ECONSTOR
Make Your Publications Visible.

A Service of
ZBW
Leibniz-Informationszentrum Wirtschaft
Leibniz Information Centre for Economics

Goldberg, Linda S.; Klein, Michael W.

Working Paper

Establishing credibility: Evolving perceptions of the European Central Bank

Staff Report, No. 231

Provided in Cooperation with:
Federal Reserve Bank of New York

Suggested Citation: Goldberg, Linda S.; Klein, Michael W. (2005) : Establishing credibility: Evolving perceptions of the European Central Bank, Staff Report, No. 231, Federal Reserve Bank of New York, New York, NY

This Version is available at:
https://hdl.handle.net/10419/60665

Standard-Nutzungsbedingungen:

Die Dokumente auf EconStor dürfen zu eigenen wissenschaftlichen Zwecken und zum Privatgebrauch gespeichert und kopiert werden.

Sie dürfen die Dokumente nicht für öffentliche oder kommerzielle Zwecke vervielfältigen, öffentlich ausstellen, öffentlich zugänglich machen, vertreiben oder anderweitig nutzen.

Sofern die Verfasser die Dokumente unter Open-Content-Lizenzen (insbesondere CC-Lizenzen) zur Verfügung gestellt haben sollten, gelten abweichend von diesen Nutzungsbedingungen die in der dort genannten Lizenz gewährten Nutzungsrechte.

Terms of use:

Documents in EconStor may be saved and copied for your personal and scholarly purposes.

You are not to copy documents for public or commercial purposes, to exhibit the documents publicly, to make them publicly available on the internet, or to distribute or otherwise use the documents in public.

If the documents have been made available under an Open Content Licence (especially Creative Commons Licences), you may exercise further usage rights as specified in the indicated licence.

WWW.ECONSTOR.EU

Mitglied der Leibniz-Gemeinschaft

Federal Reserve Bank of New York
Staff Reports

Establishing Credibility: Evolving Perceptions of the European Central Bank

Linda S. Goldberg
Michael W. Klein

Staff Report no. 231
November 2005

This paper presents preliminary findings and is being distributed to economists and other interested readers solely to stimulate discussion and elicit comments. The views expressed in the paper are those of the authors and are not necessarily reflective of views at the Federal Reserve Bank of New York or the Federal Reserve System. Any errors or omissions are the responsibility of the authors.

Establishing Credibility: Evolving Perceptions of the European Central Bank

Linda S. Goldberg and Michael W. Klein

Federal Reserve Bank of New York Staff Reports, no. 231
November 2005
JEL classification: F3, E5, E6

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 evolved significantly during the first six years of the Bank's operation, 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 U.S. yield curve to news offer no comparable evidence of any change in market perceptions of the inflation aversion of the Federal Reserve.

Key words: central banking, European Central Bank, Federal Reserve, inflation, exchange rate, monetary policy, credibility, yield curve

---

Goldberg: Federal Reserve Bank of New York and NBER (e-mail: linda.goldberg@ny.frb.org). Klein: Fletcher School, Tufts University and NBER (michael.klein@tufts.edu). The authors have benefited from comments from Donald Andrews, Roc Armenter, Menzie Chinn, Sylvester Eijffinger, Graham Elliott, Arturo Estrella, Francesco Giavazzi, Ulrich Müller, Zhongjun Qu, Richard Portes, John Rogers, Jim Stock, Cédric Tille, and participants at the January 2005 Allied Social Science Associations meeting, the March 2005 Trans-Atlantic Finance Institute-*Journal of International Money and Finance* conference, the Spring 2005 NBER International Finance and Macroeconomics Program meeting, and the Federal Reserve System Committee on International Economic Analysis. The views expressed in this paper are those of the authors and do not necessarily reflect the position of the Federal Reserve Bank of New York or the Federal Reserve System

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."[2]

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.

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, 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-

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

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 test 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 market perception of the anti-

[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, 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.

