1 percent, 5 percent, and 10 percent is denoted by ***, **, and *, respectively.

ISM survey enter with insignificant coefficients. Searches on Google, however, do not increase significantly after a news surprise.

4.6 The Asymmetric Response of Information Demand

In the specifications discussed before, we used the absolute news surprise, as positive and negative surprises should equally raise the demand for information about monetary policy. We now distinguish between positive and negative surprises in order to assess whether the effect is symmetric.

The announcement literature contains several examples of negative surprises, i.e., a surprising contraction of the economy, having a stronger effect on asset prices compared with a surprise expansion of the economy of identical absolute magnitude (Andersen et al. 2003, 2007; Hautsch and Hess 2007). Though the focus of the analysis is on the demand for information triggered by macro surprises, not the adjustment of market prices, we aim to find out whether the demand for policy-relevant information is also characterized by an asymmetry. Do negative news spark a higher demand for Fed information than positive news?

We construct a dummy variable that is one for negative surprises, i.e., $D_t^{neg} = 1$, and zero otherwise. The interaction of this dummy with the news surprise thus allows negative surprises to have an effect on website views that is different from positive surprises.

The modified regression reads as follows:

$$V_t^i = \beta_0 + \beta_1 |Surp_t^j| + \beta_2 T_t^{after} + \beta_3 V_{t-1}^i + \beta_4 D_t^{neg} + \beta_5 D_t^{neg} \times |Surp_t^j| + \Gamma' X_t + \varepsilon_t. \quad (3)$$

We are particularly interested in the coefficient β_5. The estimated coefficients are reported in Table 7. We restrict ourselves to non-farm payroll (Bloomberg) surprises. For the "Monetary Policy" section and the "FOMC" subsection, news surprises remain a highly significant determinant of website views. The demand for policy information, however, is asymmetric: the news surprise enters both regression equations with a positive sign. Hence, larger absolute surprises increase website views. The negativity dummy enters with a positive coefficient. Hence, negative news raise website views independent from the magnitude of the surprise. The interaction term with the

Table 7. Response of Website Views to Non-farm Payroll Surprises: Asymmetries and Dispersion

	Website Section			
	Monetary Policy		FOMC	
Absolute Surprise	6.515 (1.554***)	5.243 (1.213***)	7.211 (1.543***)	5.809 (1.359***)
Dummy Negative		14.281 (4.911***)		17.020 (5.534***)
Interaction Term		–7.425 (2.042***)		–7.328 (2.619***)
Dispersion	–17.358 (5.931***)		–15.809 (6.874**)	
Days after FOMC	Yes	Yes	Yes	Yes
Dummies	Yes	Yes	Yes	Yes
Lagged Views	Yes	Yes	Yes	Yes
# Obs.	1,464	1,464	1,464	1,464
R^2	0.91	0.91	0.93	0.93

Note: The dependent variable is the daily (log, multiplied by 100) number of views of the specific section of the Federal Reserve's website. The series of macroeconomic surprises are drawn from Bloomberg. We include the following variables: (i) the dispersion measured by the absolute difference between the Bloomberg and the FX Street surprises, (ii) a dummy that is one if the surprise is negative and zero if it is positive, and (iii) an interaction term between the negativity dummy and the absolute surprise. "Days after FOMC" counts the number of days elapsed since the last FOMC statement. The regression also includes a constant. The sample period covers October 2, 2015 to October 8, 2019. Robust standard errors are in parentheses. A significance level of 1 percent, 5 percent, and 10 percent is denoted by ***, **, and *, respectively.

negativity dummy is negative. Thus, the marginal effect of the size of the surprise is smaller for negative surprises.

This result has to be interpreted against the backdrop of the interest rate cycle over much of the sample period. After the lift-off from the zero lower bound in December 2015, the Fed raised the target federal funds rate in several steps, before it implemented the first rate cut in July 2019. Hence, through most of the sample period, markets reflected on a faster or slower tightening of monetary conditions, but not on an easing of policy. According to our results, markets believed that favorable labor market news required gathering additional information about monetary policy, possibly because they make a further tightening more likely. Negative news, in contrast, might have been interpreted as an indicator of the Fed leaving monetary conditions unchanged.

4.7 The Role of Forecast Dispersion

Unfortunately, we do not have access to the survey responses underlying the construction of the two alternative labor market surprise series used in this paper, so we cannot use the underlying cross-sectional dispersion of forecasts. However, we can use the difference between both surprises as a proxy for the dispersion of forecasts. Remember that both surprise series are constructed as the difference between realization and median forecast. The difference between both surprises, hence, should be proportional to the difference in median forecasts in two alternative surveys of market participants, as both surprise measures share the same realized value. There is, however, one important limitation to keep in mind: the two surveys may have been conducted at slightly different days in a given month. As a result, the difference between the two alternative surprise series also reflects different information sets. Hence, the dispersion measure is indicative only.

In this spirit, we construct the simple measure of forecast dispersion as the absolute difference of news surprises,

$$dis_t^{nonfarm} = |\widehat{Surp_t}^{nonfarm,Bloomberg} - \widehat{Surp_t}^{nonfarm,FXStreet}|, \quad (4)$$

where the hat over the variables denotes that we divide both surprise series by their standard deviation to make sure that both surprises are equally volatile. Note that this measure reflects forecast dispersion before the release of the labor market figures. The appendix shows that periods with a wider dispersion of forecasts are also periods with large surprises. This supports the notation that forecast dispersion is a useful but noisy measure of uncertainty.

We include $dis_t^{nonfarm}$ as an additional variable in our empirical model. A positive coefficient on forecast dispersion would indicate that markets' need for information increases if the release has been preceded by more heterogeneous views on the state of the economy. We find that forecast dispersion reduces the demand for Federal Reserve information on announcement days; see Table 7 for the set of coefficients. The positive impact of the absolute news surprise on the number of views remains significant. This result is puzzling: to the extent a more dispersed range of forecasts of non-farm payroll employment ahead of the official data release reflects uncertainty about the state of the economy, we should expect a higher demand for Fed information. One potential explanation is that a wider dispersion of forecasts reflects uncertainty of the Fed itself about the economy, thus calling for a less activist policy response to the labor market news and, as a consequence, a smaller demand for information by the public.

5. Conclusions

In this paper, we used the number of views of the Federal Reserve Board's website to quantify the demand for information about U.S. monetary policy. This allows us to shed light on the demand for information around macroeconomic news surprises, which are widely regarded as the main source of public information about the state of the business cycle. The literature has firmly established that interest rates are very sensitive to news surprises, as investors recalibrate their expectations of future policy in light of the new information.

We add to this literature by showing that market participants not only update their beliefs about monetary policy, but also actively search for new information. The number of website views of the policy-relevant sections of the Fed's website increases after macroeconomic surprises. Our findings are in line with the literature on the allocation of attention of market participants and offer new insights into the demand for central bank information.

The paper also offers lessons for the design of central bank communication. We stress the role of the website as an important source of information on days of scheduled macroeconomic data releases. This implies that information provided on these specific dates should be relatively more effective than on other days. On these days, central banks such as the Federal Reserve could provide information they believe is particularly important to steer market expectations.

A visit to the current home page of the Fed is not very informative. The front page does not include information on the current level of the federal funds rate target, the date of the next FOMC meeting, the current inflation rate, the inflation target, or the definition of the Fed's mandate. Making this information quickly and easily accessible, without the need to click through the site, would enhance the role of the website as a main tool of communication with the public.[16] A better understanding of the demand for information about monetary policy should contribute to the design of a more effective central bank communication.

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Which Monetary Shocks Matter in Small Open Economies? Evidence from Canada*

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We investigate the monetary policy transmission in a representative small open economy—Canada—over the 2000–17 period. By using a novel set of external instruments, we identify the impacts of domestic (Canada) and foreign (United States) monetary shocks on financial and macroeconomic variables in Canada in a unified structural VAR framework. Our results first confirm that domestic monetary policy transmission operates through interest rate, foreign exchange, and credit channels in Canada. The results further suggest that U.S. monetary policy shocks also have sizable effects on financial conditions in Canada, in line with the credit and risk-taking channels of international monetary spillovers. That said, the U.S. monetary spillovers into Canadian macroeconomic variables could be offset by the fluctuations in exchange rates and net exports in line with the trilemma hypothesis. Finally, our results are robust to various types of instrumental variables on monetary policy shocks.

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1. Introduction

In a Mundell-Fleming world, with a flexible exchange rate, domestic monetary autonomy and open capital accounts are simultaneously compatible. As global financial markets become increasingly integrated and global factors become crucial drivers of local financial market developments, however, there have been extensive debates on the effectiveness of domestic monetary policy in small open economies with floating exchange regimes (Obstfeld 2015; Rey 2015, 2016; Aizenman, Chinn, and Ito 2016; among many others).[1] To shed some more light on this issue, this study reexamines domestic and cross-border monetary policy transmission in a novel structural vector-autoregressive (SVAR) model. To our knowledge, this paper is one of the first studies that investigate the impacts of local and foreign monetary policy shocks in a unified framework.

The different views on the Mundell-Fleming "trilemma" may reflect different perspectives on monetary policy independence and inconsistent empirical frameworks across studies. First, the existing studies on international monetary spillovers often focus on a single type of transmission channel—for instance, either the financial or the trade channel. The dilemma hypothesis predicts that monetary shocks originating from the center country determine the international financial conditions, and spill over to financial markets (and domestic demand) in other open economies, thereby leading to synchronized macroeconomic and financial conditions across the economies. On the other hand, in line with the trilemma, the fluctuations in exchange rates and net exports may buffer the impact of the financial spillovers onto macroeconomic outcomes. It is thus crucial to consider various transmission channels to truly argue the implications of foreign monetary policy shocks.

Second, studies have often focused on a single type of financial market, without considering the consequence on a variety of financial and credit markets that play different roles in monetary policy transmission (Dedola, Rivolta, and Stracca 2017; Gai and Tong 2019).[2] In addition, there is still no consensus on the role

[1] See Section 2.3 for a more detailed review of the literature on the debates.

[2] For example, when assessing monetary policy autonomy, Rey (2015, 2016) pays attention to the covarying general financial conditions among countries and

of foreign exchange rates, a key variable in the trilemma debate. Many earlier studies pointed out that conditional movements of foreign exchange rates exhibit puzzling deviations from the predictions of Dornbusch (1976)'s overshooting hypothesis (Eichenbaum and Evans 1995; Grilli and Roubini 1996; and Cushman and Zha 1997).[3] In contrast, more recent studies find evidence of more consistent movements of exchange rates following monetary policy shocks (Bjørnland 2009; Kim, Moon, and Velasco 2017; Rogers, Scotti, and Wright 2018; Inoue and Rossi 2019).

Third, the literature has typically not considered the consequence of domestic and foreign monetary policy shocks in a unified framework. Bernanke (2017) points out that the standard Mundell-Fleming model does not predict that small open economies can completely insulate their economy from policy shifts in a center country. The model instead implies that, under a flexible exchange regime, countries can insulate the domestic macroeconomic situation from external shocks by steering interest rates. This calls for a balanced view in understanding the trilemma debates: investigating the effectiveness of domestic monetary policy is equally as important as examining the transmission of international monetary shocks. However, it is econometrically challenging to distinguish the impact of foreign and domestic monetary policy shocks—in particular, if the monetary policies are correlated across countries.

Against this background, we contribute to the literature by investigating both international spillovers and domestic transmission of monetary policy shocks into various financial markets, trade, and macroeconomic variables in Canada, based on a single, open-economy SVAR framework. The model allows us to compare the impact of domestic (for which monetary spillovers from the center country are controlled) and foreign (for which the United States is

concludes that non-U.S. central banks lose their control over local financial conditions. Conversely, Obstfeld, Ostry, and Qureshi (2019) focus on the movements in short-term interest rates by implicitly assuming frictionless transmission of monetary shocks to the macroeconomy through capital and financial markets.

[3]The Mundell-Fleming model predicts that the cross-border transmission of the shocks is attenuated by adjustments in exchange rates. However, despite many results on the puzzling movements of exchange rates in response to monetary policy shocks, only a few studies reconcile the empirical results with the theory (Bruno and Shin 2015a).

the proxy) monetary shocks on multiple market interest rates at a variety of maturities, currency rates, credit costs, and capital flows, as well as effective exchange rates, export and import, in Canada.[4]

Our empirical findings are summarized as follows. First, as the trilemma hypothesis predicts, domestic monetary policy transmission appears to operate through a variety of channels in Canada. Both short- and long-term rates react significantly to domestic monetary policy shocks, confirming the role of the conventional interest rate channel. Foreign exchange rates in this process also respond significantly to monetary policy shocks, as the overshooting theory would predict. Contrary to a group of earlier findings that report counterevidence for the overshooting theory by Dornbusch (1976), we find that an increase in local policy rates causes the nominal exchange rate to appreciate instantaneously, and then to depreciate gradually. The shocks generate an increase in credit spreads in Canada, consistent with the predictions of the credit channel of monetary policy transmission.

In tandem, international spillovers of monetary policy shocks also play an important, and possibly stronger, role in driving financial conditions in Canada. Following U.S. monetary tightening, market interest rates (with both short- and long-term maturities) significantly rise, and the impacts persist for a prolonged period. More interestingly, overnight rates, which are monetary policy instruments in Canada, also respond significantly to U.S. monetary policy shocks. Following a contractionary U.S. monetary policy shock, credit spreads increase substantially, along with an immediate outflow of international capital investments. This is consistent with the predictions by the credit and risk-taking channels of international monetary policy transmission (Rey 2015, 2016; Hofmann, Shim, and Shin 2016). The correlated movements of U.S. and Canadian financial asset prices are also consistent with the international portfolio rebalancing channel as in Blanchard et al. (2016) and Alpanda and Kabaca (2020).

Finally, the response of macroeconomic variables in Canada diverges across the two types of monetary policy shocks. According

[4]As will be seen in Section 5, our results based on a counterfactual analysis suggest that the omission of U.S. monetary policy shocks can bias the estimated impact of domestic monetary policy transmission.

to our empirical results, following a domestic monetary tightening, output and price levels decline significantly, as New Keynesian theory would predict. On the other hand, U.S. monetary tightening leads to expansionary consequences on both output and prices in Canada. This result may have been partly driven by the expenditure-switching effects of U.S. monetary shocks, as the subsequent currency depreciation and the improvement in net exports offset the negative impacts of tightened financial conditions on the real economy in Canada. In line with the predictions by Mundell-Fleming's trilemma, the insulation of foreign monetary spillovers appears to have contributed to the Bank of Canada's successful policy operation for macroeconomic stabilization.[5]

Our empirical results offer a somewhat nuanced perspective on the trilemma debate. The results suggest that domestic monetary policy is still effective in stabilizing domestic financial and macroeconomic conditions when exchange rates freely float (i.e., the trilemma hypothesis has not been violated in Canada, at least for the sample period of this paper). At the same time, the effects of U.S. monetary spillovers on domestic financial and credit conditions are also evident, although the macroeconomic consequences of the spillovers could be offset by the fluctuations in exchange rates and external transactions. The upshot is that, depending upon the extent to which different types of transmission channels operate in the open economy, the macroeconomic consequences of foreign monetary policy spillovers could be quite different. This also highlights the crucial role of exchange rate in small open economies, in a financially globalized world: it can be either a shock amplifier or a shock absorber.

In this paper, we seek to avoid the potential simultaneity issue involving monetary policy actions and other macroeconomic or financial variables. To do so, we use a novel set of external instruments, instead of imposing arbitrary assumptions about

[5] Admittedly, however, this result can be mainly specific to the Canadian economy, which heavily depends on bilateral trades with the United States—in particular, trade in commodities. As argued by some earlier studies—including Iacoviello and Navarro (2019) and Ca' Zorzi et al. (2020)—the macroeconomic consequence of international spillovers could be highly dependent on country characteristics, as well as the sample period that governs the degree of various channels of transmission.

causal relationships among endogenous variables.[6] The identification scheme, initially proposed by Stock and Watson (2012) and Mertens and Ravn (2013), has considerable appeal because it exploits the attractive features of SVARs while addressing the identification issues raised above by using information from external instruments. Recent studies on monetary policy transmission combine an SVAR setup with such an identification scheme, exploiting high-frequency external instrument variables, obtained from futures rates on monetary policy instruments (Gertler and Karadi 2015; Nakamura and Steinsson 2018; Jarociński and Karadi 2020). Unlike the findings in the literature on the United States, this method has not yet been widely applied to the cases of other economies, including Canada.[7]

To overcome the issue, we test three different types of novel instrumental variables for monetary policy shocks in Canada: so-called (i) market based, (ii) model based, and (iii) narrative based. These types of instrumental variables have been widely employed in the literature to identify U.S. monetary policy shocks, but not for the case of Canada with only a few exceptions. First, we measure daily changes in spot overnight interest rates on monetary policy decision dates as a monetary policy surprise (IV1). Second, we calculate the conditional expectations for future short-term interest rates in Canada using a standard affine term structure model and take the changes around monetary policy announcements as the proxy for monetary policy shocks (IV2). Finally, benchmarking the approach in Romer and Romer (2004) and Champagne and Sekkel (2018), we use residuals in the forward-looking Taylor-rule equation as a proxy for monetary policy shocks (IV3). To the extent that each instrumental variable may deliver different information about the

[6]See Faust et al. (2003), Bjørnland (2009), and Gertler and Karadi (2015) on the endogeneity issues.

[7]One critical reason for this omission may be that there are no futures markets with active trading for the operating targets of monetary policy in those countries, and thus high-frequency identification of monetary surprise is not easily applicable. However, there are a few recent studies that employ high-frequency external instruments in other countries, including the case of the United Kingdom as in Cesa-Bianchi, Thwaites, and Vicondoa (2020). Using high-frequency movements in three-month sterling futures rates in the United Kingdom as a proxy for monetary policy surprise, the paper finds a significant impact from U.K. monetary policy shocks on its own economy.

monetary shocks, we employ these instrumental variables together in our analysis.[8]

We focus on the case of Canada as the best candidate country for this study. First, Canada is an advanced economy equipped with highly developed financial and credit markets. This allows us to test a variety of channels of monetary transmission. Second, Canada has adopted inflation targeting since the 1990s, when it adopted the flexible exchange rate regime and used short-term interest rates as the operating target for monetary policy. The records of monetary policy reports since then enable us to extract information on the Bank of Canada's own expectations of future macroeconomic situations and monetary policy stances. These features help obtain different types of instrumental variables on monetary policy shocks. Finally, compared with other small open economies, Canada shows the greatest macroeconomic and financial connections with the United States, which helps validate our selection of the United States as a center country (e.g., Cushman and Zha 1997).

Our research is closely related in its methodology and empirical results to the fast-growing body of recent studies on international spillovers of monetary policy (Rey 2015, 2016; Rogers, Scotti, and Wright 2018; Kearns, Schrimpf, and Xia 2020). Expanding on these studies, but diverging from them, our paper seeks to find commonality and heterogeneity in the impact of domestic and foreign monetary policy shocks on domestic financial and goods markets, and policy implications for the effectiveness of monetary policy implementation. Our study is also different from earlier studies in that our sample periods cover both pre- and post-global financial crisis periods, and, to that end, we resort to multiple sets of instrumental variables for domestic monetary policy shocks. Our paper is closely related, regarding domestic monetary policy transmission in Canada, to the

[8]For instance, movements in financial prices (IV1) are the closest in nature to the high-frequency futures data. However, they can include information shocks (Jarociński and Karadi 2020). In addition, when interest rates were close to their effective lower bounds or when unconventional policies were implemented, high-frequency variables may reveal merely limited information about the policy. The other type of IVs may exploit the policy information embedded in the yield curve irrespective of monetary policy regime (IV2) or distinguish the central bank information shocks (IV3).

analyses by Roldós (2006) and Champagne and Sekkel (2018). However, this paper differs from the earlier studies in that our focus is on the monetary transmission through multiple financial markets in Canada. Our work expands the discussion on the various channels of international monetary spillovers as in Iacoviello and Navarro (2019) and Ca' Zorzi et al. (2020), although we focus on a focal economy rather than multiple economies. Finally, this paper supplements a large group of studies on the trilemma debates. Our study is differentiated in the sense that we discuss hypotheses on both views of trilemma and dilemma within a single country framework while earlier studies mostly provide cross-country evidence that supports either of the two opposing views. In addition, while many earlier studies resort to regression analysis or event-study framework, we employ SVAR models with a novel identification scheme to examine dynamic impacts of local and foreign shocks.

The rest of this paper is organized as follows. In Section 2, we provide an overview of the theoretical channels of domestic and international monetary policy transmissions in the context of open-economy structural models. In Section 3, we specify an SVAR model and its identifying restrictions. Section 4 summarizes the empirical results. Section 5 presents the results of the robustness exercises and Section 6 concludes.

2. Monetary Policy Transmission in an Open Economy

In this section, we justify our SVAR framework by reviewing the theoretical channels of domestic and international transmission of monetary policy shocks in a small open economy. Our main focus is to understand the role of each transmission channel on the monetary policy independence of the economy in the context of the trilemma or dilemma hypothesis.

2.1 Domestic Monetary Policy Transmission

We first unravel the channels of domestic monetary policy transmission. Standard New Keynesian models, which assume sticky prices and frictionless financial markets, posit that monetary policy shocks are transmitted to credit costs and thus to aggregate spending operations via yield curves. Given the expectations hypothesis of the term structure, the effect of monetary policy decisions on the paths of current and expected short-term interest rates is summarized in (1):

$$r_t^m = m^{-1} E_t \left[\sum_{j=0}^{m-1} r_{t+j} \right] + \xi_t^m, \tag{1}$$

where r_t^m is an m-period zero-coupon government bond yield at time t, r_t is a short-term interest rate (e.g., the central bank policy rate), and ξ_t^m is an m-period term premium.

The term premium captures additional compensation for the interest rate (duration) risk inherent in medium- or long-term bond positions, as well as the residual effects of idiosyncratic market factors. If the premium is assumed to be constant over time, changes in the path of short-term rates will dominate changes in long-term rates, and allow central banks to influence output and inflation (*interest rate channel*).

When there is a degree of financial friction, credit markets would play an important role in the transmission of monetary shocks into financial and macroeconomic conditions (Bernanke and Gertler 1995). For instance, corporate bond yields (r_t^{cb}) usually exceed sovereign bond rates with the same maturity (r_t^m) to compensate for external finance premium (x_t^m), as in (2):

$$r_t^{cb} = r_t^m + x_t^m. \tag{2}$$

The *credit channel* particularly highlights the accelerating effect of monetary policy shock; for instance, contractionary monetary policy shocks tighten financial constraints in the private credit market and thus raise credit spreads (e.g., Gertler and Karadi 2015).

Finally, monetary policy shifts in a small open economy affect the value of domestic currency as indicated in the uncovered parity condition in (3).

$$r_t = r_t^* + E_t [\Delta e_t] + \rho_t, \tag{3}$$

where e_t is the nominal exchange rate vis-à-vis U.S. dollar and ρ_t is the currency risk premium in open economies at time t. Changes in foreign exchange rates then bring about changes in the relative price of tradable goods and in the value of assets denominated in foreign currency, and finally foreign demand for domestic products (*exchange rate channel*).

2.2 International Monetary Policy Spillovers

2.2.1 Financial Channel

The impact of foreign monetary shock on the domestic economy is another key issue in understanding monetary policy independence. This is because the extent of a central bank's control over macroeconomic developments, especially when looking through the lens of a small open economy in a financially integrated world, is controversial; policy and other monetary shocks can migrate from other countries under financial globalization, possibly causing monetary spillovers even when exchange rates float freely (Bruno and Shin 2015a; Passari and Rey 2015; Rey 2016). Taking this into account, we first consider international monetary transmission mechanisms which operate through short- and long-term yield structures.

With a high level of capital and financial market integration, a country's manipulation of short-term rates (r_t^*), especially if it is a large open economy such as the United States, directly affects short-term rates (r_t) in the other country following the interest-parity relationship represented in (3). Although, according to the Mundell-Fleming model, changes in the interest rate differential between the two countries are assumed to be absorbed mainly by adjustments in exchange rates, market interest rates in an open country are likely to be influenced by foreign monetary policy shocks, depending on the behavior of the exchange rate and the risk premium. For instance, the international connection between each country's long-term bond yields can be navigated in the form of (4) which combines Equations (1) and (3):

$$r_t^m = \underbrace{r_t^{*m}}_{(i)} + m^{-1} E_t \left[\sum_{j=0}^{m-1} \left(\underbrace{\Delta e_{t+j}}_{(ii)} + \underbrace{\rho_{t+j}}_{(iii)} \right) \right] + \underbrace{\xi_t^m - \xi_t^{*m}}_{(iv)}. \qquad (4)$$

Equation (4) implies that unexpected monetary policy shocks in a center country at first adjust market interest rates in a certain open economy (i). They also put additional pressure on market rates depending upon the responses of exchange rates and risk (term) premiums. If the balance sheets of borrowers and lenders in the open economy are denominated in U.S. dollars, for instance, the strong dollar caused by contractionary U.S. monetary shocks can tighten credit conditions in the open economy as well (*ii*). This is because a debtor's balance sheet becomes weak due to high liabilities relative to assets, and a creditor's lending capacity also drops. This retards economic activity and deteriorates the government fiscal position in the open economy. U.S. monetary tightening may also raise perceived risk and uncertainty in international financial markets. Consequently, the tightening can boost tail risks for the sovereign bonds of small open economies (*iii*), and compress capital flows into those bonds (*iv*), thereby leading to potentially unintended procyclical dynamics in their bond markets (*risk-taking channel*; Bruno and Shin 2015a; Hofmann, Shim, and Shin 2016, 2020). These channels work in the reverse in the case of monetary policy easing in the United States.

Finally, in a highly integrated financial market, particularly where the U.S. dollar is predominant as a funding and an investing currency, U.S. monetary policy shocks can also affect the net worth of agents through corporate bond markets in small open economies, thus making their financial conditions co-move (*international credit channel*; Rey 2016; Cesa-Bianchi and Sokol 2017).[9]

2.2.2 Trade and Aggregate Demand Channels

On top of the aforementioned financial channel, there are other types of transmission channels of U.S. monetary policy into open economies via trade and aggregate demand.[10]

The trade channel is based on the predictions of demand substitution between home- and foreign-produced goods and services, followed by the shifts of monetary shocks and the subsequent changes in the terms of trade. For instance, U.S. monetary tightening is

[9] As in Bernanke (2017), if r_t denotes a shadow price of credit, Equation (3) or (4) captures foreign credit availability in an open economy and ρ_t reflects the external finance premium.

[10] Note that some other studies refer to these channels as exchange rate channel and trade channel, respectively.

expected to lead to an appreciation (depreciation) of the U.S. dollar (other currencies) and, in turn, to enhance the competitiveness of open economies with flexible exchange rates. Output in the open economy will then rise, boosted by cheaper exports.[11] Thus the trade channel implies that the effects of monetary policy shocks on domestic and foreign economies are in an opposite direction in the case of open economies with flexible exchange rates such as Canada.

The aggregate demand channel rests on the idea of cross-border real spillover through trade. For an open economy which trades actively with the rest of the world, a substantial part of its aggregate demand is affected by its trading partners' business cycle. For instance, higher U.S. interest rates reduce incomes and expenditures in the United States, leading to lowered U.S. demand for both domestically produced and imported goods, and reducing activity and GDP abroad (Erceg, Guerrieri, and Gust 2005). Unlike the trade channel, the aggregate demand channel thus induces the effects of monetary policy shocks on domestic and foreign economies in the same direction.

Overall, the relative strength of each channel should depend on the share of exports and imports in economic activity, especially with the United States. In addition, more recent studies focus on the exchange rate pass-through to import (and export) prices in determining the strength of the trade channel. In this theory, currency invoicing in import and export prices plays a crucial role in the transmission of foreign monetary policy shocks into trades, output, and inflation in open economies (Cao, Dong, and Tomlin 2015; Gopinath 2015; Devereux, Dong, and Tomlin 2017; Gopinath et al. 2020).[12]

In the case that export items are priced mostly in the currency of the producer (commonly referred to as *producer-currency pricing*), as the Mundell-Fleming model predicts, a strong effect of expenditure switching is expected. For instance, a U.S. monetary tightening widens U.S. trade deficit (while exports fall, imports

[11]By contrast, a country that pegs its exchange rate against the U.S. dollar could experience the pressure of a currency appreciation that ultimately lowers its gross domestic product (GDP).

[12]For a detailed literature review of the relation between currency invoicing and exchange rate pass-through, see Ha, Stocker, and Yilmazkuday (2019) and Ca' Zorzi et al. (2020).

expand towards foreign goods in the United States) and trade surplus in its trading partners. On the contrary, if exports are priced in the currency of the importer (*local-currency pricing*), U.S. monetary policy spillovers through expenditure switching and exchange rate pass-through to inflation could be largely muted in non-U.S. economies. Finally, when all exports are priced in a single currency (*dominant-currency pricing*), the effects of U.S. dollar appreciation are inconsequential in the United States, specifically on its imports from abroad, since the prices of imported goods are unchanged. In a non-U.S. economy, however, a widespread rise in import prices is expected because of currency depreciation against the U.S. dollar, which induces expenditures switching away from imports and towards domestically produced goods.[13]

2.3 Dilemma vs. Trilemma Debates

The different nature of theoretical channels of international monetary policy spillovers leads to an active debate over the effects of foreign monetary shocks on the domestic financial market and macroeconomic conditions and the effectiveness of domestic monetary policies in open economies.

A group of recent studies emphasizing the role of global factors in driving domestic monetary policies has received much attention (Rey 2015, 2016). They argue that flexible foreign exchange regimes do not necessarily guarantee monetary policy independence in a world of open financial and capital markets. This is because monetary policy decisions in large economies inevitably affect global financial conditions, in turn affecting small open economies which typically have a high dependency on foreign currency borrowing. To the extent that market interest rates in small open economies are significantly affected by global financial conditions, their movements often deviate from a central bank's policy stance (Turner 2013). More recent literature highlights this aspect by focusing on international credit or the risk-taking channel (Bruno and Shin 2015a, 2015b; Passari and

[13]Note that the dominant-currency paradigm relies on the key assumptions that exporters have a substantial degree of monopoly power and U.S. dollar prices are sticky. Recent studies raise questions if these assumptions can hold outside the United States. See McLeay and Tenreyro (2020), for example.

Rey 2015),[14] or the international portfolio rebalancing channel as in Blanchard et al. (2016) and Alpanda and Kabaca (2020). As a result, central banks in small open economies can face a dilemma, rather than a trilemma, if the fluctuations in the exchange rate cannot fully insulate domestic economy from the impacts of external shocks, and if the policies designed considering only domestic conditions can result in unintended results including some trade-offs between output and inflation or between macroeconomic stabilization and financial stability.

Another strand of studies maintains that the trilemma remains alive (Obstfeld 2015; Aizenman, Chinn, and Ito 2016; Bekaert and Mehl 2019; Obstfeld, Ostry, and Qureshi 2019). These studies argue that exchange rate flexibility is still crucial in preserving the independence of monetary policy; as the Mundell-Fleming model predicts, the effects of foreign monetary spillovers on a small open economy are expected to be mitigated when adjustments in exchange rates change the terms of trade and trade balance in the economy. The studies thus argue that changes in monetary policy still can steer domestic inflation and output gap targets regardless of external developments (Bernanke 2017). Increased co-movement of interest rates across countries may be largely attributable to business cycle synchronization rather than intensified financial interconnection across jurisdictions (Klein and Shambaugh 2015; Aizenman, Chinn, and Ito 2016; Caceres et al. 2016). According to their views, financial integration can even enhance the effectiveness of monetary policy because currency appreciation after policy rate rises debases the value of foreign assets, thereby having an aggregate-demand-reducing negative wealth effect (Georgiadis and Mehl 2015).

Finally, a group of recent studies (among them, Han and Wei 2018; Cheng and Rajan 2020) focuses on asymmetric effects of foreign (or global) shocks and suggests the hypothesis of “2.5-lemma” or something between a trilemma and a dilemma.

[14]These studies suggest that changes in credit condition or risk appetite in international financial markets translate into local financial markets in open economies through the global financial factor or global financial intermediaries.

3. Estimation of SVAR Model

3.1 SVAR Modeling

We assume the economy is described by a structural form equation (5):

$$AY_t = \sum_{i=1}^{g} B_i Y_{t-i} + \varepsilon_t, \tag{5}$$

where Y_t is an $n \times 1$ vector of macroeconomic and financial variables. A and $B_i (\forall i \geq 1)$ are non-singular coefficient matrices. ε_t is an $n \times 1$ structural disturbances vector and serially uncorrelated. $E(\varepsilon_t \varepsilon_t') = I$ where I is the identity matrix (i.e., structural disturbances are assumed to be mutually uncorrelated). g denotes the optimal number of VAR lags, which can be set based on the information criteria.[15] For notational brevity, the specification in (5) omits deterministic terms and exogenous regressors.

Pre-multiplying each side of the equation by A^{-1}, we obtain a reduced-form representation (6):

$$Y_t = \sum_{i=1}^{P} \alpha_i Y_{t-i} + e_t, \tag{6}$$

where $\alpha_i = A^{-1} B_i$, and e_t are the reduced-form residuals which are related to the structural shocks by (7):

$$e_t = \begin{bmatrix} e_t^p \\ e_t^q \end{bmatrix} = S\varepsilon_t = [s^p s^q] \begin{bmatrix} \varepsilon_t^p \\ \varepsilon_t^q \end{bmatrix} \tag{7}$$

with $S = A^{-1}$. e_t^p are the residuals of domestic and foreign monetary policy instruments (i.e., $e_t^p = \left[e_t^{MP*} e_t^{MP}\right]'$) and e_t^q is a vector for the residuals of the other variables. The analogous definition applies to structural shocks ε_t^p and ε_t^q. s^p and s^q denote the column in matrix S that corresponds to the impact of structural policy

[15]We consider the Schwarz information criteria or Hannan-Quinn information criteria. In our estimation, we set two lags ($g = 2$) and find no serial correlation of VAR residuals.

shocks ε_t^p and ε_t^q, respectively, on the vector of reduced-form residuals (e_t). The variance-covariance matrix of the reduced-form VAR is $\Sigma = E[e_t e_t'] = E[SS']$.

The structural moving average representation as a function of structural shock is given as (8):

$$Y_t = \sum_{j=0}^{\infty} C_j S \varepsilon_{t-j} = \sum_{j=0}^{\infty} C_j s^p \varepsilon_{t-j}^p + \sum_{j=0}^{\infty} C_j s^q \varepsilon_{t-j}^q, \quad (8)$$

where C_j denotes the coefficients of the structural moving average (MA) form. Accordingly, if the endogenous variable responds to monetary policy innovations, the impulse response function (IRF), which is the dynamic response of the k-th element of vector Y (Y_k) to a unit shock of ε_t^p at time $t + j$, can be obtained by (9):

$$IRF_{k,j} = \frac{\partial Y_{k,t+j}}{\partial \varepsilon_{k,t}^p} = C_{k,j} s^p, \quad (9)$$

where $C_{k,j}$ is the k-th row of C_j.

3.2 Data

Open-economy monetary SVAR models in the literature typically consider short-term interest rates, foreign exchange rates, and macroeconomic variables such as output and price as endogenous variables for domestic economy in studying the monetary policy transmission (Kim 2001; Bjørnland 2009; Passari and Rey 2015; Rey 2015, 2016). Expanding on this, we employ 11 monthly macroeconomic and financial variables in the SVAR, reflecting the theoretical setup described in Section 2: logs of seasonally adjusted U.S. and Canadian consumer price index (P* and P), logs of seasonally adjusted U.S. industrial production (Y* and Y), U.S. and Canadian policy interest rates (MP* and MP), three-month and five-year Canadian government bond yields (R3m, R5y),[16] short- and long-term credit spreads (CS3m, CS3y), capital inflows to Canada (in Canadian dollar; CF), and logs of the nominal foreign exchange rate

[16]The variables are specified in levels to implicitly determine any potential co-integrating relationship between them; see Hamilton (2020).

Table 1. List of Data

Category	Variables
United States	Y*: Industrial Production (S.A.) P*: PCE Inflation (S.A.) MP*: Effective FFR
Canada	Y: Industrial Production (S.A.) P: Consumer Price Index (S.A.) MP: Money Market Financing Rates R3m: TB (Three-Month) Yields R5y: TB (Five-Year) Yields CS3m: Corporate Paper Rate – TB Rate (Three-Month) CS3y: Mortgage Bond Rate – TB Rate (Three-Year) CF: Net Capital Inflow to Canada (in Mil. Canadian Dollar) FX: Nominal Foreign Exchange Rate per U.S. Dollar
Control Variables	Commodity Price Index; U.S. Dollar Index; CBOE VIX Crisis Dummy Variable with 1 for the Period between September 2008 and June 2009

against one unit of the U.S. dollar (FX).[17] Following the prior literature, four external variables are included in the SVAR system to isolate exogenous latent factors that can affect endogenous variables simultaneously: the international commodity price index, a dummy variable for the global financial crisis, the Chicago Board Options Exchange (CBOE) volatility index (VIX), and the dollar index.[18] The sample period is January 2000–December 2017.[19] Table 1 summarizes the description of the data.

[17]In Section 5, we additionally test the robustness of main results by replacing foreign policy rates (effective federal funds rates) with shadow overnight rates as proposed in Wu and Xia (2016) and U.S. T-bill yields with one-year maturity.

[18]Unlike Rey (2015, 2016), we consider the VIX as an external variable because we focus more on the direct spillovers of U.S. monetary shocks. However, our robustness test (as shown in Section 5) which includes the VIX as an endogenous variable confirms that the main results are not sensitive to this alteration.

[19]Note that part of the sample considered in the paper was the period when U.S. interest rates were stuck at or close to the zero lower bound and when U.S. monetary policy involved unconventional measures, including quantitative easing (QE) and various forms of forward guidance. Thus, we compare the result with the ones in pre- and post-crisis periods to verify the robustness. See Section 4.2.2 for the results.

3.3 Identification of Monetary Policy Shocks

We employ the external instrument identification strategy, which avoids imposing any strong assumptions on the contemporaneous interactions among endogenous variables. Expanding on the recent studies, we recover structural parameters related to monetary policy shocks using a variety of instrumental variables. The novel part of our analysis is that we consider the transmission of domestic and foreign monetary policy shocks together in a single framework while avoiding the simultaneity problem. This enables us to evaluate and compare the overall impacts of each shock in an open economy. More specifically, this unified empirical setup also helps analyze the impact of domestic monetary shocks while we isolate the impact of foreign shocks, and vice versa. Without considering both types of monetary policy shocks, especially in a highly open economy such as Canada, the identification of monetary policy shocks and their dynamic impacts may be biased due to the omitted-variables problem. In identifying domestic and foreign monetary policy shocks, we follow Mertens and Ravn (2013)'s approach to orthogonalize the two shocks by assuming that U.S. monetary policy shocks have a contemporaneous impact on local (Canada) monetary policy but not vice versa.[20] The procedures for the identification of monetary policy shocks are summarized in Appendix A.2.

3.4 Instrumental Variables

A valid instrument for monetary policy shocks should satisfy the following two conditions as in (10) and (11):

$$rank(E[Z_t \varepsilon_t^p]) = L \qquad \text{(relevancy)} \qquad (10)$$

$$E[Z_t \varepsilon_t^q] = 0 \qquad \text{(orthogonality)}, \qquad (11)$$

where L is the number of endogenous variables, and Z is the instrumental variables.

[20] In this two-country VAR model, we also impose a block exogeneity restriction in Equation (6). In other words, we assume that a small open economy, Canada, does not have any feedback effects on a foreign country or the world economy, the United States. See for example, Cushman and Zha (1997), Kim and Roubini (2000), Cesa-Bianchi and Sokol (2017), and Dedola, Rivolta, and Stracca (2017).

Given that the effects of monetary policy on the economy are determined by the reaction of market participants to monetary policy shocks, the literature has extensively used the changes in short-term futures rates around the announcements of monetary policy decision as a proxy for monetary policy surprise (Kuttner 2001; Gürkaynak, Sack, and Swanson 2005; Gertler and Karadi 2015; Miranda-Agrippino 2016; Cesa-Bianchi, Thwaites, and Vicondoa 2020).[21] Such variations reflect changes in the expectations of market participants regarding future interest rates (or monetary policy stance).

Following Gürkaynak, Sack, and Swanson (2005) and Gertler and Karadi (2015), in identifying U.S. monetary policy shocks, we use changes in the federal fund futures rates and Eurodollar futures rates (MP1, FF4, ED2, ED3, and ED4) with a variety of maturities, within a narrow (30-minute) window around FOMC meetings. These variables are now extensively employed in the literature as proxies for U.S. monetary surprises in that they capture exogenously the revisions in market participants' expectation around the monetary policy announcements. We extend the high-frequency series of U.S. monetary policy shocks from that of Gertler and Karadi (2015) to 2017.[22]

On monetary policy shocks in Canada, since the country is not yet equipped with derivative markets for monetary policy instruments with ample depth, we instead use an alternative set of high-frequency measures of short-term interest rates (repo rates, prime rates, overnight rates, three-month government bond yields) (classified as IV1). However, the recent literature argues that the high-frequency data may contain information not only about policy

[21]This includes high-frequency movements (e.g., 30-minute window) of short-term futures rates (federal funds futures rates and three-month sterling futures rates) around monetary policy decision meetings.

[22]Here we use the instrumental variables by Gertler and Karadi (2015) considering their relevancy. Nakamura and Steinsson (2018) and Jarociński and Karadi (2020) differentiate central bank information shocks with pure monetary policy effects. As a robustness check, we test the instrumental variables, decomposing the series into the pure monetary policy shocks and central bank information shocks, similar to Jarociński and Karadi (2020). The results are not qualitatively different. The details are provided in Section 5.

Table 2. Instrumental Variables

Country	Category	Description
United States	MP1	Changes in the expectations of current-month federal funds futures rates (FFFRs)
	FF4	Changes in 3-month-ahead FFFRs
	ED2	Changes in 6-month-ahead Eurodollar futures rates (EDs)
	ED3	Changes in 9-month-ahead EDs
	ED4	Changes in 12-month-ahead EDs
Canada	Market Based (IV1)	MP surprise = daily change in the short-term spot rates on MP decision date (overnight, repo, and prime rates)
	Model Based (IV2)	Change of expected sum of short-term rates (EH) (computed by the affine term structure model)
	Narrative Based (IV3)	Residuals from policy reaction function of the Canada central bank (Romer and Romer 2004; Champagne and Sekkel 2018)

but also about central banks' assessment on the economic outlook (Jarociński and Karadi 2020) and that they provide only limited information when unconventional policies were adopted. To overcome these limitations, we also test other types of instrumental variables: changes in expected future short-term rates implied by the term structure model (IV2), and residuals of central bank policy reaction functions (IV3). The instrumental variables tested for Canada and the United States are explained in detail below and summarized in Table 2.[23]

[23]Following Gertler and Karadi (2015), we construct monthly-frequency instrumental variables by taking the following steps. First, for each day of the month, we compute cumulative monetary surprises over the recent 31 days; and second, we compute monthly surprise series using the sum of the cumulated daily surprises within each month.

3.4.1 IV1: Daily Short-Term Spot Rate Changes Around Monetary Policy Announcements

Following Cochrane and Piazzesi (2002) and others, we first consider the daily movements of short-term interest rates around monetary policy decision announcements, by defining the daily change in the spot rates as a monetary policy surprise. Financial market participants anticipate monetary policy decisions before actual policy announcements, and short-term rates may have already been adjusted beforehand. Conversely, if the monetary policy announcement is a mere surprise, market rates will adjust only after the announcements.[24]

This approach rests on the following two assumptions. First, asset prices move according to the efficient market hypothesis. In such market conditions, new information, including a monetary policy decision, is reflected in the asset prices as soon as it is released. Second, short-term rates are more sensitive to monetary policy news than long-term rates because central banks typically adjust short-term rates to steer macroeconomic variables. This indicates that news other than a monetary policy decision on the dates can be regarded as white noise. Figure A.1 in Appendix A shows the movements of representative short-term rates for the United States and Canada. Short-term rates deviate significantly from policy targets when markets expect adjustments in the monetary policy stance.

3.4.2 IV2: Monetary Policy Surprise Implied in Term Structure Model

We calculate the conditional expectation for short-term interest rates using a standard affine term structure model, and take its changes around monetary policy decisions as the proxies of monetary policy shocks. The expectations hypothesis assumes that long-term interest rates consist of the expected path of short-term rates and term premium, as illustrated in Equation (1). Given that the

[24] Another possible market-based instrument for monetary surprise is a measure of the shifts in overnight index swap (OIS) rate in Canada. We test these instruments as an alternative to our IV1 and find that overall results do not change despite the wider confidence bands with the use of these instruments that mostly reflect the shorter sample periods. See Section 5 for the details.

current and future paths of short-term interest rates are directly linked to the effects of the interest rate channel and forward guidance, changes in the expected future path of short-term interest rates around monetary policy decisions will mirror the changes in a market participant's expectations of the monetary policy stance of central banks (Chari, Stedman, and Lundblad 2017; Curcuru et al. 2018; etc).[25]

We compute the changes in the expected future path of short-term interest rates from zero-coupon bonds with the maturities of 3, 6, 9, and 12 months. Data for zero-coupon rates are obtained from Bank of Canada. For the estimation, we follow Adrian, Crump, and Moench (2013) considering that the methodology has some computational advantages over typical estimation strategies such as maximum likelihood especially when yields of high frequency are used. See Appendix A.3 for technical details of the estimation.

3.4.3 IV3: Residuals from Policy Reaction Functions

Benchmarking the approach in Romer and Romer (2004), and the extension of the methodology to Canada as in Champagne and Sekkel (2018), we use residuals in the forward-looking Taylor-rule equation as a proxy variable for Canadian monetary policy shocks (IV3). The main idea is that by using internal forecast information in the central bank, we can extract a measure of unanticipated movement in monetary policy target rates (or surprise component) which is orthogonal to information about past, current and future economic developments.

We follow Champagne and Sekkel (2018) and take two steps in estimating the Taylor-rule equation. First, using the minutes of

[25]Conventional monetary policies are believed to affect bond yields mostly through current and expected future paths of short-term rates. Unconventional policies, including balance sheet policies, affect the expected path of short-term interest rates by signaling that short-term rates will remain low for long (signaling channel), and the term premium by influencing supply-demand imbalances in bond markets (portfolio balance channel). When measuring the impacts of monetary policy shifts using the moves of expected short rates, we consider the fact that the Bank of Canada had not adopted balance sheet policies over the sample period. In light of this, our approach is different from that of Inoue and Rossi (2019), who focus on the effects of monetary policy on the whole yield curve.

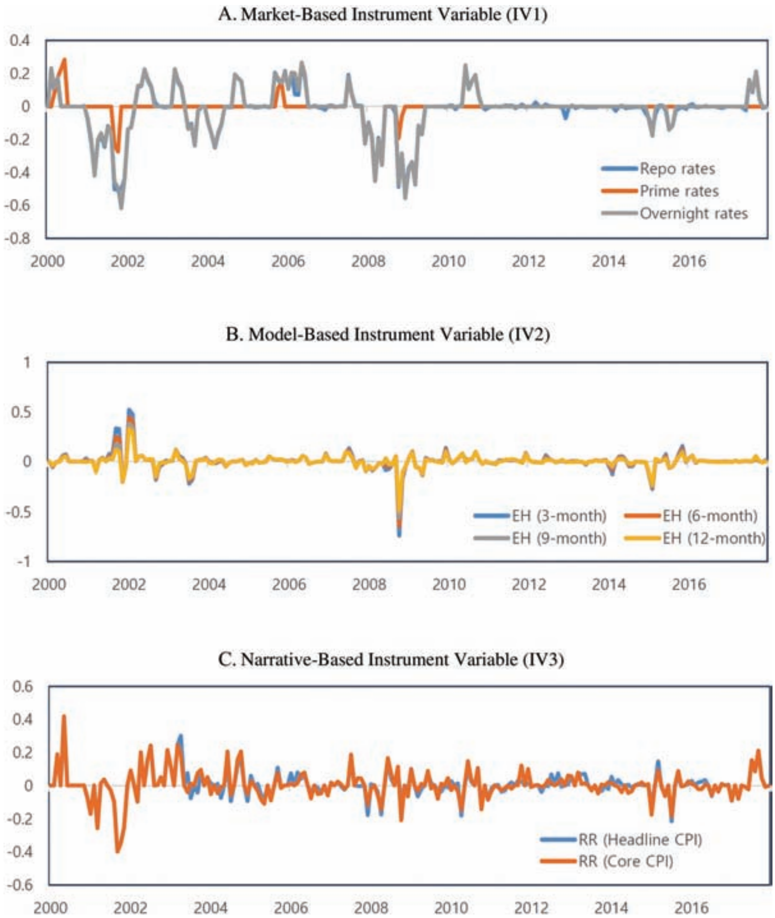
monetary policy reports (source: Bank of Canada Monetary Policy Reports), we collect real-time forecasts for output and inflation in Canada. We use both headline and core consumer price index (CPI) inflation for the inflation forecast. Second, we regress changes in monetary policy target rates from the previous monetary policy decision meeting to the current meeting (Δr_m) on a set of explanatory variables that purge the intended policy rate. Technical details of the estimation are presented in Appendix A.4.

