Out-of-Sample Exchange Rate Predictability with Taylor Rule Fundamentals

Tanya Molodtsova* and David H. Papell**

University of Houston

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Abstract

An extensive literature that studied the performance of empirical exchange rate models following Meese and Rogoff's (1983a) seminal paper has not yet convincingly overturned their result of no out-of-sample predictability of exchange rates. This paper extends the conventional set of models of exchange rate determination by investigating predictability of models that incorporate Taylor rule fundamentals. Using Clark and West's (2006) recently developed inference procedure for testing the equal predictability of two nested models, we find evidence of short-term predictability for 11 out of 12 currencies vis-à-vis the U.S. dollar over the post-Bretton Woods float. The evidence of predictability is much stronger with Taylor rule models than with conventional interest rate, purchasing power parity, or monetary models.

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* Department of Economics, University of Houston, Houston, TX 77204-5019. Tel: +1 (832) 858-2076 Email: tvmolodt@mail.uh.edu

**Department of Economics, University of Houston, Houston, TX 77204-5882. Tel/Fax: +1 (713) 743-3807/3798. Email: dpapell@uh.edu

1. Introduction

The failure of open-economy macro theory to explain exchange rate behavior using economic fundamentals has prevailed in the international economics literature since the seminal papers by Meese and Rogoff (1983a, 1983b), who examine the out-of-sample performance of three empirical exchange rate models during the post-Bretton Woods period. Using monthly data from March 1973 through November 1980 for generating one-to-twelve month horizon predictions, they find that the random walk model produces consistently more accurate forecasts than the empirical exchange rate models on the basis of the root mean squared error (RMSE) comparison. The authors conclude that economic models of exchange rate determination of the 1970's vintage do not perform better than a naïve "no change" model.

Starting with Mark (1995), a number of studies have found evidence of greater predictability of economic exchange rate models at longer horizons. Chen and Mark (1996) assess the out-of-sample performance of the three alternative fundamentals proposed in the literature: those implied by the purchasing power parity (PPP), the uncovered interest rate parity (UIRP) and the flexible-price monetary model. They find evidence of greater predictability of economic models at long horizons, with monetary fundamentals having the most predictive power. Chinn and Meese (1995), using an error correction version of the model, also find evidence of long-term exchange rate predictability.

The findings that macroeconomic fundamentals have predictive power relative to the random walk and out-of sample performance of economic models increases with the forecast horizon have been questioned in subsequent research by Kilian (1999), Faust, Rogers and Wright (2003), and Berkowitz and Giorgianni (2001). The recent comprehensive study by Cheung, Chinn and Pascual (2005) examines the out-of-sample performance of the interest rate parity, monetary, productivity-based and behavioral exchange rate models and concludes that none of the models consistently outperforms the random walk at any horizon.

There is a disconnect between most research on out-of-sample exchange rate predictability, which is based on empirical exchange rate models of the 1970s, and the literature on monetary policy evaluation, which is based on some variant of the Taylor (1993) rule. A recent literature uses Taylor rules to model exchange rate determination. The Taylor rule specifies that the central bank adjusts the short-run nominal interest rate in response to changes in inflation and the output gap. By specifying Taylor rules for two countries and subtracting one from the other, an equation is derived with the interest rate differential on the left-hand-side and the inflation and output gap differentials on the right-hand-side. If one or both central banks also target the purchasing power parity (PPP) level of the exchange rate, the real exchange rate will also appear on the right-hand-side. Positing that the interest rate differential equals the expected rate of depreciation by UIRP and solving expectations forward, an exchange rate equation is derived.

Engel and West (2005) use the Taylor rule model as an example of present value models where asset prices (including exchange rates) will approach a random walk as the discount factor approaches one. Engel and West (2006) construct a "model-based" real exchange rate as the present value of the difference between

home and foreign output gaps and inflation rates, and find a positive correlation between the "model-based" rate and the actual dollar-mark real exchange rate. Mark (2007) considers Taylor rule interest rate reaction functions for Germany and the U.S. and estimates the real dollar-mark exchange rate path assuming that the exchange rate is priced by uncovered interest rate parity. He provides evidence that the interest rate differential can be modeled as a Taylor rule differential and the real dollar-mark exchange rate is linked to the Taylor rule fundamentals, which may provide a resolution for the exchange rate disconnect puzzle.

In this paper, we examine out-of-sample exchange rate predictability with Taylor rule fundamentals. The starting point for our analysis is the same as for the Taylor rule model of exchange rate determination, the Taylor rule for the foreign country is subtracted from the Taylor rule for the United States (the domestic country). 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. In Taylor's (1993) original formulation, the rule posits 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 foreign central bank follows a similar rule, we construct a *symmetric* model with inflation and the output gap on the right-hand-side. Following Clarida, Gali, and Gertler (1998), (hereafter CGG), we can also posit that the foreign central bank 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 the lagged interest rate differential appears on the right-hand-side. Alternatively, we can derive a model with *no smoothing* that does not include the lagged interest rate differential. Models with and without smoothing can be symmetric or asymmetric.

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

4. 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*. The models with and without a constant can be either symmetric or asymmetric, with or without smoothing.

The models we have specified all have the interest rate differential on the left-hand-side. The most straightforward way to construct an exchange rate forecasting equation is, using UIRP, to replace 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 either inflation or the output gap would produce a forecast of exchange rate depreciation. This approach, however, is unsatisfactory for three reasons. First, an extensive literature has shown that regressing exchange rate changes on interest rate differentials not only does not produce coefficients equal to one, as predicted by UIRP, it often produces negative coefficients. Second, the recent "carry trade" literature indicates that countries with high interest rates appear to have appreciating currencies. Third, as argued by Clarida and Waldman (2007), if an unexpected increase of the inflation rate above its target creates the expectation that the central bank will respond by raising the interest rate, the exchange rate will appreciate, rather than depreciate, in response to the news.[1] We therefore use Taylor rule fundamentals, the variables that enter various specifications of the Taylor rule, to forecast exchange rate changes, but do not impose restrictions on the direction of the forecasts.

The relevant literature on exchange rate predictability compares out-of-sample predictability of two models (linear fundamental-based model and a random walk) on the basis of different measures. The most commonly used measure of predictive ability is mean squared prediction error (MSPE). In order to evaluate out-of-sample performance of the models based on the MSPE comparison, tests for equal predictability of two non-nested models, introduced by Diebold and Mariano (1995) and West (1996), are most often used (henceforth, DMW tests).[2]

While the DMW tests are appropriate for non-nested models, the simulation evidence in Clark and McCracken (2001, 2005), McCracken (2007) and Corradi and Swanson (2007) demonstrates that when comparing MSPE's of two nested models mechanical application of the DMW procedures leads to non-normal test statistics and the use of standard normal critical values usually results in very poorly sized tests, with far too few rejections of the null.[3] This is a problem for out-of-sample exchange rate predictability because, since the null is a random walk, all tests with fundamental-based models are nested. In addition to being severely undersized, the standard DMW procedure demonstrates very low power, which makes this statistic ill-suited for detecting departures from the null. Rossi (2005) documents the existence of size distortions of the DMW tests by revisiting the Meese and Rogoff puzzle. While her paper suggests a possible way to solve this problem by adjusting critical value of the tests, the resulting statistic has low power.

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[1] Engel (2007) shows that this result appeared earlier in Engel and West (2006).

[2] Under the null of equal predictive accuracy, the DMW statistic is assumed to be zero and has an asymptotic standard normal distribution. One of the most highly cited studies that use the DMW statistic for testing for equal predictive ability of the models is Mark (1995).

