Taylor Rules and the Euro

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Abstract

This paper uses real-time data to show that inflation and either the output gap or unemployment, the variables which normally enter central banks' Taylor rules for interest-rate-setting, can provide evidence of out-of-sample predictability and forecasting ability for the United States Dollar/Euro exchange rate from the inception of the Euro in 1999 to the end of 2007. We also present less formal evidence that, with real-time data, the Taylor rule provides a better description of ECB than of Fed policy during this period. The strongest evidence is found for specifications that neither incorporate interest rate smoothing nor include the real exchange rate in the forecasting regression, and the results are robust to whether or not the coefficients on inflation and the real economic activity measure are constrained to be the same for the U.S. and the Euro Area. The evidence is stronger with inflation forecasts than with inflation rates and with real-time data than with revised data. Bad news about inflation and good news about real economic activity both lead to out-of-sample predictability and forecasting ability through forecasted exchange rate appreciation.

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1. Introduction

The behavior of exchange rates between Europe and the United States, either via multiple currencies until 1999 or via the Dollar/Euro exchange rate thereafter, has been one of the most studied topics in international economics. The results of this research have been less than stellar. The inability to connect exchange rates with macroeconomic fundamentals, characterized as the "exchange rate disconnect puzzle", has produced pessimism regarding the usefulness of empirical exchange rate models and focused attention on unquantifiable speculative and psychological factors.

A major contributing factor to this exchange rate pessimism has been the inability of empirical exchange rate models, starting with the seminal paper of Meese and Rogoff (1983), to forecast nominal exchange rates out-of-sample better than a naïve no-change, or random walk, forecast. While Mark (1995) provided hope that the models would forecast better at long horizons, more recent work such as Cheung, Chinn, and Pascual (2005) concludes that no model consistently does better than a random walk.

This literature still employs the empirical exchange rate models of the 1970s used by Meese and Rogoff. A money market equilibrium equation, or LM curve, for the foreign country is subtracted from a similar equation for the domestic country, producing an equation with the interest rate differential on the left-hand-side and money supply, income, and price level differentials on the right-hand-side. Using Uncovered Interest Rate Parity (UIRP) and long-run Purchasing Power Parity (PPP), and solving expectations forward, a monetary exchange rate model is derived which can be used for out-of-sample forecasting. Alternatively, the two building-blocks of the monetary model, UIRP and PPP, can be used to derive forecasting equations.

The monetary exchange rate model does not reflect how monetary policy is currently conducted or evaluated. Starting with Taylor (1993), the interest rate reaction function known as the Taylor rule, where the nominal interest rate responds to the inflation rate, the difference between inflation and its target, the output gap, the equilibrium real interest rate, and (sometimes) the lagged interest rate and the real exchange rate, has become the dominant method for evaluating monetary policy.[1] Following Clarida, Gali, and Gertler (1998), (hereafter CGG), Taylor rules have been estimated for a number of countries and time periods.

The evolution of monetary policy evaluation from LM curves and money supply reaction functions to Taylor rules has influenced exchange rate modeling. Clarida, Gali, and Gertler (2002) and Clarida (2009) investigate the derivation and implications of a two-country optimizing model with an open economy IS curve, Phillips curve, and Taylor rule. Cavallo and Ghironi (2002) and Benigno and Benigno (2008) construct two-country optimizing agent models and analyze the path of nominal exchange rates in response to various types of shocks. Gali (2008) embeds Taylor rules in open economy dynamic stochastic general equilibrium models and traces out the effects of monetary policy shocks on nominal exchange rates. Engel and West

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[1] Asso, Kahn, and Leeson (2007) examine the intellectual history of the Taylor rule and its influence on macroeconomic research and monetary policy.

(2006), Mark (2009), and Engel, Mark, and West (2007) examine the empirical performance of Taylor-rule based exchange rate models.

A major focus of Taylor rule estimation, pioneered by Orphanides (2001), is the use of real-time data that reflects the information available to central banks when they make their interest-rate-setting decisions. Interest rate reaction functions using real-time data have been estimated by Orphanides (2003, 2004) and Rudebusch (2006) for the United States, Nelson (2003) for the United Kingdom, Clausen and Meier (2005) and Gerberding, Worms, and Seitz (2005) for Germany, and Sauer and Sturm (2007), Gerdesmeier and Roffia (2004), Gorter, Jacobs, and de Haan (2008), and Sturm and Wollmershauser (2008) for the Euro Area.

Although the argument for using real-time data seems at least as compelling for exchange rate forecasting as for Taylor rule modeling, virtually all existent literature on exchange rate predictability uses fully revised data to assess the out-of-sample performance of empirical exchange rate models. The first, and until recently only, paper to use real-time data to evaluate nominal exchange rate predictability is Faust, Rogers and Wright (2003). They examine the predictive ability of Mark's (1995) monetary model using real-time data for Japan, Germany, Switzerland and Canada vis-à-vis the U.S. dollar and conclude that, while the models consistently perform better using real-time data than fully revised data, they do not perform better than the random walk model.

Molodtsova and Papell (2009), exploiting recent econometric work by Clark and West (2006), test the out-of-sample predictability of nominal exchange rate changes using Taylor rule fundamentals for 12 countries from 1973 to 2006. While real-time data is not available during the post-Bretton Woods period for most of the countries, they construct output gaps as deviations from "quasi-revised" trends in potential output, where the trends, while incorporating data revisions, are updated each period so as not to incorporate *ex post* data. Although they find strong evidence of short-run predictability with quasi-revised data for most of the considered currencies using Taylor rule fundamentals, they do not produce forecasts with real-time data.[2]

In Molodtsova, Nikolsko-Rzhevskyy, and Papell (2008), we integrate research on monetary policy evaluation and out-of-sample exchange rate predictability with real-time data. We estimate Taylor rule interest rate reaction functions with real-time data for the United States and Germany from 1979, the beginning of the European Monetary System (EMS), through 1998, the advent of the Euro, and use these specifications as fundamentals for evaluating out-of-sample predictability of the United States Dollar/Deutsche Mark nominal exchange rate. We find that evidence of predictability increases with the use of real-time, rather than revised, data and with models that allow differential inflation and output coefficients in the Federal Reserve and Bundesbank reaction functions and include the exchange rate in the Bundesbank reaction function.

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[2] Engel, Mark and West (2007) use a more constrained version of the Molodtsova and Papell (2009) specification with fully revised data. They find less evidence of short-horizon predictability, but more evidence of long-horizon predictability, than Molodtsova and Papell.

This paper uses real-time data to evaluate out-of-sample forecasting and predictability of the United States Dollar/Euro exchange rate from the inception of the Euro in 1999 to the end of 2007. The end of the sample was originally chosen for data availability when we wrote the first draft of the paper. We did not update the date through 2008 because of the zero bound on the nominal interest rate. Once the Federal Funds rate approaches zero, it cannot be lowered further and future interest rate setting cannot be predicted by the Taylor rule. We first ask whether Taylor rules appear to be a reasonable approximation of interest rate setting for the United States and the Euro area during this period. Since estimation of Taylor rules with (at most) eight years of data did not seem compelling, we start with visual evidence from a standard Taylor rule specification, similar to that presented by Taylor (1993). We find that simple Taylor rules generally track the direction of interest rate movements for both the Federal Reserve System (Fed) and the European Central Bank (ECB), although the fit is not nearly as close as in Taylor (1993). In particular, the shortfall of the Federal Funds Rate below the Taylor rule rate for the United States for 2002 to 2006, emphasized by Taylor (2007) as a cause of the housing price bubble, is also evident with real-time data.[3]

Having established that Taylor rules provide, at the least, some information that is useful for understanding Fed and ECB monetary policy, we proceed to see if they are useful for out-of-sample exchange rate forecasting and predictability. At the onset, we need to make clear the distinction between forecasting and predictability.[4] If we were evaluating forecasts from two non-nested models, we could compare the mean squared prediction errors (MSPE) from the two models, scaled to produce the tests of Diebold and Mariano (1995) and West (1996) (DMW statistic), and determine whether one model forecasts better than the other. In our case, the null hypothesis is a random walk and all alternative models are nested, resulting in severely undersized tests with standard normal critical values. In order to evaluate forecasting ability, whether the MSPE from the alternative model is significantly smaller than the MSPE from the null model, we use critical values from McCracken (2007). To evaluate predictability, whether the vector of coefficients on the Taylor rule fundamentals is jointly significantly different from zero in a regression with the change in the exchange rate on the left-hand-side, we use the Clark and West (2006) adjustment of the DMW statistic.

The starting point for our analysis is the same as for the Taylor rule model of exchange rate determination, the Taylor rule for the Euro Area is subtracted from the Taylor rule for the United States. There are a number of different specifications that we consider. While each specification has the interest rate differential on the left-hand-side, there are a number of possibilities for the right-hand-side variables.

1. Taylor posited that the Fed sets the nominal interest rate based on the current inflation rate, the inflation gap - the difference between inflation and the target inflation rate, the output gap - the difference between GDP and potential GDP, and the equilibrium real interest rate. Assuming that the ECB follows a

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[3] Poole (2007) presents a similar picture for the U.S. with real-time data from 1987:09 to 2006:06. Taylor (2009) depicts residuals from a Eurozone policy rule with the same coefficients, but does not use real-time data.

[4] This distinction has been emphasized by Inoue and Kilian (2004) and Rogoff and Stavrakeva (2008).

similar rule, we construct a *symmetric* model with inflation and the output gap on the right-hand-side. Following the results in CGG for Germany, we can also posit that the ECB includes the difference between the exchange rate and the target exchange rate, defined by PPP, in its Taylor rule and construct an *asymmetric* model where the real exchange rate is also included.

2. It has become common practice, following CGG, to posit that the interest rate only partially adjusts to its target within the period. In this case, we construct a model with *smoothing* so that lagged interest rates appear on the right-hand-side. Alternatively, we can derive a model with *no smoothing* that does not include lagged interest rates. Models with and without smoothing can be symmetric or asymmetric.

3. If the Fed and ECB respond identically to changes in inflation and the output gap, so that the coefficients in their Taylor rules are equal, we derive a *homogeneous* model where relative (domestic minus foreign) inflation and the relative output gap are on the right-hand-side. If the response coefficients are not equal, a *heterogeneous* model is constructed where the domestic and foreign variables appear separately. The homogeneous and heterogeneous models can be either symmetric or asymmetric, with or without smoothing.[5]

The models we have specified all have the interest rate differential on the left-hand-side. If UIRP held with rational expectations, an increase in the interest rate would cause an immediate appreciation of the exchange rate followed by forecasted (and actual) depreciation in accord with Dornbusch's (1976) overshooting model. In that case, we could derive an exchange rate forecasting equation by replacing the interest rate differential with the expected rate of depreciation and use the variables from the two countries' Taylor rules to forecast exchange rate changes, so that an increase in inflation and/or the output gap would produce a forecast of exchange rate depreciation. There is overwhelming evidence, however, that UIRP does not hold in the short run, as summarized by empirical research on the forward premium puzzle, as in Chinn (2006), and the delayed overshooting puzzle, as in Eichenbaum and Evans (1995). We assume that investors use this evidence for forecasting, so that an increase in inflation and/or the output gap will increase the country's interest rate, cause immediate exchange rate appreciation, and produce a forecast of further exchange rate appreciation.

Using real-time data with Taylor rule fundamentals, we find very strong evidence of out-of-sample predictability and forecasting ability for the Dollar/Euro exchange rate. The strongest evidence comes from symmetric specifications without interest rate smoothing. The results are robust to whether the specification is homogeneous or heterogeneous and to whether the output gap is constructed by Hodrick-Prescott (HP) filtering, taken from OECD estimates, or proxied by the difference between the unemployment rate and the natural rate of unemployment. For these specifications, the evidence of forecasting ability is nearly as strong

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[5] If, in addition to having the same inflation response and interest rate smoothing coefficients, the two central banks have identical target inflation rates and equilibrium real interest rates, there is *no constant* on the right-hand-side. Otherwise, there is a *constant*. Since the restrictions necessary to eliminate the constant seem very unlikely to be fulfilled, we only estimate models with a constant.

as the evidence of predictability. Asymmetric specifications which include the real exchange rate in the forecasting regression and/or specifications with smoothing provide less evidence of predictability and forecasting ability than symmetric specifications without smoothing.

Does our evidence for the Dollar/Euro exchange rate come from Taylor rule fundamentals, or is it driven by either inflation or the output gap, but not both? In order to answer this question, we estimate forecasting regressions where inflation, the HP filtered output gap, the OECD measured output gap, and the unemployment rate enter separately. The results are similar to those where inflation and a measure of economic activity enter jointly, indicating that both components of Taylor rule fundamentals are important for out-of-sample predictability and forecasting.

We also investigate "predictability" and "forecasting ability" with revised inflation and real economic activity data, recognizing that we are no longer replicating the environment experienced by market participants. In contrast to many applications of real-time data, we expect that, because the exchange rate is an asset price, predictability and forecasting ability can decrease with revised data because information is being used that was unavailable both when the forecasts were made and when the forecasted exchange rate was realized. We find that both predictability and forecasting ability decrease when revised data is in the forecasting regression. The decreases are more dramatic with the OECD measured output gap, which is consistently larger than the real-time OECD measured output gap, than with the revised HP filtered output gap or the unemployment rate, neither of which are as substantially different from their real-time counterparts.

When estimating Taylor rules with real-time data, it is common practice to use forecasted, rather than realized, values of inflation and real economic activity.[6] We first use one-year-ahead forecasted inflation and realized values of economic activity, and find that both predictability and forecasting ability increase compared to the results with realized inflation. We then use forecasted inflation with either the forecasted OECD output gap or the forecasted unemployment rate. The results for both predictability and forecasting ability are even stronger than with forecasted inflation and realized values of economic activity and constitute the strongest results in the paper.