3.4.4 Properties of Instrumental Variables

Figure 1 depicts the movements of selected instrumental variables over the sample period. In panel A, we show the monthly series of changes in representative three short-term rates in Canada—repo rates, overnight rates, and prime rates—around monetary policy decision dates (IV1). The series shows distinct movements in principle after the dot-com crash in early 2000 and the global financial crisis around 2008–09; however, the prime rates exhibit relatively less variation around the events than the other two rates. Panel B describes the instrumental variables related to the changes in the sum of expected short-term rates for the maturities of 3, 6, 9, and 12 months (IV2). All the variables follow a similar path, and the changes for the three-month bond move with larger variation over time. Panel C exhibits the residuals from the central bank's policy reactions functions, using the headline and core CPI as anchoring price measures (IV3). It is notable that the variables show comparatively less reaction during major episodes such as the global financial crisis, indicating that some of the variation in policy rates is already anticipated by economic agents.

Figure A.2 in Appendix A.1 summarizes the cross-correlations of the instrumental variables in the format of a heatmap. The figure suggests that the instrumental variables are highly and positively correlated to each other but with different degrees, which validates our use of multiple instrumental variables. The positive correlations are relatively higher among the same types of variables, and among the variables with the same countries (the average coefficients of the instrumental variables: the United States 0.74, Canada 0.42). The cross-correlations of instrumental variables between the United

Figure 1. IVs for Canadian Monetary Policy Shocks



States and Canada are 0.10 on average, which are smaller but not trivial.[26]

[26]By reporting strongly positive correlations among cross-country monetary policy shocks obtained from 280 macroeconomic models, Georgiadis and Jančoková (2020) argue that the multilateral monetary policy shocks can co-vary due to strong financial spillovers. In this context, given the close economic relationship between the two countries, the positive cross-correlation among shocks is not surprising. Reflecting this aspect, we also consider the monetary policy spillovers from the United States to Canada in identifying the monetary policy shocks, as explained further in Appendix A.2.

Next, we report t-statistics, F-statistics (in the case of multiple instrumental variables), and R^2s from the first-stage regression of residual of policy indicators projected on the instrumental variables to test the relevance of the instrumental variables. The results are summarized in Table 3 for the United States (panel A) and Canada (panel B).[27] To the extent that "F-statistics > 10" is commonly regarded as a rule-of-thumb criterion to protect against the weak instrumental-variable problem in practice, the instrumental variables for which F-statistics in the first-stage regression are higher than 10 are strongly relevant to the exogenous monetary policy shocks. The instrumental variables are finally chosen considering their relevancy and the type of data used for the construction.

4. Empirical Results

To investigate how the domestic and international monetary policy transmissions operate in Canada, we now present the impulse responses of financial, capital flow, and macroeconomic variables to domestic monetary policy shocks, and then to U.S. monetary policy shocks.

4.1 Effects of Canadian Monetary Policy Shocks on the Economy

Figure 2 displays the impulse response of Canadian variables to a contractionary monetary policy shock that increased overnight rates in Canada by 1 percentage point. Panels A through C in the figure sequentially report the empirical results using three different types of instrumental variables, as explained in the previous section. Figure 3 shows the results where we employ the three types of instrumental variables together. As the results are consistent across different types of instrumental variables, we focus on those with all types of instrumental variables that report the highest explanatory power (i.e., R^2) in the first-stage regression.

[27]Staiger and Stock (1997) suggest that the F-statistics of the instrumental variables should be greater than 10 to ensure that the maximum bias in the IV estimators is less than 10 percent. In the case of a single instrumental variable, the F-statistics should be replaced by t-statistics.

Table 3. Relevancy Test Results

A. U.S. IV						
		t				
F	R^2	$MP1$	$FF4$	$ED2$	$ED3$	$ED4$
17.51	0.30	5.81	–1.53	–0.63	0.98	–1.05
B. Canada IV						
Market-Based Instrument Variable (IV1)						
				t		
		F	R^2	*overnight*	*repo*	*prime*
$IV_{CA,1} =$						
{overnight}		58.31	0.22	7.64	—	—
{repo}		55.04	0.21	—	7.42	—
{overnight, repo}		**29.64**	**0.22**	**1.89**	**–0.99**	—
{overnight, repo, prime}		21.05	0.23	1.94	–1.15	1.80
Model-Based Instrument Variable (IV2)						
			t			
	F	R^2	EH_{3m}	EH_{6m}	EH_{9m}	EH_{12m}
$IV_{CA,2} =$						
$\{EH_{3m}\}$	7.50	0.03	2.74	—	—	—
$\{EH_{6m}\}$	11.36	0.05	—	3.37	—	—
$\{EH_{3m}, EH_{6m}\}$	**31.59**	**0.23**	**–7.02**	**7.34**	—	—
$\{EH_{3m}, EH_{6m}, EH_{9m}, EH_{12m}\}$	17.50	0.25	1.71	–1.75	1.78	–1.79

(*continued*)

Table 3. (Continued)

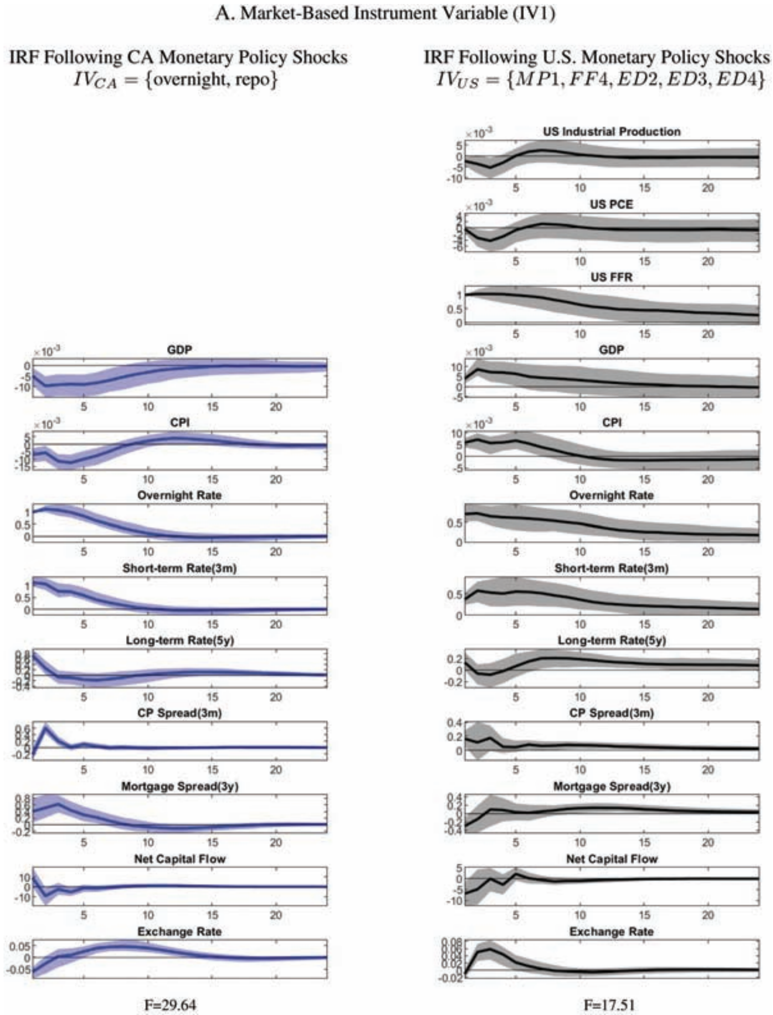
B. Canada IV

Narrative-Based Instrument Variable (IV3)

			t			
	F	R^2	RR_{hcpi}	RR_{ccpi}	CS_{new}	CS_{full}
$IV_{CA,3} =$						
$\{RR_{hcpi}\}$	81.10	0.28	9.01	—	—	—
$\{RR_{ccpi}\}$	92.55	0.30	—	9.62	—	—
$\{CS_{new}\}$	17.44	0.08	—	—	4.18	—
$\{CS_{full}\}$	27.87	0.08	—	—	—	5.28
$\{RR_{hcpi}, RR_{ccpi}\}$	**46.23**	**0.30**	**0.48**	**2.91**	—	—
$\{CS_{new}, CS_{full}\}$	14.86	0.14	—	—	1.32	3.37

Note: F, R^2, and t denote F-statistics, R^2, and t-statistics based on the first-stage regression of the instrumental variable(s) on the VAR residuals for the U.S. or Canada MP variable. Bold parts are the finally chosen IV sets for each category. Each instrumental variable name represents the type of data used for construction, which is explained in Table 2. For instance, "$IV_{CA,1}$\{overnight\}" and "$IV_{CA,1}$\{repo\}" indicate market-based instrumental variables (IV1) for Canadian monetary policy shocks based on overnight rates and repurchase rates, respectively. Similarly, "$IV_{CA,2}\{EH_{3m}\}$" indicates model-based instrument variables (IV2) based on the changes in expected sum of short-term (three-month) rates. "$IV_{CA,3}$" indicates narrative-based instrument variables, i.e., residuals from policy reaction function of the Canada central bank, benchmarking Romer and Romer (2004) ("*RR*") or Champagne and Sekkel (2018) ("*CS*").

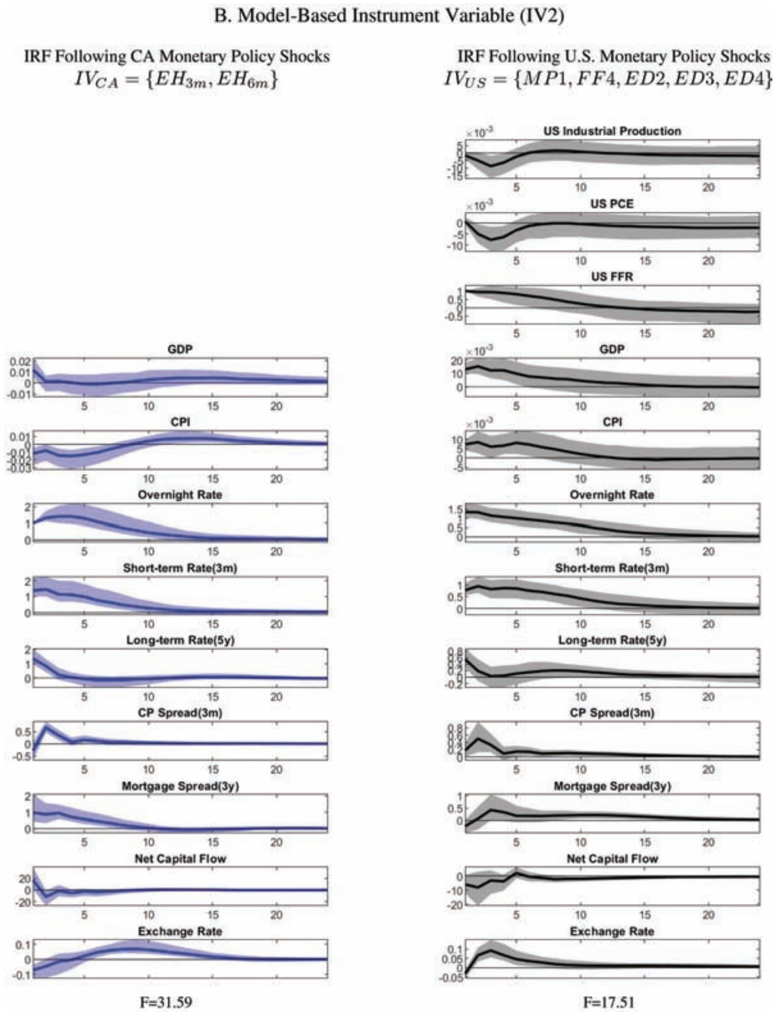
Figure 2. Impulse Response of Canada Variables to Monetary Shocks



(*continued*)

Market Interest Rates. An interest rate channel of monetary policy operates in Canada; following a contractionary domestic monetary policy shock, Canadian market interest rates respond significantly, although the magnitude and persistence of the impact

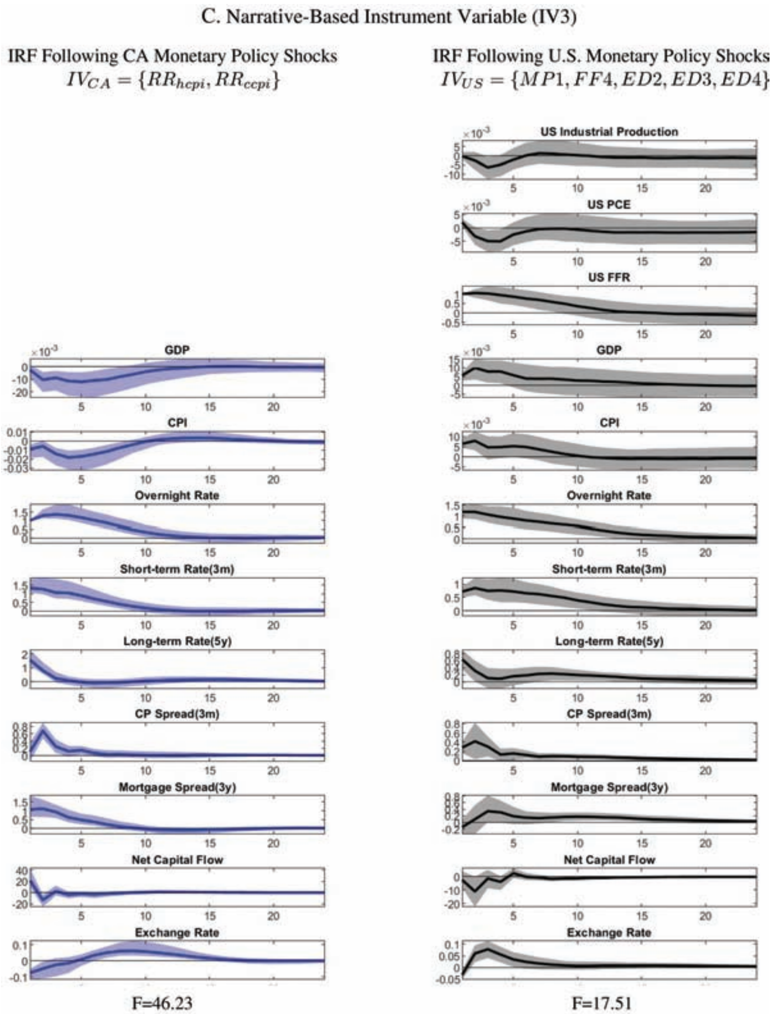
Figure 2. (Continued)



(*continued*)

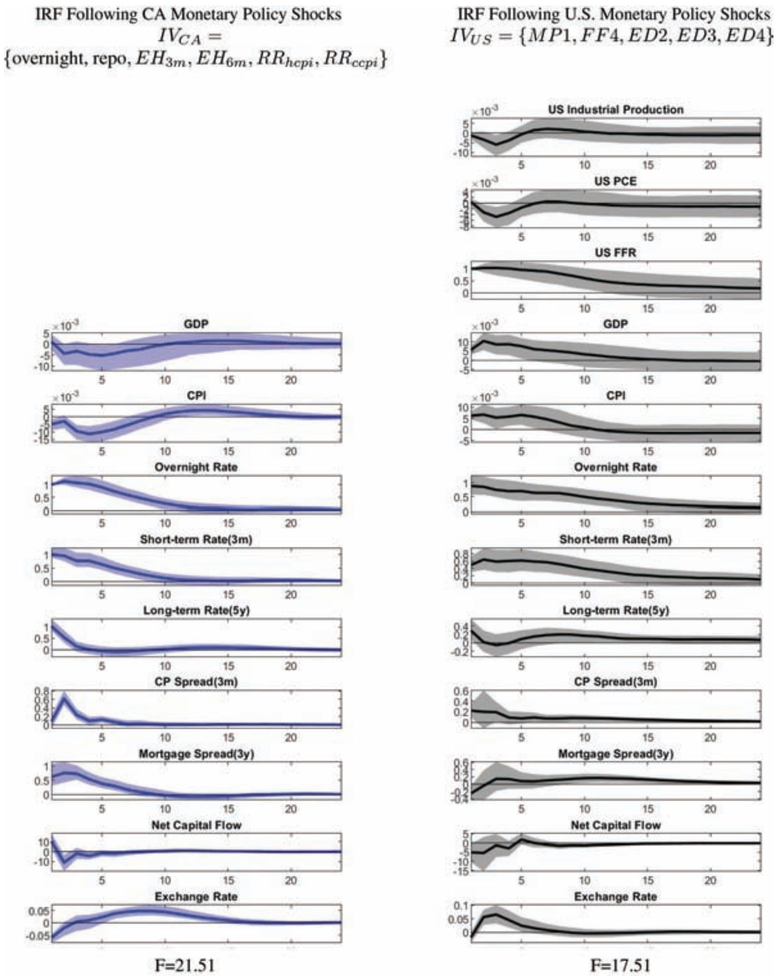
weaken with the bond yields with longer maturity. Short-term interest rates (three-month T-bill rates) move in tandem with overnight rates, by rising around 100 basis points (bps) on impact, and the effects persist for a year. Long-term rates (five-year bond yields)

Figure 2. (Continued)



Note: Y-axis indicates percent (or percentage point for interest rates, or mil. CAD for capital flow). X-axis indicates months after shock. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively. Shaded areas are the 16 and 84 percentiles of the empirical distribution based on wild-bootstrapped samples. "F" denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set.

Figure 3. Impulse Response of Canada Variables to Monetary Shocks



Note: Y-axis indicates percent (or percentage point for interest rates, or mil. CAD for trade). X-axis indicates months after shock. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively. Shaded areas are the 16 and 84 percentiles of the empirical distribution based on wild-bootstrapped samples. "F" denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set.

also increase sizably, but the impact dies out quickly—within two to three months after the shock.

More short-lived responses of long-term interest rates may reflect the offsetting effects of various factors that determine the level of long-term interest rates, as illustrated in Section 2. First, the impact of monetary tightening could be dampened due to the subsequent exchange rate appreciations and weakened future inflation expectations. Also, as a country with a high level of foreign currency debt, especially in U.S. dollars, currency appreciations can enhance the borrowing capacity in Canada, thereby reducing the tail risks associated with currency risk premium (Hofmann, Shim, and Shin 2016).[28]

Exchange Rates. Following the contractionary (+100 bps) monetary policy shock, the Canadian dollar immediately appreciates by around 5 percent, and then depreciates gradually until it reaches the original level. This is in line with the predictions of the overshooting theory in Dornbusch (1976), which is based on the uncovered interest rate parity (UIP) condition. Earlier studies often find that, following a contractionary monetary policy shock, domestic currency either depreciates (exchange rate puzzle; see Grilli and Roubini 1996), or if it appreciates, it does so for a prolonged period of up to three years, thereby exhibiting hump-shaped behavior that violates the UIP condition (delayed overshooting; see Eichenbaum and Evans 1995; Cushman and Zha 1997; Kim, Moon, and Velasco 2017). Unlike the foregoing findings, our results show that the initial appreciation of Canadian currencies on a contractionary monetary shock is not followed by long and persistent appreciation.

Capital Flows. Net foreign capital inflows to Canada increase immediately following a contractionary domestic monetary policy shock, partially reflecting the subsequent increase in domestic-foreign interest rate differentials as well as the appreciation of the domestic currency. This impact quickly dissipates, however, as the domestic currency starts to depreciate, and the negative impact of

[28]The ratio of foreign liability, including portfolio investment and loans, over nominal GDP has persistently increased in Canada (2000:Q4: 123.5% → 2018:Q4: 199.3%; source: Statistics Canada). U.S.-dollar-denominated debt is one of the dominant foreign-currency debts, which is 32.3 percent of total foreign liability.

monetary tightening is transmitted to macroeconomic variables over time.

Credit Spreads. A credit channel of monetary policy transmission also appears to operate in Canada, notably, through both short- and long-term financing premiums. Following a contractionary monetary policy shock, credit spreads increase up to 60 bps in short-term instruments (three-month CP spreads), with statistical significance. The credit spread under long-term instruments (three-year mortgage bond spreads) demonstrates a similar magnitude of response, but the impact persists for a prolonged period, seven to nine months after the shock.

Output and Prices. Following a contractionary monetary policy shock, both output (monthly GDP) and consumer prices in Canada significantly decrease by up to 1 percent. The impacts are maximized around two quarters after the shock, and persist for around a year. These results are consistent with New Keynesian theories that highlight the role of financial and credit markets in the monetary policy transmission.[29] The results are also in line with earlier empirical studies on monetary policy transmission in Canada, including Roldós (2006) and Champagne and Sekkel (2018), despite the different sample periods and different identifying assumptions of monetary policy shocks.

4.2 Effects of U.S. Monetary Policy Shocks on the Canada Economy

Next, we examine the effects of U.S. monetary policy shocks on the financial markets, trade, and macroeconomic variables in Canada. On the right-hand sides of Figures 2 and 3, we plot the impulse responses of the variables to a 1 percentage point increase in U.S. federal funds rates.[30] Again, our focus is mainly on the results with all types of instrumental variables for Canada monetary policy shocks.

[29]The response of output is weaker than expected and less statistically significant than prices, partly reflecting the capital inflows caused by exchange rate appreciation.

[30]As shown in Figures 2 and 3, U.S. monetary tightening is unambiguously followed by declines in U.S. output and prices levels, which are standard in the literature. In this section, we focus on the transmission of the shocks into the Canadian economy.

4.2.1 Financial Market Indicators

Policy Rates in Canada. There appears to be monetary policy synchronization between the United States and Canada. Overnight interest rates in Canada significantly increase following a contractionary U.S. monetary policy shock, and the impact persists longer than a year. The response is comparable in light of the magnitude and persistence of what follows domestic monetary policy shock.

What does this imply? Does it mean that the Canadian central bank does not have monetary autonomy? Given our significant results on the effects of domestic monetary policy, as shown in the previous subsection, the answer will be no. This result may instead reflect the economic dependence of Canada on the United States in the aspects of trade and financial transactions. On the one hand, the consequences of U.S. monetary policy shocks on the economy could spill over to Canada. Alternatively, this result may reflect synchronized monetary policy actions in Canada and the United States, to neutralize the impact of U.S. monetary policy shocks on Canadian financial markets by reducing the volatility in exchange rates and capital flows (Turner 2013). It is also likely that as an inflation-targeting country, Bank of Canada responds to inflation fluctuations driven by U.S. monetary shocks.

Market Interest Rates. In line with the interest rate parity condition, as illustrated in Equation (4) in Section 2, U.S. monetary tightening raises interest rates in Canada with both short- and long-term maturities; a 1 percentage point increase in the federal funds rates raises three-month T-bill rates and five-year bond yields in Canada by up to around 60 and 30 basis points, respectively, consistent with the findings in Ehrmann, Fratzscher, and Rigobon (2011). The increase in the Canadian market interest rates could be driven by several factors, as discussed in Section 2: the increase in U.S. bond rates and correlated movements in term premiums, the consequent increase of currency risk premium in Canada, or correlated movements in Canadian overnight rates.

The above results are different from some earlier findings in the literature, such as those of Kim (2001), who finds that short-term interest rates in non-U.S. G-6 countries do not react strongly to U.S. monetary policy shocks. The difference in the results may partly reflect structural changes over time, including the integration of financial markets. Our results suggest that the endogenous reaction of monetary policy instruments and market interest rates in Canada to U.S. monetary surprises is substantial and lasts longer than the response following domestic monetary shocks, consistent with what is found in Faust et al. (2003) and, more recently, in Rey (2015, 2016), Cesa-Bianchi, Thwaites, and Vicondoa (2020), and Miranda-Agrippino (2016) in the United Kingdom and Germany.[31]

Exchange Rates. Following a contractionary U.S. monetary policy shock, the Canadian dollar depreciates (i.e., U.S. dollar appreciates) up to 6 percent within three to four months, and it appreciates gradually, reverting toward long-term levels. Again, such a response, consistent with what was found in the response of the Canadian dollar following domestic monetary policy shocks, is compatible with the predictions of the overshooting hypothesis without any evidence of exchange rate puzzle or delayed overshooting. In line with Kim and Roubini (2000), Rogers, Scotti, and Wright (2018), and Inoue and Rossi (2019), this result suggests that the inappropriate identification of monetary policy shocks may account for the puzzles observed in the previous literature.

This finding is also important in the sense that the trilemma-related debates are closely related to different views on the role of the exchange rate as a shock absorber. In particular, the conventional Mundell-Fleming model predicts that the international monetary spillovers into a small open economy are mitigated due to adjustments in local-currency values, which in turn change the terms of trade and trade balance in the economy. The flexible reactions of the exchange rate to U.S. monetary policy shocks therefore suggest that the variable can act as a shock absorber. The effects on Canadian trade are discussed further in the next section.

Capital Flows. Net capital inflows to Canada decline following contractionary U.S. domestic monetary policy shocks. Despite the increase in market interest rates in Canada following U.S. monetary tightening, the depreciation of Canadian currency and increase in credit costs in international financial markets could play a negative role in capital inflows to Canada.

[31]The results are also consistent with Rogers, Scotti, and Wright (2018) in the context of the size and persistence of U.S. monetary policy shocks, although the study focuses on the zero lower bound period.

Credit Costs. Short-term credit spreads significantly increase by around 20 basis points following a U.S. monetary tightening shock. Long-term credit spreads also increase to a similar degree, but with some lags, around 10 months after the shock. Indeed, a growing number of studies report empirical evidence that an international credit channel operates significantly, as there is a rapidly growing dependence on U.S.-dollar-denominated liabilities by small open economies, especially after the global financial crisis (Passari and Rey 2015; Rey 2015). Credit conditions in Canada are thus expected to be significantly affected by U.S. monetary tightening given that the Canadian economy demonstrates high reliance on the United States and a considerable portion of foreign debt is raised in U.S. dollars.

4.2.2 Trade Variables

We now turn our focus to the impact of U.S. monetary policy shocks on trades in Canada. Compared with the baseline model in the previous section, we substitute bilateral nominal exchange rate per U.S. dollar and net capital flows with nominal effective exchange rates (NEER) and trade balance (or net export). This exercise aims to test the role of exchange rates (NEER) at the occurrence of external monetary shocks as predicted by the Mundell-Fleming trilemma. The conventional Mundell-Fleming model predicts that spillovers from foreign monetary policy are dampened in a small open economy largely by the sequential adjustments in exchange rates and net exports.

Consistent with the results for bilateral exchange rates, as shown in panel A of Figure 4, the response of NEER does not exhibit any puzzling movements that deviate from the predictions of the overshooting hypothesis.[32] Following a contractionary U.S. monetary policy shock, Canadian NEER depreciates by around 4 percent within three months, and the impacts dissipate in a few months. The depreciation in domestic currency leads to an improvement in the competitiveness of Canadian goods and services in the

[32]Note that, by definition, the increase of NEER indicates an appreciation of currency while a decrease indicates depreciation, which is opposite to the bilateral exchange rates (CAD/USD).

Figure 4. Impulse Response of Canada Variables to Monetary Shocks

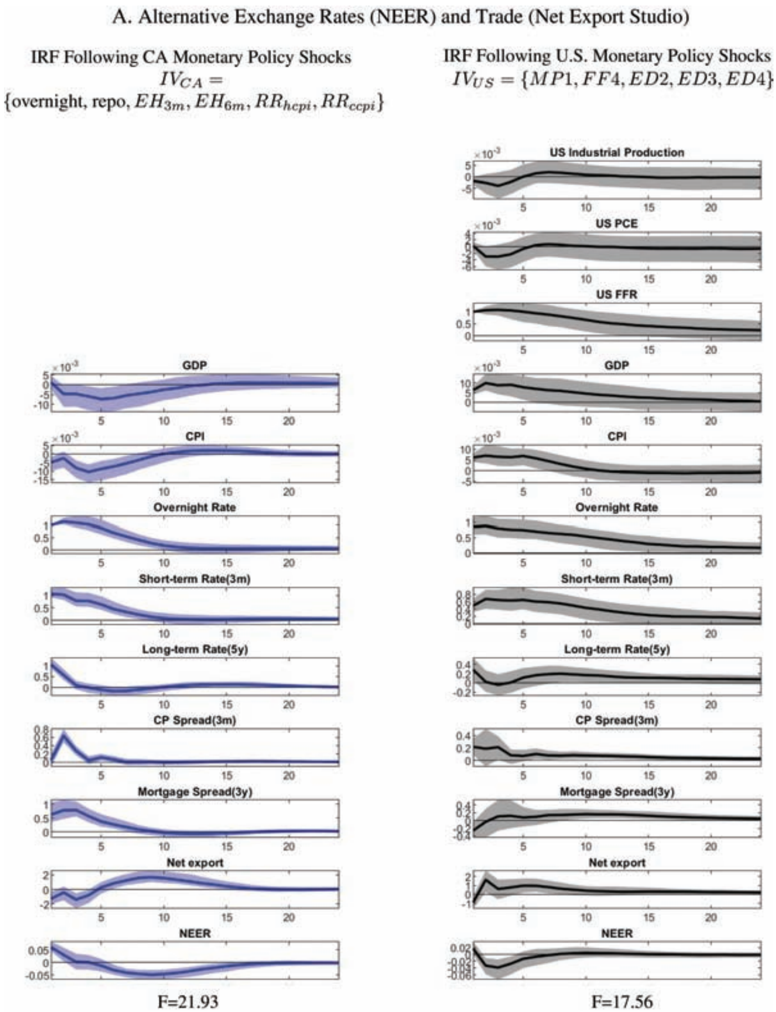


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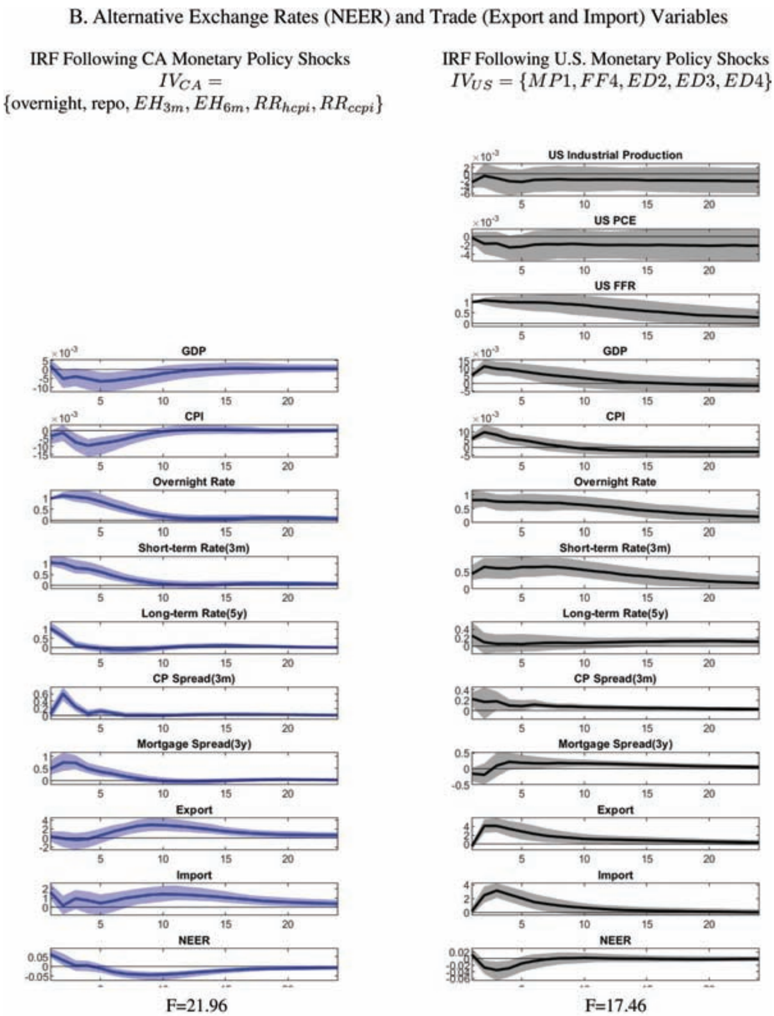
Figure 4. (Continued)



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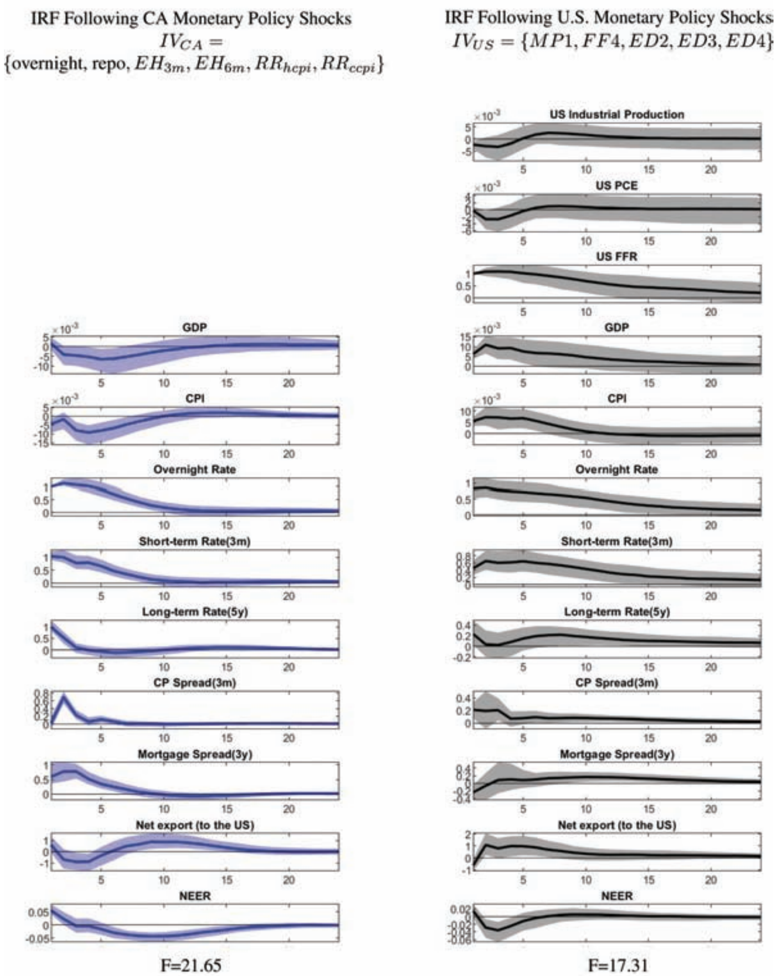
international market. This finally raises net exports, defined as exports minus imports, and its impacts persist for around a year.[33]

[33]Interestingly, Canadian GDP rises on impact, while net exports instantly drop, followed by a gradual rise. Such an initial gap may suggest that some other potential channels than the expenditure switching are at play. For instance,

Figure 4. (Continued)

C. Alternative Exchange Rates (NEER) and Trade (Export and Import with the United States)



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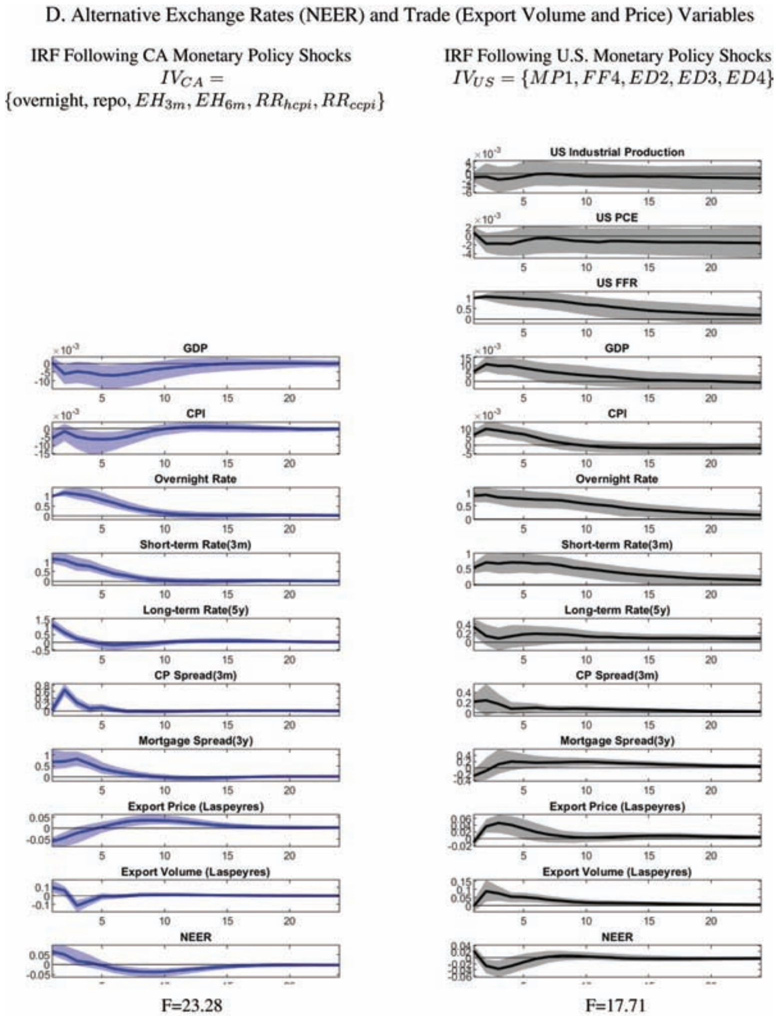
wealth effects may work in the very short term due to valuation effects in Canada's holdings of U.S.-dollar-denominated foreign assets. Alternatively, the immediate decline of the real interest rates, if any, could also lead to such responses.

Figure 4. (Continued)



Note: Y-axis indicates percent (or percentage point for interest rates). X-axis indicates months after shock. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively. Shaded areas are the 16 and 84 percentiles of the empirical distribution based on wild-bootstrapped samples. "F" denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set.

Meanwhile, following a contractionary domestic monetary policy shock, NEER appreciates and net exports significantly decline as expected.[34]

The increase in trade balance following U.S. monetary tightening appears to be mainly driven by an increase in exports rather than the decline in imports, as exhibited in panel B of Figure 4. The trade channel of monetary transmission is more pronounced when we consider the trades between the United States and Canada. Panel C of the figure reports that the positive impact of U.S. monetary tightening on net exports in Canada is statistically significant (i.e., confidence bands do not include zero throughout the forecasting horizons) and the impacts are somewhat more persistent compared with the total net export.

Given that a large part of Canada's imports from the United States is invoiced in U.S. dollars (Gopinath 2015; Gopinath et al. 2020), our results on the inflationary effects of U.S. monetary tightening seem quite consistent with the theory. In the dominant currency regime, domestic currency depreciation is expected to pass through into domestic import and consumer prices. Meanwhile, the expansionary effects of U.S. monetary tightening on Canadian exports and outputs seems somewhat puzzling because an appreciation of the U.S. dollar would not stimulate U.S. imports from Canada due to relative price movements, but instead, it would only depress Canadian imports from the United States.

However, it would be worth noting that the dominant currency paradigm relies on the key assumptions that exporters have a substantial degree of monopoly power and U.S. dollar prices are sticky. If the assumptions do not hold, as McLeay and Tenreyro (2020) argue, domestic currency depreciation could trigger a large increase in export volume rather than export prices. Considering that Canada is a small open economy that heavily depends on commodity exports, and that commodity prices are quite flexible while demands for commodities are elastic in global markets, the assumptions behind

[34]We also conduct a counterfactual analysis which blocks the reaction of exchange rates to monetary shocks (the results are available upon request). The responses of the variables following a U.S. monetary policy shock in this model are, by and large, greater than the ones obtained from the baseline results, especially for the financial variables. Again, this is in line with the Mundell-Fleming theory that exchange rates play an important role in mitigating the international monetary spillover.

the dominant currency invoicing may not hold strongly in Canada. Hence, if Canadian exporters are price takers with elastic demands in global markets, the expansionary consequence of U.S. monetary tightening through the trade channel could be stronger despite the dominant-currency invoicing. In light of this, panel D of Figure 4 provides the results of the model which includes Canadian export volume and export prices. U.S. monetary tightening is indeed followed by a large increase in export volume rather than export prices.[35]

4.2.3 Macroeconomic Variables

Finally, let us explore the impacts of U.S. monetary policy shocks on Canadian macroeconomic variables (output and price levels). The net impacts of U.S. monetary tightening on the variables could be, ex ante, either contractionary or expansionary. Put another way, if the negative impacts of U.S. monetary shocks on Canadian financial conditions and domestic demands (*financial channel* and *aggregate demand channel*) overshadow the positive impacts on exchange rate pass-through and net exports (*trade channel*), then output and price levels in Canada would decline after the U.S. monetary policy tightening (Iacoviello and Navarro 2019).[36] On the other hand, if the trade channel is at play more effectively than the financial and aggregate demand channels, the macroeconomic consequences of U.S. monetary tightening would be instead *expansionary*, consistent with the empirical results in Rey (2016) (for Canada) and Mirando-Agrippino and Rey (2020) (for the global economy), and the predictions by Jones, Kulish, and Rees (2018), where industrial production and price levels in Canada significantly increased following contractionary U.S. monetary policy shocks.[37]

[35]Alternatively, it could be the case that U.S. contractionary monetary spillover brings about the improvement of Canadian trade balance due to simultaneous factors such as global trade collapse around global economic slowdowns. To empirically test this possibility, we also carried out the estimation with the model which includes global trade volume as an additional control variable and confirmed that the main results did not change.

[36]Bluedorn and Bowdler (2011) report that a U.S. monetary tightening induces the expenditure-reducing effects rather than expenditure-switching effects. Di Giovanni and Shambaugh (2008) also show that high foreign interest rates have a contractionary effect on annual real GDP growth in the domestic economy, but that this effect is centered on the countries with fixed exchange rates.

[37]Similarly, Blanchard et al. (2016) show that a contractionary foreign monetary policy shock or, equivalently, capital outflows tend to result in expansionary

As was already shown in Figures 2, 3, and 4, following a contractionary U.S. monetary shock, both output and price levels in Canada unambiguously increase, and the impacts persist for around a year after the shock. The results suggest that the expansion of external demands by the trade channel outweighs the decline in domestic demands by financial and real spillovers, thus having positive net effects on the output in Canada. In addition, the positive response of price levels in Canada seems to reflect the effects of currency depreciation that are passed on to import prices, and finally to consumer prices.

To elaborate more on this, we estimate an additional model where net export and domestic demand are scaled by a ratio of GDP. In so doing, we can test whether the macroeconomic outcomes following U.S. monetary tightening result in the expenditure switching (from domestic demand to foreign demand). As shown in panel A of Figure 5, the ratio of net export per nominal GDP in Canada increases by around 2 percentage points following a U.S. monetary tightening. This indicates that the degree of increase in the net export is greater than that of total output in Canada. On the contrary, domestic demand significantly declines following a U.S. monetary policy shock, as shown in panel B. The latter point is confirmed when we employ alternative proxy variables for domestic consumption and investment (here we employ retail sales volume in panel C and industrial production in the construction sector in panel D).

4.3 Summary

Our results presented so far provide several implications for the "dilemma-versus-trilemma" debate. First, we present evidence for the significant transmission of domestic monetary policy shocks into financial and macroeconomic conditions in Canada. Our results confirm that the transmission of domestic monetary policy shocks works through various types of channels, including interest rate, exchange rate, and credit channels, which is in line with the prediction of the trilemma hypothesis.

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outcomes. Kearns and Patel (2016) provide empirical evidence that the trade channel is offset in part by the financial channel in emerging economies but the effect is smaller in advanced economies due to the weak financial channel.

That said, our empirical results find strong and persistent financial and macroeconomic spillovers from the United States to Canada, as argued by Feldkircher and Huber (2016), Georgiadis (2016), and Dedola, Rivolta, and Stracca (2017). Furthermore, the transmission of U.S. monetary policy shocks appears to operate through

Figure 5. Impulse Response of Canada Variables to Monetary Shocks

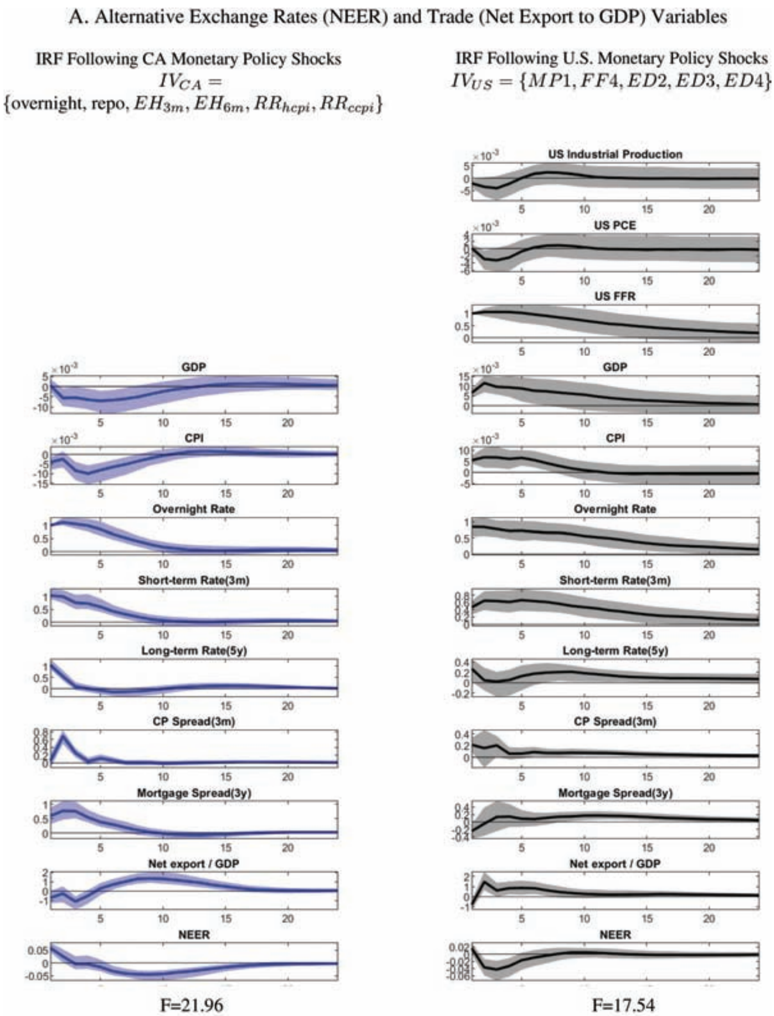

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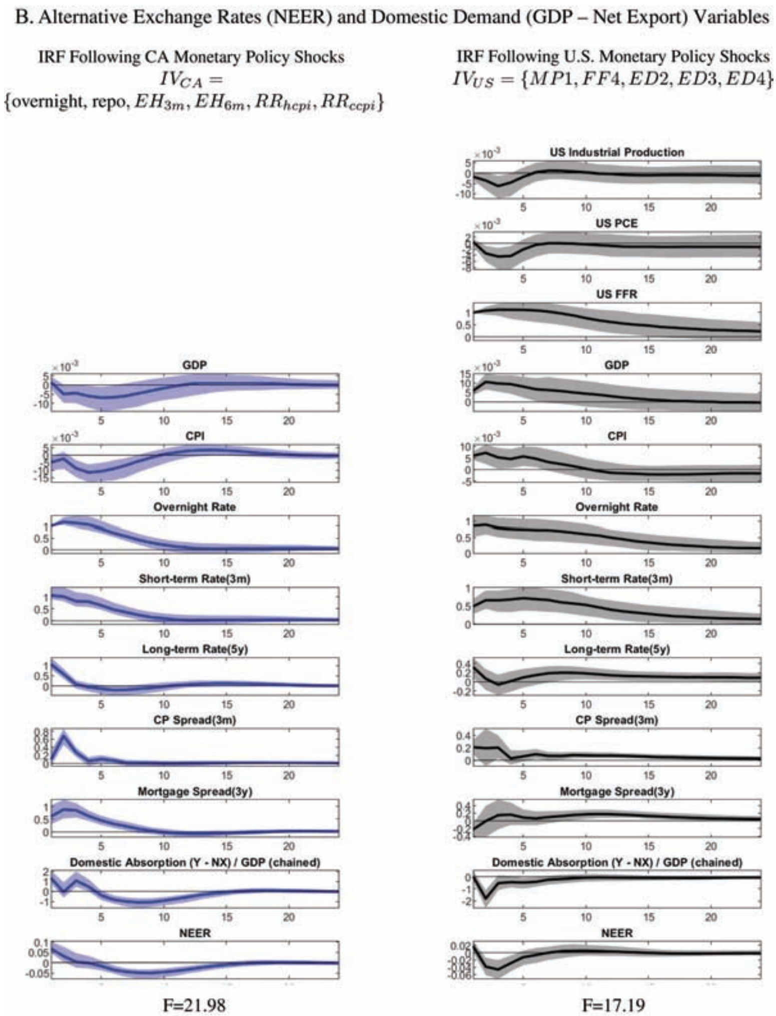
Figure 5. (Continued)



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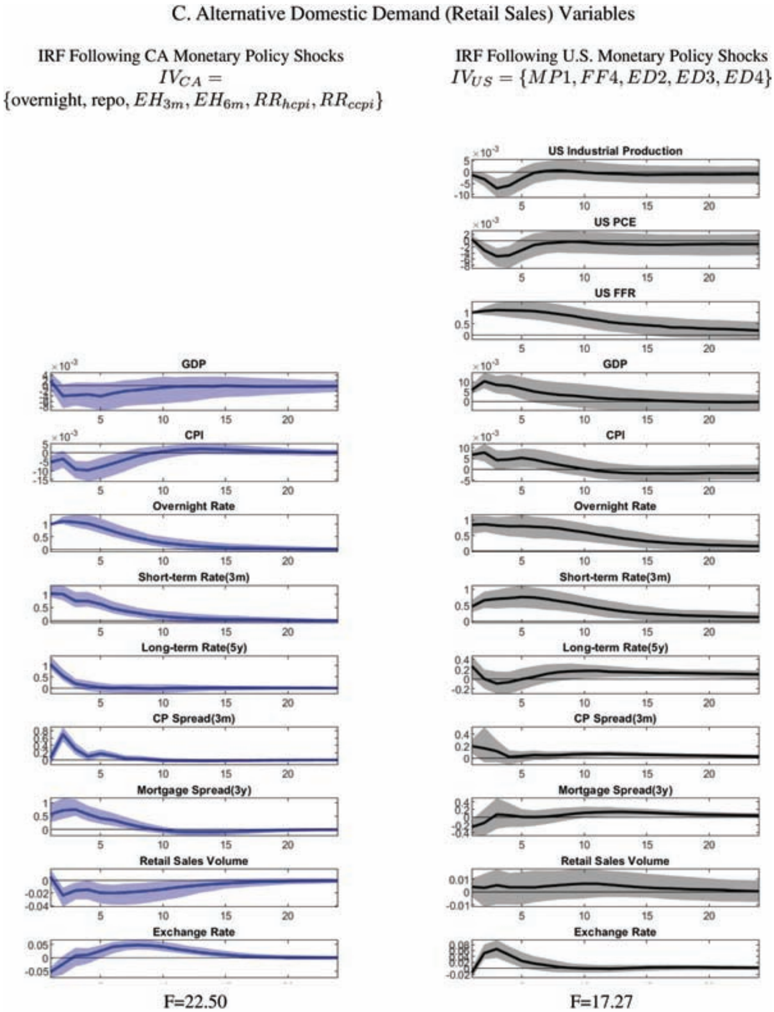
the channels with different natures—financial, aggregate demand, and trade channels—that can have somewhat opposing effects on macroeconomic variables in Canada.

