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Establishing Credibility: Evolving Perceptions of the European Central Bank

Linda S. Goldberg
Federal Reserve Bank of New York
NBER

Michael W. Klein
Fletcher School
Tufts University
N.B.E.R.

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Abstract: The credibility of a central bank's anti-inflation stance, a key determinant of its success, may reflect institutional structure or, more dynamically, the history of policy decisions. The first years of the European Central Bank (ECB) provide a natural experiment for considering whether, and how, central bank credibility evolves. In this paper, we present a model demonstrating how the high-frequency response of asset prices to news reflects market perceptions of the anti-inflation stance of a central bank. Empirical tests of this model on high frequency data, regressing both the change in the slope of the German yield curve and the change in the euro/dollar exchange rate on the surprise component of price news, suggest significant instability in the market's perception of the policy stance of the ECB during its first five years of operation. Estimated smoothed paths of the coefficients linking news to asset prices show that these coefficients change with policies undertaken by the ECB. In contrast, there is no evidence of parameter instability for the response of the slope of the United States yield curve to price news during this period, suggesting no comparable evolution in the market perceptions of the commitment to inflation fighting by the Federal Reserve.

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

JEL Classification: F3, E5, E6

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1. Introduction

The credibility of the anti-inflation stance of a central bank plays a key role in determining whether the 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 credibility is considered vitally important and "helps keep inflation low."

This consensus on the importance of credibility 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 attaining credibility. 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 structure of a central bank as compared to its conduct for attaining and maintaining its credibility.

A majority of respondents to Blinder's survey believe that central bank credibility is based more 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, in cross sections of countries, that institutional structure is associated with economic performance, 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 the pursuit of longer-run gain.[2] 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 credibility are especially relevant for a new central bank. An analysis of the experience of the European Central Bank

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

[2] 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).

(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.[3] 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 asset prices contain information about the likely future course of inflation, and brings this insight to high frequency data in order to isolate the change in market perceptions of a central bank policy stance. 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 policy. This test uses high frequency asset price data and the surprise components of economic data releases to estimate whether the market perception of the anti-inflation credibility of a central bank changes over time.[4] 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 having a weaker anti-inflationary policy stance. 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

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

[4] 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).

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 break in the relationship between news and the yield curve as credibility evolves. [5]

In Section III we apply this test 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 data on the term structure of European bonds, we find evidence that the market's perception of the anti-inflation credibility of the ECB has evolved over time and responded to ECB policy actions. As a benchmark for our analysis, and also to identify whether the results we found for the ECB could be attributed to changes in the economic environment rather than in specific views of its credibility, we also test for changes in the market's perception of the anti-inflation policy of the Federal Reserve over the same sample period. In contrast to our results for the ECB, we find no evidence of changing perceptions of the policy of the Federal Reserve, a result that is not surprising 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 central bank credibility affects the relationship between news and asset prices. We then discuss the empirical implementation of this model to the yield curve, so that we can use high frequency data to isolate the effects of news on short and longer term inflation expectations under different market perceptions of central bank policy.

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[5] 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.

2.1 Empirical Specification

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

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

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.[6] 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), as a first step, takes into account market expectations about the policy response to news. As a second step, below, we also broaden the interpretation of q_t so that it measures the slope of the yield curve. First, 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) $$q_{t^+} - q_{t^-} = \alpha + \beta\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \phi\left(E_{t^+}M_{t^+} - E_{t^-}M_{t^-}\right) + \varepsilon_{t^+},$$

captures the possibility that $q_{t^+} - q_{t^-}$ responds to economic news directly and indirectly 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

[6] In the empirical estimation, we allow for the possibility that the error term is heteroskedastic.

$E_{t^+}M_{t^+}$. 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) $$M_t = V_t - \lambda_i(x_t - \bar{x}_i)$$

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 can be inflation data, and V represents an indicator of other economic conditions, for example the output gap or unemployment. A higher value of λ represents a stronger anti-inflationary policy of the central bank.

