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ESTABLISHING CREDIBILITY: EVOLVING PERCEPTIONS OF THE EUROPEAN CENTRAL BANK

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ABSTRACT

The perceptions of a central bank's inflation aversion may reflect institutional structure or, more dynamically, the history of its policy decisions. In this paper, we present a novel empirical framework that uses high frequency data to test for persistent variation in market perceptions of central bank inflation aversion. The first years of the European Central Bank (ECB) provide a natural experiment for this model. Tests of the effect of news announcements on the slope of yield curves in the euro-area, and on the euro/dollar exchange rate, suggest that the market's perception of the policy stance of the ECB during its first six years of operation significantly evolved, with a belief in its inflation aversion increasing in the wake of its monetary tightening. In contrast, tests based on the response of the slope of the United States yield curve to news offer no comparable evidence of any change in market perceptions of the inflation aversion of the Federal Reserve.

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1. Introduction

The perception of the inflation aversion of a central bank plays a key role in determining whether its goal of low inflation is attained. This point is, by now, a standard theoretical result.[1] It is also received wisdom among practitioners. 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."

This consensus on the importance of the perception of inflation aversion naturally leads to the question of how it is achieved, 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 strong 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 majority of respondents to Blinder's survey believe that central bank credibility is based more on its history of actions than on the construction of institutional structures that insulate a central bank from political concerns and afford it independence. Nonetheless, there is also a consensus among respondents that structure matters. This latter view is consistent with empirical research that has found that institutional structure is associated with economic performance in cross sections of countries, perhaps because it indicates the ability of an institution to "tie its hands" and commit to a policy that may cause short-term pain in

[1] Seminal contributions on the role of credibility include Kydland and Prescott (1977), Calvo (1978), and Barro and Gordon (1983).

[2] 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).

the pursuit of longer-run gain.[3] There is less evidence, however, on whether and how the credibility of a particular central bank evolves over time in response to the conduct of policy. The goal of our paper is to use high-frequency data from asset markets to address this issue. Thus, this work contributes to a growing literature that focuses on shifts in the market assessment of the policy stance of a central bank and central bank efforts to obtain and keep policy credibility.

Questions relating to the achievement and the maintenance of inflation aversion credibility are especially relevant for a new central bank, or a central bank that has a change in leadership. An analysis of the experience of the European Central Bank (ECB) during its early years of operation provides a natural experiment for considering this question. The architects of the institutional structure of the ECB were mindful of lessons from economic theory concerning the importance of independence from political considerations.[4] The role of conduct was also clearly apparent. As indicated by the survey results in Blinder (2000), the directors of central banks are vitally aware that their policies are closely scrutinized for indications of general tendencies. This may be especially true with a new central bank where each policy choice can lead to a larger updating of market priors than would be the case for a long-established central bank.

This paper starts, in Section II, with the insight that the response of asset prices to inflation news reflects market perceptions of the policy stance of the central bank.[5] In particular, Ellingsen and Söderström (2001) have demonstrated that a surprise increase in

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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 (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)

[5] Forward market information has been used in other tests of policy regime credibility. For example, Svensson (1991) shows that forward exchange rates were not within the target zone band of the European Monetary System (EMS) in the 1980s, a result he interprets as indicating that the EMS generally did not offer credible bands on its members' currencies. Svensson (1993) presents a similar set of tests to determine whether the inflation targets of Canada, New Zealand and Sweden were consistent with market yields.

inflationary pressures should result in a greater increase in a long interest rate relative to a short interest rate when a central bank is perceived as being more tolerant of inflation and less credible as an inflation fighter.[6] These results suggest that if the credibility of a central bank is earned through its conduct of policy, one would expect to find a significant change in the relationship between inflation news and the term structure, as well as news and the exchange rate, as credibility evolves.[7] Alternatively, if the public's view of a central bank's anti-inflation stance arises from a static institutional structure, then there should be no evidence of a change in the responsiveness of the term structure or of the exchange rate to news.[8]

In Section III we study evolving central bank credibility by applying newly developed tests for persistent time variation in regression coefficients (from Elliott and Müller 2006) to high frequency financial market data. Specifically, we explore the evolution of the credibility of the European Central Bank from the time it began its operations in January 1999 through mid-2005. The regressions utilize hourly data on the term structure of bonds of euro-area countries and the United States, and on the euro-dollar exchange rate, and regression analysis tracks associated reactions to economic news announcements. The new econometric techniques applied are especially relevant for this context since they allow a gradual evolution of parameters in a manner consistent with results from theoretical models supporting gradual expectations updating by market participants.

