The Uncovered Interest Parity Puzzle, Exchange Rate Forecasting, and Taylor Rules

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Abstract

Recent research has found that the Taylor-rule fundamentals have power to forecast changes in U.S. dollar exchange rates out of sample. Our work casts some doubt on that claim. However, we find strong evidence of a related in-sample anomaly. When we include U.S. inflation in the well-known uncovered interest parity regression of the change in the exchange rate on the interest-rate differential, we find that the inflation variable is highly significant and the interest-rate differential is not. Specifically, high U.S. inflation in one month forecasts dollar appreciation in the subsequent month. We introduce a model in which a Taylor rule determines monetary policy, but in which not only monetary shocks but also liquidity shocks drive nominal interest rates. This model can potentially account for the empirical findings.

Recent research has found consistent success forecasting changes in nominal exchange rates, especially for the U.S. dollar relative to other high-income countries, using variables that help to determine monetary policy. This work posits that central bank policy can be described by a Taylor rule that determines the short-term nominal interest rate as a function of the inflation rate, the output gap and other variables. These variables have been dubbed the "Taylor rule fundamentals" and evidence has been amassed that these variables can be used to produce forecasts of the exchange rate that outperform the random-walk forecast (of no change in the log of the exchange rate) even at short horizons of around one month.

Engel and West (2006) and Mark (2009) introduce empirical single-equation models of the exchange rate based on general equilibrium macroeconomic models in which monetary policymakers commit to an instrument rule such as a Taylor rule. The models in these papers build on the New Keynesian models, for example, Clarida, et. al. (1998) and Benigno (2004). Subsequent work in this area has pursued the question of whether these models can be used to forecast changes in the exchange rate outside of the sample of estimation. A key work in this area is Molodtsova and Papell (2009), who find that the Taylor-rule fundamentals provide significantly lower mean-squared errors of out-of-sample forecasts relative to the random walk model, using the Clark and West (2006) test for comparing model predictions. Other contributions include Engel, et. al. (2008), Molodtsova et. al. (2008), Molodtsova et. al. (2011), Wang and Wu (2012), and Binici and Cheung (2012).

In this paper, we attempt to reconcile three related puzzles, two of which arise from the empirical work linking Taylor rule fundamentals and the exchange rate. The first puzzle, however, is well known and has been studied extensively – the uncovered interest parity (UIP) puzzle. If uncovered interest parity held, the optimal forecast of the change in the exchange rate between time *t* and time $t+1$ is the interest differential between the home and foreign country at time *t*. Under uncovered interest parity (UIP) between the U.S. and another country:

(1) $$E_t s_{t+1} - s_t = i_t^{US} - i_t^*,$$

where i_t^{US} is the one-period nominal interest rate in the U.S., i_t^* is the one-period interest rate in a foreign country, s_t is the log of the dollar price of foreign currency, and E_t represents the expectation conditional on all information known at time *t*. In a simple regression of the change in the log of the exchange rate on the interest-rate differential:

(2) $$s_{t+1} - s_t = a_0 + a_1\left(i_t^{US} - i_t^*\right) + \zeta_{t+1},$$

the literature has consistently found an estimate of the slope parameter a_1 that is negative (while the intercept coefficient a_0 is usually estimated to be close to zero.) If UIP holds, we should find the estimate of a_1 is close to one. These studies very often reject the null that $a_1 = 1$, and less often even find a_1 is significantly less than zero.

The second puzzle is that the Taylor rule fundamentals help predict the change in the exchange rate, but in the opposite direction than would arise in a model in which UIP holds. In other words, suppose that the Federal Reserve followed a policy rule for the short-term interest rate,

(3) $$i_t^{US} = \gamma_0^{US} + \gamma_1^{US}\pi_t^{US} + \gamma_2^{US}\tilde{y}_t^{US},$$

and suppose the foreign country had a similar Taylor rule, except with possibly different parameters:

(4) $$i_t^* = \gamma_0^* + \gamma_1^*\pi_t^* + \gamma_2^*\tilde{y}_t^*.$$

If uncovered interest parity, (1), held, we would have:

$$E_t s_{t+1} - s_t = \gamma_0^{US} + \gamma_1^{US}\pi_t^{US} + \gamma_2^{US}\tilde{y}_t^{US} - \left(\gamma_0^* + \gamma_1^*\pi_t^* + \gamma_2^*\tilde{y}_t^*\right).$$

For example, higher U.S. inflation should predict the dollar will depreciate ($E_t s_{t+1} - s_t > 0$). However, Molodtsova and Papell (hereinafter, referred to as MP) and others find that the coefficient signs tend to be the opposite of those implied by uncovered interest parity and Taylor rules - the Taylor rule fundamentals forecast changes in the exchange rate in the "wrong direction". They note especially that the coefficient on U.S. inflation is negative and on foreign inflation is positive. That reversal of sign appears to be consistent with the finding in the empirical work that has led to the so-called "uncovered interest parity puzzle."

It seems that the UIP puzzle and the finding that the Taylor-rule fundamentals predict in the wrong direction are simply manifestations of the same phenomenon. Indeed, if equations (3) and (4) were exactly the monetary policy rules, the interest rates would be perfectly correlated with the Taylor-rule fundamentals and so if one is negatively correlated with $s_{t+1} - s_t$, so must be the other. However, once we recognize that (3) and (4) do not exactly capture monetary policy, it is hard to reconcile these findings with a third puzzle: the Taylor-rule fundamentals do a better job of predicting the exchange rate change than does the relative interest rate differential.

To see this problem, assume that there are some variables that influence monetary policy that are not measured by the researcher, or there are monetary policy shocks. We can augment the rules posited above:

(5) $$i_t^{US} = \gamma_0^{US} + \gamma_1^{US}\pi_t^{US} + \gamma_2^{US}\tilde{y}_t^{US} + u_t^{US}$$

(6) $$i_t^* = \gamma_0^* + \gamma_1^*\pi_t^* + \gamma_2^*\tilde{y}_t^* + u_t^* .$$

Here, u_t^{US} and u_t^* are exogenous random variables. When we substitute for i_t^{US} and i_t^* in equations (5) and (6) into equation (2), we get:

(7) $$s_{t+1} - s_t = b_0 + b_1\pi_t^{US} + b_2\tilde{y}_t^{US} + b_3\pi_t^* + b_4\tilde{y}_t^* + \theta_{t+1},$$

where $b_0 = a_0 + a_1\left(\gamma_0^{US} - \gamma_0^*\right)$, $b_1 = a_1\gamma_1^{US}$, $b_2 = a_1\gamma_2^{US}$, $b_3 = -a_1\gamma_1^*$, $b_4 = -a_1\gamma_2^*$ and $\theta_{t+1} = \zeta_{t+1} + a_1\left(u_t^{US} - u_t^*\right)$. In essence, the finding of MP is that equation (7) has a better fit than the UIP regression (2). If it were the case that equation (2) could be interpreted as implying:

(8) $$E_t s_{t+1} - s_t = a_0 + a_1\left(i_t^{US} - i_t^*\right),$$

then this finding of MP would be impossible to reconcile with the statement that equations (5) and (6) correctly describe monetary policy. These latter equations in essence say that the Taylor rule fundamentals (given by the right-hand-sides of equations (3) and (4)) measure interest rates with noise. If equation (8) were true, the interest differential is the best predictor of the change in the log of the exchange rate, and so it could not be the case that the Taylor rule fundamentals do a better job than the interest rate differential in forecasting currency movements.

