NEW EVIDENCE ON THE FORWARD PREMIUM PUZZLE

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Abstract

The forward premium anomaly—exchange rate changes are negatively related to interest rate differentials—is one of the most robust puzzles in financial economics. We recast the underlying parity relation in terms of lagged forward interest rate differentials, documenting a reversal of the anomalous sign on the coefficient in the traditional specification. We show that this novel evidence is consistent with recent empirical models of exchange rates which imply exchange rate changes depend on two key variables—the interest rate differential and the magnitude of the deviation of the current exchange rate from that implied by purchasing power parity.

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I. Introduction

Well over one hundred papers document, in some form or another, the forward premium anomaly—namely, that future exchange rate changes do not move one-for-one with interest rate differentials across countries. In fact, they tend to move in the opposite direction (e.g., see Hodrick (1987) and Engel (1996) for survey evidence). This anomaly has led to a plethora of papers over the last two decades that develop possible explanations with only limited success. It is reasonable to conclude that the forward premium anomaly is one of the more robust puzzles in financial economics. Parallel to work on the forward premium puzzle, another literature has developed, starting with Meese and Rogoff (1983), documenting an equally startling puzzle—exchange rates do not seem to be related to fundamentals.[1] The random walk model has proven almost unbeatable, even against models with a variety of finance and macro variables.

This paper looks at the forward premium anomaly, and the fundamental determinants of exchange rates, in a novel way by recasting the uncovered interest rate parity (UIP) relation in terms of future exchange rate movements against forward interest rate differentials across countries. We study a subset of the G10 currencies, for which we have sufficient interest rate data, over the time period from 1980-2010. In stark contrast to current research on uncovered interest rate parity, past forward interest rate differentials have strong forecasting power for exchange rates. R^2s at some horizons exceed 10% for annual exchange rate changes relative to about 2% for the traditional specification. Moreover, the direction of these forecasts coincides with the theoretical implications of UIP.

We provide a plausible explanation for these findings. Specifically, we can explain why uncovered interest rate parity fails, why it appears to work better using lagged forward interest rate differentials, and why the explanatory power for exchange rates increases with the horizon, i.e., more lagged and stale information. The key insight is that, while interest rate differentials lead to capital flows from the carry trade, associated currency movements, PPP violations and the rejection of UIP, the build-up of these violations generally gets reversed, that is, there is a reversion back to PPP. Our explanation is consistent with recent empirical exchange rate models, which argue exchange rate changes are a function of two key state variables—the interest rate

[1] Meese and Rogoff (1983) find that the literature's typical structural models of exchange rates cannot outperform a naïve random walk model, even when one uses ex-post values of the variables of interest such as money supply, real income, inflation and interest rates. These findings are revisited and confirmed by Cheung, Chinn, and Garcia Pascual (2003) using updated data. For a theoretical analysis of this issue, see Engel and West (2005).

differential and the magnitude of the deviation of the current exchange rate from that implied by PPP. We show that these two variables separate the relevant explanatory information into two offsetting components, which, if used separately, significantly increase the explanatory power for exchange rates. The interest rate differential captures violations of UIP associated with carry trade related capital flows, while the deviation from PPP captures the reversal of this effect in the longer term as exchange rates revert to fundamentals.

We regress annual exchange rate changes of the G10 currencies on the interest rate differential and the real exchange rate. The results are striking and consistent with the story. Controlling for the real exchange rate, the coefficient on the interest rate differential becomes more negative and is identified more precisely. Moreover, together the variables generate R^2s that range up to 37% across the 9 exchange rates. These results are reconciled with the aforementioned UIP regressions that use forward interest rate differentials.

This paper is organized as follows. Section II introduces the data and presents new empirical evidence on the exchange rate parity relation in terms of forward interest rate differentials. In Section III, we provide a simple story for exchange rate determination, additional empirical evidence in support of this story, and a reconciliation of this evidence with our novel forward interest rate results. Section IV concludes.

