Evolving Perceptions of Central Bank Credibility: The ECB Experience

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revised July 2010

Abstract: We present a novel empirical framework that uses high-frequency data to test for persistent variation in market perceptions of the reaction function of the European Central Bank (ECB) during its first six years of operation, from January 1999 to mid-2005. This episode serves as a natural experiment for considering the evolution of central bank credibility since the ECB began in an atmosphere of uncertainty, and controversy, as to its policy stance. We argue that these perceptions significantly evolved over this period, a conclusion based on tests of the effect of economic news announcements on the yield curves in the euro area and on the euro-dollar exchange rate. An increase in the perceived weight on inflation in the policy decisions of the ECB, or a view that inflation targets were lower than previously perceived, occurred in the wake of ECB monetary tightening. In contrast, there is no comparable evidence of any change in market perceptions of the reaction function of the Federal Reserve over this period.

Keywords: Central Banking, European Central Bank, Federal Reserve, inflation, exchange rate, monetary policy, credibility, yield curve.

JEL Classification: F3, E5, E6.

[!] An early version of some of the material in this paper appeared in our unpublished manuscript "Establishing Credibility: Evolving Perceptions of the European Central Bank". We are grateful for the comments provided by Veronica Guerrieri, Richard Clarida, and participants in the International Seminar on Macroeconomics 2010, Amsterdam. The views expressed in this paper are those of the individual authors and do not necessarily reflect the position of the Federal Reserve Bank of New York or the Federal Reserve System. Address correspondences to Linda S. Goldberg, Federal Reserve Bank of NY, Research Department, 33 Liberty St, New York, N.Y. 10045, email: Linda.Goldberg@ny.frb.org, or Michael W. Klein, Fletcher School, Tufts University, Medford MA 02155, email: michael.klein@tufts.edu.

1. Introduction

What are the preferences of a central bank over inflation and output-gap stabilization objectives, and what is its preferred long-run inflation rate? While statements of priorities and goals are important, the credibility of such statements, and the market perception of the policy reaction function of a central bank, play a key role in determining economic outcomes. This point, early on described as the "credibility" of central bank policies, is a standard theoretical result with recent interpretation in the new-Keynesian paradigms.[1] It is also received wisdom among practitioners.

The importance of the market perception of the central bank's acceptable tradeoffs between inflation and output goals as well its specific targets naturally leads to the question of how the market acquires this perception and whether and how it evolves over time. One view is that establishing an appropriate institutional structure is the key element in insulating the monetary authority from political pressure and thereby convincing markets that a central bank has a strong and unvarying aversion to inflation. A second, more dynamic, view focuses on the role that actual policy conduct plays in building the reputation of a central bank. These two different views have distinct implications for the relative importance of the institutional structure of a central bank, as compared to its conduct, for attaining and maintaining its credibility.[2]

A survey of the heads of central banks and prominent monetary economists reflects a belief that the credibility of a central bank is based more on its past actions than on institutional structures that afford it independence by insulating it from political concerns although there is also a consensus that structure matters (Blinder 2000). Empirical research has found that institutional features related to central bank independence are associated with

[1] Seminal contributions on the role of credibility include Kydland and Prescott (1977), Calvo (1978), and Barro and Gordon (1983). More recently, Woodford (2009) discusses optimal monetary policy when there are distortions of private sector beliefs.

[2] In a survey of the heads of 84 central banks, as well as 52 prominent academic monetary economists, Blinder (2000) finds that anti-inflation credibility is considered vitally important and "helps keep inflation low." Blinder (2000) points out that the term "central bank credibility" can mean inflation aversion, incentive compatibility or pre-commitment. He reports that, among these three concepts, "…central bankers identify inflation aversion with credibility far more closely than do [academic] economists." (p. 1424) Using a five-point scale, nearly 90 percent of his central bank respondents identified the concepts "credibility" and "dedication to price stability" as "quite closely related" or "virtually the same," while just over half of the academic respondents replied that these two terms were either "unrelated," slightly related," or "moderately related." In the title and body of this paper, we use the term "credibility" to mean inflation aversion. Theoretical contributions in which credibility is synonymous with inflation aversion include Rogoff (1985, 1987) and Backus and Driffill (1985).

economic performance in cross sections of countries, perhaps because these features indicate the ability of an institution to "tie its hands" and commit to a policy that may cause short-term pain in the pursuit of longer-run gain.[3] There is less evidence, however, as to whether and how the credibility of a particular central bank's policy stance evolves over time in response to the conduct of policy and other related decisions. This is a particularly timely issue. Questions were raised about the commitment of the Federal Reserve to price stability after its response to the financial crisis of 2007-2009. Even more recently, similar concerns were raised about the ECB to its primary mandate of price stability after it introduced a Securities Markets Program (SMP) to purchase euro-area debt securities to "ensure depth and liquidity in those market segments which are dysfunctional" in May 2010.

These episodes raise a relevant question of how to determine whether market perceptions of central bank policies are changing. We show how high-frequency data from asset markets can be used to address this by providing a methodology for tracking the evolution of related market perceptions. Our analysis focuses on the experience of the European Central Bank (ECB) during its early years of operation, which is a natural experiment for this issue.

The ECB offers an example of both structure and conduct aimed towards achieving policy credibility. Its architects were mindful of lessons from economic theory concerning the importance of a structure that provided independence from political considerations.[4] The role of conduct was also clearly apparent. At its inception, the directors of the ECB were acutely aware that their policies were closely scrutinized for indications of general tendencies. This is, of course, a specific example of a more general tendency for relatively large updating of market priors when there is the establishment of a new central bank, an

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[3] For example, Cukierman (1992) analyzes the charters of central banks and shows, in a cross-country panel, that average inflation is lower in countries in which laws afford central banks greater independence. Alesina and Summers (1993) also find cross-country evidence that the level of inflation, as well as its variability, is negatively associated with indicators of central bank independence, but there is no association between central bank independence and real variables. Questions have been raised, however, about whether the *de jure* structure is closely linked to the *de facto* behavior of institutions as in Forder (1999).

[4] Article 108 of the treaty establishing the European Community discusses insulating monetary policy decisions from political influence. More recently, the May 2006 ECB publication "The European Central Bank, the Eurosystem, the European System of Central Banks" states "When performing Eurosystem related tasks, the ECB and the national central banks must not seek or take instructions from Community institutions or bodies, or from any government of an EU country or from any other body. Likewise, the Community institutions and bodies and the governments of the Member states must not seek to influence the members of the decision making bodies of the ECB or of the NCBs [national central banks] in the performance of their tasks." (p. 14)

adoption of new policies (like inflation targeting or extraordinary emergency lending at the time of a crisis), or a change in leadership. (Blinder 2000).[5]

We begin by developing insights about the response of asset prices to inflation news in Section 2. The key point is that these responses reflect market perceptions of the policy reaction function of the central bank, as shown in the work of Gürkaynak, Sack and Swanson (2005) and Gürkaynak, Levin, Marder and Swanson (2005).[6] Their calibration exercises demonstrate the responses of short interest rates, long interest rates, and the yield curve to output and inflation shocks. We build on their work by demonstrating that the patterns in these calibrations are closely tied to the public's perception of the policy reaction function. In particular, we use this model to illustrate the change in the relationship between economic news and the term structure of interest rates with changes in the perceived anti-inflation stance of the policy reaction function parameters. These results also hold for changes in the perceived inflation target.[7]

In Section 3 we propose and implement a novel method for measuring the market's view of evolving central bank "credibility." The method applies newly developed econometric tests for persistent time variation in regression coefficients (from Elliott and Müller, 2006) to high-frequency financial market data.[8] Specifically, we use these tests to explore the evolution of the effects of news announcements on the yield curve for euro area countries for the period beginning January 1999, the time the ECB began its operations, through mid-2005, using hourly data on the term structure of bonds of euro area countries and the United States, as well as the euro-dollar exchange rate. These econometric techniques are especially informative in this context since they allow for a gradual evolution of estimated parameters rather than an abrupt change at a single moment. This evolution will capture the consequences of an ongoing updating by market participants of their views of the

[5] For example, on May 19, 2010 Jean-Claude Trichet was interviewed by the Frankfurter Allgemeine Zeitung (FAZ). The interviewers stated: "…by purchasing government bonds, you've crossed a red line. Has the credibility of the ECB suffered as a result?" The Federal Reserve also had its independence and credibility questioned in light of some emergency lending facilities implemented in 2008, during the financial crisis.
[6] Forward market information has been used in other tests of policy regime credibility, as in, Svensson (1991) on the European Monetary System (EMS) in the 1980s, and Svensson (1993) on the inflation targets of Canada, New Zealand and Sweden.
[7] Indeed, this argument also holds even where there is an updating of the inflation target within the Gürkaynak, Sack and Swanson (2005) framework.
[8] Our exploration of high-frequency data builds on other studies that similarly consider drivers of exchange rates and international asset prices, including Andersen, et al (2003), Ehrmann and Fratzcher (2005), and Faust, Rogers, Wang and Wright (2006).

central bank reaction function. This type of updating can occur as market participants gradually learn through observing central bank actions and communications.