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 break in the relationship between news and the yield curve as credibility evolves. [6]

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 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 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 inflation aversion, we also test for changes in the market's perception of the monetary reaction function 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 reaction function 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 Volcker.

[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.

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, Andersen, 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

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

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. As emphasized in Andersen 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) $$q_{t^+} - q_{t^-} = \alpha + \gamma\left(x_{t^+} - E_{t^-} x_{t^+}\right) + \phi\left(E_{t^+} M_{t^+} - E_{t^-} M_{t^-}\right) + \varepsilon_{t^+},$$

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

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

[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.

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.

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 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. 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

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

[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 $\lambda_i(E_{t^+}\bar{x}_i - E_{t^-}\bar{x}_i)$ 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.

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), 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 $\lambda_H > \lambda_D$. Evolving perceptions of central bank inflation aversion would be reflected in the coefficient on inflation news. For a central bank that 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.

---

[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.

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

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.[11] 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. However, 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 $(\gamma - \phi\lambda_i)$. Instead, this econometric technique, developed by Elliott and Müller (2005) tests 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 associate changes in the estimated values of $(\gamma - \phi\lambda_i)$ with policy changes of the European Central Bank.

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 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 shown in Section 3.3, we find

[11] For example, see Backus and Driffill (1985) or, more recently, Athey, Atkeson, and Kehoe (2005).

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 $(\gamma - \phi\lambda_i)$ using the Müller – Petalas procedure, and link these time paths to actual policy changes undertaken by the ECB. Finally, in Section 3.5, to demonstrate the robustness of our results, 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 will be 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 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. The regressand $q_{t^+} - q_{t^-}$ is the change in the term premia, which under the Fisher relationship would reflect the change in the differential between the expected ten-year and two-year inflation rates due to the inflation news. Thus, when using these four bond 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 the expected long-run relative to short-run inflation rates.

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 euro 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.[12] In this case, evidence that $(\gamma - \phi\lambda_t)$ decreases over the sample period reflects a view among market participants that there is relatively strengthening of the anti-inflation stance of the ECB as compared to the U.S. Federal Reserve, due to 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. Consideration of tests on bond yields of the euro-area countries and the United States help isolate the source of the change of the responsiveness of the euro/dollar exchange rate to news.[13]

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, 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. [14]

Although our primary emphasis is on evolving credibility of the ECB, and we introduce European bonds for constructing yield curves, our choice of inflation report is the U.S. core CPI release, not European inflation series. 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. We choose U.S. inflation data because there is extensive evidence showing the rich news content of this data for euro area asset prices. This evidence stands in stark contrast with weak effects of inflation reports from Germany, Italy, France, or Europe as a whole.[15] United States macroeconomic news affects both United States and

[12] 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.
[13] 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.
[14] For a nice overview of the evidence and related literature, see Clark (2001).
[15] Recent relevant studies include Andersen et al (2003), Goldberg and Leonard (2003), Faust, Rogers, and Wright (2003), Chinn and Frankel (2004) and Ehrmann and Fratscher (2004).

European asset prices, with little or no significant effect of European macroeconomic news on asset prices in either Europe or America. Market participants argue that the news content of some German and other European price announcements has at times been questionable because of issues of data quality and episodes of data leaks prior to official announcement times.[16]

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.[17] 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.

3.2 Econometric Methods

New econometric 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.[18] 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 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

[16] 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.

[17] 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.

[18] 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)

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 $(\gamma - \phi\lambda_i)$ may be consistent with notions of markets updating perceptions of the inflation aversion 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 $(\gamma - \phi\lambda_i)$ 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. A change in γ would imply a change in the direct effect of inflation on asset prices over the sample period. While theoretically possible, we will present two types of evidence more supportive of evolving market perceptions of the ECB reaction function.

First, we compare the test results across euro area countries, the euro-dollar exchange rate, and the benchmark case of the yield curve for U.S. assets. Given the integration of European and U.S. financial markets, as recently documented by Bartolini, Hilton, and Prati (2005), it would be unusual for a γ change to occur across euro area bond markets but not for U.S. bond markets.