The estimates presented in Section III suggest that the market's perception of the inflation aversion of the ECB has evolved over time. We find evidence of significant persistent parameter instability for European term structures and the euro-dollar exchange rate. As a benchmark, we also look for evidence of persistent parameter instability in the response of the US term structure to news, but find that this relationship was stable during this period, a result consistent with the view that the Fed's long-standing commitment to

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[6] A quantitative assessment is presented in Ellingsen and Söderström (2004).

[7] Of particular relevance for this paper, which studies the effect of announcements on the slope of the yield curve, is work that links the empirical failure of the pure expectations hypothesis of the term structure of interest rates and the excess sensitivity of long maturity yields to changes in monetary policy, such as Fuhrer (1996), Gürkaynak, Sack and Swanson (2005), and Kozicki and Tinsley (2005).

[8] Klein, Mizrach and Murphy (1991) develop a similar type of analysis concerning differences in the responsiveness of asset prices to news as policy evolves in their study of the changing responsiveness of dollar exchange rates to news about the United States current account. They find the 1985 Plaza Accord altered perceptions of the degree to which American policy was concerned with the U.S. current account deficit.

price stability anchored market expectations at this time.[9] We also find no significant change in the linkages between U.S. and euro area inflation during this time.

These combined findings are consistent with market participants updating their views of the anti-inflation stance of the European Central Bank. We provide additional support for this conclusion based on the smoothed time path of the estimated parameters of the coefficient on the news announcement. For this work, we implement a new econometric technique from Müller and Petalas (2005). The results on rotations of the yield curve in response to news show that monetary tightening by the ECB altered the estimates of the market's perception of the ECB's anti-inflation stance. Discrete structural break tests (from Andrews 1993) verify the robustness of these results on the presence and dating of a change in market perceptions of the anti-inflationary stance of the ECB.

Thus, overall, this paper provides support for a dynamic view of the evolution of the perceived policy stance of the European Central Bank, with enhanced perceptions of credibility as the ECB tightened policy in the early years of its existence. The evidence presented here is consistent with the view voiced by many of Blinder's respondents that actions, and not just institutional structure, influence perceptions of the policy stance of a new central bank. The tools we apply to high frequency data from bond markets and news announcements provide a set of methods to track central bank credibility and market perceptions as it evolves.

[9] We do not find evidence of parameter instability for the response of the term structure of US interest rates to news for our sample period, January 1999 to June 2005. Those papers that find a role for a shift in Fed policy focus on longer and earlier sample periods. For example, the sample for Fuhrer (1996) is 1966 to 1994, the sample for Gürkaynek, Sack and Swanson (1995) is 1990 to 2002, and the samples used by Kozicki and Tinsley (2005) begin in 1946, with the longest one ending in 1997.

2. Central Bank Policy and Market Responses to News

In this section we present a specification that allows for a change in perceptions about the policy stance of the central bank to alter the response of asset prices to news. The motivation for this specification begins with the long-established fact that asset prices respond to news. But the magnitude of this effect, and even its direction, may depend upon the perceived policy stance of the central bank. This is especially true with respect to the term structure of interest rates. For example, a number or empirical studies have shown that the failure of the Pure Expectations Hypothesis of the term structure of interest rates for the United States can be attributed to changing views of the policy stance of the Federal Reserve (e.g. Fuhrer 1996, Gürkaynek, Sack and Swanson 1995, Kozicki and Tinsley 2005).