On the one hand, financial transmission of U.S. monetary policy seems apparent. The effect of foreign (U.S.) interest rates on

Figure 5. (Continued)



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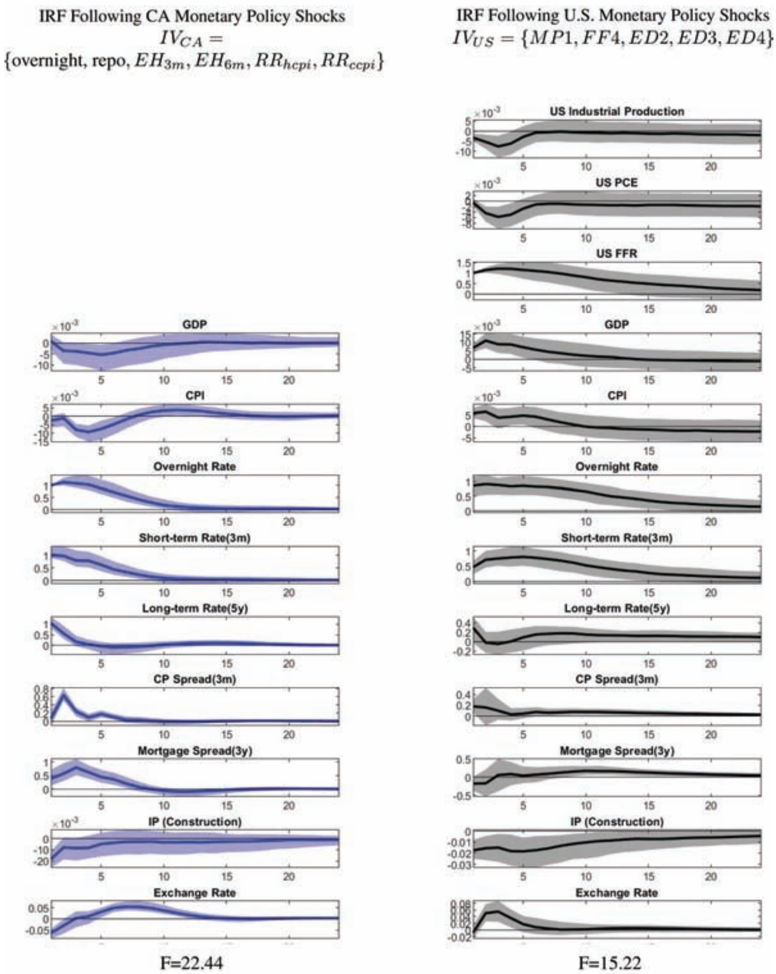
domestic interest rates is the most plausible financial channel as illustrated in Di Giovanni and Shambaugh (2008) and Rey (2015, 2016). More interestingly, Canadian interest rates of all maturities demonstrate a significant and persistent response to surprises in U.S. monetary policy. This international spillover also operates through credit conditions in Canada in both short- and long-term instruments,

Figure 5. (Continued)

D. Alternative Domestic Demand (Retail Sales) Variables



Note: Y-axis indicates percent (or percentage point). X-axis indicates months after shock. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively. Shaded areas are the 16 and 84 percentiles of the empirical distribution based on wild-bootstrapped samples. "F" denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set.

and capital inflows to domestic financial markets (Dahlhaus and Vasishtha 2014). These results collectively indicate that financial conditions in Canada are subject to the impact of monetary policies from the center country, as Rey (2015, 2016) concludes regarding the possible dilemma that central banks in open economies may confront.

On the other hand, our results point to the predictions by Mundell-Fleming's trilemma hypothesis that depreciation of exchange rates leads to the improvement in terms of trade and an increase in net exports. More interestingly, the results suggest that despite the financial tightening and decline in domestic demand, expansion of net exports and currency depreciation boost domestic output and prices.

Admittedly, these results can be mainly specific to the Canadian economy which heavily depends on trades, in particular that of commodities. Alternatively, even within the economy, the consequences of U.S. monetary spillover could have changed over time along with the economic and financial developments.[38] Our findings suggest that more cautions should be exercised to understand the nature of transmission of foreign monetary shocks by considering the relative importance of different transmission channels.

The upshot is that, depending upon the extent to which different types of transmission channels operate in the open economy, the macroeconomic consequences of foreign monetary policy spillovers could also be quite different. This also highlights the crucial role of exchange rate in small open economies, in a financially globalized world: it can be either a shock amplifier or a shock absorber.

5. Additional Econometric Considerations and Robustness Checks

This section summarizes the results of several sensitivity and robustness checks to verify the validity of our empirical results. The exercises encompass a battery of checks: estimations with (i) pure

[38]The findings in some earlier studies suggest that the macroeconomic consequence of international spillover effects could be highly dependent upon country characteristics, as well as the sample periods that govern the degree of various channels of transmission.

monetary shocks that are orthogonal to central bank information shocks, (ii) a counterfactual closed-economy model, (iii) alternative sets of instrumental variables, (iv) the Canadian monetary policy shocks following Kearns, Schrimpf, and Xia (2020), (v) different sample periods, and (vi) alternative endogenous variables, including the one which employs the VIX as an endogenous variable instead of an external control variable. We detail each of these robustness checks below.

5.1 U.S. Pure Monetary Shocks vs. Fed Information Shocks

One important finding in the recent literature is that U.S. monetary policy surprises are driven not only by shifts in the monetary policy stance but also by central-bank-specific information on the economic outlook. For instance, Jarociński and Karadi (2020) disentangle the high-frequency-identified U.S. monetary policy shocks into pure monetary policy and the Federal Reserve (Fed) information shocks by introducing sign restrictions in addition to the exogenous instrument.

In panel A of Figure A.3, we estimate the dynamic responses of the variables in Canada to a U.S. pure monetary policy shock following Jarociński and Karadi (2020).[39] Overall, the results are consistent with the baseline results: contractionary U.S. pure monetary shocks are followed by declines of output and prices in the United States, and financial tightening, but improvement in trade balance, and expansionary effects on output and prices in Canada. This may partly reflect the greater explanatory power of the monetary policy shock for U.S. monetary policy instruments than the information shock. The responses of some variables become less statistically significant compared with the baseline results. That said, the dynamic responses of the key variables—exchange rates, short-term interest rates, and macroeconomic variables—are quite significant.

Next, we report the impulse responses to the Fed information shocks in panel B. Following the shocks, as expected, the dynamic

[39] In particular, the federal funds futures surprises which have the opposite sign to the stock price (S&P 500) surprises are used as the proxy for pure monetary policy shocks. This identification approach corresponds to the *poor man's approach* of Ca' Zorzi et al. (2020) and Jarociński and Karadi (2020).

responses of U.S. macro variables are the opposite of the pure monetary shocks: expansions in output and price levels in the United States. This is because the decision of the Fed reflects expectations on higher future economic growth (and inflation) in the U.S. economy. The effects of the shocks on financial conditions in Canada are contractionary while those on macroeconomic variables are expansionary; these effects are virtually the same as those of the pure monetary shocks.

In sum, this exercise suggests that the instrumental variables employed in the baseline estimation mainly reflect monetary policy shocks, although not wholly excluding the impacts of the central bank information shock.

5.2 A Counterfactual Analysis: Model without U.S. Variables

What are the implications of correlated monetary policy shocks, if any, between the two economies? What if we do not consider the impact of monetary shocks in the center country in quantifying the impact of domestic monetary policy shocks? To shed some more light on this issue, this subsection considers an alternative model that omits the impacts of U.S. variables on the Canadian economy. Put differently, we counterfactually consider only domestic (Canadian) variables in the SVAR system.

For easier comparison, IRFs based on the closed model are plotted along with our benchmark (open-economy model) results in Figure A.4. The solid lines indicate the benchmark IRFs, and broken lines are IRFs based on the alternative specification. The comparison of the IRFs suggests that the impacts of domestic monetary policy shocks on overnight and short-term interest rates in Canada are estimated to be significantly greater throughout the forecasting horizon when we choose the alternative closed-economy model. On the contrary, IRFs of capital inflows and exchange rates become quite muted when considering the alternative model. The response of Canadian output and prices on domestic monetary shocks is also weaker until around a year after the shock with the alternative model.

A careful examination of our benchmark results (as shown in Figure 3) suggests that the IRF of a variable based on the alternative model can be approximated by a linear combination of the two types of IRFs of the variable following domestic and U.S. monetary policy shocks, respectively, based on the benchmark model. That is, the IRFs based on the alternative model could be overestimated (or underestimated) partly because the model fails to control the impacts of U.S. monetary policy shocks. For instance, in Figure 3, we find that the responses of overnight and short-term interest rates are stronger and more persistent following U.S. monetary policy shocks than those following domestic monetary policy shocks. Meanwhile, the responses of exchange rates and capital inflows seem to be in opposite directions following the two types of policy shocks, a result which is quite intuitive. The responses of macroeconomic variables (output and prices) are also somewhat heterogeneous following the two monetary policy shocks. Finally, the responses of other variables (including long-term interest rates and credit spreads) are comparable to each other across the two monetary shocks.

As explained in Section 3 and Appendix A.2, we follow Mertens and Ravn (2013) and identify domestic (Canadian) monetary policy shocks that are orthogonal to the identified U.S. monetary policy shocks. More specifically, the correlated parts of domestic monetary policy shocks with U.S. monetary shocks are regarded as stemming from U.S. policy implementations. Thus, the estimated responses of the variables based on the alternative model are likely to include the (omitted) impacts of U.S. monetary policy implementations, which are correlated with domestic policy actions. This counterfactual analysis thus implies that an omission of the variable of U.S. monetary policy instrument can significantly overestimate or underestimate the impact of domestic monetary policy transmission.

5.3 Alternative Combinations of Instrumental Variables

In Section 4, we reported the impulse response of variables using each of the three types of instrumental variables proposed. We then estimated the baseline model using the three types of instrumental variables {IV1, IV2, IV3} altogether in order to maximize the explanatory power of the instrumental variables. To test the sensitivity of the results to the selection of instrumental variables, we estimate the VAR model using alternative sets of instrumental variables: {IV1 and IV2}, {IV1 and IV3}, and {IV2 and IV3}.[40]

The corresponding IRFs are given in panels A, B, and C of Figure A.5 in Appendix A.1, and can be compared with the baseline results in Figure 3. As is evident, the impulse response of variables to domestic and U.S. monetary policy shocks is overall consistent across the different sets of instrumental variables.[41]

5.4 Alternative Instrumental Variables for IV1

We test the robustness of our benchmark results by using alternative sets of instrumental variables (IV1) in the literature. Specifically, a measure of the shifts in overnight index swap (OIS) rate is tested as an alternative market-based instrument for monetary surprise in Canada. The OIS can be used to hedge the exposure to the volatility of short-term rates or to take a speculative position on the future movement of the policy rates. Due to these functions of OIS, its change surrounding the policy decisions can be regarded as the shifts of market expectations for the future path of policy rates (Jarociński and Karadi 2020; Kearns, Schrimpf, and Xia 2020). As shown in panel D in Figure A.5, our overall results do not change much despite the wider confidence intervals with the use of these instruments that mostly reflect the shorter sample periods (June 2002–December 2015).

5.5 Alternative Sample Periods

The United States has implemented unconventional monetary policies since the onset of the global financial crisis in 2008–09. In order to check the robustness of our baseline results against the policy regime changes, we compare the estimates for different sample periods with the following two approaches. First, we test the robustness of the results using the pre-crisis (January 2000–August 2008) and

[40] Moreover, empirical results based on other types of instrumental variables that are not chosen as a benchmark are all available upon request.

[41] In addition, we do find significant results for any variables when these instrumental variables are combined with other types of instrumental variables we propose.

post-crisis (January 2010–December 2017) sample period. As displayed in panels A and B of Figure A.6, the overall results, with a few exceptions, are consistent with the full sample results in Figure 3, although the results using the pre-crisis and post-crisis sample are often less significant, as reflected in the lower F-statistics of the instrumental variables.

Second, we examine the time-varying impulse response functions of the endogenous variables by estimating a rolling-sample SVAR model. Specifically, moving windows of eight years (96 months) are chosen as subsample periods, with the first window ending in January 2008. Panel B of Figure A.6 summarizes the IRFs based on the rolling-window sample periods. The results suggest that although the magnitude of responses varies across subsample periods, they are largely consistent with the results based on the full-sample period.[42]

5.6 Alternative Endogenous Variables

In Figure A.7, we test alternative data for endogenous variables to verify the robustness of our baseline results. In panel A of the figure, we test the alternative variable of monthly real GDP with 2012 as the base year (instead of chained monthly GDP). As is evident, the results are not sensitive to the type of GDP series.

We also employ U.S. shadow federal funds rates (as in Wu and Xia 2016) or U.S. government bond yields with the maturity of one year (as in Gertler and Karadi 2015) in place of effective federal funds rates considering the fact that since the global financial crisis, the variations in overnight rates are limited. The results are summarized in panels B and C of Figure A.7, respectively. Again, the overall empirical results do not change much, although the explanatory power of the instruments becomes somewhat weaker with both of the alternative policy interest rates.

Finally, we include the VIX to control external factors that can simultaneously affect both the United States and Canada. A group of recent studies, including Rey (2015, 2016), suggests that monetary policy shocks in the center country have a significant impact

[42]Notably, IRFs for some variables, including macro variables, exhibit more amplified responses after 2016. This may partly reflect the policy normalization in both economies.

on global financial market sentiments or, more generally, the global financial cycle, and the changes in the global financial markets are transmitted through financial and macroeconomic conditions in other open economies. Considering these findings, we embody the VIX as an endogenous variable here, and examine the response of the variable as well as the overall results for the other variables. Figure A.8 reports that the VIX increases within a couple of months after the contractionary U.S. monetary policy shock, consistent with the findings in earlier studies. The response of other variables in Canada does not show any notable differences.

6. Conclusion

Mundell-Fleming's trilemma has been a central building block in conventional international macroeconomics. However, as global financial markets are increasingly integrated and global factors become crucial drivers of the developments in domestic financial markets, there are extensive debates on the effectiveness of domestic monetary policy and the roles played by exchange rates in small open economies. In this context, this paper investigates the channels of monetary policy transmission in a small open economy within and across borders, by estimating an open-economy SVAR model with novel external instruments.

Our empirical findings are summarized as follows. On the one hand, the transmission of domestic monetary policy shocks operates through a variety of channels. First, both short- and long-term rates react significantly to domestic monetary policy shocks, confirming the role of the conventional interest rate channel. Second, foreign exchange rates in this process are seen to respond significantly to monetary policy shocks, as the overshooting hypothesis by Dornbusch (1976) predicts. Contrary to the findings in earlier studies that report counterevidence for the overshooting hypothesis, we find that an increase in local policy rates causes the nominal exchange rate to quickly appreciate and then depreciate gradually. Third, contractionary domestic monetary policy shocks generate an increase in credit spreads in Canada. Macroeconomic conditions (output and price levels) respond significantly to monetary policy shocks, as New Keynesian theory predicts, reflecting the pass-through of monetary policy shocks into financial and credit markets.

On the flip side, international spillovers of monetary policy shocks also play an important role in financial and credit conditions in Canada. Following a contractionary U.S. monetary policy shock, market interest rates in Canada, in both short- and long-term maturities, are significantly increased by the impact and persist for a prolonged period. More interestingly, overnight rates in Canada, which are used as a monetary policy tool, also respond to U.S. monetary policy shocks. Following a contractionary U.S. monetary policy shock, credit spreads increase substantially along with an immediate outflow of international capital investments. This is consistent with the predictions of the credit and risk-taking channels of monetary policy transmission from international perspectives (Rey 2015, 2016; Hofmann, Shim, and Shin 2016).

Finally, the response of macroeconomic variables is divergent across the two types of monetary policy shocks; U.S. monetary tightening has expansionary and inflationary consequences on the variables in Canada unlike domestic monetary shocks. This may partly reflect the expenditure-switching effects of contractionary U.S. monetary shocks, as well as pass-through of currency depreciation to domestic prices, offsetting the negative impacts of tightened financial conditions on aggregate demand in Canada.

Our empirical results indicate that as the global financial markets become more integrated, the financial and credit conditions in a small open economy can be driven not only by domestic monetary policy but also by global factors such as U.S. monetary policy. In Canada, the impacts of international monetary spillovers could be buffered by the flexible reactions of exchange rates. Admittedly, however, this result can be mainly specific to the Canadian economy. For the countries where flexible exchange rates cannot completely insulate their economies from external shocks, monetary policy decisions made by focusing only on domestic conditions can result in policy errors in achieving macroeconomic stability. The consequences of external shocks, including monetary policy shocks in the center country, should be carefully considered when implementing monetary policies as well as other policy tools.[43]

[43]From this perspective, the International Monetary Fund proposed recently the integrated policy framework which guides how a mix of additional policy tools—macroprudential measures, FX market interventions, and capital flow managements—should be deployed to achieve policy objectives.

Appendix

A.1 Additional Figures

Figure A.1. Policy Rate and Short-Term Spot Rates

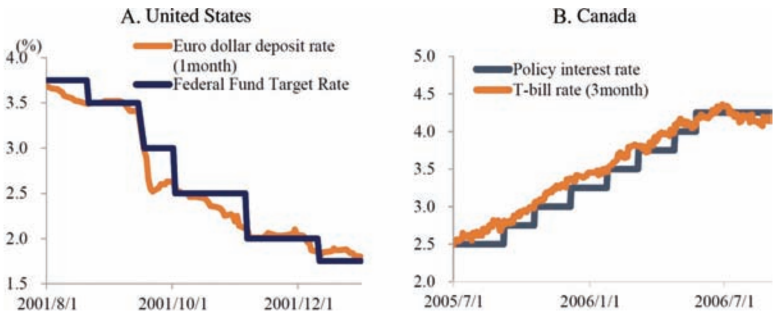
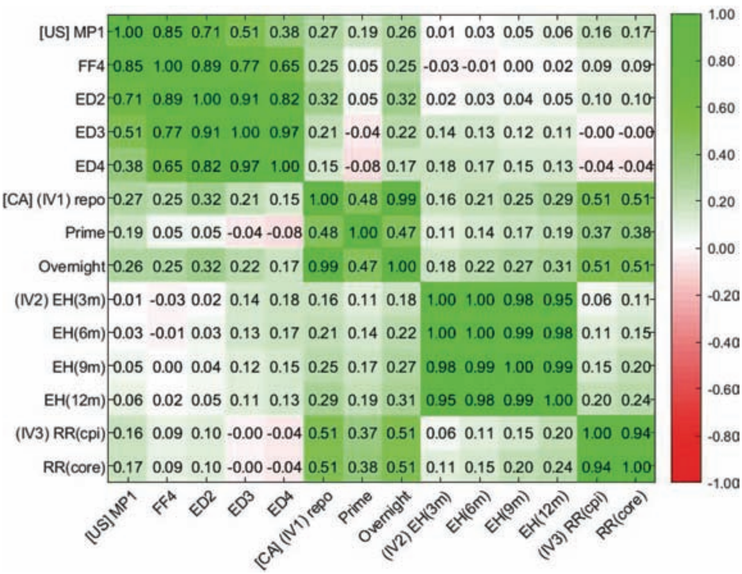
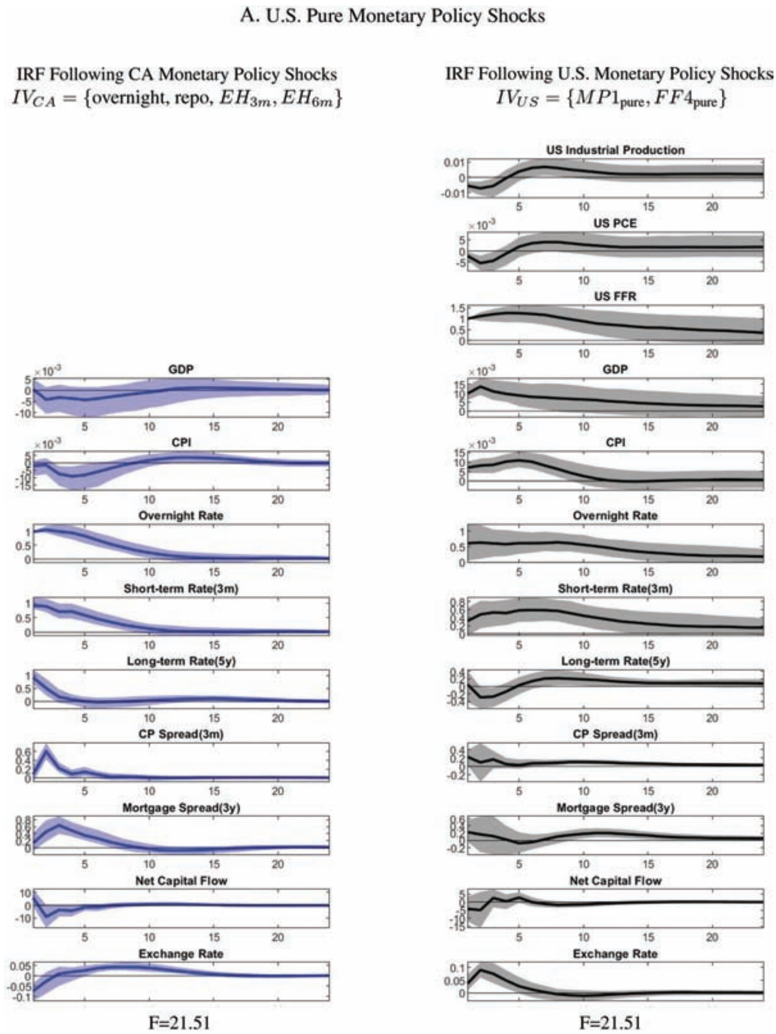


Figure A.2. Correlation Heatmap for the Instrumental Variables



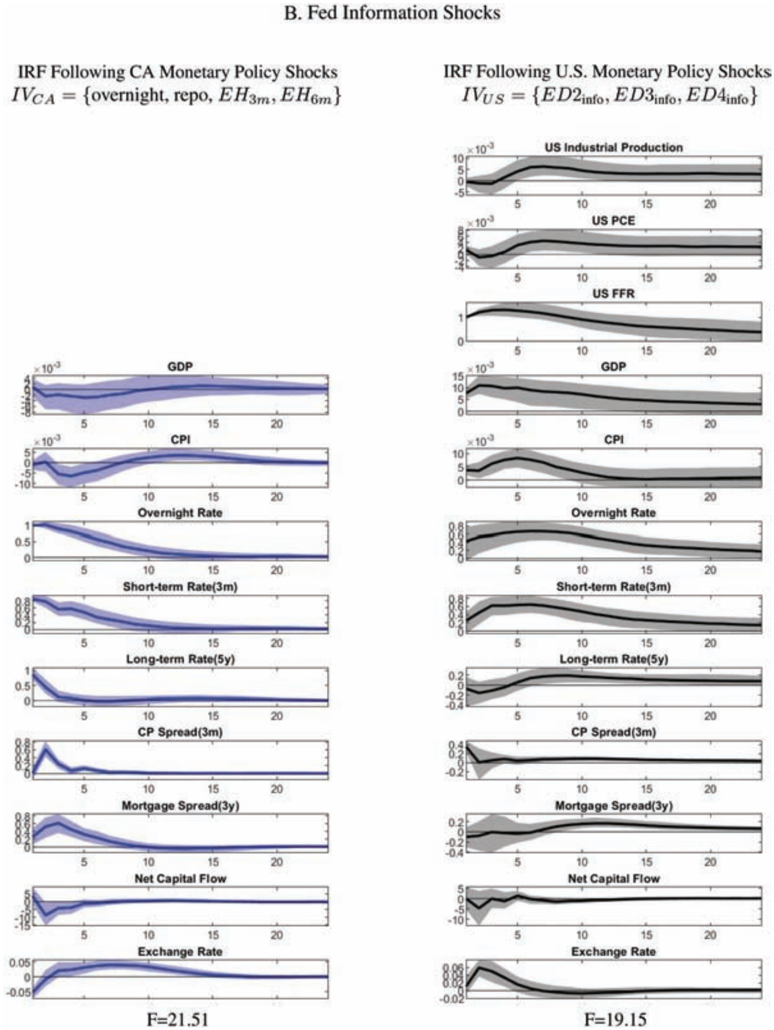
Note: Numbers in cells are correlations among the instrumental variables. Thicker colors indicate higher positive (green) or negative correlation (red) than others.

Figure A.3. IRFs with Different U.S. IV Sets



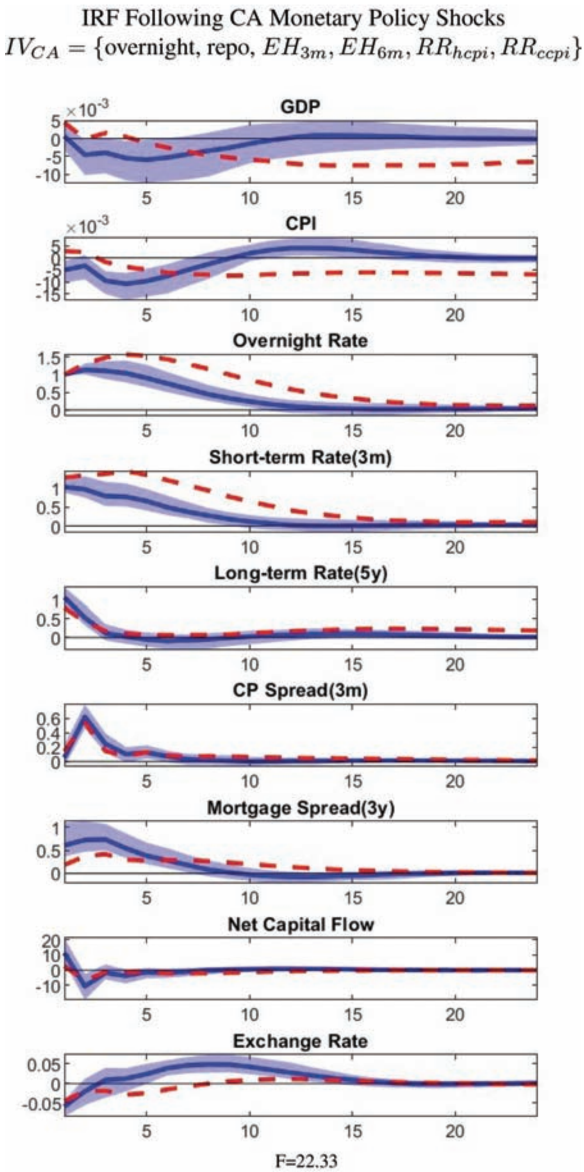
(continued)

Figure A.3. (Continued)



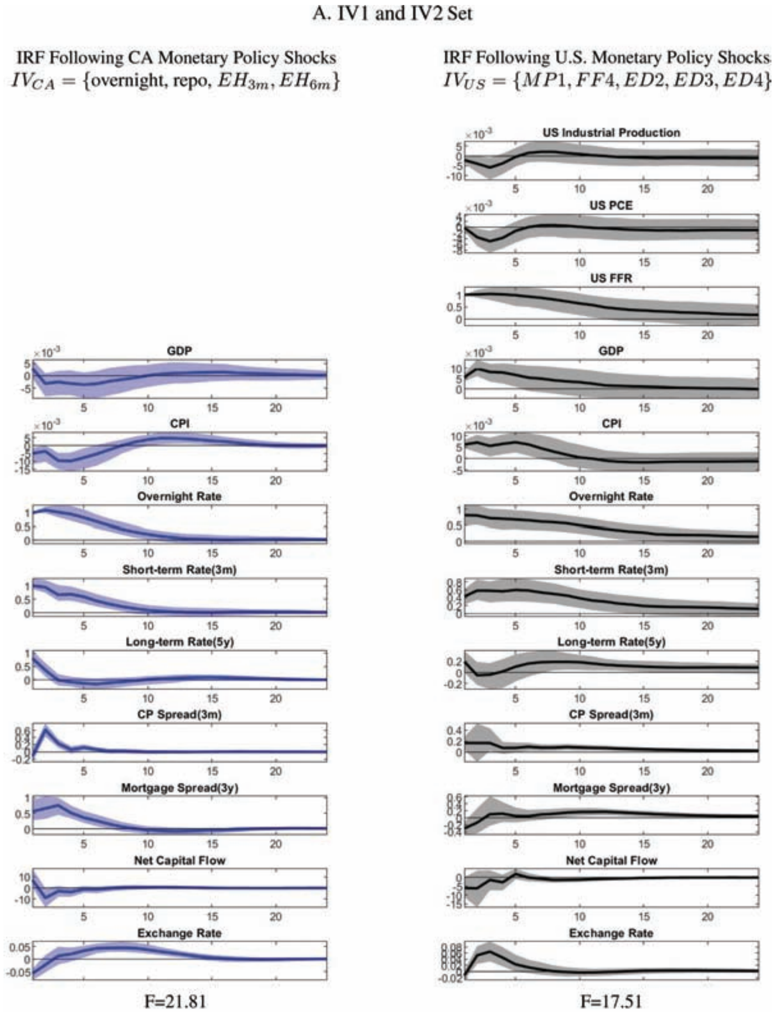
Note: Y-axis indicates percent (or percentage point for interest rates, or mil. CAD for capital flow). X-axis indicates months after shock. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively. Shaded areas are the 16 and 84 percentiles of the empirical distribution based on wild-bootstrapped samples. "F" denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set.

Figure A.4. IRFs with Alternative Specification: Model without U.S. Variable



Note: Y-axis indicates percent (or percentage point for interest rates, or mil. CAD for capital flow). X-axis indicates months after shock. Based on contractionary (+100 bps) Canadian monetary policy shocks. Solid lines (confidence intervals in shaded areas) are based on benchmark model (with U.S. monetary policy instruments) and broken lines are based on alternative model without U.S. monetary policy instruments. “F” denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set (altenative model).

Figure A.5. IRFs with Different Canadian IV Sets



(*continued*)

Figure A.5. (Continued)

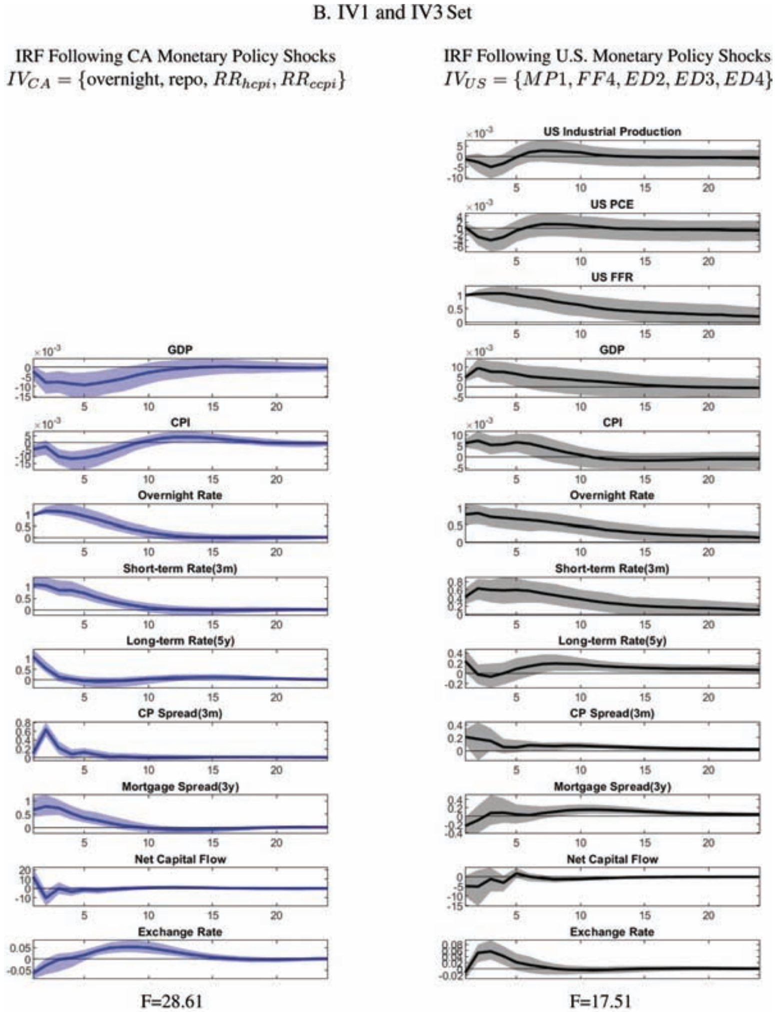


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Figure A.5. (Continued)

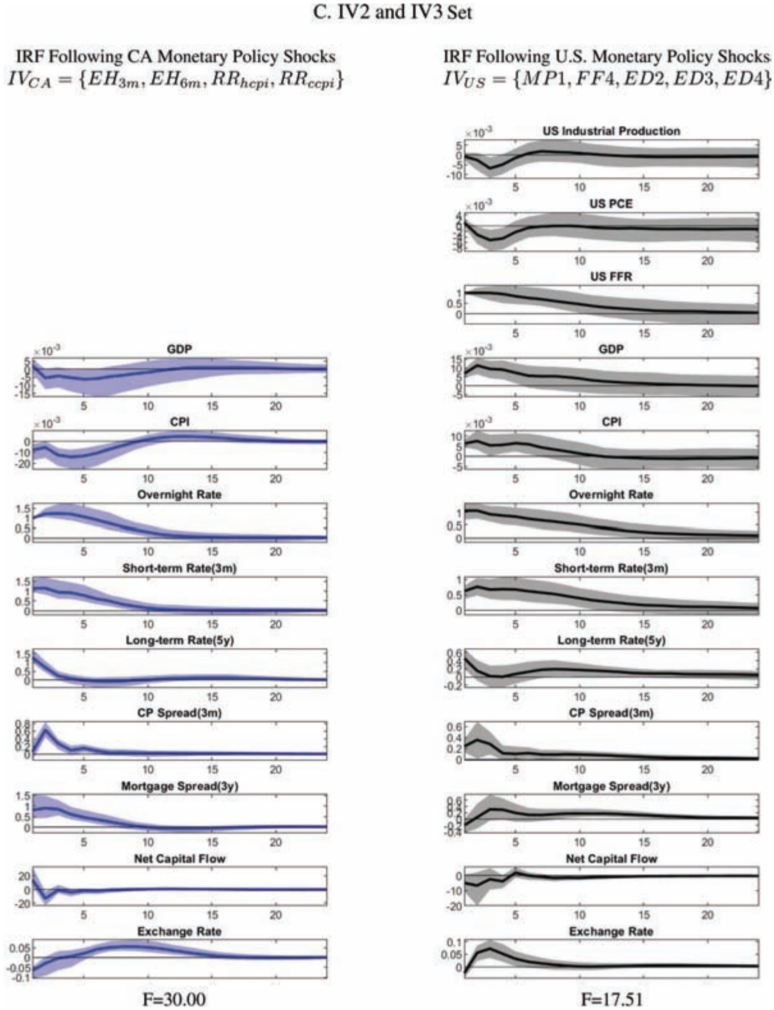

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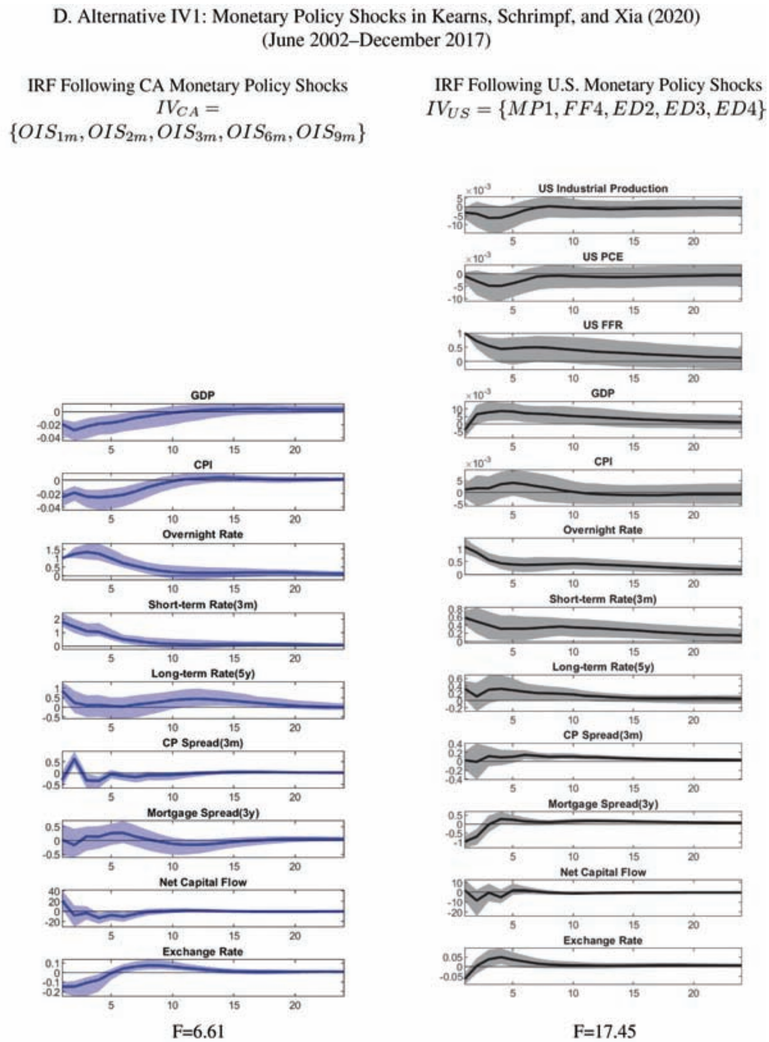
Figure A.5. (Continued)



Note: Y-axis indicates percent (or percentage point for interest rates, or mil. CAD for capital flow). X-axis indicates months after shock. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively. Shaded areas are the 16 and 84 percentiles of the empirical distribution based on wild-bootstrapped samples. "F" denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set.

Figure A.6. IRFs with Alternative Sample Periods

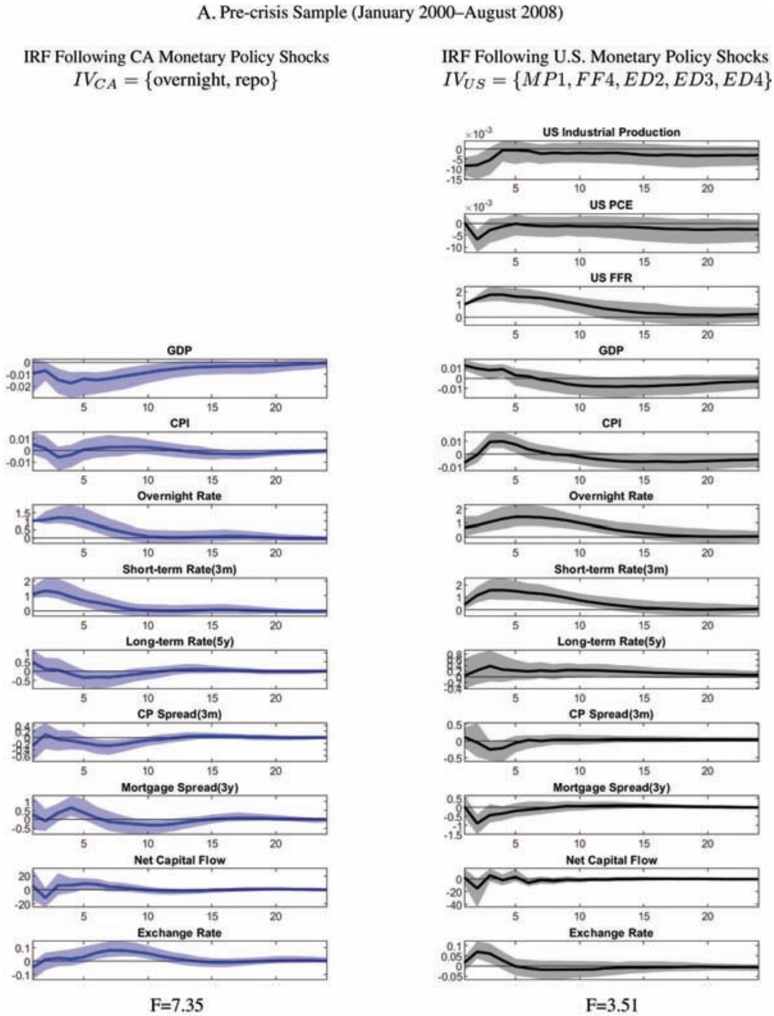


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Figure A.6. (Continued)

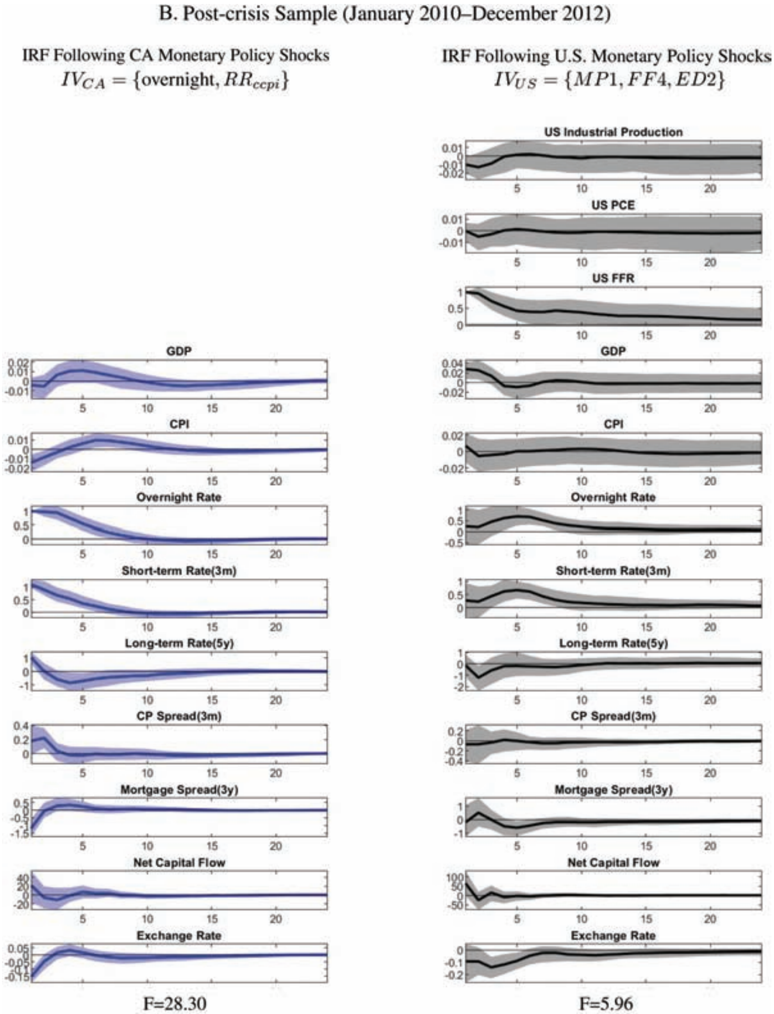

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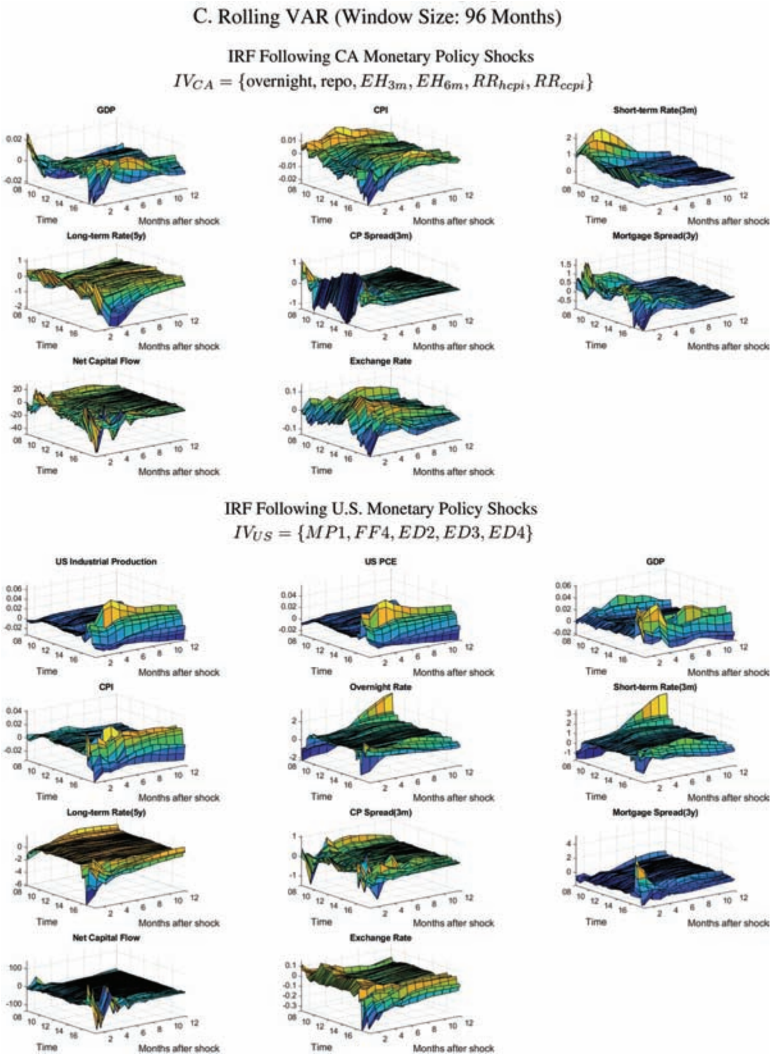
Figure A.6. (Continued)



Note: Y-axis indicates percent (or percentage point for interest rates, or mil. CAD for capital flow). "Time" indicates the sample periods, while "Months after shock" indicates the forecasting horizons after monetary policy shocks. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively.

