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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 includes 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 questions of the achievement and the maintenance of inflation aversion credibility are especially relevant for a new central bank. 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 with the insight that the responses of asset prices to economic news embed market perceptions of the policy reaction function of central banks. The relationship between asset prices and news evolves with the change in market perceptions of a central bank's monetary reaction function and its associated degree of inflation aversion. As argued by Bernanke (2004), "successful monetary policies should stabilize, or 'anchor,' inflation expectations so as to prevent them from becoming a source of instability in their own right". Therefore, with asset prices as a starting point, in Section II we present a framework for a novel test of the evolution of market perceptions of central bank inflation aversion. This framework uses high frequency asset price data on the slope of yield curves and on exchange rates, with the surprise components of economic data releases, to estimate whether the

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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] Despite these lessons, some politicians continued to try to influence policy direction. For example, 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.

market perception of the anti-inflation credibility of a central bank changes over time.[5] The key insight from this model is that a given surprise increase in inflationary pressures will result in a greater increase in a long interest rate relative to a short interest rate, and a larger exchange rate depreciation, when a central bank is perceived as being more tolerant of inflation and less credible as an inflation fighter. If unvarying institutional structure is the dominant determinant of a new central bank's credibility, then one would not expect to find a change in the high frequency relationship between economic news and the slope of the yield curve over time. But if credibility for a new central bank is earned through the conduct of its policy, one would find a significant change in the relationship between news and the yield curve as credibility evolves.[6]

In Section III we use a newly developed test for persistent time variation in regression coefficients (from Elliott and Müller (forthcoming)) to study the evolution of the credibility of the European Central Bank from the time it began its operations in January 1999 through mid-2005. Using hourly data on the euro-dollar exchange rate and on the term structure of bonds of euro-area countries, we find evidence that the market's perception of the inflation aversion of the ECB has evolved over time. We estimate the evolution of the market's view of the credibility of the ECB, using a technique developed by Müller and Petalas (2005), and find that monetary tightening by the ECB altered the estimates of the market's perception of its anti-inflation stance. The robustness of these results on the presence and dating of a change in market perceptions of the anti-inflationary stance of the ECB is demonstrated by additional tests for a discrete structural break (from Andrews 1993). Our conclusion is that these results reflect changing perceptions of the anti-inflation stance of the ECB rather than a changing economic environment. We support this interpretation by estimates demonstrating that over this period there was no change in the market's perception

[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. These tests, while informative, require the presence of an explicit target, like an exchange rate band or an inflation target, to judge credibility. Other related empirical analyses on the policy credibility of an exchange rate target zone use intervention data to estimate perceived target zone bands (Klein and Lewis 1993 and Lewis 1995).

[6] 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.

of the monetary reaction function of the Federal Reserve and that there was no significant change in the linkages between U.S. and euro area inflation. The combined findings are consistent with an explanation based on evolving views of the anti-inflation stance of the European Central Bank. By contrast, during this period market views of the Federal Reserve's stance were stable, given the Fed's long-standing commitment to price stability under the chairmanship of both Alan Greenspan and Paul Volker.

2. Central Bank Policy and Market Responses to News

In this section we present a model that shows how changes in perceptions about a policy stance can alter the response of asset prices to news. We begin with the standard framework used in empirical works that study the effect of news on asset prices, as in, for example, Anderson, Bollerslev, Diebold and Vega (2003). We then introduce a policy reaction function, and show how an evolving view of the credibility of central bank inflation aversion affects the relationship between news and asset prices. We discuss the empirical implementation of this model to the yield curve and exchange rates, so that we can use high frequency data to isolate the effects of news on asset prices under different market perceptions of central bank inflation aversion.

2.1 Empirical Specification

The standard linear specification linking the surprise component of news to the change in an asset price is

$$q_{t^+} - q_{t^-} = \alpha + \gamma\left(x_{t^+} - E_{t^-} x_{t^+}\right) + \varepsilon_{t^+} \tag{1}$$

where $q_{t^+} - q_{t^-}$ is the change in an asset price over the short period of time between t^-, just before an announcement, and t^+, just after that announcement, 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 that $x_{t^+} - E_{t^-} x_{t^+}$ is the surprise component of the announcement, and ε_{t+} is a white-noise error term. As emphasized in Anderson et al. (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. The actual set of variables that constitute x depends upon the asset studied but, in general, any variable that markets construe as revealing information about current and future economic activity may be appropriate for study.