Rogoff and Stavrakeva (2008) have criticized recent studies that employ the CW statistic, including Engel, Mark, and West (2007), Gourinchas and Rey (2007), and Molodtsova and Papell (2009), because it does not necessarily satisfy their first criterion for a "good" forecast – a forecast with a MSPE smaller than the MSPE of a driftless random walk.[7] In our terminology, the CW statistic is a test for predictability, not forecasting ability. The results in this paper, however, apply to both. For every symmetric specification

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[6] Orphanides and Wieland (2008), for example, find that the fit of Taylor rule models is improved by using inflation and unemployment forecasts.

[7] Rogoff and Stavrakeva's (2008) second criterion for a good forecast involves robustness of the out-of-sample test statistics over different forecast windows. Given the short span of data since the advent of the Euro, we cannot address this criterion.

without interest rate smoothing using real-time data, the MSPE of the model with Taylor rule fundamentals is smaller than the MSPE of the random walk model and the random walk null can be rejected using McCracken's (2007) critical values for the DMW statistic.

It is often asked whether the experience of the Bundesbank during the EMS period provides a good predictor for the actions of the ECB. The answer from this paper is clearly no. In our earlier work on the Mark/Dollar exchange rate with real-time data, we found evidence of predictability only with heterogeneous coefficients and asymmetric specifications, with or without smoothing. For the Euro/Dollar rate, we find that the evidence of predictability and forecasting ability is much stronger with symmetric specifications without smoothing, and doesn't depend on whether the coefficients are homogeneous or heterogeneous.

Clarida and Waldman (2008), using an event study methodology, find evidence that a surprise increase in inflation causes the exchange rate to appreciate in the very short run for inflation targeting countries. We find strong support for our variant of their proposition that "bad news about inflation is good news for the exchange rate", which we characterize as "bad news about inflation is good news for the forecasted exchange rate." Using the most successful homogeneous and symmetric specification, an increase in U.S. inflation relative to Euro Area inflation causes forecasted dollar appreciation whatever measure of real economic activity is included in the forecasting regression. We also find that "good news about output or unemployment is good news for the forecasted exchange rate." An increase in the U.S. output gap relative to the Euro Area output gap causes forecasted dollar appreciation while an increase in U.S. unemployment relative to Euro Area unemployment causes forecasted dollar depreciation.

2. Taylor Rule Fundamentals

We examine the linkage between the exchange rate and a set of fundamentals that arise when central banks set the interest rate according to the Taylor rule. Following Taylor (1993), the monetary policy rule postulated to be followed by central banks can be specified as

$$i_t^* = \pi_t + \phi(\pi_t - \pi^*) + \gamma y_t + r^* \tag{1}$$

where i_t^* is the target for the short-term nominal interest rate, π_t is the inflation rate, π^* is the target level of inflation, y_t is the output gap, or percent deviation of actual real GDP from an estimate of its potential level, and r^* is the equilibrium level of the real interest rate. It is assumed that the target for the short-term nominal interest rate is achieved within the period so there is no distinction between the actual and target nominal interest rate. Alternatively, as in Blinder and Reis (2005), the difference between the natural rate of unemployment and the unemployment rate can replace the output gap.[8]

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[8] While we do not explicitly incorporate time-varying inflation and/or equilibrium real interest rates, the use of rolling regressions allows for changes in the constant.

According to the Taylor rule, the central bank raises the target for the short-term nominal interest rate if inflation rises above its desired level and/or output is above potential output. The target level of the output deviation from its natural rate y_t is 0 because, according to the natural rate hypothesis, output cannot permanently exceed potential output. The target level of inflation is positive because it is generally believed that deflation is much worse for an economy than low inflation. Taylor assumed that the output and inflation gaps enter the central bank's reaction function with equal weights of 0.5 and that the equilibrium level of the real interest rate and the inflation target were both equal to 2 percent.

The parameters π_t^* and r^* in equation (1) can be combined into one constant term $\mu = r^* - \phi\pi^*$, which leads to the following equation,

$$i_t^* = \mu + \lambda\pi_t + \gamma y_t \tag{2}$$

where $\lambda = 1 + \phi$. Because $\lambda > 1$, the real interest rate is increased when inflation rises and so the Taylor principle is satisfied.

While it seems reasonable to postulate a Taylor rule for the United States that includes only inflation and the output gap, it is common practice to include the real exchange rate in specifications for other countries,

$$i_t^* = \mu + \lambda\pi_t + \gamma y_t + \delta q_t \tag{3}$$

where q_t is the real exchange rate for the Euro Area. The rationale for including the real exchange rate in the Taylor rule is that the central bank sets the target level of the exchange rate to make PPP hold and increases (decreases) the nominal interest rate if the exchange rate depreciates (appreciates) from its PPP value. Based on the evidence in CGG and Molodtsova, Nikolsko-Rzhevskyy, and Papell (2008) that the real exchange rate entered the Taylor rule for the Bundesbank during the European Monetary System period, we allow for the possibility that it should be included in the ECB's Taylor rule.

It has also become common practice to specify a variant of the Taylor rule which allows for the possibility that the interest rate adjusts gradually to achieve its target level. Following CGG, we assume that the actual observable interest rate i_t partially adjusts to the target as follows:

$$i_t = (1-\rho)i_t^* + \rho i_{t-1} + v_t \tag{4}$$

Substituting (3) into (4) gives the following equation,

$$i_t = (1-\rho)(\mu + \lambda\pi_t + \gamma y_t + \delta q_t) + \rho i_{t-1} + v_t \tag{5}$$

where $\delta = 0$ for the United States.

To derive the Taylor-rule-based forecasting equation, we construct the interest rate differential by subtracting the interest rate reaction function for the Euro Area from that for the U.S.:

$$i_t - \tilde{i}_t = \alpha + \alpha_{u\pi}\pi_t - \alpha_{e\pi}\tilde{\pi}_t + \alpha_{uy}y_t - \alpha_{ey}\tilde{y}_t - \alpha_q\tilde{q}_t + \rho_u i_{t-1} - \rho_e\tilde{i}_{t-1} + \eta_t \tag{6}$$

where ~ denotes Euro Area variables, subscripts u and e denote coefficients for the United States and the Euro Area, α is a constant, $\alpha_\pi = \lambda(1-\rho)$ and $\alpha_y = \gamma(1-\rho)$ for both central banks, and $\alpha_q = \delta(1-\rho)$ for the ECB.[9]

Suppose that U.S. inflation rises above target. If there is no smoothing, all interest rate adjustments are immediate. The Fed will raise the interest rate by $\lambda\Delta\pi$, where $\Delta\pi$ is the change in the inflation rate. If there is smoothing, the adjustment is gradual. The Fed will raise the interest rate by $(1-\rho)\lambda\Delta\pi$ in the first period. In the second period, the interest rate will be $(1-\rho^2)\lambda\Delta\pi$ above its original level, followed by $(1-\rho^3)\lambda\Delta\pi$, and so on. The maximum impact on the interest rate will be approximately $\lambda\Delta\pi$, the same as with no smoothing. Engel and West (2006) and Clarida and Waldman (2008) show that, assuming the Taylor principle is satisfied, a surprise increase in U.S. inflation will appreciate the exchange rate.

How will the increase in the interest rate differential affect exchange rate forecasts? According to the results in Eichenbaum and Evans (1995), a positive shock to the federal funds rate leads to delayed overshooting, both immediate and sustained exchange rate appreciation, with the maximal impact occurring between 22 and 33 months after the shock, followed by depreciation. This is clearly inconsistent with UIRP with rational expectations, which constrains the maximal impact to occur immediately. Faust and Rogers (2003) and Scholl and Uhlig (2008), while questioning their identification procedure, also find large UIRP deviations in response to monetary policy shocks. Assuming that investors use these empirical results for forecasting, we postulate that an increase in U.S. inflation above target will cause forecasted dollar appreciation in the short run. A similar argument would imply that an increase in Euro Area inflation above its target would lead to forecasted euro appreciation.[10]

The link between higher inflation and exchange rate appreciation potentially characterizes any country where the central bank uses the interest rate as the instrument in an inflation-targeting policy rule. In the context of the Taylor rule, three additional predictions can be made. First, if the U.S. output gap increases, the Fed will raise interest rates and cause the dollar to appreciate. Similarly, an increase in the Euro Area

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[9] As shown by Engel and West (2005), this specification would still be applicable if the U.S. had an exchange rate target in its interest rate reaction function.

[10] Cavallo and Ghironi (2002) show how Taylor-rule type monetary policy and asset dynamics can generate delayed overshooting in response to (relative) productivity shocks. Gourinchas and Tornell (2004) show that an increase in the interest rate can cause sustained exchange rate appreciation if investors systematically underestimate the persistence of interest rate shocks. Molodtsova and Papell (2009) use their results to provide theoretical link between higher inflation and forecasted exchange rate appreciation.

output gap will cause the euro to appreciate. Second, if the real exchange rate for the Euro Area depreciates and it is included in the ECB's Taylor rule, the ECB will raise its interest rate, causing the Euro to appreciate and the dollar to depreciate. Third, if there is interest rate smoothing, a higher lagged interest rate will increase current and expected future interest rates. Under UIRP and rational expectations, any event that causes the Fed or ECB to raise its interest rate will produce immediate exchange rate appreciation and forecasted depreciation. Based on the evidence discussed above, however, we believe it is more reasonable to postulate that these events will produce both immediate and forecasted appreciation.

These predictions can be combined with (6) to produce an exchange rate forecasting equation:

$$\Delta s_{t+1} = \omega - \omega_{u\pi}\pi_t + \omega_{e\pi}\tilde{\pi}_t - \omega_{uy}y_t + \omega_{ey}\tilde{y}_t + \omega_q\tilde{q}_t - \omega_{ui}i_{t-1} + \omega_{ei}\tilde{i}_{t-1} + \eta_t \tag{7}$$

The variable s_t is the log of the U.S. dollar nominal exchange rate determined as the domestic price of foreign currency, so that an increase in s_t is a depreciation of the dollar. The reversal of the signs of the coefficients between (6) and (7) reflects the presumption that anything that causes the Fed and/or ECB to raise the U.S. interest rate relative to the Euro Area interest rate will cause the dollar to appreciate (a decrease in s_t). Since we do not know by how much a change in the interest rate differential (actual or forecasted) will cause the exchange rate to adjust, we do not have a link between the magnitudes of the coefficients in (6) and (7).

A number of different models can be nested in Equation (7). If the ECB doesn't target the exchange rate $\delta = \omega_q = 0$ and we call the specification symmetric. Otherwise, it is asymmetric. If the interest rate adjusts to its target level within the period $\omega_{ui} = \omega_{ei} = 0$ and the model is specified with no smoothing. Alternatively, there is smoothing. If the coefficients on inflation, the output gap, and interest rate smoothing are the same in the U.S. and the Euro Area, so that $\omega_{u\pi} = \omega_{e\pi}$, $\omega_{uy} = \omega_{ey}$, and $\omega_{ui} = \omega_{ei}$, inflation, output gap, and lagged interest rate differentials are on the right-hand-side of Equation (7) and we call the model homogeneous. Otherwise, it is heterogeneous.

3. Taylor Rules, the Fed, and the ECB

If we were writing this paper in 2019 instead of 2009, we would start by estimating Taylor rules using real-time data for the Fed and ECB to provide a guide to the factors that might affect out-of-sample exchange rate predictability. Since that option is precluded and we are skeptical that much can be learned from estimating Taylor rules with eight years of data, we start with a more descriptive method. We first describe the real-time data available for the U.S. and Euro Area since 1999, and then provide visual evidence that Taylor rules with real-time data provide a useful characterization of interest rate setting by the Fed and ECB.[11]

3.1 Real-Time and Revised Data

We use real-time quarterly data from 1999:Q4 to 2007:Q4 for the United States and the Euro Area. The data is from the OECD Original Release and Revisions Database.[12] It has a triangular format with the vintage date on the horizontal axis and calendar dates on the vertical. The term vintage denotes the date in which a time series of data becomes known to the public.[13] For each subsequent quarter, the new vintage incorporates revisions to the historical data, thus providing all information known at the time. The revised data is constructed from the 2007:Q4 vintage in the real-time dataset for both countries.

For each forecasting regression, we use 34 quarters to estimate the historical relationship between the Taylor rule fundamentals and the change in the exchange rate, and then use the estimated coefficients to forecast the exchange rate one-quarter-ahead. The data for the first vintage starts in 1991:Q1. We use rolling regressions to predict 32 exchange rate changes from 2000:Q1 to 2007:Q4. Since we use vintage data, the estimated coefficients are based on revised data, but the forecasts are conducted using real-time data.[14]

We use the GDP Deflator to measure inflation for the U.S. and the Harmonized Index of Consumer Prices (HICP) to measure inflation for Euro Area. Following Taylor (1993), the inflation rate is the rate of inflation over the previous four quarters. We use two different measures of the output gap. First, we construct quarterly measures of the output gap from internal OECD estimates. This data comes from the semi-annual issues of OECD Economic Outlook. Each issue contains past estimates as well as future forecasts of annual values of the output gap for OECD countries including the Euro Area. Since both estimates and forecasts are annual, we used quadratic interpolation to obtain quarterly estimates.[15] The second measure of the output gap uses HP detrended real industrial production.[16] While applying the HP filter, we take account of the end-of-sample problem by forecasting and backcasting the industrial production series by 12 quarters in both directions assuming that growth rates follow an AR(4) process. The unemployment rates are from the OECD real-time database.

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[11] Recent papers that compare Taylor rules for the Fed and ECB, including De Lucia and Lucas (2007) and Gerdesmeier, Mongelli, and Roffia (2007), use either Bundesbank or synthetic Euro Area data to extend the sample back to 1993, and therefore cannot use real-time data.

[12] An alternative would be to use Euro Area Business Cycle Network data, but it does not start until 2001.

[13] There is typically a one-quarter lag before data is released, so real-time variables dated time t actually represent data through period $t-1$.