Figure A.7. IRFs with Alternative Endogenous Variables

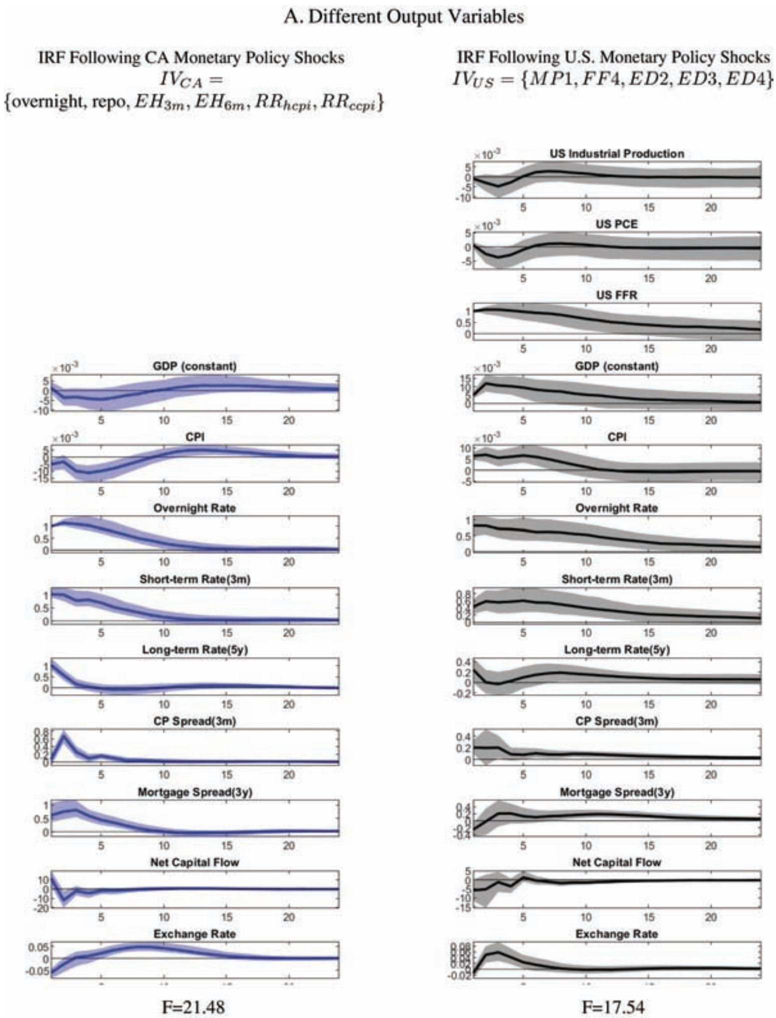


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Figure A.7. (Continued)

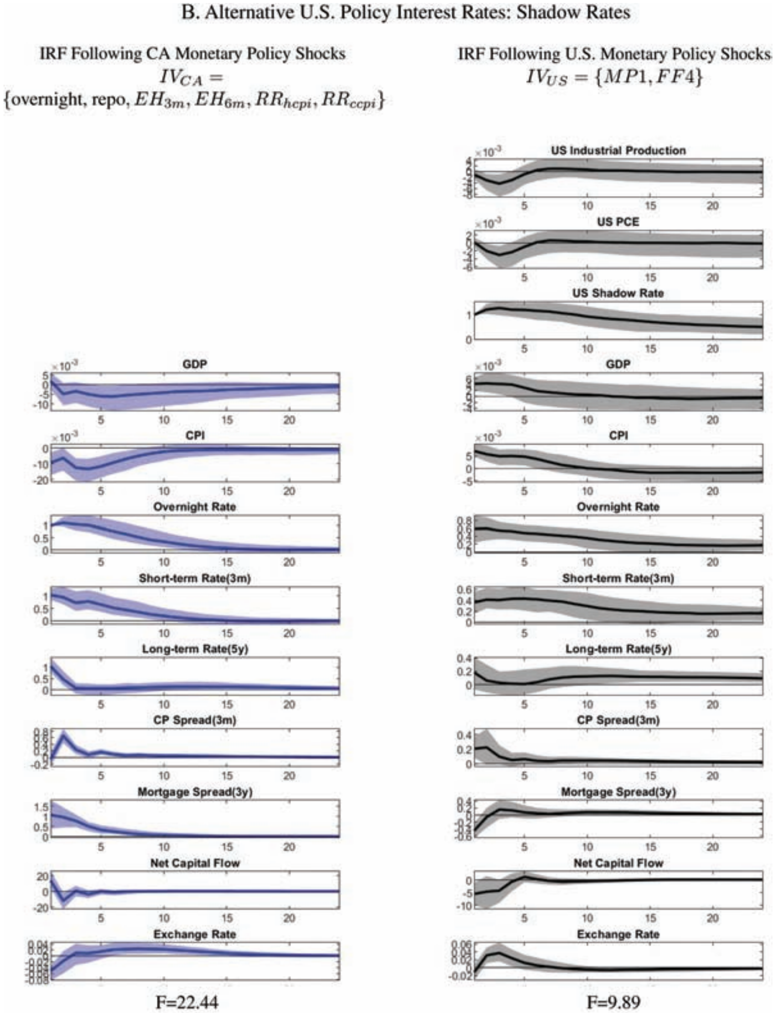

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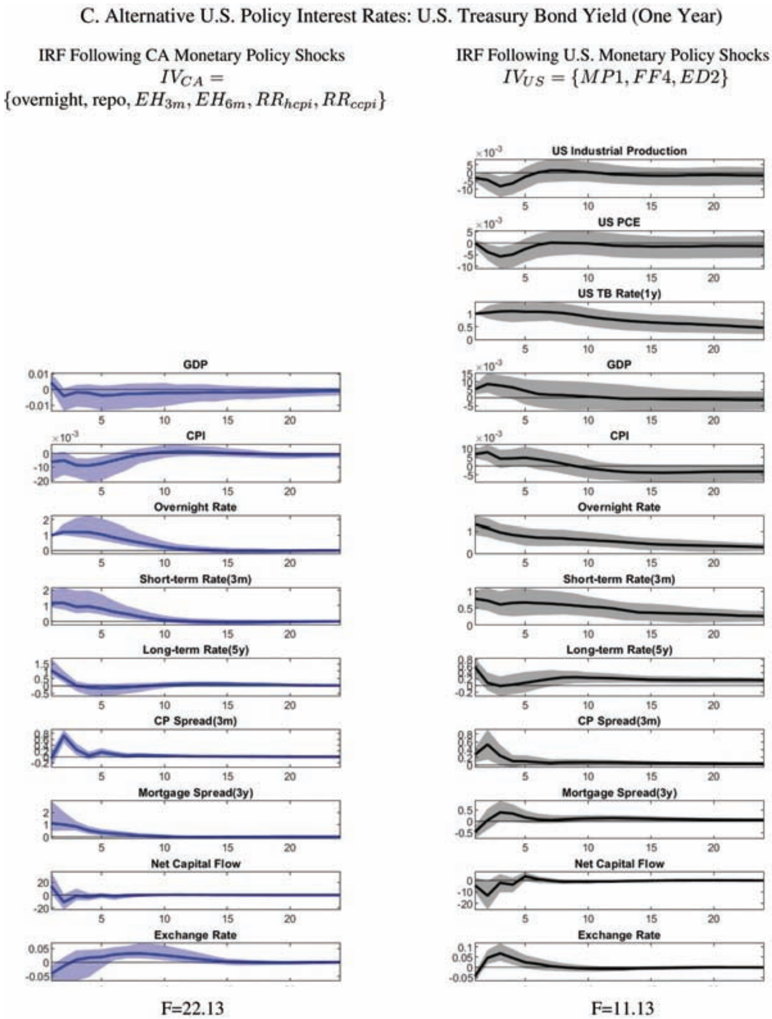
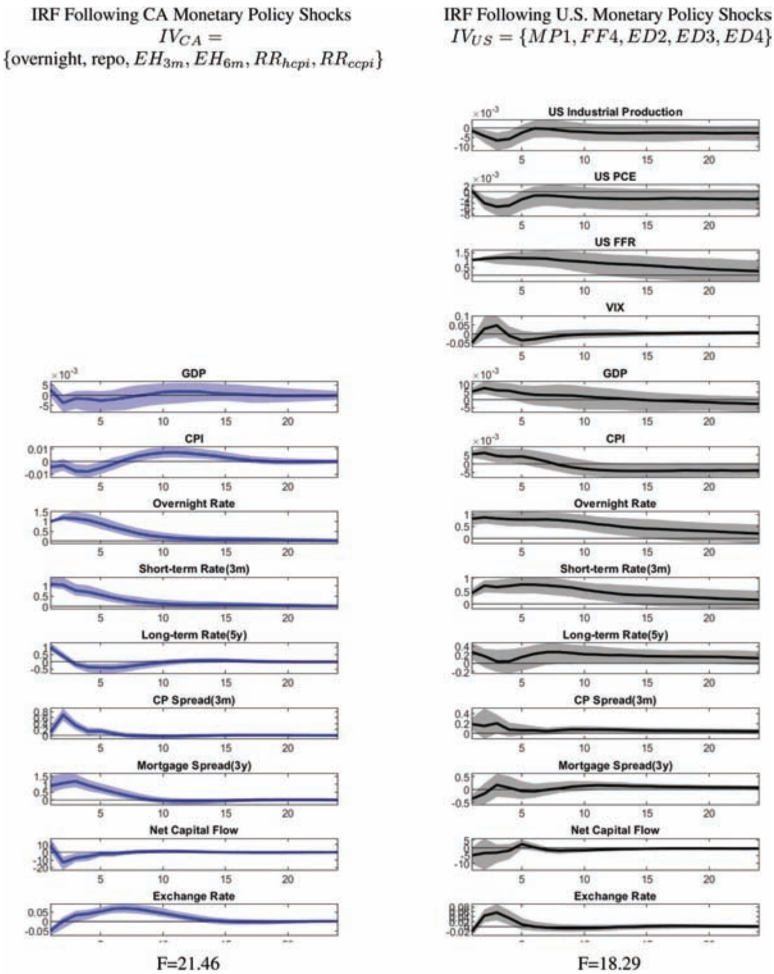
Figure A.7. (Continued)



Note: Y-axis indicates percent (or percentage point for interest rates, or mil. CAD for capital flow). X-axis indicates months after shock. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively. Shaded areas are the 16 and 84 percentiles of the empirical distribution based on wild-bootstrapped samples. “F” denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set.

Figure A.8. VIX Included as an Endogenous Variable



Note: Y-axis indicates percent (or percentage point for interest rates, or mil. CAD for capital flow). X-axis indicates months after shock. Based on contractionary (+100 bps) U.S. and Canadian monetary policy shocks, respectively. Shaded areas are the 16 and 84 percentiles of the empirical distribution based on wild-bootstrapped samples. "F" denotes F-statistics for the first-stage regression of residuals of policy indicators on each IV set.

A.2 Technical Details on External Instrument Identification Scheme

The relationship between residuals of reduced-form VAR (e_t) and structural shocks (ε_t) in Equation (7) can be rearranged as (A.1):

$$\begin{bmatrix} e_t^p \\ e_t^q \end{bmatrix} = \begin{bmatrix} s_{11} & s_{12} \\ s_{21} & s_{22} \end{bmatrix} \begin{bmatrix} \varepsilon_t^p \\ \varepsilon_t^q \end{bmatrix} = \begin{bmatrix} s_{11}\varepsilon_t^p + s_{12}\varepsilon_t^q \\ s_{21}\varepsilon_t^p + s_{22}\varepsilon_t^q \end{bmatrix}, \tag{A.1}$$

where s_{11} represents the response of the residuals of the monetary policy instrument to its own shock and s_{21} represents the responses of residual series of the other variables to the structural monetary policy shock. Since we are interested in how variables respond to monetary policy shocks, s_{11} and s_{21} are the only two parts of the impact matrix (S) to be identified.

Next, VAR residuals e_t^p and e_t^q can be expressed by the other reduced-form residuals and structural shocks ε_t^p or ε_t^q because those are composites of structural shocks as in (A.2) and (A.3):

$$e_t^p = \eta e_t^q + C_1\varepsilon_t^p \tag{A.2}$$

$$e_t^q = \theta e_t^p + C_2\varepsilon_t^q, \tag{A.3}$$

where $\eta = s_{12}s_{22}^{-1}$, $\theta = s_{21}s_{11}^{-1}$, $C_1 = s_{11} - s_{12}s_{22}^{-1}s_{21}$, and $C_2 = s_{22} - s_{21}s_{11}^{-1}s_{12}$. In particular, the 2×2 matrix C_1 represents variance–covariance between two structural monetary policy shocks.[44] It has the following relationship with s_{11} and s_{21}:

$$\begin{bmatrix} s_{11} \\ s_{21} \end{bmatrix} = \begin{bmatrix} \left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)^{-1} \\ s_{21}s_{11}^{-1}\left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)^{-1} \end{bmatrix} C_1 \tag{A.4}$$

$$C_1C_1' = \left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right) s_{11}s_{11}' \left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)'. \tag{A.5}$$

Thus, obtaining s_{11} and s_{21} requires identification of two parts: One is $s_{21}s_{11}^{-1}(=\theta)$, which can be estimated by two-stage least

[44] C_1 can be rearranged as $C_1 = s_{11} - s_{12}s_{22}^{-1}s_{21} = \left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)s_{11}$ and thus $s_{11}C_1^{-1} = \left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)^{-1}$. Similarly, C_2 can be expressed in terms of partitions of S matrix as the following form: $s_{21}C_1^{-1} = s_{21}s_{11}^{-1}s_{11}C_1^{-1} = s_{21}s_{11}^{-1}\left(I - s_{12}s_{22}^{-1}s_{21}s_{11}^{-1}\right)^{-1}$.

squares (2SLS) estimation, and the others are $s_{11}s'_{11}$ and $s_{12}s_{22}^{-1}$, which can be calculated by restrictions from the covariance matrix.[45]

Restriction from 2SLS Estimation: $\boldsymbol{s_{21}s_{11}^{-1}(=\theta)}$. Consider first the regression of Equation (A.2). Since the reduced-form residual for monetary policy instrument ($e_t^p(= s_{11}\varepsilon_t^p + s_{12}\varepsilon_t^q)$) is correlated with $C_2\varepsilon_t^q$, denoting it as u_t hereafter, we can obtain consistent estimates of θ of regression e^q on e^p from 2SLS, employing appropriate instrumental variables that satisfy the following moment conditions, as in (A.6) and (A.7):

$$E\left[Z_t u_t\right] = 0 \qquad \text{or} \qquad E\left[Z_t \varepsilon_t^q\right] = 0 \tag{A.6}$$

$$E\left[Z_t e_t^p\right] = \pi(\pi \neq 0) \qquad \text{or} \qquad E\left[Z_t \varepsilon_t^p\right] = \phi(\phi \neq 0). \tag{A.7}$$

Restriction from Covariance Matrix: $\boldsymbol{s_{11}s'_{11}}$ and $\boldsymbol{s_{12}s_{22}^{-1}}$. In addition to the restrictions derived from IV estimation, identification of s_{11} and s_{21} requires the additional restrictions from the covariance matrix. Consider the following reduced-form variance–covariance and its partitioning in (A.8):

$$\Sigma = E[SS'] \Rightarrow \begin{bmatrix} \Sigma_{11} & \Sigma_{12} \\ \Sigma_{21} & \Sigma_{22} \end{bmatrix} = \begin{bmatrix} s_{11}s'_{11} + s_{12}s'_{12} & s_{11}s'_{21} + s_{12}s'_{22} \\ s_{21}s'_{11} + s_{22}s'_{12} & s_{21}s'_{21} + s_{22}s'_{22} \end{bmatrix}. \tag{A.8}$$

Then, $s_{11}s'_{11}$, $s_{12}s_{22}^{-1}$ is obtained by the following closed-form solution in (A.9) and (A.10):

$$s_{11}s'_{11} = \Sigma_{11} - s_{12}s'_{12} \tag{A.9}$$

$$s_{12}s_{22}^{-1} = \left(s_{12}s'_{12}\theta' + (\Sigma_{21} - \theta\Sigma_{11})'\right)\left(s_{22}s'_{22}\right)^{-1}, \tag{A.10}$$

where $s_{12}s'_{12} = (\Sigma_{21} - \theta\Sigma_{11})' Q^{-1} (\Sigma_{21} - \theta\Sigma_{11})$, $s_{22}s'_{22} = \Sigma_{22} + s_{21}s_{11}^{-1}\left(s_{12}s'_{12} - \Sigma_{11}\right)\left(s_{21}s_{11}^{-1}\right)'$, and $Q = \Sigma_{22} - (\Sigma_{21}\theta' + \theta\Sigma_{21}) + \theta\Sigma_{11}\theta'$.[46]

[45]The different types of instruments may deliver quite distinct information as discussed in Section 3. Taking this into account, we identify each shock using multiple proxy variables in the 2SLS estimation.

[46]Consider first the fact that $\Sigma_{21} - \theta\Sigma_{11} = C_2 s'_{12}$ because $\Sigma_{21} - \theta\Sigma_{11} = s_{21}s'_{11} + s_{22}s'_{12} - s_{21}s_{11}^{-1}(s_{11}s'_{11} + s_{12}s'_{12}) = s_{22}s'_{12} - s_{21}s_{11}^{-1}s_{12}s'_{12} = \left(s_{22} - s_{11}^{-1}s_{12}\right)s'_{12}$. The derivation of $s_{12}s_{22}^{-1}$ is straightforward, noticing that

These restrictions from 2SLS and VAR residual covariance allow for the identification of C_1C_1' and the covariance of $C_1\varepsilon_t^p$. If structural shocks to domestic monetary policy are uncorrelated with foreign monetary policy shocks and vice versa, C_1 is a diagonal and can be directly identified up to a sign convention from Equation (A.6).[47] However, if we cannot impose zero cross-correlations between structural shocks, we must make an arbitrary assumption regarding how domestic monetary policy shocks respond contemporaneously to unanticipated movements in foreign monetary policy instruments and vice versa in order to disentangle the causal effects of shocks on both monetary policy shocks. To the extent that the model considers two countries, the United States and a small open economy, Cholesky decomposition of C_1C_1', supposing that the foreign monetary policy shock is ordered before the domestic monetary policy shock, permits economically meaningful results in this analysis. Finally, by plugging the identified C_1 back into (A.5), s_{11} and s_{21} are uniquely pinned down.

A.3 Technical Details on Instrumental Variable II: Monetary Policy Shock Implied in Term Structure Model

The affine model we consider is described below. The prices of zero-coupon bonds are derived from the pricing kernel as in (A.11):

$$P_t{}^{\tau} = E_t\left[m_{t+1}P_{t+1}{}^{\tau-1}\right], \tag{A.11}$$

where $P_t{}^{\tau}$ is the zero-coupon bond price with a maturity τ, m_{t+1} is a stochastic discount factor. E_t is a conditional expectation on

$s_{12}s_{22}' = s_{12}s_{12}'\theta' + (\Sigma_{21} - \theta\Sigma_{11})'$. $Q = Q'$ because Q is symmetric, and it is same as u_tu_t' or C_2C_2'. Using this fact, $s_{12}s_{12}'$ can be obtained by the following form: $s_{12}s_{12}' = s_{12}C_2'C_2'^{-1}C_2^{-1}C_2s_{12}' = s_{12}C_2'Q^{-1}C_2s_{12}' = (\Sigma_{21} - \theta\Sigma_{11})'\,Q^{-1}\,(\Sigma_{21} - \theta\Sigma_{11})$. And from the covariance matrix, $s_{22}s_{22}' = \Sigma_{22} - s_{21}s_{21}'$, and it can be expressed as the above because $s_{21}s_{21}' = s_{21}\left(s_{11}^{-1}s_{11}s_{11}'s_{11}^{-1\prime}\right)s_{21}' = \left(s_{21}s_{11}^{-1}\right)(\Sigma_{11} - s_{12}s_{12}')\left(s_{11}^{-1\prime}s_{21}'\right)$.

[47] If so, a simpler identification approach, such as Gertler and Karadi (2015) employ, can be directly applied to identify s_{11} and $s_{21}C_1$.

the information set up to time t. We specify the discount factor as (A.12):

$$m_{t+1} = exp\left(-r_t - \frac{1}{2}\Lambda_t'\Omega\Lambda_t - \Lambda_t'\nu_{t+1}\right) \tag{A.12}$$

with the assumption that the risk-free short-term rate (r_t) and time-varying market prices of risk (Λ_t) are linear functions of factors as in (A.13):

$$r_t = \delta_0 + \delta_1' X_t \qquad \text{and} \qquad \Lambda_t = \lambda_0 + \lambda_1 X_t, \tag{A.13}$$

where δ_0 is a constant term; δ_1 and λ_0 are $N \times 1$ vectors; and λ_1 is a $N \times N$ matrix, respectively. We assume that the transition equation for state variable X_t follows the first-order vector-autoregressive process as in (A.14):

$$X_t = \mu + \Phi X_{t-1} + \nu_t, \tag{A.14}$$

where factor shocks ν_t follows iid normal distribution $N(0, \Omega)$.

Combining Equations (A.11)–(A.14) yields the bond price and yield for maturity τ as the following affine functions of the state variables:

$$P_t{}^{\tau} = exp[A_\tau + B_\tau X_t] \tag{A.15}$$

$$R_t{}^{\tau} \approx -\frac{1}{\tau} log P_t{}^{\tau} = -\frac{1}{\tau}\left(A_\tau + B_\tau X_t\right), \tag{A.16}$$

where A_τ and B_τ are obtained in the recursive equations as in (A.17) and (A.18).

$$A_{\tau+1} = A_\tau + B_\tau'\left(\mu - \Omega\lambda_0\right) + \frac{1}{2}B_\tau'\Omega B_\tau - \delta_0 \tag{A.17}$$

$$B_{\tau+1}' = B_\tau'\left(\Phi - \Omega\lambda_1\right) - \delta_1' \tag{A.18}$$

In estimating coefficients and pricing factors, we follow the approach of Adrian, Crump, and Moench (2013), which comprises three-step regressions. First, a VAR of order one in Equation (A.14) is estimated via ordinary least squares (OLS). This step decomposes pricing factors X_t into a predictable component and estimated factor innovations v_t. Second, we regress excess bond returns on a constant, lagged pricing factors and the pricing factor innovations. By doing so, we have predictive coefficients and factor risk exposures. Third, the market price of risk parameters λ_0 and λ_1 are estimated by cross-sectionally regressing predictive coefficients on factor risk exposures.

By estimating the term structure model with this approach, we can decompose daily bond yields with a variety of maturities into the sum of expected short-term rates and term premiums. Other things being equal, monetary shocks around policy decisions directly cause changes in the future path of the short-term interest rate. Thus, we consider the daily change of short-term expectations around monetary policy announcements as monetary surprises.

A.4 Technical Details on Instrumental Variable III: Residuals from Policy Reaction Functions

We follow Champagne and Sekkel (2018) and take two steps in estimating the Taylor-rule equation. First, using the minutes of monetary policy reports (source: Bank of Canada Monetary Policy Reports), we collect real-time forecasts for output and inflation in Canada. We use both headline and core CPI inflation for the CPI inflation forecast. Second, we regress changes in monetary policy target rates from the previous monetary policy decision meeting to the current meeting (Δr_m) on a set of explanatory variables that purge the intended policy rate. The explanatory variables include (i) levels of policy rates (two weeks prior to the monetary policy meeting, r_{t-14}), (ii) forecasts of real GDP growth ($y^f_{m,j}$) and inflation ($\pi^f_{m,j}$); we here include the one- and two-month-ahead forecasts as well as the forecasts for the contemporaneous period and the forecast made one month before the meeting, (iii) changes of the variables selected in (ii) from the previous period, and (iv) other variables that could potentially reflect economic developments between meetings. The terms in (iii) reflect revisions to the forecasts relative to the previous round of forecasts. The last variable (iv) includes real-time rates of unemployment for the previous three months and the levels and changes of U.S. federal fund rates and the logs of the USD/CAD nominal exchange rate two weeks before the meeting. The estimated regression is summarized as (A.19).

$$\Delta r_m = \alpha + \underbrace{\beta_1 r_{t-14}}_{(i)} + \underbrace{\sum_{j=-1}^{2} \beta_{2,j} y^f_{m,j} + \sum_{j=-1}^{2} \beta_{3,j} \pi^f_{m,j}}_{(ii)}$$
$$+ \underbrace{\sum_{j=-1}^{2} \beta_{4,j} \left(y^f_{m,j} - y^f_{m-1,j} \right) + \sum_{j=-1}^{2} \beta_{5,j} \left(\pi^f_{m,j} - \pi^f_{m-1,j} \right)}_{(iii)}$$
$$+ \underbrace{\gamma Z}_{(iv)} + \varepsilon_m, \tag{A.19}$$

where Δr_m, changes in policy rates, is measured at the frequency of monetary policy meetings, as indicated by the subscript m. The subscript j denotes the quarter of the real-time data or forecast relative to the meeting date. Z includes other control variables.

The regression coefficients for Equation (A.19) are summarized in Table A.1 (using headline CPI inflation forecast in panel A and the core CPI inflation forecast in panel B). Consistent with the findings in Champagne and Sekkel (2018), the results indicate that changes in policy rate are significantly and positively associated with levels or changes in the forecast of inflation and/or output growth. The results also provide evidence that monetary policy decisions in Canada reflect both levels and changes in monetary policy target rates in the United States. A higher real-time unemployment level is associated with a decrease in policy rates with less statistical significance after controlling for GDP growth and inflation forecasts. R^2 of the regressions is over 0.8. This suggests that explanatory variables in the regressions which proxy the intended component of policy changes in Canada explain around 80 percent of variations in monetary policy target rates in Canada over the sample period.

Table A.1. Regression of Monetary Policy Rates on Explanatory Variables

Variable	Coefficient	Std. Error	*t*-statistic	Prob.
A. Regression with Headline CPI Inflation Forecast				
Initial Policy Rates in Canada	–0.1926	0.0259	–7.4286	0.0000
Policy Rates in United States (FFR)				
Level	0.1306	0.0182	7.1896	0.0000
Changes	0.0807	0.0273	2.9603	0.0040
CAD/USD Rates				
Level	–0.3594	0.2242	–1.6036	0.1126
Changes	–0.0062	0.0071	–0.8668	0.3886
Unemployment in Canada				
One Month Before	–0.0368	0.0726	–0.5065	0.6139
Two Months Before	–0.0177	0.0881	–0.2012	0.8410
Three Months Before	0.0557	0.0682	0.8177	0.4159
Forecasted Output Growth				
One Quarter Before	0.0268	0.0240	1.1175	0.2671
Contemporaneous	0.0455	0.0371	1.2253	0.2240
One Quarter Ahead	–0.0787	0.0378	–2.0824	0.0404
Two Quarters Ahead	0.0533	0.0327	1.6314	0.1066
Changes in Forecasted Output Growth				
One Quarter Before	–0.1498	0.0552	–2.7145	0.0081
Contemporaneous	0.1383	0.0685	2.0182	0.0468
One Quarter Ahead	0.0931	0.0629	1.4802	0.1427
Two Quarters Ahead	–0.1185	0.0498	–2.3820	0.0195
Forecasted Headline CPI Inflation				
One Quarter Before	–0.0371	0.0219	–1.6929	0.0943
Contemporaneous	0.0640	0.0369	1.7332	0.0868
One Quarter Ahead	–0.0260	0.0466	–0.5584	0.5781
Two Quarters Ahead	0.0129	0.0486	0.2664	0.7906
Changes in Forecasted Headline CPI Inflation				
One Quarter Before	0.0097	0.0479	0.2030	0.8397
Contemporaneous	0.0213	0.0591	0.3597	0.7200
One Quarter Ahead	0.1018	0.0569	1.7892	0.0773
Two Quarters Ahead	–0.1115	0.0603	–1.8485	0.0681
Constant	0.0827	0.1878	0.4404	0.6608
R^2: 0.83				

(*continued*)

Table A.1. (Continued)

Variable	Coefficient	Std. Error	*t*-statistic	Prob.
B. Regression with Core CPI Inflation Forecast				
Initial Policy Rates in Canada	–0.2169	0.0282	–7.6814	0.0000
Policy Rates in United States (FFR)				
Level	0.1434	0.0189	7.5926	0.0000
Changes	0.0748	0.0296	2.5297	0.0134
CAD/USD Rates				
Level	–0.3728	0.2549	–1.4628	0.1474
Changes	–0.0136	0.0077	–1.7637	0.0816
Unemployment in Canada				
One Month Before	–0.0458	0.0793	–0.5778	0.5650
Two Months Before	–0.0217	0.0991	–0.2187	0.8275
Three Months Before	0.0459	0.0745	0.6157	0.5398
Forecasted Output Growth				
One Quarter Before	0.0337	0.0243	1.3846	0.1700
Contemporaneous	0.0248	0.0457	0.5424	0.5890
One Quarter Ahead	–0.0533	0.0455	–1.1719	0.2447
Two Quarters Ahead	0.0375	0.0376	0.9985	0.3210
Changes in Forecasted Output Growth				
One Quarter Before	–0.1188	0.0534	–2.2232	0.0290
Contemporaneous	0.0487	0.0689	0.7062	0.4821
One Quarter Ahead	0.1431	0.0726	1.9721	0.0521
Two Quarters Ahead	–0.1156	0.0556	–2.0801	0.0407
Forecasted Headline CPI Inflation				
One Quarter Before	–0.0203	0.0729	–0.2786	0.7813
Contemporaneous	–0.1870	0.1034	–1.8078	0.0744
One Quarter Ahead	0.1376	0.1627	0.8458	0.4002
Two Quarters Ahead	–0.0523	0.1289	–0.4057	0.6861
Changes in Forecasted Headline CPI Inflation				
One Quarter Before	–0.0447	0.1201	–0.3724	0.7106
Contemporaneous	0.3166	0.1984	1.5959	0.1145
One Quarter Ahead	–0.2862	0.2274	–1.2585	0.2119
Two Quarters Ahead	0.2309	0.1557	1.4832	0.1420
Constant	0.5186	0.3299	1.5719	0.1199
R^2: 0.81				

Note: This table reports regression coefficients of changes in Canada policy rates from previous MPC meeting on various independent variables.

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Shifts in ECB Communication: A Textual Analysis of the Press Conferences*

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This paper investigates how European Central Bank (ECB) communication, made during the press conference, affects stock market volatility. First, the ECB press conferences are dissected into topics using Latent Dirichlet Allocation (LDA). Then turning points in ECB communication are captured using the estimated topic probabilities. The proposed approach does not rely on subjective interpretation of topical content. The paper finds that (i) the topics reveal communication patterns that match the ECB monetary policy stance, (ii) the content of the ECB press conference is informative for the market, and (iii) market uncertainty increases when the ECB switches to a different communication regime.

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1. Introduction

This paper considers the problem of quantifying communication of the European Central Bank (ECB) during the press conferences on the Governing Council meeting days. Communication became a key tool for central banks to maintain transparency, manage market expectations, and achieve policy goals in a zero lower bound environment, where the room for maneuvering interest rates was limited (Blinder et al. 2008). Statements explaining monetary policy decisions are scrutinized by financial market participants; however, there is a great deal of subjectivity for a human reader trying to glean information by spotting patterns in multiple long text documents.

The ECB uses various channels to communicate its monetary policy stance: press conferences, monetary policy accounts, monthly bulletins, speeches, and interviews. The press conference that takes place on the same day as the Governing Council decision announcement is the primary communication device. It provides explanations for the monetary policy decision, the core assessment of the economic and monetary situation, and the forward guidance. Two main parts of a typical press conference are (i) an introductory statement, which is agreed upon by the members of the Governing Council, and (ii) a question-and-answer (Q&A) session, when journalists have the opportunity to ask clarification questions. This structure makes the ECB press conference a case study of both prepared and extemporaneous remarks.

A growing body of economic literature applies tools from computational linguistics to analyze central bank communication. The focus of this paper is to use these tools to study how the dynamics of topical composition of the ECB press conference affects stock market volatility on the Governing Council meeting days. The analysis proceeds in two stages. The first stage provides a low-dimensional representation of the transcripts by dissecting the ECB press conferences into topics. The second stage constructs a topic-based measure that captures any changes in the ECB communication regime.

To identify topics, we apply Latent Dirichlet Allocation (LDA) (Blei, Ng, and Jordan 2003), a generative model for text that enables extraction of multiple themes that are not specified in advance. In the analysis, the transcripts are represented by a document-term matrix, with each row representing a single document and each column corresponding to a unique word. The idea is to decompose the document-word relationships into topic probabilities in each document and word probabilities in each topic. Topics are interpreted as latent dimensions underlying the text.

The second part of the analysis is motivated by the communication patterns discovered with LDA. The model can identify phases when a single topic dominates in ECB communication and when a variety of topics is discussed. A novel aspect of our research is the construction of a score based on variations in the probability of the most dominant topic on a given conference day to capture substantial textual changes in the press conferences. The score is derived separately for the decision summary, the economic analysis, the monetary analysis, and the answers provided in the Q&A session during the tenures of Jean Claude Trichet and Mario Draghi. We examine the performance of the measure in explaining stock market volatility with event-based regressions.

The key findings are as follows. First, content exploration with LDA shows clustering of similar topics in each section of the press conference over time. This is expected, as the ECB strives to send a consistent message over time and similar speeches are easier to interpret. Of primary interest, therefore, are fundamental updates to the ECB wording, i.e., periods when one topic dies out and is replaced with a different topic. Comparison of the topic proportions over time with ECB monetary policy decisions shows that the changes in different sections of the introductory statement reflect the changes in the monetary policy regime. Analyzing the Q&A section, LDA identifies a discontinuity in topic probabilities, corresponding to the first press conference held by Mario Draghi.

Second, market volatility is higher on the conference days that the ECB introduces major revisions to the monetary analysis section, as compared with those conference days when the ECB sends a relatively homogeneous message.[1] This suggests that major revisions to the content of the introductory statement are more difficult to digest for the market, even when they occur in conjunction with changes in the monetary policy stance.

[1]Throughout this paper, we use the term "homogeneous message" to refer to statements that are primarily focused on a single topic.

This paper makes three distinct contributions to the field of analyzing central bank communication with computational linguistics tools. First, to our knowledge ours was the first study to apply LDA to the ECB press conferences, although the framework was successfully employed to analyze the statements, minutes, and transcripts of the Federal Open Market Committee (FOMC) (Hansen and McMahon 2016; Fligstein, Brundage, and Schultz 2017; Hansen, McMahon, and Prat 2017; Jegadeesh and Wu 2017).[2] Common alternatives to quantify text in economic literature are hand-coding (Jansen and De Haan 2005; Rosa and Verga 2007) or automated methods that rely on keyword counting (Tetlock 2007; Loughran and McDonald 2011). These approaches are deductive, as they typically capture meaning along a single, predefined dimension, like expansion-contraction or hawkish-dovish. LDA offers several advantages in that it satisfies the following conditions (DiMaggio, Nag, and Blei 2013): (i) it is reproducible; (ii) it is automated, so that it is easily updated when new documents arrive; (iii) it is inductive, to enable content discovery without imposing prior beliefs about what to look for in the text; and (iv) and it recognizes that terms may have different meanings in different contexts.

Second, this paper proposes a new content measure that is derived from LDA output but does not rely on subjective labeling of topics. A persistent challenge when using textual analysis is how to exploit the output to extract information that is relevant for financial market participants or improves the understanding of central bank decision makers. Current applications of LDA to central bank communication often rely on assigning substantive interpretations to topics based on the top most probable words in a topic (Hansen and McMahon 2016; Jegadeesh and Wu 2017). In contrast, our proposed measure captures the degree of discussion homogeneity, circumventing the need for assigning subjective topic labels. To facilitate content exploration, we employ automated measures of topic interpretability in the model selection procedure. By providing a summary of the whole document collection, the model

[2]We subsequently became aware of a small but growing literature that applied topic models to ECB press conferences in other contexts; see, e.g., Diessner and Lisi (2019), Dybowski and Kempa (2019), and Vo (2019). Since the initial writing of our paper, the use of topic models to analyze ECB communication has become quite popular.

not only enables study of the extent to which consecutive speeches are similar but also (i) what wording makes the speeches similar, (ii) whether the topics are recurring, and (iii) how long the transition period to a new topic is.

The third contribution is methodological. LDA is a hierarchical Bayesian model, where the hyperparameters that index prior distributions on a set of latent variables are found to substantially influence model inference (Asuncion et al. 2009; Wallach, Mimno, and McCallum 2009; George and Doss 2018). We adopt a fully Bayesian approach to formally infer the values of hyperparameters. In contrast, to date, textual analyses in economics commonly has chosen the values of the hyperparameters in an ad hoc manner (Griffiths and Steyvers 2004).

The structure of the paper is as follows. Section 2 reviews strategies to quantify text in economic research, and Section 3 presents the methodology of LDA. Section 4 describes the data and text preprocessing steps. Section 5 investigates the estimated topics and the shifts in ECB communication. Section 6 concludes. An extensive set of appendices, containing estimation details and documenting other decisions made in conducting the analysis, then follows.

2. Related Literature

Our work lies in the intersection of two strands of literature: the impact of central bank communication on the financial market, and natural language processing (NLP), in particular topic modeling. This section provides an overview of the methods for mapping words to meaningful quantities within economic literature, with a focus on central bank communication.

The literature on central bank communication uses three approaches to gauge the effect of communication: an indirect approach, manual coding, and automated textual analysis. The automated methods are most relevant for this paper. The indirect approach does not quantify verbal information. Instead, using high-frequency data, it measures financial market movements within a narrow window of the decision announcement and surrounding communication. A stylized fact following from indirect analyses is that the market reaction to central bank communication is more pronounced than the reaction to monetary policy decisions (Gürkaynak, Sack, and Swanson 2005; Ehrmann and Fratzscher 2009; Brand, Buncic, and Turunen 2010). Furthermore, for the ECB the market reaction to the press conference is stronger for less anticipated decisions, indicating that the introductory statement provides relevant clarifications (Ehrmann and Fratzscher 2009). The reasoning behind this result is that in times of high uncertainty (when the surprise component in a policy decision is large), the reaction to the actual decision is muted, as the market expects a subsequent explanation and instead responds to that.

A step further is to identify pieces of information that move the markets. The information can come either in the form of topics or tone. To extract the information content, one can follow a manual or an automated approach. The manual approach involves hand-coding the statements on an ordinal scale or classifying verbal expressions to predefined categories. For example, Ehrmann and Fratzscher (2009) manually classify real-time newswire reports during the ECB press conference via the following content categories: economic outlook, inflation, second-round effects, money growth, and interest rates. Statements on inflation and interest rates turn out to be the most important market movers. By hand-coding each ECB introductory statement on a scale ranging from −2 (very dovish) to 2 (very hawkish), Rosa and Verga (2007) find that ECB words are complementary to data on macroeconomic variables in predicting the moves in the key ECB interest rate, showing that market expectations react to the unexpected component of the press conference content. The main criticism of the manual approach is high subjectivity and low reproducibility. Furthermore, because the manual coding is done ex post, coders might unintentionally mitigate the unexpected component in the statement and fail to capture how the financial markets understood the message at the release time (Blinder et al. 2008).

To overcome these issues, a strand of literature has turned to automated approaches to ensure that the analysis is transparent and scalable. Overall, within the automated methods one can either define a priori dimensions to look for in the text, or apply an algorithm to discover dimensions. In the former case, the most intuitive and relatively simple technique is a dictionary method, where a researcher predefines a list of keywords describing meanings of interest. Documents are then summarized by the number of occurrences of each word in the keyword list. In principle, by defining word lists that separate multiple categories, it is possible to capture multiple dimensions in text (Tetlock 2007); however, typically only two opposing concepts are considered. For example, the word counts can be converted to a single communication measure of incremental changes in hawkish and dovish monetary policy inclinations (Apel and Grimaldi 2012), positive and negative tone (Tetlock, Saar-Tsechansky, and Macskassy 2008; Jegadeesh and Wu 2013; Born, Ehrmann, and Fratzscher 2014) or periods of greater or less uncertainty (Jegadeesh and Wu 2017).

One of the main difficulties with the dictionary approach is developing a word list that accurately captures the meaning for a specific application. Since words often carry different sentiment or meaning in different contexts, dictionaries developed in one domain of study can lead to word misclassification when used in other disciplines (Loughran and McDonald 2011). This necessitates development of methods that are customized to central bank communication. Previous literature has gone beyond the generic dictionaries by capturing contextual information with (i) field-specific dictionaries (ii) sentence-level scores, and (iii) intensity of specific themes. Looking at the ECB, Picault and Renault (2017) manually develop a field dictionary based on the introductory statements and (similarly to our paper) investigate the European stock market reaction to the press conference. They find that market volatility increases (decreases) when the statements about monetary policy are hawkish (dovish) and the tone about the economic outlook is negative (positive). Instead of considering word occurrences in isolation, Lucca and Trebbi (2009) devise a sentence-based score and show that discourse orientation in the FOMC statements explains a large portion of the federal funds rate variation. Finally, analysis of thematic content highlights the importance of shifts in discourse reflected by changing topic intensity over time—for example, the increasing role of financial stability in monetary policy considerations (Peek, Rosengren, and Tootell 2016; van Dieijen and Lumsdaine 2018).

In this paper we employ a probabilistic topic model, specifically LDA, to identify the most important dimensions in the ECB press conferences. The central application of topic models is summarizing a large collection of documents and discovering patterns in textual data. However, the topics themselves are rarely the final objective of the analysis. Although there are examples where topic models mainly augment descriptive analysis (Quinn et al. 2010; Fligstein, Brundage, and Schultz 2017), recent applications to central bank communication attempt to derive communication measures using estimated topics, often in combination with dictionary methods in order to understand how the information in central bank communication affects market returns, volatility, and interest rate expectations (Hansen and McMahon 2016; Jegadeesh and Wu 2017). A closely related work to our paper is Jegadeesh and Wu (2017) who use LDA to investigate how the U.S. stock market reacts to proportions of discussion on, and tone adopted in, different topics in the FOMC minutes. They find that the Federal Reserve's discussion of its policy stance and inflation is most informative for the market, whereas topics like trade and consumption are not informative. Unlike the above implementations, we avoid deriving conclusions that depend on subjective interpretations of topics, instead focusing on the properties of the estimated document-topic probabilities, similar to Hansen, McMahon, and Prat (2017).

3. Methodology

3.1 Latent Dirichlet Allocation

Latent Dirichlet Allocation (LDA) introduced by Blei, Ng, and Jordan (2003) is a mixed membership model for text. The basic idea is that observations (words) are grouped into documents and each of these groups (documents) is modeled with a mixture of distributions. The components of the mixture are topics, which are multinomial probability distributions over a fixed vocabulary. The topics are shared across all documents (each document is built from the same components), but the proportions of topics in documents vary.

LDA ignores both the document order and the word order within the documents. A document is represented as a "bag of words." The inference is based on the notion of word co-occurrence. Words that often appear together across documents are likely to belong to the same topic. Intuitively, LDA trades off two conflicting goals in finding a good topical representation for a collection of documents, that of assigning words in each document to few topics versus assigning a high probability in each topic to few words (DiMaggio, Nag, and Blei 2013).

The central inferential problem in LDA is to determine the posterior distribution of topic proportions in documents ($\boldsymbol{\Theta}$), word proportions in topics ($\boldsymbol{\Phi}$), and word-topic assignments ($\mathcal{Z}$), given the hyperparameters ($\boldsymbol{\alpha}$, $\boldsymbol{\beta}$) and corpus of documents, $\mathcal{W}$. The formal statement of this idea, details on the notation used, and derivations are contained in Appendix A. A corpus is a collection of documents, where D is the number of documents, N_d is the number of words in document d, and V is the number of distinct words in the collection of documents. It is assumed that each document is composed of K topics in different proportions. The posterior distribution is proportional to the complete data likelihood function times the prior:

$$p(\boldsymbol{\Phi}, \boldsymbol{\Theta}, \mathcal{Z}|\mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}) \propto \prod_{d=1}^{D} \underbrace{p(\boldsymbol{\theta_d}|\boldsymbol{\alpha})}_{Dirichlet} \prod_{k=1}^{K} \underbrace{p(\boldsymbol{\phi_k}|\boldsymbol{\beta})}_{Dirichlet} \left(\prod_{d=1}^{D} \prod_{i=1}^{N_d} \underbrace{p(w_i^{(d)}|z_i^{(d)}, \boldsymbol{\Phi})}_{Multinomial} \underbrace{p(z_i^{(d)}|\boldsymbol{\theta_d})}_{Multinomial} \right). \quad (1)$$

Implementation of LDA involves important model specification and selection decisions. In particular, the estimation results vary according to the number of topics (K) and hyperparameter settings ($\boldsymbol{\alpha}$, $\boldsymbol{\beta}$). The goal is to obtain $p(\boldsymbol{\Phi}|\mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}), p(\boldsymbol{\Theta}|\mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta})$ and $p(\mathcal{Z}|\mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta})$, that is, the word probabilities, topic probabilities, and word-topic assignments. These distributions cannot be computed in closed form. As estimation is straightforward, we omit details here and note that a thorough discussion of the estimation method and how it builds on more common strategies for approaching LDA is contained in Appendix A.

3.2 Model Evaluation

Choosing the number of latent topics and assessing their quality is a long-studied problem in unsupervised topic modeling. Typically, there is a trade-off between predictive accuracy of the model and topic interpretability (Chang et al. 2009).

Metrics of predictive performance, such as held-out likelihood or perplexity, are conventionally used to assess model quality (Blei, Ng, and Jordan 2003; Wallach et al. 2009).[3] However, the predictive metrics have limitations. Usually fine-grained, highly specific topics yield the best model fit, but they are not easy to interpret or to generalize (Boyd-Graber, Mimno, and Newman 2014). Furthermore, predicting the content of the preprocessed text is rarely the objective of research in political, economic, or social science applications.

Roberts et al. (2014) argue that a semantically interpretable topic has two qualities: (i) it is coherent—the highest probability words for the topic tend to co-occur within documents, and (ii) it is exclusive—the words that have high probability under one topic have low probabilities under other topics.

Our model selection procedure prioritizes interpretation over prediction. We first discard solutions with the lowest coherence or exclusivity, akin to Roberts et al. (2014), and then select the solution with the lowest perplexity among the remaining models.

3.2.1 Coherence and Exclusivity

Automated measures of coherence usually assume that co-occurrence frequency of terms within documents is informative about semantical relatedness of the terms and are based on averaging some measure of pairwise association between the most probable words in a topic (Newman et al. 2010).

The models estimated on the corpus of the ECB press conferences are evaluated with a semantic coherence score of Mimno et al. (2011). The score is shown to match well with human judgments and is defined as

$$\text{Coherence}_k = \sum_{j=2}^{N}\sum_{i=1}^{j-1} \log \frac{D(w_i^{(k)}, w_j^{(k)}) + 1}{D(w_i^{(k)})}, \tag{2}$$

where $D(\cdot)$ is a function that returns the number of documents containing all of the words provided as arguments, and $w_i^{(k)}$ denotes a word from the list of top N words with the highest probability in topic k. Intuitively, the measure is related to the conditional probability of observing a word given another higher-ranked word. The

[3]Perplexity is defined as the inverse of the geometric mean per-word likelihood of the test data; see Appendix A for discussion.

semantic coherence of Mimno et al. (2011) relies on the word frequencies in documents being modeled, hence it is more intrinsic in nature.

Coherence measures inform inference about internal consistency of topic representation, but they do not penalize topics that are similar (Roberts et al. 2014). A counterpoint to semantic coherence is topic exclusivity that captures inter-topic similarity by comparing the usage rate of words with high probability in a topic relative to other topics. Exclusivity of term v in topic k is defined as (Bischof and Airoldi 2012; Airoldi and Bischof 2016):

$$\text{Exclusivity}_{v,k} = \frac{\phi_{k,v}}{\sum_{i=1}^{K} \phi_{i,v}}. \tag{3}$$

In other words, the exclusivity score is the probability of a word in a topic divided by the sum of probabilities of this word in all topics. Exclusivity of topic k is computed as an average of the scores for the top N words in that topic. A high exclusivity score indicates that the most common words in a particular topic are not common to other topics.

3.2.2 Topic Cardinality and Word Ranking

Topic-based measures of coherence and exclusivity operate on a ranking of the top N words with the highest probability. The standard practice is to select N arbitrarily (usually $N = 10$). To achieve more stable evaluation, we compute semantic coherence (2) and exclusivity (3) for different topic cardinalities: $N = 5, 10, 15, 20$ and average them (Lau and Baldwin 2016).

The word ranking based on term probability in a topic favors terms with high frequency in a corpus, but the most common words might not carry any semantically useful information, and can be used similarly in every topic. We use a FREX (frequency-exclusivity) score of Bischof and Airoldi (2012) to represent a topic with words that are both frequent and exclusive. The score combines these two dimensions via the harmonic mean of frequency and exclusivity:

$$\text{FREX}_{v,k} = \left(\frac{\omega}{\text{ECDF}(\text{Exclusivity}_{v,k})} + \frac{1-\omega}{\text{ECDF}(\phi_{k,v})} \right)^{-1}, \tag{4}$$

where ECDF is the empirical cumulative distribution function and ω is a weight given to exclusivity (set to 0.5).

3.3 Measuring Tone

We additionally compare the estimated topics with lexicon-based tone-measures.[4] Following Shapiro and Wilson (2021), we calculate the net negativity score as the difference between the fraction of negative and positive words after text preprocessing, based on the list of positive and negative words provided in the Loughran and McDonald (2011) dictionary (LM dictionary). The computations ignore negations. The score can be computed on different levels: across sections or for the whole statement.