But equation (8) is not implied by the familiar UIP regression, (2). The conditional mean of the regression error may not be zero – we cannot claim that $E_t\zeta_{t+1} = 0$.

Here we introduce a small variation on the model presented in Engel (2016) in order to explain these empirical puzzles. Engel (2016) builds a simple New Keynesian open-economy model, in which there is a deviation from UIP that arises because U.S. short-term dollar assets pay an unobserved liquidity return. The actual, observed monetary return on foreign assets must rise as the liquidity return on U.S. dollar assets rise in order for investors to be willing to hold foreign assets in equilibrium. We introduce monetary policy shocks into that model. We can show the model is able to account for the empirical puzzles that the Taylor rule fundamentals forecast the change in the exchange rate better than the interest rate differential, and with the "wrong" sign.

In section 1, we replicate the findings of MP, who find out-of-sample predictive power for the Taylor rule fundamentals in 1973-2006 monthly data. We find that when we extend their sample through 2016, the Taylor-rule fundamentals still have a significantly lower mean-squared error than the random walk forecast, using the Clark-West statistic. Next, we note that the exchange rate used in MP is a monthly average of daily exchange rates. If the daily exchange rate followed a random walk, then the monthly average exchange rate would not, and the change in the monthly exchange rate should be serially correlated. We find that if we replace the monthly average exchange rate with the end-of-month exchange rate to correct this problem, the results about the forecasting power of the Taylor rule fundamentals relative to the random walk model are not as strong as in MP. Indeed, using end-of-month exchange rates, the Taylor rule fundamentals do not generally outforecast the random walk, using the Clark-West statistic.

Moreover, we show that the Clark-West "correction" is very important in reaching the conclusion that the Taylor-rule fundamentals have significantly lower root-mean-squared-errors than the random-walk forecast. Without the correction, the Taylor rule fundamental predictions would not be significantly better. The motivation for the Clark-West correction is that the Taylor-rule model requires estimation of some parameters, while the random-walk model is nested in the Taylor rule model and in fact requires no parameter estimates. Hence, if the Taylor rule model were the true model, the estimated forecast of the change in the exchange rate would be less accurate than the true forecast of the model because of parameter estimation error. The Clark-West statistic corrects for this problem so that it gives a valid test of the out-of-sample forecasting power of an estimated model relative to a nested model. The parameters of the model are estimated by regressions in rolling samples. We plot the parameter estimates over these rolling windows and find considerable variability in the estimates. We argue that the changes over time in the parameter estimates are too large to be accounted for by estimation error of a model with fixed parameters. This finding calls into question the validity of the Clark-West correction, which is designed for models with fixed parameters.

In section 2, we take a different tack. We focus on the in-sample predictability of interest rates and Taylor-rule fundamentals. The regression, (2), that generates the empirical UIP puzzle is a form of testing for in-sample forecasting power. If UIP held, the interest-rate differential should have a coefficient that is significantly greater than 0 (a_1 should equal 1.) In fact, the UIP literature has tended to find a_1 significantly less than one, with a point estimate less than zero, implying that an increase in the U.S. interest rate portends a greater future appreciation of the dollar. We attempt to replicate this finding using monthly data from 1999-2016. We choose January, 1999 as the starting date for our analysis for two reasons. First, the euro came into existence in that month, and the euro-dollar exchange rate is an especially important one on world markets. Second, the parameter instability in the rolling regressions noted above is much more readily apparent in the long span of data that begins in 1973. There is much less parameter instability in shorter samples, including the post-1999 sample.

We are not able to confirm the UIP puzzle in our sample for most countries when we estimate equation (2), but we find a surprising result when we add U.S. and foreign inflation to that equation as independent variables. In that case, U.S. inflation at time t (that is, the increase in prices between month $t-1$ and month t) is a significant predictor of the change in the log of the exchange rate between time t and time $t+1$, for almost all countries. Foreign inflation generally is not helpful for forecasting the change in the exchange rate. Moreover, when the inflation rates are included in the regression, the interest-rate differential is not significant. This finding is very similar to the puzzle we originally pose – the inflation rate of the U.S. is significant in forecasting the change in the log of the exchange rate, but not the interest rate differential.

We then consider the possibility that the U.S. inflation rate for month t really is not useful in predicting exchange rates at the end of month $t+1$, because the consumer price index is not known at the end of month t. It is not announced until the middle of month $t+1$. It may be that when the market learns the level of the CPI at the time of the announcement of that statistic, the exchange rate reacts. Our predictive regression may be mistakenly picking up the reaction of the exchange rate to news about the month t CPI that is announced in the middle of month $t+1$. However, market participants surely have a good idea of what the CPI rate will be, because the CPI is based on observable prices of goods and services. We can examine this possibility by looking at the Bloomberg survey of market participants on their expectation of the value of the CPI. That is, at the end of month t, Bloomberg asks the market what they expect the value of the month t CPI will be when it is announced in the middle of month $t+1$. We find that it is the expected component of the CPI, rather than the surprise, that predicts the change in the exchange rate.

Section 3 then presents a slight extension of the model of Engel (2016) and shows how it may be capable of explaining these findings.

1. Forecasting the Exchange Rate using Taylor Rule Fundamentals

MP propose several variants of equation (7) to forecast changes in the exchange rate. We will focus exclusively on the specification that appeared to be most successful. It augments (7) with lagged interest rates, based on a version of the Taylor rule with interest-rate smoothing:

(9) $$s_{t+1} - s_t = b_0 + b_1\pi_t^{US} + b_2\tilde{y}_t^{US} + b_3\pi_t^* + b_4\tilde{y}_t^* + b_5 i_{t-1}^{US} + b_6 i_{t-1}^* + \theta_{t+1}.$$

This is called the "heterogeneous, symmetric" Taylor rule model. It is "symmetric" because it is based on Taylor rules that have the same variables in each country (inflation, the output gap, and the lagged interest rate), but it is "heterogeneous" because the coefficients in the Taylor rules are not assumed to be equal but of opposite sign (so it does not impose $b_1 = -b_3$.)

First, we replicate the findings of MP, which uses monthly data from March 1973 through June 2006. Some of the data comes from the IMF's *International Financial Statistics* (IFS). The price level data used to construct inflation rates is the consumer price index from line 64 of the IFS. The interest rate data is the money market rate from IFS line 60B. We use industrial production as a proxy for GDP, taken from line 66 in the IFS. We construct the output gap as deviations from the Hodrick-Prescott filtered trend output rate, following the methods described in MP. That method constructs the HP-filtered output level for date *t* using industrial production data only through date *t*, rather than the whole sample. The exchange rate is the monthly-average exchange rate from the Federal Reserve Bank of St. Louis (FRED) database.