The results in Section 4 show significant and persistent parameter instability in the effects of economic news on European term structures and on the euro-dollar exchange rate. The identified patterns are consistent with market participants updating their views of the policy reaction function of the ECB. Additional support for our updating hypothesis is provided by considering the smoothed time path of the estimated parameters of the coefficient on the news announcement, estimated through another new, and related, econometric technique (Müller and Petalas, forthcoming). Parameter values evolved in a manner consistent with the perception of an increasing aversion to inflation by the ECB as it tightened its monetary policy, or alternatively as consistent with a perceived decline in the inflation target of the ECB. These results on time-varying consequences of economic news for the yield curve are complemented by results of discrete structural break tests (Andrews 1993) that demonstrate the robustness of our findings.

Overall, these empirical results support the view that actions, and not just institutional structure, influence market perceptions of the policy stance of a new central bank. Benchmark test results for the term structure of U.S. interest rates present no evidence of persistent parameter instability in the response of the U.S. term structure to news, a result consistent with stable perceived weights in the Fed's reaction function over this same interval.[9] The results demonstrate that the tools introduced can capture evolving views of central bank preferences and credibility.

2. Central Bank Policy Reaction Functions and Market Responses to News

In this section we present a model in order to demonstrate the effects of changing perceptions on the actual response of interest rates to economic news.[10] The insights from this model inform our interpretation of the empirical results on the evolution of the response of the yield curve to news that are presented in section 4. The basic argument is that market

[9] Other work finds evidence of a shift in Fed policy focus in earlier (and longer) periods. For example, the sample for Fuhrer (1996) is 1966 to 1994, for Gürkaynak, Sack and Swanson (2005) it is 1990 to 2002, and for Kozicki and Tinsley (2005) it begins in 1946, with the longest one ending in 1997.

[10] Earlier empirical studies have also considered the possibility that the effects of news on asset prices of different maturities reveals information about market participant beliefs about central bank policies, including Huizinga and Mishkin (1986), and Fleming and Remolona (1999).

perceptions of a central bank's stance on policy have an important impact on the performance of an economy and the consequences of policy decisions. While these market perceptions are, by nature, unobservable, we argue that actions such as changing views of the relative weight a central bank places on inflation versus the output gap in its monetary policy reaction function should be identifiable through analysis of the high-frequency response of asset prices to economic news.

The model that we use to frame our analysis follows from Gürkaynak, Sack and Swanson (2005, hereafter GSS). This standard New Keynesian approach allows for a significant fraction of backward-looking agents (who, equivalently, can be assumed to act in a rule-of-thumb manner) and also allows for forward-looking agents.[11] The model consists of 6 basic equations

$$(1)\ \pi_t = \mu E_t \pi_{t+1} + (1-\mu) A_\pi(L) \pi_t + \gamma y_t + \varepsilon_t^\pi$$

$$(2)\ y_t = \mu E_t y_{t+1} + (1-\mu) A_y(L) y_t - \beta(i_t - E_t \pi_{t+1}) + \varepsilon_t^y$$

$$(3)\ i_t = (1-c)\left[\overline{\pi}_t + a\left(\overline{\pi}_t - \pi_t^*\right) + b y_t\right] + c i_{t-1} + \varepsilon_t^i$$

$$(4)\ \pi_t^* = \pi_{t-1}^* + \theta\left(\overline{\pi}_{t-1} - \pi_{t-1}^*\right) + \varepsilon_t^{\pi^*}$$

$$(5)\ \hat{\pi}_t^* = \hat{\pi}_{t-1}^* + \theta\left(\overline{\pi}_{t-1} - \hat{\pi}_{t-1}^*\right) - \kappa\left(i_t - \hat{i}_t\right)$$

$$(6)\ \ i_t(m) = \frac{1}{m}\sum_{j=1}^{m} f_j\left(i,1\right)$$

The first two equations represent the macroeconomic structure of the economy. Equation (1) specifies current inflation, π_t, as a function of expected future inflation and lagged inflation, which contributes to inflation persistence through the lag function $A_\pi(L)$. The parameter μ describes the balance of these forward-looking and backward-looking pressures. Current inflation also depends upon y_t, which captures the stance of current output relative to its

[11] See Woodford (2001) for a discussion of the relationship between the standard Taylor Rule and optimal monetary policy considerations. Woodford (2003) shows that under optimal monetary policy, the weight on inflation and output relative to targets are functions of deep parameters such as elasticities of substitution and shock persistence.

potential (i.e. the output gap). Equation (2) describes this output gap as also having forward-looking component and a persistent lagged effect. The output gap (if negative) declines as monetary policy is more expansionary, as reflected in declining real interest rates. The latter are introduced via the difference between the nominal rate i_t and inflation expectations over a comparable maturity horizon.

Shocks to inflation and output are represented by ε_t^{π} and ε_t^{y}, in equations (1) and (2), respectively. The source of these shocks may be either domestic or foreign in origin. Due to international transmission, either foreign or domestic shocks could be a source of domestic inflation and lead to policy reactions (Clarida, Galí and Gertler, 2002).[12]

The third equation specifies the policy reaction function of the central bank. The equation augments concerns about the output gap, which enters with weight *b*, and inflation, $\overline{\pi}_t$, relative to its target, $\overline{\pi}^*$, which enters with weight *a*, with policy rate smoothing. Greater values of the parameter *c* indicate a relative unwillingness of the central bank to deviate sharply from the prior period's policy rate. Equation (3) also affords a role to recent inflation history with $\overline{\pi}_t$ representing a four-quarter moving average of inflation and includes ε_t^i, an i.i.d. shock.

Equations (4) and (5) describe updating of views of the central bank inflation target, π_t^*, by the central bank (equation 4) and by private agents as denoted by the $\wedge$ notation of equation (5). As shown by Erceg and Levin (2003) and GSS, the inflation target of the central bank has a tendency to rise when inflation goes above the prior target, with this effect depending on the size of the θ parameter and can be subject to exogenous changes captured by $\varepsilon_t^{\pi^*}$. Private sector agents infer such changes by observing the deviation of i_t from their prior expectation for policy, and update their view of the target according to the strength of a Kalman gain parameter κ and when the observed policy rate is higher or lower than what they would have expected given the prior perceived target. This approach provides a mechanism for central banks to update their policy reaction function and for the private section to assess and learn about this change. The model also imposes the expectations

[12] Clarida and Waldman (2008) focus on exchange rate dynamics in which the central bank follows an interest rate rule to implement an inflation target.

hypothesis, as in Gürkaynak, Levin, Marder and Swanson (2005), in which long interest rates of maturity m, $i_t(m)$ are the cumulated sum of short interest rates captured by one year forward rates j years ahead $f_j\left(i,1\right)$.

Consider the consequences of varying values of *a* and *b* for the long rate, the short rate, and the slope of the yield curve (i.e. the difference between the long rate and the short rate). The impulse examined is news about inflation, which we occurs through a realization of either ε_t^{π} or ε_t^{y}.[13] The experiment explores the effects of this news over time, and the changes in the effects under parameterizations of the central bank reaction function. In particular, suppose there is an increase in *a* relative to *b* in Equation 3, which Fuhrer and Hooker (1993) describe as greater central bank credibility in fighting inflation. There are marked differences in the response of interest rates to news with changes in the perceived values of *a* and *b*. Indeed, by varying *a* and *b*, the entire time path of adjustment to shocks is altered in the model.[14] Clarida and Waldman (2008) show that the larger the weight on the output gap, the slower the economy's convergences to the central bank's output and inflation targets.

The graphs presented in Figure 1 present the impact effects on the long interest rate, the short interest rate, and the difference between the two (i.e. the slope of the yield curve) to a positive shock to inflation, ε_t^{π} (in the top three graphs), and a positive shock to output, ε_t^{y} (in the bottom three graphs), for a wide set of pairs of the parameters *a* and *b*. The surfaces in the third graph in each of the two rows show that an increase in the relative importance of the weight the central bank places on fighting inflation (*a*) versus the weight it places on stabilizing output (*b*) decreases the yield curve slope (i.e. short rates rise by more than long rates rise) after an inflationary shock. The central bank moves more aggressively to combat inflation, placing less importance on output and employment goals. Under higher *a* (or lower *b*), there still are consequences for both $i_t^s = i_t$ and $i_t^L = i_t(m)$, since a positive realization of either shock will have a smaller positive effect on the long rate as compared to the short rate.

[13] GSS and related studies use parameterizations based on Rudebusch (2002), which assume $c = 0.73$, $\theta = 0.02$ and $\kappa = 0.1$. We will use these parameter values in our simulation while allowing for varying values of *a* and *b*.
[14] Calibration exercises showing the path of alternative variables under parameterizations different from the GSS baselines are available from the authors.

Indeed, the model's quantitative result is that the yield curve response is negative for all cases except when output stabilization has a very high weight compared to inflation stabilization in the central bank's objective function. Another set of quantitative results also can be constructed for exchange rates, as analyzed more broadly in Clarida and Waldman (2008), Engel and West (2005) and much earlier by Hardouvelis (1988).

