Futures Contract Rates as Monetary Policy Forecasts*

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The prices of futures contracts on short-term interest rates are commonly used by central banks to gauge market expectations concerning monetary policy decisions. Excess returns—the difference between futures rates and the realized rates—are positive, on average, and statistically significant, both in the euro area and in the United States. We find that these biases are significantly related to the business cycle only in the United States. Moreover, the sign and the significance of the estimated relationships with business-cycle indicators are unstable over time. Breaking the excess returns down into risk-premium and forecast-error components, we find that risk premia are countercyclical in both areas. On the contrary, ex post prediction errors, which represent the greater part of excess returns at longer horizons in both areas, are negatively correlated with the business cycle only in the United States.

JEL Codes: E43, E44, E52.

1. Introduction

In order to infer market expectations about the future course of monetary policy, central banks commonly use prices of financial assets and survey data. The former are available at high frequencies, but they also incorporate risk and term premia, which may distort their

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information content in terms of expected future interest rates, while the latter are likely not to be affected by premia but are available at a relatively low frequency. Both measures might be biased estimators of ex post realized interest rates to the extent that they incorporate systematic forecast errors.

Recent studies for the United States have compared the information content of several financial instruments, finding that yield curves and futures contracts on short-term interest rates are good predictors of the future path of monetary policy decisions both in the short and medium term (Piazzesi and Swanson 2004; Gürkaynak, Sack, and Swanson 2006). Nevertheless, another strand of the literature has provided evidence that ex post excess returns—namely, the differences between short-term interest rates implied in the price of Eurodollars futures and the ex post realized spot rates—are, on average, positive and statistically significant (Krueger and Kuttner 1996; Sack 2002; Durham 2003). Recently, Piazzesi and Swanson (2004) have shown that this bias is time varying, countercyclical, and predictable by means of business-cycle indicators. This finding suggests that policymakers should look at adjusted measures of futures rates in order to assess the efficacy of their communication more accurately.

The label "risk premia" is often used in the financial literature to refer to predictable excess returns on the short-term interest rate (Piazzesi and Swanson 2004; Cochrane 2006). However, risk premia and predictable excess returns do not necessarily coincide. For example, in the presence of structural breaks, economic agents may need time to learn about the new environment: in the early stages of this process, previously held beliefs could lead to a long series of errors all in the same direction until forecasters finally learn about the structural break. In this case ex post excess returns may incorporate two predictable components. One is the ex ante risk premium, defined as the difference between the futures rates and the market expectation of future spot interest rates, which is required by investors when they buy or sell the financial contract. The other is a systematic prediction error.

In this respect, this paper reassesses the predictive power of short-term interest rate futures by extending the analysis of Piazzesi and Swanson (2004) along two dimensions. First, we use futures contracts on short-term interest rates in euros and investigate the size and the magnitude of ex post excess returns in the euro area, allowing a comparison with those in the United States. Second, we rely on professional forecast surveys in order to disentangle the risk premium and forecast-error components of ex post excess returns and to study their behavior over the business cycle.

Our empirical investigation reveals that euro-area ex post excess returns are of the same sign and magnitude as those in the United States, but they do not appear to be significantly related to the business cycle. In addition, the relation between excess returns and the business cycle appears to be unstable over time in both areas. This evidence is in contrast with the findings of the recent strand of the literature that studies term-structure models, which suggests that the implied risk premia should be strongly affected by business-cycle fluctuations.

We show that these puzzling results essentially depend on the common assumption that ex post excess returns coincide entirely with risk premia. Our proposed empirical breakdown of ex post excess returns suggests that risk premia are, on average, not significantly different in the United States and in the euro area, and are significantly countercyclical in both areas. Interestingly, the predictive regressions involving risk premia and business-cycle indicators are stable over time. By contrast, ex post prediction errors, which represent the largest fraction of the whole excess return at longer horizons in both areas, are significantly and negatively related to the business cycle only in the United States.

We argue that our excess returns decomposition has important implications for central banks when they assess financial markets' expectations regarding the future path of monetary policy decisions. Even though interest rate futures adjusted for both components provide the best forecast of future spot interest rates, they no longer coincide with financial markets' view. Policymakers should assess markets' expectations about future interest rates by looking at quoted futures rates adjusted by the premia component only, as the ex post prediction error reflects part of the expectations formation process.

The remainder of the paper is organized as follows. In section 2 we describe the data set used in the analysis. In section 3 we provide evidence on the size and predictability of ex post excess returns on short-term interest rates in euros, allowing a comparison with those in dollars. In section 4 we decompose ex post realized excess returns into risk premia and systematic prediction errors and investigate their relation with the business cycle. In section 5 we point out the main implications of our proposed breakdown for policymakers. Section 6 concludes.

2. The Data Set

We define the ex post excess return realized from holding the n-quarter-ahead contract to maturity as

$$x_{t+n}^{(n)} = f_t^{(n)} - r_{t+n}, \tag{1}$$

where $f_t^{(n)}$ denotes the average of the futures contract rates quoted on the first ten days of the last month of quarter t for a contract expiring at the end of quarter $t + n$ and r_{t+n} is the corresponding realized spot interest rate prevailing on the day of expiration of the futures contract.[1]

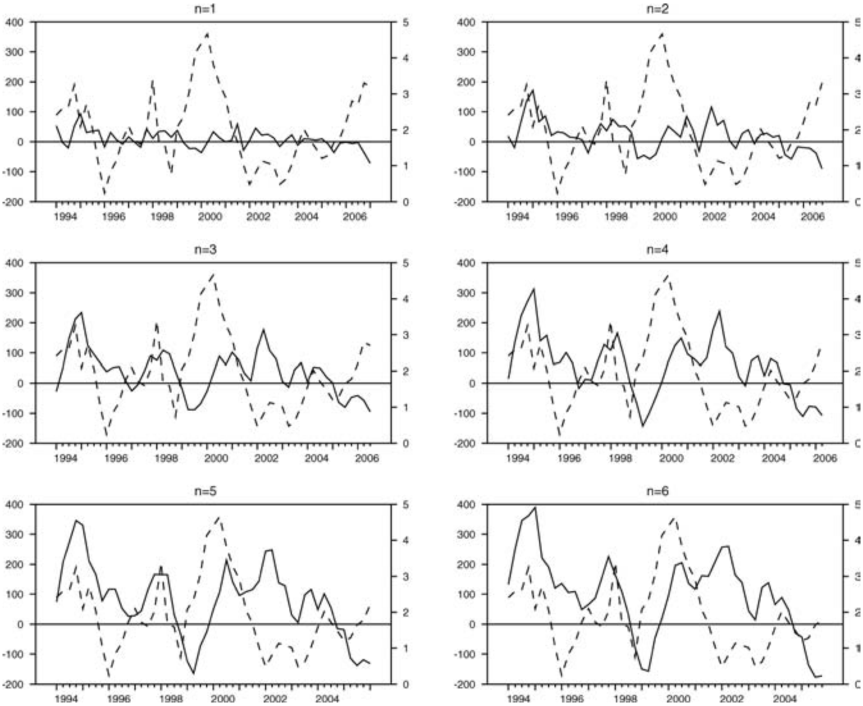
Regarding the euro area, we restrict our attention to futures contracts on short-term interest rates traded on the London International Financial Futures Exchange (LIFFE), which mature two business days prior to the third Wednesday of the delivery month. At each point in time we focus on the first six (unexpired) contracts.[2] The choice of the sample period, 1994–2007, reflects the limited availability of survey data used for the excess returns decomposition, which is the core of our analysis. In particular, for the pre-EMU period (1994:Q1–1998:Q4), we consider futures contracts linked to the British Bankers' Association offered rate (BBA LIBOR) for three-month Eurodeutschmark deposits. The idea is that the institutional features and anti-inflationary objective of the European Central Bank's (ECB's) monetary policy largely resemble those of the German Bundesbank.[3] For the EMU period (1999:Q1–2007:Q1) we

[1] Results do not change significantly using the futures contract rate quoted on the last trading day of quarter t.

[2] By far, the most actively traded futures contracts on three-month deposits are those with delivery in March, June, September, and December.

[3] Buiter (1999) suggests that the ECB adheres to a "priestly" view of central banking in that it adopts "many of the procedures and practices of the old Bundesbank."

Figure 1. Ex Post Excess Returns (Solid Line) and Real GDP Growth (Dashed Line) in the Euro Area



Notes: The sample period is 1992:Q1–2007:Q1. Ex post excess returns are measured in basis points.

focus on contracts whose underlying asset is the European Banking Federation's Euribor Offered Rate (EBF Euribor) for three-month euro deposits. For the United States we compute the ex post excess returns using futures contracts on three-month LIBOR Eurodollar deposit rates, which are quoted on the Chicago Mercantile Exchange.

Figure 1 plots the time series of the ex post realized excess returns on futures contracts in euros expiring up to six quarters ahead. Three basic features emerge. First, independently from the forecasting horizon, these returns are generally positive, suggesting that futures rates are, on average, higher than ex post realized spot rates. Second, they increase with the forecast horizon, consistently with the view that agents demand larger term premia on contracts with longer expiration dates. Third, they move significantly over time (see also Piazzesi and Swanson 2004).

3. Reassessing Ex Post Excess Returns

3.1 Constant Excess Returns

We start our analysis by checking whether futures contract rates are unbiased predictors of spot short-term interest rates. To this end, we follow Piazzesi and Swanson (2004) and regress the computed ex post excess returns on a constant term

$$x^{(n)}_{t+n} = \alpha^{(n)} + \epsilon^{(n)}_{t+n} \tag{2}$$

for the forecast horizons $n = 1, 2, 3, \ldots, 6$ quarters and test in each equation whether the estimated coefficients $\alpha^{(n)}$ are different from zero.