A more general version of equation (1) takes into account market expectations about the policy response to news. Consider the path, from time t forward, of a policy M_t, which has an effect on q_t. We can augment (1) to include the effect on $q_{t^+} - q_{t^-}$ of the change in the perception, between time t^- and t^+, of the path of policy. This specification,

$$(2) \qquad q_{t^+} - q_{t^-} = \alpha + \gamma\left(x_{t^+} - E_{t^-} x_{t^+}\right) + \phi\left(E_{t^+} M_{t^+} - E_{t^-} M_{t^-}\right) + \varepsilon_{t^+},$$

captures the possibility that $q_{t^+} - q_{t^-}$ responds to economic news directly through the coefficient γ and indirectly through the coefficient ϕ due to the effect of the news on the expected course of policy, where the perceived policy path before the announcement occurs is $E_{t^-} M_{t^-}$ and its perceived path after the announcement is $E_{t^+} M_{t^+}$. The parameter ϕ may be positive or negative, depending upon the policy and the asset. Over a short window of time, the only reasonable source of a change in the perceived path of the policy over the short time span t^- to t^+ is the surprise component of the data announcement during this window. This link arises because of a perception of the existence of a policy reaction function, such as

$$(3) \qquad M_t = V_t - \lambda_i (x_t - \bar{x}_i)$$

where V represents other variables that affect the choice of M and the subscript i on the coefficient λ allows for the possibility of a different levels of responsiveness of the central bank to the value of x at different times in the sample period. Likewise, this formulation allows for more than one target level of the variable, $\bar{x}_i$, during the sample period.

To make the discussion of the policy reaction function more concrete, consider the case where (3) represents a Taylor Rule. In this case, M represents monetary policy (such that an increase in M represents a more expansionary monetary policy), x is inflation data, $\bar{x}_i$ is an inflation target, and V represents an indicator of other economic conditions, for example the output gap or unemployment. A higher value of λ represents more inflation aversion in the reaction function of the central bank.[7]

In the presence of the perceived policy reaction function, the surprise component of the perceived change in policy, $\left(E_{t^+}M_{t^+} - E_{t^-}M_{t^+}\right)$, is defined by

(4) $$\left(E_{t^+}M_{t^+} - E_{t^-}M_{t^+}\right) = (E_{t^+}V_t - E_{t^-}V_t) - \lambda_i(x_{t^+} - E_{t^-}x_{t^+}) + \lambda_i(E_{t^+}\overline{x}_i - E_{t^-}\overline{x}_i)$$

This expression shows that expected policy can depart from its prior path due to a change in the output gap, due to the surprise component of an inflation data release, or due to a change in the target value of inflation.

Provided that there is some news in an inflation report relative to market expectations, $(x_{t^+} - E_{t^-}x_{t^+})$ is not equal to zero. In a short window of time around the inflation report, for example an hour window, in the absence of a simultaneous announcement of a change in target inflation, or of news on real economic variables such as the output gap, the other two right hand side terms of equation (4) can be set equal to zero.[8] Within this equation, the parameter λ_i associated with market views of central bank inflation aversion may change over time, and, indeed, testing for time variation in λ_i is the central empirical task in this paper.

Consider the response of bond prices to economic announcements. The effects of news announcements, such as inflation reports, on bond prices have been extensively studied. At each horizon, nominal returns are comprised of the real return, an inflation expectation, and a risk premium. Bond returns and inflation expectations are linked through the Fisher relationship. With news effects on equilibrium real interest rates common to returns at all

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[7] While this is the most common formulation of the policy reaction function, alternative formulations can of course be specified. For example, λ can be modeled as asymmetric in that it is larger when inflation exceeds targeted values, or inflation deviations from target can be entered nonlinearly, so that monetary reactions are strongest when inflation is furthest from target values.

[8] We have $(E_{t^+}V_t - E_{t^-}V_t) = 0$ if x and V are uncorrelated, but, even if this is not the case, the qualitative effects discussed below are not affected if x and V are negatively correlated. If x and V are positively correlated, we would expect the sign on λ to be positive in the policy reaction function. We would not expect, in a sample with many observations, many instances where $\lambda_i(E_{t^+}\overline{x}_i - E_{t^-}\overline{x}_i)$ does not equal zero since a nonzero value for this term would mean that the news announcement itself alters the view of the target value $\overline{x}_i$ in the short time interval t^- to t^+. Even if, say, an unusually large value of the surprise component of the news alters market participants' perceptions of the target value in one or two instances in a sample with many observations, this would leave $\lambda_i(E_{t^+}\overline{x}_i - E_{t^-}\overline{x}_i) = 0$ for the vast majority of cases.

