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Inflation News and Euro-Area Inflation Expectations*

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The protracted period of low inflation in the euro area since 2013 has triggered a significant decline in long-term inflation compensation. This paper re-investigates changes in the sensitivity of inflation compensation to inflation and macroeconomic news and expands existing work in two key dimensions: (i) we analyze all available (advanced) inflation releases for country and euro-area-wide inflation; (ii) we use daily frequency, time-varying sensitivity, and intraday regressions to reach more robust conclusions. Our key findings are twofold. First, timeliness is crucial in inflation markets: it is the early inflation news (flash estimates) which led to revisions in long-term compensation. Second, the anchoring of euro-area inflation expectations has weakened significantly since 2013.

JEL Codes: E31, E52, E58.

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1. Introduction

The monetary transmission mechanism is most effective when long-term inflation expectations are strongly anchored. The analysis of inflation expectations has therefore become a crucial element of modern monetary policy and a recurrent topic for research. Indeed, most central banks' public statements, speeches (e.g., Draghi 2014, 2019; Powell 2017), and specialized press and market commentary (e.g., *The Economist* 2014, 2017; *Financial Times* 2019) nowadays provide a detailed account on the evolution of inflation expectations.

In the case of the euro area, interest in long-term inflation expectations has become more relevant in recent years. Since early 2013, inflation has remained very low and almost constantly below the level consistent with the European Central Bank's (ECB's) quantitative definition of price stability. Such a protracted period of below-target inflation is unprecedented since the launching of the single monetary policy in 1999. While there is consensus that the ECB managed to attain a significant degree of credibility and long-term inflation expectations were strongly anchored ahead of the global financial crisis (GFC henceforth), and even during the subsequent European debt crisis, the situation over more recent years is less clear. Since well-anchored expectations are key for a sustained return of inflation to the ECB's goal, the anchoring of inflation expectations is a fundamental question for the current monetary policy discussion in the euro area.

Against this background, the goal of this paper is to investigate whether the anchoring of euro-area inflation expectations has weakened in recent years and which macroeconomic releases may help explain that weakening. We assess changes in the anchoring of inflation expectations by the reaction of long-term inflation expectations to macroeconomic data releases. The rationale is that if expectations are well anchored, far-ahead inflation expectations should not be significantly affected by news about the current state of the economy, which should only be informative to update the short-term outlook. As a measure of long-term inflation expectations we use forward inflation-linked (IL) swap rates, in line with existing literature for the euro area and other major economies (e.g., Gürkaynak, Levin, and Swanson 2010; Galati, Poelhekke, and Zhou 2011). The euro-area IL swap market is arguably the most liquid in the world, and IL swap rates offer some important advantages over other indicators of inflation expectations. IL swaps are actively traded at much higher frequency than survey data, which makes them more appropriate to analyze the reaction to macroeconomic news.[1] IL swap rates, however, comprise both the level of inflation expectations and the inflation risk premium associated with the risks surrounding them. Since inflation (and macroeconomic news in general) may trigger changes in the perceived risk surrounding future inflation, the inflation risk premium required by investors is also an important dimension of the anchoring of inflation expectations for monetary policy.

To assess whether the anchoring of euro-area inflation expectations has weakened since 2013, we implement three different but complementary pieces of analysis. First, we use daily regressions over different subsamples, expanding the approach in most previous literature on the anchoring of euro-area inflation expectations in two key aspects. By using data up to 2017, we have a sufficient sample for testing whether the below-target period since 2013 has triggered a change in the reaction of long-term inflation compensation. In order to find the relevant macroeconomic news, we provide a detailed description of the information flow of news about inflation in the specific case of the euro area. The euro-area market for IL products has the unique feature of comprising several different economies. As a result, country-specific data releases add a crucial additional dimension to the information flow, in particular regarding inflation, that needs to be taken into account when assessing market reactions to news. Specifically, advance estimates of inflation—the so-called *flash*

[1]Survey measures of euro-area long-term inflation expectations are available at quarterly frequency from the ECB's Survey of Professional Forecasters and Consensus Economics, while market-based inflation expectations are available at daily or even intraday frequency. In addition to lower frequency, survey measures have several additional shortcomings: panelists have litttle incentive to continuously update or even reveal their true forecasts, there is substantial disagreement among panelists, and survey responses may not be internally consistent (e.g., García and Manzanares 2007) and may have become disconnected from actual inflation over the most recent years (Coibion and Gorodnichenko 2015; Chan, Clark, and Koop 2018).

estimates for German, Spanish, and Italian inflation[2]—have been released since 2005 *ahead* of the euro-area-wide flash estimate, the advance estimate of euro-area-wide inflation, and Harmonised Index of Consumer Prices (HICP henceforth) inflation releases. Failing to account for those country flash estimates would therefore ignore the potentially most relevant pieces of news on euro-area inflation, and may cast doubts on the conclusions about the anchoring of inflation expectations. Indeed, we show that those early inflation releases—and in particular the first one of them, the German flash estimate—do have a statistically significant impact on euro-area short-term inflation compensation measures over our whole sample.

Second, we complement our subsample analysis with time-varying estimation of the response of inflation compensation to inflation news. As most inflation (and macroeconomic) releases only occur once a month, to overcome the short-sample problem inherent in rolling regressions we adapt the empirical approach recently used by Swanson and Williams (2014) to the flow of news on euro-area inflation, and gauge evidence of changes in sensitivity to news from rolling regressions at daily frequency.

Finally, we also provide novel additional evidence on intraday reactions of inflation compensation. Daily-frequency regressions are always subject to influences from additional news. Even a careful selection of release dates and additional control variables may not completely isolate that reaction in the data-rich environment in which financial markets nowadays operate. Intraday data on euro-area inflation compensation measures are unfortunately only available since late 2008, but we can cross-check their reaction to news since 2013 and compare it with the previous years using intraday trading evidence.

Our results point to a deterioration in the anchoring of euro-area long-term inflation expectations since 2013. Daily-frequency regressions corroborate existing findings of little sensitivity of long-term inflation compensation to news releases before 2013. In contrast, we find evidence of a significantly stronger reaction of long-term inflation compensation from 2013. Importantly, such statistically

[2]A flash estimate for HICP inflation in France has also been released since early 2016. However, as a crucial dimension of our analysis is the comparison with earlier periods, we exclude it from our analysis.

significant reactions mainly refer to the release of flash estimates at the country level, especially the German, and also the Spanish, flash estimates, which are, respectively, the first two pieces of news about each month's inflation in the euro area. These findings are robust to controlling for a wide range of additional market and macroeconomic information.

Empirical evidence allowing for time-varying sensitivity to news also points to a significantly higher reaction of long-term inflation compensation to inflation news since the euro-area inflation rate has remained below target. Moreover, while previous periods of high sensitivity—for example, over the spring of 2008 when surging oil prices pushed actual inflation to the highest levels in euro-area history—were short-lived, the responsiveness of long-term inflation compensation to inflation news has remained stubbornly high since mid-2013, even well after the launching of the ECB's sovereign bond purchases program in 2015.

Novel analysis using intraday data provides further support for a weakening in the anchoring of euro-area long-term inflation expectations in the later part of our sample. Between 2013 and 2017, we find statistically significant reactions to the release of the flash estimates for Germany and Spain, as well as of the Spanish HICP. While other variables—in particular, economic activity and monetary policy surprises—are also significant, inflation surprises display a much higher explanatory power. Over a shorter time window of 15 minutes, the reaction is somewhat more muted but nonetheless statistically significant for the German flash and also the euro-area flash estimate.

This paper belongs to a large literature investigating the reaction of financial markets to incoming news, and contributes to the analysis of euro-area long-term inflation compensation in three specific aspects: by employing a larger set of news than most existing papers and emphasizing the crucial role of the timing of inflation news (particularly the early releases of inflation at the country level), by combining three different pieces of empirical analysis (daily frequency, time-varying sensitivity, and for the first time intraday regressions) and by using a larger sample, January 2005–March 2017. Those three pieces of analysis provide substantial robustness to our key finding of a significant increase in the reaction of euro-area long-term inflation compensation to news in the latter part of our sample.

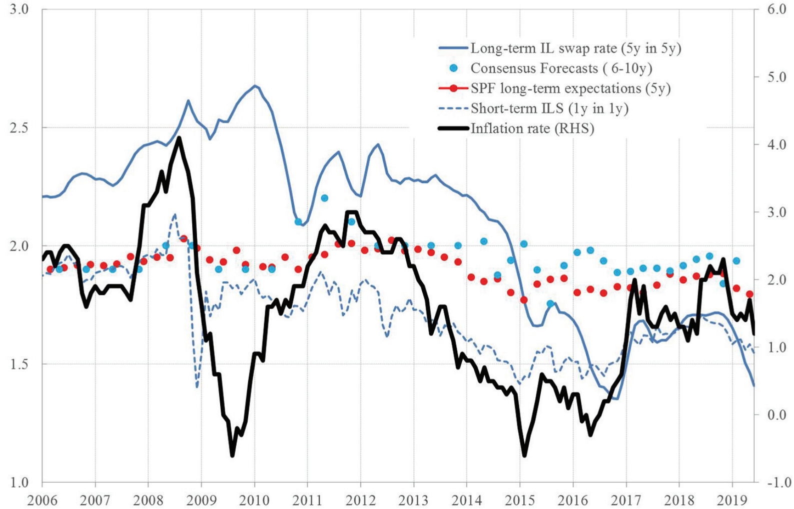
The remainder of the paper is organized as follows. Section 2 reviews the literature on the reaction of inflation compensation to news and motivates the case for further investigation after 2013. Section 3 discusses the data we employ in our analysis. Specifically, we explain in detail the flow of macroeconomic news in the euro area—in particular regarding inflation—and the characteristics of inflation compensation in the euro area. Section 4 reports our empirical results. The main findings from daily frequency, time-varying sensitivity, and intraday data regressions are described in detail. Section 5 contains an additional discussion and robustness checks of some of our findings, and places them in the context of the other results employing alternative indicators of inflation expectations. Finally, section 6 concludes.

2. Related Literature and the Case for Reassessing Euro-Area Inflation Expectations

This paper belongs to an extensive literature investigating the reaction of financial markets to incoming news, from the early works of Dwyer and Hafer (1989), Fleming and Remolona (1997), and Kuttner (2001), among others, on U.S. data. The significant development of financial markets for inflation-linked products and the lumpy manner in which inflation and other macroeconomic data are released provide a great opportunity for researchers to carry out those studies on inflation expectations, as the search for robust evidence of causality is significantly eased using (high-frequency) financial market data (e.g., Gürkaynak and Wright 2013). Indeed, the sensitivity of market-based indicators of inflation expectations, either based on bond-market instruments or from the inflation-linked (IL) swap market, to macroeconomic releases has already been studied using euro-area (and international) data following Gürkaynak, Sack, and Swanson (2005a), among others. This section provides a selective overview of that literature, focusing on the different questions it tried to address over time.

To help put those contributions in context, figure 1 depicts actual inflation dynamics in the euro area (black line), together with some key indicators of euro-area inflation expectations. Two survey measures of long-term inflation expectations, from Consensus Economics (6 to 10 years ahead) and from the ECB's Survey of Professional

Figure 1. Actual Inflation, Market-Based and Survey Long-Term Inflation Expectations



Notes: The figure shows developments in indicators of long-term inflation expectations (both market and survey based) and the actual year-on-year rates of inflation in the euro area. Specifically, it depicts two benchmark forward IL swap rates, for short-term (1-year forward IL swap rate in 1 year, dashed gray/blue line) and long-term expectations (5-year forward IL swap rate in 5 years, gray/blue line); 6–10 years Consensus Forecasts (gray/blue dots) and 5-year-ahead expectations from the ECB's Survey of Professional Forecasters (black/red dots), and the realized inflation rate (thick black line). (Figures in color can be seen in the online version of the paper at http://www.ijcb.org.)

Forecasters (SPF) (5 years ahead), are denoted by red and blue dots. Short-term (one-year in one year) and long-term (five-year in five years) forward IL swap rates are denoted by a dashed and a continuous gray/blue line, respectively, reflecting their availability at higher frequency.[3]

When interpreting differences between survey and market-based inflation expectations, it has to be borne in mind that IL swap rates, and break-even inflation rates (BEIRs), incorporate inflation risk

[3]Given the relatively limited issuance of IL bonds in the euro area—only France, Italy, Germany, Spain, and Greece have issued those bonds—the IL swap market has arguably become the most developed in the world, and can therefore provide a cleaner measure of inflation compensation than bond-based BEIRs.

premiums and should better be interpreted as the overall *inflation compensation* requested by investors, also including risks surrounding future inflation. Changes in inflation compensation could therefore reflect changes in the level of expected inflation, changes in the perceived risks about future inflation, or a combination of both. From a central bank's perspective, both components are of relevance. A credible commitment to price stability should anchor the level of expected inflation to its policy objective. In turn, the requested risk compensation provides relevant information about how firmly inflation expectations may be anchored. Inflation compensation measures are therefore a very rich source of information for testing the reaction to macroeconomic news.[4] In addition, survey measures of inflation expectations have in contrast remained relatively more stable in the euro area, but a decline away from (below but close to) the 2 percent level they exhibited since the early 2000s can also be observed at the end of our sample (e.g., Draghi 2019).

Before the GFC, the sensitivity of inflation expectations to macroeconomic news was mainly explored as a measure of the degree of anchoring of inflation expectations in the euro area (e.g., Coffinet and Frappa 2008; Beechey, Johannsen, and Levin 2011; Ehrmann et al. 2011).[5] Overall, no evidence of a significant reaction of long-term inflation compensation measures to news has been found for the euro area. Moreover, euro-area inflation expectations have been often found to be more strongly anchored than in other countries like the United States, the United Kingdom, or Sweden (e.g., Gürkaynak, Levin, and Swanson 2010).[6]

[4]We focus here on the literature looking at the reaction of inflation compensation measures; section 5 discusses alternative measures of inflation expectations.

[5]Coffinet and Frappa (2008) test the five-year forward IL rate in five years (our benchmark measure) against the euro-area flash estimate and country-specific HICP releases as well as various other macro announcements. Beechey, Johannsen, and Levin (2011) test for the one-year forward IL rate in nine years and similarly use country-specific HICP and other macroeconomic releases, but they do not consider flash estimates. Ehrmann et al. (2011) instead compare the mean and variance of one-year forward *nominal interest rates* in nine years in the nominal yield curves of euro-area countries prior to and during the single-currency period, and interpret the decline and convergence of both moments across countries as signaling a stronger anchoring of inflation expectations.

[6]Research on other countries (see De Pooter et al. 2014 for Brazil, Chile, and Mexico and Gürkaynak et al. 2007 for Canada, Chile, and the United States)

Assessing whether the anchoring of euro-area inflation expectations survived the GFC was the main question for a new wave of research (e.g., Galati, Poelhekke, and Zhou 2011, Autrup and Grothe 2014).[7] Those studies explicitly acknowledge that the analysis in the aftermath of the Lehman collapse years is subject to some uncertainty given the extreme market volatility and limited liquidity during the crisis, and added additional controls (e.g., VIX, bond market volatility, oil prices) to the regressions to account for them. Overall, their findings suggest that the anchoring of euro-area long-term inflation expectations remained strong after the GFC, despite the significant rise in long-term forward inflation compensation at the time.

In light of the weak inflation since 2013 and the launching of additional unconventional monetary policy (UMP) measures in 2015, interest in the reaction of inflation compensation to macroeconomic news has gained momentum again. Long-term forward inflation compensation has declined significantly in the euro area since actual inflation rates started surprising negatively in 2013. Inflation compensation declined in several markets, but the correction was particularly strong in the euro area, and attracted substantial attention among policymakers (Draghi 2014, 2019), as well as in specialized press and market commentary (e.g., *The Economist* 2014; *Financial Times* 2019).

To assess the reactions to news in the low inflation environment since 2013, Miccoli and Neri (2015) look at the reaction of IL swap spot and forward rates to the euro-area HICP releases at different maturities separately, comparing the average of 10 business days prior to and after the date of release. Speck (2016) instead jointly regresses the two-year spot, the three-year forward in two years, and

also finds evidence of the crucial role of inflation targeting in the anchoring of inflation expectations.

[7]Galati, Poelhekke, and Zhou (2011) use a sample for the United States, the euro area, and the United Kingdom from June 2004 until March 2009, and find that the euro-area and the U.K. inflation expectations seem to be more stable than those in the United States. To gauge the reaction of inflation compensation, they only use country-specific HICP releases as inflation news and found that the VIX has a statistically significant impact on expectations. Autrup and Grothe (2014) use data up to 2012 and find the euro-area expectations to be more stable than the U.S. ones. The paper also shows that bond-based inflation compensation, in contrast to IL swap rates, is statistically influenced by a liquidity risk premium.

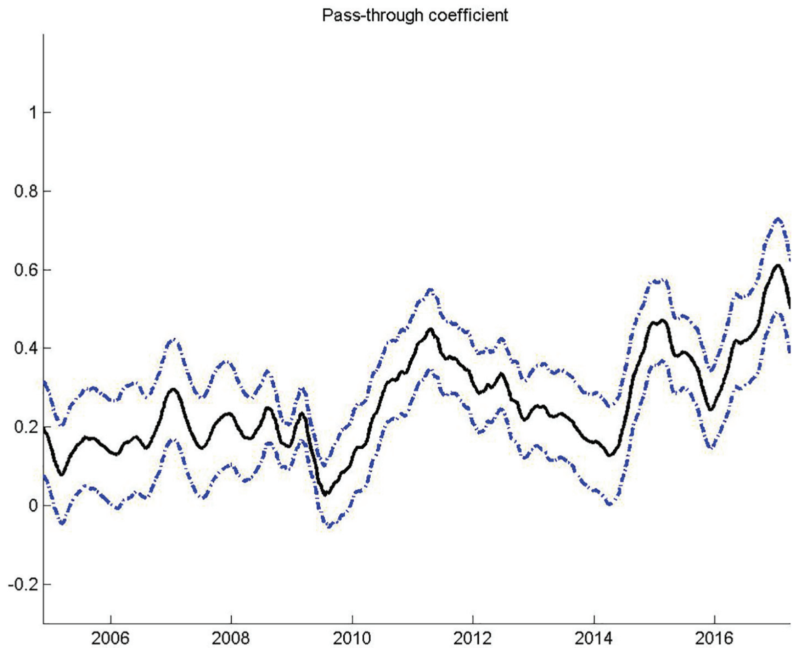
the five-year forward in five years, and tests the time variation in their sensitivity to news using country HICP estimates and a few confidence indicators. Both studies find a somewhat stronger sensitivity of medium-to-long-term inflation compensation to news in the latter part of their samples, but they disagree on whether that evidence is sufficiently conclusive for the presence of a significant de-anchoring of inflation expectations or just a longer-than-normal correction from low inflation levels. Other contributions (e.g., Fracasso and Probo 2017, Nautz, Pagenhardt, and Strohsal 2017) however find evidence of a de-anchoring of euro-area inflation expectations as early as late 2011 using structural break tests on the reaction of long-term inflation compensation to the euro-area HICP and some other macroeconomic data releases.

The lack of conclusive evidence on whether the anchoring of inflation expectations has weakened in the euro area since 2013 may reflect several potential shortcomings of existing literature. First, a challenge for the more recent research is that, for a weaker anchoring of euro-area inflation expectations to be economically meaningful, it has to be sustained over a certain period of time. A minimum sample is therefore required to reach robust conclusions. Figure 1 shows that the decline in long-term inflation compensation took place over a relatively long period between 2013 and most of 2016. Moreover, the decline was in line with that in actual inflation in a way that was not common before, suggesting there were important changes in the pricing of inflation compensation.

Additional evidence on the potential deterioration on the anchoring of inflation expectations can be found by assessing the pass-through from short-term to longer-term inflation compensation. If long-term inflation expectations are well anchored, revisions in short-term inflation expectations—reacting, for example, to actual inflation readings—should not trigger strong revisions in longer-term inflation expectations. Indeed, evidence from a standard linear regression[8] points to a much stronger impact of changes in

[8]Figure 2 reports the estimates of the posterior mean for β_t and standard quantiles in $\Delta y_t^{5y5y} = \alpha + \beta_t \Delta y_t^{1y1y} + \epsilon_t$, where β_t evolves following a random walk. The framework allows for stochastic volatility to account for potential changes in market conditions and varying volatility as suggested by Jochmann, Koop, and Potter (2010) and Chan (2013).

Figure 2. Pass-Through from Short- to Long-Term Inflation Expectations



Notes: The figure shows the estimates of the posterior median for β_t and the 16th and 84th percentiles of the confidence interval in the regression $\Delta y_t^{5y5y} = \alpha_t + \beta_t \Delta y_t^{1y1y} + \epsilon_t$, where β_t evolves following a random walk and allowing the stochastic volatility to account for potential changes in market conditions and time-varying volatility following Chan (2013) and Jochmann, Koop, and Potter (2019). For presentational purposes, results are smoothed over a 100-day window.

short-term expectations on longer-term expectations in the latter part of our extended sample (see figure 2).

The goal of this paper is to assess the extent to which the response to inflation news was responsible for that decline, and to provide robust evidence on whether the observed reactions have been different from those observed in the past. While our analysis is close in spirit to the literature reviewed above, this paper also expands existing evidence to help reach more robust conclusions. Specifically, we will analyze in greater detail the unique flow of inflation releases in the euro area, expanding the set of macroeconomic news—particularly inflation releases—with respect to most existing literature, and paying particular attention to their timing.

Evidence that news released earlier in time tends to have a greater impact on financial assets than news about the same variable released later has been supported by early event studies (e.g., Fleming and Remolona 1997; Andersen et al. 2003). More recently Hess and Niessen (2010) and Gilbert et al. (2017), among others, have explored the tradeoffs between timeliness and precision of macroeconomic news in asset price responses. Both studies find that financial markets favor timeliness to more comprehensive macroeconomic news.[9]

The timing of macroeconomic releases has not been emphasized in the literature on the reaction of inflation expectations to news. In most countries—for example, in the large number of studies employing U.S. data—inflation news mostly comes from two variables, namely CPI and PPI indexes. The relevant flow of inflation news in the euro area is instead much richer. We discuss in detail the flow of inflation releases in the euro area and how they can be incorporated into event studies on the euro area in the next section. In addition, we also provide novel evidence of euro-area inflation expectations on the reaction to news using intraday frequency data.

3. News and Inflation Expectations in the Euro Area

A central argument of this paper is that a meaningful event-study analysis on euro-area data requires taking into account some specific

[9]Gilbert et al. (2017) estimate the intrinsic value of a macroeconomic announcement (defined as the ability to nowcast gross domestic product (GDP) growth, inflation, and the federal funds target rate), in terms of relation to fundamentals, timing, and revision noise, and find that timing is the most important characteristic in explaining asset price impact. Similarly, Hess and Niessen (2010) test the tradeoff between early availability and information quality of news on two German economic indicators—the ZEW, with an earlier release, and the IFO indicator, comprising more information—and report evidence of a larger ZEW impact due to its earlier release. Similarly, Hess (2004) studies the impact of a large and diverse set of macroeconomic announcements using high-frequency analysis and finds that the response to announcements that are released in month $m+1$ is significantly stronger than the impact of announcements released a month later $(m + 2)$.

characteristics of euro-area macroeconomic data releases and financial market indicators. We therefore discuss them in detail before embarking on our empirical analysis.

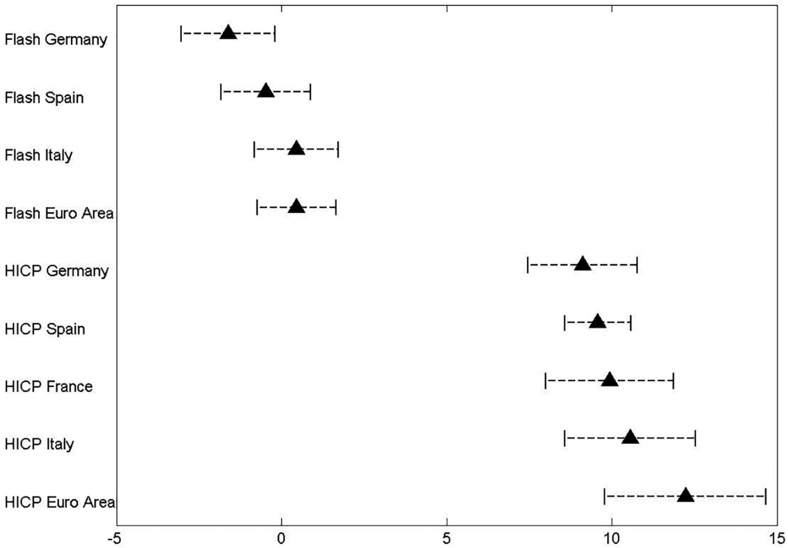
3.1 The Flow of Euro-Area Releases

Financial markets nowadays operate in a data-rich environment (Bernanke and Boivin 2003), and indeed evidence suggests that market participants do monitor a large number of macroeconomic indicators (e.g., Bartsch et al. 2014; Swanson and Williams 2014). Given its unique characteristics as a monetary union comprising a relatively large number of countries, the presence of a data-rich environment is even more extreme in the case of euro-area financial markets.

In addition to the amount of relevant news, the timing of the macroeconomic releases is another important aspect to take into account. In the euro area, particularly regarding inflation news but also some other macroeconomic series like the GDP, there are a number of key country-specific macroeconomic releases that precede the release of euro-area aggregates. A key motivation of our analysis is that failing to account for those country-specific releases can impair a proper understanding of the reaction of inflation expectations to macroeconomic news and bias the empirical results.

Since we will focus on the reaction of long-term inflation expectations, we will pay particular attention to the news surrounding the releases of the HICP. There are two kinds on HICP releases in the euro area. *Flash* releases are *advance* estimates regularly issued by statistical offices. Eurostat releases a flash estimate for the euro-area-wide HICP, but flash estimates are also issued for Germany, Spain, and Italy (and, since 2016, also France) by the corresponding national statistical institutes. Those country flash estimates are normally limited to a figure for the year-on-year inflation rate for the current month and do not contain a breakdown of the different subcomponents (e.g., service or non-energy industrial goods inflation rates). The euro-area flash estimate, for example, has only included a breakdown of main HICP subcomponents since late 2012. The relevance of the flash estimates instead stems from the fact that they are released well ahead of (about two to three weeks before) the full release for the corresponding country or the euro-area-wide HICP.

Figure 3. Schedule of Flash and HICP Inflation Data Releases



Notes: The figure shows the average calendar timing of flash estimates and HICP data releases for European countries (Germany, Spain, France, and Italy) and euro-area-wide inflation. The y-axis lists the inflation releases; the x-axis indicates the average release days (triangle), and whiskers reflect the distribution over our sample (January 2005 to March 2017). 0 denotes the last business day of the month to which the release refers; 5 denotes five business days after the last business day of the month, etc. For example, the flash HICP for Germany was released on average 2 days before the last business day of the corresponding month and the HICP release for Italy on average more than 10 working days into the following month.

Figure 3 illustrates the flow of inflation releases in the euro area. The x-axis indicates the number of business days between the release dates of the monthly flash estimates and the corresponding HICP with respect to the last business day of the reference month m. A zero on the x-axis therefore indicates that a flash estimate was released on the last day of the month to which it refers, negative numbers indicate the number of days before the end of the month (the most common situation), while positive numbers point to days into the month following the one to which the flash release refers. More formally, $flash_{m,t}$ provides an estimate of $HICP_{m,t}$ for the corresponding country or the euro area in month m.

Dates of the inflation releases nonetheless vary from month to month across countries, depending on data collection and processing. A triangle in figure 3 indicates the average business day of the inflation data release, and the whiskers represent the distribution in the release dates. For example, the German flash estimate is released on average two business days within the reference month, the earliest among the flashes (the Spanish flash is released one day before the end of the month, while the Italian and the euro-area-wide flash tend to be released one day into the following month). All flash estimates are nonetheless released ahead of the corresponding HICP inflation releases, which take place in the month following the reference period: 9 days for the German, 10 for the Spanish and the French, 11 for the Italian, and about 13 days for the euro area after the end of the reference month.

Besides inflation data releases, we also consider data releases for other major euro-area macroeconomic series to provide an ample coverage of factors shaping the macroeconomic situation in the euro area and therefore potentially having an impact on inflation expectations. Specifically, we consider producer price index (PPI), GDP, purchasing manager's index (PMI), consumer confidence, industrial confidence, and production (new orders, unemployment rate, growth of the monetary aggregate (M3), retail sales, and trade balance). A complete list of the series can be found in table 1.

Financial markets are forward-looking and, under the efficient market hypothesis, all available and expected information should be priced in, so only the "unexpected" or "news" component of macroeconomic data releases should lead to changes in inflation compensation. We measure the news component of each release using the macroeconomic expectations collected by Bloomberg L.P. from a selection of professional economists until soon before the release takes place. In our econometric analysis, we use euro-area-wide and some key country releases for which Bloomberg has collected expectations over a large part of our sample.

We compute the surprise component for each data release as the difference between the actual values of the official release and the Bloomberg survey expectations for each variable k for reference period t. Using the surprise component of the releases also removes any issues of endogeneity arising from inflation expectations feeding back to the macroeconomy, because any such effects, to the

Table 1. Long-Term Inflation Compensation (5y5y, Daily): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	0.57**	(2.19)	0.34	(0.89)	1.08***	(4.91)
Spain	0.39*	(1.85)	0.62**	(2.09)	−0.04	(−0.16)
Italy	−0.06	(−0.30)	−0.08	(−0.30)	−0.17	(−0.56)
Euro Area	0.10	(0.38)	0.15	(0.38)	0.00	(0.00)
HICP						
Germany	0.10	(0.54)	0.05	(0.26)	0.43	(1.15)
Spain	0.23	(0.99)	0.24	(0.90)	0.11	(0.33)
France	−0.08	(−0.58)	−0.13	(−0.86)	0.14	(0.40)
Italy	0.19	(1.06)	0.38*	(1.81)	−0.40	(−1.27)
Euro Area	0.06	(0.43)	0.08	(0.45)	0.04	(0.24)
PPI	−0.05	(−0.25)	0.00	(−0.02)	−0.14	(−0.55)
GDP	0.40***	(2.69)	0.45***	(2.80)	0.19	(0.60)
PMI	−0.03	(−0.15)	−0.03	(−0.16)	−0.02	(−0.04)
Consumer Confidence	0.16	(1.15)	0.12	(0.93)	0.36	(0.83)
Industrial Confidence	−0.13	(−0.63)	−0.17	(−0.76)	−0.03	(−0.09)
Industrial Production	0.47***	(2.65)	0.48**	(2.27)	0.30	(1.27)
New Orders	−0.19	(−0.79)	−0.17	(−0.71)	—	—
Unemployment Rate	−0.28	(−1.16)	−0.19	(−0.58)	−0.55**	(−2.01)

(continued)

Table 1. (Continued)

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
M3	−0.13	(−0.80)	−0.14	(−0.81)	−0.06	(−0.17)
MRO	—	—	—	—	0.40***	(3.42)
Retail Sales	0.13	(0.81)	0.11	(0.62)	—	—
Trade Balance	−0.03	(−0.13)	−0.04	(−0.13)	−0.01	(−0.03)
Mon. Policy Shock	4.59	(0.63)	7.99	(0.93)	−13.33	(−0.80)
CB Info. Shock	25.75**	(2.18)	24.97*	(1.85)	13.52	(0.69)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.02		0.00		0.03	
R^2	0.03		0.03		0.07	
$H_0 : \beta = 0$ (p-value)	0.00		0.06		0.00	
Chow Test (p-value)	0					

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in long-term inflation compensation (5y5y, in basis points). The surprise component of macroeconomic releases, $\mathbf{X}_t$, is normalized by their historical standard deviations; coefficients there represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero (joint significance test). Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

extent that they are predictable, should be already incorporated into market expectations.

Among the set of variables, monetary policy surprises deserve particular attention. Bloomberg L.P. also collects market expectations about the ECB's policy rates, but as the ECB was highly predictable, there were very few monetary policy surprises, particularly before 2013.[10] We nonetheless include them in our set of surprises. Recent literature has however emphasized that central banks' announcements reveal information not just about policy rates but also about the central bank's assessment of the economic outlook. Jarociński and Karadi (2018) uses a wider range of market price reactions to identify two structural shocks embodied in central bank announcements—a purged monetary policy shock and a central bank information shock—and find evidence of a significant impact of information shocks on macroeconomic data, including on long-term inflation expectations. Moreover, central bank information shocks appear to be more frequent in the ECB's announcements than in the U.S. Federal Reserve's Federal Open Market Committee's announcements, for example. In order to assess the response of long-term inflation compensation measures on inflation surprises, we also add those monetary policy and central bank information shocks to our set of surprises.

3.2 Measuring Inflation Expectations

Our empirical analysis will focus on long-term forward inflation compensation. Forward IL swap rates provide a very useful means of interpreting markets' inflation compensation at medium-to-long-term horizons. Standard macroeconomic models predict that inflation should return to its steady state between 5 and 10 years after a typical shock (e.g., Gürkaynak, Sack, and Swanson 2005b). We will focus our analysis on the five-year forward IL swap rate five years ahead, which has become the most widely used measure to assess

[10]The expectation collection was also extended beyond policy rates following the introduction of some UMP measures, but they are available over too short a period for our analysis.

developments in euro-area long-term inflation expectations for policy (e.g., Draghi 2014) and research purposes.[11]

While euro-area IL swap contracts have been very actively traded since 2004 over maturities from 1 to 30 years, available evidence and market intelligence suggest that the 5- and 10-year maturities have tended to concentrate around 50 percent of all trade activity in the market.[12] Using swap rates also alleviates difficulties in controlling for the sovereign and the liquidity risk embodied in the prices of IL bonds since the GFC, which includes the main period of interest in our analysis.

4. Assessing the Anchoring of Inflation Expectations

This section reports the main findings of our empirical analysis. To reach robust conclusions, we investigate the response of long-term inflation compensation to inflation and macroeconomic news by means of three different but complementary pieces of analysis.

4.1 Evidence from Daily Regressions

Table 1 reports estimation results from the standard linear regression:

$$\Delta y_t = \alpha + \beta \mathbf{X}_t + \epsilon_t, \tag{1}$$

[11]Formally, the price of a spot zero-coupon swap with a 10-year maturity, s_t^{10y}, reflects the average inflation compensation over the next 10 years. It is similar for the five-year spot rate, s_t^{5y}. In contrast, by construction, 5-year forward inflation compensation 5 years ahead, f_t^{5y5y}, reflects the average inflation compensation priced in between 5 and 10 years ahead, a medium-to-long-term period that captures well the movements in inflation compensation we are interested in. Formally, the long-term forward IL rates implicit in the term structure of IL swap rates can be calculated from the 5- and 10-year spot rates as $f_t^{5y5y} = \left[\frac{(1+s_t^{10y})^{10}}{(1+s_t^{5y})^5}\right]^{\frac{1}{10-5}} - 1$.

[12]Evidence from other indicators used in the related literature, like the 1-year forward in 9 years for example, involves using 9- and 10-year spot IL swap rates, with the former having only around 7 percent of trading activity of the latter. The forward rate should therefore be interpreted with caution. As a robustness check and link to previous literature, the appendix nonetheless shows that our main findings are robust to different measures of long-term inflation compensation (see tables A.1 and A.2 in the appendix).

where t indexes days, and $\Delta y_t \equiv y_t - y_{t-1}$ is the change in the inflation swap rate over a day; $\mathbf{X}_t$ is a vector of macroeconomic news in data releases, including all macroeconomic releases and the monetary policy surprises (as described above), and ϵ_t denotes the regression residual capturing the influence of other information or financial factors on inflation compensation that day. Surprises are normalized by its standard deviation, and the estimated coefficients can be interpreted as basis points per standard deviation in the release.

The estimated impact coefficients and their standard deviations are reported over three different subsamples to gain some insights about their changes over time. The first two columns report results over the full sample 2005–17. The releases of the German and the Spanish flash inflation estimates as well as the euro-area activity (GDP and industrial production) and the central bank information shock are statistically significant at standard significance levels. Inflation news in particular leads to a change in long-term inflation compensation of around 0.6 and 0.4 basis point on average on the day of their release. These results underscore the importance of accounting for the timing of the inflation news to assess the reaction of long-term inflation compensation: not only do flashes tend to trigger larger (and statistically significant) reactions than the HICP releases, but also the strongest reactions are those for the flash releases that take place earlier in time and are therefore the earliest pieces of news about euro-area inflation.

The remaining columns of table 1 report results over two subperiods: the pre-disinflation period, 2005–12, and the period over which actual euro-area inflation was consistently below target, 2013–17, the main period of interest here. Results point to some important changes in the news that triggered statistically significant reactions in long-term inflation compensation. Over the pre-disinflation period, the economic activity news (GDP and industrial production) and the monetary policy information shock are identified as the most relevant non-inflation news ones. Regarding inflation news, statistically significant reactions on inflation expectations were triggered by the Spanish flash and the Italian HICP before 2013, possibly reflecting the market focus on periphery countries during the European debt crisis period.

Over the below-target period, however, the relevant surprises are quite different. Among inflation news, it is the German flash estimate which leads to large reactions in long-term inflation compensation (its impact coefficient rises to 1.08, more than three times its value before 2013) and becomes strongly significant. In contrast, market reactions to economic activity news seemed to shift from the GDP and industrial production towards the unemployment rate. In addition, monetary policy shocks continue to be significant news, now captured by unexpected changes in the main refinancing rate, whose level changes did trigger significant surprises among market participants since 2013.[13]

Although the estimated responses per standard deviation to news reported in table 1 may appear to be limited and the R^2 low, it is important to bear in mind that daily changes in euro-area long-term inflation compensation have been, on average, relatively limited (around 2 basis points). Indeed, in normal circumstances—with a strong anchoring of inflation expectations—one should observe that macroeconomic data releases trigger no change at all in euro-area long-term inflation compensation. Therefore, low coefficient estimates are common in this type of analysis given the low signal-to-noise ratio of any single monthly data release for the true underlying state of economic activity and inflation, particularly over long horizons (Swanson and Williams 2014). The statistical significance of some coefficients in table 1, particularly inflation news, instead suggests that regressions like (1) are extremely informative to gauge the sensitivity of long-term inflation expectations to economic news and the anchoring of inflation expectations.

The comparison of relevant surprises in the pre-disinflation period (before 2013) and in the below-target period (after 2013) provides two main insights. First, while there are some differences between the significant news in the two periods, the overall joint test of no significance cannot be rejected before 2013 (p-value of 0.06 for the pre-disinflation period) but is strongly rejected for the below-target period (p-value of 0.00). Second, the Chow forecast

[13] The standard joint regression commonly used in the related literature, including table 1 here, associates a zero with the days on which there are no releases for the variable. This standard practice implies no distinction to the value associated to a fully anticipated release, that is a no-surprise in the announcement. This is particularly relevant for the monetary policy surprises regarding the ECB's main refinancing rate and its strong predictability before unconventional monetary policy measures were introduced (see Hartmann and Smets 2018).

test also rejects the null hypothesis of overall stability of the coefficients over the two periods.[14] The remainder of the section provides additional empirical evidence in support of the robustness of those conclusions.[15]

In our sample, releases of individual macroeconomic series only occur once a month. Within a joint regression for any given news surprise, all other news is set to zero, though strictly speaking no news release has occurred for any other variable. To gain additional evidence on the main drivers of changes in inflation compensation, we also estimate daily regressions using the surprises for individual variables:

$$\Delta y_t = \alpha + \beta x_t + \epsilon_t, \tag{2}$$

where x_t denotes the vector of releases for a single variable. This type of analysis allows for regression estimates to be based only on data for those days on which a release for each variable takes place therefore complementing joint regressions above (e.g., Kilian and Vega 2011).

Table 2 corroborates the results from the joint variable regressions. News on German and Spanish flash estimates has a significant impact over the full sample, leading to a change in long-term inflation compensation of around 0.6 and 0.5 basis point in the day of its release, respectively. However, subsample analysis suggests the statistical significance is mainly related to the latter part of the sample. Prior to 2013, there had been no strong significant reaction to inflation releases, while the reactions to surprises in flash estimates turned strongly significant over the below-target period 2013–17. Specifically, the reaction to the German flash estimate rose

[14] The Chow forecast test selects a subsample to check the validity across the full sample. Since we assume that the below-target (BT) period has a more significant parameter vector, we use it to estimate the parameter vector $\widehat{\beta_{BT}} = (\mathbf{X}'_{t,BT}\mathbf{X}_{t,BT})^{-1}\mathbf{X}'_{t,BT}\Delta y_{t,BT}$ as well as obtain the residual sum of squares RSS_{BT}. The estimated parameter vector $\widehat{\beta_{BT}}$ and the surprises of the full sample $\mathbf{X}_{t,All}$ are used to estimate $\Delta y_{t,All}$ and subsequently obtain the RSS_{All}. The F statistic is then computed: $F(T_{PD}, T_{BT} - K) = \frac{RSS_{All} - RSS_{BT}/T_{PD}}{RSS_{BT}/(T_{BT} - K)}$.

[15] The appendix also confirms these results employing the same regression using a two-day window (tables A.4 and A.5) as well as using the 1y9y forward rate as an alternative measure of long-term inflation compensation (tables A.1 and A.2).

Table 2. Long-Term Inflation Compensation (5y5y, daily): Individual Variable Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample			Pre-disinflation			Below Target		
	β	t-stats	R^2	β	t-stats	R^2	β	t-stats	R^2
Flash									
Germany	0.60**	(2.36)	0.05	0.34	(0.91)	0.01	1.10***	(5.48)	0.26
Spain	0.48**	(2.35)	0.06	0.58*	(1.79)	0.08	0.33	(1.38)	0.03
Italy	0.04	(0.26)	0.00	0.00	(−0.02)	0.00	0.12	(0.37)	0.00
Euro Area	0.15	(0.82)	0.01	0.14	(0.52)	0.00	0.16	(0.72)	0.01
HICP									
Germany	0.09	(0.53)	0.00	0.06	(0.32)	0.00	0.46	(1.19)	0.02
Spain	0.26	(1.16)	0.02	0.31	(1.17)	0.02	0.02	(0.06)	0.00
France	−0.06	(−0.40)	0.00	−0.07	(−0.50)	0.00	0.08	(0.22)	0.00
Italy	0.19	(1.01)	0.01	0.36	(1.66)	0.03	−0.33	(−1.24)	0.03
Euro Area	0.06	(0.39)	0.00	0.06	(0.35)	0.00	0.03	(0.18)	0.00
PPI	−0.02	(−0.12)	0.00	0.02	(0.06)	0.00	−0.10	(−0.45)	0.00
GDP	0.42***	(2.98)	0.04	0.46***	(2.89)	0.06	0.26	(0.83)	0.01
PMI	−0.02	(−0.08)	0.00	−0.02	(−0.12)	0.00	0.34	(0.68)	0.00
Consumer Confidence	0.13	(0.93)	0.00	0.11	(0.74)	0.00	0.34	(0.80)	0.01
Industrial Confidence	−0.10	(−0.46)	0.00	−0.09	(−0.39)	0.00	0.01	(0.03)	0.00
Industrial Production	0.45***	(2.71)	0.05	0.46**	(2.38)	0.05	0.37	(1.57)	0.03
New Orders	−0.17	(−0.74)	0.01	−0.17	(−0.74)	0.01	—	—	—
Unemployment Rate	−0.30	(−1.28)	0.02	−0.25	(−0.82)	0.01	−0.30	(−1.11)	0.03
M3	−0.14	(−0.86)	0.01	−0.16	(−0.91)	0.01	−0.05	(−0.14)	0.00
MRO	—	—	—	—	—	—	0.35**	(2.50)	0.08
Retail Sales	0.13	(0.81)	0.00	0.09	(0.48)	0.00	—	—	—
Trade Balance	−0.02	(−0.09)	0.00	−0.02	(−0.05)	0.00	−0.02	(−0.09)	0.00
Mon. Policy Shock	−2.70	(−0.40)	0.00	4.37	(0.47)	0.00	−19.62**	(−2.36)	0.13
CB Info. Shock	25.31**	(2.42)	0.08	26.63*	(1.98)	0.08	25.22***	(3.19)	0.11

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for each single variable in an individual regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in long-term inflation compensation (5y5y, in basis points). The surprise component of macroeconomic releases, x_t, is normalized by their historical standard deviations; coefficients β therefore represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. R^2 reports the explanatory variable for each individual regression. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

to 1.1 basis points—with a t-statistic of 5.48—almost twice as much as over the sample as a whole (0.6) and more than three times the average response (0.34) over the pre-disinflation period 2005–12. The impact of the Spanish flash estimate instead seems to be somewhat stronger before 2013, which is consistent with the early start of the decline in Spanish inflation after its post-GFC peak compared with the euro-area-wide inflation, and our interpretation that market participants paid more attention to inflation developments in periphery countries during the onset of the European debt crisis. Individual variable regressions also confirm the relevance of economic activity (GDP, industrial production) mainly before 2013, and monetary policy variables over the whole sample for long-term inflation compensation.

Importantly, the individual variable regressions underscore the strong role of the German flash estimate, the first piece of inflation news in the euro area every month, during the below-target period. Surprises in the German flash releases explain 26 percent of the variation of long-term inflation compensation in their release days, at least twice as much as any other individual variable, and almost four times as much as the R^2 for the joint regression shown above. Moreover, over the pre-disinflation period the explanatory power of all variables, including the statistically significant ones on economic activity, remain in single digits. Finally, the same pattern of stronger responses in the below-target period (2013–17) compared with the previous pre-disinflation period (2005–12) emerges after controlling for a large set of additional indicators of market conditions.[16] Joint and individual variable regressions at daily frequency provide evidence consistent with an increase in the sensitivity of long-term inflation compensation to inflation news since 2013. While actual inflation developments provide a strong case for splitting the sample in 2013, we investigate the time variation in the sensitivity of long-term inflation compensation to news in greater detail next.

[16]Table A.3 in the appendix consider an extended regression in which additional controls for market conditions have been included in the regression. Specifically, changes in the EA Citigroup surprise index, oil and gold prices, the EUR/USD exchange rate, VIX and VSTOXX, and CDS premiums are considered as controls.

4.2 Evidence on Time-Varying Sensitivity to News

The subsample analysis suggests that the impact of macroeconomic data releases on long-term inflation compensation has changed over time and, more specifically, has strongly risen since 2013. To obtain more precise evidence on how those reactions have evolved over time, equations (1) and (2) could be estimated over rolling windows. However, as most macroeconomic announcements only occur once a month, any rolling window based on individual variable release would suffer from small-sample problems. We therefore provide some evidence of time variation in the response of inflation compensation to macroeconomic releases using the approach recently introduced by Moessner and Nelson (2008) and Swanson and Williams (2014, SW henceforth) to study the evolution of the reaction of U.S. yields to macroeconomic news. The first step as proposed by Moessner and Nelson (2008) is to specify a nonlinear equation. A set of multiplicative time dummies δ allows for the time variation in the response of the variables of interest to macroeconomic news to be measured by the combined effect $\delta\beta$. Such a specification therefore measures whether the response to macroeconomic news has changed over time even if the relative magnitude of the elements of β are constant over time.

Formally, the regression (1) used in the previous section is extended to a nonlinear least squares specification as follows:

$$\Delta y_t = \gamma^{\tau_i} + \delta^{\tau_i}\beta\mathbf{X}_t + \epsilon_t^{\tau_i}, \qquad (3)$$

where the parameters γ^{τ_i} and δ^{τ_i} take on different values in each calendar year $i = 2005, 2006, \ldots, 2017$. The second step is the key motivation for regression (3). It allows us to reduce the small-sample problem associated with allowing every element of X_t to vary over time. More specifically, by pooling information across data releases, we can obtain more detailed evidence on the time variation in the response of long-term inflation compensation to macroeconomic news through the estimation of daily rolling regressions of the form

$$\Delta y_t = \gamma^{\tau} + \delta^{\tau}\hat{\mathbf{X}}_t + \epsilon_t^{\tau}, \qquad (4)$$

where $\hat{\mathbf{X}}_t \equiv \hat{\beta}\mathbf{X}_t$ is the generic surprise regressor defined using $\hat{\beta}$ from regression (3).

To make this approach operational, two additional assumptions are needed. First, a normalization for δ^{τ_i} needs to be chosen to separately identify β and δ^{τ_i}. The estimated response of long-term inflation compensation to macroeconomic news will then be relative to the response over the normalization period. Following SW, we normalize δ^{τ_i} to average 1 over a given period—in our baseline case, over the period of below-target period 2013–17. Intuitively, if there has been a deterioration in the anchoring of inflation expectations since 2013, by normalizing the response between 2013–17 to 1, we should observe a response of inflation compensation to macroeconomic news over the rest of our sample that is, on average, *below* 1.[17]

Second, τ in regression (4) denotes the period over which the rolling regressions are carried out. To identify a weakening of the anchoring of inflation expectations that is taking place over a substantial period of time rather than capturing a stronger reaction to news that may be triggered by just a few influential data releases in a few months in a row, we run regression (4) over a two-year window.[18] Following SW, we account for the two-stage sampling uncertainty by scaling the standard error σ_{δ^τ} of δ^τ in (4) by the weighted average of the derived standard errors of δ^{τ_i} in regression (3).[19]

Table 3 reports the estimated coefficients of nonlinear regression (3). The German flash estimate, the Spanish HICP, industrial production, and the central bank information shock are found to be significant at least at the 10 percent level and carry the expected sign, with the two inflation surprises significant at the 1 percent level and the 5 percent level, respectively. Figure 4 depicts the estimated sensitivity of long-term inflation compensation to macroeconomic news δ^{τ_i} together with their uncertainty bands using all variables

[17]Figure A.1 in the appendix offers results for different normalization periods and shows that our main findings do not depend on the specific period chosen for the normalization of δ^{τ_i}.

[18]The subsample analysis in the previous section points to an increase in the impact of inflation news since 2013. We look for supporting evidence by assessing the responses on rolling regressions over one-sided windows, while SW report results over two-sided windows centered around the business date.

[19]Specifically, σ^τ is adjusted as $\sigma^\tau_{adj} = \sigma^\tau \left(\sum_i^{I_\tau} w_i \zeta^{\tau_i}\right)$, where I_τ specifies the calendar years the rolling window covers and w_i the number of days in calendar year i devided by the total number of days of rolling window τ.

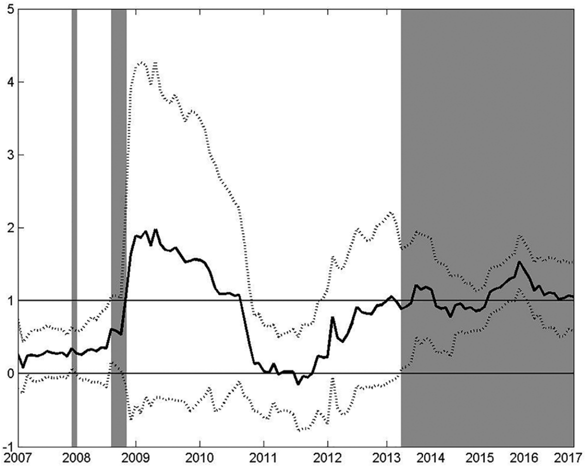
Table 3. Long-Term Inflation Compensation (5y5y, daily): Nonlinear Regression $\Delta y_t = \gamma^{\tau_i} + \delta^{\tau_i}\beta\mathbf{X}_t + \varepsilon_t$

	All Variables		Selected Variables	
	β	t-stats	β	t-stats
Flash				
Germany	0.89***	(4.14)	0.87***	(4.03)
Spain	0.12	(0.69)	0.11	(0.53)
Italy	−0.20	(−1.22)		
Euro Area	−0.05	(−0.16)		
HICP				
Germany	0.32	(1.34)		
Spain	0.68**	(2.43)	0.63**	(2.26)
France	−0.01	(−0.08)		
Italy	0.37	(1.62)	0.27	(1.49)
Euro Area	−0.01	(−0.11)		
PPI	−0.24	(−1.37)		
GDP	0.09	(0.54)		
PMI	0.05	(0.34)		
Consumer Confidence	0.08	(0.57)		
Industrial Confidence	−0.19	(−1.15)		
Industrial Production	0.50*	(1.85)	0.49*	(1.79)
New Orders	0.13	(0.92)		
Unemployment Rate	−0.16	(−0.62)		
M3	0.09	(0.69)		
Retail Sales	0.04	(0.20)		
Trade Balance	−0.07	(−0.49)		
Mon. Policy Shock	6.17	(1.14)		
CB Info. Shock	19.89*	(1.81)	19.96*	(1.87)
Observations	1,767		687	
R^2	0.05		0.11	
$H_0 : \beta$ Constant, p-value	0.99		0.99	
$H_0 : \delta$ Symmetric (p-value)	0.92		0.92	
$H_0 : \delta$ Constant (p-value)	0		0	

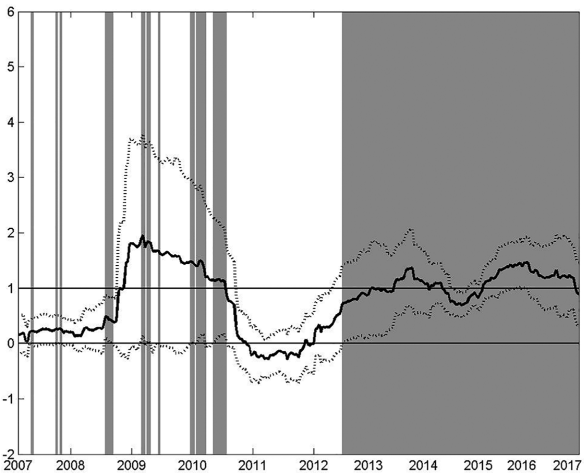
Notes: The table reports estimated β coefficients from a regression on days of releases from January 2005 to March 2017, using surprises for all variables in a joint (nonlinear) regression (Swanson and Williams 2014). Coefficients indexed τ_i may take on different values in different calendar years. Heteroskedasticity-consistent t-statistics are in parentheses; and ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. H_0 : β *constant* implies that the coefficient vector β in the regression is constant over time and only the scalar coefficients δ^{τ_i} vary against an alternative in which every element of β is permitted to vary independently across calendar years. $H_0 : \delta$ *symmetric* assumes that the δ^{τ_i} in the regression are the same for positive and negative surprises $\beta\mathbf{X}_t$, against an alternative in which we allow separate coefficients $\delta_+^{\tau_i}$ and $\delta_-^{\tau_i}$ for positive and negative values of $\beta\mathbf{X}_t$ in each calendar year i. $H_0 : \delta$ *constant* assumes that the time-varying sensitivity coefficients δ^{τ_i} in the regression are constant over time. That is, we test whether $\delta^{\tau_i} = 1$ for each calendar year $i = 2005, \ldots, 2015$.

Figure 4. Long-Term Inflation Compensation (5y5y, daily): Time-Varying Sensitivity Coefficients δ^τ from Nonlinear Regression

A. δ^τ Using All Variable Regression Results



B. δ^τ Using Selected Variable Regression Results



Notes: This figure plots the coefficients δ^τ estimated from a regression $\Delta y_t = \gamma^\tau + \delta^\tau \hat{\mathbf{X}}_t + \epsilon_t^\tau$. Panels A and B depict δ^τ using the all and selected variable regression results (see table 3), respectively. Dotted gray lines depict heteroskedasticity-consistent $\pm$ two-standard-error bands, adjusted for two-stage sampling uncertainty in δ^τ. Shaded regions denote δ^τ being significantly above 0 and suggest a weaker anchoring of inflation expectations.

employed in the daily regressions above. While the sensitivity to news displays some variation in the responses between 2005 and 2017, and several episodes of responsiveness significantly above zero can be identified over the sample, the level and significance of the sensitivity of inflation compensation to news is (in relative terms) clearly higher, and more persistent, from 2013. Indeed, since euro-area inflation has remained (mostly) below target, the estimates of δ^{τ_i} rose from a level statistically not different from zero since mid-2010 to be most of the time above 1 since 2014.[20]

Our results also identify a sharp deterioration in the anchoring of inflation expectations in the second half of 2008. This is consistent with the substantial increase in oil prices, actual inflation, and inflation compensation over 2008 ahead of the collapse of Lehman Brothers. The results depicted in figure 4 suggest that the 2008 episode was more short-lived, but the period 2008–09 should be interpreted with caution because euro-area inflation markets were under significant stress during that time, and it cannot be ruled out that the stronger sensitivity to news may just be the result of higher volatility of inflation compensation measures. Indeed, there is a significant widening of uncertainty in the estimation around that time and δ^{τ} becomes statistically insignificant despite its relatively high level.

Including many non-significant macroeconomic news in (3) might however add unnecessary noise to the regression. As a robustness check, we also provide the estimation of δ^{τ_i} restricting the set of variables to those whose news is found to trigger significant reactions in regressions above (results for the restricted nonlinear specification for those six variables are reported in table 3). The regression results and the rolling parameter δ^{τ} are very much in line with those from the previous regression. Similarly, the estimates of the time-varying

[20]Following SW, tests for several additional dimensions of the time variation in the responses are reported at the bottom of table 3. First, we test whether the relative response coefficients β are constant over time (only δ^{τ_i} varies) against an alternative in which every element of β varies. p-values close to 1 support the restricted specification (3). Second, we test whether the δ^{τ_i} coefficients are the same for positive and negative surprises. With a p-value close to 1, the hypothesis of a symmetric response is not rejected. Finally, we test the hypothesis that the time-varying coefficient δ^{τ_i} is constant over time—that is, for each calendar year $i = 2005, \ldots, 2017$ $\delta^{\tau_i} = 1$. With a low p-value, the data clearly reject that restriction.

sensitivity of long-term inflation compensation to news further single out the relevance of the period beyond 2013 in terms of statistical and economic significance over the sample as a whole.

Overall, the time-varying regressions point to a deterioration in the anchoring of euro-area inflation expectations since 2013. While our findings are robust to different normalization periods and the inclusion of different inflation releases in the analysis, it is also true that their estimation requires nonetheless pooling together a number of variables. The next section provides additional evidence on the responses to news on specific data releases.

4.3 Evidence from Intraday Data

We expand existing analysis on the reaction of euro-area inflation expectations to news by conducting regressions using high-frequency (intraday) data. Arguably, event-study analysis, as the one conducted in previous sections and the related literature, is preferably undertaken with intraday data: the smaller the window around the news arrival, the less likely anything but the news under investigation affects asset price changes, reinforcing the case for direct causality and making the event study resemble as close as possible a controlled natural experiment (Gürkaynak and Wright 2013).

To our knowledge, this is the first event-study analysis using intraday data for euro-area inflation compensation measures, and there is some uncertainty about the speed of reaction of the euro-area IL swap market to news. We consider the changes in the spot and forward inflation compensation over a benchmark window of 120 minutes after the data release, allowing for 10 minutes before the data release for the formation of expectations ahead of the release. Over that window, at least in our sample, there are no important additional releases that interfere with market trading of inflation compensation. Therefore, the changes should be directly attributable to the specific data release under study. International evidence on intraday data on inflation compensation is also limited. Beechey and Wright (2009), for example, look at the U.S. market but focus on bond-based BEIRs rather than IL swaps, which allows them to consider a shorter window of 15 minutes.

A shortcoming of our high-frequency data is that their collection is only available from late 2008. We use the specific time of the data release as reported in Bloomberg, and our intraday trades of euro-area IL swap rates from Reuters. While we cannot use our intraday data sample to investigate the anchoring of inflation expectations prior to the GFC, we can however assess the reaction of inflation compensation measures in the most important period, namely the severe disinflation that took place in the euro area since 2013, and compare it with the 2009–12 period.

Intraday evidence corroborates our previous findings of a deterioration in the anchoring of long-term inflation expectations in the later part of our sample. Between 2013 and 2017, over a 120-minute window from the data release, long-term inflation compensation has significantly reacted to the news component of the flash releases for Germany and Spain (1 percent and 10 percent), as well as of the Spanish HICP (5 percent) (see table 4). Significant responses (at 10 percent) were also found for PPI, the unemployment rate, and policy rate announcements, in line with the evidence from daily regressions. In the 2009–12 period instead we find some significant reactions to the German flash estimate (at 10 percent) and the PPI (with a negative sign), but those reactions seem to be quantitative and qualitatively weaker than those after 2013. First, before 2013 we cannot reject the joint hypothesis of nonsignificance (p-value 0.8), while since 2013 we can safely reject the joint nonsignificance with a p-value of 0.01. In addition, the regression coefficients indicate a structural break with the p-value of the Chow forecast test of 0.059.

As argued in section 4.1, the joint regressions might add additional noise to the regression by treating all other variables as zero surprise though only one variable has a news release in a given day. As a robustness check, we also run variable-by-variable regressions over the same window length (t–10 minutes, t+120 minutes) to gain additional insights on the variables driving changes in inflation compensation. German and Spanish flash news is strongly significant (at 1 percent) and explains 38 percent and 12 percent of the variability of long-term inflation compensation, significantly above the Spanish HICP, the PPI, and the policy rate surprise, which are also found to be significant. In stark contrast, no single inflation news surprise is found to trigger a significant

Table 4. Long-Term Inflation Compensation (5y5y, intraday): 120-Minute Window (minus 10 minutes)

	Joint Linear Regression				Individual-Variable Linear Regressions					
	2009–12		Below Target		2009–12			Below Target		
	β	t-stats	β	t-stats	β	t-stats	R^2	β	t-stats	R^2
Flash										
Germany	0.53*	(1.75)	0.56***	(3.63)	0.58	(1.58)	0.08	0.62***	(3.90)	0.38
Spain	0.09	(0.41)	0.2*	(1.67)	0.07	(0.24)	0.00	0.38**	(2.66)	0.12
Italy	−0.29	(−1.11)	−0.14	(−0.54)	−0.33	(−1.58)	0.09	0.02	(0.08)	0.00
Euro Area	−0.06	(−0.33)	−0.09	(−0.43)	−0.28	(−1.31)	0.04	0.01	(0.05)	0.00
HICP										
Germany	−0.14	(−0.89)	0.07	(0.16)	−0.17	(−0.98)	0.02	0.08	(0.19)	0.00
Spain	0.06	(0.40)	0.29**	(2.24)	0.20	(1.14)	0.01	0.27**	(2.51)	0.03
France	−0.07	(−0.45)	0.37*	(1.95)	−0.18	(−1.18)	0.03	0.32	(1.50)	0.03
Italy	−0.14	(−0.63)	0.04	(0.17)	−0.24	(−1.04)	0.07	0.02	(0.09)	0.00
Euro Area	−0.29	(−0.78)	0.05	(0.26)	−0.47	(−1.10)	0.10	0.01	(0.09)	0.00
PPI	−0.47**	(−2.36)	−0.18*	(−1.83)	−0.62**	(−2.05)	0.10	−0.20*	(−2.00)	0.05
GDP	0.13*	(1.69)	−0.24	(−1.51)	0.16**	(2.44)	0.04	−0.11	(−0.78)	0.00
PMI	−0.06	(−0.31)	−0.17	(−0.72)	−0.11	(−0.42)	0.00	−0.03	(−0.12)	0.00
Consumer Confidence	−0.01	(−0.14)	0.01	(0.06)	−0.07	(−0.69)	0.00	0.02	(0.17)	0.00
Industrial Confidence	−0.08	(−0.44)	0.01	(0.06)	−0.05	(−0.27)	0.00	0.05	(0.22)	0.00
Industrial Production	0.11	(0.92)	−0.1	(−0.72)	0.18	(1.34)	0.03	−0.14	(−0.97)	0.01
New Orders	0.26	(0.75)	—	—	0.39	(1.03)	0.05	—	—	—
Unemployment Rate	0.18	(0.34)	−0.42*	(−1.93)	0.00	(0.00)	0.00	−0.28	(−1.42)	0.04
M3	0.02	(0.16)	0.28	(1.33)	0.07	(0.61)	0.00	0.27	(1.24)	0.04
MRO	—	—	0.16*	(1.67)	—	—	—	0.23**	(2.48)	0.06
Retail Sales	−0.03	(−0.21)	—	—	−0.03	(−0.18)	0.00	—	—	—
Trade Balance	−0.11	(−0.87)	0.05	(0.73)	0.18	(1.57)	0.03	0.10	(0.75)	0.01
Observations	516		562							
$\bar{R}^2$	0.00		0.03							
R^2	0.03		0.07							
$H_0 : \beta = 0$ (p-value)	0.80		001							
Chow Test (p-value)	0									

Notes: The table reports estimated β coefficients from regressions on days with data releases over two periods: from April 2009 to December 2012 (pre-disinflation) and from January 2013 to March 2017 (below target), using surprises for each variable in a joint regression (left columns) and for each single variable in an individual regression (right columns). Changes in long-term inflation compensation are calculated over a 120-minute window. The surprise component of macroeconomic releases k is normalized by their historical standard deviations; coefficients represent a 10^{-2} basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

reaction before 2013, when only the PPI and the GDP are found to be significant.

We also investigate the robustness of our conclusions to the speed of reaction to news in the euro-area IL swap market by looking at two different windows around our benchmark, 15 minutes and 240 minutes after the release, and computing changes from 10 minutes before the release. See tables 5 and 6. Results over a shorter time of a 15-minute window confirm the strong impact of the German flash release after 2013, and also of other inflation news (Italian, with a wrong sign, and euro-area flashes, and the French CPI, at 10 percent significance levels) as well as the unemployment rate and M3 releases. Overall, there is evidence of statistical significance at individual and joint level after 2013, while, despite some significant impact, the hypothesis of joint insignificance of news before 2013 cannot be rejected (p-value 0.46), while the stability of the regression coefficients can be rejected (p-value 0).

Reactions over a longer window of t+240 minutes after the release (from t–10 minutes) confirm that impact identified over shorter windows persists over time, and further corroborate the main message found so far. Specifically, while very few significant reactions to news are found prior to 2013, since then German and Spanish flashes, the Spanish HICP, the PPI (with a negative sign though), and news about the ECB's main refinancing rate are found to trigger statistically significant reactions. Intraday analysis therefore provides strong support to the importance of early releases, as well as the presence of a weakening of the anchoring in inflation expectations since 2013.

5. Discussion

Previous sections have reported robust evidence of an increase in the response of long-term inflation compensation to inflation and macroeconomic news in the euro area in recent years by means of three different but complementary pieces of analysis, namely daily, time-varying, and intraday regressions. This section highlights some additional dimensions of our findings, and relates them to some additional evidence and recent literature.

Table 5. Long-Term Inflation Compensation (5y5y, intraday): 15-Minute Window (minus 10 minutes)

	Joint Linear Regression				Individual-Variable Linear Regressions					
	2009–12		Below Target		2009–12			Below Target		
	β	t-stats	β	t-stats	β	t-stats	R^2	β	t-stats	R^2
Flash										
Germany	0.08	(0.94)	0.21**	(2.50)	0.10	(0.83)	0.03	0.19**	(2.11)	0.16
Spain	0.07	(0.52)	−0.01	(−0.10)	0.04	(0.27)	0.01	0.07	(0.87)	0.02
Italy	0.05	(0.47)	−0.18**	(−1.98)	−0.07	(−0.70)	0.02	−0.04	(−0.56)	0.00
Euro Area	−0.23**	(−2.35)	0.19*	(1.77)	−0.18*	(−1.94)	0.09	0.19*	(1.85)	0.07
HICP										
Germany	0.01	(0.05)	0.06	(0.52)	0.02	(0.13)	0.00	0.09	(0.74)	0.01
Spain	−0.17	(−0.78)	0	(0.08)	−0.24	(−0.86)	0.02	0.01	(0.11)	0.00
France	0.02	(0.26)	0.12*	(1.74)	0.01	(0.16)	0.00	0.13	(1.55)	0.02
Italy	0.03	(0.22)	0.05	(0.50)	0.01	(0.08)	0.00	0.07	(0.62)	0.01
Euro Area	−0.12	(−1.19)	0.04	(0.35)	−0.11	(−0.91)	0.02	0.04	(0.43)	0.01
PPI	−0.18	(−1.41)	0.06	(1.25)	−0.21	(−1.23)	0.02	0.06	(1.10)	0.01
GDP	−0.06	(−0.87)	−0.09	(−0.63)	−0.11	(−0.93)	0.03	−0.04	(−0.34)	0.00
PMI	−0.15**	(−1.97)	−0.13	(−0.68)	−0.18**	(−2.05)	0.07	−0.16	(−0.81)	0.00
Consumer Confidence	−0.03	(−0.41)	0.02	(0.22)	−0.04	(−0.40)	0.00	0.01	(0.12)	0.00
Industrial Confidence	0.02	(0.36)	0.21	(1.47)	0.01	(0.18)	0.00	0.26	(1.67)	0.10
Industrial Production	−0.06	(−0.76)	−0.06	(−0.63)	−0.08	(−0.92)	0.02	−0.08	(−0.74)	0.01
New Orders	0.24	(1.37)	—	–	0.28	(1.43)	0.06	—	—	—
Unemployment Rate	−0.08*	(−0.58)	−0.17*	(−1.78)	−0.17	(−1.44)	0.05	−0.17**	(−2.27)	0.04
M3	0.07	(0.56)	−0.18**	(−1.96)	0.10	(−0.76)	0.01	−0.19**	(−2.14)	0.09
MRO	—	—	0.04	(0.79)	—	—	—	0.04	(0.77)	0.01
Retail Sales	0.08	(1.70)	—	—	0.08	(1.42)	0.04	—	—	—
Trade Balance	−0.11	(−0.87)	0.05	(0.73)	−0.15	(−0.92)	0.02	0.05	(0.72)	0.01
Observations	562		516							
$\bar{R}^2$	0		0.02							
R^2	0.04		0.05							
$H_0 : \beta = 0$ (p-value)	0.46		0.04							
Chow Test (p-value)	0									

Notes: The table reports estimated β coefficients from regressions on days with data releases over two periods: from April 2009 to December 2012 (pre-disinflation) and from January 2013 to March 2017 (below target), using surprises for each variable in a joint regression (left columns) and for each single variable in an individual regression (right columns). Changes in long-term inflation compensation are calculated over a 15-minute window. The surprise component of macroeconomic releases k is normalized by their historical standard deviations; coefficients represent a 10^{-2} basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table 6. Long-Term Inflation Compensation (5y5y, intraday): 240-Minute Window (minus 10 minutes)

	Joint Linear Regression				Individual-Variable Linear Regressions					
	2009–12		Below Target		2009–12			Below Target		
	β	t-stats	β	t-stats	β	t-stats	R^2	β	t-stats	R^2
Flash										
Germany	0.98	(1.61)	0.91***	(3.19)	1.05	(1.42)	0.07	1.04***	(3.46)	0.32
Spain	0.10	(0.33)	0.42*	(1.71)	0.09	(0.21)	0.00	0.69**	(2.62)	0.10
Italy	−0.63	(−1.49)	−0.10	(−0.19)	−0.59*	(−1.76)	0.10	0.09	(0.16)	0.00
Euro Area	0.10	(0.29)	−0.38	(−0.92)	−0.38	(−0.96)	0.02	−0.17	(−0.58)	0.00
HICP										
Germany	−0.29	(−1.15)	0.08	(0.09)	−0.37	(−1.32)	0.03	0.07	(0.08)	0.00
Spain	0.28	(0.98)	0.57**	(2.33)	0.65	(1.60)	0.04	0.54**	(2.52)	0.04
France	−0.15	(−0.52)	0.62	(1.62)	−0.37	(−1.28)	0.04	0.52	(1.25)	0.02
Italy	−0.3	(−0.84	0.02	(0.05)	−0.49	(−1.28)	0.09	−0.03	(−0.08)	0.00
Euro Area	−0.46	(−0.71)	0.06	(0.22)	−0.83	(−1.10)	0.10	−0.01	(−0.04)	0.00
PPI	−0.75**	(−1.96)	−0.43**	(−2.29)	−1.03*	(−1.82)	0.09	−0.45**	(−2.50)	0.10
GDP	0.31*	(1.94)	−0.38	(−1.52)	0.43**	(2.73)	0.08	−0.19	(−0.75)	0.00
PMI	0.03	(0.09)	−0.22	(−0.63)	−0.04	(−0.07)	0.00	0.10	(0.28)	0.00
Consumer Confidence	0.01	(0.05)	0	(−0.02)	−0.09	(−0.44)	0.00	0.03	(0.15)	0.00
Industrial Confidence	−0.19	(−0.49)	−0.19	(−0.77)	−0.12	(−0.31)	0.00	−0.17	(−0.53)	0.01
Industrial Production	0.28	(1.24)	−0.14	(−0.50)	0.43*	(1.75)	0.05	−0.20	(−0.70)	0.01
New Orders	0.28	(0.47)	—	—	0.50	(0.77)	0.03	—	—	—
Unemployment Rate	0.43	(0.46)	−0.66	(−1.60)	0.17	(0.12)	0.00	−0.40	(−1.02)	0.02
M3	−0.03	(−0.11)	0.75*	(1.86)	0.03	(0.12)	0.00	0.74*	(1.74)	0.09
MRO	—	—	0.29	(1.42)	—	—	—	0.42**	(2.26)	0.06
Retail Sales	−0.14	(−0.52)	—	—	−0.14	(−0.43)	0.01	—	—	—
Trade Balance	0.35	(1.25)	0.15	(0.71)	0.50*	(1.95)	0.04	0.15	(0.69)	0.01
Observations	562		516							
$\bar{R}^2$	0.00		0.03							
R^2	0.03		0.07							
$H_0 : \beta = 0$ (p-value)	0.82		0.02							
Chow Test (p-value)	0									

Notes: The table reports estimated β coefficients from regressions on days with data releases over two periods: from April 2009 to December 2012 (pre-disinflation) and from January 2013 to March 2017 (below target), using surprises for each variable in a joint regression (left columns) and for each single variable in an individual regression (right columns). Changes in long-term inflation compensation are calculated over a 240-minute window. The surprise component of macroeconomic releases k is normalized by their historical standard deviations; coefficients represent a 10^{-2} basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

5.1 The Impact of Inflation News and Its Timing

The statistical evidence of a significant increase in the impact of inflation surprises on inflation compensation since 2013 points to a weakening of the anchoring of euro-area long-term inflation expectations. Our analysis sheds additional evidence on the role of the timing of the inflation news to assess that de-anchoring: not only does inflation news have a stronger impact in the below-target period, but that evidence is consistent with a stronger role of the early inflation releases, particularly the German but also the Spanish flash estimates, instead of the HICP releases (country and euro-area-wide). We believe this evidence is fully consistent with an optimal pricing of news among market participants: earlier inflation news is fully priced in, making additional information released later on somewhat less likely to cause a significant impact.

The importance of early inflation releases in the pricing of inflation compensation is indeed a crucial contribution of this paper. We have emphasized their role in explaining changes of long-term inflation compensation, but it is important to note that flash estimates also have a crucial role in the formation of inflation expectations at shorter horizons. Indeed, regression results along the lines of the ones reported in the previous section confirm the relevance of the flash estimates for short-term measures of inflation compensation at one-year (spot IL swap rate) and the one-year forward IL swap rate in one year (see tables A.6 and A.7 in the appendix). Specifically, early inflation releases—in particular, the German flash but also the Italian and the euro-area flashes—are found to be statistically significant and have a much stronger impact than the corresponding HICP releases. Moreover, among the HICP releases only the French HICP, for which there was no flash estimate before 2016, appears to trigger strong significant reactions in short-term inflation compensation. This evidence provides strong support for the crucial importance of early released inflation news in the expectation formation in inflation markets.

That observation is consistent with recent literature investigating the tradeoff between early availability and information quality, namely that it is difficult to achieve substantial improvements that outweigh the disadvantage of having formed price expectations using the early, even if less precise, indicator (see Hess and Niessen 2010, Gilbert et al. 2017). Hence, once there is a statistically significant reaction to the early-released flash estimates, the fact that the euro-area flash estimate or the HICP releases have not triggered significant market reactions should not be interpreted as signaling strong anchoring. The crucial statistical role of the early-released country flashes in the variable-by-variable and joint regressions reported above is clear evidence of their economic relevance. Moreover, evidence from different windows for intraday movements, daily regressions, and two-day regressions does confirm that the effects of early news are long-lasting.

Previous sections also report statistically significant reactions to other macroeconomic surprises both before (e.g., GDP, industrial production) and after (e.g., unemployment rate) 2013. Their statistical significance, however, tends to be less consistent in joint and variable-by-variable regressions. It changes over subsamples, and—in contrast to inflation news in the below-target period—tests for joint significance tend to be rejected at standard significance levels. Similarly, while monetary policy surprises may also have an impact on long-term inflation expectations, their inclusion in the joint regressions does not diminish the significance of early inflation news. Indeed, the joint regressions point to a lower relevance of central bank communication (as measured by the central bank information shock) beyond surprises in the interest rate announcement themselves in the below-target period.

The strong relevance of inflation news is also consistent with empirical evidence pointing at the superior performance of univariate models when forecasting inflation (e.g., Faust and Wright 2013 or Chan, Clark, and Koop 2018 in two very different settings). Non-inflation surprises, which rather relate to the current state of the business cycle, may be arguably more relevant for the anchoring of inflation expectations in countries like the United States, where the Federal Reserve has a double mandate of inflation stability and maximum employment, than to the euro area, where the ECB's target is solely defined in terms of price stability.[21]

[21]Speck (2016), for example, also finds evidence of significant surprises using German and Italian business confidence as well as French manufacturing confidence, but concludes there is no clear evidence of a de-anchoring of inflation expectations over his sample.

5.2 What Is Behind the Reaction of Inflation Compensation?

The two components of inflation compensation, the level of long-term inflation expectations and the inflation risk premium, reflect different dimensions of anchoring, and are therefore very relevant for our goal in this paper. While it is difficult to disentangle these two (unobserved) components,[22] we argue that the stronger sensitivity of long-term forward inflation compensation to macroeconomic news since 2013 is indicative of a weaker anchoring of inflation expectations. In this section we use additional information from survey data to discuss whether that weakening may reflect changes in expected inflation or the inflation risk premium, or both.

It is also important to note that our analysis of inflation compensation does not rely on the expectations theory of the term structure. Since long-term forward IL swap rates capture the compensation that investors demand both for expected inflation and for the risks associated with inflation at that horizon, changes in inflation compensation need not be due solely to shifts in the conditional expected level of inflation at long horizons: if the anchoring of inflation expectations weakens, economic news might well also shift the inflation risk premium in long-term inflation compensation, either because investors' perceptions regarding the distribution of long-run inflation outcomes changes or because the economic news triggered a change in pricing of long-run inflation risks.

The significant decline in long-term inflation compensation and its very low levels since 2013 (see figure 1) suggest that a decline in the level of inflation expectations embodied in long-term inflation compensation measures, above and beyond a potential decline in the inflation risk premiums,[23] is the most plausible interpretation.

[22] Approaches to correct BEIRs for liquidity premium—based on relative traded volumes or asset-swap spreads, for example—necessarily involve some crucial assumptions on model specification or its presence across maturities. In turn, inflation risk premium estimation usually takes place using term structure models. Estimates vary significantly across specifications even for a single country, but they generally point to significant variation in inflation risk premium over time and across maturities (see, for example, Kim, Walsh, and Wei 2019).

[23] Term structure models regularly monitored by the Federal Reserve or the ECB point to a compression in inflation risk premium in the United States and the euro area, but cannot fully explain the decline of long-term inflation compensation measures (Board of Governors of the Federal Reserve System 2015).

Additional results beyond the main focus of this paper are consistent with that interpretation. For example, trend inflation estimates along the lines of Stock and Watson (2007) point to a significant decline in long-term optimal conditional forecast since 2013. Moreover, that finding holds even allowing for forward-looking information into the term structure estimation, For example, García and Poon (2018) show that the protracted decline in long-term inflation compensation since 2013 is explained by both a significant decline in the level of long-term inflation expectations—from around 2 percent to lows around 1.5 percent—and a decline in inflation risk premiums—from around 40 basis points in mid-2013 to zero and even temporarily negative levels in the second half of 2016.

Survey data do not incorporate inflation risk premiums, and offer a somewhat longer history (since the start of the euro area in 1999) but at much lower frequencies (quarterly; see figure 1). Yet, Łyziak and Paloviita (2017) tested the sensitivity of long-term inflation expectations to changes in short-term ones and, by splitting the sample at the second quarter of 2008, found evidence of a de-anchoring of long-term inflation expectations based on the significant impact of short-term expectations on longer-term ones over the second part of their sample.

5.3 Evidence from Additional Inflation Derivatives

We have focused on the literature investigating the reaction of inflation compensation to news because it is the closest to the analysis carried out in this paper. A growing literature has however investigated changes in the anchoring of euro-area long-term inflation expectations using alternative data sources from the expanding IL derivatives market, mainly the growing inflation options market.[24] Here we review its findings, which overall tend to support a weakening in the anchoring of inflation expectations in the latter part of our sample.

[24] Inflation options, in similar fashion to standard options for stocks or interest rates, provide financial protection when inflation, the underlying asset here, moves above or below (cap and floor options, respectively) a given threshold (i.e., the strike price or rate). As in the case of IL swap rates, inflation options are nowadays actively traded over the counter, and their market is more developed in the euro area than in most other countries (Gimeno and Ibañez 2018).

Galati et al. (2016) investigate the reaction the deflation risk implied by risk-neutral densities (RNDs) on year-on-year inflation to oil price changes, and find evidence consistent with a subtle, but nonetheless statistically significant, weakening in the anchoring of long-term inflation expectations. Gimeno and Ibañez (2018) estimate inflation RNDs using IL swap and inflation options (caps and floors) across a large number of horizons—including at the five-year forward in five years benchmark reference for monetary policy—and find that the priced probability associated with negative inflation values at that long horizon, after declining from 2012, rose significantly again from early 2014 and almost doubled in less than two years to around 12 percent (above 30 percent at the two-year forward in two years). While those probabilities may not appear to be very high, their pricing at fairly long horizons suggests that they do not reflect just temporary factors, which is consistent with a weakening of the anchoring of euro-area long-term inflation expectations since 2013. Those perceived risks can explain why investors could be willing to forgo some return in the form of risk premiums in circumstances in which nominal bonds may turn from inflation bets to "deflation hedges" (Campbell, Sunderam, and Viceira 2017). A downward revision in inflation risk premium in light of a rising probability of persistent low inflation, even if the perceived deflation risks somewhat recede, is also a potential concern for any central bank, and part of a de-anchoring of long-term expectations (and associated risks) that the long-term inflation compensation measures used in this paper are able to capture.

Natoli and Sigalotti (2017) investigate tail correlations using inflation swaps and options data and introduce a new indicator based on the odds that strong negative shocks to short-term expectations are connected to large declines in long-term expectations using a logistic regression framework. That paper reports an increase in the risk of de-anchoring since late 2014 for the euro area, while evidence for the United States and the United Kingdom instead points to firmly anchored inflation expectations. Scharnagl and Stapf (2015) also find low sensitivity in euro-area inflation options to news before early 2013, and conclude that inflation expectations remained well anchored in their sample.

5.4 Additional Robustness Checks

Our main findings are also robust to increasing the number of controls in the joint regressions (see table A.3 in the appendix). Interestingly, the relevance of market conditions also changes from the pre-disinflation period to the below-target period. While bond and stock market volatility, and index of CDS premiums in euro-area sovereign bond markets, or even an index of economic surprises for the U.S. economy, triggered statistically significant reactions in euro-area long-term inflation compensation before 2013, in the below-target period in contrast there was an important shift in the relevant news for inflation compensation, with early inflation news (i.e., the German flash estimate) emerging as the main driver, above and beyond oil price changes as the most relevant variable.

Liquidity considerations have led our choice of the five-year IL forward rate in five years as a benchmark inflation compensation measure. Alternative measures employed in the related literature, such as the one-year forward in nine years, are also broadly supportive of our main findings (see tables A.1 and A.2 in the appendix). While the strongly significant reactions triggered by the German flash estimate in the below-target period are corroborated, results over the full sample period are instead somewhat puzzling. Limited liquidity in the euro-area IL swap market at this maturity however calls for some caution when interpreting these empirical results.

The linear specification of our regressions may also trigger some concerns about the irrelevance of changes in the distribution of data surprises, in particular if the distribution of data surprises X_t in the period 2013–17 were very different from that in the earlier part of the sample. In our empirical regressions, the surprise component of data releases X_t is regarded as strictly exogenous, as the expectations data used in their calculation are assumed to incorporate all relevant information available prior to the release. The inflation surprises used in our analysis are therefore considered independent of all past and future values of the changes in long-term inflation compensation on the left-hand side of these regressions, and strict exogeneity implies that the empirical distribution of the data surprises is irrelevant for our estimates of the response coefficients in our linear empirical regressions (1), (2), and (3). The distribution of inflation surprises—in particular, those for the surprise component in country flash estimates—has however been quite similar across time. Such similarity in the size of surprises may be puzzling at first sight, but it is consistent with a gradual downward revision to short-term inflation expectations in line with the decline in actual inflation among financial market participants. In those circumstances, the surprises in inflation releases should not look very different from those over earlier periods. Moreover, given the significant downward revision in short-term inflation expectations since 2013, it is plausible that the protracted period of low inflation also led to downward revisions in the level of long-term inflation compensation, as shown in figure 1, and its responsiveness to inflation news as shown in previous sections.

6. Conclusions

Following the unprecedented period of low (and mostly below-target) inflation in the euro area since 2013, whether the anchoring of euro-area inflation expectations has weakened in recent years is a crucial question. The monetary transmission mechanism is most effective when long-term inflation expectations are strongly anchored, and, indeed, evidence on earlier samples has found that euro-area inflation expectations were well anchored around the ECB's quantitative definition of price stability of (below but close to) 2 percent.

This paper provides two main findings to shed new light on those considerations. First, by expanding the standard analysis to all available early inflation releases for country and euro-area-wide inflation, we find a crucial role for the timing of inflation releases. Indeed the main reactions in inflation compensation are triggered by early data releases, the country flash estimates regularly released ahead of the euro-area-wide data, and official country HICP releases. Second, taking into account those early releases, we find evidence of stronger reactions of long-term inflation compensation measures to news since 2013. Evidence from daily, time-varying, and intraday regressions consistently shows a significant impact of inflation news over recent years that had not been observed before in the euro area, and points to a significant deterioration in the anchoring of inflation expectations in the euro area.