The second way we attempt to identify the source of changes in $(\gamma - \phi\lambda_i)$ is by estimating its path over time and considering whether its movements correlate to actual changes in ECB policy. We do this estimation using the method developed by Müller and Petalas (2005) to estimate the smoothed time path of a time-varying parameter.

Finally, we test for the robustness of both the finding of parameter instability and the timing of changes in $(\gamma - \phi\lambda_i)$ by presenting sup-Wald tests for parameter stability (see Andrews 1993, 2003). 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_1 is a dummy variable that equals 0 for the first *n* observations of the sample and equals 1 for the remaining $T - n$ observations.[19] This generates a set of $0.7 \times T$ β_1's and $0.7 \times T$ associated test statistics. The sup-Wald test compares the largest F-value for all of the β_1's with critical values presented in Andrews (2003) and, if this sup-Wald statistic exceeds the critical value, the date associated with that β_1 is the statistically significant estimated break date.[20]

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 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.[21] 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.

[19] 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.

[20] Estrella and Rodriguez (2005) propose an alternative technique, where the hypothesis tested is a one-sided directional change, increase or decrease, in the estimated parameters.

[21] As suggested by Elliott and Müller (2005), 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.

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 and intercept	***-21.52***	-9.08	-7.11	-8.75	**-16.69**
number of observations	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 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. 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 $(\gamma - \phi\lambda_i)$ 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.[22]

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.[23]

---

[22] 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.