horizons along the yield curve, when we difference across the returns of long and short-dated bonds we abstract from the effect of news on equilibrium real returns. Over short windows around news announcements, differencing also abstracts from the effect of news on term premia or liquidity premia. In this case, the regressand $q_{t^+} - q_{t^-}$ is the expected change in the differential in the long horizon (for example, 10 years) and short horizon (for example 2 years) inflation rates due to the news announcement

$$q_{t^+} - q_{t^-} = (E_{t^+}\pi_{t^+}^{10} - E_{t^+}\pi_{t^+}^{2}) - (E_{t^-}\pi_{t^-}^{10} - E_{t^-}\pi_{t^-}^{2}) \tag{5}$$

where π_t^{10} is the average expected inflation rate over 10 years and π_t^2 is the average expected inflation rate over 2 years.[9] The change in the slope of the yield curve over short horizons reflects the evolution of inflation expectations across the term structure.

Substituting (4) into (2), and interpreting $q_{t^+} - q_{t^-}$ as the slope of the yield curve, we get

$$q_{t^+} - q_{t^-} = \alpha + (\gamma - \phi\lambda_i)(x_{t^+} - E_{t^-}x_{t^+}) + \varepsilon_{t^+} . \tag{6}$$

This equation suggests that, in the presence of changing perceptions about the central bank's policy, the estimated coefficient $(\gamma - \phi\lambda_i)$ on inflation news in regressions of the slope of the yield curve is unstable.[10]

We can be more precise about the instability of $(\gamma - \phi\lambda_i)$ by considering the simple case where a policy action undertaken by the central bank, such as a major tightening, changes the views of market participants. Suppose that before this action, market participants thought that the central bank policy was dovish (D), that is, that the central bank generally accommodated inflationary shocks, while after the action there was the view that the central bank would be more aggressive or hawkish (H) in combating inflation. These policies are distinguished by the condition that $\lambda_H > \lambda_D$. Evolving perceptions of central bank inflation aversion would be reflected in the coefficient on inflation news. For a central bank that

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[9] Fleming and Remolona (1999) also argue that the effects of news on asset prices of different maturities reveals information about market participant beliefs about central bank reaction functions. Early research by Huizinga and Mishkin (1986) applied to monthly data for the United States recognizes the sensitivity of the slope of the yield curve to perceptions of monetary policy regimes in the 1970s and 1980s.

[10] If news alters the perception of the target level of *x*, we could find evidence of a time-varying intercept as well since, in that case, we would have $q_{t^+} - q_{t^-} = (\alpha + \phi\lambda_i(E_{t^+}\bar{x}_i - E_{t^-}\bar{x}_i)) + (\gamma - \phi\lambda_i)(x_{t^+} - E_{t^-}x_{t^+}) + \varepsilon_{t^+}$.

gained market credibility as an inflation fighter, we would expect to find a larger value for the estimated coefficient in the earlier period as compared to the later period since $(\gamma - \phi\lambda_D) > (\gamma - \phi\lambda_H)$. If perceptions of inflation aversion were unvarying, perhaps because these perceptions were solely driven by the initial and unvarying institutional structure, we would expect to find a stable relationship between the slope of a central bank's yield curve and inflation news.[11]

This example, pointing to a discrete change in perceptions of λ and the regression coefficient, offers a particularly stark view of shifts in market perceptions of the central bank reaction function. An evolving view of central bank policy, one reflecting a gradual learning process, may be more consistent with reality and with theories of central bank credibility formation.[12] In either case, a more complete depiction of this model would specify the way in which the market's view of the stance of the central bank evolves over time in response to policy. The advantage of the main econometric technique that we use for identifying the changes over time in the slope of the yield curve is that it does not require us to specify this learning process nor the associated evolution of $(\gamma - \phi\lambda_i)$. Instead, this new econometric technique developed by Elliott and Müller (forthcoming) allows us to test for a very general form of persistent parameter instability over the sample period. An associated method of estimating the smoothed parameter path under very general assumptions, from Müller and Petalas (2005), enables us to identify the timing of changes in the estimated values of $(\gamma - \phi\lambda_i)$, which we then proceed to link to the policy changes of the European Central Bank.

3. Evolving Perceptions of European Central Bank Policy

In this section we present the data, methodology, results of our tests for changes in $(\gamma - \phi\lambda_i)$, and argue that the changes we observe are the results of changing market perceptions of the anti-inflation stance of the European Central Bank during its first six-and-

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[11] An alternative explanation for persistent changes in $(\gamma - \phi\lambda_i)$ could be changes in the economic environment, as represented by changes in γ or ϕ. We will show in the next section, however, that there is no statistical support for the contention that changes in these coefficients are the source of the persistent time variation of $(\gamma - \phi\lambda_i)$.