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

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

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

In the short window surrounding a release of inflation numbers, typically $(x_{t^+} - E_{t^-}x_{t^+})$ is not equal to zero. 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.[7] [8] Within this

[7] 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.

equation, the parameter λ_i 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. Bond returns and inflation expectations are linked through the Fisher relationship. If we consider a short dated bond, for example a 2 year, and a long dated bond such as a 10 year, the Fisher relationship shows that $i_t^{10} = E_t\pi_{10} + r_t^{10}$ and $i_t^2 = E_t\pi_2 + r_t^2$, where the long and short real interest rates are r_t^{10} and r_t^2, respectively. With news affects on equilibrium real interest rates common to returns at all 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. We also are abstracting from the effect of news on term premia or liquidity premia, a very reasonable approach over the short time horizon between times t^- and t^+. The regressand $q_{t^+} - q_{t^-}$ is the expected change in the differential in the long horizon and short horizon inflation rates due to the news announcement since

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

where π_{10} is the average expected inflation rate over 10 years and π_2 is the average expected inflation rate over 2 years.[9]

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

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[8] We would not expect, in a sample with many observations, many instances where $\lambda_i(E_{t^+}\bar{x}_i - E_{t^-}\bar{x}_i)$ 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 $\bar{x}_i$ 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 $\lambda_i(E_{t^+}\bar{x}_i - E_{t^-}\bar{x}_i) = 0$ for the vast majority of cases.

[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. Research by Estrella, Rodriguez and Schich (2002) focuses on the relationship between the slope of the yield curve and subsequent real activity and inflation. Using monthly data from 1967 to December 1998, the German yield curves were found to be most informative at inflation prediction horizons beyond two years, while the US term structure was most informative at somewhat shorter horizons. Such analyses, building on Estrella and Mishkin (1997), also argue that inversions in the slope of the yield curve have been successful as recession predictors. The measured strength of these relationships are not stable over time, and vary with the maturity horizons of the bonds examined. 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. These studies do not examine the effects of news announcements on the slope of the yield curve, or use high frequency data as in our study.

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

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

We can be more precise about the instability of $(\beta - \phi\lambda_i)$ 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), meaning one of accommodating inflation 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 $\lambda_H > \lambda_D$. We would expect to find a larger value for the estimated coefficient in the earlier period as compared to the later period since $(\beta - \phi\lambda_D) > (\beta - \phi\lambda_H)$. In contrast, suppose perceptions of policy were unvarying, perhaps because these perceptions were driven by the initial and unvarying institutional structure. We would not expect to find a significantly different estimated coefficient on the response of the slope of a central bank's yield curve to price news across the sample period studied, reflecting stability over time in the anchoring of inflation expectations.

This example, pointing to a discrete change in perceptions of λ, offers a particularly stark view of expectations formation since one episode alters market expectations. An evolving view of central bank policy, one reflecting a gradual learning process, may be more consistent with reality. 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 main econometric technique we use for identifying the changes over time in the slope of the yield curve does not require us to specify this learning process nor the associated evolution of $(\beta - \phi\lambda_i)$. As we develop in section 3.2, our application of new econometric techniques by Elliott and Müller (2005) allows for tests for general parameter instability, and estimation of the smoothed parameter path under very general assumptions.

[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 $q_{t^+} - q_{t^-} = (\alpha + \phi\lambda_i(E_{t^+}\bar{x}_i - E_{t^-}\bar{x}_i)) + (\beta - \phi\lambda_i)(x_{t^+} - E_{t^-}x_{t^+}) + \varepsilon_{t^+}$.

3. Evolving Perceptions of European Central Bank Policy

In this section we present the data, methodology, and results of our tests for changes in the market's perception of the anti-inflation stance of the European Central Bank during its first six-and-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 these 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 or exchange rate series may not, in fact, reflect an evolving perception of ECB credibility but, rather, some structural change common to financial markets across all four of these industrial countries. However, as 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 using the Elliott – Müller test. In Section 3.4 we present estimates of the time path of $(\beta - \phi\lambda_i)$ using the Müller – Petalas procedure, and link these time paths to actual policy changes undertaken by the ECB. Finally, in Section 3.5, 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 inflation announcements and related market expectations, and various asset prices, where the assets are government bonds and foreign exchange. We begin this section with a description of our construction of inflation surprises. We then discuss the five different asset prices used as dependent variables in our estimation.