To construct one-month-ahead forecasts of the change in the exchange rate between time *t* and time $t+1$, we estimate equation (9) using data through time *t*, and then produce the forecast using the estimated equation. Our first forecast is for March 1982, so we use a 120 month sample (March 1973 – February 1983) to forecast the parameters. We then employ rolling regressions – keeping the estimation sample at 120 months – to update our parameter estimates[1]

When we use the data available on Papell's website, we are able to replicate the results in MP very closely. However, when we collect the data from the sources cited in MP, we find a few

[1] The forecasts for the European countries that eventually adopted the euro go through December 1998, as the euro was adopted in January 1999.

differences. Table 1 reports the *p*-values for the Clark and West (2006) statistic that we calculate, and compares our calculations to those presented in MP. This model provides an exchange rate forecast that has a significantly lower mean-squared error than the random walk forecast of no change in the exchange rate for eight of the twelve currencies. One difference in our data is that our measure of the output gap in some cases differs from MP's near the end of the sample, which may be attributable to data revisions. There were three more significant divergences. Our data for the Portuguese exchange rate differs substantially from MP's. Our data for the interest rate in Switzerland diverges from MP's after 1984. And we find data for industrial production for Sweden in the IFS starting only in 1997. With our dataset, we still find that the Taylor rule model has significant out-of-sample forecasting power in the original MP sample period.

Table 1 also extends the sample through December 2015 for the countries that are not in the Eurozone. We can see that, if anything, the predictive power of the Taylor rule fundamentals has increased in the updated sample. The *p*-values for the test of the forecasting power relative to the random walk null are lower for the Australian dollar, Canadian dollar, Danish krone, Japanese yen, Swiss franc and U.K. pound, compared to the sample in MP. These are all the non-Euro currencies. This is an especially striking finding because generally when the literature finds a model that out-predicts the random walk, the results are fragile and do not extend when the sample period is changed.[2]

Table 2 compares the out-of-sample RMSE of the Taylor-rule model to that of the random walk. The notable point is that the Taylor-rule model, in almost all cases, has a higher RMSE. That is, if one were to use the estimates from this model to forecast exchange rates, one would do worse than using the forecast of no change in the exchange rate. The Taylor-rule model is found to produce better forecasts than the random walk – indeed, statistically significantly better – because of the Clark-West "correction". Clark and West note that even if the Taylor-rule model were true, the econometrician may do worse at forecasting the exchange rate because the parameters may be estimated with error. Whenever we are comparing one model nested in another, the nested model has fewer parameters to estimate, so there is less estimation error. Clark and West propose a simple way to take account of the estimation error – a correction to the RMSE of the more general model that allows us to compare it to the RMSE of the nested model.

In Table 1, in parentheses under the *p*-values for the Clark-West test are reported the *p*-

[2] On this point, see Faust, et al. (2003) and Cheung, et al. (2005, 2016)

values for the Diebold-Mariano-West test that does not correct for the fact that the Taylor-rule model requires estimation of parameters, while the random walk model is nested and actually requires no parameter estimation. Note that these *p*-values are quite large, and the Taylor rule model would not be significantly better than the random walk using this test. However, if the conditions of the Clark-West model are satisfied, these test statistics are incorrect because they do not correct for estimation error as the Clark-West test does. We return to this point below.

As noted above, the exchange rate used in MP is the monthly average exchange rate from the FRED database. Typically studies of exchange-rate forecasting use end-of-period exchange rates rather than period averages. If the null hypothesis were that the daily exchange rate followed a random walk, then the monthly average would not follow a random walk, but instead have a high-order moving average component. The fact that the Taylor-rule fundamentals can outforecast the random-walk model for the monthly average exchange rate may not be that interesting, because changes in the monthly average exchange rate may actually be serially correlated.

In Table 3, we repeat the exercise of Table 1, but using an end-of-month exchange rate. These exchange rates are from the Federal Reserve database, H.10 release, and are measured as noon buying rates in New York on the last trading day of each month. We show results with a sample period that is the same as in MP, and for our extended sample period. The table reports the *p*-values for the test of the null hypothesis that the random walk forecast is no worse than the Taylor-rule model. In both samples, with few exceptions we find that the Taylor-rule model does not have significantly better forecasting power than the random walk model. This is the reverse of the finding when we use monthly-average exchange rates, as in MP. These statistics embed the Clark-West correction, but the out-of-sample forecasting power of the Taylor-rule model found by MP is apparently an artifact of using monthly-average exchange rates.

Another interesting aspect of the forecasts is the behavior over time of the coefficient estimates from the rolling regressions for the Taylor-rule model. Figure 1 plots the coefficient on U.S. inflation, from our regressions using end-of-month exchange rate data. We see that there is considerable variation of the parameter estimates over time. Could this variation arise as a result of estimation error of a constant parameter? This seems improbable. The Figure also plots the 95 percent confidence interval for the parameter estimate at each point in time. We can see by inspection that if we pick almost any point in time *t*, that the parameter estimate for most of the other time periods lie outside the 95-percent confidence interval for the parameter estimate at time t. This strongly suggests that over the full sample, the parameters of the model are not constant. We note that we find the same thing when we plot the coefficient estimates using the monthly-average exchange rate. Indeed, MP's Figure 1 displays their estimates of this coefficient, and it shows similar time variation.

Potentially this is a concern because the Clark-West correction is developed under the assumption that there is a parameter – a constant, not a time-varying parameter – that is estimated, but with estimation error. If the coefficients in this regression vary over time, then the constant-parameter model is misspecified, and so the Clark-West correction is not valid.

In this section, we have seen a couple of reasons to be dubious about the out-of-sample forecasting power of the Taylor-rule model. When we use end-of-period rather than monthly average exchange rates, the Taylor-rule model no longer has significant forecasting properties relative to the random walk, using the Clark-West correction. And, in any case, the validity of the Clark-West procedure is questionable because the parameters of the Taylor-rule forecasting model appear to move greatly over time.

2. In-sample Forecasting and an Extended UIP Test

In this section, we focus on in-sample forecasting power of the interest differential and the Taylor-rule fundamentals. The well-known test for UIP, the regression (2), can be considered an example of an in-sample test. If one finds that $a_1 \neq 0$, one can conclude that the interest rate differential, $i_t^{US} - i_t^*$, has forecasting power for the change in the exchange rate, $s_{t+1} - s_t$. Under the null of the UIP test, $a_1 = 1$, so if uncovered interest parity holds, the interest differential should indeed forecast exchange rate changes. The UIP puzzle is the empirical finding that the slope coefficient is generally found to be significant, but negative. The interest rate differential has forecasting power, but in the opposite direction of the UIP hypothesis.

The first column of Table 4 reports the slope coefficient estimates for this test of UIP.[3] Our exchange rate data is the end-of-month data described above: from the Federal Reserve

[3] We do not use Australia in these tests because inflation data is not available monthly. We add Norway, for which the relevant data is available.

database, H.10 release, noon buying rates in New York on the last trading day of each month, for the Canadian dollar, Danish krone, the euro, Japanese yen, Norwegian krone, Swiss franc, Swedish krona, and U.K. pound. In this table, we also use one-month interest rates measured on the last day of each month. They are the midpoint of bid and offer rates for one-month Eurocurrency rates, as reported on Intercapital from Datastream. We begin the sample in January, 1999, which corresponds with the advent of the euro, and use data through December, 2015. Our choice of start date is dictated by our concern about parameter stability. We have noted above that in the out-of-sample forecasting exercises, the parameters move considerably over the long sample.

In contrast to the usual test for UIP, we do not generally find a significantly negative slope coefficient on the interest rate differential. The point estimate is negative for only four of the eight currencies. In no case is the slope coefficient significantly different from zero, in fact, indicating that the interest rate differential does not have in-sample forecasting power for the change in the exchange rate. Moreover, we cannot reject the UIP null that the slope coefficient is equal to one for any of the currencies. In short, in this data, the UIP puzzle does not hold.