[12] For example, see Backus and Driffill (1985) or, for a more recent contribution, Athey, Atkeson, and Kehoe (2005).

one-half years of operation, from January 1999 through June 2005. We begin, in Section 3.1, with a description of the data we use for the econometric tests. Five different dependent variables are examined: the change in the term spread (alternatively called the change in the slope of the yield curve) for German, French and Italian government bonds, the change in the Euro/dollar exchange rate, and the change in the term spread of United States government bonds. The tests for possible parameter instability of the United States term spread 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. Likewise, an absence of parameter instability in the euro-dollar exchange rate regression could support a common structural change across U.S. and euro-area markets. However, as is shown in Section 3.3, we find no evidence of parameter instability for the regressions using the United States term spread series while we do find significant evidence of a change in the term spread for the tests using the other four series. In Section 3.4 we present estimates of the time path of $(\gamma - \phi\lambda_i)$. After eliminating competing explanations for observed changes in the time path, we show that the timing of changes in the path correspond to actual policy changes undertaken by the ECB. Finally, in Section 3.5, 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 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. The change in the term spread between 10-year and 2-year interest rates, $q_t = r_t^{10} - r_t^2$, for French, Italian, German, or United States government bonds are four of the dependent variables. The regressand $q_{t^+} - q_{t^-}$ is the change in the term premia, which under the Fisher relationship would have the interpretation as the change in the differential between the expected ten-year and two-year inflation due to the inflation news. Thus, when using these bond yield series, $(\gamma - \phi\lambda_i)$ can be interpreted as a function of the direct effect (represented by γ) and the indirect effect (via a policy response, as represented by $\phi\lambda_i$) of current inflation news on long-run relative to short-run expected inflation.

There is a similar interpretation of $(\gamma - \phi\lambda_i)$ in the regressions that use the fifth dependent variable, where $q_{t^+} - q_{t^-}$ represents the change in the logarithm of the euros per U.S. dollar exchange rate, thirty minutes before and thirty minutes after the news announcement. In this case, a positive value of $q_{t^+} - q_{t^-}$ indicates a depreciation of the euro, reflecting either an increase in expected inflation in Europe relative to the in the United States or an expectation of relatively more tightening of monetary policy in the United States than in Europe.[13] In this case, evidence that $(\gamma - \phi\lambda_i)$ decreases over the sample period can reflect a view among market participants that there is a relatively stronger anti-inflation stance of the ECB as compared to the U.S. Federal Reserve during this period, which can result from a strengthening of the anti-inflation policy stance of the ECB, a weakening of the anti-inflation stance of the Fed, or some combination of the two. We conduct tests on bond yields of the euro-area countries and the United States to help isolate the source of the change of the responsiveness of the euro/dollar exchange rate to news.[14]

Inflation Announcements and Expectations: Across the euro area and the United States, candidate inflation announcement measures for our study include measures of consumer price inflation for the full euro area, for individual countries in the euro area, and

[13] The change in the exchange rate over this short time horizon is a reflection of a change in the expected value of the exchange rate at some longer horizon which, in turn, reflects some form of long-run purchasing power parity.
[14] A similar joint use of interest rate data and exchange rate data was used by Engel and Frankel (1984) to analyze the response of interest rates to monetary announcements.

for the United States. For constructing "news" to which financial markets react, the study also requires measures of market expectations for the full period of analysis starting with the inception of the euro.

The *news* or surprise component of an economic data release is the difference between the actual release 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.[15] The ECB discusses policy in terms of the total harmonized index of consumer prices (HICP), and when core inflation is viewed as a signal about future HICP, this refers to the total index excluding energy and unprocessed foods. From Europe, candidate series for our analysis are the eurozone CPI excluding energy food and tobacco, the eurozone harmonized CPI, or specific inflation series from individual countries, like the French final total CPI, or German producer price index or wholesale price indices. From the United States, candidate inflation series are the consumer price index or the core CPI.[16] Although our primary analytical emphasis is on evolving credibility of the ECB, and we introduce European bonds for constructing yield curves, our analysis utilizes the U.S. core CPI release, not European inflation reports.

A number of reasons underlie this choice of a news variable. First, the euro area inflation series – *both* announcements and expectations of the announcements -- were not available at the introduction of the euro. Using these series would lead us to miss the critical early years when market participants were forming expectations of the monetary policy preferences of the ECB. Second, the alternative of using individual euro area country inflation data is not supported by financial market outcomes. Market participants argue that the news content of some European inflation announcements has at times been questionable because of issues of data quality and episodes of data leaks prior to official announcement

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[15] 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. Gurkaynek 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.