[14] An alternative method of constructing real-time data is to use "diagonal" data that does not incorporate historical revisions. With that method, the estimated coefficients would also use real-time data. Since the vintages are not available before 1999 and we only have 32 forecast periods, we do not have that option for this paper.

[15] Since the data is updated semi-annually, we assume that, in the quarter following the period in which the estimates are released, the public uses the estimates and forecasts from the previous quarter. We interpolate between the current and immediate past release of the Economic Outlook.

[16] The industrial production data starts in 1990:Q1. We use industrial production instead of GDP because the latter does not start until 1995 for the Euro Area in the OECD database.

The forward-looking specifications for the U.S. also use the Philadelphia Fed Survey of Professional Forecasters (SPF) data, which consists of annualized quarter-over-quarter GDP deflator inflation and unemployment forecasts at different horizons. We convert them into year-over-year rates by taking the average of four consecutive forecasts. The data is available for the entire sample. For the Euro Area, the ECB publishes Euro Area SPF forecasts for the one-year-ahead HICP inflation rate. The first round of the survey was conducted in 1999:Q1. This means that we do not have the same forecast for 1991:Q1, which is the starting point for our "vintage" regressions. To deal with this issue, we note that the first "vintage" regression which the public could have run using OECD real-time data was in 1999:Q4 when the first OECD vintage was published. At that time, inflation data for 1990:Q1-1999:Q3 was available. To construct the t+4 inflation forecast for any vintage, we use the realized t+4 values of inflation until 1998:Q4 and real-time Euro Area SPF forecasts from 1999:Q1 to 2007:Q4. The data for t+4 SPF forecasts of unemployment for Euro Are is constructed by the same method.[17]

The exchange rate, defined as the quarterly-averaged US dollar price of a Euro, and the short-term nominal interest rates, defined as the interest rate in the third month of each quarter, are taken from OECD Main Economic Indicators (MEI) database. The short-term interest rate is the money market rate (EONIA) for Euro Area and the Federal Funds Rate for the U.S. Since interest data for the Euro Area does not exist prior to 1994:Q4, we use the German money market rate from the IMF International Financial Statistics Database (line 60B) for the earlier period. The real Euro/USD exchange rate is calculated as the percentage deviation of the nominal exchange rate from the target defined by Purchasing Power Parity, where the two countries' price levels are measured by the CPI for the U.S. and the HICP for the Euro Area.

The real-time and revised data are depicted in Figure 1. In line with all research in this area since Orphanides (2001), the differences are much larger for output gaps than for inflation, reflecting the changes in potential output, as well as the data revisions themselves, for the former but not the latter. Revisions in unemployment are much smaller than revisions in the output gaps for both the U.S. and the Euro Area. For the U.S., the revisions in the HP filtered output gap are larger than the revisions in the OECD estimated output gap, while the opposite is true for the Euro Area. The largest revisions are for the OECD real-time estimates of the output gap for the Euro Area, which are substantially below the revised estimates from 1999:Q4 to 2004:Q2.

These points are illustrated in Table 1 that provides summary statistics of revised and real-time data. The average U.S. real-time inflation and unemployment rate are virtually the same as the revised inflation and unemployment rate and the average Euro Area real-time inflation and unemployment rate differ from their revised counterparts insignificantly. The largest differences are found between the average U.S. real-time and revised HP filtered output gap and Euro Area OECD output gap. These differences are very close in size and

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[17] Since the first forecasts are conducted in 1999:Q4, they are real-time forecasts even though realized t+4 values of inflation and unemployment are used through 1998:Q4.

equal to 1.14 percentage points for the U.S. and 1.13 percentage points for Euro Area. The average U.S. real-time and revised OECD output gap and Euro Area real-time and revised HP filtered output gap differ by 0.05 and 0.11 percentage points, respectively. These differences suggest that policy recommendations based on HP filtered output gap for the U.S. and OECD estimates of the output gap for the Euro Area may be substantially different with revised and real-time data.

Table 2 shows the descriptive statistics on data revisions in our sample. A positive and significant value for the mean of the revision indicates that the variable was on average revised upwards, so that the existence of measurement errors or the availability of new information (or both) made the statistical agency realize that the inflation rate and/or the output gap was higher than perceived in real-time. We can see that the mean revision for inflation and unemployment is essentially zero for the U.S., and insignificant for Euro Area. Both HP filtered output gap for the U.S. and OECD estimates of the output gap for the Euro Area are on average revised upwards.

To explore the nature of data revisions in our sample, we examine the correlations between the data revisions, defined as $X_{revised}$-$X_{real\text{-}time}$, and the real-time and revised series. According to Mankiw and Shapiro (1986), if data revisions represent pure noise, they should be uncorrelated with the revised data but correlated with the real-time series. The opposite should be true if data revisions represent pure news. The correlations in Table 2 indicate that revisions in the Euro Area HP filtered output gap represent pure news and revisions in Euro Area OECD output gap are dominated by news. The revisions in Euro Area inflation and unemployment represent mostly noise. The properties of news are more pronounced in the U.S. revisions of inflation and the HP filtered output gap, while revisions of the U.S. OECD output gap and unemployment are dominated by noise.

3.2 Taylor Rules for the Fed and ECB

We provide visual evidence of how closely interest rate setting by the Fed and the ECB can be characterized by a Taylor rule with real-time data. In panel A of Figure 2, we depict the actual U.S. and Euro Area interest rate and the counterfactual interest rate implied by a Taylor rule with a coefficient of 1.5 on inflation, 0.5 on the output gap, an inflation target of 2 percent, an equilibrium real interest rate of 2 percent, and no smoothing. Except for using real-time rather than revised data and a different time period, this is exactly the exercise conducted in Taylor (1993). We use GDP real-time inflation for the US, HICP real-time inflation for the Euro Area, and OECD estimates of the output gap.[18]

The results for the U.S. show that, while the Federal Funds rate and interest rate implied by the Taylor rule are clearly positively correlated, the Federal Funds rate is consistently below the rate implied by the Taylor rule from 2002:Q4 to 2007:Q1, nearly exactly replicating the results reported by Taylor (2007) with

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[18] Figures for HP filtered data (not reported) are similar.

revised data.[19] Taylor (2007) argues that the gap between the actual Federal Funds rate and the rate implied by the Taylor rule was an important contributing factor to the housing price bubble in the U.S. Our results show that, in the context of this argument, the discrepancy is not an artifact of using revised data that were not available to the Fed at the time that interest-rate-setting decisions were made, but also appears in real-time data. For the Euro Area, while the overall fit is closer, the actual Money Market Rate is below the rate implied by the Taylor rule from 2003:Q1 to 2007:Q1. While this is similar to the pattern found for the U.S., the magnitude of the gap is much smaller for the Euro Area than for the U.S.

It is often argued that monetary policy evaluation should be conducted with forward-looking data. In panel B of Figure 2, we depict forward-looking specifications, for which we use the t+4 SPF inflation forecasts for both the Euro Area and the U.S. Everything else, including the coefficients on inflation and the output gap, the inflation target of 2 percent, and the equilibrium real interest rate of 2 percent, is the same as with contemporaneous inflation. For the U.S., the pattern is similar to that found with contemporaneous inflation except that it starts in 2002:Q3 and ends in 2005:Q4. In addition, the gap between the actual Federal Funds rate and the rate implied by the Taylor rule is smaller with forecasted inflation. For the Euro Area, the actual Money Market rate with forecasted inflation is very close to the rate implied by the Taylor rule for almost the entire period, with the actual rate higher in 2000-2001 and the implied rate higher in 2004-2006.

With forecasted inflation, we can construct *ex ante* real interest rates as the nominal interest rate minus the expected rate of inflation, and calculate the equilibrium real interest rate, 1.45 percent for the U.S. and 1.33 percent for the Euro Area, as the average real interest rate over the period.[20] The results with a forward-looking specification and calculated equilibrium real interest rate are shown in panel C. According to Equation (1), the equilibrium real interest rate has a point-for-point affect on the nominal interest rate, so this lowers the interest rate implied by the Taylor rule by 0.55 percent for the U.S. and 0.67 percent for the Euro Area. For the U.S., the gap between the actual Federal Funds rate and the rate implied by the Taylor rule starts in 2002:Q3 and ends in 2005:Q4 and is smaller than that found with an equilibrium real interest rate of 2 percent. For the Euro Area, the actual Money Market rate with forecasted inflation is above the rate implied by the Taylor rule for most of the period, and they are very close from 2003:Q2 to 2006:Q3.

Using real-time data visual methods that make no attempt to produce a good fit between the actual and implied interest rates, we have shown that the Taylor rule provides a good approximation of interest rate setting by both the Fed and the ECB since 1999. For the U.S., the major deviation occurred between 2003 and 2006 and, as described by Taylor (2007), was produced because the actual Federal Funds rate was consistently below the rate implied by the Taylor rule. For the ECB, the differences between the actual and

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[19] In Taylor (2007), the actual and implied paths diverge in 2002:Q2 and merge again in 2006:Q3.

[20] Because the equilibrium real interest rate is calculated by averaging over the period, this is not literally a real-time exercise even though it uses real-time inflation and output gap data.

implied interest rates are smaller, and the actual Money Market rate was neither consistently above nor below the rate implied by the Taylor rule.

4. Forecast Comparison Based on MSPE

We have provided some informal evidence that, using real-time data, interest rate setting by the Fed and the ECB is broadly consistent with Taylor rule. We now turn to the central question of our paper, whether Taylor rule fundamentals can provide evidence of out-of-sample forecasting and/or predictability for the United States Dollar/Euro exchange rate. Before addressing this issue, we need to summarize some econometric results.

We are interested in comparing the mean square prediction errors from two nested models. The benchmark model is a zero mean martingale difference process, while the alternative is a linear model.

Model 1: $y_t = \varepsilon_t$

Model 2: $y_t = X_t^{'}\beta + \varepsilon_t$, where $E_{t+1}(\varepsilon_t) = 0$

Suppose we have a sample of T+1 observations. The last P observations are used for predictions. The first prediction is made for the observation *R*+1, the next for *R*+2, …, the final for *T*+1. We have *T*+1=*R*+*P*, where *R*=34, and *P*=32 quarters. To generate prediction for period *t*=*R, R*+1*, …, T*, we use the information available prior to *t.* Let $\hat{\beta}_t$ is a regression estimate of β_t that is obtained using the data prior to t. The one-step ahead prediction for model 1 is 0, and $X_{t+1}\hat{\beta}_t$ for model 2. The sample forecast errors from the models 1 and 2 are $\hat{e}_{1,t+1} = y_{t+1}$ and $\hat{e}_{2,t+1} = y_{t+1} - X_{t+1}\hat{\beta}_t$, respectively. The corresponding MSPEs for the two models are $\hat{\sigma}_1^2 = P^{-1}\sum_{t=T-P+1}^{T} y_{t+1}^2$ and $\hat{\sigma}_2^2 = P^{-1}\sum_{t=T-P+1}^{T}(y_{t+1} - X'_{t+1}\hat{\beta}_t)^2$.

We want to test the null hypothesis that the MSPEs are equal against the alternative that the MSPE of Model 2 is smaller than the MSPE of Model 1. Under the null, the population MSPEs are equal. We need to use the sample estimates of the population MSPEs to draw the inference. The procedure introduced by Diebold and Mariano (1995) and West (1996) uses sample MSPEs to construct a t-type statistics which is assumed to be asymptotically normal. To construct the DMW statistic, let

$$\hat{f}_t = \hat{e}_{1,t}^2 - \hat{e}_{2,t}^2 \qquad \text{and} \qquad \bar{f} = P^{-1}\sum_{t=T-P+1}^{T}\hat{f}_{t+1} = \hat{\sigma}_1^2 - \hat{\sigma}_2^2$$

Then, the DMW test statistic is computed as follows,

$$DMW = \frac{\bar{f}}{\sqrt{P^{-1}\hat{V}}}, \qquad \text{where} \qquad \hat{V} = P^{-1}\sum_{t=T-P+1}^{T}(\hat{f}_{t+1} - \bar{f})^2 \tag{8}$$

McCracken (2007), among others, shows that application of the DMW statistic with standard normal critical values to non-nested models results in severely undersized tests. Clark and West (2006) demonstrate analytically that the asymptotic distributions of sample and population difference between the two MSPEs are not identical, namely the sample difference between the two MSPEs is biased downward from zero. They show that the sample difference between the two MSPEs is uncentered under the null.

$$(9)\qquad \hat{\sigma}_1^2 - \hat{\sigma}_2^2 = P^{-1}\sum_{t=T-P+1}^{T}\hat{f}_{t+1} = P^{-1}\sum_{t=T-P+1}^{T}y_{t+1}^2 - P^{-1}\sum_{t=T-P+1}^{T}(y_{t+1} - X'_{t+1}\hat{\beta}_t)^2 = 2P^{-1}\sum_{t=T-P+1}^{T}y_{t+1}X'_{t+1}\hat{\beta}_t - P^{-1}\sum_{t=T-P+1}^{T}(X'_{t+1}\hat{\beta}_t)^2$$

Under the null, the first term in (9) is zero, while the second one is greater than zero by construction. Therefore, under the null we expect the MSPE of the naïve no-change model to be smaller than that of a linear model. The intuition behind this result is the following. If the null is true, estimating the alternative model introduces noise into the forecasting process because it is trying to estimate parameters which are zero in population. Use of the noisy estimate will lead to a higher estimated MSPE and, as a result, the sample MSPE of the alternative model will be higher by the amount of estimation noise.

In order to test for predictability, we construct the corrected test statistic as described in Clark and West (2006) by adjusting the sample MSPE from the alternative model by the amount of the bias in the second term of equation (9). This adjusted CW test statistic is asymptotically standard normal. When the null is a martingale difference series Clark and West (2006, 2007) recommend adjusting the difference between MSPEs as described above and using standard normal critical values for inference.[21]

It is important to understand the distinction between predictability and forecasting ability. We use the term "predictability" as a shorthand for "out-of-sample predictability" in the sense used by Clark and West (2006, 2007), rejecting the null of a zero slope in the predictive regression in favor of the alternative of a nonzero slope. The CW methodology tests whether the regression coefficient β is zero rather than whether the sample MSPE from the model-based forecast is smaller than the sample MSPE from the random walk forecast.