[3] McCracken (2007) shows that using standard normal critical values for the DMW statistic results in severely undersized tests, with tests of nominal 0.10 size generally having actual size less than 0.02.

Why are we concerned with undersized tests? In many cases, undersized tests are much less of a problem than oversized tests. In the case of exchange rate predictability, however, the typical result is that the random walk null cannot be rejected in favor of the model-based alternative. Using undersized tests, such as the unadjusted DMW statistic with standard normal critical values, could lead to the incorrect conclusion that the random walk forecasts better than the economic model when the model has statistically significant predictive power.

We apply a recently developed inference procedure for testing the null of equal predictive ability of a linear econometric model and a martingale difference model proposed by Clark and West (2006, 2007), which we call the CW procedure. This methodology is preferable to the standard DMW procedure when the two models are nested. The test statistic takes into account that under the null the sample MSPE of the alternative model is expected to be greater than that of the random walk model and adjusts for the upward shift in the sample MSPE of the alterative model. The simulations in Clark and West (2006) suggest that the inference made using asymptotically normal critical values results in properly-sized tests.[4]

To our knowledge, four previous papers use the CW statistic to investigate exchange rate predictability. Clark and West (2006) examine exchange rate predictability using interest rate differentials for 4 countries vis-à-vis the U.S. dollar: Japan, Switzerland, Canada and the U.K. Using the CW statistic they find that the economic model significantly outpredicts the random walk for Canada and Switzerland at a 1-month horizon. Using the CW inference procedure, Gourinchas and Rey (2007) find that the ratio of net exports to net foreign assets forecasts movements in FDI-weighted and trade-weighted exchange rates better than the no-change model at 1 to 16 quarter horizon. Alquist and Chinn (2007) examine out-of-sample performance of the sticky-price monetary model, UIRP model and a measure of external imbalances suggested by Gourinchas and Rey. The CW procedure rejects the null of no predictability for UIRP at long horizons and a transformation of net exports variable performs well at short horizons. Engel, Mark, and West (2007), using a particular variant of the Taylor rule specification derived in an earlier version of this paper as well as models with monetary and PPP fundamentals, find little evidence of exchange rate predictability at short horizons, but more at longer horizons.

We evaluate the out-of-sample exchange rate predictability of models with Taylor rule fundamentals using the CW statistic for 12 OECD countries vis-à-vis the United States over the post-Bretton Woods period starting in March 1973 and ending in December 1998 for the European Monetary Union countries and June 2006 for the others. In order to construct Taylor rule fundamentals, we need to define the output gap, and we use deviations from a linear trend, deviations from a quadratic trend, and the Hodrick-Prescott filter. Recent work on estimating Taylor rules for the United States, notably Orphanides (2001), has emphasized the importance of using real-time data, the data available to central banks at the point that policy

[4] West (2006) provides a summary of recent literature on asymptotic inference about forecasting ability.

decisions are made. Since real-time data is not available for most of these countries over this period, we define potential output using "quasi-real time" trends which, although using revised data, are updated each period so that *ex-post* data is not used to construct the trends.[5] Orphanides and van Norden (2002), using a variety of detrending techniques, show that most of the difference between fully revised and real-time data comes from using *ex post* data to construct potential output and not from the data revisions themselves.

The results provide strong evidence of short-run exchange rate predictability using Taylor rule fundamentals. At the 1-month horizon, we find statistically significant evidence of exchange rate predictability at the 5 percent level for 11 of the 12 currencies. The models with heterogeneous coefficients, smoothing, and/or a constant provide substantially more evidence of predictability than the models with homogeneous coefficients, no smoothing, and/or no constant. The symmetric models (no exchange rate targeting) provide more evidence of predictability than the asymmetric models for the specifications that include a constant, but less evidence for the specifications that do not include a constant. Overall, the specification that produced the most evidence of exchange rate predictability was a symmetric model with heterogeneous coefficients, smoothing, and a constant. For that model, the no predictability null was rejected at the five percent level for 10 of the 12 countries for at least one of the three output gap measures, and at the 10 percent level for at least two of the three measures.

One issue concerning these results is that, because we are estimating numerous models, inference based on the p-values of the most statistically significant models is likely to be overstated. This is particularly important because we use three output gap measures for each specification. In order to correct for data snooping, we implement Hansen's (2005) test of superior predictive ability. While, as expected, the level of statistical significance falls, there is still substantial evidence of exchange rate predictability.

In order to compare our results with Taylor rule fundamentals with other models, we use the CW statistic to evaluate the out-of-sample performance of exchange rate models based on interest rate differentials, as in Clark and West (2006), purchasing power parity fundamentals, and three variants of monetary fundamentals, as in Engel, Mark, and West (2007), for the same currencies and time period. The evidence of predictability is much weaker for these models than for the models with Taylor rule fundamentals. At the 1-month horizon, we find statistically significant evidence of exchange rate predictability at the 5 percent level for only 3 of the 12 currencies using at least one of the models and at the 10 percent level for one additional currency. For all four currencies, the strongest evidence is provided by the model based on interest rate differentials that includes a constant term.[6]

[5] While real-time OECD data is available since 1999, this period is too short for comparability with previous work over the post-Bretton Woods period.

[6] We also investigated longer (6, 12, and 36 month) horizons, and found no evidence of exchange rate predictability for either the Taylor rule or the other models.

2. Exchange Rate Models

2.1 Taylor Rule Fundamentals

We examine the linkage between the exchange rates 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

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

where i_t^* is the target for the short-term nominal interest rate, π_t is the inflation rate, π_t^* 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.

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 $\mu = r^* - \phi\pi_t^*$, which leads to the following equation,

(2) $$i_t^* = \mu + \lambda\pi_t + \gamma y_t$$

where $\lambda = 1 + \phi$.

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, following CGG, to include the real exchange rate in specifications for other countries,

(3) $$i_t^* = \mu + \lambda\pi_t + \gamma y_t + \delta q_t$$

where q_t is the real exchange rate. 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.

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:

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

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

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

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 foreign country from that for the U.S.:

(6) $$i_t - \tilde{i}_t = \alpha + \alpha_{u\pi}\pi_t - \alpha_{f\pi}\tilde{\pi}_t + \alpha_{uy}y_t - \alpha_{fy}\tilde{y}_t - \alpha_q\tilde{z}_t + \rho_u i_{t-1} - \rho_f\tilde{i}_{t-1} + \eta_t$$

where ~ denotes foreign variables, u and f are coefficients for the United States and the foreign country, α is a constant, $\alpha_\pi = \lambda(1-\rho)$ and $\alpha_y = \gamma(1-\rho)$ for both countries, and $\alpha_q = \delta(1-\rho)$ for the foreign country.[7]

The most direct way to derive a forecasting equation is to postulate that the expected rate of depreciation is proportional to the interest rate differential:

(7) $$E(\Delta s_{t+1}) = \omega(i_t - \tilde{i}_t)$$

The variable s_t is the log of the nominal exchange rate determined as the domestic price of foreign currency. Assuming that UIRP holds, ω = 1 and (6) can be substituted into (7) to produce a forecasting equation.