Inflation Announcements and Expectations: The news variable we use 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,[11] 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. [12]

Although our primary emphasis is on evolving credibility of the ECB, and we will introduce European bonds for constructing yield curves, the analysis relies on U.S. inflation data, not European inflation series. This is a deliberate choice, motivated by extensive evidence on the news content of United States price announcements as contrasted with comparable announcements of prices for Germany, Italy, France, or Europe as a whole. Recent studies, for example by Andersen et al (2003), Goldberg and Leonard (2003), Faust, Rogers, and Wright (2004), Chinn and Frankel (2004) and Ehrmann and Fratscher (2004), support United States macroeconomic news affecting both United States and European asset prices, with little or no significant effect of European macroeconomic news on asset prices in either Europe or America.[13] Market participants argue that the true news content of some German and other European price announcements are questionable because of issues of data quality and episodes of data leaks prior to official announcement times.[14]

The *news* or surprise component of an economic data release represents the difference between the actual release and the markets' 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] 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

[11] See Anderson et al (2003), Goldberg and Leonard (2003), Faust, Rogers and Wright (2004) and Ehrmann and Fratzscher (2004).
[12] For a nice overview of the evidence and related literature, see Clark (2001).
[13] Gadzinski and Orlandi (2004) argue that inflation persistence is comparable across the United States and the euro area.
[14] 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.
[15] Money Market Services were the source of these data through December 2003. Haver Analytics provided continuous expectations and announcement data through 2004 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.

the inflation reports. In creating 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.

Asset Price Data: Five different dependent variables will be used in the regressions. In each case, the dependent variable, $q_{t^+} - q_{t^-}$, represents the change in q between 30 minutes before and 30 minutes after the monthly inflation announcement. Four of the dependent variables are 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. By the Fisher relationship and equation (5) the regressand $q_{t^+} - q_{t^-}$ is the expected change in the differential in the 10-year and 2-year inflation rates due to the inflation news.[16] Thus, when using these four bond series, $(\beta - \phi\lambda_i)$ can be interpreted as a function of the direct effect (represented by β) and the indirect effect (via a policy response, as represented by $\phi\lambda_i$) of current inflation news on the expected long-run relative to short-run inflation rates.

There is a similar interpretation of $(\beta - \phi\lambda_i)$ in the regressions that use the fifth dependent variable, where $q_{t^+} - q_{t^-}$ represents the change in the logarithm of the euro per U.S. dollar exchange rate, thirty minutes before and thirty minutes after the news announcement. With this regressand, and applying purchasing power parity concepts, an

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[16] The Fisher relationship shows that $r_t^{10} = E_t \pi_{10} + i_t^{10}$ and $r_t^2 = E_t \pi_2 + i_t^2$, where the ten-year and two-year real interest rates are i_t^{10} and i_t^2, respectively. News effects on equilibrium real interest rates are common to returns at all horizons along the yield curve. As a consequence, when we difference across the returns of long and short-dated bonds we abstract from the effect of news on equilibrium real returns. We also are abstracting from the effect of news on term premia or liquidity premia, which seems reasonable given the short time horizon between times t^- and t^+. In a similar vein, Fleming and Remolona (1999) argue that the effects of news on asset prices of different maturities reveals information about market participant beliefs about central bank reaction functions. Research by Estrella, Rodriguez and Schich (2002) focuses on the relationship between the slope of the yield curve and subsequent real activity and inflation. The German application was over monthly data from 1967 to December 1998. In inflation prediction, the German yield curves were found to be most informative at prediction horizons beyond two years, while the US term structure was most informative at somewhat shorter horizons. Such analyses, building on Estrella and Mishkin (1997), also argue that inversions in the slope of the yield curve have been successful as recession predictors. The measured strength of these relationships are not stable over time, and vary with the maturity horizons of the bonds examined. 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. These studies do not examine the effects of news announcements on the slope of the yield curve, or use high frequency data as in our study.

increase in $q_{t^+} - q_{t^-}$ due to United States core CPI news indicates an increase in the expected inflation rate in Europe relative to the expected inflation rate in the United States (Klein, Mizrach and Murphy 1991). In this case, $(\beta - \phi\lambda_i)$ decreases across time if, during that period, market participants view 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 such that there is the anti-inflation stance of the ECB becomes more hawkish, relative to its prior history and relative to the policy of the U.S. Federal Reserve.

3.2 Econometric Methods

New tests developed by Elliott and Müller (2005) 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.[17] This feature of their test makes it well suited for our purposes since we do not need to test for a particular type of learning by market participants about central bank policy. The Elliott and Müller (2005) "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 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.