4. Data

This section introduces the ECB press conference and describes the steps to convert text to numerical data. It also presents the financial data used to measure the market reaction to the topic dynamics of the press conference.

4.1 The ECB Press Conference

The ECB's monetary policy decisions are published at 13:45 CET on the day of the Governing Council monetary policy meeting. The press conference starts at 14:30 CET on the same day. It begins with an introductory statement by the ECB President, who explains the monetary policy decision.

The press conference consists of six major sections: (i) summary of the ECB's monetary policy decision; since July 2013 this summary also includes forward guidance; (ii) economic analysis; (iii) monetary analysis; (iv) "cross-check" paragraph; (v) fiscal policy and structural reforms; (vi) question-and-answer (Q&A) session.

The economic analysis and monetary analysis sections are the two pillars by which the Governing Council evaluates the risks to price stability. The economic analysis part looks at the short- to medium-term outlook, whereas the monetary analysis part assesses

[4]We are grateful to an anonymous referee for suggesting this analysis.

medium- to long-term trends. The cross-check paragraph was introduced in 2003 and its role is to compare signals from the two pillars.[5]

The ECB held monetary policy meetings and related press conferences on a monthly basis until December 2014. From 2015 the frequency of the meetings was changed to a six-week cycle. Our analysis considers all ECB press conferences between January 2004 and April 2018, covering 91 speeches from Jean-Claude Trichet (whose eight-year term expired at the end of October 2011) and 65 speeches from Mario Draghi. The textual data have been scraped from the ECB website.[6]

4.2 Preparing the Documents

The ECB press statements have a standardized structure, with sections that are fixed over time. Each section defines a main theme that is easily captured at the preprocessing stage (i.e., via the section headings), but latent topics within the theme are more difficult to capture by the human reader and vary over time. Importantly, the sections are sufficiently long to enable us to run LDA on them separately. For our purposes, therefore, a document is defined at the section level and a separate model is estimated for each section. The main motivation for treating the sections separately is to allow us to track the topics within sections and compare the changes across sections. In addition, focusing on the sections separately provides more confidence about the context in which words should be understood, alleviating drawbacks of the "bag-of-words" representation, as well as being able to assess whether substantial updates occur to just a single part of the statement or to multiple sections simultaneously. Running the LDA on the whole introductory statement would likely lead to topics that are dominated by the general sectional themes that are already known before running the model.

[5]In May 2003 the ECB introduced the new structure of the introductory statement in which economic analysis is discussed first and monetary analysis is put second. The ECB motivated this decision by stating that "the Governing Council wishes to clarify communication on the cross-checking of information in coming to its unifed overall judgement on the risks to price stability" (European Central Bank 2003).

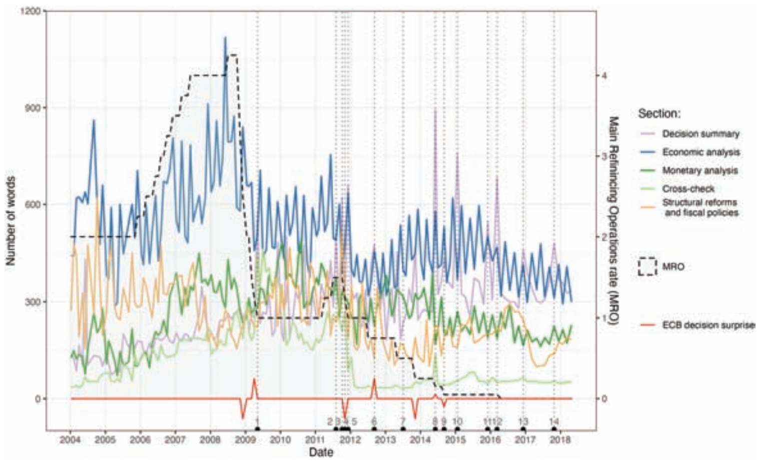
[6]See https://www.ecb.europa.eu/press/pressconf (accessed April 2018).

For each press conference we (i) break the transcript into individual paragraphs; (ii) assign each paragraph to a section; and (iii) extract answers from the Q&A session. We use keywords that are defined as bold word sequences in the HTML code of the press conference to record the section where each paragraph is located. For example, a paragraph that contains the keyword "key ECB interest rates" is identified as the first paragraph of the decision summary, and a paragraph containing the keyword "economic analysis" marks the beginning of the economic analysis section.

Figure 1 shows how the number of words per section of the introductory statement evolved over time, along with the main refinancing operations rate (MRO, dashed line), monetary policy surprise, and decisions regarding non-standard monetary policy measures. The surprise component (red line) is measured by subtracting the Bloomberg® survey median forecast from the ECB rate announcement (Bloomberg L.P. 2018). Based on the raw word counts, economic analysis is given broader coverage than monetary analysis. In addition, the ECB appears to have communicated relatively more on the economic outlook when it was raising the interest rate (until about the mid-2008) than when it cut the interest rate. We base this inference on a comparison of the relative number of words devoted to the economic analysis section versus the monetary analysis section.[7] The spikes in the number of words in the decision summary can be matched with ECB announcements about new monetary policy tools and implementation details. In addition, since Mario Draghi became the ECB President in November 2011, the coverage of the cross-check part has sharply decreased and currently contains only a single sentence that the cross-check of the monetary analysis and the economic analysis confirms the need for the undertaken monetary policy action. Because of LDA's deficiency in handling documents that are too short and the low informational value of the cross-checking over Draghi's tenure, the cross-check section is not considered in the estimation.

[7]We note, however, that this pattern may just reflect the time period associated with these episodes, with the cuts being driven by the European sovereign debt crisis rather than the overall state of the economy, and hence may not hold generally.

Figure 1. Number of Words per Section of the ECB Introductory Statement and the Main Refinancing Operations (MRO) Rate



Note: The figure shows the raw word counts in five sections of the introductory statement: decision summary, economic analysis, monetary analysis, cross-check, structural reforms, and fiscal policies. The dashed line represents the level of the main refinancing operations rate. The red line shows the policy decision surprise which is measured by subtracting the Bloomberg® survey median forecast from the ECB rate announcement (Bloomberg L.P. 2018). The timeline markers represent the following events: 1. Announcement of the first covered bond purchase program (CBPP1) and one-year longer-term refinancing operation (LTRO); 2. Announcement of six-month LTRO; 3. Announcement of CBPP2; 4. Announcement of three-year LTRO, collaterals, and reserve ratio. 5. The first introductory statement by Mario Draghi; 6. Introduction of the forward guidance; 7. Announcement of the outright monetary transactions (OMT); 8. Announcement of targeted longer-term refinancing operations (TLTROs); 9. Announcement of CBPP3 and the asset-backed securities purchase program (ABSPP); 10. Announcement of the expanded asset purchase program (APP, known as quantitative easing); 11. Announcement about extension of APP; 12. Announcement of the corporate sector purchase program (CSPP) and TLTRO2; 13. Announcement about extension of APP; 14. Announcement about unwinding of the stimulus.

4.3 Vocabulary Selection

Text preprocessing choices can substantially affect model output (Boyd-Graber, Mimno, and Newman 2014; Denny and Spirling 2018). Common text treatments are removing punctuation and numbers, lowercasing, stop word removal, term normalization (stemming or lemmatization), n-gram inclusion, and removing very common or very rare words.

First, we remove neutral sentences or parts of sentences that introduce the next section and are repeated in every speech, for example: “Ladies and gentlemen, the Vice President and I are very pleased to welcome you to our press conference”, “Let me now explain our assessment in greater detail, starting with the economic analysis”, “We are now at your disposal for questions”. The complete list of expressions that were removed is provided in Appendix B. We also clean from the Q&A section the answers in French, because English translations of these answers (that are included in the analysis) immediately follow.

The second step is to convert all words to lowercase, remove punctuation, stop words, and month names. Stop words are common function words like “the” or “and” with no inherent useful information, and their overwhelming presence in all documents can produce spurious associations between content words (Roberts et al. 2014).[8] We also remove all words containing non-alphabetic characters, except for abbreviations for money aggregates (M1, M2, M3) and groups of countries (G3, G7, G8, etc.).

The third step is term normalization: each term is classified into its part of speech (POS) using the Stanford POS tagger (Collobert et al. 2011) and reduced to its dictionary form by lemmatization.[9] An alternative approach to reduce inflectional and derivational word forms is stemming. We opt to use a lemmatizer because it is more accurate than a stemmer and it is unlikely to over-conflate (Schofield and Mimno 2016). See Appendix B for discussion.

Finally, we identify collocations and create multiword expressions, called n-grams, which allow one to capture the broader context of a word and reduce ambiguities resulting from the “bag-of-words” assumption.[10] The list of n-grams, available from the authors on

[8]The stop word list is from http://snowball.tartarus.org/algorithms/english/stop.txt (accessed April 2018). It includes pronouns, articles, prepositions, and conjunctions.

[9]The Stanford POS-tagging algorithm is used to provide auxiliary information about the part of speech for the WordNet lemmatizer in Python.

[10]Specifically, we manually rate bigrams and trigrams occurring at least 50 times to filter out n-grams that do not add new meaning beyond the meaning of

Table 1. Data Dimensionality Reduction After Preprocessing Steps

	Raw	Stop Words Removal and Lemmatization	Creating n-grams
Total Words	776,112	365,930	326,343
Average Section Length	829	391	349
Unique Words (Vocabulary Size, Overall)	9,224	6,120	6,216
Decision Summary	1,798	1,252	1,337
Economic Analysis	1,805	1,263	1,346
Monetary Analysis	1,589	1,042	1,094
Cross-Check	901	654	699
Structural Reforms, Fiscal Policies	2,381	1,675	1,741
Q&A	8,827	5,935	6,021

Note: This table reports descriptive statistics of the vocabulary in the ECB press conferences before and after implementing the preprocessing steps: stop words removal, lemmatization, and creating n-grams. Generally, using n-grams increases the vocabulary size; for example, one might have the bigram "monetary_policy" and the unigrams "monetary" and "policy" in the vocabulary list. The last column presents the vocabulary size used in the estimation of Latent Dirichlet Allocation (LDA).

request, includes technical terms used by the ECB such as "full allotment" or "covered bond", expressions providing context for very common words, like "key ecb interest rate unchanged", as well as long-used statements specific to ECB communication, such as the premise to "never pre commit" to any future policy action.

Table 1 reports descriptive statistics of the vocabulary before and after implementing the preprocessing steps.

4.4 Financial Data

We use daily closing values of the VSTOXX index to measure investors' reaction to ECB communication patterns on press

the constituting words (i.e., "regard second", "question say"). We then sort the n-grams by pointwise mutual information (Bouma 2009) to select those n-grams that occur less often but where the association between words is high (i.e., "tail risk" or "banca italia"). Selected n-grams are treated as separate words in the analysis. Words that appear without collocation stay as separate words.

conference days. The VSTOXX index represents the implied volatility of the Euro Stoxx 50 index (EURO STOXX 50 real-time option prices) and is designed to reflect market expectations of near-term volatility. The index was also investigated in the context of ECB communication and monetary policy actions by Fratzscher, Lo Duca, and Straub (2016) and Picault and Renault (2017), and is often used as a proxy for uncertainty in the euro area. The daily closing values of the VSTOXX index for stock market volatility are sourced from Bloomberg®. The series is log-transformed and differenced to approximate the daily percentage change.

A number of control variables are considered in the empirical investigation: the surprise component of the ECB interest rate decision, a dummy variable for the announcements regarding non-standard monetary policy measures (the complete list of the announcements is presented in Figure 1), German two-year government bond yields, and the surprise component of the U.S. jobless claims. The data on German government bond yields, the MRO rate, and released values of the U.S. jobless claims are collected from Bloomberg®. The sample period for the financial variables is from January 2004 to April 2018. All surprise components are constructed by deducting the Bloomberg® survey median expectations of professional forecasters from the released value.

5. Results

This section describes the main findings. It starts with general remarks about model selection and properties of the estimated topic-word and document-topic distributions. Next, it investigates the changing attention to different topics over time.

5.1 Estimated Topics

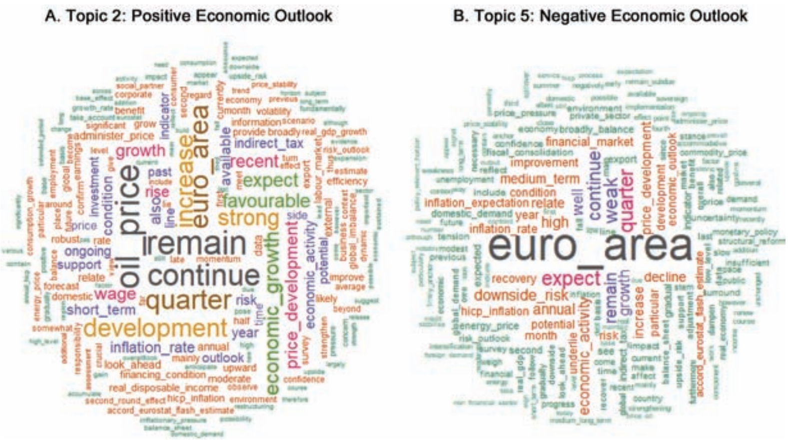
The model (1) is estimated for the four separate sections of the press conference—decision summary, economic analysis, monetary analysis, and Q&A—using a different number of topics for each. In line with the findings of Chang et al. (2009), higher model complexity (more topics) results in lower perplexity, but also in lower average coherence. Exclusivity does not seem to be related to semantic coherence, confirming that the two measures capture distinct aspects of topic interpretability. The set of solutions with the highest coherence and exclusivity is dominated by relatively parsimonious models. The selected dimensionality varies across sections, but it does not exceed 10 topics. Diagnostic plots illustrating model selection are presented in Appendix C.

We find that document-topic distributions are generally sparse in all sections, i.e., few topics comprise a document. The conclusion about sparsity of the document-topic distributions does not change if a different number of topics is specified. Furthermore, LDA groups the press conferences in time although no information about the order of documents is incorporated in the estimation procedure. The sparsity of document-topic distributions and the similarity of consecutive documents lead to identification of different phases of ECB communication. Although the sections of the press conference were considered separately in the estimation, the algorithm identifies a rise of a new topic in each section at approximately the same time.[11]

As expected, frequent words in the corpus often end up scattered across the top most likely words in many topics. The term re-ranking using the FREX score downgrades general terms and corpus-specific stop words and reveals intuitive topic interpretations based on keywords that are both frequent and exclusive to a specific topic. For example, Figure 2 presents word clouds that show relative (raw) frequency of words in the two most popular topics in the economic analysis section. The more frequent a term is in a topic, the larger it becomes in a cloud. If topics are represented solely by their most frequent terms, they would be described by non-exclusive words and many topics in the economic analysis section would appear to be similar. On the other hand, re-ranking words by the FREX score captures important differences between words in the topics, suggesting

[11] It is worth stressing that the topic sparsity in the ECB press conferences is not detected if one follows the heuristics about Dirichlet prior parameters (Griffiths and Steyvers 2004), widely applied in economic research, instead of estimating the hyperparameters. The heuristic ($\alpha = \frac{50}{K}$) imposes that the document-topic distribution is smooth for $K < 50$. In line with the heuristic regarding the Dirichlet prior parameter for topic-word distributions ($\beta = 0.1$) the estimated word-topic distributions are sparse: there is a limited number of words with relatively high probability in each topic.

Figure 2. Distributions over Terms Represented as Word Clouds



Note: The word clouds show the top 200 most frequent terms in two topics of the economic analysis section. The topics were estimated using Latent Dirichlet Allocation (LDA) algorithm. In the word clouds the size of a term is proportional to the term probability. If topics were represented in terms of their most frequent terms, they would be described by non-exclusive, high-frequency words (for example, "euro area", "continue", "remain") and many topics in this section would appear to be similar. However, the FREX score gives high ranks to keywords that are both frequent and distinctive for a specific topic. Top terms ranked by the FREX score in topic 2 (left) are *side, robust, economic growth, earnings, favourable, efficiency, lie, short term, consumption growth,* whereas top terms ranked by the FREX score in topic 5 (right) are *weak, low level, economic outlook, gradual, public, expected, modest, insufficient, global demand, slow.*

one of them be labeled as "Positive economic outlook", and the other one as "Negative economic outlook" (see Table 2).

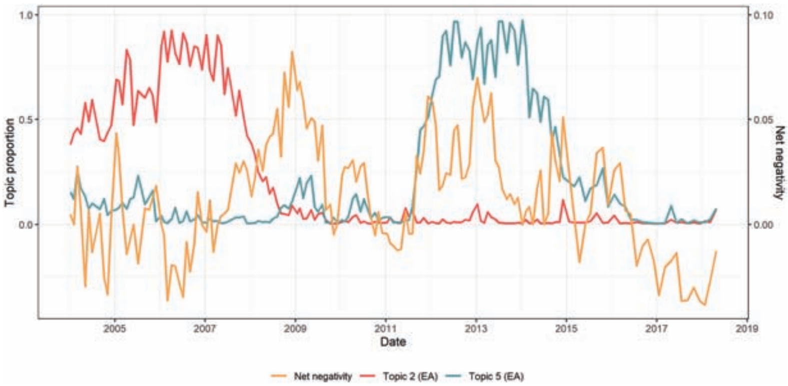
Topic labels are consistent with the tone that is associated with the occurrence of those topics. Figure 3 shows the proportion of the two topics of the economic analysis section, topic 2 (labeled as "Positive economic outlook") and topic 5 (labeled as "Negative economic outlook", along with the net negativity score (right scale) computed for the economic analysis (EA) section. During the time that topic 2 dominates, the net negativity score is lower. Conversely, the net negativity score is higher when topic 5 dominates the discussion. Importantly, these topics represent more than just

Table 2. Top 10 Terms Describing Topics of the Economic Analysis Section, Ranked by the FREX Score

1 "Projections"	2 "Positive Economic Outlook"	3 "Wage-Price Spiral"
ecb	side	scheme
eurosystem	economic_growth	avoid
range	robust	party
staff_macroeconomic_projection	earnings	food_price
projection	efficiency	sound
revise	favourable	behaviour
staff_projection	oil_price	shock
foresee	lie	constraint
upwards	short_term	power
downwards	consumption_growth	call
4 "Stimulus"	**5 "Negative Economic Outlook"**	**6 "Recovery"**
correction	weak	private_consumption
function	low_level	economic_recovery
stimulus	economic_outlook	exchange_rate
macroeconomic	gradual	measure
owing	public	structural_reform
financial_system	expected	monetary_policy
commodity_price	underlie	closely
aim	modest	geopolitical_risk
keep	insufficient	pick
protectionist_pressure	global_demand	household

Note: The FREX score gives high ranks to terms that are both frequent and exclusive.

Figure 3. Proportion of Topic 2 (labeled as "positive economic outlook") and Topic 5 (labeled as "negative economic outlook"), Along with the Net Negativity Score, for the Economic Analysis (EA) Section



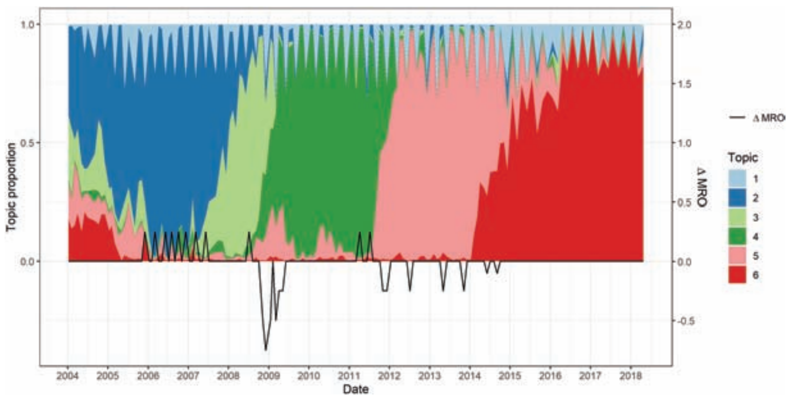
Note: This figure plots the proportion of the economic analysis section devoted to two topics along with the net negativity score for the section. The topics were estimated using the LDA algorithm (Blei, Ng, and Jordan 2003). The net negativity score is the difference between the fraction of negative and positive words in the section, based on the list of positive and negative words provided in the Loughran and McDonald (2011) dictionary.

the tone, however. Topic 2 is not consistently active in every time period that net negativity is lower, and topic 5 is not consistently active in every time period that the net negativity is higher.

5.2 Interpreting Topical Content

As external validation of the ECB communication patterns identified by LDA, we compare the attention to different topics with changes in the main refinancing operations rate in order to analyze how different communication regimes correspond to the phases of the ECB monetary policy stance. For the interpretation of textual themes, we focus on the economic analysis section and the Q&A section. The results obtained for the remaining sections are provided in Appendix D.

Figure 4. Topics in the Economic Analysis Section over Time



Note: This figure plots the proportion of the economic analysis section devoted to each topic along with the ECB MRO rate decisions. The topics were estimated using the LDA algorithm (Blei, Ng, and Jordan 2003). The sample comprises 156 transcripts of the section from the ECB press conferences between 2004 and 2018.

Figure 4 graphs topic proportions over time in the economic analysis section. The key terms of topic 1 ("staff macroeconomic projection", "range", "revise", "upwards", "downwards"; see Table 2) appear to capture a discussion about macroeconomic projections. The topic is especially active on the press conference days in March, July, September, and December when the quarterly staff macroeconomic projections are presented.

The remaining topics in the economic analysis section can be reasonably associated with various phases of the ECB monetary policy stance. Topic 2 remains strong during the tightening phase 2005–07. The topic is mostly characterized by both frequent and exclusive terms such as "robust", "favourable", and "efficiency", emphasizing a positive economic outlook. It declines shortly after the sequence of the rate hikes; its proportion falls permanently below 50 percent on the meeting in December 2007, whereas the last rate hike in the sequence occurred in June 2007.

Topic 3 is the most prominent during the first phase of policy responses to the financial turmoil that started in August 2007 (Stark 2009). In that period the ECB particularly often used the keyword "scheme" to express the concern about a wage-price spiral, but in general the fundamentals of the euro-area economy were described as "sound".[12]

The bankruptcy of Lehman Brothers in September 2008 marks the intensification of the crisis and precedes an abrupt change in ECB communication. Topic 4 surges in November 2008, exactly on the first conference day the ECB cut its key interest rate by 50 basis points after the Lehman collapse.[13] Distinctive for this phase is a discussion about "financial system" and "stimulus". This phase ended with two interest rate increases in April and July 2011, which turned out to be premature (Constâncio 2018).

The rise of topic 5 marks the start of the recession in the third quarter of 2011 that lasted until the first quarter of 2013, according to the Euro Area Business Cycle Network.[14] This phase is associated with the easing cycle where the language used by the ECB ("weak", "low level", "modest", "insufficient", "slow") reflected the weakness of the economy.

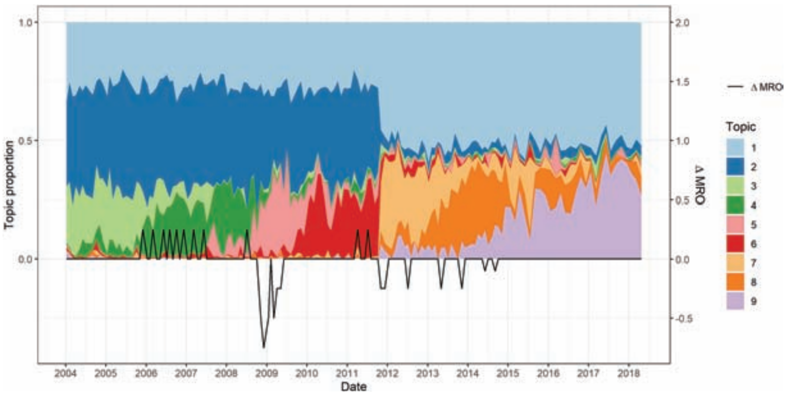
The discourse represented by topic 6 was emerging gradually, as the interest rates were approaching the zero lower bound. The timing coincides with the ECB's introduction of its unconventional monetary policy instruments and hence predominant for topic 6 is the keyword "monetary policy measure", but the other frequent and exclusive terms are "economic recovery", "structural reform", "exchange rate", "household", and "private consumption". Interestingly, a reading of the statements confirms that the ECB expressed concerns about exchange rate developments, discussed the structural reforms, private consumption, and the situation of the households as a part of its economic analysis solely in the statements where topic 6 is active (2004–05 and 2013–17) and never in between. What

[12]The ECB has repeatedly used the term "scheme" and "shock" in the following context: "the Governing Council is concerned about the existence of schemes in which nominal wages are indexed to consumer prices. Such schemes involve the risk of upward shocks in inflation leading to a wage-price spiral" (press conference, July 3, 2008).

[13]The first press conference after the Lehman collapse was held on October 2, 2008 and the decision was to keep the interest rates unchanged. The first interest rate cut in response to the financial crisis was unscheduled. It took place on October 8, 2008 as a part of the coordinated action with other major central banks.

[14]See https://eabcn.org/dc/chronology-euro-area-business-cycles (accessed April 2023).

Figure 5. Topics in the Q&A Section over Time



Note: This figure plots the proportion of the Q&A section devoted to each topic along with the ECB MRO rate decisions. The topics were estimated using the LDA algorithm (Blei, Ng, and Jordan 2003). The sample comprises 156 transcripts of the section from the ECB press conferences between 2004 and 2018.

is common to these two periods is that both concern the phase of the economic recovery. The recovery discussed in 2004–05 followed the protracted period of economic slowdown experienced from mid-2001 to mid-2003 (European Central Bank 2009). This suggests that there might exist some recurring textual patterns of central bank communication, although the current sample is too short to draw definitive conclusions between communication patterns and the business cycle.

Turning to the Q&A session, recall that following the introductory statement the ECB has the opportunity to clarify its messages and emphasize its point of view about the economic outlook. Therefore, the questions may reveal ambiguities in ECB communication or indicate topics that reporters find important. In contrast to the introductory statement, which is prepared by the whole Governing Council, the answers provided by the ECB president during the Q&A session are non-prompted. Therefore, we can expect differences between the Q&A session and the other sections, as well as between the wording used by Jean-Claude Trichet and Mario Draghi.

Figure 5 shows the topical representation of answers provided during the Q&A session. Several interesting points emerge. A spontaneous section, in comparison with the prepared sections, appears to have a larger proportion of words that do not contribute to the informational value of the president's response, for example "mean", "come", "particular", or "already" (see Table 3). As expected, LDA with an asymmetric prior on the document-topic distribution was able to handle these very common words in an appropriate fashion and sequester them into topics 1 and 2.

There is a discontinuity in the topics' probabilities occurring at the first conference held by Mario Draghi in November 2011. The discontinuity in the time series of topics 1 and 2 may reflect different speaking styles of both presidents, but there is also a clear split among specialized topics discussed during the tenures of Trichet and Draghi.

Starting with the answers of Trichet, the attention to the topic "Vigilance" was dominating in advance of and during the tightening phase in 2005–07. This observation is in line with Jansen and De Haan (2007), who found that the term "vigilance"/"vigilant" was used extensively in ECB communication starting in March 2004 and continued to be mentioned after the tightening cycle, but less often. The code word "vigilance" used to be a clear signal for financial markets that the ECB would pre-announce any interest rate hike.[15] Topic 5 also has a natural label. It clusters terms related to various liquidity-injecting operations provided to the banking sector (main refinancing operations and longer-term refinancing operations (LTROs)).

The Q&A sessions held by Draghi seem to be richer in content. The focal points are the forward guidance and non-standard monetary policy measures (LTROs—long-term refinancing operations, OMTs—outright monetary transactions, the asset purchase program), the Greek crisis, and ELA (emergency liquidity assistance, on which the Greek banks have been highly dependent since being cut off from standard ECB funding options).

5.3 Shifts in ECB Communication

We exploit the feature of the ECB statements that topics emerge and disappear over time to investigate whether the transition periods in

[15] According to Reuters, June 22, 2011: "The ECB used the phrase 'strong vigilance' in March before increasing rates in April. It also used the phrase repeatedly during its 2005–07 rate hike cycle, typically one month before it raised rates, although there were exceptions to that rule."

Table 3. Top 10 Terms Describing Topics of the Q&A Section, Ranked by the FREX Score

1 "General Terms"	2 "General Terms"	3 "Vigilance"	4	5 "Liquidity"
one	particular	vigilant	correction	commercial_bank
point	already	vigilance	episode	refinance
way	line	body	labour_productivity	supply
let	present	homework	counter	decrease
time	respect	diagnosis	social_partner	refinance_operation
first	observe	banca_italia	dynamic	bold
second	anchor	erm_ii	economist	exceptional
mean	mention	financial_environment	salary	money-market
come	necessary	invite	counterpart	unlimited
also	credible	favourable	m1	main_refinance_operation
6	**7 "LTRO/OMT"**	**8 "Low Inflation"**	**9 "QE"**	
head	esm	abs	asset_purchase	
restore	ltros	weak	eurozone	
doctrine	access	low_inflation	sustained	
phase	omts	ssm	objective	
ahead	omt	weakness	npls	
peer	compact	supervisor	mostly	
governance	ltro	forward_guidance	path	
advanced_economy	risk_aversion	lending	qe	
commensurate	ela	energy_price	constâncio	
public_finance	contraction	subdue	towards	

Note: The FREX score gives high ranks to terms that are both frequent and exclusive. We do not find a clear interpretation for topics 4 and 6, therefore the topics have no labels.

ECB communication increase market volatility. The market reaction is measured with the VSTOXX index.

The aim is to derive a simple topic-based measure that captures the phases of ECB communication when the message was relatively homogeneous, that is, primarily focused on a single topic. An intuitive approach is to define a summarizing communication measure for each section as a proportion of the topic with the highest probability on a conference day:

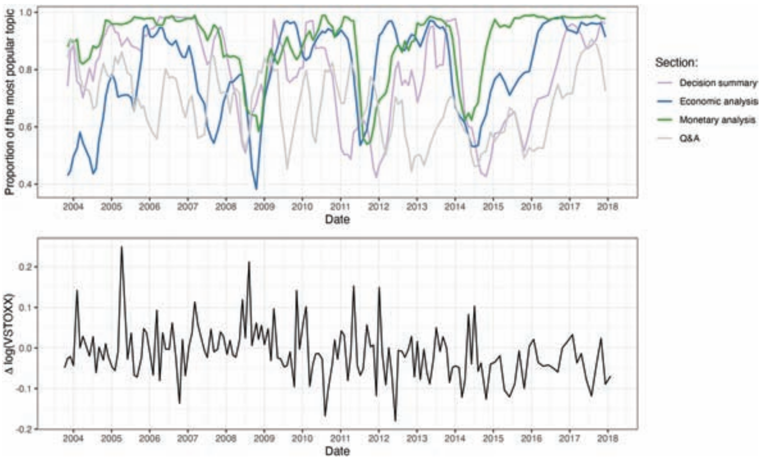
$$I_d = \max_{k \in 1,\ldots,K} (\hat{\theta}_{d,k}). \tag{5}$$

Large values (near one) imply that ECB communication is dominated by a single topic, whereas small values represent a situation where a variety of topics is discussed.

The analysis is constrained to four sections: decision summary, economic analysis, monetary analysis (the latter two referring to the two pillars of the ECB decision making), and the Q&A session (because of its unique clarification role). Because the purpose of the communication measure is to analyze how topics change over time, we ignore topics that constitute a fixed part of discussion. First, we omit the topic on macroeconomic projections, because the words in this section mainly capture the vocabulary used to describe the numerical projections but not their meaning (see Table 2). While the new rounds of quarterly macro projections may affect the ECB narrative, these narrative changes should be reflected in different topics in the other sections. Second, topics 1 and 2 in the Q&A section are omitted, because they place high probability on general terms (corpus-specific stop words). The stop words are frequent, but not content-bearing (see Table 3). In the sections mentioned above, the probability of the dominating topic from the set of remaining topics is then normalized by dividing by the sum of topic probabilities in this set.[16] Figure 6 graphs the communication measures derived from LDA document-topic distributions.

[16] For the economic analysis section the measure is $I_d^{EA} = \max_{k \in \{2,\ldots,6\}} (\hat{\theta}_{d,k}) \times \frac{1}{\Sigma_{i \in \{2,\ldots,6\}} \hat{\theta}_{d,i}}$. For the Q&A section it is $I_d^{QA} = \max_{k \in \{3,\ldots,9\}} (\hat{\theta}_{d,k}) \times \frac{1}{\Sigma_{i \in \{3,\ldots,9\}} \hat{\theta}_{d,i}}$.

Figure 6. Topic-Based Communication Measures for Four Sections of the ECB Press Conference: Decision Summary, Economic Analysis, Monetary Analysis, and Q&A

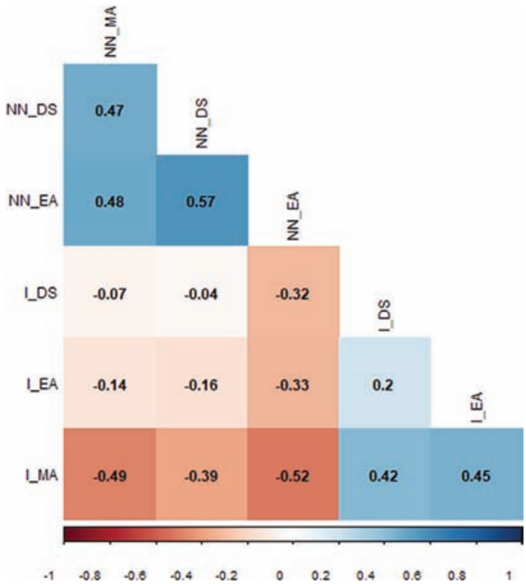


Note: This figure plots the topic-based communication measures, smoothed using a three-point moving-average filter (top panel), and the daily percentage change (close to close) of the VSTOXX on the day of the ECB press conference (bottom panel). Topic-based communication measures are constructed as the probability of the dominant topic in a specific section.

5.3.1 Co-movement of Topic-Based Communication Measures and Tone

The topic-based communication measures in Figure 6 appear to co-move, indicating that the ECB updates the different sections at approximately the same time. The rise of new topics can be linked to the intensification of financial market tensions and changes in policy stance. Updates in topics can also be accompanied by a change in tone. The association between topics and tone is illustrated in Figure 7, which shows the correlation matrix of the net negativity (NN) and topic-based communication measures (I, for sections of the introductory statement; DS—decision summary, EA—economic analysis, MA—monetary analysis). The strongest correlations of the tone-based measures are with the topics of the monetary analysis section. A lower focus on a single topic (or transition to a different topic) in each section is associated with higher net negativity

Figure 7. Correlation Matrix for the Topic-Based Communication Measures (I_{-}) and Net Negativity (NN_{-}) Scores in the Decision Summary (DS), Economic Analysis (EA), and Monetary Analysis (MA) Sections



Note: This figure shows the correlation between the topic-based communication measures and the net negativity scores for each section. The net negativity score is the difference between the fraction of negative and positive words in the section, based on the list of positive and negative words provided in the Loughran and McDonald (2011) dictionary. A negative correlation between I_{-} and NN_{-} indicates that a lower focus on a single topic (or transition to a different topic) in a section is associated with higher net negativity expressed in this section.

expressed in the monetary analysis. Figure 8 gives further insights into topic concentration in the monetary analysis and economic analysis sections. It shows that a change in topics (lower topic concentration) is usually accompanied by a spike in the net negativity score. This observation indicates that the change in narrative in the statements is often accompanied by a change in tone.

5.3.2 The Impact of ECB Communication

We analyze the impact of ECB communication through event-based regressions, where only statement days are considered. The empirical

Figure 8. Topic-Based Communication Measures for the Economic Analysis and the Monetary Analysis Sections, Along with the Respective Net Negativity Scores for Each Section

investigation is complicated by the fact that the ECB press conference always takes place on the same day that a monetary policy decision is announced. To control for the effects of policy actions, we include the absolute surprise component in the ECB monetary policy decision, as in Rosa and Verga (2007) and Ehrmann and Fratzscher (2009).[17]

[17]In this paper we measure the surprise component using the median expected monetary policy rate from the Bloomberg® survey. We also considered the monetary policy surprise as captured by high-frequency interest rate changes in the press release window using the Euro Area Monetary Policy Event-Study Database developed by Altavilla et al. (2019), of which we were subsequently made aware. The results are qualitatively similar and are not reported here but are available from the authors on request. We thank Peter Tillmann for the suggestion to use the alternative surprise measure.

Following Coenen et al. (2017), to account for non-standard policy measures we include a dummy variable for the days when such measures were announced and the absolute change in German two-year government bond yields, which is intended to capture the absolute surprise component in decisions about unconventional monetary policy tools. To control for other macroeconomic news, the surprise component of the U.S. jobless claims releases on Thursdays at 8:30 ET is included, as it coincides with the ECB press conference.[18] Appendix E provides descriptive statistics and correlations.

The event-based regressions are nested in the following equation:

$$\Delta V_t = \alpha + \beta_1|s_t^{MRO}| + \beta_2 D_t^A + \beta_3|r_t^{DE}| + \beta_4|s_t^{JC}| + \beta_5\Delta V_{t-1} + \beta_6 I_t^{DS} + \beta_7 I_t^{EA} + \beta_8 I_t^{MA} + \beta_9 I_t^{QA} + \beta_{10} I_t^{QA} \times D_t^{Draghi} + \beta_{11} NN_t + \varepsilon_t, \qquad (6)$$

where ΔV_t denotes the daily percentage change in the VSTOXX index on the conference day t relative to the previous day; s_t^{MRO} and s_t^{JC} are surprise components of the MRO rate and the U.S. jobless claims, respectively; D_t^A is an indicator for announcements regarding non-standard monetary policy measures; r_t^{DE} is a daily change in German two-year government bond yields; and $I_t^{DS}, I_t^{EA}, I_t^{MA}, I_t^{QA}$ denote the index values that capture changes in communication by section: decision summary, economic analysis, monetary analysis, Q&A. The communication score for the Q&A section is also interacted with an indicator variable for presidency (D_t^{Draghi}). In addition, we control for the net negativity score (NN_t) calculated jointly for the economic analysis and monetary analysis sections (the difference between the fraction of negative and positive words in those sections based on LM dictionary). Table 4 presents the estimation results.[19]

[18] In the sample period there were seven press conferences that took place on Wednesday instead of Thursday. Also, four times a year the ECB/Eurosystem macroeconomic projections are published on the ECB website. Including an indicator variable to control for the timing of these macroeconomic projections does not affect our qualitative conslusions and hence the results are omitted.

[19] The model selection necessarily involved human judgment in balancing multiple criteria (exclusivity, coherence, predictive power). As a robustness check, we construct the communication variable for each section using the output of a topic model with the number of topics greater by one or less by one than in the baseline

Table 4. Regression Results

	Dependent Variable: ΔV_t			
	(1)	**(2)**	**(3)**	**(4)**
$\lvert s_t^{MRO} \rvert$	–0.115	–0.150	–0.137	–0.134
	[0.333]	[0.209]	[0.239]	[0.261]
ΔV_{t-1}	–0.017	0.017	0.071	0.071
	[0.876]	[0.881]	[0.518]	[0.518]
r_t^{DE}	–0.176**	–0.167**	–0.156*	–0.156*
	[0.039}	[0.049]	[0.060]	[0.060]
$\lvert s_t^{JC} \rvert$	0.001	0.0004	0.0003	0.0003
	[0.111]	[0.298]	[0.452]	[0.443]
D_t^A	–0.023	–0.023	–0.014	–0.014
	[0.240]	[0.253]	[0.483]	[0.489]
I_t^{DS}		0.030	0.020	0.020
		[0.374]	[0.538]	[0.547]
		(0.008; 0.043)	(–0.001; 0.033)	(0; 0.033)
I_t^{EA}		–0.031	0.0001	–0.0001
		[0.375]	[0.998]	[0.998]
		(–0.052; –0.015)	(–0.026; 0.014)	(–0.025; 0.014)
I_t^{MA}		–0.096*	–0.116**	–0.121**
		[0.059]	[0.021]	[0.035]
		(–0.119; –0.052)	(–0.133; –0.067)	(–0.14; –0.061)
I_t^{QA}		0.028	0.025	0.024
		[0.483]	[0.517]	[0.531]
		(0.008; 0.045)	(0.005; 0.042)	(0.004; 0.043)
$I_t^{QA} \times D_t^{Draghi}$			–0.050***	–0.051***
			[0.005]	[0.006]
			(–0.053; –0.044)	(–0.053; –0.044)
NN_t				–0.049
				[0.869]
Constant	–0.019**	0.052	0.069	0.074
	[0.020]	[0.262]	[0.133]	[0.181]
Observations	156	156	156	156
Adjusted R^2	0.034	0.058	0.103	0.097

Note: p-values are in brackets and the sampling uncertainty (5th to 95th percentile) is in parentheses. The sampling uncertainty is computed based on 400 draws from the posterior distribution. *p < 0.1; **p < 0.05; ***p < 0.01.

model. We also test for a topic model that strictly dominates other models in terms of both coherence and exclusivity. Table F.1 in Appendix F reports the estimates. The results are qualitatively similar to the baseline specification.

In the regressions, we use the values of the communication measures derived from the matrix of document-topic distributions averaged across 400 draws from a Markov chain. As a result, there is uncertainty arising from the sampling algorithm used to estimate topics. The regression analysis is repeated for each draw to obtain a distribution of the effect, similarly to Hansen, McMahon, and Prat (2017). Table 4 therefore also reports the range of the 5th to 95th percentiles of these distributions.

The major transitions in ECB communication regarding monetary analysis contain incremental information about the ECB monetary policy decisions not already incorporated in market expectations, after controlling for announcements about non-standard monetary policy measures. The uncertainty proxied by the VSTOXX index is on average lower when the ECB sends a homogeneous message by focusing primarily on a single topic than in times of transitions to a different topic.[20] The results in Table 4 show that the implied volatility on the conference day decreases by approximately 1.16 percentage points when the proportion of the most dominant topic in the monetary analysis section increases by 10 percentage points. This effect is significant at the 5 percent level. The results suggest that increased focus on a single topic in the monetary analysis section at any point in time is associated with less market volatility.

Conversely, we also see episodes where the lack of focus (as seen by a period of transition between topics) was associated with higher market volatility. Looking across the four topics in the monetary analysis section over time, the three key dates when the topics shifted permanently are October 2, 2008; October 6, 2011; and May 8, 2014 (see Figure D.2 in Appendix D). The shift in topics is reflected in substantial updates to the statements. In each case, prior to the shift, certain sentences appear repeatedly in the statements, and after the

[20]This finding is reminiscent of that of Ehrmann and Talmi (2020), who show that market volatility decreases when consecutive central bank communications are semantically similar; for the ECB, similarity has increased over time and the effect has similarly strengthened. We emphasize, however, that in this paper we focus on the within-statement coherence and do not explicitly consider cross-statement similarity. We thank an anonymous referee for drawing our attention to this paper.

shift new elements appear or replace the previous sentences and are subsequently repeated.[21]

October 2008 marked a key turning point in the global financial crisis, as the month began with great uncertainty following the September 29 stall where the U.S. House of Representatives voted down the plan put forth by the U.S. Department of Treasury to address the unfolding crisis. After five days of turmoil, the Troubled Assets Relief Program (TARP) was passed on October 3, bringing financial markets around the world back from the brink of collapse. Looking at 10 statements preceding the statement on October 2, 2008, the ECB was repeatedly pointing at prevailing upside risks to price stability at medium to longer-term horizons, the underlying strength of monetary expansion, and temporary factors which may overstate the impact of monetary expansion.[22] There was little evidence that the financial market turbulence since early August had strongly influenced the availability of bank credit in Europe. The ECB in its statements confirmed that the borrowing by non-financial corporations had remained strong. A large drop in the probability of topic 1 occurred at the meeting on October 2, just one day before the TARP passage. The ECB explained that the latest available data (from August) had not embodied the impact of the intensification of the financial market turmoil, and spent more time explaining substitution effects. From the next meeting onward (November 6, 2008) the ECB was pointing at the diminishing impact of upside risks to price stability and an identifiable impact of financial market tensions. The average change in the volatility index rose by 4.8 percentage points, from 1.6 percent to 6.4 percent, on the press conference days during the time when the concentration index fell below 0.6, compared with 10 meetings before October 2008 when the concentration index was above 0.8.

October 2011 marked the point when the European sovereign debt crisis was intensifying and threatening the banking sector. The

[21] These findings are consistent with those of Ehrmann and Talmi (2020), who find an increase in government bond yield volatility when statements change greatly relative to the previous statement and a larger effect the longer the string of similar statements that preceded it.

[22] This is evident in the list of terms under Topic 1 in Table E.2 in Appendix E.

ECB meeting on October 6, 2011 was Jean-Claude Trichet's last meeting before Mario Draghi took the helm. The common parts of the statements preceding the statement on October 6, 2011 were the dampening impact of the steepening yield curve on M3, the stable size of the balance sheets of the banks with the expanding provision of credit to the private sector, and strengthened growth of loans to non-financial corporations. Overall, the monetary expansion was described as moderate and inflationary pressures were contained. The ECB was calling on banks to take appropriate measures to further strengthen their capital. Starting with the statement from November 3, 2011 (the first under Draghi) the ECB introduced a new narrative, focusing on factors related to the intensification of financial market tensions and their negative effects on monetary developments. Prior to the October 6 meeting, the monetary analysis concentration index hovered above 0.95; it then fell below 0.6 and remained at that level until February 9, 2012. The average change in the volatility index on the press conference days during this time (October 2011–February 2012) rose by 2.6 percentage points from –3.6 percent to –1 percent. In comparison, during the 10 meetings before October 2011, the concentration index was above 0.9 (excluding the meeting on August 4, 2011, when the exceptionally high volatility resulted from overall anxiety in both Europe and the United States about deepening economic problems and the uncertainty over the ECB purchasing bonds of Italy and Spain).

May 2014 marked the last month that deposit rates were positive. In the statements preceding the statement on May 8, 2014, the ECB was acknowledging subdued monetary and credit dynamics, reflecting the state of the business cycle, heightened credit risk, and the ongoing adjustment of financial and non-financial sector balance sheets. Up to May 2014, the ECB was repeatedly expressing its concerns about the transmission of monetary policy to the financing conditions in euro-area countries, the fragmentation of euro-area credit markets, and the resilience of the banking sector. Starting in June, a comprehensive package of non-standard policy measures was gradually introduced in order to improve credit conditions. Prior to the May 8 meeting, the monetary analysis concentration index hovered above 0.9; it then fell below 0.7 and remained at that level until November 2014. The average change in the volatility index rose by

0.7 percentage points, from –2.7 percent to –2 percent, on the press conference days during the time when the concentration index was below 0.7, compared with 10 meetings before May 2014, when the concentration index was above 0.9.

The changing composition of the decision summary is not significant. This is expected, as any effect of this section should be already subsumed into the effect of announcements about the policy rate and non-standard monetary policy measures. Although stock market volatility was on average higher under the leadership of Trichet than under Draghi, the changing composition of the Q&A session is on its own not informative for the market. This result agrees with the findings of Ehrmann and Fratzscher (2009), who analyze the reaction of three-month EURIBOR (euro-area interbank offered rate) futures and find that the Q&A session does not systematically add information beyond that contained in the introductory statement, suggesting that in most cases the introductory statement provides sufficient clarity.

5.3.3 Consideration of Other Market Variables

Because we are interested in the influence of ECB communication on market uncertainty, we have to this point focused on the VSTOXX, as it is designed to reflect market participants' expectations of near-term volatility and is often used as a proxy for uncertainty in the euro area. Our interest in uncertainty is primarily motivated by the observed pattern in the press conferences, namely that at distinct points in time the ECB introduces substantial updates to the statements. Specifically, we are interested in how market participants perceive and digest these new narratives and whether these changes are abrupt and visible to market participants in a way that increases their uncertainty. New content in the statements can raise uncertainty if it is unexpected or insufficient (i.e., it does not contain sufficient explanations of the economic or monetary situation that motivated the change), compared with the content that was repeated in the statements before the major update and was familiar to market participants. It is perhaps not surprising that the monetary analysis section is where we find a significant effect on market volatility, as this section usually discusses monetary and financial conditions in the euro area. In this subsection, we consider whether there are similar effects for other market variables.

We estimate the event-based regressions using the daily percentage change in a number of other market variables, namely the Euro Stoxx Index, the Euro 50 Stoxx (on which the VSTOXX is based), the DAX, and the CAC 40 (see Table 5). As expected, the communication variables also affect the euro-zone market indices themselves, particularly the monetary analysis section index and the Q&A interaction with the indicator variable for Draghi's tenure. The positive and significant coefficients are suggestive of the idea that a higher concentration in one topic is good for stock markets.[23] Consistent with our intuition, the effect on the FTSE 100 is significant but weaker; a similar result is obtained using changes in the Nikkei, but there it is the economic analysis index rather than the monetary analysis one that appears to have a modest effect.