(8) $$\Delta s_{t+1} = \alpha + \alpha_{u\pi}\pi_t - \alpha_{f\pi}\tilde{\pi}_t + \alpha_{uy}y_t - \alpha_{fy}\tilde{y}_t - \alpha_q\tilde{z}_t + \rho_u i_{t-1} - \rho_f\tilde{i}_{t-1} + \eta_t$$

There are two problems with this specification. First, UIRP does not hold in the short run, so we would not expect the coefficients in (8) to match the estimated Taylor rules. Based on empirical work on UIRP and (more recently) carry trade, it is not even clear whether or not ω, which equals one by UIRP, is positive or negative. Second, there is very strong evidence that interest rates do not completely adjust to their target levels within the period. Suppose that the U.S. inflation rate (actual or forecasted) rises above its target. This causes the Fed to raise the interest rate but also creates an expectation that the Fed will further raise the interest rate in the future. Since the increase in the interest rate may or may not cause expected depreciation of the exchange rate, and the expectation of further increases in the interest rate may cause expected

[7] 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.

appreciation of the exchange rate, we do not have a strong prior that even the signs of the coefficients in (8) are correct. Since a similar logic applies to an increase in the foreign inflation rate above its target, as well as to the other variables, we estimate Equation (8) as an exchange rate forecasting equation with Taylor rule fundamentals without restricting the signs or magnitudes of the coefficients.[8]

A number of different models can be nested in Equation (8). If the foreign central bank doesn't target the exchange rate $\delta = \alpha_q = 0$ and we call the specification symmetric. Otherwise, it is asymmetric. If the interest rate adjusts to its target level within the period $\rho_u = \rho_f = 0$ 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 foreign country, so that $\alpha_{u\pi} = \alpha_{f\pi}$, $\alpha_{uy} = \alpha_{fy}$, and $\rho_u = \rho_f$, inflation, output gap, and lagged interest rate differentials are on the right-hand-side of Equation (8) and we call the model homogeneous. Otherwise, it is heterogeneous. Finally, if the coefficients on inflation, interest rate smoothing coefficients, inflation targets, and equilibrium real interest rates are the same between the U.S. and the foreign country, $\alpha = 0$. Otherwise, a constant term is included in Equation (8).

2.2 Interest Rate Fundamentals

Under the UIRP condition (7), the expected change in the log exchange rate is equal to the nominal interest rate differential. If we were willing to assume that UIRP held, we could use (7) as a forecasting equation. Since empirical evidence indicates that, while exchange rate movements may be consistent with UIRP in the long-run, it clearly does not hold in the short-run, we need a more flexible specification.[9] Following Clark and West (2006), we use the interest rate differential as fundamentals in a forecasting equation,

$$\Delta s_{t+1} = \alpha + \omega(i_t - \tilde{i}_t) \tag{9}$$

Since we do not restrict $\omega = 1$, or even its sign, (9) is consistent with UIRP, where a positive interest rate differential produces forecasts of exchange rate depreciation, and the carry trade literature, where a positive interest rate differential produces forecasts of exchange rate appreciation.[10]

2.3 Monetary Fundamentals

Following Mark (1995), most widely used approach to evaluating exchange rate models out of sample is to represent a change in (the logarithm of) the nominal exchange rate as a function of its deviation from its fundamental value. Thus, the h-period-ahead change in the log exchange rate can be modeled as a function of its current deviation from its fundamental value.

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[8] Clarida and Waldman (2007) construct a model that combines a Taylor rule with a Phillips curve to derive conditions under which a surprise increase in U.S. inflation will appreciate the exchange rate, and use event study methodology to test the model.

[9] See Chinn and Meredith (2004) and Chinn (2006).

[10] See Burnside, Eichenbaum, Kleshchelski, and Rebelo (2006) for a discussion of carry trade.

(10) $$s_{t+h} - s_t = \alpha_h + \beta_h z_t + \nu_{t+h,t},$$

where $$z_t = f_t - s_t$$

and f_t is the long-run equilibrium level of the nominal exchange rate determined by macroeconomic fundamentals.

We select the flexible-price monetary model as representative of 1970's vintage models. The monetary approach determines the exchange rate as a relative price of the two currencies, and models exchange rate behavior in terms of relative demand for and supply of money in the two countries. The long-run money market equilibrium in the domestic and foreign country is given by:

(11) $$m_t = p_t + ky_t - hi_t$$

(12) $$m_t^* = p_t^* + k^* y_t^* - h^* i_t^*,$$

where m_t, p_t, and y_t are the logs of money supply, price level and income and i_t is the level of interest rate in period t; asterisks denote foreign country variables.

The monetary model assumes purchasing power parity, according to which exchange rates in the two countries will move to balance the prices:

(13) $$s_t = p_t - p_t^*,$$

where s_t is the log of nominal exchange rate determined as the domestic price of foreign currency.

Subtracting equation (12) from equation (11), using the PPP condition (13) to solve for the exchange rate, and assuming that countries are homogenous in terms of income elasticities and interest rate semi-elasticities of money supplies, we obtain the following equation,

(14) $$s_t = (m_t - m_t^*) - k(y_t - y_t^*) + h(i_t - i_t^*)$$

According to equation (14), an increase in the money supply differential between the domestic and foreign countries leads to a depreciation of the domestic currency. On the other hand, an increase in relative income of the domestic country creates additional demand for domestic money. In response to this, domestic residents cut their consumption forcing prices to fall, which leads through PPP to an appreciation of the domestic currency.

Assuming that UIRP holds and substituting the one-period-ahead interest rate differential in (14), the following expression for the exchange rate is obtained,

(15) $$s_t = (m_t - m_t^*) - k(y_t - y_t^*) + hE(\Delta s_{t+1})$$

Iterating equation (15) forward and assuming no rational speculative bubbles, the fundamental value of the exchange rate is

(16) $$f_t = (m_t - m_t^*) - k(y_t - y_t^*)$$

We construct the monetary fundamentals with a fixed value of the income elasticity, k, which can equal to 0, 1, or 3. We substitute the monetary fundamentals (16) into (10), and use the resultant equation for forecasting.

2.4 Purchasing Power Parity Fundamentals

As a basis of comparison, we examine the predictive power of PPP fundamentals. There has been extensive research on PPP in the last decade, and a growing body of literature finds that long-run PPP holds in the post-1973 period[11]. Since the monetary model is build upon PPP but assumes additional restrictions, comparing the out-of-sample performance of the two models is a logical exercise. Mark and Sul (2001) use panel-based forecasts and find evidence that the linkage between exchange rates and monetary fundamentals is tighter than that between exchange rates and PPP fundamentals.

Under PPP fundamentals,

(17) $$f_t = (p_t - p_t^*)$$

where p_t is the log of the national price level. We use the CPI as a measure of national price levels. We substitute the PPP fundamentals (17) into (10), and use the resultant equation for forecasting.

3. Forecast Comparison

3.1 Data

The models are estimated using monthly data from March 1973, the beginning of the floating exchange rate period, through December, 1998 for European Monetary Union countries and June, 2006 for the rest of the countries[12]. The currencies we consider are the Japanese yen, Swiss franc, Australian dollar, Canadian dollar, British pound, Swedish kronor, Danish kroner, Deutsche mark, French franc, Italian lira, Dutch guilder, and Portuguese escudo. The exchange rates defined as the US dollar price of a unit of foreign currency are taken from the Federal Reserve Bank of Saint Louis database.[13]

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[11] See Papell (2006) for a recent example.

[12] Some of the models are estimated using shorter spans of data because of data unavailability. The footnotes for the tables list these exceptions.

[13] The UIRP model is estimated using data from January, 1975 for Canada, September, 1975 for Switzerland and January, 1983 for Portugal because the interest rate data are not available prior to those periods. The monetary model is estimated using the data from December 1977 for France, December 1974 for Italy, and December 1979 for Portugal because of the lack of money supply data prior to those periods. Also, the money supply series end in December, 2004 for Sweden and April, 2006 for United Kingdom.