II. Uncovered Interest Rate Parity: Evidence

A. Data

We use monthly data from Datastream on exchange rates, price levels, and interest rates for the countries corresponding to the G10 currencies. The choice of sample period for each country is based on the availability of interest rate data. A subset of four countries (the United States, the United Kingdom, Switzerland, and Germany) is used extensively due to the availability of term structure data at annual maturities out to five years going back to 1976. In particular, data for the term structure of zero-coupon interest rates are derived from LIBOR data (with maturities of six and twelve months) and swap rates (two-, three-, four-, and five-year semi-annual swap rates).[2]

[2] Cubic spline functions are fitted each month for each country to create a zero curve for maturities of 6, 12, 18, …, 60 months. Our spline function fits the available data exactly, namely LIBOR rates for the 6-month and 12-month maturities, and semi-annual swap rates for maturities of 24 months, 36 months, 48 months, and 60 months. Therefore, the only maturities we need to spline are 18 months, 30 months, 42 months, and 54 months. We maximize the smoothness of the spline function over these unknowns by minimizing the sum of squared deviations.

Since swap data only become available in the late 1980s, we augment our zero curve data with data from Philippe Jorion. Jorion and Mishkin (1991) collect and derive data for zero coupon bonds from one month to five years for this subset of countries.[3] Swap and LIBOR data is preferred to typical government bond data because the quotes are more liquid and less prone to missing data, supply and demand effects, and tax-related biases. To the extent that there is a swap spread (i.e., the difference between the swap and government bond rates) embedded in the data, its effect is diminished in our analysis by our use of interest rate differentials across countries. Using the zero curve data, we compute continuously compounded, one-year spot interest rates and one-year forward interest rates from years 1 to 2, 2 to 3, 3 to 4, and 4 to 5. For the remainder of the countries we compute one-year, continuously compounded, spot interest rates starting in January 1980, or later as dictated by data availability.

Using the exchange rate data, we compute annual changes in log exchange rates with the U.S. dollar as the base currency, starting in January 1980 (or later as dictated by the availability of interest rate data) and ending in December 2010, i.e., we examine changes in the USD/FX rates for the G10 countries. Given the monthly frequency of the underlying data, adjacent annual changes have an 11-month overlap. The choice of the start date reflects the fact that our analysis of the subset of 4 countries with extensive term structure data matches the j to $j+1$ year forward interest rate at time $t-j$ with the subsequent exchange rate change from time t to time $t+1$.[4] Thus, the 4 to 5 year forward interest rate in January 1976, the first observation, is matched with the annual exchange rate change from January through December 1980.

To ensure that we use exactly the same exchange rate series for all regressions for these countries, we use calendar year 1980 as the first observation throughout, truncating the interest rate series accordingly. We use the same sample period for the exchange rates of the other countries if there is sufficient interest rate data. Finally, we also combine this exchange rate data with CPI data to construct real exchange rates for all country pairs. Further discussion of these series is postponed until Section III.

To summarize, the final dataset consists of annual exchange rate changes, with the first observation corresponding to calendar year 1980 and the last to calendar year 2010 (361

[3] We thank Philippe Jorion for graciously providing us with the data.

[4] Throughout the paper we use annual exchange rate changes and annual interest rates and forward rates; thus, for ease of exposition, all periods are denoted in years with the exception of the simulation analysis in Section II.D. However, as noted above, these annual quantities are calculated on a monthly overlapping basis to maximize the information content of the empirical analysis.

observations sampled monthly) for 5 of the 9 exchange rates, with start dates ranging from February 1986 to January 1993 for the other 4. For all countries we also have matched 1-year, spot interest rates covering a sample whose dates corresponds to the beginning of the period of each annual exchange rate change, e.g., from 1/1980-1/2010 for the 5 countries with the full sample. For the 4 countries with term structure data, we also have forward interest rates over the periods 1/1979–1/2009, 1/1978–1/2008, 1/1977–1/2007, and 1/1976–1/2006 for horizons j = 1,…,4, respectively (all with 361 observations). Table 1, Panels A and B contain descriptive statistics for these variables.

