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## The Role of Asset Prices in Euro Area Monetary Policy: Specification and Estimation of Policy Rules and Implications for the European Central Bank

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# **The Role of Asset Prices in Euro Area Monetary Policy: Specification and Estimation of Policy Rules and Implications for the European Central Bank\***

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**Abstract:** This paper estimates forward-looking and forecast-based Taylor rules for France, Germany, Italy, as well as the euro area, using both final revised data and real-time data. We are particularly interested in the impact of adding asset prices to the standard Taylor rule specification. Since forward-looking Taylor rules are usually estimated via GMM, we perform extensive tests for over-identifying restrictions and instrument relevance, the latter generally eschewed in previous work. We find that asset prices can be highly relevant as instruments in policy rules. Forecast-based rules perform best using the root mean squared error metric. However, forecast-based rules are best estimated in difference form. Encompassing tests are used to select the “best” policy rule. Finally, we perform a series of counterfactual experiments and ask whether the ECB’s monetary policy resembles that of any particular euro area core country, as well as asking what interest rates would have been like in the core countries had the ECB conducted monetary policy prior to 1999. Finally, while estimation using real-time data results in a deterioration in the forecasting performance of standard policy rules we do not find that they provided seriously misleading advice about the appropriate stance of monetary policy.

**Keywords:** Monetary policy reaction functions, asset prices, instruments, European Central Bank

**JEL:** E52, E58, C52

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## **1. Introduction**

The creation of the European Central Bank (ECB) resulted in a historic transfer of responsibility in monetary policy from many central banks to a single supra-national authority. The creation of a single monetary policy also took place at a time when central banks were perhaps becoming increasingly pre-occupied with the behavior of financial asset prices. By virtue of its structure, the ECB is a creature of its predecessors' experiences, and, hence, it is also worthwhile considering what monetary policy in the euro area looks like relative to what the national central banks implemented prior to 1999.

It is unlikely that all economic shocks had the same impact in each of the prospective euro area member states (e.g., as with German reunification, or the EMS crisis of 1992) and, until the end of 1998, these might have elicited different monetary policy responses. The behavior of short-term interest rates for three core euro area members, namely France, Germany, and Italy, and a synthetic euro area-wide short-term interest rate prior to 1999 together with the actual euro area repo rate since 1999, illustrates that while convergence in interest rates is apparent, it actually becomes visually most noticeable only the year before the ECB took over responsibility for monetary policy. In preceding years, interest rate spreads are sizeable and variable.

It is now standard practice to evaluate monetary policy based on Taylor rules, a device that explains central banks' reactions to inflation and the output gap. Whereas inflation performance used to be interpreted through the behavior of consumer prices alone, more questions are being asked about whether asset prices may have possibly also played an indirect role in the conduct of monetary policy. Policy makers debate whether devoting more explicit attention to asset price movements could improve the quality (or effectiveness) of

monetary policy. Indeed, the transition to European Monetary Union may have prompted some of the national central banks to voice concerns over specific asset prices, depending on their perceived importance in potentially influencing inflation or output growth in their respective national economies. For example, real exchange rate considerations may have mattered relatively more in the conduct of monetary policy in some countries, while housing or equity price developments loomed larger in others.

It has been suggested that central banks pay attention to asset prices (e.g., Bernanke 2004, European Central Bank 2001, 2005). Central bankers are unsure about whether, and how much, information is contained in asset prices beyond that can be derived from the past history of inflation and output.<sup>1</sup> Nevertheless, it is unclear whether the monetary authorities worry over perceived excesses in some asset price relative some equilibrium or fundamental value (e.g., as in the “irrational exuberance” statement made famous by Alan Greenspan), or rather the volatility of asset prices. The best that can be said about the link between asset prices and interest rates is that central bankers are conflicted about their role in influencing policy.<sup>2</sup>

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<sup>1</sup> Bernanke (2005) suggests that “Central bankers naturally play close attention to interest rates and asset prices,...[they] are potentially valuable sources of timely information about economic and financial conditions...[and] should embody a great deal of investors’ collective information and beliefs about the future course of the economy.” He does not actually state that asset prices actually influence the Fed’s monetary policy. Other central bankers, including Alan Greenspan, have also suggested either that asset prices have only an indirect effect on interest rates or were largely ignored in the past (Norris 2005).

<sup>2</sup> “It is far from obvious that bubbles, even if identified early, can be preempted at a lower cost than a substantial economic contraction and possible financial destabilization – the very outcomes we would be seeking to avoid.” (Greenspan 2004).

In any event, whether any of these considerations play any role in interest rate developments in the euro area prior to and since the creation of the ECB are empirical questions. Not to be forgotten is the role of monetary aggregates. The ECB has been criticized about the emphasis placed on its two pillars of monetary policy (e.g., Gerlach and Svensson 2003, but see Scheller 2004 and von Hagen and Hofmann 2005). However, it can be argued that a monetary authority that focuses on the “longer run” objective of inflation control would do well to take “money seriously” (Laidler 1988) whereas evincing a concern for other asset prices might signal a form of “tunnel vision” in the conduct of monetary policy (e.g., European Central Bank 2005).

This paper addresses a series of interrelated questions. First, can we detect empirically whether euro area central banks took account of asset price developments, motivated perhaps by the need to maintain financial stability, or to ensure that monetary conditions would be adequate for the launch of the euro? Second, if there is considerable diversity in the relative importance of certain asset prices in guiding the stance of monetary policy across countries ostensibly committed to a single currency, does this potentially pose a problem for the conduct of a one size fits all monetary policy for the euro area? To explore this question, we conduct a counterfactual experiment that asks what the monetary policy of the ECB looks like relative to the hypothetical interest rate that might have been obtained if individual euro area members’ set interest rates after 1999. We may also turn the question around and ask what monetary policy might have looked like in each euro area member state had the ECB been responsible for setting interest rates prior to 1999?

Third, it is widely assumed that policy-making is forward-looking. Therefore, estimated policy rules should reflect the process by which central banks set interest rates. This requires

variables that are unobservable. Current best practice involves estimating reaction functions using GMM with relatively little attention paid to choosing instruments, and still less whether they are statistically relevant in the estimated reaction function. We are the first to report an extensive series of instrument relevance tests for monetary policy rules.

Alternatively, we can rely on proxies for expectations of key variables. Clearly, central banks and financial markets typically rely on forecasts. Such forecasts, at least indirectly, may incorporate asset price developments. Since good conduct in monetary policy has also been described as requiring a forecast-based response (e.g., Svensson 2003) it is somewhat surprising that there have been few attempts to estimate and empirically evaluate such rules. For example, relying on U.S. data, Fuhrer and Tootell (2004) show that estimated forward (and backward) looking policy rules admit asset prices because policy makers rely on real time data and forecasts in setting interest rates. Since forward- and forecast-based policy rules are non-nested hypotheses of interest rate determination, we finally resort to encompassing tests to decide whether one type of rule is capable of dominating the other.

A more recent concern about estimates of reaction functions is that they are generally based on final-revised data. Yet, as shown by Orphanides (2001), when data available in real time are used, policy implications can be substantially different. While progress on developing real time data sets for the euro area, and member countries, has been made (e.g., Herrmann, Orphanides, and Siklos 2005), it is not yet possible to systematically evaluate how flawed is the evaluation of policy rules based on the latest revised data for most euro area member countries. Using data constructed by Paloviita and Mayes (2005), and Gerberding, Worms, and Seitz (2005) an attempt is made to assess the importance of this problem. The real-time data problem suggests another avenue through which asset prices might have

played a more important role in monetary policy than heretofore argued. With some exceptions (e.g., real exchange rates, housing prices), asset price information is less likely to suffer from the measurement problems associated with, say, output gap proxies in real-time. The relevant data are available more quickly and suffer from relatively fewer biases. Nevertheless, reliance on asset prices in policy rules raises other challenges in evaluating policy rules, as we shall see.

The plan of the paper is as follows. The next two sections discuss the potential role of asset prices in a Taylor rule. A separate section describes the data and presents the empirical evidence. Next, we estimate Taylor rules for three core euro area countries. They are: France, Germany, and Italy.<sup>3</sup> We examine the plausibility of the various estimated rules. Our aim is to examine how well such rules perform under alternative treatment of various types of asset prices in policy rules. The paper concludes with a summary and lessons learned.

Briefly, if we include assets prices as additional instruments in forward-looking policy rules, estimates become not only more plausible but produce policy rules that achieve a better fit. Some asset prices are found to be statistically highly relevant instruments in forward-looking policy rules. The asset price that is most useful in helping explain nominal interest rate movements are not the same across the three countries considered. Nevertheless, more often than not, a monetary aggregate performs well. We interpret these results to mean that, whereas the central banks of France, Germany, and Italy did not directly respond to asset

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<sup>3</sup> Policy rules for smaller euro area members including Austria, Belgium, Finland, and the Netherlands were also examined but, as these did not materially affect the conclusions reported below, we do not discuss them any further. The real GDP of the three core area countries considered in this paper accounts for roughly three-quarters of euro area-wide real GDP.

price developments, these did influence expectations of inflation and the output gap. Indeed, it would be difficult for the ECB to know which asset price to respond to since, prior to 1999, different types of asset prices appeared to influence monetary policy performance in individual euro area countries. Turning to the role of real-time data, we find that there is a deterioration in the performance of policy rules relative to reliance on final revised data. However, we do not find, as did Orphanides (2001), that using output gap information in real-time produces seriously misleading advice, possibly because the information content of a variety of asset prices mitigates the influence of mis-measurement in the degree of economic slack.

## **2. Taylor Rules and Asset Prices**

While tradition, and a considerable amount of empirical evidence, points to inflation, output, and a desire to minimize interest rate volatility as chief among the concerns of most central banks (e.g., Favero and Rovelli 2003, but see Rudebusch 2002 and Lansing 2002), recent events have prompted some to consider the possibility that asset prices may also be included in a policy reaction function. While some of the debate has turned on whether the monetary authority should target asset price developments, central banks have argued against this position because they treat asset prices as forward-looking indicators of inflation and/or the output gap.