How can we test for forecasting ability with nested models? While there is not a universally accepted answer to this question, we use McCracken's (2007) asymptotic critical values for the DMW test with nested models to produce correctly sized tests. The critical values depend on the ratio P/R, which is 0.94 with our data, and the number of additional estimated parameters in the unrestricted model.[22] Since the null hypothesis is a random walk, with zero estimated parameters, the number of additional estimated parameters is given in

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[21] Because the null hypothesis for the CW statistic is a zero mean martingale difference process, we can only test the null that the exchange rate is a random walk, not a random walk with drift. Clark and West (2006, 2007) argue that standard normal critical values are approximately correct and advocate using them instead of bootstrapped critical values. Clark and McCracken (2009) consider the impact of data revisions on tests of equal predictive ability. Because the nominal exchange rate is unrevised and a random walk under the null, even predictable real-time data revisions do not have an impact on the asymptotic distributions and the Clark and West results can be used.

[22] The critical values are tabulated for P/R = 0.8 and 1.0, and we construct the values for P/R = 0.94 by interpolation.

Equation (7), ranging from three for the symmetric model with homogeneous coefficients and no smoothing to eight for the asymmetric model with heterogeneous coefficients and smoothing.[23]

One disquieting aspect of both tests is that it is possible to find evidence of predictability and/or forecasting ability when the MSPE of the random walk forecast is smaller than the MSPE of the linear model forecast. The issue arises because, whether good size is achieved by adjusting the DMW statistic, as in Clark and West (2006), or by adjusting the critical values, as in McCracken (2007), the distribution of the critical values is not centered around the point where the two MSPEs are equal. This is not problematic in the context of testing for predictability, which is a test of whether the regression coefficient β is significantly different from zero. It is problematic, however, in testing for forecasting ability, which is a test of whether the MSPE from the model is smaller than the MSPE from the random walk. In general, using the critical values in McCracken (2007), specifications with larger P/R and larger number of additional estimated parameters are more likely to produce rejections of the random walk null when the MSPE from the random walk is smaller than the MSPE from the model. In the context of our specifications, with P/R equal to 0.94, it is not possible to reject the random walk when the MSPE from the random walk is smaller than the MSPE from the model at the 10 percent significance level or higher with three additional estimated parameters. With four-to-seven additional parameters, it is possible to reject the random walk at the 10 percent, but not the 5 percent, significance level and, with eight additional parameters, it is possible to reject the random walk at the 5 percent level.

5. Empirical Results

5.1 Taylor Rule Fundamentals

We now turn to our empirical results. Tables 3-6 present results for one-quarter-ahead forecast comparisons. The first row of each panel reports the ratios of the out-of-sample MSPE from the linear model with Taylor rule fundamentals to that of the random walk model. The second row reports p-values that test predictability - can the null hypothesis that the true model is a random walk be rejected in favor of the alternative of the model with Taylor rule fundamentals, using the CW test with asymptotic critical values.

The third row reports p-values that test forecasting ability – can the null hypothesis that the random walk model has the same MSPE as the Taylor rule model be rejected in favor of the alternative hypothesis that the MSPE of the Taylor rule model is smaller than the MSPE of the random walk model, using the DMW test with McCracken's (2007) critical values.

Table 3 presents the central results of the paper. With a symmetric specification that does not include the real exchange rate in the forecasting regression and no interest rate smoothing, the MSPE of the Taylor

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[23] Since we have a relatively small number of observations, we also calculated bootstrapped critical values for the CW test, which made virtually no difference in the results and are not reported. We did not calculate bootstrapped critical values for the DMW test because, as described by McCracken (2007), the asymptotic distribution is not well-approximated by a standard normal distribution and so bootstrapping cannot be used to correct finite sample size distortions relative to an asymptotic standard normal distribution.

rule model is smaller than the MSPE of the random walk model, so that the ratio is less than unity, for all six cases. We use the CW statistic to test for predictability. With homogeneous coefficients, the random walk (no predictability) null hypothesis is rejected in favor of the alternative hypothesis of out-of-sample predictability for the Euro/Dollar exchange rate with Taylor rule fundamentals at the 1 percent level when inflation and either the HP filtered output gap or the unemployment rate, and at the 5 percent level when inflation and the OECD estimated output gap, is included in the forecasting regression. With heterogeneous coefficients, the null hypothesis is rejected at the 5 percent level for inflation and all three real economic activity measures.

The results in Table 3 also provide strong evidence of forecasting ability. Using the DMW test, the random walk (equal MSPE) null hypothesis is rejected in favor of the alternative hypothesis of out-of-sample forecasting ability of the symmetric model with homogeneous coefficients and no smoothing at the 1 percent level with inflation and the HP filtered output gap, the 5 percent level with inflation and the unemployment rate, and at the 10 percent level with inflation and the OECD output gap. For the symmetric model with homogeneous coefficients and no smoothing, the null hypothesis is rejected at the 5 percent level for all three real economic activity measures.

We illustrate these results in Figure 3, which depicts actual and predicted Dollar/Euro exchange rate changes for all six specifications of the symmetric model with no smoothing. Positive values represent depreciation of the dollar while negative values represent appreciation of the dollar. Since the "zero" line represents the random walk (no change) forecast, periods for which the actual and predicted values are both either positive or negative lower the MSPE ratio, while periods for which the actual and predicted values are of opposite sign raise the MSPE ratio. The predicted exchange rate changes display considerable variability, although not as much as the actual changes. While the predicted changes do not pick up the sharp quarter-by-quarter spikes of the actual changes, they track the pattern of dollar appreciation in 2000 and 2001, depreciation in 2002 – 2004, appreciation in 2005, and depreciation in 2006 and 2007 found in the data. Given the generally accepted view that exchange rates are unconnected with macroeconomic fundamentals, the figure provides strong collaboration of the statistical evidence that Taylor rule fundamentals can forecast changes in the Dollar/Euro rate.

The results for the symmetric specifications with interest rate smoothing are not as strong. The MSPE of the Taylor rule model is smaller than the MSPE of the random walk model, so that the ratio is less than unity, for three out of six cases. While the no predictability null can be rejected for all six specifications, the significance level is 1 percent for one model, 5 percent for four models, and 10 percent for one model. The evidence of forecasting ability is weaker, 5 percent for three models and 10 percent for two models. The results for the asymmetric specification (that includes the real exchange rate in the forecasting regression) without smoothing are somewhat weaker than for the symmetric specification with smoothing. While the MSPE ratio is again less than unity for three of the six cases and the no predictability null can be rejected for all six specifications, the significance levels are 5 percent for three models and 10 percent for the three additional models, and evidence of forecasting ability can only be found for three of the six specifications. The weakest results are found for the asymmetric specification with smoothing. The MSPE ratio is greater than unity for all six cases, and evidence of either predictability or forecasting ability can be found for only one specification at the 10 percent level.

It is clear from the results in Table 3 why the DMW statistic, but not the CW statistic, satisfies Rogoff and Stavrakeva's (2008) criterion for a good forecast. With the DMW statistic, the random walk null hypothesis is never rejected in favor of the alternative of forecast ability of the Taylor rule model at the 5 percent (or higher) level when the MSPE of the Taylor rule model is larger than the MSPE of the random walk model, so that the MSPE ratio is greater than unity. There are several rejections at the 10 percent level, but these are for specifications with smoothing that contain seven or eight additional estimated parameters. With the DMW statistic, all rejections of the random walk null in favor of the alternative of predictability at the 10 percent level are for specifications where the MSPE ratio is greater than unity, all rejections at the 1 percent level are for specifications where the MSPE ratio is less than unity, and the rejections at the 5 percent level are mixed.

Rogoff and Stavrakeva (2008) advocate using DMW tests with bootstrapped critical values to evaluate forecasting ability. In order to evaluate the implications of this approach, we conducted simulations to determine rejection frequencies using McCracken's (2007) critical values with the data and specifications in Table 3, where the series being forecasted was a zero mean martingale difference process rather than the first-difference of the log exchange rate. For the symmetric model with homogeneous coefficients and no smoothing, with three additional parameters, the rejection frequencies (averaged over the three real economic activity measures) are 7.4, 3.7, and 0.8 percent for tests of nominal 10, 5, and 1 percent size. As additional parameters are added, the rejection frequencies fall substantially. This does not, however, mean that it would be preferable to use bootstrapped critical values. Since the null hypothesis for the simulations is that the series being forecasted is a random walk, not that the random walk and Taylor rule models produce forecasts with equal MSPEs, the use of bootstrapped critical values is appropriate for investigating predictability, not forecasting ability. We experimented using the same simulations to calculate bootstrapped critical values, and the number of rejections increased substantially, particularly for the models with a larger number of additional parameters. These rejections, however, come entirely from specifications where the MSPE ratio is greater than unity, which violates Rogoff and Stavrakeva's criterion for a good forecast. In the context of our particular specification, it seems preferable to use McCracken's asymptotic critical values of the DMW test to evaluate forecasting ability than to use bootstrapped critical values.

We have presented evidence that, using symmetric specifications that do not include the real exchange rate in the forecasting regression, the random walk (no predictability) null hypothesis can be consistently rejected in favor of the alternative hypothesis of out-of-sample predictability and the equal MSPE hypothesis can be consistently rejected in favor of the alternative hypothesis of out-of-sample forecast ability for the Euro/Dollar exchange rate with Taylor rule fundamentals. Since the specifications include inflation and either the HP filtered output gap, the OECD estimated output gap, or the unemployment rate in the forecasting regression, it is not clear, however, whether the source of the rejections comes from inflation, a measure of real economic activity, or both.

This question is addressed in Table 4 by reporting CW and DMW statistics when either inflation or a measure of real economic activity, instead of both, is included in the forecasting regressions. For the symmetric specifications without smoothing, the random walk (no predictability) null can be rejected in favor of the alternative at the 1 percent level with the unemployment rate and at the 5 percent level for the HP filtered output gap, the OECD output gap estimate, and inflation with both homogeneous and heterogeneous coefficients. Because the null can be rejected when either inflation or any of the real economic activity measures are included in the forecasting regression, this constitutes evidence of out-of-sample exchange rate predictability from a specification with Taylor rule fundamentals rather than a specification that is solely focused on either inflation or real activity. The evidence of forecast ability is also strong, with the equal MSPE null rejected at the 1 percent level for two of the three real economic activity measures and at the 5 percent level for the additional economic activity measure and inflation with both homogeneous and heterogeneous coefficients. As with the specifications that include both variables, the evidence of both predictability and forecast ability weakens with symmetric specifications with smoothing and asymmetric specifications without smoothing, and disappears with asymmetric specifications with smoothing.[24]

The next topic that we consider is "predictability" and "forecast ability" with revised data. While we subscribe to the view that, because revised data was not available to market participants at the time forecasts were made, only real-time data should be used to evaluate predictability and forecast ability, the use of revised data is so ubiquitous in the out-of-sample literature that we choose to examine it. The results with revised data are reported in Table 5. For the symmetric specifications without smoothing, the evidence of out-of-sample exchange rate predictability is less than with real-time data in four out of six cases and equal to that with real-time data for the two additional specifications. For the same specifications, the evidence of forecasting ability is less with revised data than with real-time data for all six cases. The largest decrease in the evidence of both predictability and forecasting ability occur when inflation and the OECD estimated output gap are included. This is consistent with the visual evidence in Figure 1 that the differences between the revised and real-time data are larger for the OECD estimated output gap than for the either the HP filtered output gap or the unemployment rate. As with real-time data, the evidence of both predictability and forecasting ability declines with the symmetric specifications with smoothing and the asymmetric specifications without smoothing, and disappears entirely with the asymmetric specifications with smoothing.

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[24] If the random walk null was not rejected with either inflation or the real economic activity measures, that would also have constituted evidence of predictability/forecast ability with Taylor rule fundamentals. If, however, the null was rejected for either inflation or the real activity measures, but not both, that would not have been evidence of predictability/forecast ability with Taylor rule fundamentals.

It is often argued that forward-looking monetary policy rules provide a superior description of central banks' behavior than rules based on the most recent estimates of inflation. Following Orphanides (2001, 2003), most of this literature uses Greenbook forecasts for the U.S. Since Greenbook forecasts are not publicly available past 2002 and there is no equivalent for the ECB, we use SPF forecasts for both. Two sets of results are depicted in Table 6. In the first three rows of each panel, the specifications from Table 3 are replicated with the current inflation rate replaced by forecasted inflation four quarters ahead.[25] We find evidence that out-of-sample exchange rate predictability is improved by using forecasted rather than actual inflation. For the most successful specification, the symmetric model without smoothing, the MSPE ratio declines with forecasted inflation in four out of six cases, rises in one, and is virtually unchanged in the remaining case. The evidence of both predictability and forecasting ability is stronger with forecasted inflation than with realized inflation. The bottom two rows of each panel report specifications with both forecasted inflation and four-quarter-ahead OECD output gap and unemployment rate forecasts. These specifications depict the strongest results in the paper. Again focusing on the most successful specification, the symmetric model without smoothing, the four MSPE ratios constitute four of the five lowest ratios in Tables 3 – 6. (The lowest ratio, also in Table 6, is with the unemployment rate for the specification with homogeneous coefficients.) For these four specifications, the random walk null hypothesis can be rejected in favor of predictability using the CW statistic and forecasting ability using the DMW statistic at the 1 percent level. As with realized data in Table 3, the evidence of both predictability and forecasting ability declines with the symmetric specifications with smoothing and the asymmetric specifications without smoothing, and disappears almost entirely with the asymmetric specifications with smoothing.