B. Existing Evidence

The expectations hypothesis for exchange rates (forward parity) is commonly written as

$$E_t s_{t+j} = f_t^{\,j}, \tag{1}$$

where s_{t+j} is the log of the spot price of foreign currency at time $t+j$, and $f_t^{\,j}$ is the log of the j-year forward exchange rate at time t. Assuming no arbitrage and covered interest rate parity (i.e., $f_t^{\,j} - s_t = j(i_{t,j} - i^*_{t,j})$, where $i_{t,j}$ is the domestic, j-year, continuously compounded (log), annualized interest rate at time t and the superscript * denotes the corresponding foreign interest rate), the expected change in the exchange rate equals the interest rate differential. Thus, one standard way of testing equation (1) for annual changes in exchange rates is to estimate the regression

$$\Delta s_{t,t+1} = \alpha + \beta(i_{t,1} - i^*_{t,1}) + \varepsilon_{t,t+1}, \tag{2}$$

where $\Delta s_{t,t+1} \equiv s_{t+1} - s_t$. Under uncovered interest rate parity (UIP), α and β should be 0 and 1, respectively. That is, high interest rate currencies should depreciate and low interest rate currencies should appreciate in proportion to the interest rate differential across the countries. Intuitively, expected (real) returns on bonds in the two countries should be equal. This hypothesis has been resoundingly rejected, and, most alarming, β tends to be negative, i.e., exchange rates move in the opposite direction to that implied by the theory. In the context of equation (1), the forward premium, $s_{t+j} - f_t^{\,j}$, has a systematic bias and is predictable.[5]

[5] See, e.g., Engel (1996) and Lewis (1995) for surveys of this literature. Interestingly, some evidence suggests that the forward premium anomaly may be confined to developed economies and may be asymmetric or state dependent even in those economies (Bansal and Dahlquist (2000), Wu and Zhang (1996)).

One possible explanation for these findings is the existence of a risk premium in exchange rates. However, in order for this omitted variable in the regression in equation (2) to cause the coefficient β to change signs, this risk premium must exhibit significant time-variation and be negatively correlated with the interest rate differential, as noted in Fama (1984). While such a risk premium could explain the results from a statistical perspective, from an economic standpoint the key challenge is to identify what risk this premium is providing compensation for. So far, attempts to match the implied risk premium to economic risks have proven unsuccessful.[6]

As a first look at equation (2), Table 1, Panel C reports estimates from regressions of annual exchange rate changes of the G10 currencies on interest rate differentials on a monthly overlapping basis. The β coefficients are all negative for the USD/FX exchange rates, confirming the well-known negative relation between exchange rates and interest rate differentials. While the estimates are fairly noisy, tests of the null hypothesis that the coefficients equal 1 can be resoundingly rejected for seven of the nine US-G10 currency pairs.

The low R^2s in most of the regressions are also notable, and this feature is both disappointing and puzzling. The key fundamentals underlying expected exchange rate movements are interest rate differentials between countries. For example, these interest rate differentials may represent expected inflation rate differentials. Since inflation is fairly predictable (see, e.g., Fama and Gibbons (1984)), and inflation differentials are a fundamental driver of exchange rates under purchasing power parity, one would have expected the model to explain a much larger degree of the variation in these exchange rates.

C. Information about Exchange Rate Changes in Long-Maturity Forward Rates

Equation (1), UIP, is almost always cast in terms of interest rate differentials and then tested using equation (2). In this subsection, we present a novel way to analyze UIP by recasting the parity relation in terms of future exchange rate movements against forward interest rate differentials across countries.

Specifically, we can also use equation (1) to define expected changes in future exchange rates as the difference between two forward exchange rates. That is,

$$E_t[\Delta s_{t+j,t+k}] = f_t^k - f_t^j, \qquad (3)$$

[6] See, e.g., Bekaert and Hodrick (1993), Bekaert (1995, 1996), Bekaert, Hodrick and Marshall (1997), Mark and Wu (1998) and Graveline (2006).

where $k > j$. Under the expectations hypothesis of exchange rates, the period t expected depreciation from $t+j$ to $t+k$ equals the difference in the corresponding forward exchange rates at time t. Under covered interest rate parity, we can replace the forward exchange rates in equation (3) with the interest rate differentials between the two countries, i.e.,

$$E_t[\Delta s_{t+j,t+k}] = k\,(i_{t,k} - i^*_{t,k}) - j\,(i_{t,j} - i^*_{t,j}). \tag{4}$$