The primary source of data used to construct macroeconomic fundamentals is the IMF's *International Financial Statistics* (IFS) database[14]. We use M1 to measure the money supply for most of the countries. We use M0 for the U.K. and M2 for Italy and Netherlands, because M1 data is unavailable for these countries. Using M2 as a measure of the money supply provides similar results. We use the seasonally adjusted industrial production index (IFS line 66) as a proxy for countries' national income since GDP data are available only at the quarterly frequency.[15] The price level in the economy is measured by consumer price index (IFS line 64). The inflation rate is the annual inflation rate, measured as the 12-month difference of the CPI.[16] We use money market rate (or "call money rate", IFS line 60B) as a measure of the short-term interest rate that the central bank sets every period. Our choice of countries reflects our intention to examine exchange rate behavior for major industrialized economies with flexible exchange rates over the sample.

The output gap depends on the measure of potential output. Since there is no presumption about which definition of potential output is used by central banks in their interest rate reaction functions, we consider percentage deviations of actual output from a linear time trend, a quadratic time trend, and a Hodrick-Prescott (1997) (HP) trend as alternative definitions.[17] In order to mimic as closely as possible the information available to the central banks at the time the decisions were made, we use quasi-real time data in the output gap estimation. For a given period t, we use only the data points up to t-1 to construct the trend. Thus, in each period the OLS regression is re-estimated adding one additional observation to the sample.[18]

3.2 Estimation and Forecasting

We construct 1- to 36-month ahead forecasts for the linear regression models with each of the fundamentals described above. We use data over the period March 1973 - February 1982 for estimation and reserve the remaining data for out-of-sample forecasting. Let us concentrate for simplicity on one-step-ahead predictions. To evaluate the out-of-sample performance of the models, we estimate them by OLS in rolling regressions and construct CW statistics. Each model is initially estimated using the first 120 data points and the one-period-ahead forecast is generated. Then, we drop the first data point, add an additional data point at the end of the sample, and re-estimate the model. A one period-ahead forecast is generated at each step.[19]

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[14] The complete Data Appendix and data files are available at the author's web-site: www.uh.edu/~dpapell

[15] The industrial production series for Australia and Switzerland, and the CPI series for Australia, which were available only quarterly, are transformed into monthly periodicity using the "quadratic-match average" option in Eviews 4.0.

[16] An important focus of Taylor rule estimation for the U.S. has been the forward-looking nature of policymaking, either by using *ex post* realized values of inflation as in CGG or by using Greenbook forecasts as in Orphanides (2001). Since, for the purpose of evaluating out-of-sample predictability, it is inappropriate to use *ex post* data and central bank forecasts are not available for other countries, we use actual inflation rates.

[17] While it would be desirable, following Orphanides (2001) for the U.S., to use central bank generated estimates of the output gap, these are neither available for our entire sample nor available for other countries.

[18] We call this quasi-real time data because, while the trend is updated each period, the data incorporate revisions that were not available to the central banks at the time decisions were make. True real time data is not available for most of the countries that we study over the entire floating rate period.

[19] We use out-of-sample rather than in-sample methods and estimate rolling rather than recursive regressions for comparison with the extensive literature following Meese and Rogoff (1983a), and choose a rolling window of 120 observations to estimate alternative forecast models following the empirical exercise in Clark and West (2006). Inoue

3.3 Forecast Comparison Based on MSPE

Each model's out-of-sample predictability is compared to that of the martingale difference process using an adjusted test statistic, which is constructed as described in Clark and West (2006). We are interested in comparing the mean square prediction errors from the two nested models. The first model is a zero mean martingale difference process, while the other 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*, *R*=120, *P*=260 for non-EU countries and 190 for EU countries. 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 MSPE's 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 are interested in testing the null hypothesis of no predictability against the alternative that exchange rates are linearly predictable.[20] Thus,

$$H_0 : \sigma_1^2 - \sigma_2^2 = 0$$

$$H_1 : \sigma_1^2 - \sigma_2^2 > 0$$

Under the null, the population MSPE's are equal. We need to use the sample estimates of the population MSPE's to draw the inference. The procedure introduced by Diebold and Mariano (1995) and West (1996) uses sample MSPE's to construct a t-type statistics which is assumed to be asymptotically normal. To construct the DMW statistic, let

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and Kilian (2004) advocate using in-sample rather than out-of-sample methods and using recursive methods for out-of-sample forecasting.

[20] 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.

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

Then, the DMW test statistics is computed as follows,

$$(18) \qquad DMW = \frac{\bar{f}}{\sqrt{P^{-1}\hat{V}}}, \qquad \text{where} \qquad \hat{V} = P^{-1} \sum_{t=T-P+1}^{T} (\hat{f}_{t+1} - \bar{f})^2$$

Clark and West (2006) demonstrate analytically that the asymptotic distributions of sample and population difference between the two MSPE's are not identical, namely the sample difference between the two MSPE's is biased downward from zero. This means that using the test statistic (18) with standard normal critical values is not advisable.

It is straightforward to show that the sample difference between the two MSPE's is uncentered under the null.

$$(19) \quad \hat{\sigma}_1^2 - \hat{\sigma}_2^2 = P^{-1} \sum_{t=T-P+1}^{T} \hat{f}_{t+1} = P^{-1} \sum_{t=T-P+1}^{T} y_{t+1}^2 - P^{-1} \sum_{t=T-P+1}^{T} (y_{t+1} - X'_{t+1}\hat{\beta}_t)^2 = 2P^{-1} \sum_{t=T-P+1}^{T} y_{t+1} X'_{t+1}\hat{\beta}_t - P^{-1} \sum_{t=T-P+1}^{T} (X'_{t+1}\hat{\beta}_t)^2$$

Under the null, the first term in (19) 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. In finite samples, use of the noisy estimate of the parameters will lead to higher estimated MSPE. As a result, the sample MSPE of the alternative model will be higher by the amount of estimation noise.

To properly adjust for this shift, 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 (19). 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 MSPE's as described above and using standard normal critical values for inference.[21]

It is important to understand the distinction between predictability and forecasting content. The CW methodology tests whether the regression coefficient β is zero rather than whether the model-based forecast is more accurate than the random walk forecast. Since the CW statistic is constructed by adjusting the sample MSPE from the alternative model by the amount of bias under the null, it is entirely possible for the null

[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, event though the statistics are non-normal according to Clark and McCracken (2001), and advocate using them instead of bootstrapped critical values.

hypothesis that $\beta = 0$ to be rejected even when the sample MSPE from the random walk forecast is smaller than the sample MSPE from the model-based forecast.

4. Empirical Results

4.1 Predictability with the CW Statistic

To illustrate how the CW statistic is constructed, consider the following exercise. Suppose we estimate the symmetric Taylor rule model in Table 1 with no smoothing, homogeneous coefficients, no constant, and the output gap defined as a deviation from a linear time trend. Let us pick Canada for illustration purposes. The unadjusted MSPE of the model is 8.10, which is higher than the MSPE of the random walk of 7.83. The DMW statistic would be negative. To obtain a test statistic that is centered around zero, we need to adjust the MSPE of the economic model downward. This adjustment term, $P^{-1} \sum_{t=T-P+1}^{T} (X'_{t+1}\hat{\beta}_t)^2$, is equal to 0.52 in our example. After subtracting this adjustment term from the MSPE of the economic model, we obtain the adjusted MSPE which is smaller that that of the random walk and equal to 7.58. The CW statistic is positive with a t-statistic equal to 1.45. This means that the test based on adjusted difference rejects the null of no exchange rate predictability with the Taylor rule model at the 10% significance level.