Empirically, however, it has long been recognized that there is excess sensitivity of the long end of the yield curve to news, as discussed in Ellingsen and Söderström (2001). While Gurkaynak, Sack and Swanson (2005) provide a range of explanations for this excess sensitivity in the data, the implication is that the unexplained excess sensitivity of the long end of the yield curve would shift up the contours shown in the far right graphs of Figure 1, potentially locating that plane in more positive space. For our purposes, the main point is not whether quantitatively there is a net positive or a net negative impact effect of news for a given set of values of a and b. Rather, the key issue is that an increase in central bank credibility results in an inflationary shock moving the economy to a different point on the contour, so that credibility improvements reduce the response of the long interest rate to a greater degree than they reduce the response of the short rate. In particular, an increase in credibility alternatively could be interpreted as a decrease in the expected inflation target of the central bank, as in Erceg and Levin (2003) and GSS. Our approach can be viewed as complementary, although it is noteworthy that the pattern of yield curve response to news does not vary substantively with the κ or θ parameters indicating updating of the perceived central bank inflation target, π_t^*.

3. Empirical Approach

3.1 Testing Strategy

The model presented in the previous section provides a framework interpreting a time-varying response of the immediate effect of economic news on the term structure of interest rates. An empirical application requires a method for testing for instability in this relationship. In this section we discuss the method we use to test for persistent variation in the immediate response of the term structure to news.

We use a linear specification linking the surprise component of news to the change in an asset price, as is standard in research on the high-frequency response of asset prices to news (see, for example, Anderson, Bollerslev, Diebold and Vega, 2003). The specification for the effect of news on any asset price q_t, allowing for the possibility of a time-varying coefficient on the news variable, is

(7) $$q_{t^+} - q_{t^-} = \alpha + \gamma_i \left(x_{t^+} - E_{t^-} x_{t^+} \right) + \varepsilon_{t^+}$$

where $q_{t^+} - q_{t^-}$ is the change in the term structure over the short period of time between t^-, just before an announcement, and t^+, just after that announcement (i.e. $d(i_t^L - i_t^S)$), x_{t^+} represents the announced value of a variable, which is known at time t^+, $E_{t^-} x_{t^+}$ represents the expected value of that variable before the announcement (so $x_{t^+} - E_{t^-} x_{t^+}$ is the surprise component of the announcement), and ε_{t^+} is a white-noise error term. This parsimonious specification is most appropriate when the time horizon between t^- and t^+ is short, for example, when it is measured in minutes rather than days, and when news about the variable x does not become available at the same time (that is, within the span t^- to t^+) as announcements about some other relevant variable.

In our application, we define the term structure as $q_t = i_t^L - i_t^S$, with L and S denoting long term and short term interest rates respectively. Our theoretical motivation argued that a significant evolution in the response of the difference between 10-year and 2-year European interest rates would be expected for a central bank with changing market perceptions of its policy reaction function priorities.

The challenge, in this context, is that there is not a single, widely-recognized dramatic change in policy that was clearly a watershed that led to a change in the public's perceptions of the ECB. Thus, we cannot perform a simple Chow test over γ_i. Moreover, it is unlikely that there was a discrete change in expectations, either in response to a single event or a small number of events. Rather, it is more reasonable to think of perceptions changing gradually as market participants learned about the ECB through its pronouncements and, especially, its actions.[15] Thus, we would like to test for persistent change in the estimated parameter γ_i.[16]

Elliott and Müller (2006) have developed a test for the presence of persistent time variation in one or more regression coefficients. Their *quasi-Local Level* (*qLL*) statistic provides asymptotically equivalent tests for a large class of persistent breaking processes against the alternative of structural stability. The test does not require the specification of an exact breaking process, such as breaks that occur in a random fashion, serial correlation in the changes of coefficients, or a clustering of break points.[17] This feature of their test makes it well suited for our purposes since we do not need to test for a particular type of updating by market participants of their views on central bank inflation aversion. The *qLL* statistic takes a negative value, and a value smaller (i.e. more negative) than the critical value implies a failure to reject time variation in one or more coefficients for the entire sample period. This procedure tests for persistent time variation over the entire sample and, as such, does not identify a particular date as the one most likely to represent a discrete break-point.[18]

As will be shown below, we do, in fact, find evidence that there has been persistent time variation in the slope coefficient in term spread regressions for Germany, France and Italy, but not for the United States. We also find that there is significant parameter instability in the response of the bilateral euro/dollar exchange rate to news.[19] We interpret this combined finding of parameter instability for European rates and the euro-dollar rate and parameter stability for U.S. rates as reflecting an evolving view of the inflation aversion of the ECB rather than as some structural change common to financial markets across all four of these industrial countries.

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[15] Theoretical analysis of learning about central bank behavior is consistent with a gradual evolution of perceptions rather than a one-time shift. For example, see Backus and Driffill (1985) or Athey, Atkeson, and Kehoe (2005).

[16] In the interest of robustness, however, we will also present results from a test for an unknown break-point, from Andrews (1993).

[17] Elliott and Müller write that, for their tests, "…the precise form of the breaking process [of the coefficients] is irrelevant for the asymptotic power of the tests." (p.927) An implication of this is that "From a practical perspective… the researcher does not have to specify the exact path of the breaking process in order to be able to carry out (almost) efficient inference." (p. 914)

[18] The specification (3) allows for time variation in γ. In the interest of offering a more general set of tests, we will also consider the possibility of time variation in α.

[19] The Appendix presents the theoretical underpinnings of the exchange rate specification. If we failed to find persistent parameter instability in the euro-dollar exchange rate regression we would be concerned that there may have been a common structural change across U.S. and Euro Area markets over the sample period.

The Elliott and Müller *quasi-Local Level* (*qLL*) statistic says nothing about the direction of change of γ_i. Yet, following from the results presented in section 2, a decrease in γ_i can be interpreted as an increase in the perceived relative inflation aversion of the central bank. For this purpose, we rely on methods from Müller and Petalas (2005), who show how to calculate the smoothed time path of γ_i. We present these estimated time paths. For robustness, the results from these smoothed estimates are supported by the sup-Wald tests for parameter stability (see Andrews, 1993 and 2003) which offer a break date for γ_i that roughly corresponds to the peak value of the estimated smoothed time path.

As will be shown, we find that γ_i decreases over the sample period for the cases where there is evidence of a significant persistent change in γ_i (i.e. for the three European yield curves and for the euro-dollar exchange rate). Even more tellingly, the reduction in γ_i tends to occur in the wake of monetary tightening by the ECB. It is unlikely that other candidate explanations for changes in the responsiveness of the slope coefficients that are not linked to the perception of the ECB policy stance would map as closely to actual ECB policy changes.

3.2 Data

The three types of data used in our analysis are various asset prices, where the assets are government bonds and foreign exchange, inflation announcements, and related market expectations of inflation. We begin this section with a discussion of the five different asset prices used as dependent variables in our estimation. We then describe our construction of inflation surprises.

Asset Price Data: Five different dependent variables are used in the regressions. In each case, the dependent variable, $q_{t^+} - q_{t^-}$, represents the change in q between thirty minutes before and thirty minutes after each monthly inflation announcement over the period January 1999 to June 2005. The change in the term spread between 10-year and 2-year interest rates for French, Italian, German, or U.S. government bonds are four of the dependent variables, with a robustness section (section 4.4) considering the 2-year and 10-year rates separately.

The fifth dependent variable is where $q_{t^+} - q_{t^-}$ represents the change in the logarithm of the euro-dollar exchange rate, thirty minutes before and thirty minutes after the news announcement. Through short-run interest rate parity, the exchange rate move should reflect the relative effects of news on interest rates in the euro area versus in the United States (see appendix). In this case, a positive value of $q_{t^+} - q_{t^-}$ rather than indicating an increase in the premium of the long rate relative to the short rate, indicates a depreciation of the euro. Evidence that γ_i decreases over the sample period in the exchange rate specification is consistent with a situation of more of an increase in the perceived anti-inflation stance of the ECB, as compared to the U.S. Federal Reserve.

Inflationary Announcements and Expectations: To capture the economic news η_t that lead to asset price updating, we restrict our attention to inflation announcement measures. Candidate data releases for our study potentially include indicators of consumer price inflation for the full euro area, for individual countries in the euro area, and for the United States. The construction of the "news" variable, which is the appropriate variable to be employed in the specification, also requires measures of market expectations for the full sample period. While some earlier studies use VARs to generate measures of shocks, such as Engel and West (2006), we require a more high frequency measure for our analysis of the asset price consequences of news.

Ideally, the "news" variable should capture the inflation surprise for the specific country associated with the yield curve data. While euro area economic data are direct candidates for an empirical application to high-frequency financial data for that region, data availability limits some of their usefulness for our purposes and over the interval we examine. The two necessary components of euro area inflation series – *both* the announcements of inflation and market expectations of the inflation releases - were not available at the time of the introduction of the euro in January 1999. Thus, these euro area series cannot be used to study the critical early years of the ECB when market participants were forming expectations of its monetary policy preferences. Additionally, the actual news content and market impact of inflation announcements from individual European countries were of mixed value to markets in the early part of the sample because of issues related to data quality and data leaks prior to official announcement times (Ehrmann and Fratscher, 2005).