In the absence of arbitrage opportunities, this analysis is also considered a test of the validity of the (pure) rational-expectations hypothesis—namely, that futures contract rates are, on average, equal to the expected spot interest rates.[4] We notice that in the financial literature (Fama 1984; Campbell and Shiller 1991; Campbell 1995) the validity of this hypothesis has also been tested by running predictive regressions of the type

$$r^{(n)}_{t+n} = \alpha^{(n)} + \beta^{(n)} f^{(n)}_t + \epsilon^{(n)}_{t+n} \tag{3}$$

and performing the joint test of the null hypothesis that $\alpha^{(n)} = 0$ (zero-mean term premia) and $\beta^{(n)} = 1$ (no time-varying term premia).[5] However, some drawbacks of this second approach have been recently stressed. First of all, standard errors in regressions of this

[4]In the weaker version of the forward-rate expectation hypothesis, the constant term is allowed to be nonzero.

[5]Interestingly, Gürkaynak, Sack, and Swanson (2006) find that the hypothesis that $\beta = 1$ cannot be rejected for a number of U.S. market interest rates. This evidence, they say, suggests only that the time-varying excess returns are not correlated enough with the ex post spot interest rates spreads to drive the estimated coefficients far from one. It does not rule out the possibility that they are correlated with other variables, such as business-cycle indicators.

type are typically large enough that the expectations hypothesis cannot be rejected, as regression tests are not powerful enough to distinguish between the expectations hypothesis and alternative hypotheses in a sample of the length considered here (Kim and Orphanides 2005). Moreover, equation (3) may raise concerns regarding spurious correlation among variables, insofar as spot interest rates and futures contract rates are nonstationary variables. Although the results could be strongly sample dependent, there is some evidence that various international nominal short- and long-term interest rates may contain a unit root in the levels of the series (e.g., Rose 1988; Rapach and Weber 2004).[6]

Results for the estimated coefficients of equation (2) are summarized in table 1, where standard errors are computed by means of the Newey-West heteroskedasticity and autocorrelation-consistent procedure, in order to take into account the futures contracts overlapping. In the euro area the average ex post realized excess returns are significantly positive over the sample period, ranging from about 10 basis points at the one-quarter horizon to 100 basis points at the six-quarter horizon.

A corresponding analysis for the United States suggests that ex post excess returns have likewise been significantly positive and also slightly larger than those obtained for the euro area, ranging from about 20 basis points at the one-quarter horizon to 110 basis points at the six-quarter horizon.[7]

3.2 Time-Varying Excess Returns

Relying on previous studies for the U.S. Treasury market (Fama and Bliss 1987; Cochrane and Piazzesi 2002) and, more recently, for quoted futures rates (Piazzesi and Swanson 2004), we assess whether

[6] In order to deal with nonstationarity, the validity of the expectations hypothesis is usually tested by subtracting the current level of spot rates or first-differencing the variables in equation (3) (Jongen, Verschoor, and Wolff 2005; Gürkaynak, Sack, and Swanson 2006).

[7] In annualized terms, excess returns in the euro area range from 34 basis points at the one-quarter horizon to 68 basis points at the six-quarter horizon; in the United States they range from 73 to 75 basis points. In the sample period 1985:Q1–2005:Q4, Piazzesi and Swanson (2004) find that the average annualized excess returns range from 60 basis points at the one-quarter horizon to 100 basis points at the six-quarter horizon.

Table 1. Constant Excess Returns

Euro Area						
n	1	2	3	4	5	6
$\alpha^{(n)}$	8.4** (4.4)	20.5** (9.8)	37.7** (16.8)	59.1** (23.3)	80.7** (28.9)	102.2** (32.9)
United States						
$\alpha^{(n)}$	18.3** (6.0)	33.3** (14.7)	51.7* (25.5)	73.6** (34.5)	93.6** (42.8)	112.2** (49.6)

Notes: The sample period is 1994:Q1–2007:Q1. Ex post excess returns are measured in basis points. Predictive regressions are estimated by OLS. Newey-West standard errors are reported in parentheses. * denotes significance at the 10 percent confidence level; ** denotes significance at the 5 percent level.

the term structure of interest rates implied in futures contracts in euros is also characterized by time-varying and predictable excess returns. The predictability of excess returns is explored by running the following regressions,

$$x_{t+n}^{(n)} = \alpha^{(n)} + \beta^{(n)} z_t + \gamma^{(n)} f_t^{(n)} + \epsilon_{t+n}^{(n)}, \tag{4}$$

which involve a business-cycle indicator observable at time t—namely, z_t—and the level of the futures rate itself. Under the assumption that excess returns can be interpreted as risk premia, their predictability using business-cycle indicators finds theoretical foundation in standard asset pricing models (Cochrane 2006), while the broader specification in (4), which includes the futures rate as an additional regressor, essentially relies on the recent strand of the financial literature that uses the affine structure to model the yield curve and the price of risk. These studies typically employ Gaussian affine term-structure models in which time-varying risk premia depend on two latent factors usually identified, respectively, with the level of the short-term interest rate and the slope of the yield curve. The significant relationship between the yield curve and observable state variables reflecting business-cycle fluctuations has been amply documented in Ang and Piazzesi (2003), Ang, Dong, and Piazzesi (2004), Rudebusch and Wu (2004), Ang, Piazzesi, and

Wei (2006), Hördal, Tristani, and Vestin (2006), and Pericoli and Taboga (2006).[8]

Results for the euro area are reported in the top part of table 2 and refer to two business-cycle indicators. For each maturity, the first column shows the estimated coefficients obtained using the annual growth rate of real GDP, which is commonly considered the most natural proxy for the business cycle. Because official real GDP data are released with a lag and frequently revised, there may be significant differences between the data used in the regression and the data available to market participants at the time contract prices were settled. To avoid this problem, we perform real-time predictive regressions using real GDP lagged one quarter and alternative business-cycle indicators. In particular, we use indices from the European Commission's survey of manufacturing industry, household consumption, construction, and retail trade. In order to select a narrower set of variables from the large volume of available survey data, we performed a preliminary cross-correlation analysis at business-cycle frequencies between each of them and real GDP. Among the variables with greater contemporaneous correlation, we find that "employment expectations for the months ahead" in manufacturing industry has the best properties in terms of significance and goodness of fit in regression (4).[9] As the survey is available at monthly frequency, in our quarterly regressions we include the data for the second month of the quarter considered, in order to avoid the use of data not available when agents form their expectations. Moreover, in order to compare the results obtained with different variables and between the two areas, we normalize the regressors to have zero mean and unit variance. Excess returns on futures contracts in euros do not appear to be significantly related to the business cycle.

Table 2 allows us to compare the predictability of excess returns in the two areas in the same sample period. For the United States, we

[8]For a survey, see Diebold, Piazzesi, and Rudebusch (2005).

[9]The contemporaneous correlation of this variable with real GDP at business-cycle frequencies is 0.6. We also run regressions including simultaneously two or more business-cycle indicators and involving one or more estimated common factors obtained from a dynamic factor model based on all the considered business-cycle indicators. Results in terms of goodness of fit are not better than those obtained with employment expectations. The results obtained with other survey data are available from the authors upon request.

Table 2. Time-Varying Excess Returns

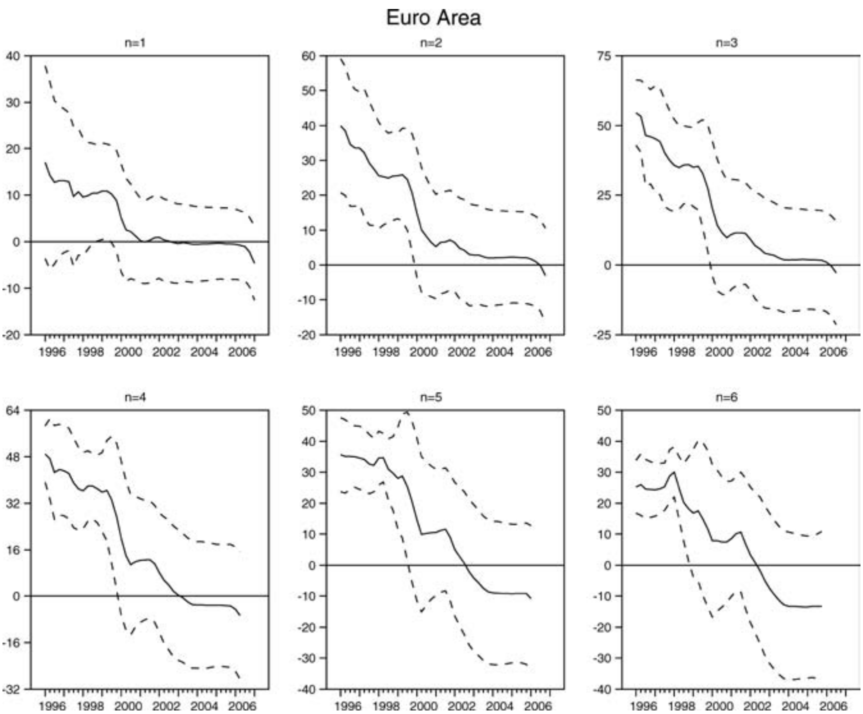
	$n = 1$		$n = 2$		$n = 3$		$n = 4$		$n = 5$		$n = 6$	
Euro Area												
Constant	8.4** (3.8)	8.4** (4.1)	20.5** (8.1)	20.5** (8.7)	37.7** (13.0)	37.7** (14.0)	59.1** (16.3)	59.1** (17.3)	80.7** (18.5)	80.7** (19.2)	102.2** (19.8)	102.2** (20.2)
RGDP	−13.5** (4.1)	—	−18.2 (11.1)	—	−20.0 (17.6)	—	−22.5 (18.5)	—	−22.5 (16.7)	—	−20.8 (14.8)	—
E(empl)	—	−4.7 (4.6)	—	−3.0 (9.2)	—	−2.8 (13.9)	—	−6.7 (15.3)	—	−10.7 (15.1)	—	−13.2 (14.3)
Future	16.7** (4.9)	10.4** (4.6)	32.6** (13.0)	22.9** (9.6)	52.2** (17.5)	41.8** (12.7)	79.0** (17.5)	68.3** (13.7)	100.9** (16.7)	91.4** (14.4)	115.9** (16.8)	108.0** (15.8)
R^2	0.22	0.09	0.24	0.16	0.32	0.27	0.46	0.43	0.57	0.55	0.63	0.62
United States												
Constant	18.3** (5.6)	18.3** (5.7)	33.3** (12.9)	33.3** (12.9)	51.7** (21.4)	51.7** (19.2)	73.6** (28.0)	73.6** (21.8)	93.6** (33.4)	93.6** (22.8)	112.2** (37.3)	112.2** (24.2)
RGDP	−14.7* (8.2)	—	−35.2** (14.1)	—	−57.9** (20.8)	—	−76.5** (25.7)	—	−91.5** (30.5)	—	−103.2** (32.9)	—
NFP	—	−22.8** (10.6)	—	−73.8** (23.0)	—	−127.2** (28.9)	—	−177.4** (27.2)	—	−211.8** (24.6)	—	−224.6** (24.9)
Future	11.6* (6.8)	22.8** (10.7)	30.8** (12.1)	74.5** (23.4)	54.7** (18.1)	131.2** (28.8)	81.5** (24.1)	189.5** (30.1)	109.2** (28.5)	234.9** (31.1)	134.6** (32.9)	261.5** (32.9)
R^2	0.06	0.04	0.12	0.19	0.18	0.33	0.24	0.48	0.31	0.61	0.37	0.66