Appendix

Table A.1. Alternative Long-Term Inflation Compensation (1y9y, daily): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	0.34	(0.89)	−0.11	(−0.21)	1.38***	(3.27)
Spain	0.92**	(2.58)	1.34**	(2.56)	0.10	(0.26)
Italy	0.11	(0.27)	0.26	(0.50)	−0.14	(−0.22)
Euro Area	−0.74	(−1.61)	−1.22*	(−1.83)	0.14	(0.28)
HICP						
Germany	0.40	(1.25)	0.32	(0.95)	1.22	(1.41)
Spain	0.11	(0.16)	0.07	(0.08)	0.47	(0.58)
France	−0.64	(−0.88)	−0.77	(−0.96)	0.56	(0.97)
Italy	0.41	(1.09)	0.42	(0.90)	0.42	(0.89)
Euro Area	−0.05	(−0.13)	−0.04	(−0.09)	−0.06	(−0.16)
PPI	−0.26	(−0.79)	−0.28	(−0.68)	−0.23	(−0.44)
GDP	0.71**	(2.52)	0.71**	(2.24)	0.76	(1.45)
PMI	0.02	(0.06)	0.08	(0.26)	−1.36*	(−1.85)
Consumer Confidence	0.28	(0.98)	0.32	(1.04)	0.17	(0.23)
Industrial Confidence	−0.60*	(−1.71)	−0.72*	(−1.79)	−0.15	(−0.24)
Industrial Production	0.88**	(2.30)	1.03**	(2.20)	0.40	(0.91)
New Orders	−0.69	(−1.28)	−0.67	(−1.25)	—	—

(continued)

Table A.1. (Continued)

	Full Sample		**Pre-disinflation**		**Below Target**	
	β	**t-stats**	β	**t-stats**	β	**t-stats**
Unemployment Rate	−0.58	(−1.31)	−0.50	(−0.79)	−0.58	(−1.32)
M3	−0.19	(−0.60)	−0.01	(−0.04)	−0.97	(−1.62)
MRO	—	—	—	—	0.04	(0.35)
Retail Sales	0.05	(0.17)	0.30	(0.73)	—	—
Trade Balance	−0.05	(−0.17)	−0.12	(−0.29)	0.07	(0.19)
Mon. Policy Shock	1.51	(0.13)	5.17	(0.39)	−16.52	(−0.47)
CB Info. Shock	21.59	(1.52)	19.75	(1.21)	9.10	(0.29)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.01		0.01		0.01	
R^2	0.02		0.03		0.05	
$H_0 : \beta = 0$ (p-value)	0.02		0.10		0.30	
Chow Test (p-value)	0					

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in long-term inflation compensation (1y9y, in basis points). The surprise component of macroeconomic releases, X_t, is normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero (joint significance test). Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.2. Alternative Long-Term Inflation Compensation (1y9y, daily): Individual Variable Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample			Pre-disinflation			Below Target		
	β	t-stats	R^2	β	t-stats	R^2	β	t-stats	R^2
Flash									
Germany	0.40	(1.08)	0.01	−0.13	(−0.26)	0.00	1.35***	(3.41)	0.14
Spain	0.89**	(2.55)	0.04	1.08*	(1.98)	0.05	0.59	(1.59)	0.03
Italy	−0.20	(−0.75)	0.00	−0.39	(−1.30)	0.01	0.29	(0.52)	0.01
Euro Area	−0.54	(−1.57)	0.02	−0.99**	(−2.13)	0.04	0.44	(1.17)	0.03
HICP									
Germany	0.39	(1.24)	0.01	0.32	(0.98)	0.01	1.18	(1.29)	0.04
Spain	0.15	(0.20)	0.00	0.17	(0.19)	0.00	0.34	(0.44)	0.01
France	−0.54	(−0.76)	0.01	−0.67	(−0.86)	0.02	0.64	(1.10)	0.02
Italy	0.35	(0.93)	0.01	0.35	(0.74)	0.01	0.30	(0.67)	0.01
Euro Area	−0.07	(−0.18)	0.00	−0.09	(−0.20)	0.00	0.02	(0.06)	0.00
PPI	−0.24	(−0.75)	0.00	−0.26	(−0.67)	0.00	−0.19	(−0.39)	0.00
GDP	0.72***	(2.67)	0.04	0.71**	(2.36)	0.04	0.88	(1.56)	0.04
PMI	0.04	(0.12)	0.00	0.11	(0.32)	0.00	−0.35	(−0.50)	0.00
Consumer Confidence	0.19	(0.66)	0.00	0.23	(0.72)	0.00	0.20	(0.26)	0.00
Industrial Confidence	−0.57	(−1.58)	0.01	−0.67	(−1.63)	0.02	0.03	(0.05)	0.00
Industrial Production	0.86**	(2.31)	0.05	0.96**	(2.18)	0.06	0.30	(0.63)	0.01
New Orders	−0.68	(−1.29)	0.02	−0.68	(−1.29)	0.02	—	—	—
Unemployment Rate	−0.69	(−1.62)	0.02	−0.73	(−1.23)	0.02	−0.40	(−1.00)	0.02
M3	−0.21	(−0.65)	0.00	−0.05	(−0.12)	0.00	−0.94	(−1.58)	0.04
MRO	—	—	—	—	—	—	0.09	(0.42)	0.00
Retail Sales	0.04	(0.12)	0.00	0.25	(0.58)	0.00	—	—	—
Trade Balance	−0.06	(−0.19)	0.00	−0.11	(−0.25)	0.00	0.02	(0.06)	0.00
Mon. Policy Shock	−4.64	(−0.41)	0.00	2.51	(0.17)	0.00	−20.97	(−1.07)	0.07
CB Info. Shock	21.84*	(1.69)	0.01	20.10	(1.25)	0.01	25.41	(1.45)	0.05

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for each single variable in an individual regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in long-term inflation compensation (1y9y, in basis points). The surprise component of macroeconomic releases, x_t, is normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. R^2 reports the explanatory variable for each individual regression. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g, new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.3. Long-Term Inflation Compensation (5y5y, daily): Joint Linear Regression (with controls) $\Delta y_t = \alpha + \beta X_t + \beta Z_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	0.60**	(2.33)	0.41	(1.12)	0.99***	(4.61)
Spain	0.41**	(2.10)	0.64**	(2.32)	−0.01	(−0.03)
Italy	−0.04	(−0.22)	−0.08	(−0.31)	−0.17	(−0.58)
Euro Area	0.10	(0.40)	0.19	(0.48)	−0.01	(−0.03)
HICP						
Germany	0.12	(0.66)	0.08	(0.41)	0.37	(0.79)
Spain	0.24	(1.05)	0.23	(0.85)	0.25	(0.74)
France	−0.08	(−0.49)	−0.15	(−0.88)	0.20	(0.59)
Italy	0.12	(0.67)	0.28	(1.44)	−0.38	(−1.28)
Euro Area	0.05	(0.34)	0.05	(0.29)	0.02	(0.12)
PPI	−0.02	(−0.09)	0.05	(0.18)	−0.14	(−0.52)
GDP	0.38**	(2.55)	0.42**	(2.54)	0.13	(0.39)
PMI	−0.06	(−0.30)	−0.05	(−0.28)	−0.01	(−0.03)
Consumer Confidence	0.10	(0.67)	0.03	(0.18)	0.36	(0.88)
Industrial Confidence	−0.14	(−0.73)	−0.19	(−0.85)	0.14	(0.41)
Industrial Production	0.41**	(2.57)	0.41**	(2.17)	0.33	(1.34)
New Orders	−0.27	(−1.17)	−0.25	(−1.06)	—	—
Unemployment Rate	−0.26	(−1.08)	−0.19	(−0.57)	−0.55*	(−1.84)
M3	−0.24	(−1.60)	−0.33*	(−1.94)	−0.04	(−0.14)

(*continued*)

Table A.3. (Continued)

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
MRO	—	—	—	—	0.35**	(2.48)
Retail Sales	0.11	(0.69)	0.09	(0.50)	—	—
Trade Balance	0.04	(0.19)	0.08	(0.26)	0.06	(0.21)
Mon. Policy Shock	7.00	(0.90)	11.52	(1.22)	−10.86	(−0.63)
CB Info. Shock	18.77	(1.57)	17.02	(1.24)	13.81	(0.66)
Citigroup EA Surprise	0.00	(0.55)	0.01	(0.93)	−0.02*	(−1.68)
Citigroup U.S. Surprise	0.02*	(1.85)	0.03*	(1.93)	0.00	(0.20)
Oil Price	0.06	(1.49)	0.00	(0.08)	0.23***	(3.17)
Gold Price	0.00	(−0.17)	0.00	(−0.09)	0.00	(−0.71)
FX EUR/USD	0.05	(0.29)	0.03	(0.14)	0.28	(0.66)
EA Bond Market Vol.	0.40*	(1.88)	0.31	(1.29)	0.35	(0.81)
VSTOXX	−0.10**	(−2.01)	−0.10*	(−1.77)	−0.06	(−0.64)
VIX	0.04	(0.80)	0.07	(1.22)	−0.11	(−1.05)
CDS	−0.09***	(−5.71)	−0.10***	(−5.61)	−0.04	(−1.21)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.06		0.06		0.06	
R^2	0.08		0.09		0.11	
$H_0 : \beta = 0$ (p-value)	0.00		0.00		0.00	
Chow Test (p-value)	0.1					

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target). Dependent variable is the change (basis points) in long-term inflation compensation (5y5y); the covariates consist of the surprise component of macroeconomic releases, $\mathbf{X}_t$, which is normalized by their historical standard deviations, and control variables $\mathbf{Z}_t$ that are observed at daily frequency; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.4. Long-Term Inflation Compensation (5y5y, two days): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	0.62**	(2.34)	0.56	(1.56)	0.82**	(2.19)
Spain	0.33	(1.10)	0.62	(1.45)	−0.12	(−0.34)
Italy	0.02	(0.05)	0.07	(0.15)	−0.25	(−0.46)
Euro Area	−0.07	(−0.15)	−0.16	(−0.23)	−0.04	(−0.09)
HICP						
Germany	−0.02	(−0.07)	−0.09	(−0.30)	0.56***	(3.09)
Spain	0.17	(0.63)	0.04	(0.14)	0.91	(1.62)
France	−0.12	(−0.47)	−0.26	(−1.00)	1.05**	(2.28)
Italy	0.39	(1.45)	0.58**	(2.11)	−0.14	(−0.24)
Euro Area	0.08	(0.36)	0.13	(0.44)	0.03	(0.15)
PPI	−0.37	(1.35)	−0.44	(−1.35)	−0.18	(−0.39)
GDP	0.49*	(1.91)	0.48	(1.62)	0.58	(1.38)
PMI	0.50	(1.53)	0.47	(1.37)	1.31**	(2.35)
Consumer Confidence	0.11	(0.47)	0.15	(0.61)	−0.14	(−0.36)
Industrial Confidence	−0.25	(−0.53)	−0.27	(−0.46)	−0.31	(−0.68)
Industrial Production	0.54	(2.08)	0.64**	(2.13)	0.20	(0.48)
New Orders	0.09	(0.26)	0.10	(0.28)	—	—

(continued)

Table A.4. (Continued)

	Full Sample		**Pre-disinflation**		**Below Target**	
	β	**t-stats**	β	**t-stats**	β	**t-stats**
Unemployment Rate	−0.30	(−0.86)	−0.04	(−0.07)	−0.99***	(−2.92)
M3	−0.20	(−0.73)	−0.52*	(−1.67)	1.21**	(2.52)
MRO	—	—	—	—	0.76***	(4.37)
Retail Sales	0.04	(0.14)	−0.25	(−0.76)	—	—
Trade Balance	0.10	(0.40)	0.11	(0.30)	0.11	(0.31)
Mon. Policy Shock	20.26	(1.80)	28.27**	(2.25)	−2.86	(−0.13)
CB Info. Shock	44.83***	(2.62)	40.20***	(2.08)	44.46	(1.69)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.01		0.01		0.02	
R^2	0.03		0.03		0.06	
$H_0 : \beta = 0$ (p-value)	0.00		0.04		0.00	
Chow Test (p-value)	0.00					

Notes: The table reports estimated β coefficients from regressions on days with data releases from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is changes (basis points) in long-term inflation compensation (5y5y) across a two-day window, $\Delta y_t \equiv y_{t+1} - y_{t-1}$. The surprise component of macroeconomic releases, $\mathbf{X}_t$, is normalized by their historical standard deviations; coefficients therefore represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero (joint significance test). Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.5. Long-Term Inflation Compensation (5y5y, two days): Individual Variable Linear Regression$\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample			Pre-disinflation			Below Target		
	β	t-stats	R^2	β	t-stats	R^2	β	t-stats	R^2
Flash									
Germany	0.68**	(2.58)	0.05	0.56	(1.54)	0.03	0.91**	(2.66)	0.09
Spain	0.41	(1.38)	0.02	0.45	(0.96)	0.02	0.32	(0.98)	0.02
Italy	0.04	(0.15)	0.00	0.00	(0.01)	0.00	0.12	(0.24)	0.00
Euro Area	0.05	(0.16)	0.00	−0.08	(−0.15)	0.00	0.22	(0.62)	0.01
HICP									
Germany	−0.04	(−0.15)	0.00	−0.09	(−0.31)	0.00	0.46***	(2.91)	0.01
Spain	0.22	(0.84)	0.01	0.14	(0.49)	0.00	0.68	(1.17)	0.03
France	−0.08	(−0.33)	0.00	−0.19	(−0.70)	0.00	0.73	(1.53)	0.03
Italy	0.37	(1.53)	0.02	0.53**	(2.07)	0.03	0.00	(0.01)	0.00
Euro Area	0.08	(0.36)	0.00	0.10	(0.36)	0.00	−0.05	(−0.22)	0.00
PPI	−0.33	(−1.24)	0.01	−0.41	(−1.24)	0.02	−0.12	(−0.28)	0.00
GDP	0.48**	(2.02)	0.03	0.45	(1.59)	0.02	0.76*	(1.79)	0.04
PMI	0.51	(1.53)	0.02	0.47	(1.37)	0.02	1.64**	(2.50)	0.04
Consumer Confidence	0.08	(0.30)	0.00	0.13	(0.45)	0.00	−0.10	(−0.27)	0.00
Industrial Confidence	−0.24	(−0.50)	0.01	−0.18	(−0.30)	0.00	−0.42	(−0.83)	0.01
Industrial Production	0.50	(1.99)	0.03	0.58	(2.03)	0.04	0.19	(0.43)	0.00
New Orders	0.06	(0.17)	0.00	0.06	(0.17)	0.00	—	—	—
Unemployment Rate	−0.33	(−0.92)	0.01	−0.17	(−0.37)	0.00	−0.64*	(−1.93)	0.06
M3	−0.23	(−0.80)	0.01	−0.56	(−1.80)	0.04	1.24**	(2.63)	0.12
MRO	—	—	—	—	—	—	0.56	(1.94)	0.06
Retail Sales	−0.01	(−0.02)	0	−0.45	(−1.38)	0.02	—	—	—
Trade Balance	0.12	(0.46)	0.00	0.12	(0.34)	0.00	0.10	(0.32)	0.00
Mon. Policy Shock	8.17	(0.75)	0.01	22.92*	(1.71)	0.04	−27.18*	(−1.96)	0.11
CB Info. Shock	38.20**	(2.41)	0.10	39.42*	(1.92)	0.10	42.57	(2.95)	0.14

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for each single variable in an individual regression. Dependent variable is changes across a two-day window, $\Delta y_t \equiv y_{t+1} - y_{t-1}$, in long-term inflation compensation (5y5y, in basis points). The surprise component of macroeconomic releases, x_t, is normalized by their historical standard deviations; coefficients β therefore represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; and ***, **, * denote a 1, 5, and 10 percent level of significance, respectively. R^2 reports the explanatory variable for each individual regression. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.6. Short-Term Inflation Compensation (1-year spot IL swap rate, daily): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	2.14***	(5.67)	1.58***	(4.44)	3.20***	(3.83)
Spain	0.80*	(1.87)	0.71	(1.64)	0.51	(0.55)
Italy	1.21***	(2.90)	0.85**	(1.96)	1.29	(1.48)
Euro Area	1.70***	(2.43)	1.67***	(2.22)	1.06	(0.91)
HICP						
Germany	0.25	(1.36)	0.25	(1.23)	0.03	(0.06)
Spain	−0.52	(−1.63)	−0.43	(−1.25)	−0.99*	(−1.70)
France	0.87***	(3.69)	0.83***	(3.28)	1.11	(1.46)
Italy	−0.05	(−0.22)	0.15	(0.57)	−0.76	(−1.37)
Euro Area	0.52**	(2.23)	0.62**	(2.45)	0.13	(0.26)
PPI	−0.22	(−0.61)	−0.36	(−0.76)	0.38	(0.84)
GDP	0.28	(1.32)	0.26	(1.08)	−0.03	(−0.06)
PMI	3.56***	(2.84)	3.69***	(2.89)	1.25	(0.35)
Consumer Confidence	−0.41*	(−1.71)	−0.44	(−1.63)	−0.06	(−0.12)
Industrial Confidence	−0.23	(−0.72)	−0.09	(−0.25)	−1.14	(−1.43)
Industrial Production	0.29	(0.72)	0.33	(0.69)	−0.06	(−0.10)
New Orders	0.03	(0.06)	0.07	(0.14)	—	—

(continued)

Table A.6. (Continued)

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Unemployment Rate	0.76	(0.91)	2.60***	(3.04)	−3.31**	(−2.08)
M3	0.20	(0.90)	0.19	(0.91)	0.20	(0.29)
MRO	—	—	—	—	1.09	(1.19)
Retail Sales	−0.35	(−1.41)	−0.59*	(−1.68)	—	—
Trade Balance	−0.27	(−0.93)	−0.36	(−0.94)	−0.15	(−0.38)
Mon. Policy Shock	−30.91	(−1.22)	−38.83	(−1.34)	−26.48	(−0.68)
CB Info. Shock	27.05	(1.38)	37.87*	(1.93)	8.60	(0.22)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.06		0.08		0.07	
R^2	0.08		0.09		0.01	
$H_0 : \beta = 0$ (p-value)	0		0		0	
Chow Test (p-value)	0					

Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in short-term inflation compensation (1-year spot IL swap rate, in basis points). The surprise component of macroeconomic releases, $\mathbf{X}_t$, is normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Table A.7. Short-Term Inflation Compensation (1-year forward IL swap rate 1-year ahead, daily): Joint Linear Regression $\Delta y_t = \alpha + \beta X_t + \varepsilon_t$

	Full Sample		Pre-disinflation		Below Target	
	β	t-stats	β	t-stats	β	t-stats
Flash						
Germany	1.35***	(5.32)	1.32***	(4.17)	1.37***	(3.27)
Spain	0.20	(0.67)	0.29	(0.73)	0.05	(0.10)
Italy	0.59**	(2.06)	0.40	(1.06)	0.86**	(1.96)
Euro Area	0.81**	(2.14)	0.90*	(1.87)	0.49	(0.85)
HICP						
Germany	0.33*	(1.65)	0.32	(1.49)	0.34	(0.78)
Spain	−0.57*	(−1.82)	−0.60	(−1.63)	−0.45	(−1.11)
France	0.87***	(3.92)	0.89***	(3.75)	0.58	(0.82)
Italy	−0.07	(−0.29)	−0.01	(−0.03)	−0.31	(−0.64)
Euro Area	−0.02	(−0.09)	0.05	(0.16)	−0.26	(−0.64)
PPI	−0.13	(−0.52)	0.06	(0.17)	−0.46	(−1.28)
GDP	−0.38	(−1.57)	−0.46*	(−1.83)	0.09	(0.23)
PMI	0.03	(0.08)	0.00	(−0.01)	0.66	(0.65)
Consumer Confidence	−0.63**	(−1.99)	−0.75**	(−2.07)	0.02	(0.05)
Industrial Confidence	−0.40	(−1.10)	−0.44	(−1.03)	−0.27	(−0.62)
Industrial Production	0.09	(0.23)	0.12	(0.27)	−0.04	(−0.08)
New Orders	−0.14	(−0.50)	−0.13	(−0.46)	—	—

(*continued*)

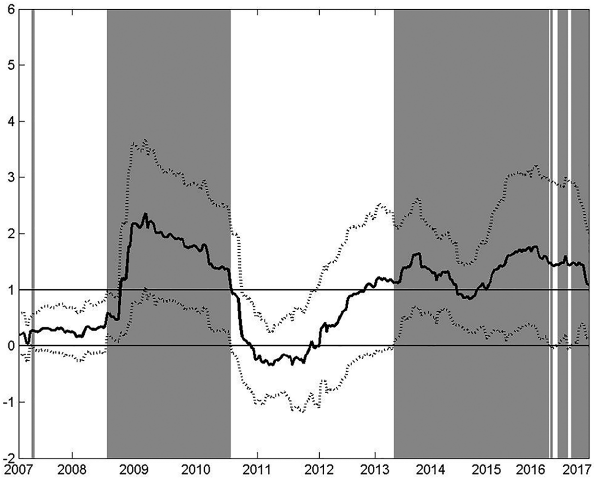
Table A.7. (Continued)

	Full Sample		**Pre-disinflation**		**Below Target**	
	β	**t-stats**	β	**t-stats**	β	**t-stats**
Unemployment Rate	−0.26	(−0.89)	0.16	(0.52)	−1.15**	(−2.29)
M3	0.21	(0.93)	0.21	(0.85)	0.17	(0.34)
MRO	—	—	—	—	0.38**	(2.29)
Retail Sales	−0.19	(−0.75)	−0.31	(−0.86)	—	—
Trade Balance	−0.34	(−1.39)	−0.47	(−1.42)	−0.09	(−0.31)
Mon. Policy Shock	−24.87	(−0.97)	−29.37	(−0.99)	−22.12	(−0.74)
CB Info. Shock	25.27	(1.12)	30.89	(1.36)	13.25	(0.48)
Observations	1,767		1,209		525	
$\bar{R}^2$	0.04		0.04		0.04	
R^2	0.05		0.05		0.08	
$H_0 : \beta = 0$ (p-value)	0		0		0	
Chow Test (p-value)	0					

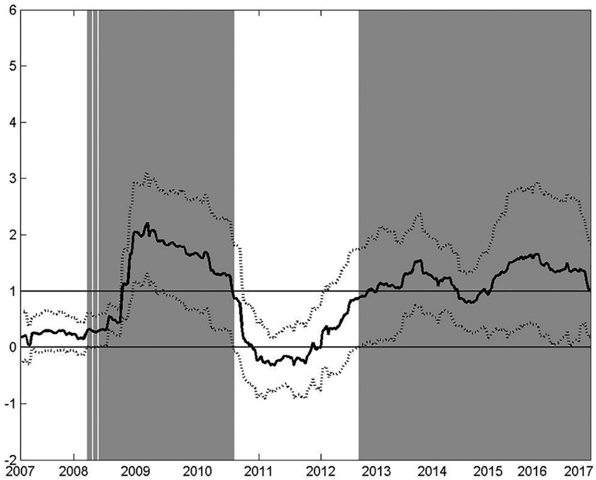
Notes: The table reports estimated β coefficients from regressions on days with data releases over three periods: from January 2005 to March 2017 (full sample), from January 2005 to December 2012 (pre-disinflation), and from January 2013 to March 2017 (below target), using surprises for all variables in a joint regression. Dependent variable is one-day change, $\Delta y_t \equiv y_t - y_{t-1}$, in short-term inflation compensation (one-year forward IL swap rate one year ahead, in basis points). The surprise component of macroeconomic releases, $\mathbf{X}_t$, is normalized by their historical standard deviations; coefficients represent a basis point per standard deviation response. Heteroskedasticity-consistent t-statistics are in parentheses; ***, **, and * denote a 1, 5, and 10 percent level of significance, respectively. $H_0 : \beta = 0$ p-value is for the test that all elements of β are zero. Chow test (p-value) is for the H_0 of parameter constancy between the pre-disinflation and below-target period. (—) indicates an omission either due to the discontinuity of the expectations collection (e.g., new orders) or for lack of surprises over most of the period of the regression (e.g., MRO).

Figure A1. Long-Term Inflation Compensation (5y5y, daily, alternative normalization): Time-Varying Sensitivity Coefficients δ^{τ} from Nonlinear Regression

A. δ^{τ} Normalized across 2005–08



B. δ^{τ} Normalized across 2008–12



Notes: This figure plots the coefficients δ^{τ} estimated from a regression $\Delta y_t = \gamma^{\tau} + \delta^{\tau}\hat{\mathbf{X}}_t + \epsilon_t^{\tau}$. δ^{τ_i} has been normalized to an average value of 1 in panel A for the years $i = 2005, \ldots, 2008$ and in panel B for the years $i = 2008, \ldots, 2012$. Dotted gray lines depict heteroskedasticity-consistent $\pm$ two-standard-error bands, adjusted for two-stage sampling uncertainty in δ^{τ}. Shaded regions denote δ^{τ} being significantly above 0 and suggest a weaker anchoring of inflation expectations.

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Monetary Policy Credibility and Exchange Rate Pass-Through*

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A long-standing conjecture in macroeconomics is that declines in exchange rate pass-through over the past three decades are in part due to improved monetary policy performance. In a large sample of emerging and advanced economies, we find evidence that a relatively more credible monetary policy regime—measured by better-anchored inflation expectations—is associated with lower exchange rate pass-through to consumer prices. The results are robust to controlling for the level and variability of nominal variables and for the import content of the consumption basket.

JEL Codes: E31, E52, F41.

1. Introduction

The empirical literature has reported wide variation in the rate at which changes in the nominal exchange rate pass through to domestic prices, both across countries and over time. Many empirical papers document a generalized decline in pass-through rates over the past three decades (Campa and Goldberg 2005; Choudhri and

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Hakura 2006). Taylor (2000) conjectures that the pass-through rate is endogenous to the monetary policy framework and its credibility, such that improvements in the latter—reflected in stronger nominal anchors and a track record of price stability—could be responsible for falling exchange rate pass-through to consumer prices. The argument behind Taylor (2000)'s conjecture is that the extent to which a firm decides to pass along an increase in its costs is lower when their inflation expectations are better anchored.

This paper provides an empirical test of Taylor's hypothesis. We find evidence that greater monetary policy credibility, as measured by better-anchored inflation expectations, has significantly reduced the exchange rate pass-through to consumer prices.

We begin by estimating models of exchange rate pass-through to consumer prices in a sample of 62 emerging and advanced economies. In line with the existing literature, we confirm that the degree of exchange rate pass-through to consumer prices has fallen over the past few decades, with the largest decline registered among emerging economies. But we also document substantial heterogeneity in exchange rate pass-through coefficients across countries. We then estimate exchange rate pass-through to prices of imported goods at the border—or pure traded goods—and use input-output tables to compute the import content of domestic consumption as in Burstein, Eichenbaum, and Rebelo (2005) and Gopinath (2015).[1] We argue that heterogeneity in pass-through rates across countries and over time cannot be explained by differences in the response of import prices or in the composition of consumption baskets, but rather is related to changes in price-setting behavior for domestically produced goods and services.

We then focus on the role of monetary policy credibility—proxied by how well-anchored inflation expectations are—in explaining the

[1]Burstein, Eichenbaum, and Rebelo (2005) show that the usual decomposition of consumer prices into tradable and nontradable components that relies on retail prices can be misleading for pass-through analysis. The reason is that the retail price of tradable goods includes two sizable nontradable components: distribution costs—including wholesale and retail services, marketing, advertising, and local transportation services—and local goods that are produced only for the local market. These components reflect the pricing of locally produced goods and services that are unlikely to be arbitraged in international markets, while prices of imported goods at the border better capture the pricing behavior of "pure" traded goods.

heterogeneity in estimates of exchange rate pass-through to consumer prices. We find that the metrics used in earlier studies to test Taylor's hypothesis, such as the level and volatility of inflation (e.g., Gagnon and Ihrig 2004; Choudhri and Hakura 2006), continue to be related to the degree of exchange rate pass-through to consumer prices in our sample.[2] However, we argue that these proxies may be imperfect measures of monetary policy credibility, since their variation may also capture properties of underlying shocks over which central banks have little control (Devereux and Yetman 2010).

We argue that inflation forecasts offer relevant complementary information. As shown by Svensson (1997), the credibility of monetary policy is a forward-looking concept that should be reflected in the degree to which inflation expectations are anchored (see also Demertzis, Marcellino, and Viegi 2012). We exploit two empirical characteristics of inflation forecasts that are associated with well-anchored inflation expectations (Capistrán and Ramos-Francia 2010; Dovern, Fritsche, and Slacalek 2012; Ehrmann 2015; Kumar et al. 2015): average inflation forecasts should be relatively stable over time, and the dispersion of individual inflation forecasts should be small. All else equal, professional forecasters should tend to agree more on the future path of the price level relative to an economy where there is not enough clarity on the objectives and degree of commitment of the central bank to those objectives.

We estimate pass-through rates from country-specific rolling regressions, and regress these estimates on a set of metrics of monetary performance, including the variability of average inflation forecasts and the extent of disagreement across individual forecasters. We find that monetary regimes where inflation expectations are relatively better anchored exhibit lower exchange rate pass-through to consumer prices. The results are robust to conditioning on the level and volatility of inflation and the exchange rate, the import content of consumption, and time and country fixed effects.

[2]Theoretical work has also argued that as average inflation or inflation volatility increase, firms adjust prices more frequently and this leads to higher exchange rate pass-through to domestic prices (Devereux and Yetman 2010). Bouakez and Rebei (2008) estimate a dynamic general equilibrium model for the Canadian economy and conclude that the decline of consumer price pass-through can be largely attributed to the adoption of inflation targeting.

The rest of the paper is organized as follows. Section 2 presents some conceptual considerations underpinning our empirical analysis. Section 3 presents the empirical approach to estimate exchange rate pass-through to consumer prices, describes our data, and documents the extent of cross-country and time variation in pass-through estimates. Our main contribution is in section 4, where we explore the role of monetary policy performance and credibility in explaining exchange rate pass-through to consumer prices. Section 5 concludes.

2. Exchange Rate Pass-Through and Monetary Policy Credibility: Conceptual Considerations

The main hypothesis to be tested in this paper is that countries whose monetary policy is less credible exhibit a higher degree of exchange rate pass-through to consumer prices. To set the stage for our empirical analysis, we start with a discussion of two aspects that underpin our hypothesis and empirical approach. The first is an argument for why the extent of exchange rate pass-through is linked to monetary policy performance, and specifically to the credibility of monetary policy. The second is a discussion of empirical proxies for the degree of monetary policy credibility. Then we present a simple schematic model to motivate the empirical specification.

2.1 Conceptual Framework

We start from the simple framework presented by Devereux and Yetman (2010), which features nominal rigidities à la Calvo (1983). In their framework, forward-looking firms set their price as a function of the current and expected future exchange rates, such that the persistence of an underlying shock to the exchange rate influences prices. Since the price level is an aggregation of prices set by individual firms, this mechanism creates a link between the exchange rate and inflation.

How does monetary policy affect this link? For a given persistence of the underlying shock, monetary policy has no effect on exchange rate pass-through when the frequency at which firms adjust prices is assumed to be exogenous. Following a shock, tighter monetary policy can dampen the responses of prices and of the exchange rate, but does not affect the pass-through rate. In this setting, however, the pass-through rate does depend on the persistence of the underlying perturbation. Intuitively, the more protracted the shock is, the larger the response of inflation, which leads to a higher pass-through rate in the short run. Differences across countries and over time in the level and the volatility of inflation and the volatility of the exchange rate may thus reflect differences not only in monetary policy performance but also in the time-series properties of the underlying shocks.

However, when Devereux and Yetman (2010) extend the model to allow firms to choose how frequently to set prices, monetary policy can affect the degree of exchange rate pass-through. In this alternative setting, the volatility of inflation and the exchange rate would reflect not only the properties of the underlying shocks but also the influence of monetary policy via its effect on the frequency of price adjustments by firms (Devereux and Yetman 2002).

What does this mean for empirically testing the link between the degree of exchange rate pass-through and monetary policy credibility? It implies that controlling for cross-country differences in average inflation, the volatility of inflation, and the volatility of the exchange rate—as previous empirical studies have done (i.e., Gagnon and Ihrig 2004)—is correct but not sufficient. We go one step further and test the hypothesis by considering the role of monetary policy credibility while controlling for realized inflation and exchange rate volatility.

How can we measure the credibility of monetary policy in practice? Our rationale for using metrics from surveys of inflation expectations is based on Svensson (1997). He argues that the central bank does not have perfect control over inflation, since unexpected shocks hit the economy all the time. Given the lag in the transmission of monetary policy, these shocks will inevitably move inflation temporarily away from the target. Therefore, deviations of inflation from target or inflation volatility do not necessarily reflect the credibility (or lack thereof) of monetary policy, as they include the effect of unexpected shocks. Rather, the credibility of the central bank is given by how well inflation expectations remain anchored at the central bank's policy horizon, given the available information set. In fact, Svensson (1997) states that "the central bank should be held accountable for the forecast deviations from the target rather than the realized inflation deviations . . . Adjusting the instrument so that the inflation forecast equals the target is the best the central bank can do." Applying Svensson's criteria literally would mean focusing only on inflation targeters. Instead, we will proxy monetary policy credibility with measures of the stability and coordination of inflation forecasts, which the literature has associated with well-anchored inflation expectations (Capistrán and Ramos-Francia 2010; Dovern, Fritsche, and Slacalek 2012; Ehrmann 2015; Kumar et al. 2015). These measures are strongly correlated with Svensson's proposed metric in a sample of inflation targeters. We present the empirical strategy based on the above theoretical considerations in the following section.

2.2 Rationale for the Econometric Specification

The domestic price level of any country i reflects an aggregation of prices of tradable ($P_t^{T,i}$) and nontradable ($P_t^{NT,i}$) goods and services. Using, for simplicity, a Cobb-Douglas aggregator with σ denoting the weight of tradable goods and services in the domestic consumption basket, the domestic price level is given by

$$P_t^i = (P_t^{T,i})^{\sigma}(P_t^{NT,i})^{1-\sigma}, \tag{1}$$

where $P_t^{T,i} = E_t^i P_t^{x,i}, E_t^i$ is the nominal exchange rate expressed as units of domestic currency per U.S. dollar, and $P_t^{x,i}$ denotes the export prices of country i's trading partners expressed in U.S. dollars. For simplicity, this two-country setting assumes that all tradable consumption is imported.

We further consider that there is a fraction γ of tradable goods in the consumption basket for which importers and exporters are price takers (i.e., international commodities). Denoting natural logarithms with lowercase letters, the price of these commodities, $p_t^{com,i}$, which include oil and food, are determined in global markets and are denominated in U.S. dollars:

$$p_t^{x,i} = \gamma p_t^{com,i} + (1-\gamma)\, p_t^{non-com,i} \tag{2}$$

$$p_t^{com,i} = d_1 oil_t + d_2 food_t. \tag{3}$$

As in Campa and Goldberg (2005), we posit that most export prices are a markup ($mkup_t^x$) over the exporter's marginal cost (mc_t^x).

Thus, we assume that the export price of noncommodity tradable goods and services is given by

$$p_t^{non-com,i} = mkup_t^x + mc_t^x. \tag{4}$$

We proxy marginal costs of country i's trading partners with producer price indexes, $ppi_t^{x,USD}$ (expressed in U.S. dollars). We posit that markups are sensitive to demand conditions in the destination market, which are proxied by the output gap, gap_t:

$$p_t^{non-com,i} = b_1 gap_t + b_2 ppi_t^{x,USD}. \tag{5}$$

As the producer price index of country x in U.S. dollars, $ppi_t^{x,USD}$, is a function of the producer price index in country x's domestic currency and its bilateral exchange rate with respect to the U.S. dollar, e_t^x, the export price of noncommodity tradable goods and services can be expressed as

$$p_t^{non-com,i} = b_1 gap_t - b_2 e_t^x + b_2 ppi_t^{x,LC}. \tag{6}$$

Finally, we assume that domestic prices of nontradable goods are a markup over domestic marginal costs, mc_t^i, which are themselves a function of domestic demand conditions, gap_t:

$$p_t^{NT,i} = mkup_t^i + mc_t^i = a_1 gap_t. \tag{7}$$

Thus, the domestic price level (equation (1)) expressed in natural logarithms becomes

$$\begin{aligned} p_t^i = (1-\sigma)\, p_t^{NT,i} + \sigma \Big[\gamma \left(e_t^i + p_t^{comm,i} \right) \\ + (1-\gamma) \left(e_t^i + p_t^{non-com,i} \right) \Big]. \end{aligned} \tag{8}$$

Substituting for $p_t^{NT,i}$, $p_t^{comm,i}$, and $p_t^{non-com,i}$ using equations (3), (6), and (7) into (8) gives

$$\begin{aligned} p_t^i = (1-\sigma)\, a_1 gap_t + \sigma\gamma \left(e_t^i + d_1 oil_t + d_2 food_t \right) \\ + \sigma (1-\gamma) \left(e_t^i + b_1 gap_t - b_2 e_t^x + b_2 ppi_t^{x,LC} \right). \end{aligned}$$

Rearranging terms yields

$$\begin{aligned} p_t^i &= \sigma e_t^i - [\sigma(1-\gamma) b_2] e_t^x + (\sigma\gamma d_1)\, oil_t + (\sigma\gamma d_2)\, food_t \\ &\quad + [\sigma(1-\gamma) b_2]\, ppi_t^{x,LC} \\ &\quad + [(1-\sigma) a_1 + \sigma(1-\gamma) b_1]\, gap_t, \end{aligned} \tag{9}$$

which can be simplified to the following expression:

$$p_t^i = \beta_1 e_t^i - \beta_2 e_t^x + \rho oil_t + \delta food_t + \varphi ppi_t^{x,LC} + \vartheta gap_t. \tag{10}$$

3. Exchange Rates and Consumer Prices: Empirical Strategy

We begin our empirical analysis with an estimation of the overall impact of a currency movement on consumer prices in a sample of 32 advanced and 30 emerging market economies using data from January 1995 to June 2019. The reduced-form specification is a variant of standard empirical models (Campa and Goldberg 2005; Gopinath 2015), along the lines described in the previous section (equation (10)) when generalized to N trading partners. We estimate cumulative response functions in country-specific and panel settings using Jordà's (2005) local projection method.

In line with the above model, the panel specification is given by

$$\begin{aligned} p_{i,t+h-1} - p_{i,t-1} &= \alpha^h + \sum_{j=0}^{J} (\beta_j^h \Delta neer_{i,t-j} + \gamma_j^h \Delta oil_{i,t-j} \\ &\quad + \delta_j^h \Delta food_{i,t-j} + \vartheta_j^h gap_{i,t-j} + \varphi_j^h \Delta mPPI_{i,t-j}) \\ &\quad + \sum_{j=1}^{J} \rho_j^h \Delta p_{i,t-j} + \mu_i^h + \varepsilon_{i,t}^h, \end{aligned} \tag{11}$$

where $p_{i,t}$ denotes the natural logarithm of the price level in country i during period t (such that the dependent variable measures cumulative inflation between $t-1$ and $t+h$); *neer* is the natural logarithm of the import-weighted nominal effective exchange rate, as in Gopinath (2015) (see below for details); *oil* and *food* are the natural logarithm of international oil and food prices in U.S. dollars; *gap*, is the output gap, proxied by the cyclical component of industrial production;

and $mPPI$ is the natural logarithm of the import-weighted producer price index of countries from which country i imports, which proxies for the cost of production in trading partners (see below for details). The Δ operator denotes a first difference; μ_i are country fixed effects; and $\varepsilon_{i,t}$ is a random disturbance.

We include six lags in our specification, which is estimated by ordinary least squares using data at monthly frequency. It is worth noting, however, that there is considerable uncertainty surrounding the timing of the inflationary effects from depreciation, due to differences in microstructures across sectors and countries, including different degrees and nature of nominal rigidities. This can be reflected in nonlinear responses of consumer prices to depreciations or in different lag structures across countries. Since we conduct panel and country-specific regressions, and for simplicity use the same specification, a flexible estimation method that is robust to misspecification is desirable. Thus, the choice of using Jordà's (2005) local projections method—rather than a vector autoregressive model, for instance—follows primarily from this objective.[3]

Since we have defined the dependent variable in our regression equation (11) in cumulative terms, the estimate of β_0^h is the cumulative impact of an innovation in the nominal effective exchange rate on the consumer price index.[4] We do not take a stand on the underlying source of variation in the exchange rate. Therefore, the responses we report should be interpreted as conditional on the average constellation of shocks that moved the exchange rate during the sample period. Identification of the structural shock behind the currency depreciation is critical for questions related to the optimal policy response to the shock (e.g., Albagli, Naudon, and Vergara 2015; Forbes, Hjortsoe, and Nenova 2018). Our focus, instead, is on the role of monetary policy credibility in explaining cross-country and time differences in average pass-through coefficients.

[3] Jordà (2005) and Teulings and Zubanov (2014) present Monte Carlo simulations showing that the local projections method is more robust to misspecification than autoregressive models.

[4] We follow Jordà's (2005) suggestion and include the residual from the estimation corresponding to horizon $h-1$ as an additional regressor in the estimation for horizon h to improve the efficiency of the estimator. Adding the residual from the regression for horizon $h-1$ also addresses a potential bias identified in Teulings and Zubanov (2014).

Table 1. Economies in Sample

Emerging		Advanced	
Argentina	Malaysia	Australia	Korea
Bolivia	Mexico	Austria	Latvia
Brazil	Pakistan	Belgium	Lithuania
Bulgaria	Panama	Canada	Luxembourg
Chile	Paraguay	Czech Republic	New Zealand
China	Peru	Denmark	Norway
Colombia	Philippines	Estonia	Portugal
Costa Rica	Poland	Finland	Singapore
Ecuador	Romania	France	Slovak Republic
El Salvador	Russia	Germany	Slovenia
Guatemala	South Africa	Greece	Spain
Honduras	Thailand	Hong Kong S.A.R.	Sweden
Hungary	Turkey	Ireland	Switzerland
India	Ukraine	Israel	The Netherlands
Indonesia	Uruguay	Italy	United Kingdom
		Japan	United States

3.1 Data Description

Tables 1 and 2 report the sample coverage and summary statistics, respectively. The dependent variable in most of the analysis corresponds to the headline consumer price index, as reported in the International Monetary Fund's (IMF's) International Financial Statistics and by Haver Analytics.[5] In estimations of pass-through at the border, the dependent variable is import prices expressed in local currency.[6]

[5]The sample for Argentina uses data from January 1995 to December 2010, before a gap between the official and the parallel exchange rate emerged. Consumer price index (CPI) data after December 2006 corresponds to private analysts' estimates.

[6]Import price series are less widely available than consumer price indexes, and their comparability across countries and over time is fraught with difficulties due to important methodological differences (Burstein and Gopinath 2014). Moreover, commonly used data sources do not always indicate the currency of denomination requiring careful inspection and manipulation. The sample of countries with available import price data, the data sources, and the currency in which the original data is reported are documented in table A.1 in the appendix.

Table 2. Summary Statistics: January 1995–June 2017

Variables	Emerging Economies				Advanced Economies			
	N	**Mean**	**Std. Dev.**	**Quartiles**	**N**	**Mean**	**Std. Dev.**	**Quartiles**
$\Delta p_{i,t}$	8,075	0.7	2.0	$(0.1, 0.9)$	8,616	0.2	0.5	$(-0.1, 0.4)$
$\Delta NEER_{i,t}$	8,052	0.3	3.4	$(-0.8, 0.9)$	8,568	–0.1	1.4	$(-0.7, 0.5)$
$gap_{i,t}$	7,557	0.1	4.5	$(-1.9, 2.1)$	8,587	0.0	4.4	$(-2.0, 2.2)$
$\Delta mPPI_{i,t}$	8,052	0.3	0.6	$(0.0, 0.5)$	8,568	0.2	0.5	$(-0.0, 0.5)$

Source: Authors' calculations.

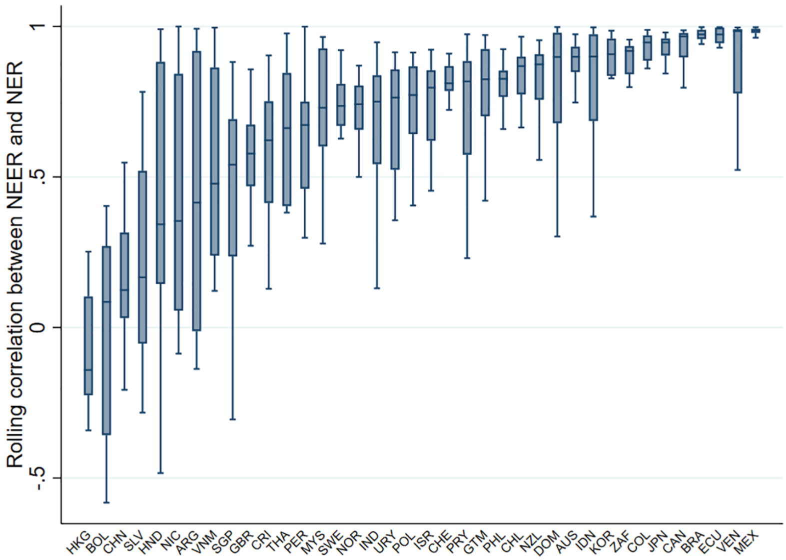
International oil and food prices correspond to composite indexes in U.S. dollars reported in the IMF's *World Economic Outlook*. The output gap is approximated by the cyclical component of industrial production when available, and to an interpolated quarterly GDP (gross domestic product) series otherwise. The cyclical component of industrial production is computed using a Hodrick-Prescott (HP) filter with smoothing coefficient equal to 129,600 on monthly data. We mitigate the endpoint bias in our filter because the estimations include data for control variables through June 2017, despite the data set extending to June 2019.[7]

We estimate the pass-through from changes in the nominal effective exchange rate (*neer*), rather than the bilateral exchange rate against the U.S. dollar. As argued by Caselli and Roitman (2019), the *neer* summarizes more closely the complete set of relative price adjustments that can be expected to affect the consumer price index. This choice is not innocuous, since bilateral exchange rate dynamics often diverge significantly from those of the nominal effective exchange rate, and their degree of co-movement varies a great deal over time and across countries. Figure 1 shows the distribution of country-specific pairwise correlations between monthly changes in the nominal effective exchange rate and the bilateral exchange rate against the U.S. dollar over three-year rolling windows starting in 1993. These correlations have frequently been far from unity in several economies, with the median correlation being smaller than 0.5 in many cases.

Following Gopinath (2015), the multilateral nominal effective exchange rate ($neer_{i,t}$) is constructed as a weighted average of the bilateral exchange rate of each trading partner vis-à-vis the U.S. dollar, weighted by their import shares. This approach is more appropriate than using total trade as weights, as it better captures the impact of a currency depreciation on domestic prices, since the composition of exports by destination can differ substantially from the composition of imports by origin. More precisely, the monthly percentage change for country i at time t is given by

[7] This is done to ensure that estimations at all horizons up to 24 months reflect the same underlying data.

Figure 1. Correlation between Bilateral and Multilateral Nominal Effective Exchange Rates



Sources: Bloomberg, L.P.; Haver Analytics; and authors' calculations.
Notes: The figure shows the distribution of rolling correlations between the multilateral nominal effective exchange rate (NEER) and the bilateral exchange rate against the U.S. dollar (NER) for each country over three-year rolling windows starting in 1993. The rectangles and the central marker denote the interquartile range and the median of its distribution, respectively.

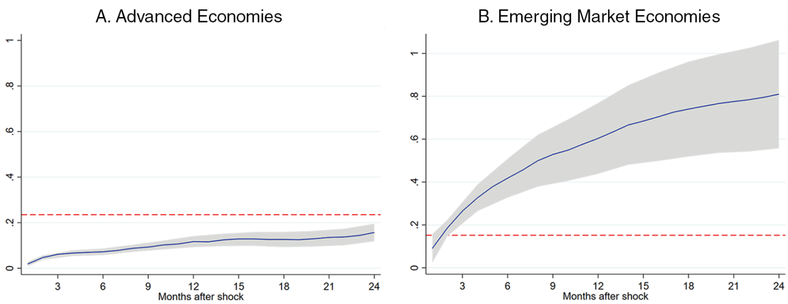
$$\Delta neer_{i,t} = \sum_{j=1}^{J} \omega_{ij,t} \left(\Delta e_{i,t} - \Delta e_{j,t} \right), \quad i \neq j, \tag{12}$$

where $e_{i,t}$ is the natural logarithm of country i's bilateral exchange rate (in local currency per U.S. dollar); Δ is the first-difference operator; and $\omega_{ij,t}$ is the share of exports from country j to country i in country i's total imports, as reported in the IMF's Direction of Trade Statistics. Weights have been lagged one year and are measured at annual frequency.

Using the same import weights $\omega_{ij,t}$, the monthly change in the cost of production in country i's import partners is proxied by

$$\Delta mPPI_{i,t} = \sum_{j=1}^{J} \omega_{ij,t} \Delta PPI_{j,t}, \quad i \neq j, \tag{13}$$

Figure 2. Cumulative Impulse Response of Consumer Prices Following a Nominal Effective Depreciation of 1 Percent (in percentage points)



Notes: Cumulative impulse response of headline consumer prices (in percentage points) to a 1 percent innovation in the nominal effective exchange rate estimated using local projection methods. Shaded bands correspond to 95 percent confidence intervals. Dashed lines denote the share of household final consumption that is imported (including direct imports and the import content of domestically produced goods) averaged over the sample period and across countries in each group.

where $PPI_{j,t}$ is the natural logarithm of country j's producer price index.

3.2 Estimation Results

Figure 2 reports the cumulative impulse responses of headline consumer prices to a 1 percent innovation in the nominal effective exchange rate and the corresponding 95 percent confidence bands, under the baseline sample that uses data from January 1995 to June 2019. We start by reporting panel estimates with countries pooled according to their income-based designation of advanced versus emerging market economies. We focus our discussion on pass-through coefficients corresponding to the cumulative percentage increase in the headline CPI one and two years after each percentage-point increase in the nominal effective exchange rate. There are important differences between estimates when countries are pooled by income group. In line with earlier studies (e.g., Choudhri and Hakura 2006), we find that the pass-through rate is lower for advanced economies than for emerging markets, with the cumulative impact after two years reaching 0.16 for the former and 0.81 for the latter (table 3).

A first step to shed light on the difference in pass-through rates to consumer prices among countries is to explore whether there are also important differences in the response of prices of imported goods at the border—which can be interpreted as the response of pure traded goods.[8] Under perfectly competitive markets, it is generally the case that the exchange rate pass-through to import prices is complete. However, empirical studies have documented evidence of incomplete pass-through to import prices (e.g., Campa and Goldberg 2005 and Gopinath 2015), and many mechanisms have been proposed to support this possibility. For instance, under imperfect competition, market power allows exporting firms to "price to market" by adjusting their profit margins in response to the wealth and substitution effects triggered by the currency movement. Alternatively, firms may choose the currency of invoicing to minimize costs incurred from price adjustment.[9]

Figure 3 shows the pooled estimates of pass-through to import prices for advanced economies and emerging markets. A key observation is that pass-through to import prices is close to complete both in advanced and emerging markets, though slightly higher for emerging markets.[10]

The higher pass-through rate to consumer prices in emerging markets despite a more similar response of import prices could reflect differences in their consumption baskets. In particular, there may be differences in the import content in households' consumption between the two country groups. We use data from the Eora multiregion input-output tables (Lenzen et al. 2012, 2013) and compute

[8]The literature has documented that the prices of locally produced goods and services and those of purely traded goods (e.g., import prices at the border) respond differently following a change in the exchange rate (Burstein, Eichenbaum, and Rebelo 2005).

[9]Devereux, Engel, and Storgaard (2004) argue that agents will choose to price their goods in the currency that most reliably holds its value. Accordingly, delivering price stability could lead to an endogenous fall in the pass-through of the exchange rate to import prices.

[10]This finding is supported by results from individual economy regressions similar to equation (2) not reported here but available upon request.

Table 3. Local Projections Panel Estimation Results (1995–2019)

Horizon	NEER	P_oil	P_food	Output Gap	mPPI	Residual_{h–1}	N	Countries	R^2 (within)
A. Exchange Rate Pass-Through to Consumer Prices									
Advanced Economies									
6	0.07*** (0.007)	0.01*** (0.003)	0.06*** (0.009)	0.02*** (0.004)	0.39*** (0.07)	1.02*** (0.05)	8,410	32	0.62
12	0.12*** (0.012)	0.02*** (0.04)	0.06*** (0.014)	0.03*** (0.01)	0.5*** (0.09)	1.02*** (0.059)	8,398	32	0.65
18	0.13*** (0.017)	0.02*** (0.007)	0.08*** (0.016)	0.03*** (0.01)	0.529*** (0.12)	1*** (0.07)	8,386	32	0.61
24	0.16*** (0.02)	0.03*** (0.007)	0.06*** (0.018)	0.03** (0.01)	0.65*** (0.15)	1.01*** (0.07)	8,374	32	0.61
Emerging Economies									
6	0.42*** (0.05)	0 (0.009)	0.04** (0.02)	0.01 (0.019)	1.01*** (0.13)	1.13*** (0.09)	7,376	30	0.76
12	0.6*** (0.08)	0.01 (0.01)	0.04 (0.03)	0.09** (0.041)	1.44*** (0.29)	1.07*** (0.09)	7,376	30	0.75
18	0.74*** (0.113)	0.02 (0.02)	0.03 (0.04)	0.16** (0.06)	1.5*** (0.41)	1.05*** (0.1)	7,376	30	0.71
24	0.81*** (0.129)	0.03 (0.03)	0.02 (0.046)	0.23*** (0.08)	1.94*** (0.54)	1.03*** (0.12)	7,341	30	0.71

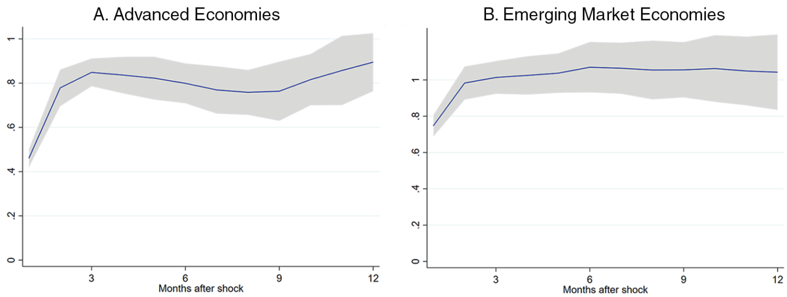
(*continued*)

Table 3. (Continued)

Horizon	NEER	P_oil	P_food	Output Gap	mPPI	Residual_{h–1}	N	Countries	R^2 (within)
B. Exchange Rate Pass-Through to Import Prices									
Advanced Economies									
6	0.8***	0.04***	0.23***	0.05***	1.89***	1.05***	6,283	28	0.61
	(0.05)	(0.01)	(0.05)	(0.02)	(0.23)	(0.08)			
12	0.89***	0.1***	0.03	0.02	1.29***	1.03***	6,277	28	0.58
	(0.07)	(0.02)	(0.05)	(0.03)	(0.33)	(0.06)			
18	0.71***	0.09***	0.2***	−0.06**	0.75*	1.01***	6,271	28	0.56
	(0.07)	(0.02)	(0.04)	(0.03)	(0.39)	(0.06)			
24	0.74***	0.04**	0.17***	−0.12***	2.24***	1.01***	6,257	28	0.56
	(0.08)	(0.02)	(0.04)	(0.03)	(0.47)	(0.07)			
Emerging Economies									
6	1.07***	0.03*	0.01	0.07	2.62***	1.08***	3,163	15	0.57
	(0.071)	(0.02)	(0.04)	(0.05)	(0.41)	(0.07)			
12	1.04***	0.01	0.03	0.2***	3.42***	1.03***	3,145	15	0.54
	(0.106)	(0.03)	(0.06)	(0.07)	(0.77)	(0.06)			
18	1.13***	0.03	0.01	0.22***	1.97**	1.02***	3,127	15	0.53
	(0.092)	(0.03)	(0.06)	(0.08)	(0.91)	(0.07)			
24	1.3***	0.06	−0.01	0.26***	2.49**	1.03***	3,109	15	0.54
	(0.117)	(0.04)	(0.08)	(0.1)	(0.99)	(0.07)			

Source: Authors' calculations.
Notes: The table shows estimation results based on equation (11) but, for brevity, the coefficients on the additional lags of the dependent variable and regressors are omitted. The results are available upon request.

Figure 3. Cumulative Impulse Response of Import Prices Following a Nominal Effective Depreciation of 1 Percent (in percentage points)



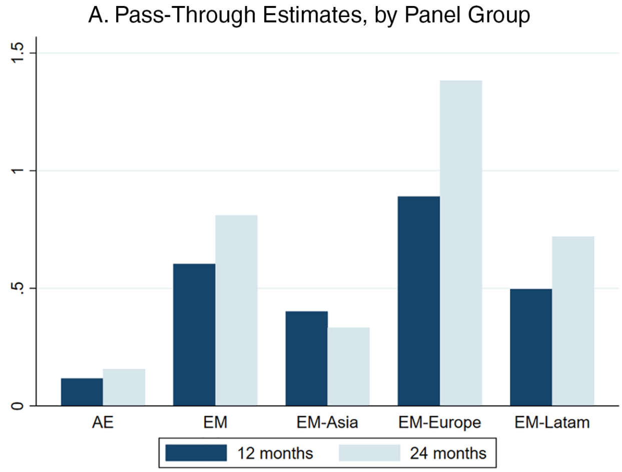
Notes: Cumulative impulse response of import prices in local currency (in percentage points) to a 1 percent innovation in the nominal effective exchange rate estimated using local projection methods. Shaded bands correspond to 95 percent confidence intervals.

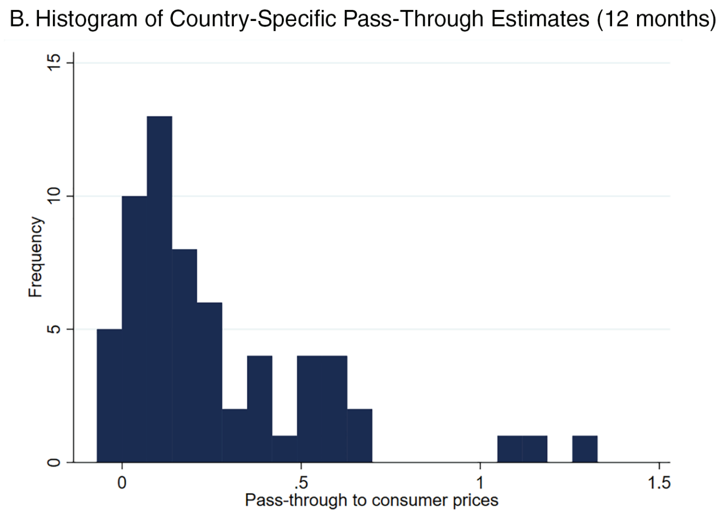
the import content of final consumption as in Gopinath (2015).[11] The horizontal dashed lines in figure 2 show that the import content of consumption is larger in advanced economies; therefore, it cannot explain a higher pass-through rate in emerging markets. Hence, the explanation for the latter result must lie within the price-setting behavior for domestically produced goods and services (including the distribution and retail margins on imported goods).

Before turning to what may explain such differences in pass-through rates to consumer prices, we explore the extent of heterogeneity across income groups, individual countries, and over time. In figure 4, panel A reports exchange rate pass-through estimates for emerging market economies by region, including Asia, Europe, and Latin America, and for advanced economies as a reference. We

[11]The import content of households' consumptions is constructed as the sum of two components: (i) a *direct* component that corresponds to imports of final consumption goods, and (ii) an *indirect* component that accounts for the value of imported inputs used in the production of domestic goods absorbed by local households (computed as the product of the value of output in each domestic sector that is absorbed by resident households and the share of imported inputs in that sector's output value).

Figure 4. Exchange Rate Pass-Through to Consumer Prices, Cross-Country Heterogeneity





Sources: Authors' calculations.
Notes: Cumulative response of headline consumer prices (in percentage points) to a 1 percent innovation in the nominal effective exchange rate after one and two years. AE = advanced economies. EM = emerging market economies.

estimate impulse response functions using equation (11) separately for each panel of these country groups. There is regional variation across emerging markets, with the pass-through rate after one year being lower in Asia (0.4) and Latin America (0.5) than in Eastern Europe (0.9). There are also differences in the speed of adjustment of consumer prices, with notably strong marginal effects in the second year among emerging markets in Eastern Europe.

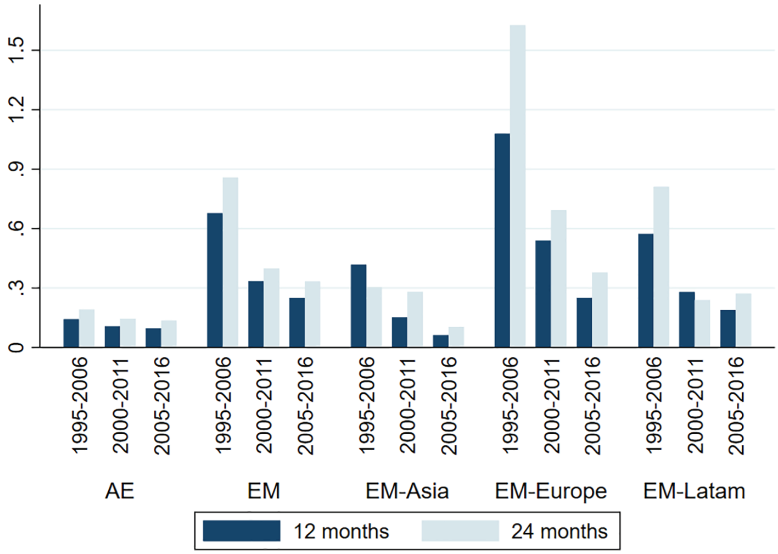
To explore the heterogeneity in pass-through rates across individual economies, we estimate country-specific versions of equation (11) over the same baseline sample period (1995–2019):

$$p_{t+h-1} - p_{t-1} = \alpha^h + \sum_{j=0}^{J} (\beta_j^h \Delta neer_{t-j} + \gamma_j^h \Delta oil_{t-j} + \delta_j^h \Delta food_{t-j} + \vartheta_j^h gap_{t-j} + \varphi_j^h \Delta mPPI_{t-j}) + \sum_{j=1}^{J} \rho_j^h \Delta p_{t-j} + \varepsilon_t^h. \quad (14)$$

Figure 4, panel B, reports the histogram of the estimated one-year cumulative exchange rate pass-through to consumer prices for the economies in our sample. In line with the literature, we find substantial heterogeneity in the magnitude of the estimated exchange rate pass-through coefficients across countries: point estimates at a one-year horizon range from negative (although not statistically significant) to larger than unity for a handful of emerging economies.

To put our results in context, in their extensive review of the literature, Goldberg and Knetter (1997) have documented an exchange rate pass-through to import prices around 0.6 for advanced economies with flexible exchange rate regimes, and somewhat lower for the United States. Campa and Goldberg (2005) find that exchange rate pass-through to aggregate import prices averages close to 0.5 in the short run and about 0.64 in the long run, also focusing on advanced economies—with a wide range of variation from 0.16 (Ireland) to 0.79 (Netherlands). Disaggregating across goods, they find a pass-through to import prices ranging from 0.62 in the manufacturing sector to 0.85 for raw materials. For nonmanufactured goods, the pass-through is 0.78. Our results are somewhat higher. For advanced economies, our estimates of exchange rate

Figure 5. Exchange Rate Pass-Through Coefficients over Time



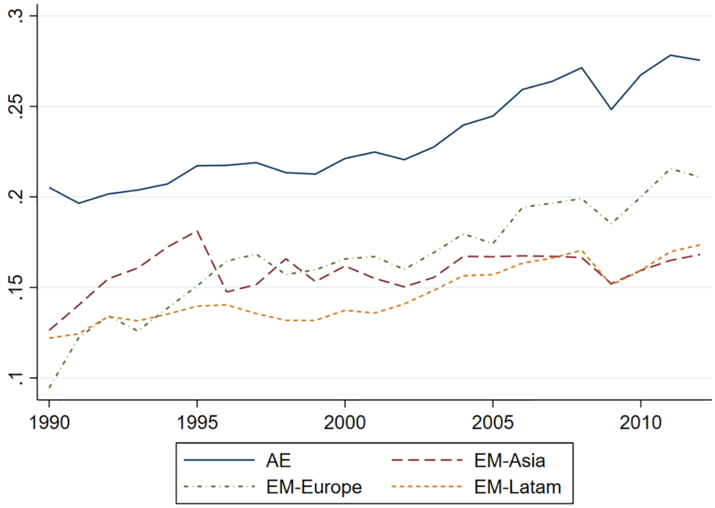
Sources: Authors' calculations.
Notes: Cumulative response of headline consumer prices (in percentage points) to a 1 percent innovation in the nominal effective exchange rate after one and two years. AE = advanced economies. EM = emerging market economies.

pass-through to import prices are about 0.5 on impact, but rise to around 0.8 in the medium term.

Unlike Goldberg and Knetter (1997) and Campa and Goldberg (2005), we also provide estimates of exchange rate pass-through to import prices for emerging markets. We find that pass-through to import prices in these countries is higher, starting around 0.75 on impact and rising above 1.0 in the medium term. Although our point estimates for advanced economies are slightly higher than the other papers mentioned, our sample of countries and the time period of estimation differ substantially. We consider our results to be in line with the literature, in the sense that previous estimates lie within the confidence intervals of our estimations at most horizons.

We finally explore whether there is evidence of declining exchange rate pass-through over time by running panel regressions in three subsamples of 12 years starting in 1995, 2000,

Figure 6. Import Content of Private Consumption (percent of total household consumption)



Sources: Eora MRIO; and authors' calculations.

and 2005. Figure 5 reports the results for the one- and two-year cumulative exchange rate pass-through to consumer prices of a 1 percent innovation in the nominal effective exchange rate. Consistent with other studies (e.g., Choudhri and Hakura 2006), we find that the exchange rate pass-through to consumer prices has systematically decreased in all country groups over the past two decades. This has been the case in both advanced and emerging economies, though the decline has been more pronounced among the latter. It is also true across regions, with important reductions in pass-through rates for emerging Asia, Europe, and Latin America. In Latin America, for example, the average pass-through rate over 2005–16 had fallen to one-third of its 1995–2006 level.

Note that the differences in pass-through rates over time do not seem related to changes in the share of imports in final consumption. Figure 6 shows that the import content of consumption expenditure has steadily increased in these economies over the past 20 years.

4. Exchange Rate Pass-Through and Monetary Policy Credibility

In the previous sections, we documented a substantial amount of cross-country and time variation in exchange rate pass-through coefficients which does not appear related to differences in the response of prices at the border or to differences in the share of imports in final consumption expenditure. In this section, we explore the role of monetary policy credibility in determining exchange rate pass-through coefficients. The conceptual support for the following estimations has been presented in section 2.

Previous studies have shown that a lower level and volatility of inflation are associated with lower exchange rate pass-through (Gagnon and Ihrig 2004; Choudhri and Hakura 2006). As discussed in section 2, these metrics may reflect increased performance of monetary policy but could also capture the variability in the time-series properties of the underlying shocks. The specific hypothesis we test here is whether a more credible monetary framework, as captured by more stable and better-coordinated inflation expectations, can lead to a reduction in exchange rate pass-through to consumer prices, in the spirit of Svensson (1997) and Taylor (2000).

To test this hypothesis, methodologically, we follow a two-stage procedure along the lines of Campa and Goldberg (2005). In a first stage, we gather our time-varying estimates of exchange rate pass-through for each economy, generated by estimating equation (14) in rolling 12-year windows starting in January of each year since 1995. In a second stage, we explore whether these estimates are related to proxies of monetary policy performance. To this end, we regress $\hat{\beta}^{12}_{0,i,\tau}$ on two alternative metrics for monetary policy credibility (**Cred**) and a set of control variables (X), measured as the average for the corresponding window τ and country i:[12]

$$\hat{\beta}^{12}_{0,i,\tau} = \gamma \mathbf{Cred}_{i,\tau} + \theta \boldsymbol{X}_{i,\tau} + \delta_i + \varsigma_\tau + \varepsilon_{i,\tau}. \tag{15}$$

[12]We use the inverse of the variance of the estimated pass-through coefficient as weights to give more weight to those coefficients estimated more precisely in the first-stage regressions. We restrict the sample to those countries that have data for all variables in $X_{i,\tau}$.

The vector $\boldsymbol{X}$ includes average inflation ($\bar{\pi}$), inflation volatility ($\sigma(\pi)$), average depreciation ($\overline{\Delta neer}$), and exchange rate volatility ($\sigma(neer)$). Beyond these factors, there are likely to be many others that affect the degree of exchange rate pass-through, including structural characteristics of local markets, such as the degree of competition among importers and domestic producers, and common external factors. Their exclusion from the specification could give rise to omitted-variable bias if these factors are correlated with monetary policy performance. To address these concerns, we include country fixed effects (δ_i) and time fixed effects (ς_τ) in all specifications.[13]

Table 4 reports the estimation results for the determinants of cumulative pass-through rates to consumer prices after 12 months. Columns 1–4 show the role of the level and volatility of key nominal variables when these are introduced sequentially. Standard measures of price stability such as the mean and standard deviation of the inflation rate are positively related to our estimates of exchange rate pass-through, in line with the findings of Gagnon and Ihrig (2004) and Choudhri and Hakura (2006). The coefficients are statistically significant and economically important: a 1 percentage point increase in the mean rate of inflation is associated with a pass-through coefficient that is higher by 0.42. Following Campa and Goldberg (2005), we also consider the average rate of depreciation and standard deviation of the multilateral exchange rate, finding that both are associated with higher exchange rate pass-through.

As noted in section 2, variables such as the average depreciation, inflation, and the volatility of inflation and the exchange rate may be affected by monetary policy but may also reflect cross-country and time differences in the time-series properties of the underlying shocks that triggered the depreciation. Therefore, we do not interpret the statistical significance of these variables as sufficient evidence of the role of monetary policy credibility in explaining exchange rate pass-through coefficients. Instead, we focus on the extent of inflation expectations' anchoring as a proxy for credibility. Better-anchored inflation expectations could make firms less likely to

[13]Given the dimensionality of the data, the within-group estimator is quite conservative: country fixed effects account for a large share of overall variation in the data, as there are only nine estimation windows for each country in the sample and these windows overlap substantially.

Table 4. Determinants of Exchange Rate Pass-Through

	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Mean Inflation	0.42***						0.15	0.19*
	(0.05)						(0.11)	(0.11)
Inflation Volatility		0.26***					−0.01	0.00
		(0.03)					(0.06)	(0.06)
Mean Depreciation			0.17***				0.05	0.04
			(0.03)				(0.03)	(0.03)
Volatility of Exchange Rate				0.00***			0.00	0.00
				(0.00)			(0.00)	(0.00)
Volatility of Near-Term Inflation Forecasts					0.03***		0.01	0.00
					(0.00)		(0.01)	(0.01)
Log of Mean Disagreement						0.18***	0.09***	0.09***
						(0.02)	(0.03)	(0.03)
Import Share of Consumption								1.05***
								(0.38)
Constant	0.05	0.04	0.15***	0.09**	0.11***	0.35***	0.21***	0.13**
	(0.04)	(0.04)	(0.04)	(0.04)	(0.03)	(0.04)	(0.06)	(0.06)
Observations	318	318	318	318	318	318	318	318
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Countries	40	40	40	40	40	40	40	40
Time Windows	11	11	11	11	11	11	11	11
Adjusted R^2	0.780	0.782	0.766	0.742	0.788	0.794	0.806	0.811

Source: Authors' calculations.
Notes: The dependent variable is the estimate of exchange rate pass-through to consumer prices for each country i and window τ at a 12-month horizon. The regressions are estimated by weighted least squares, with observations weighted by the inverse variance of the country-specific exchange rate pass-through estimates. Standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

pass along movements in the exchange rate to domestic consumers—as is the case in the model of Devereux and Yetman (2010) as firms choose to update prices less frequently.

We measure inflation expectations with inflation forecasts of individual professional forecasters that are compiled monthly by Consensus Economics and are available for 40 countries in our sample. These forecasts are published as fixed-event forecasts for inflation in the current and upcoming calendar year. Since uncertainty about fixed-event forecasts decreases over the calendar year as the forecast horizon shrinks, we follow the common approach of combining these forecasts linearly into a synthetic 12-month fixed horizon.[14] The one-year-ahead synthetic fixed-horizon forecast is constructed as the weighted average of the forecasts for the current ($\hat{X}_{t+k|t}$) and next calendar year ($\hat{X}_{t+12+k|t}$), with weights that vary according to the date the forecast was produced:

$$\tilde{\hat{X}}_{t+12|t} = \frac{k}{12}\hat{X}_{t+k|t} + \frac{12-k}{12}\hat{X}_{t+12+k|t}, \tag{16}$$

where k is the months remaining in the year at the time the forecast was produced ($k \epsilon \{1, 2, \ldots, 12\}$) at period t.

Kumar et al. (2015) argue that if inflation expectations are well anchored, inflation forecasts by individual forecasters—and hence average forecasts—should be relatively stable over time. As a first anchoring measure, we consider the variability over time of the average inflation forecast at a synthetic 12-month horizon.[15] We find that, indeed, where forecasts of inflation are more volatile, exchange rate pass-through coefficients are significantly higher (column 5).

If the monetary framework is credible and inflation expectations are well anchored, the dispersion of individual inflation forecasts should also be small (Kumar et al. 2015). Dovern, Fritsche, and Slacalek (2012) propose that lower disagreement among professional forecasters of inflation reflects greater credibility of monetary policy,

[14]See, for instance, Mankiw, Reis, and Wolfers (2004) and Dovern, Fritsche, and Slacalek (2012).

[15]The variability of the average inflation forecast for country i and window τ is given by $\sqrt{\frac{1}{T-1}\sum_{t=1}^{T}(\pi_t^e - \overline{\pi^e})^2}$ for $t \in \tau$, where π_t^e is the average (across forecasters) synthetic 12-month-ahead inflation forecast at time t and $\overline{\pi^e}$ denotes its average over all time t within τ.

and they find that it is related to measures of central bank independence among G-7 economies.[16] Relatedly, Capistrán and Ramos-Francia (2010) find that the adoption of inflation targeting reduces the degree of forecast disagreement among developing economies. They argue that this result reflects a more predictable monetary policy that is able to better coordinate expectations.

Following this literature, we measure disagreement as the standard deviation of inflation forecasts of individual professional forecasters at the synthetic 12-month fixed horizon. The results in table 4, column 6, show a strong and significant positive relationship between the log of mean disagreement and exchange rate pass-through coefficients, suggesting that better-anchored inflation expectations make firms less likely to pass along movements in the exchange rate to domestic prices.[17]

While a predictable and credible monetary policy should lead to closer coordination among forecasters regarding the future path for inflation, a valid concern is that the inverse is not necessarily true: agents could agree that the central bank will miss its target, such that there is low disagreement but also a lack of credibility. Promisingly, figure 7 shows a strong correlation between forecast disagreement and the average (squared) deviation of inflation from central bank targets.[18]

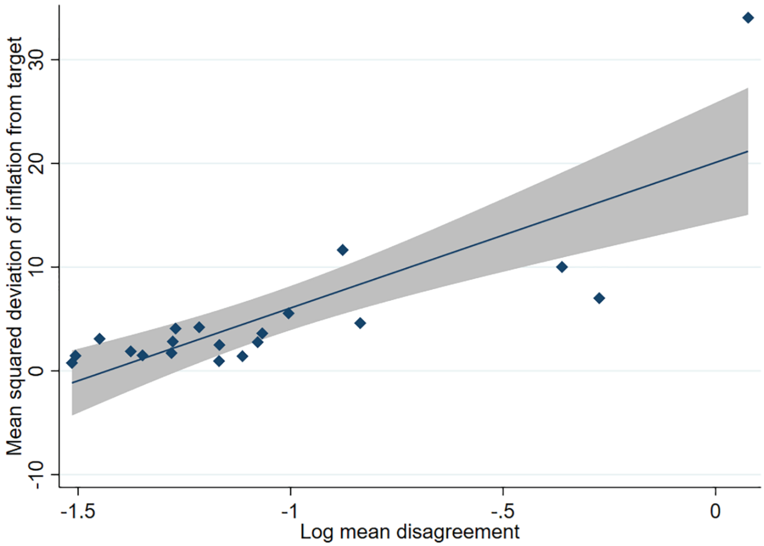
The alternative measures of monetary performance considered are correlated among themselves, since they capture related aspects of price stability. For instance, Mankiw, Reis, and Wolfers (2004) and Capistrán and Ramos-Francia (2010) document that forecast disagreement is increasing in the level of inflation. In column 7, we

[16]Disagreement among forecasters also captures factors besides monetary policy performance. The variability of shocks affecting the economy is also expected to increase disagreement among forecasters. However, Dovern, Fritsche, and Slacalek (2012) show that the relationship between disagreement and central bank independence is robust to controlling for macroeconomic volatility.

[17]The log of mean disagreement for country i in window τ is defined as the natural logarithm of $\frac{1}{T}\sum_{t=1}^{T}\left(\sqrt{\frac{1}{J-1}\sum_{j=1}^{J}\left(\pi_{j,t}^{e}-\pi_{t}^{e}\right)^2}\right)$ for $t \in \tau$, where $\pi_{j,t}^{e}$ denotes the synthetic 12-month-ahead inflation forecast of agent j at time t and π_t^e is the average across forecasters.

[18]Since only inflation-targeting central banks announce explicit targets for inflation, the sample used for figure 7 shrinks considerably.

Figure 7. Disagreement among Professional Forecasters of Inflation and Mean Deviation of Inflation from Target



Source: Authors' calculations.
Notes: The vertical axis shows mean squared deviations of inflation from inflation targets over 2000–19. The horizontal axis shows log mean disagreement among professional forecasters of inflation over the same sample. They gray area denotes a 95 percent confidence interval.

include all explanatory variables in the same specification. Interestingly, disagreement seems to act as a good summary measure for the role of price stability in the determination of pass-through. Conditional on the first and second moment of nominal variables, lower forecast disagreement remains associated with smaller exchange rate pass-through to consumer prices.[19] The coefficient on disagreement also remains economically meaningful: a decline in disagreement from the top to the bottom of the interquartile range would be associated with a decline in the pass-through rate of about 0.06—which is

[19]The inclusion of the volatility of actual and expected inflation in the specifications further mitigates reverse causality concerns, since difficulties with forecasting inflation would likely be associated with higher actual and expected inflation.

almost one-third of the average cumulative pass-through rate in the sample after 12 months. By including the variability of inflation and the exchange rate, we indirectly control for a variety of factors that determine these outcomes, which may include, for example, foreign exchange interventions by the central bank. As such, our findings stress that it is the credibility of the central bank that lowers the extent of exchange rate pass-through.

As discussed in section 3, a varying degree of exchange rate pass-through could also reflect changes in the composition of the consumption basket. Yet, after controlling for the import content of households' consumption, the results in column 8 show that the exchange rate pass-through remains positively related to the degree of forecast disagreement, with a coefficient that remains large and statistically significant.[20]

5. Concluding Remarks

We revisit a long-standing question in macroeconomics on the determinants of the response of consumer prices to currency depreciations. Using data for a broad sample of 62 emerging and advanced economies since 1995, we start by documenting a widespread decline in the degree of exchange rate pass-through despite an increase in the import content of domestic consumption. But we also document substantial heterogeneity in pass-through coefficients across countries.

We then explore the role of monetary policy performance and credibility in explaining cross-country and time variation in the extent of exchange rate pass-through. We first confirm earlier results

[20]Some studies in the literature have used more widely available metrics of import content such as measures of trade openness including the ratio of imports to GDP (e.g., Gagnon and Ihrig 2004; Choudhri and Hakura 2006) and find no statistical link between pass-through estimates and this import share. However, an aggregate openness metric such as the ratio of imports to GDP also includes imports of non-consumption goods, consumption goods that are absorbed by the government, and goods that are re-exported to other final destinations, thus providing an imprecise proxy for the assessment of exchange rate pass-through to consumer prices.

that pass-through coefficients are positively and significantly associated with the level and variability of inflation, i.e., enhanced performance of monetary policy. But because these variables could also be capturing the time-series properties of the underlying shock, we then study how the credibility of monetary policy—as captured by the degree of anchoring of inflation expectations—affects exchange rate pass-through coefficients.

We find that more stable and better-coordinated inflation expectations of professional forecasters are associated with significantly lower exchange rate pass-through to consumer prices. Moreover, the extent of disagreement about future inflation across individual professional forecasters remains significant after controlling for the first and second moment of nominal variables and the import content of consumption. The effect is also economically important in explaining the degree of exchange rate pass-through. An increase in disagreement from the 25th to the 75th percentiles of our sample is associated with an increase in the estimated pass-through coefficient by 0.06. This is a sizable change, given that the average cumulative pass-through after 12 months is about 0.2 in the sample. This contribution provides novel evidence in support of the long-standing conjecture that the improvement of monetary policy frameworks—specifically, the credibility of monetary policy—has led to lower exchange rate pass-through to consumer prices by establishing stronger nominal anchors.

Appendix

Table A.1. Import Price Data Sources

Economy	Source	Local Currency
Austria	Haver Analytics/Eurostat	Yes
Belgium	Haver Analytics/Eurostat	Yes
Brazil	Haver Analytics/FUNCEX	No
Canada	Haver Analytics/StatCan	Yes
Colombia	Haver Analytics/BANREP	Yes
Czech Republic	Haver Analytics/CSO	Yes
Denmark	Haver Analytics/IFS	Yes
El Salvador	Haver Analytics/BCR	No
Estonia	Haver Analytics/Eurostat	Yes
Finland	Haver Analytics/Eurostat	Yes
France	Haver Analytics/Eurostat	Yes
Germany	Haver Analytics/Eurostat	Yes
Greece	Haver Analytics/Eurostat	Yes
Hong Kong SAR	Haver Analytics/HKCSD	Yes
Hungary	Haver Analytics/CSO	Yes
Indonesia	Haver Analytics/BPS	Yes
Ireland	Haver Analytics/Eurostat	Yes
Italy	Haver Analytics/Eurostat	Yes
Japan	Haver Analytics/JPCSD	Yes
Korea	Haver Analytics/NSO	Yes
Latvia	Haver Analytics/Eurostat	Yes
Lithuania	Haver Analytics/Eurostat	Yes
Luxembourg	Haver Analytics/Eurostat	Yes
Malaysia	Haver Analytics/DSM	No
Mexico	Haver Analytics/BMEX	No
Paraguay	Haver Analytics/BCP	Yes
Peru	Haver Analytics/BCRP	No
Poland	Haver Analytics/CSO	Yes
Portugal	Haver Analytics/Eurostat	Yes
Singapore	Haver Analytics/DoS	Yes
Slovak Republic	Haver Analytics/Eurostat	Yes
Slovenia	Haver Analytics/Eurostat	Yes
Spain	Haver Analytics/Eurostat	Yes
Sweden	Haver Analytics/SCB	Yes
Switzerland	Haver Analytics/SFSO	No
Thailand	Haver Analytics/MoC	Yes
The Netherlands	Haver Analytics/Eurostat	Yes
Turkey	Haver Analytics/TRSTAT	No
United Kingdom	Haver Analytics/IFS	Yes
United States	Haver Analytics/BLS	Yes

Source: Authors' compilation.
Note: Local currency denotes whether the import price data is directly reported in local currency by the source.

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Fiscal Transfers without Moral Hazard?*

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Recent debate has focused on the introduction of a central stabilization capacity as a completing element of the Economic and Monetary Union. Its main objective would be to contribute to cushioning country-specific economic shocks, especially when national fiscal stabilizers are run down. There are two main potential objections to such schemes proposed so far: first, they may lead to moral hazard, i.e., weaken the incentives for sound fiscal policies and structural reforms. Second, they may generate permanent transfers among countries. Here we present a scheme that is relatively free from moral hazard, because the transfers are based on changes in world trade in the various industrial sectors. These changes can be considered as largely exogenous, hence independent from an individual government's policy. Therefore, the scheme is better protected

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against manipulation compared with other schemes based on domestic variables (e.g., unemployment). Our scheme works as follows: if a sector is hit by a negative shock at the world market level, then a country with an economic structure that is skewed toward this sector receives a temporary transfer from the other countries. We show that the transfers generated by our scheme are countercyclical. In addition, since transfers are based on deviations from trends in sectoral export value-added, the danger of permanent transfers from one set of countries to the other countries is effectively ruled out. Finally, we show that the transfers are robust to different sectoral aggregation and revisions in the underlying export data.

JEL Codes: E32, E62, E63.

1. Introduction

The global economic and financial crisis that started in 2007 and the ensuing euro-zone debt crisis have shown the painful consequences of having an incomplete monetary union. In response to these developments, substantial effort has been made to improve the euro zone's fiscal and financial architecture with the introduction of the European semester, a strengthening of the Stability and Growth Pact (through the "Six-Pack" and the "Two-Pack"), the "Fiscal Compact," and the introduction of the first elements of a banking union. More recently, European governments have agreed to launch the "Next Generation EU" (NGEU) to mitigate the adverse consequences of the COVID-19 crisis. Still, Europe's Economic and Monetary Union (EMU) remains unfinished. As part of the process toward the completion of the EMU, further steps need to be taken. In particular, a fiscal union is deemed as needed. Indeed, in contrast to other monetary unions, the EMU lacks a central fiscal capacity, which could help cushion country-specific and common shocks.[1] Fiscal policy remains decentralized, implying that the potential for macroeconomic stabilization through area-wide fiscal policies is under-exploited. Even though the crisis was triggered within the global financial system, the lack of a euro-area central

[1]The NGEU has a temporary nature, and it is supposed to be phased out after the end of the pandemic crisis.

fiscal capacity may have contributed to its severity and, looking ahead, would not help to alleviate the impact of future crises.

Discussions about fiscal centralization already started some years ago. In a report in June 2012, the then president of the European Council, Van Rompuy (2012), identifies an integrated budgetary framework as one of four building blocks to consolidate the EMU. Shortly after, in December 2012, the "Four Presidents' Report" (Van Rompuy et al. 2012) discusses the gradual creation of a central fiscal capacity aimed at both promoting structural reforms and mitigating asymmetric shocks. The "Five Presidents' Report" (Juncker et al. 2015) sketches the steps toward completion of the EMU and, more specifically, also toward a fiscal union as one of its main building blocks. It discusses the notion of a euro-area stabilization function with the guiding principles that it should not lead to permanent transfers, which would be avoided through the convergence of economic structures beforehand, and not undermine the incentives for sound fiscal policy.[2] The capacity would also not be intended as a crisis management tool, but rather thought to improve the economic resilience to temporary shocks of the euro zone and its individual members. Most recently, the European Commission (2017)'s reflection paper sketches the main concrete options for a macroeconomic stabilization function for the euro area. One would be a scheme to protect investment in the case of a downturn. Another would be an unemployment reinsurance scheme to support national unemployment schemes. Importantly, the former scheme is generally conceived as a mechanism to cushion area-wide (aggregate) shocks, while the latter would address country-specific (idiosyncratic) shocks.