In light of these constraints, we construct proxies for euro area country inflation surprises by starting with the observation that, empirically, U.S. inflation affects the relative price of European exports as well as the cost of imported inputs and consumption goods in Europe (Campa and Goldberg, 2010). Thus, the U.S. core CPI release can map to inflation news for euro area countries. Conceptually, e a number of important reasons support this choice. First, this is an economically valid measure since it is well established that U.S. economic news for output and inflation are consequential indicators of demand and price pressures on European markets. As earlier indicated, these consequences can arise as direct effects transmitted through international trade or can be the result of common global shocks. Andersen *et al.* (2003), Goldberg and Leonard (2003), Faust, Rogers, Wang, and Wright (2007), and Ehrmann and Fratscher (2005), all establish that U.S. announcements have strong news content with large and significant effects on European asset prices. Financial market studies document this phenomenon as well.[20] The observation that U.S. inflation news contains information that is perceived as relevant for European inflation is consistent with recent research efforts empirically decomposing cross-country patterns in inflation, as in Ciccarelli and Mojon (2008).[21]

To construct proxy inflation news in the euro area, we work with the U.S. core CPI measure, for which we have both actual release information and prior market expectations data, and establish the relationship between the U.S. inflation series and respective inflation measures in euro area countries. The U.S. core CPI is one of the most closely followed inflation measures by the market, both in terms of significance in econometric studies and as reflected by its importance in policy discussions.[22] First, we perform linear regressions over these inflation series and find the degree of co-movement suggested by a regression model. As a second step we scale U.S. inflation news using the regression beta's so that we have a proxy for the unit of news to euro area country inflation rate for the respective economies.

[20] See Citigroup (2006). A range of other studies consider exchange rate consequences of news, such as Clarida and Waldman (2008) and Bartolini, Goldberg, and Sacarny (2008).
[21] Of course, an impact of U.S. inflation news for European outcomes does not preclude an ECB reaction to European inflation series as well. But, over the hour-long period representing the time before and after a U.S. inflation announcement (the time period we study), this news is the dominant information reaching the market.
[22] See Clark (2001) for evidence on this point and for an overview of related literature.

Table 1 provides the results of the time series regressions of U.S. inflation on inflation series for Germany, France, Italy, and the euro area in respective columns. The regression betas show that the co-movement is strongest between French and U.S. monthly inflation series, but also highly statistically significant with other countries and the euro area aggregate. These respective betas imply the unit of news for local inflation rates that will be associated with a 1 unit measure of news in U.S. inflation.[23]

One potential concern with our use of news in U.S. inflation is that our findings of persistent parameter instability might be due to a changing relationship between inflation in the U.S. and inflation in the euro area countries over the sample period. Perceptions of this change could then be a source of changes in the responsiveness of European interest rates to news. We test for such evidence using *qLL* statistics to identify persistent time variation in the relationship between inflation in the U.S. and in the euro area over this period. The results presented in the qLL for β row of Table 1, however, show that this is not the case. There is no evidence of significant time variation of the coefficient on monthly U.S. inflation for the period January 1998 to December 2005 in any of these four regressions. The smallest *qLL* statistic is -6.9 (the critical value for the 90 percent level of confidence is -7.14). In the econometric analysis that follows, this result of parameter stability helps isolate inflation aversion as the source of the persistent time variation in γ_i.[24]

The *news* or surprise component of core CPI is defined as the difference between the actual release value and the markets' prior expectation of the contents of the release. The expectations data we use are median responses from weekly surveys of market participants conducted by Money Market Services, a division of Standard & Poor's, and more recently from Action Economics.[25] A regression of the 75 median monthly survey responses on the actual monthly inflation reports generates a coefficient of 0.68, with p-value of 0.026, with the regression unable to reject unbiasedness of the survey as a predictor of the actual value of the inflation reports. In creating the inflation "news" variable, we normalize news by the

[23] The inflation news for a euro area country is the estimated beta times the inflation news for the United States.
[24] We also test for separate persistent time variation in the intercept term, α_i, and for joint persistent time variation in both the slope coefficient and the intercept.
[25] Money Market Services were the source of these data through December 2003. Haver Analytics provided continuous expectations and announcement data through 2005 using data from Action Economics. Gürkaynak and Wolfers (2007) show that these data have been among the best performing expectations series for important macroeconomic variables over the sample period that we analyze.

sample standard deviation of the difference between the reported and the expected values of the announcements so that the variable introduced as *news* driving the yield curve and exchange rates in our empirical methods has mean 0 and standard deviation 1.

4. Evolving Market Perceptions of the ECB Reaction Function

The results of the Elliott and Müller (2006) test applied to the slopes of the yield curves are presented in Section 4.1. These tests show evidence of significant persistent time variation in the slope coefficient for term spreads of German, French and Italian bonds, as well as for the euro-dollar exchange rate, but not for U.S. bonds. In Section 4.2 we show that the timing of changes in the estimates of the smoothed time path of γ_i corresponds to actual policy changes undertaken by the ECB. In Section 4.3 to demonstrate the robustness of our results concerning the presence and timing of a persistent change in the market's perception of the ECB's anti-inflation stance, we present sup-Wald tests for a discrete break in the regression relationship (from Andrews, 1993) and the dates associated with those breaks. In Section 4.4 we provide separate results for the effects of economic news announcements on the short and long ends of the yield curve to provide further economic insights into the process behind the evolving perceptions of the ECB reaction function.

4.1 Time Variation in the Effects of News on the Yield Curve

In this section we apply the tests for time variation in the slope of the specifications, reporting the results of the Elliott and Müller (2006) *qLL* statistic for regressions using the five series discussed above as dependent variables. As mentioned, the *qLL* statistics are negative. Evidence of an evolving view of the policy stance of the ECB over time would be implied by a value of the q*LL* statistic smaller (i.e. more negative) than its critical value for regressions using the change in the term spread for German, French and Italian government bonds, as well as for the change in the euro-dollar exchange rate. In a regression using the U.S. term spread the value of the *qLL* statistic larger than its critical value would be evidence against persistent time variation in the perception of the anti-inflation stance of the Federal Reserve over this period.

Results of this test are presented in Table 2.[26] The first row is a test of the general persistent variation in the slope coefficient only. The second row is a test of the general persistent variation in the intercept only. The third row is a joint test of the general persistent variation in both the slope and the intercept coefficients. Critical values are included in the bottom row of the table. Entries in bold and italic represent a *qLL* statistic that is significant at better than the 99 percent level of confidence, bold entries represent a *qLL* statistic that is significant at between the 95 percent and 99 percent levels of confidence, and italic entries represent a *qLL* statistic that is significant at between the 90 percent and 95 percent levels of confidence.

The results in Table 2 provide evidence of persistent time variation in γ_i in regressions of inflation news on the change in the term spread of German government bonds and French government bonds, and in the euro-dollar exchange rate, at greater than the 95 percent level of confidence, and on the change in the term spread of Italian bonds at between the 90 percent and the 95 percent level of confidence. In contrast, there is no significant evidence of persistent time variation in the slope coefficient in a regression of *news* on the change in the term spread of U.S. government bonds over this same period of time. As robustness checks to alternative specifications, we also provide tests for persistent time variation in the intercept, as well as jointly over the slope and intercept of these specifications. There is also no significant evidence of any persistent time variation in the intercept α_i of these regressions, with none of the test statistics significant at the 95 percent level of confidence.

All of these results are consistent with the model in Section 2 in which γ_i varies with an evolving view of the inflation aversion of the ECB in the period after its inception. There is not a corresponding evolution in the view of the inflation preferences of the Federal Reserve during this period, which followed almost fifteen years of observations of the policy actions of the Federal Reserve Board of Governors under the leadership of Chairman Greenspan.

[26] As suggested by Elliott and Müller (2006), we allow for the possibility of heteroskedasticity in the variance-covariance matrix of the score series $\left\{(x_{t+} - E_t x_{t-}) \times \varepsilon_{t^+}\right\}$ by using the Newey-West (1987) correction. We have written a Stata program for conducting the Elliott – Müller test which is available on request.

The finding of a significant persistent variation for the slopes of the German, French and Italian yield curves, as well as for the euro-dollar exchange rate, and the rejection of significant persistent time variation for the slope of the U.S. yield curve, suggest that results are being driven by variation in the market's perception of the anti-inflation stance of the ECB, rather than some overall change in inflation dynamics affecting the U.S. and European countries.

4.2 Estimated Paths of γ_i

The central hypothesis in this paper is that the perceived anti-inflation stance of the ECB evolved with its policy actions. The *qLL* statistics presented in the previous section suggest that there was, in fact, persistent time variation in the term spreads of German, French, and Italian bonds, as well as in the euro-dollar rate-, over this sample period, but there was no similar significant variation in the U.S. term spread. While these results support our hypothesis, an even stronger case can be made by considering the estimated time paths of the estimated γ_i's for each country's inflation news in light of the policy moves by the ECB and the economic performance of the Euro-12 area during this period. In this section we show that the estimated parameter paths of γ_i from regressions for each of the three European term spread regressions and the euro-dollar rate regression followed a pattern consistent with an evolving credibility of the ECB, given its policy moves and the economic environment in the Eurozone over this time.

Figure 2 presents the estimated smoothed parameter paths of γ_i for each of the four term spreads contingent on news in the country's inflation and Figure 3 presents the estimated smoothed time path for the euro-dollar exchange rate and, to provide context, repeats the presentation of the time paths for the U.S. and German term spreads from Figure 2. The time paths are calculated using the technique developed by Müller and Petalas (forthcoming), who show how to estimate the parameter path for general unstable time series models by minimizing a weighted-average risk criterion, a procedure that is akin to a smoothing problem. This procedure requires only general assumptions about the true persistent time variation of the coefficients.[27] The shaded area in these figures indicates the period of the seven interest rate increases, from November 1999 through October 2000.