Notes: The sample period is 1994:Q1–2007:Q1. RGDP is real GDP growth rate. E(empl) is employment expectations for the months ahead from the industrial survey by the European Commission. NFP is the growth rate of nonfarm payrolls. Ex post excess returns are measured in basis points. All predictive variables are standardized. Predictive regressions are estimated by OLS. Newey-West standard errors are reported in parentheses. * denotes significance at the 10 percent confidence level; ** denotes significance at the 5 percent confidence level.

use as business-cycle indicators the annual growth of real GDP and the real-time year-on-year change in nonfarm payrolls. In this case, our estimates confirm the results obtained by Piazzesi and Swanson (2004) for the sample period 1985–2005. The slope coefficients are, in general, highly significant and negative, and their size increases with the forecast horizon. However, some concerns may arise with these estimates.

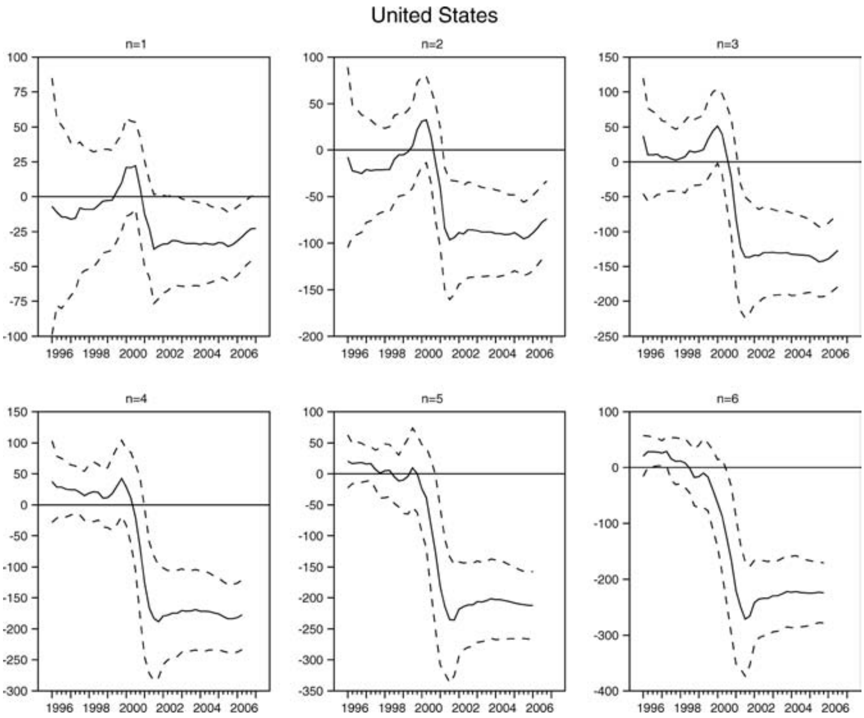
A first issue is the stability of the estimated coefficients. In figure 2 we plot the recursive estimates of coefficients of the business-cycle indicator used in equation (4). Interestingly, the coefficients decreased significantly over time both in the euro area and in the United States. In particular, we cannot exclude that the coefficients

Figure 2. Recursive Coefficients for the Business-Cycle Indicator



(continued)

Figure 2. (Continued)



Notes: Recursive least-squares estimates. The initial estimate is obtained using the sample 1994:Q1–1996:Q1. Employment expectations for the months ahead are used in predictive regressions for the euro area. Nonfarm payrolls are used in predictive regressions for the United States. Dotted lines represent the two standard-error bands around the estimated coefficients.

were positive in the period 1994–2000 and became negative afterwards. The cumulative sum (CUSUM) tests for overall stability of the estimated regressions show significant departures of the computed test statistics from their expected value, thus providing evidence for the presence of parameter or variance instability in the predictive regressions (figure A1 in the appendix).

Another important concern is that excess returns may be nonstationary in the sample period. To the extent that the regressor variables are also nonstationary, the interpretation of the previous estimated predictive regressions may prove erroneous. In table 3 we investigate the time-series properties of the variables used in the

Table 3. Unit-Root Test

	Euro Area		**United States**	
	No Trend	**Linear Trend**	**No Trend**	**Linear Trend**
$x_t^{(1)}$	−3.687**	−5.188**	−2.734**	−3.057*
$x_t^{(2)}$	−3.106**	−3.677**	−2.828**	−3.103*
$x_t^{(3)}$	−3.014**	−3.620**	−1.735*	−2.041
$x_t^{(4)}$	−2.613**	−2.983**	−1.723*	−1.928
$x_t^{(5)}$	−2.726**	−3.067**	−1.591	−1.704
$x_t^{(6)}$	−2.573**	−2.960**	−1.382	−1.914
$f_t^{(1)}$	−1.395	−2.027	−1.639	−2.568
$f_t^{(2)}$	−1.671	−2.368	−1.271	−1.411
$f_t^{(3)}$	−1.446	−1.881	−1.385	−1.630
$f_t^{(4)}$	−1.536	−2.098	−1.477	−1.848
$f_t^{(5)}$	−1.580	−2.307	−1.568	−2.077
$f_t^{(6)}$	−1.520	−2.410	−1.634	−2.287
z_t	−1.180	−2.346	−1.807	−2.109

Notes: The sample period is 1994:Q1–2007:Q1. The lag order p has been selected using a Schwarz information criterion with the maximum lag length of 8. * denotes the rejection of the null hypothesis at the 10 percent level; ** denotes the rejection of the null hypothesis at the 5 percent confidence level.

predictive regressions by means of the modified augmented Dickey-Fuller test (DF-GLS) for unit root (Elliot, Rothenberg, and Stock 1996).[10]

While excess returns on futures contracts in euros appear to be stationary at all maturities, for those in dollars we cannot reject the hypothesis that they contain a unit root, at least at horizons longer than two quarters. Strong evidence of nonstationarity is also found for futures rates in both areas, while for the business-cycle indicators the evidence is less clear-cut and needs to be treated with

[10]In order to discriminate whether the variables of interest are stationary around a deterministic trend, we also show the results by including in the test regression both the constant term and a linear trend.

caution because of the relatively low power of tests in small samples. These findings suggest that the significant relation between excess returns and the business cycle in the United States may simply reflect a common long-run trend but not short-run co-movements among variables.[11]

To the extent that we interpret excess returns as proxies for risk premia, the results of the previous predictive regressions are puzzling. Why in the overall sample do risk premia behave so differently in the two areas? Why has the relation between the business cycle and the risk premia changed over time?

4. Understanding Excess Returns: A Decomposition

First of all, we argue that the previously estimated regressions provide correct measures of the risk premia only under the crucial assumption that the agents are perfectly rational—namely, that they do not make systematic errors in their predictions.[12] In that case, prediction errors are orthogonal to the information set and the only predictable part of the excess return is the risk premium.

However, the financial literature suggests that prices may differ systematically (at least for a period of time) from what people expected them to be for different reasons: (i) prices reflect information to the point where the marginal benefits of acting on information do not exceed the marginal cost (Fama 1991); (ii) agents may rationally process only a limited amount of information because of capacity constraints (Sims 2003); (iii) even if forecasts are formed rationally, allowing for large interest rate movement with small probability, the forecast will appear biased when judged ex post (the so-called peso problem) (Bekaert, Hodrick, and Marshall 2001); (iv) in a changing environment, agents in the market form expectations

[11]We have also estimated the predictive regressions using techniques that take into account the nonstationarity of time series, such as dynamic OLS (e.g., Stock and Watson 1993), fully modified least squares (e.g., Phillips and Hansen 1990), and the vector error correction model (e.g., Johansen 1991, 1995). We find the long-run relationships between excess returns and predictive variables to be significant at horizons longer than one quarter.