The second panel of Table 4 ("Specification 2") includes U.S. and foreign inflation in the standard UIP regression. That is, we estimate the following equation:

(10) $$s_{t+1} - s_t = b_0 + b^i \left(i_t^{US} - i_t^* \right) + b^{US} \pi_t^{US} + b^* \pi_t^* + \zeta_{t+1}$$

We find that for almost every currency, the coefficient on U.S. inflation, b^{US}, is negative and significantly different than zero, the only exception being in the regression for the dollar/Japanese yen rate. In that case, the point estimate of b^{US} is negative, but it is insignificant. The coefficient on foreign inflation is generally insignificant. There are a few exceptions: for the euro, Swiss franc, and Swedish krona, b^* is significantly positive, and for the Norwegian krone, it is significantly negative. The coefficient on the interest differential, b^i, is insignificant in all cases, except marginally for the euro. In no case is b^i significantly different than one.

These findings remarkably overturn the UIP puzzle. It is no longer the case that the interest differential predicts the change in the exchange rate, but with the wrong sign. Instead, the U.S. inflation rate has explanatory power. When U.S. inflation is high in one month, it appears that we can reliably predict that the dollar will appreciate in the subsequent month.

One possible explanation for this finding is that the U.S. inflation rate for month *t* is not really known at the end of month *t*. The CPI inflation rate is announced with a two-week lag after the end of the month. It may be the case that the news of the month *t* inflation rate incorporated in the CPI announcement in the middle of month $t+1$ causes the exchange rate to move during month $t+1$. So, our measured inflation for month *t* might actually not be known at the end of month *t*, and therefore is not legitimately a predictor of the currency depreciation in period $t+1$.

To examine this hypothesis, we make use of the Bloomberg survey of commercial and investment banks that collects forecasts of the announcement of inflation. To be clear, these are not forecasts of inflation, but instead they are forecasts of what the Bureau of Labor Statistics will announce. For example, in mid-April, the BLS may announce the measure of the CPI inflation rate for March. Bloomberg surveys in-house economists at the beginning of April, and asks what they think inflation was for March – what they forecast the BLS will announce. We take the median of the Bloomberg survey as our measure of what markets think inflation was for month *t*, as of the end of month *t*. The actual inflation data is released in the middle of the month for all of the countries in our dataset, and the survey is taken four to eleven days prior to the release of the data. We call these measures of expected inflation π_t^{USe} and π_t^{*e}, and take them to be proxies for what the market though month *t* inflation was at the end of month *t*. We estimate the equation:

$$(11) \qquad s_{t+1} - s_t = b_0 + b^i\left(i_t^{US} - i_t^*\right) + b^{US}\pi_t^{USe} + b^*\pi_t^{*e} + \zeta_{t+1}$$

Table 5 reports the estimates of equation (11), and compares it to the estimates of equation (10). On the whole, there is very little change. For three of the countries in which the coefficient was negative and marginally significant, we find the coefficient is still negative but marginally insignificant. There is not much change in the estimated magnitude of the effect, but a small increase in the standard error of the coefficient estimate. For all of the countries except Japan, the estimated coefficient on U.S. inflation is negative.

3. A Proposed Solution to the Puzzles

Here we would like to develop a model in which the U.S. inflation rate predicts the change in the exchange rate (high inflation predicts a dollar appreciation subsequently), and that when we control for the U.S. inflation rate, the interest rate differential is not helpful in forecasting the rate of change of the exchange rate. The second fact requires that UIP be violated, because if UIP holds, only the interest differential can forecast the change in the log of the exchange rate. However, it must be the case that the U.S. inflation rate contains information not contained in the interest differential that is helpful for predicting the exchange rate.

We can extend slightly the model in Engel (2016). That paper assumed a Taylor rule with interest rate smoothing, while the model here does not, but it does allow for serially correlated monetary policy shocks, which are very similar in their effect to including a lagged interest rate. The advantage of the model here is that there is a simple, closed-form algebraic solution. The superscript *R* refers to the value of a variable in the U.S. relative to its value in the foreign country. For example, π_t^R is U.S. minus Foreign inflation, or i_t^R is U.S. minus foreign interest rate. In all of the equations, we assume the parameters are the same for the U.S. and the foreign country, which allows us to simplify the system and write the equations in relative terms. The disadvantage of this simplification is that, in the end, it will imply the coefficient on foreign inflation in equation (10) should be equal in absolute value, but of the opposite sign, to the coefficient on U.S. inflation. This model is clearly too simple to fully explain the data, but we view it as providing intuition to the elements that might belong in a more complete model.

The dynamic model has three equations. First, there is the Taylor rule for setting monetary policy. We assume the that each country targets its own inflation rate, and there is a serially correlated error term.

(12) $\quad i_t^R = \sigma\pi_t^R + u_t \ , \ \sigma > 0 \, .$

We assume the error term is serially correlated:

$$u_t = \rho u_{t-1} + \nu_t \ , \ 0 < \rho < 1$$

where ν_t a mean-zero, i.i.d. random variable.

Engel (2016) derives a model, based on Nagel (2016), in which the expected returns on U.S. bonds falls relative to the return on foreign bonds as the U.S. interest rate rises. That is, if the U.S. bond has some value for its liquidity, then the foreign bond must be expected to pay a higher monetary return. Engel (2016) shows that when the U.S. interest rate is relatively high, the U.S. bond's liquidity return will be relatively high. That is because the U.S. interest rate increases under a monetary tightening. The money supply is reduced, so agents value U.S. bonds more for their liquidity. If the U.S. bond pays an intangible liquidity return, its monetary return will be lower than that of the foreign country. This means that the excess monetary return on the foreign bond will be positively related to the difference between the U.S. and foreign interest rate. We let α denote the sensitivity of the excess monetary return on the foreign bond to the $i_t^{US} - i_t^*$ interest differential. In addition, η_t is a mean-zero, i.i.d. random shock to the liquidity return, such that the Home bond is more liquid as η_t is larger. We have:

(13) $i_t^* + \left(E_t s_{t+1} - s_t\right) - i_t^{US} = \alpha\left(i_t^{US} - i_t^*\right) + \eta_t$, $\alpha > 0$.

The expected return differential between U.S. and foreign bonds is $i_t^{US} - i_t^* - \left(E_t s_{t+1} - s_t\right)$.

At first glance, this equation seems like it could not possibly deliver the UIP puzzle, because we have assumed $\alpha > 0$. Rearranging (13), we have $E_t s_{t+1} - s_t = (1+\alpha)\left(i_t^{US} - i_t^*\right) + \eta_t$, so it seems as if we regress $s_{t+1} - s_t$ on $i_t^{US} - i_t^*$, we must get a coefficient greater than one, and certainly not negative. However, $i_t^{US} - i_t^*$ is an endogenous variable, and it responds to liquidity shocks, η_t, so $i_t^{US} - i_t^*$ and η_t are correlated. Engel (2016) shows that the model is capable of explaining a negative slope parameter in the UIP regression (2). In any case, our regressions do not find evidence of the standard UIP puzzle in data since 1999.