2.1 Theoretical Underpinnings

The theoretical work of Ellingsen and Söderström (2001) shows how the response of the term structure of interest rates to news varies with a change in the perceived policy stance of the central bank. The model assumes central bank private information on the state of the economy or the inflation target of the central bank. A dynamic aggregate demand – aggregate supply model with Svensson (1997)-style inertia in inflation and output is assumed, with forward-looking monetary policy that is set optimally with respect to a quadratic loss function for the central bank. Ellingsen and Söderström augment this model with a term structure equation. The focus of their work is on how monetary policy decisions may both reveal private information about the state of the economy and reflect the unknown policy stance of the central bank.

The key result in Ellingsen and Söderström (2001) for our work is their finding that with "… a more inflation-averse central bank, short rates will respond more, but long rates less to a given shock." (p. 1601). In their model, both a positive demand shock and a positive supply shock leads to a smaller effect on the slope of the yield curve when the central bank is perceived to be strongly averse to inflation than when it puts relatively greater weight on output deviations rather than inflation deviations in its loss function.[10] The intuition behind this result is that a positive inflation shock leads to an expectation of an increase in the short-term interest rate which causes a decrease in economic activity. A central bank that is more inflation averse (that is, one that places relatively more weight in its loss function on inflation, and less on deviations of output from its target) responds more strongly to a positive inflation shock, and creates a deeper recession, than a central bank that is less inflation averse. In the long run, a given inflationary shock will persist for less time in the economy if the central bank is more inflation averse, so future short rates will increase less with a given shock if the central bank is more inflation averse. Thus, the slope of the yield curve responds less to a given inflationary shock when a central bank is more inflation averse.[11]

This theoretical result provides the foundation of the specification used in our empirical analysis.[12] To translate the result from this model to an empirical specification, we begin by writing the differential of the term structure as

$$(1) \qquad d(i_t^L - i_t^S) = (\frac{di_t^L}{d\eta} - \frac{di_t^S}{d\eta})d\eta$$

where i_t^L is a long interest rate, i_t^S is a short interest rate, and $d\eta$ is the shock to either demand or supply.

The effect of a change in the inflation aversion of the central bank on the responsiveness of the term structure to the shock turns on the relative size of $d\left(\frac{di_t^L}{d\eta}\right)\Big/ d\lambda$

[10] The supply shock is the innovation to an equation in which inflation depends upon its lagged value as well as lagged output. The demand shock is the innovation to an equation in which output depends upon its lagged value and the lagged real interest rate. Both shocks are assumed to be i.i.d. with mean zero, which is relevant for our analysis since the shocks we study are the surprise component of news announcements.

[11] Ellingsen and Söderström write that, due to the inertia in inflation and output "even a large dose of forward-looking behavior does not destroy our qualitative results." (p. 1598) and they note that purely forward looking models are at odds with the data, citing Estrella and Fuhrer (2002). Ellingsen and Söderström also investigate how differences in the information available to the public and to the central bank affects the response of the slope of the yield curve to monetary policy shocks, but the information structure does not affect the response of the slope of the yield curve to inflation news.

[12] 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. See Huizinga and Mishkin (1986) and Fleming and Remolona (1999).

and $d\left(\frac{di_t^S}{d\eta}\right)\Big/ d\lambda$ where λ is the relative weight on deviations of output from its target, compared with deviations of inflation from its target, in the central bank loss function in Ellingsen and Söderström (2001) (and therefore, a central bank with a larger value of λ is less inflation averse). Proposition 2 of that paper shows that

(2) $$d\left(\frac{di_t^L}{d\eta}\right)\Big/ d\lambda > d\left(\frac{di_t^S}{d\eta}\right)\Big/ d\lambda$$

and, therefore, a positive inflationary shock will elicit more of a response at the long end of a yield curve, with the slope of the yield curve changing accordingly, when a central bank is perceived to be less inflation averse (that is, when it has a larger value of λ) than when it is perceived to be more inflation averse.

This result can be mapped to the standard linear specification linking the surprise component of news to the change in an asset price. The specification for the effect of news on asset prices, allowing for the possibility of a time-varying coefficient on the news variable, is

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

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. As emphasized in Andersen, Bollerslev, Diebold and Vega (2003) 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.