4.2 Taylor Rule Fundamentals

With a choice between symmetric and asymmetric, homogeneous and heterogeneous, with and without smoothing, and with and without a constant, we estimate 16 models with three measures of the output gap, for a total of 48 models for each country.[22] Two overall results are apparent. First, models with heterogeneous coefficients provide stronger evidence of exchange rate predictability in all eight cases. Second, models with a constant provide stronger evidence of exchange rate predictability in six of the eight cases. We therefore focus on the models with heterogeneous coefficients that include a constant. Table 1 presents the results for 1-month-ahead forecasts of exchange rates using asymmetric Taylor rule fundamentals with no smoothing, with linear, quadratic and HP trends to estimate potential output. The model significantly outperforms the random walk for 4 out of 12 countries with a linear trend (Italy at the 1% significance level, Canada and Sweden at the 5% significance level, and the U.K. at the 10% significance level), for 2 out of 12 countries with a quadratic trend (Italy at the 1% and Canada at the 5% significance level), and for 7 out of 12 countries with an HP trend (Canada at the 1%, Italy, Japan, Sweden, and the U.K. at the 5%, and Netherlands and Switzerland at the 10% significance level). The model significantly outperforms the random walk in 13 out of 36 cases and with at least one of the output gap specifications for 7 out of 12 countries.

[22] With heterogeneous coefficients, it would require a particular combination of coefficients, target inflation rates, and equilibrium real interest rates for the terms that comprise the constant to cancel out. Nevertheless, the constant could be small if the smoothing coefficients were large, and so we include the heterogeneous model without a constant.

Table 2 depicts the results for the asymmetric Taylor rule model with smoothing. The model significantly outperforms the random walk for 4 out of 12 countries with a linear trend (Italy at the 1% significance level, Canada and Japan at the 5% significance level, and Australia at the 10% significance level), for 6 out of 12 countries with a quadratic trend (Canada, Italy, and Japan at the 1%, and Netherlands, Switzerland, and the U.K. at the 10% significance level), and for 8 out of 12 countries with an HP trend (Italy and Japan at the 1%, Canada, Netherlands, and Switzerland at the 5%, and Australia, France, and the U.K. at the 10% significance level). The model significantly outperforms the random walk in 18 out of 36 cases and with at least one of the output gap specifications for 8 out of 12 countries.

Short-term predictability increases when we use the Taylor rule where the foreign country does not target the exchange rate. Table 3 presents the results for the symmetric Taylor rule model with no smoothing. The model with Taylor rule fundamentals significantly outperforms the random walk for 8 out of 12 countries with a linear trend (Canada at the 1% significance level, Australia, Denmark, France, Italy, Sweden, and the U.K. at the 5% significance level, and Germany at the 10% significance level), for 6 out of 12 countries with a quadratic trend (Canada and Italy at the 1%, France, Germany, and the U.K. at the 5%, and Switzerland at the 1% significance level), and for 6 out of 12 countries with an HP trend (Canada at the 1%, France, Italy, Sweden, and the U.K. at the 5%, and Switzerland at the 10% significance level). The model significantly outperforms the random walk in 20 out of 36 cases and with at least one of the output gap specifications for 9 out of 12 currencies.

The strongest results are found for the symmetric Taylor rule model with smoothing. As depicted in Table 4, the model with Taylor rule fundamentals significantly outperforms the random walk for 10 out of 12 countries with a linear trend (Canada and Italy at the 1% significance level, Australia, France, Japan, Netherlands, and the U.K. at the 5% significance level, and Denmark, Germany, and Switzerland at the 10% significance level), for 9 out of 12 countries with a quadratic trend (Canada, Italy, and Japan at the 1%, Australia, France, Germany, Netherlands, and the U.K. at the 5%, and Switzerland at the 1% significance level), and for 9 out of 12 countries with an HP trend (France, Italy, and Netherlands at the 1%, Australia, Canada, Denmark, Japan, Switzerland, and the U.K. at the 5% significance level). The model significantly outperforms the random walk in 28 out of 36 cases and with at least one of the output gap specifications for 10 out of 12 currencies.[23]

Combining the four Taylor rule models, evidence of short-term predictability is found for 11 out of 12 countries, five countries at the 1% level and six additional countries at the 5% level. No evidence of predictability is found for Portugal. More evidence is found with symmetric models than with asymmetric models and with models with smoothing than with models with no smoothing. Overall, the strongest results

[23] We investigate robustness of the results by splitting the sample in half. The symmetric specification with heterogeneous coefficients and a constant, but no smoothing, provides the strongest evidence of predictability in both subsamples. There is evidence of predictability in each subsample, which is relatively stronger in the earlier subsample.

are found with the symmetric Taylor rule model with heterogeneous coefficients, smoothing, and a constant. For that model alone, evidence of short-term predictability is found for 10 out of 12 countries, four countries at the 1% level and six additional countries at the 5% level.[24]

4.3 Interest Rate, Monetary, and PPP Fundamentals

Table 5 contains the results for 1-month-ahead forecasts of the exchange rates using the interest rate, monetary, and PPP fundamentals described in Section 2. We do not find much evidence of exchange rate predictability with any of the models. The strongest evidence comes from interest rate fundamentals with a constant, where the model significantly outperforms the random walk for 4 out of 12 countries (Japan at the 1% significance level, Switzerland at the 5% significance level, and Australia and Canada at the 10% significance level). Without a constant, the model with interest rate fundamentals significantly outperforms the random walk for 2 countries (Australia and Canada at the 10% significance level).

The evidence is weaker for monetary fundamentals. With the coefficient on relative output k equal to 0, the model significantly outperforms the random walk for 2 out of 12 countries with a constant (Canada and Japan at the 5% significance level) and 1 country without a constant (Japan at the 10% significance level). The evidence with $k = 1$ and $k = 3$ is weaker with a constant and the same without a constant. The weakest evidence is found with PPP fundamentals, where the model significantly outperforms the random walk for 1 country (Japan at the 10% significance level) without a constant and for no countries with a constant.

4.4 Testing for Superior Predictive Ability

Since we are simultaneously testing multiple hypotheses, inference based on conventional p-values is likely to be contaminated. This issue arises because we have 58 different models of 12 bilateral exchange rates yielding 696 test statistics. As a result of an extensive specification search, it is possible to mistake the results that could be generated by chance for genuine evidence of predictive ability. To increase the reliability of our results, 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.[25]

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

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[24] Engel, Mark, and West (2007), using a specification of Taylor rule fundamentals from an earlier version of this paper, find little evidence of predictability. They use an asymmetric model with no smoothing, a constant, homogeneous coefficients, and HP filtered output which, in Table 1, produces only four rejections at the 5 percent level. In addition, they impose $\phi = \gamma = 0.5$ for both countries and $\delta = 0.1$ for the foreign country, which further restricts the forecasts.

[25] Hansen (2005) provides details on the construction of the test statistic and confirms the advantages of the test by Monter 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

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. Therefore, rejecting the null indicates that at least one linear model is strictly superior to the random walk. Tables 6-8 report the SPA p-values that 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.

The SPA test is designed to guard against "evidence" of predictability obtained by estimating a large number of models and focusing on the one with the most significant results. With Taylor rule fundamentals, the most arbitrary choice is the measure of the output gap, and we need to evaluate how estimating models with linear, quadratic, and HP detrending for each specification affects our evidence of predictability. The Taylor rule specifications themselves, in contrast, are not arbitrary. The choice among constant/no constant, homogeneous/heterogeneous, symmetric/asymmetric, and smoothing/no smoothing are guided by economic theory and previous empirical research.