We can add and subtract expected Home relative to Foreign inflation to write this expected return differential as $-\left(i_t^R - E_t\pi_{t+1}^R\right) + E_t q_{t+1} - q_t$, where q_t is the real exchange rate: $q_t = s_t - p_t^R$, and $\pi_{t+1}^R = p_{t+1}^R - p_t^R$. $i_t^R - E_t\pi_{t+1}^R$ is the difference in the real interest rate in the Home country and the Foreign country.

The model of liquidity described above then says:

(14) $-\left(i_t^R - E_t\pi_{t+1}^R\right) + E_t q_{t+1} - q_t = \alpha i_t^R + \eta_t$, $\alpha > 0$.

The third equation in the model is the Phillips curve that relates the relative inflation rates in the two countries to the real exchange rate. This is a standard New Keynesian Phillips curve that says that Home inflation will tend to be higher when q_t is high (which means relative prices are low in the Home country):

(15) $\pi_t^R = \delta\left(q_t - \bar{q}_t\right) + \beta E_t\pi_{t+1}^R$, $\delta > 0$, $0 < \beta < 1$.

In practice, it is reasonable to assume δ is small (that is, close to zero, so perhaps something like 0.05 if a time period is one-quarter long) and β is very close to one. $\bar{q}_t$ is an exogenously given "long-run" value for the real exchange rate, and it follows the serially-correlated process:

$$\bar{q}_t = \mu \bar{q}_{t-1} + \varepsilon_t, \ 0 < \mu < 1,$$

where ε_t is a mean-zero, i.i.d. random variable.

Before considering the general solution to this model, it is helpful to examine a simple special case. Set $\text{var}\left(\bar{q}_t\right) = 0$, so there are only two shocks, η_t and u_t. This simple case has the unattractive feature that $E_t s_{t+1} - s_t$ should be perfectly explained by π_t^R and i_t^R in the model. That does not mean that a regression of $s_{t+1} - s_t$ on π_t^R and i_t^R would have a perfect fit, however, because the regression error would just equal the forecast error, $s_{t+1} - E_t s_{t+1}$. Also assume the monetary shock, u_t, is i.i.d., so $\rho = 0$. We have already assumed that η_t is i.i.d.

In this case, we can write the solutions for $E_t s_{t+1} - s_t$, π_t^R and i_t^R as:

$$i_t^R = \frac{1}{1+\sigma\delta\left(1+\alpha\right)} u_t - \frac{\sigma\delta}{1+\sigma\delta\left(1+\alpha\right)} \eta_t$$

$$\pi_t^R = -\frac{\delta\left(1+\alpha\right)}{1+\sigma\delta\left(1+\alpha\right)_2} u_t - \frac{\delta}{1+\sigma\delta\left(1+\alpha\right)} \eta_t$$

$$E_t s_{t+1} - s_t = \frac{1+\alpha}{1+\sigma\delta\left(1+\alpha\right)} u_t + \frac{1}{1+\sigma\delta\left(1+\alpha\right)} \eta_t$$

By inspection, we see $E_t s_{t+1} - s_t = -\frac{1}{\delta}\pi_t^R$, which means that the relative inflation rate will predict the change in the log of the exchange rate, but the interest differential will have no additional predictive power. It is easy to see where this is coming from. Add and subtract expected relative inflation to $E_t s_{t+1} - s_t$, so we have:

$$E_t s_{t+1} - s_t = E_t q_{t+1} - q_t + E_t \pi_{t+1}^R.$$

Since shocks are i.i.d., we must have $E_t q_{t+1} = E_t \pi_{t+1}^R = 0$, so $E_t s_{t+1} - s_t = -q_t$. The Phillips curve in this case is given by $\pi_t^R = \delta q_t$ since $\bar{q}_t = 0$ and $E_t q_{t+1} = 0$. But then, $E_t s_{t+1} - s_t = -q_t = -\frac{1}{\delta}\pi_t^R$.

When π_t^R rises, the central banks raise i_t^R, which leads to a real appreciation of the U.S. dollar. That appreciation causes an expectation of a nominal depreciation to restore the real exchange rate to its equilibrium value (which is zero), because prices are not expected to adjust. The appreciation leads to current inflation, by the Phillips curve. So current inflation predicts the depreciation of the currency. The fit is perfect in this case. The interest differential, on the other hand, is not perfectly correlated with $E_t s_{t+1} - s_t$ because of the risk premium.

The intuition can be deepened by looking at the solution for the real exchange rate:

(16) $$q_t = -\frac{(1+\alpha)(1-\beta\rho)}{1+\sigma\delta(1+\alpha)}u_t - \frac{1}{1+\sigma\delta(1+\alpha)}\eta_t .$$

Both a monetary tightening in the U.S. (an increase in u_t) and an increase in the liquidity value of U.S. bonds (an increase in η_t) lead to a real appreciation of the dollar, and a subsequent expected nominal depreciation. Both of these shocks also lower inflation in the U.S. relative to the foreign country. It is clear that the monetary tightening would have that effect. The liquidity shock also has that effect because the real appreciation leads to greater relative U.S. inflation through the Phillips curve. The upshot is that both shocks lower U.S. inflation, and they both cause a real U.S. appreciation which foretells a nominal depreciation.

However, the two shocks have opposite effects on the relative U.S. to foreign interest rate. Interest rates can rise either because monetary policy has tightened or because there is a shock that makes U.S. bonds less liquid. Those two events have different effects on the value of the currency – a U.S. monetary tightening appreciates the dollar, but when U.S. bonds are less valued for liquidity, the dollar depreciates. In turn, the expected path of future exchange rates is different. High interest rates predict a future depreciation if there has been a monetary tightening, but an appreciation if there has been a denigration of the liquidity value of U.S. bonds. As a result, the interest differential is not as useful in forecasting the change in the exchange rate as is the inflation differential.

It is easy to believe that this model delivers a positive coefficient on the inflation differential in regression (10), if one ignores the fact that inflation and interest rates are endogenous and respond to the shocks. From the Taylor rule, we have $i_t^R = \sigma\pi_t^R + u_t$. The risk premium definition is given by $-i_t^R + E_t(s_{t+1} - s_t) = \alpha i_t^R + \eta_t$. This gives us $E_t(s_{t+1} - s_t) = (1+\alpha)i_t^R + \eta_t$, which can be written as $E_t(s_{t+1} - s_t) = \sigma(1+\alpha)\pi_t^R + (1+\alpha)u_t + \eta_t$. This equation might give the impression that if we regress $s_{t+1} - s_t$ on π_t^R, we would get a positive coefficient. But that is wrong, because π_t^R is negatively correlated with $(1+\alpha)u_t + \eta_t$.