This paper proposes a novel "export-based stabilization capacity" (ESC) that allows for cross-border transfers in response to

[2]A future central stabilization capacity is more broadly discussed, for example, by European Central Bank board member Cœuré (2016) and the president of the Dutch central bank (Knot 2016) before the European Parliament. In their report, D'Alfonso and Stuchlik (2016) explore the potential options concerning a centralized fiscal capacity for the European Parliament. Recently, a motion was put to vote in the European Parliament which sets out a roadmap toward a budgetary capacity for the euro zone (European Parliament 2017). The case for enhanced fiscal risk sharing in the EMU is also made by a recent study of the International Monetary Fund (Berger, Dell'Ariccia, and Obstfeld 2018) and by the European Fiscal Board (2018). For a conceptual discussion, see De Grauwe (2018).

exogenous changes in the world trade in the various sectors. A stabilization capacity may be particularly beneficial in the presence of asymmetric shocks, which the ECB can by definition not address, as monetary policy is formed on the basis of aggregate developments in the euro zone,[3] while private insurance through cross-border capital flows remains limited, because asset holdings are notoriously home biased. Our ESC works in a very simple and intuitive way: suppose that world trade in a specific sector is hit by a negative shock, as reflected in total euro-zone export in that sector falling below its trend. Then, euro-zone members that are relatively more export intensive in this sector receive a transfer from the members that are relatively less intensive in this sector.

This ESC has a number of advantages, which should enhance its political acceptability when compared with many existing proposals, although quite naturally, as is the case with any cross-border transfer scheme, the prospect of having to pay a transfer to other countries at some point may generate political resistance. First, the transfers respond to exogenous developments in the world market, which are largely outside of the control of individual governments. As such, the scheme is relatively free from moral hazard, given that it would not weaken the incentives of governments to run virtuous fiscal policies and implement structural reforms. Indeed, as pointed out in the recent contribution by the group of 14 French and German economists (Bénassy-Quéré et al. 2018), the political feasibility of expanding arrangements for risk sharing in the euro area requires that moral hazard issues be taken seriously. Second, since it is based on deviations from trends in world trade in individual sectors and the estimated trend is updated each year, the scheme would by construction avoid permanent transfers from a set of countries to another. Third, our scheme does not need to rely on a long-run process of convergence of economic structures before it can be implemented. Fourth, the scheme is designed such that each period all the cross-border transfers (almost) add up to zero. Fifth, the scheme

[3]With perfectly flexible markets, asymmetric shocks can be handled easily, because production factors move quickly to those parts of the union where undercapacity prevails. However, there is an abundance of evidence that European markets are highly rigid. In particular, labor mobility is low, both within, but even more so, across countries.

is shown to be very robust to revisions in the underlying export data and to different sectoral disaggregations. Finally, it is important to realize that the transfers are not specifically earmarked for sectors in decline. They are intergovernmental[4] and, hence, under the simplest configuration of the scheme, a net recipient government can determine its use. However, in order to minimize possible adverse incentives that could emerge if a government delays the transition to emerging sectors, transfers could also be granted under certain conditionality: they could be disbursed only if a government actively supports the transition of activity toward upcoming sectors. In general, however, how transfers are best utilized goes beyond the present work. A discussion on this can be tackled in future research.

We perform a simulation of our ESC using sectoral export data from the Organisation for Economic Co-operation and Development (OECD) for all 19 euro-zone countries over the period 2002–14. This allows us to estimate how the transfers would have materialized over this period, if the ESC were in place. In the baseline version of our scheme, a country receives a transfer associated with a given sector if that sector is hit by a negative shock at the world market level, and if the export share of the country in that sector is higher than the total euro-area exports share of the country (i.e., the country has a relative specialization in that sector). In the baseline scheme, the income loss that a country experiences after a negative shock to the export sectors in which it is relatively more specialized is compensated by a transfer from the other EMU countries, i.e., the ones less affected by that shock. We then consider variants to the baseline scheme. More specifically, we first assume that the transfer is limited only to the labor income loss, which would be the more appropriate scheme if the holdings of the equity in the exporting firms were spread beyond national borders. The second variation imposes that the transfer that a country pays after experiencing on net a positive shock to its export sectors cannot be bigger than the net increase in government revenues resulting from the increased economic activity following the shock.

[4]Exactly the same allocation as under our mechanism could be achieved by sectors directly paying transfers to or receiving transfers from a central fund on the basis of the same shocks to sectoral net exports value-added we consider.

We find that the net transfers received (or paid) by the countries in our sample are countercyclical: they are more positive (or less negative) when the output gap is lower. Over the full sample, cumulative transfers tend to stabilize and to return toward zero toward the end of the sample, thus suggesting that permanent transfers are ruled out under this scheme. We also show that the transfers are robust to the use of preliminary rather than ex post data and to the reaggregation of the set of sectors into a smaller number of sectors.

To put our scheme into perspective, the transfers it generates tend to be non-negligible but limited in (cumulative) magnitude. However, transfers can reach larger magnitudes when this is most needed, i.e., during a severe recession. These properties may make our scheme more politically palatable than schemes that expose countries to the risk of large payments to other countries. Of course, the scheme cannot address all idiosyncratic shocks. However, if needed, it could be combined with auxiliary arrangements that address other sources of shocks or common shocks, though probably at a greater moral hazard risk (e.g., Feld and Osterloh 2013). As is the case for any scheme featuring in the public debate, practical obstacles also need to be overcome. The main obstacle is the timely availability of the data that serve as inputs for the calculation of the transfers. While we show that our results are robust to the use of nonrevised export data, even these data become available with quite some delay. Yet, the purpose of this paper is not to provide a blueprint for a system that can be implemented right away. Rather, we aim at demonstrating that a scheme like ours has the potential to generate plausible transfers with a number of desirable properties. As data provision becomes better and faster, practical implementation comes within sight. Viewed from a different angle: by exposing the data needs of the practical implementation of a plausible transfer scheme, we may encourage statistical agencies to work on fulfilling these data needs.

The remainder of this paper is organized as follows. Section 2 reviews the related literature. Section 3 lays out the design of our baseline ESC and its proposed variants, which is followed by a discussion of the data sources in section 4. Section 5 reports and discusses the transfers based on actual data. The robustness of the baseline scheme is investigated in section 6. Finally, section 7 concludes the main body of this paper.

2. Literature Review

2.1 Risk-Sharing Channels in Federations and across Countries

Eventually, the need for a centralized stabilization capacity in the euro zone will be determined by the amount of cross-border risk sharing that already exists. Over the past two decades, there has been a substantial amount of work, using a variety of empirical approaches, analyzing the magnitude of risk sharing across countries and across regions. A large fraction of it focuses on interregional risk sharing in the United States and other federal countries. However, there also exist a number of studies focusing on the euro zone.

Risk sharing of asymmetric shocks in federations can take place through a variety of private and public channels. For example, individuals may hold equity stakes in companies from different regions. In their seminal contribution, Asdrubali, Sorensen, and Yosha (1996) explore the importance of the various channels through which consumption risk sharing takes place among the states in the United States. They demonstrate that there exists substantial risk sharing through cross-state asset holdings. Among the public channels, a federal tax-transfer scheme may be important. Von Hagen (1999) summarizes the estimates in the early literature of the share of state-specific shocks insured through the federal tax-transfer system in the United States. It ranges from 7 to 40 percent, as found by Sala-i-Martin and Sachs (1991), although most of the estimates are on the order of 10 to 15 percent. Other countries for which insurance through the tax-transfer mechanism has been estimated are Canada, France, Germany, the United Kingdom, and Italy. For Canada, this source of insurance is quite consistently estimated to be close to 15 percent, although more recent work by Balli, Basher, and Jean Louis (2012) gets to an estimate of 27 percent. The estimated degree of implicit insurance among German and French regions is 35–40 percent (e.g., see Pisani-Ferry, Italianer, and Lescure 1993 and Mélitz and Zumer 1998), while for the United Kingdom it is around 20 percent. The lowest degree of implicit insurance seems to prevail in Italy, for which Obstfeld and Peri (1998) arrive at an estimate of only 3 percent. An important complication of this line of research is that it is hard to distinguish pure insurance against asymmetric shocks from redistribution, which takes place at the same time if state incomes differ on average. Recent work by Poghosyan, Senhadji, and Cottarelli (2016) for the United States, Canada, and Australia distinguishes between interregional fiscal transfers smoothing idiosyncratic versus permanent shocks. They find that 4–11 percent of the idiosyncratic shocks are smoothed (i.e., risk sharing), as opposed to 13–24 percent of permanent shocks (i.e., redistribution). Work by Feld, Schaltegger, and Studerus (2018) for the Swiss federation suggests that the stabilization of short-term income fluctuations through the tax-transfer system is less than 10 percent, while redistribution amounts to about 20 percent.

The literature suggests that cross-border consumption risk sharing through private asset holdings plays only a limited role in Europe (for example, see Sorensen and Yosha 1998 and, more recently, European Central Bank 2018). Nevertheless, there exists evidence that overall consumption risk sharing has increased over time. Cimadomo, Fortuna, and Giuliodori (2017) estimate an increase from about 40 percent at the start of EMU to about 65 percent in 2015. Both increased cross-border holdings of financial assets and international official assistance contribute to this improvement (see also Milano 2017).[5,6] Farhi and Werning (2017) provide a rationale for this latter finding: they show theoretically that some degree of public intervention is helpful even in the presence of complete markets which would allow insurance against idiosyncratic shocks. Therefore, they make a strong case for fiscal insurance as a necessary complement to risk sharing via private channels.

2.2 Analysis of Proposed Stabilization Schemes

The debate around a supranational automatic stabilization mechanism for Europe dates back to the 1970s (see, e.g., Marjolin et al.

[5] Hepp and von Hagen (2013) estimate an increase in the role of factor markets in interstate consumption risk smoothing in Germany after its unification. However, risk sharing through the government sector continues to be important by smoothing around 10 percent of shocks after the unification.

[6] Not all private channels may have contributed to the increase in risk sharing. Hoffmann et al. (2018) find that the contribution via the credit channel collapsed with the breakdown of the interbanking market in 2008 (see also Beetsma et al. 2018).

1975) and reemerged in the 1990s (see, e.g., Padoa-Schioppa et al. 1987). However, as highlighted in Beblavý and Lenaerts (2017), proposals remained unexecuted for two main reasons. On the one hand, there was a common belief that market adjustment mechanisms alone would lead to macroeconomic stabilization. On the other hand, the launch of the EMU was expected to be accompanied by stronger business cycle synchronization for member states, and therefore by fewer and weaker asymmetric shocks (Allard et al. 2013). The recent crisis suggests that business cycle convergence is far from achieved. In addition, if national fiscal buffers are run down completely, shocks remain unsmoothed or are even amplified. This explains a renewed attention in the post-crisis debate on a centralized fiscal capacity which could help attenuate the effects of macroeconomic shocks in the euro area.

Recent proposals mainly build on schemes addressing either country-specific (i.e., idiosyncratic) shocks or aggregate shocks, i.e., shocks common to all members of the currency union. As regards the first category, studies have typically focused on shocks hitting country-specific GDP, the output gap, or employment. For example, Enderlein, Guttenberg, and Spiess (2013) propose a "European fund" calibrated on country-specific output gaps: member states would contribute to the fund when their cyclical position is better than the euro-area average, and they would receive a net transfer when they are in a worse position. Another scheme recently proposed is the one by Furceri and Zdzienicka (2015), which focuses on country-specific GDP shocks. The authors simulate a supranational fiscal stabilization mechanism for the euro area, financed by a gross contribution of 1½–2½ percent of countries' gross national product (GNP). The scheme would imply transfers to countries hit by negative GDP shocks. The authors show that such a scheme could provide significant stabilization for the euro area, comparable to the level of fiscal risk sharing observed in Germany and other federally organized countries. In general, the main criticism of schemes based on the output gap is that this variable is unobservable and subject to large revisions. These revisions complicate the implementation of such schemes "in real time," while the unobservability of the output gap may lead to disagreement about its measurement, in particular when cross-border transfers are based on it. Schemes based on GDP, on the other hand, are less likely to be subject to big revisions.

However, the estimation of country-specific GDP shocks is not trivial and the outcomes would be subject to the deployed methodology.

Many of the current proposals have focused on a European unemployment insurance scheme. The main reason is that unemployment expenditure is the main category of public spending that moves automatically (although typically with a lag) with the business cycle. Therefore, a common unemployment insurance based on cross-country transfers could work well to reinforce national automatic stabilization mechanisms. Several proposals for such a scheme were brought into the political debate. For example, Dolls et al. (2015) model transfers based on household-level data for euro-zone economies and find that about 10 percent of the income fluctuations caused by transitions into and out of unemployment could be absorbed by means of a common unemployment insurance scheme.[7] The main advantage of a scheme based on unemployment is that it would be strongly countercyclical. In addition, unemployment data are subject to small revisions. However, unemployment insurance schemes are especially prone to moral hazard, as unemployment spending not only depends on cyclical developments but also crucially on structural characteristics of labor markets, on which economic policy has a decisive influence. Awareness of the fact that transfers received are on average higher when average unemployment is higher may weaken the government's incentive to conduct politically costly structural reforms. For this reason, it has been proposed that transfers should be triggered conditionally on the fulfilment of a so-called double condition: unemployment should exceed its historical average over, say, the past 15 years and it should be increasing significantly at, say, more than 1 percentage point in a year (see, e.g., Martinez Mongay 2019).

Other recent proposals focus on a euro-area "investment capacity." Such a scheme would finance national investment projects in downturns. This discourages countries from cutting public investment, thus reducing their growth potential, when faced with the need of fiscal consolidation. At the same time, it would contribute

[7] Other examples are Beblavý, Gros, and Maselli (2015), Abraham et al. (2017), Carnot, Kizior, and Mourre (2017), and Arnold et al. (2018). For surveys and analyses of different schemes, see Beblavý and Lenaerts (2017), Favaque and Huart (2017), and the German Council of Economic Experts (2018, chapter IV).

to stabilizing the economic cycle. For example, the German Ministry of the Economy and the French Treasury developed proposals for a common budget for infrastructure and stabilization as one of the main new elements of a reformed euro-area fiscal framework (see Zettelmeyer 2016; Bara, Castets, and Zakhartchouk 2017). Such schemes are typically designed to address aggregate shocks hitting the whole euro area, especially when monetary policy is constrained by the zero lower bound (see, e.g., International Monetary Fund 2016). By contrast, the European Investment Stabilisation Function (EISF) proposed by the European Commission (2018) intends to address asymmetric shocks. However, because of resistance of some European Union (EU) member states, stabilization facilities currently have a rather low priority on the EU policy agenda. The Budgetary Instrument for Convergence and Competitiveness (BICC) is the only central instrument likely to be implemented in the near future (see European Commission 2019). However, the BICC is not intended for stabilization. The stalemate in the area of central stabilization capacities creates room for new proposals that could become attractive in the future.

3. The Design of Our ESC

Ideally, our transfer scheme would optimize a formal microfounded welfare criterion.[8] However, this would be beyond the scope of this paper. Hence, our objective is more modest, and a useful starting point for the discussion of the design of our ESC is formed by the desiderata listed by Von Hagen and Hammond (1995) for a central stabilization capacity. First, insurance should be provided primarily against asymmetric shocks, because for these shocks the loss of an independent monetary policy is most important. Second, transfers should be based on serially uncorrelated shocks only. Transfers in response to persistent shocks might reduce policymakers' incentives to undertake politically costly reforms to overcome the structural

[8]An example of an analysis of transfers in a dynamic stochastic general equilibrium maximizing framework is Bandeira (2018), who investigates the welfare and economic stabilization properties of a fiscal transfer scheme between members of a monetary union in response to changes in sovereign spreads. Potential moral hazard on the side of governments resulting from such transfers and a numerical analysis like ours based on actual data are not addressed.

problems that form the source of the persistence of the shocks. Third, the scheme should be simple and automatic for it to be acceptable for the general public. Fourth, over time net transfers should be zero on average. Fifth, the scheme should be financially balanced at the supranational level. Finally, setting up such a scheme is only worthwhile if it is able to offset a substantial part of the asymmetric shocks.

Clearly, a central stabilization capacity that fulfills all these conditions simultaneously will be difficult to design and, in practice, tradeoffs seem unavoidable. For example, a larger scheme likely increases the danger of moral hazard, ceteris paribus. However, a suitable design of the scheme may mitigate this tradeoff. Also, more than two decades of experience with a common currency may have affected the general perspective on these tradeoffs. In particular, the experience of the global financial recession and the ensuing European sovereign debt crisis, with ECB policy rates stuck at the lower bound, may have strengthened the case for a central stabilization capacity to also provide stabilization in response to (highly adverse) common shocks. The divisions among EU member states on whether and how to proceed with EU integration have, even more than before, demonstrated that the viability of a central stabilization capacity requires broad political acceptability. Finally, allowing for some persistence of the shock underlying the capacity, in particular for the duration of a business cycle downturn or upturn, seems desirable. Indeed, all stabilization schemes proposed so far, such as those based on unemployment, allow for some degree of persistence in the underlying shocks (e.g., Dolls et al. 2015). However, when transfers flow in one direction for too long, political resistance from the net contributors will become so large that the scheme is bound to collapse. Hence, for a central stabilization capacity to be politically viable over a long time span, average annual transfers should be roughly equal to zero. The objective of the scheme we propose is to generate countercyclical transfers with only a limited effect on moral hazard, while fulfilling as much as possible other properties required to make it politically palatable.

To limit potential moral hazard, the cross-border transfers associated with a central stabilization capacity should be conditioned on estimates of shocks that are as much as possible beyond the control of the individual governments. The ESC we propose below conditions cross-border transfers on changes in world market conditions in the individual exporting sectors of the economy. In contrast to, for example, the output gap or unemployment, these conditions are, at most, to a minor extent affected by an *individual* government's policies. Hence, the scope for moral hazard associated with our ESC should be limited.

3.1 Baseline Scheme: Equalizing Income Shifts as a Fraction of the Value-Added of Exports

In the following, we present the main building blocks of our workhorse scheme, which aims at compensating the full income losses (relative to the other participating countries) following shocks to exports. Suppose that there are $j = 1, \ldots, S$ sectors trading on the world market. The euro area is formed by N countries. Denote by x_{ijt} the period-t value-added of exports by sector j in country i toward the rest of the world. One can write

$$x_{ijt} = w_{ijt} x_{jt},$$

where $x_{jt} = \sum_i x_{ijt}$ is the total value-added of euro-area exports of sector j products, while w_{ijt} is country i's share in this total. In particular, it includes also the export value-added by euro-zone members to other euro-zone members.[9] An alternative would be to exclude intra-euro-zone trade. However, because most of the trade of euro-zone countries is among themselves, such an approach would miss a large fraction of the shocks to the value-added of exports hitting the countries in the system and the transfers based on the remaining shocks will become largely irrelevant. Now, consider the following decomposition:

$$\begin{aligned} x_{ijt} - x_{ij,t-1}\left(1+g_j\right) &= w_{ijt}x_{jt} - w_{ij,t-1}x_{j,t-1}\left(1+g_j\right) \\ &= \left(\Delta w_{ijt}\right)x^*_{jt} + \left(\Delta w_{ijt}\right)\left(x_{jt} - x^*_{jt}\right) \\ &\quad + w_{ij,t-1}\left(x_{jt} - x^*_{jt}\right), \end{aligned} \tag{1}$$

[9] Double-counting is avoided by considering the value-added of exports instead of the total value of exports.

where g_j is the trend growth rate of the euro-area value-added exports of sector j and $x_{jt}^* \equiv x_{j,t-1}(1+g_j)$, i.e., the euro-area value-added exports of sector j we would expect in period t on the basis of the trend. Hence, the shock to be addressed by the transfer scheme is the deviation from the sector-specific trend value, $x_{jt} - x_{jt}^*$. We choose this shock definition because total exports tend to grow over time, but trend growth rates may systematically differ for different sectors. For example, upcoming high-tech sectors likely grow at a systematically higher rate than traditional sectors generating products with little technological content. Since countries feature different sectoral specializations, not taking account of these different sectoral trends may generate transfers that flow systematically in one direction. With our shock definition, we expect shocks to be temporary and to fluctuate around the trend. Hence, we expect them to have at most moderate serial correlation and to be close to zero on average. Below, we will explain in detail how we compute the trend growth rates of the individual sectors.

The first component on the right-hand side of equation (1), $(\Delta w_{ijt})\, x_{jt}^*$, could be negative, because country i's productivity grows more slowly than the EMU-average productivity in this sector or because the quality of its products improves more slowly than the EMU average in this sector, thus resulting in $\Delta w_{ijt} < 0$. The component could also be positive due to improvements in competitiveness relative to other EMU countries. Changes in the weight w_{ijt} are likely to be at least partly the result of differences in government policies, business climate, investment behavior, fiscal devaluations, and so on, and would not justify any cross-border transfers, because they are to a significant extent the result of national choices. The same is true for the second term on the right-hand side of (1), $(\Delta w_{ijt}) \left(x_{jt} - x_{jt}^*\right)$. However, this term is only of a second-order magnitude and is, therefore, likely to be relatively small. Finally, the term $z_{ijt} \equiv w_{ij,t-1} \left(x_{jt} - x_{jt}^*\right)$, which is mainly driven by changes in total euro-zone exports in sector j, and which is based on sectoral weights for the previous period t–1, is largely beyond the control of *national* policymakers.[10] Hence, if moral hazard is to be

[10]Common policies at the EMU level may well have an effect on $x_{jt} - x_{jt}^*$. For example, ECB policy could lead to a fall in the external value of the euro, thereby boosting exports to the rest of the world. However, the influence of an *individual*

minimized, intra-European cross-border transfers could be based on the component z_{ijt} in (1). Of course, z_{ijt} is not perfectly insulated from potential moral hazard. While the weight $w_{ij,t-1}$ is given in period t, future weights can still be affected by current policies. However, we expect future weights to be less relevant for current policies than current weights, because of time discounting and because of the chance that another government will be in office next period.

The ESC requires a choice regarding the component of the income change beyond the government's direct control that is compensated by the transfer. Our baseline scheme aims at compensating the *full income loss*, which would be a natural choice if all the compensation to the capital providers went to domestic inhabitants. An alternative, which we study below, is to compensate the loss of labor income, which would be a natural choice if the shares in the companies producing in each country are perfectly spread over all the euro-zone inhabitants. A natural objective is that the change in the component of the value-added of exports that is beyond the direct control of the government, i.e., z_{ijt}, plus the transfer T_{ijt} implied by this change, is constant for each country as a *fraction* of its total value-added of exports, that is,

$$\frac{w_{kj,t-1}\left(x_{jt}-x_{jt}^{*}\right)+T_{kjt}}{X_{k,t-1}}=\frac{w_{ij,t-1}\left(x_{jt}-x_{jt}^{*}\right)+T_{ijt}}{X_{i,t-1}}, \forall k \neq i, \quad (2)$$

where $X_{it} \equiv \sum_j x_{ijt}$ is defined as the total value-added of country i's exports in period t. Further, we want to impose that aggregated over all countries, the transfers associated with sector j are zero in period t, i.e.,

$$\sum_i T_{ijt}=0. \quad (3)$$

This restriction obviates the need for a central budget capacity to implement the transfer scheme.[11] Imposing the above requirements,

government on $x_{jt}-x_{jt}^{*}$ and, hence, the incentive for moral hazard, would be limited at most.

[11]However, as discussed below, this restriction will be relaxed in the case in which a country that is supposed to contribute to the scheme in a given year is also affected by negative GDP growth in the same year. This may imply that the scheme will be unbalanced in those years.

we can now calculate the transfers as follows. Using (2), the equal percentage net (i.e., including the transfers) effects as a fraction of total value-added of exports for two countries i and k imply that

$$\begin{aligned} T_{kjt} &= \frac{X_{k,t-1}}{X_{i,t-1}} \left\{ \left[w_{ij,t-1} - \frac{X_{i,t-1}}{X_{k,t-1}} w_{kj,t-1} \right] \left(x_{jt} - x^*_{jt} \right) + T_{ijt} \right\} \\ &= \left[\frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} - w_{kj,t-1} \right] \left(x_{jt} - x^*_{jt} \right) + \frac{X_{k,t-1}}{X_{i,t-1}} T_{ijt}, \forall k \neq i. \end{aligned} \tag{4}$$

There are $N-1$ such equations. Using the restriction (3) that the sum of the transfers across countries be zero, we have

$$\begin{aligned} \left[1 + \sum_{k \neq i} \left(\frac{X_{k,t-1}}{X_{i,t-1}} \right) \right] T_{ijt} = \sum_{k \neq i} \left[w_{kj,t-1} - \frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} \right] \\ \times \left(x_{jt} - x^*_{jt} \right). \end{aligned} \tag{5}$$

Using the condition that the weights $w_{ij,t-1}$ sum to one over the countries, we obtain (see appendix A) for the transfers associated with sector j:

$$T_{ijt} = \left[\frac{X_{i,t-1}}{X_{t-1}} - w_{ij,t-1} \right] \left(x_{jt} - x^*_{jt} \right). \tag{6}$$

Expression (6) has a very simple intuitive interpretation. Recall that $X_{i,t-1}/X_{t-1}$ is country i's value-added of exports as a share of aggregate euro-wide value-added of exports, while $w_{ij,t-1} = x_{ij,t-1}/x_{j,t-1}$ represents the corresponding share for sector j only. If the difference between the two is negative, country i's value-added of exports share in sector j exceeds its overall value-added of exports share in the euro zone, i.e., its exports are relatively intensive in sector j (say, the Netherlands in agriculture) compared with its overall exports package. This implies that when there is, for instance, a *positive* shock in the total euro-area export in agriculture ($x_{jt} - x^*_{jt} > 0$), then the Netherlands will be more favorably affected than the other countries. In this case, the net transfer for the Netherlands (associated with this specific sector) would be negative according to (6), meaning that the country would be a contributor. The transfer would go to countries which are relatively less exposed to agriculture (say, to Finland). Conversely, if a *negative* shock hits the same sector, the Netherlands will receive a transfer from countries in which agriculture is relatively less important (for instance, from Austria).

The mechanism can be illustrated further with a simple numerical example: suppose that there are only two countries, Germany and Greece, with shares of total euro-zone exports of, respectively, 90 percent and 10 percent. Additionally, suppose that there is only one sector, say tourism (e.g., hotels and restaurants), and that Greece's exports are relatively more specialized in that sector (e.g., $w_{ij,t-1} = 20\%$). Suppose now that a negative shock hits the total euro-area export of that sector. For example, $x_{jt} - x^*_{jt} = -€1000$ million. Then, given that Greece's exports are relatively more exposed to that sector, it will receive a transfer from Germany amounting to $T_{GR,j,t} = (0.10 - 0.20) * (-1000) = +€100$ million.

From (6), we obtain country i's *total* transfer *from* (or *to*) the rest of the euro zone as

$$T_{it} = \sum_j \left[\frac{X_{i,t-1}}{X_{t-1}} - w_{ij,t-1} \right] \left(x_{jt} - x^*_{jt} \right), \tag{7}$$

where $X_t \equiv \sum_i X_{it}$ is aggregate (i.e., across all sectors) euro-area-wide value-added of exports in period t. When T_{it} is positive, there is a net resource flow from the rest of the euro zone to country i, while when it is negative, there is a net resource flow into the opposite direction.

Some remarks are warranted. First, equation (2) implies that the *sum* over all sectors j of the term $w_{ij,t-1} \left(x_{jt} - x^*_{jt} \right) + T_{ijt}$ as a share of the total value-added of a country's exports is the same for all countries. In other words, the total change in national income (including transfers associated with all the individual sectors) as a share of a country's value-added of exports is the same for all countries. Second, because of the period-by-period balanced budget of the scheme, all the transfers paid will find their way to the other countries in the same year, so there is no saving or dissaving at the aggregate level. Third, while for a given weight $w_{ij,t-1}$ transfers are a function of the temporary deviations $x_{jt} - x^*_{jt}$ from the trend, over a sufficiently long period of time these temporary deviations are zero on average. Therefore, a sector that is on a declining trend should not lead to systematic transfers. Fourth, the derivative of T_{it} in (7) with respect to $w_{ij,t-1}$, given by $-\left(x_{jt}-x^*_{jt}\right)$, is positive in the case of a negative exports shock in sector j. The positive effect of a higher weight $w_{ij,t-1}$ on transfers precisely when such a negative shock hits could give the government an incentive to avoid policies that lead to a decline in the weight $w_{ij,t-1}$. However, because deviations from trend should cancel out on average, a priori we would not expect this to be a source of moral hazard. In the case of a long-lasting negative deviation from the trend, the transfer scheme could undermine the incentive of the government to enter into a politically costly restructuring of the economy toward activities with more long-run viability. Hence, as such, our scheme would not be fully immune to moral hazard. However, as time moves forward, the estimate of the trend is updated further (as we explain below), which reduces the likelihood of the described situation. A solution that further reduces the scope for moral hazard would be to earmark transfers explicitly for compensating losers from structural reforms toward activities with a better future.

Finally, we do not control for the source of the common sectoral shock $\left(x_{jt}-x^*_{jt}\right)$, which may be driven by world market developments, but may also be partly affected by euro-zone-wide (trade) policies. However, there is a priori no reason not to compensate countries if they are relatively badly hurt by the euro zone's own policies, as long as these policies are not affected by moral hazard at the *individual* country's level. The latter is unlikely, since the influence of an individual country on common euro-zone-wide policies is only limited.

We introduce two modifications to scheme (7) in order to increases its political acceptability. First, as suggested by Hebous and Weichenrieder (2015a, 2015b), a transfer scheme with a balanced budget requirement like the current one is likely to produce as a fraction of GDP more volatile transfers for small than for large countries. To see whether this may be the case here as well, we rewrite (7) as

$$T_{it}=\frac{X_{i,t-1}}{X_{t-1}}\sum_j\left[1-\frac{x_{ij,t-1}/X_{i,t-1}}{x_{j,t-1}/X_{t-1}}\right]\left(x_{jt}-x^*_{jt}\right). \tag{8}$$

The part in square brackets depends on the sectoral diversification of the country, and is positive when country i is less exposed to sector j than the euro area as a whole, and vice versa. The fraction $X_{i,t-1}/X_{t-1}$ in front of the summation in (8) shows that transfers are proportional to the size of the country's value-added of exports relative to the euro-area aggregate. Due to their openness, the export shares of small countries in the euro-area aggregate are higher on average than their GDP shares in the euro-area aggregate. Hence, small and open countries will on average inevitably experience larger transfers in absolute terms as a percentage of GDP than larger and more closed economies. To better align the volatility of the transfers as a fraction of GDP across the participating countries, we modify the baseline transfer scheme into

$$T_{it} = \frac{Y_{i,t-1}}{Y_{t-1}} \sum_j \left[1 - \frac{x_{ij,t-1}/X_{i,t-1}}{x_{j,t-1}/X_{t-1}} \right] \left(x_{jt} - x^*_{jt} \right), \tag{9}$$

where $Y_{i,t-1}$ is the GDP of country i and Y_{t-1} is euro-area GDP.[12]

The second modification to arrive at our baseline transfer scheme is that we assume that a country cannot contribute to the scheme in a given year if its GDP growth is negative in the same year. Such a situation would be most likely to occur during a common and severe downturn coinciding with a collapse in world trade as we have seen during the global financial crisis. Widespread negative GDP growth coincides with a decline in essentially all export sectors. However, because sectoral structures and the magnitudes of the sectoral shocks differ, with zero euro-area aggregate transfers, some countries would still be required to make a net transfer payment even though their GDP growth is negative. As this would lead to procyclical transfers and, therefore, would unlikely be politically acceptable in practice, we rule this out by imposing a lower bound of zero on a country's transfers when GDP is contracting

As a result of these two modifications, the aggregate of transfers over the euro zone in a given year may be positive. However, our simulations show that the resulting imbalance when averaged over

[12]An alternative would be to replace $X_{i,t-1}/X_{t-1}$ with the population fraction of country i in the euro zone. However, we find that the differences with the case in which we replace $X_{i,t-1}/X_{t-1}$ with country i's GDP fraction are minimal and, hence, we do not report the case with population fractions.

the sample period will be very small. Therefore, we will not explicitly address in our simulations how this imbalance can be eliminated. There are several possible ways in which this can be done. For example, the scheme's participants could contribute a (very) small fraction of their GDP to a dedicated central budget when their growth is above a certain threshold. Or the central capacity may be allowed to issue debt to finance the imbalance, after which the participating countries share in the debt-servicing costs in proportion to their GDP, again as long as their growth is positive. Such financing of the central budget during periods when growth is (sufficiently) positive would in fact strengthen the countercyclical character of the scheme.

3.2 Compensation for Labor Income Loss

Suppose that asset holdings are perfectly diversified over all individuals in the euro area. Then, it is natural to assume that the ESC should cover unexpected changes in labor income only. Let p_{ijt} be the average productivity in sector j in country i measured as the value of production per worker expressed in euros. Hence, the change in employment in sector j in country i associated with z_{ijt} equals $w_{ij,t-1}\left(x_{jt}-x_{jt}^{*}\right)/p_{ijt}$. Furthermore, let s_{ijt} be the average salary in sector j in country i. Then, expressed in euros, the amount of labor income associated with z_{ijt} equals $w_{ij,t-1}\left(x_{jt}-x_{jt}^{*}\right)\left(s_{ijt}/p_{ijt}\right)$. Notice that $lsh_{ijt} \equiv s_{ijt}/p_{ijt}$ is the *labor share* in value-added in country i in sector j.

Now, imposing equal net (i.e., after transfers) income changes as a share of the value-added of exports for the two countries i and k implies

$$\frac{w_{kj,t-1}\left(x_{jt}-x_{jt}^{*}\right)lsh_{kjt}+T_{kjt}}{X_{k,t-1}} = \frac{w_{ij,t-1}\left(x_{jt}-x_{jt}^{*}\right)lsh_{ijt}+T_{ijt}}{X_{i,t-1}}, \forall k \neq i. \tag{10}$$

Rewriting, using the fact that there are N–1 such equations and the restriction that the sum of the transfers be zero, we obtain after some manipulation (see appendix B):

$$T_{ijt} = \sum_{k \neq i} \left[\frac{X_{i,t-1}}{X_{t-1}} w_{kj,t-1} lsh_{ijt} - \frac{X_{k,t-1}}{X_{t-1}} w_{ij,t-1} lsh_{kjt} \right] \left(x_{jt} - x^*_{jt} \right). \tag{11}$$

The transfers are extremely easy to calculate. However, simplification of this expression is only possible in the simple case in which the labor shares of value-added in a given sector j are identical across the countries, i.e., $lsh_{jt} \equiv lsh_{ijt}, \forall i$. In that case,

$$T_{it} = \sum_j lsh_{jt} \left[\frac{X_{i,t-1}}{X_{t-1}} - w_{ij,t-1} \right] \left(x_{jt} - x^*_{jt} \right). \tag{7$'$}$$

Hence, in this case, up to the proportionality factor lsh_{jt} for sector j, the transfers are the same as when the full income effect from the shock is equalized across the countries.

In line with our baseline scheme, we align the volatilities of the transfers across countries by setting

$$T_{it} = \frac{Y_{i,t-1}}{Y_{t-1}} \sum_j lsh_{jt} \left[1 - \frac{x_{ij,t-1}/X_{i,t-1}}{x_{j,t-1}/X_{t-1}} \right] \left(x_{jt} - x^*_{jt} \right). \tag{9$'$}$$

Again, we will cap negative transfers at zero when a country's GDP growth is negative.

3.3 Compensation for Losses of Tax Revenues

According to the proposed ESC scheme, a country that experiences an improvement in the world trade of its relatively export-intensive sectors will have to make a net transfer that will benefit less fortunate countries. However, it may not be so easy to free up the resources for making these transfers: the extra export value-added resulting from the positive shock is spent on compensating the providers of labor and capital. Yet, the additional income also generates additional tax revenues for the government of the lucky country, and these can be used for transfers to unlucky countries that are confronted with a shortfall in tax revenues. A complication with conditioning transfers on tax revenues is that countries have different tax rates: countries with relatively low tax rates would experience relatively small transfers in absolute magnitude, and vice versa for countries with relatively high tax rates. The transfer scheme should be designed in such a way that it avoids potential incentives to manipulate tax rates in order to extract more transfers. Hence, we assume that all transfers in a given period are based on a common tax rate τ_t. Further, to ensure that the governments that are net payers into the system have the resources available to pay their transfers, τ_t cannot be too high.

The transfer scheme based on the effect of sectoral value-added exports shocks for government revenues is a direct variation of the baseline scheme. Denote the tax rate of country i by τ_t. Hence, the transfer received from or paid to the other countries in the system would be

$$T_{it} = \tau_t \frac{Y_{i,t-1}}{Y_{t-1}} \sum_j \left[1 - \frac{x_{ij,t-1}/X_{i,t-1}}{x_{j,t-1}/X_{t-1}} \right] \left(x_{jt} - x^*_{jt} \right), \qquad (12)$$

and $T_{it} = 0$ if both the right-hand side of (12) and country-i GDP growth in period t are negative. Again, we align the cross-border volatility of the transfers by making them proportional to the GDP share, and we exclude negative transfers in case of a recession.

4. The Data

We obtain yearly data on x_{ijt}, the value-added content of exports by country and by sector toward the rest of the world (including the other EMU countries), from the OECD (2017b) Trade in Value Added (TiVA) database. The sample covers all the 19 euro-area (EA19) countries and runs from 1995 to 2014. The industrial sectors into which exports are subdivided correspond to those of the third revision of the International Standard Industrial Classification (ISIC Rev. 3). Using these data, we can calculate $x_{jt} = \sum_i x_{ijt}, x_t = \sum_j x_{jt}$, and $w_{ijt} = x_{ijt}/x_{jt}$. We exclude the sector "Finance and Insurance" from our analysis. The sector is generally tightly regulated by the authorities and often faces specific tax treatment. Moreover, its disproportionate presence in some relatively small countries would result in very large cross-border transfer payments (or receipts) when shocks hit that sector, undermining the political acceptability of the proposed arrangement. Hence, our data set comprises the 32 sectors listed in table 1.

Table 1. Country Shares of Value-Added Sectoral Exports (w_{ij})

	DE	FR	IT	ES	NL	GR	PT	IE	CV	LV	AT	BE	EE	FI	LT	LU	MT	SK	SI
Agriculture, Hunting, Forestry, and Fishing	12.3	25.9	8.4	17.5	18.7	4.0	1.4	1.4	0.1	0.6	1.7	3.4	0.3	1.4	0.9	0.3	0.1	1.2	0.4
Mining and Quarrying	13.3	7.7	4.9	5.2	57.4	1.0	1.8	1.2	0.0	0.1	2.9	2.0	0.2	0.8	0.3	0.1	0.0	1.0	0.1
Food Products, Beverages, and Tobacco	20.3	19.4	11.9	10.5	15.4	2.0	1.8	7.1	0.2	0.3	2.8	5.6	0.2	0.8	0.5	0.3	0.1	0.6	0.3
Textiles, Textile Products, Leather, and Footwear	15.6	14.6	40.1	9.3	2.2	2.1	5.6	0.5	0.1	0.3	2.5	3.6	0.3	0.6	0.6	0.3	0.1	0.9	0.7
Wood and Products of Wood and Cork	22.7	8.9	9.9	5.4	3.0	0.4	6.5	1.1	0.0	3.8	13.2	4.1	2.1	13.5	1.2	0.3	0.0	2.1	1.8
Pulp, Paper, Paper Products, Printing . . .	34.0	12.2	8.4	6.9	6.9	0.4	2.4	3.0	0.0	0.1	5.9	3.7	0.2	13.6	0.2	0.3	0.1	1.0	0.7
Coke, Refined Petroleum Products, and Nuclear Fuel	18.5	15.4	10.5	9.0	22.8	2.7	0.8	1.3	0.0	0.0	1.5	10.3	0.2	2.6	3.0	0.0	0.0	1.1	0.0
Chemicals and Chemical Products	34.2	19.6	9.9	6.7	9.2	0.7	0.6	8.1	0.0	0.1	2.1	6.2	0.1	1.3	0.1	0.1	0.0	0.4	0.6
Rubber and Plastics Products	37.4	15.4	16.7	8.0	5.7	0.7	1.7	0.9	0.0	0.1	4.0	4.0	0.1	1.6	0.3	0.9	0.1	1.3	1.0
Other Non-metallic Mineral Products	25.7	12.0	24.5	13.4	4.0	1.4	3.4	1.0	0.0	0.2	4.8	5.0	0.2	1.5	0.2	0.7	0.0	1.2	0.7
Basic Metals	34.1	16.0	14.3	9.8	4.8	1.7	1.1	0.3	0.0	0.3	5.4	6.5	0.0	3.2	0.0	0.8	0.0	1.4	0.5
Fabricated Metal Products Except . . .	33.5	11.1	24.3	7.3	6.1	0.7	1.6	0.6	0.0	0.1	5.2	4.1	0.2	1.8	0.2	0.3	0.0	1.5	1.3
Machinery and Equipment, nec	43.7	11.7	25.5	3.9	3.9	0.3	0.6	0.5	0.0	0.0	4.0	1.7	0.0	2.7	0.1	0.1	0.0	0.6	0.5
Computer, Electronic, and Optical Products	39.9	19.3	10.3	3.7	5.6	0.4	0.9	7.1	0.0	0.1	3.2	1.5	0.2	5.8	0.1	0.1	0.2	1.1	0.4
Electrical Machinery and Apparatus n.e.c.	46.8	15.9	14.2	6.2	2.1	0.6	1.4	1.1	0.0	0.1	4.4	2.2	0.2	2.5	0.1	0.1	0.1	1.3	0.7
Motor Vehicles, Trailers, and Semi-trailers	54.6	14.9	9.0	10.5	1.7	0.1	1.0	0.1	0.0	0.0	2.7	3.0	0.0	0.4	0.0	0.0	0.0	1.5	0.4

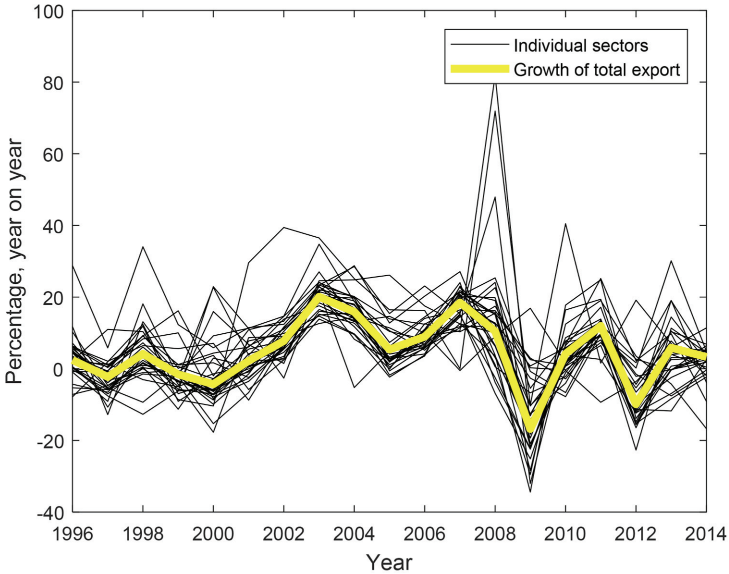
(*continued*)

Table 1. (Continued)

	DE	FR	IT	ES	NL	GR	PT	IE	CV	LV	AT	BE	EE	FI	LT	LU	MT	SK	SI
Other Transport Equipment	33.9	37.6	11.7	6.5	3.5	0.5	0.5	0.4	0.0	0.1	1.8	1.0	0.1	1.7	0.2	0.0	0.0	0.3	0.5
Manufacturing n.e.c.; Recycling	23.4	17.3	26.1	7.7	7.6	1.0	1.7	1.3	0.2	0.2	4.3	4.6	0.4	1.1	0.7	0.1	0.2	1.0	1.1
Electricity, Gas, and Water Supply	68.1	8.9	2.0	2.5	5.6	0.4	0.3	0.2	0.0	0.1	3.3	5.9	0.3	0.8	0.1	0.3	0.0	0.3	0.8
Construction	12.0	34.4	6.2	2.5	8.1	2.3	2.9	0.6	2.1	0.3	7.3	11.6	0.6	5.3	0.4	0.9	0.1	1.3	1.1
Wholesale and Retail Trade; Repairs	27.0	20.1	17.1	8.4	7.3	1.1	1.8	3.1	0.2	0.2	4.1	5.3	0.3	1.6	0.3	0.8	0.1	1.0	0.4
Hotels and Restaurants	8.8	13.3	22.9	25.9	2.4	6.0	3.2	1.4	1.0	0.1	8.7	2.7	0.3	0.8	0.2	0.8	0.4	0.6	0.6
Transport and Storage	20.9	18.6	12.8	12.2	10.0	5.8	2.3	1.7	0.4	0.6	3.0	6.6	0.5	1.5	0.7	1.0	0.2	0.7	0.6
Post and Telecommunications	18.9	12.7	14.9	8.3	10.5	1.8	1.9	3.7	1.0	0.3	5.5	11.3	0.4	1.3	0.3	5.4	0.3	0.9	0.7
Real Estate Activities	15.1	13.6	28.8	12.8	5.5	2.7	4.7	0.8	2.7	0.1	7.0	2.2	0.2	0.7	0.5	1.1	0.2	0.6	0.6
Renting of Machinery and Equipment	24.0	31.8	4.1	6.3	3.9	1.1	1.0	16.9	0.4	0.2	2.7	3.3	0.3	1.9	0.1	1.8	0.2	0.2	0.0
Computer and Related Services	28.0	4.1	6.6	12.2	7.8	0.6	0.6	23.2	0.4	0.1	2.9	5.2	0.2	5.4	0.1	1.6	0.2	0.5	0.3
Research and Development and Other . . .	28.9	16.6	10.2	11.2	7.7	0.6	0.8	2.8	0.5	0.2	4.3	11.4	0.2	2.6	0.1	0.8	0.2	0.5	0.4
Public Admin. And Defense; Compulsory . . .	1.3	25.7	1.1	30.9	18.7	1.3	3.8	0.0	0.0	0.0	4.6	7.8	0.2	2.7	0.0	1.1	0.1	0.4	0.2
Education	11.8	16.3	5.6	8.5	25.0	1.8	2.6	2.7	2.1	0.1	3.7	14.4	0.3	0.8	0.2	1.4	0.3	0.8	1.7
Health and Social Work	24.6	28.8	2.5	9.3	9.0	2.1	2.9	3.1	0.2	0.1	9.1	2.0	0.1	2.7	0.3	1.0	0.1	1.3	0.9
Other Community, Social, and Personal Services	21.7	20.5	10.6	10.0	7.2	4.5	1.8	1.7	1.4	0.2	6.9	4.9	0.4	1.3	0.6	2.7	0.9	1.6	1.3
Country's Total in Euro-Area Export Value-Added	31.2	17.4	15.0	9.0	7.4	1.6	1.6	3.0	0.2	0.2	3.8	4.8	0.2	2.2	0.3	0.5	0.1	0.9	0.5

Notes: This table reports the share (in percent) of each euro-zone country's export value-added over total euro-zone export value-added for each of the 32 sectors included in the sample. Numbers are averages over the 1995–2014 period. The sum of each row is therefore 100 percent, while the last row indicates the share of that country's total export value-added over the total euro-zone export value-added. AU = Austria, BE = Belgium, EE = Estonia, FI = Finland, FR = France, DE = Germany, GR = Greece, IE = Ireland, IT = Italy, LV = Latvia, LU = Luxemburg, NL = Netherlands, PT = Portugal, SK = Slovakia, SI = Slovenia, ES = Spain, CY = Cyprus, LT = Lithuania, and MT = Malta. Green (red) are the three largest (smallest) value-added export shares, for each country.

Figure 1. Growth Rate of Value-Added of Sectoral Exports at the Euro-Area Level



Notes: The figure shows the annual growth rate of the individual sectoral export value-added at the euro-area level. Sectoral labels are not shown for sake of space.

Figure 1 depicts the annual growth rate of total euro-zone exports (gray or yellow line in the print and online version, respectively) and of euro-zone exports by individual sector (not labeled for simplicity). The annual growth rate of total exports averages 5.4 percent over the full sample. Export growth is generally positive, but several years are also characterized by negative growth rates. Most notably, the 2009 global economic and financial crisis exhibits a very severe fall of more than 15 percent in total exports, with some sectors dropping by as much as 35 percent in that year.

Table 1 also reports, for each sector and country, the country's average (over time) exports share in that sector's total euro-zone exports. At the bottom of the table we report, for each country, its share in total euro-zone exports. For each country, we have marked the three smallest (shaded dark gray) and the three largest (shaded light gray) sectors in terms of euro-zone share. Obviously, given that Germany is the largest economy and the largest exporter, it is also the largest exporter in a substantial number of sectors. In some sectors, it is very dominant, such as "Electrical Machinery and Apparatus n.e.c.," "Motor Vehicles, Trailers, and Semi-trailers," and "Electricity, Gas, and Water Supply." Hence, large negative shocks in these sectors could potentially lead to large transfers to Germany that need to be financed by all the other countries. This effect is mitigated by the fact that Germany is a relatively diversified economy over the various sectors and that, in the case of our baseline, the transfers are essentially driven by the *difference* between Germany's share in individual sectors with Germany's share in total exports (recall equation (7)), which is high not only because of its size but also because of its openness, and its share in the exports of the specific sectors.

Data on nominal and real GDP of the EA19 countries are retrieved from the OECD (2017a) and from the World Bank (2017). The output gap is taken from the OECD (2017a). Using data on compensation of employees as a fraction of value-added by industry and country from the OECD (2019) structural analysis (STAN) database, we calculate the labor share of gross value-added in sector j (lsh_{jt}) as

$$lsh_{jt} = \sum_i \left[\frac{\frac{WL_{ijt}}{GVA_{ijt}} Y_{it}}{\sum_i Y_{it}} \right],$$

where WL_{ijt} is the total compensation of employees and GVA_{ijt} is gross value-added in country i, sector j, and year t. Hence, lsh_{jt} is a weighted average of the labor shares in sector j in the different countries. Data are available over the full sample period 1995–2014.

The tax rate τ_t is the EA19 value for "Total Receipts from Taxes and Social Contributions (including imputed social contributions) after Deduction of Amounts Assessed but Unlikely to be Collected" as a percentage of euro-area GDP. It is retrieved from Eurostat (2017). It ranges from 39.0 percent in 2010 to 41.3 percent in 2014.

Average revisions of export data over time are computed using different editions of the AMECO database (2017).

All data are annual, and expressed either in million US$ or percentages. U.S. dollars were chosen for all amounts expressed in currency so as to avoid exchange rate complications.

5. Transfers Calculated on the Basis of Actual Data

In this section, we present the results for our baseline scheme, described in subsection 3.1, and of its variants (presented in subsections 3.2–3.3) simulated over the period 2002–14 and based on the data described above.

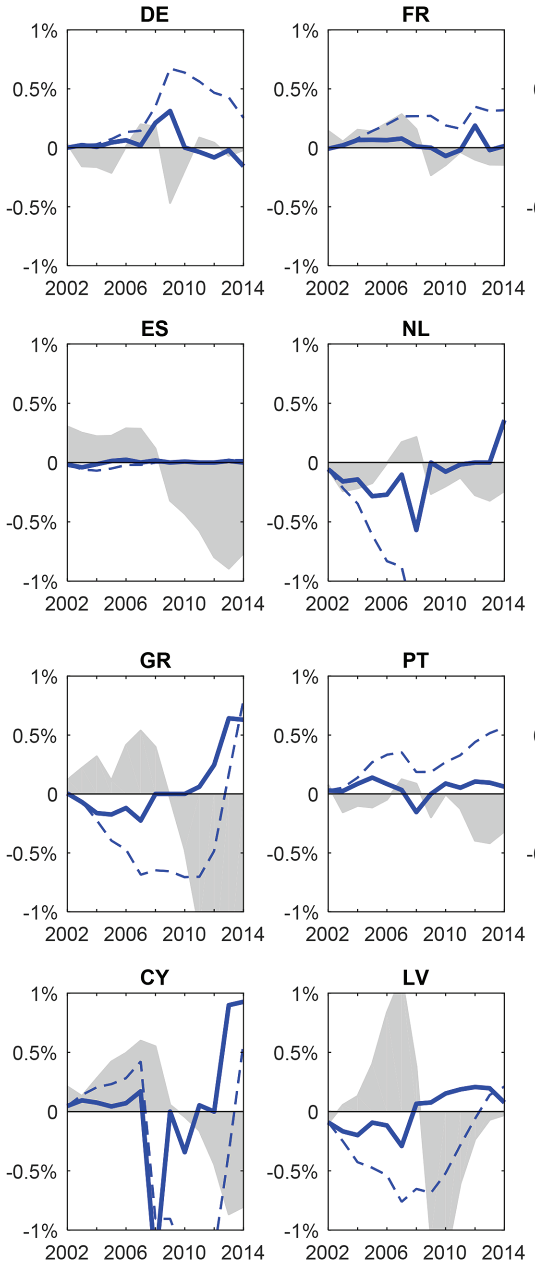
The first step consists of estimating the sectoral trends for each sector j, which underpin the equation for the baseline scheme (9). To that aim, we adopt a “pseudo real-time” approach: we assume that we are in year t, and that we observe data from t_0 (the beginning of the sample) until t. Then, we apply the Hodrick-Prescott filter to that sample and come up with an estimate of the trend value for sector j, i.e., x^*_{jt}. Finally, we estimate the shock for period t and sector j as $x_{jt} - x^*_{jt}$, and include the estimate in equation (9). We repeat these steps from 2002 until the end of the sample, i.e., $t = 2002, 2003, \ldots, 2014$, while keeping the first available year as fixed (1995). Hence, for each year t, we will therefore come up with a (slightly) different estimate of the trend, given that available information would be updated based on the incoming year.[13] This “recursive” approach addresses the end-period problem of the Hodrick-Prescott filter and avoids the use of future observations in performing the trend-cycle decomposition. This procedure will result in a time series of shocks for each sector j from 2002 until 2014. It turns out that these shocks are on average close to zero and symmetrically distributed around it.

5.1 Baseline Scheme: Equalizing Income Shifts as a Fraction of the Value-Added of Exports

Figure 2 depicts, for each euro-zone country, the simulated annual transfers in percent of GDP of the same year (the solid line). In addition, the figure plots the cumulative transfers (the dashed line). The latter are calculated as $x_{jt} = \sum_{\tau=2002}^{t} T_{i,\tau}/y_{it}$, i.e., the sum of the transfers up to year t over GDP in year t. Finally, the figure also shows the output gap as the gray filled area. We first report the “big-5” euro-zone countries (Germany, France, Italy, Spain, the

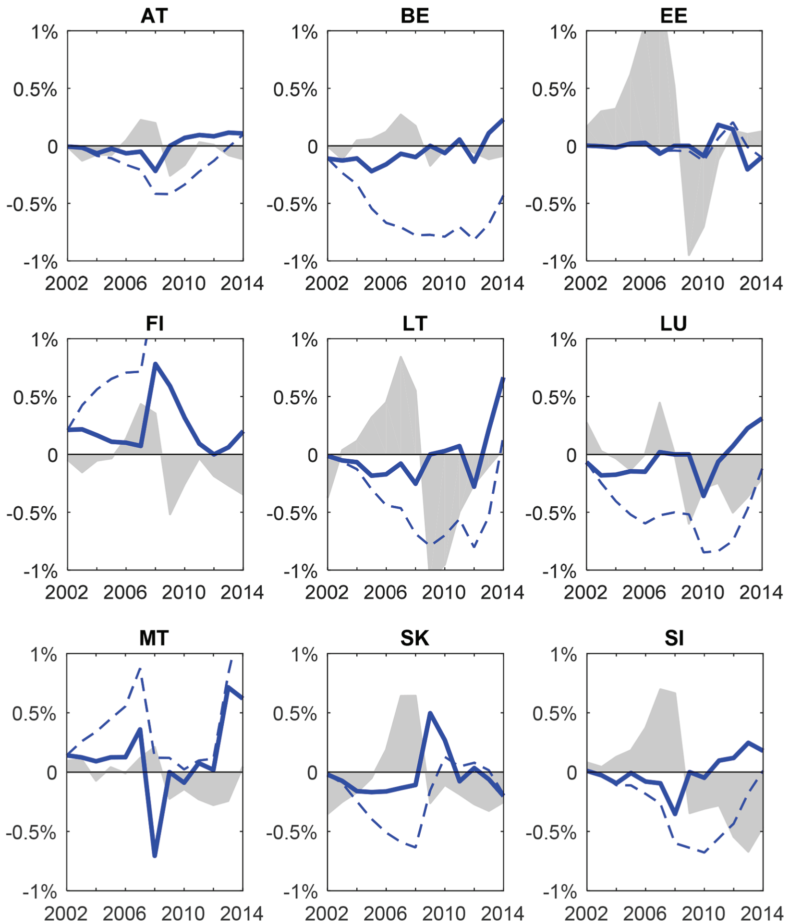
[13] We have experimented with a different starting year, 2001 and 2003, and results are very similar (not reported, available upon request).

Figure 2. Annual and Cumulated Transfers Implied by the Baseline Scheme in Percent of the Country's GDP



(*continued*)

Figure 2. (Continued)



Notes: The figure shows the annual and cumulated transfers for each country, generated by the baseline scheme (equation (9)), as a percent of a country's GDP in that year. The solid lines are the annual transfers, while the dashed lines are the cumulated transfers. The gray shaded areas are output gaps published by the OECD (2017a). For all countries, we report data in the scale +1 to −1 percent of GDP, for consistency. In the case of the Netherlands, the cumulated transfers are at around 1 percent of GDP in 2014. For Finland, they are about 2.5 percent of GDP. For country codes, see the notes to table 1.

Netherlands), then the countries that received financial aid during the 2008–09 global financial crisis and the ensuing 2010–12 sovereign debt crisis (Greece, Portugal, Ireland, Cyprus, Latvia), followed by the remining nine euro-area countries.

We start by discussing the results for the "big-5" euro-zone countries. In the depth of the global financial crisis, in 2009, in particular Germany is a major receiver of transfers—in that year it receives about 0.3 percent of GDP in net transfers. Indeed, this year is characterized by sizable and negative shocks in sectors, such as "Motor Vehicles, Trailers, and Semi-trailers" and "Machinery and Equipment n.e.c.," in which Germany is relatively more specialized.[14] That would have been compensated by small but negative transfers (contributions to other countries) in the following years, bringing down cumulative transfers by the end of the sample (dashed line). According to the baseline scheme, France receives small transfers (amounting to less than 0.1 percent of GDP) in the pre-crisis period, and a somewhat larger transfer in 2012, driven by the performance of the sectors in which it is more exposed (transport equipments and construction). The case of Italy, being a large country going through a period of low and even negative growth after the start of the sovereign debt crisis, provides a particularly interesting illustration. In each of the years 2009–13 it receives net transfers, although those for 2009 are very small. These net transfers cannot be attributed to any sector specifically, although there are a number of sectors, "Textiles, Textile Products, Leather, and Footwear," "Fabricated Metal Products except Machinery and Equipment," "Machinery and Equipment n.e.c.," "Manufacturing n.e.c.; Recycling," and "Electricity, Gas, and Water Supply," responsible for a relatively substantial fraction of the incoming transfers. Except for "Electricity, Gas, and Water Supply," all these sectors switch from contributing to transfer outflows before the crisis to generators of transfer inflows after the crisis. Interesting, there is also a set of sectors, "Chemicals and Chemical Products," "Motor Vehicles, Trailers, and Semi-trailers," "Computer, Electronic, and Optical Products," and "Transport and

[14]The euro-area value-added shares of Germany's exports in "Motor Vehicles, Trailers, and Semi-trailers" and "Machinery and Equipment n.e.c." are, respectively, 54.6 percent and 46.8 percent, which compares with an overall share—of total German exports over euro-area exports—of 31.2 percent (see table 1).

Storage," that switch from contributing to transfer inflows before the crisis to transfer outflows after the crisis. Hence, for a large, sectorally diversified country like Italy, net transfers are the result of many offsetting contributions coming from the individual sectors. This is, in particular, also the case for Spain. Net transfers to and from Spain are very close to zero in each period. More detailed inspection shows that the overall low transfers are due to the fact that some dominant sectors are hit by shocks that go into opposite directions, thus generating offsetting effects. If we aggregate the absolute values of the transfers associated with the individual sectors, the result would be similar to that for Italy. Finally, the Netherlands contribute to the scheme in the first part of the sample, due to the good performance of sectors in which it is highly exposed (mining and fuel) and low exposure to declining sectors (textile). In the last part of the sample, however, it becomes a net receiver of transfers.

Moving to the set of "crisis countries," figure 2 shows that Greece, the EU country hit hardest by the sovereign debt crisis, would after the start of the crisis benefit in particular from substantial incoming transfers associated with "Transport and Storage" and "Hotels and Restaurants," while it would experience outgoing transfers associated with "Chemicals and Chemical Products," "Fabricated Metal Products except Machinery and Equipment," "Machinery and Equipment n.e.c.," and "Computer, Electronic and Optical Products," sectors that were hit relatively hard, but in which it has relatively little export presence. The case of Ireland is also of specific interest. The country has negative growth in 2008 and 2009 and is one of the major victims of the sovereign debt crisis. However, the problems seem to have been largely absorbed in the public sector, by taking over failing banks. As of 2010, the country grows again at a rather substantial speed and it becomes a net payer of transfers. For Portugal, we see a substantial turnaround of transfers associated with "Transport and Storage," which switch from being relatively large and negative up to 2008 to being relatively large and positive as of 2009. "Textiles, Textile Products, Leather, and Footwear" also switches from being a subtractor to a contributor to the net transfer in 2009 and this is also the case for "Hotels and Restaurants." However, Portugal's net transfer is the sum of many sector-related contributions that can be positive and negative. In particular, the contributions by "Machinery and Equipment n.e.c.," "Computer, Electronic and Optical Products," and "Motor Vehicles, Trailers, and Semi-trailers" are quite strongly negative shortly after the start of the crisis. This is the consequence of the combination of these (rather large) sectors being hurt particularly severely by the crisis and Portugal having relatively little export presence in these sectors. Finally, for Latvia we observe a switch in 2009 from negative to positive contributions to the net transfer by the sectors "Wood and Products of Wood and Cork" and "Transport and Storage." This switch dominates a simultaneous switch into the other direction caused by "Motor Vehicles, Trailers, and Semi-trailers," a badly hurt sector in which Latvia has negligible export presence. In size, these three switches are the dominant ones for Latvia.

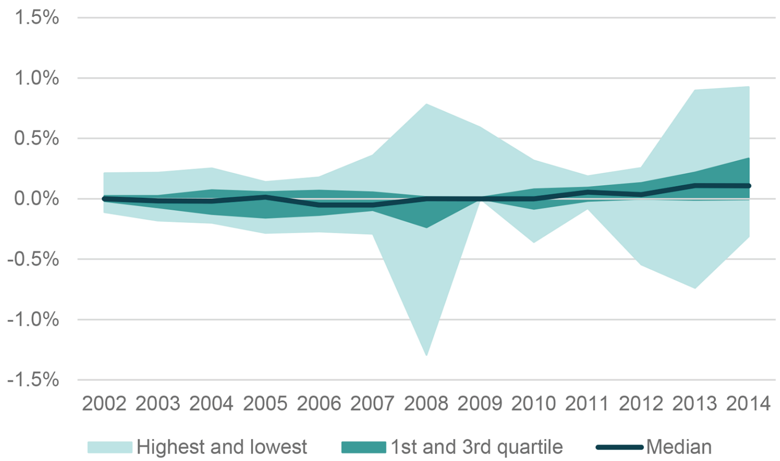
In general, we note that—even if the baseline scheme would have prescribed negative transfers for some countries in some years—these were put to zero due to the condition that a country cannot contribute to the scheme in times of negative GDP growth. This is, for instance, the case of France in 2009, Greece in 2008 and 2009, and Portugal in 2009.

Overall, the graphs suggest that, first, annual transfers are *on average* rather small in absolute magnitude, amounting in most cases to less than 0.2 percent of GDP. Figure 3A, which depicts for each year the cross-country dispersion of the baseline transfers in percent of GDP, shows that the median transfer is zero or close to zero in all years, while the 25th and 75th percentiles are generally also close to zero. However, occasionally, annual transfers can become quite substantial, reaching levels on the order of 0.5–1 percent of GDP. The dispersion is widest during the economically difficult years of 2008 and 2013–14. Generally speaking, transfers tend to be somewhat larger in absolute magnitude for the smaller economies. A priori one might expect this to be the result of a smaller degree of sectoral diversification of these countries' exports. The so-called Herfindahl index, calculated as $H_{it} = \left(\sum_j \left(w_{ijt}/\sum_j w_{ijt}\right)\right)^2$, is increasing in the degree of sectoral export specialization of a country i.[15] Figure 4

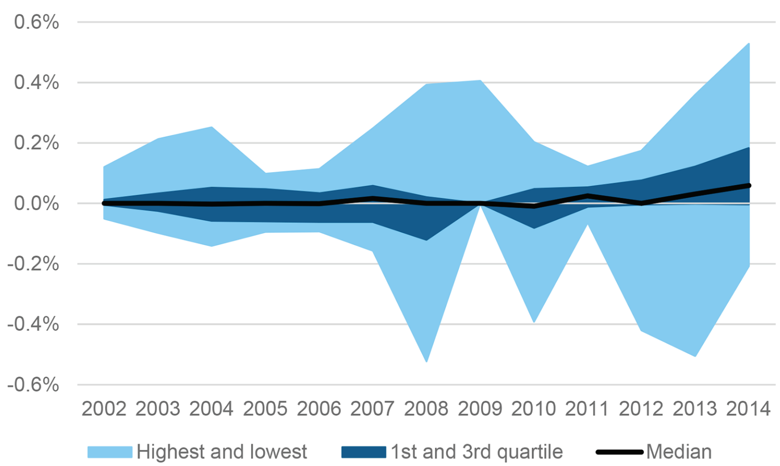
[15]The idea of the Herfindahl index is that if an economy is weakly diversified, it features some sectors with large export weights, which drive up the index, because the latter is based on the sum of the quadratic values of the weights.

Figure 3. Dispersion of Transfers for the Different Schemes (% of GDP)

A. Baseline



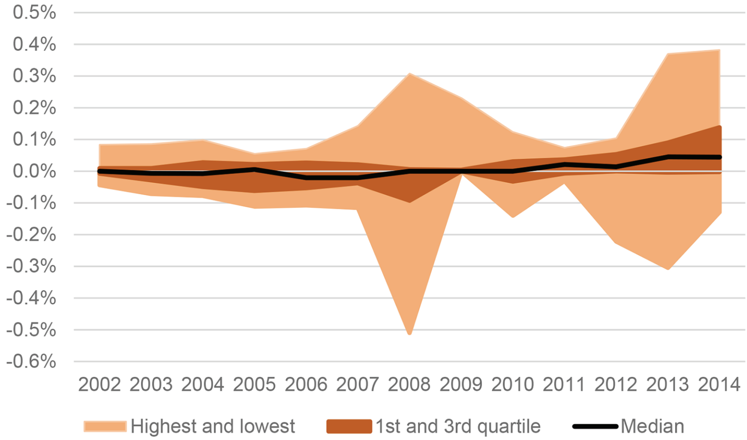
B. Scheme Based on Labor Shares



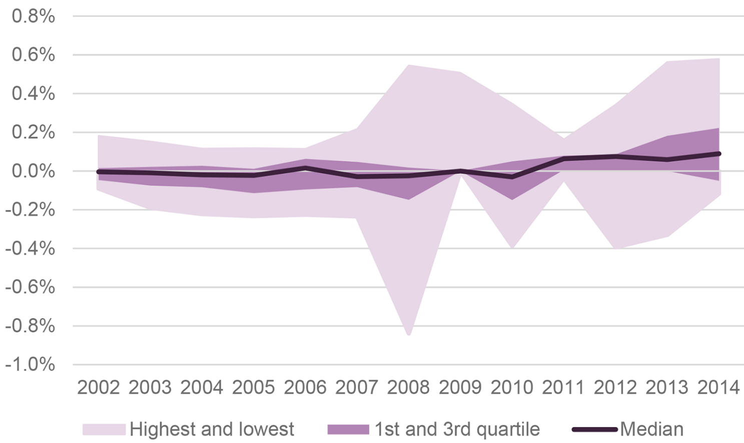
(*continued*)

Figure 3. (Continued)

C. Scheme Based on Tax Revenues

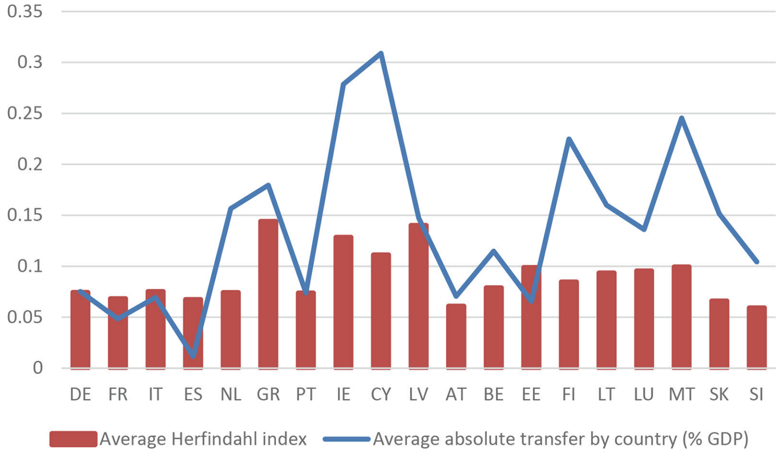


D. Scheme Based on Reaggregation of Sectors



Notes: The figure reports the dispersion of the transfers generated by (A) the baseline scheme, (B) the scheme based on labor shares, (C) the scheme based on tax rate revenue, and (D) the scheme based on reaggregation of sectors. In particular, the charts show, for each year, the highest and lowest values, the first and third quartile, and the median value.

Figure 4. Sectoral Diversification and Size of Transfers



Notes: The figure depicts by country the average of the absolute values of the transfers across the years and the average of the Herfindahl index across the years. Values are in percent of that country's GDP.

suggests that the index, calculated for each country as the average over the sample years, is indeed slightly larger on average for the smaller than for the larger economies, thus confirming some positive correlation (precisely, 0.59) between the absolute magnitude of the transfers and the Herfindahl index.

Second, transfers tend to be countercyclical, i.e., they are generally positive (negative) when the output gap is negative (positive), as is also visible from the charts, which combine transfers with the output gap for each country. As mentioned above, for example, Germany was hit relatively more by the 2009 crisis, thus receiving a transfer in that year. Countries relatively less exposed to sectors that are hit particularly hard by the crisis would, in principle, have to make net transfer payments, but due to the widespread negative growth, these net payments are capped at zero. As shown in figure 2, however, in spite of these net receipts in the depth of the crisis, German cumulative transfers are almost identical to zero at the end of the sample.

The period 2012–14, which contains the double-dip recession in the wake of the EU sovereign debt crisis, is also characterized by strong countercyclical net transfers: Greece receives net transfers of 0.64 percent and 0.63 percent of GDP in 2013 and 2014, respectively, while Italy receives net transfers of 0.26 percent and 0.14 percent in 2012 and 2013, respectively. Other examples of substantial net transfers are Cyprus with 0.90 percent and 0.93 percent of GDP and Malta with 0.71 percent and 0.62 percent of GDP in 2013 and 2014, respectively. Other examples are the crisis countries Latvia and Portugal. Overall, while transfers are on average of limited magnitude, at specific moments of severe economic circumstances they can be quite substantial, meaningfully helping to ameliorate a country's cyclical situation. This has in particular been the case for most crisis countries at moments when they suffered from substantial economic slack.

Indeed, crucial for the transfer scheme's economic usefulness and political viability is the degree to which the transfers correlate with overall economic activity: it is desirable that transfers are positive when the economy is doing relatively poorly and vice versa when it is doing relatively well. Table 2 reports the coefficient estimates of various panel regressions. We perform separate regressions of transfers on the net value of exports, its lag, and various measures of activity, namely the output gap and its first lag, and GDP growth and its first lag. We also estimate a variant in which we split observations into positive and negative output gaps.[16] The table reports just the coefficient associated with these explanatory variables, for simplicity. We find that transfers are significantly countercyclical to the net value of exports and to all activity measures, as indicated by the negative coefficients. The lagged value of exports enters with a negative sign, but is not significant, presumably because the number of observations is rather limited due to the rather short sample period, which results from the need to initialize the trend. The degree of countercyclicality is stronger and more significant in the case of a negative output gap than in the case of a positive output gap, indicating that transfers act as stabilizers precisely when this is most desirable, i.e., when the economy is relatively depressed.

Third, transfers tend to revert to zero or to switch sign after a few years (e.g., see Germany). This is a direct consequence of the

[16] We include country fixed effects and allow standard errors to be clustered by country.

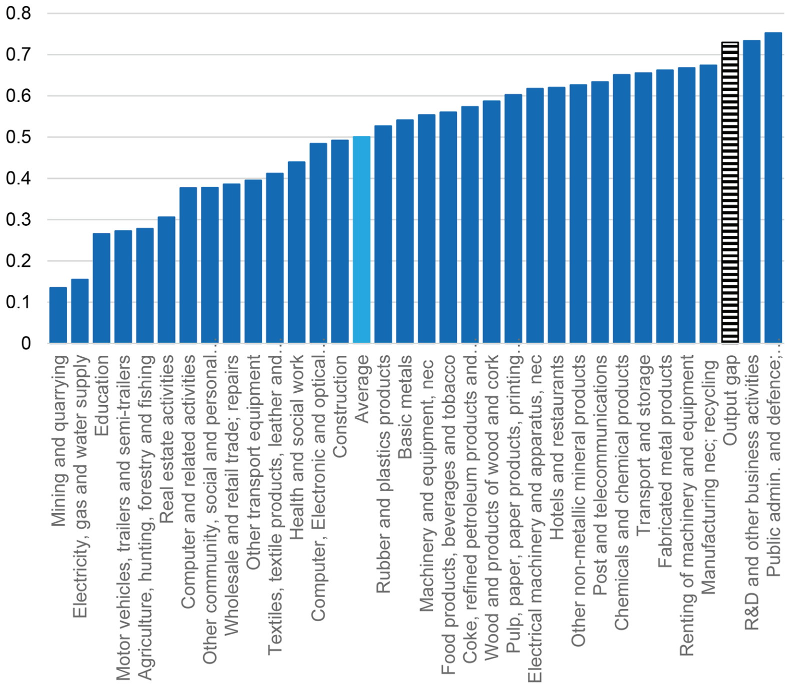
Table 2. Testing the Countercyclicality of the Schemes

Dependent →	**Transfers/GDP for Variant**			
Explanatory ↓	**Baseline**	**Labor Shares**	**Taxes**	**Reagg-regation**
Exports	−0.32*	−0.14	−0.13*	−0.27**
	(0.18)	(−0.13)	(0.073)	(0.12)
Lagged Exports	−0.20	−0.022	−0.083	−0.16*
	(0.17)	(0.10)	(0.066)	(0.088)
Output Gap	−0.019***	−0.0077**	−0.0078***	−0.014**
	(0.0060)	(0.0029)	(0.0025)	(0.0052)
Lag Output Gap	−0.016**	−0.0054	−0.0064**	−0.011*
	(0.0074)	(0.0039)	(0.0030)	(0.0056)
Positive Output Gap	−0.017*	−0.0093*	−0.0069	−0.0068
	(0.0099)	(−0.0048)	(0.0040)	(0.0061)
Negative Output Gap	−0.028**	−0.015**	−0.011**	−0.020**
	(0.012)	(0.0065)	(0.0051)	(0.0099)
GDP Growth	−0.011**	−0.0051**	−0.0045**	−0.0083**
	(0.0044)	(0.0024)	(0.0018)	(0.0035)
Lag GDP Growth	−0.0113*	−0.0035	−0.0046*	−0.0070
	(0.0059)	(0.0030)	(0.0024)	(0.0047)

Notes: The table reports regression coefficients from panel regressions which are run on the explanatory variables reported in the first column, country fixed effects, and a constant, over all countries in the sample and period 2002–14. The regressions are conducted on one explanatory variable at a time, except for the positive and negative output gaps, which are entered jointly as explanatory variables in the same regression. The standard errors, reported in the parentheses, are based on clustering over the countries. ***, **, and * indicate significance at the 1 percent, 5 percent, and 10 percent level, respectively.

design of the transfer scheme, whereby transfers depend on deviations of sectoral value-added from sectoral trends. These deviations cannot be permanently positive or negative. Moreover, net transfers depend on the relative performances of sectors to the extent that sectoral exports structures differ across countries. Generally, cumulative transfers are close to zero at the end of the sample period or they stabilize at relatively low levels. Only for the Netherlands, Finland, and Malta do cumulative transfers in absolute value end above 1 percent of GDP. Moreover, the only country for which cumulative transfers are monotonic over our sample period is Finland. Obviously, even under a pure insurance scheme ex post cumulative transfers are likely

Figure 5. Autoregressive Coefficients of Sectoral Shocks

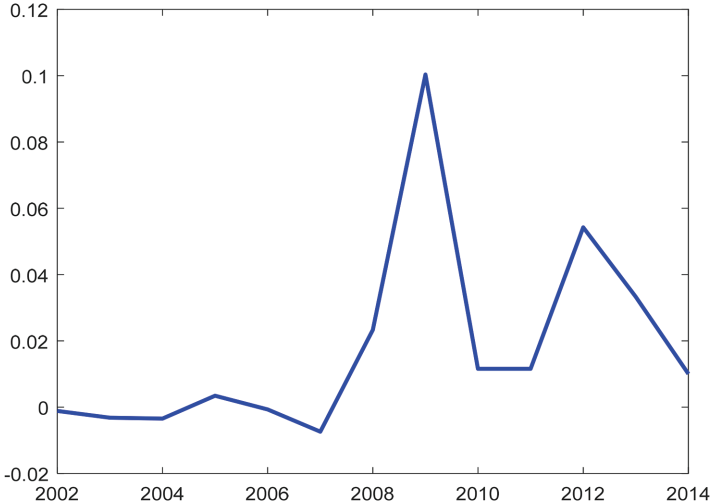


Notes: The figure reports the coefficient of a regression of sectoral shocks on their first lag. The regression is done for each sector separately. The lightest bar reports the average of the estimated regression coefficients. The bar with horizontal lines in the figure reports the AR(1) coefficient of a (year, country) panel regression of the output gap on its first lag.

to differ from zero at the end of the sample period, because the specific shock materializations are unlikely to be fully symmetric at the individual country level.

Cumulated transfers are driven by the degree of persistence of the shocks. Therefore, we formally test this persistence with a panel regression for each sector, in which we regress the sectoral shocks on a country fixed effect and their own lag, again clustering standard observations by country. Figure 5 depicts the coefficient estimates by sector. The coefficient estimate is always significantly positive. Hence, the sectoral shocks exhibit persistence. Indeed, persistence is a necessary condition to provide enough countercyclical force,

Figure 6. Sum of Transfers across Countries, Each Year, as Percent of Euro-Area GDP



Note: The figure shows the sum of transfers across countries, each year, as percent of euro-area GDP in that year.

because output gaps are also persistent reflecting the duration of business cycles.[17] However, the coefficient estimates are also always significantly smaller than one, and generally substantially so, indicating that shocks die out within a reasonable amount of time. Ideally, to exert maximal countercyclical effect, the sectoral shocks and thus the transfers exhibit an amount of persistence that reflects a business cycle downturn (or upturn). Indeed, the average persistence of the sectoral shocks turns out to be only somewhat smaller than that of the output gap.

As explained earlier, the scheme is generally not completely balanced at the aggregate level. Figure 6 depicts the sum of the transfer payments each year, across all countries. The scheme is in practice close to balance before the 2008–09 crisis. The imbalance is largest during the global financial crisis in 2009 and the widespread recession in 2012–13 associated with the debt crisis that forced governments to

[17] A panel regression of the output gap on its first lag yields a coefficient of 0.73, which is highly significantly different from zero.

consolidate in order to retain capital market access. These are precisely the sample years of widespread negative growth, hence during which the transfer scheme is at its most useful. The average annual imbalance amounts to only 0.02 percent of GDP.

5.2 Transfers Based on Compensation of Loss of Labor Income

This subsection explores transfers that are intended to compensate for the loss of labor income. Conceptually, this would be the more natural scheme to consider when there is no home bias in asset holdings in companies, i.e., all stakes in equity, corporate bonds, and other corporate financing vehicles are perfectly evenly spread over the entire euro zone. Transfers are now governed by expression (7′). The transfers generated under this scheme are qualitatively very similar to the baseline transfers: they are largely a scaled-down version of the transfers in the previous subsection. Hence, for the sake of space, we do not show the charts for each individual country, but only report the results in the summary charts and tables.[18] Figure 3B summarizes the information in the individual country figures by showing for each year the dispersion in the cross-country transfers' distribution. As expected, the dispersion is smaller than under the baseline. As before, the coefficient estimates in table 2 reveal countercyclicality of the transfers. Finally, table 3 reports a correlation coefficient between the new transfers and the baseline transfers of 0.93.

5.3 Transfers Based on Tax Revenues

As we argued earlier, a country that is required to pay a transfer because of a positive shock in the main export market(s) has already spent part of the resources generated by the additional activity in the form of wages to the workers and compensation for the capital hired to produce the extra output. Hence, these resources are not readily available to the government. However, the government obtains additional tax revenues because of the taxes paid on the generated additional income. In this subsection, we assume that the

[18]Results for individual countries are available upon request.

Table 3. Correlation between Annual Transfers Generated by Different Schemes

	Baseline	Labor Share	Taxes	Reaggregation	Leaving Out One Country
Baseline	1.00	0.93	1.00	0.93	0.99
Labor Share	0.93	1.00	0.93	0.84	0.92
Taxes	1.00	0.93	1.00	0.93	0.99
Reaggregation	0.93	0.84	0.93	1.00	0.92
Leaving Out One Country	0.99	0.92	0.99	0.92	1.00

Notes: The table reports the correlation coefficients between transfers of each pair of schemes. The numbers are the average of these pairwise correlations, across all 19 countries.

transfers are based on the change in tax revenues, i.e., they are calculated using expression (12). Figure 3C shows the dispersion of the transfers by year. Obviously, the dispersion is smaller than under the baseline. As reported in table 3, the new transfers are almost perfectly correlated with the baseline transfers.[19] This is also borne out by the regressions in table 2, which continue to show a high degree of countercyclicality of the transfers under this alternative scheme.

5.4 Summary

The preceding discussion of the different transfer schemes warrants a number of conclusions. First, in general, transfers are (strongly) countercyclical. Second, cumulative transfers tend to stabilize over time or end at values not too far from zero. Third, the annual transfers are generally of limited size, but reach in some instances quite large values, suggesting that they can exert quite a strong countercyclical force, in particular during highly adverse economic circumstances. That is, they reach relatively large values when they matter most. In absolute magnitude, transfers tend to be larger for

[19]The correlation is not entirely perfect, because the applied common tax rate varies slightly over time.

the smaller economies than for the larger economies. The rather limited *average* size of the transfers should be conducive to the political feasibility of the scheme. Starting on a small scale also seems to have been the strategy of the European Commission when in 2018 it proposed the EISF to be included in the new EU multiannual financial framework 2021–27.

6. Robustness

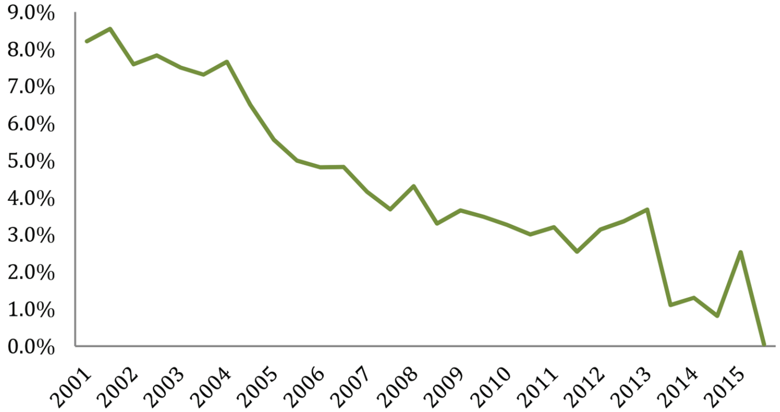
This section explores the robustness of the baseline results to revisions in the underlying data, the country composition of the panel, and a reaggregation of the sectors.

6.1 Data Revisions

Data on exports are subject to revisions over time, as better information becomes available and definitions and computation procedures change. As our transfer scheme would have to make use of real-time data for its implementation, it is important to assess whether it is sufficiently robust to data revisions.

The lagged variables entering equation (7) are likely to be relatively stable, as they have been already subject to one round of revision. Moreover, the shares x_{it}/x_t are generally quite robust to revisions, as both the numerator and the denominator tend to be revised in the same direction and with comparable proportions. Inspection of the data for subsequent vintages shows that large parts of the data revisions are common across all the countries in the sample. This is not surprising, as changes in the common methodology of constructing figures, for example, will apply to all the sample countries. However, in particular the data for x_{jt} may be affected by revisions. Does this significantly affect the level of annual transfers? While real-time data for exports *by sector* are not available in existing data sets, the AMECO database reports real-time values for *total* exports in each country. Figure 7 shows the simple average of revisions for this variable. Revisions are calculated as the difference between the "ex post" values, as published in the winter 2017 edition of the database, and the real-time estimates. Revisions are reported

Figure 7. Average Absolute Value of Revisions (in % of real-time value)



Notes: The figure depicts the average (across countries) percent difference of the total export value-added from the winter 2017 edition of the AMECO database and the real-time estimate of the year indicated on the horizontal axis.

in absolute value as a percentage of the real-time estimate.[20] The figure shows that revisions generally increase with the time passed since the first publication of the figures. They are close to 9 percent at their maximum and around half of this on average over the whole sample period.