Before turning to the evolution of the time paths during the full sample period, we first note that Figure 2 shows that the estimated value of γ_i at the beginning of the sample is substantially larger for the three European government bonds than for the U.S. bonds.[28] Another immediately apparent characteristic of the four time paths in Figure 2 is the relative variability of the three European γ_i 's as compared to that of the United States, which, of course, is a reflection of the results of the Elliott – Müller tests presented in the previous section.

The time variation of the estimated paths of γ_i, viewed in light of both the actions undertaken by the ECB and contemporaneous published views of its conduct, bolster our contention that the variation in this parameter is due to changing views of ECB policy stance. In order to make this point, we offer in Table 3 an overview of the policy of the ECB from the time it began its operations in January 1999 until the end of our sample period in June 2005, and the economic environment in which these policy moves took place. This table includes the prevailing interest rate for refinancing operations set by the ECB (which is its policy interest rate), the dates when new interest rates took effect, the year-on-year Harmonized Index of Consumer Price (HICP) inflation for the euro area in the month immediately before the policy move, the unemployment rate for the Euro-12 countries in the month preceding the policy move, and the growth of real GDP in the quarter preceding the policy move.[29] We also refer to conclusions on ECB conduct presented in various annual

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[27] Müller and Petalas (2005) describe their procedure as an extension of the Kalman smoothing formulae with the optimal smoother for the true path of the time varying coefficient as a function of the score sequence $\{(x_{t+} - E_t x_{t-}) \times \varepsilon_{t^+}\}$. See their paper for details, and for an outline of how to implement their procedure. We have written a Stata program for implementing the Müller - Petalas procedure, which is available on request.

[28] This could reflect the efforts by politicians to weigh in on the conduct of ECB policy in the period before it began operations. For example, Austrian Chancellor Viktor Klima said, at a summit in Pörtschach, Austria, in October 1998, “There are good conditions for low interest rates in the euro zone. Stable prices, growth and employment are not contradictory.” Oskar Lafontaine, appointed Finance Minister of Germany in the Autumn of 1998, called for the new ECB to lower interest rates from the time of his appointment until his resignation in March 1999. In response, Wim Duisenberg, the first president of the ECB, stated in November 1998 that it was a “normal phenomenon” for politicians to offer their views on the conduct of monetary policy, but “it would be very abnormal if those suggestions were to be listened to.” See “Wim Duisenberg, Banker to a New Europe,” *The Economist*, November 26, 1998.

[29] On June 8 2000, the ECB announced that starting June 28, 2000, the main refinancing operations of the Eurosystem would switch from fixed rate tenders to variable rate tenders. Thereafter, the key interest rate set by

volumes of the Centre for Economic Policy Research (CEPR) publication, *Monitoring the European Central Bank*.

Table 3 shows that the policy interest rate, the rate on main refinancing operations, was 3.00% in January 1999. The inflation rate in December 1998 was 0.79%, and the unemployment rate was 9.71% in that month, while the growth for the Euro-12 in the last quarter of 1998 was 1.90%.[30] The ECB lowered the policy interest rate by 50 basis points three months after it began operations. There were no further policy moves until a 50 basis point increase in November 1999 that returned the policy interest rate to 3.00%. At that time, the year-on-year HICP inflation rate had risen from under 1% in the early Spring of 1999 to 1.36% in October, real GDP growth had picked up and the unemployment rate had fallen. Reflecting on this period and the initial policy stance of the ECB, the June 2000 issue of *Monitoring the European Central Bank* (Favero, Freixas, Persson, and Wyplosz, 2000) presented the view that the ECB ran a looser monetary policy than the one that would have been expected from the Federal Reserve or the Bundesbank had these central banks faced a similar economic environment. The authors of this publication concluded that there was "some market evidence that the ECB's credibility has indeed been wavering, at least in the second part of 1999."

Indeed, the estimated values of γ_i for the European term spreads initially rise, reaching a peak at the time of the May 1999 core CPI announcement for the French and Italian bond yields and at the time of the June 1999 announcement for the German bond yields. These peaks followed in the wake of the April 1999 interest rate cut, but preceded the series of interest rate hikes that began in November of that year. The peak value of γ_i for the euro-dollar exchange rate occurs in April 2000, in the midst of the seven interest rate increases by the ECB between November 1999 and October 2000.

Real GDP growth continued to rise and the unemployment rate continued to decrease during the period between November 1999 and October 2000 when the ECB raised interest rates seven times, with a cumulative change in the interest rate of 175 basis points to 4.75% by October 2000. The March 2001 issue of *Monitoring the European Central Bank* (Alesina,

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the ECB was the minimum bid rate of the variable rate tenders for the main refinancing operations. See www.ecb.int/stats/monetary/rates.

Blanchard, Galí, Giavazzi and Uhlig, 2001) concluded that these interest rate increases marked a departure of ECB policy from its earlier pattern.[31] Accordingly, our estimates of γ_i for the European term spreads and the euro-dollar exchange rate declined through this period.

This decline in γ_i for the German and Italian term spreads continued, while estimated γ_i for the French term spread and the euro-dollar rate remained largely unchanged, in the wake of the four interest rate cuts between May 2001 and November 2001. At this time, GDP growth continued to slow and unemployment began to rise. The fact that these rate cuts were not viewed as the ECB backsliding from its hawkish stance is supported by the surprise at the continued tightness of ECB monetary policy expressed in the April 2002 volume of *Monitoring the European Central Bank* (Begg, Canova, De Grauwe, Fátas, and Lane, 2002), with calls for a shift in policy in light of the softening economic conditions in the eurozone. In fact, the ECB cut its policy interest rate three times over the next half-year, in December 2002, March 2003 and June 2003. By the time of the last of these interest rate cuts, the estimated values of γ_i for the French and Italian term spreads were higher than their values immediately in the wake of the interest rate cut in November 2001. Subsequently, there was a reduction in the estimated γ_i for the Italian and French term spreads, which is consistent with unchanged monetary policy in the face of continued weak economic performance and quiescent inflation.

Thus, the evidence presented by the estimated smoothed time paths of the parameters of the news regressions, along with the timing of ECB monetary policy changes and the narrative descriptions of ECB policy, supports the conclusion that the public perception of the anti-inflation stance of the ECB evolved during the period under study in response to actions taken by the ECB.

[31] This volume of *MECB* also demonstrated some frustration with a continuing lack of policy transparency, as shown by its recommendation that "The Bank should stop leaving markets and policy analysts to guess what it is really attempting to accomplish as by doing this it runs the risk of a breakdown in communications."

4.3 Sup-Wald Statistics

Sup-Wald tests for discrete changes in γ_i, based on Andrews (1993, 2003), show the robustness of both the Elliott – Müller *qLL* tests and of the smoothed paths of the γ_i coefficients obtained through the Müller – Petalas procedure. These sup-Wald tests are based on a more restricted assumption concerning the break point than the *qLL* test but, since a break-point rather than the overall stability of the parameter is estimated, the sup-Wald tests also provide a date for the break. We compare these dates to the smoothed parameter paths presented in Section 4.2.

The sup-Wald tests are conducted by running a series of regressions of the form

$$(8) \qquad q_{t^+} - q_{t^-} = \alpha + \beta\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \beta_I D_I \left(x_{t^+} - E_{t^-}x_{t^+}\right) + \varepsilon_{t^+}$$

where D_I is a dummy variable that equals 0 for the first *n* observations of the sample and equals 1 for the remaining $T - n$ observations. The sup-Wald test requires running a set of $0.7 \times T$ regressions (if one imposes a 15 percent trimming of observations, as is suggested by Andrews, 1993) which generates a set of $0.7 \times T$ β_I's and $0.7 \times T$ associated test statistics. The sup-Wald test compares the largest F-value for all of the β_I's with critical values presented in Andrews (2003) and, if this sup-Wald statistic exceeds the critical value, the date associated with that β_I is the statistically significant estimated break-date.

Table 4 presents the sup-Wald statistics based on sets of the five different dependent variables that take the form of (8), among which four have as the dependent variable the change in one of the term spreads, and one has as the dependent variable the change in the euro-dollar exchange rate. The statistics presented in the top section of this table show evidence of a significant break, at better than the 99 percent level of confidence, for the regressions using the change in the term spread for German government bonds and for the euro/dollar exchange rate, and at between the 95 and 99 percent level of confidence for the change in the term spread of Italian government bonds. According to the sup-Wald tests, there is no evidence of a significant discrete break for the regression using the change in the term spread of French or U.S. government bonds.

It is interesting to compare the dates obtained through the sup-Wald tests with the smoothed parameter paths obtained using the Müller – Petalas method, where the latter can also be viewed as encompassing either discrete or gradual break-points. The dates presented by sup-Wald tests for the significant estimated break-points for the term spread regressions, November 16, 2000 for the German case and June 15, 2001 for the Italian case, occur about mid-way between the peak and the trough of the respective time paths of γ_i in the period between mid-1999 and late-2001. There is also a consistency between the estimated break-points of February 21, 2001 for the euro-dollar regression and the Müller – Petalas estimated time path for the coefficient in that regression. These break-dates, based upon sup-Wald tests, occur after the seven interest rate increases by the ECB between November 1999 and October 2000, and before any (for the German term spread and the euro-dollar rate) or all but one (for the Italian term spread) of the subsequent interest rate cuts. Thus, these tests support the robustness of the results presented in the previous sections.