While an unspecified evolution of $(\beta - \phi\lambda_i)$ may be more consistent with our notions of learning about the policy stance of a new central bank, the identification of a break date would be useful for our purposes (if, in fact, this is how perceptions changed) because our tests do not allow us to distinguish between time variation in $(\beta - \phi\lambda)$ due to variation over time in λ rather than, say, variation over time in β, 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 creation of the ECB in January 1999, it is reasonable to expect that

[17] Elliott and Müller write that, for their tests, "…the precise form of the breaking process [of the coefficients] is irrelevant for the asymptotic power of the tests." (p.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)

the most likely cause of a time variation in $(\beta - \phi\lambda)$ is changes in λ rather than changes in β or ϕ. One way we can attempt to identify the source of changes in $(\beta - \phi\lambda)$ is to estimate its path over time and consider how its movements correlate to actual changes in ECB policy. We do this using the method developed by Müller and Petalas (2005) to estimate the smoothed time path of a time-varying parameter.

We also test for the robustness of both the finding of parameter instability and the timing of changes in $(\beta - \phi\lambda_i)$ by presenting sup-Wald tests for parameter stability (see Andrews 1993 and Bai 1997). These tests are predicated on the existence of one or more discrete break points. They are conducted by running $0.7\times T$ regressions, where T is the number of observations in the data set, that take the form

$$(6) \qquad q_{t^+} - q_{t^-} = \alpha + \beta\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \beta_I D_I\left(x_{t^+} - E_{t^-}x_{t^+}\right) + \varepsilon_{t^+}$$

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.[18] This 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 (1993) and, if this sup-Wald statistic exceeds the critical value, the date associated with that β_I is the statistically significant estimated break date. Bai (1997) shows how to test for multiple break dates using a similar methodology.

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 (2005) *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

[18] The sup-Wald tests require trimming, that is, fewer than T regressions are run for a sample with T observations. Andrews (1993) suggests the use of symmetric 15% trimming and therefore, in this case, the test would involve $0.7\times T$ regressions. There is no comparable trimming in the Elliott and Müller test.

bonds before and after the inflation announcements, as well as for the change in the euro / dollar exchange rate before and after the inflation announcements, 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 before and after the news announcements.

Results of this test are presented in Table 1.[19] 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.

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*	-5.40	**-8.79**
Slope & Intercept	***-21.52***	-9.08	-7.11	-8.93	**-16.69**
No. of obs.	74	74	72	67	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 show that there is evidence of persistent time variation in $(\beta - \phi\lambda)$ 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.

[19] As suggested by Elliott and Müller (2005), we allow for the possibility of heteroskedasticity in the variance-covariance matrix of the score series $\{news_t \times \varepsilon_{t^+}\}$ by using the Newey-West (1987) correction.

All of these results are consistent with the model presented above in which $(\beta - \phi\lambda_i)$ varies as λ_i changes with an evolving view of the anti-inflation credibility 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 11 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. But 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 evidence in this table 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 $(\beta - \phi\lambda_i)$ presented in the next section.

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

In this section we present the estimated parameter paths of $(\beta - \phi\lambda_i)$, 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.[20]