For example, Bernanke and Gertler (1999) make the case against including a reaction to asset prices whereas Cecchetti et al. (2000) argue the opposing view. One difference between both studies is that Cecchetti et al.'s paper investigates a relatively wider range of policy rules. They do search for the optimal rule. Hence, their result is, at best, based on an



approximately optimal rule. As argued in Filardo (2004), if non-fundamental asset price movements can have real economic effects, then they should be incorporated into the central bank's policy rule. The debate partly centers on the information content of asset prices concerning the future course of the economy. In a related fashion, Goodhart (2001) advocates a measure of inflation that goes beyond merely incorporating the effects of changing prices for goods and services to also include the impact of asset prices (e.g., equity and housing prices).

Nevertheless, we are left with the problem of asking which asset prices central banks might be most concerned with. In the case of most euro area economies one might well argue that policy makers are concerned with real exchange rate movements (e.g., Leitemo 1999, Leitemo and Røisland 1999, and Medina and Valdés 2002). Lately, however, attention has turned to the behavior of equity and housing prices, or some other financial indicator, as central banks are increasingly seen as responsible for stemming the cycle of booms and busts in asset prices. A difficulty which we do not tackle here in a systematic fashion is that, based on the evidence in International Monetary Fund (2000) and Bordo and Jeanne (2004), central banks ought to respond and, perhaps, measure booms and busts in equity and housing markets differently. Alternatively, central banks might separately react to some weighted average or linear combination of asset prices (e.g., Smets 1997).

Even if it is deemed desirable to incorporate a role for asset prices, the investigator faces a number of difficulties. In particular, there is no widespread agreement on how best to define equilibrium real exchange rate or asset price levels though these problems seem no less intractable than defining the "trend" in output used in deriving an output gap measure. In the case of equity and possibly housing prices, matters are complicated still further because

there is an element of “irrational exuberance” or a “bubble” component that is difficult to measure empirically. Of course, this paper is unable to address all of these issues. Instead, we rely on variations of “standard” filters to approximate the variables of interest. Nevertheless, in one experiment, we consider identifying the parameters of interest using information based on the volatility of monetary policy indicators. As pointed out in International Monetary Fund (2000) housing prices, for example, are less volatile than equity prices and this could have implications for macroeconomic performance more generally. Indeed, reliance on a volatility measure to help identify central bank reactions to inflation and output gap shocks is similar to the identification strategy proposed by Rigobon (2003).

One asset for which there is perhaps more agreement on the equilibrium outcome, but not how to measure it, is a monetary aggregate. The difficulty is that it is unclear whether financial innovations since the 1980s, and the shift in emphasis in the conduct of monetary policy toward reliance on an interest rate instrument, has reduced the predictive role of monetary aggregates for inflation in the “long-run”, and output in the short-run. Nevertheless, owing to the role of monetary aggregates in the operational framework of the ECB, and its key progenitor prior to 1999, namely the Deutsche Bundesbank, we also evaluate its role in the conduct of monetary policy, as do Clarida, Gali, and Gertler (1998), and Gerdesmeier and Roffia (2004).

Table 1 presents a selection of published estimates of Taylor rules, principally for euro area countries. The estimates tend to be consistent with the “Taylor principle”, according to which a larger than unit nominal interest rate response is required for any unit increase in inflation. Some are based on estimates of forward-looking reactions functions and others are of backward-looking variety. Ordinarily, they leave out a role for asset prices with the

notable exception of Bernanke and Gertler (1999) who examine how stock prices affect interest rate determination in the U.S. and Japan. Cecchetti (2003) also reports that the Fed reacted to stock market developments relying again on a Taylor rule type equation.

Table 1 about here

There is considerable variation in the estimated weights on the inflation and output gap objectives, the role of asset prices is not extensively investigated, nor is the robustness of results or the relative suitability of forward or forecast-based models extensively analyzed.<sup>4</sup> Finally, except for Germany (e.g., Faust, Rogers, and Wright 2001), there is little discussion of what ECB policy looks like for the euro area in relation to monetary policy in the member countries prior to the start of European Monetary Union. Instead, a few papers (e.g., Gerdesmeier and Roffia 2004, Gerlach-Kristen 2003) have relied on synthetic euro area data created for a sample before the ECB became solely responsible for monetary policy.

### **3. Estimation and Testing Strategy**

One motivation for estimating a variety of reaction functions is to determine how wide was the variation in policy rules behavior in selected euro area countries in the lead up to the start of European Monetary Union and the potential diversity in the potential influence of asset prices among euro zone members. We begin with a standard version of the reduced form version of Taylor's rule, which can be written in regression form as:

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<sup>4</sup> Gerdesmeier and Roffia (2004) perform an extensive comparison of forward versus backward-looking Taylor rules for the euro area.

$$i_t = \bar{i} + \gamma_{\tilde{\pi}} \tilde{\pi}_t + \gamma_{\tilde{y}} \tilde{y}_t + \rho i_{t-1} + \nu_t \quad (1)$$

where  $i_t$  is the (nominal) interest rate instrument of monetary policy,  $\bar{i} = (1 - \rho)\alpha$  the sum of the steady-state real interest rate and the annual inflation target, for  $\alpha$  see equation (2) below.  $\tilde{\pi}$  and  $\tilde{y}$  denote, respectively, the inflation and output gaps.  $\rho$  is the interest rate persistence or smoothing term, and  $\nu_t$  denotes a residual term.

If the reaction function is forward looking, the inflation gap is simply the difference between expected and targeted inflation rates,  $E_t(\pi_{t+n}) - \pi^*$ . In a similar fashion, the output gap is the percentage deviation of real GDP from its potential level. The coefficients  $\gamma_{\tilde{\pi}} = (1 - \rho)\beta$  and  $\gamma_{\tilde{y}} = (1 - \rho)\theta$  reflect the weights policy makers place on inflation versus the output gap. A central bank's reaction to inflation and output is captured via estimates of  $\beta$  and  $\theta$ , respectively. Both parameters measure the steady-state responses of interest rates to inflation and the output gap, though they must be “identified” since the economy responds to the same variables.<sup>5</sup> The coefficients in the reaction function are obtained from an expression that summarizes the interest rate targeting policy of the central bank, namely:

$$i_t^* = \alpha + \beta_n E_t(\pi_{t+i}) + \theta_n E_t(\tilde{y}_{t+k}) \quad (2)$$

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<sup>5</sup> In an optimizing framework (e.g., Clarida, Gali, and Gertler 2000), these weights also reflect to some extent the underlying structure of the economy as well as the persistence of economic shocks.

where  $i_t^*$  is the interest rate target,  $E_t(\tilde{\pi}_{t+i})$  and  $E_t(\tilde{y}_{t+k})$  are the conditional expectations of inflation and of the output gap,  $i$  or  $k$  periods ahead,  $i, k \geq 0$ .<sup>6</sup>

Since  $E_t(\tilde{\pi}_{t+i})$  and  $E_t(\tilde{y}_{t+k})$  are unobservable, instruments serve as proxies. Alternatively, a forward-looking central bank might be best thought of as acting on the basis of forecasts of inflation and output growth. In this case, central bank forecasts or private sector forecasts serve as proxies. Hence, a version of (1), where  $i = k = n$ , using forecasted values would be written as:

$$i_t = \bar{i} + \gamma_{\tilde{\pi}}^* \tilde{\pi}_{t+n,t}^f + \gamma_{\tilde{y}}^* \tilde{y}_{t+n,t}^f + \rho i_{t-1} + v_t, \quad (1a)$$

where most of the variables have been previously defined, and  $\tilde{\pi}_{t+n,t}^f$  and  $\tilde{y}_{t+n,t}^f$  are, respectively, published forecasts of inflation (or the inflation gap) and the output gap  $t+n$  periods ahead made at time  $t$ . Since we rely below on quarterly data, in what follows we shall focus our attention on the case where  $n = 4$ , which is typically what is assumed in the literature.

A difficulty with the resort to forecasts is that it is not immediately clear whether the published forecasts are based on the assumption that interest rates are unchanged.<sup>7</sup> A second difficulty is that, depending on the source of the forecast, updating is done at different

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<sup>6</sup> In equation (2) the inflation target has been normalized to zero. Equation (1) is then derived from the relation  $i_t = \rho i_{t-1} + (1 - \rho) i_t^* + v_t$ . Moreover, while there is no requirement that  $i = k$ , this is the general practice followed in empirical work.

<sup>7</sup> Fuhrer and Tootell (2004) point out that if the correlation between the contemporaneous asset price in question and lagged interest rate changes is significant, this will tend to bias the coefficient on the asset price variable in the estimated policy rule away from zero. Goodhart (2005) shows for UK data that using “ex post” forecasts, that is, ones made after an interest rate decision, results in a potentially serious misspecification.

intervals, namely monthly, quarterly, or even semi-annually. A third problem is that forecasts are for inflation (or real GDP growth) for a particular calendar year. Yet it is conceivable that, as the forecast period approaches, the arrival of additional data are used to change forecasts. As a result, we assume that the data available to the forecaster is augmented as far as possible to reflect the data that would be used to generate inflation and real GDP growth forecasts for the relevant calendar year. Clearly, we do not know exactly the information set used by the various forecasters so our attempt to control for the changing span of time between the forecast period and the data that would be available in generating the forecast is an approximation only.

As noted previously, there is some evidence about the desirability of incorporating a mix of forward or even forecast-based elements. A natural way of considering such elements, not heretofore considered, is to ask whether one type of reaction function estimate can “encompass” another. In this fashion, forward and forecast-based Taylor rules are tested against each other to determine whether, statistically speaking, it might be preferable to estimate such rules as linear combinations of each other.