In addition, we verified that the communication variables do not seem to affect an unrelated volatility indicator by performing the same regression as in the paper using the Japan VIX index (VXJ), the implied volatility index of the Nikkei. As suspected, the communication indices do not affect the Japan VIX, strengthening the argument that the communication indices matter primarily for euro-area uncertainty.

6. Conclusions

In this paper we empirically search for the main communication patterns in the ECB press conferences and analyze how shifts in those patterns affect a key stock market volatility index on the Governing Council meeting days. Using a generative model for text, we decompose each section of the press conference into a set of coherent and exclusive topics. This approach has the potential to reveal previously understudied dimensions in the transcripts.

[23]We thank an anonymous referee for raising a series of questions that led us to the results in this subsection and for suggesting this interpretation. We are intrigued by this possibility but note that at this point it is a somewhat speculative conclusion and warrants a more comprehensive investigation.

Table 5. Regression Results Using the Euro Stoxx Index (SXXE), the Euro 50 Stoxx (SX5E), the DAX, the CAC 40, the FTSE (UKX), the Nikkei 225 (NKY), and the Japan VIX (VXJ)

	Dependent Variable						
	$\Delta SXXE$ (1)	**$\Delta SX5E$ (2)**	**ΔDAX (3)**	**ΔCAC (4)**	**ΔUKX (5)**	**ΔNKY (6)**	**ΔVXJ (7)**
$\lvert s_t^{MRO} \rvert$	0.091***	0.091***	0.105***	0.096***	0.053**	0.015	–0.117
	[0.001]	[0.001]	[0.0003]	[0.0005]	[0.019]	[0.572]	[0.188]
Lagged Dependent Var.	0.192*	0.192*	0.119	0.183*	0.132	–0.070	0.028
	[0.051]	[0.051]	[0.227]	[0.058]	[0.124]	[0.445]	[0.730]
r_t^{DE}	0.042**	0.042**	0.035*	0.045**	0.055***	0.059***	–0.087
	[0.020]	[0.020]	[0.064]	[0.018]	[0.0005]	[0.002]	[0.160]
$[s_t^{JC}]$	–0.0001	–0.0001	–0.0001	–0.0001	–0.0001	–0.0002**	0.0003
	[0.379]	[0.379]	[0.262]	[0.389]	[0.414]	[0.039]	[0.356]
D_t^A	–0.003	–0.003	–0.006	–0.004	–0.003	0.007	–0.008
	[0.449]	[0.449]	[0.212]	[0.411]	[0.392]	[0.107]	[0.594]
I_t^{DS}	–0.001	–0.001	–0.003	–0.0003	–0.002	–0.002	0.004
	[0.871]	[0.871]	[0.669]	[0.972]	[0.700]	[0.823]	[0.861]
I_t^{EA}	–0.007	–0.007	–0.006	–0.008	–0.001	0.014*	–0.017
	[0.345]	[0.345]	[0.436]	[0.332]	[0.826]	[0.073]	[0.538]
I_t^{MA}	0.022*	0.022*	0.022*	0.022*	0.015	0.012	–0.006
	[0.077]	[0.077]	[0.087]	[0.086]	[0.147]	[0.325]	[0.881]
I_t^{QA}	0.001	0.001	0.005	0.001	0.005	0.013	–0.030
	[0.950]	[0.950]	[0.601]	[0.903]	[0.494]	[0.139]	[0.306]
$I_t^{QA} \times D_t^{Draghi}$	0.007*	0.007*	0.008**	0.007*	0.004	–0.002	0.013
	[0.059]	[0.059]	[0.046]	[0.095]	[0.242]	[0.587]	[0.312]
NN_t	–0.049	–0.049	–0.034	–0.053	–0.010	0.049	0.446**
	[0.437]	[0.437]	[0.618]	[0.423]	[0.853]	[0.469]	[0.045]
Constant	–0.014	–0.014	–0.017	–0.015	–0.015	–0.027**	0.024
	[0.238]	[0.238]	[0.193]	[0.228]	[0.161]	[0.031]	[0.557]
Observations	156	156	156	156	156	156	156
Adjusted R^2	0.153	0.153	0.134	0.146	0.116	0.093	0.030

Note: *p < 0.1; **p < 0.05; ***p < 0.01.

The results show that similar ECB press conferences are clustered in time even though nothing in our approach imposes such clustering. The main topics surge, die out over time, and rarely reappear in the analyzed sample period, 2004–18. Market volatility increases when the ECB substantially updates its wording in the monetary analysis section, as compared with keeping it rather static relative to the previous period, controlling for the unexpected components in standard and non-standard monetary policy measures. The revisions to the ECB narrative in general accompany the changes in policy direction, but the results suggest that shifts in ECB communication introduce incremental volatility above and beyond that created by a change in policy stance. Although in our paper we do not consider similarity of consecutive statements explicitly, but rather within-statement homogeneity, our findings corroborate the results of Ehrmann and Talmi (2020), who show that market volatility increases when substantial changes occur following a sequence of similar statements.

The main contribution to the current literature that applies computational linguistics tools to analyze central bank communication is a new topic-based communication measure that does not depend on subjective interpretations of topics. Furthermore, the topic model is estimated using a fully Bayesian approach rather than making ad hoc choices about model hyperparameters. Estimating hyperparameters reveals specific features of modeled transcripts. Although we use a “bag-of-words” algorithm that does not incorporate document ordering, the results demonstrate the ability of Latent Dirichlet Allocation to identify series of documents that change the current discourse. We emphasize, however, that the model does not fully eliminate the need to read statements in order to understand what they are about or to guide modeling decisions. Nonetheless it sheds light on how a central bank introduces new policy narratives and to what extent markets are sensitive to those transitions.

Appendix A. Estimation Details

This appendix first derives the posterior distribution of latent variables given the observed words, in the context of Latent Dirichlet Allocation. It then provides a discussion of choices in model specification and an overview of two popular strategies to approximate the posterior distributions in LDA: Markov chain Monte Carlo (MCMC) methods—in particular, collapsed Gibbs sampling (Griffiths and Steyvers 2004). The Metropolis-within-Gibbs sampling approach, which extends upon collapsed Gibbs sampling, is then presented as the preferred estimation method.

A.1 Latent Dirichlet Allocation

Latent Dirichlet Allocation (LDA), introduced by Blei, Ng, and Jordan (2003), is a mixed membership model for text. The basic idea is that observations (words) are grouped into documents and each of these groups (documents) is modeled with a mixture of distributions. The components of the mixture are topics, which are multinomial probability distributions over fixed vocabulary. The topics are shared across all documents (each document is built from the same components), but the proportions of topics in documents vary.

To formalize this idea, let D be the number of documents, N_d is the number of words in document d, V is the number of distinct words (vocabulary size) in a collection of documents (a corpus), K is the number of topics. The corpus is denoted as $\mathcal{W} = \{\boldsymbol{w}^{(1)}, \ldots, \boldsymbol{w}^{(D)}\}$, where $\boldsymbol{w}^{(d)} = \{w_i^{(d)}\}_{i=1}^{N_d}$ is the collection of words in document d and $w_i^{(d)} \in \{1 : V\}$ is the i-th word in document d. Let $\mathcal{Z} = \{\boldsymbol{z}^{(1)}, \ldots, \boldsymbol{z}^{(D)}\}$ denote topic assignments, where $\boldsymbol{z}^{(d)} = \{z_i^{(d)}\}_{i=1}^{N_d}$ and $z_i^{(d)} \in \{1 : K\}$ is a topic assignment for word $w_i^{(d)}$.[24] Let $\boldsymbol{\Theta}$ be a $D \times K$ matrix of topic proportions in documents and $\boldsymbol{\Phi}$ is a $K \times V$ matrix of word probabilities. $\boldsymbol{\theta}_d$ is a K-dimensional vector of topic proportions in document d where $\theta_{d,1}, \cdots, \theta_{d,K}$ are positive random variables that sum to 1. Similarly, topic k, $\boldsymbol{\phi}_k$, is a V-dimensional vector where $\phi_{k,1}, \cdots, \phi_{k,V}$ are positive random

[24] Blei, Ng, and Jordan (2003) define $z_i^{(d)}$ and $w_i^{(d)}$ as unit vectors of length K and V, respectively, that contain a single 1 in the k-th or v-th element, respectively, and zero otherwise, $k = 1, \ldots, K$ and $v = 1, \ldots, V$. Such defined multidimensional variables have the multinomial distribution. In general, a multinomial vector contains counts that sum to n. Because in our case $n = 1$, $z_i^{(d)}$ and $w_i^{(d)}$ can be defined as unidimensional variables with $p(z_i^{(d)}|\boldsymbol{\theta}_d) = \prod_{k=1}^{K} \theta_{d,k}^{I(z_i^{(d)}=k)}$ and $p(w_i^{(d)}|\boldsymbol{\phi}_k) = \prod_{v=1}^{V} \phi_{k,v}^{I(w_i^{(d)}=v)}$.

variables that sum to 1. It is assumed that K and V are known and fixed. The generative process for text is as follows (Blei, Ng, and Jordan 2003):

(i) For document $d = 1, \ldots, D$ choose the topic proportions $\boldsymbol{\theta}_d \backsim Dirichlet(\boldsymbol{\alpha})$, where $\boldsymbol{\alpha}$ is a K-dimensional hyperparameter.

(ii) For topic $k = 1, \ldots, K$ choose the word distribution $\boldsymbol{\phi}_k \backsim Dirichlet(\boldsymbol{\beta})$, where $\boldsymbol{\beta}$ is a V-dimensional hyperparameter.

(iii) For document $d = 1, \ldots, D$: for word $i = 1, \ldots, N_d$:
   (a) choose the topic $z_i^{(d)} \backsim Multinomial(\boldsymbol{\theta}_d)$;
   (b) choose the word $w_i^{(d)} \backsim Multinomial(\boldsymbol{\phi}_{zij})$.

We only observe the set of documents, $\mathcal{W}$. The underlying topic assignments $\mathcal{Z}$, word probabilities $\boldsymbol{\Phi}$, and topic proportions in documents $\boldsymbol{\Theta}$ are latent; $\boldsymbol{\alpha}$, $\boldsymbol{\beta}$ are concentration hyperparameters that are selected in advance.

The central inferential problem in LDA is to determine the posterior distribution of topic proportions in documents ($\boldsymbol{\Theta}$), word proportions in topics ($\boldsymbol{\Phi}$), and word-topic assignments ($\mathcal{Z}$).

The joint posterior density is

$$p(\boldsymbol{\Phi}, \boldsymbol{\Theta}, \mathcal{Z}|\mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}) = \frac{p(\boldsymbol{\Phi}, \boldsymbol{\Theta}, \mathcal{Z}|\mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta})}{p(\mathcal{W}|\boldsymbol{\alpha}, \boldsymbol{\beta})} \propto p(\mathcal{W}, \mathcal{Z}|\boldsymbol{\Phi}, \boldsymbol{\Theta}, \boldsymbol{\alpha}, \boldsymbol{\beta}) p(\boldsymbol{\Theta}|\boldsymbol{\alpha}) p(\boldsymbol{\Phi}|\boldsymbol{\beta}). \tag{A.1}$$

The following priors are assumed for model parameters $\boldsymbol{\Phi}$ and $\boldsymbol{\Theta}$:

$$p(\boldsymbol{\Theta}|\boldsymbol{\alpha}) = \prod_{d=1}^{D} p(\boldsymbol{\theta_d}|\boldsymbol{\alpha}) = \prod_{d=1}^{D} Dirichlet(\boldsymbol{\theta_d}; \boldsymbol{\alpha}), \tag{A.2}$$

$$p(\boldsymbol{\Phi}|\boldsymbol{\beta}) = \prod_{k=1}^{K} p(\boldsymbol{\phi_k}|\boldsymbol{\beta}) = \prod_{k=1}^{K} Dirichlet(\boldsymbol{\phi_k}; \boldsymbol{\beta}). \tag{A.3}$$

To derive the joint likelihood function of $\mathcal{W}$ and $\mathcal{Z}$, we first consider the density of data $\mathcal{W}$ given topic assignments $\mathcal{Z}$ and model parameters:

$$p(\mathcal{W}|\mathcal{Z}, \mathbf{\Phi}, \mathbf{\Theta}, \boldsymbol{\alpha}, \boldsymbol{\beta}) = p(\mathcal{W}|\mathcal{Z}, \mathbf{\Phi}) = \prod_{d=1}^{D}\prod_{i=1}^{N_d} p(w_i^{(d)}|z_i^{(d)}, \mathbf{\Phi}). \quad \text{(A.4)}$$

The probability $p(w_i^{(d)}|z_i^{(d)}, \mathbf{\Phi}) = \phi_{z_i^{(d)}, w_i^{(d)}}$ is an element of matrix $\mathbf{\Phi}$ located in $z_i^{(d)}$-th row and $w_i^{(d)}$-th column. The density function of $\mathcal{Z}$ is

$$p(\mathcal{Z}|\mathbf{\Phi}, \mathbf{\Theta}, \boldsymbol{\alpha}, \boldsymbol{\beta}) = p(\mathcal{Z}|\mathbf{\Theta}) = \prod_{d=1}^{D}\prod_{i=1}^{N_d} p(z_i^{(d)}|\boldsymbol{\theta}_d). \quad \text{(A.5)}$$

The probability $p(z_i^{(d)}|\mathbf{\Theta}) = \theta_{d, z_i^{(d)}}$. The joint density of data and latent variable $\mathcal{Z}$ (the complete data likelihood function) is

$$p(\mathcal{W}, \mathcal{Z}|\mathbf{\Phi}, \mathbf{\Theta}, \boldsymbol{\alpha}, \boldsymbol{\beta}) = \prod_{d=1}^{D}\prod_{i=1}^{N_d} p(w_i^{(d)}|z_i^{(d)}, \mathbf{\Phi}) p(z_i^{(d)}|\boldsymbol{\theta}_d). \quad \text{(A.6)}$$

The posterior distribution is proportional to the complete data likelihood function times the prior:

$$p(\mathbf{\Phi}, \mathbf{\Theta}, \mathcal{Z}|\mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}) \propto \prod_{d=1}^{D} \underbrace{p(\boldsymbol{\theta}_d|\boldsymbol{\alpha})}_{Dirichlet} \prod_{k=1}^{K} \underbrace{p(\boldsymbol{\phi}_k|\boldsymbol{\beta})}_{Dirichlet} \left(\prod_{d=1}^{D}\prod_{i=1}^{N_d} \underbrace{p(w_i^{(d)}|z_i^{(d)}, \mathbf{\Phi})}_{Multinomial} \underbrace{p(z_i^{(d)}|\boldsymbol{\theta}_d)}_{Multinomial} \right). \quad \text{(A.7)}$$

A.2 Choices in Model Specification and Estimation

Implementation of LDA involves important model specification and selection decisions. In particular, the estimation results vary according to the number of topics (K) and hyperparameter settings ($\boldsymbol{\alpha}$, $\boldsymbol{\beta}$). This section discusses the decisions made with respect to both of these dimensions.

A.2.1 The Number of Topics

For the number of topics, there is no "right" answer to this choice; rather, the choice depends on interpretability and goals of the analysis (Grimmer and Stewart 2013; Roberts et al. 2014). DiMaggio, Nag, and Blei (2013) note that "the test of the model as a whole is its ability to identify a number of substantively meaningful and analytically useful topics, not its success in optimizing across all topics."

Typically, in choosing the number of topics, there is a trade-off between predictive accuracy of the model and topic interpretability (Chang et al. 2009). Metrics of predictive performance, such as held-out likelihood or perplexity, are conventionally used to assess model quality (Blei, Ng, and Jordan 2003; Wallach et al. 2009). Perplexity is defined as the inverse of the geometric mean per-word likelihood of the test data.

To evaluate the model fit, one can ask how well the model predicts words in a testing set. Noisy topics will fail to replicate test data, resulting in low held-out likelihood and high perplexity. However, the predictive metrics have limitations. Usually fine-grained, highly specific topics yield the best model fit, but they are not easy to interpret or to generalize (Boyd-Graber, Mimno, and Newman 2014). Furthermore, predicting the content of the preprocessed text is rarely the objective of research in political, economic, or social sciences, especially since the preprocessing steps substantially simplify the original documents (Grimmer and Stewart 2013).

One strand of literature focuses on evaluating topic quality from the perspective of interpretability using automated measures that correlate well with human ratings and thus are better able to serve real-world objectives such as discerning meaningful themes or augmenting the subsequent causal analysis with human-interpretable textual information.

Topics are usually interpreted based on top words with the highest probability (Blei, Ng, and Jordan 2003; Griffiths and Steyvers 2004). Roberts et al. (2014) argue that a semantically interpretable topic has two qualities: (i) it is coherent—the highest probability words for the topic tend to co-occur within documents, and (ii) it is exclusive—the words that have high probability under one topic have low probabilities under other topics. The metrics of coherence and exclusivity that we use for the model selection are described in the paper.

A.2.2 The Concentration Hyperparameters

The concentration hyperparameters determine the amount of smoothing or sparsity of the topic-word and the document-topic distributions. The parameter $\boldsymbol{\alpha}$ informs the model about the concentration of topics within the document. Low α_k (less than 1) means that a document is more focused (i.e., a single topic dominates); high α_k (greater than 1) means that discussion is less focused and several topics occur with similar intensity in the document. Similarly, $\boldsymbol{\beta}$ informs about the concentration of words within a topic. Low beta means a few words are characteristic of the topic. A large $\boldsymbol{\beta}$ implies more uniform topic-word probabilities and leads to similar topics.

Several studies demonstrate that selection of the hyperparameters has a strong influence on both prior and posterior distributions of $\boldsymbol{\Theta}$ and $\boldsymbol{\Phi}$ (Asuncion et al. 2009; Wallach, Mimno, and McCallum 2009; George and Doss 2018). Implementations of LDA typically assume that Dirichlet priors are symmetric ($\beta_1 = \cdots = \beta_V = \beta$ and $\alpha_1 = \cdots = \alpha_K = \alpha$). It is expected that $\beta < 1$ so that many words have low probabilities in a topic.

Following the recommendation of Wallach, Mimno, and McCallum (2009), we implement a combination of priors which is found to be superior: an asymmetric Dirichlet prior over $\boldsymbol{\Theta}$ and a symmetric Dirichlet prior over $\boldsymbol{\Phi}$. First, an asymmetric Dirichlet prior over the document-topic distributions allows some topics to be more likely. These topics may place high probability on words that appear more frequently than other words in every document. Second, asymmetry increases stability of the results as the number of topics increases: if additional topics are redundant, they will be seldom used.

Another decision point is determining the values for hyperparameters. An ad hoc specification of the hyperparameters dominates in the economic literature. Griffiths and Steyvers (2004) provide the most widely applied recommendation: $\alpha = \frac{50}{K}$, $\beta = 0.1$ (Tirunillai and Tellis 2014; Hansen and McMahon 2016; Fligstein, Brundage, and Schultz 2017; Hansen, McMahon, and Prat 2017). This choice is not based on any particular principle.

In this paper we infer the values of concentration parameters in a fully Bayesian setting by placing proper prior distributions on $\boldsymbol{\alpha}$

and $\boldsymbol{\beta}$ (Jacobs, Donkers, and Fok 2016).[25] We use Metropolis-within-Gibbs sampling, which extends upon collapsed Gibbs sampling, for estimation.

A.3 Collapsed Gibbs Sampling

The Gibbs sampler is one technique to produce a sample from the posterior distribution by sequentially drawing samples of a random variable from its conditional distribution given the current values of all other variables. Application of the standard Gibbs algorithm would consider the following sample scheme:

(i) Sample $\boldsymbol{\phi}_k | \boldsymbol{\Phi}_{-k}, \boldsymbol{\Theta}, \mathcal{Z}, \mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}$ for $k = 1, \ldots, K$

(ii) Sample $\boldsymbol{\theta}_d | \boldsymbol{\Phi}, \boldsymbol{\Theta}_{-d}, \mathcal{Z}, \mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}$ for $d = 1, \ldots, D$

(iii) Sample $z_i^{(d)} | z_{-i}^{(d)}, \mathcal{Z}^{(-d)}, \boldsymbol{\Theta}, \boldsymbol{\Phi}, \mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}$ for $d = 1, \ldots, D$; $i = 1, \ldots, N_d$,

where the notation $-d$, $-i$, $-k$ refers to all elements except the d^{th}, i^{th}, and k^{th}, respectively. As iterations continue, the sample values converge to the target posterior distribution in Equation (1) in the paper. The Gibbs sampler is inefficient, because $\boldsymbol{\Theta}$ and $\boldsymbol{\Phi}$ strongly depend on topic assignments $\mathcal{Z}$ and the chain is highly autocorrelated. The classical procedure can be improved upon using the conjugacy of the Dirichlet distribution and the multinomial distribution. Parameters $\boldsymbol{\Theta}$ and $\boldsymbol{\Phi}$ are integrated out from the full conditional posterior distribution for $z_i^{(d)}$. The collapsed Gibbs sampler considers simulating

$$z_i^{(d)} | z_{-i}^{(d)}, \mathcal{Z}^{(-d)}, \mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta} \text{ for } d = 1, \ldots, D; i = 1, \ldots, N_d. \qquad \text{(A.8)}$$

To derive the sampling distribution, let $c_{k,d,v} = \sum_{i=1}^{N_d} I(z_i^{(d)} = k, w_i^{(d)} = v)$ denote the number of words of type v assigned to

[25] Another principled way of setting hyperparameters is iterating between Gibbs sampling and a gradient-based optimization for hyperparameters (Wallach 2006) or finding the hyperparameters by grid search (Asuncion et al. 2009).

topic k in document d. In the following, an asterisk means that the corresponding index is summed out, that is

$$c_{k,*,v} = \sum_{d=1}^{D} c_{k,d,v};\ c_{k,d,*} = \sum_{v=1}^{V} c_{k,d,v};\ c_{k,*,*} = \sum_{d=1}^{D}\sum_{v=1}^{V} c_{k,d,v}. \quad \text{(A.9)}$$

As $z_i^{(d)}$ takes only K different values, the sampling distribution is multinomial with probabilities (Griffiths and Steyvers 2004):

$$p(z_i^{(d)}|z_{-i}^{(d)}, \mathcal{Z}^{(-d)}, \mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta} \propto \frac{\left(c^{-(d,i)}_{z_i^{(d)},d,*} + \alpha_{z_i^{(d)}}\right)}{\left(\sum_{k=1}^{K} c^{-(d,i)}_{k,d,*} + \alpha_k\right)} \times \frac{\left(c^{-(d,i)}_{z_i^{(d)},*,w_i^{(d)}} + \beta_{w_i^{(d)}}\right)}{\left(\sum_{v=1}^{V} c^{-(d,i)}_{z_i^{(d)},*,v} + \beta_v\right)}, \quad \text{(A.10)}$$

where $c^{-(d,i)}$ denotes a count that does not include word i in document d. See the next subsection for the derivation.

For a single draw we can estimate $\mathbf{\Phi}$, $\mathbf{\Theta}$ from the counts:

$$\theta_{d,k} = \frac{\alpha_k + c_{k,d,*}}{\sum_{k=1}^{K}(\alpha_k + c_{k,d,*})};\ \phi_{k,v} = \frac{\beta_v + c_{k,*,v}}{\sum_{v=1}^{V}(\beta_v + c_{k,*,v})}. \quad \text{(A.11)}$$

Posterior mean estimates are obtained by averaging over the draws. However, the posterior inference is complicated by a label switching problem (Stephens 2000). The problem emerges because the complete data likelihood (A.6) is invariant to permutations of the topics' labels (there are $K!$ permutations), hence the posterior will inherit the invariance of the likelihood if priors are symmetric. Various relabeling algorithms can be applied to undo label switching before averaging over the draws. Many off-the-shelf solutions provide posterior estimates based on a single iteration of Gibbs sampling. For example, `R` package `lda` (Chang 2015) uses the state at the last iteration of Gibbs sampling and `R` package `topicmodels` (Hornik and Grun 2011) by default returns the sample with the highest posterior likelihood.

A.4 Full Conditional Distribution of Topic Assignments

This section presents the derivation of the full conditional distribution of topic assignments $z_i^{(d)}$ for word i in document d, which is required to construct a Gibbs sampler.

$$
\begin{aligned}
&p(z_i^{(d)}|z_{-i}^{(d)}, \mathcal{Z}^{(-d)}, \mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}) \propto p(\mathcal{Z}, \mathcal{W}|\boldsymbol{\alpha}, \boldsymbol{\beta}) \\
&\quad = \int\int p(\mathcal{Z}, \mathcal{W}, \boldsymbol{\Theta}, \boldsymbol{\Phi}|\boldsymbol{\alpha}, \boldsymbol{\beta}) d\boldsymbol{\Theta} d\boldsymbol{\Phi} \\
&\quad = \int \prod_{d=1}^{D} \left(p(\boldsymbol{\theta}_d|\boldsymbol{\alpha}) \prod_{i=1}^{N_d} p(z_i^{(d)}|\boldsymbol{\theta}_d) \right) d\boldsymbol{\Theta} \\
&\qquad \times \int \prod_{k=1}^{K} p(\boldsymbol{\phi}_k|\boldsymbol{\beta}) \prod_{d=1}^{D} \prod_{i=1}^{N_d} p(w_i^{(d)}|z_i^{(d)}, \boldsymbol{\Phi}) d\boldsymbol{\Phi} \\
&\quad = \prod_{d=1}^{D} \int p(\boldsymbol{\theta}_d|\boldsymbol{\alpha}) \prod_{i=1}^{N_d} \prod_{k=1}^{K} \theta_{d_k}^{I(z_i^{(d)}=k)} d\boldsymbol{\theta}_d \\
&\quad \times \prod_{k=1}^{K} \int p(\boldsymbol{\phi}_k|\boldsymbol{\beta}) \prod_{d=1}^{D} \prod_{i=1}^{N_d} \prod_{v=1}^{V} \phi_{k,v}^{I(z_i^{(d)}=k, w_i^{(d)}=v)} d\boldsymbol{\phi}_k \\
&\quad \propto \int p(\boldsymbol{\theta}_d|\boldsymbol{\alpha}) \prod_{k=1}^{K} \theta_{d,k}^{c_{k,d,*}} d\boldsymbol{\theta}_d \times \prod_{k=1}^{K} \int p(\boldsymbol{\phi}_k|\boldsymbol{\beta}) \prod_{v=1}^{V} \phi_{k,v}^{c_{k,*,v}} d\boldsymbol{\phi}_k \\
&\quad = \int \frac{\Gamma(\sum_{k=1}^{K} \alpha_k)}{\prod_{k=1}^{K} \Gamma(\alpha_k)} \prod_{k=1}^{K} \theta_{d,k}^{c_{k,d,*}+\alpha_k-1} d\boldsymbol{\theta}_d \\
&\qquad \times \prod_{k=1}^{K} \int \frac{\Gamma(\sum_{v=1}^{V} \beta_v)}{\prod_{v=1}^{V} \Gamma(\beta_v)} \prod_{v=1}^{V} \phi_{k,v}^{c_{k,*,v}+\beta_v-1} d\boldsymbol{\phi}_k \\
&\quad \propto \frac{\prod_{k=1}^{K} \Gamma(c_{k,d,*}+\alpha_k)}{\Gamma(\sum_{k=1}^{K} c_{k,d,*}+\alpha_k)} \underbrace{\int \frac{\Gamma(\sum_{k=1}^{K} c_{k,d,*}+\alpha_k)}{\prod_{k=1}^{K} \Gamma(c_{k,d,*}+\alpha_k)} \prod_{k=1}^{K} \theta_{d,k}^{c_{k,d,*}+\alpha_k-1} d\boldsymbol{\theta}_d}_{=1,\text{ integrating over the entire support of Dirichlet}} \\
&\qquad \times \prod_{k=1}^{K} \frac{\prod_{v=1}^{V} \Gamma(c_{k,*,v}+\beta_v)}{\Gamma(\sum_{v=1}^{V} c_{k,*,v}+\beta_v)} \\
&\qquad \times \underbrace{\int \frac{\Gamma(\sum_{v=1}^{V} c_{k,*,v}+\beta_v)}{\prod_{v=1}^{V} \Gamma(c_{k,*,v}+\beta_v)} \prod_{v=1}^{V} \phi_{k,v}^{c_{k,*,v}+\beta_v-1} d\boldsymbol{\phi}_k}_{=1,\text{ integrating over the entire support of Dirichlet}}
\end{aligned}
$$

$$= \frac{\prod_{k=1}^{K}\Gamma(c_{k,d,*}+\alpha_k)}{\Gamma(\sum_{k=1}^{K}c_{k,d,*}+\alpha_k)} \times \prod_{k=1}^{K}\frac{\prod_{v=1}^{V}\Gamma(c_{k,*,v}+\beta_v)}{\Gamma(\sum_{v=1}^{V}c_{k,*,v}+\beta_v)}$$

$$\propto \frac{\prod_{k=1}^{K}\Gamma(c_{k,d,*}+\alpha_k)}{\Gamma(\sum_{k=1}^{K}c_{k,d,*}+\alpha_k)} \times \prod_{k=1}^{K}\frac{\Gamma(c_{k,*,w}+\beta_{w_i^{(d)}})}{\Gamma(\sum_{v=1}^{V}c_{k,*,v}+\beta_v)}. \qquad \text{(A.12)}$$

The next step is to separate terms, which depend on the current term (d,i). This involves splitting the counts into the part that does not count the current position and the part counting only the current position. We also use that $\Gamma(x+1) = x\Gamma(x)$.

$$p(z_i^{(d)}|z_{-i}^{(d)}, \mathcal{Z}^{(-d)}, \mathcal{W}, \boldsymbol{\alpha}, \boldsymbol{\beta}$$

$$\propto \frac{\prod_{k=1;k\neq z_i^{(d)}}^{K}\Gamma\left(c_{k,d,*}^{-(d,i)}+\alpha_k\right) \times \Gamma\left(c_{z_i^{(d)},d,*}^{-(d,i)}+\alpha_{z_i^{(d)}}+1\right)}{\Gamma\left(1+\sum_{k=1}^{K}c_{k,d,*}^{-(d,i)}+\alpha_k\right)}$$

$$\times \prod_{k=1;k\neq z_i^{(d)}}^{K}\frac{\Gamma\left(c_{k,*,w_i^{(d)}}^{-(d,i)}+\beta_{w_i^{(d)}}\right)}{\Gamma\left(\sum_{v=1}^{V}c_{k,*,v}^{-(d,i)}+\beta_v\right)} \times \frac{\Gamma\left(1+c_{z_i^{(d)},*,w_i^{(d)}}^{-(d,i)}+\beta_{w_i^{(d)}}\right)}{\Gamma\left(1+\sum_{v=1}^{V}c_{z_i^{(d)},*,v}^{-(d,i)}+\beta_v\right)}$$

$$\propto \frac{\prod_{k=1}^{K}\Gamma\left(c_{k,d,*}^{-(d,i)}+\alpha_k\right) \times \left(c_{z_i^{(d)},d,*}^{-(d,i)}+\alpha_{z_i^{(d)}}\right)}{\Gamma\left(\sum_{k=1}^{K}c_{k,d,*}^{-(d,i)}+\alpha_k\right) \times \left(\sum_{k=1}^{K}c_{k,d,*}^{-(d,i)}+\alpha_k\right)}$$

$$\times \prod_{k=1}^{K}\frac{\Gamma\left(c_{k,*,w_i^{(d)}}^{-(d,i)}+\beta_{w_i^{(d)}}\right)}{\Gamma\left(\sum_{v=1}^{V}c_{k,*,v}^{-(d,i)}+\beta_v\right)} \times \frac{\left(c_{z_i^{(d)},*,w_i^{(d)}}^{-(d,i)}+\beta_{w_i^{(d)}}\right)}{\left(\sum_{v=1}^{V}c_{z_i^{(d)},*,v}^{-(d,i)}+\beta_v\right)}$$

$$\propto \frac{\left(c_{z_i^{(d)},d,*}^{-(d,i)}+\alpha_{z_i^{(d)}}\right)}{\left(\sum_{k=1}^{K}c_{k,d,*}^{-(d,i)}+\alpha_k\right)} \times \frac{\left(c_{z_i^{(d)},*,w_i^{(d)}}^{-(d,i)}+\beta_{w_i^{(d)}}\right)}{\left(\sum_{v=1}^{V}c_{z_i^{(d)},*,v}^{-(d,i)}+\beta_v\right)}$$

$$\propto \left(c_{z_i^{(d)},d,*}^{-(d,i)}+\alpha_{z_i^{(d)}}\right) \times \frac{\left(c_{z_i^{(d)},*,w_i^{(d)}}^{-(d,i)}+\beta_{w_i^{(d)}}\right)}{\left(\sum_{v=1}^{V}c_{z_i^{(d)},*,v}^{-(d,i)}+\beta_v\right)}. \qquad \text{(A.13)}$$

A.5 Metropolis-within-Gibbs Sampling

MCMC methods have the advantage of being asymptotically exact, but collapsed Gibbs sampling requires ad hoc hyperparameter specification. The approach adopted in this paper deviates from the common strategies in order to achieve asymptotically exact results and formally infer concentration hyperparameters. The estimation is based on collapsed Gibbs sampling mixed with a Metropolis-Hastings step. In marketing research Jacobs, Donkers, and Fok (2016) implement Metropolis-within-Gibbs sampling to predict purchases with LDA, where a product purchase corresponds to a word and a customer corresponds to a document.

The basic LDA model is extended by adding one more layer to the hierarchical structure where log-normal prior distributions are imposed on the Dirichlet concentration parameters. Based on the considerations in Section 3.2, the Dirichlet prior on the topic-document distributions is asymmetric, whereas the Dirichlet prior on the topic-word distributions is symmetric. The posterior distribution (marginalized over $\mathbf{\Theta}$ and $\mathbf{\Phi}$) is rewritten as

$$p(\mathcal{Z}, \boldsymbol{\alpha}, \beta | \mathcal{W}) \propto \left(\prod_{d=1}^{D} \prod_{i=1}^{N_d} \underbrace{p(w_i^{(d)} | z_i^{(d)}, \beta)}_{Multinomial} \underbrace{p(z_i^{(d)} | \boldsymbol{\alpha})}_{Multinomial} \right) \underbrace{\pi(\beta)}_{Lognormal} \prod_{k=1}^{K} \underbrace{\pi(\alpha_k)}_{Lognormal} . \tag{A.14}$$

The choice of the parameters for the prior distributions is guided by heuristics proposed by Griffiths and Steyvers (2004) for text modeling. The mode of the prior distribution for β is set to 0.1 and the variance is such that 95 percent of the probability mass is under 1. This specification reflects a prior belief that the word-topic distributions are sparse. The mode of the prior distribution for α_k, $k = 1, \ldots, K$, is set to $\frac{50}{K}$ and the variance is chosen such that 95 percent of the probability mass is under $\frac{50}{3}$. This prior specification favors more uniformly distributed document-topic probabilities, although it remains rather uninformative.

A.5.1 Metropolis-Hastings Step

In each sampling step of the Metropolis-within-Gibbs sampling procedure the topic assignments $\mathcal{Z}$ are drawn from the collapsed full posterior distribution (A.10). However, the full conditional distributions of $\boldsymbol{\alpha}$ and β are non-standard and cannot be obtained using the Gibbs sampler.

The full conditional posterior distribution of β is

$$p(\beta|\mathcal{Z},\mathcal{W},\boldsymbol{\alpha}) \propto \pi(\beta)\prod_{k=1}^{K}\left(\frac{\Gamma(V\beta)}{\Gamma(V\beta+\sum_{v=1}^{V}c_{k,*,v})}\prod_{v=1}^{V}\frac{\Gamma(\beta+c_{k,*,v})}{\Gamma(\beta)}\right). \tag{A.15}$$

The full conditional posterior distribution of α_k, $k = 1,\ldots,K$ is

$$p(\alpha_k|\mathcal{Z},\mathcal{W},\boldsymbol{\alpha}_{-k},\beta) \propto \pi(\alpha_k)\prod_{d=1}^{D}\left(\frac{\Gamma\left(\sum_{k=1}^{K}\alpha_k\right)}{\Gamma\left(\sum_{k=1}^{K}\alpha_k+c_{k,d,*}\right)}\times\frac{\Gamma\left(\alpha_k+c_{k,d,*}\right)}{\Gamma(\alpha_k)}\right). \tag{A.16}$$

The samples from the conditional posterior distributions (A.16) and (A.15) are obtained using the random-walk Metropolis-Hastings sampler. In general, the sampler makes use of proposal distributions of a known functional form for each Dirichlet concentration parameter. The proposal distributions specify the probability of moving to another "candidate" point in the parameter space in the next iteration, given the sample value in the current iteration. The candidate sample is then accepted or rejected, based on the acceptance probability. Specifically, the random-walk Metropolis-Hastings step is composed of the following parts:

(i) The candidate values β^* are sampled from $\log N(\beta, s_\beta^2)$, where $\beta^{(m)}$ denotes the current value of the parameter and s_β^2 is the variance of the proposal distribution, and the candidate values for α_k^* are sampled from $\log N(\alpha_k^{(m)}, s_{\alpha k}^2), k = 1,\ldots,K$.

(ii) For each univariate Metropolis-Hastings sampler, we compute the log acceptance probability (log transformation is applied to evaluate the gamma functions). For example, for the parameter β:

$$\log\delta = \min(r, 0) \tag{A.17}$$
$$r = \log(p(\beta^*|\mathcal{Z},\mathcal{W},\boldsymbol{\alpha})) + \log(q(\beta^{(m)}|\beta^*) - \log(p(\beta^{(m)}|\mathcal{Z},\mathcal{W},\boldsymbol{\alpha})) - \log(q(\beta^*|\beta^{(m)}),$$

where $q(\beta|\beta^{(m)})$ is a proposal density.

(iii) Set $\beta^{(m+1)} = \beta^*$ with probability δ.
Set $\beta^{(m+1)} = \beta^{(m)}$ with probability $1 - \delta$.

A.5.2 Calibration of the Proposal Distribution

Variances of the proposal distributions are calibrated within the first 500 iterations of the Metropolis-within-Gibbs sampling. The procedure for calibrating the proposal distributions closely follows Jacobs, Donkers, and Fok (2016). The target acceptance rate is 50 percent. The calibration window size is 10. For each calibration window the number of accepted samples (n_A) is stored. If n_A falls outside the 95 percent confidence bounds of the binomial distribution $B(10,\ 0.5)$, then the variance is decreased by $\max\left(\sqrt{\frac{n_A}{10\times0.5}}, \frac{1}{2}\right)$ or increased by $\min\left(\sqrt{\frac{n_A}{10\times0.5}}, 2\right)$. See Jacobs, Donkers, and Fok (2016) for details. The initial Metropolis-Hastings standard deviations are $s_\beta = 0.9$, $s_{\alpha_k} = 0.5$.

A.5.3 Posterior Analysis

The estimation is conducted using the whole vocabulary.[26] The sampler runs for 6,000 iterations. Some portion of the initial sample must be discarded as the burn-in period, because the starting values are not sampled from the target posterior distribution. We discard the first 2,000 draws as the burn-in. Every tenth draw is stored. This results in 400 samples from the posterior distribution. We repeat the above procedure for different numbers of latent topics: $K = 3, \ldots, 20$.

Stephens's algorithm (Stephens 2000) is implemented to verify whether label switching has occurred and to perform relabeling if

[26]The multiple starts procedure shows that the chain is not sensitive to the starting values; therefore, in estimation that procedure is omitted.

necessary. The posterior mean estimates of the model parameters are obtained by averaging over the draws.

To achieve a robust evaluation, we compute the measures of semantic coherence and topic exclusivity for different topic cardinalities: $N = 5, 10, 15, 20$, where N denotes the number of words with the highest probability in a topic (Lau and Baldwin 2016). A single score for the model with K components is obtained by averaging the topic-level scores.

A.5.4 Multiple Random Starts

Standard MCMC methods, such as the Metropolis-Hastings algorithm, are known to slowly traverse the support of highly multimodal distributions (Jasra, Holmes, and Stephens 2005). Therefore it is important to investigate the influence of initialization on the solution. In topic modeling, perplexity is a standard measure to examine the model fit and the convergence of the Markov chains initialized from different starting points (Asuncion et al. 2009; Airoldi et al. 2014). Lower perplexity indicates better performance of the model in predicting out-of-sample words. To evaluate perplexity, we split words into a training set (75 percent of words per document) that is used to estimate model parameters and a testing set (25 percent of words per document). Words in the testing set serve to evaluate the generalization capability of the model and therefore are not used in the parameter estimation. Perplexity is defined as the inverse of the geometric mean per-word likelihood of the test data (Hornik and Grun 2011):

$$Perplexity = exp\left(-\frac{\sum_{d=1}^{D}\sum_{v=1}^{V} c_{*,d,v}^{test}\log(\sum_{k=1}^{K}\phi_{k,v}\theta_{d,k})}{\sum_{d=1}^{D} N_d^{test}}\right), \tag{A.18}$$

where $\phi_{k,v}$ and $\theta_{d,k}$ are estimated on the training data.

The sampler is run from five multiple random starts on the training data. For each starting point the sampler runs for 2,000 iterations. In each iteration topic assignments are simulated with a collapsed Gibbs sampling step and concentration parameters for the Dirichlet distributions are simulated with a Metropolis-Hastings step. In each run of the sampler and in each iteration, we use the parameters estimated on the training data to measure the goodness of fit for the test data. In addition, some portion of the initial sample must be discarded as the burn-in period, because the starting values are not sampled from the target posterior distribution. We therefore discard the first 1,000 draws as the burn-in portion.

It is assumed that the MCMC chain has converged to the posterior distribution when the values of perplexity across iterations stabilize.[27]

After convergence, differences in the estimated perplexities for multiple runs turned out to be marginal, indicating that the estimated results are stable across initializations (see Table A.1).

A.5.5 Implementation

We implement the estimation procedure and model diagnostics in `C++` integrated with `R` using application programming interface (API) enclosed in the `Rcpp` package (Eddelbuettel and François 2011). The full code is provided in the replication files.

[27]The absolute percentage change in perplexity score between every tenth iteration is less than 1 percent.

Table A.1. Perplexity Scores for Five Chains with Different Initializations (iterations 1,000–2,000)

Section	Mean					Standard Deviation				
	1	2	3	4	5	1	2	3	4	5
DS	357.87	362.74	363.39	363.65	361.44	2.26	2.32	2.40	2.44	2.29
EA	383.23	380.24	383.34	383.37	380.74	1.54	1.51	1.50	1.44	1.44
MA	293.57	297.42	293.52	293.49	293.72	1.36	1.28	1.31	1.29	1.37
QA	1,039.17	1,039.39	1,041.37	1,039.19	1,038.50	1.30	1.51	1.32	1.27	1.39

Note: Sections (number of topics in parentheses): DS – Decision summary (6); EA – Economic analysis (6); MA – Monetary analysis (4); QA – Q&A (10).

Appendix B. Vocabulary Selection

The analysis presented in the paper required many text preparation steps, such as removing punctuation and numbers, lowercasing, stop word removal, term normalization, or n-gram inclusion. Preparing documents involves preprocessing decisions such as which normalization technique to use or how to reduce vocabulary size, taking into account individual characteristics of our data set. This appendix discusses additional text preprocessing details.

We remove formulaic phrases that are often used to introduce a section of the ECB press conference. These phrases are repeated in many speeches, but have low informational value. The list of removed expressions is provided in Table B.1.

Two term normalization approaches are usually distinguished—stemming and lemmatization. In the paper we opt to use a lemmatizer, although both techniques aim to reduce inflectional and derivational word forms to a common base form.

Stemming refers to applying a set of rules to remove the affixes (for example, it reduces "increasing" to "increas," "stability" to "stabil," "financial" to "financi"). The most widely used methods are algorithmic stemmers (i.e., Porter 2001), which operate without a lexicon and thus ignore word meaning.

Table B.1. List of Expressions Removed from the Corpus of the ECB Press Conferences

"Ladies and gentlemen, the Vice President and I are very pleased to welcome you at the press conference."
"I will now report on the outcome of today's meeting of the Governing Council of the ECB, which was also attended by (. . .)"
"Based on its regular economic and monetary analysis the Governing Council decided"
"Let me now explain our assessment in greater detail, starting with the economic analysis."
"Turning to the monetary analysis"
"We are now at your disposal for questions."

In contrast to algorithmic stemmers, lemmatization requires specifying the part of speech of a word in a sentence in order to reduce the word to its dictionary form (lemma). A lemmatizer transforms all plurals into singular forms and past-tense verbs to present-tense verbs (e.g., “left” to “leave,” “developments” to “development,” but “stability” and “financial” are unaffected).

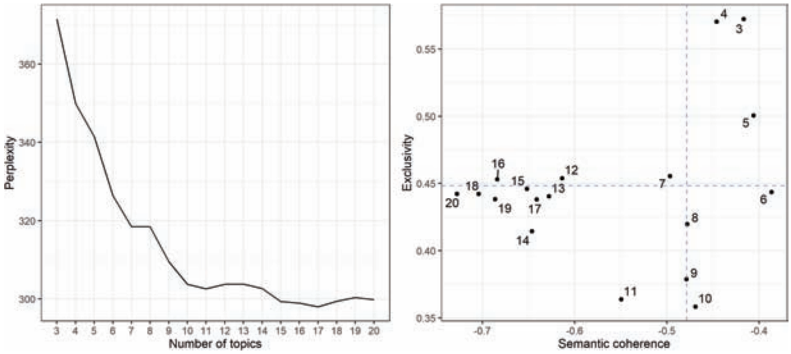
A lemmatizer is more accurate than a stemmer, and it is unlikely to over-conflate (Schofield and Mimno 2016). First, a lemmatizer finds a common form for irregular verbs and nouns (“analyses” — “analysis”, “indices” — “index”), which an algorithmic stemmer cannot do. Second, a stemmer may remove too many endings and conflate terms with different meanings. For example, a stemmer (e.g., the Porter 2001 stemmer) would view the following pairs of words as equivalent while lemmatization would not: “import” and “important”, “income” and “incoming”, “emerging” and “emergency”, “future” and “futures”, “maturity” and “mature”, “consistent” and “consist”, “positive” and “position”, “accounts” and “accountability”.

A lemmatizer increases precision at the expense of recall. In contrast to a stemmer, it is not able to conflate semantically related words belonging to different parts of speech. For example, in the sentence: “With regard to fiscal policies, the Governing Council sees continued reasons for concern”, the term “continued” is tagged as an adjective and its lemma is “continued”. The Porter stemmer conflates “continue”, “continuing”, and “continued” to the same stem, “continu”.

Appendix C. Model Selection

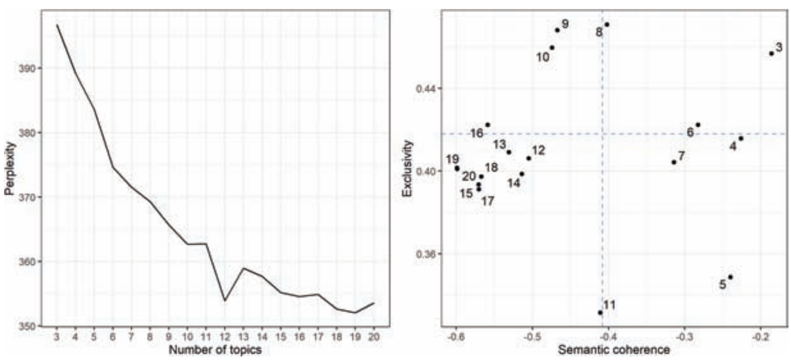
In Figures C.1–C.4 the left panel shows the average perplexity and the right panel presents the average semantic coherence versus the average exclusivity for models with different numbers of topics specified. Lower perplexity indicates better predictive performance of the model, while higher coherence and exclusivity indicate more interpretable topics, on average. The dashed lines mark the 2/3 quantile along each dimension (exclusivity, semantic coherence). We first discard the least performing solutions along the two dimensions separately to remove solutions that have, for example, extremely

Figure C.1. Selection of Number of Topics in the Decision Summary



Note: Selected number of topics is five.

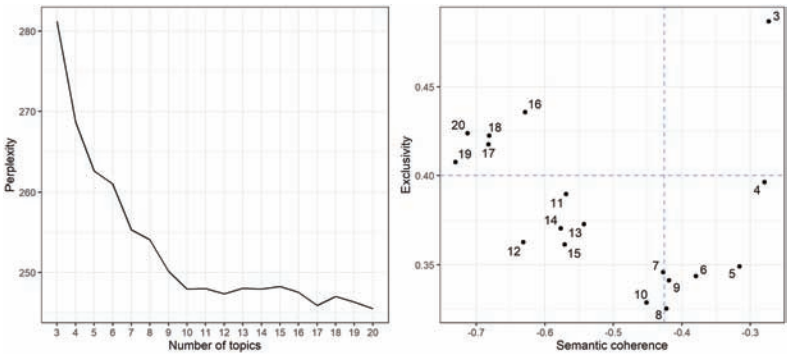
Figure C.2. Selection of Number of Topics in the Economic Analysis Section



Note: Selected number of topics is six.