[23] 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

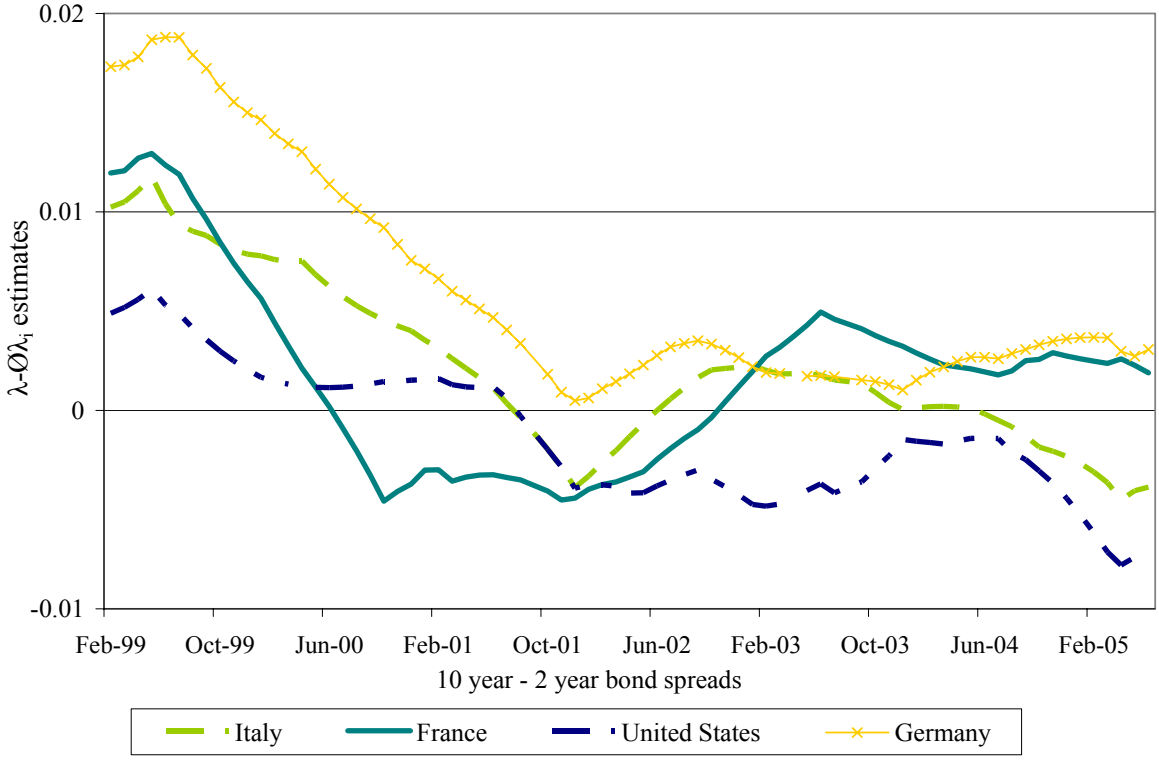
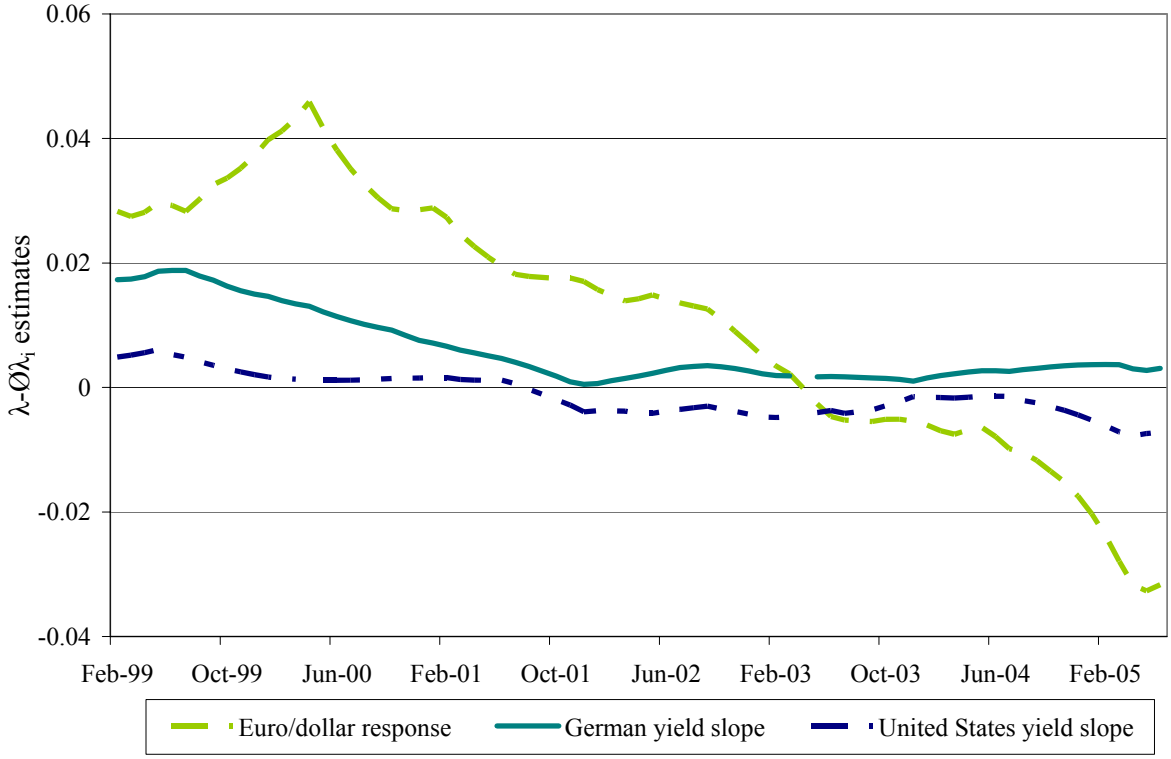


Figure 2 Time Profile of Euro/Dollar Response to News



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.[24] 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.[25] Four additional interest rate cuts by the ECB occurred on August 30, 2001 (a cut by 25 basis points), on September 17, 2001 (a cut by 50 basis points), on November

[24] 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.

[25] 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.