The results with forecasted inflation are illustrated in Figure 4, which depicts actual and predicted Dollar/Euro exchange rate changes for the 10 specifications of the symmetric model with no smoothing. Panels A – C use inflation forecasts and realized measures or real economic activity, while Panels D – E use forecasted inflation and either unemployment rate or OECD output gap forecasts. As in Figure 3, the "zero" line represents the random walk (no change) forecast, periods for which the actual and predicted values are both either positive or negative lower the MSPE ratio, and periods for which the actual and predicted values are of opposite sign raise the MSPE ratio. The predicted exchange rate changes with forecasted coefficients display considerable variability, although again not as much as the actual changes, and generally track the initial appreciation and subsequent depreciation of the dollar against the euro.

5.2 Testing for Superior Predictive Ability

Since we are testing simultaneously hypotheses that involve 24 different alternative models, conventional p-values can be misleading. As a result of extensive specification search, we may mistake significant results generated by chance for genuine evidence of predictive ability. To address the issue of

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[25] Given the one-quarter lag in data releases, we use forecasts of inflation made in period *t* with data through period $t-1$ for period $t+3$.

multiple hypothesis testing, we perform the test of superior predictive ability (SPA) proposed by Hansen (2005). The SPA test is designed to compare the out-of-sample performance of a benchmark model to that of a set of alternatives. This approach is a modification of the reality check for data snooping developed by White (2000). The advantages of the SPA test are that it is more powerful and less sensitive to the introduction of poor and irrelevant alternatives.[26]

We are interested in comparing the out-of-sample performance of linear exchange rate models to a naïve random walk benchmark. The SPA test can be used for comparing the out-of-sample performance of two or more models. It tests the composite null hypothesis that the benchmark model is not inferior to any of the alternatives against the alternative that at least one of the linear economic models has superior predictive ability. In the context of using the CW statistic to evaluate out-of-sample predictability, the null hypothesis is that the random walk has an MSE which is smaller than or equal to the adjusted MSE's of the linear models, as described by the last term in Equation (9). Therefore, rejecting the null indicates that at least one linear model is strictly superior to the random walk. SPA p-values take into account the search over models that preceded the selection of the model being compared to the benchmark. A low p-value suggests that the benchmark model is inferior to at least one of the competing models. A high p-value indicates that the data analyzed do not provide strong evidence that the benchmark is outperformed.[27]

Table 7 reports SPA p-values for nine sets of forecasts based on symmetric and asymmetric Taylor rule specifications that are compared to a random walk forecast. The first four rows of Table 8 have three measures of economic activities as alternatives. The next four rows report SPA p-values with a larger set of alternatives for the symmetric and asymmetric Taylor rule specifications. These statistics test the random walk benchmark against six alternatives. For example, "homogenous" would denote smoothing and no smoothing for the three economic activity measures. The ninth raw, denoted "all", tests the random walk benchmark against 12 alternatives: homogenous with smoothing, homogenous with no smoothing, heterogeneous with smoothing, and heterogeneous with no smoothing for the three measures of economic activity.

The SPA p-values strongly confirm the results in Table 3. Every symmetric specification is significant at the 5 percent level and no asymmetric specification is significant at the same level.[28] Within the class of symmetric specifications, the p-values are lower for the homogeneous and no smoothing specifications than for the heterogeneous and smoothing specifications and, not surprisingly, are lowest for the homogeneous specifications without smoothing.

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[26] Hansen (2005) provides details on the construction of the test statistic and confirms the advantages of the test by Monte Carlo simulations. We use the publicly available software package MULCOM to construct the SPA-consistent p-values for each country. The code, detailed documentation, and examples can be found at http://www.hha.dk/~alunde/mulcom/mulcom.htm

[27] We cannot use the SPA test to evaluate forecasting ability because the critical values for the DMW tests are not based on a standard normal distribution.

[28] For two of the nine specifications, the null can be rejected at the 10 percent level.

We conclude this section by contributing to the debate over whether the policies followed by the Bundesbank during the EMS are a good predictor of the policies followed by the ECB. While we do not estimate policy rules and cannot answer the question directly, we can ask whether the Taylor rule specifications using real-time data that were successful in providing evidence of out-of-sample predictability for the Dollar/Mark exchange rate continue to be successful for the Dollar/Euro rate. In Molodtsova, Nikolsko-Rzhevskyy, and Papell (2008), we find the strongest evidence of predictability for the Dollar/Mark rate with heterogeneous coefficients, no smoothing, and an asymmetric specification with the real exchange rate in the forecasting regression. For the Dollar/Euro rate in this paper, this specification does not provide any evidence of predictability when the results for all three measures of real economic activity are jointly evaluated. More generally, predictability for the Dollar/Mark exchange rate is only achieved with a heterogeneous and asymmetric specification while predictability for the Dollar/Euro exchange rate is stronger with a homogeneous and symmetric specification, with the latter more important than the former. These results are consistent with the view that, like the Fed but unlike the Bundesbank, the ECB does not put much weight on the exchange rate when setting interest rates.

5.3 Is Bad News About Inflation Good News for the Forecasted Exchange Rate?

The final topic that we consider is to explore what we are forecasting when we find evidence of out-of-sample exchange rate predictability. In Figure 5, we depict the dynamics of the coefficients on inflation and real economic activity differentials for the symmetric model with homogeneous coefficients and no smoothing which, as described in Table 7, has the lowest p-values among all specifications using the CW statistic. As reported in Table 3, it produces significant evidence of predictability when inflation and either the HP filtered output gap or the unemployment rate are included in the forecasting regression at the 1 percent level and when inflation and the OECD estimated output gap are included at the 5 percent level.

The coefficients on the inflation differentials, reported in Figure 5 along with 90% confidence interval bands, are virtually always negative and consistently significantly different from zero for all three measures of real economic activity.[29] Since the inflation differential equals U.S. inflation minus Euro Area inflation and the exchange rate is dollars per euro, a negative coefficient means that when U.S. inflation rises relative to Euro Area inflation, out-of-sample exchange rate predictability is achieved by forecasting dollar appreciation. This is consistent with the argument of Clarida and Waldman (2008) that "bad news about inflation is good news for the exchange rate" for inflation targeting countries. It is not consistent with using long-run PPP to forecast exchange rates, in which case an increase in U.S. inflation relative to Euro Area inflation would lead to forecasted dollar depreciation rather than appreciation. While there are many differences between the two studies - they consider multiple currencies while we examine only the Dollar/Euro rate, they use an event study methodology with a very short window while we utilize a longer

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[29] The only exception is the specification with the HP filtered output gap until mid-2001.

one-quarter-ahead horizon, they define "news" as unexpected changes in inflation while we use actual inflation differentials, and they examine the impact of inflation news on realized exchange rate changes while we examine the effect of inflation on forecasted exchange rate changes – we reinforce their findings using a very different methodology.

Figure 5 also depicts the coefficients on the three real economic activity differentials in the same forecasting regressions. The coefficients on the HP filtered output gap differentials are always negative, the coefficients on the OECD estimated output gap differentials are negative starting in 2001:Q3 and the coefficients on the unemployment differentials are positive starting in 2002:Q1, and generally are significant between 2003 and 2007.[30] Since the output gap represents the percentage by which output exceeds potential, a positive relative output gap differential between the U.S. and the Euro Area is "good news" for the U.S. and a positive unemployment differential is "bad news" for the U.S. We find that "good news about output or unemployment is good news for the forecasted exchange rate." The negative coefficients on the U.S. output gap relative to the Euro Area output gap reflect forecasted dollar appreciation while the positive coefficients on U.S. unemployment relative to Euro Area unemployment reflect forecasted dollar depreciation.

6. Conclusions

It has become standard practice for monetary policy evaluation of the Fed and ECB to be conducted via some variant of a Taylor rule where the short-term nominal interest rate responds to inflation and a measure of real economic activity.[31] While neither the Fed nor the ECB follow a mechanical rule and there is much disagreement over the coefficients and variables that enter the rule that best describes their behavior, even a cursory reading of FOMC press releases and the ECB Monthly Bulletin makes it clear why Taylor rules have become so ubiquitous. This is clear from both the Fed's dual mandate and the concern by the Governing Council of the ECB with real economic activity as well as price stability.

In this paper, we analyze whether the variables that normally enter central banks' interest-rate-setting rules, which we call Taylor rule fundamentals, can provide evidence of out-of-sample predictability and/or forecasting ability of the Dollar/Euro exchange rate. We use real-time data that was available to market participants at the point that their exchange rate forecasts were conducted and are careful to distinguish between predictability and forecasting.

The major result of the paper is that the null hypotheses of no predictability/no forecasting ability can be rejected against the alternative hypotheses of predictability/forecasting ability with Taylor rule fundamentals for a wide variety of specifications that include inflation and a measure of real economic activity in the forecasting regression. The strongest evidence comes from the simplest specifications that closely

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[30] The mixed results at the beginning of the sample may reflect initial uncertainty regarding how the ECB was going to conduct policy.

[31] The actions of the Fed in response to the 2008 financial crisis once the zero bound on nominal interest rates was approached is clearly an exception to this statement, which is why we did not extend our sample past 2007.

resemble the original Taylor rule, where the interest rates set by the Fed and the ECB respond only to inflation and a measure of real economic activity. The results are strengthened when either inflation or inflation and real economic activity forecasts, rather then realized values, are included in the forecasting regression, and are robust to whether the coefficients in the two Taylor rules are homogeneous or heterogeneous. The evidence is stronger with real-time data than with revised data. Bad news about inflation and good news about real economic activity both lead to out-of-sample predictability through forecasted exchange rate appreciation.

Meese and Rogoff (1983) compare the out-of-sample forecast accuracy of the random walk model to various structural and time series exchange rate models for three bilateral and one trade-weighted exchange rate, and find that the random walk model performs as well as any estimated model. While recent work using the CW statistic has reported rejections of the random walk null against various alternatives, Rogoff and Stavrakeva (2008) argue that these results constitute (using our terminology) evidence of predictability, but not evidence of forecasting ability. While we do not claim any generality for our results, we have shown that, using what is arguably the most important exchange rate of the twenty-first century, the U.S. Dollar/Euro rate, models with Taylor rule fundamentals with real-time data that only include inflation and a measure of real economic activity out-perform the random walk model using out-of-sample tests of both predictability and forecasting ability. The result that the model with Taylor rule fundamentals outperforms the random walk model out-of-sample holds whether the metric is comparison of the MSPEs between the two models, as in Meese and Rogoff (1983), or rejection of the random walk null by the DMW test, as in Rogoff and Stavrakeva (2008).

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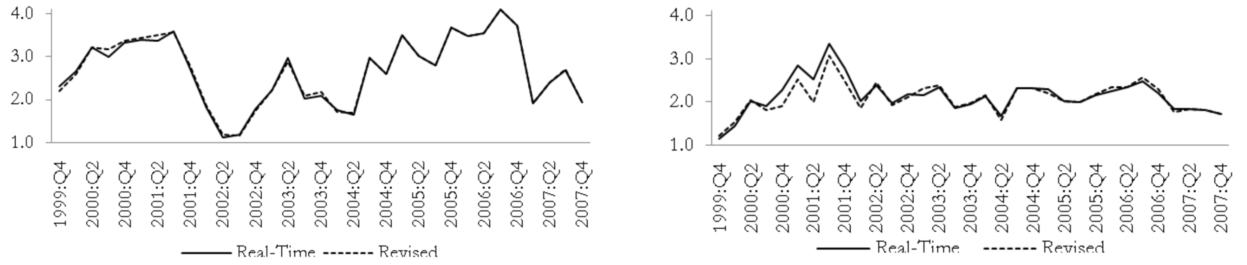
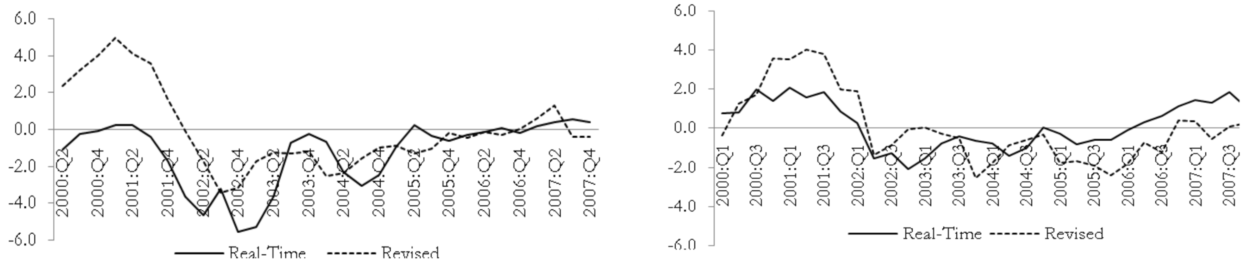