Rearranging the interest rate differential terms in equation (4), and using the definition of forward interest rates,[7] we get

$$\begin{aligned} E_t[\Delta s_{t+j,t+k}] &= (k\,i_{t,k} - j\,i_{t,j}) - (k\,i^*_{t,k} - j\,i^*_{t,j}) \\ &= (k-j)(if_t^{\,j,k} - if_t^{\,j,k^*}) \end{aligned}, \tag{5}$$

where $if_t^{\,j,k}$ and $if_t^{\,j,k^*}$ are the continuously compounded, annualized, forward interest rates at time t from $t+j$ to $t+k$ for domestic and foreign currencies, respectively. Equation (5) is the basis for the empirical analysis to follow. It says that, under UIP, the expected depreciation in future exchange rates is equal what we call the *forward interest rate differential*.

Equation (5) extends the classical approach to characterizing and testing the expectations hypothesis presented in equations (1) and (2). It implies a more general specification of the expectations hypothesis,

$$\Delta s_{t,t+1} = \alpha_j + \beta_j\,(if_{t-j}^{\,j,j+1} - if_{t-j}^{\,j,j+1^*}) + \varepsilon_{t-j,t+1}. \tag{6}$$

Under the expectations hypothesis of exchange rates, the annual exchange rate change from t to t+1 should move one-for-one with the forward interest rate differential from j to j+1 that was set at time t–j. That is, α_j and β_j should equal 0 and 1 respectively. Equation (2) is a special case of equation (6) for $j = 0$. Note that the specification in equation (6) is identical to that in equation (5), but we have chosen, for ease of exposition, to fix the period over which exchange rate movements are measured and lag the forward interest rate differentials rather than fix the point in time at which we measure forward interest rate differentials and lead the change in the exchange rate.

Using regression equation (6), Table 2, Panel A provides estimates over different horizons and across a subset of the G10 currencies for tests of the expectations hypothesis of

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[7] The annualized forward interest rate is defined as $if_t^{\,j,k} \equiv \dfrac{k\,i_{t,k} - j\,i_{t,j}}{k-j}$

exchange rates.[8] This analysis requires a history of long-term forward interest rates, and, as described in Section II.A above, we have such data for the United States, the United Kingdom, Switzerland, and Germany. In contrast to Table 1, Panel C and the conclusions in much of the literature, Table 2 shows that forward interest rate differentials can predict changes in future exchange rates. At least as important is that their predictive power has the right sign. The U.S./Germany forward interest rate differentials at horizons of one to four years yield coefficients of 0.68, 0.76, 2.02, and 3.17 for the USD/DEM exchange rate. The results for the USD/GBP and USD/CHF exhibit similar patterns, with coefficients of 0.92, 3.41, 1.94, and 2.54 and -0.16, 0.42, 1.40, and 2.07 looking forward one to four years, respectively. These results are quite different from the significant negative coefficients that plague Table 1, Panel C (i.e., -0.84, -0.71 and -1.26 for USD/DEM, USD/GBP, and USD/CHF, respectively).

The coefficient estimates exhibit two features in addition to the fact that they are positive. First, they tend to increase in the horizon. Second, for longer horizons they seem to exceed the theoretical value of 1. However, these coefficient estimates are noisy, especially at longer horizons, so more formal tests are warranted. Table 2, Panel B reports tests that the coefficients are equal and that the coefficients are equal to one. The second column lists the horizon over which the test is conducted, either $j = 0, 1, 2, 3,$ and 4 or $j = 1, 2, 3,$ and 4 (but not $j = 0$). The third column indicates whether we are testing for equality across horizons (labeled "=") or whether we are testing the tighter restriction that the coefficients are equal to one (labeled "=1"). In the former case, the fourth and fifth columns provide the restricted coefficient estimate under the Lagrange multiplier test and the associated standard error. Turning to the results, the Lagrange multiplier (LM) tests yield only two rejections at the 10% level, both for the hypothesis that the coefficients equal one at all horizons. In contrast, the Wald tests yield rejections in all but four cases.[9] Thus, there is definitely evidence, though perhaps not overwhelming, of horizon-dependent coefficients and rejections of UIP.

Note that equation (6) exploits the information in the entire forward curve. However, the error term is now a j-year ahead forecast, and is serially correlated up to $(j+1)12-1$ observations,

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[8] Using different specifications, Chinn and Meredith (2005), Bekaert, Min and Ying (2007) and Chinn and Quayyum (2012) also analyze implications for uncovered interest rate parity at short- and long-horizons.