Next, the question arises how to proxy the inflation target  $\pi^*$  necessary to evaluate the inflation gap  $\pi_{t+n} - \pi^*$ . Usually,  $\pi^*$  is assumed to be a constant (say 2%). Since our conclusions were unaffected by the type of inflation gap proxy used, we only report results which assume a constant inflation objective.<sup>8</sup> This has the slight advantage of allowing for a

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<sup>8</sup> Stock and Watson (2003) recommend a one-sided HP filter. In the present paper we also use the regular HP filter with a standard smoothing parameter (1600) and a larger smoothing parameter (4800). Next, we also assumed that the inflation target is the mid-point of the spread between the average annual inflation rate in the euro area countries and the average annual inflation rate in the three lowest inflation rate countries in the euro area plus 1.5%, as

comparison with the bulk of the literature on Taylor rules. Similarly, estimation of the output gap has been problematic, especially since subsequent research has shown that the interpretation of monetary policy actions in a historical context is very much dependent on whether policy makers at the time based their decisions on a mis-measurement of the output gap (e.g., Orphanides (2001) for the US). As in the case of the inflation gap, we utilize an HP filter as there are relatively few alternative proxies for the output gap in a cross-section of countries.<sup>9</sup>

The last consideration hints at potential problems which stem from reliance on final revised data instead of real-time data. Other than for Germany, we are not aware of the availability of a real-time data set for Italy and France.<sup>10</sup> For the euro area the available data (Gerdesmeier and Roffia 2005) are, obviously, a too short sample for our purposes. However, Paloviita and Mayes (2005) create a real-time data set based on OECD Economic Outlook publication sampled at the annual frequency while Gerberding, Worms and Seitz (2005) have done the same for Germany.<sup>11</sup> The problem noted earlier in dealing with the delay between the arrival of new information and the forecast period also poses a difficulty for our real-time specified in the Maastricht Treaty (article 121, ex article 109j). The conclusions are robust to these alternatives.

<sup>9</sup> Much of the literature uses a two sided HP filter for convenience, or an alternative measure of the economy's capacity, but comparable time series are not available for the vast majority of euro area countries. We also generate, but do not report here, estimates of the output gap based on a Blanchard-Quah type decomposition with no impact on our conclusions.

<sup>10</sup> After this paper was written we became aware of Golinelli and Parigi (2005) who have constructed a real-time data set for GDP for Italy.

<sup>11</sup> Quarterly German data from Gerberding, Worms, and Seitz (2005), available at [www.bundesbank.de/vf2/vf2\\_daten.en.php](http://www.bundesbank.de/vf2/vf2_daten.en.php), are not compatible with the data for Italy and France developed in Paloviita and Mayes (2005).

data set.<sup>12</sup> However, it is the only data set that contains comparable information for all three of the euro area countries considered in this paper.

Following much of the empirical literature we estimate (1), and its variants, via GMM which is the standard estimation approach adopted by most in the literature, in spite of some drawbacks (Mavroeidis 2004, Jondeau, Le Bihan, and Galles 2003). An important issue, which is only now receiving increased scrutiny, concerns the choice of instruments and, more importantly, their relevance. For example, the *J*-test for goodness of fit is usually reported without comment.<sup>13</sup> Nevertheless, focusing on such a test alone also poses problems because one risks choosing a model with theoretically implausible coefficients in the rule. Therefore, we augment the tests for over-identifying restrictions by reporting the information criterion GMMIC due to Andrews (1999), used to determine whether the chosen instruments are orthogonal to the error term, and Hall and Peixe's (2003) test for instrument relevance CCIR. Both of these tests are based on GMM estimation (also see Hall 2005).

We also consider an additional test for instrument relevance based on the two stage least squares estimation approach. Shea (1997) points out that regressing the endogenous variables against the chosen instruments can be misleading when there are several endogenous variables in the estimated specification. Jondeau, Le Bihan, and Galles (2004) report simulation evidence that finds that even small mis-specifications can lead to implausibly

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<sup>12</sup> The authors rely on updates published by the OECD each December.

<sup>13</sup> Authors typically report the *J*-statistic to determine the horizon used by the policy makers. However, it is difficult to reject the null of the validity of chosen instrument sets and, hence, it is difficult to discriminate among competing versions of the same estimated policy rule. They do not, however, report whether plausible reaction function estimates were obtained for all the combinations reported. As we shall see, estimates at different horizons do not always produce plausible coefficients.



large coefficients for the forward looking variables in GMM estimated equations. Godfrey (1999) develops a simple measure for computing instrument relevance based on a two stage least squares estimation specification.<sup>14</sup> With this in mind, we use asset prices, among other variables, as instruments in the GMM phase of estimation and test both their impact on the fit of estimated policy rules as well as their relevance as instruments.<sup>15</sup>

Our strategy is to obtain estimates of the main parameters of interest, namely,  $\gamma_{\tilde{\pi}}$ ,  $\gamma_{\tilde{y}}$ ,  $\theta$ , and  $\beta$ , for select euro area members prior to 1999. Then, relying on historical estimates, we ask to what extent the behavior of the euro repo rate since 1999 reflects pre-ECB monetary policy as well as the potential role of asset prices since the adoption of the single currency. We not only consider whether asset price departures from some proxy for their equilibrium value influences interest rate determination but also whether the volatility of asset prices serve as superior instruments in forward-looking rules.

Once the reaction functions are estimated we can back-out the implied target interest rates over time, i.e.,  $i_t^*$ , to determine how well a reaction function “fits” with actual interest

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<sup>14</sup> While the test for instrument relevance is based on two stage least squares it has the advantage that it can accommodate more than one endogenous variable. Recent tests for instrument relevance can be considerably more complex when there is more than one endogenous variable. Jim Stock’s weak instruments web page updates information on this topic [ksghome.harvard.edu/~jstock/ams/websupp/index.htm](http://ksghome.harvard.edu/~jstock/ams/websupp/index.htm).

<sup>15</sup> An objection that can be raised is that GMM is a non-linear estimation technique while the tests of instrument relevance used here are based on two stage least squares estimates. Since GMM is the estimation technique of choice we retain its use. Furthermore, relying on other tests of instrument relevance (e.g., the F-test as in Stock, Wright, and Yogo 2002) we obtain comparable results. Gerdesmeier and Roffia (2004) report finding few substantive differences between their GMM and two stage least squares estimates for the euro area.

rate developments in the individual countries considered. Since (1) and (2), in particular, are non-nested, we can also ask, via encompassing tests, which measure provides the best “fit” overall (Chong and Hendry 1986, and Collins and Siklos 2004).

Finally, we perform a series of counterfactual experiments. First, we ask what interest rates would have been, in the three countries considered if European Monetary Union had not taken place. Second, we examine what interest rates would have been had the euro been introduced earlier. As usual, all such counterfactuals, while speculative, at least give an indication of whether ECB monetary policy has been too tight for some countries. Such a strategy provides an indication of how tight monetary policy had to be, prior to 1999, in order to generate conditions favorable to the introduction of the euro.

## **4. Data and Empirical Evidence**

### **4.1 Data**

All data were either quarterly at the source or converted to the quarterly frequency by taking monthly averages where necessary. In the case of real GDP, seasonally unadjusted data were used and adjustment was made using X-11. To generate the output gap we use the HP filter with various smoothing parameters but eventually settled on the standard value of 1600 as our conclusions were unaffected by alternative values up to 9600 considered. For the euro we also rely on estimates of the output gap used in the euro area-wide model.<sup>16</sup>

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<sup>16</sup> Data were obtained from the Bank of Finland. However, members of the Euro Area Business Cycle Network at [www.eabcn.org](http://www.eabcn.org) may also access the relevant time series.

We also experimented with a one-sided HP filter due to the well-known endpoint problem with such filters with no discernible impact on our conclusions.<sup>17</sup> As the BIS made available its aggregate asset price index (Borio and Lowe 2002) we can get some idea of the potential impact of relying on individual asset prices instead of some weighted average of key asset prices.<sup>18</sup> We consider the following individual asset prices: housing prices, equity prices, the real exchange rate, a measure of financial conditions proxied by the financial conditions index<sup>19</sup>, and a broad monetary aggregate. Next, we transform all variables, except the nominal interest rate, measured in percent, into 100 times the fourth order log difference, that is,  $(\log X_t - \log X_{t-4})$ , where  $X$  is the variable of interest.

Turning to the forecast-based rules, we employ one-year ahead forecasts of inflation, real GDP growth, or the output gap. The relevant forecasts were obtained from Consensus Economics, the OECD, and The Economist. Both the Consensus and Economist forecasts are

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<sup>17</sup> Rennison (2003) reports that the HP filter (though he seems to prefer a version augmented with the residuals from a reduced form Phillips curve), combined with estimates from a Blanchard-Quah structural vector autoregression, is the best output gap estimator. He generated data from an artificial economy and uses a Monte Carlo experiment to determine which from among several estimators of the output gap are able to mimic the “true” output gap. Turning points and amplitudes of the resulting output gap data published by the Bank of Canada are very similar to the residuals from HP filtered data.

<sup>18</sup> In the empirical results we rely on the nominal measure. The BIS’s asset price index is essentially a weighted average of equity, residential and commercial property prices, where the weights are their respective shares in private sector wealth. The calibration of weights has changed over time since the index was first introduced (Borio, Kennedy, and Prowse 1994).

<sup>19</sup> The weights estimated in Goodhart and Hofmann (2000) were used. The financial conditions index can be thought of as an extension of the monetary conditions index, representing a linear combination of interest rates and exchange rates, to include housing and stock prices.

monthly and were converted into quarterly via simple averaging. OECD forecasts are semi-annual and simple interpolation was used to generate quarterly forecast equivalents. Forecasts are for the CPI (or the HCPI in the case of the euro area) and real GDP growth except OECD forecasts where output gap forecasts were employed.

As mentioned earlier, real-time data are from Paloviita and Mayes (2005) and Gerberding, Worms, and Seitz (2005), and were constructed on an annual basis from the OECD's Economic Outlook. Rather than use interpolation we rely on the current year's real time output as a proxy that the central bank would use during the entire year. Clearly, this is a simplification. Alternatively, we can proxy real-time effects by changing the instrument set used in GMM and two stage least squares estimation of reaction function as follows. We rely on only three quarters and longer lagged values of final revised data to account for lags in revising data, as well as forecasts of the current period inflation and real GDP growth and output gaps. This approach also serves another function, namely to determine the deterioration, if any, in instrument relevance when the most recent two quarters of data are not available to policy makers.<sup>20</sup>

Figure 1 plots gaps in housing, stock prices and in the BIS's aggregate asset price index. Also highlighted are periods of "booms and bust" cycles, as defined by Bordo and Jeanne (2004) and Detken and Smets (2004). These are based on a type of moving average of the relevant prices shown. Sharp and, occasionally, persistent departures from the equilibrium proxy are evident. There appear to be few differences between HP filtering and quadratic or cubic de-trending, the most commonly used techniques in this context, though the amplitudes

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<sup>20</sup> Since forecasts also serve as instruments, and presumably are at least partly influenced by real-time data considerations, some information one and two quarters lagged is thereby incorporated into the estimation.

are greater when de-trending is used or, of course, when a larger smoothing parameter is chosen in the case of the HP filter. No “bust” period in equity prices were identified by Bordo and Jeanne (2004) for Germany and France, at least for the period covered here. Yet, both HP filtering as well as cubic detrending reveal substantial stock price declines on more than one occasion. For Italy, the two bust periods identified by Bordo and Jeanne (2004) appear broadly consistent with either filter used. However, comparable price declines at the beginning of the sample as well as during the later 1990s did not make the grade in this setup.