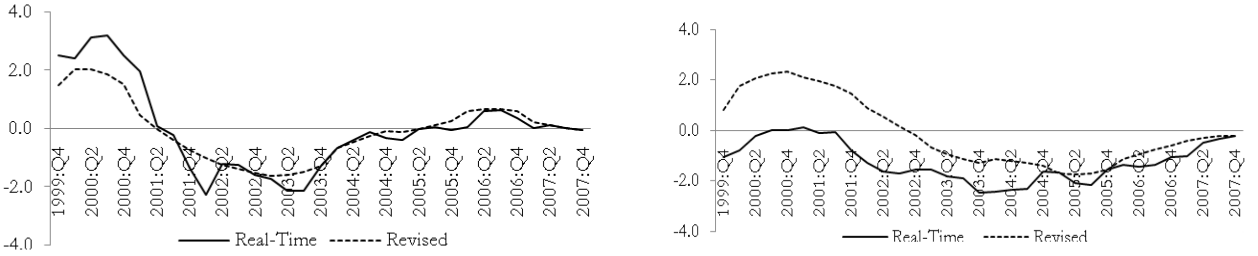
United States Euro Area

A. Real-Time and Revised Inflation



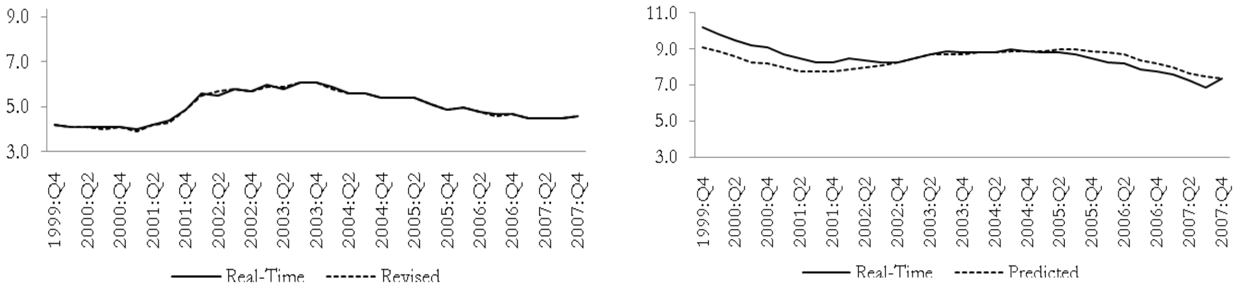
United States Euro Area

B. Real-Time and Revised HP Filtered Output Gap



United States Euro area

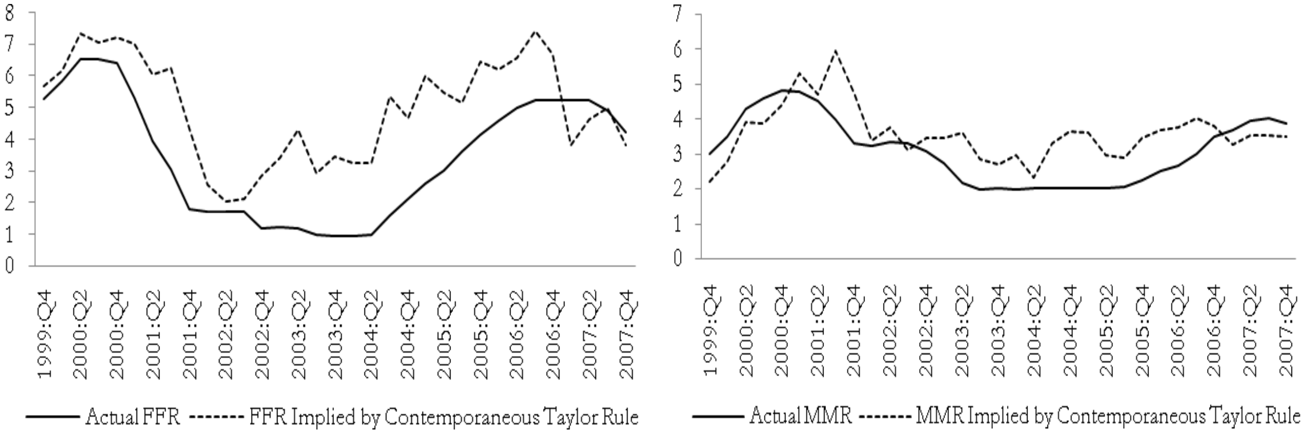
C. Real-Time and Revised OECD Estimates of Output Gap



United States Euro Area

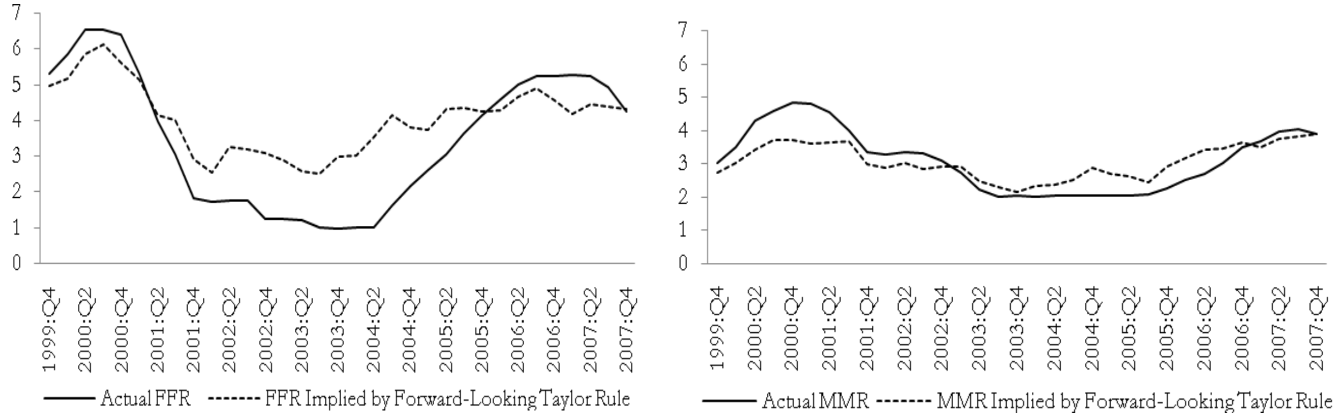
D. Real-Time and Revised Unemployment Rate

Figure 1. Real-Time and Revised Data for United States and Euro Area



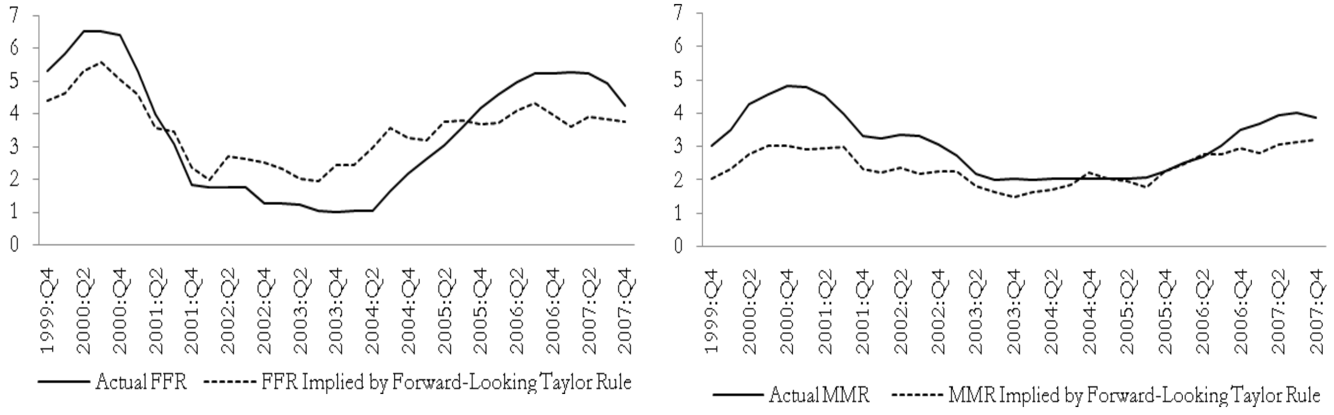
United States Euro Area

A. Contemporaneous Taylor Rule with the Equilibrium Real Interest Rate Set at 2%



United States Euro Area

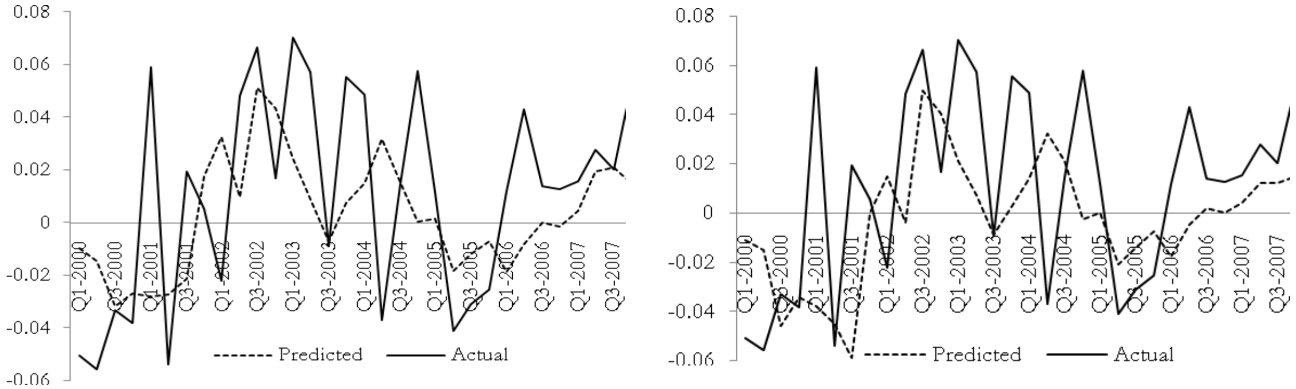
B. Forward-Looking Taylor Rule with the Equilibrium Real Interest Rate Set at 2%



United States Euro Area

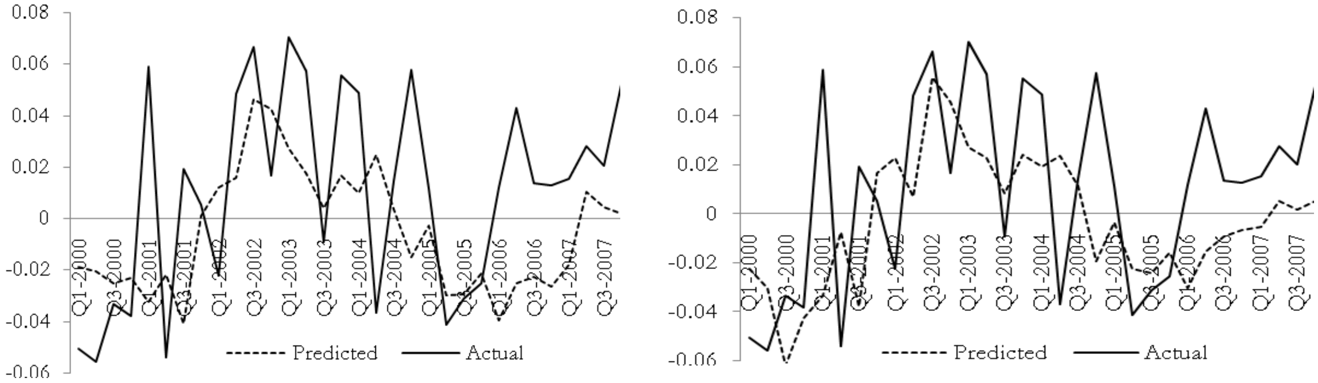
C. Forward-Looking Taylor Rule with a Calculated Equilibrium Real Interest Rate

Figure 2. Actual and Counterfactual Interest Rates for the U.S. and Euro Area



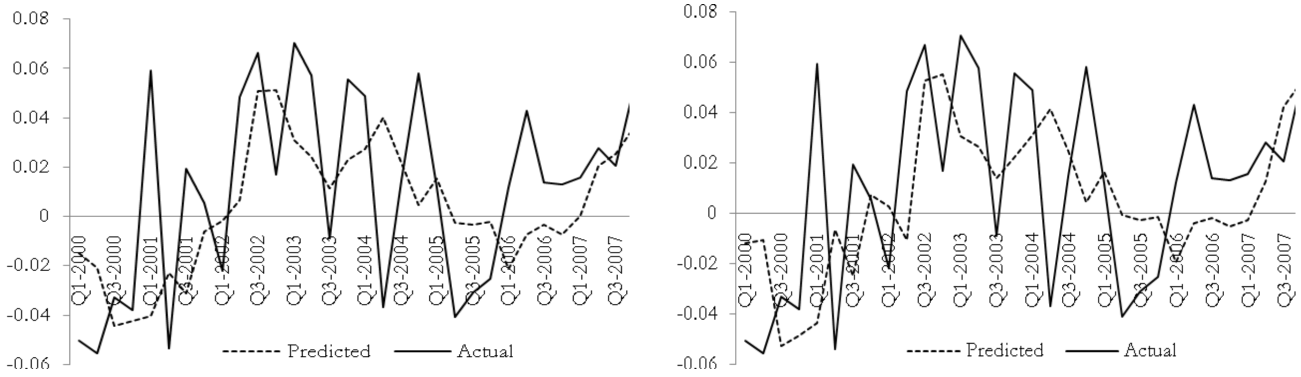
Homogeneous Coefficients Heterogeneous Coefficients

A. HP Filtered Output Gap



Homogeneous Coefficients Heterogeneous Coefficients

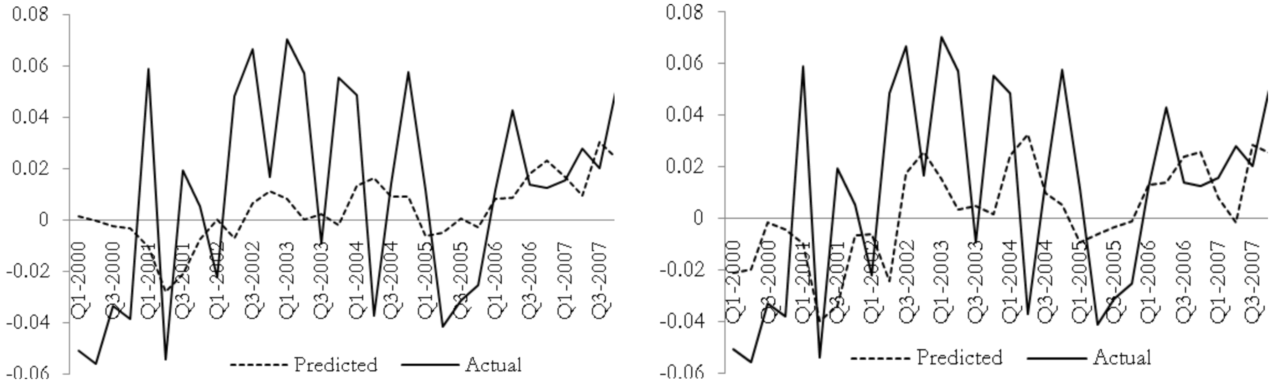
B. OECD Estimates of Output Gap



Homogeneous Coefficients Heterogeneous Coefficients

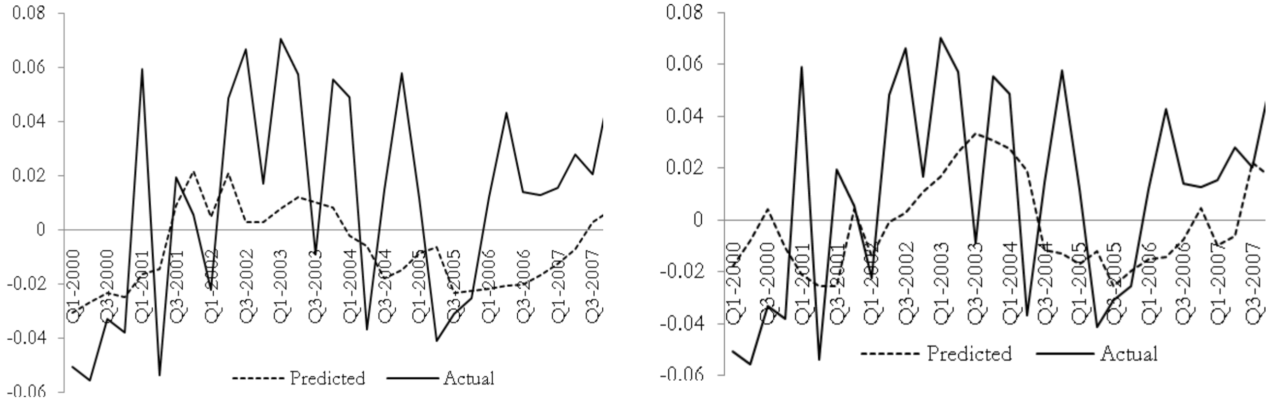
C. Unemployment Rate

Figure 3. Actual and Predicted Changes in the Dollar/Euro Exchange Rate Based on the Symmetric Model with No Smoothing