The full solution to the model is given by:

$$(17) \qquad i_t^R = \frac{-\sigma\delta\left(1-\mu\right)}{D_1}\bar{q}_t + \frac{\left(1-\rho\right)\left(1-\beta\rho\right)-\delta\rho}{D_2}u_t - \frac{\sigma\delta}{D_3}\eta_t$$

$$(18) \qquad \pi_t^R = \frac{-\delta\left(1-\mu\right)}{D_1}\bar{q}_t - \frac{\delta\left(1+\alpha\right)}{D_2}u_t - \frac{\delta}{D_3}\eta_t$$

$$(19) \qquad E_t s_{t+1} - s_t = \frac{-\left(1-\mu\right)\delta\sigma\left(1+\alpha\right)}{D_1}\bar{q}_t + \frac{\left(1+\alpha\right)\left[\left(1-\rho\right)\left(1-\beta\rho\right)-\rho\delta\right]}{D_2}u_t + \frac{1}{D_3}\eta_t .$$

$$(20) \qquad q_t = \frac{\delta\left[\sigma\left(1+\alpha\right)-\mu\right]}{D_1}\bar{q}_t - \frac{\left(1+\alpha\right)\left(1-\beta\rho\right)}{D_2}u_t - \frac{1}{D_3}\eta_t ,$$

where

$$D_1 = \delta\left[\left(1+\alpha\right)\sigma-\mu\right]+\left(1-\beta\mu\right)\left(1-\mu\right)$$

$$D_2 = \delta\left[\left(1+\alpha\right)\sigma-\rho\right]+\left(1-\beta\rho\right)\left(1-\rho\right)$$

$$D_3 = 1+\sigma\delta\left(1+\alpha\right).$$

This simple three-equation model cannot be expected to replicate the moments of many different variables in the open economy. But we can see that it will tend to deliver our finding that the inflation rate is a better predictor of the change in the exchange rate than the interest differential under certain assumptions. First, if the variance of the equilibrium real exchange rate is relatively low, as in the example above, then monetary and liquidity shocks are the key drivers of inflation, interest rates and exchange rates. The persistence of monetary shocks also plays a role. In the example above, we assumed they were i.i.d. More generally, if the persistence is low, the intuition of the example goes through. It is possible, however, that when monetary policy shocks are very persistent, a monetary tightening actually lowers nominal interest rates. That could occur because the effect on inflation of a very persistent monetary tightening is to lower inflation immediately by a substantial amount. However, the more plausible case is the one in which monetary tightening raises the nominal interest rate. Then the intuition of the simple example goes through.

To reiterate the point, interest rates have an ambiguous effect on currency values. On the one hand, if the U.S. interest rates rises because of a monetary tightening, the dollar will appreciate but the subsequently is expected to depreciate. But if the interest rate rises because U.S. interest bearing assets have a lower liquidity return, the dollar depreciates, with an expected ensuing appreciation. On the other hand, shocks that raise U.S. inflation unambiguously lead to a dollar depreciation on impact, but a subsequent expectation of an appreciation. That is, both a monetary easing and a reduction in the liquidity return lead to higher inflation, currency depreciation and an expectation of an appreciation.

4. Conclusions

The key findings of this paper are contained in Table 4. When the standard UIP regression is augmented with U.S. and foreign inflation, we find consistently across all currencies that higher U.S. inflation predicts dollar appreciation in the subsequent month. Section 1 of this paper casts some doubt on the evidence that Taylor-rule fundamentals can consistently outforecast the random walk model of exchange rates out of sample, but the in-sample significance of U.S. inflation is intriguing. There is actually no internal contradiction between the claim that an economic fundamental, like U.S. inflation, is not useful in producing a superior forecast relative to the random walk, but is significant in regression (10). As Engel and West (2005) demonstrate, this is exactly the outcome that arises in present-value models of the exchange rate, when the discount factor is close to one.

We illustrate a model in which U.S. bonds pay a liquidity return that could potentially account for our empirical findings. The model is extremely simple, and intended to be illustrative. We believe our empirical conclusions present a challenge for open-economy macroeconomists.

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Table 1: Replicating Table 4 of Molodtsova-Papell (2009)
Specification: Heterogenous symmetric Taylor rule model with smoothing and constant using HP filter for potential output construction

Country	Our Results (A) (data end at Dec 2015)^	Our Results (B) (data end at Jun 2006)^	Molodtsova-Papell (2009) (data end at Jun 2006)
Non Euro Zone			
Australia	0.002*** (0.616)	0.044** (0.834)	0.038**
Canada	0.000*** (0.631)	0.007*** (0.548)	0.021**
Denmark	0.032** (0.962)	0.077** (0.995)	0.032**
Japan	0.012** (0.886)	0.019** (0.759)	0.011**
Switzerland	0.069* (0.981)	0.209 (0.981)	0.016**
Sweden	^^	^^	0.593
U.K.	0.016** (0.739)	0.133 (0.866)	0.033**
Pre-Euro Zone **(**data end at Dec 1998)^			
France	0.034** (0.913)		0.008***
Germany	0.209 (0.917)		0.126
Italy	0.000*** (0.432)		0.001***
Netherlands	0.019** (0.794)		0.009***
Portugal	0.558 (0.974)		0.898

Notes: The table reports p-values for 1-month-ahead CW tests of equal predictive ability between the null of a martingale difference process and the alternative of a linear model with Taylor rule fundamentals. DMW p-values without CW correction are reported in parenthesis. The alternative model is the model with symmetric Taylor rule fundamentals with smoothing, which is estimated with heterogenous inflation and output coefficients, and with a constant using HP trends to estimate potential output. *, **, and *** indicate that the alternative model significantly outperforms the random walk at 10, 5, and 1% significance level, respectively, based on standard normal critical values for the one-sided test. Estimation window is 120 months.

^The models are estimated using data from January 1975 for Canada, September 1975 for Switzerland, February 1983 for Portugal, January 1989 for UK and March 1973 for the rest of the countries. The sample ends in December 1998 for Euro Area countries and Decemeber 2015 for column A (and June 2006 for column B) for the rest of the countries.

^^In this exercise, we restrict our self by using the same data source as Molodtsova-Papell (2009). i.e. nominal exchange rate data from FRED, all other data from IMF IFS. The recent IMF IFS data has Sweden Industrial Production data only start from 1997. Therefore, we are unable to do a meaningful comparison with the p-value in Molodtsova-Papell (2009).

Table 2: Comparison of the out-of-sample RMSE of the Taylor-rule model and the random walk Specification: Heterogenous symmetric Taylor rule model with smoothing and constant using HP filter for potential output construction.

Country	Taylor-rule model RMSE (data end at Dec 2015)^	Taylor-rule model RMSE (data end at Jun 2006)^	Random walk RMSE (data end at Dec 2015)^	Random walk RMSE (data end at Jun 2006)^
Non Euro Zone				
Australia	0.000747	0.000634	0.000735	0.000601
Canada	0.000260	0.000162	0.000255	0.000160
Denmark	0.000693	0.000724	0.000650	0.000671
Japan	0.000753	0.000799	0.000718	0.000774
Switzerland	0.000801	0.000847	0.000744	0.000776
U.K.	0.000471	0.000416	0.000448	0.000378
Pre-Euro Zone (data end at Dec 1998)^				
France	0.000755		0.000691	
Germany	0.000818		0.000766	
Italy	0.000709		0.000715	
Netherlands	0.000797		0.000762	
Portugal	0.000514		0.000506	

Notes: The table reports root-mean-square error (RMSE) for 1-month-ahead forecasting with the Taylor rule fundamentals model and the random walk. Estimation window is 120 months.
^ The models are estimated using data from January 1975 for Canada, September 1975 for Switzerland, February 1983 for Portugal, January 1989 for UK and March 1973 for the rest of the countries. The sample ends in December 1998 for Euro Area countries and Decemeber 2015 for column 1,3 (and June 2006 for column 2,4) for the rest of the countries.