We identify γ_i, $q_{t^+} - q_{t^-}$, and $\left(x_{t^+} - E_{t^-} x_{t^+}\right)$ in (3) with the terms in equation (1) $(\frac{di_t^L}{d\eta} - \frac{di_t^S}{d\eta})$, $d(i_t^L - i_t^S)$, and $d\eta$, respectively. The inclusion of a subscript for the coefficient (that is, γ_i rather than γ) indicates that it can vary over time in a manner consistent with the theoretical result in (2) that the responsiveness of the term structure to inflation news changes as perceptions of the inflation aversion of the central bank evolve. Also note that the perceived inflation aversion, rather than the actual inflation aversion, matters for γ_i because the very short-run response of an asset price to news is driven purely by private expectations and perceptions of policy. In particular, given the result in (2), γ_i decreases as the public perception of a central bank's aversion to inflation gaps relative to output gaps rises (in terms of the Ellingsen and Söderstrom (2001) model, as the public's perceived value of λ decreases).

2.2 Empirical Strategy

To illustrate our empirical strategy we first present a particularly stark example and discuss what we would expect to find in a regression on the term structure. We then discuss a more realistic situation and describe how the econometric tests we employ allow us to test for the presence of time variation in λ_i in that case.

Suppose a dramatic policy action undertaken at time *T* by a central bank significantly increases its perceived aversion to inflation from that moment on. Before time *T*, the public perception of the relative weight placed on output in the central bank's quadratic loss function is λ_D (for dovish) and after the policy action it is λ_H (for hawkish) where $\lambda_D > \lambda_H$. The coefficients in (3) associated with these parameters are γ_D and γ_H , respectively. The result from Ellingsen and Söderström (2001) presented in (2) shows that $\gamma_D > \gamma_H$. Therefore, with knowledge of the date *T*, one could simply perform a Chow test for a statistically significant difference between γ_i before and after this date.

While this example shows how we might expect the coefficient in (3) to change over time, it offers an unrealistic picture of the likely evolution of perceptions of the anti-inflation stance of the European Central Bank. The two key differences between this example and a more realistic framework are the lack of a single, widely-recognized dramatic change in policy that was clearly a watershed, and the likelihood that perceptions changed gradually rather than once and for all.[13] A more articulated model could specify the manner in which the market's view of the stance of the central bank evolves over time in response to policy.

But we need not constrain ourselves to one type of model or another to implement the main econometric technique that we use for identifying the changes over time in the slope of the yield curve. 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.[14] 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 (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.[15]

We would like to ensure that the persistent time variation in γ_i is due to variation over time in the perceived inflation aversion of the ECB rather than, say, variation over time in the direct responsiveness of the change in asset prices to the surprise component of news. In the wake of the creation of a new central bank, such as the ECB in January 1999, it is reasonable to expect that the most likely source of a time variation in γ_i is changes in the perceived inflation aversion. In fact, we do find evidence that there has been persistent time

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[13] 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, for a more recent contribution, Athey, Atkeson, and Kehoe (2005).

[14] 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)

[15] 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.

variation in the slope coefficient in term spread regressions for Germany, France and Italy, and we interpret this as reflecting an evolving view of the inflation aversion of the ECB.

This interpretation is bolstered by some other tests. Evidence will be presented that shows no persistent time variation in γ_i for the slope of the U.S. term spread. This test using the U.S. yield curve is offered as a benchmark; were we to find evidence of parameter instability for regressions based on this series, we would be concerned that evidence of parameter instability using European bond yields may not, in fact, reflect an evolving perception of ECB inflation aversion but, rather, some structural change common to financial markets across all four of these industrial countries. As a further robustness check, we test for the changing effects of news announcements on the bilateral euro/dollar exchange rate.[16] Furthermore, we will also present evidence that the relationship between U.S. inflation and inflation in the euro-area has not exhibited persistent time variation; given the use of U.S. Core CPI news, this then helps isolate inflation aversion as the source of the persistent time variation in γ_i.[17]

Our interpretation that the persistent time variation in the slope coefficient for European term spreads and the euro/dollar exchange rate reflects evolving perceptions of the inflation aversion of the ECB is also bolstered by comparing estimates of the smoothed time path of γ_i (using the method developed by Müller and Petalas 2005) to actual changes in ECB policy. We will show that the decrease in the estimated values of slope coefficients occurs in wake of monetary tightening by the ECB. The results from these smoothed estimates are supported by the sup-Wald tests for parameter stability (see Andrews 1993, 2003) which offer a break date for γ_i that roughly corresponds to the peak value of the estimated smoothed time path. Furthermore, the presence of a statistically significant break date supports the conclusion of the Elliott and Müller (2006) test concerning the persistent time variation of the coefficient.