4.4 News Consequences along the Yield Curve

As a final set of exercises, we consider the separate paths of the 2-year and 10-year bond yields for each of the Euro Area countries. The simple model exposited in Section 2 argues that, for example, higher weights on inflation gap aversion would generate declining γ_i for long term yields and a higher γ_i on short rates (as shown in Figure 1). The *qLL* tests for parameter instability present evidence of instability γ_i on 10-year rates for Germany, Italy and France, but do not identify instability in the 2-year rates.[32] The estimated smoothed parameter paths for 2- and 10-year interest rates for each country, shown in Figures 4a, 4b, and 4c, indicate more of an evolution of γ_i on 10-year rates consistent with enhanced inflation aversion of the ECB over its early years of operation, with less conclusive findings over the 2-year rates.

5. Conclusions

The importance of the reputation of a central bank for the success of its operations is stressed in theory and is evident from practical experience. An important question is whether a central bank gains credibility in its anti-inflation and low inflation target stance through its

[32] The *qLL* statistics for the 10-year rates are -7.53 for Germany, -7.99 for France, and -7.20 for Italy, all of which are significant at better than the 90% level of confidence. The respective *qLL* statistics for the 2-year rates are -3.76 for Germany, -4.04 for France, and -3.90 for Italy.

institutional structure or through the conduct of policy. This question is especially relevant for a newly established central bank that faces the challenge of establishing its reputation, sometimes in the face of political controversy over the appropriate conduct of monetary policy. Likewise, such questions are important for changing leadership at established central banks, and critical for understanding the implications for central bank credibility as choices are made over alternative monetary regimes, for example, regarding inflation targeting.

The evolution of the markets' perceptions of the central bank policy function of the ECB in the first few years since it began operations in January 1999 is interesting for a number of reasons. One of these reasons is the inherent interest of the economic experience of the eurozone. A second reason is that the establishment of the ECB provides a natural experiment for considering how the perception of a central bank's priorities evolves over time. This episode is a particularly interesting one because of the controversy surrounding the conduct of monetary policy in Europe as the ECB began its operations.

The issue of market perceptions of the credibility and priorities of central banks returned at various times during and following the financial crisis that began in 2007 and spread across markets. The expansion of the Federal Reserve balance sheet, and the range of special liquidity programs led to discussions of central bank priorities. Financial reform debates and legislation also led to discussions about the importance of maintaining central bank independence in the United States, and repeated defense of this independence, along with transparency and accountability (Bernanke 2010). In Europe, the ECB has likewise argued that recent policy initiatives do not change the priorities or independence of the central bank (Trichet 2010). In both cases, however, market perceptions still matter for the consequences of shocks and the effectiveness of policy responses.

In this paper, we have proposed and executed a novel test for central bank "credibility" changes via the study of the evolution of yield curve response to economic news announcements. We show conceptually that the use of high-frequency data can provide useful insights into evolving market perceptions about the inflation aversion and priorities of a central bank. This empirical methodology and the use of high-frequency data provide a unique window into the evolution of market perceptions of a monetary policy reaction function, an issue more typically and less precisely addressed using lower frequency data. Our evidence is consistent with an evolution of perceptions of the policy stance of the ECB, one linked to its interest rate policy over its early years. During this same period, there was not a similar shift in the market's perception of the policy stance of the Federal Reserve, a period marked by the stability in its leadership, the consistency of its stated goals and broad support for its conduct of policy. Overall, the tools we have presented and applied are relevant for ongoing questions of the changing effects of news on market activity, and of the changing policies and perceptions of monetary authorities worldwide.

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Table 1: Elliott-Müller Test for Persistent Time Variation in Effects of U.S. Inflation Inflation in Germany, France, Italy and the Euro Area

$\pi_i = \alpha + \beta\pi_{US.} + \varepsilon$	Monthly Inflation in			
	Germany	France	Italy	Euro Area
α	0.049	**0.055**	***0.163***	***0.107***
(s.e.)	(0.034)	(0.026)	(0.0137)	(0.027)
β	***0.194***	***0.362***	***0.117***	***0.275***
(s.e.)	(0.089)	(0.069)	(0.036)	(0.072)
qLL for β	-4.74	-4.29	6.91	-1.86

Table provides estimated coefficients and associated standard errors (s.e.) for regressions of monthly inflation in the United States on monthly inflation across specific euro area countries and a euro area aggregate c. Estimation period is January1998 – December 2005 (96 observations but for Germany which does not include observations for Dec. 2001, Dec. 2002, Dec. 2004)

Critical Values for *qLL*: 1% ***-11.05***; 5% **-8.36**; 10% *-7.14*

Bold and Italic = significant at 99% level, **Bold** = significant at 95% level, *Italics* = significant at 90% level.

Table 2: Elliott-Müller Test for Persistent Time Variation in Yield Curve Specifications

Test of Time Variation of	Change in Term Spread of Government Bonds of				Change in euro-dollar
	Germany	France	Italy	United States	
Slope, γ	***-11.11***	**-8.75**	*-7.43*	-5.61	**-8.79**
Intercept, α	*-7.84*	-3.14	*-7.44*	*-7.85*	-3.21
Joint Slope & Intercept	***-21.72***	-9.08	-7.13	-8.75	***-16.69***
No. of obs.	74	74	72	73	75

Critical Values: 1 coefficient (Slope alone) 1% ***-11.05***; 5% **-8.36**; 10% *-7.14*
2 coefficients (Slope & Intercept) 1% ***-17.57***; 5% **-14.32**; 10% *-12.80*

Table provides *qLL* statistics for estimated coefficients from regressions of the form $q_{t^+} - q_{t^-} = \alpha + \gamma_i\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \varepsilon_{t^+}$ where $q_{t^+} - q_{t^-}$ is the change in the term structure $d(i_t^L - i_t^S)$ between t^-, just before an announcement, and t^+, just after that announcement (i.e. $d(i_t^L - i_t^S)$), x_{t^+} represents the announced value of U.S. core inflation and $E_{t^-}x_{t^+}$ represents its expected value. Estimation period January 1999 to June 2005.

Critical Values for *qLL*: 1% ***-11.05***; 5% **-8.36**; 10% *-7.14*

Bold and Italic = significant at 99% level, **Bold** = significant at 95% level,
Italics = significant at 90% level.

Table 3: Dates of Changes in Interest Rate by European Central Bank and HICP Inflation in Previous Month January 1, 1999 – June 2005

With Effect from	Policy Interest Rate (Main Refinancing Operation)*	Year-on-year HICP inflation, in previous month	Unemployment Rate, Euro-12, in previous month	Real GDP growth in previous quarter
January 1, 1999	3.00	0.79	9.71	1.90
April 9, 1999	2.50	0.98	9.46	2.11
November 5, 1999	3.00	1.36	8.82	3.05
February 4, 2000	3.25	1.85	8.59	4.08
March 17, 2000	3.50	1.94	8.54	4.08
April 28, 2000	3.75	1.93	8.47	4.38
June 9, 2000	4.25	1.73	8.21	4.38
September 1, 2000	4.50	2.02	8.09	4.68
October 6, 2000	4.75	2.50	8.08	3.82
May 11, 2001	4.50	2.75	7.72	2.87
August 31, 2001	4.25	2.55	7.77	2.08
September 18, 2001	3.75	2.34	7.80	2.08
November 9, 2001	3.25	2.25	7.86	1.66
December 6, 2002	2.75	2.29	8.42	1.14
March 7, 2003	2.50	2.37	8.56	1.07
June 6, 2003	2.00	1.81	8.67	1.00

Source: European Central Bank, http://www.ecb.int/stats/monetary/rates/html/index.en.html

All data are in percent. HICP inflation, Unemployment, GDP growth from ECB webpage. See http://sdw.ecb.int

* Interest rate is for Main Refinancing Operations. On June 8, 2000, the ECB announced that, starting from the operation to be settled on June 28, 2000, the main refinancing operations of the Eurosystem would switch from fixed rate tenders to variable rate tenders. The minimum bid rate for these variable rates refers to the minimum interest rate at which counterparties may place their bids.