[9] We employ both the Lagrange multiplier and Wald statistics for testing the joint hypotheses. As shown by Berndt and Savin (1977), there is a numerical ordering between these statistics, which may lead to different inferences being drawn. For an especially relevant discussion, see Bekaert and Hodrick (2001) in the context of testing the expectations hypothesis of the term structure. In their context, the Wald test over-rejects while the Lagrange multiplier test under-rejects, results that are consistent with our simulation evidence discussed later.

for monthly overlapping data. Therefore, one of the difficulties in studying multi-step ahead forecast regressions like those specified in equation (6) is the availability of data. While sophisticated econometrics have somewhat alleviated the problem (Hansen and Hodrick (1980) and Hansen (1982)), the benefits are still constrained by the number of independent observations. There are two sources for the serial correlation of the error term. The first arises from sampling annual exchange rate changes on a monthly basis, leading to a moving average structure out to 11 months. Sampling at the monthly frequency improves the efficiency of the estimators, but only to a degree (Boudoukh and Richardson (1994) and Richardson and Smith (1992)). The second potential source arises directly from the *j*-year ahead forecast. For the regression in equation (6), however, the degree of serial correlation in the errors depends upon the relative variance of exchange rates versus interest rate differentials, and the correlation of unexpected shocks to these variables. There are strong reasons to suspect that these factors mitigate the serial correlation problem. Table 1, Panel A shows that exchange rates are much more variable than interest rate differentials, and they are also relatively unpredictable (see Table 2, Panel A). Therefore, because the forecast update component of the residual in equation (6) is likely to be small relative to the unpredictable component as we move forward in time, the induced serial correlation in the errors will be correspondingly small, and the overlap will not substantially reduce the effective number of independent observations. This intuition is confirmed through a Monte Carlo simulation described in Section II.D below. Section II.D also provides a comparison with alternative long-horizon methodologies for forward premium regressions (e.g., see Chinn and Meredith (2005) and Chinn and Quayyum (2012)).

For now, Table 2, Panel A reports statistics from the simulation model of Section II.D. We report the cross-sectional standard deviation (across replications) of the relevant parameter estimate (in the column "SD"), and the two-sided simulated P-value for the test that $\beta = 1$ (in the column "P-value"), i.e., the percentage of the replications in which the absolute magnitude of deviation of the estimated coefficient from one equals or exceeds the deviation for the estimated coefficient from the actual data. For these calculations, we simulate under the null hypothesis of $\beta = 1$ and use the resampled exchange rate changes for the relevant exchange rate, but simulating under normality produces similar results. The cross-sectional standard deviations tend to exceed the reported standard errors, especially at longer horizons, suggesting that these standard errors may be somewhat understated. However, the inferences drawn from the P-values are consistent with those from standard hypothesis tests of the individual coefficients. Specifically, the short-horizon (j = 0) coefficients are statistically significantly different from one, as is the coefficient for j = 2 for the USD/GBP.

As a final comment on the evidence, note that in Table 2 the regression R^2s have a tendency to increase with the horizon. While the dependent variable, i.e., annual exchange rate changes, is the same, the forecasting variable differs. For all three exchange rates, the R^2s are higher for the forward interest rate differential regressions (equation (6)) at horizon j = 4 than for the interest rate differential regression (equation (2)). What is remarkable about this result is that the information in the former regressions is (i) old relative to current interest rates, and (ii) more subject to measurement error due to the calculation of forward rates. We argue below that this finding is an important clue to understanding the fundamental relation between exchange rates, inflation, and interest rates, and, more importantly, the forward premium anomaly.

D. Information in Long Maturity Forward Rates: A Monte Carlo Simulation Analysis

One potential concern with the results reported in Table 2 is that the standard errors are spuriously low and the R^2s are spuriously high due to small sample problems in the regressions. We argue in Section II.C that the overlap problem is not that serious due to the relatively low predictability of exchange rate changes, but it is still important to verify this conjecture. Consequently, we construct a Monte Carlo experiment in which we employ a VAR for the relevant forward interest rate differentials, spot interest rate differentials, and changes in exchange rates, imposing the expectations hypotheses of interest rates and using two different models for exchange rates. In one experiment we impose the expectations hypothesis for exchange rates, i.e., we assume uncovered interest rate parity holds, and in the other experiment we assume exchange rates follow a random walk, i.e., exchange rate changes are unpredictable.