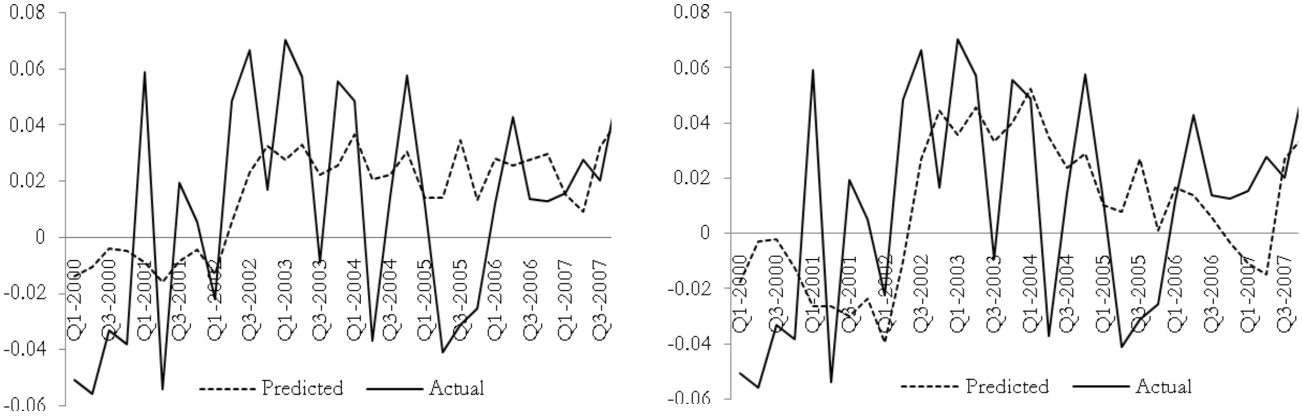
Homogeneous Coefficients Heterogeneous Coefficients

A. HP Filtered Output Gap



Homogeneous Coefficients Heterogeneous Coefficients

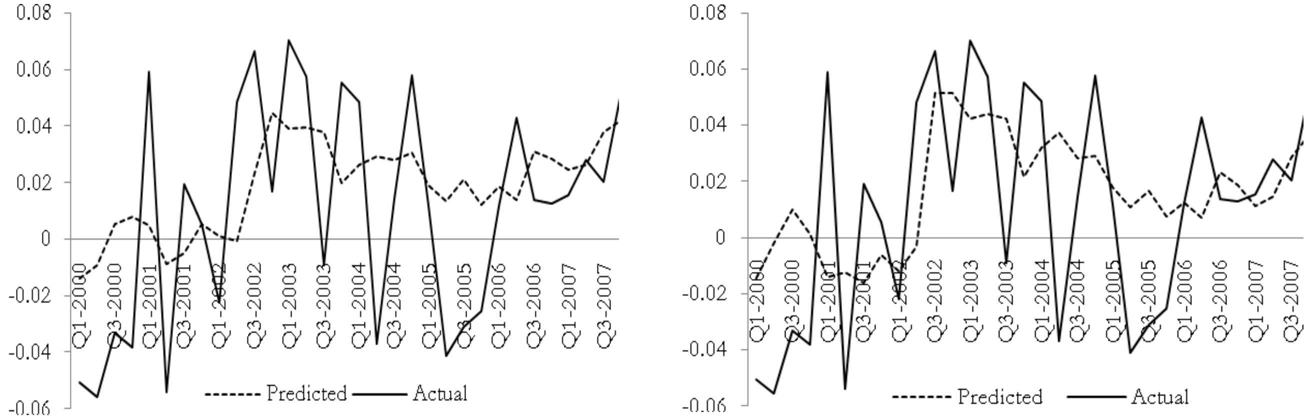
B. OECD Estimates of Output Gap



Homogeneous Coefficients Heterogeneous Coefficients

C. Unemployment Rate

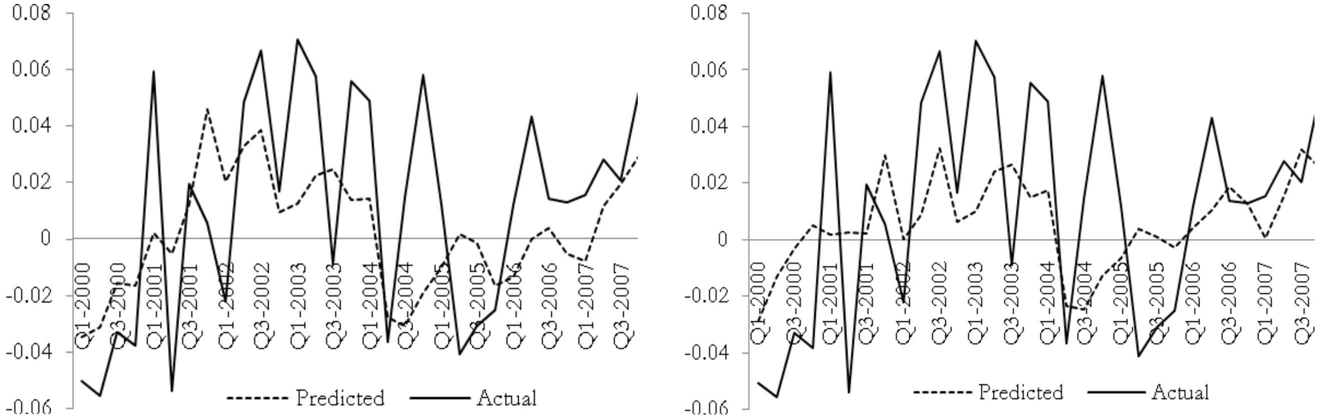
Figure 4. Actual and Predicted Changes in the Dollar/Euro Exchange Rate Based on the Symmetric Model with t+4 Inflation Forecasts and No Smoothing



Homogeneous Coefficients

Heterogeneous Coefficients

D. t+4 Unemployment Rate Forecasts

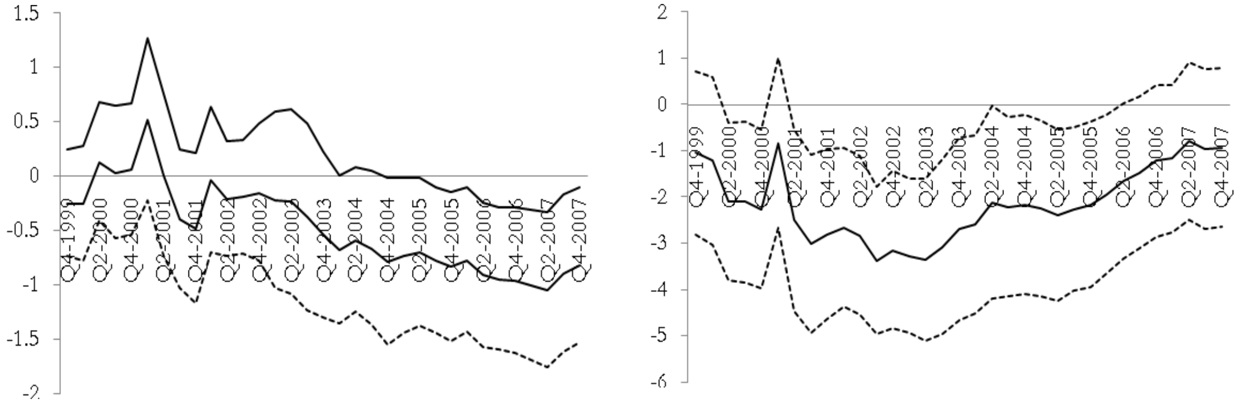


Homogeneous Coefficients

Heterogeneous Coefficients

E. t+4 OECD Output Gap Forecast

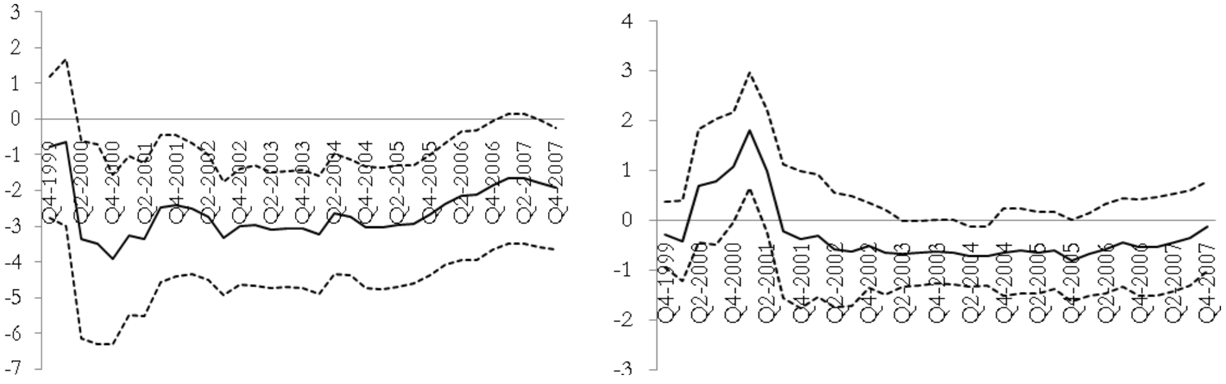
Figure 4. Actual and Predicted Changes in the Dollar/Euro Exchange Rate Based on the Symmetric Model with t+4 Inflation Forecasts and No Smoothing (continued)



Inflation Differential Coefficient

Output Gap Differential Coefficient

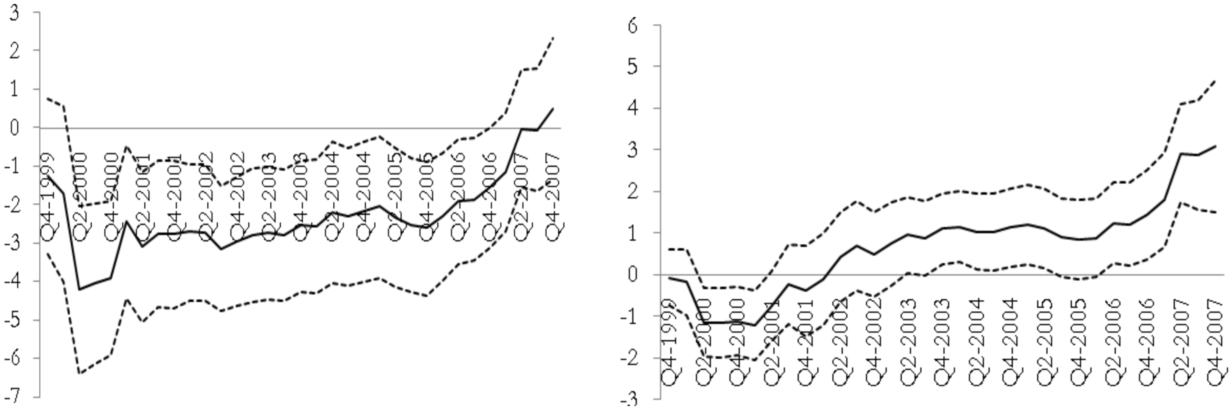
A. Homogeneous Specification with HP Filtered Output Gap



Inflation Differential Coefficient

Output Gap Differential Coefficient

B. Homogeneous Specification with OECD Estimates of Output Gap



Inflation Differential Coefficient

Unemployment Differential Coefficient

C. Homogeneous Specification with Unemployment Rate

Figure 5. Dynamics of Forecasting Equation Coefficients

Table 1. Summary Statistics

	U.S.				Euro Area			
	Mean	*SD*	*Min*	*Max*	*Mean*	*SD*	*Min*	*Max*
A. Revised Data								
Revised Inflation	2.68	0.78	1.16	4.08	2.09	0.35	1.23	3.06
Revised HP Filtered Output Gap	-0.05	2.11	-3.41	4.96	0.07	1.80	-2.54	4.00
Revised OECD Output Gap	-0.03	1.05	-1.61	2.05	-0.08	1.36	-1.76	2.33
Revised Unemployment Rate	5.00	0.69	3.90	6.10	8.40	0.49	7.40	9.10
B. Real-Time Data								
Real-Time Inflation	2.67	0.78	1.12	4.08	2.14	0.40	1.16	3.34
Real-Time HP Filtered Output Gap	-1.19	1.76	-5.55	0.56	0.18	1.20	-2.11	2.05
Real-Time OECD Output Gap	0.02	1.48	-2.27	3.20	-1.22	0.79	-2.46	0.12
Real-Time Unemployment Rate	5.01	0.68	4.00	6.10	8.54	0.68	6.90	10.20

Note: The statistics reported for each variable are: Mean, the mean, SD, the standard deviation, Min, and Max, the minimum and maximum values. The data is for 1999:Q4-2007:Q4.