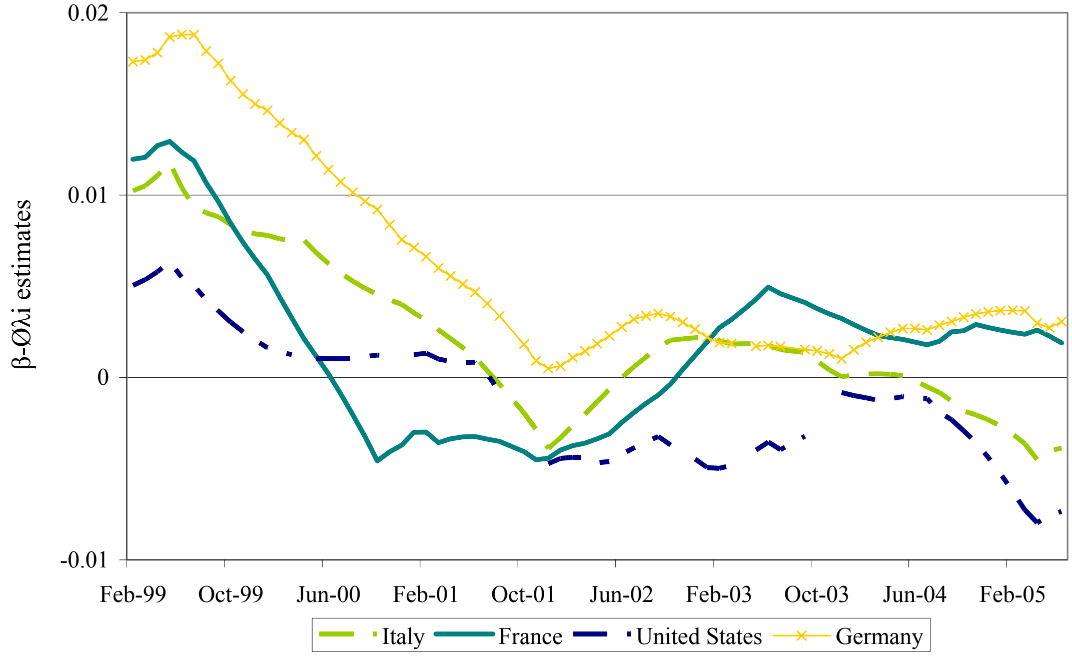
[20] Müller and Petalas (2005) describe their procedure as an extension of the Kalman smoothing formulae with the optimal smoother for the true path of the time varying coefficient a function of the score sequence $\{news_t \times \varepsilon_{t^+}\}$. See their paper for details, and for an outline of how to implement their procedure.

Figures 1 presents the estimated parameter paths of $(\beta - \phi\lambda_i)$ 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 $(\beta - \phi\lambda_i)$ 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 $(\beta - \phi\lambda_i)$ 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 it as having a weaker anti-inflation stance 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 $(\beta - \phi\lambda_i)$'s as compared to that of the United States. The standard deviations of the estimated $(\beta - \phi\lambda_i)$'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.

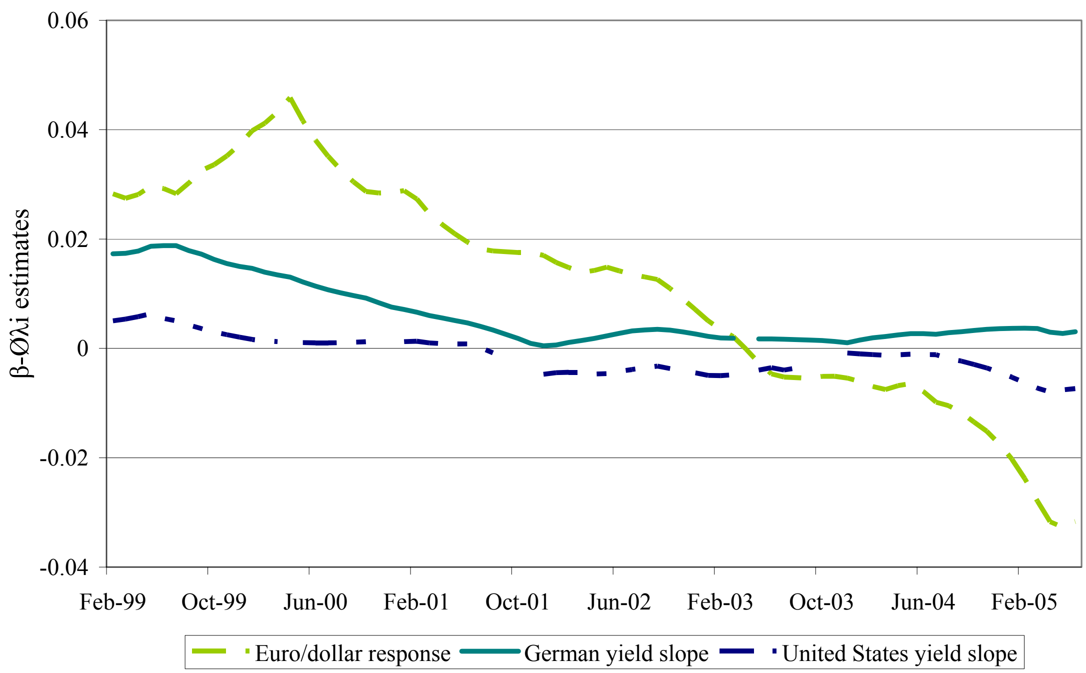
The time variation of the estimated paths of $(\beta - \phi\lambda_i)$ in light of the actions undertaken by the European Central Bank bolster our contention that the variation in this parameter is due to changing views of its policy stance (as reflected in λ_i) rather than, say, changing values in β.