[12]The concept of rational expectations, as described in Sargent (1986) asserts that outcomes should not differ systematically (i.e., regularly or predictably) from what people expected them to be.

by learning from past experience (Timmermann 1993) or they are subject to irrational exuberance (Shiller 2000).[13]

In all these cases, ex post excess returns realized from holding the n-quarter-ahead futures contract to maturity may embody two predictable components:

$$x_{t+n}^{(n)} = \theta_t^{(n)} + \sigma_{t+n}^{(n)}, \tag{5}$$

where

$$\theta_t^{(n)} = f_t^{(n)} - E(i_{t+n}|I_t) \tag{6}$$

and

$$\sigma_{t+n}^{(n)} = E(i_{t+n}|I_t) - i_{t+n}. \tag{7}$$

The first component, $\theta_t^{(n)}$, is the ex ante risk premium, defined as the difference between the futures rates and the market expectation of future spot interest rates, conditional on the information set available to the agents at time t. The second component, $\sigma_{t+n}^{(n)}$, is the ex post prediction error made by market participants in forecasting future spot rates and is measured as the difference between the conditional expectation on future interest rates and ex post realized spot rates. As in absence of perfect rationality this second component may be (at least in the short run) systematically different from zero, ex post excess returns can differ substantially from risk premia.

As a proxy for market expectations, $E(i_{t+n}|I_t)$, we consider the mean of short-term interest rate forecasts from the Consensus Forecast survey. This survey has the advantage of providing a long time series on a quarterly basis regarding expectations on future short-term interest rates at horizons up to eight quarters ahead.

The use of survey forecasts may raise concerns for several reasons. The most important one in our context is that, in principle, survey respondents may just use the unadjusted futures contract rates in

[13]There is growing empirical evidence, based mainly on survey data, that the perfect-rationality assumption is violated for expectations on many macroeconomic and financial variables and for many industrialized countries, including the United States and members of the EMU (e.g., Froot 1989; Gourinchas and Tornell 2004; Jongen, Verschoor, and Wolff 2005; Bacchetta, Mertens, and van Wincoop 2006).

order to provide their own forecasts on future spot short-term interest rates. In this case, the forecast would also incorporate the premia component, and the ex post forecast error would be observationally equivalent to the original excess return. Since most of the respondents to the Consensus Forecast survey are professional forecasters who work for institutions operating in the financial markets, even though they may differ from people operating directly in the market, it is likely that they share their information. Therefore, it seems reasonable to assume that respondents to the survey are able to separate the premium component from the forecast component. This hypothesis is also supported by evidence presented by Kim and Orphanides (2005) for the United States that shows that survey expectations on short-term interest rates based on the Blue Chip Financial Forecast incorporates the premium correction.

The estimates of the average value of the two components are obtained by running the regressions[14]

$$\sigma_{t+n}^{(n)} = \alpha_{\sigma}^{(n)} + \epsilon_{t+n}^{(n)} \tag{8}$$

$$\theta_{t}^{(n)} = \alpha_{\theta}^{(n)} + \eta_{t}^{(n)}. \tag{9}$$

Results are reported in table 4. The estimates show that in the euro area, average risk premia are significant at all forecast horizons and are smaller than the corresponding systematic forecast errors at horizons longer than two quarters. In particular, the ex ante risk premium ranges from about 10 to 35 basis points, while the systematic prediction error is between 0 and 70 basis points (see also figure A2 in the appendix).[15] The former

[14]Consensus Economics receives the answers of the survey the first Friday of the last month of the quarter in which it publishes the results of the survey. Since the risk premia are computed using the averages of the market prices of futures contracts quoted on the first ten trading days of the month in which the quarterly Consensus Forecast survey is published, the information sets of respondents to the Consensus Forecast survey and market operators should not be significantly different. In order to verify that the information sets of market participants are not too different, the predictive regressions have been also estimated using spot data from various days on either sides of the first Friday of the last month of the quarter. The results are robust to this modification.

[15]In annualized terms, the risk premia range from 36 to 23 basis points and prediction errors from about 0 to 45 basis points.

Table 4. Excess Returns Decomposition

Euro Area						
n	**1**	**2**	**3**	**4**	**5**	**6**
$\theta^{(n)}$	9.1** (2.3)	13.6** (4.5)	17.7** (6.3)	24.5** (8.2)	30.4** (9.5)	34.2** (10.6)
$\sigma^{(n)}$	−0.7 (4.4)	6.9 (9.8)	19.9 (16.1)	34.6 (21.4)	50.3** (25.6)	68.0** (29.5)
United States						
$\theta^{(n)}$	12.2** (3.8)	17.6** (5.4)	25.1** (6.6)	32.2** (7.2)	37.9** (8.4)	42.0** (9.0)
$\sigma^{(n)}$	6.0 (5.4)	15.7 (10.0)	26.6* (14.8)	41.5** (18.6)	55.7** (22.0)	70.2** (24.8)
Estimated Coefficients for Risk Premia (tbill3m–LIBOR3m)						
$\phi^{(n)}$	28.6** (2.6)	28.7** (2.7)	28.6** (2.7)	28.3** (2.7)	28.1** (2.8)	28.1** (2.8)
$\gamma^{(n)}$	10.9** (2.0)	11.1** (2.1)	11.3** (2.1)	11.7** (2.2)	11.9** (2.2)	12.2** (2.3)

Notes: The sample period is 1994:Q1–2007:Q1. $\theta_1^{(n)}$ and $\sigma_1^{(n)}$ refer to the subsample period 1994:Q1–1998:Q4; $\theta_2^{(n)}$ and $\sigma_2^{(n)}$ refer to the subsample period 1999:Q1–2007:Q1. Ex ante risk premia and ex post forecast errors are measured in basis points. Predictive regressions are estimated by OLS. Newey-West standard errors are reported in parentheses. * denotes significance at the 10 percent confidence level; ** denotes significance at the 5 percent confidence level.

component accounts for more than 60 percent of the overall predictable excess returns at the two-quarter horizon, for about 50 percent at the three-quarter horizon, and for about 40 percent at longer horizons.

For the United States, the Consensus Forecast survey reports expectations on the three-month Treasury-bill rate, which may differ from the three-month LIBOR because of the existence of different premia (Campbell and Shiller 1991; Cochrane and Piazzesi 2002). Therefore, the ex ante risk premium, $\widehat{\alpha}_{\sigma}^{(n)}$, is obtained by adjusting the Consensus Economics forecast for an estimated time-varying premium

$$PR_t \equiv i_t - tb_t = \phi + \tau x_t + e_t, \tag{10}$$

where i_t is the money-market rate (three-month LIBOR) and tb_t is the three-month Treasury-bill rate.[16] In table 4 we report the results of the nonlinear least-squares joint estimation of the two different premia:

$$\theta_t^{(n)} \equiv f_t^{(n)} - E_t[tb_{t+n}] - PR_t = \alpha_\sigma^{(n)} + \epsilon_{t+n}^{(n)} \quad (11)$$

$$PR_t = \phi^{(n)} + \tau^{(n)} x_t + e_t^{(n)}. \quad (12)$$

Average risk premia in the United States, $\theta_t^{(n)}$, range between 10 and 40 basis points; they are not significantly different from those in the euro area at all horizons and they account for about 50 percent of the overall excess return at the two-quarter and three-quarter horizons and for about 40 percent at longer horizons.[17] Systematic prediction errors started to increase significantly in 2000 (see figure A3 in the appendix), when the Federal Reserve stopped announcing its expected future policy stance ("policy bias"), and returned to the lowest level in 2003, when the Federal Open Market Committee reintroduced a direct indication about its future inclinations, suggesting that the systematic error may be strongly related to the communication strategy of the central bank.

In order to investigate the business-cycle properties and the predictability of the two different components $\theta_t^{(n)}$ and $\sigma_{t+n}^{(n)}$, we report in table 5 the results obtained from the following regressions for both the euro area and the United States:

$$\sigma_{t+n}^{(n)} = \alpha_\sigma^{(n)} + \beta_\sigma^{(n)} z_t + \gamma_\sigma^{(n)} f_t^{(n)} + \epsilon_{t+n}^{(n)} \quad (13)$$

$$\theta_t^{(n)} = \alpha_\theta^{(n)} + \beta_\theta^{(n)} z_t + \gamma_\theta^{(n)} f_t^{(n)} + \eta_t^{(n)}. \quad (14)$$

In both areas risk premia vary significantly along the business cycle. The coefficients of the business-cycle indicators are negative at all horizons and highly significant, and their magnitude increases with the forecast horizon. In periods of faster growth, risk premia in

[16] We use the same premium at all forecast horizons, assuming that $E_t[PR_{t+n}] = PR_t$ for $n = 1, \ldots, 6$.

[17] In order to investigate whether risk premia are, on average, different in the two areas, we run a regression, pooling the data of the two areas, on a constant and a country dummy. The estimated coefficients for the country dummy are not significantly different from zero at the 10 percent confidence level.

Table 5. Time-Varying Risk Premia

	$n=1$		$n=2$		$n=3$		$n=4$		$n=5$		$n=6$	
Euro Area												
Constant	9.1** (1.9)	9.1** (1.9)	13.6** (3.3)	13.6** (2.9)	17.7** (4.3)	17.7** (3.9)	24.5** (5.4)	24.5** (5.0)	30.4** (5.6)	30.4** (5.1)	34.2** (4.3)	34.2** (5.9)
RGDP	−2.0 (1.9)	—	−4.6 (3.8)	—	−8.0* (4.4)	—	−12.4** (5.4)	—	−14.5** (5.8)	—	−13.7** (5.8)	—
E(empl)	—	−3.6* (2.0)	—	−8.4** (3.5)	—	−11.4** (4.0)	—	−15.2** (4.5)	—	−17.0** (4.9)	—	−17.8** (4.9)
Future	7.7** (2.3)	7.6** (2.0)	16.0** (4.3)	15.6** (3.3)	23.5** (5.1)	21.8** (3.6)	32.3** (6.1)	28.8** (4.1)	38.6** (6.1)	34.2** (3.9)	41.0** (6.6)	37.1** (4.4)
R^2	0.14	0.18	0.31	0.40	0.42	0.51	0.48	0.56	0.56	0.63	0.57	0.64
United States												
Constant	12.2** (3.9)	12.2** (3.9)	17.6** (5.4)	17.6** (5.4)	25.1** (6.3)	25.1** (6.0)	32.2** (6.3)	32.2** (5.8)	37.9** (6.7)	37.9** (5.9)	42.0** (6.6)	42.0** (5.7)
RGDP	−4.7 (4.8)	—	−13.1** (6.6)	—	−20.8** (7.6)	—	−23.5** (7.5)	—	−28.0** (7.8)	—	−26.0** (7.5)	—
NFP	—	−10.0** (7.8)	—	−27.7** (10.4)	—	−44.7** (11.1)	—	−49.8** (10.3)	—	−56.0** (9.7)	—	−52.4** (8.8)
Future	10.1** (4.3)	13.5* (7.6)	21.4** (6.0)	34.6** (10.0)	32.9** (6.8)	55.7** (10.6)	41.2** (6.9)	66.6** (9.9)	52.2** (7.0)	78.8** (9.2)	58.9** (6.7)	82.8** (8.3)
R^2	0.28	0.26	0.24	0.27	0.31	0.39	0.43	0.53	0.52	0.64	0.62	0.73