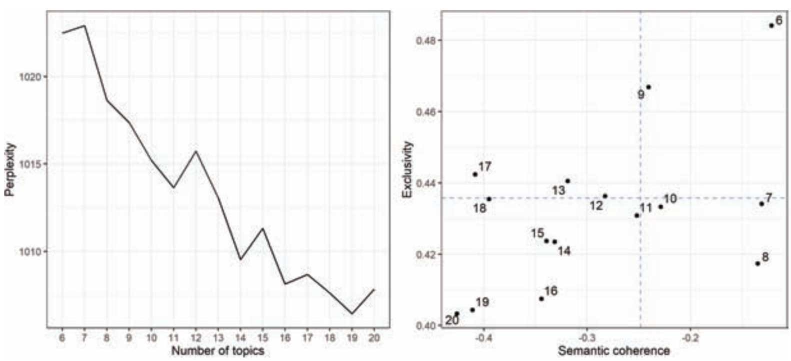
high coherence but very low exclusivity and vice versa. Therefore, we select one of the models located in the top right corner of the graph.

Figure C.3. Selection of Number of Topics in the Monetary Analysis Section



Note: Selected number of topics is four.

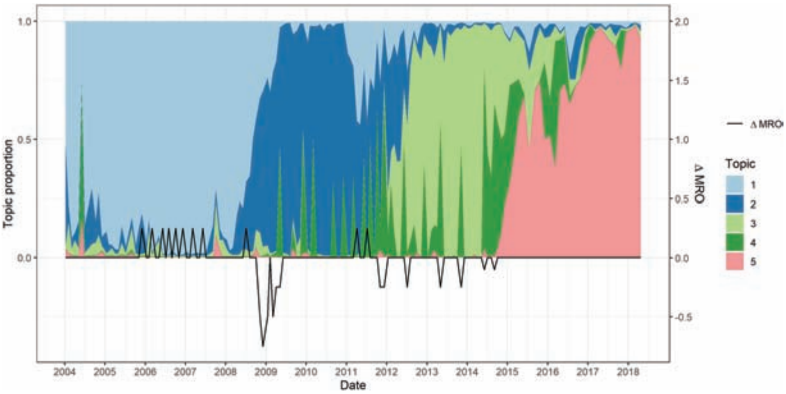
Figure C.4. Selection of Number of Topics in the Q&A



Note: Selected number of topics is nine.

Appendix D. Estimated Topics

Figure D.1. Topics in the Decision Summary over Time



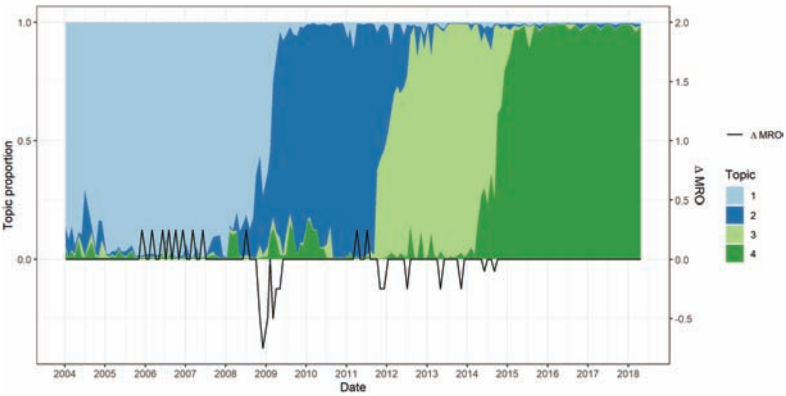
Note: This figure plots the proportion of the decision summary devoted to each topic along with the ECB MRO rate decisions. The topics were estimated using the LDA algorithm (Blei, Ng, and Jordan 2003). The sample comprises 156 transcripts of the section from the ECB press conferences between 2004 and 2018.

Table D.1. Top 10 Terms Describing Topics of the Decision Summary, Ranked by the FREX Score

Topic 1	Topic 2	Topic 3	Topic 4	Topic 5
upside_risk	likely	subdue	conduct	asset_purchase
vigilance	temporary	picture	operation	sustained
prerequisite	purchase_power	prolong	fixed_rate	path
monitor_closely_development	nature	government	procedure	net
solidly	non_standard_measure	weakness	maintenance_period	case
ongoing	construction	extend	tender	monthly
vigorous	inflationary_pressure	impact	full_allotment	beyond
nominal	moderate	dynamic	mros	app
real	allotment	base	three_month	run
timely_manner	acute	financial	least	financial_condition

Note: The FREX score gives high ranks to terms that are both frequent and exclusive

Figure D.2. Topics in the Monetary Analysis Section over Time



Note: This figure plots the proportion of the monetary analysis section devoted to each topic along with the ECB MRO rate decisions. The topics were estimated using the LDA algorithm (Blei, Ng, and Jordan 2003). The sample comprises 156 transcripts of the section from the ECB press conferences between 2004 and 2018.

Table D.2. Top 10 Terms Describing Topics of the Monetary Analysis Section, Ranked by the FREX Score

Topic 1	Topic 2	Topic 3	Topic 4
price_stability	challenge	resilience	recovery
upside_risk	fund	heightened	place
horizon	recapitalisation	country	across
medium_long_term	government	transmission	firm
expansion	full	adjusted	condition
monetary_expansion	advantage	step	annual_rate
price	address	fragmentation	begin
strength	outside	inflow	improvement
trend	different	establish	narrow
influence	steep	adequate	put

Note: The FREX score gives high ranks to terms that are both frequent and exclusive.

Appendix E. Descriptive Statistics

Table E.1. Descriptive Statistics

	Mean	**Std.**	**Min.**	**Max.**
ΔV_t	−0.012	0.066	−0.180	0.249
$\lvert s_t^{MRO}\rvert$	0.009	0.045	0.000	0.250
D_t^A	0.077	0.267	0.000	1.000
r_t^{DE}	−0.003	0.063	−0.204	0.288
$\lvert s_t^{JC}\rvert$	14.144	12.367	0.000	64.000
I_t^{DS}	0.784	0.185	0.399	0.988
I_t^{EA}	0.790	0.170	0.361	0.984
I_t^{MA}	0.889	0.126	0.519	0.993
I_t^{QA}	0.679	0.141	0.419	0.960

Note: This table displays descriptive statistics of variables used in the event-based regressions. ΔV_t denotes the daily percentage change in the VSTOXX index on the conference day t relative to the previous day, $|s_t^{MRO}|$ and $|s_t^{JC}|$ are absolute surprise components of the Main Refinancing Operations (MRO) rate and the U.S. jobless claims, respectively, D_t^A is an indicator for announcements regarding non-standard monetary policy measures, r_t^{DE} is a daily change in German two-year government bond yields and I_t^{DS}, I_t^{EA}, I_t^{MA}, I_t^{QA} denote the index values that capture changes in communication by section: decision summary, economic analysis, monetary analysis, Q&A.

Table E.2. Correlation Matrix

	ΔV_t	I_t^{DS}	I_t^{EA}	I_t^{MA}	I_t^{QA}
ΔV_t	1.000				
I_t^{DS}	0.040	1.000			
I_t^{EA}	−0.140	0.197	1.000		
I_t^{MA}	−0.166	0.420	0.448	1.000	
I_t^{QA}	0.104	0.281	−0.130	0.105	1.000

Note: This table displays Pearson correlation coefficients of our four topic-based communication variables and the VSTOXX index between 2004 and 2018 at a monthly frequency.

Appendix F. Robustness Check

Table F.1. Robustness Check: Different Number of Topics

	Dependent Variable: ΔV_t		
	(1)	**(2)**	**(3)**
$\lvert s_t^{MRO}\rvert$	−0.112	−0.144	−0.154
	[0.337]	[0.220]	[0.193]
δV_{t-1}	0.064	0.065	0.053
	[0.558]	[0.558]	[0.623]
r_t^{DE}	−0.159*	−0.162**	−0.146*
	[0.054]	[0.049]	[0.077]
$\lvert s_t^{JC}\rvert$	0.0004	0.0004	0.0003
	[0.372]	[0.316]	[0.522]
D_t^A	−0.008	−0.020	−0.004
	[0.691]	[0.320]	[0.858]
I_t^{DS}	0.027	−0.012	0.034
	[0.433]	[0.681]	[0.397]
I_t^{EA}	−0.015	0.007	−0.039
	[0.644]	[0.858]	[0.295]
I_t^{MA}	−0.130***	−0.094**	−0.130***
	[0.010]	[0.029]	[0.006]
I_t^{QA}	0.031	0.040	0.026
	[0.525]	[0.278]	[0.493]
$I_t^{QA} \times D_t^{Draghi}$	−0.053***	−0.051***	−0.042**
	[0.006]	[0.008]	[0.016]
Constant	0.090*	0.058	0.103*
	[0.098]	[0.173]	[0.084]
Observations	156	156	156
Adjusted R^2	0.112	0.103	0.117

Note: Column 1 reports the results where the baseline dimensionality is decreased by 1; column 2 reports the results where the baseline dimensionality is increased by 1; column 3 reports the regression results where the number of topics is selected first by discarding solutions below the 2/3 quantile along dimensions: coherence, exclusivity, and then selecting the model that strictly dominates other models in terms of both coherence and exclusivity; p-values are in brackets. *$p < 0.1$; **$p < 0.05$; ***$p < 0.01$.

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Central Banks in Parliaments: A Text Analysis of the Parliamentary Hearings of the Bank of England, the European Central Bank, and the Federal Reserve*

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This paper investigates whether parliamentary hearings are effective in holding central banks accountable against their mandates. To this end, it applies text analysis on the hearings of the Bank of England, the European Central Bank, and the Federal Reserve from 1999 to 2019. It finds that central bank objectives play a crucial role in determining the topic of the hearings. It also shows that sentiments are more negative when the distance between inflation and the central bank's

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inflation aim increases. These results suggest that parliamentary scrutiny serves its intended purpose. However, topics and sentiment react more to inflationary rather than deflationary deviations of inflation away from target.

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1. Introduction

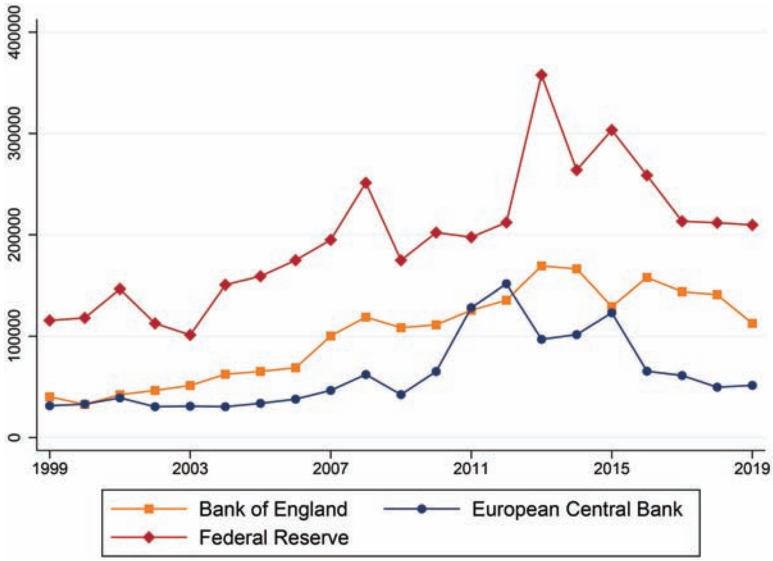
Delegation of responsibilities to unelected institutions might give rise to a perceived democratic deficit over time, even when they originate from a democratic decision. For such delegation to be acceptable in a constitutional democracy, unelected officials need to be accountable to democratically elected institutions, which represent the view of the people.

This fundamental norm is an essential basis of the delegation of monetary policy to an independent institution, the central bank. Governments delegate monetary policy to central banks that can conduct policies independently from pressures in order to achieve lower levels of inflation, as shown theoretically and empirically by Barro and Gordon (1983), Alesina (1989), and Grilli, Masciandaro, and Tabellini (1991). As they do so, they put in place a series of arrangements that allow elected representatives to monitor the central bank's achievement of its objective. The most common of these arrangements across central banks is parliamentary hearings (Bank for International Settlements 2009), i.e., the requirement for the central bank (generally the governor) to explain and justify its policy decisions before the parliament on a regular basis.

For a long time, this principle had been hardly a subject of discussion, either in the academic or public debate. However, with the recent financial crisis the trade-off between independence and accountability has become more complex. On the one hand, the key role of central banks during the crisis led to increased public attention being paid to their policies compared with the pre-crisis period (see Figure 1). On the other hand, the adoption of non-standard measures made the scrutiny of monetary policy more complex (Coeuré 2018).

This revived the debate around the legitimacy of granting independence to unelected powers in constitutional democracies (Tucker

Figure 1. Number of Newspaper Articles Citing the Bank of England, the European Central Bank, and the Federal Reserve, 1999–2019



Sources: Authors' elaboration on data from Factiva as of December 2019. The data used cover newspapers in all the languages available on Factiva.

2018). Moreover, with the emergence of populism during the crisis, the institutional tenets of central banks have been increasingly challenged. The literature emphasize this change of public perception toward central banks, arguing that the rise of populism might put their independence at risk (Buiter 2016, De Haan and Eijffinger 2017, Goodhart and Lastra 2017, Rodrik 2018; for a review, see Merler 2018). In contrast to the past, critical voices toward central bank independence now dominate (Masciandaro and Romelli 2015, Issing 2018).

As a result, central bankers around the world now see preserving central bank independence as a challenging task. This is also confirmed by a survey we conducted among 30 experts working on institutional matters in their respective central banks worldwide: the majority of the respondents identified the preservation of central bank independence as the main challenge for central banks in

2019 (Figure A.1 in the appendix). These results are in line with those of a similar expert survey in which 39 of the 70 respondents agree with the statement that there will be significant changes in the independence of monetary policy in the United Kingdom and the euro zone in the foreseeable future (Den Haan et al. 2017).

In this context, it is therefore crucial to understand how independent central banks interact with their elected counterparts. Over the recent years, the latter have started discussing the desirability and possible ways to better exercise parliamentary control over central bank activities. The "Audit the Fed" bill was presented by U.S. Senator Rand Paul to strengthen Congress's control over and access to the Federal Reserve's information and possibly make meeting-by-meeting monetary policy decisions subject to congressional review (Bernanke 2016, 2022). On the other side of the Atlantic, the European Parliament has requested several inputs to academics to benchmark the European Central Bank's accountability against other central banks and assess possible avenues to reinforce the parliamentary control over its activities (Lastra et al. 2020).

If the academic discussion has so far mainly focused on whether central banks have become too independent (Balls, Howat, and Stansbury 2018), it is equally important to give attention to the aspect on which the legitimacy of central bank independence rests—namely central bank accountability and, in particular, on how this is ensured. Moreover, the limited literature on central bank accountability focuses on how to enhance legitimacy in the statute of the central bank, limiting its considerations to understanding which arrangements are best suited to hold the central bank accountable (Tucker 2018), and not on what actually happens in a given arrangement.

This leaves open the fundamental question on how elected representatives actually monitor the central bank in a given arrangement. In other words, it is not clear whether accountability works in practice. This broad question can be narrowed down to two queries related to parliamentary hearings: (i) what topics are discussed? and (ii) what drives the tone of the hearings? The answers to these questions are not trivial. The topic of the discussion is meant to be the fulfillment of the objective(s) of the central bank. However, scholars argue that often this is not the case (Schonhardt-Bailey 2013; Claeys, Hallerberg, and Tschekassin 2014a). Politicians may find monetary policy too technical or simply not appealing before the electorate, and may prefer to discuss other topics. Similarly, we expect the tone of the discussion to turn more negative when the central bank diverges from its objective. At the same time, sentiments may be driven by negative economic conditions, regardless of the central bank's ability to cope with them. Moreover, politicians may assume a more aggressive tone toward the central bank for electoral reasons, regardless of its performance in fulfilling the objective (Goodhart and Lastra 2017).

In this paper we intend to fill this gap and answer these questions empirically. To do so, we apply text analysis techniques to the transcripts of the parliamentary hearings of three central banks, the Bank of England (BoE), the European Central Bank (ECB), and the Federal Reserve (Fed), for the period 1999–2019. In particular, we use *topic* and *sentiment* analysis to inspect, respectively, what drives the focus and the tone of the hearings. By doing so, we are able to test through panel data regressions whether the focus and the tone of the hearings are associated with the objective of the central bank or whether other factors play a more relevant role.

Our contribution to the literature is threefold. First, we provide a new empirical methodology to assess parliamentary hearings, an essential aspect of central bank accountability, as well as new findings on the three cases we examine. This is relevant compared with the existing empirical literature on central bank accountability, which focuses on de jure accountability, i.e., accountability as enshrined in laws and regulations (see De Grauwe and Gros 2008 for a review), rather than on de facto accountability, i.e., the actual interaction between the central bank and elected bodies in a given framework.

Second, we enrich the literature on central bank communication. While existing research mostly looks at central bank announcements to the public through press conferences (Altavilla et al. 2019; Lamla and Vinogradov 2019), publications (Born, Ehrmann, and Fratzscher 2014; Bholat et al. 2019; Hansen, McMahon, and Tong 2019), speeches (Hansen, McMahon, and Tong 2019; Neuhierl and Weber 2019; Moschella, Pinto, and Diodati 2020), minutes of their meetings (Apel and Blix-Grimaldi 2012; Hansen, McMahon, and Prat 2017), our work is the first to explore the communication between central banks and parliaments in a comparative setting.

Third, our work adds to the emerging literature that applies text mining to central banking (for a review, see Bholat et al. 2015). While existing works analyze the text of central bank policy announcements and speeches (Lucca and Trebbi 2009; Born, Ehrmann, and Fratzscher 2014; Tobback, Nardelli, and Martens 2017; Hansen, McMahon, and Tong 2019; Gorodnichenko, Pham, and Talavera 2021), the minutes of their meetings (Apel and Blix-Grimaldi 2012; Hansen, McMahon, and Prat 2017; Shapiro and Wilson 2019), or of news and tweets related to central banks (Bianchi, Kung, and Kind 2019; Binder 2021; Ehrmann and Wabitsch 2022), we provide evidence on a type of central bank text which has been largely left unexplored, i.e., the transcripts of central banks' parliamentary hearings. Few exceptions in the political science literature analyzed these text sources, but they all focused on specific case studies (Schonhardt-Bailey 2013; Sanders, Lisi, and Schonhardt-Bailey 2018; Bisbee, Fraccaroli, and Kern 2022; Ferrara et al. 2022; Fraccaroli et al. 2022a, 2022b).

Three important caveats apply to our findings. First, our analysis focuses on the monetary policy functions of the central banks—thus leaving aside the supervisory functions and the accountability provisions applicable to them. Second, we look at one specific arrangement of central bank accountability, namely parliamentary hearings. While this is the most diffused and, generally, the most relevant tool to hold central banks accountable, there exist other provisions too (Fraccaroli, Giovannini, and Jamet 2018). Third, in some jurisdictions the executive plays an important role in holding the central bank accountable together with the parliament. As we focus on parliamentary hearings, our study does not encompass the relationship between the central bank and the government.

For these reasons, our analysis is limited to the accountability of the central bank vis-à-vis the parliament for monetary policy matters. While this represents one of the most relevant and widely used accountability practices, our evidence is not necessarily valid for other accountability practices, such as those between the central bank and the executive, or the hearings on non-monetary policy matters.

The remainder of this paper is structured as follows. In the next section we define central bank accountability in a principal-agent framework and discuss the limitations of existing measures that aim to capture and assess accountability. Section 3 briefly describes the parliamentary hearings of the BoE, the ECB, and the Fed, explaining in particular their objectives and functioning. Section 4 outlines our database and text-based methodology to account for accountability practices. Moreover, it presents the empirical model we use to explore the topic and sentiments of the hearings. In Section 5 we present and discuss the empirical results. The final section concludes.

2. Central Bank Accountability: Theory and Measurement

2.1 Theoretical Framework

Central bank accountability (CBA) can be understood as the legal and political obligation for a central bank to explain and justify its decisions to citizens and their elected representatives. According to the Bank for International Settlements (2009), accountability encompasses three main characteristics: (i) scrutiny by others; (ii) regular accounting for one's actions; and (iii) the risk of negative repercussions, if performance is considered unsatisfactory.

The rationale for CBA can be envisaged in a principal-agent framework, where powers are delegated to an agent to be exercised independently of its principal (Fratianni, Hagen, and Waller 1997; Gailmard 2014). In this setup, as noted by Fischer (1995), accountability is needed for two main reasons. First, it sets incentives for the central bank to meet its goals; and second, it provides democratic oversight of its policies. CBA is indeed key to ensure that independence does not lead to arbitrariness and that the mandate is fulfilled, while preserving the benefits of independence.

In a nutshell, this principal-agent framework can be described as follows. Assume that there are two principals, A and B, with divergent preferences over inflation, i.e., A is more inflation averse than B. The two principals are elected representatives: they could be two contending political parties or, in the special case of a monetary union, the representatives of two countries. When they delegate monetary policy to an independent agent (the central bank), they agree on a mandate, or objective, which is equidistant from their preferences. If the central bank were to drift away from the objective agreed by the two principals, it would benefit one of the principals to the detriment of the other. To avoid this, the two principals establish (ex ante) a commonly agreed objective, independence from external influence, and an accountability framework. The latter aims to provide a set of arrangements that allow them to scrutinize whether the central bank is respecting its mandate.

As accountability centers on an evaluation of performance, this is translated in practical terms in the establishment of a legal obligation for the central bank to testify before its principal(s). The latter is eventually the people as represented by the parliament or the government (or other institutions) according to the jurisdiction in which they operate.

According to the theory, therefore, the focus of parliamentary hearings should be the objective of the central bank, and whether the central bank has been able to fulfill it. Nevertheless, scholars raised doubts around the ability of the parliamentary hearings to actually assess the performance of the central bank, as monetary policy is highly technical (Schonhardt-Bailey 2013; Claeys, Hallerberg, and Tschekassin 2014a, 2014b) and may therefore have a low political appeal to the electorate than other matters, as for example issues related to the transparency of the central bank.

Moreover, a number of political and economic drivers may divert the focus of the discussion away from the objective and affect the tone of the debate. First, macroeconomic conditions could influence both the focus and the tones of the hearings. For example, an increase in unemployment may divert the discussion away from price stability considerations.[1] The same might hold for financial distress, which would shift the focus from price stability to financial stability. While we might expect negative economic conditions to worsen the tone of the discussion, the opposite could also be true. In times of financial distress, the interactions between the central bank

[1]This example holds for the cases of the BoE and of the ECB, where price stability is a statutory objective whereas employment is not. In the case of the Fed, this would not represent a divergence from the objective, as its mandate includes the promotion of maximum employment. The example still applies to all three central banks if we substitute unemployment with another macroeconomic variable that is not included in the objective(s) of the central bank. For a more detailed discussion on the objectives of the three central banks, see the next section.

and parliamentarians could intensify, as they did in Europe during the euro crisis (Fraccaroli, Giovannini, and Jamet 2018), since both bodies, under different roles, cooperated to tackle the euro area's problems (Torres 2013; Collignon and Diessner 2016).

A second factor is elections. According to the political business cycle theory, as elections approach, politicians tend to exert higher pressures on central banks, calling for a more expansionary monetary policy which would result in short-term gains at the expenses of higher inflation in the long run (Nordhaus 1975; Alesina 1989). For this reason, the occurrence of an election in the near future may divert the discussion away from price stability to issues related to employment. However, the opposite could also be true: as elections approach, politicians want to signal to their voters that they are effective scrutineers, and might therefore increase their focus on the objective of the central bank. In both cases, we might expect tones to become more negative. On the other hand, tones might be more positive if the incumbent exploits the hearings to praise existing economic conditions in order to get reelected.

A third element is uncertainty. Baker, Bloom, and Davis (2016) find that greater economic policy uncertainty is associated with both political (e.g., tight presidential elections) and economic (e.g., failure of Lehman Brothers) events and has negative repercussions on the economy, such as greater stock price volatility and reduced investment. Uncertainty can also affect negatively perceptions toward the central bank's policies. Using data on citizens' perceptions toward the BoE, the ECB, and the Bank of Japan, Istrefi and Piloiu (2020) show that shocks to economic policy uncertainty deteriorate public trust in central banks. Uncertainty is therefore likely to be associated with more negative tones.

2.2 Measurement Issues in the Empirical Literature

It follows that from a theoretical standpoint it is not clear which factors drive in practice the topics and the tones of the hearings, nor how these factors may influence them. These gaps in the theory motivate an empirical analysis.

However, the existing empirical literature on CBA mostly focuses on the design of accountability arrangements, and not on how accountability is discharged.[2] By looking at a number of aspects in the statutes of central banks (e.g., the possibility for the government to override a decision of the central bank), scholars created CBA indices to rank and compare the degree of de jure accountability of different central banks across the world (Briault, Haldane, and King 1998; De Haan, Amtenbrink, and Eijffinger 1999; Bini-Smaghi and Gros 2000; see De Grauwe and Gros 2008 for a review). These indices, which are summarized in Table A.1 in the appendix, are similar to the widely used indices of central bank independence (e.g., the ones constructed by Grilli, Mascandiaro, and Tabellini 1991 and Cukierman, Webb, and Neyapti 1992, which was updated by Garriga 2016). While these measures can be useful to compare the legal provisions in place in different countries for the principal(s) to scrutinize the central bank (de jure accountability), they do not describe whether this scrutiny serves its intended purpose (de facto accountability). This shortcoming is even more problematic considering that the absence of changes in de jure accountability[3] has been seen by some as a factor that negatively affect public opinion towards central banks, increasing threats toward their independence (Goodhart and Lastra 2017; Merler 2018).

Whether accountability frameworks actually work remains therefore an open question. To fill this gap, we propose a new methodology based on text analysis of the parliamentary hearings, one of the most common and relevant tools to hold central banks accountable. The next section describes why parliamentary hearings offers a good basis for analysis across several jurisdictions and provides a brief overview of the hearings of the BoE, the ECB, and the Fed.

3. The Parliamentary Hearings and the Cases of the BoE, the ECB, and the Fed

While there exist other accountability practices (for a review of the accountability practices of the ECB, see Fraccaroli, Giovannini,

[2]In the political science jargon, we could say that the empirical literature tend to focus mostly on CBA from an *input legitimacy* perspective rather than from a *throughput* one (Schmidt 2013).

[3]We refer to changes in CBA for the monetary policy functions. Reforms have been implemented for the new function of banking supervision, as we discuss later in the paper.

and Jamet 2018), parliamentary hearings provide a good basis to examine the practice of central bank accountability for two main reasons.

First of all, according to the Bank for International Settlements (2009), most central banks are accountable to parliaments. Out of a sample of 47 countries, in 64 percent of them central banks are accountable to parliament, in 30 percent to the minister of finance, in 21 percent to the government or its head, in 9 percent to the head of state, and in 17 percent to other bodies (e.g., cantons in Switzerland or private shareholders in the Republic of South Africa and other cases). Moreover, the transcripts of the hearings are generally publicly accessible online. For these reasons, the methodology we propose in this work is applicable to a wider number of central banks allowing for cross-country comparisons.

Secondly, the hearings are the direct expression of CBA. This characteristic can be appreciated in comparison with other methodologies adopted to study the relationship between central banks and politicians. For example, Binder (2021) studies the pressures of the executive on the central bank using the text of news reports, whereas Bianchi, Kung, and Kind (2019) analyze the tweets of U.S. President Trump against the Fed. While only the first of these methodologies has the advantage of being comparable across countries, both approaches provide fundamental information on the relationship between the central bank and the executive. This is particularly relevant, as the executive can be influential over the central bank's policy since in many jurisdictions it holds the power to remove the central bank governor.[4] However, this data is unidirectional, as it does not incorporate information on how the central bank responds to these pressures. On the contrary, parliamentary hearings are based on a question-and-answer (Q&A) session where the staff of the

[4]This is not the case for the President of the ECB. The governor of the BoE can be removed only by the Bank's Court of Directors, whose members are appointed by the Crown, with the exception of the Chair of the Court, who is appointed by the Chancellor of the Exchequer. To do so, the Court first needs the consent of the Chancellor of the Exchequer (UK Parliament 2016). In the United States, the President can remove a member of the Board of Governors for inefficiency, neglect of duty, or malfeasance in office. However, it is not clear whether the U.S. President has the authority to fire the Chair of the Fed's Board of Governors (Conti-Brown 2015, 2019).

central bank and parliamentarians interact in real time. Moreover, and more importantly, as previously described, the hearings rest on an explicit legal requirement to scrutinize the central bank. Moreover, it can be argued that the information on the executive's policy preferences toward the central bank is indirectly captured in our data through the participation to the hearings of parliamentarians from the governing parties, who are likely to share the policy preferences of the government.

Thirdly, although they have different electoral and party systems, parliaments tend to reflect a more plural picture of the political environment the central bank is exposed to, as they generally include both parties in support and against the existing government. This is an advantage compared with approaches that look exclusively at the relationship between the central bank and the executive, such as that of Bianchi, Kung, and Kind (2019) and Binder (2021).

An important caveat to our analysis is that we examine only those hearings that are related to monetary policy. This is relevant since, following the crisis, the increased involvement of central banks in financial stability and banking supervision led in some cases to the establishment of separate hearings for these functions.

The United Kingdom established separate hearings for the members of the newly created Financial Policy Committee to discuss the Financial Stability Report. In Europe, the creation of the Single Supervisory Mechanism in 2014 included the establishment of the hearings of the Chair of the Supervisory Board on the topic of banking supervision. In the United States, the 2010 Dodd-Frank Act created the Financial Stability Oversight Council (FSOC), which testifies on an annual basis before the Senate Committee on Banking, Housing, and Urban Affairs on its Annual Report. However, while the Chairman of the Federal Reserve is a voting member of the FSOC, its chair is the Secretary of the Treasury (analogous to the minister of finance), who is also the one that testifies before Congress.

While these hearings offer an interesting data source, they are relatively recent compared with the ones on monetary policy, and therefore leave little room for comparison due to their short time series. Moreover, the three cases we analyze have very different institutional structures to deal with banking supervision and financial stability more broadly, making the comparison on this function more cumbersome. For example, while in the United Kingdom the creation of the Financial Policy Committee was accompanied by a change in the statute of the BoE to include a financial stability objective, the statutory objectives of the ECB and of the Fed were left unchanged (for a recent discussion on the case of the FSOC, see Kashyap and Siegert 2020).

We acknowledge, though, that monetary policy and financial stability can be interlinked, as noted by Smets (2014). Theoretically, this link leaves room for discussions on financial stability during the hearings for monetary policy too. Therefore, while we do not investigate this issue directly, as it goes beyond the scope of our research, we include the topic of financial stability in our analysis.

Following these considerations, in the next subsection we describe the hearings envisaged for the monetary policy functions of the three central banks.

3.1 The Regular Hearings of the BoE, the ECB, and the Fed

As previously discussed, parliamentary hearings are meant to be a tool for elected representatives to scrutinize whether and how the central bank is achieving its mandate. One of the advantages of comparing the BoE, the ECB, and the Fed is that for all three price stability is a primary objective.

The Bank of England Act states that "in relation to monetary policy, the objectives of the Bank of England shall be to maintain price stability" and "subject to that, to support the economic policy of Her Majesty's Government, including its objectives for growth and employment" (Part II, Article 11). The definition of price stability is a task of the British Treasury (Art. 12), which set the inflation target at 2 percent.[5] Similarly, the primary objective of the ECB is "to maintain price stability" as enshrined in Article 2 of the Statute of the European System of Central Banks and of the European Central Bank.[6] In 1998 the Governing Council of the ECB provided a

[5]The full text of the Act is available at the following link: https://www.bankofengland.co.uk/-/media/boe/files/about/legislation/1998-act.

[6]The statute is available at the following link: https://www.ecb.europa.eu/ecb/legal/pdf/oj_c_2016_202_full_en_pro4.pdf. In the statute the price stability objective applies to all the European System of Central Banks (ESCB), which

quantitative definition of this objective: inflation rates of below, but close to, 2 percent over the medium term.[7] The price stability objective of the Fed is enshrined in Section 2A of the Federal Reserve Act, which states that "the Board of Governors of the Federal Reserve System and the Federal Open Market Committee shall maintain [. . .] stable prices."[8] The Federal Open Market Committee then stated that inflation at the rate of 2 percent is consistent with the Fed's statutory mandate.[9]

However, there are also relevant differences. A first crucial difference is the higher independence enjoyed by the ECB and the Fed in defining price stability compared with the BoE. As described in the previous paragraph, while the ECB and the Fed have the autonomy to provide a quantitative definition of their price stability objective (which is decided by the Governing Council and by the Federal Open Market Committee, respectively), this is not the case for the BoE, whose inflation target is set by the government. Second, while for the BoE and the ECB price stability is the main monetary policy objective, the Fed has also the objective to promote the goal of maximum employment, which is in no way subordinated to the price stability mandate. This is an important difference compared with the BoE and the ECB, where employment is a secondary objective, i.e., an objective that is subject to the achievement of price stability.[10]

In our empirical analysis we exploit these commonalities and differences to investigate how the mandates democratically assigned to

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extends also to those national central banks that are members of the European Union but not of the euro area.

[7]The Governing Council of the ECB is composed of the President, the Vice-President, the other members of the ECB Executive Board, and the governors of the National Central Banks that are part of the euro area. The precise definition of price stability provided by the Governing Council is the following: "Price stability is defined as a year-on-year increase in the Harmonised Index of Consumer Prices (HICP) for the euro area of below 2%."

[8]The Federal Reserve Act is available at this link: https://www.federalreserve.gov/aboutthefed/section2a.htm.

[9]The statement is available at this link: https://www.federalreserve.gov/newsevents/pressreleases/monetary20120125c.htm.

[10]In the case of the BoE this subordination is explicit in Article 11b of the Bank of England Act. In the case of the ECB, this subordination is set in the requirement for the ECB (Article 2 of the Statute), without prejudice to the objective of price stability, to contribute to the achievement of the objectives set in Article 3 of the Treaty on European Union. These objectives include, among others, full employment.

the central banks can influence the focus of the discussion. Before doing so, we briefly describe the arrangements that set the interactions between each central bank and its respective parliament.

Bank of England. The BoE is held accountable by the House of Commons Treasury Committee through regular hearings. The members of the Treasury (Select) Committee are elected representatives of the House of Commons, the lower chamber of the U.K. Parliament. They belong to different parties and are appointed by the House of Commons, which also elects the chair of the Committee. The BoE's hearings typically take place when the Bank of England Inflation Report is published.[11] In these reports, the BoE explains its inflation projections on which the BoE's Monetary Policy Committee (MPC) bases its policy decisions. The report is a tool to scrutinize whether and how the BoE reaches its inflation target, which is set at 2 percent by the government (specifically by the Treasury). The BoE then discusses the Inflation Report with the Treasury Committee, which is responsible for overseeing the spending, policies, and administration of the BoE. Differently from the ECB and the Fed, the BoE Governor participate to the hearings together with other members of the MPC. While the Treasury Committee has sole statutory authority to scrutinize the BoE, also the Economic Affairs Committee of the House of Lords holds hearings with the BoE (Schonhardt-Bailey 2015; Sanders, Lisi, and Schonhardt-Bailey 2018). The textual data we collect is, however, dominated by hearings before the House of Common's Treasury Committee, which are 58, against only 8 hearings before the House of Lords' Economic Affairs Committee, which are the only available transcripts online for the period of our study. We include both sets of hearings, as the separation of tasks between the two committees is "not necessarily clear," as argued by Russell (2013). However, Russell (2013) also notes that while the Treasury Committee is officially responsible to hold the BoE accountable for its policy, the Economic Affairs Committee focuses more on issues related to administration, clarification, and simplification. Our database on the BoE comprehends 66 transcripts of the hearings from 1999 to 2018, including the mandates of three governors, namely

[11] The BoE is required to publish a report on inflation by Art. 18.2b of the Bank of England Act.

those of Edward George (1993–2003), Mervyn King (2003–13), and Mark Carney (2013–20).

European Central Bank. The ECB's accountability obligations are set out explicitly in primary EU law. Article 284(3) of the Treaty on the Functioning of European Union (TFEU) and Article 15.3 of the Statute of the European System of Central Banks and of the European Central Bank provide that the ECB is primarily accountable to the European Parliament, as the representative of EU citizens. A cornerstone of this accountability framework is the "monetary dialogue," i.e., the ECB President's participation in the regular public quarterly hearings before the Committee on Economic and Monetary Affairs (ECON committee), where he delivers a statement on the ECB's actions and answers questions from members of the European Parliament (MEPs) attending the hearing. The members of the ECON Committee are MEPs appointed by the political groups and the non-attached members of the European Parliament. All political groups are represented in ECON, as the committees are required to reflect as far as possible the political composition of the parliament.[12] Moreover, MEPs are from different EU member states, including those countries which are not part of the euro. Our text data for the case of the ECB hence relies on the transcripts of the monetary dialogues for the period 1999–2018. This time span covers three ECB presidencies, including those of Wim Duisenberg (1998–2003), Jean-Claude Trichet (2003–11), and Mario Draghi (2011–19).

Federal Reserve. The Fed is accountable to the public and the U.S. Congress. Although the formalization of the hearings took place in the Humphrey-Hawkins Act in 1978 (Full Employment and Balanced Growth Act of 1978 (P.L. 95-523)), the Fed appeared before Congress since 1976. The Federal Reports Elimination and Sunset Act of 1995 provided for the cessation of the legal requirements for the Humphrey-Hawkins Act reports to Congress after 1999, but the Fed and Congress agreed to continue their reporting arrangements (Schonhardt-Bailey 2013). According to these practices, the Chair of the Board of Governors of the Federal Reserve System appears each

[12]Pursant of Rule 209 of the Rules of Procedure of the European Parliament: https://www.europarl.europa.eu/doceo/document/RULES-9-2019-07-02_EN.pdf.

year twice before the Senate Committee on Banking, Housing, and Urban Affairs and twice before the House Committee on Financial Services. In such hearings the Fed Chairman reports to Congress on its semiannual Monetary Policy Report, which focuses on recent economic developments and on the Fed's plans for monetary policy, and replies to congressmen's questions. Each committee is composed of a chairman, who is generally the majority party member with the greatest seniority, a vice-chairman, and a ranking member, the latter being the most senior member from the opposition party. In the practice of recent years, the assignment of congressmen to the committee takes place during party conferences, where each conference prepares a roster of party members.[13] Our database for the Fed consists therefore of four hearings per year, two before the Senate and two before the House, from 2000 to 2018, covering the chairmanships of Alan Greenspan (1987–2006), Ben Bernanke (2006–14), Janet Yellen (2014–18), and Jerome Powell (2018–). A part of the oversight hearings, Fed Chairmen appear before Congress for reconfirmation hearings. This was the case for Volcker (1983), Greenspan (1992, 1996, 2000, 2004), and Bernanke (2009). However, also in this case we comprehend in our textual database only semiannual hearings to ensure consistency.

4. Methodology

We apply topic and sentiment analysis to the transcripts of central banks' parliamentary hearings in order to capture, respectively, the focus and the tone of the discussions. In this section we first briefly describe the text data preprocessing and then the text analysis methodology we implement.

4.1 Text Data and Preprocessing

For each central bank we collect the transcripts of their parliamentary hearings from 1999 to 2019, which are available in all three cases on the websites of the respective parliaments. In all three cases, transcripts are available in English. However, 10 of the transcripts of the

[13]For more details, see https://www.senate.gov/artandhistory/history/common/briefing/Committees.htm.

Table 1. Data Description of the Transcripts for the ECB, Fed, and BoE Hearings

	ECB	Fed	BoE
Number of Transcripts	81	64	66
Average Number of Words per Transcript	6,783	14,647	8,366
Total Number of Words	549,423	937,408	552,156

Note: Values relative to the average number of words and to the total number of words refer to the transcripts after cleaning the data from stop words, numbers, and white spaces.

ECB are not available fully in English, as some parts are reported in the original language used by MEPs. We translate in English the non-English text in this subset of transcripts using Google Translate. Our method is motivated by De Vries, Schoonvelde, and Schumacher (2018) who, by comparing different translating methodologies on the corpus of debates in the European Parliament, find that Google Translate performs well for text analysis models based on bag-of-words, as the ones we intend to apply.

Then, we preprocess the text in each transcript. This implies tokenizing the text, i.e., splitting raw character strings into individual elements, removing English stop words (e.g., “the,” “for,” “and”), numbers, punctuation, and white spaces. Text preprocessing is a common method in text analysis to reduce the data dimensionality, which is beneficial for both the computation and the interpretability of the model (Gentzkow, Kelly, and Taddy 2019). Descriptive statistics of the three databases following the preprocessing are summarized in Table 1.

4.2 Topic Analysis

First, we use topic analysis to investigate whether central banks and parliamentarians focus on the central bank’s objectives during the debate. We apply a dictionary technique, which consists in creating a list of key words related to a specific topic and in matching these words with those present in the transcripts.[14] The number of matches in each transcript is then divided by the total number of words in each transcript to avoid the score being inflated by the length of the text. In this way, we obtain a measure of the intensity of the focus on a specific topic at transcript level based on the frequency of key words for each document.

We create multiple text bags to account for different topics. To investigate whether parliamentary debates actually focused on the central banks' monetary policy objective(s), we first create a list of key words related to the topic of price stability, which is a primary objective for all three central banks. The advantage of applying this method to the cases of the ECB and of the BoE is that they both have a clearer prioritization of price stability as their primary objective. To compare price stability with the evolution of other topics, we create two other lists of text related to major topics of discussions, namely employment, which is the other primary objective of the Fed and a highly relevant macroeconomic variable, and financial stability. All the key words selected for the three lists are available in Section A.3 of the appendix.

The lists on price and financial stability are based on common English words related to the two topics and which abstract from the specific language features of each country. They hence have the advantage of being applicable to transcripts in English of other central banks, providing an overview of the evolution of topics in other countries. The cost associated to generality stems from the omission of those words used to address central-bank-specific monetary policy programs of the three central banks (e.g., the term "APP" that refers to the ECB's Asset Purchase Programme). However, since our aim is to compare the discussion around price stability across central banks and over time, a parsimonious and general dictionary better suits the purposes of our research question.

The dictionary approach is also more suitable to our database than the Latent Dirichlet Allocation (LDA) approach introduced by Blei, Ng, and Jordan (2003). LDA is an unsupervised method that proved successful in extracting the topics in different types of texts of individual central banks (Hansen and McMahon 2016; Hansen,

[14]For an application of dictionary techniques to extract the topics of central bank communication, see Hansen and McMahon (2016).

McMahon, and Prat 2017; Hartmann and Smets 2018; Hansen, McMahon, and Tong 2019). Compared with the dictionary approach, LDA has the advantage of identifying the topics discussed in each document without requiring any prior input from the researcher (a part of the selection of a fixed number of topics to identify).[15] However, since LDA defines topics based on the distribution of words across documents (i.e., each hearing in our case), it rests on the assumption that the set of terms used to discuss a certain topic is comparable across documents. While this assumption is reasonable when the textual database is composed of speeches of the same central bank (as in the literature mentioned above), it is problematic when applied to our database, where common terms related to the same topics co-occur with terms that are country specific.

For example, while the term "rate" features in all three cases as it may refer to "interest rate" or "inflation rate", in the case of the Fed it often appears as part of the trigram "federal funds rate". This leads LDA to identify a topic based on terms that co-occur with such trigram, a limitation also highlighted in Thomas, McNaught, and Ananiadou (2011). While such topic is helpful to detect when monetary policy is discussed in the Fed's hearings, it is not general enough to capture monetary policy discussions in the hearings of the BoE and ECB as well. For this reason, LDA does not identify latent topics that are general enough to allow a comparison across the three cases. The issue persists if we estimate the distribution of topics separately for each central bank. By doing so, the latent topics extracted by the LDA would not be comparable across central banks, as they would be based on different combinations of words. For these reasons, while the dictionary approach necessarily relies on a subset of terms chosen ex ante, it is preferable to LDA, as it provides comparable indicators of topics.[16]

[15] More precisely, LDA considers each document as a mixture of latent topics, where the topic distribution is assumed to have a Dirichlet prior.

[16] As a test, we apply the LDA approach to our database, setting the number of topics, K, equal to seven. We choose seven topics based on the output of the pre-estimation test developed by Cao et al. (2009), which selects the optimal number of topics based on topic density (formally, it identifies the K that minimizes the average cosine distance of topics). Before running the pre-estimation test and the LDA, we remove the names of the heads of the three central banks, to reduce

We estimate the following linear regression in order to identify which factors are more likely associated with changes in the focus on the central bank objective:

$$Y_{it} = \alpha + \delta O_i + \lambda(|\pi_{it} - \pi^*_{it}|) + \gamma[(|\pi_{it} - \pi^*_{it}|) \times D_{it}] + \eta E_{it} + \zeta \mathbf{X}_{it} + \phi \mathbf{W}_{it} + e_{it},$$

where Y_{it} is the score of a topic text bag for central bank i during hearing t. Since we aim to see whether the objective of the central bank is a relevant driver of the debate on a specific topic, we include a dummy O_{it} which equals 1 if i has O as main statutory objective at time t. In our main specification Y_{it} is the topic of price stability and O_i equals 1 for the cases of the BoE and for the ECB. If the objective of the central bank is a crucial driver of the focus on a topic, we expect the coefficient δ to be positive and significant.

As pointed out in the theoretical framework section, policy drifts can be relevant drivers of the discussion too. We therefore include $|\pi_{it} - \pi^*_{it}|$, which captures the absolute distance of the actual rate of inflation, π, from the targeted rate of inflation, π^*, which we set equal to 2 percent, as it approximates the aim of all three central banks.[17] Importantly, we look at the absolute distance between the two values to account for both inflationary and deflationary deviations from the aim.

The relevance of policy drifts for the topic of discussion could vary depending on whether deviations of inflation from π^* are positive (i.e., inflationary) or negative (deflationary). To capture this difference, we interact $|\pi_{it} - \pi^*_{it}|$ with a dummy, D_{it}, that equals 1 when the rate of inflation is higher than the 2 percent target, and 0 otherwise. The interaction allows us to analyze how the relationship between the focus of the hearings and the policy drift changes when inflation deviates from π^* above ($D_{it} = 1$) and below ($D_{it} = 0$) the target.

the probability of generating central-bank-specific topics. Despite this adjustment, the latent topics identified fail to capture general subjects of discussions that are applicable to all three central banks for comparison. For brevity, we do not report here the results of the LDA estimation, which are available upon request.

[17] The inflation target of the BoE, as set by the British government, and of the Fed, as set by the FOMC, is 2 percent. The ECB aims at inflation rates close to but below 2 percent.

E_{it} is a dummy equal to 1 if hearing t precedes an election in the country of central bank i. For the case of the BoE we look at general elections, for the ECB at European elections,[18] and for the Fed at presidential elections.

X_{it} is a vector of macroeconomic controls including unemployment, GDP growth, and credit-to-GDP, which is a good proxy for financial stability (Schularick and Taylor 2012).[19] In particular, we employ quarterly data on total credit to private non-financial sector in the United Kingdom, the euro area, and the United States. This variable displays a strong correlation with the scores of our financial stability text bag, as shown in Figure A.3 in the appendix.

W_{it} is a vector of text-based variables including uncertainty and a text-based index of hawkish-dovish ratio. Our measure of uncertainty is similar to the one built by Baker, Bloom, and Davis (2016) and is based on the matches of the terms "uncertainty(-ies)" and "uncertain", which are then weighted by the number of words in the text. The hawkish-dovish ratio is taken from Apel and Blix-Grimaldi (2012) and is detailed in Section 5.

4.3 Sentiment Analysis

We apply a similar methodology to measure the tone of hearings. Following the literature on sentiment analysis applied to texts, it is possible to obtain a quantitative estimate of the tone of a document by matching the words in the text with predefined lists of positive and negative terms (Loughran and McDonald 2011; Kearney and Liu 2014).

Differently from the topic analysis, in this case we do not create our own dictionary, but rely on the lists of positive and negative

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[18] We look at European elections since they are the elections for the legislators involved in the parliamentary hearings of the ECB.

[19] While credit growth is a good predictor of financial crises (Schularick and Taylor 2012), we acknowledge that there can be other measures to proxy for financial stability, such as the occurrence of a systemic crisis in a specific year (Laeven and Valencia 2012) or bank-level indicators (e.g., non-performing loans, Tier 1 capital, and so on). Data on credit growth have the advantage of being at quarterly level, differently from data on crises which are on a yearly basis, and of being harmonized and adjusted for breaks by the Bank for International Settlements, differently from bank-level data which often refer to different accounting standards and cannot always be compared across countries.

sentiments created by Hu and Liu (2004) (HL, henceforth). The lists contain 2,006 positive terms and 4,791 negative terms. We choose this lexicon instead of other sentiment dictionaries, such as the Harvard General Inquirer Dictionary (GI) used by Tetlock (2007) and the lexicon built by Loughran and McDonald (2011) (LM), for two main reasons.