We use this information to conduct a counterfactual experiment. In particular, based on the findings about the size of the revisions of total exports, we assume that sectoral exports are affected by randomly drawn annual revisions that are uncorrelated across countries, sectors, and years. The revisions are drawn from a uniform distribution and range between −10 percent and +10 percent of the "true" data (with an average absolute magnitude of 5 percent, therefore broadly comparable with the average revisions reported in

[20]Data for Malta and Cyprus are excluded from the average because the revisions are extremely large. The small size of the two economies justifies this choice of excluding them in order to avoid biases.

figure 7).[21] Concretely, we generate artificial samples in which each actual value x_{ijt} (export by country i, sector j, year t) reported in our ex post data is multiplied by a different random number extracted from a uniform distribution between 0.9 and 1.1 around one. The resulting randomly generated $\hat{x}_{ijt}$ is then used to compute all the variables needed for the transfers in (7).

Figure 8 shows the annual transfers implied by our baseline scheme based on the actual data (the dark solid line) together with the transfers based on 1,000 artificial data sets generated randomly as explained above. Shaded areas represent the 90 percent confidence region of the distribution of simulated transfers. It can be seen how, regardless of the artificially constructed revisions, the simulated annual transfers exhibit the same qualitative pattern as that of the "actual" transfers. This is also evident for the cumulated transfers (the figure is omitted for space reasons, but available upon request), suggesting that the scheme maintains the same features and remains countercyclical. Indeed, repeating the previous regressions with the simulated series yields medians of the coefficient estimates that are all negative, thus confirming the countercyclicality of the simulated transfers.

6.2 Leaving Out One Country at a Time

The countries in our transfer scheme vary substantially in size, while they are also quite different in terms of sectoral structure. In this robustness check, we explore whether the baseline transfers are affected by leaving out individual countries. In particular, we calculate the transfers to and from the other countries in the system, while we leave out one country at a time. The remaining 18 countries form a "closed system" with transfers calculated exactly as under the baseline with just one country less. Hence, all the shares are expressed in terms of the total for the euro area minus that country, while the transfers among the remaining countries (almost) add up to zero each year. Since there are 19 countries, we do this 19 times. Deviations from the baseline transfer patterns could be expected if

[21] If anything, assuming that all these correlations are zero stacks the odds of this experiment against us, since the noise introduced in the fluctuations in sectoral exports of a country relative to other countries as a result of the revisions is likely larger than in reality.

Figure 8. Baseline Annual Transfers under Actual and Simulated Revisions Data in Percent of the Country's GDP

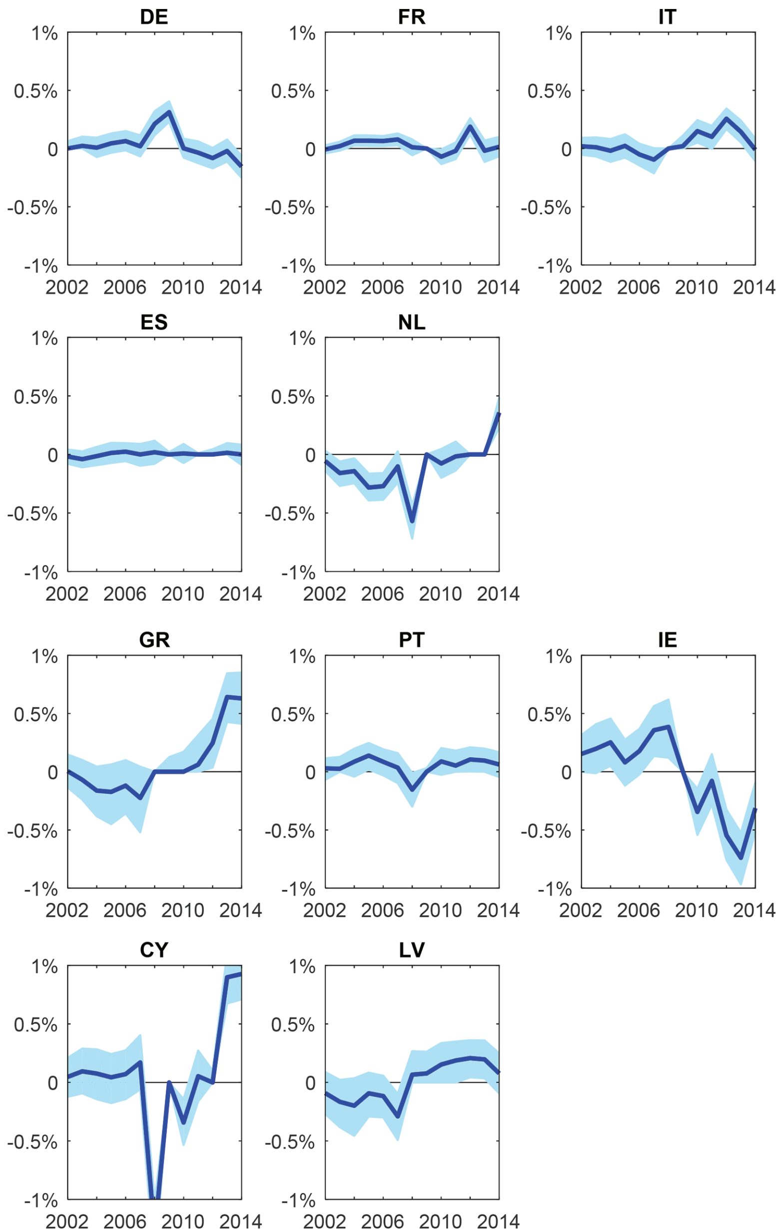


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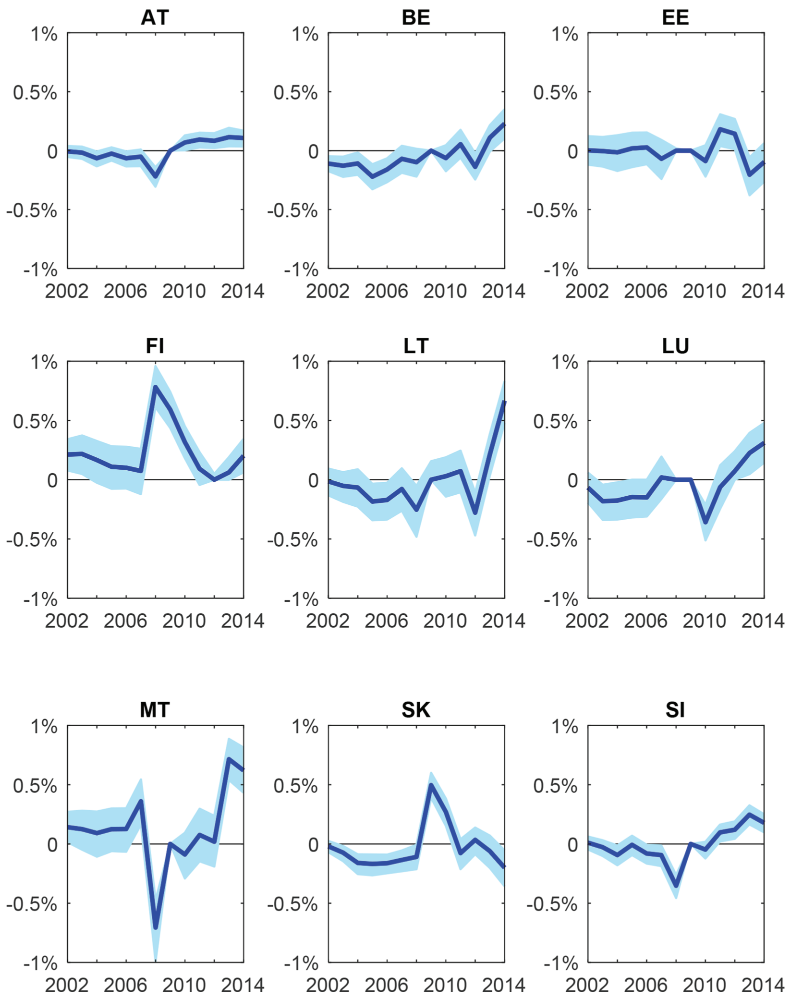
Figure 8. (Continued)



Notes: The figure reports the annual transfers calculated for 1,000 simulated series using the baseline scheme in equation (9). Transfers are in percent of GDP. The simulations are based on the assumption that each value of x_{ijt} is subject to a random revision drawn from the uniform density ranging from −10 percent to +10 percent around the actual value in the data. Shaded areas represent the 90 percent confidence bands. For country codes, see the notes to table 1.

Table 4. Reaggregation of Original Sectors

Original Aggregation	New Aggregation
Agriculture, Hunting, Forestry, and Fishing	Agriculture, Hunting, Forestry, and Fishing
Mining and Quarrying	Mining and Quarrying
Food Products, Beverages, and Tobacco	Food Products, Beverages, and Tobacco
Textiles, Textile Products, Leather, and Footwear	Textiles, Textile Products, Leather, and Footwear
Wood and Products of Wood and Cork	Wood, Paper, Paper Products, Printing, and Publishing
Pulp, Paper, Paper Products, Printing, and Publishing	
Coke, Refined Petroleum Products, and Nuclear Fuel	Chemicals and Non-metallic Mineral Products
Chemicals and Chemical Products	
Rubber and Plastics Products	
Other Non-metallic Mineral Products	
Basic Metals	Basic Metals and Fabricated Metal Products
Fabricated Metal Products except Machinery and Equipment	
Machinery and Equipment n.e.c.	Machinery and Equipment n.e.c.
Computer, Electronic, and Optical Products	Electrical and Optical Equipment
Electrical Machinery and Apparatus n.e.c.	
Motor Vehicles, Trailers, and Semi-Trailers	Transport Equipment
Other Transport Equipment	
Manufacturing n.e.c.; Recycling	Manufacturing n.e.c.; Recycling
Electricity, Gas, and Water Supply	Electricity, Gas, and Water Supply
Construction	Construction
Wholesale and Retail Trade; Repairs	Wholesale and Retail Trade; Hotels and Restaurants
Hotels and Restaurants	

(*continued*)

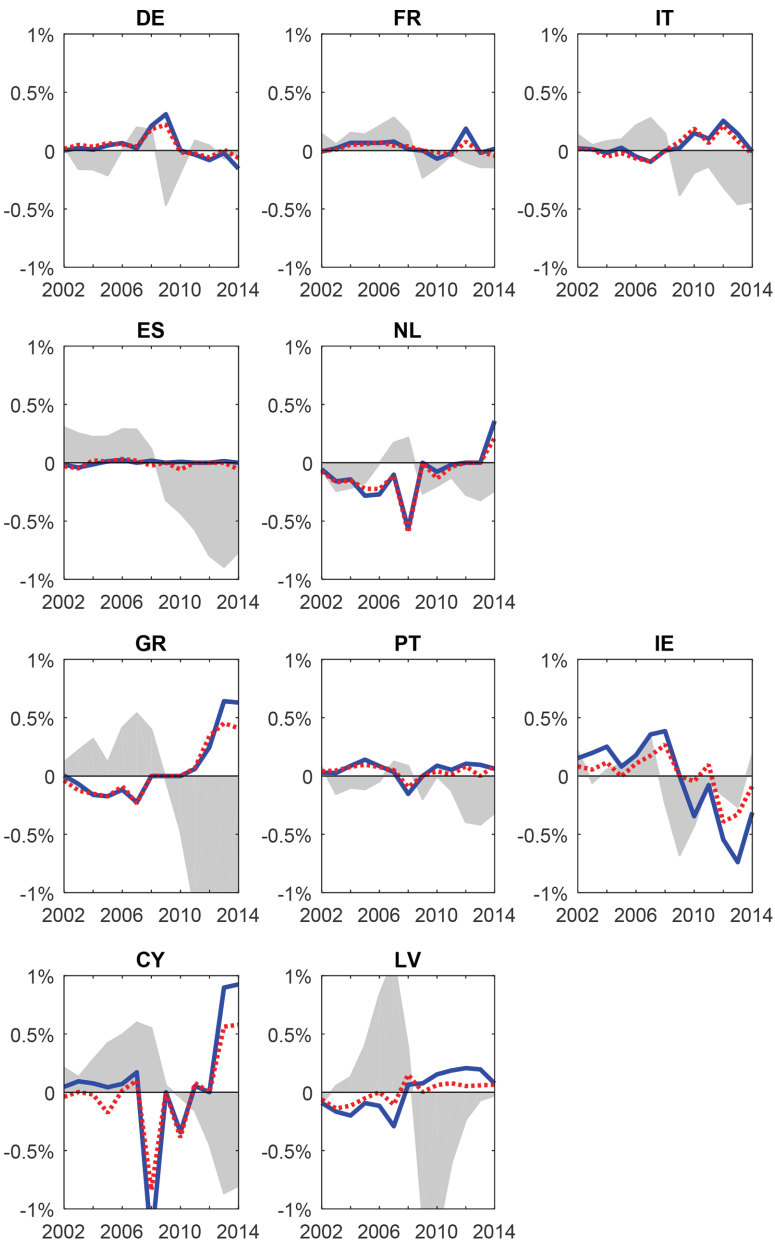
Table 4. (Continued)

Original Aggregation	New Aggregation
Transport and Storage	Transport and Storage, Post and Telecommunication
Post and Telecommunications	
Real Estate Activities	Real Estate, Renting, and Business Activities
Renting of Machinery and Equipment	
Computer and Related Activities	
Research and Development and Other Business Activities	
Public Admin. and Defense; Compulsory Social Security	Community, Social, and Personal Services
Education	
Health and Social Work	
Other Community, Social, and Personal Services	

Notes: The original aggregation is based on the OECD 33 industry list (TiVA 2016); the new aggregation is based on the OECD TiVA 2013 classification. The sector "Finance and Insurance" has been removed from the original data set. The sector is generally tightly regulated by national authorities and often faces specific tax treatment. Moreover, its disproportionate presence in some relatively small countries would result in very large cross-border transfer payments (or receipts) when shocks hit that sector.

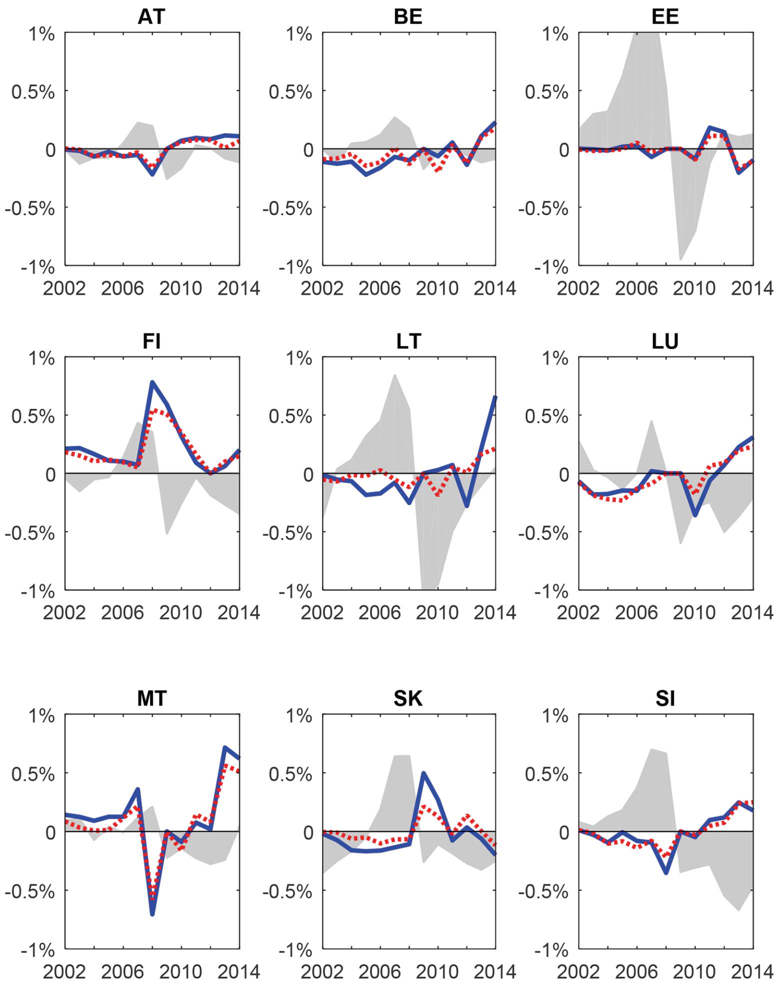
individual countries dominate specific export sectors and those sectors account for a substantial fraction of the overall transfers flowing to or from individual countries. However, for each of the 19 sample countries, the 18 plots of the transfers (each one corresponding to leaving out one of the other countries) coincide to a very substantial extent. To save space, we do not report the figures here. However, they are available upon request. The robustness of the transfers for leaving out individual countries is also clear from table 3. The average correlation with the baseline transfers over all (year, country) combinations and averaged over all 19 cases of dropping one country at a time is 0.99.

Figure 9. Annual and Cumulated Transfers under Baseline Scheme for Reaggregated Sectors in Percent of the Country's GDP



(*continued*)

Figure 9. (Continued)



Notes: The figure shows the annual transfers for each country, generated by the scheme where we include 17 reaggregated sectors, as in table 4 (dotted line), together with the baseline transfers as of figure 2 (solid line). Transfers are in percent of a country's GDP in that year. The gray shaded areas are output gaps published by the OECD (2017a). For all countries, we report data in the scale +1 to −1 percent of GDP, for consistency. In the case of the Netherlands, the cumulated transfers are at around 1 percent of GDP in 2014. For Finland, they are about 2.5 percent of GDP in 2014. For country codes, see the notes to table 1.

6.3 Sectoral Reaggregation

In this subsection, we investigate the robustness of our findings to reaggregating the exports sectors. In particular, we impose a higher level of aggregation by merging related sectors into a new set of sectors and recalculating transfers analogously to how we calculate them under the baseline. Table 4 lists the aggregation of the original 32 sectors into the new, smaller set of 17 sectors. Figure 9 depicts the transfers under the new sectoral aggregation. They remain very similar to the baseline transfers: the correlation of the cross-border transfers generated by this scheme with the baseline transfers is 0.93 (see table 3). If anything, the variability of transfers within each country (see, e.g., Ireland and Cyprus) and the cross-country dispersion decreases somewhat. This is the consequence of the fact that we now have fewer sectors, therefore fewer shocks which are also more smoothed due to an averaging effect. Finally, the regressions reported in table 2 confirm that the newly computed transfers remain strongly countercyclical.

7. Concluding Remarks and Discussion

Asymmetries in shocks and transmission mechanisms are major obstacles to the proper functioning of a monetary union. The current range of possibilities to deal with such shocks in the EMU is rather limited. Cross-border labor mobility is low, although it may increase in the future as European economic integration proceeds and national institutions become more alike. Risk sharing at the private level through capital markets is rather limited too, although again one might expect this channel to become more important as impediments to the cross-border trade of assets are reduced and the capital market union draws to a completion. Finally, the use of fiscal policy is restricted because of the rules imposed through the Stability and Growth Pact. The latter calls for a government budget that is close to balance or in surplus in the medium run, to enable automatic stabilizers to do their work. However, reaching a situation in which all the EMU participants have eliminated their structural deficits will be difficult.

Because the cross-border sharing of asymmetric shocks will remain limited for the foreseeable future, we analyze the adoption of a cross-border transfer scheme. Obviously, a major source of (political) resistance to such a scheme is the potential for moral hazard. However, we propose a scheme that goes a substantial way toward avoiding moral hazard by conditioning transfers on (exogenous) world market developments for the relevant sectors in the euro area. Our scheme has other advantages as well: transfers are imposed to (almost) add up to zero on an annual basis and they are based on deviations of exports from sectoral trends, implying that they can only go into the same direction for a limited amount of time. In particular, a permanent reduction in a sector's exports cannot lead to permanent transfers. Our baseline scheme aimed at equalizing income shifts as a fraction of exports yields highly countercyclical transfers. Moreover, the cumulative transfers return toward zero over time or they end the sample period at a low level. These findings are robust for different variants of our transfer scheme and in particular also for the case in which we allow for revisions in the real-time figures of the sectoral export value-added.

Of course, before our scheme can be made fully operational, practical obstacles would need to be overcome. Although we have demonstrated the robustness of our scheme to data revisions, we still view the timely availability of the data that serve as input for the calculation of the transfers as the main practical obstacle. This is in particular the case for data on sectoral activity. However, when sufficient practical need is perceived for the timely availability of such data, governments and statistical agencies may invest more resources in achieving this objective.

Another issue concerns the question of how transfers received by governments should be put to best use. Because a transfer receipt comes on top of regular resource flows, it could be politically easier to earmark it for ameliorating the consequences of structural reforms or help in transforming the economy toward activities with a more prosperous future. In general, however, the concern that our scheme may delay the transition to more productive sectors is mitigated by the fact that transfers are *on average* relatively small (apart from when they are most needed, i.e., in recessions). Therefore, countries would continue to be confronted with the cost (in terms of higher unemployment, less tax receipts, etc.) of delaying the restructuring of declining sectors.

A final issue concerns the possibility that some countries may benefit more from the stabilizing effects of the transfer scheme than other countries. Our baseline scheme limits the differences in the average magnitudes of the transfers across the countries. Still, to enhance the political acceptability of the scheme further, one could, for example, envisage that countries that benefit more from the sharing of asymmetric shocks pay an "insurance premium" that goes into a collective fund that can compensate countries that benefit less or that can absorb small temporary aggregate imbalances in the scheme. As an alternative, the transfer scheme could start with a subset of countries that are expected to experience roughly equal variance in their transfers. The fact that transfers add up to (almost) zero and that they are robust against dropping countries would facilitate this option. However, further investigation of these options is left for future research.

To investigate further whether our proposed scheme does not distort incentives, one could explore whether countries with more rigid labor and product markets would on average receive more transfers. Thorough investigation of this issue would require comparable panel data on market rigidities. It would also require a sufficiently long sample period. Such an analysis is beyond the scope of the present paper but would be an interesting avenue for further research.

Appendix A. Derivation of (6)

We can rewrite (2) as

$$w_{kj,t-1}\left(x_{jt}-x_{jt}^{*}\right)+T_{kjt}=\frac{X_{k,t-1}}{X_{i,t-1}}\left[w_{ij,t-1}\left(x_{jt}-x_{jt}^{*}\right)+T_{ijt}\right]\Leftrightarrow$$

$$T_{kjt}=\frac{X_{k,t-1}}{X_{i,t-1}}\left[w_{ij,t-1}\left(x_{jt}-x_{jt}^{*}\right)+T_{ijt}\right]-w_{kj,t-1}\left(x_{jt}-x_{jt}^{*}\right)\Leftrightarrow$$

$$T_{kjt}=\left[\frac{X_{k,t-1}}{X_{i,t-1}}w_{ij,t-1}-w_{kj,t-1}\right]\left(x_{jt}-x_{jt}^{*}\right)+\frac{X_{k,t-1}}{X_{i,t-1}}T_{ijt}.$$

Hence,

$$\sum_{k\neq i}T_{kjt}=\sum_{k\neq i}\left[\frac{X_{k,t-1}}{X_{i,t-1}}w_{ij,t-1}-w_{kj,t-1}\right]\left(x_{jt}-x_{jt}^{*}\right)+\sum_{k\neq i}\frac{X_{k,t-1}}{X_{i,t-1}}T_{ijt}.$$

Using that transfers add up to zero:

$$-T_{ijt} = \sum_{k \neq i} \left[\frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} - w_{kj,t-1} \right] \left(x_{jt} - x^*_{jt} \right) + \sum_{k \neq i} \frac{X_{k,t-1}}{X_{i,t-1}} T_{ijt} \Leftrightarrow$$

$$\left[1 + \sum_{k \neq i} \frac{X_{k,t-1}}{X_{i,t-1}} \right] T_{ijt} = \sum_{k \neq i} \left[w_{kj,t-1} - \frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} \right] \left(x_{jt} - x^*_{jt} \right) \Leftrightarrow$$

$$\frac{X_{t-1}}{X_{i,t-1}} T_{ijt} = \left[1 - w_{ij,t-1} - \frac{X_{t-1} - X_{i,t-1}}{X_{i,t-1}} w_{ij,t-1} \right] \left(x_{jt} - x^*_{jt} \right) \Leftrightarrow$$

$$T_{ijt} = \left[\frac{X_{i,t-1}}{X_{t-1}} - w_{ij,t-1} \right] \left(x_{jt} - x^*_{jt} \right).$$

Appendix B. Derivation of (11)

Start from (10) and rewrite:

$$T_{kjt} = \frac{X_{k,t-1}}{X_{i,t-1}} \left\{ \left[w_{ij,t-1} lsh_{ijt} - \frac{X_{i,t-1}}{X_{k,t-1}} w_{kj,t-1} lsh_{kjt} \right] \times \left(x_{jt} - x^*_{jt} \right) + T_{ijt} \right\}$$

$$= \left[\frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} lsh_{ijt} - w_{kj,t-1} lsh_{kjt} \right] \left(x_{jt} - x^*_{jt} \right) + \frac{X_{k,t-1}}{X_{i,t-1}} T_{ijt}, \forall k \neq i.$$

Hence,

$$\sum_{k \neq i} T_{kjt} = \sum_{k \neq i} \left[\frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} lsh_{ijt} - w_{kj,t-1} lsh_{kjt} \right] \left(x_{jt} - x^*_{jt} \right) + \sum_{k \neq i} \frac{X_{k,t-1}}{X_{i,t-1}} T_{ijt}.$$

Using that transfers add up to zero:

$$-T_{ijt} = \sum_{k \neq i} \left[\frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} lsh_{ijt} - w_{kj,t-1} lsh_{kjt} \right] \left(x_{jt} - x_{jt}^* \right)$$

$$+ \sum_{k \neq i} \frac{X_{k,t-1}}{X_{i,t-1}} T_{ijt} \Leftrightarrow$$

$$\left[1 + \sum_{k \neq i} \frac{X_{k,t-1}}{X_{i,t-1}} \right] T_{ijt} = \sum_{k \neq i} \left[w_{kj,t-1} lsh_{kjt} - \frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} lsh_{ijt} \right]$$

$$\times \left(x_{jt} - x_{jt}^* \right) \Leftrightarrow$$

$$\frac{X_{t-1}}{X_{i,t-1}} T_{ijt} = \sum_{k \neq i} \left[w_{kj,t-1} lsh_{kjt} - \frac{X_{k,t-1}}{X_{i,t-1}} w_{ij,t-1} lsh_{ijt} \right] \left(x_{jt} - x_{jt}^* \right) \Leftrightarrow$$

$$T_{ijt} = \sum_{k \neq i} \left[\frac{X_{i,t-1}}{X_{t-1}} w_{kj,t-1} lsh_{kjt} - \frac{X_{k,t-1}}{X_{t-1}} w_{ij,t-1} lsh_{ijt} \right] \left(x_{jt} - x_{jt}^* \right).$$

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The Determinants of European Banks' Capital Structure: Is There a Difference between Public and Private Banks?*

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In this paper we examine whether the determinants of European banks' capital structure depend on the type of the institution (private or public). Using an international sample of 586 banks from 21 European countries for the period of 2000 to 2016, we find that when compared with private banks, the determinants of public banks' capital structure are more closely aligned with those that affect nonfinancial firms. Furthermore, this paper provides evidence that these differences can result in certain consequences with regards to the access to the market, which implies that public banks are more subject to market discipline. This is a very topical question, particularly when studied in the context of the introduction of more demanding capital requirements through Basel III and, in particular, in the context of the new resolution regime which imposes additional capital requirements, where banks' access to equity and debt markets plays a pivotal role.

JEL Codes: G21; G32.

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1. Introduction

According to modern financial literature, the study of capital structure began with Modigliani and Miller (1958), who stated that in a frictionless world with full information and complete markets, the value of firms is independent of their capital structure (Santos 2001), leaving room for further research regarding the impact of disregarding the "frictionless assumption," i.e., by adding taxes, costs of financial distress, imperfections in the product market, transactions costs, asymmetry of information, and agency costs.

The study of the above-mentioned frictions has originated several theories, such as the trade-off theory, the pecking order theory, and the market timing theory (Flannery and Rangan 2006).

The great majority of the published empirical evidence over the last decades which is dedicated to the topic of capital structure was developed for nonfinancial firms. For instance, Rajan and Zingales (1995) justify the exclusion of financial firms from their sample, because their leverage would be strongly influenced by explicit (or implicit) investor insurance schemes, such as deposit insurance. Furthermore, regulation such as minimum capital requirements can directly affect the capital structure of financial firms.

Nevertheless, more recent empirical research regarding bank capital structure has contributed evidence which supports that the attributes that affect bank capital structure are not very different from those that influence the capital structure of nonfinancial firms. Studies by Flannery (1994), Flannery and Rangan (2008), and Allen, Carletti, and Marquez (2011) have shown that market discipline (carried out by subordinated creditors, or even by depositors) has played an important role in banks' capital structure. By the same token, other empirical papers published on this matter, such as Barber and Lyon (1997), Brewer Iii, Kaufman, and Wall (2008), Gropp and Heider (2010), and De Jonghe and Öztekin (2015), have found that bank-specific indicators also explain, to a large extent, bank capital structure, i.e., most banks seem to optimize their capital structure in much the same way as firms, except when their capital ratios are close to the regulatory minimum.

Nevertheless, most of the empirical studies developed so far regarding banks' capital structure are focused on large, public banks (Gropp and Heider 2010) and also on the main determinants of the speed of adjustment toward the banks' target capital ratio (De Jonghe and Öztekin 2015), tending to omit an analysis of the relationship between the type of banks and their behavior regarding capital structure decisions. It is exactly this gap that we aim to fill with this paper.

Accordingly, we attempt to empirically answer the following question: Are the determinants of banks' capital structure different in public versus private banks? For the purpose of this paper, public banks are considered to be those that are quoted/listed in capital markets, whereas private banks are those that are not quoted/listed in capital markets.

The answer to this question is relevant from an empirical as well as a policy point of view. From an empirical point of view, it is important to challenge the conclusions drawn from the above-mentioned studies, which use samples that comprise only large and public banks—which are subject to a different intensity of market discipline and/or use a pooled data set of banks, extrapolating the results for the whole sample, which in turn could hide some non-negligible differences regarding the determinants of the capital structure of public and private banks.

From a policy point of view, when examined in the context of the requirements for additional capital, which has led to banks increasingly having to resort to own funds through the issue of, among others, equity or debt instruments with certain specific characteristics (e.g., subordinated debt) where the access to the market plays a pivotal role, it is important to investigate how the determinants of banks' capital structure have influenced banks' access to the market (which is represented by the share of subordinated debt for total assets).[1]

Using an international sample of 586 banks from 21 European countries for the period of 2000 to 2016, our results for the

[1] Beyond the new capital requirements encompassing the new capital buffers—namely the conservation capital buffer, countercyclical capital buffer, and systemically important institutions buffer (introduced with the Basel III framework)—the Directive on Banking Recovery and Resolution that entered into force in 2016 implies banks' compliance with the additional capital requirements so-called minimum requirement for own funds and eligible liabilities (MREL) which should equal 2 × (pillar 1 requirements + pillar 2 requirements and the combined buffer requirement) – 125 basis points.

whole sample confirm, to some extent, those of Gropp and Heider (2010), where size is positively related with leverage, whereas profits, market-to-book value, and risk all have a negative impact on leverage, which is broadly consistent with the pecking order and market timing theories.

Furthermore, considering the argument that banks manage their capital or leverage ratios based on achieving a target, we show that the speed of adjustment is material in banks, which enables these institutions to converge toward their long-run target.

Interestingly, our results have identified important differences between public and private banks. Whereas the determinants of the capital structure of public banks present a set of similarities with nonfinancial firms (which is in line with the results presented by Gropp and Heider 2010), for private banks we find that those determinants which are typical of the market timing and pecking order theories fail to evidence the same relevance as in the case of public banks.

Besides the higher asymmetry of information which characterizes private firms, including financial ones, this paper provides empirical evidence that those banks whose capital structure has been driven by the determinants envisaged in the literature (public banks) have been subjected to more market discipline (which is represented by the share of funding through subordinated debt).

The consequences of the above-mentioned differences observed between public and private banks for both the empirical literature and financial stability are twofold. From an empirical point of view, this is the first study which shows that "one size does not fit all," i.e., that the determinants of banks' capital structure vary between private and public banks. From a financial stability point of view, during an initial phase, the differences observed in this paper could negatively affect the access of private banks to capital and debt markets and thus compound the difficulties in complying with the more demanding capital requirements envisaged in Basel III, as well as in the new resolution regime—which implies the issuance of bail-in-able instruments. However, in the long term, it is expected that the implications of the new resolution regime (such as the issuance of MREL, composed of bail-in-able instruments, such as subordinated debt) will contribute to broadening market discipline to all types of banks and to the alignment of the determinants of their capital structure. Additionally, the supervisors and regulators may well be able to reap the benefits of a wider market discipline—which is detailed below in this paper.

The remainder of the paper is organized as follows: Section 2 presents a literature review, which is focused on capital structure determinants (in general for nonfinancial firms, and specifically for banks). Section 3 develops hypotheses about how the determinants of capital structure in nonfinancial firms play a role in the case of banks. Section 4 introduces the data set, including descriptive statistics for bank-specific indicators and a cross-country analysis. The methodological approach is discussed in section 5. Section 6 reports the results and robustness tests, and the main conclusions are presented in section 7.

2. Literature Review

2.1 The Determinants of Capital Structure

Modern capital structure theory began with Modigliani and Miller (1958), advocating that the value of firms is independent of its capital structure in a frictionless world with full information and complete markets (Santos 2001), leaving room to exploit certain frictions.

To this end, Berger, Herring, and Szegö (1995) summarize that the main frictions that cause capital decisions to depart from Modigliani and Miller (1958) are (i) taxes—as interest payments are tax deductible, whereas dividends are not, and that by substituting debt for equity, firms are able to pass on larger returns to investors by reducing payments to the government, and, therefore, all else being equal, shareholders prefer to fund firms almost entirely with debt; (ii) financial distress—where more leverage can lead to an increase in the likelihood of incurring the costs of financial distress; (iii) asymmetric information—which stems from the fact that managers generally have access to more information about their own earnings prospects and financial condition than the market; (iv) transactions costs of new issues—which when combined with asymmetric information can also influence the relative costs of internal versus external finance, and the relative costs of debt versus equity, and; (v) agency costs between shareholders and creditors or between the former and managers (Jensen and Meckling 1976).

A sizable amount of research has been carried out since then, which can be summarized into three main theories (Flannery and Rangan 2006): the trade-off theory, the pecking order theory, and the market timing theory. According to the trade-off theory, firms select target debt-equity ratios to trade off the costs and benefits of leverage, contradicting the unrealistic assumption that firms are always at their equilibrium leverage (Öztekin and Flannery 2012). The benefits of leverage include the tax deductibility of the debt service (the well-known "interest tax shield") and the agency benefits of debt associated with conflicts of interest between managers and stockholders. The costs of debt can be identified as costs of bankruptcy or financial distress, agency costs due to misalignment of interests between stockholders and creditors, and trading costs (according to Öztekin 2015, if a country's characteristics make the issuance of debt and equity expensive, firms tend to exhibit slower adjustment speeds). The pecking order theory, which was first suggested by Myers and Majluf (1984), is based on the assumption that information asymmetries between insiders and outsiders can lead managers to perceive that the uninformed market would generally underprice their firm's equity—and for this reason, managers have a preference for investments to be financed first with internal funds, second with secured debt, and lastly with equity—which is only used as a last resort. This order of preference between internal versus external funds, and between debt versus equity, is a justification for firms to maintain a certain financial slack, particularly those firms that operate in industries which are particularly opaque and are subject to asymmetry of information between insiders and outsiders. The market timing theory argues that a firm's leverage reflects its cumulative ability to sell overpriced equity shares, i.e., share prices fluctuate around their "true" value, and managers tend to issue shares when the firm's market-to-book value is high. Therefore, by exploiting asymmetric information, firms increase the wealth of their current shareholders.

Recent research on capital structure has been developed around these three competing, but not necessarily mutually exclusive, theories. Many empirical studies have tried to identify the correct measures for characterizing the attributes present in these theories.

Amongst the main studies of interest in this context, we highlight Titman and Wessels (1988), Rajan and Zingales (1995), Flannery and Rangan (2006), Lemmon, Roberts, and Zender (2008), and Frank and Goyal (2009). The measures used in these papers comprise (i) collateral, which is defined as being the sum of liquid assets and fixed assets; (ii) firm size, which is represented by total assets; (iii) market-to-book ratio, which is defined as being the ratio between the market value of assets and the book value of total assets; (iv) firm risk, which is usually computed as the standard deviation of a firm's market stock returns; and (v) profits. Collateral and firm asset risk can be included in the trade-off theory, whereas profits and market-to-book value are more easily aligned with the pecking order theory and market timing, respectively. To this end, Harris and Raviv (1991) showed that leverage increases with fixed assets (collateral), as well as with nondebt tax shields, growth opportunities, and firm size, and that it decreases with volatility, advertising expenditures, bankruptcy probability, and the uniqueness of the products.

With regards to the relationship between the above-mentioned variables and firms' capital structure, Rajan and Zingales (1995) concluded that (i) the relationship between the ratio of fixed assets to total assets and leverage is positive—the rationale underlying this factor being that tangible assets are easy to collateralize and they thus reduce the agency costs of debt; (ii) leverage is negatively correlated with market-to-book value for two reasons: first, in theory, firms with high market-to-book ratios have higher costs of financial distress and, second, this negative correlation stems from firms' tendency to issue stock when their stock price is high in relation to earnings or book value (which thus supports the assumptions underlying the market timing theory); (iii) size is positively correlated with leverage, as, in theory, larger firms are better diversified and have a lower probability of being in financial distress; and (iv) profitability is negatively correlated with leverage (which supports the pecking order theory). To a large extent, these results are consistent with those obtained by Frank and Goyal (2009), who used a different sample of firms as well as a different time span.

Flannery and Rangan (2006), Lemmon, Roberts, and Zender (2008), Öztekin and Flannery (2012), and Öztekin (2015) all focused on testing firms' willingness to move toward their target leverage ratio, and found that firms converge toward their long-run targets at a rate of more than 30 percent per year, which depends, to a larger extent, on the institutional environment. An example is those countries which ease the issuance of debt and/or equity contribute to accelerate the speed of adjustment toward the optimal leverage ratio (Öztekin and Flannery 2012 and Öztekin 2015).

2.2 The Special Case of Banks

In many ways, banks are different from nonfinancial firms, which are able to induce differences in their capital structure decisions. Cocris and Ungureanu (2007) surmise that the main differences are that a high proportion of bank failures lead to negative externalities; agency problems are enhanced by the inefficient monitoring of banks by depositors and other stakeholders; information flow is complex, due to the opaque environment in which banks operate; banks are heavily regulated; sometimes the regulator itself is a bank stakeholder; the diversification of activities within a bank conglomerate intensifies agency problems between corporate insiders and small shareholders; in general, banks have a more concentrated equity ownership than nonfinancial firms, which makes it more difficult for small equity holders to exert influence over the management of banks; there is less competition regarding financial products and takeover activity; and banks have a safety net available, which affects the stakeholders' incentives to monitor banks. In addition, Harding, Liang, and Ross (2013) highlight the high levels of leverage which characterizes banks' capital structure, which was arguably responsible for the failure of the majority of investment banks during the recent global financial crisis.

To some extent, the identification of these special features of the banking sector are shared by Berger, Herring, and Szegö (1995) and Santos (2001), who both mention that the main frictions in the banking sector which can justify the departure from Modigliani and Miller (1958) are the following: the existence of a safety net—which is defined as all government actions which are devised to improve the safety and soundness of the banking system and thus shield banks' creditors (mainly depositors) from the full consequences of bank risk-taking; and the capital requirements stated by the regulators to protect themselves against the costs of financial distress, agency problems, the reduction in market discipline (caused by the safety net), and the systemic risk posed by the banking sector.

These characteristics explain why the majority of the empirical research on firm capital structure was carried out with nonfinancial firms (see, for example, Rajan and Zingales 1995).

Nevertheless, the empirical research regarding banks' capital structure has evidenced that those attributes that affect bank capital structure are not so far removed from the attributes which influence nonfinancial firms' capital structure. Thus, Modigliani and Miller's theory holds, within limits, for banks as well (Miles, Yang, and Marcheggiano 2013). Research carried out by Flannery (1994), Flannery and Sorescu (1996), Morgan and Stiroh (2001), Flannery and Rangan (2008), and Allen, Carletti, and Marquez (2011) has shown that market discipline demonstrated by subordinated creditors, and even by depositors (as in Martinez Peria and Schmukler 2001), has played an important role in explaining banks' capital structure.

Regarding the existence of similarity, or not, between the determinants of banks' capital structure and nonfinancial firms and whether the traditional capital determinants also hold for financial institutions, an empirical work carried out by Barber and Lyon (1997) and Brewer Iii, Kaufman, and Wall (2008) established that bank-specific indicators also contribute to explain banks' capital structure. Barber and Lyon (1997) concluded that the relationships between size, market-to-book value, and security returns are positive and similar for both financial and nonfinancial firms. Brewer Iii, Kaufman, and Wall (2008) found that leverage is positively and significantly correlated with banks' risk; however, the remaining variables, such as return on assets and size, are not significantly related to leverage. Nevertheless, bank-specific variables collectively explain banks' leverage, taking into account the result from the Wald test. With regards to speed of adjustment, the value attained was 12 percent.

The more recent studies on banks' capital structure were developed by Gropp and Heider (2010)—who expound on the influence of bank-specific indicators on banks' capital structure, comparing these with those evidenced by nonfinancial firms—and also De Jonghe and Öztekin (2015), whose focus was to investigate the adjustment process for targeting capital. Using a sample of large public banks from the United States and Europe (from 1991 to 2004), the first study evidences that regulation is not the main feature that causes the divergence of banks' capital structure from that which was argued by Modigliani and Miller (1958). For these authors, most banks seem to optimize their capital structure in much the same way as firms do, except when their capital ratios approach the levels of the regulatory minimum. Additionally, as demonstrated by Lemmon, Roberts, and Zander (2008), banks move toward their target leverage ratios at a speed of adjustment of 45 percent. This evidence contradicts the "regulatory view" of banks, whereby they should converge toward a common target, namely the minimum requirement set under Basel I, which gives support to the results obtained by Berger et al. (2008), who argue that banks actively manage their capital ratios.

The second study, that of De Jonghe and Öztekin (2015), using a sample of banks from 64 countries during the 1994–2010 period, based on the studies developed for nonfinancial firms by Öztekin and Flannery (2012) and Öztekin (2015), found that in the case of banks, the speed of adjustment is heterogeneous, depending on the institutional environment—which is consistent with the results obtained with nonfinancial firms. That is to say that the speed of adjustment toward the target capital ratio increases in those countries that have more stringent capital requirements, better supervisory monitoring, and more-developed capital markets—which in turn decreases the costs of debt and/or equity issuance. In addition, and consistent with those studies developed so far, this study found that smaller, more profitable, and cost-efficient banks have higher capital ratios.

3. Research Question and Hypothesis Development

This paper aims to empirically answer the following question: Are the determinants of banks' capital structure different in public versus private banks? Public banks are those that are quoted/listed in capital markets, whereas private banks are those that are not quoted/listed in capital markets.

On the one hand, we expect that the determinants of public banks' capital structure and their respective decisions regarding the optimal capital are different from those of private banks, due to the fact that the former has easier access to capital and debt markets, as a result of having less of the information asymmetry which characterizes public firms, whether they be financial or nonfinancial firms. On the other hand, considering that the banking sector is one of the most regulated in the world (Santos 2001), supplemented by the fact that all banks (public or private) are obliged to disclose information through pillar 3 of the regulatory framework, the determinants of public banks' capital structure, as identified by Gropp and Heider (2010), may well be not so different from those of private banks.

Accordingly, we test the hypothesis that the determinants of public banks' capital structure are different from the determinants of private banks against the hypothesis that the determinants of banks' capital structure are the same for both types of banks.

A detailed overview of the explanatory variables and their sources is provided in table 1. Table 2 presents the expected signs of the effects of bank-specific indicators on banks' capital structure, as well as a summary of the arguments that support such expectations.

4. Data

The data for this paper include banks which have their headquarters in 21 European countries, as well as subsidiaries of foreign banks (mainly from the United States), which add up to 586 banks, for the period of 2000–16. The data give rise to a panel of 6,065 bank-year observations. As mentioned in table 1, all data were collected from the Bankscope and SNL databases.

As shown in figure 1, the sample covers a high percentage of the European banking system's assets, which represents a share of 90 percent in 2012, decreasing slightly to 85 percent from 2012 onward—which is due to the change of the database from Bankscope to SNL.

We obtain an unbalanced panel data set, on account of a data gap and entry/exit in the sample. It is worth noting that the occurrence of certain mergers and acquisitions, together with the emergence of new banks, caused changes in the constitution of the sample during the period under analysis (table 3). It should be mentioned that these mergers and acquisitions events are included in banks' fixed effects.

Table 1. Variables: Description and Sources

Variables	Description	Source
BOOK VALUE OF ASSETS	The sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment.	Bankscope/SNL
BOOK VALUE OF EQUITY	The sum of the following items: "capital," "reserves," and "net income."	Bankscope/SNL
BOOK LEVERAGE (BL)	Computed as 1 – (book value of equity/book value of assets).	Bankscope/SNL
MARKET LEVERAGE (ML)	Calculated as 1 – [market value of equity (computed as number of outstanding shares*end of year stock price)/market value of bank (computed as the market value of equity + book value of liabilities)].	Bankscope/SNL
SIZE	The log of book value of assets.	Bankscope/SNL
PROFITABILITY (PROF)	Computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets.	Bankscope/SNL
MARKET-TO-BOOK RATIO (MBR)	Calculated as the ratio between market value of assets and book value of assets.	Bankscope/SNL
COLLATERAL (COLL)	Computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets.	Bankscope/SNL
RISK	Computed as the annualized standard deviation of monthly stock price returns (market*value of equity/market value of bank) or the log of the standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year.	Bloomberg/Bankscope/SNL
SUBORDINATED DEBT	A specific type of debt that ranks below the other types of debt such as deposits and other debt securities issued by an institution.	Bankscope/SNL
SUBORDINATED DEBT/ASSETS	Calculated as the ratio between total subordinated debt liabilities and total assets.	Bankscope/SNL

(*continued*)

Table 1. (Continued)

Variables	Description	Source
NONDEPOSITS/ASSETS	Computed as the ratio between nondeposits liabilities and total assets.	Bankscope/SNL
SECURITIES/ASSETS	The ratio between securities and total assets.	Bankscope/SNL
LOANS/ASSETS	The ratio between net loans and total assets.	
TOTAL CAPITAL RATIO	The ratio between own funds (as defined in Directive 2000/12/EC until 2006, in Directive 2006/49/EC between 2007 and 2013, and in Regulation 575/2013 from 2014 onwards) and risk-weighted assets.	Bankscope/SNL
TIER 1 CAPITAL RATIO	The ratio between own funds of tier 1 quality (as defined in Directive 2000/12/EC until 2006 and in Directive 2006/49/EC, in Directive 2006/49EC between 2007 and 2013, and in Regulation 575/2013 from 2014 onwards—essentially equity and reserves) and risk-weighted assets.	Bankscope/SNL
TIER 2 CAPITAL RATIO	The ratio between own funds of tier 2 quality (as defined in Directive 2000/12/EC until 2006, in Directive 2006/49/EC between 2007 and 2013, and in Regulation 575/2013 from 2014 onwards—essentially subordinated debt) and risk-weighted assets.	Bankscope/SNL
PUBLIC DUMMY	Dummy variable that takes the value of 1 if the bank is listed in market, and 0 otherwise.	Bankscope/SNL
SPEED OF ADJUSTMENT	Computed as 1 – coefficient on book leverage or market leverage lagged by one period.	Bankscope/SNL
CAPR	Dummy variable drawn from the study carried out by Cerutti et al. (2017) which takes the value of 1 if any tight capital measure was taken by each regulatory authority, and 0 otherwise.	Cerutti et al. (2017)

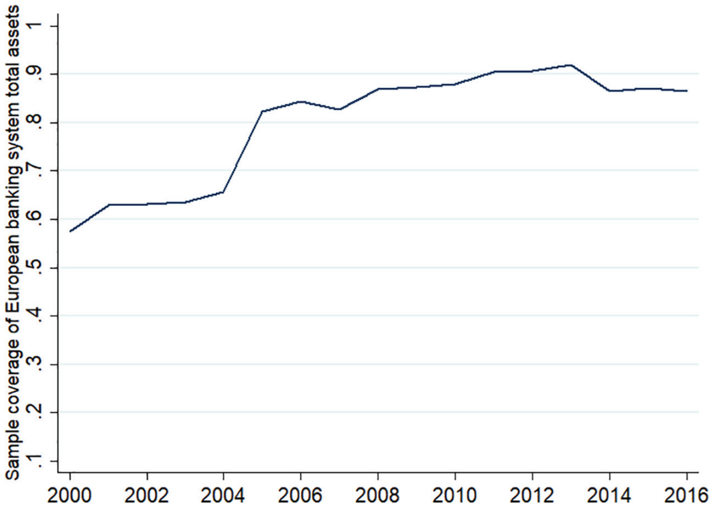
Note: This table presents the definitions and sources of the main variables used in this study.

Table 2. Expected Effects of Each Variable on Banks' Capital Structure

Variables		Expected Effects and Rationale
MARKET-TO-BOOK RATIO (MBR)	–	Based on Gropp and Heider (2010), leverage is negatively correlated with market-to-book ratio. In theory, firms with high market-to-book ratios have higher costs of financial distress.
PROFITABILITY (PROF)	–/+	Based on empirical literature (Gropp and Heider 2010 and De Jonghe and Öztekin 2015) leverage has a negative relationship with profits. These results underpin on the concept that firms that face a lower cost of raising equity in the short term (i.e., profitable firms) tend to hold significantly more capital. This is true mainly when debt financing is the dominant source of external financing (Rajan and Zingales 1995).
SIZE	+	In theory, larger firms are better diversified and have a lower probability of being in financial distress coupled with the fact that, being larger, they are better known by the market and find it easier to issue more debt than smaller firms. The empirical literature (Gropp and Heider 2010 and De Jonghe and Öztekin 2015) supports these arguments.
COLLATERAL (COLL)	+/–	Based on Gropp and Heider (2010), leverage has a positive relationship with collateral. The rationale underlying this factor is that tangible assets are easy to collateralize and thus they reduce the agency costs of data.
RISK	–	According to Gropp and Heider (2010), risk reduces leverage. Firms with more volatile cash flows face higher expected costs of financial distress and should use less debt.
SPEED OF ADJUSTMENT	+	According to Gropp and Heider (2010) and De Jonghe and Öztekin (2015), we expect that banks, like nonfinancial firms, converge toward bank-specific targets.

Note: This table presents the effects of each variable on bank capital structure, where MBR ratio is calculated as the ratio between market value of assets and book value of assets; PROF is computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets; SIZE is defined as the log of total assets which is computed as the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; COLL is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets; RISK is calculated as the log of annualized standard deviation of monthly stock price returns (market*value of equity/market value of bank) or through the log of standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year (in the case of private banks); and SPEED OF ADJUSTMENT is computed as 1 – beta on book leverage or market leverage lagged by one period.

Figure 1. Sample Coverage of Countries' Banking Systems' Total Assets



4.1 Descriptive Statistics

Table 3 (panels A and B) presents descriptive statistics of the banks analyzed in our sample.

In our sample, the total assets of the average bank were 81 billion euros. We observed that the total assets of the bank in the 90's percentile was 211 billion euros, whereas the total assets of the banks in the 10's percentile was almost 800 million euros, which thus points to a large variability in the size of banks operating in Europe. For the average bank in the sample, the book value of equity and the book value of deposits was approximately 4 and 40 billion euros, respectively, which, when considered together with the book or market leverage ratios (92 percent and 91 percent on average, respectively), gives a clear picture of the higher leverage ratio that characterizes banks' capital structure when compared with most other industries.

As depicted in table 3 (panel A), banks' profitability (PROF) was affected by the 2008 crisis, as was the market-to-book ratio (MBR). The profitability of the average bank decreased from 5 percent in 2000 to 4 percent in 2008, and the market-to-book ratio

Table 3. Descriptive Statistics of Bank-Specific Indicators

Variables	2000	2008	2016	Total	Of Which Public	Of Which Private
Book Value of Assets (Millions)						
No. of Observations	273	477	177	6,065	1,517	4,548
Mean	50,183	78,406	163,124	80,911	186,959	45,539
Std. Dev.	114,513	255,208	335,003	229,117	391,000	118,461
Percentile 10%	888	686	8,114	797	1,579	698
Percentile 50%	7,983	8,279	28,810	9,697	24,196	7,815
Percentile 90%	176,844	184,572	527,859	211,103	636,133	126,621
Book Value of Equity (Millions)						
No. of Observations	273	477	177	6,065	1,517	4,548
Mean	2,367	2,847	9,667	3,962	8,931	2,304
Std. Dev.	5,560	7,774	18,127	11,091	17,230	7,353
Percentile 10%	88	61	647	69	162	62
Percentile 50%	512	533	2,411	694	1,794	534
Percentile 90%	6,099	6,248	29,587	19,988	29,030	4,691
Book Value of Deposits (Millions)						
No. of Observations	272	463	177	5,939	1,491	4,448
Mean	43,858	33,777	70,774	39,604	86,090	24,021
Std. Dev.	103,277	100,844	138,039	101,417	166,668	58,815
Percentile 10%	866	443	3,420	565	1,212	499
Percentile 50%	8,385	4,600	13,710	6,339	14,521	4,876
Percentile 90%	117,369	71,351	250,920	103,118	319,060	61,903
PROF						
No. of Observations	270	474	176	6,031	1,506	4,525
Mean	0.05	0.04	0.16	0.05	0.06	0.05
Std. Dev.	0.03	0.04	0.13	0.08	0.10	0.08
Percentile 10%	0.03	0.02	0.04	0.01	0.01	0.01
Percentile 50%	0.04	0.04	0.16	0.03	0.04	0.03
Percentile 90%	0.07	0.06	0.30	0.12	0.17	0.08

(continued)

Table 3. (Continued)

Variables	2000	2008	2016	Total	Of Which Public	Of Which Private
MBR						
No. of Observations	100	128	71	1,826	1,826	—
Mean	2	1	1	1	1	—
Std. Dev.	2	1	1	7	7	—
Percentile 10%	0.0	0.0	0.23	0.01	0.01	—
Percentile 50%	1.5	0.6	0.57	0.93	0.93	—
Percentile 90%	3.2	1.6	1.49	2.41	2.41	—
RISK						
No. of Observations	273	477	177	6,064	1,517	4,547
Mean	0.00	0.01	0.00	0.00	0.00	0.00
Std. Dev.	0.01	0.01	0.00	0.01	0.01	0.01
Percentile 10%	0.00	0.00	0.00	0.00	0.00	0.00
Percentile 50%	0.01	0.00	0.00	0.00	0.00	0.00
Percentile 90%	0.01	0.01	0.00	0.01	0.01	0.01
BL						
No. of Observations	273	477	177	6,065	1,517	4,548
Mean	0.92	0.92	0.91	0.92	0.91	0.92
Std. Dev.	0.06	0.07	0.04	0.07	0.07	0.07
Percentile 10%	0.88	0.86	0.86	0.87	0.86	0.86
Percentile 50%	0.94	0.93	0.92	0.93	0.93	0.93
Percentile 90%	0.97	0.97	0.96	0.97	0.96	0.97
ML						
No. of Observations	100	128	71	1,826	1,826	—
Mean	0.92	0.92	0.91	0.92	0.92	—
Std. Dev.	0.07	0.04	0.07	0.07	—	—
Percentile 10%	0.88	0.86	0.86	0.87	0.87	—
Percentile 50%	0.94	0.93	0.92	0.93	0.93	—
Percentile 90%	0.97	0.97	0.96	0.97	0.97	—

(continued)

Table 3. (Continued)

Variables	2000	2008	2016	Total	Of Which Public	Of Which Private
COLL						
No. of Observations	264	458	175	5,876	1,473	4,403
Mean	0.22	0.28	0.26	0.28	0.26	0.28
Std. Dev	0.13	1.12	0.14	0.95	0.14	1.09
Percentile 10%	0.09	0.08	0.13	0.08	0.11	0.07
Percentile 50%	0.21	0.20	0.23	0.21	0.23	0.20
Percentile 90%	0.45	0.44	0.47	0.45	0.46	0.45

Notes: This table presents descriptive statistics of banks' specific indicators during the period 2000–16. The table presents the number of observations (N), the mean, standard deviation (Std. Dev.), and some percentiles of the following variables: BOOK VALUE of ASSETS, which is the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; BOOK VALUE of EQUITY, which is the sum of the following items: "capital," "reserves," and "net income"; BOOK VALUE of DEPOSITS, which is total deposits from clients; PROF, when computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets; MBR ratio, which is calculated as the ratio between market value of assets and book value of assets; RISK, which is computed as the log standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year; BL, which is computed as 1 – (book value of equity/book value of assets); ML, which is calculated as 1 – [market value of equity (computed as number of outstanding shares*end of year stock price)/market value of bank (computed as the market value of equity + book value of liabilities)]; and COLL, which is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets. Public banks are the ones that have their equity quoted/listed in the capital market whereas private banks are banks that do not have their equity quoted/listed in the capital market.

dropped from 2 to 1 during the same period. While banks' profitability has recovered to a level of around 16 percent up until 2016, market-to-book ratio has never recovered to its pre-crisis level.

Table 3 also shows that there are significantly more private banks than public banks, which appears to indicate that the European Banking System is mainly composed of small and medium-sized private banks, with the sample of public banks being more heterogeneous than that of private banks—as the standard deviation in several bank-specific indicators is higher in the case of public banks. Furthermore, public banks are larger, on average (with average assets of 187 billion euros, when compared with 46 billion euros for private banks), although they are more profitable (1 percentage point above the average profitability of private banks) but less leveraged (1 percentage point below the average leverage shown by private banks) than private banks. These differences are statistically significant when the Wilcoxon signed-rank test is carried out, as shown in table 4. In addition, public banks seem to rely more on nondeposits debt and subordinated debt than private banks. According to Flannery (1994), Flannery and Sorescu (1996), Morgan and Stiroh (2001), Flannery and Rangan (2008), and Allen, Carletti, and Marquez (2011), this kind of debt is more subject to market discipline than deposits.[2]

4.2 Cross-Country Comparison Throughout the Period 2009–16

This subsection explores the main differences between countries, in particular the distribution of public and private banks and their capital composition.

In table 5 we can observe that the percentage of public banks across European countries for the period of 2000 to 2016, in both number and percentage of total assets of each country's banking systems, varies significantly, ranging from 0 percent to 75 percent, and 0 percent to 92 percent, respectively. It should be added that the majority of banks are private at the European level.

[2]Some of the differences between the data presented in table 3 and those shown in table 4 are due to the fact that the variables presented in table 4 were windsorized, whereas those presented in table 3 were not.

Table 4. Descriptive Statistics of Bank-Specific Indicators: Public vs. Private

Variables	Observations	Mean
BL		
Private	4,548	0.922
Public	1,517	0.919
Difference		0.5
BOOK VALUE OF ASSETS (Million Euros)		
Private	4,548	41,110
Public	1,517	107,438
Difference		0.00
PROF		
Private	4,525	0.05
Public	1,506	0.06
Difference		0.00
COLL		
Private	4,403	0.23
Public	1,473	0.25
Difference		0.00
RISK		
Private	4,547	0.0029
Public	1,517	0.0031
Difference		0.00
SECURITIES/ASSETS		
Private	4,517	0.239
Public	1,514	0.230
Difference		0.00
LOANS/ASSETS		
Private	4,548	0.565
Public	1,516	0.585
Difference		0.36
NONDEPOSITS/ASSETS		
Private	4,448	0.304
Public	1,491	0.357
Difference		0.00

(*continued*)

Table 4. (Continued)

Variables	Observations	Mean
SUBORDINATED DEBT/ASSETS		
Private	3,547	0.017
Public	1,337	0.018
Difference		0.02

Notes: This table presents the main differences between public and private banks regarding the following variables: BL, which is computed as 1 – (book value of equity/book value of assets); BOOK VALUE of ASSETS, which is the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; PROF, which is computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets; COLL, which is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets; RISK, which is computed as the log standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year; SECURITIES/ASSETS, which is the ratio between securities and total assets; LOANS/ASSETS, which is the ratio between set loans and total assets; NONDEPOSITS/ASSETS, which is computed as the ratio between nondeposits liabilities and total assets; and SUBORDINATED DEBT/ASSETS, which is calculated as the ratio between total subordinated debt liabilities and total assets. Each variable was windsorized at 0.05 on both the left and right tail. We test for the differences between the two types of banks using a nonparametric test of significance knows as Wilcoxon rank-sum test. Public banks are the ones that have their equity quoted/listed in the capital market, whereas private banks are banks that do not have their equity quoted/listed in the capital market.

Table 5 also shows that a high percentage of public banks (both in terms of number and with regards to the share of the total assets of countries' banking systems) are located in northern and central European countries. With a few exceptions, these countries represent those which have the most capitalized banks and also those which have shown more access to tier 2 capital instruments—mostly composed of subordinated debt. This feature is in alignment with the recent analysis published by the European Central Bank (Gaiduchevici and Zochowski 2017), which stresses that Germany and France have the largest bank debt markets. In another research

Table 5. List of Public Banks and Their Representativeness per Country

Countries	% of Public Banks in the Sample (in number)	% of Public Banks in the Sample (in total assets)	TCR	Tier 1 CR	Tier 2 CR
Austria (AT)	16	11	14	11	3
Belgium (BE)	7	41	16	13	3
Czech Republic (CZ)	17	73	14	13	1
Denmark (DK)	27	83	15	13	2
Estonia (EE)	0	0	22	15	7
Finland (FI)	43	57	16	11	5
France (FR)	19	52	14	11	3
Germany (DE)	9	34	15	11	4
Greece (GR)	75	90	12	11	1
Hungary (HU)	11	51	16	12	4
Ireland (IE)	13	32	13	11	2
Italy (IT)	25	63	13	11	2
Luxembourg (LU)	11	2	17	13	4
Netherlands (NL)	14	56	17	15	2
Poland (PL)	58	84	14	13	1
Portugal (PT)	20	41	13	10	3
Slovakia (SK)	0	0	14	13	1
Slovenia (SI)	25	90	12	10	2
Spain (ES)	10	54	13	10	3
Sweden (SE)	33	92	15	13	2
United Kingdom (UK)	9	49	16	12	4
Average	21	46	15	12	3

Notes: This table presents the representativeness (on average) of public banks per country in number and total assets for the period 2000–16. Public banks are those that have their equity quoted/listed in the capital market, whereas private banks do not have their equity quoted/listed in the capital market. Total capital ratio (TCR), tier 1 capital ratio (Tier 1 CR), and tier 2 capital ratio (Tier 2 CR) are computed according to table 1.

note published by BBVA Research (Garcia and Rocamora 2018), the banks of Greece, Portugal, Ireland, and Spain all display the greatest shortfalls of MREL—which proves that countries with a small number of public banks are more likely to be able to overcome difficulties in complying with MREL requirements.

Attention should be paid to Greece—which recorded a high percentage of public banks and yet a lower number of tier 2 capital instruments—which is an exceptional case, originating from the dramatic reduction of the number of total banks after the onset of the banking and sovereign crisis in that country, which led to a consolidation of Greece's banking system. However, as they are public, these public banks overcame many difficulties by issuing bail-in-able instruments, as a result of concerns regarding the sustainability of both Greek banks and the Greek economy as a whole. Estonia is at the other end of the spectrum, as it has no record of listed banks, although its banking system has the highest percentage of tier 2 capital instruments in the sample. This feature could be explained by the fact that this country is inundated by subsidiaries of foreign banks, whose listed holding companies have much easier access to capital markets to fund their activities, and they thus gain from the benefit of belonging to reputable and well-known banking groups.

5. Methodology

To address the research question presented in section 3, we carry out the following empirical analysis for year, country, and bank's fixed effects:

$$
\begin{aligned}
L_{ict} = \beta_0 &+ \beta_1 PROF_{ict-1} + \beta_2 SIZE_{ict-1} + \beta_3 COLL_{ict-1} \\
&+ \beta_4 RISK_{ict-1} + C_i + C_t + C_c + e_{it},
\end{aligned}
$$

where L_{ict} is book leverage (BL), and the remaining variables are those described in table 1. i, c, and t represent bank, country, and time, respectively, and C_i, C_t, and C_c represent bank, time, and country fixed effects. Banks' specific variables are lagged by one year (as used by Gropp and Heider 2010), in order to mitigate possible endogeneity problems.[3] We ran the model above for all the banks in the sample, and subsequently distinguished between public and private banks in the subsamples. As a first step, we apply exactly the same model to private and public banks, adopting the same dependent and explanatory variables for both types of banks (table 6, regressions 2 and 3). Apart from the variables considered above, we next add the MBR in the case of public banks, and also the market leverage, rather than use book leverage, as expressed below:

$$L_{ict} = \beta_0 + \beta_1 MBR_{ict-1} + \beta_2 PROF_{ict-1} + \beta_3 SIZE_{ict-1} + \beta_4 COLL_{ict-1} + \beta_5 RISK_{ict-1} + C_i + C_t + C_c + e_{it}.$$

At a first glance, this second step could undermine the comparison between public and private banks, as we can observe regressions with one different dependent variable and one additional explanatory variable. In order to properly assess the determinants of public banks' capital structure, we opt to consider market values instead of book values in terms of leverage ratio, as the former are the most frequently used to make decisions in all types of public firms (including banks). Additionally, it is essential to control for market-to-book ratio (MBR) in order to assess the capital structure of listed firms/banks, as well as to evaluate the behavior of the other determinants. Furthermore, by not considering MBR, which is a special feature of public firms or banks, we choose to neglect the empirical work which has been carried out to date. In particular, the results obtained by Gropp and Heider (2010), whose sample was only composed of public and large banks, showed that the regressions with a higher R^2 (and thus with the greatest explanatory power) were those which use market leverage as a dependent variable and MBR as one of the explanatory variables. In this paper, the particularities of public banks gain even more importance and need to be taken into consideration, as the sample is more heterogeneous than that considered by Gropp and Heider (2010), which only includes large, public banks.

As in Gropp and Heider (2010), we consider the regulation and supervision frameworks in year and country's fixed effects. It should

[3]We do not include dividends as used by Gropp and Heider (2010) due to availability issues.

Table 6. Regression Results: The Determinants of Banks' Capital Structure

	Regression 1 (All Banks) BL	**Regression 2 (Public Banks) BL**	**Regression 3 (Private Banks) BL**	**Regression 4 (Public Banks) ML**	**Regression 5 (Gropp and Heider 2010) ML**
MBR	—	—	—	−0.0289***	−0.118***
	—	—	—	(0.0039)	(0.039)
PROF	−0.0231	−0.0143**	−0.0451*	−0.1087**	−0.392***
	(0.0252)	(0.0383)	(0.0249)	(0.0508)	(0.079)
SIZE	0.0167***	0.0106**	0.0187***	0.0224**	0.013**
	(0.0022)	(0.0043)	(0.0024)	(0.0094)	(0.006)
COLL	0.0096	0.0179*	0.0068	−0.0113	0.006
	(0.0071)	(0.0098)	(0.0091)	(0.0204)	(0.013)
RISK	−0.0015***	−0.0072***	−0.0018***	−0.0100**	−0.016***
	(0.0004)	(0.0023)	(0.0014)	(0.0023)	(0.003)
Year Fixed Effects	Yes	Yes	Yes	Yes	No
Bank's Fixed Effects	Yes	Yes	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes	Yes	No
R-square	0.86	0.88	0.86	0.87	0.88
Observations	4,876	1,217	3,559	1,210	2,415

(continued)

Table 6. (Continued)

Notes: This table reports coefficients of the following regression models:

$$L_{ict} = \beta_0 + \beta_1 PROF_{ict-1} + \beta_2 SIZE_{ict-1} + \beta_3 COLL_{ict-1} + \beta_4 RISK_{ict-1} + C_i + C_t + C_c + e_{it}$$

$$L_{ict} = \beta_0 + \beta_1 MBR_{ict-1} + \beta_2 PROF_{ict-1} + \beta_3 SIZE_{ict-1} + \beta_4 COLL_{ict-1} + \beta_5 RISK_{ict-1} + C_i + C_t + C_c + e_{it},$$

where L_{ict} is the ML (regressions 4 and 5) or BL as presented in the table, BL is computed as 1 – (book value of equity/book value of assets); ML is calculated as 1 – [market value of equity (computed as number of outstanding shares*end of year stock price)/market value of bank (computed as the market value of equity + book value of liabilities)]; MBR is calculated as the ratio between market value of assets and book value of assets; PROF is computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets; SIZE is defined as the log of total assets which is computed as the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; COLL is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets; and RISK is computed as the annualized standard deviation of monthly stock price returns (market* value of equity/market value of bank) – regressions 2 and 4 or the log of the standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year – regressions 1 and 3. Regressions 2 and 4 only consider public banks, whereas regression 3 only considers private banks. We use the explanatory variables lagged by one period to mitigate possible endogeneity problems. Regression 5 presents the results obtained by Gropp and Heider (2010) using a sample consisting of the 200 largest traded banks in the United States and EU from the Bankscope database from 1991 to 2004. Public banks are those that have their equity quoted/listed in capital markets, whereas private banks are those that do not have their equity quoted/listed in capital markets. The sample is from 2000 to 2016. Robust standard errors clustered at the bank and country-year levels are in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

be mentioned that we relax this assumption as a robustness check, by including a variable in the model which attempts to proxy the regulation framework of each country. The variable in question is a dummy variable, which assumes a value of 1 if a country's banking regulatory authority has implemented a measure which influences banks' capital requirements, and consequently banks' capital structure, and 0 otherwise. The data for this variable were sourced from the study of Cerutti et al. (2017).

Considering that the standard deviation of stock returns does not apply to private banks, we also adjust this variable, using the standard deviation of return on assets, which is computed as being the standard deviation during the last three years of the ratio between net income and average assets (as described in table 1). This measure has already been adopted by other researchers, such as Beaver, Kettler, and Scholes (1970), Miller and Bromiley (1990), and Bromiley (1991), and also Titman and Wessels (1988), albeit for a context of capital structure.

As a final step, we follow a common practice in the empirical literature on capital structure and opt to use a partial adjustment framework (Flannery and Rangan 2006, Lemmon, Roberts, and Zender 2008, Gropp and Heider 2010, and De Jonghe and Öztekin 2015), which states that in a frictionless world, both banks and firms always maintain their target capital ratio. The speed of this adjustment depends on the tradeoff between adjustment costs and the costs of operating with suboptimal leverage.

Accordingly, in a partial adjustment model, a bank's current capital ratio, $K_{ij,t}$, is a weighted average of its target capital ratio (with weight $\lambda \in [0, 1]$), $K^*_{ij,t}$, and the previous period's capital ratio, $K_{ij,t}$, as well as a random shock, ξ_{ijt}:

$$K_{ij,t} = \lambda K^*_{ij,t} + (1 - \lambda) K_{ij,t-1} + \xi_{ijt}.$$

Every year, banks try to close a proportion λ of the gap between their actual and target capital levels. The smaller the λ, the more rigid bank capital is; that is to say, banks take more time to reach their target. Therefore, λ is interpreted as being the speed of adjustment. As banks' target capital ratio is not manifested, we have to model each bank's target capital level as a function of observed bank characteristics $X_{ij,t-1}$, banks', year, and countries' fixed effects:

$$K^*_{ij,t} = \beta X_{ij,t-1}.$$

Substituting the equation of target leverage in equation of partial adjustment yields the following specification:

$$K_{ij,t} = \lambda\beta X_{ij,t-1} + (1-\lambda)K_{ij,t-1} + \xi_{ijt},$$

which, when translated to our model, produces the following regression:

$$L_{ict} = \beta_0 + \lambda\beta_1 X_{ict-1} + (1-\lambda)L_{ict-1} + C_i + C_t + C_c + e_{ict},$$

where $X_{ij,t-1}$ represents banks' determinants such as profitability, market-to-book ratio, size, collateral, and risk.

Despite acknowledging that the generalized method of moments (GMM) panel analysis proposed by Blundell and Bond (1998) is the most suitable estimator for handling dynamic panel data (with lagged dependent variables), we choose to use a fixed-effects estimator—as our sample has a relatively long time-series dimension, and after taking into consideration the bias caused by the presence of both lagged dependent variables and fixed effects (Nickell 1981) and also bearing in mind that our objective is not to estimate the true speed of adjustment, but rather to test whether these variables and fixed effects are similar for both public and private banks, together with the fact that the above-mentioned bias reduces proportionally as the number of time-series observations increases (Blundell and Bond 1998).

In addition, the panel GMM methods are prone to flaws such as the weak instrument problem (Bun and Windmeijer 2010).[4]

The results are presented with and without the lag dependent variable. Acknowledging the fact that the leverage ratio is characterized by a high time persistence—which, to a certain degree, hampers the contributions of the other identified determinants of capital structure—we opt to present and analyze the contributions of the other determinants, with and without this variable, which is

[4] We also carry out the analysis using GMM, which proves not to impair the results presented in the paper.

in line with the research methodology of both Lemmon, Roberts, and Zender (2008) and Gropp and Heider (2010).

It is worth mentioning that the results obtained here are presented alongside those of Gropp and Heider (2010), in order to facilitate the comparison between them both.

6. Results

This section provides empirical evidence for the research question and the hypothesis presented in section 3.

6.1 Banks' Capital Determinants

Table 6 outlines that banks' capital structure definitively does not depend exclusively on regulation, in contrast to the argument of Gropp and Heider (2010) that the coefficients associated with each determinant of banks' capital structure are statistically nonsignificant. We find that leverage is indeed positively correlated with size and is negatively related with risk. These results are highly aligned with those obtained by Rajan and Zingales (1995), Frank and Goyal (2009), and Gropp and Heider (2010), except for collateral and profitability—whose relationships are not statistically significant, even though they manifest the same signs in the vast majority of the regressions.

Nevertheless, some differences exist between the coefficients originating from the research carried out by Gropp and Heider (2010) and also from those described in this paper (regressions 4 and 5 of table 6)—which are mostly related with the magnitude of the coefficients. These differences result from different samples and time span used by both studies, as well as the use of time fixed effects in this paper, as detailed in table 6.[5] For example, the economic importance (magnitude) of the market-to-book ratio (MBR) coefficient in Gropp and Heider (2010) when compared with that obtained in this paper is probably due to the stress experienced in capital markets between 2008 and 2012 in several European countries, which results

[5] Gropp and Heider (2010) use a sample comprising the 200 largest traded banks in the United States and EU from the Bankscope database, from 1991 to 2004.

in market timing being a capital determinant with less economic relevance. Additionally, the differences encountered with regards to regressions 1, 2, and 3 are hardly surprising, as they include private banks, whereas the sample used by Gropp and Heider (2010) only includes large public banks.

The split of the analysis between private and public banks yields some interesting results. When we use regressions 2 and 3, with the same dependent and explanatory variables, to assess the differences between public and private banks, we find that the determinants of both types of banks' capital structures are broadly similar. Nonetheless, we can observe that profitability and collateral count more in public banks' capital structure than in the case of private banks. Profitability and collateral coefficients have a higher statistical significance in public banks, when compared with private banks. In the case of public banks, profitability is statistically significant at 5 percent, and collateral is statistically significant at 10 percent, whereas in the case of private banks, profitability is significant at 10 percent, and collateral is not statistically significant at all.

Although the above-mentioned analysis has some drawbacks—such as assuming that market timing (measured by market-to-book ratio) does not apply to public banks and that the book values of assets, liabilities, and equity are taken into account in banks' capital decisions to the same degree as market value—it neglects to a large extent the empirical studies which have been carried out over the past years, in particular those carried out by Brav (2009) in the case of nonfinancial firms, and Gropp and Heider (2010) in the case of banks.

In fact, regression 4 reveals that the role carried out by the market timing theory is important for public banks, and that market values are important for capital structure, which highlights the differences between public and private banks.

Taking into account the arguments presented above, we base our analysis on regressions 3 and 4, in order to explain the main differences between public and private banks, and we opt to only subject these specifications to robustness checks.

That said, the comparison between the results from regressions 3 and 4 of table 6 all indicate that the determinants of public and private banks' capital structure are different. As such, the results of this paper evidence that, in the case of private banks, only size and risk are statistically significant at 5 percent and these maintain the expected signs, whereas in the case of public banks, their leverage is negatively related with market-to-book ratio (thus supporting the market timing theory), as well as profitability (associated with the pecking order theory) and risk, and are positively correlated with size, with all the coefficients being statistically significant at a 5 percent level. These results are in line with those obtained by Flannery and Rangan (2008), Lemmon, Roberts, and Zender (2008), Frank and Goyal (2009), and Gropp and Heider (2010).

So far, we can state that in the case of private banks, there are two theories which do not apply in the same manner as in the case of public banks. First, the market timing theory by nature does not apply to private banks, as previously observed in the case of nonfinancial firms (Brav 2009), because the equity of private banks is not marked-to-market. Additionally, the pecking order theory does not apply to private banks with the same significance and magnitude as it is applicable to public banks. According to regression 3, the beta for banks' profitability reaches –0.045 and is only statistically significant at a 10 percent level, whereas in the case of public banks (regression 4), this measure amounts to –0.109, and it is statistically significant at a 5 percent level. It is logical that profitability counts more in the case of public banks, as the pecking order theory gains more relevance against the cost of issuing equity in the market—which is only available to public banks. Economically speaking, we can see that an increase of 1 percentage point of profitability for the average public bank (table 3, panel A) results in a decline of 1 basis point in banks' leverage, whereas for private banks, this increase results in a reduction of 0.2 basis point—in other words, profitability does not have an economic impact on banks' leverage in the case of private banks. Regarding other determinants, we also find differences in magnitude of size and risk; however, these do not result in a significant impact for the average bank in both types of banks.

Considering the argument that banks manage their capital or leverage ratios toward a target, table 7 shows that the speed of adjustment remains stable at around 40 percent whether we consider either the full sample or just private banks—which means that the banks of our sample converge toward their long-run target at a speed of adjustment which is approximately the same as that obtained by Gropp and Heider (2010) (46 percent) and greater

Table 7. Regression Results: The Determinants of Banks' Capital Structure

	Regression 1 (All Banks) BL	Regression 2 (Public Banks) BL	Regression 3 (Private Banks) BL	Regression 4 (Public Banks) ML	Regression 5 (Gropp and Heider 2010) ML
L_{t-1}	0.6128***	0.6077***	0.6028***	0.7263***	0.532***
	(0.0327)	(0.0694)	(0.0371)	(0.0536)	(0.045)
MBR	—	—	—	0.0089**	0.032**
	—	—	—	(0.0035)	(0.015)
PROF	0.0185	0.0286	0.0060	−0.0078	−0.096**
	(0.0154)	(0.0220)	(0.0179)	(0.0386)	(0.045)
SIZE	0.0038**	0.0046**	0.0041**	0.081	−0.005**
	(0.0015)	(0.0021)	(0.0019)	(0.0062)	(0.002)
COLL	0.0066	0.0114	0.0060	−0.0135	0.007
	(0.0047)	(0.0072)	(0.0061)	(0.0131)	(0.012)
RISK	−0.0012***	−0.0021**	−0.0013***	−0.0004	−0.003
	(0.0003)	(0.0009)	(0.0003)	(0.0019)	(0.002)
Year Fixed Effects	Yes	Yes	Yes	Yes	No
Bank's Fixed Effects	Yes	Yes	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes	Yes	No
R-square	0.91	0.92	0.91	0.90	0.92
Observations	4,876	1,217	3,559	1,210	2,059

(*continued*)

Table 7. (Continued)

Notes: This table reports coefficients of the following regression models:

$$L_{ict} = \beta_0 + \lambda\beta_1 X_{ict-1} + (1-\lambda)L_{ict-1} + C_i + C_t + C_c + e_{ict},$$

where L_{ict} is the ML (regressions 4 and 5) or BL as presented in the table, BL is computed as 1– (book value of equity/book value of assets); ML is calculated as 1–[market value of equity (computed as number of outstanding shares*end of year stock price)/market value of bank (computed as the market value of equity + book value of liabilities)]; $X_{ij,t-1}$ stands for banks' determinants such as MBR, PROF, SIZE, COLL, and RISK. MBR is calculated as the ratio between market value of assets and book value of assets; PROF is computed as the ratio between the sum of pre-tax profit and interest expenses and the book value of assets; SIZE is defined as the log of total assets which is computed as the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; COLL is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets; and RISK is computed as the annualized standard deviation of monthly stock price returns (market*value of equity/market value of bank) – regressions 2 and 4 or the log of the standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year – regressions 1 and 3. We use the explanatory variables lagged by one period to mitigate possible endogeneity problems. Regression 5 presents the results obtained by Gropp and Heider (2010) using a sample consisting of the 200 largest traded banks in the United States and EU from the Bankscope database from 1991 to 2004. Public banks are those that have their equity quoted/listed in capital markets, whereas private banks are those that do not have their equity quoted/listed in capital markets. The sample is from 2000 to 2016. Robust standard errors clustered at the bank and country-year levels are in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

than that obtained by De Jonghe and Öztekin (2015) (29 percent). Despite the similar significance of the lag coefficient for both types of banks, it is interesting to observe that private banks, which evidence a speed of adjustment of 40 percent, are slightly faster than public banks in reaching their capital target ratios, whose speed of adjustment is around 27 percent. This feature is probably due to the fact that public banks are more exposed to market volatility than private banks. It is not surprising that public banks show a lower speed of adjustment than private banks, bearing in mind that the time period considered in this study encompasses the global financial crisis, as well as the sovereign debt crisis experienced in Europe—both of which negatively affected capital markets, and especially those banks that detained sovereign debt and were under the spotlight of the markets.

It should be mentioned that the effect of the lag of the leverage ratio on the other determinants does not put into question the differences found between public and private banks. This result is due to the acknowledged high persistence of the leverage ratio, over time, which jeopardizes the effect of other variables. To this end, we commence our analysis by ignoring the lag of the leverage ratio and instead include the persistence of the leverage ratio in our analysis, as carried out by Lemmon, Roberts, and Zender (2008) and Gropp and Heider (2010), in order to better assess the slight differences between public and private banks.

6.2 Access to the Subordinated Debt and the Influence of Its Discipline

The results presented so far point out some differences between public and private banks, both in statistical significance and in magnitude, with regards to their decisions in relation to the optimization of their capital structure, with particular relevance for their main determinants. These results are hardly surprising, given that it is natural for public banks to have easier access to capital and debt markets, which in turn can highly influence their decisions regarding capital structure. It is therefore important to investigate the implications of these differences in terms of access by banks to capital and debt markets and consequently the space given to market discipline.

Market discipline can be defined as the "process by which informed market investors gather and monitors firm's activities and prospects as well as their risk" (Flannery and Sorescu 1996). The importance of market discipline has been recognized by supervisors and regulators since the implementation of Basel II (Basel Committee on Banking Supervision 2006). It is important to mention that Basel II introduced a third pillar called "Market Discipline"—whose aim is to encourage market discipline by introducing the disclosure of a broader range of information, in order to enable market participants to assess key pieces of information regarding capital, risk exposures, risk-assessment procedures, and the capital adequacy of the institution (BCBS 2006).

Even though one stream of the literature argues that banks' capital structure and risk-taking are heavily determined by regulators and supervisors, rather than by markets (Berger, Herring, and Szegö 1995, Rajan and Zingales 1995, Santos 2001, and Calomiris and Wilson 2004), another, more recent, strand of the literature advocates that the attributes that affect banks' capital structure and target ratios are not that different from those which influence nonfinancial firms' capital (Flannery 1994, Flannery and Rangan 2008, Gropp and Heider 2010, Allen, Carletti, and Marquez 2011, and De Jonghe and Öztekin 2015). This view stresses that banks are, to a certain extent, subject to market discipline, and that more space should be given to the market discipline of banks.

In this context, in order to shed some light on the differences between public and private banks with regards to access to capital and debt markets and ultimately the influence of market discipline on both types of banks' capital determinants, it is useful to test whether public banks rely more on nondeposits debt and subordinated debt than private banks. According to Flannery (1994), Flannery and Sorescu (1996), Morgan and Stiroh (2001), Flannery and Rangan (2008), and Allen, Carletti, and Marquez (2011), this kind of debt is more subject to market discipline than deposits.

Taking the share of assets funded by subordinated debt as a proxy for market access and market discipline, table 8 shows that public banks have been relying more on subordinated debt to fund their assets than private banks—which could imply that public banks have been capturing the preference of the market, as their capital determinants are more similar to those observed in the case

Table 8. Regression Results: The Differences between Public and Private Banks

	Regression 1 (All Banks)	Regression 2 (All Banks)
	Nondeposits	Subordinated Debt
PROF	0.4821**	0.0071
	(0.2374)	(0.0089)
SIZE	0.2699***	−0.0078***
	(0.0283)	(0.0014)
COLL	0.0044	0.0015
	(0.0728)	(0.0035)
RISK	−0.0057	0.0001
	(0.0052)	(0.0002)
PUBLIC	−0.1219	0.0102**
	(0.1190)	(0.0050)
Year Fixed Effects	Yes	Yes
Bank's Fixed Effects	Yes	Yes
Country Fixed Effects	Yes	Yes
R-square	0.52	0.59
Observations	4,815	4,028

Notes: This table reports coefficients of the following regression model:

$$L_{ict} = \beta_0 + \beta_1 PROF_{ict-1} + \beta_2 SIZE_{ict-1} + \beta_3 COLL_{ict-1} + \beta_4 RISK_{ict-1} + \beta_5 PUBLIC_{ict} + C_i + C_t + C_c + e_{it},$$

where L_{ict} is the nondeposits liabilities (regression 1) or subordinated debt (regression 2) as presented in the table, where "Nondeposit" is computed as the ratio between nondeposits liabilities and total assets; "Subordinated Debt" is calculated as the ratio between total subordinated debt liabilities and total assets; PROF is computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets; SIZE is defined as the log of total assets which is computed as the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; COLL is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets; RISK is computed as the log of the standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year; and PUBLIC is a dummy variable that assumes 1 if the bank is quoted/listed and 0 otherwise. The sample period is from 2000 to 2016. Robust standard errors clustered at the bank and country-year levels are in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

of nonfinancial firms—which is commonly stated in the extant literature.