Table 6 reports the results for the 16 Taylor rule specifications, where the benchmark model is the random walk and the alternatives are the three output gap measures. The SPA p-values strongly confirm the results in Tables 1-4. Combining the 16 models, evidence of short-term predictability is again found for 11 out of 12 countries (Canada at the 1% significance level, Australia, France, Italy, Japan, Netherlands, Sweden, and the U.K. at the 5% significance level, and Denmark, Germany, and Switzerland at the 10% significance level). The models with heterogeneous coefficients provide more evidence of exchange rate predictability than the models with homogeneous coefficients and the models with a constant provide more evidence of predictability than the models without a constant, with the most evidence provided by models with both heterogeneous coefficients and a constant. As above, the strongest results are found with the symmetric Taylor rule model with heterogeneous coefficients, smoothing, and a constant. For that model alone, evidence of short-term predictability is again found for 10 out of 12 countries (Canada at the 1% significance level, Australia, France, Italy, Japan, and Netherlands at the 5% significance level, and Denmark, Germany, Switzerland, the U.K. at the 10% significance level). While, as expected, the SPA p-values are higher than the most significant single-output-gap p-values, the results show that the evidence of exchange rate predictability reported above is not an artifact of picking the output gap specification with the lowest p-value for each model.

Table 7 reports SPA p-values with a larger set of alternatives for the Taylor rule specifications with heterogeneous coefficients and a constant. While these specifications are the ones for which the most evidence of predictability was found, there seems to be no compelling reason to think that the Fed and

foreign central banks followed the same quantitative interest rate reaction function in response to inflation and output deviations, much less, in addition, had the same inflation targets and equilibrium real interest rates. The first four columns test the random walk benchmark against six alternatives. For example, "symmetric" would denote smoothing and no smoothing for the three output gap measures. The SPA p-values again confirm our previous results. Combining the 4 models, evidence of short-term predictability is found for 10 out of 12 countries (Canada, France, Italy, Japan, and Netherlands at the 5% significance level and Australia, Denmark, Sweden, Switzerland, and the U.K. at the 10% significance level). The symmetric models provide more evidence of out-of-sample exchange rate predictability than the asymmetric models (9 versus 3 out of 12 countries at the 10 percent significance level or higher) and the models with smoothing provide more evidence of predictability than the models with no smoothing (9 versus 4 out of 12 countries at the 10 percent significance level or higher). The fifth column, denoted "all", tests the random walk benchmark against 12 alternatives: symmetric with smoothing, symmetric with no smoothing, asymmetric with smoothing, and asymmetric with no smoothing for the three output gap measures. While, as expected, the SPA p-values decline with the inclusion of the asymmetric and no smoothing specifications, evidence of short-term exchange rate predictability is found for 7 out of 12 countries (Canada and Japan at the 5 percent significance level and Australia, France, Italy, Netherlands, and Sweden at the 10 percent significance level).

For the purpose of comparison, Table 8 reports SPA p-values for the interest rate, PPP, and monetary models. There are two alternatives for the interest rate and PPP models, with and without a constant, and six alternatives for the monetary models, $k=0$, $k=1$, and $k=3$ with a constant and no constant. Evidence of short-run exchange rate predictability is found for 3 out of 12 countries with the interest rate model (Canada and Japan at the 5 percent significance level and Switzerland at the 10 percent significance level), one country (Canada at the 10 percent significance level) for the monetary model, and no countries for the PPP model. This is in accord with the results reported in Table 5, and provides further confirmation that the evidence of short-run out-of-sample exchange rate predictability is much stronger for models with Taylor rule fundamentals than for conventional models.

5. Conclusions

Research on exchange rate predictability has come full circle from the "no predictability at short horizons" results of Meese and Rogoff (1983a, 1983b) to the "predictability at long horizons but not short horizons" results of Mark (1995) and Chen and Mark (1996) to the "no predictability at any horizons" results of Cheung, Chinn, and Pascual (2005). We come to a very different conclusion, reporting strong evidence of out-of-sample exchange rate predictability at the one-month horizon for 12 OECD countries vis-à-vis the United States over the post-Bretton Woods period.

We find very strong evidence of exchange rate predictability with Taylor rule fundamentals. Using the CW statistic, we reject the no predictability null hypothesis at the 5 percent level for 11 of 12 countries.

While, as expected, the significance level falls when we calculate p-values using Hansen's (2005) test of superior predictive ability to control for estimating multiple specifications, we still find substantial evidence of predictability. The strongest results are found with the symmetric Taylor rule model with heterogeneous coefficients, smoothing, and a constant. The result that predictability increases when the coefficients are not restricted to be identical between the United States and the foreign country and when interest rate smoothing is incorporated is consistent with evidence from estimation of Taylor rules. We find much less evidence of short-run predictability using models with interest rate, monetary, and PPP fundamentals.

These results suggest a number of directions for future research. Engel, Mark, and West (2007) use the CW statistic and find some evidence of predictability for a variety of models. Their evidence is stronger for panel data than single-equation models for monetary and PPP fundamentals, with the opposite result for Taylor rule fundamentals, and is stronger with 16-quarter-ahead than with one-quarter-ahead forecasts. Molodtsova, Nikolsko-Rzhevskyy, and Papell (2007) estimate Taylor rules using real-time data for Germany and the United States, and find strong evidence of predictability of exchange rate changes at the one-quarter horizon using real-time, but not revised, data. Molodtsova (2007) uses real-time OECD data, available starting in 1999, to evaluate short-horizon exchange rate predictability with Taylor rule fundamentals for 9 OECD currencies, plus the Euro, vis-à-vis the U.S. dollar and finds strong evidence of exchange rate predictability at the 1-month horizon for 8 out of 10 exchange rates. As in this paper, the strongest results are found with a symmetric Taylor rule model with heterogeneous coefficients, smoothing, and a constant.

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Table 1. Asymmetric Taylor Rule Model with No Smoothing

Country	*Linear Trend*	*Quadratic Trend*	*HP Filter*	*Linear Trend*	*Quadratic Trend*	*HP Filter*
	w/o Constant			*w/ Constant*		
A. Homogenous Coefficients						
Australia	0.411	0.734	0.578	0.679	0.700	0.728
Canada	0.018**	0.036**	0.095*	0.036**	0.023**	0.025**
Denmark	0.329	0.464	0.895	0.432	0.493	0.831
France	0.319	0.348	0.098*	0.248	0.277	0.021**
Germany	0.504	0.396	0.255	0.798	0.624	0.444
Italy	0.132	0.045**	0.025**	0.013**	0.005***	0.013**
Japan	0.058*	0.190	0.072*	0.174	0.195	0.084*
Netherland	0.675	0.673	0.780	0.259	0.288	0.458
Portugal	0.586	0.434	0.458	0.821	0.733	0.741
Sweden	0.295	0.245	0.476	0.493	0.295	0.521
Switzerland	0.746	0.740	0.257	0.899	0.747	0.280
U.K.	0.234	0.158	0.023**	0.215	0.188	0.011**
B. *Heterogenous Coefficients*						
Australia	0.218	0.445	0.612	0.184	0.385	0.496
Canada	0.008***	0.006***	0.010***	0.023**	0.015**	0.003***
Denmark	0.062*	0.236	0.343	0.109	0.284	0.126
France	0.066**	0.049**	0.026**	0.179	0.125	0.303
Germany	0.099*	0.185	0.822	0.121	0.188	0.303
Italy	0.016**	0.007***	0.024**	0.004***	0.002***	0.018**
Japan	0.435	0.477	0.133	0.550	0.404	0.026**
Netherlands	0.172	0.123	0.268	0.325	0.250	0.089*
Portugal	0.940	0.875	0.331	0.562	0.555	0.661
Sweden	0.024**	0.175	0.052*	0.043**	0.269	0.034**
Switzerland	0.354	0.254	0.114	0.226	0.173	0.081*
U.K.	0.102	0.207	0.141	0.082*	0.101	0.039**

Notes to Tables 1 through 4:

1. The models without interest rate smoothing are estimated using data from March, 1973 for all countries. The models that include lagged interest rates are estimated from January, 1975 for Canada, September, 1975 for Switzerland and January, 1983 for Portugal.
2. To detrend the monthly output series using HP filter, we use the value of smoothing parameter equal to 14400.
3. The output gap in period t is calculated using all the currently available data. The output gap for the first period is calculated using output series from 1971:1 to 1973:3.
4. The p-values for a one-sided test for equal predictive ability of the two models are based on standard normal critical values.
5. In all of the tables *, **, and *** denote test statistics significant at 10, 5, and 1% level, respectively, based on critical values for one-sided test.