[16] The available series are: EuroZone: CPI ex Energy, Food & Tobacco Y/Y %Chg, As First Reported (NSA, %); EuroZone: Harmonised CPI M/M %Chg, As First Reported (NSA, %); France: Final Total CPI M/M %Chg, As First Reported (NSA, %); Germany: Producer Price Index: Mfg M/M %Chg, As First Reported (NSA, %); Germany: Wholesale Price Index M/M %Chg, As First Reported (NSA, %); U.S.: Consumer Price Index, Month/Month Change, Last Actual (SA, %) ; U.S.: Core CPI, M/M %Chg, As First Reported (SA, %) .

times.[17] This widely discussed issue is consistent with regression findings that show that the European price announcements have low news content and have weak effects on asset prices from Germany, Italy, France, or Europe as a whole.[18]

By contrast, U.S. news announcements have strong news content, and have large effects on both United States and European asset prices.[19] The use of U.S. inflation news in asset price regressions in regressions of euro area data also is justified by its importance as an indicator of pending euro area inflation.[20] Over our estimation horizon, U.S. inflation provides statistically significant information for euro area inflation.

The news variable we have selected is the difference between the monthly announcement of the Core Consumer Price Index for the United States and the expected value of this announcement prior to its release, as measured by survey responses. The closely watched core CPI is the best inflation measure for this analysis, as evidenced by the impact of related news on markets, the theoretical literature on prices and monetary policy, and Humphrey Hawkins testimony by Alan Greenspan in recent years, where the CPI excluding food and energy is typically the only measure of price inflation discussed.[21] 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 *news* has mean 0 and standard deviation 1.

[17] This argument has been made with respect to German inflation series. As robustness checks, we also examine the implications of alternative U.S. price news, and of a range of measures of European price news. Our results are robust to other measures, but these other measures sometimes have small and volatile effects on bond yields or the slope of the yield curve.
[18] Recent studies include Goldberg and Leonard (2003),and Ehrmann and Fratscher (2004).
[19] See Andersen *et al.* (2003), Goldberg and Leonard (2003), Faust, Rogers, and Wright (2004), Chinn and Frankel (2004) and Ehrmann and Fratscher (2004)
[20] Taking the monetary reaction function literally, this choice would imply that the ECB reacts to U.S. inflation news. Our specification does not preclude an ECB reaction function to other more local inflation series as well. It does, however, make explicit an assumption that U.S. inflation reports contain information perceived as relevant for Europe and euro-area monetary policy.
[21] For a nice overview of the evidence and related literature, see Clark (2001).

3.2 Econometric Methods

New econometric tests developed by Elliott and Müller (forthcoming) allow one to test for the presence of persistent time variation in one or more regression coefficients over the sample period without specifying the 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.[22] 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 Elliott – Müller (forthcoming) "quasi-Local Level" (*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.

We would like to ensure that the persistent time variation in $(\gamma - \phi\lambda_i)$ for European yield curves, and for the euro/dollar exchange rate, is due to variation over time in λ rather than, say, variation over time in γ, 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 cause of a time variation in $(\gamma - \phi\lambda_i)$ is changes in λ rather than changes in γ or ϕ. This view is bolstered by some other tests. Evidence will be presented that shows no persistent time variation in $(\gamma - \phi\lambda_i)$ for the slope of U.S. yield curves, suggesting no change in the overall environment across countries that has led to a change in γ or ϕ rather than in λ. 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 λ as the source of the persistent time variation in $(\gamma - \phi\lambda_i)$.

Comparing our estimates of the smoothed time path of $(\gamma - \phi\lambda_i)$ (using the method developed by Müller and Petalas (2005)) to actual changes in ECB policy will also, in a less formal way, help isolate λ as the source of the persistent time variation in that we will show

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[22] 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.10) 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. 4)

a decrease in $(\gamma - \phi\lambda_i)$ (reflecting an increase in λ) corresponding to 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 $(\gamma - \phi\lambda_i)$ 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 test concerning the persistent time variation of the coefficient.[23]

3.3 Time Variation in the Effects of News on the Slope of the Yield Curve

In this section we report the results of the Elliott and Müller (forthcoming) *qLL* statistic for the five asset price series discussed above. These statistics are negative, and a smaller (i.e. more negative) value of the statistic allows one to reject the null hypothesis of a lack of persistent time variation in the effect of news on inflation expectations. Thus, suppose there was an evolving view of the policy stance of the ECB over time, but not of the Fed over this same period. We would expect to see a smaller q*LL* statistic than some 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, but a *qLL* statistic larger than this critical value for a regression in which the dependent variable is the change in the term spread on United States government bonds.