Figure 1 about here

Turning to housing prices reported in the middle panel in Figure 1, downturns identified via various filters broadly match those reported in Bordo and Jeanne (2004) but only one appears in the sample booms shown here across the three countries considered. Finally, if we examine the BIS’s nominal aggregate asset price index, and compare with the so-called high and low cost booms reported in Detken and Smets (2004), we find general agreement with the application of an HP filter to the data.

Rolling standard deviations of the asset prices considered here were also examined (not shown), for each of the three core countries in our data set. Calculations were based on the HP filtered version of the gap measures described above. Alternative de-trending procedures seem to make little difference to the result and are, consequently, not shown. Real exchange rate volatility behavior is similar for France and Germany, while there is a noticeable spike in real exchange rate volatility in Italy during the 1992 - 96 period. This period overlaps, of

course, with the 1992 EMS crisis. There is much more diversity in the volatility of housing prices across the three countries considered. Each of the core countries experiences spikes in volatility though at very different times. The volatility in stock prices is fairly common across the three countries shown here, though Italy experiences a spike in stock price volatility in the early 1980s not seen in the other two countries. The rise in stock price volatility is also noticeable in the last few years of the sample, after the ECB took over responsibility for implementing monetary policy.

The BIS's aggregate asset price index, which averages the effects of selected asset prices, i.e., equity, residential and commercial property prices, reveals distinct volatility "regimes" for each of the three core countries. France experiences three sharp rises in asset price volatility, namely in 1984 - 85, in 1989 - 91, and then again after 1998. In the case of Germany, there are also at least three volatility regimes, in 1983 - 85, 1989 - 90, and 1997 - 2000. Finally, Italy experiences two noticeable increases in volatility, the first in 1980 - 81, and then again in 1985 - 90. Obviously, this method of identifying volatility regimes is somewhat ad hoc and there are other methods that could be used to identify volatility regimes, e.g., Markov switching, though it seems unlikely that vastly different results would be obtained.

Figure 2 illustrates the differences between real-time and final revised estimates of the output gap. As is now well-known, the discrepancies between the first release of data and their ultimately revised value can be considerable. Nevertheless, overall differences appear somewhat less pronounced for German data than either for France or Italy. Moreover, differences between HP filtered data measurement of the output gap and ones based on the OECD's Economic Outlook (Paloviita and Mayes (2005) for the OECD's methodology) are

considerable. Unfortunately, the number of available observations for the OECD's proxy is small and this limits somewhat empirical testing relying on this data source.

Figure 2 about here

Since estimates of (1a) rely on published forecasts a potentially misleading role for asset prices could be obtained if forecasts are inefficient (Fuhrer and Tootell 2004). A simple test of forecast efficiency consists in estimating the following regression:

$$z_{t+i} = \delta_0 + \delta_1 z_{t+i}^f + \sum_{k=1}^j \delta_{2k} x_{t-k} + \xi_t, \quad (3)$$

where  $z_{t+i}$  is inflation or the output gap,  $z_{t+i}^f$  are the forecasts from Consensus Economics, the OECD Economic Outlook, and The Economist. All forecasts that appear in the forecast-based Taylor rule estimates discussed below set  $i = 4$ .<sup>21</sup>  $x_{t-k}$  ( $k$  is set either to 1 or 4) is proxied either by the BIS's aggregate asset price index, or the rate of change in housing prices, equity prices and the real exchange rate.

Estimation results (not shown but available on request) of (3) reveal that, for inflation, lags in the BIS's index do not provide additional information beyond Consensus or OECD forecasts. When asset prices are examined separately, we find that housing prices do not improve the explanatory power of equation (3) while the results are more mixed for the equity and real exchange rate variables. For example, The Economist's forecasts for inflation in France and Italy appear inefficient, otherwise all other forecasts are efficient. When all

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<sup>21</sup> Consensus forecasts are available for a longer horizon but we only consider the additional case for  $i = 8$ .

three asset prices are considered jointly, only OECD forecasts display some inefficiency. Turning to the output gap, there is almost no evidence of forecast inefficiency regardless of the forecast source.<sup>22</sup> Our findings show strong evidence of forecast efficiency for the BIS's index and when all three asset prices enter equation (3) jointly.

Equation (3) is specified in the conventional manner. However, the discussion in Filardo (2004) suggests that, beyond  $x_{t-k}$ , predictions for  $z_{t+i}$  should also not be improved upon by the addition of lags in inflation, output, or both. Indeed, when we re-estimate (3) adding lags of  $z_t$  the results (not shown but available on request) are mixed. Therefore, conditional on  $x_{t-k}$ , forecasts of  $z_{t+i}$  could be improved by the addition of the past history of  $z_t$ . This could reflect a non-fundamental component. Hence, central banks perhaps ought to devote more effort at detecting how and when to respond to asset prices directly. We leave this possibility for future research.

#### **4.2 The Choice of Instruments and Instrument Relevance**

Table 2 panel A shows the non-rejection rates for the null that the over-identifying restrictions cannot be rejected, based on the  $J$ -test. Since it is unclear how forward-looking central banks are, equation (2) is estimated for a range of horizons ranging from the contemporaneous up to eight quarters ahead. Examining the results for France, Germany, and Italy, it is difficult to discriminate among instrument sets and horizons in spite of the relatively large number of instrument sets considered and the inclusion of a wide variety of

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<sup>22</sup> Other than for the OECD, forecasts are for real GDP growth. Whether (3) uses GDP growth or a proxy for the output gap created in the manner previously described does not change the outcome of the forecast efficiency tests.



asset prices as candidates. Interestingly, the situation changes dramatically when euro area-wide data are considered. While the non-rejection rates are modest at short horizons, the  $J$ -test rejects the over-identifying restrictions at horizons of a year or longer.

Table 2 about here

Panel B of Table 2 considers Andrews' GMM information criterion GMMIC for final revised data whereas panel C replicates the results for the available real-time data. To conserve space results are shown only for the cases where inflation and the output gap enter either contemporaneously or four quarters ahead. Regardless of the specification considered, or whether final revised or real-time data are employed, the instrument set most orthogonal to the error term consists of a combination of individual asset prices, namely the real exchange rate, housing prices, and equity returns. Nevertheless, if an investigator were to choose among individual asset prices then, in the case of final revised data, a monetary aggregate is the best choice most of the time. Turning to real-time data, the BIS' asset price index and equity returns appear to be the best choices overall.

As argued in Hall and Peixe (2003), Andrews' test represents a necessary but not sufficient condition for determining the optimality of the chosen instrument set. They introduce a test of instrument relevance based on the canonical correlation coefficient CCIC. Panel D and E of Table 2 present the relevant test statistics and these largely support the conclusions based on Andrews' information criterion. Indeed, if a more parsimonious instrument set is preferred, the evidence suggests that, when relying on final revised data, the BIS' asset price index is preferred for Germany and the euro area while housing price and a

monetary aggregate would be chosen, respectively, for France and Italy. However, when real-time data are used, either the BIS' index, or, what is a comparable indicator, a financial conditions index represents the best second choices.

#### **4.3 Taylor Rule Coefficient Estimates and the Role of Asset Prices**

Table 3 presents estimates of the steady-state parameters in forward-looking Taylor rules for a variety of instrument set combinations. While previous testing ostensibly narrowed down the likely candidates as instruments these tests are based on GMM estimation and these are known to be sensitive to the chosen specification. Accordingly, the earlier tests are supplemented in the tests for instrument relevance based on two stage least squares estimation of forward-looking rule (2). Finally, as an additional indicator of goodness of fit, we show the root mean squared error (RMSE) for the difference between actual and in-sample forecasts of nominal interest rates for each specification considered. Results shown in panel A are based on final revised data while those in panel B are based on available real-time data.

Table 3 about here

Generally, the various instrument relevance tests and the goodness of fit measure based on the RMSE find evidence in favor of the instrument set consisting of a combination of individual asset prices. This result is certainly suggestive of the notion that central banks look at a variety of indicators when setting the instrument of monetary policy. Not surprisingly, in

view of the role of the exchange rate in the European context, the real exchange rate is a close second, at least based on the RMSE criterion.

The role of instrument choice is clearly seen from the steady-state parameter estimates on the inflation and output gap variables. The responses to inflation range from 0.94 to 1.93 across the three countries considered with responses smallest for France and largest for Germany. Indeed, for Germany and Italy, the responses are typically significantly above one, as required by the Taylor principle. Interestingly, if the policy rule is chosen on the basis of instrument relevance, the coefficient on inflation is generally smaller than when the “standard” set of instruments are chosen. These estimates are also well within the range of earlier published results as can be seen in Table 1. Turning to the output gap, the steady-state coefficients are generally insignificant, at least if a 5% critical value is adopted, but not so when the instrument set is the preferred one, accordingly to the various other tests considered. Hence, the output gap remains an important variable in the policy rules of all three countries.

As shown in panel B of Table 3, however, replacing final revised output gap data with real-time data does not greatly change the estimates of the steady-state parameters for either inflation or the output gap, that is, when the preferred instrument set (“combined”) is chosen. Nevertheless, the additional instrument relevance tests do not point as clearly and unambiguously to a particular instrument set as being the most preferred over others. In addition, differences in RMSE across the various specifications considered are negligible. Therefore, while policy rules appear similar as between final revised data and real-time data, some of the tests that permit discrimination among different information sets are less helpful. Recall that there was no selection problem when GMM-based tests are relied upon.