Notes: The sample period is 1994:Q1–2007:Q1. Ex ante risk premia are measured in basis points. Predictive regressions are estimated by OLS. Newey-West standard errors are reported in parentheses. * denotes significance at the 10 percent confidence level; ** denotes significance at the 5 percent confidence level.

the euro area may range between 10 basis points (for the one-quarter horizon) and 40 points (for the six-quarter horizon); in periods of slower (or negative) growth, they are between 20 and 80 basis points. In the United States, risk premia range from 10 to 25 basis points in periods of faster growth and from 25 to 95 basis points in periods of slower (or negative) growth.

The recursive estimates of the risk-premia equation (figure 3) and the corresponding CUSUM tests (figure A4 in the appendix) suggest that the sign and the significance of the estimated relationships between risk premia and the business cycle (and, more in general, of the estimated regression) are stable over time in both areas. Moreover, as shown in table A1 in the appendix, unit-root tests suggest that risk premia are stationary at all horizons considered.

Figure 3. Recursive Coefficients for the Business-Cycle Indicator in Risk Premia Regressions

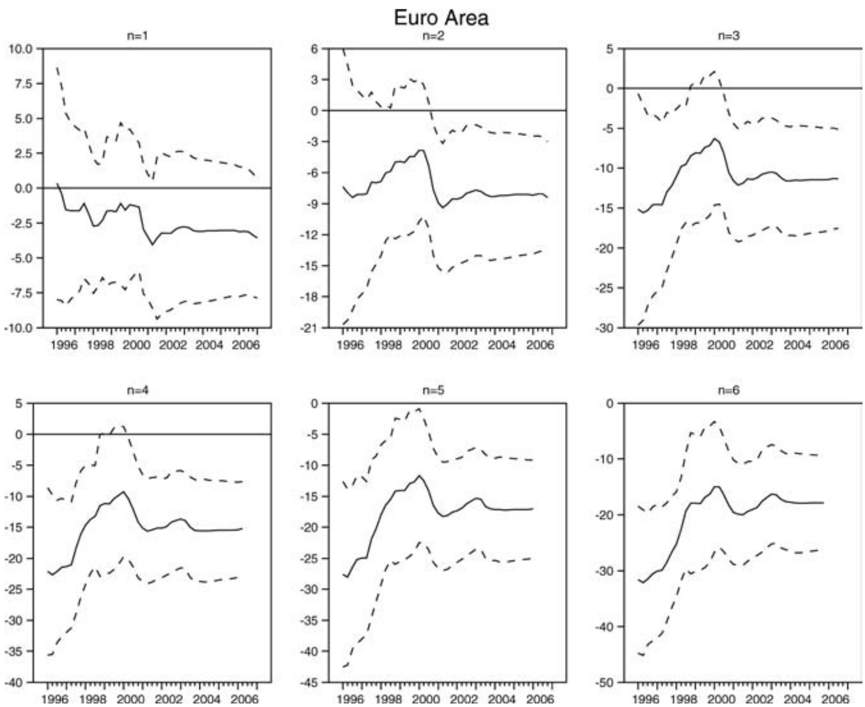


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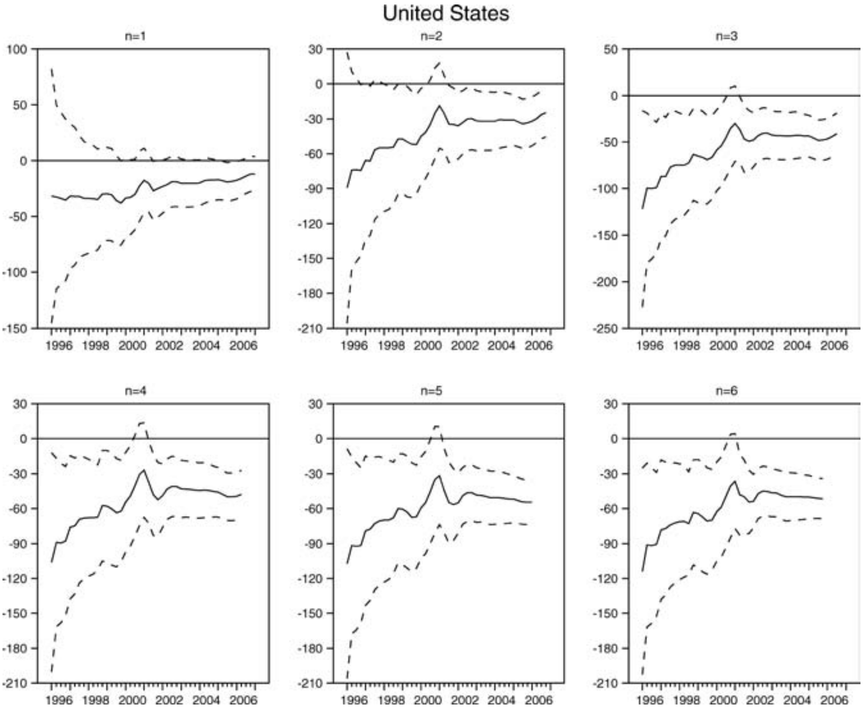
Figure 3. (Continued)



Notes: Recursive least-squares estimates. The first estimate is obtained using the sample 1994:Q1–1996:Q1. Employment expectations for the months ahead and nonfarm payrolls are used respectively in predictive regressions for the euro area and for the United States. Dotted lines represent the two standard-error bands around the estimated coefficients.

As a robustness check for the euro area, we consider the shorter sample period 1999:Q1–2007:Q3 (table 6). The estimates suggest that with stage 3 of the EMU the risk premia have diminished in the euro area but have still remained statistically significant at all forecast horizons. Moreover, the coefficients of employment expectations are negative and highly significant at horizons beyond one quarter and they are of the same magnitude of those obtained in the overall sample.

The predictability of ex post prediction errors along the business cycle is assessed in table 7. The estimated relationships between forecast errors and business-cycle indicators largely resemble those

Table 6. Time-Varying Risk Premia in the Euro Area After the Start of Stage 3 of EMU

n	1	2	3	4	5	6
Constant	6.3**	6.5*	8.1*	10.8**	12.1**	12.4*
	(2.2)	(3.4)	(4.4)	(5.4)	(5.8)	(6.6)
E(empl)	−5.3	−9.4*	−14.0**	−18.9**	−19.7**	−17.0**
	(4.1)	(5.4)	(6.6)	(7.7)	(8.2)	(7.2)
Future	8.1**	12.4**	18.0**	23.3**	27.2**	27.7**
	(3.9)	(3.9)	(5.0)	(5.6)	(6.5)	(5.8)
R^2	0.08	0.13	0.23	0.33	0.39	0.40

Notes: The sample period is 1999:Q1–2007:Q3. Newey-West standard errors are reported in parentheses. * denotes significance at the 10 percent confidence level; ** denotes significance at the 5 percent level.

of total excess returns. In the euro area, employment expectations are not significantly correlated with forecast errors, while in the United States the estimated coefficients are significantly negative at all horizons.[18]

A theoretical analysis of the reasons behind the presence of forecast errors that are predictable and significantly countercyclical only in the United States lies beyond the scope of this paper. However, it should be noted that in the presence of structural changes, economic agents may need time to learn about the new environment: in the early stages of this process, previously held beliefs could lead to systematic biased predictions. To the extent that learning behaviors converge to rational expectations, the prediction bias would be a temporary phenomenon (see, for example, Evans and Honkapohja 2001). Therefore, it is not surprising that in the sample analyzed here the properties of the ex post prediction error are different in the two areas and change over time.

In this respect, a possible explanation for the empirical evidence described in this section regarding prediction errors in the United

[18]Bacchetta, Mertens, and van Wincoop (2006) analyze excess returns and forecast errors in the foreign exchange market and find that, in general, the predictability of the two measures are strictly related, in the sense that a variable that is successfully used in predicting expectation errors is also helpful for predicting the total excess returns.