Table 3: Replicating Table 4 of Molodtsova-Papell (2009) using end of month exchange rate data
Specification: Heterogenous symmetric Taylor rule model with smoothing and constant using HP filter for potential output construction.

Country	Our Results (data end at Dec 2015)^	Our Results (data end at Jun 2006)^	Molodtsova-Papell (2009) (data end at Jun 2006)
Non Euro Zone			
Australia	0.033**	0.149	0.038**
Canada	0.161	0.272	0.021**
Denmark	0.287	0.328	0.032**
Japan	0.155	0.119	0.011**
Switzerland	0.368	0.434	0.016**
Sweden	^^	^^	0.593
U.K.	0.048**	0.272	0.033**
Pre-Euro Zone (data end at Dec 1998)^			
France	0.139		0.008***
Germany	0.631		0.126
Italy	0.001***		0.001***
Netherlands	0.027**		0.009***
Portugal	0.292		0.898

Notes: The table reports p-values for 1-month-ahead CW tests of equal predictive ability between the null of a martingale difference process and the alternative of a linear model with Taylor rule fundamentals. Numbers without CW correction are reported in parenthesis. The alternative model is the model with symmetric Taylor rule fundamentals with smoothing, which is estimated with heterogenous inflation and output coefficients, and with a constant using HP trends to estimate potential output. *, **, and *** indicate that the alternative model significantly outperforms the random walk at 10, 5, and 1% significance level, respectively, based on standard normal critical values for the one-sided test. Estimation window is 120 months.

^The models are estimated using data from January 1975 for Canada, September 1975 for Switzerland, February 1983 for Portugal, January 1989 for UK and March 1973 for the rest of the countries. The sample ends in December 1998 for Euro Area countries and Decemeber 2015 for column 1 (and June 2006 for column 2) for the rest of the countries.

^^In this exercise, we restrict our self by using the same data source as Molodtsova-Papell (2009). i.e. nominal exchange rate data from FRED, all other data from IMF IFS. The recent IMF IFS data has Sweden Industrial Production data only start from 1997. Therefore, we are unable to do a meaningful comparison with the p-value in Molodtsova-Papell (2009).

Table 4: UIP regression with inflation included

Specification 1 - UIP regression:

$$s_{t+1} - s_t = a_0 + a_1\left(i_t^{US} - i_t^*\right) + \zeta_{t+1}$$

Specification 2 - UIP regression with inflation included:

$$s_{t+1} - s_t = b_0 + b^i\left(i_t^{US} - i_t^*\right) + b^{US}\pi_t^{US} + b^*\pi_t^* + \zeta_{t+1}$$

Country	Specification 1	Specification 2		
	$\hat{a}_1$	$\hat{b}^i$	$\hat{b}^{US}$	$\hat{b}^*$
Canada	0.08	2.05	-4.78*	4.32
	(2.76)	(2.97)	(2.74)	(3.78)
Denmark	-0.76	-0.76	-4.92*	2.30
	(1.95)	(1.95)	(2.50)	(3.74)
Euro zone	-1.96	4.30*	-17.42***	22.22***
	(1.95)	(2.29)	(3.61)	(5.13)
Japan	0.51	0.54	-0.57	-1.78
	(1.11)	(1.34)	(2.13)	(2.28)
Norway	0.15	-0.16	-4.75**	-5.95**
	(1.50)	(1.52)	(2.14)	(2.56)
Switzerland	-1.71	-1.18	-8.69***	9.79**
	(1.86)	(1.98)	(3.15)	(4.40)
Sweden	-1.32	1.34	-9.62***	5.87*
	(1.64)	(1.95)	(2.93)	(3.26)
UK	0.45	-0.41	-4.34**	-1.10
	(1.94)	(1.95)	(1.70)	(1.95)

Notes: The standard errors are reported in parenthesis. *, **, and *** indicate that the alternative model significantly different from zero at 10, 5, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. Sample period are monthly data from January 1999 to December 2015 (204 observations).

Table 5: UIP regression with inflation survey data included

Specification 1 – UIP regression with inflation included:

$$s_{t+1} - s_t = b_0 + b^i \left(i_t^{US} - i_t^* \right) + b^{US} \pi_t^{US} + b^* \pi_t^* + \zeta_{t+1}$$

Specification 2 - UIP regression with inflation survey data included:

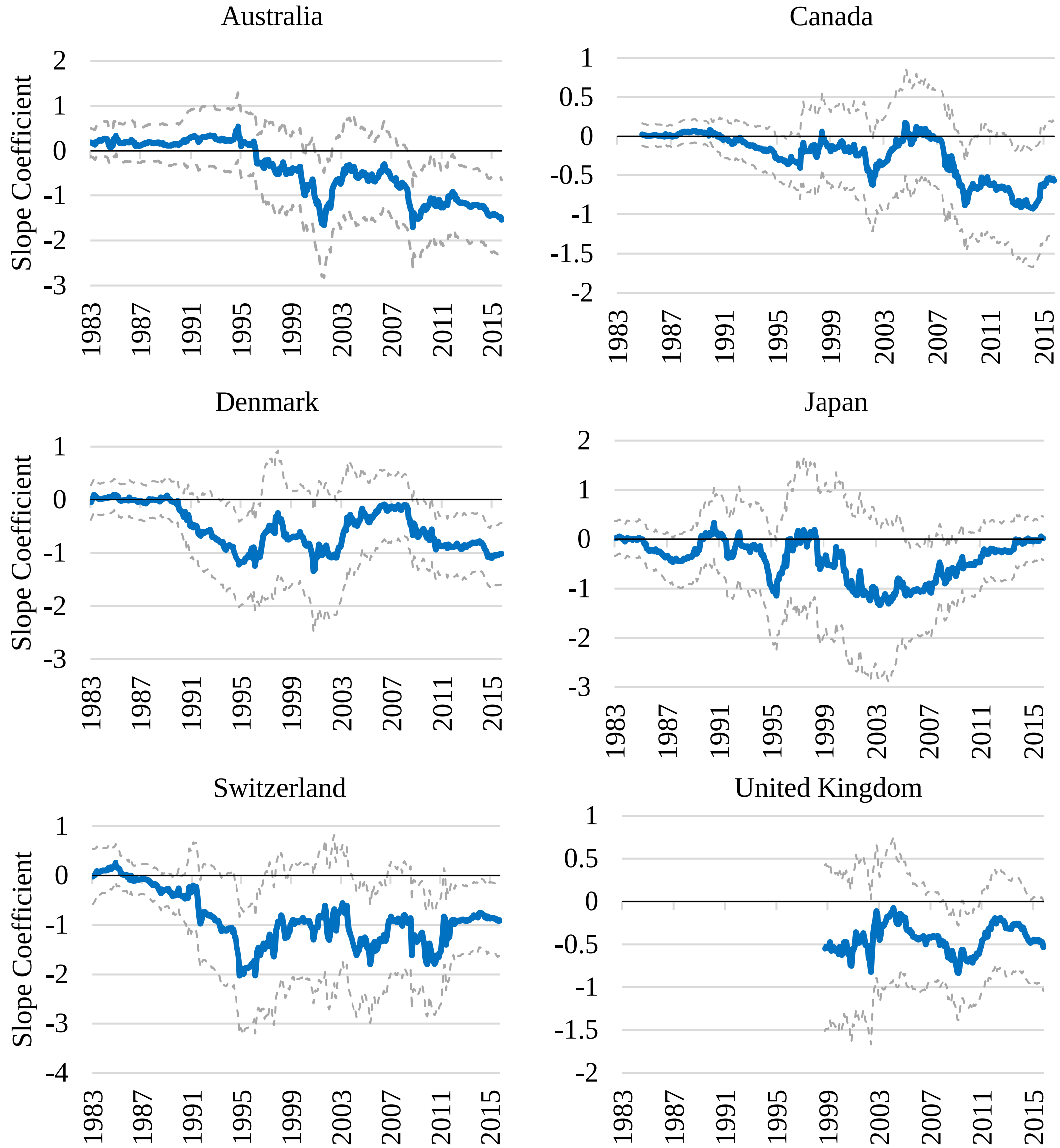
$$s_{t+1} - s_t = b_0 + b^i \left(i_t^{US} - i_t^* \right) + b^{US} \pi_t^{USe} + + b^* \pi_t^{*e} + \zeta_{t+1}$$