Results of this test are presented in Table 1.[24] The first row is a test of the general persistent variation in the slope coefficient only. The second 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.

[23] The sup-Wald tests are discussed in Section 3.5.
[24] As suggested by Elliott and Müller (forthcoming), we allow for the possibility of heteroskedasticity in the variance-covariance matrix of the score series $\{(x_{t+} - E_t x_{t-}) \times \varepsilon_{t^+}\}$ by using the Newey-West (1987) correction. We have written a Stata program for conducting the Elliott – Müller test which is available on request.

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	**-10.95**	**-8.81**	*-7.43*	-5.61	**-8.79**
Slope & Intercept	***-21.52***	-9.08	-7.11	-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 $(\gamma - \phi\lambda_i)$ 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 United States government bonds over this same period of time.

All of these results are consistent with the model presented above in which $(\gamma - \phi\lambda_i)$ varies as λ_i changes 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 second row of *qLL* statistics in Table 1 present results of a test of the joint persistent time variation of the slope and intercept terms of a regression that takes the form of specification (6) (footnote 10 presents a discussion of the possibility of persistent time variation in the intercept as well as the slope). There is even stronger evidence of persistent time variation in this joint test for the change in the term spread for German bonds and the change in the euro / dollar exchange rate, but weaker evidence for persistent time variation when the dependent variable is the change in the term spread of either French or Italian government bonds. Again, the benchmark regression, of the change in the term spread of

United States government bonds, fails to offer evidence of persistent time variation in the coefficients of the regression.

The finding of a significant persistent variation for the slopes of the German, French and Italian yield curves, as well as for the Euro, 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. But one could also argue that these results are consistent with a situation where there was a change in the relationship between U.S. and European inflation rates over this period since the announcements are, in all regressions, of U.S. core CPI. 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 $(\gamma - \phi\lambda_i)$ presented in the next section.

3.4 Estimated Paths of $(\gamma - \phi\lambda_i)$

In this section we present the estimated parameter paths of $(\gamma - \phi\lambda_i)$, using the technique developed by Müller and Petalas (2005). They 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.[25]

Figures 1 presents the estimated parameter paths of $(\gamma - \phi\lambda_i)$ for the regressions using the four bond term spreads. Figure 2 presents the parameter path for the slope coefficient in the regression on the euro/dollar exchange rate and, to provide comparability to the first figure, the slope coefficients for the German and United States term spreads that are presented in Figure 1.

The first thing to note from Figures 1 and 2 is that the estimated value of $(\gamma - \phi\lambda_i)$ for each of the term spreads for the three European government bonds, as well as the euro/dollar exchange rate, is greater than the estimated value of $(\gamma - \phi\lambda_i)$ for the term spread of 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. Another immediately apparent characteristic of the four time paths in Figure 1, and the time path of the euro/dollar exchange rate in Figure 2, is the relative variability of the three European $(\gamma - \phi\lambda_i)$'s as compared to that of the United States.[26]

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[25] 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.

[26] The standard deviations of the estimated $(\gamma - \phi\lambda_i)$'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.