[16] 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 United States and euro-area markets over the sample period.
[17] We will 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.

3. Implementation to Evolving Market Perceptions of the European Central Bank

In this section we first discuss, in Section 3.1, the data we use for our tests. The results of the Elliott and Müller (2006) test are presented in Section 3.2. 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 United States bonds. In Section 3.3 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. Finally, in Section 3.4, 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 (from Andrews 1993) for a discrete break in the regression relationship and the dates associated with those breaks.

3.1 Data

The two types of data used in our analysis are various asset prices, where the assets are government bonds and foreign exchange, and inflation announcements and related market expectations. 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 United States government bonds are four of the dependent variables. The fifth dependent variable is where $q_{t^+} - q_{t^-}$ represents the change in the logarithm of the euro /U.S. dollar exchange rate, thirty minutes before and thirty minutes after the news announcement. 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. Combined with the yield curve results, the γ_i decrease on the euro/ U.S. dollar exchange rate reflects a stable perception of the stance of the Fed combined with a perception of an increasingly anti-inflation stance of the ECB .

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.

Although our primary analytical emphasis is on evolving credibility of the ECB, and we use data on European yield curves, our analysis utilizes data on the U.S. core CPI release, not European inflation reports. There are three well-established reasons for this choice. First, the euro-area inflation series – *both* the announcements and expectations of the announcements -- were not available at the time of the introduction of the euro in January 1999 and 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. Second, the actual news content and market impact of inflation announcements from individual European countries has been of limited value to markets because of issues related to data quality and data leaks prior to official announcement times. This widely documented finding is consistent with regression findings that show that the European price announcements have low news content and have weak effects on asset prices in Germany, Italy, France, the United Kingdom, and Europe as a whole.[18] A third reason for the use of U.S. rather than European data is based on a positive, rather than a negative, finding. As documented in Andersen *et al.* (2003), Goldberg and Leonard (2003), Faust, Rogers, Wang, and Wright (2006), and Ehrmann and Fratscher (2005), U.S. announcements have strong news content, and they have large and significant effects on both United States and European

[18] Recent studies include Goldberg and Leonard (2003),and Ehrmann and Fratscher (2005).

asset prices. Financial market studies document this phenomenon as well.[19] One reason exposited for this impact is that U.S. inflation news contains information that is perceived as relevant for European inflation.[20]

We have selected the United States Core CPI as the series for our analysis of the interaction between news, asset prices, and changing central bank credibility or inflation aversion. The core CPI data are among the one most closely followed by the market, both in terms of significance in econometric studies and as reflected by the fact that, in recent years, this series, as compared with a range of alternative price news, was typically the only inflation series discussed by Alan Greenspan in his Humphrey – Hawkins testimony.[21]

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.[22] 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 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.

3.2 Time Variation in the Effects of News on the Slope of 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 above, the *qLL* statistics

[19] A recent market studies supporting these conclusions is Citigroup (2006).
[20] 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.
[21] See Clark (2001) for evidence on this point and for an overview of related literature.
[22] 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. Gürkaynek and Wolfers (2005) show that these data have been among the best performing expectations series for important macroeconomic variables over the sample period that we analyze.

are negative. Evidence of an evolving view of the policy stance of the ECB over time is given 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 US term spread the value of the *qLL* statistic larger than its critical value is evidence against persistent time variation in the perception of λ, the anti-inflation stance of the Federal Reserve over this period.

Results of this test are presented in Table 1.[23] 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.

Table 1: Elliott-Müller Test for Persistent Time Variation

Test of Time Variation of	Change in Term Spread of Government Bonds of				Change in Euro/$
	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*

The results in Table 1 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

[23] 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.

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 United States 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 for significant at the 95 percent level of confidence.

All of these results are consistent with the model presented above in which γ_i varies with an evolving view of the inflation aversion of the European Central Bank 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.