We also consider two different distributional assumptions for the shocks to exchange rate changes. In the first analysis, we assume that the shocks across all equations follow a multivariate normal distribution. In the second analysis, we resample the shocks to exchange rates from the series of monthly exchange rate changes observed in the data. We then simulate these models, generating 100,000 replications of 432 monthly observations. For each replication, we aggregate the data to an annual frequency, as in the empirical analysis, and we then estimate equation (6). For comparison purposes, we also estimate the long-horizon regression version of equation (2) following Chinn and Meredith (2005) and Chinn and Quayyum (2012). Thus we can assess the small sample properties of our specification and also compare them to those of the alternative long-horizon regressions.

Specifically, for the first experiment, we assume that the expectations hypotheses of exchange rates and interest rates hold at a monthly frequency, and that the longest maturity forward rate differential (the forward rate from month 59 to month 60) follows an AR(1) process:[10]

$$
\begin{aligned}
&\Delta s_{t,t+1} = i_{t,1} - i^{*}_{t,1} + \varepsilon^{s}_{t,t+1} \\
&i_{t+1,1} - i^{*}_{t+1,1} = if^{1,2}_{t} - if^{1,2^*}_{t} + \varepsilon^{1}_{t,t+1} \\
&if^{1,2}_{t+1} - if^{1,2^*}_{t+1} = if^{2,3}_{t} - if^{2,3^*}_{t} + \varepsilon^{2}_{t,t+1} \\
&if^{2,3}_{t+1} - if^{2,3^*}_{t+1} = if^{3,4}_{t} - if^{3,4^*}_{t} + \varepsilon^{3}_{t,t+1} \\
&\vdots \\
&if^{58,59}_{t+1} - if^{58,59^*}_{t+1} = if^{59,60}_{t} - if^{59,60^*}_{t} + \varepsilon^{59}_{t,t+1} \\
&if^{59,60}_{t+1} - if^{59,60^*}_{t+1} = \rho(if^{59,60}_{t} - if^{59,60^*}_{t}) + \varepsilon^{60}_{t,t+1}
\end{aligned}
\tag{7}
$$

where

$$
\varepsilon_{t,t+1} \equiv \begin{bmatrix} \varepsilon^{s}_{t,t+1} \\ \varepsilon^{1}_{t,t+1} \\ \varepsilon^{2}_{t,t+1} \\ \vdots \\ \varepsilon^{59}_{t,t+1} \\ \varepsilon^{60}_{t,t+1} \end{bmatrix} \sim MVN\left(0, \Sigma\right)
\tag{8}
$$

We impose the following structure on the covariance matrix of the shocks:

$$
\Sigma = \left[\begin{array}{c|ccccc}
\sigma^2_s & 0 & \cdots & 0 & \cdots & 0 \\ \hline
 & \sigma^2_i & \cdots & \upsilon^{t-1}_{ij}\rho_{ij}\sqrt{\upsilon^{t-1}_{i}}\sigma^2_i & \cdots & \upsilon^{59}_{ij}\rho_{ij}\sqrt{\upsilon^{59}_{i}}\sigma^2_i \\
 & & \ddots & & & \vdots \\
 & & & \upsilon^{t-1}_{i}\sigma^2_i & & \upsilon^{59-t}_{ij}\rho_{ij}\sqrt{\upsilon^{t+58}_{i}}\sigma^2_i \\
 & & & & \ddots & \vdots \\
 & & & & & \upsilon^{59}_{i}\sigma^2_i
\end{array}\right]
\tag{9}
$$

Specifically, we impose that the variances of the shocks to forward interest rate differentials decline in maturity and that the correlations between the shocks to forward interest rate

[10] Throughout this subsection, periods are measured in months.

differentials decline in the difference between the maturities, at fixed rates determined by the parameters υ_i and υ_{ij}, respectively. We also impose zero correlation between the shock to exchange rate changes and the shocks to forward interest rate differentials. In the data, these correlations are relatively small and negative. However, these negative correlations are another manifestation of the violations of UIP that result in negative coefficients in the forward premium regressions in Tables 1 and 2. Therefore, we set the correlations to zero for the purposes of the Monte Carlo analyses.