Figure 1 Time Profile of Yield Curve Slope Response to News

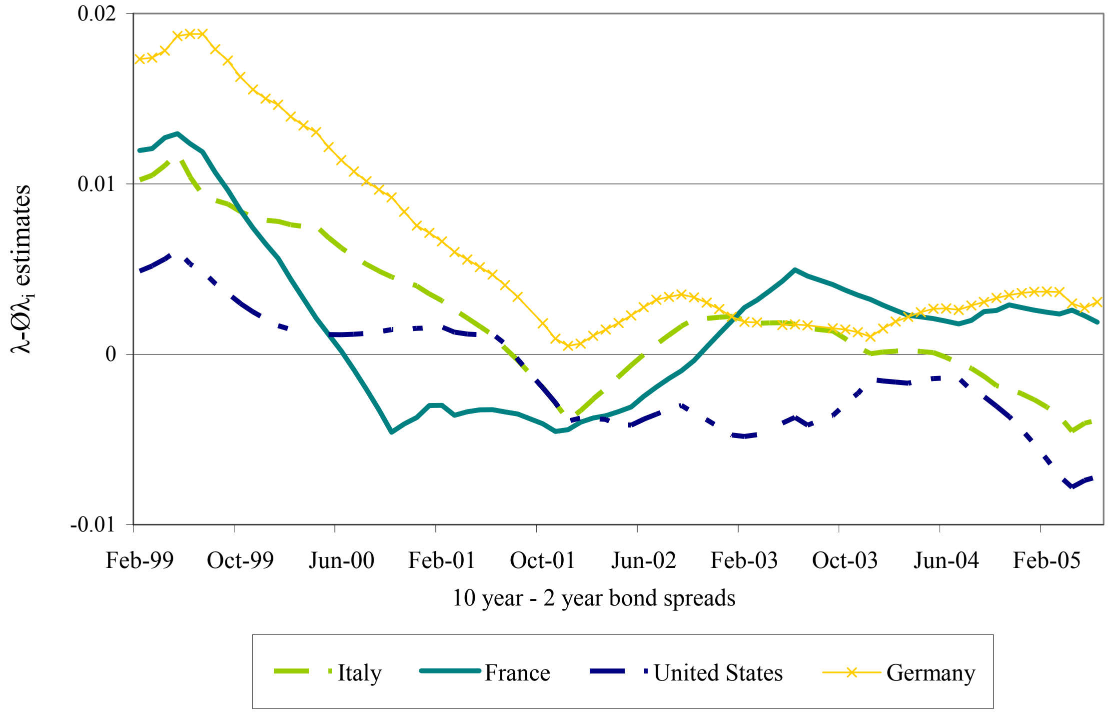
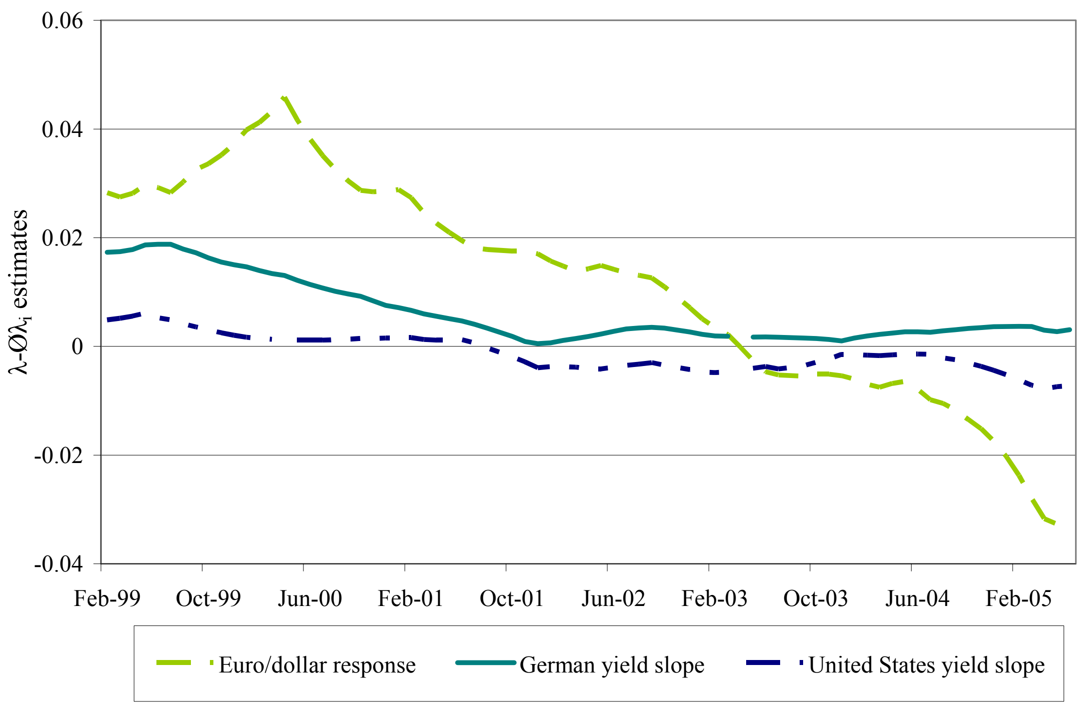