Next, we turn to estimates for the euro area in Table 4. Owing to an absence of real-time data for the samples considered only estimates using final revised data are shown. The most striking results are the differences between the estimates that alternatively include and exclude the era since the ECB became responsible for the conduct of monetary policy in the euro area. Based on the full sample estimates, the ECB appears to react more aggressively to inflation and the output gap than for a sample that ends in 1998. Indeed, the ECB appears to react most aggressive toward the output gap when either the standard set of instruments are used or are augmented with lags in the money gap. In contrast, the ECB reacts positively and significantly to the output gap and somewhat less aggressively to inflation when a combination of different asset prices form the instrument set. Nevertheless, it should be noted that, based on the additional instrument relevance tests, and the RMSE criterion, either a euro area-wide wealth measure or the money gap are the additional instruments that appear to improve estimates of the policy rule. As before, results based on GMM estimation were less ambiguous.

Table 4 about here

Table 5 presents RMSE estimates for the forecast-based rule represented by equation (1a). The results are quite clear-cut. First, paralleling the results for final revised data, specifications which rely on one year forecasts of the output gap and inflation, and are augmented by lags in a combination of asset prices, provide the best overall fit among the various specifications considered. This is true for each of the three countries considered, as well as for the euro area-wide data. Furthermore, forecasts from The Economist outperform

slightly those based on Consensus forecasts while OECD forecasts, which are published only semi-annually, perform most poorly of all. Lastly, at least based on Consensus Forecasts, there is little deterioration in the in-sample forecasting performance between one and two year-ahead forecasts. A comparison between the RMSE either for final revised data, real-time data, or forecast-based data suggest the latter provide superior estimates.

Table 5 about here

To gauge the comparative advantage of one type of policy rule over another, Table 6 presents encompassing tests that rely on the best performing policy rule estimates based on final revised and forecast-based data. At the 1% level of significance the results reveal a clear-cut preference for forecast-based rules over forward-looking estimated rules. This result carried over at more modest significance levels with the possible exception of France at the 5% level.

Table 6 about here

#### **4.4 Counterfactual Experiments**

The results in Table 4 suggest that European Monetary Union produced a significant change in policy response parameters. Given the small number of observations available for the ECB era, we can obtain additional insights by performing some counterfactual experiments. The results are shown in Figure 3. Figure 3 (a) forecasts interest rates if policy rules estimates prior to European Monetary Union carried over past 1998. Until 2001, euro

area-wide interest rates are considerably lower than if France, Germany, or Italy were able to continue to set interest rates. Indeed, there is no evidence that euro area interest rates are higher than they might otherwise be.

Figure 3 (b) compares interest rates implied by the preferred estimated rule (Table 3) against a synthetic estimate of a euro area-wide interest rate. The ECB, had it been responsible for monetary policy prior to 1999, most resembles Banque de France policy. The ECB counterfactual interest rate is too high for Germany and too low for Italy. Note, however, that differences virtually disappear by about 1995.

The final counterfactual experiment, shown in Figure 3 (c), consists in asking what interest rates are implied by policy rules estimated using real-time data that rely on the preferred instrument set. When a standard set of instruments is augmented by lags in the money gap it suggests that the euro area-wide interest rate was too high but only until 1999, that is, when the ECB became responsible for monetary policy. Indeed, ECB policy looks just about right according to either implied policy rule. Indeed, when euro area-wide financial wealth is part of the instrument set the resulting implied interest rates closely resemble the actual and synthetic euro area-wide interest rate. Hence, there is little indication that real-time data misled policy makers in the euro area since 1978.

## **5. Conclusions**

It is commonplace to estimate forward-looking rules via GMM. Relatively little effort has been devoted to ascertaining the performance of different instrument sets and their relevance. The appropriate tests suggest that asset prices, especially a combination that consists of a real exchange rate, equity returns, and housing prices, play a significant role in determining the

responses to inflation and the output gap in the euro area before and since the ECB became solely responsible for monetary policy. In contrast to Orphanides (2001), we do not find that policy rules estimated with real-time output gap data mislead policy makers, though central banks appear relatively less aggressive toward inflation and the output gap than based on policy rules estimated with final revised data.

Lastly, a series of counterfactual experiments suggest that interest rates in France, Germany, and Italy would be higher in the absence of EMU. In general, however, ECB policy appears neither too tight nor too loose relative to a monetary policy that might have been carried out had the central banks in France, Germany and Italy been able to continue to independently set interest rates.

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**Table 1: Selected Estimates of Taylor Rule Parameters**

Author	Type <sup>5</sup>	Inflation	Output Gap	Other	Sample
Bernanke and Gertler (1999)	FL	1.12 - 2.21 1.60 - 1.71	0.20 - 0.33 0.14 - 0.20	0.19 - 0.29 -.08	Japan 1968 - 89 US 1968 - 89
Cecchetti and Krause (2001) <sup>1</sup>	BL	0.99 (0.99) 0.99 (0.99) 0.47 (0.98) 0.96 (0.99) 0.25 (0.88) 0.98 (0.99)	0.00 (0.00) 0.00 (0.00) 0.53 (0.02) 0.04 (0.00) 0.75 (0.12) 0.02 (0.00)	NA NA NA NA NA NA	Austria 1982 - 89 (1990 - 97) Belgium France Germany Italy UK
Cecchetti (2003)	FL	0.34 - 0.67	0.41 - 0.50	-0.65 s <sup>2</sup> -0.23 bank <sup>3</sup>	US 1990 - 2003
Clarida, Gali and Gertler (1998)	FL	1.10 - 1.37 1.81 - 2.04 1.05 - 2.20 0.48 - 0.98 0.59 - 1.33 0.59 - 0.91	0.25 - 0.35 0.03 - 0.10 0.14 - 0.52 0.17 - 0.28 -0.07 - 0.88 -0.03 - 0.22	NA NA NA	Germany Japan US UK France Italy 1979 - 89
Smets (1997)	FL FL	2.83 - 3.09 0.85 - 2.91	0.26 - 0.36 1.00 - 2.01	0.00 -0.14 - 0.22 <sup>8</sup>	Australia Canada 1989 -96
Clausen and Hayo (2003)	FL	2.28 2.89 2.02	1.73 0.49 2.46	NA NA NA	France Germany Italy 1979 - 96
Gerlach and Schnabel (2000)	BL FL	1.51 - 1.58 0.98 - 1.62	0.45 - 0.49 0.22 - 0.32	NA -0.03 - (-0.56) <sup>4</sup>	1990 - 98 Euro-11
Faust, Rogers and Wright (2001)	FL	1.31	0.18	NA	Germany
Gerlach-Kristen (2003)	FL	2.73	1.44	NA	Euro area 1988 - 2002
Gerdesmeier and Roffia <sup>6</sup> (2004)	BL FL FL BL	1.53 1.82 1.86 -0.01	1.39 0.77 0.26 2.18	NA NA 0.41 <sup>4</sup> 0.28 <sup>4</sup>	Euro area 1985 - 2002
Hayo and Hofmann (2003)	FL	1.25	0.32	NA	Germany 1990 - 98
Hetzel (2000)	BL	1.56 1.16	0.62 0.14	NA NA	Greenspan 1987 - 99 Volcker 1979 - 87
Chadha, Sarno, Valente (2003)	FL	1.69 - 2.39 0.98 - 1.46 1.10 - 2.37	-0.72 - 1.07 -0.61 - 0.60 -0.02 - 0.05	0.07/0.02 0.01/0.89 0.02/0.002	US UK Japan 1979 - 2000

### **Table 1: Selected Estimates of Taylor Rule Parameters (Continued)**

Note: <sup>1</sup> Coefficients are such that they are constrained to sum to 1 and represent measures of aversion to inflation and output variability; <sup>2</sup> measure of equity premium risk; <sup>3</sup> measure of stress in the banking system; <sup>4</sup> money growth, lagged inflation, fed funds rate, or real euro/\$ exchange rate (separate coefficients not reported; NR); \* simulations; <sup>5</sup> type of reaction function: FL = forward-looking, BL = backward-looking; <sup>6</sup> the first two sets of estimates are for their equations (20), (21), Table 3; the second set is for their equations (13a), (13b), Table 6; <sup>7</sup> Data are for stock returns for Germany apparently since authors sometimes refer to the euro area and the EU3 as adequately represented by German data (op. cit., p. 16); <sup>8</sup> Average of three asset prices: a trade-weighted exchange rate, 10 year government bond yields a broad stock market index.



**Table 2: Testing Over-Identifying Restrictions**

Panel A: Rejection Frequencies<sup>1</sup>

$i=0, k=0 \text{ to } 8$			$i=0 \text{ to } 8, k=0$			$i,j=4$			$i,j=3$			$i,j=5$					
France	Germany	Italy	France	Germany	Italy	France	Germany	Italy	France	Germany	Italy	France	Germany	Italy			
25.4	23.8	19.0	23.8	22.2	25.4	14.3	14.3	28.6	71.4	28.6	42.9	14.3	28.6	14.3			
Euro area <sup>2</sup>			Euro area			Euro area			Euro area			Euro area					
47.6			41.2			36.5			41.3			85.7			100		
57.1			100			100			100			100					

Panel B: Andrews' GMM Information Criterion

k = 0									
i=	s	m	e	r	h	a	f	C	
FR <sup>1</sup>	0	-51.3	-68.1	-66.9	-66.0	-68.1	-67.3	-67.7	-100.9
	4	-51.3	-67.3	-65.5	-68.4	-67.7	-66.7	-68.7	99.1
GE <sup>2</sup>	0	-48.7	-65.7	-69.5	-65.6	-65.2	-63.8	-64.8	-99.7
	4	-51.8	-65.1	-67.1	-67.0	-65.9	-64.9	65.2	-99.2
IT <sup>3</sup>	0	-52.6	-69.3	-68.2	-67.4	-65.8	-66.7	-65.4	-98.0
	4	-50.8	-66.6	-65.8	-66.3	-64.5	-63.2	-65.3	-97.7
EA <sup>4</sup>	0	-54.4	-65.1	-67.2	-65.2	NA	-70.1	-69.5	-100.4
	4	-55.0	-66.0	-68.8	-68.7	NA	-70.0	-70.3	-101.0
k = 4									
FR	4	-50.3	-66.1	-67.5	-66.2	-67.9	-67.8	-66.9	-100.0
GE	4	-50.8	-67.2	-66.4	-67.2	-64.4	-64.7	-64.6	-99.0
IT	4	-51.2	-67.4	-65.9	-67.4	-64.9	-62.9	-66.6	-97.8
EA	4	-54.8	-68.7	-71.2	-69.5	NA	-72.8	-73.1	-104.4