Therefore it seems that the main differences found so far regarding the determinants of capital structure of public and private banks have been influencing the access to market financing, fostering market discipline in the case of public banks. This empirical evidence is also in line with the descriptive analysis carried out in subsection 4.2, as well as with the conclusions presented by the European Central Bank (Gaiduchevici and Zochowski 2017) and other research notes published by investment banks (Garcia and Rocamora 2018), which all highlight the fact that the southern European countries, which have a lower percentage of public banks, present material shortfalls regarding MREL requirements.

Regarding the fact that it is apparently easier for public banks to gain access to capital markets when compared with private banks (which is represented by the share of funding by subordinated debt), it is interesting to explore whether the market makes a distinction within the public bank sector between the levels of risk, size profitability, and collateral, by interacting each bank-specific variable with the dummy variable which identifies public banks. Additionally, it was decided to test whether the characteristics of the capital market of the country where the bank is headquartered influences access to the market for the case of public banks. This test was carried out by applying a measure of market efficiency which had previously been developed by Svirydzenka (2016), and then interacting this measure with a dummy variable which identifies public banks.

Taking into account the results outlined in table 9, it appears that public banks have equal access to market funding; however, this access is facilitated in more efficient markets, as the coefficient resulting from the interaction between public banks and market efficiency is both statistically significant and positively related with the share of subordinated debt.

6.3 Robustness Checks

We are tempted to be of the opinion that the differences observed are driven by other characteristics, rather than being private or public. Acknowledging this, we have split the original sample into two

Table 9. Regression Results: The Differences between Public and Private Banks

	Regression 1 Subdebt	**Regression 2 Subdebt**	**Regression 3 Subdebt**	**Regression 4 Subdebt**	**Regression 5 Subdebt**
PROF	0.0062	0.0070	0.0062	0.0073	0.0073
	(0.0094)	(0.0088)	(0.0090)	(0.0090)	(0.0089)
SIZE	−0.0078***	−0.0075***	−0.0079***	−0.0078***	−0.0078***
	(0.0014)	(0.0015)	(0.0014)	(0.0014)	(0.0014)
COLL	0.0015	0.0016	0.0064*	0.0014	0.0018
	(0.0035)	(0.0035)	(0.0036)	(0.0035)	(0.0035)
RISK	0.0001	0.0000	0.0000	0.0002	0.0001
	(0.0002)	(0.0002)	(0.0002)	(0.0003)	(0.0002)
PUBLIC	0.0100*	0.0218	0.0143***	0.0073	0.0085*
	(0.0051)	(0.0162)	(0.0055)	(0.0052)	(0.0044)
FME	—	—	—	—	−0.0004
	—	—	—	—	(0.0021)
PUBLIC*PROF	0.0027	—	—	—	—
	(0.0094)	—	—	—	—
PUBLIC*SIZE	—	−0.1219	—	—	—
	—	(0.0017)	—	—	—
PUBLIC*COLL	—	—	−0.0160*	—	—
	—	—	(0.0084)	—	—
PUBLIC*RISK	—	—	—	−0.0005	—
	—	—	—	(0.0003)	—
PUBLIC*FME	—	—	—	—	0.0046*
	—	—	—	—	(0.0026)

(*continued*)

Table 9. (Continued)

	Regression 1 Subdebt	Regression 2 Subdebt	Regression 3 Subdebt	Regression 4 Subdebt	Regression 5 Subdebt
Year Fixed Effects	Yes	Yes	Yes	Yes	Yes
Bank's Fixed Effects	Yes	Yes	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes
R-square	0.59	0.59	0.59	0.59	0.59
Observations	4,028	4,028	4,028	4,028	4,028

Notes: This table reports the coefficients of several regressions which exploit interactions between public banks and bank-specific variables as well as a country-specific variable that characterizes each country financial market efficiency, inspired by the work carried out by Svirydzenka (2016). "SubDebt" stands for subordinated debt, calculated as the ratio between total subordinated debt liabilities and total assets; PROF is computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets; SIZE is defined as the log of total assets, which is computed as the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; COLL is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets; RISK is computed as the log of the standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year; PUBLIC is a dummy variable that assumes 1 if the bank is quoted/listed, and 0 otherwise; and FME stands for financial market efficiency index computed as stock market turnover ratio, i.e., stocks traded to capitalization (Svirydzenka 2016). The sample period is from 2000 to 2016. Robust standard errors clustered at the bank and country-year levels are in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

subsamples which group the institutions into two clusters according to their size (as a proxy for similarity). Cluster 1 comprises small/medium banks (banks with average assets of less than 81 billion euros), and cluster 2 comprises large banks (banks with average assets equal to or greater than 81 billion euros). By using this division, we can isolate the differences between public and private, comparing similar banks by using size as the common characteristic for the two groups of banks.

Table 10 shows that, in the case of small/medium banks (cluster 1), the differences between the determinants of public and private banks' capital structure are maintained when compared with the results presented in table 6. That is to say that the determinants of public banks' capital structure are closer to those which affect nonfinancial firms. In the case of large banks (cluster 2), to a large extent these results do not confirm those initially reported. Accordingly, we can notice certain differences between public and private banks in relation to their capital structure, although these differences are not relevant for the group of large banks (cluster 2). This may be partly explained by the fact that in cluster 2 we obtained a substantially lower number of observations, which hampers the variability in the coefficients, compounded by the fact that large banks are subject to tighter market discipline than small/medium banks—which diminishes the effect of the type of banks on their capital structure determinants.

As mentioned in section 5, we relax the assumption that each country's regulatory framework is included in country and year fixed effects, in similarity to Gropp and Heider (2010), and we include a dummy variable which is adapted from the study carried out by Cerutti et al. (2017)—which assumes the value of 1 if any tight capital measure was implemented by each regulatory authority, and 0 otherwise. The inclusion of this variable does not change the results nor the conclusions presented in this paper (table 11).

7. Main Conclusions and Policy Implications

Over the last decade there has been a considerable increase in the number of empirical studies which have focused on testing how the determinants of capital structure applied to nonfinancial firms can also apply to banks. Two of the most important recent studies

Table 10. Regression Results: The Determinants of Banks' Capital Structure

	Regression 1 (Private Banks) (Cluster 1) BL	**Regression 2 (Public Banks) (Cluster 1) BL**	**Regression 3 (Private Banks) (Cluster 2) BL**	**Regression 4 (Public Banks) (Cluster 2) ML**
MBR	—	−0.0274***	—	−0.0324***
	—	(0.0045)	—	(0.0069)
PROF	−0.0565**	−0.1805***	0.0108	0.0201
	(0.0285)	(0.0614)	(0.0353)	(0.0835)
SIZE	0.0186***	0.0174	0.0226***	0.0224**
	(0.0027)	(0.0122)	(0.0043)	(0.0095)
COLL	0.0050	0.0031	0.0156	0.0101
	(0.0105)	(0.0263)	(0.0145)	(0.0368)
RISK	−0.0019***	−0.0082***	−0.0015**	−0.0129***
	(0.0006)	(0.0028)	(0.0007)	(0.0039)
Year Fixed Effects	Yes	Yes	Yes	Yes
Bank's Fixed Effects	Yes	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes	Yes
R-square	0.85	0.88	0.82	0.79
Observations	3,067	831	492	379

Notes: This table reports coefficients of the following regression model:

$$L_{ict} = \beta_0 + \beta_1 PROF_{ict-1} + \beta_2 SIZE_{ict-1} + \beta_3 COLL_{ict-1} + \beta_4 RISK_{ict-1} + C_i + C_t + C_c + e_{it},$$

where L_{ict} is the ML (regressions 2 and 4) or BL as presented in the table, BL is computed as 1 – (book value of equity/book value of assets); PROF is computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets; SIZE is defined as the log of total assets, which is computed as the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; COLL is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets; RISK is computed as the log of the standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year. Public banks are those that have their equity quoted/listed in capital markets, whereas private banks are those that do not have their equity quoted/listed in capital markets. The sample is from 2000 to 2016. Robust standard errors clustered at the bank and country-year levels are in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

Table 11. Regression Results: The Determinants of Banks' Capital Structure

	Regression 1 (All Banks) BL	Regression 2 (Private Banks) BL	Regression 3 (Public Banks) ML
MBR	—	—	−0.0249***
	—	—	(0.0040)
PROF	−0.0456*	−0.0401	−0.1151*
	(0.0274)	(0.0325)	(0.0642)
SIZE	0.0164***	0.0179***	0.0225**
	(0.0024)	(0.0026)	(0.0110)
COLL	0.0111	0.0092	−0.0132
	(0.0072)	(0.0093)	(0.0201)
RISK	−0.0014***	−0.0019***	−0.0109***
	(0.0005)	(0.0006)	(0.0025)
CAPR	0.0009	0.0006	−0.0020
	(0.0017)	(0.0019)	(0.0040)
Year Fixed Effects	Yes	Yes	Yes
Bank's Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
R-square	0.87	0.87	0.87
Observations	4,512	3,345	1,067

Notes: This table reports coefficients of the following regression model:

$$L_{ict} = \beta_0 + \beta_1 PROF_{ict-1} + \beta_2 SIZE_{ict-1} + \beta_3 COLL_{ict-1} + \beta_4 RISK_{ict-1} + \beta_5 CAPR_{ict-1} + C_i + C_t + C_c + e_{it},$$

where L_{ict} is the ML (regression 3) or BL as presented in the table, BL is computed as 1– (book value of equity/book value of assets); ML is calculated as 1 – [market value of equity (computed as number of outstanding shares*end of year stock price)/market value of bank (computed as the market value of equity + book value of liabilities)]; PROF is computed as the ratio between the sum of pretax profit and interest expenses and the book value of assets; SIZE is defined as the log of total assets which is computed as the sum of the following items: "cash and advances in other credit institutions," "claims on other credit institutions," "total loans and receivables," "financial assets classified at fair value through profit or loss," "financial assets classified as available for sale," "financial assets classified as held for trading," "financial assets classified as held to maturity," and "other assets" net of the respective impairment; COLL is computed as the ratio between the sum of the following items: "total securities," "fixed assets," and "cash and due from banks" and the book value of assets; RISK is computed as the log of the standard deviation of return on assets (which is computed as the ratio between net income and the average of book value of assets) calculated from the last three observations for the respective year; and CAPR is a dummy variable drawn from the study carried out by Cerutti et al. (2017) which takes the value of 1 if any tight capital measure was taken by each regulatory authority, and 0 otherwise. Public banks are those that have their equity quoted/listed in capital markets, whereas private banks are those that do not have their equity quoted/listed in capital markets. The sample is from 2000 to 2016. Robust standard errors clustered at the bank and country-year levels are in parentheses. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively.

are those of Gropp and Heider (2010) and De Jonghe and Öztekin (2015). Using a sample of large and public banks from Europe and the United States (from 1991 to 2004), the first study shows that regulation is not the main feature that distinguishes bank capital structure from what was argued by Modigliani and Miller (1958), as, for these authors, most banks appear to optimize their capital structure in much the same way as nonfinancial firms—except when their capital is close to the regulatory minimum.

Taking Gropp and Heider (2010) as a starting point, we assess whether the determinants of European banks' capital structure also depend on the type of the institution (i.e., whether it is private or public). We attempt to empirically answer the following question: Are the determinants of banks' capital structure different in public versus private banks? This topic has gained significant attention after the implementation of the new capital requirements of Basel III (conservation buffer, countercyclical buffer, and capital buffers for systemically important institutions, as well as leverage requirements—unweighted capital requirements) and also the MREL requirements envisaged in the new resolution framework—where access by banks to market financing plays a pivotal role in the success of compliance with these new requirements, which consequently strengthen the entire financial system and ultimately financial stability as a whole.

To a certain extent, our results for the sample as a whole confirm those of Gropp and Heider (2010), namely that leverage is positively correlated with size and is negatively correlated with profits, market-to-book value, and risk—which is broadly consistent with the pecking order and market timing theories.

Interestingly, our results have identified differences between public and private banks which influence the access to market. As such, public banks whose capital determinants are more similar to those observed in the case of nonfinancial firms have been those which have been more active in funding their assets with subordinated debt (proxy used for market financing), in accordance with the empirical literature which compares them with private banks.

In this context, this paper sheds some light on the potential success of the introduction of additional capital requirements under Basel III, and also how the entrance into force of the EU Directive on Banking Recovery and Resolution (BRRD), which regulates access to equity and debt markets, plays a pivotal role in their success. The new resolution regime envisages a "bail-in process"—where shareholders, creditors, and a certain proportion of deposits are called upon to jointly share any losses with the first- and second-mentioned stakeholders—which implies that banks have to comply with the additional capital requirements (which are known as MREL), and that, in turn, they should comply with those equity and/or debt instruments that have bail-in-able characteristics.

The observed differences between public and private banks in accessing market financing can result in different levels of implementation of the new resolution regime and the MREL requirements across banks. Therefore, according to the results reported in this paper, public banks are expected to be quicker in complying with these kinds of requirements than private banks, which are more able to withstand difficulties in this regard.

In summary, the results presented in this paper represent serious implications from a policy point of view. On the one hand, the results highlight the potential challenges facing private banks as a consequence of the implementation of the new resolution regime when compared with public banks, due to the fact that the latter are already subject to diversified liabilities and capital ratios, where subordinated debt plays an important role. The unlevel playing field across banks in terms of accessing capital and debt markets might well undermine the success of this new regime, and in fact ultimately result in negative consequences for financial stability. Furthermore, private banks, which are characterized by a weaker investor base, due to a lack of, or limited experience in, issuing equity and debt instruments, are expected to withstand a higher spread—which in turn could undermine their profitability and also the internal generation of capital. On the other hand, the combination of this new resolution regime and the requirement to issue bail-in-able debt instruments could contribute to increasing the number of banks which are subject to market discipline and align the capital structure determinants of public and private banks and thus change the capital determinants of private banks to be more similar to those of public banks.

Despite the weakness demonstrated by market forces which came to light during the last financial crisis, market discipline is welcomed by regulators, and some strands of the literature recognize that market discipline can augment the role traditionally carried out by regulatory and supervisory bodies. According to some authors, market discipline can be beneficial in several ways. First, the market can provide information to supervisors regarding the probability of default by banks (Flannery and Sorescu 1996, Gropp and Vesala 2004, Ashcraft 2008, Flannery and Rangan 2008, Distinguin, Kouassi, and Tarazi 2013, Hoang, Faff, and Haq 2014, and Oliveira and Raposo 2019)—which helps supervisors to efficiently allocate resources. Second, the market can discipline banks directly by adopting certain covenants regarding debt issues (Ashcraft 2008). Third, this type of discipline can reduce the moral hazard incentives which governmental guarantees create for banks. And finally, market discipline can improve efficiency and thus create pressure on less efficient banks to change their modus operandi (Martinez Peria and Schmukler 2001).

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The Effect of the Single Currency on Exports: Comparative Firm-Level Evidence*

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We investigate how adopting the euro affects exports using firm-level data. In contrast to previous studies, we focus on two countries, Slovakia and Estonia, which adopted the single currency individually and had different exchange rate regimes. The results highlight the importance of the transaction costs channel related to exchange rate volatility. The euro changeover has a strong pro-exports effect for a country with a floating exchange rate, while it has almost no effect for a country with a fixed exchange rate to the euro. The export effect manifests itself mainly through the intensive margin and is heterogeneous across firms, with more productive firms and smaller exporters benefiting the most.

JEL Codes: F14, F15.

1. Introduction

Assessment of the potential benefits of any currency union relies predominantly on the savings that come from eliminating nominal exchange rate volatility and reducing transaction costs. These savings are expected to lead to higher exports, higher gross domestic

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product (GDP), and consequently higher living standards in the economies of the currency union. Introducing currency unions such as the euro area can affect trade through more than one channel. According to Baldwin et al. (2008) we may consider (i) trade prices being reduced as transaction costs from exchange rate volatility and foreign exchange fall; (ii) trade prices being reduced by increased competition; and (iii) opportunities opening up for newly traded goods. However, there is no consensus on which channel has a decisive role for the gains in trade.

The more years that pass since the introduction of the single European currency, the more information there naturally is on the impact of this step on international trade. While there is ample macro-level evidence built on the gravity-type models that show the euro changeover had a positive impact on trade, micro-level analyses remain limited to a small number of countries. The distribution of gains from trade and the mechanism behind this distribution are still unclear. The aim of this paper is to bring more evidence on the topic using firm-level data. The paper contributes to the literature on the effects of common-currency areas on trade, first, by studying two natural experiments where trade costs were reduced but there was no increase in competition from other countries and, second, by testing the heterogeneous effect of the euro on exports.

The data come from two relatively new euro-area members: Slovakia, which joined the common-currency area in 2009, and Estonia, which joined in 2011. The difference-in-differences methodology is applied where the euro adoption effect is identified by firm-level bilateral trade flows to European Union (EU) countries. The treatment group consists of exports to the euro-area countries, while the control group consists of exports to the non-euro-area EU countries. Building on the theoretical models of heterogeneous and multiproduct firms (Melitz 2003 and Bernard, Redding, and Schott 2011), we answer questions about whether adopting the euro has raised the probability of exports to a given destination, has increased the number of products for each destination, or has boosted average or total exports to each destination. This approach is used to examine whether the benefits of euro adoption are manifested mostly through the intensive margin or the extensive margin. The incidence of gains from adopting the euro is tested across productivity and size groups, and across other firm characteristics such as age, foreign ownership, and financing structure. The unconditional quantile regression technique of Firpo, Fortin, and Lemieux (2009) is applied to study the effect of the euro along the distribution of exports, to test whether the smallest exporters or the largest ones benefited the most.

The Slovakian and Estonian changeovers to the euro are good case studies for a number of reasons. The lion's share of the literature on how the euro affected trade is based on papers that use data from the countries that introduced the euro in 1999. All these countries switched to the euro at the same time, and transaction costs were reduced for all of them. This meant the introduction of the euro affected trade in two ways, with a positive effect from lower transaction costs and a negative effect from increased competition from other euro-area countries. Berthou and Fontagne (2013) control for the competition effect indirectly and find that the euro effect is underestimated when the increased competition is ignored. The advantage of our paper is that we use two cases where the euro was introduced in one country at a time, so that there was no effect of increased competition from other countries.

Our two-country natural experiment study has further advantages. The timing of the effect is concentrated, as the euro was introduced for electronic and cash transactions at the same time, and a much larger control group of EU destination markets is available compared with the situation when the euro was first introduced. Most importantly, the cases analyzed in this paper, Slovakia and Estonia, provide insightful comparative evidence about the channel behind the effects. Estonia had a currency board system with a strict peg to the euro prior to the changeover, while Slovakia had a floating exchange rate against the euro. Although, both countries joined the Exchange Rate Mechanism II (ERM II) with the commitment to maintain their exchange rate within the ±15 percent fluctuation band around the agreed central parity without severe tensions, in line with the ongoing strong economic convergence Slovakia was two times permitted to revalue its parity. As a result, in contrast to Estonia, Slovakia saw its transaction costs from exchange rate volatility fall significantly in its run up to the euro. We use these similarities and differences to identify the channel behind the gains.[1] To our knowledge, there is only one paper that uses the data of the recent

[1] Please see more in section 2 on the case of the two changeovers.

euro-area members to estimate the effect of euro on trade, Mika and Zymek (2018) at the macro level, but they do not ask which channel contributed to the effect, and they ignore the cross-country variation in exposure to different channels.

There is a lot of research on how common-currency areas affect trade at the macro level, less at the product level, and very little at the micro level. Abundant macroeconomic studies applying different versions of the gravity models typically find that a common currency has a positive effect on trade. Many amendments to the gravity-type estimates, which relate trade performance with the size of engaged economies and various measures of trade costs, have emerged since Rose (2000) published his encouraging results. The main contributions are critically reviewed in Baldwin (2006), who concludes that the euro trade effect varies between 5 percent and 10 percent, and also in Bun and Klaassen (2007), Baldwin et al. (2008), and Polak (2018), who suggest the effect is even smaller. Baldwin and Taglioni (2007) or Head and Mayer (2014) give evidence on more estimates of the euro effect that are frequently disputable. More recently, Glick and Rose (2015) show that the estimates of the currency union effect are sensitive to the exact econometric methodology and conclude that the euro has a smaller trade effect than other currency unions do. A possible reason for the milder effect could be the deep pre-accession integration in the common market.

Using product-level trade data helps to unveil more of the consequences of a currency union. Baldwin and Di Nino (2006) provide supportive evidence for the newly traded goods hypothesis. Flam and Nordström (2007) find a stronger trade effect for products that were not exported continuously and confirm the significant and substantial effects on the extensive margin of trade from the introduction of the Economic and Monetary Union (EMU). Simple stylized facts based on product-level data for the trade of new euro-area countries indicate that the euro promotes exports of intermediate or semi-finished products, as shown by Flam and Nordström (2007) or Rotili (2014).

A microeconomic approach offers even more aspects for study than the aggregate or product-level approach does. The theoretical approaches build on new trade theory and Melitz (2003). In his framework, a fall in export costs allows smaller and less productive firms to start exporting and increases the value of exports for each firm. Bernard, Redding, and Schott (2011) propose a multiproduct model, where a fall in trade costs leads to firm selection into the export market, with an increase in both the number of destinations for each type of product and the range of products exported by firms to a given destination.

As we have access to detailed firm-level trade data, we contribute to the smaller stream of empirical firm-level literature that uncovers processes that are usually hidden in aggregate trade figures. Baldwin et al. (2008) offer the first unconditional evidence of the euro trade effect for France and Belgium and confirm the newly traded goods hypothesis. However, the conditional estimates with a more rigorous approach are not conclusive. Berthou and Fontagne (2008a, 2008b, 2013) find that the adoption of the euro in France has a statistically significant impact in reducing trade costs. Berthou and Fontagne (2013) show that the euro changeover increased firm-level exports by 5 percent in France and that the intensive margin dominated the effect. Nitsch and Pisu (2008) estimate the euro trade effect on Belgian exporters. They find no statistically significant effect on total firm-level exports, but find that intra-euro-area trade has expanded through new markets and new product margins. De Nardis, Pappalardo, and Vicarelli (2008), employing Italian firm-level data, find that the euro had no statistically significant effect on total firm-level exports, but it had an effect along the extensive margin of new markets.

There is also no consensus on which type of firm saw its exports increase the most from the changeover to the euro. Berthou and Fontagne (2008b) find that firm efficiency and the composition effect play a role in the decision by firms to export, but the newly traded goods hypothesis is not subject to the presence of the composition effect of firm size. There is evidence that the most productive firms started to export more because of the euro changeover (Berthou and Fontagne 2013) or that less productive firms started to export more (Nitsch and Pisu 2008). It has also been found that the exports of the smallest firms increased the most due to the introduction of the euro (Nitsch and Pisu 2008 and Esteve-Pérez et al. 2011).

We find that adopting the euro had a statistically significant and strong economic impact on exports for Slovakia but almost no effect for Estonia. For Slovakia we find that the changeover to the euro increased exports by 18 percent and that the intensive margin dominated the effect. One possible explanation for this larger effect is that we are studying countries that adopted the euro separately and not in a big group of countries, and so no adverse competition effect from other countries emerges. We claim that the main mechanism behind the effect is the reduction of transaction costs from the exchange rate volatility that exporters were exposed to in Slovakia but not in Estonia.[2] It is also found that the gains in trade from the reduced transaction costs are distributed heterogeneously across firms. More productive firms benefit the most from the reduced transaction costs. These findings provide empirical evidence on theoretical models like that of Melitz (2003), as we confirm the prediction that reduced trade costs contribute to more concentrated distribution of productivity. We also find that the exports of smaller and medium-sized exporters increased the most after the changeover to the euro.

The next section introduces the main facts on the euro changeover in the analyzed countries; section 3 provides a detailed description of the data used in our econometric analyses; section 4 describes our methodology; in section 5 we present the estimation results and robustness tests; and section 6 concludes.

2. The Case of Two Changeovers, Slovakia and Estonia

The sample countries in this study, Slovakia and Estonia, are exposed differently to the reduction of transaction costs caused by the introduction of the euro, as they had different monetary policies before they adopted the euro. Slovakia had a fixed exchange rate system in the 1990s and shifted to a managed floating exchange rate with inflation targeting from 1998. The indicative target inflation was gradually lowered toward 2 percent prior to the accession to the common-currency area (Banerjee et al. 2017). Estonia had a fixed exchange rate system with a currency board from 1992, where the Estonian currency was first strictly pegged to the German mark

[2] Please note that the effect from common currency might have been even larger if the country had joined from a fully free-floating exchange rate compared with the ±15 percent fluctuation band of ERM II. In the same vein, the effect could have been smaller if the country had strictly followed the ERM II without recent revaluations.

and from 2002 onward until the changeover, the Estonian kroon was pegged to the euro (*Kroon & Economy* 2008). Both of the countries joined the ERM II—Slovakia in 2005 and Estonia in 2004. While Slovakian koruna was agreed to float within the 15 percent nominal exchange rate bound and revalued its central parity by 8.5 percent in March 2007 and by 17.6 percent in May 2008, Estonian kroon continued to be fixed to the euro throughout its entire participation in the ERM II.

The variations of monthly exchange rates with euro-area countries and non-euro-area countries were quite similar before the changeover in Slovakia. The coefficient of variation of monthly exchange rates with euro-area trade partners was 2.6 percent before the changeover and dropped to zero afterwards (for more details, see Lalinsky and Meriküll 2018). However, Estonia has not experienced any exchange rate volatility with the euro-area trade partners prior to or after the accession. This implies that the benefits in trade from the removal of exchange rate volatility are different in the two countries and the gains expected in exports should be larger in Slovakia than in Estonia.

Another source of transaction costs is the reduced cost of foreign exchange. The European Commission (1990) estimated that the expected gains from foreign exchange brought by the euro were from 0.1 percent to 1 percent of GDP and were higher for small euro-area countries like our sample countries. The euro was already a dominant currency in extra-euro-area trade before the changeover in Slovakia and Estonia, as around 90 percent of extra-euro exports were invoiced in euros in Slovakia and 50 percent in Estonia one year before the changeover (European Central Bank 2012). Unfortunately, there are no public statistics on currency invoicing for intra-euro-area trade, but it is also likely to have been higher in Slovakia than in Estonia.[3] It can be expected that the gain from reduced foreign exchange transaction costs in exports was larger in Slovakia than in Estonia, though both of the countries gained. The reasoning for this is that the drop in transaction costs from foreign exchange was larger in Slovakia than in

[3] The invoicing information is available in the confidential customs data for Estonia used in this paper. Based on this source, 67 percent of the volume of exports to the euro area has been invoiced in euros before the changeover.

Estonia, as currencies other than the euro played a minor role in the exports of Slovakian firms, while the exports of Estonian firms were more frequently invoiced in other currencies also after the changeover.

The findings on the effects of the introduction of the euro suggest that the gains in trade were different across countries and that countries which were more tightly integrated before adopting the euro gained more. Baldwin and Taglioni (2004) propose a model to explain this regularity. They show that countries which have lower trade barriers before the introduction of a common currency have larger expected gains. This implies that countries with close proximity to other euro-area countries or which trade a lot with other euro-area countries have larger expected gains for exports. Both of the sample countries in this study export the majority of their products to the EU, though Slovakia is more tightly integrated in trade with the EU than Estonia is. Slovakia sent 86 percent of its exports to the EU before adopting the euro (Eurostat indicator ext_lt_intratrd from 2008) and Estonia sent 69 percent of its exports to the EU before it adopted the euro (Eurostat indicator ext_lt_intratrd from 2010). Within the EU, Slovakia is again more tightly connected to the euro area, exporting 56 percent of its EU trade to the euro area, while Estonia exports 46 percent. Slovakia is a neighbor of one euro-area country, Austria, and is close to such large euro-area countries as Germany, France, and Italy. Estonia had one euro-area neighbor, Finland, which is also one of its main trading partners, but the rest of Estonia's main trading partners were not in the euro area at the time of the changeover.

This implies that the potential gains for trade from adopting the euro are larger for Slovakia than those for Estonia, and the main reason for this is that Slovakia had a floating exchange rate before adopting the euro, and Estonia did not. Table 1 summarizes the main channels behind the gains for trade from common-currency areas. The three main channels that can reduce transaction costs all have a positive effect on trade. The increased competition channel has a negative effect on trade as export prices are reduced, export markets become more transparent, and product markups are reduced. Unlike when the euro was introduced in 1999, our sample countries did not face increased competition from other countries because they joined the euro area one country at a time.

Table 1. Expected Gains in Trade from the Euro

Channel of Euro Impact on Trade	Expected Direction of the Effect	Introduction of the Euro in 1999	Changeover to the Euro in Slovakia in 2009	Changeover to the Euro in Estonia in 2011
Transaction Costs from Exchange Rate Volatility	(+)	Strong	Strong	No
Transaction Costs from Foreign Exchange	(+)	Variable	Strong	Medium
Interaction of Transaction Costs and Importance of the Euro Area in Trade Prior to Accession	(+)	Variable	Strong	Medium
Increased Competition from Other Euro-Area Members	(–)	Strong	No	No

Source: Compiled by the authors from related literature.

3. Data

We use detailed firm-level trade and balance sheet data for Slovakia and Estonia. These two countries represent the new Central and Eastern Europe (CEE) and Baltic euro-area countries well in terms of their level of development or trade openness.[4] As with the old euro-area countries, high levels of confidentiality for the detailed transaction data mean that strict data-handling rules are required, and these prevent cross-country combination of data sets.[5]

We use customs data on all exporting firms in Slovakia and Estonia, covering the NC8 codes for products, the ISO codes for destination countries, and the FOB values of the export flows. The data represent fairly exhaustive information on the exports of the countries analyzed, running between 2006 and 2011 for Slovakia and between 2008 and 2013 for Estonia. We aggregate the eight-digit NC codes to actual six-digit HS codes to ensure better comparability of product codes over time. The data are of very high quality, as the same administrative data have been used by national statistical institutions to produce official trade statistics.

The customs and commercial register data set is combined with the balance sheet data. We use real value-added, the real book value of net capital, employment, and material inputs to calculate firm-level total factor productivity (TFP). The TFP is calculated using the GMM-based approach suggested by Wooldridge (2009).[6] The

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[4]They both represent small, highly open economies. Cyprus and Malta differ significantly in their trade openness based on trade in goods.

[5]According to Castellani and Koch (2015), firm-level trade data are in general available for all seven new euro-area member states except Cyprus, but their confidential and restrictive accessibility rules make them difficult to access.

[6]We follow Dhyne et al. (2014) and implement the approach as in Wooldridge (2009) and Galuscak and Lizal (2011) that relies on the GMM framework of the specification suggested by Levinsohn and Petrin (2003). Production functions were estimated at the industry level, i.e., all firms in the same industry (and country) were assumed to have the same marginal returns of labor and capital. The industries are defined at the two-digit NACE level. The simultaneity of productivity and production inputs is addressed by introducing the polynomial of capital and material costs and GMM-type instruments for labor. Firm-level TFP was then calculated as the difference between the actual and predicted value-added taking into account the firm's values of labor, capital, and material costs. Given the unavailability of firm-specific price deflators, only industry-level ones, the TFP used in our regressions represents a revenue-based productivity measure.

real values are derived using GDP deflators at the two-digit NACE level. Interest paid and profits are used to derive a debt burden indicator that accounts for a financial situation effect. The balance sheet data are harmonized across countries using an approach that originates from the CompNet microdata project.[7] To ensure better compatibility of Slovakian and Estonian data, we use a sample of firms with 20 or more employees and firm-destination trade flows that are 1,000 euros per year or larger. As exports are highly concentrated, we still cover 99 percent of total exports in Slovakia and 95 percent in Estonia.

Tables 2 and 3 provide descriptive statistics about all the trade margins and explanatory variables analyzed. The descriptive statistics have been provided for the treatment and control group destination countries and for the period before and after the changeover to the euro. There is no evidence that average unconditional trade margins have developed differently for the treatment and control groups, as exports have increased to both of the destination country groups and the increase has been even faster in non-euro-area markets. The sample countries are similar in terms of the probability of a firm being an exporter, firm age, and the share of foreign-owned firms, though Slovakian firms are somewhat larger than Estonian firms and export larger volumes.

Additional aggregate explanatory data on macroeconomic indicators come from publicly available databases published by the International Monetary Fund and Eurostat.

4. Methodology

The aim of this paper is to investigate how joining a common-currency area affects trade. Following the new trade theory, we consider three types of adjustment: firm selection to export; changes in product varieties, which represent extensive margins; and changes in the average value of exports, which represent an intensive

[7]See Dhyne et al. (2014) for more details on the definition of variables and outlier treatments. This source also discusses the methodology for the TFP calculation.

Table 2. Descriptive Statistics of the Main Variables (EU trade), Slovakia 2006–11

	Control Group: EU Non-Euro-Area Countries				Treatment Group: EU Euro-Area Countries			
	Before		After		Before		After	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Share of Exporters in Each Destination (n = 86,332)	0.519	0.500	0.490	0.500	0.486	0.500	0.471	0.499
Share of Exporters in Each Destination × Product (n = 632,223)	0.265	0.441	0.245	0.430	0.266	0.442	0.257	0.437
Number of HS6 Products per Destination (n = 32,991)	5.125	7.200	4.862	7.146	4.963	8.503	4.768	9.289
Average Exports per HS6 Product in Destination (ths. EUR) (n = 32,991)	394.0	2,098.7	539.6	4,137.6	589.5	2,605.0	774.8	6,423.7
Total Exports per Destination (ths. EUR) (n = 32,991)	2,128.4	15,978.5	1,733.4	10,342.1	3,420.7	25,437.0	3,179.7	27,313.1
Firm Age (Years) (n = 32,991)	11.6	4.5	13.0	5.0	11.3	4.6	12.8	5.2
Firm Employment (n = 32,991)	381.2	1,009.0	315.4	719.8	433.0	1,108.4	339.3	779.0
Share of Foreign-Owned Firms (n = 32,991)	0.430	0.495	0.512	0.500	0.486	0.500	0.559	0.497
Firm Log(TFP) (n = 32,991)	−0.049	1.616	−0.023	1.614	−0.282	1.512	−0.214	1.594
Firm Debt Burden (n = 32,991)	0.237	0.246	0.207	0.235	0.237	0.243	0.198	0.232

Source: Authors' calculations from commercial register and customs data.
Notes: Foreign-owned firms are defined as a binary variable where majority foreign-owned firms take the value "1" and the rest "0." The firm debt burden represents interest paid divided by operating profit.

Table 3. Descriptive Statistics of the Main Variables (EU trade), Estonia 2008–13

	Control Group: EU Non-Euro-Area Countries				Treatment Group: EU Euro-Area Countries			
	Before		After		Before		After	
	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.	Mean	Std. Dev.
Share of Exporters in Each Destination (n = 12,014)	0.597	0.491	0.634	0.482	0.579	0.494	0.638	0.481
Share of Exporters in Each Destination × Product (n = 73,656)	0.378	0.485	0.450	0.497	0.385	0.487	0.455	0.498
Number of HS6 Products per Destination (n = 5,979)	4.1	6.8	5.0	8.2	4.4	8.5	5.2	10.1
Average Exports per HS6 Product in Destination (ths. EUR) (n = 5,979)	215.6	621.8	268.6	1034.1	372.7	1013.8	498.1	1349.9
Total Exports per Destination (ths. EUR) (n = 5,979)	665.0	2,124.4	2,414.5	37,280.0	1,025.1	2,903.2	1,636.7	5,391.4
Firm Age (Years) (n = 5,979)	13.7	4.1	15.9	5.2	13.5	4.1	15.7	5.2
Firm Employment (n = 5,979)	119.5	193.4	126.1	183.7	123.2	182.5	138.4	210.1
Share of Foreign-Owned Firms (n = 5,979)	0.358	0.480	0.411	0.492	0.375	0.484	0.428	0.494
Firm Log(TFP) (n = 5,979)	−0.061	1.948	0.077	1.855	−0.337	2.004	−0.149	1.906
Firm Debt Burden (n = 5,979)	0.143	0.225	0.078	0.155	0.126	0.218	0.072	0.146

Source: Authors' calculations from commercial register and customs data.
Notes: Foreign-owned firms are defined as a binary variable where majority foreign-owned firms take the value "1" and the rest "0." The firm debt burden represents interest paid divided by operating profit.

margin. The unit of observation is the trade flow to a particular destination country at the firm level, or firm times destination market.

In the baseline estimation strategy, we start with the probability of a firm exporting using a within-fixed-effect estimator.[8] The dependent variable in these regressions takes the value of 1 if the firm exports to a particular destination market and 0 otherwise.

In the next step, we continue by estimating the effect of the euro on the product margin (the number of six-digit HS code products exported for each firm in a destination market), the intensive export margin (the average value of exports of a six-digit HS code product for each firm in a destination market), and the total firm exports in a destination market using a fixed-effect estimator.

We follow the methodology of Berthou and Fontagne (2013), but in addition to their approach we introduce a dynamic specification with lagged dependent variable where the persistence of the export margin is controlled for, and we introduce industry-specific year dummies at the two-digit NACE level. The euro effect is identified by a difference-in-differences style dummy variable that is equal to 1 during the period following the adoption of the euro if the destination country was a member of the euro area, and 0 otherwise. We compare exports to the euro-area countries with exports to the remaining non-euro-area EU countries, so destination markets outside the EU are removed from the control group to ensure better comparability of the treatment and control groups. The number of EU members was 27 during the sample period, so excluding the home country results in 26 countries, of which 15 were euro-area members at the time when Slovakia introduced the euro and 16 at the time when Estonia did so. Only manufacturing firms are used in the estimations, as these are responsible for the majority of trade in goods.

[8]This is in contrast to the logit approach, which estimates the effect of independent variables on the probability of the firm changing its status from non-exporter to exporter, meaning it takes into account only information for the firms that change their status (switchers), while the within-fixed-effect approach keeps all the observations, meaning it takes into account both switchers and non-switchers. The logit model with only switchers in the sample has been estimated for robustness.

The following dynamic specification is applied:

$$\begin{aligned} TM_{ijt} = {} & \alpha_{ij} + \beta_1 TM_{ijt-1} + \beta_2 Post_t \times EA_{ij} + \beta_3 \log(TFP_{ijt-1}) \\ & + \beta_4 \log(\mathrm{GDP}_{jt}) + \beta_5 \log(\mathrm{REER}_{jt}) + \beta_6 \log(\mathrm{MP}_{jt}) \\ & + \tau_t \times sector_k + e_{ijt}, \end{aligned} \tag{1}$$

where i denotes the firm, j is the destination country, t is the year, and k is the industry at the two-digit NACE level. TM_{ijt} stands for the trade margin. Two types of trade margin are used, one of which is a binary variable capturing the whether the firm exports or not. The estimation of equation (1) for this type of trade margin covers firms that have exported at least once during the sample period. Trade margins of the other type are continuous variables and are defined only for positive firm-destination-level trade flows. As these trade margins have values larger than zero, logarithm has been taken of these margins. We prefer this two-part approach over the selection model, as it is proven to be robust to endogenous selection and avoids an often fruitless search for instruments that affect the decision to export but not the value of exports (see Drukker 2017 for the formal presentation and Nitsch and Pisu 2008 for the discussion of these issues). The firm-destination fixed effects are controlled for and are denoted by α_{ij}.

$Post_t \times EA_{ij}$ represents a combination of two dummy variables: $Post_t$ is equal to 1 after the home country joined the euro area (for the period 2009–11 for Slovakia and 2011–13 for Estonia), and 0 otherwise; and EA_{ij} is equal to 1 if the destination country was a member of the euro area at the time of the changeover, and 0 otherwise. The difference-in-differences effect of adopting the euro is captured by the coefficient β_2 and has a statistically significant positive value if the common-currency area increases the export margin.

The lagged TFP at the firm level controls for the dynamics of firm-level productivity as more productive firms are expected to enter export markets more likely by new trade theory (Melitz 2003). In order to isolate the effect of the euro from other economic factors, we control for a number of macrovariables in the destination country: gross domestic product $log(GDP_{jt})$, the real effective exchange rate $log(REER_{jt})$, and import prices $log(MP_{jt})$. GDP is expected to control for demand in the destination country, the real effective exchange rate for price competitiveness in the destination country, and import prices for the potential effect of imports from third countries.[9] Here we follow Berthou and Fontagne (2013), who used the same set of macro controls and lagged TFP. To control for the remaining industry-level developments in export markets, the industry-specific year dummies $\tau_t \times sector_k$ are added. The industry-specific year dummies also capture possible developments in the domestic economy that can induce firms to export. The standard errors e_{ijt} are clustered at the firm and destination levels and are expected to have conventional properties.

Our empirical specification builds on theoretical models of heterogeneous and multiproduct exporters. We analyze the effect of reduced trade costs on various trade margins that allows testing for predictions of models with multiproduct exporters (Bernard, Redding, and Schott 2011). These models predict that introduction of the euro and reduction of transaction costs leads to new entries to euro-area destination markets and a larger number of products exported to euro-area markets, but not necessarily larger exports per firm. Our empirical specification follows the structural gravity model and we control for the time-invariant inward multilateral resistance terms by introducing firm-destination fixed effects.[10]

We apply the decomposition of Berthou and Fontagne (2013) to disentangle the euro effect into that from the product-intensive and product-extensive margins. This approach allows to test the role of the product margin or the newly traded goods channel (Baldwin and Taglioni 2004) in the effect of the common-currency area on trade. First, three separate regressions are estimated for continuous trade margins as the logarithm of total exports for each destination, the logarithm of the number of products exported to each destination, and the logarithm of the average value of exports for each product in

[9]The role of industry-level import prices at the two-digit NACE level in the destination country has also been tested, but as the results were similar to the ones with country-level import prices, the latter have been used throughout the paper.

[10]For a full control in the panel setting, a proxy for the time-varying multilateral resistance term would be needed. However, in the time span of our analysis these resistance terms change very rarely and only little. Unlike in the aggregate country- or product-level specification, we cannot control for outward multilateral resistance terms, as the database covers only one exporting country at the time.

destinations. The total effect on the value of exports is decomposed as follows:

$$\frac{\partial \log(X_{ijt})}{\partial Post_t \times EA_{ij}} = \frac{\partial \log(N_{ijt})}{\partial Post_t \times EA_{ij}} + \frac{\partial \log(\bar{x}_{ijt})}{\partial Post_t \times EA_{ij}}, \tag{2}$$

where X_{ijt} denotes total exports to each destination, N_{ijt} the number of products for each destination, and $\bar{x}_{ijt}$ the average value of exports in destinations. The first term on the right-hand side of equation (2) captures the effect from the new products exported, or the product-extensive margin, and the second term captures that from the average value of exports per product, or the product-intensive margin.

5. Results

5.1 Baseline Results and Intensive vs. Extensive Margin

The estimation results for equation (1) on all the trade margins are presented in tables 4 and 5. Our results show that the euro has a positive trade effect across all the margins except number of products for Slovakia, but only for the probability of exporting for Estonia. The finding that the euro has no statistically significant effect on overall firm-level trade in Estonia but that the decision to export to new destination markets is affected can be related to experimentation in new markets with little export value that does not stand out in the total exports of firms. The euro increased the probability of exporting into euro-area destination markets by 1.8 percent in Slovakia and by 4.7 percent in Estonia.[11] We interpret the long-run effects and not the short-run effects of our dynamic specification, which provides better reference to the related literature that usually does not use dynamic specification. These effects are in line with

[11]Given our dynamic specification, the coefficients-on-treatment variable, β_2, refers to the short-run effect, and the long-run effect is calculated as $\beta_2/(1-\beta_1)$; see equation (1) for notation. The relatively low persistence of trade margins implies that usually the long-term effects are quite close to the short-term effects.

Table 4. The Euro Effect on Firm-Level Exports, Slovakia 2006–11, Manufacturing Firms, Within-Group Estimation

	Export Decision in Each Destination	Number of Products per Destination, N_{ijt}	Average Export Value per Product in Destination, $\bar{x}_{ijt}$	Total Exports per Destination, X_{ijt}
Lagged Dependent	0.045***	0.133***	0.177***	0.228***
	(0.005)	(0.010)	(0.010)	(0.011)
$Post_t \times EA_{ij}$	0.017**	0.020	0.111***	0.130***
	(0.007)	(0.014)	(0.030)	(0.032)
Log(TFP$_{ijt-1}$)	−0.005	0.003	0.029	0.020
	(0.005)	(0.010)	(0.023)	(0.024)
Log(GDP$_{jt}$)	0.197***	0.124	0.677***	0.742***
	(0.038)	(0.086)	(0.186)	(0.193)
Log(MP$_{jt}$)	0.073	0.098	−0.427*	−0.336
	(0.054)	(0.119)	(0.255)	(0.264)
Log(REER$_{jt}$)	−0.185***	−0.158	−0.639**	−0.740***
	(0.054)	(0.118)	(0.254)	(0.267)
Year × Sector FE	Yes	Yes	Yes	Yes
Firm × Destination FE	Yes	Yes	Yes	Yes
Observations	86,332	32,991	32,991	32,991
No. of Objects	20,535	10,523	10,523	10,523
Within R^2	0.014	0.0437	0.0849	0.1140

Source: Authors' calculations from the commercial register and customs data.
Notes: *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively. Clustered standard errors are in parentheses.

Table 5. The Euro Effect on Firm-Level Exports, Estonia 2008–13, Manufacturing Firms, Within-Group Estimation

	Export Decision in Each Destination	**Number of Products per Destination, N_{ijt}**	**Average Export Value per Product in Destination, $\bar{x}_{ijt}$**	**Total Exports per Destination, X_{ijt}**
Lagged Dependent	0.108***	0.183***	0.234***	0.250***
	(0.014)	(0.025)	(0.022)	(0.024)
$Post_t \times EA_{ij}$	0.042**	0.005	0.002	0.004
	(0.020)	(0.034)	(0.066)	(0.067)
Log(TFP$_{ijt-1}$)	−0.000	−0.020	0.055	0.032
	(0.009)	(0.016)	(0.038)	(0.038)
Log(GDP$_{jt}$)	0.006	0.126	1.328***	1.433***
	(0.094)	(0.162)	(0.394)	(0.390)
Log(MP$_{jt}$)	0.384**	0.341	−0.817	−0.471
	(0.160)	(0.318)	(0.642)	(0.642)
Log(REER$_{jt}$)	0.477***	0.269	−1.552**	−1.279**
	(0.180)	(0.350)	(0.730)	(0.736)
Year × Sector FE	Yes	Yes	Yes	Yes
Firm × Destination FE	Yes	Yes	Yes	Yes
Observations	12,014	5,979	5,979	5,979
No. of Objects	3,542	2,262	2,262	2,262
Within R^2	0.045	0.111	0.129	0.150

Source: Authors' calculations from the commercial register and customs data.
Notes: *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively. Clustered standard errors are in parentheses.

previous findings, such as the increase of a couple of percent from Belgian data (Nitsch and Pisu 2008).[12]

For total exports, we find that adopting the euro had a statistically significant and relatively strong economic impact in Slovakia, but no effect in Estonia. The euro increased the exports of Slovakian manufacturing by 18 percent ($\exp(0.13/(1 - 0.228)) - 1$), which is a large effect in comparison with results published on the introduction of the euro in 1999. For example, Baldwin (2006) concludes that the feasible macro-level euro trade effect findings based on gravity-type models are between 5 percent and 10 percent, while from microdata, Berthou and Fontagne (2008a) find that the euro increased exports by 5 percent, but Nitsch and Pisu (2008) and de Nardis, Pappalardo, and Vicarelli (2008) find there to be no effect. Our results from the Slovakian data are clearly from the upper bound of feasible effects. The main reason for the large effect in Slovakia is that this country benefited strongly along all the channels that have potential for positive gain, while it did not face increased competition from the other countries.

Our results indicate that the euro effect mainly manifested itself via the intensive margin and only partially via the decision to export new products. The euro effect on the average export per products is 14 percent ($\exp(0.111/(1 - 0.177)) - 1$) in Slovakia, and it accounts for almost 80 percent of the total increase in exports. This result is in line with the findings of Berthou and Fontagne (2008a), who also find the effect of newly traded goods to be less than 20 percent, while it is in contrast to the findings of Nitsch and Pisu (2008), who find that the euro increased newly traded goods but that there was no statistically significant effect on overall firm-level trade.

The micro-level control variables have the expected signs, all the export margins tend to have some persistence, and the lagged TFP, if statistically significant, has a positive effect on the trade margin. This gives support to our dynamic specification. Among the

[12]The panel fixed effects logit model with only the export decision of switchers in the sample shows an even stronger effect of the euro changeover to trade, but the statistical significance is unchanged. In this model the probability of exporting to a new euro-area destination increased by 11.3 percent in Slovakia and 5.7 percent in Estonia.

macro-level control variables, destination market GDP has a positive effect on the trade margin, while the price competitiveness proxy (REER) and import prices (MP) have varying effects depending on the country and specification.

The introduction of the euro reduced exchange rate volatility in Slovakia but not in Estonia, which suggests that the transaction costs channel from exchange rate volatility is important in the manifestation of gains from common currencies. In sum, our findings suggest that the transaction costs channel, both from exchange rate volatility and from foreign exchange, was an important mechanism behind the gains from trade due to adoption of the euro. This is not something that has been confirmed so far. Baldwin et al. (2008), who summarize the previous literature, conclude that non-euro-area countries in the EU did not face trade diversion after the introduction of the euro and use it as evidence that the transaction costs channel was not the main one. They claim that the main mechanism was increased competition and improved market transparency in euro-area countries, and that the newly traded goods hypothesis had a potentially important role. We find that the newly traded goods hypothesis accounted for only 20 percent of the total increase in trade, while we can exclude the increased competition channel from our empirical setup and confirm the strong effect from reduced transaction costs.

5.2 Results over Firm Characteristics

It was shown that the intensive margin has dominated the effect of the euro on exports in our sample countries. This subsection tests whether the effects have been heterogeneous over firm productivity and size, and also over age, ownership, and debt. The heterogeneity of the effects has been tested by interacting the treatment dummy with firm characteristics before the euro was adopted. We start with the total factor productivity (TFP), which in theory has been the most important determinant of entry to export markets. The firms have been divided into four TFP quartiles based on their average TFP three years prior to accession. In contrast to our baseline specification, this approach helps us to address possible nonlinearity.

The results are shown in table 6, where only the interaction terms with treatment are presented, as the rest of the coefficients do not

Table 6. The Euro Effect over Firm Characteristics, Slovakia 2006–11 and Estonia 2008–13, Manufacturing Firms, Within-Group Estimation

	Slovakia		Estonia	
	Export Decision in Each Destination	**Total Exports per Destination, X_{ijt}**	**Export Decision in Each Destination**	**Total Exports per Destination, X_{ijt}**
Regression with TFP Quartiles				
$Post_t \times EA_{ij} \times TFP\_q1_i$	−0.001	0.035	−0.011	−0.207**
	(0.011)	(0.056)	(0.029)	(0.087)
$Post_t \times EA_{ij} \times TFP\_q2_i$	0.022**	0.095**	0.111***	0.114
	(0.010)	(0.044)	(0.029)	(0.098)
$Post_t \times EA_{ij} \times TFP\_q3_i$	0.017*	0.148***	0.056**	0.034
	(0.010)	(0.045)	(0.029)	(0.096)
$Post_t \times EA_{ij} \times TFP\_q4_i$	0.026***	0.191***	−0.016	0.096
	(0.010)	(0.048)	(0.033)	(0.097)
Regression with Size Groups				
$Post_t \times EA_{ij} \times Size\_1_i$	0.029**	0.103	0.051*	0.007
	(0.011)	(0.064)	(0.027)	(0.083)
$Post_t \times EA_{ij} \times Size\_2_i$	0.023***	0.157***	0.036	−0.005
	(0.008)	(0.037)	(0.023)	(0.073)
$Post_t \times EA_{ij} \times Size\_3_i$	−0.006	0.100**	0.043	0.054
	(0.010)	(0.043)	(0.040)	(0.167)
Regression with Age				
$Post_t \times EA_{ij}$	0.017***	0.125***	0.046**	−0.012
	(0.007)	(0.032)	(0.020)	(0.067)
$Post_t \times EA_{ij} \times Young_i$	−0.018	0.126	−0.050	0.194
	(0.023)	(0.113)	(0.074)	(0.266)
Regression with FDI				
$Post_t \times EA_{ij}$	0.006	0.118***	0.059***	−0.021
	(0.008)	(0.038)	(0.023)	(0.074)
$Post_t \times EA_{ij} \times FDI_i$	0.024***	0.020	−0.040	0.037
	(0.009)	(0.043)	(0.028)	(0.081)
Regression with Debt				
$Post_t \times EA_{ij}$	0.020**	0.178***	0.030	−0.010
	(0.009)	(0.045)	(0.022)	(0.071)
$Post_t \times EA_{ij} \times Debt_i$	−0.005	−0.186*	0.133**	0.089
	(0.021)	(0.106)	(0.067)	(0.221)

Source: Authors' calculations from the commercial register and customs data.
Notes: The table presents only the coefficients of the interaction terms with the treatment variable and productivity or size. The rest of the control variables, not presented, are the same as in the baseline estimations or in equation (1). Separate regressions are estimated for each trade margin, and for each firm characteristic and its interaction. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively. Clustered standard errors are in parentheses.

differ much from the baseline estimates and are not shown. The results for Slovakia show some nonlinearity of the effect in terms of the decision to export, as firms from the second and fourth productivity quartiles have gained the most. The overall effect on the value of exports is strongest for the most productive firms, while the effect is already statistically significant in the second productivity quartile. The results for Estonia show that it was firms from the second productivity quartile that started to export to new destinations after the euro was adopted. This also explains why the effect does not show up in total exports, as it was rather less productive firms that started to export.

Our results indicate that the gains from adopting the euro were more equally distributed than in previous studies. Berthou and Fontagne (2013) find that the effects were concentrated in the most productive firms from the fourth productivity quartile, while Nitsch and Pisu (2008) find that less productive firms gained the most. Our results show that firms from the second to the fourth productivity quartile gained and that the effect was the strongest for the most productive firms. In this sense our findings are in line with the theory of Melitz (2003), which predicts that a reduction in trade costs allows a wider set of more productive firms to start exporting.

A similar exercise is to test whether the effect differed across firm size. Esteve-Perez et al. (2011) claim that only small firms experienced trade gains from the introduction of the euro, a finding that Nitsch and Pisu (2008) confirm. In this paper the firms are divided into three size groups: small firms with 20 to 49 employees, medium firms with 50 to 249 employees, and large firms with 250 or more employees. As in the exercise with productivity, the average firm size three years prior to adoption of the euro is calculated, and from this firms are allocated into three size groups. The results are presented in table 6. For Slovakia, where we find the euro has a strong effect on trade, we confirm the previous findings for new destination markets, where smaller firms started to export to new markets after the introduction of the euro. However, the gains over export volumes are quite equally distributed across firm size. The results from the Estonian sample are statistically insignificant, as in the baseline estimation.

Lastly, we test whether the gains from the euro have been distributed equally over other firm characteristics such as firm age, ownership, and debt burden. There is no theoretical evidence that the reduction in trade costs has a varying effect over these firm characteristics. It is rather that these estimates indicate whether trade costs differ across firms with different characteristics. The interaction terms with the treatment dummy are mostly statistically significant.

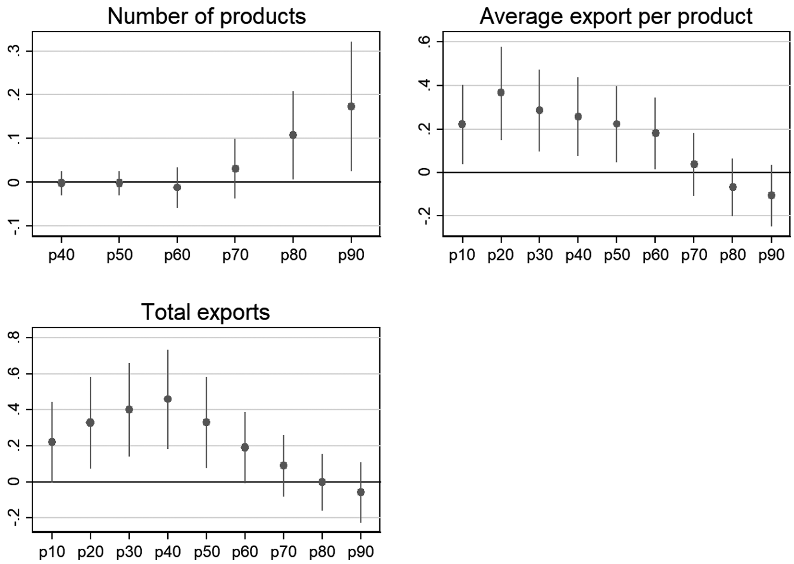
5.3 Effects over the Distribution of Exports

This subsection, like the previous one, tests the heterogeneity of gains in trade from the introduction of the euro. Here we test whether the gains in trade from the reduction of transaction costs differ across the outcome variable, which is exports. As the effects on the distribution of the outcome variable are estimated, the data with positive trade flows are used and the binary variables such as the decision to export are not analyzed.

The unconditional quantile regression by Firpo, Fortin, and Lemieux (2009) is applied and the xtrifreg command by Borgen (2016) is used to implement the panel estimations with fixed effects in Stata. This method allows us to estimate how the explanatory variables affect the unconditional distribution of the outcome variable by using the recentered influence function technique. The advantage of this method is that unlike the conventional quantile regression, where the results are interpreted in terms of the conditional distribution of the outcome variable, this approach allows much more intuitive interpretation of the results in terms of the unconditional distribution of the outcome variable. The unit of observation is firm-level exports to a destination country, as in the previous sections. The same specification as in equation (1) has been used and the estimations have been run for nine quantiles.

Figures 1 and 2 present the results for Slovakia and Estonia, respectively. Only the effects of the treatment dummy on the export margin are presented, and the rest of the coefficients are not shown. The results confirm the finding that the euro had strong effects on trade in Slovakia and no effect in Estonia. Most importantly, the distribution of effects for Slovakia is cardinally different along the extensive margin and the intensive margin. On the extensive margin it is shown that those firms that already exported a large number of products to a market started to export new products following the introduction of the euro. The newly traded goods hypothesis

Figure 1. The Distribution of the Euro Effect on Exports, Slovakia 2006–11, Manufacturing Firms



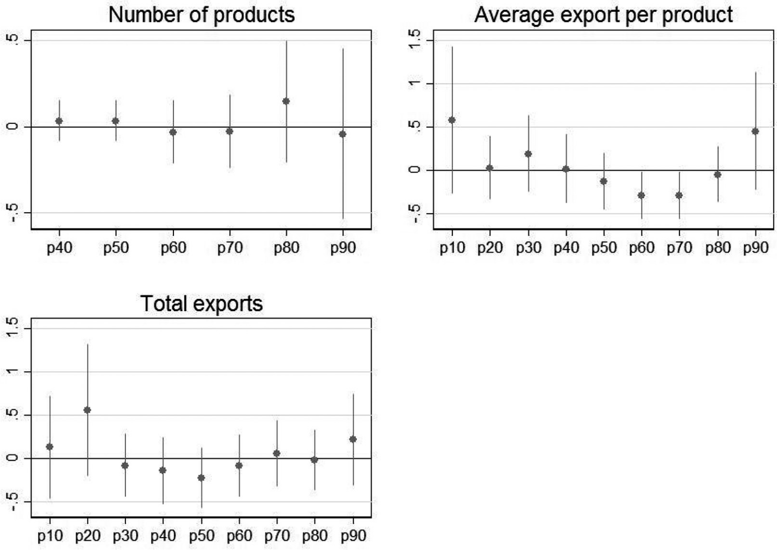
Source: Authors' calculations from the commercial register and customs data.
Notes: Each coefficient on the figure represents one regression for the particular percentile, e.g., p10 shows the effect of the euro on exports at the 10th percentile of firm-destination export flows. The confidence intervals reflect statistical significance at 10 percent. The figure presents the long-term effects, i.e., $(\exp(\beta_2/(1-\beta_1))-1)$ according to specification (1).

seems to be an important channel for benefits for firms that already export a lot of products or for destinations which are already served by many products.[13]

In contrast to the case of the extensive margin, it was rather smaller exporters or markets where small amounts were exported that gained most from the euro along the intensive margin. The effects are statistically significant up to the 60th percentile for

[13]It may be noted that the recentered influence function of the decile cannot be defined for this part of the distribution where there is no variation in the dependent variable, so the graph only starts from the 40th percentile for the product margin where only one product was exported.

Figure 2. The Distribution of the Euro Effect on Exports, Estonia 2008–13, Manufacturing Firms



Source: Authors' calculations from the commercial register and customs data.
Notes: Please refer to the notes of figure 1.

Slovakia. In the overall effect on trade, the intensive margin dominates over the extensive margin, so total exports also increased for smaller exporters or in destination markets where exports were small. The overall effect on exports is large and statistically significant up to the 50th percentile of firm-destination trade flows in Slovakia. Our results are in line with the predictions of the multiproduct models that a decline in transaction costs is related to an increase in the number of products exported, but not necessarily to an increase in the average export per product, as new products are traded in smaller volume (Bernard, Redding, and Schott 2011). We find that the total export did not increase for the largest exporters, which are usually exporting many products, but it increased for smaller exporters. The overall effect on trade is the strongest around mid-size exporters from the 40th percentile, where the effect is as large as 40 percent.

5.4 Effects over Sectors and Product Groups

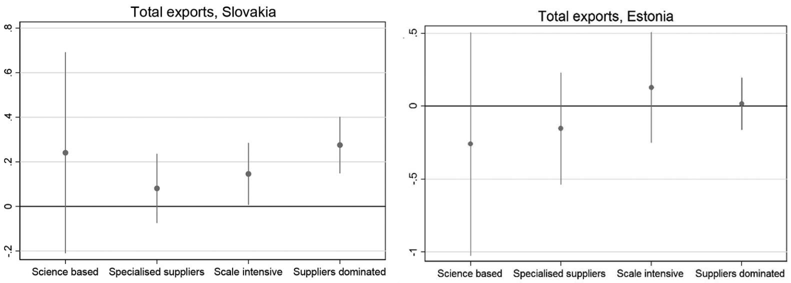
Several studies suggest that the gains from trade differ across sectors or macrosectors. Baldwin and Taglioni (2004) note that the exports of industries that are characterized by imperfect competition and increasing returns to scale increased more following the introduction of the euro than did the exports of industries working with natural resources or producing raw materials. Baldwin et al. (2008) summarize the macro-level sectoral findings by stating that the gains from trade from adopting the euro have been concentrated in a few industries, and most industries did not experience any increase in trade from the introduction of the euro. They take this argument as evidence against the transaction costs being an important channel behind the gains in trade from common currency. There is also evidence that the euro has enhanced vertical specialization and especially increased the trade in intermediate and final goods (Flam and Nordström 2007 and Martinez-Zarzoso and Johannsen 2017).

The firm-level studies do not provide much information on the sectoral distribution of the euro effects. De Nardis, Pappalardo, and Vicarelli (2008) find from Italian microdata that it was indeed the scale-intensive industries dominated by traditional goods or suppliers that experienced a boost to exports from the introduction of the euro. They use Pavitt's (1984) taxonomy to divide sectors into four groups and find that there was no effect in science-based industries and industries of specialized suppliers that produce specialized technology or inputs for other firms.

Our findings support the results of Baldwin and Taglioni (2004) and De Nardis, Pappalardo, and Vicarelli (2008). We find that scale-intensive and traditional sectors producing highly differentiated goods benefited the most from the introduction of the euro. The results for the total exports are presented in figure 3. We contribute to this discussion also by testing whether there are different euro effects for firms from different NACE two-digit industries. We observe that the euro effects on total exports are large and positive in the majority of industries for Slovakia, but are always statistically insignificant for Estonia (the results are available from the authors upon request).

In order to fully exploit the granularity of our trade data, we divide manufacturing firms into groups based on type of product

Figure 3. The Effect of the Euro on Sector Defined According to Pavitt's (1984) Taxonomy, Slovakia 2006–11 (LHS) and Estonia 2008–13 (RHS), Manufacturing Firms



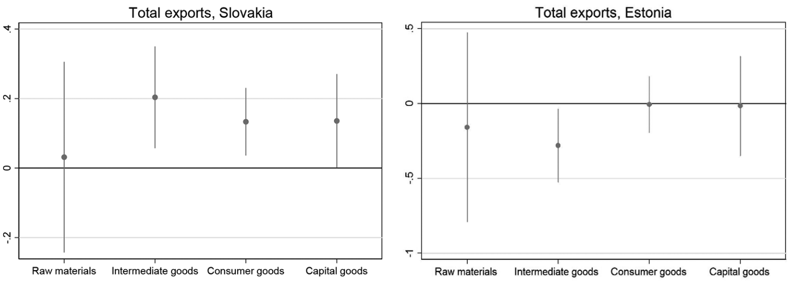
Source: Authors' calculations from the commercial register and customs data.
Notes: Each coefficient in the table represents one regression for the particular sector and trade margin. The confidence intervals reflect statistical significance at 10 percent. The figure presents the long-term effects, i.e., $(\exp(\beta_2/(1-\beta_1))-1)$ according to specification (1).

exported rather than sector of activity. We use the standard end-use product group categories of HS six-digit goods, which divide products into raw materials, intermediate goods, consumer goods, and capital goods.[14]

Figure 4 shows the results. The euro effect on total firm exports other than for raw materials is quite equally distributed across intermediate, consumer, and capital goods in Slovakia, but the effect on intermediate goods is the largest. These estimates confirm our previous findings that quite a broad range of industries benefited from the introduction of the euro, which supports our main argument that transaction costs from exchange rate volatility were driving the gains. Our results also support the macro-level findings that the trade in intermediate goods increased the most and the trade in raw materials did not increase following the introduction of the euro. This implies that the euro further increased the vertical specialization of trade in Slovakia. Both of our sample countries have among

[14] HS Standard Product Groups following UNCTAD statistical classifications of products (UNCTAD-SoP1, SoP2, SoP3, SoP4) were used. See https://wits.worldbank.org/referencedata.html for the reference

Figure 4. The Effect of the Euro on Different Goods, Slovakia 2006–11 (LHS) and Estonia 2008–13 (RHS), Manufacturing Firms



Source: Authors' calculations from the commercial register and customs data.
Notes: Each coefficient in the table represents one regression for the particular sector and trade margin. The confidence intervals reflect statistical significance at 10 percent. The figure presents the long-term effects, i.e., $(\exp(\beta_2/(1-\beta_1))-1)$ according to specification (1).

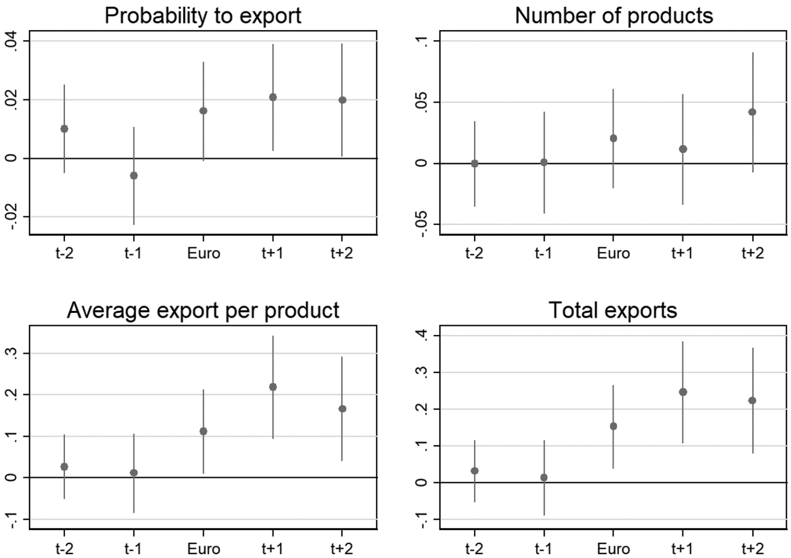
the highest degrees of vertical integration in the Organisation for Economic Co-operation and Development (OECD 2009).

5.5 Robustness Tests

In addition to the presented results over firm characteristics and distribution of exports, we run a number of robustness tests to validate the results. First, we perform the placebo tests over the time and the cross-section. Second, we test for the role of the estimation method and control for the Nickell's bias in our dynamic specification by GMM estimation.

Figures 5 and 6 present the placebo tests over time, i.e., the coefficients of the interaction of the euro-area dummy with year dummies. The placebo year effects are expected to be statistically insignificant before the euro was adopted, and they should become statistically significant after the adoption of the euro in order to support the causal interpretation of the results. The yearly effects allow also testing for the common trend assumption before the changeover. If the yearly effects are statistically insignificant before the changeover, it shows that the conditional trends in dependent variable are similar

Figure 5. Timing of the Effect of the Adoption of the Euro, Slovakia 2006–11, Manufacturing Firms

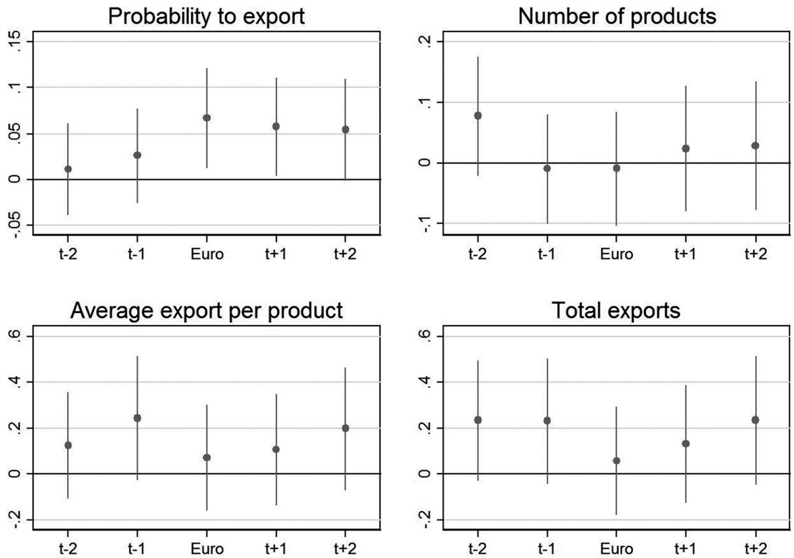


Source: Authors' calculations from the commercial register and customs data.
Notes: Each graph represents the results from one regression, where instead of one treatment dummy in equation (1), five treatment dummies have been used: the year two years before the accession time EA_{ij}; the year one year before the accession time EA_{ij}, and so on. The year three years before the accession has been used as a reference period, and this dummy is omitted from the regressions. The confidence intervals reflect statistical significance at 10 percent. The figure presents the long-term effects, i.e., $(\exp(\beta_2/(1-\beta_1))-1)$ according to specification (1).

for the treatment and control group prior to the changeover. These yearly effects also show the timing of the effect.

The results for Slovakia show that there were no differences between exporting to euro-area and non-euro-area countries before the changeover, while there was more export to euro-area countries after the changeover. The strongest effect along the intensive margin appears one year after the adoption of the euro, while the strongest effect for the number of products appears somewhat later at two years after the adoption of the euro. The results for the longer time span over five years show that the maximum effect along the product margin appeared three years after the euro was adopted. The effects

Figure 6. Timing of the Effect of the Adoption of the Euro, Estonia 2008–13, Manufacturing Firms



Source: Authors' calculations from the commercial register and customs data.
Note: Please refer to the notes of figure 3.

for all the trade margins remained in the same magnitude for five years after the euro changeover (these results are available from authors upon request). The effect is persistent for all the statistically significant cases, so it is not an on-off effect from the temporary experimentation but persists over the treatment period of three or five years. The results for Estonia are clearly statistically insignificant for the volume of trade, while like for Slovakia, the probability of exporting to new destinations increases immediately after the euro is introduced.

The placebo treatment over the cross-section is defined so that the sample is limited to non-euro-area export destinations and the treatment and control groups have been assigned randomly. The effect of this placebo treatment is expected to be statistically insignificant. Table 7 presents the results: the placebo treatment over the cross-section shows no statistically significant treatment effects.

Table 7. The Euro Effect on Total Firm-Level Exports per Destination, Manufacturing, Robustness Tests

	Slovakia			Estonia		
	Placebo Treatment, Non-Euro-Area Countries Only[a]	**Alternative Estimation Method, System GMM**	**Alternative Estimation Method, OLS**	**Placebo Treatment, Non-Euro-Area Countries Only**	**Alternative Estimation Method, System GMM**	**Alternative Estimation Method, OLS**
Lagged Dependent	0.225***	0.325***	0.859***	0.136***	0.288***	0.868***
	(0.016)	(0.017)	(0.003)	(0.031)	(0.043)	(0.007)
$Post_t \times Treatment_{ij}$	0.065	0.109***	0.016	0.020	−0.048	0.053
	(0.053)	(0.036)	(0.019)	(0.095)	(0.042)	(0.040)
Log(TFP_{ijt-1})	0.041	−0.086***	0.124***	0.054	−0.062	0.058**
	(0.036)	(0.024)	(0.011)	(0.058)	(0.049)	(0.027)
Log(GDP_{jt})	0.666***	0.479**	0.061***	0.703	1.683***	0.033***
	(0.256)	(0.216)	(0.006)	(0.472)	(0.392)	(0.011)
Log(MP_{jt})	−0.122	0.131	0.341	−0.289	−0.733	−0.738
	(0.364)	(0.300)	(0.217)	(0.843)	(0.755)	(0.454)
Log(REER_{jt})	−0.329	−0.170	−0.359*	−0.867	−0.513	−0.021
	(0.346)	(0.305)	(0.197)	(1.017)	(0.749)	(0.589)
Year × Sector FE	Yes	Yes	Yes	Yes	Yes	Yes
Firm × Destination FE	Yes	No	No	Yes	No	No
Observations	15,550	32,991	32,991	2,966	5,979	5,979
No. of Objects	4,922	10,523		1,140	2,262	
R^2	0.124		0.761	0.183		0.772
Sargan Test		5.526			1.651	
No. of Instruments		123			119	

Source: Authors' calculations from the commercial register and customs data.
Notes: [a]The treatment group consists of Denmark, Hungary, Lithuania, Sweden, and Romania; the control group consists of Bulgaria, Czechia, Latvia, Poland, and the United Kingdom. *, **, and *** denote significance at the 10 percent, 5 percent, and 1 percent level, respectively. Clustered standard errors are in parentheses.

This raises confidence that our specification is able to control for destination country-specific shocks to trade which occurred during the changeover but were not related to the changeover. This demonstrates that the euro effect is not just some residual trend in the data, but that it disappears when the treatment group is "wrongly" assigned.

The robustness of the estimation method has been tested by estimating the specification in equation (1) by ordinary least squares (OLS) and system GMM; the latter method addresses the Nickell's bias in our baseline fixed-effects estimates. Table 7 reports that the persistence of trade margins is underestimated with our default fixed-effects estimator, as expected, but the statistical significance and the size of the long-run effect, 18 percent $(\exp(0.109/(1-0.325)) - 1)$ for Slovakia, are unchanged.

6. Conclusion

This paper studies the effect of adopting the euro on firm-level exports using data from two recent euro changeovers in Slovakia and Estonia. It is the first paper to test the firm-level trade effects of the euro in countries that were not initial members of euro area. The contribution of the paper is twofold. First, the paper provides evidence of the effect of the euro on exports by studying two cases of changeovers where the trade costs of a new member were reduced, while the competition in the euro area was unaffected. By doing this we can point to the exact channel in action much better than previous studies based on the standard gravity model. Second, the paper provides new evidence for the scarce and inconclusive findings about the heterogeneous effect of the euro on exports.

Our findings for new euro-area countries using microdata show a relatively large positive trade effect from the adoption of the euro in Slovakia that has manifested itself mainly through the intensive margin, and almost no effect in Estonia. We find that joining the euro area increased Slovakian exports to the euro area by 18 percent. In contrast to the previous studies, we have the advantage of studying countries that adopted the euro separately, so we can abstract away from the effect of increased competition and consider only the channels of foreign exchange transaction costs or transaction costs related to exchange rate volatility. Taking into account the differences in the pre-euro exchange rate regimes in the countries analyzed, where Slovakia had a floating exchange rate with the euro and Estonia had a fixed rate, our results indicate the important point that the major part of the euro trade effect can be assigned to savings from the reduction in exchange rate volatility. This result, however, does not imply that countries with a fixed exchange rate to the common currency are not subject to gains from it. The gains from the transaction costs channel can arise much earlier at the time when exchange rate was fixed and the net gains during the changeover can be positive as the costs from giving up country-specific monetary policy are also lower.

The analysis of the heterogeneity of effects shows that the positive overall effect on the value of exports was the strongest for the most productive firms, but in contrast to previous studies we find gains to be more equally distributed across firm size. The euro changeover does not have a robust interaction effect with other firm characteristics such as firm age, foreign ownership, or debt burden. The results of the unconditional quantile analysis show that it was smaller and medium-sized exporters that increased their exports as a result of the changeover to the euro and they benefited mostly from the intensive margin. Our results indicate that scale-intensive and traditional sectors producing highly differentiated goods and exports of intermediate goods benefited the most from the introduction of the euro, while quite a broad range of industries and products benefited from the introduction of the euro. These results suggest that small and already very open economies can experience a wider distribution of gains and a wider distribution of exports from the reduction in trade costs. Various robustness tests, including estimation of the placebo effects along the time and cross-sectional dimension or using different estimation methods, confirm our baseline results.

Our results are encouraging for small open economies with floating exchange rates that are planning to join the euro area or any other currency union. If the reduction in trade costs is substantial, it can lead to a substantial increase in trade. The differences in the scale and the heterogeneity of the trade effect are an interesting space for further research.

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Household Wealth and Resilience to Financial Shocks in Italy*

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Financial shocks in a sector of the economy transmit to other sectors via financial linkages. This paper constructs the matrix of bilateral financial sectoral exposures in Italy over the last two decades. Using this information, it develops a method to simulate how each sector absorbs plausible financial shocks. A fall in the value of government bonds directly affects banks and indirectly affects households via equity holdings in banks. A bank bail-in is absorbed by foreigners and by households, particularly those at the top of the wealth distribution. Conversely, in a bank bailout these two groups benefit from a government transfer.

JEL Codes: G11, G32, G33.

1. Introduction

Italian household wealth is high by international standards. Total net household wealth at end-2013 was estimated at over €9 trillion, or 5½ times gross domestic product (GDP). Average wealth per household exceeds €350,000 and per capita is about €150,000 (Bank of Italy 2014). As a percent of disposable income, it is higher than in most euro-area peers, including Austria, Finland, France, Germany,

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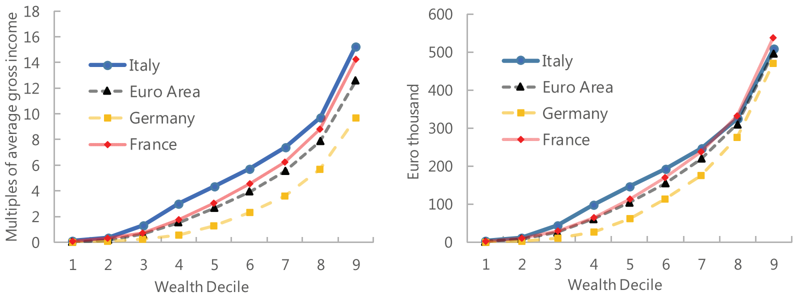
Figure 1. Average Net Wealth per Household by Wealth Decile, 2014



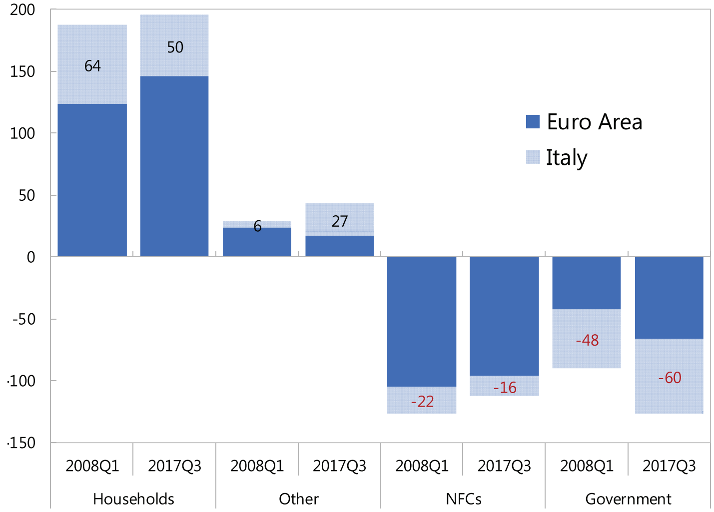
Notes: The left panel shows net wealth per decile divided by the average gross income across deciles. The dats are from the ECB Household Finance and Consumption Survey.

or Luxembourg. The middle and upper segments of the distribution are particularly wealthy compared with the euro-area average, both as a share of income and in absolute terms (figure 1). Real assets—principally dwellings—constitute almost two-thirds of total net wealth, while financial assets are mostly concentrated in cash and deposits, shares, and insurance reserves.

High net wealth of the household sector contrasts with weaker financial positions of other sectors. Sectoral net financial positions—defined as financial assets minus financial liabilities—in Italy display much larger imbalances than the euro-area aggregate (figure 2). Households stand out for their large positive balances, but the government features very negative financial wealth. The position of nonfinancial corporations is also strongly negative but more in line with the euro-area aggregate. Although real net household wealth has declined since the onset of the Global Financial Crisis (GFC), sectoral imbalances have tended to widen.

While high household net wealth or savings are a key strength of the Italian economy, negative financial wealth positions in other sectors could signal financial vulnerabilities. As a general matter, deeply indebted sectors may face difficulties raising further funding and experience higher default rates, thus turning into a source of contagion to other sectors. Indeed, the literature on sectoral fund flows has long emphasized the financial non-neutrality of sectoral

Figure 2. Net Financial Assets by Sector (percent of GDP)



Notes: "NFCs" denotes nonfinancial corporations, "Government" is the general government, and "Other" includes the rest of the world and the financial sector: monetary and financial institutions, other financial institutions, insurance companies, and pension funds. The combined height of the bars shows the position in Italy. The figures on the bars indicate the difference between Italy and the euro area. The data are from the Bank of Italy.

limits (e.g., Poterba 1987 documents the lack of a "corporate veil"). Understanding the transmission of shocks across sectors and their ultimate impact on households requires information about the intersectoral bilateral financial linkages. Shocks to the value of a given instrument issued by a given sector transmit to other sectors via direct asset exposures and via equity interlinkages. For example, a fall in the value of corporate debt can directly affect financial institutions holding that debt, and in turn transmit to households with claims on those financial institutions.

This paper constructs the matrix of bilateral financial sectoral exposures and simulates the impact of a series of illustrative financial shocks. Instrument-level intersectoral financial positions are inferred from the Bank of Italy's flow-of-funds data. The information on financial exposures is then illustratively used to infer the net financial impact across sectors of a 10 percent fall in the value of government bonds, equivalent to a two-standard-deviation change in sovereign spreads, and of a 10 percent fall in the value of bank liabilities (i.e., the assets held by other sectors in banks), equivalent to the combined liabilities of the third and fourth Italian banks. The shock to bank liabilities is successively modeled as leading to either a bail-in or a bailout. Household wealth survey data allow us to pin down the impact of these shocks across the household wealth distribution.

The matrix of bilateral exposures reveals that, since 1995, household wealth has been increasingly kept in insurance and pension fund assets and abroad. Households' direct exposure to the government has declined, although it is now intermediated by financial institutions. Government liabilities have been increasingly funded by the rest of the world and financial institutions, with an important contribution from the Bank of Italy in recent years, reflecting the ECB's quantitative easing. However, international financial diversification of Italian residents has slowed down. Since the GFC, nonfinancial corporate balance sheets have shrunk, with households being the sector that has reduced more its investment.

Within the household sector, the distribution of financial exposures is related to wealth levels. According to the Survey on Household Income and Wealth of the Bank of Italy, financial wealth and, in particular, risky assets are concentrated at the upper end of the distribution. The top two household wealth deciles accumulate more than two-thirds of financial wealth and an even larger proportion of equity and nonsecured debt. Less wealthy households own almost all their financial wealth in the form of insured bank deposits.

Given the bilateral instrument-level exposures, a fall in the value of government bonds is estimated to directly affect the financial sector and indirectly households. The Bank of Italy, private monetary financial institutions (MFIs, mostly banks), insurance and pension funds, and the rest of the world bear sizable balance sheet losses. However, as private financial institutions are ultimately owned by other sectors, primarily domestic and foreign households, these households—especially at the upper end of the wealth distribution—bear the brunt of the losses. Given the healthy financial position of households, together with the concentration of financial assets in wealthy households, the impact of the shock on the real economy should be small.

A bank bail-in is more able than a bailout to transmit to the rest of the world part of the shock to the value of bank liabilities. In a bail-in scenario, the burden of bank debt restructuring is shared by domestic and foreign households, as those are the ultimate holders of MFI equity and bonds. Here too, the burden falls mostly on the top wealth deciles.

On the contrary, in a bailout, the government would transfer resources to the wealthiest households and to the rest of the world. After a bailout, the top decile is less affected as a share of its financial wealth than the upper-middle range of the wealth distribution. From a simple arithmetic consideration, bailout interventions add to public debt, in turn imposing costs via taxes on labor income, which fall more evenly across the distribution than the wealth effects. Their countercyclical implications too are weak—as wealthier households have a lower marginal propensity to consume. Raising bank capital levels could help alleviate the need for, and cost of, a bailout. A progressive wealth tax could also undo the regressive distributional effects of the fiscal transfer implicit in a bailout.

In Italy, interest rates of bank and real-sector debt liabilities have been historically quite sensitive to changes in sovereign spreads (see Albertazzi et al. 2014). To capture the response to sovereign debt prices of the price of other *debt* securities, a correlated debt shock is simulated next. Note that this goes beyond the *equity* interlinkages already modeled above. Compared with an individual shock to government bonds, a correlated debt shock has a worse ultimate impact on the rest of the world and the central bank, while households' net financial position is less affected as their debt liabilities also depreciate.

Through these accounting exercises based on balance sheet exposures, the paper makes transparent some of the tradeoffs involved in the absorption of financial shocks. The paper simply calculates how changes in the prices of particular assets are absorbed in the balance sheets across sectors, regardless of what caused the price changes. This exercise abstracts from factors such as the real effects of shocks—which Bofondi, Carpinelli, and Sette (2018) and Cingano, Manaresi, and Sette (2016) estimate to be significant in Italy—portfolio reallocation in response to shocks, financial instability channels such as contagion across individual banks, or adjustment dynamics. On the other hand, policy decisions, particularly regarding bail-ins and bailouts, must take a comprehensive view and consider all transmission channels. Yet, the balance sheet analysis is sufficient to unambiguously show that the Italian household sector, and particularly households at the upper end of the distribution, have ample capacity to absorb plausible financial shocks. Government intervention aimed at preventing this absorption would be fiscally costly and ultimately of a regressive nature.