First of all, HL has a predictive accuracy on economic texts that is comparable to LM and higher than GI, as found by Shapiro, Sudhof, and Wilson (2019). By evaluating the performance of GI, LM, and HL on a database of economic and financial news and comparing the scores of each dictionary with the human ratings on the same articles, they find that LM and HL lexicons have a similar rank correlations with human ratings and that are larger than the correlation of the GI lexicon.

Second, HL contains a larger number of terms and of terms that are unique compared with the other two (Shapiro, Sudhof, and Wilson 2019). This is not an advantage per se. In fact, the smaller size of LM is related to the fact that it is built specifically for the economic and financial domain, as it uses words extracted from the annual reports that U.S. firms submit to the Securities Exchange Commission to summarize their financial performance. On the other hand, the terms in HL are extracted from a feature space of movie reviews, and have therefore the disadvantage of not being specific to economics. However, the specificity of LM is not necessarily beneficial for our application. Since LM terms derive from companies' reports, the sentiments they report in that context do not necessarily fit the context of the hearings. For example, "persistent", which does not have a necessarily negative connotation in parliamentary hearings, features in the negative list in LM, whereas it does not feature in the HL dictionary. Second, LM may not be able to capture the wide range of lexicon, or sentiments, that populate parliamentary debates. For example, in one hearing a parliamentarian blames the central bank for "blackmailing" his jurisdiction. The term "blackmail-" is not present in LM, which therefore does not assign any score to this word, whereas HL assigns a negative score to it. Moreover, an additional benefit of HL, which derives from its construction, is that it relies on more robust sentiment scores, as they are extracted from the rating assigned by the reviewers on their own reviews.

As the HL text bags have been created externally to evaluate tones, they do not necessarily fit with the lexicon adopted for parliamentary debates. For this reason we removed some terms that did not match with positive or negative tones in the specific context of parliaments. For instance, we remove "accommodative" from the positive text bag, as such term has a descriptive connotation when referring to monetary policy, and not necessarily a positive one as in common texts. Following our changes, the list of positive words amounts to 1,968 terms, whereas the list of negative ones amounts to 4,782.

Then, we compute positive and negative scores based on the count of words matched with each bag in each transcript. Once we have obtained these scores, we take the difference between positive and negative terms, to get an estimate of net sentiments (Twedt and Rees 2012). Moreover, we weight net sentiments by the total number of terms in each transcript, to prevent the length of hearings from inflating sentiments upward or downward due to a larger number of terms rather than due to the intensity of the tones. A similar sentiment ratio is proposed in Nyman et al. (2018) and Shapiro, Sudhof, and Wilson (2019), with the difference that the latter subtract matches of terms related to excitement to those related with anxiety to capture sentiments shifts in financial markets. Formally, for each transcript t associated with each central bank i we compute the following ratio:

$$SentimentRatio_{it} = \frac{|Positive_{it}| - |Negative_{it}|}{N_{it}},$$

where $Positive_{it}$ and $Negative_{it}$ are the number of terms matched in each transcript and N_{it} is the total number of words in each transcript. As pointed out by Shapiro, Sudhof, and Wilson (2019), one advantage of this approach is that it is simple and transparent. In addition, they note that this approach is mathematically equivalent to assigning a score of 1 to positive matches and a score of –1 to negative matches and averaging the word-specific valence scores across all words in a text.

Other works propose a different sentiment ratio, where the number of matches per sentiment is weighted by total sum of matches of both sentiments and add unity to get rid of negative values (Birz and Lott 2011; Apel and Blix-Grimaldi 2012). For robustness, we compute an alternative estimate of sentiment ratio based on this methodology. In particular, we estimate the following equation:

$$SentimentRatio_{it} = \left[\frac{Positive_{it}}{Positive_{it} + Negative_{it}} - \frac{Negative_{it}}{Positive_{it} + Negative_{it}}\right] + 1.$$

For simplicity, in the rest of the paper we discuss sentiment ratio referring implicitly to the first measure. We provide the results for the alternative measure for sentiment ratio in the appendix.

We estimate a similar regression model to the one used for topics:

$$SentimentRatio_{it} = \alpha + \lambda(|\pi_{it} - \pi^*_{it}|) + \gamma[(|\pi_{it} - \pi^*_{it}|) \times D_{it}] + \eta E_{it} + \zeta \mathbf{X}_{it} + \rho \mathbf{V}_{it} + \mu_i + e_{it}.$$

This model differs from the one used for topics in three main aspects. First, in this model we do not include the objective dummy as an explanatory variable for sentiments. Second, we now include central bank fixed effects, which are captured by μ_i, that we did not include in the topic model to avoid collinearity with the objectives' dummies.

Third, in this equation we replace the text-based indicator of uncertainty (described in Section 4.2) with a vector of variables that captures macroeconomic uncertainty, V_{it}. This choice is motivated by the concern that the text-based indicators of sentiments and uncertainty rely on similar key terms and may hence capture the same phenomenon. To circumvent this problem, we replace the text-based measure of uncertainty with the volatility of inflation, unemployment, and growth, which can be considered as exogenous to the sentiment index. In line with previous studies (Judson and Orphanides 1999; Caglayan and Xu 2016), we take the within-year variance of each variable to measure its volatility.

As a robustness check, we estimate another model where we replace the indicators of macroeconomic volatility with the same text-based indicator of uncertainty used in the topic analysis. However, in this case we remove from the sentiment text bags those terms that feature also in the uncertainty text bag. This allows us to estimate separately the relationship between sentiments and uncertainty.

5. Results

5.1 Results on Topics

We first regress the price stability score, given by the number of matches of the price stability text bag, on the presence of an inflation objective as the sole primary objective of the central bank. As mentioned in the previous section, the inflation objective dummy equals 1 for the cases of the BoE and of the ECB.

The results of the regression are displayed in Table 2.[20] We notice that the inflation objective is positively and significantly correlated with the frequency of price stability terms. This suggests that the presence of price stability as primary statutory objective is associated with a more intense focus of the hearings on the topic of price stability. This result is significant also once we control for the divergence of inflation rates from the 2 percent aim. Moreover, the coefficient of the objective dummy remains positive and significant also once we control for uncertainty, the presence of elections, and macroeconomic factors such as unemployment, GDP, and credit.

The distance of inflation from the target shows a strong positive correlation with the focus on price stability, but only when interacted with the above target dummy. This suggests that the more inflation grows above the central bank target, the higher the attention of the hearing on the issue of price stability. The same does not hold for deflationary drifts. When inflation decreases below target—i.e., when the above target dummy is equal to 0—the correlation with price stability is not significant. Therefore, the focus on price stability increases only when inflation grows above the inflation target, and not when it decreases below the target.

Moreover, it is interesting to notice that unemployment is negatively and significantly correlated with the focus on price stability. This result may indicate that, as unemployment increases, the attention shifts away from price stability. This may also reflect that the hearings react swiftly to changes in inflation or unemployment in a

[20]We apply the variance inflation factor to detect the presence of collinearity in this and the following models. The mean variance inflation factors for the topic and sentiment regression models are 2.04 and 1.51, respectively. These results indicate that our estimates are robust to multicollinearity.

Table 2. OLS Estimates on Topic Price Stability as Dependent Variable

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Inflation Objective	1.102***	1.308***	1.653***	1.595***	1.603***	1.139***	1.033***
	(0.182)	(0.174)	(0.207)	(0.204)	(0.205)	(0.331)	(0.336)
$\|\pi - 2\%\|$		–0.325***	–0.187	–0.187	–0.192	–0.208*	–0.150
		(0.122)	(0.115)	(0.118)	(0.119)	(0.113)	(0.115)
$\|\pi - 2\%\| \times D$		0.360***	0.349***	0.352***	0.349***	0.298**	0.387***
		(0.122)	(0.120)	(0.123)	(0.123)	(0.129)	(0.135)
Unemployment Rate (log)			–1.667***	–1.555***	–1.552***	–1.090**	–0.947**
			(0.364)	(0.368)	(0.368)	(0.453)	(0.454)
Uncertainty				1.177	1.193	0.362	0.364
				(1.333)	(1.334)	(1.429)	(1.472)
Elections					0.188	0.181	0.167
					(0.173)	(0.155)	(0.161)
GDP (log)						–0.280*	–0.395**
						(0.166)	(0.171)
Credit-to-GDP (%)							–0.010*
							(0.005)
Observations	162	151	151	151	151	151	151
R-squared	0.116	0.229	0.352	0.357	0.358	0.370	0.382
Central Bank FE	No	No	No	No	No	No	No

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

"Phillips curve" fashion, decreasing the attention on price stability when unemployment grows and inflation decreases, and vice versa.

It is not clear whether the mandate of the BoE foresees a hierarchy between the price stability and financial stability objectives. While this might seem surprising, there are a number of other cases where the subordination is not specified by the law, as documented in a survey of 114 central bank statutes by (Jeanneau 2011). To account for this issue, we provide a new specification, where the inflation objective dummy equals 1 for the whole time series if the central bank is the ECB, whereas it equals 0 for the BoE after 2011, when the BoE is entrusted the objective of financial stability.[21] The results, displayed in Table A.3 in the appendix, are robust to this specification. Both the objective dummy and the interaction between distance from inflation and the above-target dummy remain positively and significantly correlated with the focus on price stability under all specifications. The other regressors display similar coefficients to the ones of Table 2.

Another difficulty is related to the inflation target of the Fed. The FOMC stated its first formal and public commitment to an inflation target of 2 percent on January 24, 2012 (Powell and Wessel 2020).[22] For this reason, distance of inflation from 2 percent might not necessarily capture perceived drifts from the price stability objective in the Fed hearings preceding 2012. To address this limitation, we test the robustness of our results under three potential inflation targets for the case of the Fed. More precisely, we replace the inflation target of 2 percent (π^*) with three alternative targets, namely 1, 1.5, and 2.5 percent for the case of the Fed before January 2012. Accordingly, we change the above-target dummy D to fit each of these targets.

[21]This date refers to the Financial Services Act 2012, which amended the Bank of England Act 1998 (Tucker, Hall, and Pattani 2013). It is however not easy to set a precise date for the start of the BoE's financial stability mandate. As pointed out in Murphy and Senior (2013), the Financial Policy Committee existed in non-statutory form since 2011. Moreover, as noted by Jeanneau (2011), the details of the BoE's financial stability mandate, which is quite general in its statutory form, are spelled out in the antecedent 2009 Banking Act. We therefore adopted alternative inflation objective dummies referring to these years, finding that the results, which for simplicity we do not report in this work, do not substantially differ from the ones in Tables 2 and 9.

[22]The FOMC announced the target via the "Statement on Longer-Run Goals and Monetary Policy Strategy" published on January 24, 2012.

The results, which are displayed in Table A.4 in the appendix, are similar to the ones of the baseline model.[23] Both the inflation objective and distance of inflation above the inflation target are positively and significantly correlated with the focus on price stability under all three alternative targets.

These initial results suggest that the focus on price stability is motivated by accountability concerns. When price stability is a primary objective, the focus of the hearings on price stability is higher. Moreover, as inflation grows away from the target, the principal and the agent focus more on the topic of price stability.

To further test the relevance of the statutory objective in shaping the topic of the discussion, we focus on the employment objective of the Fed. We replace the dependent variable with the frequency of employment-related terms. We keep the objective dummy as the main regressor: since the Federal Reserve is the only central bank of the three that has employment (and not only price stability) among its primary objectives, this is equivalent to the inverse of a dummy that equals 1 if the central bank is the Federal Reserve.

Results are displayed in Table 3. The inflation objective dummy is negatively and significantly correlated with the focus on employment. This means that the focus on employment is higher in the hearings of the Federal Reserve than in those of the other central banks, which do not have full employment as their primary objective. This result provides further evidence of a significant and positive association between the statutory objective and the focus of the discussion on the topic of the objective.

The coefficient of unemployment may seem puzzling at a first glance: its negative sign suggests that, as unemployment grows, the focus on employment in the hearings decreases. This result may appear counterintuitive, as we might expect higher unemployment to trigger the concerns of the speakers on the subject, and not the opposite. However, this result might also indicate that the focus of the hearings on the topic of employment increases as employment grows. In other words, the negative correlation between the topic of employment and unemployment rates may mirror the

[23]Table A.4 presents the results for the models with the topic of price stability, the topic of employment, and the sentiment ratios as dependent variables.

Table 3. OLS Estimates on Topic Employment as Dependent Variable

Variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
Inflation Objective	–0.751***	–0.770***	–0.656***	–0.648***	–0.644***	–0.718***	–0.678***
	(0.102)	(0.098)	(0.104)	(0.107)	(0.107)	(0.155)	(0.160)
$\|\pi - 2\%\|$		–0.169***	–0.123***	–0.124***	–0.126***	–0.128***	–0.150***
		(0.048)	(0.045)	(0.044)	(0.044)	(0.044)	(0.048)
$\|\pi - 2\%\| \times D$		–0.140**	–0.144**	–0.144**	–0.145**	–0.153**	–0.187**
		(0.061)	(0.062)	(0.063)	(0.064)	(0.068)	(0.073)
Unemployment Rate (log)			–0.550***	–0.567***	–0.566***	–0.492***	–0.547***
			(0.133)	(0.149)	(0.149)	(0.180)	(0.183)
Uncertainty				–0.177	–0.171	–0.303	–0.304
				(0.477)	(0.481)	(0.554)	(0.546)
Elections					0.078	0.077	0.082
					(0.091)	(0.092)	(0.097)
GDP (log)						–0.045	–0.001
						(0.070)	(0.074)
Credit-to-GDP (%)							0.004**
							(0.002)
Observations	162	151	151	151	151	151	151
R-squared	0.299	0.338	0.412	0.412	0.414	0.415	0.425
Central Bank FE	No	No	No	No	No	No	No

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

positive correlation between the topic of employment and employment rates.[24]

To test this claim, we select the 50 parts of speech of the transcripts that display the highest scores in the employment dictionary, i.e., those that have the highest number of key terms that are matched with the employment dictionary.[25] By inspecting each of them, we notice that the great majority of the excerpts express concerns on a number of issues related to the growth of employment. Moreover, the minority of excerpts that focus on unemployment mostly discuss the level of unemployment or its side effects, rather than unemployment growth.[26] This evidence is informative on why discussions on employment vary as a reaction to changes in employment, providing an explanation for the negative and significant coefficient of the unemployment rate.

Overall, these results suggest that parliamentary scrutiny serves its intended role, as the statutory objective and deviations from the inflation target are among the main drivers of the discussion. The focus on price stability is higher where it represents the main statutory objective of the central bank. In line with this, the focus on employment is positively associated with the Fed dummy (i.e., the inverse of the inflation objective dummy).

Nevertheless, it should be noted that the inflation objective dummy risks being correlated with other unobserved variables and should therefore be interpreted with caution. For example, the focus of price stability could be lower in the hearings of the Fed due to

[24]If we run the same model of Table 3 and replace the rate of unemployment with the rate of employment, the coefficient of employment is positive and significant under all specifications. The estimates of this test are reported in Table A.2 in the appendix.

[25]To provide an illustration, we report the top 10 excerpts in Section A.5 of the appendix. All terms are in lowercase to allow matches with the terms in the dictionaries. The other excerpts are available upon request.

[26]More in detail, the majority of the 50 excerpts (74 percent) does not express disquietude about rising unemployment. On the contrary, the excerpts reflect concerns on a number of issues related to employment growth, such as the quality of jobs being created and how this is captured by employment statistics. Moreover, among the remaining 26 percent of speeches, the majority does not discuss raising unemployment in most instances. These speeches rather focus on the level of unemployment, or on the social and psychological side effects of high levels of unemployment.

unobserved political, economic, or cultural reasons which are specific to the United States, regardless of the central bank mandate.

While our model does not rule out this possibility, the significant coefficient of the distance between inflation and the inflation target lends support to the view that the mandate matters. In other words, central banks and parliaments increase their focus on price stability not only when this is the main objective but also when inflation drifts away from the central bank objective. However, the interaction term shows that this reaction is not symmetric, as the focus on price stability increases only when such deviations are inflationary, and not when they are deflationary.

5.2 Results on Sentiments

We now investigate which variables are associated with shifts in the sentiment ratio presented in Section 4. As discussed, sentiments can be a good proxy of the tone adopted in the hearings. For example, if the central bank is deviating from the objective assigned by its principal, we would expect the tone of the discussion to be more negative.

We first test separately the correlation between sentiments and three variables: distance from the inflation aim, unemployment, and economic uncertainty (Table 4, columns 1–3). Distance from the inflation aim above the target is significantly associated with a decrease in the sentiment ratio (Table 4, column 1). The negative coefficient of the interaction term indicates that the more central banks deviate upward from their inflation aim, the worse net sentiments become.

Once we control for economic uncertainty (inflation, unemployment, and GDP volatility), macroeconomic conditions, and elections, the coefficient for above-target deviations from inflation remains significant. The results under the alternative approach to compute sentiment ratio, shown in Table A.5 in the appendix, do not differ substantially from the ones presented in Table 4. On the other hand, the results for deflationary deviations are less robust. While also the coefficient of deflationary deviations is negative, it is (weakly) significant only once we control for macroeconomic conditions (columns 7 and 8). Furthermore, under the alternative sentiment ratio indicator, the coefficient of deflationary deviations is not

Table 4. OLS Estimates on Sentiment Ratio as Dependent Variable

Variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
$\lvert\pi - 2\%\rvert$	–0.119			–0.207	–0.214	–0.214	–0.302*	–0.265*
	(0.106)			(0.159)	(0.159)	(0.159)	(0.165)	(0.154)
$\lvert\pi - 2\%\rvert \times D$	–0.471***			–0.486***	–0.490***	–0.490***	–0.510***	–0.333**
	(0.131)			(0.161)	(0.162)	(0.162)	(0.160)	(0.156)
Unemployment Rate (log)		–0.295		–0.034	–0.036	–0.036	–0.040	0.723
		(0.378)		(0.408)	(0.412)	(0.412)	(0.400)	(0.472)
Inflation Volatility			–0.882***	–0.723**	–0.723**	–0.723**	–0.768**	–0.583*
			(0.287)	(0.289)	(0.292)	(0.292)	(0.295)	(0.297)
Unemployment Volatility			0.598	1.303*	1.299*	1.299*	1.426*	1.796**
			(0.651)	(0.770)	(0.773)	(0.773)	(0.746)	(0.740)
GDP Volatility			–0.040	–0.072	–0.079	–0.079	–0.056	–0.007
			(0.072)	(0.074)	(0.072)	(0.072)	(0.074)	(0.077)
Elections					0.164	0.164	0.149	0.076
					(0.312)	(0.312)	(0.315)	(0.303)
GDP (log)							1.331**	2.920***
							(0.524)	(0.689)
Credit-to-GDP (%)								–0.032***
								(0.008)
Observations	151	151	153	147	147	147	147	147
R-squared	0.457	0.423	0.436	0.477	0.478	0.478	0.503	0.553
Central Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: Robust standard errors are in parentheses. *p < .05; **p , .01; ***p < .001.

significant under all specifications. Overall, the result of the interaction suggests that sentiments tend to worsen following inflationary deviations from the target rather than deflationary ones.

A relevant issue regards the content captured by our measure of sentiments. Lower sentiment scores may proxy not only for a negative, aggressive, or confrontational tone in the discussion but also for the speakers' depictions of the negative economic outlook or uncertainty.

To account for this issue, we include among the regressors the text-based measure developed by Apel and Blix-Grimaldi (2012) to capture hawkish and dovish stances in monetary policy. Their index is built on two dictionaries able to capture hawkish and dovish stances on monetary policy. Applied to the minutes of the monetary policy meetings of the Swedish central bank, these measures proved useful to predict future policy rate decisions (Apel and Blix-Grimaldi 2012). Their dictionaries are therefore helpful to disentangle sentiments from negative economic considerations. We first remove from the sentiment dictionaries those terms that also feature in the hawkish and dovish dictionaries. Then, we apply the same dictionaries to the transcripts and obtain two scores capturing the degree of hawkish and dovish sentiments of each hearing. From these scores, we extract a hawkish-dovish ratio, based on the difference between the hawkish and dovish score divided by the number of total words in the transcript, similarly to the sentiment ratio.[27]

Table 5 displays the results including the hawkish-dovish ratio. The hawkish-dovish ratio is positively correlated with sentiments, suggesting that more hawkish stances are associated with a more positive tone during the hearing. This relationship can be explained by the fact that hawkish policy stances are generally associated with periods of economic growth, and therefore of positive economic conditions. However, the coefficient is not significantly correlated with sentiments. The inclusion of the hawkish-dovish ratio does not affect the sign nor the significance of the coefficients of the indices of policy drift. Inflationary drifts from target remain negatively correlated with sentiments under all specifications.

[27]This measure could also be defined as "net hawkishness," as suggested by Apel and Blix-Gimaldi (2012).

Table 5. OLS Estimates on Sentiment Ratio (adjusted to Hawkish-Dovish terms) as Dependent Variable

Variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**	**(8)**
Hawkish-Dovish Ratio	0.101	0.128	0.126	0.107	0.106	0.106	0.139	0.089
	(0.084)	(0.089)	(0.092)	(0.101)	(0.100)	(0.100)	(0.098)	(0.099)
$\lvert\pi - 2\%\rvert$		–0.079	–0.074	–0.157	–0.165	–0.165	–0.225	–0.220
		(0.109)	(0.114)	(0.163)	(0.163)	(0.163)	(0.164)	(0.157)
$\lvert\pi - 2\%\rvert \times D$		–0.490***	–0.486***	–0.489***	–0.494***	–0.494***	–0.512***	–0.340**
		(0.126)	(0.129)	(0.154)	(0.155)	(0.155)	(0.153)	(0.150)
Unemployment Rate (log)			–0.086	0.024	0.020	0.020	0.034	0.753
			(0.382)	(0.399)	(0.403)	(0.403)	(0.393)	(0.455)
Inflation Volatility				–0.687**	–0.687**	–0.687**	–0.722***	–0.550*
				(0.273)	(0.275)	(0.275)	(0.276)	(0.285)
Unemployment Volatility				0.992	0.989	0.989	1.025	1.507**
				(0.784)	(0.784)	(0.784)	(0.747)	(0.756)
GDP Volatility				–0.064	–0.072	–0.072	–0.052	–0.002
				(0.073)	(0.070)	(0.070)	(0.072)	(0.074)
Elections					0.188	0.188	0.172	0.104
					(0.291)	(0.291)	(0.290)	(0.280)
GDP (log)							1.240**	2.753***
							(0.517)	(0.661)
Credit-to-GDP (%)								–0.032***
								(0.008)
Observations	162	151	151	147	147	147	147	147
R-squared	0.429	0.471	0.471	0.487	0.488	0.488	0.511	0.560
Central Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

As a further test, we replace the measures of economic uncertainty with the broader text-based indicator of uncertainty used in the topic analysis. To this end, we remove from the negative text bag those terms that overlap with the uncertainty indicator. This test is motivated by the works of Baker, Bloom, and Davis (2016) and of Istrefi and Piloiu (2020), who find that text-based measures of uncertainty are correlated with factors that are not strictly economic, such as major (geo)political events, and lower public trust toward central banks.[28] As monetary policy is generally effective in reducing uncertainty in the markets (Bekaert, Hoerova, and Lo Duca 2013), it is possible that high uncertainty is seen by some parliamentarians as the sign that central banks are not doing enough. For these reasons, uncertainty may play a relevant role in worsening sentiments.

Results are presented in Table 6. Uncertainty displays a negative and significant correlation with sentiments under all specifications. Nevertheless, the sign and significance of inflationary deviations remain robust to the inclusion of uncertainty. This evidence strengthens the finding that sentiment reacts negatively to inflationary deviations from the central bank mandate.

6. Conclusion

Our results suggest that parliamentary hearings overall serve their intended purpose. We show that the hearings tend to focus on the statutory objective of the central bank and that this focus increases as the central bank deviates from the mandate. Moreover, we find that the tone of the debates worsens when inflation surpasses the central bank's inflation aim. However, topics and sentiments react more to inflationary rather than deflationary deviations of inflation away from target. This evidence is particularly relevant in light of

[28] Examples of geopolitical events highlighted by Baker, Bloom, and Davis (2016) include the 9/11 terrorist attack for the United States, and the Scottish independence referendum and the Brexit referendum for the United Kingdom. The peak for the Brexit referendum is documented in the updated index provided by the authors at this link: https://www.policyuncertainty.com/uk_monthly.html. Based on data on the BoE, the ECB, and the Bank of Japan, Istrefi and Piloiu (2020) show that shocks in uncertainty deteriorate public trust toward central banks.

Table 6. OLS Estimates on Sentiment Ratio (negative terms adjusted to uncertainty) as Dependent Variable

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
$\|\pi - 2\%\|$	−0.119			−0.130	−0.133	−0.133	−0.213*	−0.161
	(0.106)			(0.114)	(0.114)	(0.114)	(0.118)	(0.110)
$\|\pi - 2\%\| \times D$	−0.471***			−0.504***	−0.505***	−0.505***	−0.528***	−0.316**
	(0.131)			(0.135)	(0.135)	(0.135)	(0.134)	(0.139)
Unemployment Rate (log)		−0.280		−0.095	−0.094	−0.094	−0.122	0.729
		(0.376)		(0.389)	(0.391)	(0.391)	(0.373)	(0.455)
Uncertainty			−2.648**	−3.056***	−3.043***	−3.043***	−2.733***	−2.988***
			(1.158)	(1.124)	(1.127)	(1.127)	(1.034)	(0.966)
Elections					0.120	0.120	0.123	0.075
					(0.298)	(0.298)	(0.297)	(0.300)
GDP (log)							1.380***	2.938***
							(0.504)	(0.631)
Credit-to-GDP (%)								−0.033***
								(0.007)
Observations	151	151	151	151	151	151	151	151
R-squared	0.456	0.421	0.436	0.478	0.479	0.479	0.507	0.559
Central Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

the criticism that elected representatives tend to focus on topics other than the central bank's performance in fulfilling its mandate, which should be the object of scrutiny.

In exploring these questions, we introduce a new empirical approach to study central bank accountability practices based on text analysis. We apply this method on a novel database, which consists of the parliamentary hearings of the BoE, the ECB, and the Fed. The dictionary-based techniques that we present in this paper can be extended to other central banks in order to track the topic and tones of their parliamentary hearings. This approach opens new avenues for the research on central bank accountability, which so far has been largely dominated by theoretical or qualitative considerations.

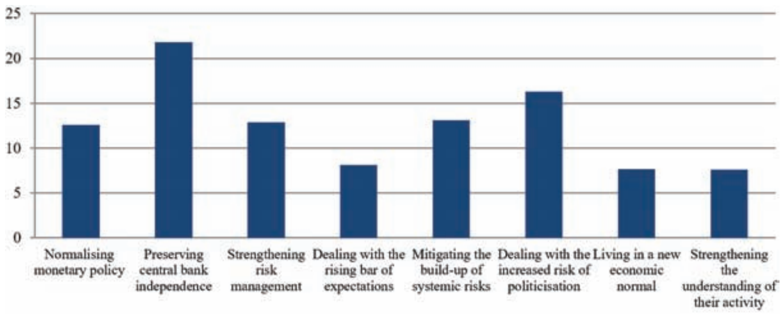
Looking forward, future works could improve this method by overcoming its existing limitations. While our approach is effective in identifying changes in the topics and sentiments in parliamentary hearings, it analyzes accountability based on the hearing as unit of analysis. This provides us with partial information on how accountability is discharged, since it presents the interactions between central banks and parliaments as a single block. In future research, we intend to develop further this rich database and look at whether shifts in sentiments are mainly driven by the central bank or by parliamentarians. This would allow us to study whether parliamentarians' individual characteristics play a role in explaining the tone and focus of their participation in hearings. For example, by analyzing news on the pressures from the governments on the central banks in a number of countries, Binder (2021) finds that pressures are more likely when the executive is left-wing or nationalist. It is worth exploring whether this applies also to the context of parliamentary hearings, where—differently from the approach based on governments—it is possible to compare how different parties interact simultaneously with the central bank. Therefore, while our contribution already provides new insights on central banks' parliamentary hearings, it also opens promising avenues for further research.

In conclusion, our work sheds new light on the central banks' parliamentary hearings, a key accountability practice, as an unexplored but rich source of data. Fraccaroli, Giovannini, and Jamet (2018) provide evidence of how other types of accountability practices could be exploited to obtain quantitative estimates of the evolution of accountability. Some of them, such as the written questions that parliamentarians address to the central bank, can be potentially assessed through text analysis tools.

Appendix

A.1 Survey Results

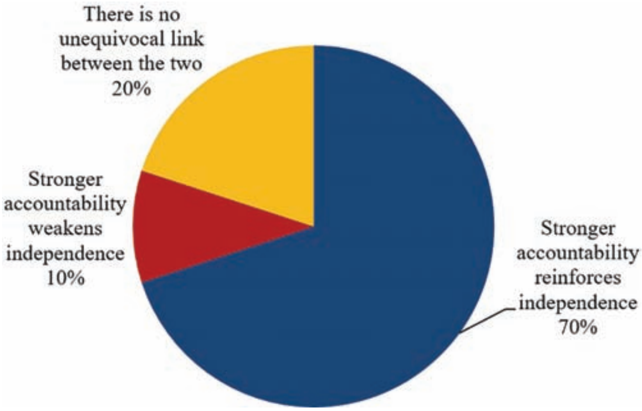
Figure A.1. The Main Challenges for Central Banks in 2019



Sources: Authors' elaboration on a survey conducted in January 2019 among 30 central bank staff working on institutional matters in their respective central banks.

Note: The following central banks participated in the survey: Central Bank of Malta, Central Bank of Luxembourg, Reserve Bank of Australia, Bank of Mexico, Federal Reserve, European Central Bank, Bank of Ghana, Central Bank of Ireland, Bank of Estonia, Croatian National Bank, National Bank of Ukraine, Central Bank of Norway, Danmarks Nationalbank, Central Bank of Brazil, Swiss National Bank, Sveriges Riksbank, National Bank of Belgium, Bank of Portugal, Deutsche Bundesbank, Netherlands Bank, Central Bank of Cyprus, Bank of England. Moreover, representatives of the International Monetary Fund (IMF) and the Financial Stability Board (FSB) also participated in the survey.

Figure A.2. The Relationship between Central Bank Independence and Accountability



Sources: Authors' elaboration on a survey conducted in January 2019 among 30 central bank staff working on institutional matters in their respective central banks.
Note: The following central banks participated in the survey: Central Bank of Malta, Central Bank of Luxembourg, Reserve Bank of Australia, Bank of Mexico, Federal Reserve, European Central Bank, Bank of Ghana, Central Bank of Ireland, Bank of Estonia, Croatian National Bank, National Bank of Ukraine, Central Bank of Norway, Danmarks Nationalbank, Central Bank of Brazil, Swiss National Bank, Sveriges Riksbank, National Bank of Belgium, Bank of Portugal, Deutsche Bundesbank, Netherlands Bank, Central Bank of Cyprus, Bank of England. Moreover, representatives of the IMF and FSB also participated in the survey.

A.2 Indices of Central Bank Accountability

Measuring central bank accountability empirically is challenging. As central banks are institutions embedded in their specific political and legal national context, they are characterized by different governance traits and legal foundations (constitutions, central bank statutes, additional regulations, etc.) that make cross-country comparisons more difficult (see Frisell, Roszbach, and Spagnolo 2008, Hasan and Mester 2008, and Bank for International Settlements 2009).

Despite these differences, some works identify a number of common criteria to evaluate the statutory accountability of central banks. Building on the example of the widely diffused central bank

Table A.1. Overview of the Most Widely Used Measures of Central Bank Accountability

Authors	Type of Index	Aspects Covered by the Index
Briault, Haldane, and King (1996)	Binary	• External monitoring by parliament • Minutes of the meetings are published • Inflation or monetary policy report are published • Government can override a decision of the central bank
De Haan, Amtenbrink, and Eijffinger (1999)	Binary	• Clarity of the monetary policy objective (e.g., quantification of the objective) • Transparency of monetary policy (e.g., publication of inflation or monetary reports) • Final responsibility of monetary policy (e.g., central bank law can be changed by simple majority)
Bini-Smaghi and Gros (2000)	Binary	• Ex ante accountability (e.g., definition of the central bank objectives) • Ex post accountability (e.g., public hearings and meetings) • Procedures (transparency of the central bank vis-à-vis the parliament and the public)

independence indexes (such as those developed by Bade and Parkin 1988; Alesina 1989; Grilli, Mascandiaro, and Tabellini 1991; Cukierman, Webb, and Neyapti 1992), researchers constructed accountability indices based on central banks' legal frameworks (Briault, Haldane, and King 1998; De Haan, Amtenbrink, and Eijffinger 1999; Bini-Smaghi and Gros 2000). These indices, summarized in Table A.1, are constructed by selecting a number of common criteria that are applicable to the statutes of most, if not all, central banks.

One limitation of these measures is their low time-variation, due to the dependence of their variability on reforms in central bank laws. For example, in its 20 years of history, the ECB has experienced no change in its accountability indices, as the relevant statutory provisions for its central banking role (i.e., excluding banking supervision) have not been reformed. The same applies to other central banks (De

Grauwe and Gros 2008) once supervisory functions, which are not the focus of this paper, are excluded.

While the stability of the indices through time cannot be considered a problem per se, as it still offers a useful cross-country comparison, it does not provide information on the continuous evolution and changes in the interactions between central banks and their principal. In other words, an analysis based on indices provides essential insights on the de jure setting of the accountability framework defined in the contact between the principal and the agent; nevertheless, it is silent on the way in which the agent de facto discharges its accountability over time and how the principal reacts to that.

In this context, it is interesting to note that while the Bank of Japan is assessed by CBA indices as the least accountable central bank when compared with the FED, the BoE, and the ECB (De Grauwe and Gros 2008), it is one of the central banks that has held by far the highest number of parliamentary appearances for accountability reasons. In 2005 and 2006, the Bank of Japan appeared before the Diet (the Japanese parliament), respectively, 33 and 35 times, hence more frequently than the Fed's appearances (21 and 15 times, respectively) and the ECB's (5 times in both years) (Shirakawa 2008; Heckel 2014).

A.3 Text Bags for Topic Analysis

Price Stability:

- price(s), inflate, inflation, inflationary, HICP, CPI, PCE, PCE index, PCE inflation, deflation, deflator, deflationary, deflate, hyperinflation, hyperinflationary.

Employment:

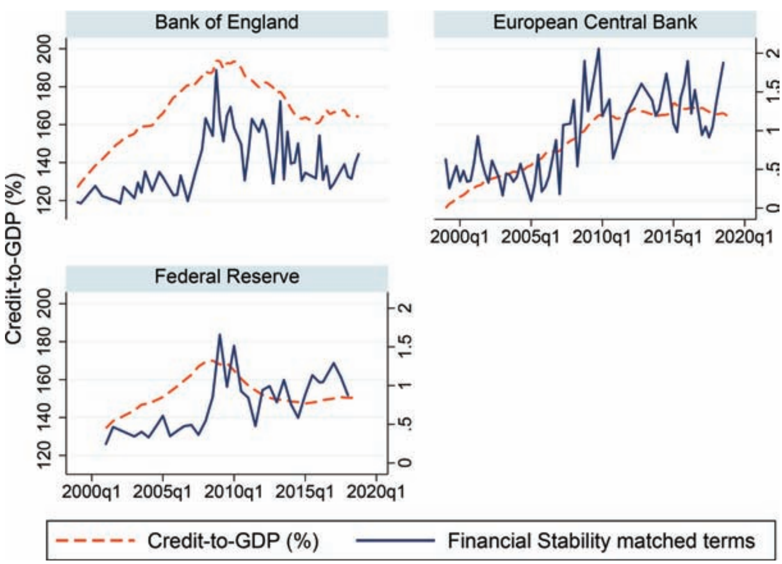
- employ(-ee/-er), (un)employment, underemployment, firing, fixed-term, full-time, part-time, inactivity, job(s), jobless, labo(u)r, labo(u)r force, labo(u)r market, self-employed, temporary, vacancy(-ies), work(er), workers, working, working (age/time), works.

Financial Stability:

- financial (in)stability, bank (in)stability, (financial) crisis, financial stress, financial risk, systemic risk, contagion, financial shocks, bubble, financial imbalance, misalignment, credit growth, banks, insurers, hedge funds, investment funds, financial markets, securities markets, leverage, capital, derivatives, off-balance sheet exposures, special purpose vehicles, off-balance sheet vehicles, payment systems, settlement systems, central securities depositories, non-performing loans, npls, non-performing exposures, foreign currency loans, correlated exposures.

A.4 Financial Stability and Credit Growth

Figure A.3. Credit-to-GDP and Focus on Financial Stability in the Parliamentary Hearings by Central Bank



Note: Credit-to-GDP is total credit to private non-financial sector, using Bank for International Settlements data. Data for the ECB refer to the euro area.

A.5 Robustness Check for the Employment Topic

A.5.1 Top 10 Speeches with the Highest Scores in the Employment Dictionary

Bank of England, 2006: "for example, the measure of unemployment on the claimant count may be an accurate measure of the unemployment statistics for those claiming unemployment benefit, but ignores the large chunk of migrant workers who, at least not yet, have qualified for unemployment benefit, and may be understating unemployment a little bit."

Federal Reserve, 2009: "the longer-run projections for output growth and unemployment may be interpreted as the committee's estimates of the rate of growth of output and unemployment that are sustainable in the long run in the united states, taking into account important influences such as trend growth rates of productivity and the labor force improvements in worker education and skills, the efficiency of the labor market and matching workers in jobs, government policies affecting technological development or the labor market and other factors."

Federal Reserve, 2013: "unemployment is one problem, but long-term unemployment and underemployment—and by "underemployment," i mean people who are either working fewer hours than they would like or possibly are working at jobs well below their skill level—are also indicative of a weak labor market."

Federal Reserve, 2016: "so there are an enormous number of job openings, and there is a certain degree of mismatch of workers who are looking for work with the job openings that are available within the federal reserve, and i personally have been looking at workforce development programs, job training programs, some of which i think are doing a very good job of trying to build the skills and that are needed to fill available jobs and work to match workers with jobs."

Bank of England, 2013: "i would remind you that there are a very large number of workers, who have jobs, who want to work much more than they are working today, and i would not underestimate the extent to which there is slack not just of people looking for a job."

Federal Reserve, 2004: "contracted workers will either be in the payroll series itself—you will remember that temporary employment is part of the payroll series, and the number of contractual types of work are there as well—but to the extent that they are proprietors, or they are essentially self-employed, they will be picked up in the denominator of the productivity estimate largely from the household survey data, which is really the sole source of self-employed."

Federal Reserve, 2004: "if you include workers who are working part-time because they cannot find a full-time job and workers who want to work but are not in the labor force, the unemployment rate is roughly 9.6 percent."

Bank of England, 2013: "what i said in my speech was that i do believe that the nature of the way in which employers and employees work together within the labour market has changed, and that one of the values employers are increasingly recognising is of the longevity of some of their employees, that they gain productivity and they gain skills through working with individual companies over a longer time frame, which makes employers reluctant to lose those employees, simply because short-term demand conditions are less than they would like, or certainly deficient in this case."

Federal Reserve, 2014: "probably the unemployment rate is the single best indicator, and it has come down to 6.1 percent, which is really notable progress, and broader indicators that include marginally attached workers, discouraged workers, and those with involuntary unemployment, parttime employment, those have come down as well."

Federal Reserve, 2015: "in addition, long-term unemployment has declined substantially, fewer workers are reporting that they can find only part-time work when they would prefer full-time employment, and the pace of quits—often regarded as a barometer of worker confidence in labor market opportunities—has recovered nearly to its pre-recession level."

A.5.2 Focus on Employment and Employment Rate

Table A.2. OLS Estimates on Topic Employment as Dependent Variable and Employment Rate as Regressor

Variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
Inflation Objective	–0.751***	–0.770***	–0.693***	–0.674***	–0.670***	–0.568***	–0.567***
	(0.102)	(0.098)	(0.102)	(0.104)	(0.105)	(0.182)	(0.191)
$\lvert\pi - 2\%\rvert$		–0.169***	–0.161***	–0.163***	–0.165***	–0.169***	–0.168***
		(0.048)	(0.043)	(0.042)	(0.042)	(0.043)	(0.048)
$\lvert\pi - 2\%\rvert \times D$		–0.140**	–0.186***	–0.190***	–0.191***	–0.189***	–0.189**
		(0.061)	(0.062)	(0.064)	(0.065)	(0.067)	(0.072)
Unemployment Rate (log)			3.164***	3.440***	3.440***	4.182***	4.197**
			(0.779)	(0.930)	(0.931)	(1.492)	(1.707)
Uncertainty				–0.421	–0.416	–0.318	–0.318
				(0.522)	(0.527)	(0.553)	(0.552)
Elections					0.086	0.089	0.089
					(0.094)	(0.094)	(0.095)
GDP (log)						0.062	0.062
						(0.089)	(0.090)
Credit-to-GDP (%)							–0.000
							(0.002)
Observations	162	151	150	150	150	150	150
R-squared	0.299	0.338	0.418	0.420	0.422	0.424	0.424
Central Bank FE	No	No	No	No	No	No	No

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

A.6 Robustness Check for the Price Stability Topic

Table A.3. OLS Estimates on Topic Price Stability as Dependent Variable and with a Second Version of the Inflation Objective Dummy

Variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Inflation Objective v2	0.773***	1.062***	1.709***	1.664***	1.669***	1.369***	1.325***
	(0.187)	(0.180)	(0.202)	(0.194)	(0.195)	(0.202)	(0.212)
$\|\pi - 2\%\|$		–0.307**	–0.097	–0.095	–0.099	–0.115	–0.082
		(0.131)	(0.117)	(0.121)	(0.120)	(0.109)	(0.110)
$\|\pi - 2\%\| \times D$		0.425***	0.489***	0.496***	0.494***	0.430***	0.481***
		(0.132)	(0.159)	(0.155)	(0.158)	(0.139)	(0.140)
Unemployment Rate (log)			–2.278***	–2.083***	–2.080***	–1.468***	–1.392***
			(0.347)	(0.351)	(0.352)	(0.357)	(0.368)
Uncertainty				2.141*	2.161**	0.076	0.102
				(1.090)	(1.091)	(1.174)	(1.199)
Elections					0.165	0.198	0.190
					(0.235)	(0.175)	(0.175)
GDP (log)						–0.452***	–0.506***
						(0.102)	(0.110)
Credit-to-GDP (%)							–0.006
							(0.005)
Observations	162	151	151	151	151	151	151
R-squared	0.081	0.210	0.408	0.424	0.425	0.492	0.497
Central Bank FE	No	No	No	No	No	No	No

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

Table A.4. OLS Estimates with Fed Inflation Target as 1%, 1.5%, and 2.5% before January 2012 on Price Stability (PriceStab), Employment (Empl), and Sentiment Ratio (SRatio)

	1%			1.5%			2.5%		
Variables	**PriceStab (1)**	**Empl (2)**	**SRatio (3)**	**PriceStab (4)**	**Empl (5)**	**SRatio (6)**	**PriceStab (7)**	**Empl (8)**	**SRatio (9)**
$\lvert\pi - 2\%\rvert$	0.192	0.176***	0.164	–0.166	–0.173***	–0.207	–0.170	–0.168***	–0.212
	(0.125)	(0.051)	(0.159)	(0.121)	(0.050)	(0.164)	(0.121)	(0.049)	(0.164)
$\lvert\pi - 2\%\rvert \times D$	0.274**	–0.192***	–0.379***	0.334***	–0.202***	–0.438***	0.334***	–0.207***	–0.433***
	(0.114)	(0.058)	(0.130)	(0.125)	(0.067)	(0.153)	(0.125)	(0.068)	(0.153)
Unemployment Rate (log)	–0.958**	–0.576***	1.433***	–0.950**	–0.561***	1.504***	–0.947**	–0.562***	1.510***
	(0.448)	(0.180)	(0.353)	(0.450)	(0.181)	(0.337)	(0.450)	(0.181)	(0.337)
Inflation Volatility			–0.922***			–0.973***			–0.984***
			(0.342)			(0.345)			(0.345)
Unemployment Volatility			0.871			1.082			1.099
			(0.861)			(0.883)			(0.882)
GDP Volatility			–0.068			–0.080			–0.081
			(0.088)			(0.090)			(0.090)
Elections	0.182	0.090	0.069	0.173	0.088	0.076	0.172	0.087	0.077
	(0.167)	(0.097)	(0.367)	(0.163)	(0.096)	(0.363)	(0.164)	(0.095)	(0.364)
GDP (log)	–0.388**	0.002	0.351***	–0.392**	0.000	0.325***	–0.393**	–0.001	0.324***
	(0.172)	(0.074)	(0.122)	(0.171)	(0.074)	(0.117)	(0.171)	(0.074)	(0.117)
Credit-to-GDP (%)	–0.009*	0.004**	–0.017***	–0.009*	0.004**	–0.016***	–0.009*	0.004**	–0.016***
	(0.005)	(0.002)	(0.006)	(0.005)	(0.002)	(0.006)	(0.005)	(0.002)	(0.006)
Inflation Objective	1.170***	–0.749***		1.115***	–0.718***		1.108***	–0.721***	
	(0.341)	(0.165)		(0.338)	(0.162)		(0.336)	(0.163)	
Uncertainty	0.333	0.390		0.365	–0.367		0.351	–0.370	
	(1.477)	(0.532)		(1.478)	(0.535)		(1.474)	(0.536)	
Observations	151	151	147	151	151	147	151	151	147
R-squared	0.376	0.441	0.421	0.379	0.437	0.422	0.380	0.438	0.421
Central Bank FE	No	No	No	No	No	No	No	No	No

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

A.7 Sentiment Analysis under the Alternative Weighting Methodology

Table A.5. OLS Estimates on Sentiment Ratio (alternative weighting method) as Dependent Variable

Variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
$\lvert\pi - 2\%\rvert$	–0.021			–0.040	–0.041	–0.041	–0.035
	(0.022)			(0.034)	(0.033)	(0.033)	(0.034)
$\lvert\pi - 2\%\rvert \times D$	–0.119***			–0.126***	–0.127***	–0.127***	–0.119***
	(0.035)			(0.042)	(0.042)	(0.042)	(0.042)
Unemployment Rate (log)		–0.085		–0.042	–0.042	–0.042	–0.010
		(0.095)		(0.102)	(0.103)	(0.103)	(0.106)
Inflation Volatility			–0.197***	–0.135*	–0.135*	–0.135*	–0.125
			(0.074)	(0.080)	(0.080)	(0.080)	(0.080)
Unemployment Volatility			0.204	0.318*	0.317*	0.317*	0.327*
			(0.151)	(0.179)	(0.179)	(0.179)	(0.180)
GDP Volatility			–0.001	–0.013	–0.014	–0.014	–0.013
			(0.017)	(0.018)	(0.017)	(0.017)	(0.017)
Elections					0.034	0.034	0.032
					(0.060)	(0.060)	(0.0060)
Credit-to-GDP (%)							–0.001
							(0.001)
Observations	151	151	153	147	147	147	147
R-squared	0.452	0.411	0.433	0.467	0.468	0.468	0.471
Central Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

Table A.6. OLS Estimates on Sentiment Ratio (adjusted to Hawkish-Dovish terms and alternative weighting method) as Dependent Variable

Variables	**(1)**	**(2)**	**(3)**	**(4)**	**(5)**	**(6)**	**(7)**
Hawkish-Dovish Ratio	0.026	0.037*	0.036	0.030	0.030	0.030	0.026
	(0.020)	(0.021)	(0.022)	(0.024)	(0.024)	(0.024)	(0.026)
$\lvert\pi - 2\rvert$		–0.007	–0.005	–0.024	–0.026	–0.026	–0.022
		(0.023)	(0.024)	(0.035)	(0.035)	(0.035)	(0.035)
$\lvert\pi - 2\%\rvert \times D$		–0.124***	–0.122***	–0.127***	–0.128***	–0.128***	–0.119***
		(0.033)	(0.033)	(0.040)	(0.040)	(0.040)	(0.040)
Unemployment Rate (log)			–0.040	–0.029	–0.030	–0.030	0.003
			(0.091)	(0.099)	(0.100)	(0.100)	(0.100)
Inflation Volatility				–0.125*	–0.125*	–0.125*	–0.115
				(0.074)	(0.074)	(0.074)	(0.074)
Unemployment Volatility				0.233	0.233	0.233	0.253
				(0.183)	(0.183)	(0.183)	(0.187)
GDP Volatility				–0.011	–0.013	–0.013	–0.012
				(0.018)	(0.017)	(0.017)	(0.017)
Elections					0.039	0.039	0.037
					(0.057)	(0.057)	(0.056)
Credit-to-GDP (%)							–0.001
							(0.001)
Observations	162	151	151	147	147	147	147
R-squared	0.431	0.467	0.467	0.478	0.479	0.479	0.482
Central Bank FE	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Note: Robust standard errors are in parentheses. *p < .05; **p < .01; ***p < .001.

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