Panel C: Real Time Data

k = 0									
i=	s	m	e	r	h	a	f	C	
FR	0	-47.1	-60.1	-60.8	-50.2	-61.9	-62.7	-60.8	-92.3
	4	-47.0	-60.6	-61.8	-58.9	-62.0	-62.9	-61.0	-92.1
GE	0	-45.0	-61.1	-62.2	-60.7	-60.7	-61.9	-61.1	-93.0
	4	-41.5	-58.5	-58.3	-59.7	-58.2	-58.7	-58.4	-90.5
GE98	0	-42.1	-55.5	-58.3	-57.6	-55.7	-54.8	-57.3	-87.8
	4	-43.0	-54.8	-57.8	-58.5	-58.1	-58.5	-58.6	-87.6
IT	0	-46.1	-60.9	-62.1	-63.6	-59.5	-61.4	-63.1	-93.6
	4	-45.7	-60.6	-61.8	-63.0	-60.0	-61.3	-63.0	-94.3
FR	4	-45.5	-57.8	-60.3	-59.8	-59.6	-63.2	-60.1	-94.0
GE	4	-45.7	-58.3	-60.2	-59.6	-61.0	-62.4	-59.9	-90.3
GE98	4	-41.6	-54.7	-57.2	-56.1	-56.2	-55.7	-55.7	-85.6
IT	4	-45.1	-60.2	-62.1	-61.3	-60.3	-60.7	-61.2	-94.0

**Table 2: Testing Over-Identifying Restrictions (Continued)**

Panel D: CCIC, Final Revised Data

	FR	GE	IT	EA
s	-60.347	-60.170	-59.495	-61.775
m	-60.356	-59.600	-59.543	-75.172
e	-60.354	-60.219	-59.525	-76.714
r	-60.403	-60.230	-59.502	-76.726
h	-60.421	-60.222	-59.525	NA
a	-60.379	-60.272	-59.510	-77.718
f	-60.388	-60.224	-59.504	-74.519
c	-95.137	-94.955	-93.756	-110.98

Panel E: CCIC, Real-Time Data

	FR	GE	IT	EA
S	-52.277	-49.214	-52.114	-52.231
m	-51.676	-49.256	-52.199	-51.644
e	-52.288	-49.372	-52.176	-52.280
r	-52.312	-49.266	-51.601	-51.307
h	-52.273	-49.272	-52.104	-52.313
a	-52.349	-49.404	-52.211	-52.257
f	-52.329	-49.293	-52.123	-52.336
c	-87.442	-82.256	-86.315	-87.530

Notes: 1.  $i$  and  $j$  refer to the number of periods ahead in inflation and the output gap, respectively, as specified in equation (2).

The frequencies refer to the fraction of times the Null hypothesis (that the over-identifying restrictions cannot be rejected) cannot be accepted at the 5% level. For France, Germany and Italy all results are for the sample 1978-1998. All data are quarterly. Based on all instrument set permutations.

2. For the euro area the first column is for the 1978-1998 sample; the second column is for the sample 1978-2003.

**Table 3: Taylor Rule Coefficient Estimates and Additional Instrument Relevance Tests**

Panel A: GMM Estimates of Steady-State Parameters

<b>FRANCE</b>								
Asset Price	$\beta$	$\theta$	$\rho$	$r_{\pi}^2$	$r_y^2$	RMSE	F-test	
							$\pi$	$\tilde{y}$
Combined	1.12 (0.52)	3.88 (0.01)	0.89	0.60	0.97	0.97	294	5.40
BIS index	0.97 (0.88)	2.50 (0.17)	0.90	0.70	0.26	0.99	292	4.06
Equities	0.98 (0.93)	-1.43 (0.54)	0.87	0.51	0.07	1.00	279	4.29
FCI	1.09 (0.00)	2.15 (0.19)	0.90	0.73	0.21	0.95	266	5.89
Housing	1.04 (0.91)	7.30 (0.24)	0.92	0.03	0.43	1.07	365	1.62
Real ER	1.02 (0.00)	1.14 (0.28)	0.88	0.81	0.34	0.95	277	6.60
Standard	0.94 (0.76)	-0.36 (0.91)	0.87	0.45	0.05	0.99	360	5.15
Money	1.47 (0.30)	0.95 (0.17)	0.86	0.12	0.42	0.54	36.17	4.28
<b>GERMANY</b>								
Combined	1.52 (0.01)	1.36 (0.00)	0.89	0.99	0.94	0.52	32.33	4.84
BIS index	1.89 (0.00)	0.39 (0.29)	0.85	0.46	0.45	0.58	38.44	6.92
Equities	1.84 (0.01)	1.44 (0.00)	0.88	0.49	0.36	0.59	43.90	6.59
FCI	1.83 (0.00)	0.41 (0.11)	0.78	0.43	0.45	0.60	35.70	6.36
Housing	1.93 (0.00)	0.89 (0.03)	0.86	0.02	0.63	0.59	38.27	2.96
Real ER	1.43 (0.00)	1.03 (0.00)	0.88	0.95	0.49	0.51	41.81	5.88
Standard	1.93 (0.00)	0.69 (0.10)	0.85	0.35	0.30	0.59	50.31	7.58
Money	1.13 (0.78)	3.99 (0.52)	0.91	0.26	0.26	0.94	233	8.23
<b>ITALY</b>								
Combined	1.25 (0.01)	1.04 (0.05)	0.81	0.18	0.56	0.94	228	3.41
BIS index	1.37 (0.04)	2.21 (0.11)	0.81	0.14	0.26	1.01	251	4.84
Equities	1.45 (0.06)	3.15 (0.09)	0.85	0.12	0.16	1.02	266	4.90
FCI	1.38 (0.00)	2.15 (0.02)	0.82	0.15	0.36	1.00	289	4.87
Housing	1.30 (0.02)	1.54 (0.10)	0.80	0.00	0.39	0.98	268	1.94
Real ER	1.26 (0.00)	1.10 (0.16)	0.79	0.14	0.37	0.95	298	4.83
Standard	1.53 (0.06)	3.85 (0.09)	0.84	0.13	0.18	1.10	336	6.42
Money	1.64 (0.27)	4.44 (0.38)	0.86	0.44	0.42	1.05	303	5.66

Panel B: Real-Time Data GMM Estimates of Steady-Date Parameters

<b>FRANCE</b>								
Asset Price	$\beta$	$\theta$	$\rho$	$r_{\pi}^2$	$r_y^2$	RMSE	F-test	
							$\pi$	$\tilde{y}$
Combined	1.48 (0.23)	2.86 (0.06)	0.89	0.29	0.76	0.92	281	9.14
BIS index	1.28 (0.36)	1.76 (0.14)	0.87	0.28	0.69	0.92	329	12.52
Equities	1.20 (0.56)	0.99 (0.51)	0.87	0.35	0.68	0.92	306	9.17
FCI	1.30 (0.41)	1.75 (0.30)	0.89	0.38	0.55	0.92	294	11.31
Housing	1.30 (0.37)	1.75 (0.15)	0.87	0.30	0.40	0.92	386	7.50
Real ER	1.22 (0.54)	1.30 (0.52)	0.89	0.33	0.48	0.92	299	10.35
Standard	1.12 (0.69)	0.62 (0.70)	0.87	0.33	0.49	0.92	268	8.73
Money	1.06 (0.87)	0.15 (0.93)	0.88	0.14	0.40	0.92	268	8.73
<b>GERMANY</b>								
Combined	1.32 (0.48)	1.64 (0.01)	0.88	0.36	0.41	0.48	37.04	6.06
BIS index	1.47 (0.09)	1.05 (0.05)	0.83	0.35	0.44	0.51	39.83	8.77
Equities	1.42 (0.18)	1.16 (0.03)	0.84	0.35	0.40	0.51	41.11	7.33
FCI	1.58 (0.07)	0.84 (0.11)	0.83	0.33	0.36	0.52	37.93	6.00
Housing	1.45 (0.12)	1.16 (0.05)	0.84	0.27	0.42	0.51	37.90	5.72
Real ER	1.37 (0.38)	1.23 (0.09)	0.87	0.24	0.31	0.48	48.52	7.85
Standard	1.53 (0.05)	0.23 (0.09)	0.82	0.34	0.39	0.52	50.59	8.11
Money	0.91 (0.89)	2.34 (0.08)	0.91	0.27	0.40	0.48	45.30	12.12
<b>ITALY</b>								
Combined	1.24 (0.34)	1.53 (0.30)	0.90	0.38	0.99	0.95	284	18.95
BIS index	1.23 (0.34)	1.08 (0.48)	0.89	0.33	0.79	0.95	327	7.67
Equities	1.24 (0.33)	1.33 (0.34)	0.89	0.30	0.80	0.95	340	8.76
FCI	1.29 (0.28)	1.87 (0.28)	0.90	0.40	0.89	0.95	382	11.61
Housing	1.22 (0.38)	0.95 (0.50)	0.89	0.35	0.97	0.95	342	4.20
Real ER	1.23 (0.35)	1.33 (0.38)	0.89	0.34	0.82	0.95	386	9.98
Standard	1.22 (0.36)	1.09 (0.45)	0.89	0.32	0.86	0.95	445	9.75
Money	1.19 (0.31)	0.46 (0.94)	0.85	0.22	0.97	0.92	388	7.36

### **Table 3: Taylor Rule Coefficient Estimates and Additional Instrument Relevance Tests**

Note: Standard refers to the Taylor rule as in (2) with a standard instrument set consisting of a constant, 3 lags of the interest rate, 4 lags of inflation and the output gap, and 4 lags of the rate of change of oil prices. See Table 2 notes for additional explanations of terms and instrument set definitions.  $\beta$  is the steady state inflation parameter and  $\theta$  is the steady state parameter on the output gap. The null could not be rejected for all of the cases shown.  $r_p^2$  is the partial R-squared measure developed in Godfrey (1999), where  $p = \pi, \theta$ . It is calculated as  $(se^{OLS}/se^{GMM}) * (RSD^{GMM}/RSD^{OLS})$  where  $se$  is the standard error estimates for the coefficients on the endogenous variables  $(\pi, \bar{y})$ , and  $RSD$  is the residual standard deviations for the regressions estimated either via OLS or GMM. Partial  $r^2$  do not add up to 1 across columns. Data are quarterly. Estimates use GMM with a Bartlett kernel, Newey-West bandwidth, and HAC weighting matrix. RMSE = root mean squared errors. Sample is quarterly, 1978-1998. p-values in parenthesis for the Null that  $\beta=1$  (except for the real exchange rate and the FCI where the Null is  $\beta=0$ ) and  $\theta=0$ . Wald test (F-version) that the coefficients on lagged values of relevant asset prices are jointly insignificant. p-values are in parenthesis.