The literature has explored sectoral financial linkages in various contexts. Doepke and Schneider (2006) study the financial impact of inflation shocks across U.S. sectors and over the income distribution. Castrén and Kavonius (2009) construct the financial exposures matrix for the euro area with a similar method as this paper. Heipertz, Ranciere, and Valla (2017) estimate sectoral valuation linkages with security-level French data. Koijen et al. (2018) focus on the impact of quantitative easing (QE) by the European Central Bank (ECB) on sectoral portfolios. Lindner and Redak (2017) document the holdings of bail-in-able instruments across the wealth distribution of European households. Cortes et al. (2018) develop a method to estimate contagion across sectors within the financial system. The International Monetary Fund (2015) reviews the use of balance sheet analysis in the context of policy evaluation.

This paper contributes to the literature by developing a method to estimate the direct, indirect, and distributional impact of specific valuation shocks. Castrén and Kavonius (2009) use data with coarser sectoral information. More importantly, they only consider a (small) finite number of equity impact rounds, while the methodological contribution of this paper is to simulate the full equity impact. Estimating the full equity impact is necessary given sizable bidirectional equity holdings between sectors. It is also crucial when calculating distributional effects, as wealthier households invest disproportionately more in equity than in bonds. Heipertz, Ranciere, and Valla (2017) have access to more disaggregated data for France, but they do not focus on bank restructuring scenarios.[1]

[1] Hüser et al. (2018) study the implications of a bail-in for different types of creditors of the largest euro-area banks. Gourinchas, Martin, and Messer (2017) model wealth transfers among monetary union members in a bailout.

The rest of the paper is organized as follows. Section 2 describes the data and the accounting method. Section 3 documents financial exposures in Italy. Section 4 describes the method to simulate the impact of financial shocks. Section 5 presents the results of the simulation. Section 6 analyzes the distributional impact of financial shocks. Section 7 concludes.

2. Data and Accounting Method

Flow-of-funds data provide information on sectoral financial exposures. The Bank of Italy publishes quarterly flow-of-funds data (sourced via Haver) covering the period 1995:Q1–2017:Q3. This data set contains information, for each economic sector, on the stock positions in different financial instruments (assets and liabilities).[2] Table 1 lists the disaggregation of sectors in the data as well as the simplified grouping applied in this paper.

The data are used to construct a matrix of cross-sectoral bilateral financial exposures. A given entry (i, j) in the matrix contains the financial asset holdings of sector i invested in sector j, or equivalently the liabilities of sector j with respect to sector i. Appendix A describes the steps and necessary assumptions to infer bilateral sectoral exposures from the Italian flow-of-funds data. The data set only includes financial assets. A sector can have a nonzero net financial asset balance, which should be matched by an opposite net balance of real assets or by own-sector net worth (in the case of BOI, GOV, HH, and RoW).

Survey data allow us to zoom in to household financial exposures as a function of household wealth. The Bank of Italy's Survey on Household Income and Wealth (2016 release) contains financial information (and sample weights) for a representative sample of about 7,000 households. These data are used to calculate the distribution of stock positions and the impact of financial shocks across household wealth deciles. The ECB's Household Finance and Consumption Survey (2014 release) allows us to compare with the distribution in other euro-area economies.

[2]The classification follows the European System of Accounts (ESA) for 2010.

Table 1. Grouping of Sectors

Original Sectors in the Data	**Coding**
Nonfinancial Corporations	NFC
Monetary Financial Institutions Excluding Central Bank	MFI
Bank of Italy	BOI
Other Financial Intermediaries Excluding Non-MMF Investment Funds Non-MMF Investment Funds Financial Auxiliaries	OFI
Insurance Companies Pension Funds	INP
Central Government Local Government Social Security Funds	GOV
Households and Nonprofit Institutions Serving Households	HH
Rest of the World	RoW

3. Sectoral Financial Exposures in Italy

Table 2 contains the matrix of sectoral financial exposures in Italy in 2017:Q3, expressed as a percent of GDP. For example, nonfinancial corporates (NFCs) own financial assets of monetary financial institutions (MFIs) worth 14 percent of GDP, and equivalently MFIs have financial liabilities of 14 percent of GDP with respect to NFCs. The rightmost column shows the net financial asset (NFA) position of each sector, equal to total assets minus total liabilities.[3]

The distribution of sectoral financial linkages is highly non-uniform. Table 2 shows that NFCs have a very negative NFA

[3]Summing up all sectors, total financial assets equal total financial liabilities, i.e., the system is closed, as it includes the position of the rest of the world vis-à-vis Italy.

Table 2. Sectoral Financial Asset Exposures, 2017:Q3 (percent of GDP)

	NFC	MFI	BOI	OFI	INP	GOV	HH	RoW	Tot. As.	NFA
NFC		14	7	1	2	6	5	27	61	−112
MFI	54		5	13	1	37	37	23	170	0
BOI	0	16		0	0	22	0	13	53	6
OFI	11	18	2		0	8	4	26	69	27
INP	4	3	1	3		20	0	25	55	−2
GOV	11	5	1	4	1		4	4	30	−124
HH	54	64	19	16	50	14		30	247	196
RoW	39	50	12	6	3	47	1		157	8
Tot. Liab.	174	170	47	42	56	154	51	149	843	

Notes: Rows indicate the creditor sector and columns indicate the debtor sector. Rows sum to total assets and columns to total liabilities. NFA: net financial assets. The source data are from the Bank of Italy.

position, with liabilities mainly to MFIs and households. The mirror image is the extremely positive NFA of households, with assets predominantly in MFIs, NFCs, and insurance and pension funds (in this order). The government is very indebted, mostly due to the financial sector and the rest of the world. The Bank of Italy is an important creditor of the government, reflecting the Eurosystem's implementation of QE via its local branch. However, the public sector also holds a significant amount of assets in private sectors, especially NFCs, suggesting it has room to divest and cut its gross liabilities. MFIs are (not surprisingly) financially balanced, with assets in NFCs and the government, and liabilities to households and the rest of the world.

The portfolio composition is visualized in table 3, which expresses financial exposures as a share of total sector assets. This nets out the effect of a sector's balance sheet size. One key message is that MFIs and insurance and pension funds are relatively more exposed to government assets, which indirectly exposes their creditors and/or shareholders, such as households and NFCs.

Next, the evolution of sectoral exposures over five data snapshots is explored. These snapshots are (i) the beginning of the sample, 1995:Q1; (ii) the deployment of the euro, 2001:Q4; (iii) the onset of the GFC, 2008:Q1; (iv) the Outright Monetary Transactions program announcement, 2012:Q2; and (v) the end of the

Table 3. Sectoral Financial Asset Exposures, 2017:Q3 (percent of sector assets)

	NFC	MFI	BOI	OFI	INP	GOV	HH	RoW	Total
NFC		23	11	2	3	10	9	43	100
MFI	32		3	7	1	22	22	14	100
BOI	1	31		1	0	42	0	25	100
OFI	16	27	3		0	11	6	37	100
INP	7	5	1	5		36	0	46	100
GOV	36	17	3	12	2		14	15	100
HH	22	26	8	6	20	6		12	100
RoW	25	32	8	4	2	30	1		100
All	21	20	6	5	7	18	6	18	100

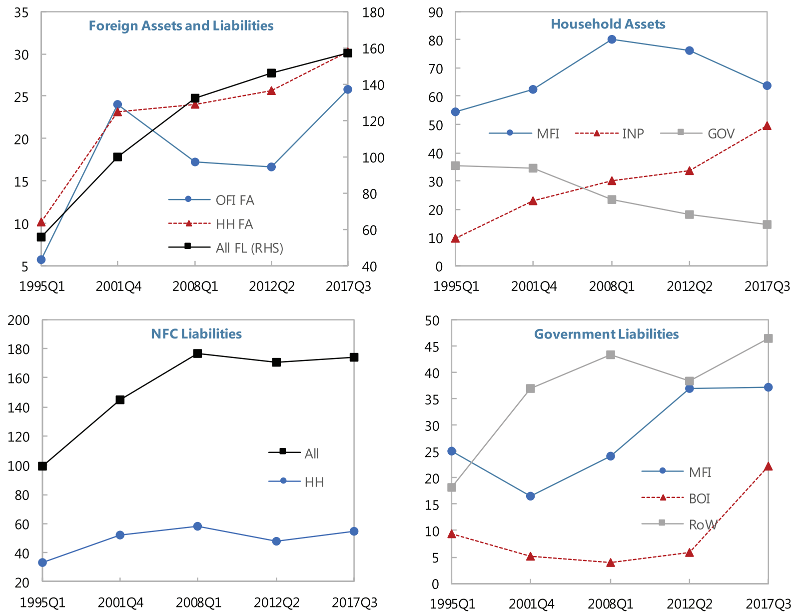
Notes: The last row shows the portfolio composition for the aggregate of all sectors. The source data are from the Bank of Italy.

sample, 2017:Q3. Appendix B contains the full matrix of bilateral exposures at each point in time, while figure 3 reports key takeaways.

The sum of financial assets in all sectors doubled from 1995:Q1 to 2017:Q3 (from 418 to 843 percent of GDP). This reflects the process of European financial integration as well as the increase in financial intermediation. The first panel of figure 3 shows the growing importance of foreign liabilities as well as the diversification of households and other financial institutions (OFIs) toward foreign asset holdings, peaking at the time of euro accession. Yet, the diversification process has slowed down, and the degree of home bias remains elevated.

The second panel shows that households shifted their asset portfolio away from the government and into insurance and pension funds. Their asset holdings in MFIs peaked in the run-up to the GFC but receded thereafter. The GFC also halted the accumulation of NFC liabilities, mostly due to a decline in household investment. On the other hand, the government became more reliant on MFI funding after the GFC, when some international creditors retreated. Since 2012, the Bank of Italy stepped in as a major creditor through the implementation of monetary policy.

Figure 3. Select Sectoral Financial Exposures (percent of GDP)



Notes: The charts are based on five (non-uniformly spaced) data snapshots. The source data are from the Bank of Italy.

4. Simulating the Impact of Financial Shocks

The sectoral exposures matrix is used to simulate the impact of financial shocks across sectors. This section describes how to calculate the impact of different shocks on sectoral net financial assets, whereas the next section discusses the results of each shock scenario.

As previously noted, such simulations capture the impact on wealth rather than on economic activity. Negative shocks to household and government wealth could depress private and public consumption, while shocks to corporate or bank balance sheets could hinder investment.[4] Instead, the focus of the paper is on gauging the

[4]Yet, financial shocks concentrated on wealthy households, such as the ones modeled in this paper, are not likely to induce a large consumption response, given their smaller marginal propensity to consume.

households' shock or loss-absorption capacity, given its importance for financial stability in Italy. Moreover, for simplicity, the simulations abstract from potential correlated price changes in other securities—beyond the change in equity values, portfolio reallocation after the shocks (e.g., prompted by regulatory constraints), or adjustment dynamics.

4.1 Valuation Shock

4.1.1 One Instrument in One Sector

A valuation shock to an instrument issued by a particular sector is considered first. A valuation shock has a *direct* impact on other sectors' NFA positions given by these sectors' holdings of the instrument suffering the shock. Formally, the direct impact on sector i's NFA of a shock to the value of its instrument-k assets in sector j is given by

$$d_{ijk} = \Delta p_{jk} a_{ijk}, \text{ if } i \neq j$$

and for sector $i = j$,

$$d_{jjk} = -\Delta p_{jk} l_{jk},$$

where Δp_{jk} is the change in value of instrument k issued by sector j, a_{ijk} the holdings of sector i in instrument-k liabilities of sector j, and l_{jk} the total instrument-k liabilities of sector j. The vector $\vec{d}_{jk}$ collects the impact for all sectors.

However, shocks also have an indirect impact given by intersectoral equity linkages. To calculate this indirect impact, sectors are divided into ultimate equity owners {BOI, GOV, HH, or RoW} and others. The four ultimate-owner sectors combined own all the equity issued by other sectors and have no equity liabilities themselves, i.e., they cannot be owned by any other entities, so they end up absorbing all shocks.[5] For the remaining sectors, the final NFA impact is zero, as any direct impact is compensated by an equal change in equity

[5] Technically, the rest of the world includes equity-issuing sectors such as foreign firms. However, data disaggregated by sector are not available for foreign residents. Hence, the small-economy assumption that changes in Italian asset prices have a negligible effect on the value of equity issued by foreigners is imposed.

liabilities to either HH, GOV, BOI, or RoW. Of course, this is not to say that financial shocks are without consequences for the non-ultimate-owner sectors, which may suffer from lower profitability, higher funding costs, default or restructuring events, market runs, and/or regulatory pressure.

Formally, the calculation involves two steps. First, the intermediate equity impact e' among the non-absorber sectors is

$$e'_{ijk} = (I - M)^{-1} \sum_{s=\{\text{NFC,MFI,OFI,INP}\}} d_{sjk} E'_{is},$$
$$\text{if } i = \{\text{NFC, MFI, OFI, INP}\},$$

where I is the identity matrix,

$$M \equiv E' - diag\,(E') - \lambda\,(E' - diag\,(E'))\,I,$$

λ is a vector of ones, and E' is a matrix showing the fraction of sector s's equity owned by sector i if $i \neq s$ and equal to the residual share otherwise, defined for $i = \{\text{NFC, MFI, OFI, INP}\}$.[6] The matrix M captures the difference between a sector's equity holdings in other sectors (off-diagonal entries) and its equity liabilities (diagonal entries). The geometric-sum term $(I - M)^{-1}$ reflects the infinite rounds of knock-on effects across sectors interlinked by mutual equity exposures.

Second, the ultimate equity impact e for the shock absorbers is calculated as

$$e_{ijk} = \sum_{s=\{\text{NFC,MFI,OFI,INP}\}} e'_{sjk} E_{is}, \text{ if } i = \{\text{BOI, GOV, HH, RoW}\},$$

where E is a matrix showing the fraction of sector s's equity owned by sector i, imposing zero ownership for $i = \{\text{NFC, MFI, OFI, INP}\}$.[7]

[6]Equity positions include both the ESA category "shares and other equity" and the fraction of non-money-market mutual fund positions which are invested in shares and other equity: 15 percent.

[7]Given the lack of perfect sector-by-sector disaggregation in the data, non-absorption by $i = \{\text{NFC, MFI, OFI, NPI}\}$ must be imposed as a constraint in the calculation.

For the non-absorber sectors, the equity impact is simply the opposite of the direct impact:

$$e_{ijk} = -d_{ijk}, \text{ if } i = \{\text{NFC, MFI, OFI, NPI}\},$$

which makes the total impact zero, as discussed above.

The total NFA impact across sectors is the sum of the direct and indirect impact of the shock:

$$\Delta NFA = \vec{d}_{jk} + \vec{e}_{jk}.$$

4.1.2 Multiple Instruments and Sectors

Generalizing the formulas to account for shocks to the prices of multiple instruments or sectors is straightforward. The direct impact for instrument k held by sector i becomes

$$d_{ijk} = \sum_{s \in \{S\} \setminus j} \Delta p_{sk} a_{isk} - \Delta p_{ik} l_{ik},$$

where S represents the subset of issuer sectors experiencing a price change in instrument k.

If price changes occur in multiple instruments, the direct impact can be simply added up. The formulas for the equity and total impacts are as for the case with one instrument in one sector above.

4.2 Bank Shock: Bail-in Scenario

The cases of a bank bail-in and bailout are slightly more complex than a simple valuation shock. These two scenarios assume that the value of other sectors' assets in a set of MFIs decline. In particular, the bail-in case assumes that all the equity of affected MFIs is wiped out and all their bonds are converted into equity, at a conversion rate of 50 percent.[8] For simplicity, it also assumes that no public

[8]This was approximately the conversion rate applied to Monte dei Paschi junior debtholders in its July 2017 bailout—the most recent one that occurred in Italy. Combined with the equity wipeout, the value of affected liabilities assumed here is above the minimum required eligible liabilities (MREL) mandated by the Bank Recovery and Resolution Directive (BRRD) to contribute public funds to a bail-in. Appendix D considers the alternative assumption of a zero conversion rate.

resolution funds are used and that other (nonbond) debtholders are not affected.

The direct impact is given by the sum of the valuation change in MFI equity and bond liabilities times the exposure of each sector to these two assets.

Since in a bail-in equity is wiped out, original shareholders do not benefit from the reduction in bond liabilities, so there is no indirect equity impact for MFIs:

$$e_{ijk} = 0, \text{ if } i = \text{MFI}.$$

For the rest of non-absorber sectors, the intermediate equity impact is

$$e'_{ijk} = (I - M'')^{-1} \sum_{s=\{\text{NFC,OFI,INP}\}} d_{sjk} E''_{is}, \text{ if } i = \{\text{NFC, OFI, INP}\},$$

where E'' and M'' are, respectively, defined like E' and M but only for sectors $i = \{\text{NFC, OFI, INP}\}$.

The ultimate equity impact for sectors other than MFIs is given by the same formulas as in the valuation shock case above.

4.3 Bank Shock: Bailout Scenario

The difference in the bailout scenario compared with the bail-in scenario is the government's intervention. In a bailout, the government makes a transfer to compensate MFI bondholders for any losses—which are assumed to be equal to those in a bail-in—so only shareholders are wiped out. Formally, the government makes a transfer t to affected MFI bondholders equal to the loss in value of their bonds:

$$t = \Delta p_{jk} \sum_i a_{ijk}, \text{ where } j = \text{MFI and } k = \text{bond}.$$

Hence, the direct impact for the government is equal to $-t$. For other sectors, the direct impact is given by their exposure to MFI equity.

Since among the troubled-bank bondholders are other MFIs, their corresponding share of the government transfer increases their equity value. Hence, the indirect equity effect for MFIs is in this case negative, equal to (minus) the government transfer to troubled-bank bondholder MFIs. Thus, the equity impact for MFIs can be calculated as the difference between the bail-in direct impact and the bailout direct impact (as the former does not include the government transfer, which is ultimately transmitted to MFI shareholders).

For the rest of non-absorber sectors,

$$e'_{ijk} = (I - M)^{-1} \sum_{s=\{\text{NFC,MFI,OFI,INP}\}} d'_{sjk} E'_{is},$$
$$\text{if } i = \{\text{NFC, OFI, INP}\},$$

where d'_{sjk} is equal to d_{sjk} if $s = \{\text{NFC, OFI, INP}\}$, and it is equal to minus the ultimate equity impact for MFIs if $s = \text{MFI}$.

The ultimate equity impact for sectors other than MFIs is given by the same formulas as in the valuation shock case.

5. Illustrative Shock Scenarios

5.1 Individual Shocks

This section presents results for a calibration of the three types of individual shocks dicussed above. A shock equivalent to a 10 percent decline in the value of government bonds (e.g., due to an increase in market perception of risk or nominal interest rates) is considered first.[9] Table 4 shows that the direct impact, given by government bond exposures, is concentrated on the financial sector—including MFIs, the Bank of Italy, and insurance and pension funds—as well as on the rest of the world. However, once equity linkages are considered, the sector with a larger total NFA decline is the rest of the world, followed by households. The impact on households as a fraction of GDP is non-negligible at about 4 percent of GDP, but only 2 percent as a fraction of their NFA. Government liabilities diminish by an equal amount, which raises the government's NFA position.

[9] This is equivalent to an increase in yields of around 220 basis points, given the duration of Italian outstanding government debt of 4.88 years (source: Bloomberg, as of May 2018), or a two-standard-deviation spread change, based on end-of-quarter year-on-year data for 1995:Q1–2018:Q3 (source: Reuters). For comparison, this is about 1.5 times the increase Italy experienced in late May 2018, and thus an economically sizable shock. Since the analysis is purely static, there is no need to specify the persistence of the shock.

Table 4. Impact of a Government Bond Value Shock (percent of GDP)

	NFA	Direct Impact	Equity Impact	ΔNFA
NFC	−112	−0.3	0.3	0.0
MFI	0	−2.1	2.1	0.0
BOI	6	−2.2	0.0	−2.2
OFI	27	−0.7	0.7	0.0
INP	−2	−2.0	2.0	0.0
GOV	−124	11.7	−0.4	11.3
HH	196	−0.8	−3.1	−3.8
RoW	8	−3.7	−1.6	−5.3

Notes: Assuming a 10 percent decline in the value of general government bonds. The direct impact on the NFA is given by the general government bond exposures. The equity impact is given by the bilateral equity linkages. ΔNFA is equal to direct impact plus the equity impact. The results are based on 2017:Q3 data from the Bank of Italy.

Next, a bank bail-in is compared with a bank bailout scenario. The two scenarios simulate a bank restructuring affecting a bank (or set of banks) constituting 10 percent of MFI liabilities, roughly equivalent to the combined size of Italy's third and fourth largest banks (by assets). This implies a 10 percent reduction in the value of MFI equity liabilities and a 5 percent reduction in the value of their bond liabilities (at a 50 percent conversion rate). In the case of a bailout, bond losses are fully compensated by the government, as explained in the previous section.

Table 5 shows that a bank bail-in mostly affects households and the rest of the world. This applies to both the direct and the total NFA impact. In fact, almost half of a bail-in's impact is absorbed by the rest of the world, which should mitigate the shock's damage to the domestic real economy. The impact on households is slightly over 1 percent of GDP, or 1/2 percent of the households' NFA. The NFA of MFIs increases as their liabilities to other sectors are reduced.

A bailout is less successful in sharing the impact with the rest of the world (table 6). The burden of a bailout falls mostly on the government, at over 1 1/2 percent of GDP, which worsens an already

Table 5. Impact of a Bank Bail-In (percent of GDP)

	NFA	Direct Impact	Equity Impact	ΔNFA
NFC	−112	−0.2	0.2	0.0
MFI	0	2.3	0.0	2.3
BOI	6	−0.1	0.0	−0.1
OFI	27	−0.1	0.1	0.0
INP	−2	−0.1	0.1	0.0
GOV	−124	−0.1	0.0	−0.1
HH	196	−0.9	−0.3	−1.2
RoW	8	−0.7	−0.1	−0.9

Notes: Assuming banks constituting 10 percent of MFI assets are bailed in. All their equity is wiped out and all their bonds converted to equity, at a 50 percent conversion rate. The results are based on 2017:Q3 data from the Bank of Italy.

Table 6. Impact of a Bank Bailout (percent of GDP)

	NFA	Direct Impact	Equity Impact	ΔNFA
NFC	−112	−0.2	0.2	0.0
MFI	0	2.9	−0.6	2.3
BOI	6	0.0	0.0	0.0
OFI	27	0.0	0.0	0.0
INP	−2	0.0	0.0	0.0
GOV	−124	−1.7	0.0	−1.6
HH	196	−0.6	0.2	−0.4
RoW	8	−0.3	0.1	−0.2

Notes: Assuming banks constituting 10 percent of MFI assets are bailed out. All their equity is wiped out, and the government compensates bondholders for 50 percent of their bond holdings value. The results are based on 2017:Q3 data from the Bank of Italy.

vulnerable financial position, and secondarily on domestic households. While foreign shareholders of domestic MFIs absorb a small fraction of the shock, the government ends up transferring resources to foreign bondholders. Overall, foreign absorption of the shock falls by 75 percent compared with a bail-in.

Table 7. NFA Impact of Bank Shocks, Higher MFI Capital Counterfactual (percent of GDP)

	Bail-In		Bailout	
	Baseline	Higher Capital	Baseline	Higher Capital
NFC	0.0	0.0	0.0	0.0
MFI	2.3	2.3	2.3	2.3
BOI	−0.1	0.0	0.0	0.0
OFI	0.0	0.0	0.0	0.0
INP	0.0	0.0	0.0	0.0
GOV	−0.1	−0.2	−1.6	−1.0
HH	−1.2	−1.3	−0.4	−0.8
RoW	−0.9	−0.8	−0.2	−0.5

Notes: The table shows the total NFA change after a bank bail-in and a bailout under two different assumptions: the "Baseline" column is based on the actual level of bank capital (like tables 5 and 6), while the "Higher Capital" column assumes a 50 percent higher MFI capital-to-total-assets ratio, proportionally distributed across shareholder sectors. The results are based on 2017:Q3 data from the Bank of Italy.

Italian banks' tier 1 capital ratios remain below the euro-area average, according to the ECB Supervisory Banking Statistics. Table 7 shows results for the bail-in and bailout simulations under the counterfactual assumption that MFIs start with a 50 percent higher aggregate capital-to-assets ratio. Such counterfactual would cut the public cost of a bailout by a third and double the foreign absorption rate. On the other hand, it would slightly increase the impact of a bail-in on domestic households, as these tend to hold a disproportionate amount of equity relative to bonds compared with foreigners. Moreover, while this exercise takes the financial shock as given, less leveraged banks may limit their risk-taking (a well-known mechanism since Jensen and Meckling 1976), potentially reducing the probability and size of shocks.

The above bank shock simulations are subject to a number of caveats. First, the conversion rate of bonds into equity in a bail-in usually depends on circumstance. Appendix C presents results assuming a zero conversion rate. Second, smaller Italian banks, which tend to feature more vulnerable balance sheets and are more likely to be restructured, are also disproportionately owned by domestic households.[10] Taking this into account would reduce the subsidy to the rest of the world associated with a bailout. Third, unlike a bail-in, a bailout may prevent contagion to other MFIs, potentially preventing detrimental knock-on effects on financial stability and investment. Finally, the bail-in and bailout scenarios must be interpreted as illustrative polar cases, since actual experiences of bank restructuring typically contain elements of both.

5.2 Correlated Shock

Shocks to the value of government debt in Italy tend to cause changes in the value of other debt securities. Albertazzi et al. (2014) estimate that a 100 basis point increase in spreads leads to a 73 basis point increase in MFI bonds rates after one quarter, a 21 basis point increase in corporate credit rates, 17 basis points in household credit, and 34 basis points in MFI term deposits. Using these estimates of the correlation between government bond rates and other securities, tables 8 and 9 recalculate the direct and total NFA impact, respectively, of a shock to the value of government bonds.[11] The tables introduce the correlations in a cumulative manner: the first column corresponds to the table 4 baseline, the second column adds a correlated response of bank bonds, the third column adds the response of corporate and household credit, and the fourth term deposits.[12]

Focusing on the direct exposures (table 8), a depreciation of MFI bonds increases the impact for households and the rest of the world, while it obviously creates a valuation gain for MFIs. Adding a depreciation of debt securities by all sectors harms the financial sector,

[10]See SNL bank-level data showing lower capitalization and profitability in several banks outside the largest two.

[11]To translate interest rate correlations into price correlations, MFI bonds and private credit are both assumed to have similar average duration as government bonds, while MFI deposits are assumed to have half the duration. Data for MFI bond average maturity are available from Bloomberg, while data on outstanding amounts of other debt securities by broad maturity bins are available from Bank of Italy statistics.

[12]The order of addition is motivated by the experience during the 2018 sovereign spread hike, when pass-through was strongest for bank bonds. In fact, the 2018 response of credit and especially term deposit rates has been muted so far, suggesting the last two columns in tables 8 and 9 should be taken as an upper bound.

Table 8. Direct Impact of a Combined Debt Shock (percent of GDP)

	Baseline	MFI Bonds	All Debt	Term Deposits
NFC	−0.3	−0.3	0.4	0.4
MFI	−2.1	−0.7	−1.4	−0.3
BOI	−2.2	−2.3	−2.2	−2.5
OFI	−0.7	−0.8	−1.1	−1.4
INP	−1.9	−2.0	−2.6	−2.6
GOV	11.7	11.7	11.6	11.6
HH	−0.8	−1.3	−0.6	−1.0
RoW	−3.7	−4.3	−4.1	−4.3

Notes: The first column assumes a 10 percent decline in the value of general government bonds, as in table 4. The next three columns cumulatively add correlated declines in the values of debt issued by other sectors. The second column adds a 7.3 percent decline in the value of bank bonds. The third column adds a 2.1 percent decline in the value of corporate debt and a 1.7 percent decline in the value of household debt. The fourth column adds a 1.7 percent decline in the value of term deposits. The results are based on 2017:Q3 data from the Bank of Italy.

Table 9. NFA Impact of a Combined Debt Shock (percent of GDP)

	Baseline	MFI Bonds	All Debt	Term Deposits
NFC	0.0	0.0	0.0	0.0
MFI	0.0	0.0	0.0	0.0
BOI	−2.2	−2.3	−2.2	−2.5
OFI	0.0	0.0	0.0	0.0
INP	0.0	0.0	0.0	0.0
GOV	11.4	11.4	11.3	11.3
HH	−3.9	−3.7	−3.6	−3.5
RoW	−5.2	−5.4	−5.5	−5.4

Notes: The first column assumes a 10 percent decline in the value of general government bonds, as in table 4. The next three columns cumulatively add correlated declines in the values of debt issued by other sectors. The second column adds a 7.3 percent decline in the value of bank bonds. The third column adds a 2.1 percent decline in the value of corporate debt and a 1.7 percent decline in the value of household debt. The fourth column adds a 1.7 percent decline in the value of term deposits. The results are based on 2017:Q3 data from the Bank of Italy.

while it relieves the losses for NFCs and households. If term deposits also depreciate to the extent observed in the past (i.e., deposit rates increase), the NFA impact on banks is quite mitigated, while deposit holders, including the Bank of Italy, other financial institutions, and households, are all worse off.

Table 9 shows that the total NFA impact under a combined shock is not substantially different from the baseline. Households' NFA tends to be slightly less affected by the shock as some of households' liabilities decrease in value, while the rest of the world and the Bank of Italy suffer a worse NFA decline as a larger chunk of their assets depreciate.

6. Distributional Impact

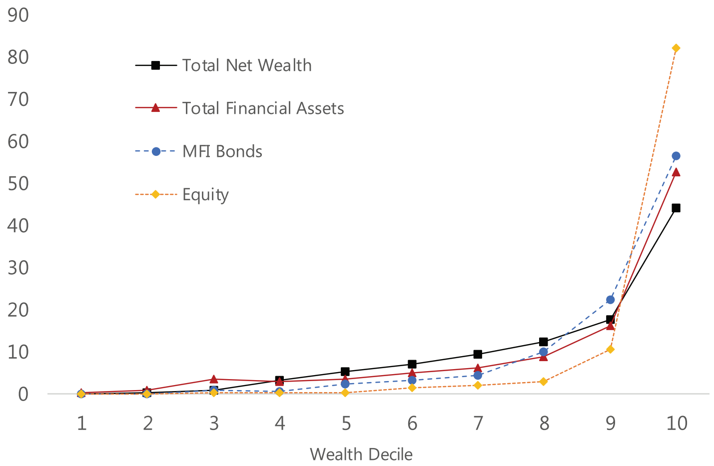
Survey data allow us further decompose the impact of shocks absorbed by each household wealth group. This section documents stylized facts of the financial wealth distribution for different financial instruments. Next, it uses this information to estimate how the burden of the financial shocks considered in the previous section is shared across household wealth deciles.

Financial wealth is concentrated at the top of the distribution (figure 4). Households at the top 20 percent of the wealth distribution hold 69 percent of financial wealth. This is particularly the case for risky assets, such as equity holdings (93 percent at the top 20), and to a lesser extent for bank bonds (79 percent at the top 20). Government bonds are distributed in line with total financial wealth.

Wealthier households hold riskier financial instruments (figure 5). The financial portfolio of less-wealthy households is almost entirely constituted by (insured) bank deposits, so they are not directly affected by government bond or bank financial shocks. Wealthier households, with a higher capacity to absorb losses, invest proportionally more in equity and nonsecured fixed-income instruments. Yet, most of their financial wealth is also in safe assets.

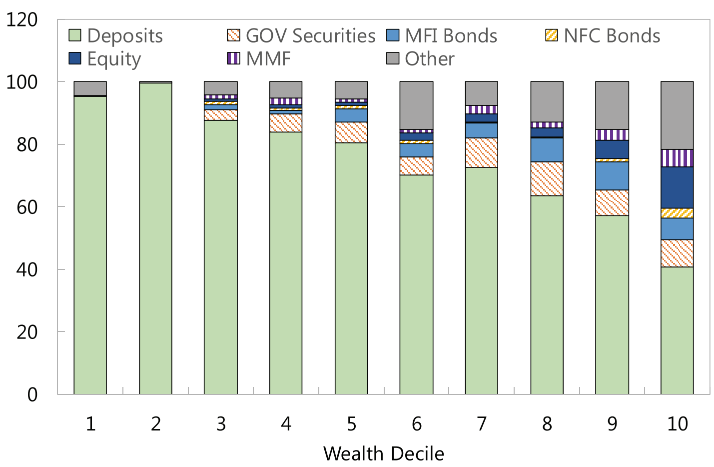
As a result, wealthy households absorb most of the financial losses after a government bond shock or a bail-in. Table 10 shows the impact of the three financial shocks by household decile, expressed both in euros per household and as a percentage of decile total

Figure 4. Distribution of Household Assets by Household Wealth, 2016:Q4 (percentage points)



Notes: Perfect equality would correspond to a flat line across deciles at 10 percent. "Equity" includes shares, equity funds, and exchange traded funds. The data are from the Bank of Italy's Survey on Household Income and Wealth.

Figure 5. Household Financial Asset Portfolio by Wealth, 2016:Q4 (percent of decile total financial assets)



Notes: "MMF" denotes money market funds. "Equity" includes shares and equity mutual funds. The category "Other" consists mostly of mutual funds which are neither equity funds nor MMF, as well as of pension funds. The data are from the Bank of Italy's Survey on Household Income and Wealth.

Table 10. Impact of Financial Shocks by Household Wealth Decile (percentage points)

	Euros per HH					% of Decile Fin. Assets		
Decile	**Wealth**	**Fin. As.**	**Gov. Bond**	**Bail-In**	**Bailout**	**Gov. Bond**	**Bail-In**	**Bailout**
1	−1,584	477	0	0	0	−0.1	0.0	0.0
2	3,326	2,438	0	0	0	0.0	0.0	0.0
3	17,651	9,488	−51	−23	−12	−0.5	−0.2	−0.1
4	65,334	8,486	−76	−18	−8	−0.9	−0.2	−0.1
5	108,434	9,928	−91	−61	−36	−0.9	−0.6	−0.4
6	146,859	13,849	−183	−92	−46	−1.3	−0.7	−0.3
7	194,062	17,044	−306	−126	−62	−1.8	−0.7	−0.4
8	254,080	24,771	−504	−277	−149	−2.0	−1.1	−0.6
9	367,042	45,014	−1,260	−650	−310	−2.8	−1.4	−0.7
10	914,511	146,949	−8,183	−2,113	−511	−5.6	−1.4	−0.3
Total	206,971	27,844	−1,066	−336	−113	−3.8	−1.2	−0.4

Notes: The first column shows household total net wealth (including real assets), while the second one shows gross financial assets.[a] The next three columns show the absolute impact absorbed by each decile in euros per household, while the last three columns show the impact as a percentage of the decile's total gross financial assets. The results assume that bank equity exposures are distributed as total equity exposures. The results are based on data from the Bank of Italy.

[a]The average level of wealth and (especially) financial assets according to the Survey on Household Income and Wealth (2016 release) is significantly lower than the 2013 aggregate reported by the Bank of Italy (see section 1). Beyond the year difference, this could be due to underreporting in the survey, especially at the right tail of the wealth distribution, where financial assets in particular are concentrated. In any case, survey data are only used to obtain the distribution. Financial wealth levels are from the flow-of-funds data.

financial assets, together with the average net real wealth and gross financial assets of each decile (first two columns). Wealthier households are more affected by financial shocks, both in absolute terms as and as a percentage of their financial assets, especially for a government bond shock (where the indirect equity impact is relatively more important). Hence, the social welfare effect of such shocks is smaller than the wealth effect.

On the contrary, the cost of a bailout is less concentrated in wealthier households. In fact, households at the top decile of the wealth distribution contribute less as a percentage of their financial assets than the next three deciles. This is because in a bailout the government compensates MFI bondholder sectors which are ultimately owned by wealthy households, such as other MFIs, OFIs, and NFCs, generating a positive indirect equity impact. Moreover, given that wealthier households have a lower marginal propensity to consume out of income shocks, the fiscal expansion associated with a bailout would probably lead to a meager GDP multiplier.[13]

Table 11 shows how the fiscal transfer implicit in a bailout is mostly distributed to the top of the wealth distribution (see "Data" column). It also computes the counterfactual transfer that would prevail absent the portfolio biases of rich households, i.e., if all wealth deciles held the same fraction of financial assets in equity, and if they held the same fraction in MFI bonds. Removing the equity bias would have a larger effect, making poorer wealth deciles receive a larger share of the transfer. Of course, such portfolio reallocation would imply that poorer deciles are more exposed to the shock in the first place.

The government could neutralize the cost of the bailout transfer implementing a tax on wealth that left each household wealth decile unaffected. Figure 6 shows the average wealth tax rate faced by each decile if such progressive wealth taxation was applied. The *proportional* tax rate on wealth needed to recoup the fiscal losses would be 0.2 percent, but a proportional tax would not completely undo the regressive effects of the bailout.

[13]See, e.g., Arrondel, Lamarche, and Savignac (2015) for an estimate of the marginal propensity to consume by wealth level in France.

Table 11. Bailout Fiscal Transfer by Household Wealth Decile (percentage points)

Decile	Euros per HH			% of Decile Fin. Assets		
	Data	No Equity Bias	No MFI Bond Bias	Data	No Equity Bias	No MFI Bond Bias
1	0	5	2	0.0	1.1	0.4
2	0	27	9	0.0	1.1	0.4
3	17	113	43	0.2	1.2	0.5
4	18	100	43	0.2	1.2	0.5
5	36	134	48	0.4	1.3	0.5
6	73	186	92	0.5	1.3	0.7
7	104	234	121	0.6	1.4	0.7
8	196	377	184	0.8	1.5	0.7
9	559	728	496	1.2	1.6	1.1
10	3,110	2,207	3,073	2.1	1.5	2.1
Total	411	411	411	1.5	1.5	1.5

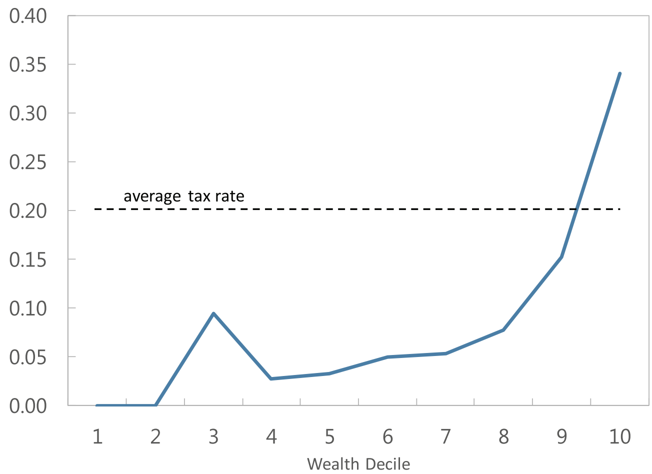
Notes: The first three columns show the absolute impact absorbed by each decile in euros per household, while the next three columns show the impact as a percentage of the decile's total gross financial assets. The results assume that bank equity exposures are distributed as total equity exposures. The results are based on data from the Bank of Italy.

7. Conclusion

Italy's sectoral asset exposures stand out for their strong household net financial positions but also for the weakness of other sectors. The household sector has been increasing its net financial position by building assets in the financial sector, which in turn has been lending to the heavily indebted public and nonfinancial corporate sectors. Reflecting quantitative easing, in recent years, the Bank of Italy has also become heavily exposed to Italian government assets.

An illustrative set of calculations reveals the diverse impact of financial shocks across sectors. The purpose is to assess the loss-absorption capacity of households rather than to fully specify the deleterious GDP and employment effects, which should also be considered in policymaking. A fall in the value of government bonds

Figure 6. Wealth Tax to Neutralize a Bailout's Fiscal Transfer (average tax rate by wealth decile, percent)



Notes: The solid line shows the average tax rate on total wealth by wealth decile needed to compensate the fiscal transfer to each wealth decile in a bailout. The dashed line shows the equivalent proportional tax rate.

affects the financial sector and is ultimately absorbed by domestic households and the rest of the world. The impact of a sizable loss in the value of government assets on net household wealth is relatively large but appears manageable, especially considering that it affects mostly the upper parts of the wealth distribution.

Regarding the analysis of bank bail-ins and bailouts, although not a comprehensive assessment of all their implications, the contribution of the paper is to highlight a series of quantitatively relevant factors that are often not transparent in the policy debate. In the case of a bail-in, the overall impact on wealth is relatively small, and the impact on welfare is even smaller, considering its incidence is almost entirely at the top end of the wealth distribution. The degree of domestic absorption in a bank bail-in is much lower than in a bailout, in which the government in effect transfers resources to foreigners and thus bears the brunt of the loss in value. These costs will need to be passed on to the domestic taxpayer, and thus across the broader parts of the income and wealth distribution, given the heavy reliance on labor income and consumption taxes as opposed to wealth taxes.

Appendix A. Identification of Sectoral Linkages in the Data

This appendix describes the necessary assumptions to infer intersectoral exposures from Italian flow-of-funds data.

For some sector-instruments, such as government bonds, identifying the counterparty sector is straightforward—the government. If the counterparty sector of an instrument is reported as "other financial institutions," the assets are distributed between OFI and INP according to the relative total liabilities of these two sectors in that particular instrument. A similar assumption applies to instruments issued by MFIs when the data does not specify whether these are issued by the Bank of Italy or other MFIs (e.g., deposits). For instrument categories "short-term loans," "medium- and long-term loans," and "insurance, pension, and guaranteed funds," the data do not provide a complete sectoral disaggregation on the asset side. Hence, the classification is based on liability-side information.

For the remaining sector-instruments, the principle of maximum entropy is applied, following previous literature (most closely Castrén and Kavonius 2009, inspired by Allen and Gale 2000). That is, asset positions of sector i on sector j are obtained multiplying the marginal distribution of sector i assets times the marginal distribution of sector j liabilities. Typically, these "unclassifiable" instruments are reported in the data as assets of sector i in "other sectors" or liabilities of sector i with respect to "other sectors."

This classification approach ensures that all instruments are allocated to both a creditor and a debtor sector. Hence, the total sum of assets and liabilities in each instrument is consistent with the balance sheet positions of the whole economy in that instrument.

Appendix B. Financial Exposures over Time

Table B.1 contains the full matrix of bilateral exposures at select points in time: (i) the beginning of the sample, 1995:Q1; (ii) the deployment of the euro, 2001:Q4; (iii) the onset of the GFC, 2008:Q1;

Table B.1. Sectoral Financial Asset Exposures over Time (percent of GDP)

1995:Q1	**NFC**	**MFI**	**BOI**	**OFI**	**INP**	**GOV**	**HH**	**RoW**	**Tot. As.**	**NFA**
NFC		7	2	1	2	5	5	10	33	−67
MFI	39		2	2	0	25	16	14	97	−1
BOI	0	2		0	0	9	0	5	17	3
OFI	7	3	0		0	5	1	6	21	0
INP	2	2	0	0		4	0	2	9	−5
GOV	9	5	1	6	1		1	2	24	−77
HH	33	54	8	11	10	35		10	162	139
RoW	11	24	1	1	1	18	0		56	8
Tot. Liab.	100	98	14	21	14	102	23	48	418	
2001:Q4	**NFC**	**MFI**	**BOI**	**OFI**	**INP**	**GOV**	**HH**	**RoW**	**Tot. As.**	**NFA**
NFC		14	1	2	3	4	6	18	48	−96
MFI	46		1	3	1	17	20	15	104	−19
BOI	0	1		0	0	5	0	6	12	4
OFI	11	7	0		0	12	4	24	58	10
INP	3	3	0	2		9	0	8	25	−7
GOV	8	5	0	5	1		3	3	26	−92
HH	52	63	5	31	23	34		23	231	198
RoW	24	31	0	4	2	37	0		100	2
Tot. Liab.	145	123	8	47	31	117	34	97	603	
2008:Q1	**NFC**	**MFI**	**BOI**	**OFI**	**INP**	**GOV**	**HH**	**RoW**	**Tot. As.**	**NFA**
NFC		15	2	1	3	6	6	19	50	−126
MFI	61		1	6	1	24	29	24	146	−13
BOI	0	1		0	0	4	0	9	15	3
OFI	12	6	0		0	4	9	17	48	18
INP	4	5	0	1		8	0	15	34	−3
GOV	10	6	1	2	1		4	2	24	−89
HH	58	80	6	13	30	24		24	235	187
RoW	32	47	1	7	2	43	0		133	22
Tot. Liab.	177	159	11	30	37	113	48	111	686	
2012:Q2	**NFC**	**MFI**	**BOI**	**OFI**	**INP**	**GOV**	**HH**	**RoW**	**Tot. As.**	**NFA**
NFC		11	4	1	1	7	5	26	54	−116
MFI	66		3	14	1	37	38	29	188	−2
BOI	0	18		0	0	6	0	14	37	7
OFI	12	21	1		0	11	5	17	67	29
INP	3	3	0	2		13	0	14	36	−2
GOV	10	5	1	2	0		4	4	27	−103
HH	47	76	14	10	34	18		26	225	171
RoW	31	57	8	9	2	38	1		146	17
Tot. Liab.	170	190	31	38	38	130	54	129	780	

(*continued*)

Table B.1. (Continued)

2017:Q3	NFC	MFI	BOI	OFI	INP	GOV	HH	RoW	Tot. As.	NFA
NFC		14	7	1	2	6	5	27	61	−112
MFI	54		5	13	1	37	37	23	170	0
BOI	0	16		0	0	22	0	13	53	6
OFI	11	18	2		0	8	4	26	69	27
INP	4	3	1	3		20	0	25	55	−2
GOV	11	5	1	4	1		4	4	30	−124
HH	54	64	19	16	50	14		30	247	196
RoW	39	50	12	6	3	47	1		157	8
Tot. Liab.	174	170	47	42	56	154	51	149	843	

Notes: The data are from the Bank of Italy.

(iv) the Outright Monetary Transactions program announcement, 2012:Q2; and (v) the end of the sample, 2017:Q3.

Appendix C. Bank Shock Counterfactual Simulations

This appendix shows the total NFA change after a bail-in and a bailout under alternative assumptions for the bond conversion rate (table C.1). The baseline shows the impact assuming that 50 percent

Table C.1. Impact of Bank Shocks, No Bond Conversion Counterfactual (percent of GDP)

	Bail-In		Bailout	
	Baseline	No Conversion	Baseline	No Conversion
NFC	0.0	0.0	0.0	0.0
MFI	2.3	3.2	2.3	3.2
BOI	−0.1	−0.2	0.0	0.0
OFI	0.0	0.0	0.0	0.0
INP	0.0	0.0	0.0	0.0
GOV	−0.1	−0.1	−1.6	−3.2
HH	−1.2	−1.6	−0.4	0.0
RoW	−0.9	−1.3	−0.2	0.0

Notes: The results are based on 2017:Q3 data from the Bank of Italy.

of the value of bonds is recovered after bonds are converted to equity, while the alternative shows the impact under a zero conversion rate.

A lower conversion rate increases total bondholder losses and thus the required government transfer in a bailout. In the bail-in case, the additional impact is roughly proportionally distributed across sectors.

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Rethinking Capital Regulation: The Case for a Dividend Prudential Target*

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Recent empirical studies have documented two remarkable patterns shown by euro-area banks in the aftermath of the Great Recession: (i) their tendency to boost capital ratios by shrinking assets (contraction in loan supply), and (ii) their reluctance to cut back on dividends (fall in retained earnings). First, I provide evidence of a potential link between these two trends. When shocks hit their profits, banks tend to adjust retained earnings to smooth dividends. This generates bank equity and credit supply volatility. Then I develop a DSGE model that incorporates this mechanism to study the transmission and effects of a novel macroprudential policy rule—that I shall call dividend prudential target (DPT)—aimed at complementing existing capital regulation by tackling this issue. Welfare-maximizing DPTs are effective (more than the CCyB) in smoothing the financial and the business cycle (by means of less volatile retained earnings) and induce significant welfare gains associated with a Basel III type of capital regulation through various channels.

JEL Codes: E44, E61, G21, G28, G35.

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1. Introduction

As a result of the pressure exerted by the private and the public sector, banks in the euro area (and elsewhere) had to increase their risk-weighted capital ratios in the aftermath of the global financial crisis. However, contrary to what happened in the rest of the world, European banks primarily improved such ratios by shrinking assets, thereby exacerbating the problem of credit supply procyclicality.

Cohen and Scatigna (2016) show that for the period 2009–13 the euro-area banking sector boosted regulatory capital ratios mainly via asset shrinking, while virtually the rest of the world did so by accumulating retained earnings. Gropp et al. (2019) conclude that European banks which had to raise their core tier 1 capital ratios in response to the European Banking Authority (EBA) 2011 capital exercise did it by shrinking assets—a reduction in total assets that has been mainly attributed to a contraction in outstanding customer loans.

As suggested in the 84th Annual Report of the Bank for International Settlements (2014) and Shin (2016), banks in the euro area failed to boost capital ratios by increasing retained earnings due to their relatively strong reluctance to cut back on dividends. According to the evidence, large and established corporations (including banks) distribute a significant percentage of their profits in the form of dividends and tend to smooth them over the cycle (see, e.g., Lintner 1956, Allen and Michaely 2003, and DeAngelo, DeAngelo, and Skinner 2009). There is, however, little agreement on why managers have such a preference for smoothing dividends and what determines their propensity to smooth (see, e.g., Leary and Michaely 2011).

The joint consideration of all available evidence on these matters points to a potential link between these two trends. Bankers' preference for smoothing dividends implies that the bulk of the adjustment to exogenous shocks that hit bank profits is mainly borne by retained earnings, thereby generating bank equity and credit supply volatility (through a balance sheet effect). Current capital legislation allows for this unintended macroeconomic effect, since it says little about the channels through which banks should adjust their capital ratios and gives such institutions "full discretion" to set their own payout policies provided that they comply with the corresponding capital requirements.[1]

The main contribution of this paper is to define a very simple framework that incorporates this mechanism to study the transmission channel and effects of a novel macroprudential rule—that I shall call dividend prudential target (henceforth DPT)—aimed at complementing existing capital regulation by tackling this issue.

In order to do so, I develop a quantitative dynamic stochastic general equilibrium (DSGE) model with a banking sector. Households (net savers), entrepreneurs (net borrowers), and bankers interact in a real, closed, decentralized, and time-discrete economy in which all markets are competitive. As in Iacoviello (2015), (i) borrowers and bankers are constrained in their capacity to borrow due to the existence of collateral constraints and regulatory capital ratios, respectively, and (ii) the relationship between the discount factors of the three types of agents is such that (i) there are financial flows in equilibrium, and (ii) the borrowing constraints are binding in a neighborhood of the steady state. These implications are crucial to focus the analysis on a possibility neither considered in the macrofinance literature (to the best of my knowledge) nor incorporated in the Basel III Accords: to regulate bank dividend policies even when credit institutions comply with their capital requirements.

As in Gerali et al. (2010), bank equity accumulates out of retained earnings with a functional form identical to the standard law of motion for physical capital. Such assumption allows for the model to account for (i) the crucial link between profit and capital generation capacity within the banking sector, and (ii) the nontrivial intertemporal decision bankers have to make when it comes to earnings distribution. The preference of the representative banker for paying large amounts of dividends and for smoothing such payouts over time (accounted for by a relatively low subjective discount factor and a constant elasticity of substitution (CES) utility function, respectively) conflicts with the obligation to retain earnings

[1]For payout policy purposes, the relevant capital adequacy ratio that should be met by credit institutions comprises, for the general case, the minimum capital requirement (8 percent); the capital conservation buffer, or CCoB (2.5 percent); and the countercyclical capital buffer, or CCyB ($\geqslant$0 percent) as an add-on to the CCoB.

and meet capital requirements as well as with the will to expand the bank's profit generation capacity (and, thus, its earnings distribution capacity) over time. In addition, credit institutions face a balance sheet constraint, by which bank assets (one-period loans extended to borrowers) must be fully financed by equity and debt (one-period deposits borrowed from savers) in each period. This allows for a simple mechanism through which (i) adjustments in retained earnings affect credit supply, and (ii) exogenous shocks that hit the real economy through the financial sector get amplified. In a nutshell, the model incorporates a mechanism through which bankers' preference for dividend smoothing in a context of borrowing limits (including capital requirements) induces suboptimally high aggregate equity and credit supply volatility.

Against this background, I design a macroprudential policy rule aimed at giving incentives for bankers to tolerate a higher degree of dividend volatility in order to sustain retained earnings and loans supply in economic downturns. The DPT is a regulatory target for bank dividend payouts that reacts to steady-state deviations of a macroeconomic indicator of the choice of the regulator (i.e., it is dynamic) and enters a quadratic penalty function whose specification is analogous to the dividend adjustment cost assumed in Jermann and Quadrini (2012) and Begenau (2020). Such specification of the sanctions regime allows to strike a balance between enforcement (it penalizes bankers who deviate from the DPT) and flexibility of the policy rule (it allows for bankers to deviate from the target conditional on the payment of a sanction).

In the baseline scenario, the only existing prudential policy instrument is static capital requirements. In alternative policy scenarios, dynamic capital requirements and dividend prudential targets are introduced to study the interactions, transmission mechanisms, and key macroeconomic effects of these policy instruments. Dynamic capital requirements are specified as the complementary of a bank debt-to-assets ratio that responds to steady-state deviations of a macroeconomic indicator of the choice of the regulator.

First, I identify the mechanism through which the DPT operates and give a first quantitative assessment of its potential to smooth the credit cycle. Then, I extend the model to carry out a welfare analysis of the regulatory scheme under consideration. I incorporate another type of borrower (impatient households), physical capital, various macroeconomic and financial exogenous shocks, and additional ingredients (e.g., GHH preferences and investment adjustment costs) that allow me to calibrate the model to quarterly data of the euro area for the period 2002:Q1–2018:Q2, and match a number of first and second moments from banking and macroeconomic aggregates (including bank assets, profits, dividends, and the payout ratio, among others).

As in Clerc et al. (2015), I assume that households own all existing firms in the economy (including banks), an assumption which has two important implications. First, there is a separation between bank ownership and bank management that allows to capture the two main channels through which dividend smoothing operates according to the evidence; bank owners' risk aversion, and managers' propensity to smooth dividends (see, e.g., Wu 2018). Second, the welfare analysis can be restricted to households without neglecting any consumption capacity generated in the economy. Optimized policy rules are obtained by maximizing a measure of social welfare—defined as a weighted average of the expected lifetime utility of the two types of households—with respect to the relevant policy parameter vector and for different welfare weighting criteria.

Optimized DPTs are countercyclical (i.e., they call for procyclical and relatively more volatile dividends in order to smooth aggregate lending and output through more stable retained earnings) and trade off the key conflictive welfare effects induced by this macroprudential instrument. On the one hand, a more responsive countercyclical DPT favors credit smoothing, which is beneficial for borrowers. On the other hand, it induces bank dividend volatility, which has a negative impact on bank owners' welfare.[2] Such welfare tradeoff primarily originates from households' risk aversion, by which such agents implicitly prefer their resources to evolve in a smooth fashion (including credit and distributed earnings). The shape of such tradeoff (and, thus, the responsiveness degree of optimized DPTs) crucially depends on bank managers' CES preferences, which account for the stylized fact of managers' propensity to smooth dividends

[2] Although less determinant, if the degree of responsiveness of the DPT is sufficiently high, a third welfare effect—by which the macroprudential rule tends to moderately restrict credit provision—comes into play.

(while remaining agnostic about the underlying drivers of such preference) and permit to accurately match the second moment of bank dividends.

Welfare-maximizing dividend prudential targets are shown to have important properties. First, they are more effective in smoothing financial and business cycles than the countercyclical capital buffer (henceforth CCyB) due to a key difference between their corresponding transmission mechanisms.[3] Second, they complement existing capital regulation and induce welfare gains associated with a Basel III type of framework through various channels: (i) they are particularly effective in mitigating the negative effects of hikes in static (or microprudential) capital requirements in terms of more restricted and volatile credit supply;[4] (ii) they reinforce the effectiveness of the CCyB in mitigating financial and economic fluctuations regardless of the nature of the shock and perform particularly better than the CCyB under nonfinancial shocks; and (iii) they allow to strike a balance between the strong preference that households who do not own banks have for the DPT and the relevance the CCyB has for bank owners.[5] Third, they mainly operate through their cyclical component, ensuring that long-run dividend payouts remain unaffected. Fourth, they are associated with a sanctions regime that acts as an insurance scheme for the real economy.[6]

[3]Both macroprudential tools smooth loan supply, but they operate through very different transmission mechanisms: Under the CCyB, bank capital readjusts in the face of shocks, permitting debt (which in the face of a negative shock represents a larger proportion of assets) to evolve in a smoother fashion. By way of contrast, the DPT directly attacks the root of the "problem" (i.e., bank dividend smoothing) by giving incentives for bank managers to optimally tolerate a higher degree of dividend volatility, thereby allowing for smoother retained earnings.

[4]Higher capital requirements translate into a higher fraction of bank loans being financed by bank capital accumulated out of "volatile" retained earnings and higher long-run profits (and dividends). The former quantitatively magnifies the problem of a higher lending supply volatility induced by adjustments in retained earnings, whereas the latter reinforces the effectiveness of DPTs in tackling the issue.

[5]While households who do not own banks have a strict preference for a countercyclical DPT (since it is more effective than the CCyB in smoothing loan supply and other aggregates of the real economy), bank owners prefer the CCyB (as it favors credit smoothing without inducing higher bank dividend volatility).

[6]During the economic downturn, deviations from the DPT are penalized with a sanction. The corresponding public revenues collected by the public authority

The paper is organized as follows. Section 2 discusses how the paper fits into the existing literature. Section 3 presents empirical evidence on bank dividends and earnings in the euro area. Section 4 describes the basic model and identifies the transmission mechanism through which a dividend prudential target smooths the credit cycle. Section 5 presents the extended model to improve the matching of the model to the data. Section 6 develops a quantitative exercise to assess the welfare effects of the DPT and its interactions with regulatory capital ratios. Section 7 concludes.

2. Related Literature

This paper relates to recent work that attempts to motivate the desirability of regulating earnings distributions under certain conditions. Based on U.S. banking data for the period 2007–09, Acharya et al. (2012) suggest the imposition of regulatory sanctions against large-scale payments of dividends that erode common equity. Similarly, Admati et al. (2013) advocate dividend restrictions and capital conservation in bad times. Goodhart et al. (2010) and Acharya, Le, and Shin (2017) provide theoretical rationale for the use of dividend restrictions for banks under various conditions, suggesting that this regulatory measure would be beneficial not only to debt holders but also to equity holders. In these two-period models, the justification for imposing dividend restrictions relates to a private equilibrium that features excessive dividends and inefficiently low bank capitalization.

This paper contributes to this strand of literature by adopting a DSGE modeling approach to assess the effectiveness of a very specific macroprudential policy rule aimed at breaking the nexus between bankers' preference for dividend smoothing and credit supply volatility. The proposed regulatory scheme plays a key role as a macroprudential tool in an environment in which banks are assumed to constantly meet their capital requirements (and there is no risk

are transferred—within the same period—to households (and/or entrepreneurs). Such transfer system acts as an insurance scheme to the real economy, as it provides economic agents of the nonfinancial private sector with a positive payoff when they need it the most (i.e., when the marginal utility of their consumption is relatively high).

of bank failure). The key mechanism through which the regulatory scheme plays such a role is by providing incentives to bank managers to tolerate a higher degree of dividend volatility, thereby smoothing retained earnings and credit supply. The proposed prudential tool is analyzed in a quantitative macro model that matches a number of first and second moments of macro and banking data of the euro area. That allows the paper to study the type of welfare tradeoffs and effects that would be induced by this prudential tool in the euro-area economy, as well as its interactions with a Basel III type of capital regulation. The design and key features of the proposed policy instrument notably differ from those of dividend restrictions presented in the literature.

The paper also connects to the banking literature that quantifies the effects of capital regulation (see, e.g., Van den Heuvel 2008, Angelini, Neri, and Panetta 2014, De Nicolò, Gamba, and Lucchetta 2014, Martinez-Miera and Suarez 2014, Mendicino et al. 2018, and Corbae and D'Erasmo 2019). A common feature of these models is that higher capital requirements lead to more restricted and volatile lending. Although the proposed model accounts for this effect, the channel through which it emerges is quite novel; a hike in capital requirements translates into a higher proportion of bank assets being financed by "volatile" equity, which induces larger fluctuations in credit supply (i.e., "balance sheet effect"). This effect is to be traded off against two other effects: (i) a "loan portfolio readjustment effect" that has an asymmetric impact on impatient households and entrepreneurs and only emerges when the corresponding hike in capital ratios is associated with a change in relative sectoral capital requirements, and (ii) a "profit-generation capacity effect" through which increased capital requirements (and cumulative retained earnings) translate into higher long-run dividend payouts. Furthermore, the DPT incorporates an important welfare tradeoff that interacts with that of capital requirements. A countercyclical DPT smooths lending while it induces higher bank dividend volatility. In order to clearly identify these tradeoffs and keep the complexity of the analysis to a minimum, the model abstracts from other effects of changing capital regulation parameters such as reducing the risk of bank failure (see, e.g., Angeloni and Faia 2013 and Clerc et al. 2015) or the risk-taking by banks (see, for instance, Admati et al. 2012 and Begenau 2020).

Finally, the proposed model builds on recent work that attempts to incorporate banking in otherwise standard DSGE models—among others, Gerali et al. (2010), Gertler and Kiyotaki (2010), Meh and Moran (2010), Gertler and Karadi (2011), Andrés and Arce (2012), Brunnermeier and Sannikov (2014), Christiano, Motto, and Rostagno (2014), and Iacoviello (2015). As in most of these papers, the main role of the banking sector in this model is to allow for resource transfers between savers and borrowers. In the tradition of Kiyotaki and Moore (1997) and Bernanke, Gertler, and Gilchrist (1999), the presence of certain frictions enables financial intermediation activities to endogenously propagate and amplify shocks to the macroeconomy.[7] However, most of this work makes assumptions that imply that bank payout policies are exogenous and/or that the payout ratio is very low and constant over time—aspects which are sharply at odds with reality and which do not permit to carry out the analysis proposed in this paper.

3. Patterns of Bank Dividends and Earnings in the Euro Area

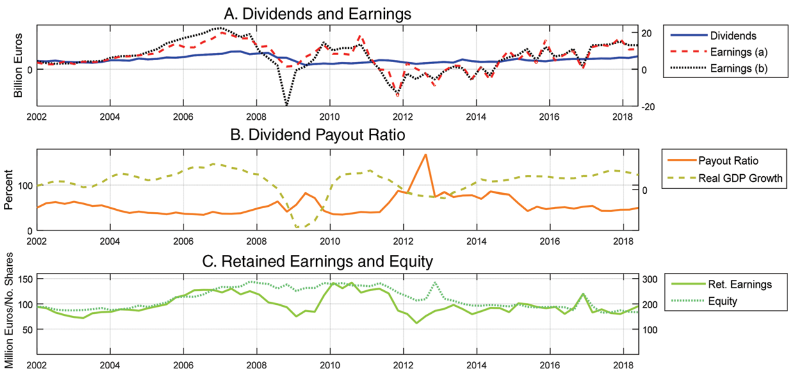
This section presents the main empirical observations that motivate the paper. Financial data plotted in figures 1 and 2 are from the Euro Stoxx Banks Index, SX7E.[8] All time series are at quarterly frequency and have been seasonally adjusted.[9] Figure 1A plots aggregate dividends in cash and earnings (net income) of the SX7E members for the period 2002:Q1–2018:Q2. While both variables are procyclical, earnings are substantially more volatile than dividends. Bank

[7]In the proposed setup, borrowing limits emerge as the key distortion that separates this equilibrium economy from its first best and allows for the proposed regulatory scheme to potentially be welfare improving.

[8]The Euro Stoxx Banks Index, SX7E, is a capitalization-weighted index in which the largest stocks in the EMU banking sector weigh in the index according to their free-float market capitalization. As of October 31, 2018, the top ten components of the index (and their corresponding weights) were Banco Santander (16.42 percent), BNP Paribas (12.90 percent), ING Group (9.89 percent), BBVA (7.90 percent), Intesa Sanpaolo (7.73 percent), Societe Generale Group (6.36 percent), Unicredit (5.81 percent), Deutsche Bank (4.01 percent), KBC Group (3.87 percent), and Credit Agricole (3.41 percent).

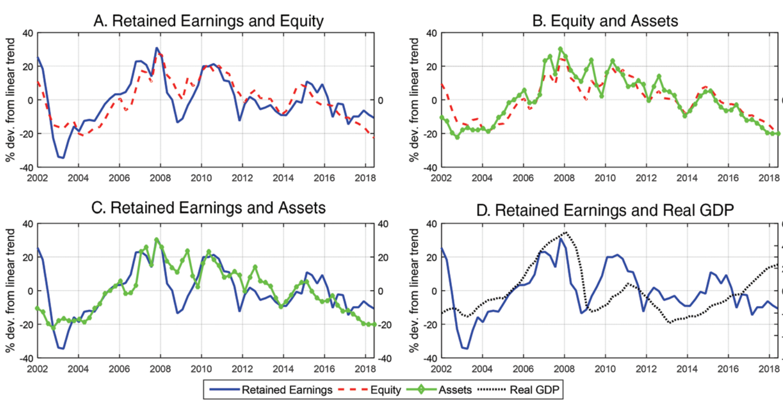
[9]See online appendix A for details on data construction. (The online appendixes can be found at http://www.ijcb.org.)

Figure 1. Bank Dividends and Earnings in the Euro Area (SX7E), 2002:Q1–2008:Q2



Data sources: Bloomberg, Eurostat, and own calculations.
Notes: SX7E refers to the Euro Stoxx Banks Index. Time series plotted in panel A have been constructed as a simple sum of the SX7E members, whereas those in panels B and C have been reported as the index itself (i.e., as a capitalization-weighted sum of the same group of banks). See the online appendixes for further details on data construction. In figure 1B the main y-axis and the secondary one differ, with the dashed line associated with the latter. In panel C the dotted line is associated with the secondary y-axis.

Figure 2. Co-movements among Bank Retained Earnings, Equity, Assets, and Real GDP



Data sources: Bloomberg, Eurostat, and own calculations.
Notes: This figure reports the cyclical component of euro-area real GDP as well as of aggregate (cumulative) retained earnings, equity, and assets of the SX7E members. In order to compute their cyclical component, the log value of seasonally adjusted and deflated time series has been linearly detrended. In panel D the main y-axis and the secondary one differ, with the dotted line associated with the latter.

managers in the euro area have a strong preference for smoothing dividends over the cycle and pay high and stable amounts of dividends in cash even in those quarters in which net income is negative. That is, the adjustment in the face of shocks that hit bank profits is mainly borne by undistributed net income. That has two important consequences. First, the dividend payout ratio of the euro-area banking sector is notably countercyclical (figure 1B), implying that bankers distribute a higher proportion of total earnings precisely when their capital positions are prone to be weaker (i.e., during the economic slowdown).[10] Second, this fact significantly affects equity dynamics, as retained earnings account for the bulk of total equity and the two variables are highly correlated (figure 1C).[11]

Figure 2 reports the cyclical component of selected aggregates to identify co-movements (i.e., patterns of positive correlation) among bank (cumulative) retained earnings, equity, assets, and real gross domestic product (GDP).[12] Figure 2A confirms that retained earnings and total equity are highly correlated, suggesting that the former is an important driver of bank equity volatility. Due to the importance of bank capital as a funding source and to the extent that the balance sheet identity always has to hold, it does not come as a surprise that the correlation between bank equity and bank assets (as a proxy for aggregate bank loan supply) is also very high and positive (figure 2B). The bottom line is that there is a high

[10] Quarterly aggregate data on payout ratios should be taken with caution for at least two reasons: (i) For each quarter, the index can only incorporate information on members whose net profits for the period are strictly positive. Otherwise the payout ratio cannot be computed. (ii) The adjustments made to raw data on net income (denominator of the payout ratio) often vary across analysts. These adjustments can be quantitatively important, especially when considering a time series that accounts for a period including a severe financial crisis and deep regulatory changes (in loan loss provisioning rules, etc.).

[11] Time series plotted in figure 1A have been constructed as a simple sum of the SX7E members, whereas those in figures 1B and 1C have been reported as the index itself (i.e., as a capitalization-weighted sum of the same group of banks).

[12] Financial data plotted in figure 2 have been constructed as a simple sum of the SX7E members. In order to compute their cyclical component, the log value of seasonally adjusted and deflated time series has been linearly detrended. These are some of the constructed time series that have been used to calibrate the extended model by matching second moments of euro-area quarterly data in section 6.

degree of co-movement among bank retained earnings, lending, and real GDP (see figures 2C and 2D).

The reluctance of bankers to cut back on dividends in the face of negative shocks that hit their profits leads to falls in retained earnings and total equity. In order to meet their capital requirements in a context of falling equity and economic slowdown (a period in which issuing new equity is often a costly or even impossible task for banks), bank managers have incentives to shrink assets by cutting back on lending. At the aggregate level, this (individual) strategy is prone to exacerbate the credit and the business cycle.

4. The Basic Model

Consider three types of agents who interact in a real, closed, decentralized, and time-discrete economy in which all markets are competitive. Households work, consume, accumulate housing, and invest their savings in one-period bank deposits. Entrepreneurs demand real estate capital and labor to produce a homogeneous final good. Due to a discrepancy in their discount factors, in the aggregate households are net savers whereas entrepreneurs are net borrowers. There are financial flows in equilibrium. Bankers intermediate financial resources by borrowing from households and lending to entrepreneurs. They devote the resulting net profit to do both: pay dividends (bankers' consumption) and meet the capital requirement by retaining earnings. For each type of agent, there is a continuum of individuals in the $[0, 1]$ interval.

In the spirit of Iacoviello (2005, 2015), entrepreneurs and bankers are assumed to face borrowing constraints that are binding in a neighborhood of the steady state. Consequently, the first best is unattainable in equilibrium. Such financial frictions play two important roles: (i) they amplify the effects of exogenous shocks through the financial sector, and (ii) they open up the possibility of a welfare-improving public intervention.

The aim of this section is to identify the transmission mechanism through which the considered policy operates. In doing so, the paper evaluates its effectiveness in favoring financial stability by smoothing the credit cycle.

4.1 Main Features

4.1.1 Households (Net Savers)

Let $C_{h,t}$, $H_{h,t}$, and $N_{h,t}$ represent consumption, housing demand, and hours worked by households in period t. The representative household seeks to maximize the objective function

$$E_0 \sum_{t=0}^{\infty} \beta_h^t \left[\log C_{h,t} + j \log H_{h,t} - \frac{N_{h,t}^{1+\phi}}{(1+\phi)} \right], \tag{1}$$

subject to the sequence of budget constraints

$$C_{h,t} + D_t + q_t(H_{h,t} - H_{h,t-1}) = R_{h,t-1} D_{t-1} + W_{h,t} N_{h,t},$$

where D_t denotes the stock of deposits, $R_{h,t}$ is the gross interest rate on deposits, q_t is the price of housing, and $W_{h,t}$ is the wage rate. $\beta_h \in (0,1)$ is the households' subjective discount factor, j is the preference parameter for housing services, and ϕ stands for the inverse of the Frisch elasticity. In each period, the representative household allocates its rents in terms of wage earnings and returns on total deposits between final consumption and investment in deposits and housing.

4.1.2 Entrepreneurs (Net Borrowers)

The representative entrepreneur chooses the trajectories of consumption $C_{e,t}$, housing $H_{e,t}$, demand for labor N_t, and bank loans $B_{e,t}$ that maximize

$$E_0 \sum_{t=0}^{\infty} \beta_e^t \log C_{e,t}, \tag{2}$$

subject to the sequence of budget constraints

$$\begin{aligned} &C_{e,t} + R_{e,t} B_{e,t-1} + q_t(H_{e,t} - H_{e,t-1}) + W_{h,t} N_t + \Phi_e(B_{e,t}) \\ &\quad = Y_t + B_{e,t}, \end{aligned}$$

with $\beta_e < \beta_h$. $B_{e,t}$ stands for bank loans, $R_{e,t}$ is the gross interest rate on loans, and $\Phi_e(B_{e,t}) = \dfrac{\phi_e}{2} \dfrac{(B_{e,t} - B_{e,t-1})^2}{B_e^{ss}}$ is a quadratic loan portfolio adjustment cost, assumed to be external to the entrepreneur as in Iacoviello (2015).[13] Y_t is final output. B_e^{ss} is the steady-state value of $B_{e,t}$ and ϕ_e is the loans adjustment cost parameter. In each period, the representative entrepreneur devotes resources in terms of produced final output and loans to consume, repay debt, remunerate productive factors, and adjust credit demand.

The homogeneous final good is produced by using a Cobb-Douglas technology that combines labor and commercial real estate as follows:[14]

$$Y_t = H_{e,t-1}^{\upsilon} N_t^{1-\upsilon}. \tag{3}$$

In addition, entrepreneurs are subject to

$$B_{e,t} \leq m_t^H E_t \left(\frac{q_{t+1}}{R_{e,t+1}} H_{e,t} \right) - m^N W_{h,t} N_t. \tag{4}$$

Expression (4) dictates that the borrowing capacity of entrepreneurs is tied to the value of their collateral. In particular, they cannot borrow more than a possibly time-varying fraction m_t^H of the expected value of their real estate stock. More precisely, $m_t^H = m^H \varepsilon_t^{mh}$ is the exogenously time-varying loan-to-value ratio, where $m^H \in [0,1]$ and ε_t^{mh} follows a zero-mean AR(1) process with autoregressive coefficient equal to ρ_{mh} and iid innovations $e_{mh,t}$ that are normally distributed and have a standard deviation equal to σ_{mh}. Moreover, the borrowing constraint indicates that a fraction $m^N \in [0,1]$ of the wage bill must be paid in advance, as in Neumeyer and Perri (2005).[15]

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[13] This cost discourages the entrepreneur from changing their credit balances too quickly, thereby contributing to match the empirical fact that bank credit varies slowly over time.

[14] The specification of a production function in which real estate enters as an input has become common practice in the macro-finance literature. See, e.g., Iacoviello (2005, 2015), Andrés and Arce (2012), and Andrés, Arce, and Thomas (2013).

[15] Without loss of generality, this assumption is made for quantitative-analysis-related reasons. It helps in shaping the steady-state levels and transition dynamics of aggregate financial variables, particularly in a reduced-form model of this kind.

4.1.3 Bankers

Let $d_{b,t}$ represent bank dividends (which are fully devoted to final consumption by bankers) in period t, and $\beta_b < \beta_h$. The representative banker seeks to maximize

$$E_0 \sum_{t=0}^{\infty} \beta_b^t \log d_{b,t}, \tag{5}$$

subject to

$$B_t = K_{b,t} + D_{b,t}, \tag{6}$$

$$d_{b,t} + K_{b,t} - (1-\delta)K_{b,t-1} = r_{e,t}B_{t-1} - r_{h,t-1}D_{b,t-1} - \Phi_b(B_t), \tag{7}$$

$$D_{b,t} \leq \gamma B_t, \tag{8}$$

where equations (6), (7), and (8) denote the balance sheet identity, the sequence of cash flow restrictions, and the borrowing constraint of the banker, respectively.

According to (6), bank assets are financed by the sum of bank equity $K_{b,t}$ (also referred to as bank capital) and debt. There is only one type of bank assets: one-period loans which are extended to entrepreneurs. Bank debt, $D_{b,t}$, is entirely composed of funds borrowed by households in the form of homogeneous one-period deposits. The model assumes full inside equity financing, in the sense that bank equity is solely accumulated out of retained earnings. Formally, the law of motion for bank capital is similar to that proposed in Gerali et al. (2010):[16]

$$K_{b,t} = J_{b,t} - \; d_{b,t} + \; (1-\delta)K_{b,t-1}, \tag{9}$$

where $J_{b,t}$ stands for bank net profits and $\delta \in [0,1]$ denotes the fraction of own resources the banker can no longer accumulate as bank capital in period t due to exogenous factors. Rearranging in expression (9), bank net profits can be decomposed into three terms:

[16]Expression (9) only differs from the law of motion for bank capital proposed in Gerali et al. (2010) in that these authors assume that net profits are fully retained, period by period (i.e., there is no bank payout policy whatsoever).

$$J_{b,t} = \underbrace{\underbrace{(K_{b,t} - K_{b,t-1})}_{\text{reinvested profits}} + \underbrace{\delta K_{b,t-1}}_{\text{eroded equity}}}_{\text{retained earnings}} + \underbrace{d_{b,t}}_{\text{distributed earnings}}, \tag{10}$$

where the term $(K_{b,t} - K_{b,t-1})$ refers to the part of profits made in period t which are reinvested in the financial intermediation business, and $\delta K_{b,t-1}$ is the fraction of bank own resources which, due to exogenous factors, cannot be further accumulated as bank capital into the next period. The term $\delta K_{b,t-1}$ can be interpreted in several manners: (i) own resources the banker devotes to manage bank capital and to play its role as financial intermediary, or (ii) equity that erodes due to a variety of factors which are not explicitly accounted for in the model and which may relate to specific characteristics of capital such as its quality.

The definition of bank equity as a stock variable that accumulates over time out of retained earnings is a crucial assumption due to empirical-related reasons. First, an important proportion of total bank equity is accumulated out of retained earnings in practice (see figure 1C). Second, expression (9) plays a key role in incorporating the empirical link between payout policies and capital ratio adjustments (discussed in section 3) in the model by connecting the profit generation capacity of the representative banker (which is essential to distribute high and stable dividends over the cycle) with its capital generation capacity (which is crucial to meet capital requirements). Third, equation (9) allows to map the model to first and second moments of data on bank dividends and earnings (see section 6).

Equation (7) is a flow-of-funds constraint which states that in each period the banker has to distribute net profits $J_{b,t}$ between dividend payouts $d_{b,t}$ and retained earnings. In the basic model, bank net profits are defined as the difference between net interest income and the corresponding credit adjustment cost.[17] $r_{e,t}$ and $r_{h,t}$ denote the net interest rates on loans and deposits, respectively.

Expression (8) stipulates that bankers are constrained in their ability to issue liabilities. For a given period t, deposits cannot exceed

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[17] As in the case of the entrepreneur, $\Phi_b(B_t) = \dfrac{\phi_b}{2}\dfrac{(B_t - B_{t-1})^2}{B^{ss}}$ is a quadratic loan portfolio adjustment cost and is assumed to be external to the banker. $\phi_b \geq 0$ is the credit adjustment cost parameter.

a proportion $\gamma \in [0,1]$ of total assets. Given that this expression is binding in a neighborhood of the steady state, $(1-\gamma)$ can be interpreted as the regulatory capital ratio.

The optimality condition for this maximization problem can be obtained after having rearranged and substituted in its three first-order conditions:

$$\frac{(1-\gamma)+\dfrac{\partial \Phi_b(B_t)}{\partial B_t}}{d_{b,t}} = \beta_b E_t \left\{ \frac{(R_{e,t+1}-\delta)-\gamma\,(R_{h,t}-\delta)}{d_{b,t+1}} \right\}. \quad (11)$$

Expression (11) stands for the optimality condition for intertemporal substitution between the part of net income devoted to the dividend payout policy (denominator on each side of equation (11)) and that dedicated to the financial intermediation activity (numerator on each side of equation (11)). The engine of the intertemporal activity of bankers is earnings retention. Importantly, bankers endogenously manage the size of their balance sheet and set the growth path of future expected profits (and, thus, of expected dividends) by controlling for retained earnings.

From the perspective of the representative banker as a consumer, in the optimum the banker is indifferent between devoting an extra unit of profits to paying dividends today and postponing such payment to the next period. From the lens of the banker as a manager, it is optimal to invest (via earnings retention) up to the point in which the marginal cost of retaining an additional unit of net profits equalizes the marginal revenue of such investment. Expressed in terms of the opportunity cost (foregone marginal utility of dividends), the right-hand side of expression (11) informs about the discounted marginal gross lending spread the banker expects to obtain tomorrow as a consequence of having invested $(1-\gamma)$ units of retained earnings today.[18]

Given the interest rate on deposits, expression (11) determines the equilibrium interest rate on loans. Hence, the assumption by which $\beta_b < \beta_h$ ensures that in the steady state $(R_{e,t+1}-\delta) - \gamma\,(R_{h,t}-\delta) > 0$.

[18]By equations (6) and (8), such decision automatically involves borrowing additional γ units of deposits and lending an extra unit of assets.

Equation (11) synthesizes the information of a powerful mechanism for transmission and amplification of shocks that hit bank profits. The preference for dividend smoothing (expression (5)) implies that the bulk of the adjustment to shocks that hit profits is going to be made via retained earnings. Due to the strong link between equity and loans (equations (6) and (8)), such fluctuations in retained earnings are going to translate into loan supply volatility.

4.1.4 Aggregation and Market Clearing

In equilibrium, all markets clear. In the case of the final goods market, the aggregate resource constraint dictates that the income generated in the production process is fully spent in the form of final private consumption, credit adjustment costs, and resources devoted to manage the capital position of the bank, $\delta K_{b,t-1}$ (also interpretable as eroded equity):

$$Y_t = C_t + \delta K_{b,t-1} + \Phi_b(B_t) + \Phi_e(B_t), \tag{12}$$

where C_t denotes the aggregate consumption of the three agent types. Formally, $C_t = C_{h,t} + C_{e,t} + d_{b,t}$. Similarly, aggregate demand for housing equalizes supply. Housing supply is specified as a fixed endowment that is normalized to unity:

$$\overline{H} = H_{h,t} + H_{e,t}.$$

4.1.5 Macroprudential Policy

Consider two prudential policy scenarios alternative to the above presented baseline case.

Dividend Prudential Target (DPT). First, assume a policy scenario in which the static capital requirement, $(1- \gamma)$, is complemented by a regulatory scheme comprising

$$d_t^* = \rho_d + \rho_\chi \left(\frac{x_t}{x^{ss}} - 1 \right), \tag{13}$$

and

$$T(d_{b,t}, d_t^*) = \frac{\kappa}{2} \left(d_{b,t} - d_t^* \right)^2, \tag{14}$$

where d_t^* refers to the dividend prudential target. ρ_d is the bank dividend payout targeted by the prudential authority in the steady state. x_t is a macroeconomic indicator of the choice of the regulator. ρ_χ, the macroprudential policy parameter of policy rule (13), measures the degree of responsiveness of d_t^* to deviations of x_t from its steady-state level.

d_t^* enters a quadratic penalty function of the type (14). $\kappa \geqslant 0$ is the penalty parameter. When $\kappa > 0$, deviating from the dividend prudential target, d_t^*, is costly to bankers. If $d_{b,t} \neq d_t^*$ in period t, the resources paid by the representative banker as a sanction for having deviated, $T(d_{b,t}, d_t^*)$, are transferred by the public authority within the same period to the nonfinancial sector of the economy (to households and/or to entrepreneurs).[19]

Expression (14) shall be interpreted as a sanctions regime the DPT is associated with, having the aim of (i) striking a balance between enforcement (it penalizes bankers who deviate from the DPT) and flexibility of policy rule (13) (it allows for bankers to deviate from the target conditional on the payment of a sanction), and (ii) penalizing large deviations relatively more than small ones.

Importantly, the transmission of the regulatory scheme mainly takes place through the optimality condition of the representative banker, which now reads

$$\frac{(1-\gamma) + \dfrac{\partial \Phi_b(B_t)}{\partial B_t}}{d_{b,t}\left[1 + \kappa(d_{b,t} - d_t^*)\right]} = \beta_b E_t \left\{ \frac{(R_{e,t+1} - \delta) - \gamma\left(R_{h,t} - \delta\right)}{d_{b,t+1}\left[1 + \kappa\left(d_{b,t+1} - d_{t+1}^*\right)\right]} \right\}. \tag{15}$$

Absent a dynamic dividend target, the banker finds it optimal to react to exogenous shocks mostly by readjusting the variables that take part in the financial intermediation activity (numerator on each side of equation (11)). Under a dividend prudential target within the class (13), the regulator aims at discouraging bankers from making adjustments via credit supply by means of more responsive bank dividends (denominator on each side of equation (15)).

[19]Note that such transfer will be reflected in the corresponding budget and flow-of-funds constraints.

Dynamic Capital Requirements (CCyB). In order to compare the transmission channel and effects of the DPT with those of the main macroprudential tool of the Basel III Accord (i.e., the countercyclical capital buffer), I consider a third scenario in which the debt-to-assets ratio, γ, is augmented with a cyclical component

$$\gamma_t = \gamma + \gamma_x \left(\frac{x_t}{x^{ss}} - 1 \right), \tag{16}$$

where γ_x is the macroprudential policy parameter associated to the regulatory capital ratio implied by equation (16), $(1 - \gamma_t)$. Note that equations (8) and (11) are directly affected by this new policy environment.

There is a strand of literature on macro-banking models that attempts to evaluate the effects of the so-called countercyclical capital buffer (CCyB) by specifying a dynamic regulatory capital ratio similar to the one associated with policy rule (16) (see, e.g., Angelini, Neri, and Panetta 2014, Clerc et al. 2015, and Mendicino et al. 2018). However, most of these models do not capture an implicit characteristic of the CCyB that is relevant for the purpose of this paper.

Since this buffer is specified as a dynamic add-on to the conservation buffer, in practice the CCyB can be interpreted as a very particular "one-sided" dynamic restriction on banks' payout policies.[20] The combination of expressions (9) and (16) accounts for the essence of this characteristic. As will become evident in the numerical exercise, countercyclical dynamic capital requirements tend to restrict bankers' capacity to distribute earnings during the upturn (since they have to meet a higher capital ratio, to some extent by accumulating more capital out of retained earnings).

4.2 Numerical Exercise

The aim of this numerical exercise is to identify the transmission mechanism through which the DPT works and to quantitatively

[20]It can be interpreted as "one-sided" because the CCyB can never be negative. Recall that according to the Basel III Accord, banks can only distribute earnings as long as they meet their minimum capital requirements plus the conservation buffer. Thus, changes in the CCyB involve changes in this restriction on equity distribution.

assess its potential to tame the credit cycle in the face of financial (collateral) shocks. In order to do so, the paper follows Angelini, Neri, and Panetta (2014), who assume the macroprudential authority seeks to minimize an ad hoc loss function with respect to the corresponding vector of policy parameters.[21]

4.2.1 Calibration

The calibration is largely based on Gerali et al. (2010) and Iacoviello (2015). The households' discount factor is set to 0.9943, implying a steady-state interest rate on deposits slightly above 2 percent (2.3 percent). The discount factor of the entrepreneur is fixed to 0.94, within the range typically suggested in the literature for constrained consumers. The banker's discount factor, β_b, is chosen to ensure that the steady-state annualized lending rate to the private sector is roughly 5.6 percent, implying an annualized lending spread of 3.4 percent.