Table 2. Asymmetric Taylor Rule Model with Smoothing

Country	*Linear Trend*	*Quadratic Trend*	*HP Filter*	*Linear Trend*	*Quadratic Trend*	*HP Filter*
	w/o Constant			*w/ Constant*		
	A. Homogenous Coefficients					
Australia	0.171	0.244	0.192	0.103	0.143	0.169
Canada	0.028**	0.055*	0.178	0.086*	0.074**	0.028**
Denmark	0.209	0.281	0.505	0.440	0.452	0.635
France	0.040**	0.037**	0.010***	0.080*	0.060*	0.013**
Germany	0.459	0.462	0.647	0.562	0.423	0.728
Italy	0.005***	0.003***	0.004***	0.005***	0.003***	0.005***
Japan	0.012**	0.004***	0.059*	0.029**	0.006***	0.076*
Netherlands	0.262	0.271	0.285	0.142	0.153	0.216
Portugal	0.462	0.430	0.238	0.919	0.906	0.637
Sweden	0.824	0.744	0.818	0.852	0.836	0.820
Switzerland	0.277	0.202	0.134	0.283	0.150	0.161
U.K.	0.354	0.291	0.146	0.342	0.307	0.064*
	B. *Heterogenous Coefficients*					
Australia	0.096*	0.151	0.174	0.054*	0.118	0.087*
Canada	0.011**	0.003***	0.076*	0.042**	0.005***	0.011**
Denmark	0.089*	0.187	0.059*	0.374	0.439	0.110
France	0.032**	0.027**	0.022**	0.195	0.125	0.056*
Germany	0.282	0.456	0.908	0.402	0.349	0.578
Italy	0.002***	0.001***	0.001***	0.009***	0.007***	0.009***
Japan	0.016**	0.002***	0.010***	0.042**	0.005***	0.004***
Netherlands	0.036**	0.016**	0.012**	0.156	0.083*	0.034**
Portugal	0.979	0.971	0.905	0.981	0.969	0.844
Sweden	0.718	0.861	0.496	0.882	0.919	0.773
Switzerland	0.173	0.147	0.055*	0.170	0.081*	0.032**
U.K.	0.172	0.280	0.176	0.108	0.095*	0.060*

Table 3. Symmetric Taylor Rule Model with No Smoothing

Country	*Linear Trend*	*Quadratic Trend*	*HP Filter*	*Linear Trend*	*Quadratic Trend*	*HP Filter*
	w/o Constant			*w/ Constant*		
A. Homogenous Coefficients						
Australia	0.536	0.615	0.588	0.155	0.383	0.478
Canada	0.048**	0.019**	0.140	0.006***	0.017**	0.025**
Denmark	0.381	0.371	0.772	0.325	0.462	0.915
France	0.174	0.852	0.234	0.406	0.447	0.152
Germany	0.591	0.599	0.400	0.555	0.412	0.210
Italy	0.098*	0.083*	0.025**	0.137	0.044**	0.026**
Japan	0.425	0.174	0.409	0.063*	0.198	0.076*
Netherlands	0.582	0.815	0.989	0.674	0.690	0.804
Portugal	0.945	0.724	0.893	0.516	0.379	0.411
Sweden	0.283	0.321	0.466	0.299	0.228	0.480
Switzerland	0.731	0.965	0.597	0.565	0.521	0.280
U.K.	0.460	0.468	0.030**	0.313	0.227	0.034**
B. Heterogenous Coefficients						
Australia	0.131	0.126	0.605	0.020**	0.140	0.328
Canada	0.003***	0.006***	0.003***	0.003***	0.003***	0.002***
Denmark	0.113	0.473	0.779	0.050**	0.185	0.305
France	0.861	0.289	0.608	0.048**	0.033**	0.023**
Germany	0.350	0.057*	0.788	0.052*	0.041**	0.211
Italy	0.115	0.014**	0.057*	0.015**	0.005***	0.020**
Japan	0.425	0.256	0.423	0.412	0.337	0.120
Netherlands	0.584	0.252	0.980	0.174	0.111	0.214
Portugal	0.736	0.574	0.064*	0.939	0.846	0.291
Sweden	0.114	0.198	0.271	0.020**	0.113	0.049**
Switzerland	0.552	0.264	0.510	0.129	0.078*	0.080*
U.K.	0.023**	0.051*	0.004***	0.020**	0.036**	0.020**

Table 4. Symmetric Taylor Rule Model with Smoothing

Country	*Linear Trend*	*Quadratic Trend*	*HP Filter*	*Linear Trend*	*Quadratic Trend*	*HP Filter*
	w/o Constant			*w/ Constant*		
A. Homogenous Coefficients						
Australia	0.059*	0.070*	0.114	0.027**	0.052*	0.043**
Canada	0.024**	0.037**	0.064**	0.030**	0.036**	0.061*
Denmark	0.299	0.269	0.722	0.325	0.277	0.536
France	0.197	0.242	0.089*	0.075*	0.075*	0.017**
Germany	0.754	0.860	0.790	0.200	0.110	0.215
Italy	0.041**	0.016**	0.007***	0.006***	0.004***	0.004***
Netherlands	0.367	0.457	0.533	0.302	0.327	0.327
Japan	0.005***	0.022**	0.125	0.013**	0.002***	0.081*
Portugal	0.469	0.342	0.186	0.421	0.349	0.194
Sweden	0.809	0.792	0.831	0.768	0.783	0.800
Switzerland	0.244	0.149	0.171	0.138	0.084*	0.099*
U.K.	0.519	0.505	0.197	0.421	0.361	0.132
B. Heterogenous Coefficients						
Australia	0.025**	0.046**	0.070*	0.015**	0.035**	0.038**
Canada	0.005***	0.002***	0.022***	0.008***	0.002***	0.021**
Denmark	0.042**	0.111	0.215	0.069*	0.138	0.032**
France	0.188	0.069*	0.078*	0.024**	0.020**	0.008***
Germany	0.118	0.108	0.668	0.066*	0.039**	0.126
Italy	0.027**	0.008***	0.015**	0.002***	0.001***	0.001***
Japan	0.016**	0.011**	0.022**	0.019**	0.001***	0.011**
Netherlands	0.063*	0.074*	0.262	0.036**	0.015**	0.009***
Portugal	0.966	0.906	0.798	0.985	0.973	0.898
Sweden	0.775	0.832	0.770	0.678	0.812	0.593
Switzerland	0.202	0.116	0.109	0.094*	0.052*	0.016**
U.K.	0.074*	0.142	0.134	0.020**	0.021**	0.033**

Table 5. p- values for the Model with Interest Rate, PPP, and Monetary Fundamentals