We calibrate the parameters of the model in order to match approximately the covariance matrix of the annual exchange rate changes and the annual spot and forward interest rate differentials, and the autocorrelation of the 4- to 5-year forward interest rate differentials. Obviously, these values differ somewhat across the three exchange rates we employ in the empirical analysis, so we target intermediate values. The inferences drawn from the Monte Carlo analysis are not sensitive to the precise choice of the parameters.

Define the state vector

$$y_{t+1} = \begin{bmatrix} \Delta s_{t,t+1} \\ i_{t+1,1} - i^{*}_{t+1,1} \\ if^{1,2}_{t+1} - if^{1,2^{*}}_{t+1} \\ \vdots \\ if^{58,59}_{t+1} - if^{58,59^{*}}_{t+1} \\ if^{59,60}_{t+1} - if^{59,60^{*}}_{t+1} \end{bmatrix} \tag{10}$$

Equations (7)-(8) imply that $y_{t+1} \sim MVN(0,\Omega)$, where Ω is a function of ρ and Σ. The simulation procedure is as follows:

1. Draw starting values y_t from the distribution $y_t \sim MVN(0,\Omega)$.
2. Draw an error vector $\varepsilon_{t,t+1}$ from the distribution $\varepsilon_{t,t+1} \sim MVN(0,\Sigma)$.
3. Compute y_{t+1} using this error vector and the lagged state vector via equation (26).
4. Return to step 2 above.

We generate 100,000 simulations of 432 monthly observations. We aggregate these monthly data to an annual frequency and construct simulated samples with the appropriate lag structure of annual, monthly overlapping data of 361 observations each, the length of our sample.

For each sample, we estimate the forward premium regressions in equation (6) and compute various test statistics. We also estimate the long-horizon versions of the forward premium regression in equation (2), after Chinn and Meredith (2005) and Chinn and Quayyum (2012).

We also conduct a second Monte Carlo exercise, which is identical to the first except that we assume that exchange rates follow a random walk:

$$\Delta s_{t,t+1} = \varepsilon^{s}_{t,t+1} \ . \tag{11}$$

Finally, we repeat the analyses above, relaxing the restriction that the shocks to exchange rate changes are normally distributed in order to incorporate the possible effects of fat tails in the relevant distribution. Instead, we resample with replacement actual monthly exchange rate changes from either the USD/GBP, the USD/DEM, or the USD/CHF series. To preserve the excess kurtosis, but to eliminate any sample-specific mean or skewness effects, we augment the two series with an equal number of observations that correspond to the negative of the observed exchange rate changes.

The second and third to last columns of Table 2, Panel A, discussed in Section II.C, and Table 3 report the key results. Table 3, Panel A compares the R^2s from the regressions in equation (6), i.e., using forward interest rate differentials, to those from the long-horizon versions of the regression in equation (2), i.e., using long-horizon spot rate differentials, under the expectation hypothesis of exchange rates (β_j = 1). The statistics in this panel, and in the remainder of Table 3, are calculated from simulations that resample from the USD/GBP exchange rate changes because this series exhibits the most excess kurtosis, but inferences from simulations under normality or using the USD/DEM or USD/CHF exchange rate changes are similar. When one uses equation (6), the biases in the R^2s are clearly less severe than in the corresponding long-horizon regressions, and they are not horizon dependent. As the horizon goes from one to four years, the bias, i.e., the difference between the mean R^2 from the simulations and the true R^2, ranges from 2.75% (5.99% simulated versus 3.24% true infinite sample R^2) to 2.63% for regressions using forward interest rates, versus an increase from 4.64% (11.43% simulated versus 6.79% true) to 6.83% for the long-horizon spot rate regressions.