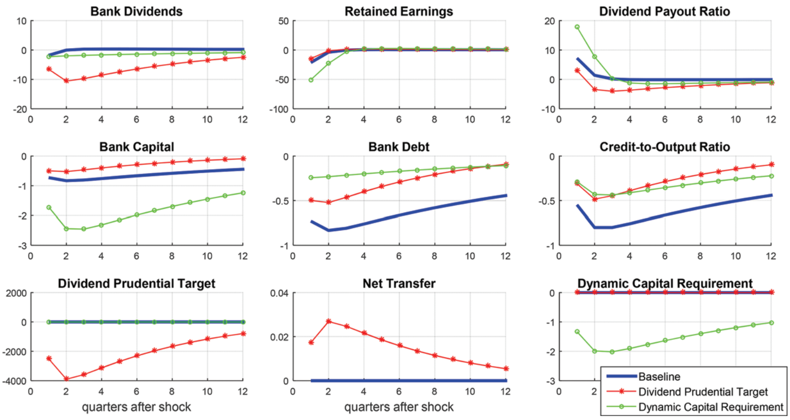
As in Iacoviello (2015), the weight of housing in the household's utility function is set to 0.075; the elasticity of production with respect to commercial housing, υ, to 0.05; the loan portfolio adjustment cost parameter of entrepreneurs and bankers to 0.25; and the leverage parameter for the bank to 0.9. The latter implies a capital-asset ratio of 0.1, implying a positive capital buffer (over the minimum capital requirement of 0.08), as the evidence suggests.

The loan-to-value ratio on housing, m_H, and the inverse of the Frisch elasticity of labor, ϕ, are set to standard values of 0.7 and 1.5, respectively.

The bank capital depreciation rate is calibrated at 0.034 so as to ensure that the steady-state dividend payout ratio is in the vicinity of 0.6, as the evidence for the SX7E index suggests. m_N is fixed to 0.5, implying a loan-to-output ratio of 1.9, as in the model estimated for the euro area in Gerali et al. (2010). The autocorrelation coefficient and the standard deviation associated with the housing

[21] In following that approach, there is no attempt in presenting such an objective function as a welfare criterion, but rather as a measure of the potential the proposed policy rule has to prevent the buildup of macrofinancial imbalances. A utility-based welfare analysis will be carried out in section 6 for the extended model.

Figure 3. The Transmission Mechanism: Impulse Response Functions to a Negative Financial Shock (basic model)



Notes: Variables are expressed in percentage deviations from the steady state. The solid line refers to the baseline scenario. The starred line corresponds to an alternative (policy) scenario in which the optimized prudential rule is a dividend prudential target. The dotted line relates to an alternative (policy) scenario in which the optimized prudential rule is a dynamic capital requirement.

collateral shock are obtained from the structural estimation of the same paper.

4.2.2 The Transmission Mechanism of Dividend Prudential Targets

Figure 3 plots the response of some key banking and financial aggregates to a negative collateral shock.[22] The shock triggers a credit crunch that negatively affects bank net profits. In line with the evidence shown in section 3, dividends and retained earnings fall during the bust (i.e., they are procyclical), with the former relatively less

[22]See Andrés, Arce, and Thomas (2013) and Iacoviello (2015) for a detailed description and presentation of the macroeconomic effects of housing collateral shocks faced by entrepreneurs in similar setups of the economy. Section 6 of this paper discusses the main macroeconomic effects of the proposed prudential instrument to a variety of shocks.

volatile than the latter. The dividend payout ratio is countercyclical since the adjustment is mainly borne by retained earnings.

The starred and dotted lines correspond to an economy in which the macroprudential authority is assumed to solve the following problem with respect to selected parameters of policy rules (13) and (16), respectively,

$$\arg\min_{\Theta} L^{mp} = \omega_z \sigma_z^2, \qquad \omega_z > 0, \tag{17}$$

where Θ refers to the vector of policy parameters with respect to which the policymaker solves the optimization problem and σ_z^2 is the asymptotic variance of a macroeconomic indicator of the choice of the regulator. Due to its relevance in macroprudential policy decisionmaking, x_t and z_t have both been chosen to be the loans-to-output ratio. Based on the literature, the preference parameter ω_z and the parameter of the penalty function (14), κ, are set to 1 and 0.426, respectively.[23]

In order to identify the optimal simple rule within the class (13) that solves (17), it has been searched over a multidimensional grid of parameter values, which can be defined as follows: $\rho_d \{0 - 1\}$, $\rho_x \{(-150) - 150\}$. The choice of the search grid deserves a thorough explanation. First, ρ_d refers to the dividend payout targeted by the prudential authority in the steady state. Taking that into account and normalizing the values for ρ_d by expressing them in terms of steady-state bank profits, it is reasonable to assume that its optimized value will lie somewhere between 0 and 1 (0 refers to the case in which all profits are retained and 1 to that in which steady-state profits are fully distributed). Second, a wide grid of values has been chosen for ρ_x, as the dynamics of this policy rule is largely unknown.

It has been searched within the baseline calibration model. The values that correspond to the optimized policy rule are the following: $\rho_d = 0.504$, $\rho_x = 66.003$. The optimal simple rule within the class (13) that solves (17) under full commitment calls for a countercyclical (i.e., $\rho_x > 0$) and highly responsive dividend prudential target

[23] 0.426 is the estimate Jermann and Quadrini (2012) provide for the parameter of a dividend adjustment cost whose functional form is identical to expression (14), and it falls within the range of values typically considered for this parameter in the literature.

and a steady-state dividend payout not far from the one targeted by bankers absent any dividend regulation.[24]

Then, I solve (17) with respect to parameter γ_x of policy rule (16) for the following grid of parameter values: $\gamma_x \{(-1) - 1\}$. Such grid is based upon the Basel III Accord and has been chosen to assess whether the optimized capital buffer in this model is countercyclical (i.e., $\gamma_x < 0$) or not. The optimized policy rule within the class (16) that solves (17) under full commitment corresponds to $\gamma_x = -0.461$.[25]

Both macroprudential policy rules are effective in smoothing loan supply and the loans-to-output ratio. However, the DPT seems to be relatively more effective than the CCyB due to the different transmission channels through which each of the policy rules operate. Under an optimized dynamic capital requirement, the target capital ratio of the representative banker readjusts. The bulk of such adjustment in the face of an exogenous shock is borne by bank capital (i.e., ultimately by retained earnings).[26] Consequently, debt—which now represents a larger proportion of total assets—evolves in a smoother fashion (than under the baseline scenario), unambiguously generating a smoothing effect on credit supply. By way of contrast, dividend prudential targets directly attack the root of the "problem" (i.e., dividend smoothing). They provide incentives for bankers to optimally tolerate a higher degree of dividend volatility, thereby allowing for smoother retained earnings, equity, and loans supply.

Table 1 reports the prudential losses and optimized policy parameter values related to the solution to problem (17), for two alternative arguments of the loss function, $\sigma_z^2 \equiv \left\{\sigma_B^2; \sigma_{B/Y}^2\right\}$; two different macroeconomic indicators, $x \equiv \{B; B/Y\}$; and three alternative

[24]Recall that the baseline calibration implies a dividend payout ratio of roughly 0.6.

[25]In order to ensure that I have found a global minimum in each of the two optimization problems, I have selected different tuples of initial conditions. Optimized parameter values remain the same regardless of the initial guess.

[26]Note that, under an optimized dynamic capital requirement, the "problem" of dividend smoothing exacerbates (i.e., the proportion of the adjustment suffered by bank profits in the face of an exogenous shock that is borne by retained earnings is larger than under the baseline scenario).

Table 1. Optimized Rules and Prudential Losses: Collateral Shock (basic model)

		$\boldsymbol{\sigma_B^{2(1)}}$	$\boldsymbol{\sigma_{B/Y}^2}$
A. $x_t = B_t$			
(i) $\{\rho_d, \rho_x\}$	Loss Variation(2)	(-83.81)	(-85.65)
	$\rho_d^{(3)}$	0.535	0.532
	ρ_x	66.811	66.786
(ii) $\{\rho_x\}$	Loss Variation	(-76.54)	(-78.28)
	ρ_x	52.755	52.755
(iii) $\{\gamma_x\}$	Loss Variation	(-65.11)	(-64.80)
	γ_x	-0.335	-0.335
B. $x_t = B_t/Y_t$			
(i) $\{\rho_d,\ \rho_x\}$	Loss Variation	(-80.40)	(-82.61)
	ρ_d	0.546	0.504
	ρ_x	67.378	66.003
(ii) $\{\rho_x\}$	Loss Variation	(-73.86)	(-76.01)
	ρ_x	54.962	54.962
(iii) $\{\gamma_x\}$	Loss Variation	(-73.71)	(-73.43)
	γ_x	-0.461	-0.461

Notes: (1) Asymptotic variance that enters the objective function of the prudential authority in problem (17). Such problem has been solved numerically by means of the osr (i.e., optimal simple rule) command in Dynare. (2) Percentage changes in the value of the loss function under the corresponding policy scenario with respect to the baseline scenario. (3) Values of the autonomous component of the dividend prudential target have been normalized by expressing them in terms of steady-state bank profits.

policy scenarios, $\Theta \equiv \{(\rho_d, \rho_x); \rho_x; \gamma_x\}$. Panels A and B refer to the cases in which $x_t = B_t$ and $x_t = B_t/Y_t$, respectively. For each part of the table, sections (i), (ii), and (iii) present the results of the solution to the mentioned problem when optimizing with respect to $\Theta = (\rho_d, \rho_x)$, $\Theta = \rho_x$, and $\Theta = \gamma_x$, respectively.

The main findings of this exercise can be summarized as follows. First, and due to the different transmission mechanisms through which they operate, the DPT is more effective in smoothing credit supply and the loans-to-output ratio than dynamic capital requirements.[27] Second, the DPT mainly reduces loan supply volatility through its cyclical component, allowing for tangible macroeconomic effects over the cycle without having to affect long-run dividend payouts.[28] Third, the transfer system defined by equation (14) can be interpreted as a sanctions regime that acts as an insurance scheme for the real economy. As noted in figure 2, the net transfer associated with the optimized DPT, $T_t^*(d_{b,t}, d_t^*)$, is countercyclical. That is, its recipients (households and/or entrepreneurs) benefit from a positive payoff when the marginal utility of their consumption is relatively high.

5. Extended Model

In order to improve the dynamics of the model and its mapping to the data, the model is extended in three main directions. First, a second type of household with a lower subjective discount factor is incorporated into the model. Thus, two types of households coexist, one being net savers (patient households) and the other one being net borrowers (impatient households). In equilibrium bank loans are now extended to credit-constrained households and entrepreneurs. Second, the model allows for variable physical capital. Capital-good producers sell their output to entrepreneurs, who use it as an input in the productive process. Third, additional shocks are considered to allow for a more comprehensive analysis of dividend prudential targets.

In this version of the model households own all existing firms (final-good-producing firms, banks, and capital-good-producing firms), which has two important implications. First, there is a separation between bank ownership and bank management that allows

[27]Note that one of the goals of this numerical exercise is to evaluate and compare the potential of policy rules (13) and (16) in taming the credit cycle rather than that of reproducing the precise effects they would generate in reality. Percentage changes induced by such rules in the asymptotic variance of credit gaps are relatively large in this numerical exercise, among other reasons, because the model assumes one-period loans (rather than long-term debt) and the asymptotic variance of the policy instrument does not enter the loss function of problem (17).

[28]Note that the differences in terms of macroprudential losses between solving the optimization problem with respect to ρ_d and ρ_x, and solving it only with respect to ρ_x are small.

to capture the two main channels through which dividend smoothing operates according to the evidence (i.e., bank owners' risk aversion and managers' propensity to smooth dividends). Second, as in Clerc et al. (2015) and Mendicino et al. (2018), the welfare analysis can be restricted to households without neglecting any consumption capacity generated in the economy.

The specification of preferences has also been revised for all types of agents: (i) Households in the extended model are assumed to have GHH preferences (see Greenwood, Hercowitz, and Huffman 1988). This type of preference has been extensively used in the business cycle literature as a useful device to match several empirical regularities. Their main difference when compared with log preferences, as assumed in the basic model, is that consumption and leisure are nonseparable and wealth effects on labor supply are arbitrarily close to zero.[29] (ii) By generalizing log utility functions of entrepreneurs and bankers to CES utility functions, corresponding elasticities of intertemporal substitution can be calibrated to match the second moments of dividends.

This section only discusses the main changes the extended model incorporates with respect to the basic version under a policy scenario in which both the DPT and dynamic capital requirements operate. The full set of equilibrium equations can be found in online appendix C.

5.1 Overview of the Model

5.1.1 Households

Impatient households discount the future more heavily than patient ones, implying $\beta_i < \beta_p$. In the extended model the representative household maximizes

[29]See Jaimovich and Rebelo (2009) for a generalization of GHH preferences and Galí (2011) for a similar specification of individual preferences that permits to control for the size of wealth effects. Schmitt-Grohé and Uribe (2012) present evidence suggesting that wealth effects on labor supply are practically zero. As in this paper, GHH preferences have been formulated by other authors when evaluating macroprudential policies, in order to prevent a counterfactual increase in labor supply during crises (see, e.g., Bianchi and Mendoza 2018).

$$E_0 \sum_{t=0}^{\infty} \beta_{\varkappa}^t \left[\frac{1}{1-\sigma_h} \left(C_{\varkappa,t} - \frac{N_{\varkappa,t}^{1+\phi}}{(1+\phi)} \right)^{1-\sigma_h} + \varepsilon_t^h j \log H_{\varkappa,t} \right], \quad (18)$$

where $\varkappa = p, i$ denotes the type of household the problem refers to. σ_h stands for the risk-aversion parameter of households and ε_t^h captures exogenous housing preference shocks. Shocks in the extended model have the same properties as the one presented in the basic version.

Patient Households (Net Savers). In the case of patient households, the maximization of (18) is restricted to the sequence of budget constraints:

$$\begin{aligned} C_{p,t} + D_t + q_t(H_{p,t} - H_{p,t-1}) = R_{d,t-1}D_{t-1} + W_tN_{p,t} + \omega_b d_{b,t} \\ + \chi T(d_{b,t}, d_t^*) + \omega_e d_{e,t}, \qquad (19) \end{aligned}$$

where $d_{e,t}$ refers to earnings distributed by entrepreneurs, $\omega_b \in [0,1]$ is the fraction of banks owned by patient households, and $\omega_e \in [0,1]$ is the proportion of entrepreneurial firms owned by the same agent type. χ is the fraction of net subsidy they receive from the prudential authority, which is considered to be equal to the stake of banks they own (i.e., $\chi = \omega_b$). That is, the degree of "insurance" received by households is assumed to be proportional to their exposure to the increased bank dividend volatility triggered by the proposed regulatory scheme. This is relevant under policy scenarios in which the DPT operates and the following inequality may hold, $T(d_{b,t}, d_t^*) \neq 0$.

Impatient Households (Net Borrowers). As a net borrower, the representative impatient household is restricted not only by a sequence of budget constraints but also by a borrowing limit,

$$\begin{aligned} &C_{i,t} + R_{i,t-1}B_{i,t-1} + q_t(H_{i,t} - H_{i,t-1}) + \Phi_i(B_{i,t}) \\ &\quad = B_{i,t} + W_tN_{i,t} + (1-\omega_b)d_{b,t} + (1-\chi)T(d_{b,t}, d_t^*) + (1-\omega_e)d_{e,t}, \end{aligned} \quad (20)$$

$$B_{i,t} \leq m_{i,t}^H E_t \left[\frac{q_{t+1}}{R_{i,t}} H_{i,t} \right]. \quad (21)$$

In each period, impatient households devote their available resources—in terms of wage earnings, loans, distributed earnings, and the corresponding net subsidy—to consume, repay their debt, demand housing, and adjust their loan portfolio. As was the case for entrepreneurs in the basic model, the borrowing capacity of impatient households is tied to the expected value of their housing property. $m^H_{i,t}$ captures exogenous shocks to such collateral.

5.1.2 Entrepreneurs

Let $\Lambda^e_{t,t+1} = \left[\omega_e \beta_p \frac{\lambda^p_{t+1}}{\lambda^p_t} + (1-\omega_e)\beta_i \frac{\lambda^i_{t+1}}{\lambda^i_t}\right]$ be the stochastic discount factor of entrepreneurs (managers), with λ^p_t and λ^i_t being the Lagrange multipliers of the patient and impatient households' optimization problems, respectively. Then, the representative entrepreneur maximizes

$$E_0 \sum_{t=0}^{\infty} \Lambda^e_{t,t+1} \quad \frac{1}{(1-\frac{1}{\sigma})} d_{e,t}^{(1-\frac{1}{\sigma})}, \tag{22}$$

subject to the sequence of budget constraints, the available technology, and the corresponding borrowing limit:

$$\begin{aligned} & d_{e,t} + R_{b,t}B_{e,t-1} + q^k_t \left[K_{e,t} - (1-\delta^k_t)K_{e,t-1}\right] + q_t(H_{e,t} - H_{e,t-1}) \\ & \quad + W_t N_t + \Phi_e(B_{e,t}) = Y_t + B_{e,t}, \end{aligned} \tag{23}$$

$$Y_t = A_t(u_t k_{e,t-1})^{\alpha} H^{\eta}_{e,t-1} N_t^{(1-\alpha-\eta)}, \tag{24}$$

$$B_t \leq m^H_{e,t} E_t \left(\frac{q_{t+1}}{R_{e,t+1}} H_{e,t}\right) - m^N W_{h,t} N_t. \tag{25}$$

Note the three differences of this optimization problem when compared with the one presented in the previous section. First, owners and managers of final-good-producing firms are no longer the same agent. Second, entrepreneurs also face technology shocks, captured by A_t. Third, in order to produce final goods, the available technology combines not only labor and commercial real estate but also variable physical capital. As in Schmitt-Grohé and Uribe (2012), the depreciation rate of physical capital is an increasing and convex function of the rate of capacity utilization. In particular,

$$\delta^k_t\,(u_t) = \delta^k_0 + \delta^k_1(u_t-1)^2 + \frac{\delta^k_2}{2}\,(u_t-1)^2. \tag{26}$$

5.1.3 Bank Managers

Similarly, $\Lambda^b_{t,t+1} = \left[\omega_b \beta_p \frac{\lambda^p_{t+1}}{\lambda^p_t} + (1-\omega_b)\beta_i \frac{\lambda^i_{t+1}}{\lambda^i_t}\right]$ stands for the stochastic discount factor of bankers. Bank managers seek to maximize

$$E_0 \sum_{t=0}^{\infty} \Lambda^b_{t,t+1} \frac{1}{(1-\frac{1}{\sigma})} d_{b,t}^{\left(1-\frac{1}{\sigma}\right)}, \tag{27}$$

subject to a balance sheet identity, a sequence of cash flow restrictions, and a borrowing constraint, respectively:

$$B_{it} + B_{e,t} = K_{b,t} + D_{b,t}, \tag{28}$$

$$\begin{aligned} d_{b,t} + K_{b,t} - (1-\delta_t)K_{b,t-1} = r_{e,t}B_{e,t-1} + r_{i,t-1}B_{i,t-1} \\ - r_{d,t-1}D_{b,t-1} - \Phi_{be}(B_{e,t}) \\ - \Phi_{bi}(B_{i,t}) - T(d_{b,t}, d^*_t), \end{aligned} \tag{29}$$

$$D_{b,t} = \gamma_{i,t}B_{i,t} + \gamma_{e,t}B_{e,t}. \tag{30}$$

As for the case of entrepreneurs, in the extended model there is a separation between ownership and management of banks. Importantly, both mechanisms through which dividend smoothing operates in the model—households' risk aversion and managers' propensity to smooth—are incorporated in the bank manager's problem via the stochastic discount factor and managers' CES preferences, respectively. The loan portfolio is composed of two types of assets, $B_{i,t}$ and $B_{e,t}$, which may differ in two aspects: (i) the complementary of their associated capital requirements, $\gamma_{i,t}$ and $\gamma_{e,t}$, and (ii) their respective adjustment cost parameters. $\delta_t = \delta\varepsilon^{kb}_t$ denotes a possibly time-varying erosion rate of bank equity, where $\delta \in [0,1]$ and ε^{kb}_t captures exogenous shocks to bank capital.[30] The solution to this

[30] ε^{kb}_t captures bank capital shocks similar to those considered in Angelini, Neri, and Panetta (2014). However, in this paper I assume that ε^{kb}_t hits eroded bank equity, $\delta K_{b,t-1}$, rather than uneroded bank capital, $(1-\delta)K_{b,t-1}$. Since the term $\delta_t K_{b,t-1}$ enters the resource constraint, this is an important consideration in order to ensure that all statistical moments of output as defined in equation (24) are identical to those of aggregate demand as defined in the resource constraint of the model economy (see the aggregate resource constraint in appendix C) and, thus, to guarantee that the model is "properly closed."

optimization problem yields two optimality conditions analogous to expression (11), one for each asset class.

5.1.4 Capital Goods Producers

At the beginning of each period, capital producers demand an amount I_t of final good from entrepreneurs which, combined with the available stock of capital, allows them to produce new capital goods. Capital producers choose the trajectory of net investment in variable capital, I_t, that maximizes

$$E_0 \sum_{t=0}^{\infty} \Lambda_{t,t+1}^{e} \left(q_{k,t} \Delta x_{k,t} - I_t \right), \tag{31}$$

subject to

$$x_{k,t} = x_{k,t-1} + I_t \left[1 - \frac{\psi_I}{2} \left(\frac{I_t}{I_{t-1}} - 1 \right)^2 \right], \tag{32}$$

where $\Delta x_{k,t} = K_{e,t} - (1 - \delta_t^k) K_{e,t-1}$ is the flow output. $S\left(\frac{I_t}{I_{t-1}}\right) = \frac{\psi_I}{2}\left(\frac{I_t}{I_{t-1}} - 1\right)^2$ is an investment adjustment cost function whose formulation has become standard in the literature (see, e.g., Christiano, Eichenbaum, and Evans 2005 and Schmitt-Grohe and Uribe 2012) due to empirical reasons.

6. Quantitative Analysis

6.1 Calibration

I follow a three-stage strategy in order to calibrate the model to quarterly euro-area data for the period 2002:Q1–2018:Q2.[31] First, several parameters are set following convention (table 2A). Some of them

[31] All time series expressed in euros are seasonally adjusted and deflated. With regards to the matching of second moments, the log value of deflated time series has been linearly detrended before computing standard deviation targets. All details on data description and construction are available in online appendix A.

Table 2. Baseline Parameter Values

Parameter	Description	Value	Source/Target Ratio
A. Preset Parameters			
φ	Inverse of the Frisch Elasticity	1	Standard
σ_h	HH Risk-Aversion Parameter	2	Standard
m_{Hi}; m_{He}	LTV Ratio on HH and NFC Housing	0.7	Standard
δ^k	Depreciation Rate of Physical Capital	0.025	Standard
δ_1^k; δ_2^k	Endogenous Depr. Rate Parameters	$r_{ke}^{ss}; 0.1 \times r_{ke}^{ss}$	Gerali et al. (2010)
ϕ_e	NFC Credit Adjustment Cost Parameter	0.06	Iacoviello (2015)
κ	Penalty Parameter	0.426	Jermann and Quadrini (2012)
B. First Moments			
β_p	Savers' Discount Factor	0.9943	$R_h^{ss} = (1.023)^{1/4}$
β_i	Borrowers' Discount Factor	0.95	$(r_b^{ss} - r_d^{ss})400 = 3.4$
j_p	Savers' Housing Weight	0.0805	$H_p^{ss}/H_i^{ss} = 1.3585$
j_i	Borrowers' Housing Weight	0.4802	$B_i^{ss}/(Y^{ss}) = 2.1403$
ω_e	Fraction of Firms Owned by HH_p	1	$B_e^{ss}/B^{ss} = 0.4510$
ω_b	Fraction of Banks Owned by HH_p	0	$B_e^{ss}/Y^{ss} = 1.7530$
α	Capital Share in Production	0.2699	$I^{ss}/Y^{ss} = 0.2119$
η	Real Estate Share in Production	0.0385	$(q^{ss}H^{ss})/(4Y^{ss}) = 2.802$
γ_e	Debt-to-Assets, NFC Risk Adjusted	0.8508	$\gamma_e/\gamma_i = 2.1176$
γ_i	Debt-to-Assets, HH Risk Adjusted	0.9295	$K_b^{ss}/B^{ss} = 0.105$
δ	Depreciation Rate of Bank Capital	0.041	$d_b^{ss}/J_b^{ss} = 0.5625$
C. Second Moments			
ψ	Investment Adj. Cost Parameter	0.092	$\sigma_I/\sigma_Y = 2.642$
ϕ_i	HH Credit Adj. Cost Parameter	0.511	$\sigma_B/\sigma_Y = 6.473$
σ	Banker EIS	2.40	$\sigma_{db}/\sigma_Y = 15.050$
σ_h	Std. Housing Pref. Shock	0.1999	$\sigma_q/\sigma_Y = 2.429$
σ_{kb}	Std. Bank Capital Depr. Shock	0.0495	$\sigma_{Kb}/\sigma_Y = 6.554$
σ_{mh}	Std. NFC Collateral Shock	0.0024	$\sigma_{Jb}/\sigma_Y = 59.102$
σ_{mk}	Std. HH Collateral Shock	0.0026	$\sigma_C/\sigma_Y = 0.748$
σ_A	Std. Productivity Shock	0.0020	$\sigma_Y = 2.138$

Notes: Parameters in panel A are set to standard values in the literature, whereas those in panels B and C are calibrated to match data targets. Abbreviations HH and NFC refer to households and nonfinancial corporations (entrepreneurs), respectively. HH_p stands for patient households.

Table 3. Steady-State Ratios

Variable	Description	Model	Data
C^{ss}/Y^{ss}	Total Consumption-to-GDP Ratio	0.7632	0.7607
I^{ss}/Y^{ss}	Gross Fixed Capital Formation-to-GDP Ratio	0.2196	0.2119
$r_b^{ss} \times 400$	Annualized Bank Rate on Loans (Percent)	6.020	5.6
$r_d^{ss} \times 400$	Annualized Bank Rate on Deposits (Percent)	2.293	2.3
$(r_b^{ss} - r_d^{ss})400$	Annualized Bank Spread (Percent)	3.727	3.4
$(1-\gamma_e)/(1-\gamma_i)$	Capital Requirement of NFC Loans-to-Mortgage Loans	2.1176	2.1176
K_b^{ss}/B^{ss}	Capital Requirements on Mortgage and NFC Loans	0.105	0.105
B_i^{ss}/Y^{ss}	HH Loans-to-GDP Ratio	2.1875	2.1291
B_e^{ss}/Y^{ss}	NFC Loans-to-GDP Ratio	1.7938	1.7530
B_i^{ss}/B^{ss}	Fraction of HH Loans	0.5494	0.5490
B_e^{ss}/B^{ss}	Fraction of NFC Loans	0.4506	0.4510
d_b^{ss}/J_b^{ss}	Bank Dividend Payout-Ratio	0.5621	0.5625
h_p^{ss}/h_i^{ss}	Savers-to-Borrowers Housing Ratio	1.4763	1.3585
$(q^{ss}H^{ss})/(4Y^{ss})$	Housing Wealth-to-GDP Ratio	2.6104	2.8018

Notes: All series in euros are seasonally adjusted and deflated. Data targets have been constructed from euro-area quarterly data for the period 2002:Q1–2018:Q2. The exceptions are the following: annualized bank rates, which have been taken from constructed series presented in Gerali et al. (2010), and the target for capital requirements, which has been based on the Basel III regime. Data sources are Eurostat, ECB, and Bloomberg.

are standard in the literature. Others are based on papers in the field of macro-finance. The inverse of the Frisch elasticity of labor is set to a value of 1, whereas the risk-aversion parameter of household preferences is fixed to a standard value of 2. Loan-to-value ratios on housing (for both households and entrepreneurs) are set equal to 0.7. These values are based on data of the big four euro-area economies and coincide with those presented in Gerali et al. (2010) and Quint and Rabanal (2014), among others. Regarding the dynamic depreciation rate of physical capital δ_t^k, δ_0^k is fixed to a standard value of 0.025 while, following convention, δ_1^k and δ_2^k are defined as specific fractions of the steady-state interest rate on physical capital. The adjustment cost parameter value for corporate loans coincides with that obtained in the structural estimation by Iacoviello (2015).

Second, another group of parameters is calibrated by using steady-state targets (tables 2B and 3). The patient households' discount factor, $\beta_p = 0.9943$, is chosen such that the annual interest rate equals 2.3 percent. The impatient households' discount factor is set to 0.95, in order to generate an annualized bank spread of 3.4 percent. Household weights on housing utility, j_p and j_i, have been calibrated to match the savers-to-borrowers housing ratio and the household loans-to-GDP ratio, respectively.

Patient households are assumed to own all the entrepreneurial and capital-producing firms of the economy, $\omega_e = 1$, while impatient households own all the banks, $\omega_p = 0$. This calibration is based on the following reasons: (i) They are chosen to match a corporate loans-to-GDP ratio of 175.3 percent and a weight of corporate loans on total credit of 0.451, respectively. (ii) It permits to limit the welfare analysis to two types of agents (henceforth referred to as savers and borrowers) while fully separating by agent types the two main types of welfare tradeoffs triggered by optimized dividend prudential targets.[32,33]

The shares in final-good production of physical capital α and commercial real estate η are set to match an investment-to-GDP ratio of 21.19 percent and an aggregate real estate wealth-to-annual output of 280.2 percent, respectively.

With regard to bank parameters, I proceed as follows. The depreciation rate of bank capital δ is set to 0.041, which is consistent with a payout ratio of 0.563, in line with the evidence of the SX7E banks'

[32]The assumption by which both patient and impatient households can potentially own banks and nonfinancial corporations in the model is empirically relevant. However, there is no evidence on what proportion of each type of firms are owned by each type of household. Thus, and given the targeted steady-state ratios in the calibration, it is desirable to assume that each type of representative household fully owns in isolation one of the two main types of firms in order to clearly identify the relevant welfare tradeoffs. Of course, that requires main results of the welfare analysis to be taken cautiously and interpreted accordingly.

[33]In addition, the proposed setup does not allow for savers to own all entrepreneurial firms and banks. Were they owners of all banks, the relationship between β_p and $\Lambda^b_{t,t+1}$ would be such that there would not be positive financial flows in equilibrium. Alternative setups have been proposed in the literature to allow savers to be owners of all firms in the economy (see, e.g., Gertler and Kiyotaki 2010, Gertler and Karadi 2011, and Clerc et al. 2015). However, in order for these approaches to be applicable, these authors have to make assumptions implying that dividend payout ratios are constant and (usually) very low, a result that is sharply at odds with reality and which does not permit to carry out the type of analysis proposed in this paper.

index presented in section 3.[34] Note that after having rearranged in the steady-state expression of equation (9),

$$\frac{d_b^{ss}}{J_b^{ss}} = 1 - \frac{\delta K_b^{ss}}{J_b^{ss}},$$

from which the influence parameter δ has on the steady-state payout ratio becomes evident. The calibrated values of the complementaries of capital requirements on household loans γ_i and corporate loans γ_e are obtained by solving a system of two linear equations:

$$0.895 = \gamma_i \frac{B_i^{ss}}{B^{ss}} + \gamma_e \frac{B_e^{ss}}{B^{ss}}, \tag{33}$$

$$(1 - \gamma_e) = 2.1176(1 - \gamma_i). \tag{34}$$

Equation (33) is the result of equating the steady-state leverage ratio to 0.895 after having normalized expression (30) to total loans. Its interpretation is straightforward. The equilibrium capital requirement is a weighted average of the two sectoral capital requirements, $(1-\gamma_e)$ and $(1-\gamma_i)$, and it has been set to 0.105. Such value has been chosen for empirical and regulatory reasons: (i) It is similar to the pre-crisis historical average of regulatory capital ratios. (ii) According to existing capital legislation, in general terms, the authority cannot impose any restriction on dividend payouts as long as the bank meets the minimum capital requirement (0.08) plus a conservation buffer of 0.025.

Expression (34) indicates that the capital requirement on corporate loans is slightly more than two times that on household loans. This is exactly the same proportion held by these two sectorial ratios according to the internal-ratings-based (IRB) calibration presented in Mendicino et al. (2018). For simplicity, a 100 percent risk weight has been assumed for each of the two asset types.[35]

[34]This result is aligned with Lintner (1956) and subsequent literature, who found that corporations target a payout ratio of roughly 55 percent.

[35]This assumption is reasonable. As the Capital Requirements Regulation (EU) stipulates, exposures to corporates with an "average" credit rating or for which no credit assessment is available shall be assigned a 100 percent risk weight. Unless certain conditions are met, exposures fully secured by a mortgage on immovable property shall also be assigned a risk weight of 100 percent.

Table 4. Second Moments

Variable	Description	Model	Data
A. Banking Data (SX7E)			
σ_{db}/σ_Y	Std. Bank Dividends	14.880	15.050
σ_{Jb}/σ_Y	Std. Bank Profits	43.037	59.102
σ_{Kb}/σ_Y	Std. Bank Capital	6.087	6.554
$\sigma_B/\ \sigma_Y$	Std. Bank Assets	6.870	6.473
B. Macro Data (EA)			
σ_q/σ_Y	Std. Housing Prices	2.133	2.429
σ_I/σ_Y	Std. Investment	3.318	2.642
σ_C/σ_Y	Std. Consumption	0.933	0.748
σ_Y	Std.(GDP) $\times$ 100	2.136	2.138

Notes: All series are seasonally adjusted and deflated, and their log value has been linearly detrended before computing standard deviation targets. Since some observations in the series "bank profits" (i.e., earnings) take negative values, in this case a constant has been added to all observations before taking logs, such that the minimum of the transformed series is equal to one. The standard deviation (Std.) of GDP is in quarterly percentage points.

Third, the size of shocks and certain adjustment cost parameters are calibrated to improve the fit of the model to the data in terms of relative volatilities (see tables 2C and 4). The investment adjustment cost parameter ψ_I is set to target a relative standard deviation of investment of 2.642 percent. The adjustment cost parameter on household loans ϕ_i is fixed to a value of 0.511, thereby (i) favoring corporate loans to be relatively more volatile than household loans, as supported by the evidence in the euro area (recall that corporate loans parameter ϕ_e has been preset to 0.06), and (ii) roughly matching the relative volatility of bank assets.

I have matched the second moments of bank dividends and earnings by calibrating the elasticity of intertemporal substitution (EIS) of bankers and the size of the bank capital shock. Several important considerations are worth noting in this regard. First, I have opted to account for the stylized fact of managers' preference for dividend smoothing by means of a CES utility function (and matched the second moment of bank dividends by calibrating parameter σ) rather than by assuming linear preferences and a dividend adjustment cost function (in the baseline scenario) of the type (14) (and attempted to match the second moment of dividends by calibrating parameter κ), as elsewhere in the literature (see, e.g., Jermann and Quadrini 2012 and Begenau 2020),[36] for two main reasons: (i) the latter specification does not permit to match the relative volatility of aggregate bank dividends with a sufficient degree of accuracy, and (ii) even though a careful microfoundation of the potential forces underlying managers' preference for dividend smoothing is beyond the scope of this paper, assuming that the origin of this phenomenon relates to individual preferences seems more reasonable than associating it with an external adjustment cost parameter.[37] Second, calibrating the size of the bank capital shock is relevant to allow for dividends and earnings to be sufficiently volatile, while fixing the value of σ permits to create a wedge between the standard deviation of earnings and that of dividends.

As in the basic model, the autoregressive parameters of the five shocks that are present in the extended model correspond to the estimates proposed in Gerali et al. (2010).

6.2 Welfare Analysis

This section investigates the main welfare consequences of complementing capital requirements with a dividend prudential target. In

[36]More precisely, the volatility of aggregate dividends in this type of DSGE model is mainly driven by two key forces: households' risk aversion, which implicitly involves a preference for smoothing available resources (including distributed earnings), and the specification of some motive for dividend smoothing in the (bank) manager's problem. For instance, Begenau (2020) assumes that household preferences are logarithmic (in consumption) and banks risk neutral (in dividends), with a dividend adjustment cost function of the type (14) aimed at accounting for dividend smoothing through the calibration of parameter κ. The calibration of the model to quarterly data of the U.S. economy suggests that households' logarithmic preferences imply a "too low" relative standard deviation of consumption (0.81 in the data versus 0.38 in the model), which translates into a "too low" relative standard deviation of bank dividends (28.01 in the data versus 13.40 in the model) even if parameter κ is set to 0.01. Jermann and Quadrini (2012) also assume that managers are risk neutral and a dividend adjustment cost function, but calibrate parameter κ to match the relative volatility of the aggregate dividends-to-GDP ratio (rather than aggregate dividends).

[37]In fact, there is no broad consensus on why managers have such a preference for smoothing dividends and what determines their propensity to smooth (see, e.g., Leary and Michaely 2011).

order to do so, a normative approach is adopted and a measure of social welfare—specified as a weighted average of the expected lifetime utility of savers and borrowers—is maximized with respect to the corresponding policy parameter/s. Formally,

$$\arg\max_{\Theta} V_0 = \zeta_p V_0^p + \zeta_i V_0^i, \tag{35}$$

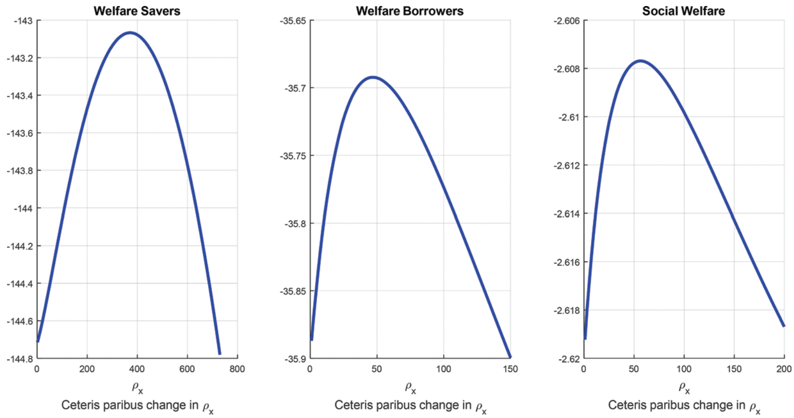
where $V_0^{\varkappa} = E_0 \sum_{t=0}^{\infty} \beta_{\varkappa}^t U \left(C_{\varkappa,t}, H_{\varkappa,t}, N_{\varkappa,t}\right)$ is the expected lifetime utility function of household type $\varkappa = p, i$; $\zeta_{\varkappa}$ denotes the utility weight of agent class $\varkappa = p, i$; and Θ refers to the vector of policy parameters with respect to which the objective function is maximized. Problem (35) is subject to all the competitive equilibrium conditions of the extended model. As in Schmitt-Grohé and Uribe (2007), welfare gains of agent type "$\varkappa$" are defined as the implied permanent differences in consumption between two different scenarios. Formally, consumption equivalent gains can be specified as a constant $\lambda_{\varkappa}$, which satisfies

$$\begin{aligned} &E_0 \sum_{t=0}^{\infty} \beta_{\varkappa}^t U \left(C_{\varkappa,t}^a, H_{\varkappa,t}^a, N_{\varkappa,t}^a\right) \\ &\quad = E_0 \sum_{t=0}^{\infty} \beta_{\varkappa}^t U \left[\left(1+\lambda_{\varkappa}\right) C_{\varkappa,t}^b, H_{\varkappa,t}^b, N_{\varkappa,t}^b\right], \end{aligned} \tag{36}$$

where superscripts a and b refer to the alternative policy scenario and the baseline case, respectively.

Since there is no widely accepted criterion to assign values to ζ_p and ζ_i, I rely on two alternative but complementary criteria that have often been used in the recent macro-finance literature to prevent an overweight of savers' welfare related to a higher discount factor (see, e.g., Lambertini, Mendicino, and Punzi 2013, Mendicino and Punzi 2014, and Mendicino et al. 2018). Welfare weighting criterion A solves problem (35) by further assuming that $\zeta_{\varkappa} = (1-\beta_{\varkappa})$, with $\varkappa = p, i$. That ensures the same utility weights across households discounting future utility at different rates. Welfare criterion B goes one step further in treating both types of agents equally and imposes additional restrictions on the solution to problem (35) according to which welfare gains have to be non-negative and identical across households (i.e., $\lambda_p = \lambda_i$ and $\lambda_p \geqslant 0$, $\lambda_i \geqslant 0$) and

Figure 4. Welfare Effects of DPTs (welfare effects of ceteris paribus changes in ρ_x)



Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of the dividend prudential target, ρ_x, while keeping the other policy parameter, ρ_d, to its baseline calibration value.

$\zeta_p + \zeta_i = 1$. Under this criterion, social welfare gains are identical to those of savers and borrowers regardless of the weights assigned to each of them.

Figure 4 plots the individual and social welfare effects of changing the value of parameter ρ_x in a policy rule of the type (13), with $\kappa = 0.426$, $\rho_d = d_b^{ss}$, and $x_t = Y_t$.[38] There is a considerable range of positive ρ_x values for which both types of agents are better off than under the baseline scenario. Interestingly, figure 4 makes clear that each type of agent faces a different tradeoff when being exposed to changes in ρ_x. Such tradeoffs primarily depend on the smoothing effects DPTs trigger on household and entrepreneurial firm loans (which have a direct positive impact on borrowers and an indirect one on savers, as owners of nonfinancial corporations) as well as on the welfare costs in terms of higher bank dividend volatility and

[38] As in Angelini, Neri, and Panetta (2014), and without loss of generality, the macroeconomic indicator x_t incorporated in the policy rule under consideration, (13), has been chosen to be final output, Y_t. Social welfare effects have been plotted under welfare weighting criterion A.

Table 5. Welfare Gains of Optimal DPTs

	Savers	Borrowers	Social
A. Welfare Criterion A (i.e., $\zeta_\kappa = 1 - \beta_\kappa$)			
$[\rho_x^* = 55.51]$	0.1725	0.3471	0.0183
B. Welfare Criterion B (i.e., $\lambda_p = \lambda_i$)			
$[\rho_x^* = 82.96]$	0.2597	0.2597	0.2597

Notes: Second-order approximation to the welfare gains associated with the optimal dividend prudential target and the corresponding optimized policy parameter for each of the two proposed welfare criteria. Welfare gains are expressed in percentage permanent consumption.

modestly more restricted credit provision. Since the latter effect only comes into play under very highly responsive countercyclical DPTs and savers do not own banks in the baseline calibration, the welfare tradeoff faced by patient households is more favorable than that experienced by borrowers.

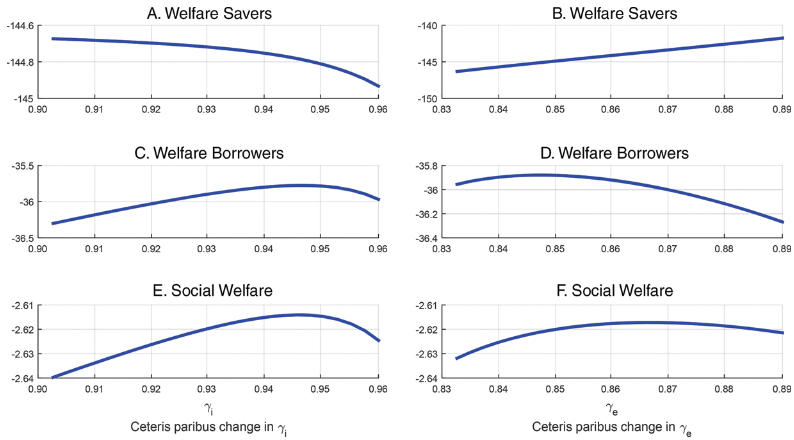
Based on the information provided by these welfare tradeoffs, I numerically solve problem (35) for the two proposed welfare criteria by searching over the following grid of parameter values: $\rho_x \{0 - 200\}$.[39] Table 5 reports the corresponding optimized parameter values and the welfare gains.

6.2.1 Interactions with Capital Requirements and Welfare Effects

Angeloni and Faia (2013) analyze optimized monetary policy rules under alternative Basel regimes. Inspired by their approach, this section examines the interactions between dividend prudential targets and existing capital regulation as well as the corresponding welfare tradeoffs and effects.

[39]In each case, the model is solved by using second-order perturbation techniques in Dynare (Adjemian et al. 2011). Unconditional lifetime utility is computed as the theoretical mean based on first-order terms of the second-order approximation to the nonlinear model, resulting in a second-order accurate welfare measure (see, e.g., Kim et al. 2008). This approach ensures that the effects of aggregate uncertainty are taken into account.

Figure 5. Welfare Effects of Capital Ratios (welfare effects of ceteris paribus changes in γ_i and γ_e)



Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the capital adequacy parameters γ_i and γ_e. Note that the static capital requirement on NFC loans is $(1-\gamma_e)$, whereas the static capital requirement on HH loans is $(1-\gamma_i)$.

Microprudential Capital Regulation. In the extended model adjustments in static capital requirements , $(1-\gamma_e)$ and $(1-\gamma_i)$, affect welfare through three main transmission channels. First, a ceteris paribus hike in a sectoral capital requirement (e.g., a reduction in γ_e) leads to a higher volatility in the corresponding type of lending (i.e., B_e) due to a "balance sheet effect" induced by banks' preference for dividend smoothing. Note that a higher capital ratio translates into a larger fraction of bank loans being financed by bank equity, a source of funding that accumulates out of "volatile" retained earnings. In particular, a ceteris paribus decrease in γ_e has a negative impact on savers' welfare through more restricted and volatile lending on entrepreneurial firms (see figure 5B). The same applies to the effect of ceteris paribus changes of γ_i on borrowers' welfare (see figure 5C).

Second, a decrease in the ratio of sectoral capital requirements, $(1-\gamma_e)\,/\,(1-\gamma_i)$ (which may be induced by a reduction in γ_i and/or by an increase in γ_e), triggers a "loan portfolio readjustment effect" by which the weight of household loans decreases in the bank's balance sheet in favor of entrepreneurial firm loans. That has a positive impact on savers' welfare (see figures 5A and 5B) and a negative effect on borrowers' welfare (see figures 5C and 5D).[40]

Note, however, that figures 5C and 5D display welfare tradeoffs. This is the case because the previously mentioned effect conflicts with a third effect; higher capital ratios require bankers to retain more earnings, thereby inducing a positive "profit generation capacity effect" by which bank owners (i.e., borrowers) benefit from higher long-run dividend payouts.

As can be shown in figures 5E and 5F, the predominance of the effects leading to more restricted and volatile lending implies that, when keeping all other parameters at their baseline values, optimal sectoral capital adequacy parameters, $\gamma_i^* = 0.9455$ and $\gamma_e^* = 0.8658$, are—under welfare criterion A—associated with capital requirements which are somewhat lower than those calibrated for the baseline scenario and based on the Basel III Accord.

Figure 6 informs about how the welfare effects of ceteris paribus changes in ρ_x vary when capital requirements change due to an equiproportional variation in sectoral capital requirements (i.e., a change in overall static capital requirements with respect to its baseline value, $(1-\gamma) = 0.105$, for which the proportion implied by expression (34) is preserved).[41] Higher capital requirements lead to lower levels of savers' welfare through their negative effect on entrepreneurial firm lending, but they do not significantly modify the effectiveness of countercyclical DPTs (proxied by the welfare tradeoff they induce).

By way of contrast, a more stringent capital scenario has a positive impact on the effectiveness of dynamic DPTs in improving

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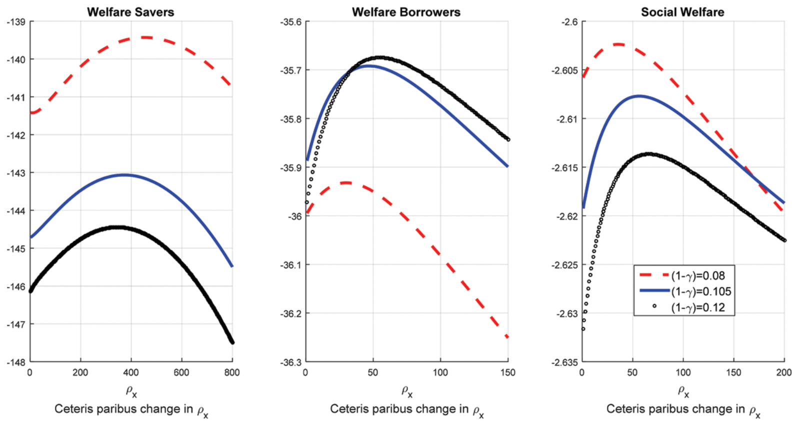
[40]The underlying reason for this readjustment in the banker's loan portfolio, and the corresponding asymmetric effect (on savers and borrowers) is that household loans become relatively more restricted and volatile than entrepreneurial firm loans (recall the "balance sheet effect").

[41]The three considered capital scenarios (including the baseline) are inspired by the Basel III Accord. 0.08 refers to the minimum capital requirement. Adding the conservation buffer (0.025) to it yields a capital ratio of 0.105. As of November 2018, all euro-area G-SIBS (global systemically important banks) were subject to a surcharge lying between 0.01 and 0.02. For that reason, the paper considers a third scenario with a capital adequacy ratio of 0.12.

Figure 6. Welfare Effects of DPTs under Alternative Capital Scenarios



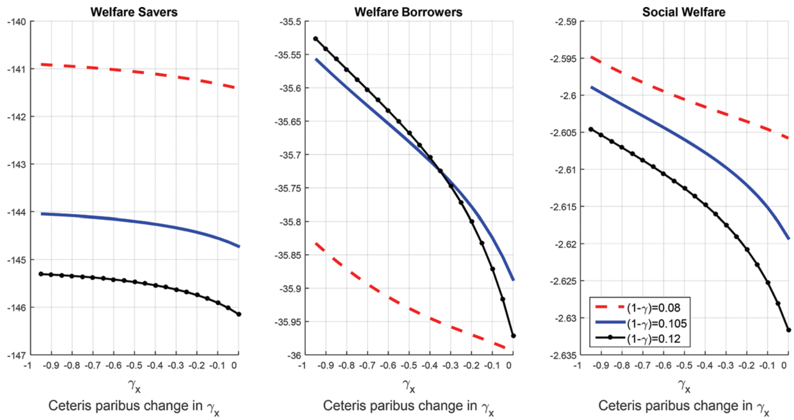
Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of the dividend prudential target, ρ_x, for three alternative capital scenarios $(1-\gamma)$.

bank owners' welfare. A larger fraction of loans being financed by "volatile" cumulative retained earnings quantitatively magnifies the problem of higher credit supply volatility triggered by dividend smoothing, and higher expected dividends improve the potential of countercyclical DPTs to mitigate the negative effects of such problem. In particular, the higher capital requirements are, the larger potentially attainable borrowers' welfare gains (through increases in ρ_x) are and the wider the range of welfare-increasing ρ_x values is.

As shown in figure 7, a similar reasoning can be followed for the case of the CCyB under alternative capital scenarios. Higher capital requirements translate into lower savers' welfare levels, while they do not materially affect the effectiveness of the CCyB. In contrast, the tighter capital requirements are, the more effective a responsive CCyB is in improving borrowers' welfare level (note that the rate at which borrowers' welfare increases with the responsiveness of the CCyB tends to increase with static capital requirements).

Table 6 reports the welfare gains from a 1 percentage point hike in static capital requirements (from 10.5 percent to 11.5 percent),

Figure 7. Welfare Effects of the CCyB for Alternative Capital Scenarios



Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of the dynamic capital requirements, γ_x, for three alternative scenarios $(1-\gamma)$.

$(1-\gamma)$, with and without introducing an optimal DPT (under the two proposed welfare criteria) in the alternative scenario (i.e., in the scenario under which $\gamma = 0.885$), with respect to the baseline scenario ($\gamma = 0.895$). Due to the above-described reasons, (i) a hike in capital requirements has a relatively more severe impact on savers' welfare than on the expected lifetime utility of borrowers and, accompanying such hike in the capital ratio with an optimal DPT has a relatively more significant positive effect on borrowers' welfare (than on the expected lifetime utility of savers).

Table 7 describes the welfare effects of introducing an optimal DPT under the three capital scenarios already considered in figure 6. As already mentioned, hikes in capital requirements exacerbate the "problem" of dividend smoothing and enhance the potential of DPTs to tackle such issue. For each of the two proposed welfare criteria, the higher capital requirements are, the more reactive optimal DPTs become and the higher individual and social welfare gains attained are.

Table 6. Welfare Gains to a 1 Percentage Point Hike in Capital Requirements

1.0 pp Increase in $(1-\gamma)$	$\boldsymbol{\rho_x}$	**Savers**	**Borrowers**	**Social**
A. Without Countercyclical DPT				
	$[\rho_x = 0]$	−0.3938	−0.1463	−0.0096
B. With Countercyclical DPT				
Welfare Criterion A	$[\rho_x^* = 61.49]$	−0.1746	0.3747	0.0177
Welfare Criterion B	$[\rho_x^* = 96.42]$	−0.0650	0.2580	(0.0965)

Notes: Second-order approximation to the welfare gains associated with a 1 percentage point hike (from 0.105 to 0.115) in the bank capital ratio (induced by an equiproportional increase in sectoral capital requirements)—with and without introducing an optimal DPT in the alternative scenario (i.e., in the scenario in which the capital ratio is set to a value of 0.115)—with respect to the baseline scenario (i.e., the scenario in which the capital ratio is set to a value of 0.105) and for the two proposed welfare criteria. Social welfare gains under welfare criterion B in panel B of the table have been proxied by the arithmetic means of savers' and borrowers' welfare gains. Welfare gains are expressed in percentage permanent consumption.

Table 7. Welfare Gains of Optimal DPTs under Alternative Capital Scenarios

Capital Scenario $(1-\gamma)$	$\boldsymbol{\rho_x}$	**Savers**	**Borrowers**	**Social**
A. $(1-\gamma) = 0.08$				
Welfare Criterion A	$[\rho_x^* = 34.54]$	0.0367	0.1006	0.0052
Welfare Criterion B	$[\rho_x^* = 47.61]$	0.0701	0.0701	0.0701
B. $(1-\gamma) = 0.105$ *(Baseline)*				
Welfare Criterion A	$[\rho_x^* = 55.51]$	0.1725	0.3471	0.0183
Welfare Criterion B	$[\rho_x^* = 82.96]$	0.2597	0.2597	0.2597
C. $(1-\gamma) = 0.12$				
Welfare Criterion A	$[\rho_x^* = 64.89]$	0.2779	0.5377	0.0285
Welfare Criterion B	$[\rho_x^* = 105.03]$	0.4030	0.4030	0.4030

Notes: Second-order approximation to the welfare gains associated with the optimal dividend prudential target under three different Basel III-based capital scenarios, and the corresponding optimized policy parameter of the DPT for each of the two proposed welfare criteria. The ratio of sectoral capital requirements imposed by equation (34) remains unchanged across the three different capital scenarios. Welfare gains are expressed in percentage permanent consumption.

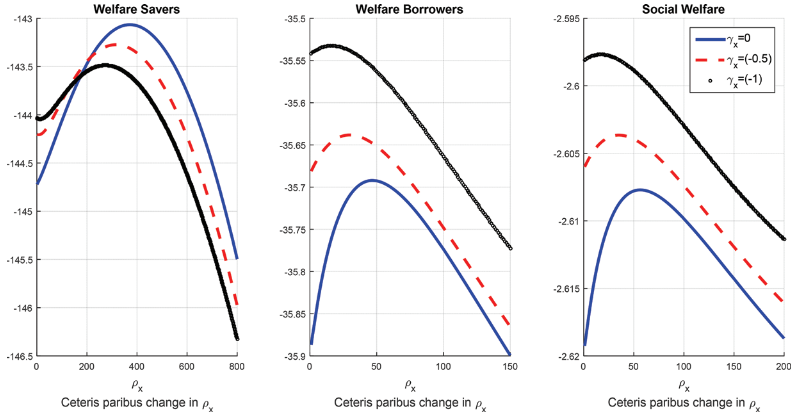
In a nutshell, higher capital requirements translate into a higher fraction of bank loans being financed by "volatile" bank capital and higher long-run profits (and dividends). The former exacerbates the problem of induced credit supply volatility, whereas the latter reinforces the effectiveness of DPTs in tackling such issue.

Macroprudential Capital Regulation. How do the DPT and the CCyB interact in this model? Is the DPT a complement or substitute for the CCyB? In order to answer these questions, I carry out several exercises. Figures 8, 9, and 10 display the welfare effects of ceteris paribus changes in ρ_x for alternative values of γ_x; the welfare effects of ceteris paribus changes in γ_x for alternative values of ρ_x; and the welfare effects of ceteris paribus changes in ρ_x and γ_x (i.e., interactions between the DPT and the CCyB), respectively.[42] There are two findings that stand out. First, if there were no boundaries to the values that ρ_x and γ_x could take, households who do not own banks (i.e., savers) would prefer to rely on a highly responsive DPT and to have no CCyB in place (since the former is more effective in smoothing lending than the latter), whereas bank owners (i.e., borrowers) would be better off with a highly responsive CCyB and no DPT (as the former allows them to benefit from credit smoothing without having to incur the cost of bank dividend volatility induced by the latter). Second, under the considered grid of parameter values, $\rho_x \{0 - 200\}$ and $\gamma_x \{(-1) - 0\}$—which are associated with what I shall refer to as "potentially implementable policy rules"—each type of household finds optimal to simultaneously have a countercyclical DPT and a CCyB in place.

The first finding suggests that, in this case, the optimal macroprudential policy mix is going to be particularly sensitive to the selected welfare weighting criterion. Figure 11 informs about the

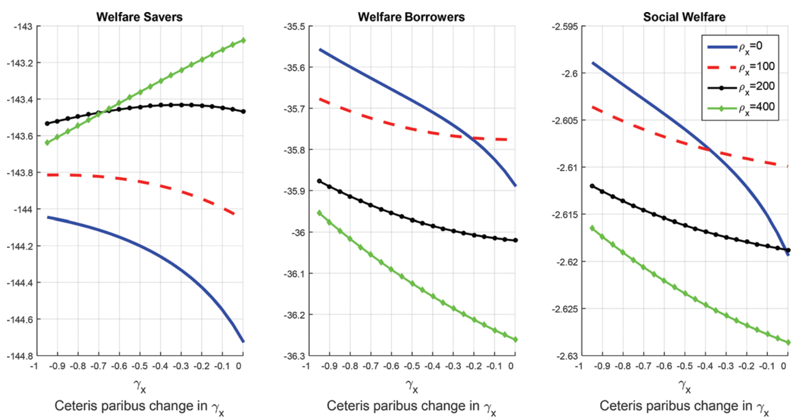
[42] I have set the grid of values that γ_x can take to $\gamma_x \{(-1) - 0\}$. This range of values is based on information related to the cap that various economies, including the European Union, have set on the CCyB in practice (see, e.g., Basel Committee on Banking Supervision 2017) as well as on a wide range of output gap estimates for the euro area based on quarterly data of real GDP for the period 2002:Q1–2018:Q2. In addition, table 8 shows that the CCyB for which the asymptotic variance of the credit gap and the loans-to-output gap are minimized relates to a value of γ_x that is in the vicinity of (–1). Note that in the limit case in which $\gamma_x = -1$, the CCyB is very highly responsive in the sense that a 1 percentage point increase in the output gap translates into a 1 percentage point increase in dynamic capital requirements.

Figure 8. Welfare Effects of DPTs for Alternative CCyBs



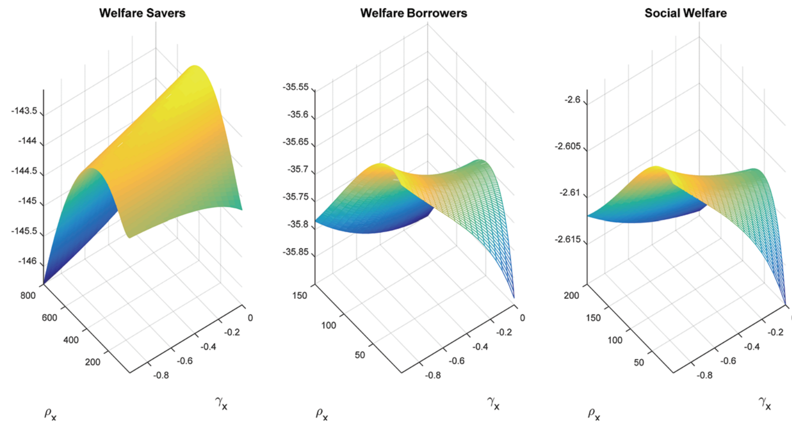
Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of the dividend prudential target, ρ_x, for alternative values of the cyclical parameter of dynamic capital requirements, γ_x.

Figure 9. Welfare Effects of the CCyB for Alternative DPTs



Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of dynamic capital requirements, γ_x, for alternative values of the cyclical parameter of the dividend prudential target, ρ_x.

Figure 10. Interactions between the DPT and the CCyB (welfare effects of ceteris paribus changes in $\rho_x - \gamma_x$)



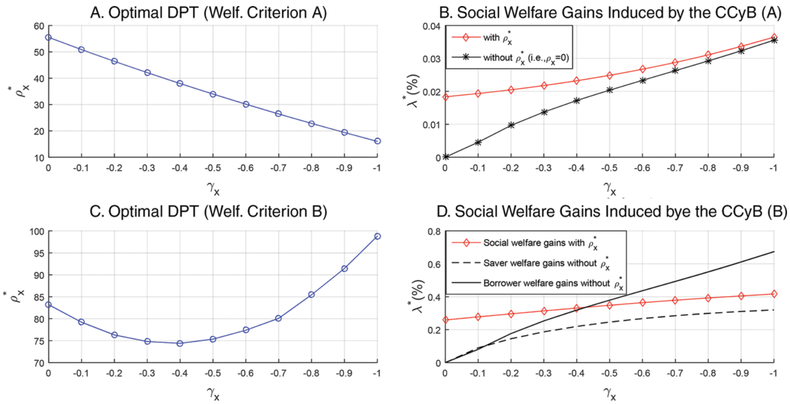
Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameters of dynamic capital requirements and the dividend prudential target, γ_x and ρ_x.

maximum contribution the DPT can make to social welfare when activating the CCyB, for different values of $\gamma_x \in [-1, 0]$. In particular, figures 11B and 11D plot the welfare gains of the CCyB, for the grid $\gamma_x \, \{(-1) - 0\}$, with and without introducing an optimal DPT (in the alternative policy scenario in which $\gamma_x < 0$), for the two proposed welfare criteria.[43] Figures 11A and 11B display the corresponding optimized ρ_x values for different values of $\gamma_x \in [-1, 0]$. Under criterion A, the more responsive the CCyB is, the larger welfare gains are and the less responsive the optimal DPT is. This result largely reflects the preferences of borrowers.

Under criterion B, there is an important subgrid of potentially implementable γ_x values (i.e., $\gamma_x \, \{(-1) - 0.4\}$) for which a more

[43]Social welfare gains of the CCyB without an optimal DPT cannot be computed under welfare criterion B since, in this case, problem (35) has no solution. In particular, there is no value of $\gamma_x \in [-1, 0]$ that satisfies $\lambda_p = \lambda_i$, as the rate at which borrowers' welfare increases with γ_x is higher than the one at which savers' welfare does, $\forall \; \gamma_x \in [-1, 0]$ (see figure 9). As an alternative, I plot the welfare gains of savers and borrowers induced by the CCyB when $\rho_x = 0$.

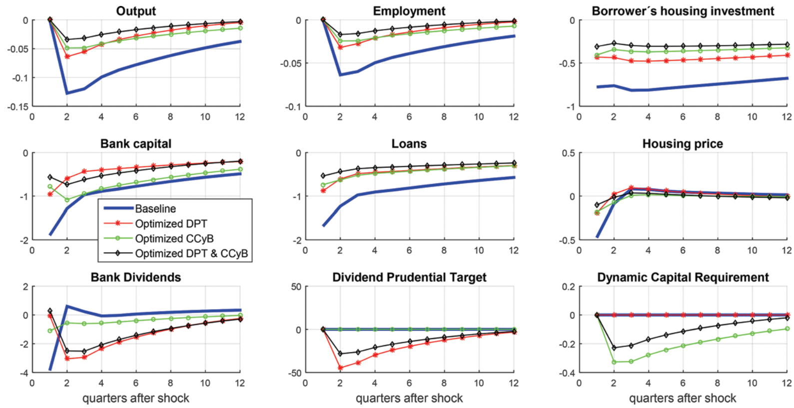
Figure 11. Optimal DPT and Welfare Gains for Alternative Calibrations of the CCyB



Notes: Panels B and D report the second-order approximation to the unconditional social welfare gains induced by the CCyB, for different values of $\gamma_x \in [-1, 0]$—with and without introducing an optimal DPT (diamond line and starred line, respectively) in the alternative policy scenario (i.e., the scenario in which $\gamma_x < 0$)—with respect to the baseline scenario (i.e., $\rho_x = 0, \gamma_x = 0$), and for the two proposed welfare criteria. Under welfare criterion B and for all values of $\gamma_x \in [-1, 0]$, there is no solution to problem (35) for the case in which an optimal DPT is not introduced. Instead, in that particular case, panel D displays the welfare gains of savers and borrowers (dashed line and solid line, respectively). Panels A and B represent—for the same values of $\gamma_x \in [-1, 0]$, the two proposed welfare criteria, and the case in which the CCyB is complemented with an optimal DPT—the corresponding optimized values of the cyclical parameter of the optimal DPT, ρ_x^*. For reporting purposes, x-axes have been reversed in all panels of the figure.

reactive CCyB calls for a more responsive DPT, a relationship aligned with the preferences of savers for the considered grid of ρ_x values (see figure 8). Even if this relationship is not the one advocated by borrowers, criterion B (i) exploits the fact that there is a wide range of $\{\rho_x > 0, \gamma_x < 0\}$ combinations for which savers and borrowers are better off than under the baseline scenario (see figure 10), and (ii) implicitly strikes a balance between this conflict and the fact that borrowers' welfare increases in the responsiveness of the CCyB at a higher rate than that of savers, $\forall\ \rho_x \in [0, 200]$ (see figure 9).

Figure 12. Impulse Responses to Negative HH Collateral Shock (extended model, macroprudential policy scenarios)



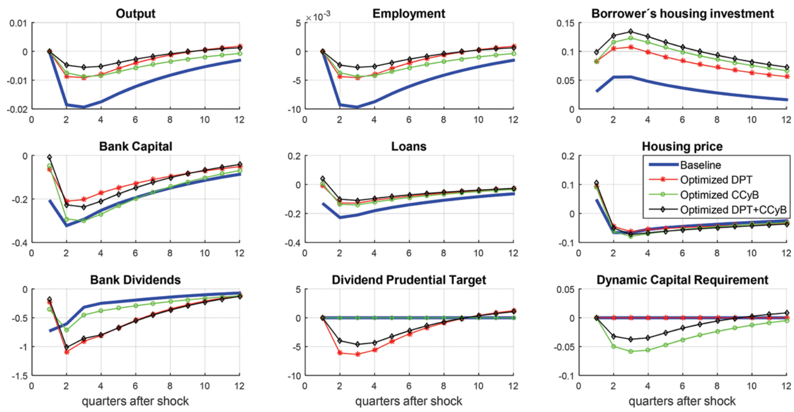
Notes: Variables are expressed in percentage deviations from the steady state. The solid line refers to the baseline scenario. The starred line corresponds to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of the dividend prudential target, ρ_x. The dotted line relates to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of dynamic capital requirements, γ_x. The diamond line makes reference to an alternative (policy) scenario in which welfare has been maximized with respect to cyclical policy parameters ρ_x and γ_x.

An important corollary of the above discussed findings is that even if the CCyB is very responsive (i.e., $\gamma_x \approx -1$) and regardless of the selected welfare criterion, it is optimal to complement such macroprudential policy with a countercyclical DPT (i.e., $\rho_x > 0$).

Figures 12–16 plot the impulse responses of key economic aggregates to the five different shocks that hit this economy. The solid line refers to the responses under the baseline scenario, while the diamond, starred, and dotted lines correspond to alternative policy scenarios in which problem (35) has been solved—under welfare criterion B—with respect to $\{\rho_x, \gamma_x\}$, ρ_x, and γ_x, respectively.[44] In the face of financial shocks, jointly optimizing with respect to $\{\rho_x, \gamma_x\}$ is more effective in smoothing financial and economic fluctuations than

[44]The policy parameter values for which the problem of social welfare solves under criterion B are, for each of the three considered macroprudential policy scenarios, $\{\rho_x^* = 98.8 \text{ and } \gamma_x^* = -1\}$, $\rho_x^* = 83.18$, and $\gamma_x^* = -1$.

Figure 13. Impulse Responses to a Negative NFC Collateral Shock (extended model, macroprudential policy scenarios)



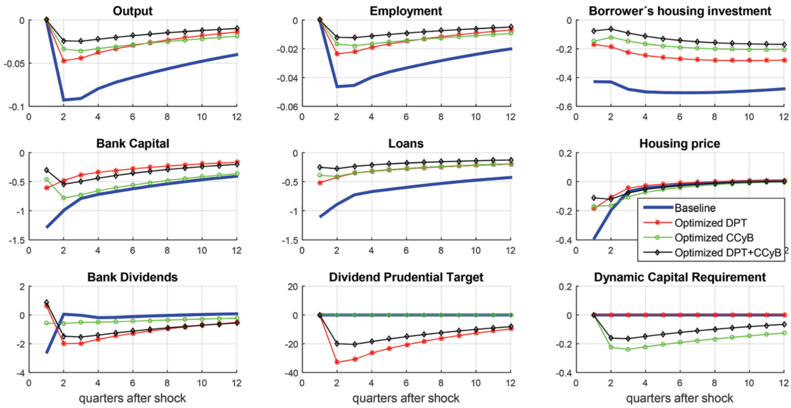
Notes: Variables are expressed in percentage deviations from the steady state. The solid line refers to the baseline scenario. The starred line corresponds to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of the dividend prudential target, ρ_x. The dotted line relates to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of dynamic capital requirements, γ_x. The diamond line makes reference to an alternative (policy) scenario in which welfare has been maximized with respect to cyclical policy parameters ρ_x and γ_x.

doing it with respect to any of the two macroprudential policy parameters separately (figures 12–14). Under nonfinancial shocks, the optimal DPT performs better than the optimal CCyB (figures 15 and 16).[45]

In order to further explore the effectiveness of the three alternative macroprudential policy scenarios in taming the credit cycle, table 8 reports the main results of solving problem (17) in the extended model, for the three considered policy parameter vectors,

[45]The finding suggesting that the CCyB is relatively more effective in taming the cycle when financial shocks hit the economy than in the presence of other types of shocks (e.g., technology shocks) has been presented in several recent studies (see, e.g., Angelini, Neri, and Panetta 2014). Thus, the comparative effectiveness of optimal DPTs in smoothing financial and economic fluctuations in the face of nonfinancial shocks should be regarded as an additional strength of this instrument as a complement to existing capital regulation.

Figure 14. Impulse Responses to a Negative Bank Capital Shock (extended model, macroprudential policy scenarios)



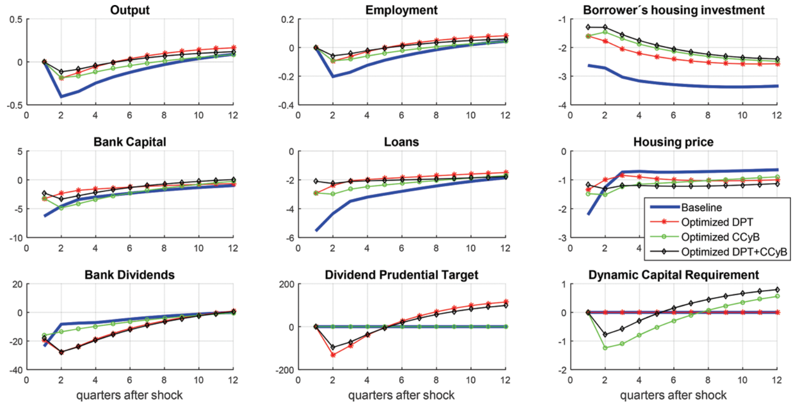
Notes: Variables are expressed in percentage deviations from the steady state. The solid line refers to the baseline scenario. The starred line corresponds to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of the dividend prudential target, ρ_x. The dotted line relates to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of dynamic capital requirements, γ_x. The diamond line makes reference to an alternative (policy) scenario in which welfare has been maximized with respect to cyclical policy parameters ρ_x and γ_x.

$\Theta \equiv \{\rho_x; \gamma_x; (\rho_x, \gamma_x)\}$, and for $z \equiv \{B; B/Y\}$. The optimized DPT is substantially more effective than the optimized CCyB in smoothing bank lending and approximately as effective as jointly optimizing with respect to $\{\rho_x, \gamma_x\}$.

In conclusion, even though both instruments are effective in taming the credit cycle, the DPT complements the CCyB in at least two dimensions. First, an optimized DPT reinforces the effectiveness of the CCyB in mitigating financial and economic fluctuations regardless of the nature of the shock and performs particularly better than the CCyB under nonfinancial shocks. Overall, an optimized DPT is more effective in smoothing lending and output than the optimized CCyB.[46] Second, due to this fact, households who do not own banks have a stronger preference for having a countercyclical

[46] As in the basic model, this is the case because DPTs directly attack the root of the "problem" (i.e., dividend smoothing).

Figure 15. Impulse Responses to a Negative Housing Preference Shock (extended model, macroprudential policy scenarios)



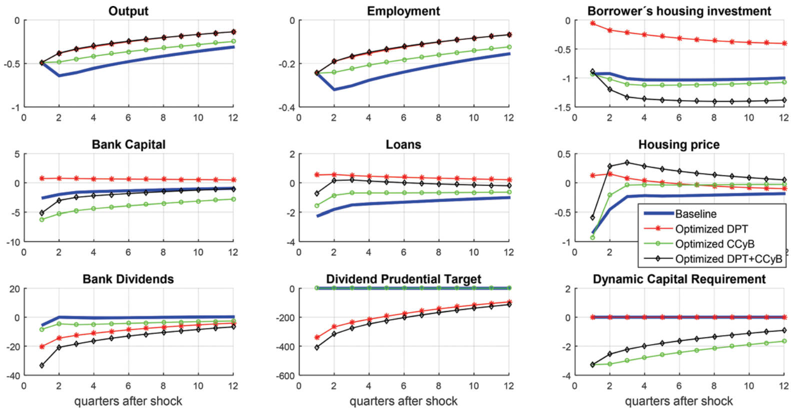
Notes: Variables are expressed in percentage deviations from the steady state. The solid line refers to the baseline scenario. The starred line corresponds to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of the dividend prudential target, ρ_x. The dotted line relates to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of dynamic capital requirements, γ_x. The diamond line makes reference to an alternative (policy) scenario in which welfare has been maximized with respect to cyclical policy parameters ρ_x and γ_x.

DPT in place, whereas those who do own banks prefer to rely on a CCyB, as the latter avoids the cost induced by DPTs in terms of higher bank dividend volatility. The bottom line is that, for a wide range of "potentially implementable policy rules," there is a variety of standard optimization criteria, suggesting that jointly calibrating both policy instruments is optimal.

6.3 Robustness Checks

In this section I first investigate the robustness of the welfare effects of the DPT to changes in key parameters. Since the main cost associated with optimized DPTs directly affects bank owners through higher bank dividend volatility, it could be the case that changes in the distribution of banks' ownership between savers and borrowers were to significantly affect the welfare tradeoff faced by each agent

Figure 16. Impulse Responses to a Negative Technology Shock (extended model, macroprudential policy scenarios)



Notes: Variables are expressed in percentage deviations from the steady state. The solid line refers to the baseline scenario. The starred line corresponds to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of the dividend prudential target, ρ_x. The dotted line relates to an alternative (policy) scenario in which welfare has been maximized with respect to the cyclical parameter of dynamic capital requirements, γ_x. The diamond line makes reference to an alternative (policy) scenario in which welfare has been maximized with respect to cyclical policy parameters ρ_x and γ_x.

class. However, figure 17 suggests that changes in the fraction of banks owned by savers, ω_b, (and, thus, in that of banks owned by borrowers) does not materially affect the shape of expected lifetime utility (as a function of ρ_x).[47]

In addition, there are two policy parameters the public authority may consider to modify, and whose values are relevant from a redistributive perspective: the penalty parameter, κ, and the fraction of net transfer that savers receive according to their bank property, χ. Due to the insurance role it plays, as parameter χ increases (and regardless of the ρ_x value), the welfare level (and tradeoff) attained by the representative saver improves, while that of the representative borrower deteriorates (see figure 18). With regards to

[47]Not surprisingly, increases in ω_b do affect the welfare level of savers (which goes up) and borrowers (which declines). This is so because the bank dividend payout received by a given household increases with the fraction of banks it owns.

Table 8. Optimized Rules and Macroprudential Losses (extended model)

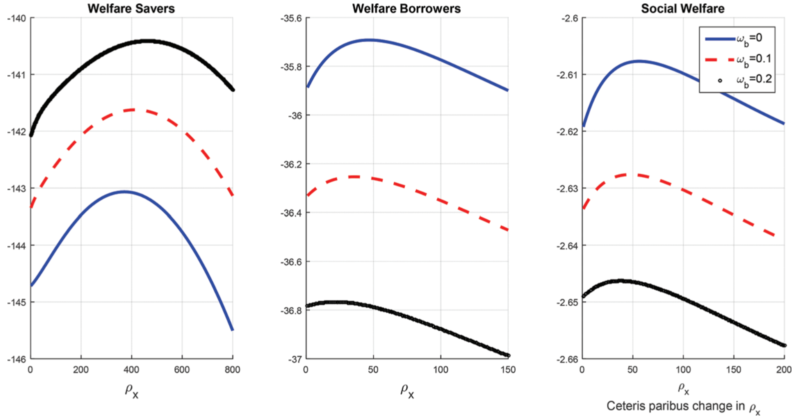
	$\boldsymbol{\sigma_B^{2(1)}}$	$\boldsymbol{\sigma_{B/Y}^2}$
A. $\{\rho_x\}$		
Loss Variation[2]	(−59.15)	(−48.82)
ρ_x	75.079	66.496
B. $\{\gamma_x\}$		
Loss Variation	(−42.92)	(−38.50)
γ_x	−1.047	−1.013
C. $\{\rho_x, \gamma_x\}$		
Loss Variation	(−59.16)	(−49.08)
ρ_x	74.793	63.805
γ_x	−0.023	−0.152

Notes: (1) Asymptotic variance that enters the objective function of the prudential authority in problem (17). Such problem has been solved numerically by means of the osr (i.e., optimal simple rule) command in Dynare. (2) Percentage changes in the value of the loss function under the corresponding policy scenario with respect to the baseline scenario.

κ, given a sensible range of values for the penalty parameter, the shape of the welfare as a function of ρ_x is not significantly affected, although as the value of κ increases, welfare tradeoffs become more pronounced and the optimal DPT becomes less responsive. As shown in figure 19, this is so because a more stringent sanctions regime makes the policy more effective (in smoothing lending through less volatile retained earnings) and more costly at the same time (e.g., higher bank dividend volatility).

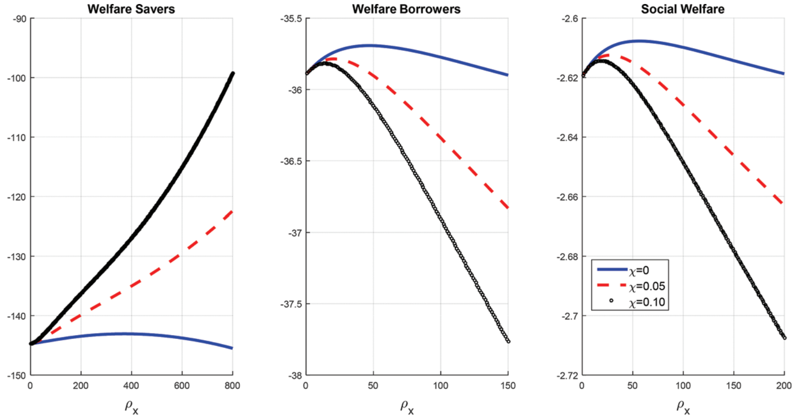
As mentioned in subsections 4.1 and 6.1, expression (9) not only permits to account for several empirical regularities (through the calibration of parameters δ and σ_{kb}) but also plays an essential role in allowing the model to reproduce the key mechanism (that triggers the main welfare tradeoff countercyclical DPTs exhibit due to individual preferences for dividend smoothing and lending smoothing), by connecting the profit generation capacity of the representative bank (which is essential to distribute high and stable dividends

Figure 17. Robustness Checks: ω_b (welfare effects of ceteris paribus changes in ρ_x)



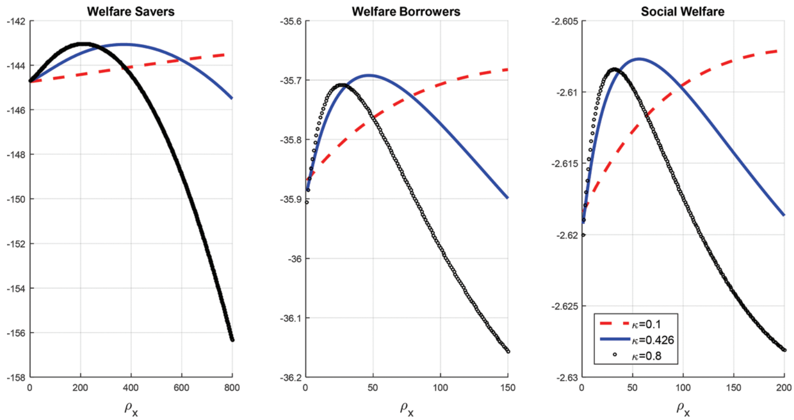
Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of the dividend prudential target, ρ_x, for alternative fractions of banks owned by savers. The solid line refers to the baseline scenario, whereas the dotted and dashed lines relate to alternative parameterization scenarios.

Figure 18. Robustness Checks: χ (welfare effects of ceteris paribus changes in ρ_x)



Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of the dividend prudential target, ρ_x, for alternative fractions χ of the net transfer that savers receive according to their bank property. The solid line refers to the baseline scenario, whereas the dotted and dashed lines relate to alternative parameterization scenarios.

Figure 19. Robustness Checks: κ (welfare effects of ceteris paribus changes in ρ_x)



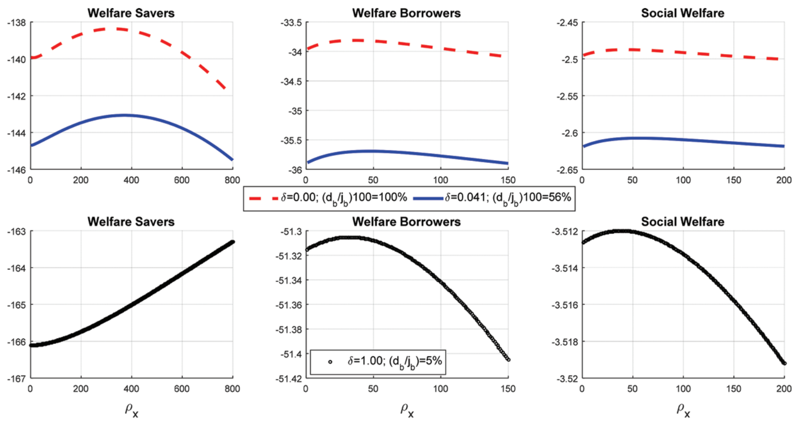
Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of the dividend prudential target, ρ_x, for alternative values of the dividend adjustment cost parameter, κ. The solid line refers to the baseline scenario, whereas the dotted and dashed lines relate to alternative parameterization scenarios.

over the cycle) with its capital generation capacity (which is crucial to meet capital requirements). Figure 20 shows that, regardless of the value taken by $\delta \in [0, 1]$, the same welfare tradeoff applies. Of course, as δ increases, the bank equity accumulated per unit of profits declines, which obliges the representative banker to reduce the size of its balance sheet by cutting on lending in order to meet capital requirements. Consequently, the welfare level of savers and borrowers declines and the welfare tradeoff faced by the latter deteriorates.[48]

Lastly, figure 21 confirms that regardless of the selected optimization criterion (from those considered in the quantitative analysis), the optimized DPT is more effective in smoothing credit supply and

[48]Interestingly, figure 20 also makes clear that the considered range of values for $\delta \in [0, 1]$ permits to match the steady-state payout ratio of the banking industry virtually regardless of the value that such ratio shall take. Note that, from expression (9), it follows that if $\delta = 0$, in the steady state bank profits are fully distributed. As in subsection 6.2, for the alternative parameterization scenarios considered in figure 20, I have assumed that $\rho_d = d_b^{ss}$.

Figure 20. Robustness Checks: δ (welfare effects of ceteris paribus changes in ρ_x)



Notes: Second-order approximation to the unconditional welfare of savers and borrowers as well as to the unconditional social welfare (under welfare criterion A) as a function of the cyclical parameter of the dividend prudential target, ρ_x, for alternative values of the depreciation rate of bank capital, δ. The solid line refers to the baseline scenario, whereas the dotted and dashed lines relate to alternative parameterization scenarios. For each scenario, the associated steady-state payout ratio is reported.

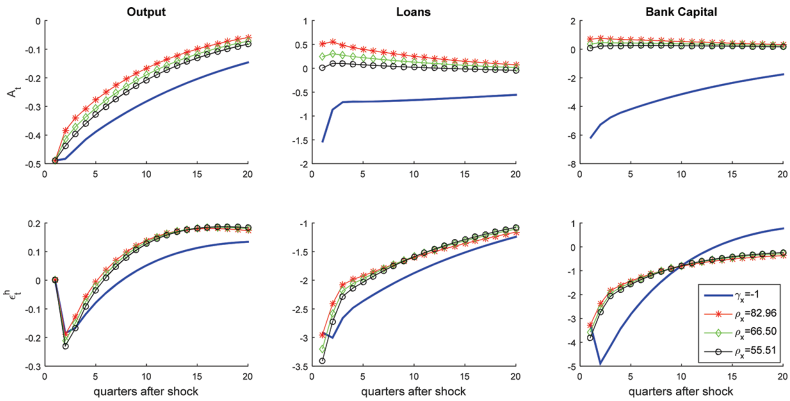
real output under technology and housing preference shocks than the optimized, highly responsive, CCyB.