Table 7. Time-Varying Forecast Errors

	$n = 1$		$n = 2$		$n = 3$		$n = 4$		$n = 5$		$n = 6$	
Euro Area												
Constant	−0.7	−0.7	6.9	6.9	19.9	19.9	34.6*	34.6*	50.3**	50.3**	68.0**	68.0**
	(4.3)	(4.7)	(9.7)	(10.3)	(14.9)	(15.6)	(18.8)	(19.5)	(21.2)	(21.7)	(23.6)	(24.0)
RGDP	−11.5**	—	−13.5	—	−12.0	—	−10.1	—	−8.1	—	−7.1	—
	(4.4)		(12.6)		(19.6)		(21.6)		(20.5)		(19.1)	
E(empl)	—	−1.1	—	5.4	—	8.5	—	8.5	—	6.3	—	4.6
		(5.1)		(10.1)		(15.1)		(16.9)		(17.3)		(17.6)
Future	9.0*	2.7	16.5	7.3	28.7	20.0	46.7**	39.5**	62.4**	57.2**	74.9**	70.9**
	(4.9)	(3.7)	(13.5)	(8.8)	(19.0)	(12.7)	(20.2)	(14.5)	(19.9)	(16.1)	(20.4)	(18.0)
R^2	0.08	−0.03	0.04	0.00	0.07	0.07	0.17	0.17	0.28	0.28	0.34	0.33
United States												
Constant	6.1	6.1	15.7*	15.7*	26.6*	26.6*	41.5**	41.5**	55.7**	55.7**	70.2**	70.2**
	(5.2)	(5.2)	(9.5)	(9.2)	(14.0)	(13.4)	(17.3)	(15.8)	(20.1)	(17.6)	(22.0)	(18.9)
RGDP	−10.2	—	−25.0**	—	−37.5**	—	−53.5**	—	−62.7**	—	−74.0**	—
	(6.5)		(12.2)		(17.6)		(21.4)		(24.2)		(22.0)	
NFP	—	−16.9	—	−57.2**	—	−88.0**	—	−131.4**	—	−155.2**	—	−169.5**
		(11.3)		(20.4)		(28.1)		(31.1)		(31.6)		(31.2)
Future	1.9	13.9	13.9	52.5**	22.4	81.8**	41.1*	127.4**	55.4**	155.4**	69.6**	174.4**
	(6.3)	(11.3)	(12.1)	(20.4)	(17.4)	(28.1)	(21.2)	(31.1)	(23.9)	(31.4)	(25.0)	(30.7)
R^2	0.25	0.15	0.17	0.17	0.16	0.21	0.18	0.30	0.20	0.39	0.25	0.45

Notes: The sample period is 1994:Q1–2007:Q1. Ex post forecast errors are measured in basis points. Predictive regressions are estimated by OLS. Newey-West standard errors are reported in parentheses. * denotes significance at the 10 percent confidence level; ** denotes significance at the 5 percent confidence level.

States may be the following. Throughout the decades of the 1990s, inflation and unemployment were trending down, while productivity was trending up. Forecasters and financial markets had a difficult time picking up on these developments in real time. As a result, they made repeated positive forecast errors in predicting inflation and negative ones in predicting output developments. As a consequence, forecast errors in predicting the future path of interest rates have been relatively small with respect to those realized in the 2000s (see figure A3 in the appendix), consistently with the assumption that the central bank sets interest rates in response to output and inflation (Taylor-type rules). Relatively small prediction errors

Figure 4. Recursive Coefficients for the Business-Cycle Indicator in the Ex Post Forecast Regressions

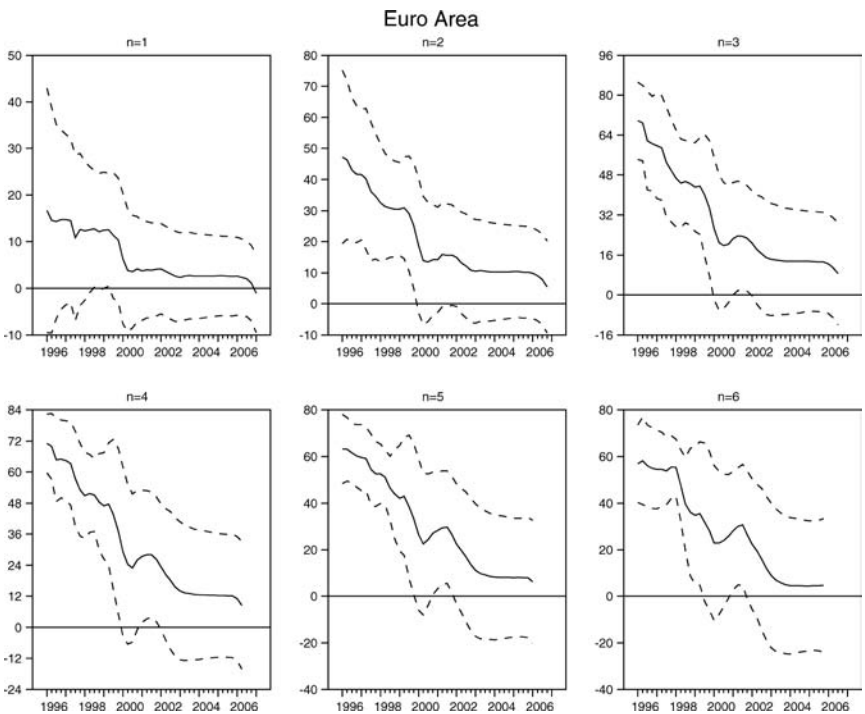


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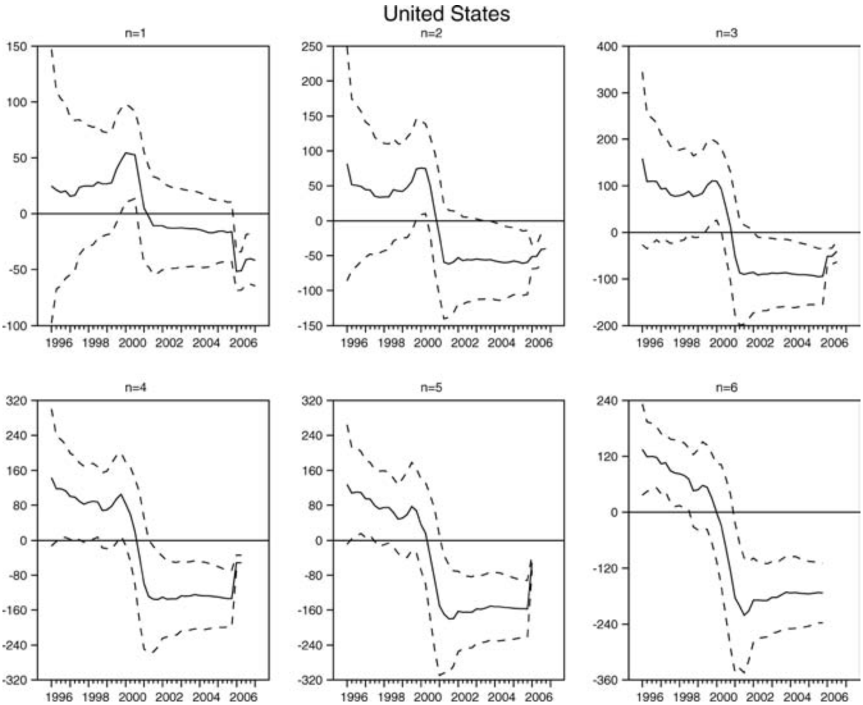
Figure 4. (Continued)



Notes: Recursive least-squares estimates. The first estimate is obtained using the sample 1994:Q1–1996:Q1. Employment expectations for the months ahead are used in predictive regressions for the euro area. Nonfarm payrolls are used in predictive regressions for the United States. Dotted lines represent the two standard-error bands around the estimated coefficients.

were, therefore, associated with relatively high economic growth in the 1990s, coherently with our estimates.[19]

Figure 4 reports the recursive estimates of the coefficients of the business-cycle indicator used in equation (13) and shows that they have significantly decreased over time both in the euro area and in the United States, thus suggesting that the instability observed in the estimates of total excess return reflects the instability of the estimates of the ex-post systematic error (see also figure A5 in the appendix).

[19]We thank an anonymous referee for suggesting this point.

5. Out-of-Sample Forecast Accuracy

Insofar as risk premia and forecast errors are predictable by means of business-cycle indicators, it is interesting to investigate whether gains are achieved in out-of-sample forecast accuracy for short-term interest rates by using adjusted futures rates.

The design of the experiment is based on rolling-endpoint regressions. An initial estimate of risk premia at different horizons is obtained using the sample period 1994:Q1–1996:Q4; we use the estimate to compute a set of out-of-sample forecasts for future interest rates up to six quarters, as follows:

$$i^f_{t+n} = f_t^{(n)} - E_t\left(\widehat{x}^{(n)}_{t+n}\right). \tag{15}$$

We then add a new observation and repeat the forecasting exercise, until the end of the sample period. Overall we collect a set of fifty-eight out-of-sample predictions at each forecast horizon. In table 8 we report the mean error (ME) and the root-mean-squared errors (RMSEs) for (i) futures rates adjusted for time-varying risk premia, (ii) constant-adjusted futures rates, and (iii) futures rates adjusted for time-varying total excess return. We perform a Diebold-Mariano test to check whether the errors obtained under the adjusted predictions are significantly different from their counterparts obtained with unadjusted futures rates.

Unadjusted futures rates perform relatively poorly in both areas. In the euro area the RMSEs of the predictions obtained with the unadjusted futures rates are larger than those obtained from a random-walk model at all horizons beyond three quarters and those obtained from the Consensus Forecast survey at all horizons beyond one quarter. Futures rates adjusted for a constant excess return already produce lower RMSEs at all forecast horizons, even if the gains in forecast accuracy are small and often not significant (RMSE is reduced by about 10 to 25 percent with respect to that obtained with unadjusted futures). Adjusting futures rates for the time-varying risk premia further improves our predictions (by about 10 percent compared with those obtained with constant-unadjusted futures). Finally, adjusting for the time-varying excess return reduces the RMSE with respect to that obtained adjusting only for the risk premia by about 5 to 25 percent at horizons longer

Table 8. Out-of-Sample Forecasts for Short-Term Interest Rates: Summary Statistics