Table 4: Sup-Wald Test for Discrete Break Point in Term Structure and Euro-Dollar Specifications

Break-Point in	Change in Term Spread of Government Bonds of				Change in euro-dollar
	Germany	**France**	**Italy**	**United States**	
Sup-Wald Statistic	***20.31***	1.78	**11.91**	4.31	***12.25***
Estimated Break-Date	Nov.16, 2000		June 15, 2001		Feb.21, 2001
No. of obs.	74	74	72	73	75

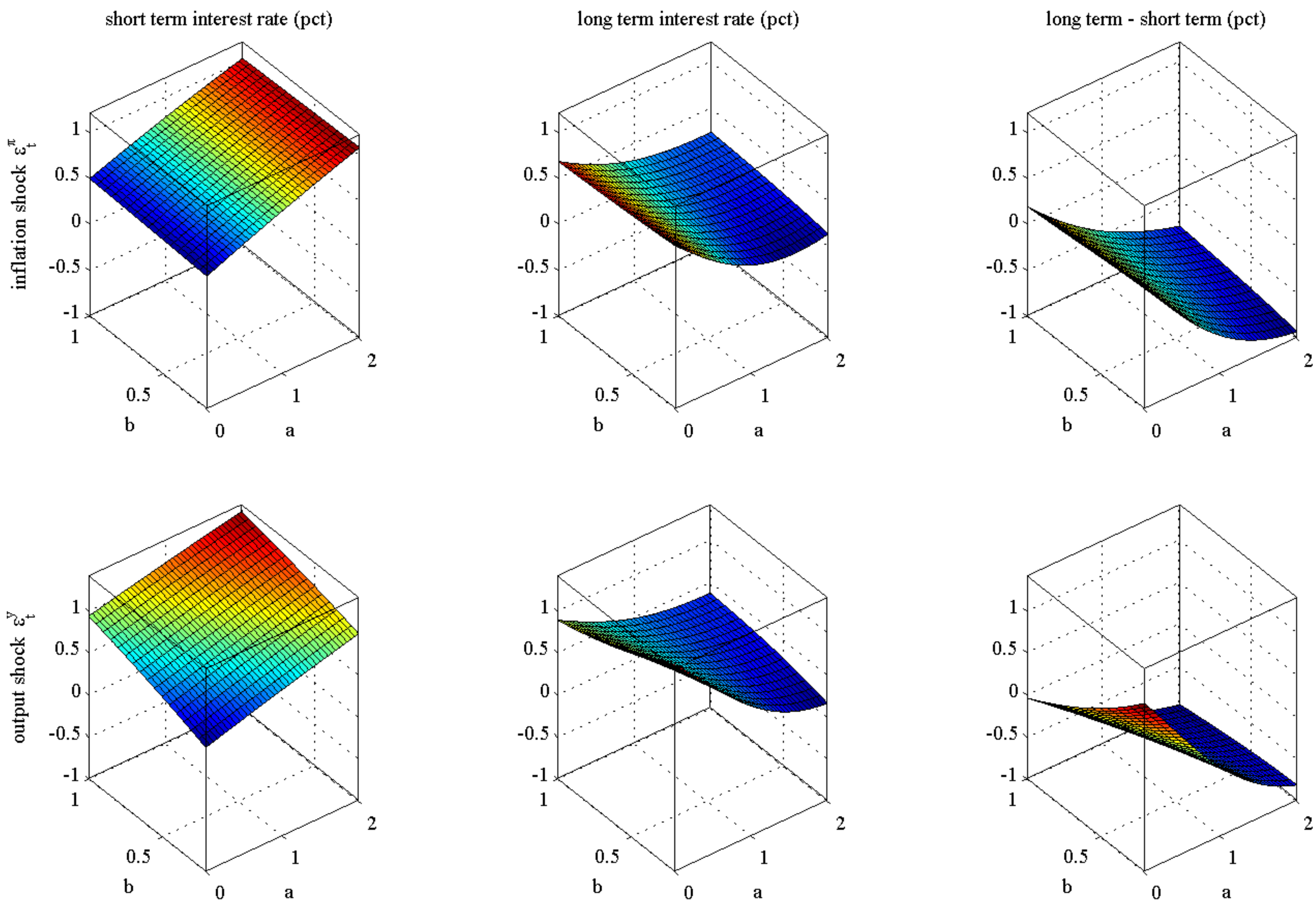
Table provides sup-Wald test statistics for estimated coefficients from regressions of the form $q_{t^+} - q_{t^-} = \alpha + \beta\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \beta_I D_I\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \varepsilon_{t^+}$

where D_I is a dummy variable that equals 0 for the first *n* observations of the sample and equals 1 for the remaining $T - n$ observations. $q_{t^+} - q_{t^-}$ is the change in the term structure $d(i_t^L - i_t^S)$ between t^-, just before an announcement, and t^+, just after that announcement (i.e. $d(i_t^L - i_t^S)$), x_{t^+} represents the announced value of U.S. core inflation and $E_{t^-}x_{t^+}$ represents its expected value. Estimation period January 1999 to June 2005.

Critical Values (from Andrews 2003) 1% ***12.16***; 5% **8.68**; 10% *7.12*

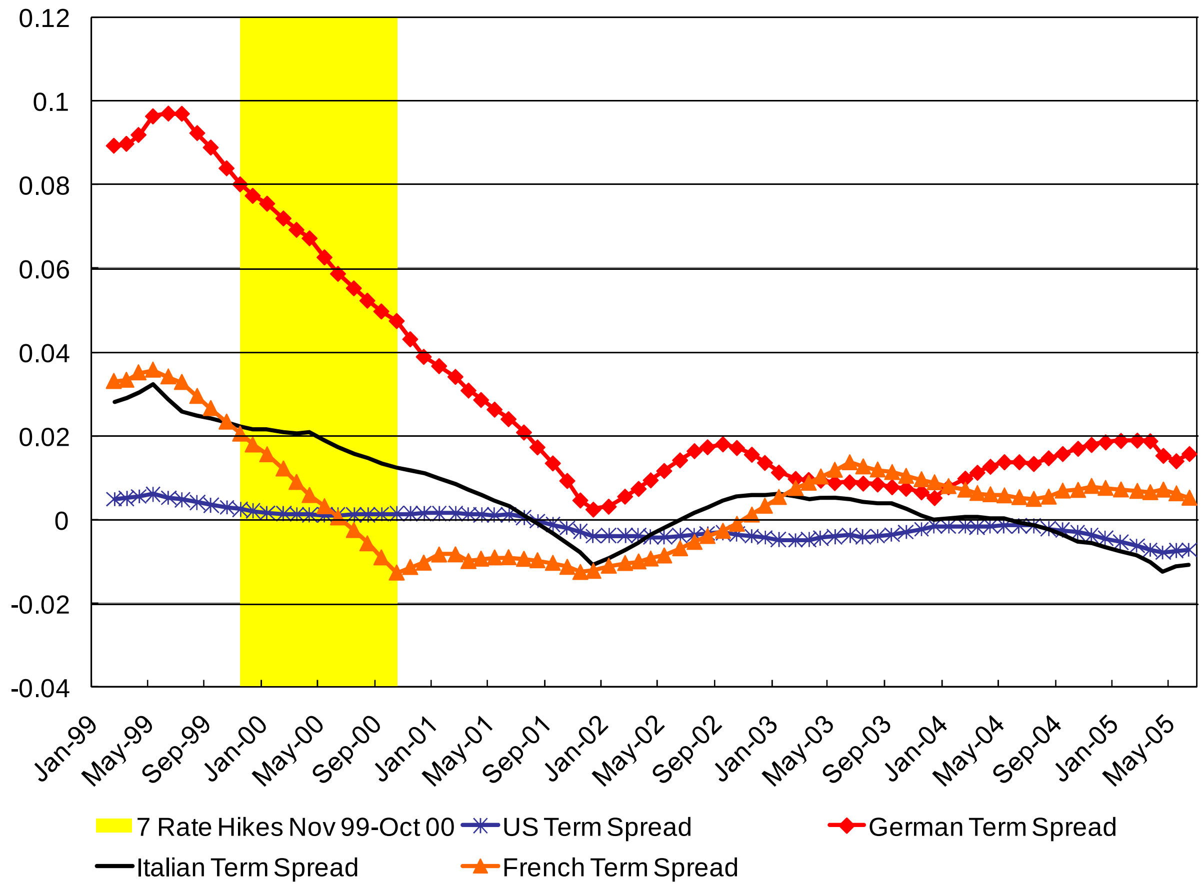
Tests conducted with 15 percent symmetric trimming of observations.

Figure 1: Impact Effects of Inflation and Output News



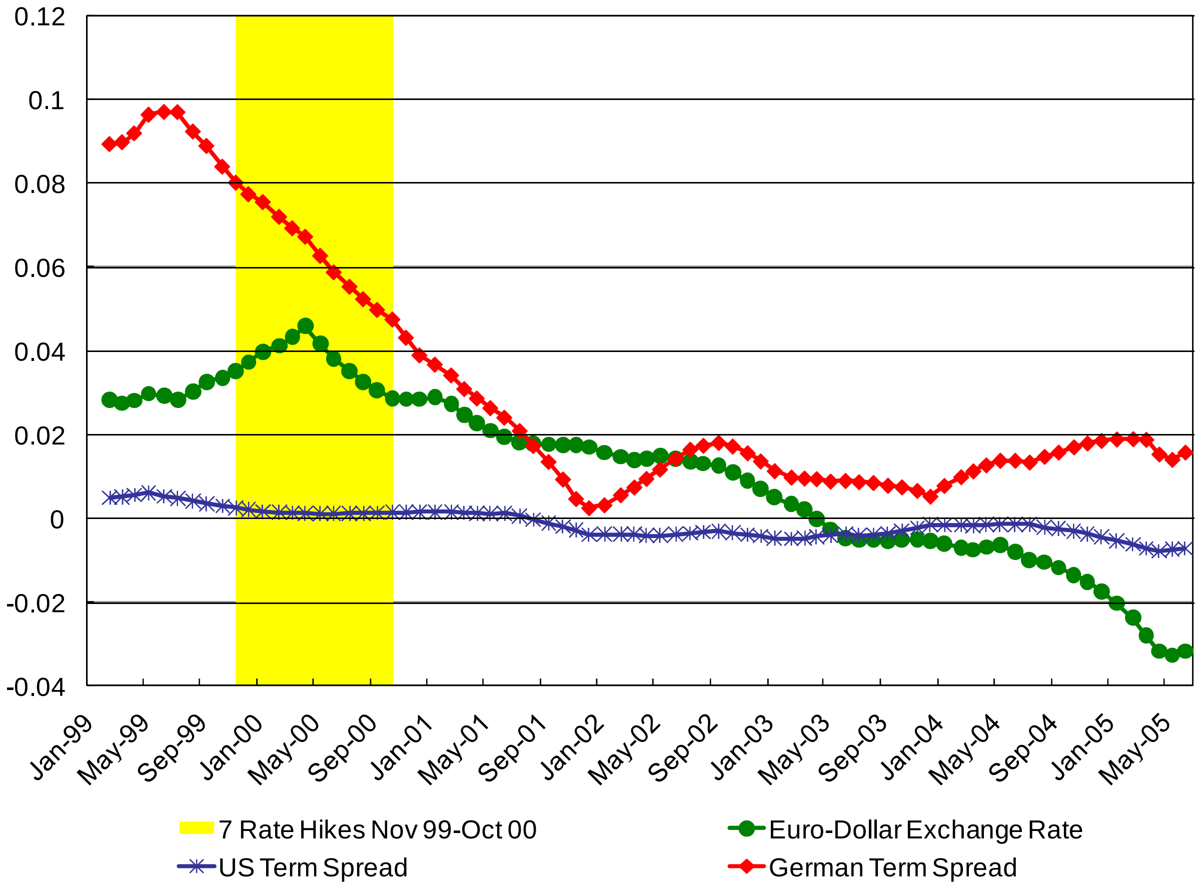
Note: 'a' represents the weight on the inflation gap while 'b' represents the weight on the output gap.