Figure 1 Time Profile of Yield Curve Slope Response to News



10 year - 2 year bond spreads

Figure 2 Time Profile of Euro/Dollar Response to News



The peak values of $(\beta - \phi\lambda_i)$ occur at the time of the May 1999 core CPI announcement for the French and Italian bond yields, at the time of the June 1999 announcement for the German bond yields, and at the time of the April 2000 announcement for the euro / dollar rate. The decline in $(\beta - \phi\lambda_i)$ 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.[21] At that time, this interest rate, the 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.

The smoothed estimate of $(\beta - \phi\lambda_i)$ 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). This is consistent with a change in the actions of the European Central Bank towards a more accommodative monetary policy. This policy move occurred on May 11, 2001 when the ECB lowered the minimum bid rate for the main refinancing operations by 25 basis points, to 4.50 percent.[22] Three additional interest rate cuts by the ECB occurred on August 30, 2001 (when the interest rate was cut by 25 basis points), on September 17, 2001 (when the interest rate was cut by 50 basis points), and on November 8, 2001 (when the interest rate was cut by 50 basis points).[23] This policy stance continued with three more rate cuts, on December 5, 2002 (a 50 basis point cut), March 6, 2003 (a 25 basis point cut) and June 6, 2003 (a 50 basis point cut). Also, early in 2003 the ECB refined their "two pillar" approach, with the

[21] 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.

[22] 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.

[23] One may be concerned that, because this period includes September 11, 2001, the effect of the terrorist attacks on that day may be responsible for the shift. However, in this figure we do not observe comparable changes in the value of $(\beta - \phi\lambda_i)$ for regressions using the U.S. term spread.

importance of M3 apparently reduced and the target for inflation 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 in the smoothed estimated path of $(\beta-\phi\lambda_i)$ for the three European government term spreads beginning in the late autumn of 2001, the estimated path of $(\beta-\phi\lambda_i)$ 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 $(\beta-\phi\lambda_i)$ 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 $(\beta-\phi\lambda_i)$'s from the bond regressions reflect a somewhat parallel evolution of perceptions for the Fed and the ECB.

3.5 Sup-Wald Statistics

Finally, to gauge the robustness of the Elliott – Müller *qLL* tests, and of the smoothed paths of the $(\beta-\phi\lambda_i)$ coefficient obtained through the Müller-Petalas method, this section presents sup-Wald tests for a discrete change in $(\beta-\phi\lambda_i)$, based on Andrews (1993) and Bai (1997). As discussed above, these sup-Wald tests are based on a more restricted assumption concerning the break point than the *qLL* test but, since a break point is estimated rather than the overall stability of the parameter, 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.

Table 2 presents the sup-Wald statistics based on sets of five different regressions that take the form of (6), four regressions in which the dependent variable is the change in one of the term spreads, and one regression in which the dependent variable is 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 95 percent level of confidence, for the regressions using the change in the term spread for German and Italian government bonds, and for the Euro / dollar exchange rate. 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. In addition, using the technique presented by Bai (1997) to search for multiple break points, there is evidence of a second break point for the change in the Euro / dollar exchange rate.

Table 2: sup-Wald Test for Discrete Break Point

Break Point in	**Change in Term Spread of Government Bonds of**				**Change in Euro/$**
	Germany	**France**	**Italy**	**United States**	
Sup-Wald statistic	***20.31***	1.78	**11.91**	3.92	**12.25**
Estimated Break Date	November 1 2000		June 15, 2001		February 21, 2001
No. of obs.	74	74	72	67	75

Critical Values (from Andrews 1993) 1% ***12.35***; 5% **8.85**; 10% *7.17*
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 and 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 $(\beta - \phi\lambda_i)$ 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 these results.

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 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 policy stance 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 policy stance 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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Institute for International Integration Studies
The Sutherland Centre, Trinity College Dublin, Dublin 2, Ireland