Country	Specification 1			Specification 2		
	$\hat{b}^i$	$\hat{b}^{US}$	$\hat{b}^*$	$\hat{b}^i$	$\hat{b}^{US}$	$\hat{b}^*$
Canada	2.05	-4.78*	4.32	5.76	-4.48	3.62
	(2.97)	(2.74)	(3.78)	(4.03)	(3.08)	(5.19)
Denmark	-0.76	-4.92*	2.30	3.98	-4.34	4.00
	(1.95)	(2.50)	(3.74)	(2.89)	(3.04)	(5.03)
Euro zone	4.30*	-17.42***	22.22***	8.09***	-15.74***	19.80***
	(2.29)	(3.61)	(5.13)	(3.03)	(3.90)	(5.52)
Japan	0.54	-0.57	-1.78	1.40	0.61	-2.44
	(1.34)	(2.13)	(2.28)	(1.77)	(2.23)	(2.42)
Norway	-0.16	-4.75**	-5.95**	3.51	-5.68**	-6.38*
	(1.52)	(2.14)	(2.56)	(2.25)	(2.27)	(3.42)
Switzerland	-1.18	-8.69***	9.79**	2.14	-7.00**	6.54
	(1.98)	(3.15)	(4.40)	(2.71)	(3.39)	(4.82)
Sweden	1.34	-9.62***	5.87*	4.29*	-8.17**	4.70
	(1.95)	(2.93)	(3.26)	(2.49)	(3.30)	(3.83)
UK	-0.41	-4.34**	-1.10	1.98	-3.07	-1.61
	(1.95)	(1.70)	(1.95)	(2.74)	(2.18)	(2.60)

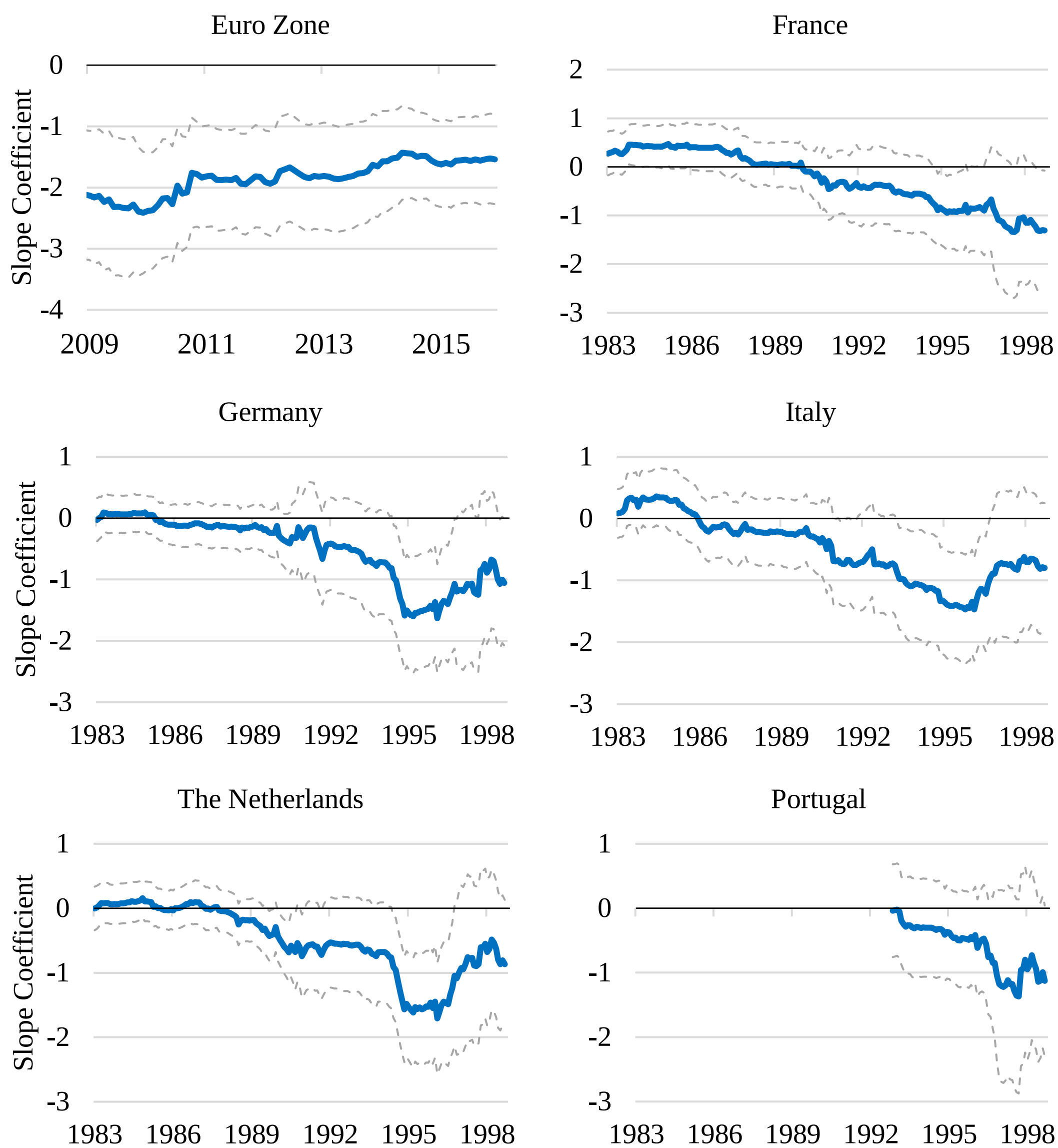
Notes: The standard errors are reported in parenthesis. *, **, and *** indicate that the coefficient is significantly different from zero at 10, 5, and 1% significance level, respectively, based on standard normal critical values for the two-sided test. Sample period for specification 1 is monthly data from January 1999 to December 2015 (204 observations). Sample period for specification 2 is monthly data from February 2004 for Denmark to December 2015 and from November 2003 to December 2015 (146 observations) for the rest of the countries.

Figure 1. Slope coefficient on U.S. inflation

A) Non-Eurozone



B) Eurozone



Note: Figures plot the slope coefficients of US inflation and 95 percent confidence interval of these regressions. Confidence intervals calculated from OLS standard errors. The first 10-year data are used to construct HP trend so forecast starts from 1983 March.