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

Finally, to gauge the robustness of the Elliott – Müller *qLL* tests, and of the smoothed paths of the $(\gamma - \phi\lambda_i)$ coefficients obtained through the Müller-Petalas method, this section presents sup-Wald tests for a discrete change in $(\gamma - \phi\lambda_i)$, based on Andrews (1993, 2003). 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.

Table 2 presents the sup-Wald statistics based on sets of five different regressions that take the form of (6), 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 2: 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
number of observations	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 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 $(\gamma - \phi\lambda_i)$ 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 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 in its inflation aversion 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 euro-zone. 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.

References

Alesina, Alberto and Lawrence H. Summers, 1993. "Central Bank Independence and Macroeconomic Performance: Some Comparative Evidence," *Journal of Money, Credit and Banking*, vol. 25, no. 2 (May), pp. 151 – 162.

Andersen, Torben G., Tim Bollerslev, Francis X. Diebold and Clara Vega, 2003. "Micro Effects of Macro Announcements: Real-Time Discovery in Foreign Exchange," *American Economic Review*, vol. 93, no. 1 (March), pp. 38 – 62.

Andrews, Donald, 1993. "Tests for Parameter Instability and Structural Change with Unknown Change Point," *Econometrica*, vol 61, no. 4 (July), pp. 821 – 856.

Andrews, Donald, 2003. "Tests for Parameter Instability and Structural Change with Unknown Change Point: A Corrigendum," *Econometrica*, vol. 71, no. 1 (January), pp. 395 – 397.

Athey, Susan, Andrew Atkeson, and Patrick Kehoe, 2005. "The Optimal Degree of Monetary Policy Discretion," *Econometrica*, vol. 73, no. 5 (September), pp. 1431-1475.

Backus, David and John Driffill, 1985. "Rational Expectations and Policy Credibility Following a Change in Regime," *Review of Economic Studies*, vol. 52, no. 2 (April), pp. 211 – 221.

Barro, Robert J. and David B. Gordon, 1983. "A Positive Theory of Monetary Policy in a Natural Rate Model," *The Journal of Political Economy*, vol. 91, no. 4 (August), pp. 589-610.

Bartolini, Leonardo, Spence Hilton, and Alessandro Prati, 2005. "Money Market Integration", Federal Reserve Bank of New York.

Bernanke, Ben, 2004. "What Policymakers Can Learn from Asset Prices," Remarks before The Investment Analysts Society of Chicago (April 15).

Blinder, Alan, 2000. "Central Bank Credibility: Why Do We Care? How Do We Build It?," *American Economic Review*, vol. 90, no. 5 (December), pp. 1421 – 1431.

Calvo, Guillermo, 1978. "On the Time Consistency of Optimal Policy in a Monetary Economy," *Econometrica*, vol. 46, no. 6 (November), pp. 1411-1428.

Chinn, Menzie and Jeffrey Frankel, 2003. "The Euro Area and World Interest Rates," Santa Cruz Center for International Economics Paper 03-17 (November).

Clark, Todd, 2001. "Comparing Measures of Core Inflation", *Economic Review*, Federal Reserve Bank of Kansas City, Quarter 2 pp.5-31.

Cukierman, Alex, 1992. *Central Bank Strategy, Credibility and Independence: Theory and Evidence,* The MIT Press, Cambridge, Massachusetts.

Elliott, Graham and Ulrich Müller, 2005. "Optimally Testing General Breaking Processes in Linear Time Series Models," Manuscript, Princeton University.

Elliott, Graham and Ulrich Müller, forthcoming. "Efficient Tests for General Persistent Time Variation in Regression Coefficients" *Review of Economic Studies*.

Engel, Charles and Jeffrey Frankel, 1984. "Why Interest Rates React to Money Announcements: An Explanation from the Foreign Exchange Market," *Journal of Monetary Economics*, vol 13, pp. 31 – 39.