	$n = 1$		$n = 2$		$n = 3$		$n = 4$		$n = 5$		$n = 6$	
	ME	RMSE	ME	RMSE	ME	RMSE	ME	RMSE	ME	RMSE	ME	RMSE
Euro Area												
Random Walk	−3.6	32.5	−7.8	54.5	−13.6	73.1	−19.4	90.0	−24.8	102.6	−30.1	112.4
Consensus	1.6	29.7	5.8	49.5	19.2	72.2	36.0	93.3	53.3	110.8	73.2	128.4
Unadjusted	9.6	25.9	23.4	51.4	42.3	80.6	66.3	113.4	92.5	141.5	118.2	165.5
Constant Adj.	−10.0	23.4	−20.1	45.4	−30.7	68.9	−43.2	94.1	−52.6	112.0	−59.2	122.4
Risk-Premia Adj.	3.2	21.6	9.2	40.5	22.9	61.3	40.6	83.3	53.6	99.0	70.6	113.4
Excess Returns Adj.	−2.1	22.3	−4.8	44.3	−5.9	65.3	2.1	79.1	11.5	85.4	20.7	89.1
United States												
Random Walk	1.5	49.5	2.4	85.4	0.2	116.4	−2.9	146.0	−6.2	172.6	−11.0	194.7
Consensus	9.5	36.7	19.6	68.8	30.9	101.5	47.3	131.2	63.5	156.2	79.8	176.6
Unadjusted	11.2	39.7	27.3	77.1	46.0	115.9	68.9	154.3	91.0	188.7	112.5	217.3
Constant Adj.	1.7	36.9	1.5	72.6	3.4	109.9	2.0	146.4	3.4	178.3	9.4	203.3
Risk-Premia Adj.	1.7	37.4	12.4	72.7	28.1	103.1	48.0	135.0	69.4	161.3	90.6	182.1
Excess Returns Adj.	−0.4	41.7	2.1	71.7	12.2	98.1	16.9	113.0	19.9	120.4	23.2	128.9

Notes: ME is the mean error; RMSE is the root-mean-squared error. Forecast errors are measured in basis points. Employment expectations for the months ahead are used in predictive regressions for the euro area. Nonfarm payrolls are used in predictive regressions for the United States.

than three quarters; however, at shorter horizons there are no significant improvements, thus confirming that in the sample analyzed here the forecast errors are not predictable by means of business-cycle indicators and are, on average, not significant at shorter horizons.

For the United States, adjusting for the time-varying excess returns improves our forecasts by up to 40 percent with respect to unadjusted futures rates, while futures rates adjusted only for the risk premia determine RMSEs between 10 and 40 percent larger than those obtained adjusting for the total excess return at horizons longer than one quarter. In this case, prediction errors are significant and predictable by means of business-cycle indicators.

These results have important implications for central banks. Even if futures rates adjusted for both risk premia and systematic prediction errors are the best predictors of future monetary policy decisions at least at longer horizons, they no longer coincide with financial markets' expectations. Therefore, for a correct assessment of the financial markets' view about future policy decisions, policymakers should use quoted futures rates adjusted only for risk premia, as systematic forecast errors represent part of agents' expectations formation process.

6. Conclusions

In this paper we show that the prices of futures contracts on three-month interest rates are biased forecasts of future short-term interest rates. We also find evidence of large and time-varying excess returns on three-month interest rate futures in the euro area, in line with the results obtained by Piazzesi and Swanson (2004) for the United States. However, unlike those in dollars, ex post excess returns on futures contracts in euros do not appear to be significantly related to business-cycle indicators, while in both areas the sign and the significance of the estimated relationships between excess returns and the business cycle are unstable over time.

We show that ex post excess returns can be divided into two components. The first is the effective ex ante risk premium demanded by investors when they buy or sell the financial contract. The second is an ex post systematic forecast error.

The empirical analysis reveals that the risk premia on futures contracts in euros and in dollars are not significantly different and, interestingly, they are significantly countercyclical in both areas. Moreover, the sign and the significance of the estimated relationships between risk premia and the business cycle turn out to be stable over time.

Finally, we find that the instability observed in the estimates of total excess returns in both areas and the lack of a significative relationship between that variable and business-cycle indicators in the euro area are determined by the instability of the estimates of the ex post systematic-error component.

The policy implication of our findings is that even though futures rates adjusted for both components are better forecasts of future monetary policy actions, in assessing markets' view about future policy decisions, it is better to use futures rates adjusted only by risk premia, as systematic forecast errors are part of agents' expectations.

Appendix. Tables and Figures

Table A1. Unit-Root Test for Risk Premia

	Euro Area		United States	
	No Trend	**Linear Trend**	**No Trend**	**Linear Trend**
$\theta_t^{(1)}$	−6.449**	−7.595**	−3.139**	−4.339**
$\theta_t^{(2)}$	−3.791**	−4.948**	−2.924**	−4.297**
$\theta_t^{(3)}$	−2.729**	−4.267**	−2.422**	−3.980**
$\theta_t^{(4)}$	−2.765**	−4.063**	−2.394**	−3.884**
$\theta_t^{(5)}$	−2.586**	−4.286**	−2.331**	−3.763**
$\theta_t^{(6)}$	−2.228**	−3.973**	−2.435**	−4.109**

Notes: The sample period is 1994:Q1–2007:Q1. The t-statistic of the augmented Dickey-Fuller (ADF) test includes in the test regression deterministic variables and p lagged difference terms of the dependent variable. The lag order p has been selected using a Schwarz information criterion with the maximum lag length of 8. * denotes the rejection of the null hypothesis at the 10 percent level; ** denotes the rejection of the null hypothesis at the 5 percent confidence level.