Figure 2 Time Profile of Euro/Dollar Response to News



The time variation of the estimated paths of $(\gamma - \phi\lambda_i)$ in light of the actions undertaken by the European Central Bank bolster our contention that the variation in this parameter is due to changing views of its policy stance (as reflected in λ_i) rather than, say, changing values in γ. The peak values of $(\gamma - \phi\lambda_i)$ occur at the time of the May 1999 core CPI announcement for the French and Italian bond yields, the June 1999 announcement for the German bond yields, and the April 2000 announcement for the euro / dollar rate. The decline in $(\gamma - \phi\lambda_i)$, consistent with markets updating their perceptions toward a more hawkish view of ECB inflation aversion, continued for the regressions using the three European bond yields until the late autumn of 2001, and for the euro/dollar rate for the remainder of the sample period. During much of two-year period beginning in the autumn of 1999, the ECB was tightening monetary policy. November 4, 1999 marked the first time that the ECB raised its key interest rate since it began operations on January 1, 1999.[27] At that time, this interest rate for main refinancing operations was raised from 2.5 percent to 3.0 percent. This was followed by another 25 basis point increase on February 3, 2000, additional 25 basis point increases on March 16 and April 27, and a 50 point basis point increase to 4.25 percent on June 9, 2000. Over this whole period, the regression coefficient on the slope of the U.S. term spread was stable.

The smoothed estimate of $(\gamma - \phi\lambda_i)$ for the three European term spreads began to rise again towards the end of 2001, up until September 2002 (for the German term spread), January 2003 (for the Italian term spread) and June 2003 (for the French term spread). In our model, this could occur if the market partially corrected their views of the degree of inflation aversion characterizing the ECB reaction function. This occurred over a period of time when the actions of the European Central Bank tilted towards a more accommodative monetary stance. On May 11, 2001 the ECB lowered the minimum bid rate for the main refinancing operations by 25 basis points, to 4.50 percent.[28] Four additional interest rate cuts by the ECB

[27] The key interest rate on fixed rate tenders was at 3.00 percent from January 1, 1999 through April 9, when it dropped by 50 basis points to 2.50 percent. On November 4, 2000 a period of monetary tightening started. For main refinancing operations, changes in the rate are effective from the first operation following the date when changes were indicated.

[28] 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

occurred on August 30, 2001 (a cut by 25 basis points), on September 17, 2001 (a cut by 50 basis points), on November 8, 2001 (a cut by 50 basis points), and on December 5, 2002 (a 50 basis point cut). Early in 2003, there were a series of additional rate cuts [March 6, 2003, 25 basis points and June 6, 2003, 50 basis points] and the ECB refined its "two pillar" approach, with the official importance of M3 apparently reduced and a more official role for an inflation goal at or slightly below two percent. The main refinancing interest rate remained at 2.00 percent from June 2003 until the end of the sample period in June 2005.

In contrast to the increase, beginning in the late autumn of 2001, in the smoothed estimated path of $(\gamma - \phi\lambda_i)$ for the three European government term spreads , the estimated path of $(\gamma - \phi\lambda_i)$ for the euro / dollar exchange rate continued to decrease through this period, and, indeed, through the rest of the sample period ending in June 2005. Of course, the behavior of the euro / dollar exchange rate depends upon the actions of the Fed as well as that of the ECB and the estimated value of $(\gamma - \phi\lambda_i)$ for the regression using the United States term spread increased along with the coefficients for the European term spreads throughout 2003; so, for this period, at least, the smoothed estimated values of $(\gamma - \phi\lambda_i)$'s from the bond regressions is consistent with a view of somewhat parallel evolution of perceptions for the Fed and the ECB.

3.5 Sup-Wald Statistics

This section presents sup-Wald tests for a discrete change in $(\gamma - \phi\lambda_i)$, based on Andrews (1993, 2003) to gauge the robustness of both the Elliott – Müller *qLL* tests and of the smoothed paths of the $(\gamma - \phi\lambda_i)$ 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 3.4.

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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.

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

$$q_{t^+} - q_{t^-} = \alpha + \beta\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \beta_I D_I\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \varepsilon_{t^+} \quad (7)$$

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 has 15 percent trimming, 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 3 presents the sup-Wald statistics based on sets of five different regressions that take the form of (7), 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. 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 3: 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. The dates presented in Table 2 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 $(\gamma - \phi\lambda_i)$ in the period between mid-1999 and late-2001, the time when these coefficients had their largest average value. There is also a consistency between the two estimated break points for the Euro / dollar regression and the Müller – Petalas estimated time paths since the first estimated break, February 21, 2001, comes at the time just before the smoothed parameter path descends from a high average value and the second estimated break, April 16, 2003, occurs immediately prior to a large decrease in the value of the estimated smoothed parameter path. Thus, there is an overall consistency between the sup-Wald results and the Müller – Petalas estimated smoothed time path, suggesting 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.

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 rich vein to mine because of the controversy surrounding the conduct of monetary policy in Europe as the ECB began its operations.

In this paper, we have proposed and executed 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 provides 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.

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