Figure 2: Term Spread Responses to News



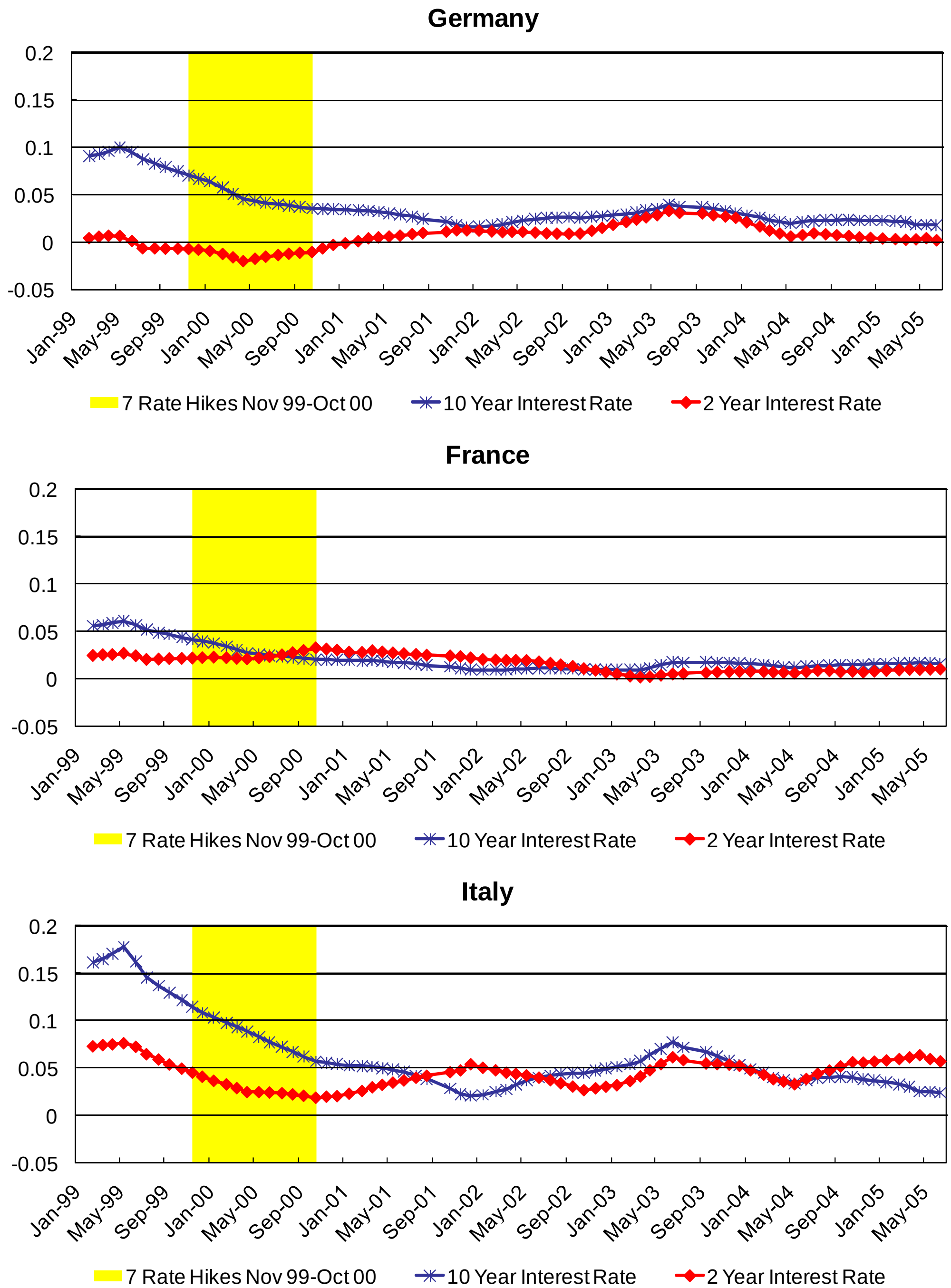
Note: The figure provides smoothed parameter estimated coefficients from regressions of core inflation news on changes in the 10-year less 2-year interest rate differentials between t^{-}, just before an announcement, and t^{+}, just after that announcement. Estimation period is January 1999 to June 2005.

Figure 3: Euro-Dollar Exchange Rate and Term Spread Responses to News



Note: The figure provides smoothed parameter estimated coefficients from regressions of core inflation news on changes in the 10-year less 2-year interest rate differentials and changes in the euro-dollar exchange rate between t^{-}, just before an announcement, and t^{+}, just after that announcement. Estimation period is January 1999 to June 2005.

Figure 4: Response of 2-year and 10-year Rates to News



Note: The figure provides smoothed parameter estimated coefficients from regressions of core inflation news on changes in the 2-year and 10-year interest rates between t^-, just before an announcement, and t^+, just after that announcement. Estimation period is January 1999 to June 2005.

Appendix

As explored in Faust, Rogers, Wang and Wright (2007) and Clarida and Waldman (2008), we use both European and American term structures to motivate further testing following the specification in (3) but with $q_{t^+} - q_{t^-}$ defined instead as the percentage change in the euro-dollar exchange rate. The link between the term structure tests and the exchange rate test can be seen by first considering the interest parity relationships

$$\text{(A1)} \quad \begin{aligned} i_t^{L,EUR} - i_t^{L,US} &= E_t e_L - e_0 \\ i_t^{S,EUR} - i_t^{S,US} &= E_t e_S - e_0 \end{aligned}$$

where $i_t^{L,EUR}$ and $i_t^{L,US}$ are the long interest rates in one of the European countries and the United States, respectively, $E_t e_L$ is the expected logarithm of the euro-dollar exchange rate at a time in the future matching the maturity of the long interest rate, e_0 is the logarithm of the current spot exchange rate, and, in the second line, the replacement of *L* with *S* reflects an interest parity relationship with shorter maturity interest rates and an expected logarithm of the exchange rate at a moment in the future matching the shorter maturity. Subtracting the short-maturity interest parity relationship from the long-maturity interest parity relationship, we get

$$\text{(A2)} \quad \left(i_t^{L,EUR} - i_t^{S,EUR}\right) - \left(i_t^{L,US} - i_t^{S,US}\right) = E_t e_L - E_t e_S$$

The same news variable will be used for both European and U.S. term structures. Thus, considering the values of (A2) before and after the news announcement, we have

$$\text{(A3)} \quad \left(E_{t^+} e_L - E_{t^+} e_S\right) - \left(E_{t^-} e_L - E_{t^-} e_S\right) = \left(\alpha^{EUR} - \alpha^{US}\right) + \left(\gamma_i^{EUR} - \gamma_i^{US}\right)\left(x_{t^+} - E_{t^-} x_{t^+}\right) + \left(\varepsilon_{t^+}^{EUR} - \varepsilon_{t^+}^{US}\right)$$

The expected exchange rate variables are unobservable, but a change in expected depreciation between the time before and the time after the announcement would affect the current spot rate. With this in mind, we can estimate

$$\text{(A4)} \quad e_{t^+} - e_{t^-} = \left(\alpha^{EUR} - \alpha^{US}\right) + \left(\gamma_i^{EUR} - \gamma_i^{US}\right)\left(x_{t^+} - E_{t^-} x_{t^+}\right) + \left(\varepsilon_{t^+}^{EUR} - \varepsilon_{t^+}^{US}\right)$$

to test for a persistent parameter variation in $\left(\gamma_i^{EUR} - \gamma_i^{US}\right)$. A finding of persistent parameter variation in the coefficient on $\left(x_{t^+} - E_{t^-} x_{t^+}\right)$ in (A4), along with a finding of persistent parameter variation in European term structure regressions but not in U.S. term structure regressions, would bolster our conclusion of an evolving perception of the anti-inflation stance of the European Central Bank.