In a nutshell, although quantitative differences may arise, the main conclusions of this exercise are robust across calibrated values of key parameters, across alternative specifications of policy scenarios, and across alternative optimization criteria. Countercyclical dividend prudential targets are very effective in smoothing financial and business cycles, and they complement and induce welfare gains associated with a Basel III type of capital regulation through various mechanisms.

7. Conclusion

Available evidence on dividends and earnings in the euro-area banking sector points to the existence of a link between payout policies and the adjustment mechanisms through which bankers opt to meet

Figure 21. Robustness Checks: Nonfinancial Shocks and the Effectiveness of Optimized DPTs



Notes: Variables are expressed in percentage deviations from the steady state. Nonfinancial shocks refer to technology and housing preference shocks. The solid line refers to a policy scenario in which welfare has been maximized with respect to the cyclical parameter of dynamic capital requirements, γ_x (which roughly coincides with the value of γ_x for which the prudential authority minimizes the asymptotic variance of the loans-to-output ratio under full commitment). The starred, dotted, and diamond lines correspond to policy scenarios in which welfare has been maximized with respect to ρ_x under the two proposed welfare criteria and the asymptotic variance of the loans-to-output ratio has been minimized under full commitment (i.e., problem (17)), respectively.

their target regulatory capital ratios. When shocks hit their profits, bank managers adjust retained earnings to smooth dividends. This generates bank equity and credit supply volatility.

I develop a quantitative DSGE model with a banking sector that incorporates this mechanism to examine the transmission and effects of a novel macroprudential policy rule—that I shall call dividend prudential target (DPT)—aimed at complementing existing capital regulation by tackling this issue. Even though welfare-maximizing DPTs are more effective in smoothing the financial and the business cycle than the CCyB, this instrument actually complements a Basel III type of framework by mitigating the negative effects of capital ratio adjustments in terms of more restricted and volatile lending and output, they operate through a transmission mechanism that is different but complementary to that of the CCyB, and they have a comparative advantage in smoothing the cycle under nonfinancial shocks when compared with the latter.

The simplicity of the model is instrumental to clearly identify the transmission mechanism through which the proposed policy rule operates. Yet, it comes at the cost of omitting ingredients which are present in reality and that could possibly change some of the results. On the one hand, assuming a positive probability of bank failure, as in Clerc et al. (2015), should further reinforce the argument in favor of this complement to existing capital regulation. In addition, a heterogeneous-agents model that accounts for the specific fraction of households who hold bank shares in practice and for the concrete weight of such shares in their asset portfolios would deliver a lower and much more realistic estimate of the costs induced by DPTs in terms of higher bank dividend volatility. On the other hand, incorporating outside equity in an environment in which bank owners can substitute their bank shares for alternative assets at a relatively low cost may make the policy proposal less attractive. In addition, the literature has shown that the approach to modeling bank risk-taking and systemic risk can notably affect macroprudential policy prescriptions (see, e.g., Martinez-Miera and Suarez 2014).

Lastly, optimal coordination between this type of prudential regulation and other macroeconomic policies should be considered as well (e.g., monetary policy). Based on the ECB annual report of 2016, one of the comments the European Parliament (2017) has recently made to the ECB relates to this issue: "The European Parliament is concerned that euro area banks did not use the advantageous environment created by the ECB to strengthen their capital bases but rather, according to the Bank for International Settlements, to pay substantial dividends sometimes exceeding the level of retained earnings."

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Unconventional Monetary Policy Shocks in the Euro Area and the Sovereign-Bank Nexus*

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We explore the effects of the ECB's unconventional monetary policy on the banks' sovereign debt portfolios. In particular, using panel vector autoregressive (VAR) models we analyze whether banks increased their domestic government bond holdings in response to nonstandard monetary policy shocks, thereby possibly promoting the sovereign-bank nexus, i.e., the exposure of banks to the debt issued by the national government. Our results suggest that euro-area crisis countries' banks enlarged their exposure to domestic sovereign debt after innovations related to unconventional monetary policy. Moreover, the restructuring of sovereign debt portfolios was characterized by a home bias.

JEL Codes: C32, E30, E52, G21, H63.

1. Introduction

The European Central Bank (ECB) responded to the global financial crisis by conducting a number of unconventional monetary policy measures in addition to lowering the policy rate. The aim of these measures was to reduce potential risks for price stability in the

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euro area by counteracting distortions on the interbank market and reducing impairments in the monetary policy transmission induced by financial market fragmentation across member states during the sovereign debt crisis (ECB 2011, 2013).[1]

In this paper, we explore the effects of the ECB's unconventional monetary policy on the balance sheet exposure of a country's banking sector to the debt issued by the national government, i.e., the so-called *sovereign-bank nexus*. The nexus is considered of primary importance for the medium-term stability of the financial system (Battistini, Pagano, and Simonelli 2014; Brunnermeier et al. 2016; Basel Committee on Banking Supervision 2017; Farhi and Tirole 2018; International Monetary Fund 2018, among others). On the one hand, a higher share of bond holdings is typically associated with a better liquidity position of banks, thus being favorable for the soundness of the banking system (Walther 2016; Richter, Schularick, and Wachtel 2017; Hoerova et al. 2018), and may even reduce the incentives for the sovereign to default (Basel Committee on Banking Supervision 2017; Gennaioli, Martin, and Rossi 2018). On the other hand, a stronger nexus may harm financial stability by making the national banking sector more sensitive to deteriorations in the sovereign's creditworthiness and may even contribute to the emergence of *diabolic loops* as repeatedly observed since the outbreak of the global financial crisis (Brunnermeier et al. 2016; Dell'Ariccia et al. 2018; Farhi and Tirole 2018). For example, many euro-area countries experienced such a loop as the market value of banks' holdings of domestic government bonds dropped due to the deterioration in the sovereign's creditworthiness, thus putting a strain on the solidity of the banking sector. Governments responded by giving safety guarantees or even implementing substantial rescue packages, which, however, might potentially increase sovereign risk further

[1]In particular, the ECB switched to regular open market operations with fixed rates and full allotment that were provided with longer maturities, relaxed collateral requirements, changed the modalities of its long-term refinancing operations, and launched outright transactions, like the Securities Market Programme (SMP) in 2010 and the extended Asset Purchase Programme (APP) in 2015, or announced these—under the Outright Monetary Transaction (OMT) program in 2012. In addition, the central bank imposed a negative interest rate on its deposit facility. See section 3 for more details.

and reinforce the impairment of banks' balance sheets.[2] Meanwhile, Battistini, Pagano, and Simonelli (2014) and Brutti and Saure (2015) consider the distortion created and intensified by the nexus as one of the core problems associated with the euro crisis.

Against this background, we estimate panel vector autoregressive (VAR) models using Bayesian methods to assess how banks across euro-area member countries adjusted their domestic sovereign debt portfolios in response to a nonstandard monetary policy shock. We focus on the period 2007–14. The ratio of banks' domestic government bond holdings increased markedly during that time.[3]

Following Canova and de Nicolo (2002), Peersman (2005), Uhlig (2005), Rubio-Ramirez, Waggoner, and Zha (2010), or Arias, Rubio-Ramirez, and Waggoner (2014), we use sign restrictions on impulse responses to identify an unconventional monetary policy shock. In particular, in our benchmark model we refer to Boeckx, Dossche, and Peersman (2017) and relate a shock to nonstandard monetary policy to an unexpected increase in the Eurosystem's total assets that is accompanied by a decrease in the spread between the euro overnight index average (EONIA) and the policy rate as well as lower financial stress. We also assess alternative identification schemes that are related to the long-term refinancing operations, changes in the composition of open market operations, and the shadow rate of monetary policy.

Our results suggest the presence of a dichotomy between the core countries of the euro area, i.e., Austria, Belgium, Germany, Finland, France, and the Netherlands, and the crisis countries Italy, Spain, and Portugal. In particular, in the years 2007–14, the banking sectors of the crisis countries significantly shifted their asset portfolios toward domestic sovereign bonds in response to expansionary unconventional monetary shocks. These shocks appear to have been quantitatively important, as they explain around 19 percent of the variation in the national banking sectors' share of domestic

[2] According to Acharya, Drechsler, and Schnabl (2014), in industrial countries there was essentially no sign of sovereign credit risk before the onset of the global financial crisis. Hence, the nexus was not considered to be problematic due to the prevailing view that sovereign credit risk was unlikely to be a concern in the near future. Rather, sovereign defaults were regarded as a problem of emerging economies.

[3] See figure A.1 in the online appendix, available at http://www.ijcb.org.

government bonds in the group of distressed euro-area economies. Moreover, banks seemed to manage their sovereign debt portfolios actively in response to sudden unconventional monetary expansions, as the rising holdings of domestic government bonds were accompanied by a reduction in the balance sheet share of bonds issued by other euro-area member states. Thus, the reaction of banks' sovereign debt portfolios in the crisis countries was characterized by a home bias. By contrast, euro-area core countries' banks seemed not to significantly increase their domestic government bond holdings relative to total assets in response to innovations related to unconventional monetary policy. Core countries' banks rather restructured their sovereign bond portfolios by lowering their holdings of rest-of-EMU government bonds. Furthermore, estimations using the approach by Jarocinski (2010) also provide support for the presence of a significant dichotomy between core and crisis economies with regard to the response of their banking sectors to unconventional monetary policy shocks. While in Italy, Spain, and Portugal the change in the domestic government bond holdings ratio occurred swiftly, in none of the core countries can a significant response of the ratio be documented. Finally, a historical decomposition suggests that the ECB's asset purchases conducted within the SMP contributed to an increase in banks' sovereign debt portfolios. In Portugal, the effect was immediately observable after May 2010, while in Italy and Spain it arose after February 2012. Furthermore, in Italy the announcement of the OMT program in September 2012 seemed also to have contributed to a higher domestic government bond holdings ratio. Overall, we conclude that the crisis countries' banks enlarged their exposure to domestic sovereign debt in response to expansionary innovations stemming from unconventional monetary policy, which possibly made them more vulnerable to sovereign distress.

Our paper is related to several recent contributions that investigate the effects of the ECB's unconventional monetary policy on the sovereign-bank nexus by means of microeconometric methods (Drechsler et al. 2016; Altavilla, Pagano, and Simonelli 2017; Crosignani, Faria-e-Castro, and Fonseca 2017; Peydro, Polo, and Sette 2017; Jasova, Mendicino, and Supera 2018, among others).[4]

[4]We discuss these contributions in greater detail in section 2 below.

These studies have the advantage of exploiting the rich information revealed by the cross-sectional and time-series dimension of the large bank-level panels. The papers usually concentrate on the contemporaneous microeconomic effects of specific policy interventions on the individual bank, while largely abstracting from dynamic macroeconomic feedback effects. We contribute to the literature by basing our empirical analysis on aggregate data, thus taking a purely macroeconomic perspective and providing direct estimates of the reaction of the aggregate banking sector in several euro-area countries to unconventional monetary policy shocks. While acknowledging that the use of aggregate data comes at the cost of losing cross-sectional information, it allows us to capture—albeit only implicitly—many macro-level interlinkages between banks' behavior and the rest of the economy.

The paper is organized as follows. Section 2 gives an overview of the related literature. In section 3, we discuss our benchmark panel VAR model setup. We outline the model framework, introduce the data, and discuss the strategy to identify an unconventional monetary policy shock. In section 4, we summarize our results. We present impulse response analysis, a decomposition of the forecast error variance, discuss alternative schemes to identify unconventional monetary policy shocks, and also discuss the results for the single euro-area countries derived from a panel of VAR models that are estimated by means of a hierarchical Bayesian panel model estimator. Section 5 provides concluding remarks.

2. Related Literature

Our work relates to a number of studies that use structural VAR models to investigate the effects of unconventional monetary policy shocks on the basis of aggregate data.[5] Most of the contributions focus on the transmission of such shocks to real activity and inflation and, in some cases, on credit market variables, e.g., different loan volumes and lending rates, or further financial market aggregates. Other papers take an explicit financial stability perspective by exploring

[5] See, for example, Baumeister and Benati (2013), Gambacorta, Hofmann, and Peersman (2014), Weale and Wieladek (2016), Boeckx, Dossche, and Peersman (2017), and Burriel and Galesi (2018).

how innovations to nonstandard monetary policy affect risk-taking behavior (risk-taking channel) or certain financial stress indicators.[6] The overall conclusion of the vast majority of these studies is that unexpected unconventional interventions induced by monetary policy tend to improve the cyclical situation as well as the refinancing conditions but might also be associated with an intensification of risk-taking in the economy. However, this literature does not discuss any possible effects on the sovereign-bank nexus.

A number of microeconometric studies explore the effects of the ECB's unconventional monetary policy on banks' sovereign debt portfolios using bank-level data. In particular, these papers focus on how certain bank-specific and/or country-specific characteristics affect the behavior of the individual bank. Most contributions focus on the episodes immediately following the introduction of the ECB's long-term refinancing operations (LTROs) conducted with extended maturities between 2011 and 2012 and between 2013 and 2014. These studies mainly test (i) the *carry-trade* hypothesis (Acharya and Steffen 2015), according to which banks go long on high-risk, high-yield sovereign debt, which they fund either by borrowing from the ECB or by going short on low-yield debt, and/or (ii) the *moral suasion* hypothesis (Ongena, Popov, and Van Horen 2019), according to which banks hold domestic government debt partly due to political pressure. Acharya and Steffen (2015) find support for the carry-trade hypothesis as, during 2007–12, particularly large banks and those with low tier 1 capital ratios and high risk-weighted assets exploited government guarantees, arbitrage in regulatory risk weights, and access to central bank funding. Ongena, Popov, and Van Horen (2019) show that during the euro-area sovereign debt crisis, banks in fiscally stressed countries were considerably more

[6]See, for example, Angeloni, Faia, and Lo Duca (2015), Neuenkirch and Nöckel (2018), or Lewis and Roth (2019). Adrian and Liang (2018) provide a comprehensive review of the risk-taking channel of monetary policy. However, most of the empirical contributions dealing with the risk-taking channel are microeconometric in nature and focus on conventional monetary policy (Maddaloni and Peydro 2011; Altunbas, Gambacorta, and Marques-Ibanez 2014; Jimenez et al. 2014, for example). Delis, Hasan, and Mylonidis (2017) and Dell'Ariccia, Laeven, and Suarez (2017) are examples of microeconometric studies also covering episodes of unconventional monetary interventions.

likely than foreign banks to increase their holdings of domestic sovereign bonds in months with relatively high domestic sovereign bond issuance. Since the effect seemed not to be triggered by the ECB's liquidity provision, they conclude that their results reflect a moral suasion behavior.[7] Furthermore, according to Altavilla, Pagano, and Simonelli (2017) publicly owned, bailed-out, and poorly capitalized banks responded to sovereign stress by scaling up their holdings of domestic sovereign debt by more relative to other banks. This pattern turns out to be especially pronounced for public-owned banks at the time of large liquidity injections by the ECB in December 2011 and March 2012. These results show support for both a carry-trade and moral suasion behavior. Evidence provided by Drechsler et al. (2016) suggests that weakly capitalized banks in particular reacted to the ECB's liquidity injections during the euro-area sovereign debt crisis by investing a substantially higher fraction of the additional liquidity in domestic government bonds.[8] Peydro, Polo, and Sette (2017) look at Italian banks and find a positive relationship between the ECB's liquidity injections and the accumulation of domestic sovereign bonds, which is mostly driven by less well-capitalized banks. However, the latter tend to increase their holdings of relatively safe bonds instead of riskier assets, which suggests that the reach for liquidity and safety are much more important drivers than risk-shifting or regulatory arbitrage. Unlike our study, these studies do not explore how the *aggregate* banking sector's sovereign bond portfolio changes in response to unconventional monetary policy surprises.

Finally, Battistini, Pagano, and Simonelli (2014) and Colangelo et al. (2017) also built empirical macroeconomic models to study the

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[7] Uhlig (2014) shows in a theoretical model that governments potentially facing refinancing difficulties typically have an incentive to allow domestic banks to accumulate more risky domestic bonds. The opposite is the case when public finances are healthy. Battistini, Pagano, and Simonelli (2014) argue that sovereign stress strengthens this incentive, leading to a positive relationship between sovereign yields and the stock of domestic government bonds held by banks.

[8] Crosignani, Faria-e-Castro, and Fonseca (2017) and Jasova, Mendicino, and Supera (2018) report that Portuguese banks also used the funds obtained via the ECB's LTROs conducted between 2011 and 2012 to buy domestic government debt which was then offered as collateral to obtain short-run liquidity. Carpinelli and Crosignani (2017) derive similar evidence for Italian banks.

sovereign-bank nexus in the euro area. Colangelo et al. (2017) employ the VAR methodology and report that after 2011 banks tended to increase the home bias in their government bond portfolios. Battistini, Pagano, and Simonelli (2014) resort to a vector error-correction model (VECM) and explore how banks' domestic sovereign debt portfolios in euro-area countries are related to changes in the *common risk* and the *country-specific risk* components of sovereign yields.[9] Nevertheless, unlike us, Battistini, Pagano, and Simonelli (2014) and Colangelo et al. (2017) do not discuss the response of banks' sovereign debt portfolios to shocks triggered by nonstandard monetary policy.

3. Panel VAR with Sign Restrictions

3.1 Benchmark Specification

Consider a panel VAR model in reduced form:

$$y_{k,t} = \sum_{j=1}^{p} B_j y_{k,t-j} + \tilde{c}_k + \varepsilon_{k,t}, \tag{1}$$

where $y_{k,t}$ is a vector of endogenous variables for country k, B_j is a matrix of autoregressive coefficients for lag j, p is the number of lags, and $\tilde{c}_k$ is a vector of country-specific intercepts, which accounts for possible heterogeneity across the units. Furthermore, $\varepsilon_{k,t}$ is a vector of reduced-form residuals. In our benchmark model, the vector $y_{k,t}$ consists of industrial production as a measure for real activity; core Harmonised Index of Consumer Prices[10] (core HICP); the Eurosystem's amount of total assets; the policy rate on the main refinancing operations; the level of financial stress, which is approximated by the Country-Level Index of Financial Stress (CLIFS); the spread between EONIA and the policy rate; and the monetary

[9] Battistini, Pagano, and Simonelli (2014) observe that in most countries an increase in the common-risk component is associated with a rise in banks' domestic exposures. In periphery countries, this positive relationship is also present in the case of the country-specific risk component.

[10] The core HICP covers all items (consumer goods) excluding energy and unprocessed food.

financial institutions' (MFI) domestic government bond holdings relative to total assets. The Eurosystem's amount of total assets, the policy rate, and the interest spread are aggregate variables, i.e., identical for all countries, while the remaining variables are country specific. Each variable is linearly detrended at the country level over the sample period. For each element of $y_{k,t}$ we use a pooled set of $M \cdot T$ observations, where M denotes the number of countries and T denotes the number of observations corrected for the number of lags p. The reduced-form residuals $\varepsilon_{k,t}$ are stacked into a vector $\varepsilon_t = [\varepsilon'_{1,t} \dots \varepsilon'_{M,t}]'$, which is normally distributed with mean zero and variance-covariance matrix Σ.

Since our sample is short, we follow Ciccarelli, Maddaloni, and Peydro (2015) by using a panel of euro-area countries that comprises the crisis countries Italy (IT), Spain (ES), and Portugal (PT) as well the core countries Austria (AT), Belgium (BE), Germany (DE), Finland (FI), France (FR), and the Netherlands (NL).[11] The panel approach allows us to pool the diverse information from the countries, while controlling for heterogeneity in the constant term. A main advantage of the approach is that it increases the efficiency of the statistical inference. However, this comes at the cost of disregarding cross-country differences by imposing the same underlying structure for each cross-section unit. We take account of this shortcoming by distinguishing between the crisis countries and the core countries. Additionally, we adopt the hierarchical prior approach of Jarocinski (2010), which allows us to explore the reaction of individual euro-area economies to innovations related to unconventional monetary policy.

3.2 Data

The data are taken from the ECB and collected on a monthly basis covering the period from 2007:M1 to 2014:M12.[12] The beginning of

[11] Note that we exclude Ireland from our analysis, because compared with the other countries the Irish series on industrial production is characterized by a marked volatility. Nevertheless, including Ireland in our panel of countries has virtually no effect on the adjustment of the MFIs' domestic government bond holdings ratio to shocks related to unconventional monetary policy, but on the response of industrial production.

[12] See the online appendix for a description of the data.

the sample period is determined by the launch of the ECB's unconventional monetary policy measures that started during 2007 before the financial crisis intensified. In particular, the central bank conducted supplementary long-term refinancing operations (LTROs) in August 2007 in response to the financial turmoil. The switch to main refinancing operations (MROs) with fixed rates and full allotment occurred in October 2008 and was accompanied by the implementation of LTROs with a maturity of 6 and 12 months, respectively, which were also offered at fixed rates with full allotment (ECB 2011). At the same time, government-guaranteed own-use bonds were accepted as collateral and the collateral rating for central bank refinancing was reduced. The SMP was launched in May 2010 in response to the sovereign debt crisis and supported the effects of two covered bond purchasing programs (CBPPs). The announcement of the OMT in September 2012 contributed to a lowering of sovereign bond yields, although the program itself was not activated (Altavilla, Giannone, and Lenza 2016).[13] Meanwhile, LTROs were offered with a maturity of up to 36 months, which were followed by a series of TLTROs (targeted LTROs) with a maturity of 45 and 48 months, respectively. Additionally, the interest rate on the deposit facility was cut to become negative. Finally, the ECB modified its communication policy by intensifying its *forward guidance.* The end of the sample period is related to the ECB's launch of the extended Asset Purchase Programme (APP) that was announced in January 2015 (Breckenfelder et al. 2016). The reason for excluding the APP from our analysis is that the path of the corresponding asset purchases was to a large degree anticipated by economic agents. In fact, the central bank published precisely the monthly volumes of its asset purchases, the structure of the purchases, as well as the duration of the program. Later information on the extension of the program as well as on the modalities of the end of the net purchases were released. Thus, the APP generated much of its effects through announcement by signaling that the future path of interest rates will be low, which is akin to forward guidance, i.e., the announcement that policy interest rates will remain at the lower bound over a

[13]See also Dell'Ariccia, Rabanal, and Sandri (2018) or Hristov et al. (2019), among others, for a discussion.

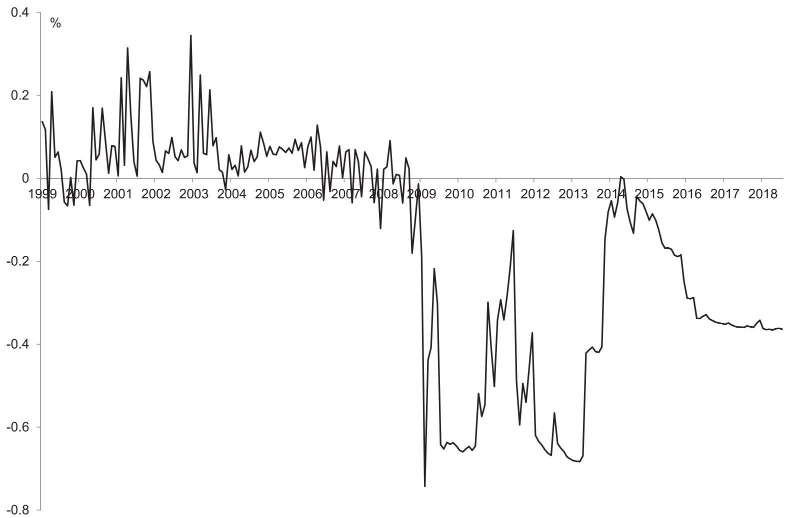
longer period (Breckenfelder et al. 2016).[14] Accordingly, if anything, the 2015–18 episode is likely characterized by, at most, negligible monetary shocks. Thus, simply extending our sample beyond 2014 could bias our estimates and inference. Although the launch of the APP came with the implementation of a second series of TLTROs, it seems that the program reflects a structural break in the conduct of unconventional monetary policy due to its differences compared with earlier nonstandard measures.

The main variable of interest is the ratio of MFIs' holdings of domestic government bonds to total assets.[15] Besides this ratio, the selection of variables refers to Gambacorta, Hoffman, and Peersman (2014) and Boeckx, Dossche, and Peersman (2017) and aims to capture the main economic interactions after the financial crisis. Industrial production and the dynamics of prices are supposed to reflect the macroeconomic development. The Eurosystem's volume of total assets serves as a measure of unconventional monetary policy (ECB 2015). Other studies such as Weale and Wieladek (2016), Gambetti and Musso (2017), or Hesse, Hofmann, and Weber (2018) use the central bank's asset purchases as a measure of nonstandard monetary policy. However, in the case of the ECB's unconventional monetary policy conducted over the period 2007–14, focusing only on asset purchases might be too narrow because it neglects the monetary policy effects that arose in the wake of open market operations with full allotment or the relaxing of collateral requirements. Alternatively, the effects of unconventional monetary policy may be measured by an expanding monetary base (Schenkelberg and Watzka 2013). However, monetary policy measures like the SMP would then not be considered, since the asset purchases conducted under this program were sterilized (Boeckx, Dossche, and Peersman 2017). The MRO

[14] A large quantity of asset purchases by the central bank under a program like the APP can be seen as a credible commitment by monetary policy to keep interest low in the future. Following Krishnamurthy and Vissing-Jorgensen (2011), this transmission channel of monetary policy is denoted as the signaling channel. Note that the OMT, in contrast to the APP, was only announced but not activated.

[15] Note that we only consider the exposures of banks to sovereign debt that are directly observable as explicit balance sheet positions. Beyond that, banks are also exposed to sovereign risk through positions in various derivatives or potentially through offshore institutions. However, we abstract from such exposure due to the very limited availability of suitable data (Koijen et al. 2017).

Figure 1. Interest Rate Spread



Notes: Data are taken from the ECB. Own calculations. Spread is the difference between EONIA and the policy rate on main refinancing operations.

rate is included to facilitate the distinction between conventional and unconventional monetary policy shocks. The interest spread is included as an additional indicator of nonstandard monetary policy. Figure 1 displays that the spread was relatively close to zero in normal times, indicating that EONIA was sticking closely to the policy rate.[16] Essentially, the provision of open market operations with full allotment widened the interest rate spread.

Finally, the CLIFS indicator is a measure of financial stress. By conditioning on it, we control for possible endogenous reactions of the Eurosystem's balance sheet to financial turbulence. More precisely, given the ECB's full allotment policy, changes in the Eurosystem's balance sheet might be demand induced in the case of elevated financial stress (Boeckx, Dossche, and Peersman 2017). Industrial production, the price level, the Eurosystem's volume of total assets,

[16] The banks' usage of the marginal lending facility at the end of a minimum reserve maintenance period mainly caused the shifts of the spread between 1999 and 2003.

and the CLIFS are in logs, while the MFIs' domestic government bond holdings ratio, the policy rate, and the interest rate spread are expressed in percent.

3.3 Identification of Unconventional Monetary Policy Shock

The VAR model (1) is estimated with Bayesian methods using—as proposed by Uhlig (2005)—a normal-Wishart prior.[17] In the model of the crisis countries, the lag order is set to $p = 3$. The selection of the lag length is based on a number of diagnostic tests on the disturbances.[18] While the Bayesian information criterion (BIC) indicates a lag length of two, the results of Lagrange multiplier (LM) tests for autocorrelation and different tests for heteroskedasticity rather suggest to use three lags. Nonetheless, our findings proved robust against alternative lag lengths, e.g., lag orders of $p = 2$ and $p = 4$, respectively. In the model of the core countries the lag order is set to $p = 4$, which is also derived on the basis of diagnostic tests on the error terms. However, to facilitate comparison with the results reported for the crisis countries, we also consider a lag length of three. Based on the outcome of the estimated models, we generate impulse responses of the variables to structural shocks η_t.[19] We identify the structural shocks through sign restrictions using the algorithm of Arias, Rubio-Ramirez, and Waggoner (2014), which allows for imposing sign restrictions as well as zero restrictions on the impulse responses to a structural shock.

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[17]In particular, the prior is rather *weak* (non-informative). Accordingly, our results are titled toward the information contained in the data itself. See also Canova (2007, chapter 10) for a discussion. As a robustness check, we also considered a Minnesota-type prior that imposes a stronger structure (shrinkage) over the parameters of the model. The results derived on the basis of that prior are similar to those reported below.

[18]The results of the diagnostic tests on the error terms are not reported here but are available upon request.

[19]Moreover, we assessed all estimated models with respect to parameter stability by conducting a number of Chow tests for different breakpoints. The results of these tests point in the direction of supporting the assumption of constant parameters. However, in spite of these findings, our results reported subsequently should be interpreted with a certain degree of caution due to possible instabilities arising over our sample period. Note, that the Chow-test results are not reported here, but they are available upon request.

The structural representation of the VAR model (1) can be expressed as

$$A_0 y_{k,t} = \sum_{j=1}^{p} A_j y_{k,t-j} + c_k + \eta_{k,t}, \tag{2}$$

with $\eta_{k,t} \sim N(0, I)$, where I is the identity matrix. The reduced-form representation of the SVAR is derived by multiplying both sides of (2) with A_0^{-1}. The structural shocks $\eta_{k,t}$ relate to the reduced-form residuals $\epsilon_{k,t}$ according to $\epsilon_{k,t} = A_0^{-1}\eta_{k,t}$, where $\epsilon_{k,t} \sim N(0, \Sigma)$. The identification of the structural parameters of the model is equivalent to finding the appropriate matrix $\tilde{A} = A_0^{-1}$, which is done by means of sign and zero restrictions. The algorithm of Arias, Rubio-Ramirez, and Waggoner (2014) uses the fact that the Cholesky decomposition of the covariance matrix of the reduced-form residuals $\Sigma = PP'$, where P' is lower triangular, can be extended by any orthogonal matrix Q as follows: $\Sigma = PP' = P'Q'QP$, where $QQ' = I$. As the algorithm further requires that Q has a uniform distribution with respect to the Haar measure, Q can be generated by means of a QR factorization of a random matrix W of proper dimensions, where each element of W follows an independent standard normal distribution. A particular Q is considered a solution to the identification problem if the impulse responses implied by $\tilde{A} = P'Q'$ satisfy a set of sign restrictions. To estimate the posterior of the structural model, we follow the steps suggested by Arias, Rubio-Ramirez, and Waggoner (2014): (i) we draw from the posterior of the reduced-form model, (ii) then we draw an orthogonal matrix Q, (iii) we keep the draw if the combination of reduced-form parameters and Q satisfies the sign and zero restrictions, and discard it otherwise, and (iv) we return to (i) until the required number of draws satisfying the restrictions is obtained. Our results are based on 10,000 draws consistent with the imposed sign restrictions. The latter are discussed subsequently.[20]

[20] It has to be noted that sign restriction relying on the Haar measure regarding the rotation matrix Q could lead to implicit priors on the impact impulse responses (Baumeister and Hamilton 2015, 2018).

3.3.1 Identification of Unconventional Monetary Policy Shocks in the Literature

Following Curdia and Woodford (2011), a number of studies have extended dynamic stochastic general equilibrium (DSGE) models by incorporating unconventional monetary policy (Chen, Curdia, and Ferrero 2012; Falagiarda 2014; Le, Meenagh, and Minford 2016; Quint and Rabanal 2017; Hohberger, Priftis, and Vogel 2019). Although these studies differ in their conclusion regarding the effectiveness of disturbances related to nonstandard measures of monetary policy, they all show that output is stimulated in response to an unconventional monetary policy shock, which is conducted in terms of a quantitative easing, i.e., through the purchase of sovereign bonds that induces the level of base money to rise. Simultaneously, they report that inflation rises after the shock, whereas the government bond yield, the spread between the bond rate and the short-term rate or the risk premium on credit, decline. Table 1 summarizes the findings.

In addition, Gertler and Karadi (2011) analyze the effects of nonstandard monetary policy conducted by the central bank in terms of a credit injection in reaction to a capital quality shock. Unconventional monetary policy significantly moderates the recession induced by an adverse shock, because it dampens the rise in the interest spread, which in turn dampens the investment decline. The central bank's balance sheet rises, but decreases slowly thereafter over time. Inflation remains largely benign.

Using VAR models, Baumeister and Benati (2013), Schenkelberg and Watzka (2013), Gambacorta, Hofmann, and Peersman (2014), Weale and Wieladek (2016), Boeckx, Dossche, and Peersman (2017), Burriel and Galesi (2018), and Hesse, Hofmann, and Weber (2018) explore empirically the macroeconomic effects across countries of an innovation related to nonstandard monetary policy. The latter is identified by imposing sign restrictions.[21] Table 2 summarizes the different identification schemes which comprise a shock to a central bank's total assets (Gambacorta, Hofmann, and Peersman 2014; Boeckx, Dossche, and Peersman 2017), a quantitative easing (QE)

[21]Weale and Wieladek (2016) also identify an unconventional monetary policy shock by adopting a Cholesky ordering.

Table 1. Theoretical Effects of an Unconventional Monetary Policy Shock

	Real Output	Inflation Rate	Government Bond Purchases	Government Bond Yield	Risk Premium	Interest Rate Spread
Falagiarda (2014)	↑	↑	↑	↓		
Chen, Curdia, and Ferrero (2012)	↑	↑			↓	
Le, Meenagh, and Minford (2016)	↑	↑	↑		↓	
Quint and Rabanal (2017)	↑	↑	↑			↓
Hohberger, Priftis, and Vogel (2019)	↑	↑	↑	↓		

Notes: Chen, Curdia, and Ferrero (2012) report impulse responses to a simulated shock to market value of long-term debt. Regarding Quint and Rabanal (2017), we report the case when the stock of assets held by the central bank follows an AR(2) process and the unconventional monetary policy shock is conducted by purchasing government bonds. The interest rate spread denotes the difference between the government bond rate and the short-term rate.

Table 2. Identification of an Unconventional Monetary Policy Shock in the Empirical Literature

Study	Model	Restrictions Imposed
Gambacorta, Hofmann, and Peersman (2014)	Panel VAR for nine advanced economies. 2008:M1–2011:M6	Output: 0 on impact; price level: 0 on impact; financial stress: ≤ 0; central bank total assets: ≥ 0.
Boeckx, Dossche, and Peersman (2017)	VAR model for the euro area. 2007:M1–2014:M12	Output: 0 on impact; price level: 0 on impact; central bank total assets: ≥ 0; financial stress: ≤ 0; spread between EONIA and policy rate: ≤ 0; policy rate: 0 on impact.
Burriel and Galesi (2018)	GVAR for euro-area countries. 2007:M1–2015:M9	Output: 0 on impact; price level: 0 on impact; central bank total assets: ≥ 0; CISS: ≤ 0; spread between EONIA and policy rate: ≤ 0; policy rate: 0 on impact; new credit growth: 0 on impact.
Hesse, Hofmann, and Weber (2018)	VAR models for the United States and the United Kingdom. 2008:M11–2014:M10	Output: 0 on impact; price level: 0 on impact; bond yield: ≤ 0; stock prices ≥ 0; asset purchase announcements: ≥ 0.
Weale and Wieladek (2016)	VAR model for the United States. 2009:M3–2014:M5	Output: unrestricted; price level: unrestricted; central bank asset purchases: ≥ 0, bond yield: ≤ 0, and real equity prices: ≥ 0.
Schenkelberg and Watzka (2013)	VAR model for Japan. 1995:M1–2010:M9	Output: 0 on impact; price level: 0 on impact and ≥ 0 thereafter; central bank reserves: > 0.
Baumeister and Benati (2013)	TVP-VAR models for the United States and the United Kingdom. 1965:Q4–2011:Q4 and 1975:Q2–2011:Q4	Output: ≥ 0; price level: ≥ 0; short-term rate: 0 on impact; interest rate spread: ≤ 0.

shock related to the central bank's asset purchases (Weale and Wieladek 2016; Hesse, Hofmann, and Weber 2018), or an increase in reserves (Schenkelberg and Watzka 2013), as well as an interest spread shock induced by nonstandard interventions of monetary policy (Baumeister and Benati 2013).[22]

In general, the results of these studies indicate that the estimated response of output to a nonstandard monetary policy shock is positive and similar to the one found in the literature on the effects of *conventional* monetary policy. Simultaneously, the price level rises in response to a disturbance to unconventional monetary policy, whereas the bond yield, the spread between the government bond rate and the short-term rate, as well as financial stress decline.[23]

3.3.2 Sign Restrictions

In our benchmark specification, we follow Boeckx, Dossche, and Peersman (2017) regarding the identification of a shock to unconventional monetary policy. Table 3 summarizes our set of restrictions, which comprises both zero restrictions and sign restrictions.[24]

Industrial production, prices, and the policy rate are restricted to a zero response on impact after the shock. Furthermore, the Eurosystem's total assets are assumed to increase, whereas the interest rate spread, i.e., the spread between EONIA and the policy rate, as well as financial stress are supposed not to rise. Finally, the reaction of

[22]Kapetanios et al. (2012) and Hristov et al. (2019) adopt the identification scheme of Baumeister and Benati (2013) to identify unconventional monetary policy shocks in the United Kingdom and the euro area.

[23]Alternatively, a number of studies seek to identify the effects of unconventional monetary policy by using an *event-study* approach that is based on high-frequency data and focuses on a narrow window around the policy announcement. See Dell'Ariccia, Rabanal, and Sandri (2018) for an overview and discussion.

[24]Recently, a controversy has surrounded the identification strategy adopted by Boeckx, Dossche, and Peersman (2017). In particular, Elbourne and Ji (2019) have argued that this strategy does not identify shocks to unconventional monetary policy. They argue that one basically obtains very similar impulse responses even without the main identifying restriction—that imposed upon the ECB's total assets. However, Boeckx et al. (2019) demonstrate through a number of empirical exercises that the approach by Boeckx, Dossche, and Peersman (2017) indeed successfully identifies unconventional monetary policy shocks.

Table 3. Identification of an Unconventional Monetary Policy Shock

	Benchmark Model
Output	0
Price Level	0
Eurosystem's Total Assets	≥ 0
Policy Rate	0
Interest Rate Spread	≤ 0
CLIFS	≤ 0
MFIs' Bond Holdings Ratio	?

Notes: Restrictions are set in accordance with Boeckx, Dossche, and Peersman (2017). Interest rate spread denotes the spread between EONIA and the policy rate. Zero restrictions are denoted by 0. Sign restrictions are imposed as ≥ 0 or ≤ 0 and are binding over a period of four months. Unrestricted responses are denoted by "?."

the MFIs' domestic government bond holdings ratio to an unconventional monetary policy shock is unrestricted. The sign restrictions are imposed as $\leq$ or $\geq$ and are binding over a period of four months.[25]

According to Boeckx, Dossche, and Peersman (2017), the set of restrictions is rich enough to ensure that the unconventional monetary policy shock is orthogonal to real economy disturbances, shocks in financial markets, and conventional innovations to monetary policy. Typically, output and prices move in opposite directions after an aggregate supply shock. An aggregate demand shock is characterized by an immediate reaction of monetary policy, whereby the spread between EONIA and the policy rate normally does not widen.[26] A shock to standard monetary policy also causes an immediate response of the policy rate, which exerts direct influence on

[25]Note that the choice of the period over which the sign restrictions are binding is arbitrary. Thus, we also considered a period of two months over which the restrictions are binding. The results are very similar to those reported below.

[26]Indeed, VAR models estimated in normal times to investigate the macroeconomic effects of a monetary policy shock frequently use the EONIA as a proxy for conventional monetary policy. See Ciccarelli, Maddaloni, and Peydro (2015) as an example.

output and the price level. Ultimately, adverse disturbances in financial markets increase financial stress, thereby causing a recession that is followed by an immediate expansionary response of monetary policy.

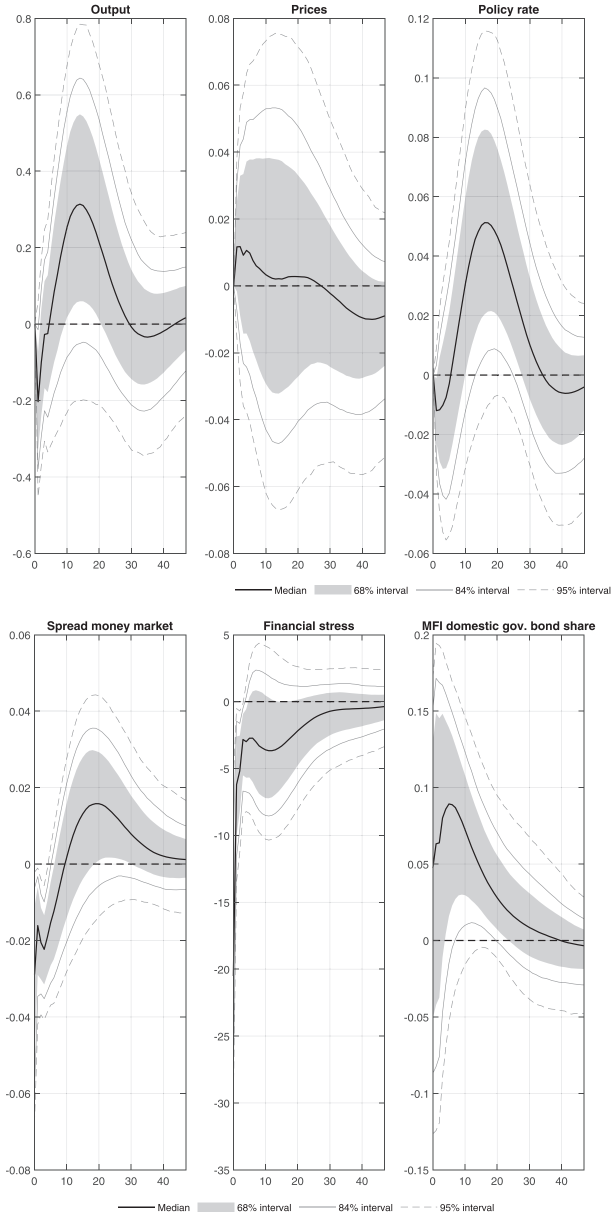
4. Results

In our analysis, we estimate separate panel VARs for each of the following two groups of euro-area member countries: (i) the *crisis countries*, comprising Italy, Spain, and Portugal, and (ii) the *core countries*, consisting of Austria, Belgium, Germany, Finland, France, and the Netherlands. In doing so, we allow for possible structural heterogeneities across the two groups. Such heterogeneities appear a priori very likely given the differences between the groups in terms of historical experience during the period of the European sovereign debt crisis, i.e., 2010–13. In particular, crisis countries' sovereigns faced difficulties in tapping international capital markets, leading to extraordinary fiscal distress. As explained in the introduction, the unfavorable situation was likely exaggerated by the emergence of *doom loops* due to the sovereign-bank nexus. All these culminated in outright financial and banking crises. As a consequence, Italy, Spain, and Portugal found themselves forced to initiate substantial recapitalizations and restructurings of their banking sectors, adopt massive cyclical and structural adjustments in fiscal policy, and start launching far-reaching structural reforms in labor and good markets. In contrast, the core countries faced moderate recessions over the years 2011–13. Since they had no crisis-like eruptions, they were not forced to undertake substantial structural adjustments, and even benefited from being accepted as *safe havens* by international investors.

4.1 Benchmark Model

Figure 2 shows the average reaction of a euro-area crisis country to an expansionary unconventional monetary policy shock. The black solid lines are the median impulse responses of the posterior distribution. The shaded areas correspond to the 68 percent posterior credibility bounds, while the thin gray lines refer to the

Figure 2. Euro-Area Crisis Countries Impulse Responses to a UMP Shock



Notes: "UMP Shock" denotes an unconventional monetary policy shock. Industrial production, prices, the Eurosystem's volume of total assets, and the CLIFS are in logs, while the remaining variables are in percent. The unconventional monetary policy shock is identified as in Boeckx, Dossche, and Peersman (2017) by using the same set of zero and sign restrictions.

84 percent and the thin gray dashed lines to the 95 percent credible sets, respectively.[27] The simulation horizon covers 48 months.

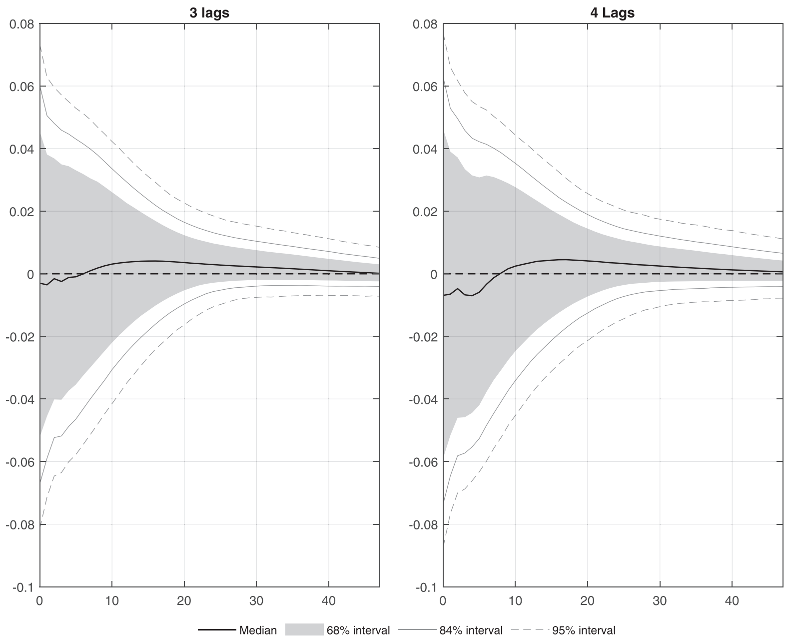
We see that the economy in the euro-area crisis countries is stimulated by a sudden unconventional monetary policy loosening. Output rises, reaching a peak around 14 months after the shock. The reaction of prices is sticky. Monetary policy starts to tighten the policy rate with a delay. These results are in line with the findings of Boeckx, Dossche, and Peersman (2017), who also report a boom of the economy after a positive central bank balance sheet innovation that is followed by a tightening of standard monetary policy. The spread between EONIA and the policy rate remains negative for about six months. The CLIFS financial stress indicator declines. Banks respond to the favorable unconventional monetary policy shock by adjusting their sovereign debt portfolios. The MFIs' domestic government bond holdings ratio exhibits a sizable increase. In particular, our results suggest a maximum median increase in the MFI ratio by 0.089 percentage point after a rise in the ECB's total assets by 1 percent. The peak effect is reached after five months. We find that the MFI ratio increases by 1.29 percent in Italy (from 6.94 percent on average to 7.03 percent), by 1.68 percent in Spain (from 5.34 percent of average to 5.43 percent), and by 2.52 percent in Portugal (from 3.56 percent on average to 3.67 percent).[28]

The reaction of the MFIs' domestic government bond holdings ratio in the core countries to an unconventional monetary policy shock is shown in figure 3, which displays the results derived from both a model estimated with three lags and four lags. The results indeed indicate that, on average, the banking sector in the core countries responds very differently from that in the crisis economies. In particular, we observe that the core countries' MFI ratio hardly responds at all to an unconventional monetary policy shock.

[27] A number of studies—for example, Peersman (2005), Uhlig (2005), Scholl and Uhlig (2008), Hofmann, Peersman, and Straub (2012), Sa, Towbin, and Wieladek (2014), Boeckx, Dossche, and Peersman (2017), and many others—report the 68 percent posterior credibility intervals.

[28] In absolute figures our results suggest that a rise in the ECB's total assets by €10 billion is followed by an increase in MFIs' domestic government bond holdings by €728,870 million in Italy, €631,416 million in Spain, and €97,192 million euro in Portugal. However, note that these figures are calculated on the basis of averages and are, thus, only an approximation.

Figure 3. Reaction of the Core Countries' MFI Ratio to a UMP Shock



Notes: See figure 2. The core countries comprise Austria, Belgium, Germany, Finland, France, and the Netherlands. "MFI Ratio" denotes the MFIs' domestic government bond holdings ratio.

4.2 Forecast Error Variance Decomposition for the Crisis Countries

In order to understand the quantitative importance of the unconventional monetary policy shock for the crisis countries, we compute the forecast error variance decomposition (FEVD), which in contrast to the impulse response analysis takes into account the estimated magnitude of the shock. Table 4 reports the forecast error variance decomposition of each variable at different forecast horizons.

We find that fluctuations in output in the crisis countries are hardly explained by the unconventional monetary policy shock. The same holds true for the evolution of prices with a maximum share of only about 3 percent. In contrast, the variation of the

Table 4. Crisis Countries Forecast Error Variance Decomposition

Month	Output	Price Level	Eurosystem's Total Assets	Policy Rate	Interest Rate Spread	CLIFS	MFIs' Bond Holdings Ratio
1st	—	—	14.13	—	8.76	30.09	13.36
2nd	0.35	0.13	13.06	0.12	9.31	25.91	13.36
3rd	0.47	0.32	12.20	0.24	10.91	23.45	13.28
6th	0.87	0.72	9.82	0.59	13.99	18.85	15.27
12th	2.16	1.46	8.05	2.39	11.45	16.41	19.32
24th	3.78	2.63	8.93	6.93	12.07	16.26	21.63
36th	4.23	3.21	9.66	7.60	12.68	16.28	21.37
48th	4.40	3.67	10.02	7.75	12.79	16.22	21.14

Notes: Quantitative importance of unconventional monetary policy shock measured in percent. Crisis countries include Italy, Spain, and Portugal. Zero restrictions set on output, prices, and the policy rate are imposed over one month. In this respect, "—" signals impact of zero restriction. Sign restrictions set on remaining variables are imposed over the first four months. The reaction of the MFIs' domestic government bond holdings ratio to an unconventional monetary policy shock is unrestricted.

MFIs' domestic government bond holdings ratio triggered by a shock to nonstandard policy is considerably larger. On average, the shock explains around 19 percent of the fluctuation. Moreover, the contribution of the fluctuation of the CLIFS indicator of financial stress caused by disturbances related to nonstandard monetary policy is also notable, fluctuating on average also around 20 percent.

4.3 Alternative Model Specifications

We assess the robustness of our results by considering three alternative models in which we adopt different schemes to identify a nonstandard monetary policy innovation.

4.3.1 Alternative A

In the first alternative specification we exclude the policy rate on the main refinancing operations from the model, and include another variable instead—the volume of liquidity provided via the MROs relative to total assets. The ECB's main refinancing operations were in normal times the most important open market operations through which the bulk of liquidity was provided. They played a pivotal role in signaling the stance of monetary policy and managing the liquidity situation in the money market. Figure 4 shows that the share of liquidity provided via the MROs relative to all liquidity-providing open market operations amounted to about 75 percent on average between 1999 and 2007, i.e., before the onset of the financial crisis.

However, the share of the MROs dropped markedly after the intensification of the financial crisis in 2008 as longer-term refinancing operations, targeted longer-term refinancing operations, and outright purchases became more and more important. Hence, the ECB's unconventional monetary policy measures are characterized not only by causing an extension in the central bank balance sheet but also by causing a change in its composition. Accordingly, to identify an unconventional monetary policy shock, we assume that it is associated with a nonpositive response of the ratio of MRO volume to the Eurosystem's total assets. The column labeled "Alternative A" in table 5 displays the set of restrictions imposed to identify the

Figure 4. Share of Main Refinancing Operations Relative to All Liquidity-Providing Open Market Operations

Notes: Data are taken from the ECB. Own calculations. Liquidity-providing open market operations include the main refinancing operations, longer-term refinancing operations, and other liquidity-providing operations such as outright purchases, structural operations, and fine-tuning operations.

unconventional monetary policy shock. The restrictions regarding the Eurosystems's total assets, industrial production, the price level, the money market spread, and the CLIFS indicator are the same as in our benchmark specification.

4.3.2 Alternative B

In our second alternative specification we replace the Eurosystem's total assets by the volume of liquidity provided by the LTROs. The latter were used by banks to fund their assets.[29] To identify an unconventional monetary policy shock, we assume that it induces an increase in the volume of LTROs. The remaining variables are

[29]The provision of the LTROs is frequently associated with the term "Sarko trade." Former French president Nicolas Sarkozy suggested that each government in the euro area may borrow from their own commercial banks, which in turn could use the LTROs to gain access to central bank refinancing.

Table 5. Alternative Identification Schemes

	Alternative A	Alternative B	Alternative C
Output	0	0	0
Price Level	0	0	0
Policy Rate	•	0	0
Eurosystem's Total Assets	≥ 0	•	•
Volume LTROs	•	≥ 0	•
Shadow Rate	•	•	≤ 0
Volume MROs to Total Assets	≤ 0	•	•
Interest Rate Spread	≤ 0	≤ 0	≤ 0
CLIFS	≤ 0	≤ 0	≤ 0
MFIs' Bond Holdings Ratio	?	?	?

Notes: "Interest Rate Spread" denotes the spread between EONIA and the policy rate. Zero restrictions on impact are denoted by 0. Sign restrictions are imposed as ≥ 0 or ≤ 0 and are binding over a period of six months. Unrestricted responses are denoted by "?". Finally "•" denotes that the variable is not included in the respective model.

subject to the same sign and zero restrictions as in our baseline specification. The column in table 5 labeled "Alternative B" summarizes restrictions imposed.

4.3.3 Alternative C

Finally, in our third alternative specification we replace the Eurosystem's total assets with a shadow rate. In particular, we resort to the shadow rate constructed by Krippner (2013), which is derived from a term structure model and attempts to proxy the true stance of monetary policy in times when conventional monetary policy is constrained by the zero lower bound and nonstandard measures of monetary policy are adopted.[30]

We identify the shock related to an expansionary unconventional monetary policy by restricting the shadow rate to decrease while

[30]Leo Krippner's shadow rate series for the euro area is available under https://www.rbnz.govt.nz/research-and-publications/research-programme/additional-research/measures-of-the-stance-of-united-states-monetary-policy/comparison-of-international-monetary-policy-measures. Figure A.2 in the online appendix displays the shadow rate for the euro area together with the policy rate.

the remaining variables are again subject to the same sign and zero restrictions as in the benchmark specification (see the column labeled "Alternative C" in table 5).[31]

4.3.4 Impulse Responses Based on the Alternative Specifications

We estimate the alternative models for the euro-area crisis countries using each time a lag order of three. Figure 5 summarizes the reactions of the MFIs' domestic government bond holdings ratio to an innovation to unconventional monetary policy.

As can be seen, the impulse responses confirm those implied by our benchmark specification, albeit in alternative A the response of the MFI ratio to the shock is less pronounced.[32] Thus, banks in the crisis countries seemed to increase their fraction of domestic sovereign bonds in their balance sheets in response to the nonstandard monetary policy measures. Regarding the core countries, each of the three alternative models indicates a reaction of the MFI ratio to an unconventional monetary policy shock that is hardly notable.[33]

4.4 Cross-Country Heterogeneity

Next, we assess possible heterogeneity across the euro-area economies.

4.4.1 Impulse Response Analysis

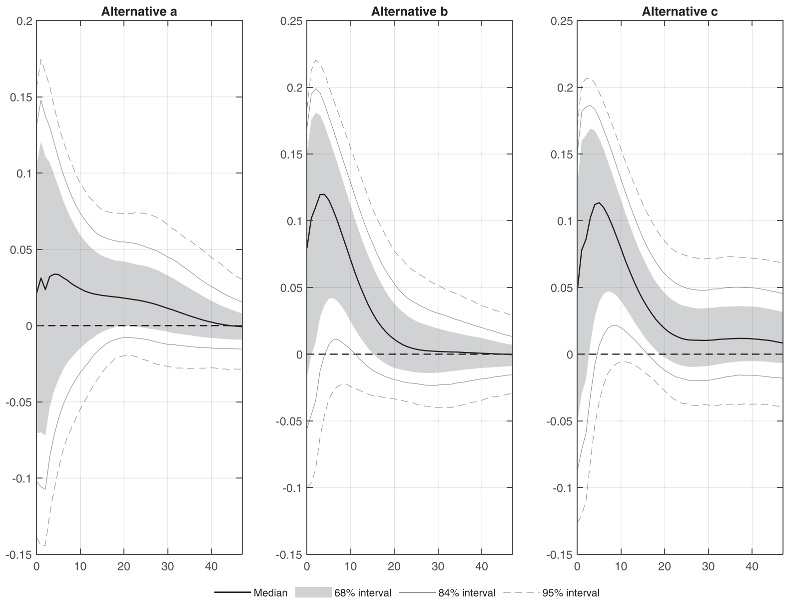
Following Jarocinski (2010), we estimate a panel of VAR models for both groups of euro-area countries, i.e., the crisis economies and

[31] In addition, we performed several further experiments with models including a shadow rate. First, replacing Krippner's shadow rate with the one proposed by Wu and Xia (2016) leaves the results unchanged. Second, irrespective of the shadow rate used, replacing the zero restrictions on industrial production and prices with "$\geq$" sign restrictions or excluding the CLIFS and the money market spread from the model delivers qualitatively similar impulse responses. The results obtained in these robustness checks are available upon request.

[32] The responses of the remaining variables in the alternative models are similar to those in the benchmark specification. They are not shown here but are available upon request.

[33] The corresponding impulse responses are available upon request.

Figure 5. Reaction of the Crisis Countries' MFI Ratios to a UMP Shock



Notes: The MFI ratio impulse responses are calculated on the basis of the alternative model specifications A–C. The black solid lines are the median impulse responses of the posterior distribution. The shaded areas correspond to the 68 percent posterior credibility bounds. The thin gray lines refer to the 84 percent and the thin gray dashed lines to the 95 percent credible sets, respectively.

the core economies, by using a hierarchical Bayesian panel model estimator. The VAR model for country k reads as follows:

$$y_{k,t} = \sum_{j=1}^{p} B_{k,j} y_{k,t-j} + \tilde{c}_k + \varepsilon_{k,t}, \tag{3}$$

where the matrices of autoregressive coefficients $B_{k,j}$ are now country specific. The Bayesian estimation is conducted with the prior that the countries in every group do not differ in terms of their macroeconomic dynamics. Hence, the countries are assumed to be special cases of the same underlying economic model, which implies that the matrices B_j tend to be similar across the economies. This prior results in the posterior which pools information across countries (Jarocinski 2010). The models for the crisis countries are estimated with a lag length of three, while for the models for the core countries four lags are used. We identify the unconventional monetary policy shock as in the benchmark model, i.e., by imposing the same zero restrictions and sign restrictions.[34]

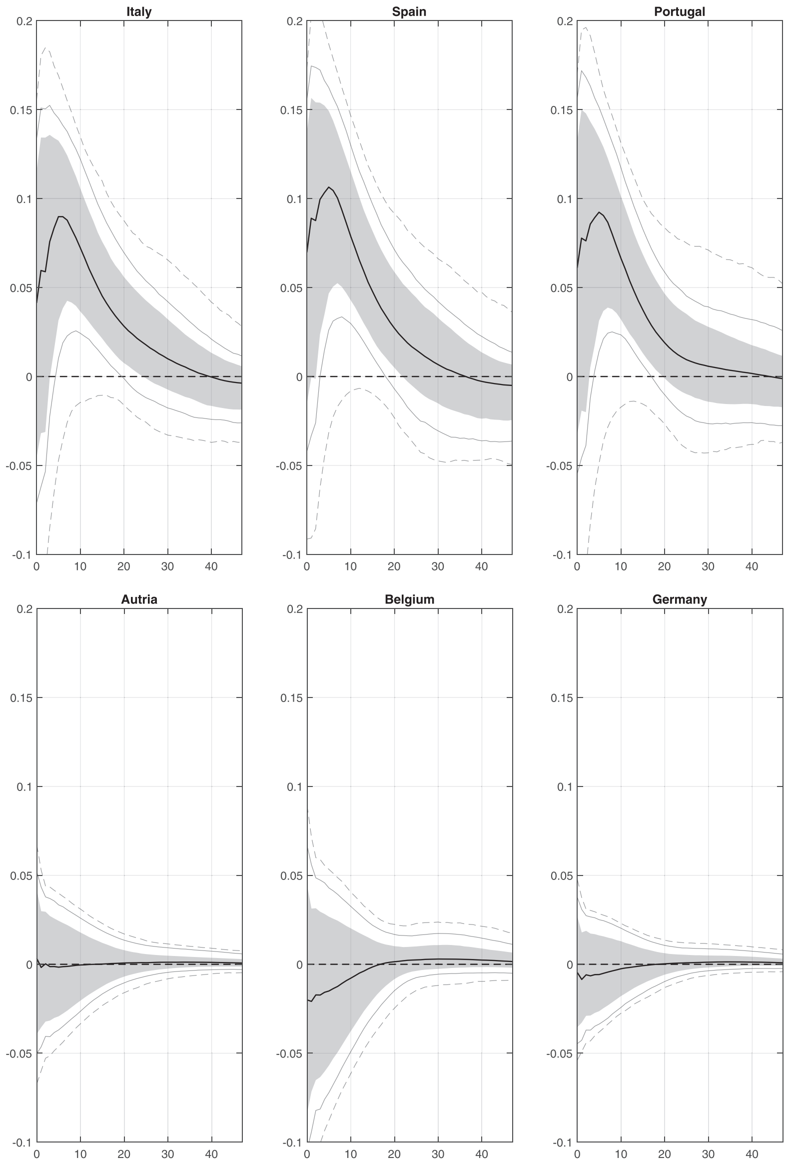
Figure 6 summarizes the responses of the MFIs' domestic government bond holdings ratios across the countries in both groups to an expansionary shock to nonstandard monetary policy. In the crisis countries, the ratios increase significantly. The adjustment across the countries appears to be qualitatively and quantitatively quite similar. By contrast, in all core countries the reaction of the ratios turns out to be insignificant.

4.4.2 Historical Decomposition

On the basis of the VAR models (3) estimated for the crisis countries, we compute a historical decomposition. While the FEVD sheds some light on the quantitative importance of the structural shocks over the entire sample period, the historical decomposition allows to figure out the relevance of a shock for each month within the sample period. We base our calculation of the historical decomposition on the models estimated in the spirit of Jarocinski (2010) because the approach allows for differences across countries by imposing only the restriction that the mean of the slope coefficients are identical for each unit.

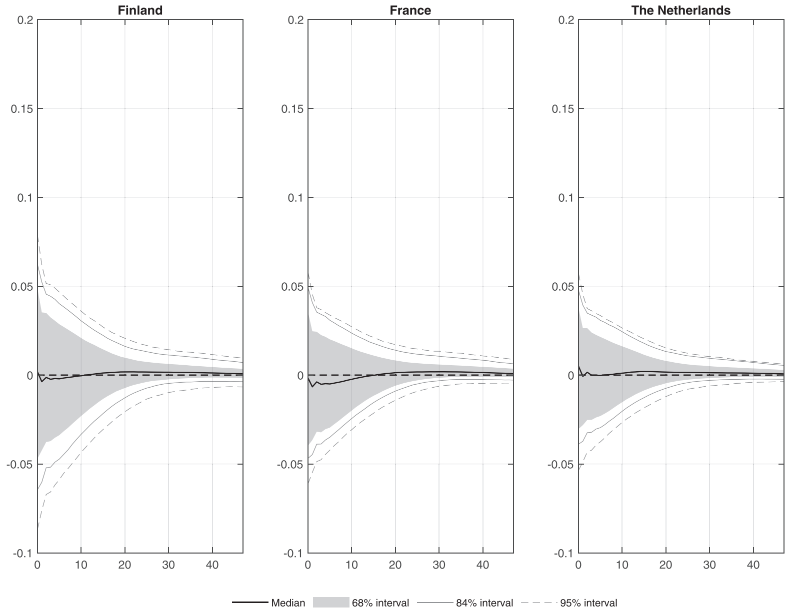
[34]We resort to the BEAR (Bayesian estimation, analysis and regression) toolbox of Dieppe, van Roye, and Legrand (2016) for estimation. The estimation requires Gibbs sampling. We perform the algorithm 10,000 times, discarding the first 5,000 samples as burn-in. The standard priors are set according to the following: For the overall tightness—a parameter strongly affecting the degree of cross-sectional heterogeneity—we set the starting value to $\lambda_1 = 0.1$, and assume the following inverse-gamma prior $\lambda_1 \sim IG(s_0/2, v_0/2)$ with $s_0 = 0.001$ and $v_0 = 0.001$. For the country-specific vector of slope coefficients we have a normal prior, i.e., $\beta_i \sim \mathcal{N}(b, \Sigma_b)$ with a diffuse prior for the overall mean b, and a Minnesota-type prior for the covariance matrix Σ_b, i.e., $\Sigma_b = (\lambda_1 \otimes I_q)\Omega_b$, where the matrix Ω_b is defined in standard Minnesota-style manner while λ_1 is the overall-tightness parameter. Furthermore, a diffuse prior is used for the covariance matrix Σ_i of the reduced-form residuals of cross-sectional unit i. Finally, we set the priors for cross-variable weighting to 1, for lag decay to 0, and for exogenous variable tightness to 100 (Jarocinski 2010).

Figure 6. Impulse Responses of MFIs' Ratios across Euro-Area Countries to a UMP Shock



(continued)

Figure 6. (Continued)

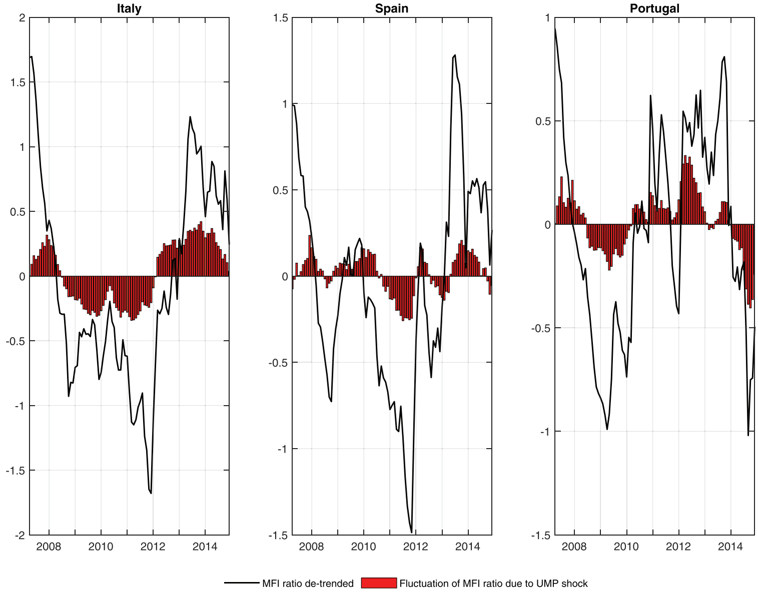


Notes: Impulse responses are derived from a panel of VAR models using a hierarchical Bayesian panel model estimator. The black solid lines are the median impulse responses of the posterior distribution. The shaded areas correspond to the 68 percent posterior credibility bounds. The thin gray lines refer to the 84 percent and the thin gray dashed lines to the 95 percent credible sets, respectively.

For every crisis country, figure 7 displays the detrended MFIs' domestic government bond holdings ratio and the historical contributions to its variation by shocks to unconventional monetary policy. A positive value signals a stimulating effect on the MFI ratios, whereas a negative value reflects a contracting impact.

Our findings suggest that the variation of banks' domestic government bond holdings ratios was affected by nonstandard monetary policy shocks. The MFI ratios fell below their trend in response to the ECB's unconventional monetary policy measures conducted between 2008 and 2009. Since the aim of these policy measures was to support the banking sector by serving as a lender of last resort, the drop in the ratios may be explained by a rise in total assets. While banks had almost unlimited access to liquidity provided through

Figure 7. Historical Decomposition



Notes: The black solid lines reflect the detrended MFIs' domestic government bond holdings ratio. The bars display the fluctuation of the MFI ratios due to an unconventional monetary policy shock.

open market operations with full allotment, they used the deposit facility as a store of liquidity due to the massive distortions on the interbank money market. Subsequently, the ECB's launch of the SMP seemed to generate a positive effect on the share of domestic government debt in banks' balance sheets between 2010 and 2012, which, however, can be decomposed into two parts. In Portugal, the asset purchases contributed to a rise in the MFI ratio in May 2010, while in Italy and Spain they initially had a contracting effect, which possibly arose because both countries were not included in the first wave of purchases. The asset purchases conducted between August 2011 and February 2012 also comprised Italian and Spanish sovereign bonds. Consequently, the MFI ratios in Italy and Spain started to rise. The ECB's announcement of the OMT program in September 2012 also seemed to stimulate the accumulation of domestic sovereign debt, possibly by removing distortions in government bond markets that were reflected by a severe rise in sovereign risk premiums. However, in Spain the effect appeared with a delay. Finally, the ECB conducted LTROs with extended maturities between 2013 and 2014, started to implement forward guidance, and set a negative interest rate on the deposit facility.[35] In response, especially in Italy and Spain, the variation of the domestic government debt ratios appeared to be stimulated by these measures.

4.5 Sovereign Debt Portfolio Adjustment

So far, our results suggest that euro-area crisis countries' banks increase their holdings of domestic government bonds relative to total assets after unconventional monetary policy shocks. However, this finding does not allow us to infer whether banks indeed increased the home bias of their sovereign bond portfolios. The reason is that the decline in the share of domestic government debt in total assets might purely mechanically reflect a reduction in the stock of outstanding loans and, thus, in total assets, without any active adjustments of the bond portfolio. In this case, the share in total assets of sovereign bonds issued by other euro-area countries and by countries outside the currency union should exhibit a similar increase in response to the monetary shock.

Therefore, we seek to shed some light on the MFIs' decisions in managing their sovereign debt portfolio. To this purpose, we estimate additional models for both groups, i.e., the crisis countries and the core countries, which differ from our benchmark specification in that the MFIs' domestic government bond holdings ratio is replaced by alternative ratios. In particular, we investigate the response to unconventional monetary policy shocks of (i) the holdings of government bonds issued in the rest of EMU relative to total assets, (ii) the ratio of domestic government bonds to those from the rest of EMU, (iii) the holdings of government bonds issued outside the EMU relative to total assets, and (iv) the ratio of domestic government bonds to those issued outside the EMU. Note that due to data availability, when working with banks' holdings of bonds from issuers outside

[35]The ECB introduced its forward guidance on July 4, 2013 when President Mario Draghi stated: "The Governing Council expects key ECB interest rates to remain at present or lower levels for an extended period of time." See also Dell'Ariccia, Rabanal, and Sandri (2018) for a discussion.

the currency union, we have to switch from a monthly to a quarterly frequency.

For the crisis countries, figure 8 summarizes the impulse responses of the alternative MFIs' ratios to an innovation related to unconventional monetary policy.

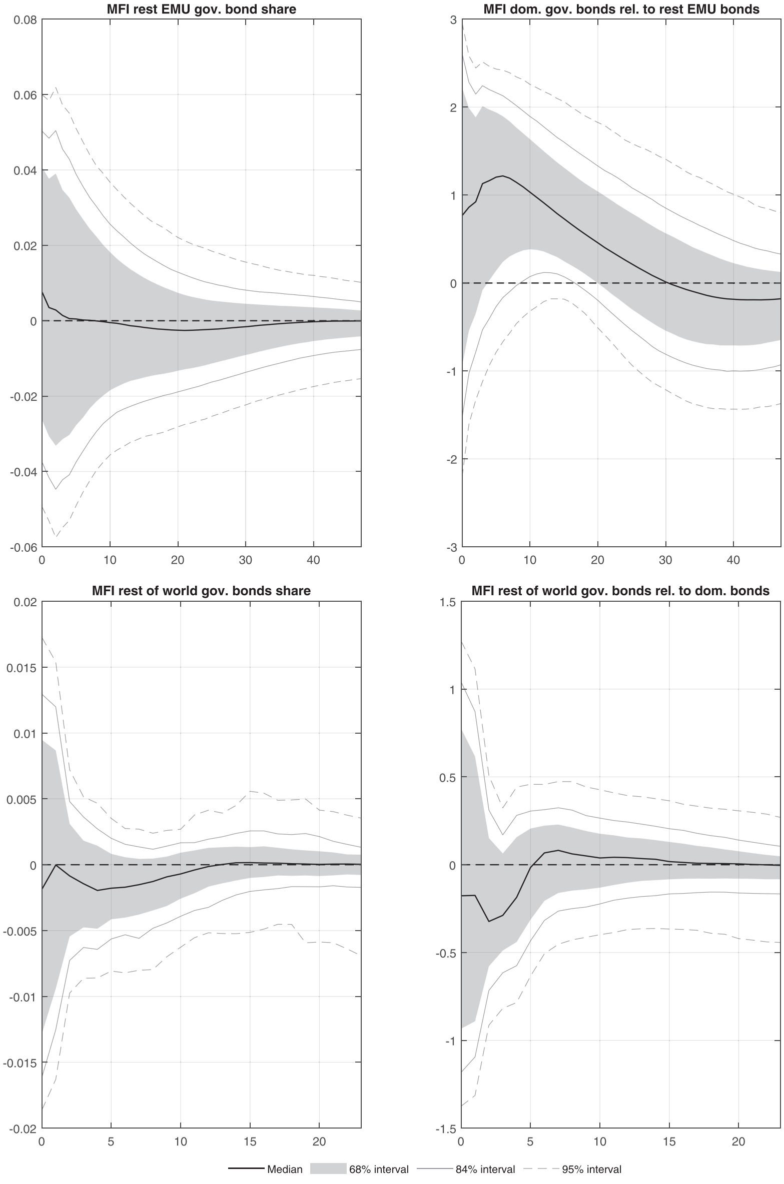
We observe that banks hardly reduced their holdings of rest-of-EMU government bonds relative to total assets. However, the management of the sovereign debt portfolios appears to be home biased, as the ratio of domestic to rest-of-EMU government bond holdings rises. Thus, the adjustment of both ratios suggests that the home bias is triggered by a notable increase in domestic government bonds rather than a reduction in the holdings of rest-of-EMU government bonds. By contrast, the share of banks' holdings of sovereign bonds issued by countries outside the EMU does not change in response to the unconventional monetary policy shock.

In the core countries, banks also seemed to adjust their sovereign debt portfolios after the nonstandard monetary policy disturbances, but in a different manner (see figure 9). Banks' holdings of rest-of-EMU bonds relative to total assets appear to decline. Additionally, the ratio of home-issued sovereign bonds to rest-of-EMU bonds rises. Hence, the rise in the ratio appears to be triggered by a drop in the holdings of rest-of-EMU bonds sovereign bonds rather than a rise in the holdings of domestic government bonds. By contrast, the holdings of sovereign bonds issued by countries outside the EMU seem not to be affected by disturbances related to unconventional monetary policy.

4.6 Extended Period: 2007–18

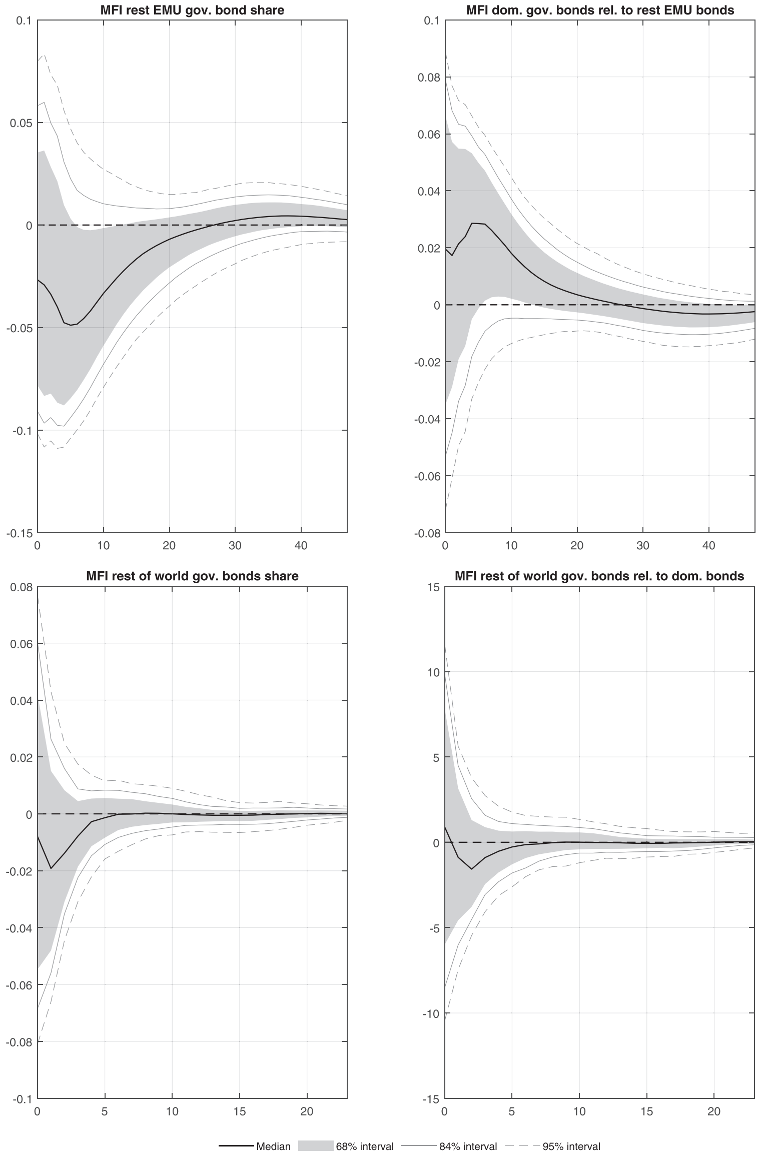
Note that we have excluded the ECB's APP from our sample because the program that started in 2015 was expected (Boeckx, Dossche, and Peersman 2017). The conditions of APP were reported in the press, following the various announcements by the ECB Executive Board members, which summarized the details regarding the conduct of the program (Gambetti and Musso 2017). In particular, the ECB announced the precise volume of the monthly asset purchases, the composition of these purchases regarding the securities issued by euro-area governments, and the maturity of the program.

Figure 8. Reaction of the Crisis Countries' Alternative MFI Ratios to a UMP Shock



Notes: MFIs' rest-of-EMU government bond share reflects the banks' holdings of rest-of-EMU government bonds relative to total assets. MFIs' rest-of-world government bond share reflects the banks' holdings of outside EMU government bonds relative to total assets. The black solid lines are the median impulse responses of the posterior distribution. The shaded areas correspond to the 68 percent posterior credibility bounds. The thin gray lines refer to the 84 percent and the thin gray dashed lines to the 95 percent credible sets, respectively.

Figure 9. Reaction of the Core Countries' Alternative MFI Ratios to a UMP Shock



Notes: MFIs' rest-of-EMU government bond share reflects the banks' holdings of rest-of-EMU government bonds relative to total assets. MFIs' rest-of-world government bond share reflects the banks' holdings of outside EMU government bonds relative to total assets. The black solid lines are the median impulse responses of the posterior distribution. The shaded areas correspond to the 68 percent posterior credibility bounds. The thin gray lines refer to the 84 percent and the thin gray dashed lines to the 95 percent credible sets, respectively.

Latter changes to the program were communicated.[36] Overall, the use of impulse response analysis might not be appropriate under these conditions because, due to a lack of surprises, the effects of an unexpected shock can hardly be identified.

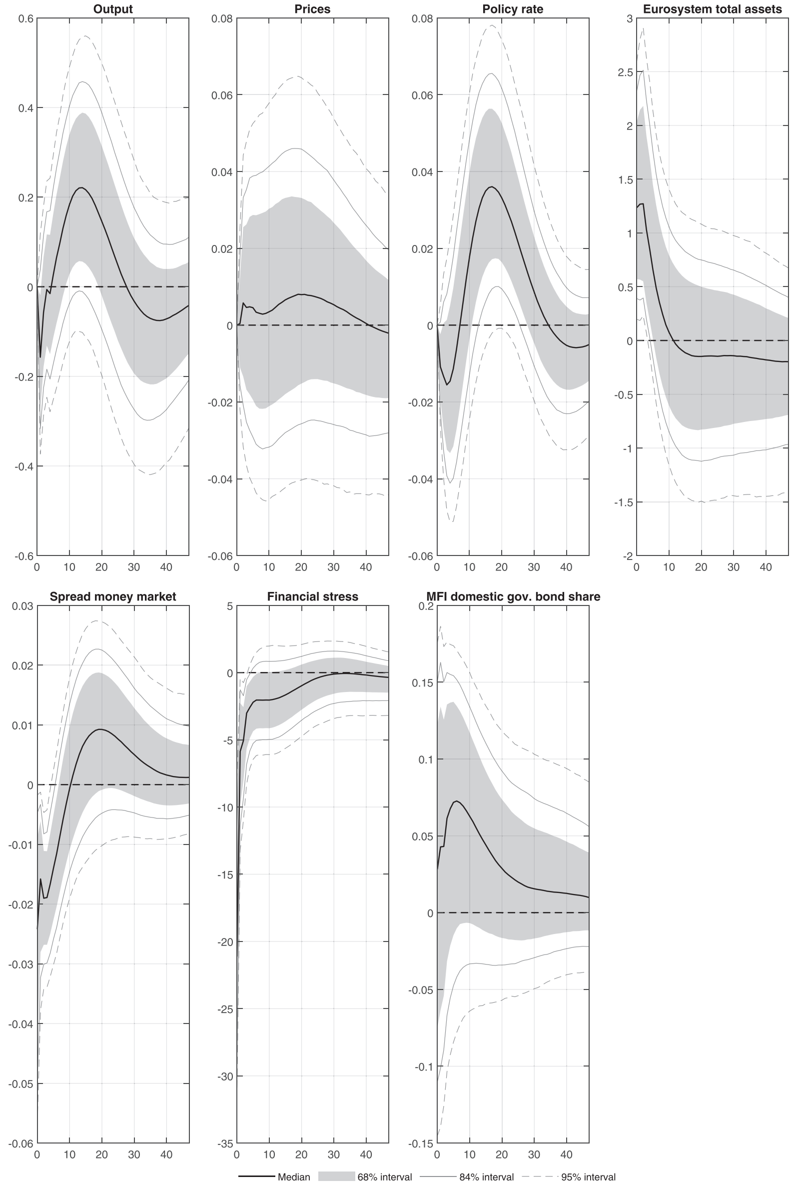
Despite the exception to embedding the episode of the ECB's APP in an impulse response analysis, we proceed by estimating our benchmark model (1) over the period 2007–18 to gain additional insights into the reaction of the MFIs' domestic government bond holdings ratio to an unconventional monetary policy shock. Referring to table 1, we identify the shock by imposing zero and sign restrictions.

For the crisis countries, figure 10 shows the impulse responses of the model variables to a favorable unconventional monetary policy shock. The results are quite similar to those reported for the period 2007–14. Output initially rises after the shock. The boom of the economy is accompanied by an increase in prices. Banks seem to adjust their domestic sovereign debt portfolios. The MFIs' domestic government bond ratio rises, hardly significantly, however, and with a substantial delay.

The delayed reaction of banks' domestic government bond holdings to innovations triggered by unconventional monetary policy might be explained by the characteristics of the ECB's APP program, whose introduction marked a break in the conduct of unconventional monetary policy measures. In particular, the size of the monthly securities purchases under the APP exceeded all existing purchases programs that had been implemented before. Thus, banks might have taken more time to adjust their domestic sovereign debt portfolio in response to unconventional monetary policy disturbances over the period 2007–18 compared with 2007–14.

[36]Initially, the ECB announced with the start of the APP a monthly amount of securities purchases of €60 billion in January 2015, which were intended to be carried out until September 2016. However, the ECB decided to expand the monthly amount of securities purchases to €80 billion in March 2016, while in April 2017 the pace of monthly purchases was reduced to €60 billion. Furthermore, the APP was expanded to run until December 2018. Finally, the ECB decided in 2018 to reduce the monthly purchases further to €30 billion and €15 billion, respectively. Overall, the ECB's accumulated purchases of sovereign securities amounted to €2,102,048 billion at the end of 2018.

Figure 10. Euro-Area Crisis Countries Impulse Responses to a UMP Shock



Notes: Impulse responses are calculated from the benchmark model that is estimated over the period 2007–18. See notes of figure 2 for further explanations. The black solid lines are the median impulse responses of the posterior distribution. The shaded areas correspond to the 68 percent posterior credibility bounds. The thin gray lines refer to the 84 percent and the thin gray dashed lines to the 95 percent credible sets, respectively.

5. Conclusion

We estimate panel VAR models for euro-area member countries to explore whether the ECB's unconventional monetary policy conducted between 2007 and 2014 induced banks to increase their domestic government bond holdings, thereby possibly promoting the sovereign-bank nexus. We impose sign restrictions on impulse responses to identify innovations related to nonstandard monetary policy. Our findings can be summarized as follows. Euro-area crisis countries' banks shifted their asset portfolios toward domestic sovereign bond holdings in response to expansionary unconventional monetary policy shocks, thus making their balance sheets more sensitive to fluctuations in sovereign risk. Nonstandard monetary policy disturbances explain around 19 percent on average of the variation in the national banking sectors' share of domestic government bonds. Moreover, banks seemed to manage their sovereign debt portfolios actively. While banks' domestic government bond holdings increased relative to total assets in response to an unanticipated unconventional monetary policy loosening, the share of their holdings of sovereign bonds issued in the rest of EMU declined. Thus, the reaction of banks' sovereign debt portfolios can be characterized by a home bias. In contrast, euro-area core countries' banks did not increase their domestic government bond holdings relative to total assets in response to disturbances related to unconventional monetary policy. However, banks rather restructured their sovereign bond portfolios by lowering their holdings of rest-of-EMU government bonds. Furthermore, estimations using the approach by Jarocinski (2010) showed that banks across crisis countries and across core countries responded in a qualitatively similar way to unconventional monetary policy shocks by adjusting their sovereign debt portfolios. While in Italy, Spain, and Portugal the change in the domestic government bond holdings ratio occurred swiftly, in none of the core countries can a significant response of the domestic government bond holdings ratio be found. Finally, our results suggest that the ECB's SMP contributed to an increase in banks' sovereign debt portfolios. In Portugal, the effect was immediately observable after May 2010, while in Italy and Spain it became apparent after February 2012. Moreover, in Italy the announcement of the OMT program in September 2012 appeared to have contributed to a higher domestic government bond holdings ratio. Overall, we conclude that euro-area crisis countries' banks enlarged their exposure to domestic sovereign debt in response to expansionary innovations stemming from unconventional monetary policy. While banks' liquidity position improves by such monetary policy measures, they might also contribute to the undesirable side effect of making the banking sector more vulnerable to sovereign distress.

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