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 US 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 US as well as European countries.

One might argue that these results are consistent with a situation where the signal value of the inflation news changed over time due to changing fundamental relationships between the eurozone and the United States. This could arise, for example, if a change in the relationship between U.S. and European inflation rates occurred over this period since the announcements are, in all regressions, of U.S. Core CPI as the η_t. The results presented in Table 2, however, show that this is not the case. This table presents tests of significant time variation in the coefficient of regressions of monthly inflation in the Euro area, Germany, France and Italy as a function of US monthly inflation. As shown in that table, there is no evidence of significant time variation of the coefficient on monthly United States inflation in 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 all cases but for Germany, there is also a highly significant relationship between U.S. inflation and the inflation rate in the European countries, or the euro area as a whole.

Table 2: Elliott-Müller Test for Persistent Time Variation in Effect of US Inflation on Inflation in euro area, Germany, France, and Italy

$\pi_i = \alpha + \beta\pi_{US} + \varepsilon$	Monthly Inflation in			
	Germany	France	Italy	Euro Area
qLL for β	-3.16	-4.29	-6.91	-1.86
α (s.e.)	0.117 (0.038)	**0.055** (0.026)	***0.163*** (0.0137)	***0.107*** (0.027)
β (s.e.)	0.021 (0.100)	***0.362*** (0.069)	***0.117*** (0.036)	***0.275*** (0.072)

Critical Values for *qLL*: 1% ***-11.05***; 5% **-8.36**; 10% *-7.14*
Regressions run using monthly inflation data for period January 1998 – Dec. 2005 (96 obs.)
Bold and Italic = significant at 99% level, **Bold** = significant at 95% level,
Italics = significant at 90% level.

The evidence in these tables is suggestive of an evolving perception of the policy stance of the European Central Bank. This conclusion is bolstered by the estimated time path of γ_i presented in the next section.

3.3 Estimated Paths of γ_i

The central hypothesis in this paper is that the perceived anti-inflation stance of the European Central Bank 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 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 1 presents the estimated smoothed parameter paths of γ_i for each of the four term spreads and Figure 2 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 United States and German term spreads from Figure 1. The time paths are calculated using the technique developed by Müller and Petalas (2005), 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.[24] The dashed vertical lines in these figures indicate dates of interest rate cuts by the ECB, and the solid vertical lines demark 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 1 shows that the estimated value of γ_i at the beginning of the sample is greater for the three European government bonds than for the United States bond. This is consistent with the view that, at the outset of the operation of the European Central Bank, the market perceived the ECB as more willing to tolerate inflation than the Federal Reserve.[25] Another immediately apparent characteristic of the four time paths in Figure 1 is the relative variability of the three European γ_i 's as compared to that of the United States which, of

[24] 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 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.

[25] 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.