**Table 4: Taylor Rule Estimates and Additional Instrument Relevance Tests for the Euro Area**

	$\beta$	$\theta$	$\rho$	$r_{\pi}^2$	$r_{\bar{y}}^2$	RMSE	F-test	
							$\pi$	$\bar{y}$
1978 - 1998								
Combined	1.44 (0.02)	3.72 (0.00)	0.87	0.34	0.30	0.60	463	6.88
BIS index	1.32 (0.19)	2.71 (0.00)	0.89	0.30	0.86	0.55	541	9.71
Equities	1.42 (0.05)	3.73 (0.00)	0.88	0.16	0.16	0.59	533	7.77
Real ER	1.38 (0.00)	3.14 (0.00)	0.87	0.34	0.29	0.58	587	8.23
Wealth	1.26 (0.00)	5.06 (0.07)	0.93	0.47	0.44	0.67	449	11.36
Standard	1.38 (0.45)	4.97 (0.34)	0.95	0.33	0.17	0.72	534	13.76
Money	-0.03 (0.86)	0.11 (0.77)	0.98	0.34	0.30	0.47	360	10.89
1978 - 2003								
Combined	1.90 (0.01)	6.49 (0.01)	0.91	0.40	0.32	0.62	500	9.84
BIS index	1.67 (0.07)	4.86 (0.03)	0.93	0.28	0.81	0.54	542	13.81
Equities	1.90 (0.07)	6.51 (0.09)	0.92	0.14	0.11	0.61	568	11.34
Real ER	1.78 (0.00)	5.20 (0.01)	0.90	0.38	0.30	0.59	600	11.64
Wealth	1.61 (0.00)	8.43 (0.23)	0.95	0.47	0.35	0.67	531	14.90
Standard	2.28 (0.63)	0.24 (0.63)	0.97	0.31	0.14	0.75	636	18.35
Money	1.20 (0.92)	0.11 (0.64)	0.98	0.49	0.37	0.47	336	17.33

Note: See the note to Table 3. Wealth is the deviation of the log of real wealth for the euro area from its HP filtered level.

**Table 5: Root Mean Squared Errors in Forecast-Based Rules**

Asset Price	FRANCE				GERMANY				ITALY				EA		
	Cons 1	Cons 2	OECD	Econ	Cons 1	Cons 2	OECD	Econ	Cons 1	Cons 2	OECD	Econ	Cons 1	OECD	Econ
Combined	0.59	0.60	0.91	0.58	0.21	0.23	0.41	0.19	0.87	0.78	0.96	0.80	INS <sup>2</sup>	0.16	INS
BIS index	0.73	0.78	0.96	0.76	0.28	0.31	0.59	0.30	1.01	0.85	1.03	1.00	0.23	0.21	0.24
Equities	0.72	0.77	0.95	0.76	0.28	0.32	0.59	0.30	0.93	0.82	1.01	0.93	0.15	0.30	0.21
FCI <sup>1</sup>	0.64	0.68	3.23	0.68	0.24	0.30	1.58	0.25	1.00	0.87	3.97	0.93	0.24	0.29	0.29
Housing	0.72	0.76	0.96	0.74	0.30	0.31	0.57	0.31	1.03	0.88	1.00	0.98	INS	INS	INS
Real ER	0.67	0.67	0.98	0.70	0.29	0.32	0.48	0.26	1.03	0.87	1.02	0.95	0.24	0.24	0.23
Standard	0.75	0.79	0.99	0.78	0.32	0.35	0.60	0.33	1.04	0.90	1.03	1.01	0.38	0.33	0.37
Money	0.70	0.76	3.42	0.73	0.26	0.28	1.39	0.27	0.98	0.87	4.10	0.94	0.21	0.66	0.32

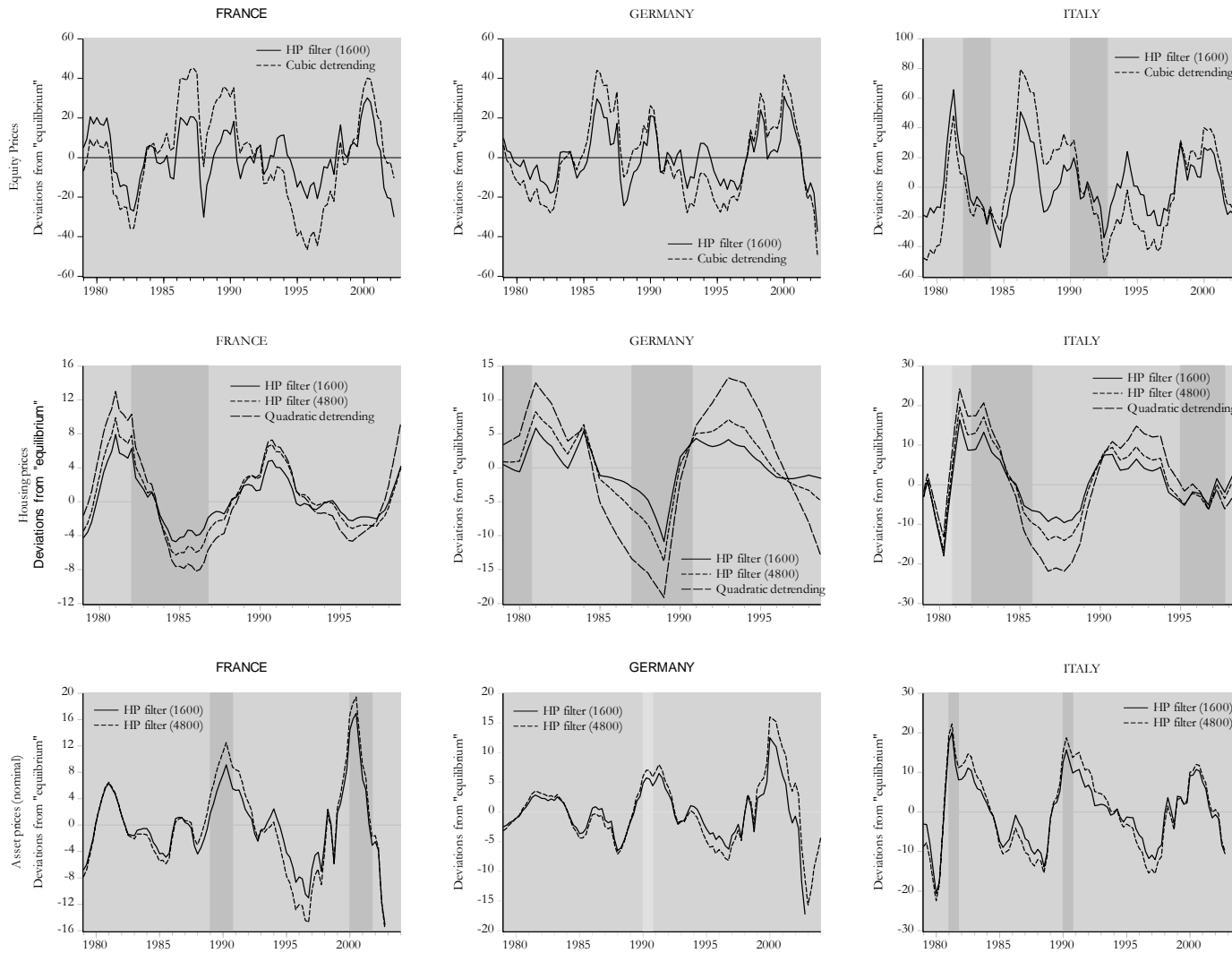
Note: The RMSE is based on equation (1a) in first differences. Asset price variables enter in levels. Cons 1 and Cons 2 are, respectively, one and two year-ahead forecasts of inflation and real GDP growth. OECD is the one year ahead forecasts of inflation and the output gap. Econ is the one year ahead forecast of inflation and real GDP growth. 1. Euro area financial wealth measure; 2. INS means insufficient or no data available.