Country	*Interest Rate*	*PPP*	*Monetary, k=0*	*Monetary, k=1*	*Monetary, k=3*	*Interest Rate*	*PPP*	*Monetary, k=0*	*Monetary, k=1*	*Monetary, k=3*
	w/o Constant					*w/ Constant*				
Australia	0.087*	0.414	0.378	0.377	0.378	0.061*	0.788	0.364	0.426	0.443
Canada	0.065*	0.295	0.396	0.415	0.446	0.026*	0.799	0.041**	0.029**	0.041*
Denmark	0.381	0.757	0.815	0.786	0.722	0.602	0.849	0.141	0.197	0.493
France	0.851	0.694	0.680	0.689	0.709	0.883	0.624	0.969	0.538	0.211
Germany	0.756	0.361	0.276	0.298	0.330	0.255	0.560	0.685	0.677	0.476
Italy	0.393	0.649	0.991	0.991	0.992	0.362	0.707	0.903	0.630	0.615
Japan	0.685	0.089*	0.063*	0.057*	0.051*	0.006***	0.103	0.039**	0.175	0.378
Netherlands	0.377	0.426	0.518	0.529	0.545	0.172	0.526	0.484	0.549	0.431
Portugal	0.132	0.915	0.520	0.571	0.638	0.272	0.989	0.388	0.304	0.197
Sweden	0.725	0.807	0.960	0.963	0.950	0.867	0.713	0.462	0.403	0.410
Switzerland	0.355	0.430	0.283	0.278	0.270	0.017**	0.783	0.177	0.191	0.246
U.K.	0.219	0.716	0.649	0.653	0.659	0.344	0.532	0.840	0.602	0.447

Notes:

1. The UIRP model is estimated using data from January, 1975 for Canada, September, 1975 for Switzerland and January, 1983 for Portugal because the interest rate data are not available prior to those periods. The monetary model is estimated using the data from December 1977 for France, December 1974 for Italy, and December 1979 for Portugal because of the lack of money supply data prior to those periods. Also, the monetary supply series end in December, 2004 for Sweden and April, 2006 for United Kingdom.
2. The p-values for a one-sided test for equal predictive ability of the two models are based on standard normal critical values.

Table 6. Tests for Superior Predictive Ability: Taylor Rule Models

Country	*No Smoothing*		*Smoothing*		*No Smoothing*		*Smoothing*	
	Sym	*Asym*	*Sym*	*Asym*	*Sym*	*Asym*	*Sym*	*Asym*
	w/o Constant				*w/ Constant*			
	A. Homogenous Coefficients							
Australia	0.739	0.562	0.106	0.234	0.285	0.794	0.128	0.154
Canada	0.069*	0.064*	0.055*	0.079*	0.028**	0.055*	0.031**	0.122
Denmark	0.616	0.504	0.398	0.294	0.506	0.575	0.298	0.514
France	0.369	0.186	0.170	0.037**	0.265	0.429	0.052*	0.050**
Germany	0.591	0.401	0.835	0.568	0.351	0.579	0.197	0.535
Italy	0.117	0.077*	0.035**	0.038**	0.081*	0.037**	0.039**	0.038**
Japan	0.332	0.154	0.025**	0.025**	0.163	0.229	0.013**	0.035**
Netherlands	0.683	0.829	0.474	0.343	0.835	0.363	0.387	0.198
Portugal	0.873	0.528	0.173	0.325	0.482	0.794	0.353	0.724
Sweden	0.497	0.407	0.821	0.784	0.384	0.466	0.805	0.828
Switzerland	0.854	0.436	0.259	0.242	0.418	0.536	0.170	0.255
U.K.	0.096*	0.053*	0.297	0.222	0.074*	0.033**	0.196	0.113
	B. *Heterogenous Coefficients*							
Australia	0.263	0.401	0.057*	0.152	0.069*	0.370	0.042**	0.103
Canada	0.020**	0.019**	0.012**	0.017**	0.009***	0.045**	0.008***	0.025**
Denmark	0.280	0.150	0.104	0.121	0.119	0.240	0.075*	0.211
France	0.517	0.061*	0.128	0.065*	0.067*	0.243	0.030**	0.130
Germany	0.132	0.217	0.194	0.419	0.099*	0.279	0.091*	0.507
Italy	0.081*	0.025**	0.032**	0.023**	0.019**	0.015**	0.022**	0.053*
Japan	0.469	0.115	0.062*	0.017**	0.272	0.736	0.017**	0.043**
Netherlands	0.353	0.174	0.113	0.036**	0.192	0.095*	0.028**	0.079*
Portugal	0.167	0.782	0.870	0.910	0.471	0.587	0.907	0.856
Sweden	0.296	0.067*	0.803	0.631	0.053*	0.048**	0.705	0.829
Switzerland	0.462	0.205	0.195	0.130	0.174	0.224	0.054*	0.098*
U.K.	0.015**	0.101	0.153	0.333	0.061*	0.165	0.058*	0.165

Notes:

1. The table reports SPA p-values for sixteen sets of Taylor-rule-based forecasts that are compared to a random walk forecast.
2. Panel A contains the results for homogenous Taylor rule fundamentals, that restrict coefficients on the inflation and output gap in the two countries to be the same, and Panel B contains the results for heterogenous Taylor rule models. The p-values are reported for the following classes of models: Sym, symmetric Taylor rule models, and Asym, asymmetric Taylor rule models, that are subdivided into Smoothing, and No Smoothing, models that include or exclude interest rate smoothing.

Table 7. Tests for Superior Predictive Ability: Heterogenous TR Models with a Constant

Country	*Sym*	*Asym*	*Smoothing*	*No Smoothing*	*All*
Australia	0.063*	0.173	0.059*	0.113	0.100*
Canada	0.013**	0.040**	0.012**	0.017**	0.020**
Denmark	0.102	0.304	0.095*	0.174	0.151
France	0.034**	0.191	0.043**	0.105	0.064*
Germany	0.120	0.372	0.123	0.157	0.190
Italy	0.063*	0.040**	0.036**	0.061*	0.064*
Japan	0.026**	0.070*	0.024**	0.457	0.042**
Netherlands	0.051*	0.128	0.037**	0.125	0.069*
Portugal	0.788	0.857	0.885	0.503	0.811
Sweden	0.060*	0.145	0.745	0.060*	0.087*
Switzerland	0.076*	0.167	0.073*	0.262	0.123
U.K.	0.081*	0.244	0.082*	0.084*	0.117

Notes:

1. The table reports SPA p-values for five sets of forecasts based on heterogenous Taylor rule fundamentals with a constant that are compared to a random walk forecast.
2. Each column contains the results for the following classes of models: All, all heterogeneous Taylor rule models with a constant, Sym, symmetric Taylor rule models, Asym, asymmetric Taylor rule models, Smoothing, and No Smoothing, models that include or exclude interest rate smoothing.

Table 8. Tests for Superior Predictive Ability: non-Taylor Rule Models

Country	*Interest Rate*	*PPP*	*Monetary*
Australia	0.126	0.508	0.515
Canada	0.049**	0.423	0.082*
Denmark	0.564	0.828	0.395
France	0.959	0.694	0.518
Germany	0.483	0.525	0.689
Italy	0.526	0.687	0.755
Japan	0.018**	0.222	0.114
Netherlands	0.333	0.577	0.673
Portugal	0.261	0.893	0.519
Sweden	0.802	0.779	0.778
Switzerland	0.088*	0.288	0.252
U.K.	0.306	0.672	0.683

Notes:

1. The table reports SPA p-values for three sets of non-Taylor-rule-based forecasts that are compared to a random walk forecast.
2. Each column contains the results for the following classes of models: Interest Rate, model with interest rate fundamentals, PPP, models with PPP fundamentals, and Monetary, models with monetary fundamentals.