Ehrmann, Michael and Marcel Fratzscher, 2004. "Equal Size, Equal Role: Interest Rate Interdependence between the Euro Area and the United States," European Central Bank working paper series no. 342 (April).

Estrella, Arturo and Frederic Mishkin, 1997. "The Predictive Power of the Term Structure of Interest Rates in Europe and the United States: Implications for the European Central Bank," *European Economic Review,* 41, pp. 1375-401.

Estrella, Arturo and Anthony Rodriguez, 2005. "One-sided test for an unknown breakpoint: Theory, computation, and application to monetary policy," Federal Reserve Bank of New York.

Estrella, Arturo, Anthony Rodriguez and Sebastian Schich, 2003. "How Stable is the Predictive Power of the Yield Curve? Evidence from Germany and the United States" *Review of Economics and Statistics*, vol. 85, no. 3 (August), pp 629-644.

Faust, Jon, John Rogers, and Jonathan Wright, 2003. "The High Frequency Response of Exchange Rates and Interest Rates to Macroeconomic Announcements", International Finance Discussion Papers 784, Board of Governors of the Federal Reserve System.

Fleming, Michael and Eli Remolona, 1999. "Price Formation and Liquidity in the U.S. Treasury Market: The Response to Public Information" *Journal of Finance* vol. LIV, no. 5 (October).

Forder, James, 1999. "Central Bank Independence and Credibility: Is there a shred of evidence?: Review," *International Finance*, vol. 3, no. 1 (April), pp. 167 – 185.

Goldberg, Linda and Deborah Leonard, 2003. "What Moves Sovereign Bond Markets? The Effects of Economic News on U.S. and German Yields," *Federal Reserve Bank of New York, Current Issues in Economics and Finance* volume 9, no. 9 (September).

Gurkaynek, Refet and Justin Wolfers, 2005. "An Initial Analysis of Market-Based Macro Forecasts, Uncertainty, and Risk," forthcoming in NBER *International Seminar on Macroeconomics 2005*.

Huizinga, John and Frederic Mishkin, 1986. "Monetary Policy Regime Shifts and the Unusual Behavior of Interest Rates," Carnegie-Rochester Conference Series on Public Policy, vol. 24 (Spring), pp. 231-274.

Klein, Michael W., and Karen K. Lewis, 1993. "Learning About Intervention Target Zones," *Journal of International Economics*, vol. 35, no. 3/4, pp. 275-295.

Klein, Michael W., Bruce Mizrach and Robert G. Murphy, 1991. "Managing the Dollar: Has the Plaza Mattered?" *Journal of Money, Credit and Banking*, vol. 23, no. 4 (November), pp. 742 – 751.

Kydland, Finn and Edward Prescott, 1977. "Rules Rather than Discretion: The Inconsistence of Optimal Plans," *The Journal of Political Economy*, vol. 85, no. 3 (June), pp. 473-492.

Lewis, Karen K., 1995. "Occasional Interventions to Target Rates," *American Economic Review*, vol. 85, no. 4 (September), pp. 691 - 715.

Müller, Ulrich and Philippe-Emmanuel Petalus, 2005. "Efficient Estimation of the Parameter Path in Unstable Time Series Models." unpublished manuscript, Princeton University.

Newey, Whitney and Kenneth West, 1987. "A Simple Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matric," *Econometrica*, vol. 55, pp 703-8.

Rogoff, Kenneth, 1985. "The Optimal Commitment to an Intermediate Monetary Target," *Quarterly Journal of Economics*, vol.100 (November) pp.1169-89.

Rogoff, Kenneth, 1987. "Reputational Constraints on Monetary Policy," *Carnegie Rochester Conference Series on Public Policy*, vol. 26. pp.141-82.

Svensson, Lars E.O., 1991. The Simplest Test of Target Zone Credibility," *IMF Staff Papers*, vol. 38, pp. 655 – 665.

Svensson, Lars E.O., 1993. The Simplest Test of Inflation Target credibility," 1993, NBER Working Paper No. 4604 (December).