Figure A1. CUSUM Test of Instability for Excess Returns Regressions

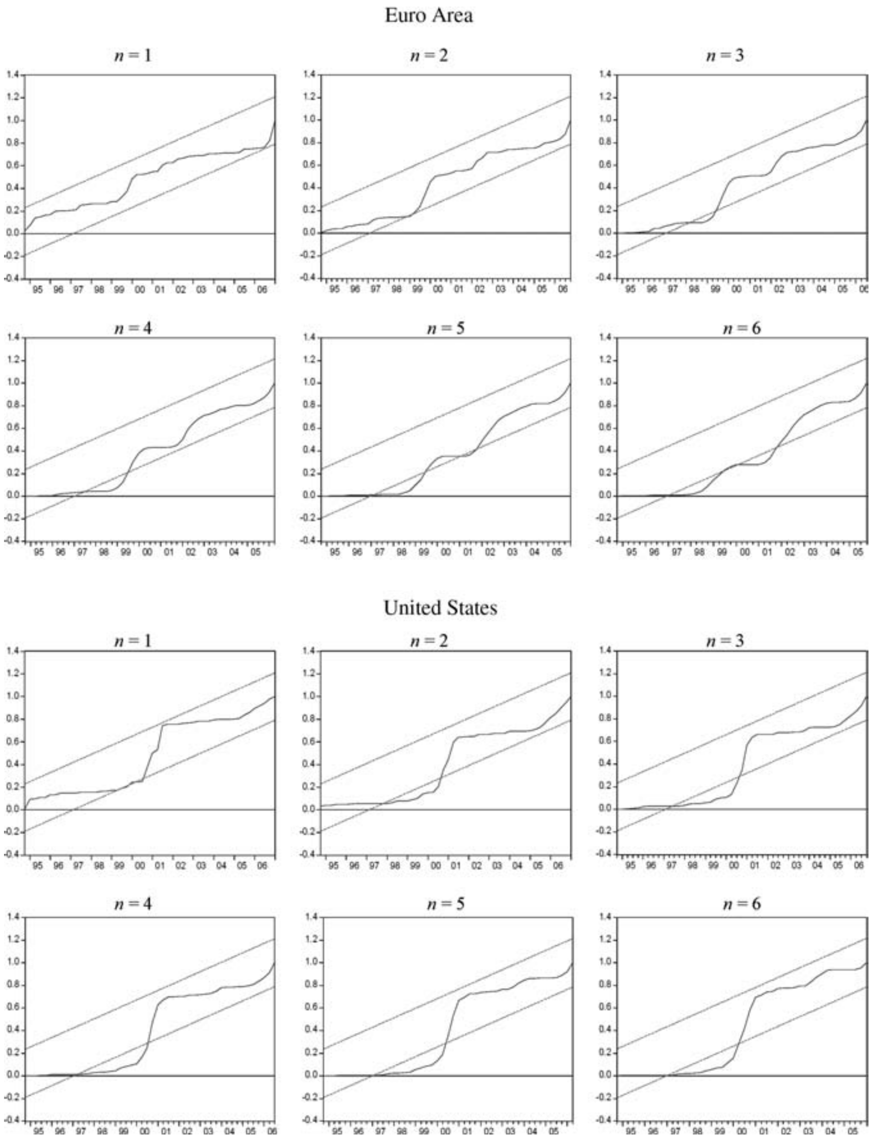


Figure A2. Risk Premia and Forecast Errors in the Euro Area

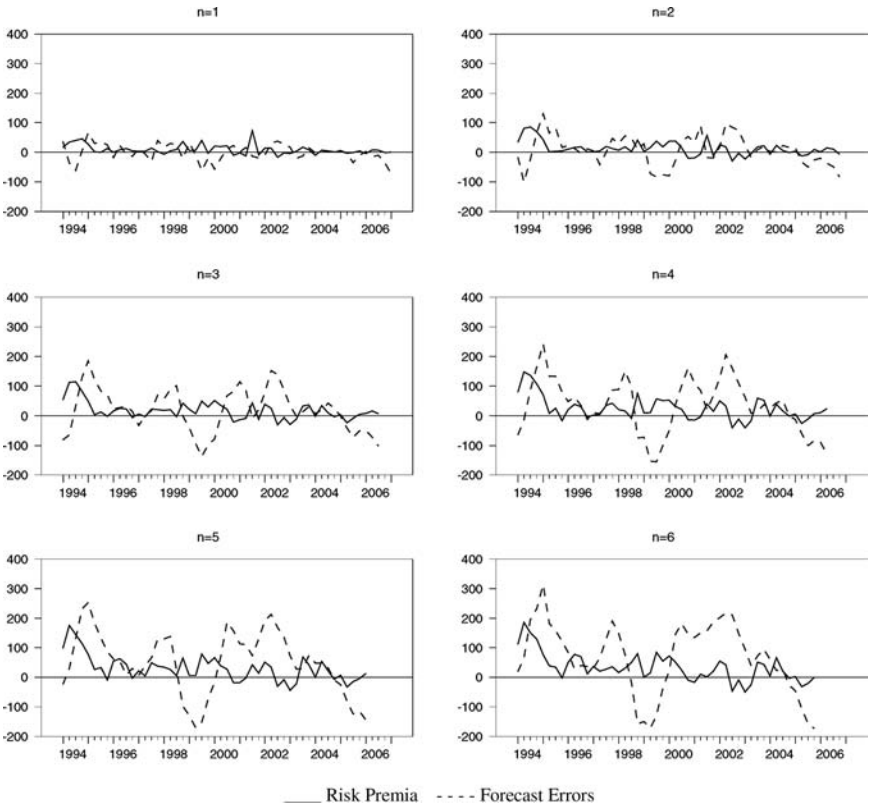


Figure A3. Risk Premia and Forecast Errors in the United States

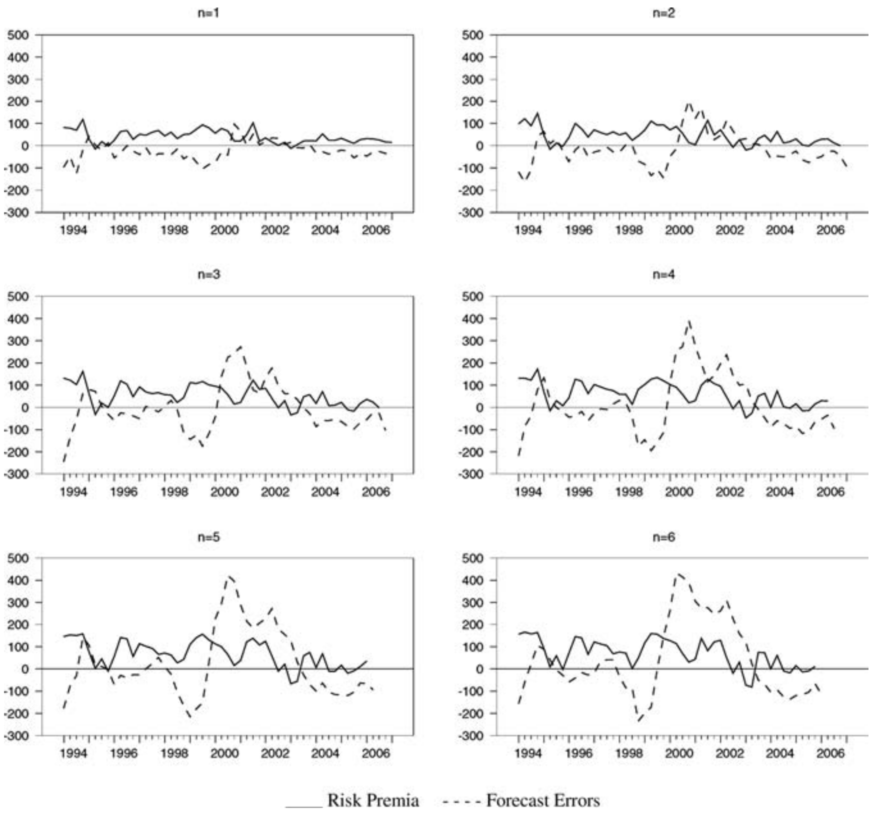


Figure A4. CUSUM Test of Instability for Risk Premia Regressions

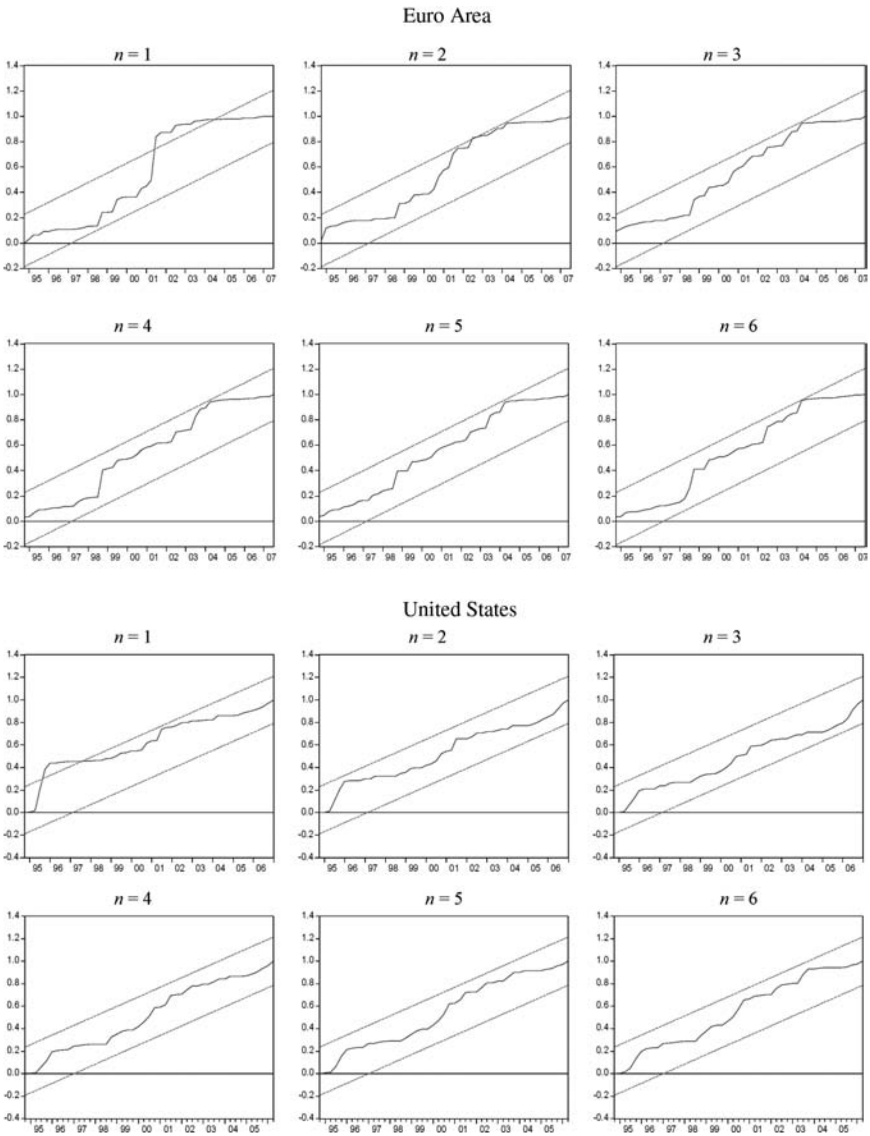
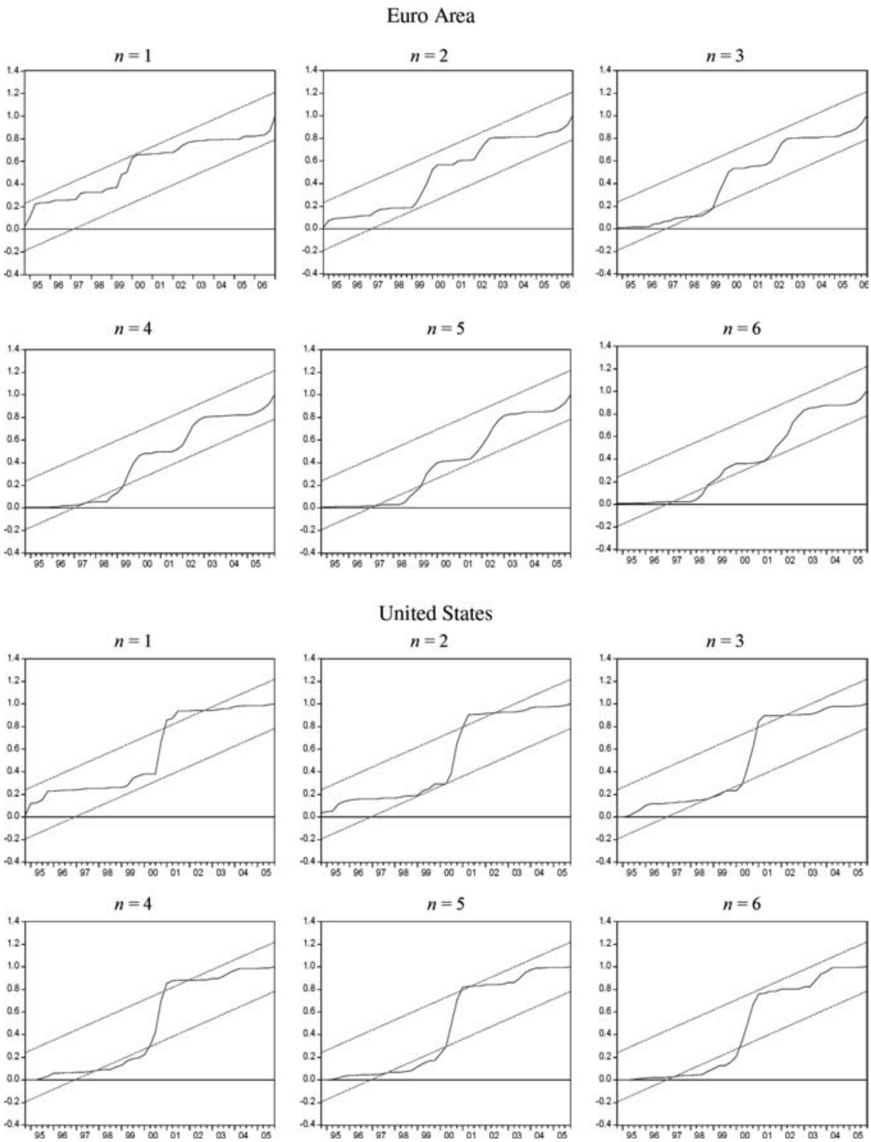


Figure A5. CUSUM Test of Instability for Forecast Errors Regressions



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