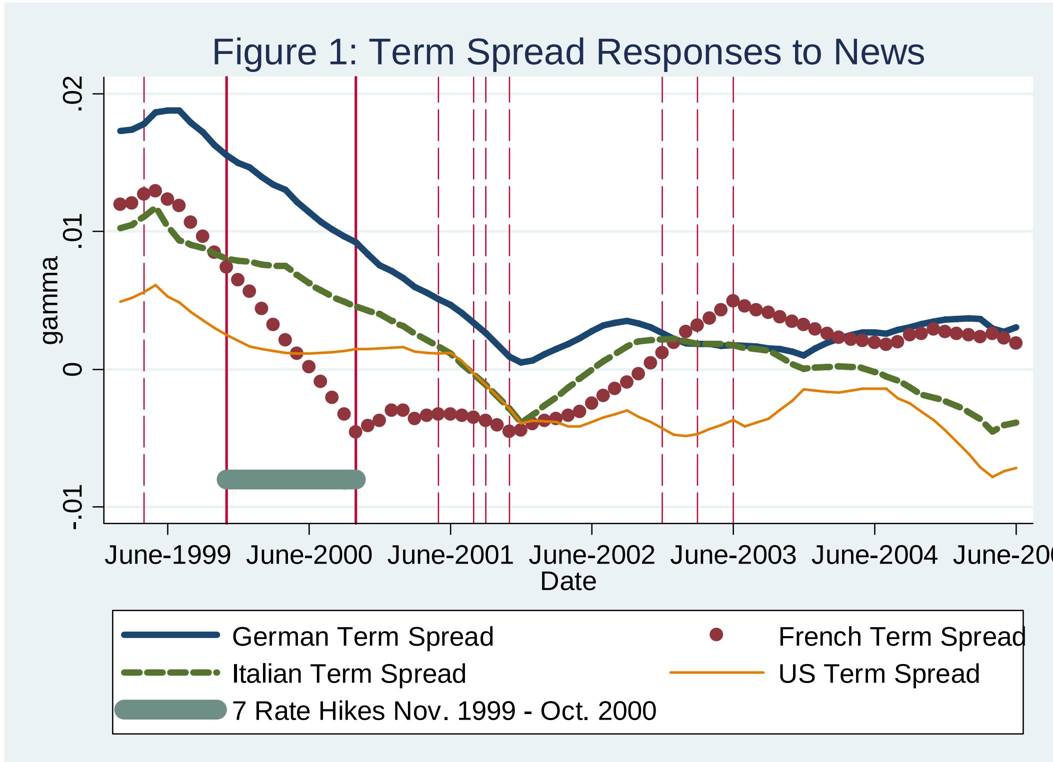
Figure 1: Term Spread Responses to News
gamma
-.01
0
.01
.02
June-1999
June-2000
June-2001
June-2002
June-2003
June-2004
June-2005
Date
German Term Spread
French Term Spread
Italian Term Spread
US Term Spread
7 Rate Hikes Nov. 1999 - Oct. 2000

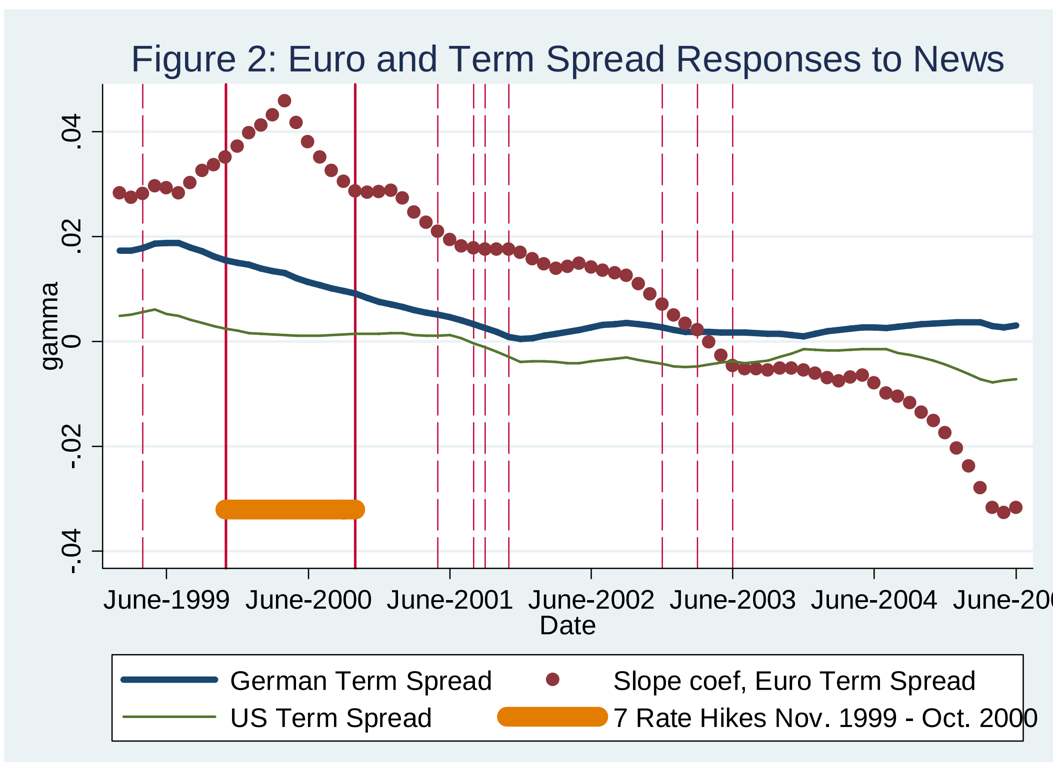
Figure 2: Euro and Term Spread Responses to News
gamma
-.04
-.02
0
.02
.04
June-1999
June-2000
June-2001
June-2002
June-2003
June-2004
June-2005
Date
German Term Spread
Slope coef, Euro Term Spread
US Term Spread
7 Rate Hikes Nov. 1999 - Oct. 2000

course, is a reflection of the results of the Elliott – Müller tests presented in the previous section.[26]