Table 2. Descriptive Statistics of Revisions

	U.S.					Euro Area				
$X_{Revised}-X_{Real\text{-}time}$	*Mean*	*SD*	*NORM*	*Corr with* $X_{Revised}$	*Corr with* $X_{Real\text{-}time}$	*Mean*	*SD*	*NORM*	*Corr with* $X_{Revised}$	*Corr with* $X_{Real\text{-}time}$
Inflation	0.01	0.06	0.03	0.15	0.08	-0.05	0.15	0.00	-0.15	-0.50
HP Filtered Output	1.14	1.96	0.21	0.63	-0.36	-0.11	1.34	0.32	0.75	0.01
OECD Output Gap	-0.05	0.58	0.11	-0.62	-0.83	1.13	0.85	0.22	0.84	0.36
Unemployment Rate	-0.01	0.06	0.00	0.16	0.25	-0.13	0.50	0.34	0.05	-0.69

Note: The statistics reported for each variable are: Mean, the mean, SD, the standard deviation, NORM, the p-values for normality test, Corr with $X_{Revised}$, Corr with $X_{Real\text{-}time}$ are correlations of revisions with revised of real-time series. Positive and significant value of the "mean" revision indicates that the variable was consistently revised upwards.

Table 3: One-Quarter-Ahead Out-of-Sample Forecasts with Real-Time Data

	w/o Smoothing		*w/ Smoothing*	
	Symmetric	*Asymmetric*	*Symmetric*	*Asymmetric*
A. Homogenous Coefficients				
HP Filtered Output Gap	0.82	0.89	0.95	1.04
CW statistic	2.415***	2.005**	2.020**	1.126
DMW statistic	1.017***	0.636**	0.234**	-0.235
OECD Estimates of Output Gap	0.97	1.05	1.09	1.26
CW statistic	1.820**	1.435*	1.301*	0.068
DMW statistic	0.150*	-0.221	-0.396	-1.382
Unemployment Rate	0.80	0.89	0.89	1.04
CW statistic	2.627***	1.981**	2.370***	1.324*
DMW statistic	0.867**	0.559**	0.427**	-0.206*
B. Heterogeneous Coefficients				
HP Filtered Output Gap	0.93	1.05	0.97	1.14
CW statistic	1.945**	1.297*	2.132**	0.934
DMW statistic	0.307**	-0.246	0.140**	-0.677
OECD Estimates of Output Gap	0.91	0.99	1.06	1.16
CW statistic	2.184**	1.738**	1.708**	0.806
DMW statistic	0.434**	0.063*	-0.242*	-0.916
Unemployment Rate	0.90	1.09	1.02	1.15
CW statistic	2.320**	1.401*	1.880**	1.175
DMW statistic	0.418**	-0.428	-0.085*	-0.736

Notes: The table reports the ratio of the out-of-sample MSPEs of the linear model to that of the random walk model and the CW and DMW statistics for the tests of equal predictability and forecasting ability, respectively, between the two models. Panel A contains the results for homogenous Taylor rule models that restrict coefficients on the inflation and output gap in the two countries to be the same, and Panel B contains the results for heterogeneous Taylor rule models. The statistics are reported for the following classes of models: Symmetric Taylor rule models that exclude real exchange rate from the forecasting regression equation, and Asymmetric Taylor rule models, that are subdivided into the models w/ Smoothing and w/o Smoothing that either include or exclude lagged interest rate. * ,**, and *** denote test statistics significant at 10, 5, and 1% level, respectively, based on standard normal critical values for the CW statistic and McCracken's (2007) critical values for the DMW statistic. We use real-time quarterly data from 1999:Q4 to 2007:Q4 for the United States and the Euro Area. The data for the first vintage starts in 1991:Q1. Rolling regressions with a 34-quarter window are used to predict 32 exchange rate changes from 2000:Q1 to 2007:Q4 based on the models with Taylor rule fundamentals.

Table 4: One-Quarter-Ahead Out-of-Sample Forecasts with Real-Time Data and Either Inflation or the Output Gap

	w/o Smoothing		*w/ Smoothing*	
	Symmetric	*Asymmetric*	*Symmetric*	*Asymmetric*
A. Homogenous Coefficients				
HP Filtered Output Gap	0.89	0.98	0.93	1.08
CW statistic	2.159**	1.158	1.972**	0.836
DMW statistic	1.291***	0.158*	0.364**	-0.472
OECD Estimates of Output Gap	0.91	1.05	1.06	1.32
CW statistic	1.974**	0.846	1.550*	-0.373
DMW statistic	0.871**	-0.425	-0.275	-1.582
Unemployment Rate	0.74	0.87	0.97	1.15
CW statistic	3.063***	2.016***	2.123**	0.720
DMW statistic	1.812***	0.747**	0.117*	-0.717
Inflation	0.91	1.02	1.02	1.16
CW statistic	1.906**	1.313*	1.585*	0.595
DMW statistic	0.546**	-0.124	-0.070	-0.861
B. Heterogeneous Coefficients				
HP Filtered Output Gap	0.87	0.99	0.95	1.13
CW statistic	2.132**	1.343*	2.004**	0.870
DMW statistic	1.077***	0.072*	0.215**	-0.640
OECD Estimates of Output Gap	0.84	0.93	1.04	1.23
CW statistic	2.220**	1.616*	1.726**	-0.013
DMW statistic	1.027***	0.418**	-0.193	-1.252
Unemployment Rate	0.85	0.97	0.97	1.11
CW statistic	2.453***	1.801**	1.934**	0.894
DMW statistic	0.611**	0.178*	0.131*	-0.603
Inflation	0.90	1.01	1.02	1.18
CW statistic	1.872**	1.258	1.542*	0.431
DMW statistic	0.551**	-0.067	-0.107*	-1.000

Notes: The table reports the ratio of the out-of-sample MSPEs of the linear model to that of the random walk model and the CW and DMW statistics for the tests of equal predictability and forecasting ability, respectively, between the two models. Panel A contains the results for homogenous Taylor rule models that restrict coefficients on the inflation and output gap in the two countries to be the same, and Panel B contains the results for heterogeneous Taylor rule models. The statistics are reported for the following classes of models: Symmetric Taylor rule models that exclude real exchange rate from the forecasting regression equation, and Asymmetric Taylor rule models, that are subdivided into the models w/ Smoothing and w/o Smoothing that either include or exclude lagged interest rate. * ,**, and *** denote test statistics significant at 10, 5, and 1% level, respectively, based on standard normal critical values for the CW statistic and McCracken's (2007) critical values for the DMW statistic. We use real-time quarterly data from 1999:Q4 to 2007:Q4 for the United States and the Euro Area. The data for the first vintage starts in 1991:Q1. Rolling regressions with a 34-quarter window are used to predict 32 exchange rate changes from 2000:Q1 to 2007:Q4 based on the models with Taylor rule fundamentals.

Table 5: One-Quarter-Ahead Out-of-Sample Forecasts with Revised Data

	w/o Smoothing		*w/ Smoothing*	
	Symmetric	*Asymmetric*	*Symmetric*	*Asymmetric*
A. Homogenous Coefficients				
HP Filtered Output Gap	0.85	0.99	1.10	1.24
CW statistic	2.198**	1.624*	1.452*	0.664
DMW statistic	0.831**	0.035*	-0.373	-1.079
OECD Estimates of Output Gap	1.04	1.13	1.22	1.41
CW statistic	1.465*	1.118	1.197	0.371
DMW statistic	-0.179	-0.580	-0.703	-1.461
Unemployment Rate	0.92	0.97	1.05	1.14
CW statistic	2.198**	1.756**	1.938**	0.923
DMW statistic	0.284*	0.105*	-0.150	-0.616
B. Heterogeneous Coefficients				
HP Filtered Output Gap	0.99	1.06	1.21	1.18
CW statistic	1.988**	1.723**	1.871**	1.248
DMW statistic	0.054*	-0.259	-0.664	-0.817
OECD Estimates of Output Gap	1.24	1.30	1.38	1.52
CW statistic	1.380*	1.021	1.297*	0.560
DMW statistic	-0.701	-0.820	-1.042	-1.225
Unemployment Rate	0.96	1.20	1.17	1.28
CW statistic	2.264**	1.474*	1.579*	1.155
DMW statistic	0.116*	-0.694	-0.580	-1.173

Notes: The table reports the ratio of the out-of-sample MSPEs of the linear model to that of the random walk model and the CW and DMW statistics for the tests of equal predictability and forecasting ability, respectively, between the two models. Panel A contains the results for homogenous Taylor rule models that restrict coefficients on the inflation and output gap in the two countries to be the same, and Panel B contains the results for heterogeneous Taylor rule models. The statistics are reported for the following classes of models: Symmetric Taylor rule models that exclude real exchange rate from the forecasting regression equation, and Asymmetric Taylor rule models, that are subdivided into the models w/ Smoothing and w/o Smoothing that either include or exclude lagged interest rate. * ,**, and *** denote test statistics significant at 10, 5, and 1% level, respectively, based on standard normal critical values for the CW statistic and McCracken's (2007) critical values for the DMW statistic. The revised data is constructed from the 2007:Q4 vintage for both the U.S. and the Euro Area. Starting in 1991:Q1, we estimate rolling regressions with a 34-quarter window to predict 32 exchange rate changes from 2000:Q1 to 2007:Q4.

Table 6: One-Quarter-Ahead Out-of-Sample Forecasts with Real-Time Data and t+4 Inflation Forecasts

	w/o Smoothing		*w/ Smoothing*	
	Symmetric	*Asymmetric*	*Symmetric*	*Asymmetric*
A. Homogenous Coefficients				
HP Filtered Output Gap	0.88	0.99	0.91	1.08
CW statistic	2.279**	1.238	2.056**	1.252
DMW statistic	1.337***	0.065*	0.397**	-0.343
OECD Estimates of Output Gap	0.91	1.06	1.08	1.21
CW statistic	1.889**	1.101	1.974**	0.861
DMW statistic	0.691**	-0.433	-0.280	-0.836
Unemployment Rate	0.69	0.79	0.89	1.17
CW statistic	3.511***	2.652***	2.336***	0.861
DMW statistic	1.826***	1.077***	0.383**	-0.737
t + 4 Unemployment Rate Forecasts	0.74	0.73	0.93	1.11
CW statistic	3.329***	2.926***	2.073**	0.945
DMW statistic	1.461***	1.391***	0.280**	-0.562
t+4 OECD Output Gap Forecasts	0.70	0.76	0.82	0.98
CW statistic	3.366***	2.966***	2.556***	1.938**
DMW statistic	1.910***	1.578***	0.978***	0.104*
B. Heterogeneous Coefficients				
HP Filtered Output Gap	0.85	1.04	0.92	1.20
CW statistic	2.391***	0.997	2.212**	1.043
DMW statistic	1.012***	-0.291	0.316**	-0.745
OECD Estimates of Output Gap	0.90	0.93	0.98	1.21
CW statistic	1.913**	1.651**	1.851**	0.494
DMW statistic	0.676**	0.610**	0.089**	-0.928
Unemployment Rate	0.84	1.15	1.03	1.33
CW statistic	2.554***	0.663	1.417*	0.010
DMW statistic	0.754**	-0.871	-0.143*	-1.466
t + 4 Unemployment Rate Forecasts	0.78	1.07	1.11	1.35
CW statistic	2.694***	0.947	1.186	-0.289
DMW statistic	1.067***	-0.403	-0.496	-1.756
t+4 OECD Output Gap Forecasts	0.73	0.84	0.89	1.17
CW statistic	3.345***	2.300**	2.310**	0.856
DMW statistic	2.252***	1.488***	0.565**	-0.777

Notes: The table reports the ratio of the out-of-sample MSPEs of the linear model to that of the random walk model and the CW and DMW statistics for the tests of equal predictability and forecasting ability, respectively, between the two models. * ,**, and *** denote test statistics significant at 10, 5, and 1% level, respectively, based on standard normal critical values for the CW statistic and McCracken's (2007) critical values for the DMW statistic. The forward-looking specifications use the Philadelphia Fed SPF forecasts for the U.S. and SPF forecasts for Euro Area. For the Euro Area, the first round of the survey was conducted in 1999:Q1. To construct the t+4 inflation and unemployment forecasts, we use the realized t+4 values of inflation before 1998:Q4 and real-time Euro Area SPF forecasts from 1999:Q1 to 2007:Q4. OECD output gap forecasts are from OECD Economic Outlook.

Table 7: Tests for Superior Predictive Ability

Models	*Symmetric*	*Asymmetric*
Homogenous w/ Smoothing	0.023**	0.175
Homogenous w/o Smoothing	0.016**	0.063*
Heterogenous w/ Smoothing	0.034**	0.227
Heterogenous w/o Smoothing	0.030**	0.109
Homogenous	0.021**	0.128
Heterogenous	0.039**	0.144
Smoothing	0.031**	0.224
No Smoothing	0.021**	0.085*
All	0.028**	0.126

Notes: The table reports SPA p-values for the CW statistic for nine sets of forecasts based on *Symmetric* (without the real exchange rate) and *Asymmetric* (with the real exchange rate) Taylor rule specifications that are compared to a random walk forecast. * ,**, and *** denote test statistics significant at 10, 5, and 1% level, respectively. Each row contains the results for the following classes of models: *All*, all Taylor rule models, *Smoothing* and *No Smoothing*, models that include or exclude interest rate smoothing, *Homogenous* and *Heterogeneous*, models that restrict or do not restrict the coefficients on inflation and measures of economic activity to be the same for the U.S. and Euro Area.