**Table 6: Encompassing Tests**

Country, Sample	Source	Test Statistic
France, 1991 - 99	Forecast	0.85 (0.07)*
	Estimated	0.17 (0.08)**
Germany, 1991 - 99	Forecast	0.98 (0.03)*
	Estimated	0.02 (0.03)
Italy, 1991 - 99	Forecast	0.92 (0.08)*
	Estimated	0.09 (0.08)
Euro Area, 1996 - 2002	Forecast	0.97 (0.03)*
	Estimated	0.04 (0.03)

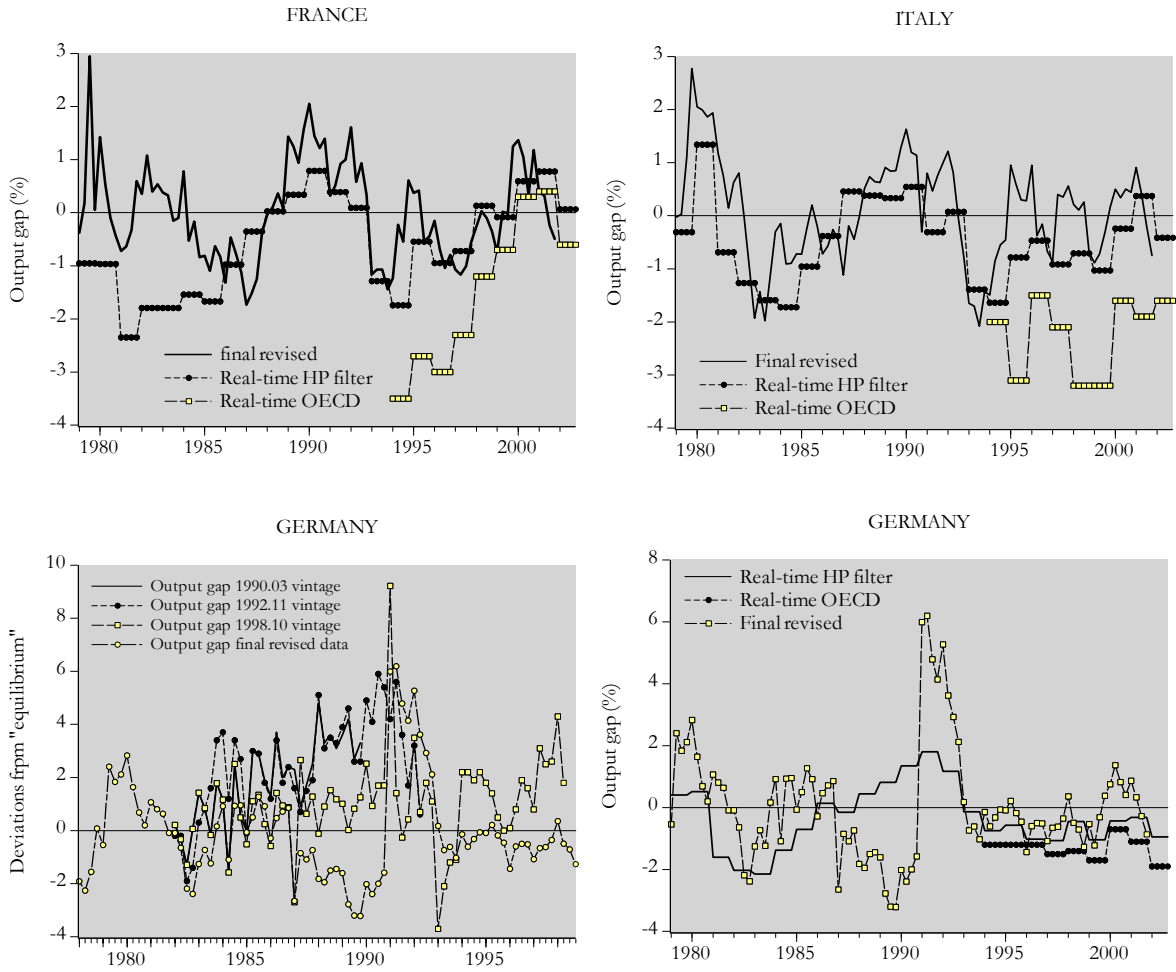
Note: The test statistic is derived from estimates of the specification  $i_t = \lambda_0 i_t^f + \lambda_1 i_t^e + \chi_t$  where  $i_t^f$  is the implied interest rate estimated from a forecast-based Taylor rule and  $i_t^e$  is the implied interest rate obtained from GMM estimation. Details are in the text. The last column gives the coefficient values and the standard error. \* indicates whether the null that  $\lambda_0 = 1$  is rejected at the 5% level, while \*\* indicates that the null that  $\lambda_1 = 0$  is rejected at the 1% level.

**Figure 1: Gaps in Asset Prices in Three Euro Area Countries**



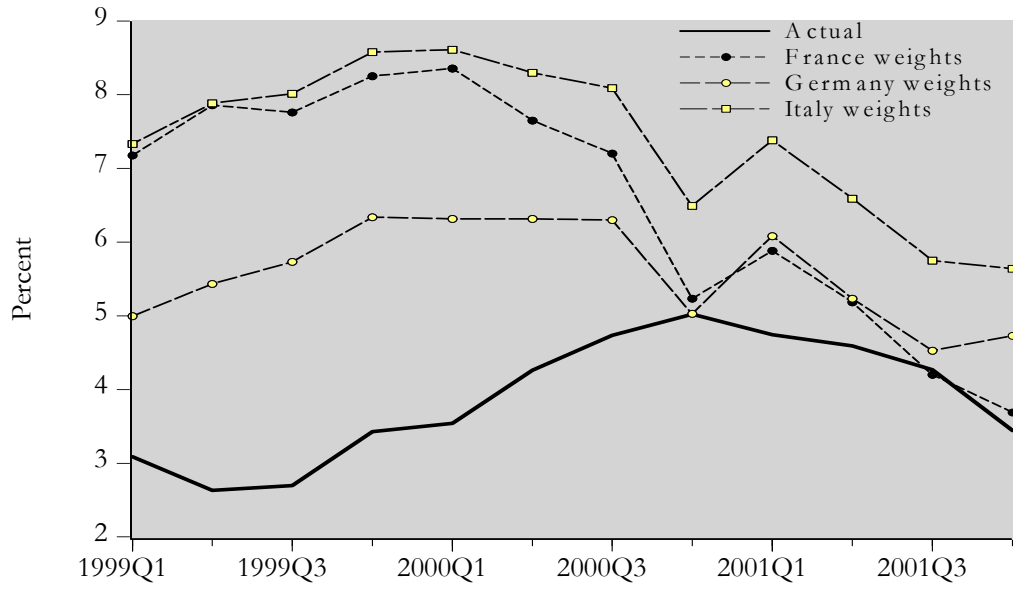


**Figure 2: Real-Time Versus Final Revised Data**

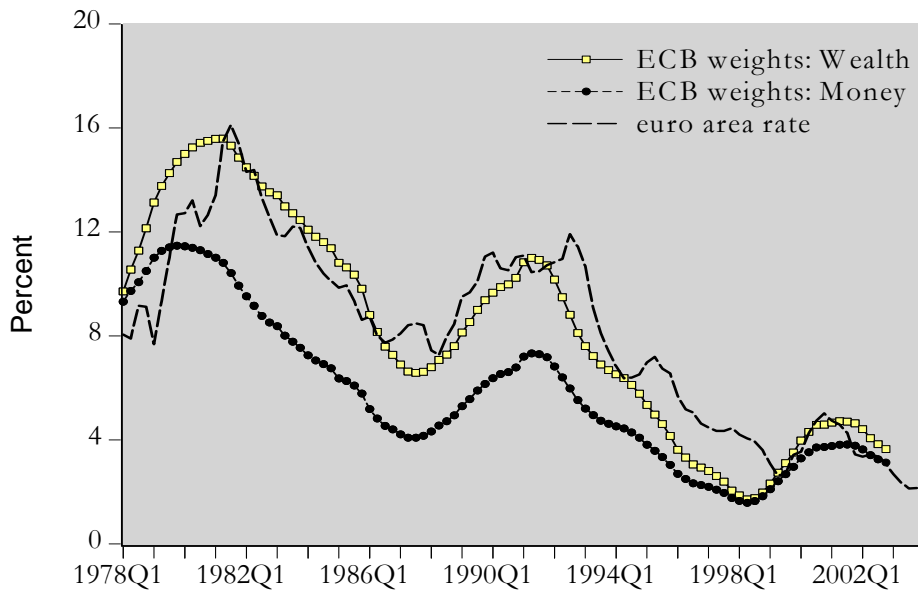


**Figure 3: Counterfactual Experiments: Final Revised Data**

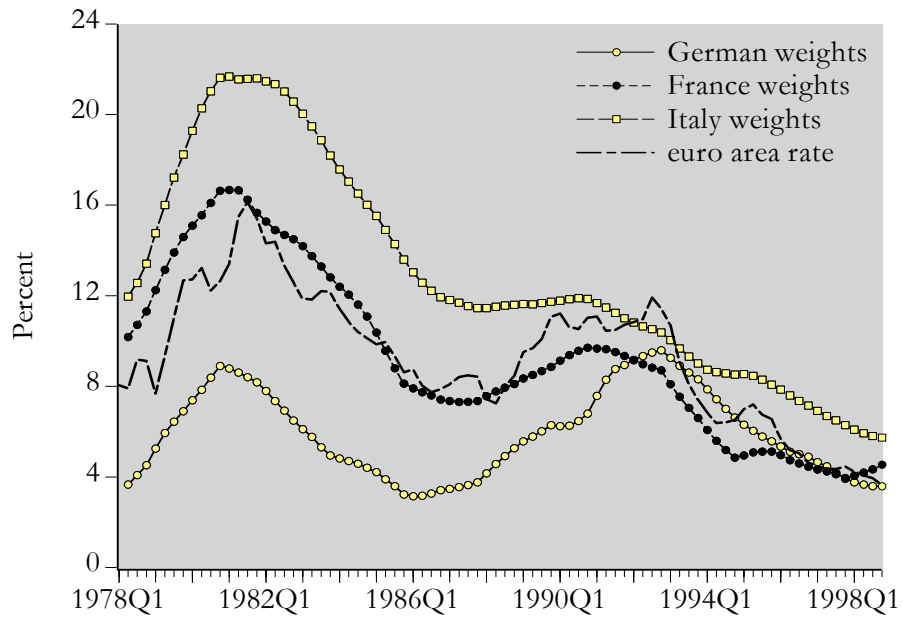
(A) Pre-ECB weights in setting euro area interest rates after 1999



(B) Post-ECB weights in setting interest rates before 1999



(C) Implied Interest Rates Based on Real-Time Data: Pre-ECB



Notes: In part (A), actual is the euro rate; France, Germany and Italy weights are the hypothetical euro area interest rates that would obtain if one of the Taylor rules shown in Table 3 carried over after 1999 (for France the FC I case; for Germany, the real ER case; for Italy the Housing case). In part (B), the synthetic euro area interest rate is shown alongside two hypothetical interest rates that use Taylor rule weights estimated in Table 4 for the euro area for the two samples shown (the case with the BIS index was used).

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