The time variation of the estimated paths of γ_i in light of both the actions undertaken by the European Central Bank 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 rather than, say, changes in the links between U.S. and European inflation. In order to make this point, we offer in Table 3 an overview of the policy of the European Central Bank 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 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 moves, and the growth of real GDP in the quarter preceding the policy moves.[27] We also will refer to conclusions on ECB conduct presented in various annual 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 the 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%.[28] 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 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)

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[26] The standard deviations of the estimated γ_i's are 0.0043 for Italy, 0.0047 for France, 0.0057 for Germany, and 0.0203 for the euro/dollar exchange rate, but only 0.0036 for the United States, all of which are consistent with the results of the Elliott – Müller *qLL* statistics presented in Table 1.

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

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	
April 28, 2000	3.75	1.93	8.47	4.38
June 9, 2000	4.25	1.73	8.21	
September 1, 2000	4.50	2.02	8.09	4.63
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	
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.

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."

Our estimates of γ_i in the early part of our sample are consistent with the view expressed in this CEPR publication. 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, Blanchard, Galí, Giavazzi and Uhlig, 2001) concluded that these interest rate increases marked a departure of ECB policy from its earlier pattern.[29] Accordingly, our estimates of γ_i for the European term spreads and the euro/dollar exchange rate declined through this period.

This decline in our estimates of γ_i for the German and Italian term spreads continued, and our estimate of 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)

This volume of *Monitoring the European Central Bank* called 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 respective values immediately in the wake of the interest rate cut in November 2001. Subsequently, there was a reduction in the

[29] 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."

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.

3.4 Sup-Wald Statistics

We conclude our analysis with sup-Wald tests for discrete changes in γ_i, based on Andrews (1993, 2003), to gauge 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 3.3.

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

$$(4) \qquad q_{t^+} - q_{t^-} = \alpha + \beta\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \beta_I D_I\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \varepsilon_{t^+}$$

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 (4), 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 United States government bonds.

Table 4: sup-Wald Test for Discrete Break Point

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

Critical Values (from Andrews 2003) 1% ***12.16***; 5% **8.68**; 10% *7.12*
Tests conducted with 15 percent symmetric trimming.

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 the 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 increase 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 suggest the robustness of the results presented in the previous sections.

4. 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 stance through its 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 inflation aversion of the European Central Bank 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 European Central bank provides a natural experiment for considering how the reputation of a central bank evolves over time. This episode is a particularly interesting because of the controversy surrounding the conduct of monetary policy in Europe as the ECB began its operations.

In this paper, we have addressed these issues by proposing and executing a novel test for the study of the evolution of market perceptions about the inflation aversion of a central bank through the use of high-frequency data. This methodology and the use of high frequency data provide a unique window into the evolution of perceptions of monetary policy rules, an issue more typically and less precisely addressed using lower frequency data. We find evidence of an evolution of perceptions of the policy stance of the ECB, one linked to its interest rate policy. There is 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 the broad support for its conduct of policy. The tools we have presented and applied are relevant for ongoing questions of changing effects of news on market activity, and changing policies and perceptions of monetary authorities worldwide.

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Appendix

The use of both European and American term structures suggests another test, following the specification in (3) but one where $q_{t^+} - q_{t^-}$ is defined 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 US, respectively, $E_t e_L$ is the expected logarithm of the euro/dollar exchange rate at a time in the future matching the 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

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

As discussed below, the same news variable will be used for both European and US 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 US term structure regressions, would bolster our conclusion of an evolving perception of the anti-inflation stance of the European Central Bank.