Equally problematic for the long-horizon regressions, there is much less independent information in these regressions compared with the forward interest rate regressions. The correlations between the coefficient estimators range from 0.69 to 0.97 across the various horizons in the long-horizon regressions, in contrast to a much lower range of correlations, from 0.37 to 0.86, in the forward interest rate regressions.[11]

Table 3, Panel B reports the results under the assumption that the exchange rate follows a random walk (β_j = 0). Again the forward interest rate regressions have smaller biases in R^2s relative to the long-horizon regressions, and there is considerably more independent information in the former regression system. The regressions using the forward interest rate differentials have a bias that ranges from 2.78% to 3.05%, while the biases in the long-horizon regressions increase with the horizon up to 9.87%. Overall, these simulation results suggest that small sample bias cannot explain the large differences in R^2s across horizons found in the data, and that the forward interest rate regressions have better statistical properties than the corresponding long-horizon regressions.

Table 3, Panel C presents simulation results for the Wald and Lagrange multiplier tests for the regressions in equation (6) across the horizons with β_j = 1. Consistent with Berndt and Savin (1997) and Bekaert and Hodrick (2001), the Wald test substantially over-rejects the null hypothesis, while the LM test tends to under-reject the null hypothesis, especially for high significance levels. For example, for the hypothesis β_j = 1 across all four horizons, the LM test rejects 5.3% and 0.2% of the time at the 5% and 1% levels, respectively, while the Wald test rejects the null hypothesis in 28.7% and 14.7% of the simulations. Moreover, while the LM test performs similarly for both the β_j = 1 and β_j equal hypotheses, the small sample properties of the Wald test are much worse for the hypothesis β_j = 1.

III. Reconciling the Forward Premium Anomaly Evidence

The results provided in Section II are important stylized facts that need to be explained in the context of recent attempts at solving the forward premium puzzle of exchange rates. In this section, we lay out a simple story for exchange rate determination that is built around evidence consistent with the existing literature. While this story is just one potential explanation for the observed behavior of uncovered interest rate parity using spot and forward interest rate differentials, we provide additional supporting empirical evidence. Specifically, we show how recent empirical exchange rate determination models in Jorda and Taylor (2012), which depend on two key state variables—the interest rate differential and the deviation of the current

[11] The coefficient estimates are slightly downward biased in both cases, but these results are omitted for brevity.

exchange rate from that implied by purchasing power parity—are consistent with the standard forward premium anomaly and our contrasting results using forward interest rate differentials..

There has been a plethora of recent papers in the area of exchange rate determination which explain the forward premium anomaly in the context of "carry trades" in which investors borrow in low interest rate currencies and invest in high interest rate currencies. This recent literature argues that relatively high expected real rates, in countries with high nominal rates, causes capital inflows and an associated appreciation of the currency (e.g., Froot and Ramadorai (2005), Burnside, Eichenbaum, Kleschelski, and Rebelo (2006), Lustig and Verdelhan (2007), Clarida, Davis, and Pedersen (2009), Farhi, Fraiberger, Gabaix, Ranciere, and Verdelhan (2009), Jorda and Taylor (2012), Jurek (2009), Berge, Jorda and Taylor (2010), Menkhoff, Sarno, Schmeling, and Schrimpf (2012), among others). In recent years these carry trades may have been undertaken primarily by hedge funds, but in earlier times long-only investors in search of high returns may have been taking one leg of the carry trade, while corporate borrowers in search of low borrowing costs may have been taking the other.

One preferred explanation is that the carry trade return resulting from the appreciation of the currency is compensation for the possibility of a crash in the currency's value – the so-called "up the stairs, down the elevator" description of high interest rate currencies (e.g., Brunnermeier, Nagel, and Pedersen (2009) and Plantin and Shin (2010)). Moreover, theories based on speculative dynamics (e.g., Plantin and Shin (2010)) and existing empirical work (e.g., Brunnermeier, Nagel, and Pedersen (2009) and Jorda and Taylor (2012)) imply that this carry risk should be increasing in the deviation from purchasing power parity (PPP). In other words, as the exchange rate moves further from its fundamental PPP relation, the tension to bring it back increases. This view is consistent with a substantial body of evidence that shows PPP holds in the long run and is therefore an important building block for exchange rates (see, e.g., Abuaf and Jorion (1990), Kim (1990), Rogoff (1996), Lothian and Taylor (1996), Taylor (2001, 2002), and Imbs, Mumtaz, Ravn, and Rey (2005)).

Thus, there are two opposing effects driving exchange rate movements—appreciation of high interest rate currencies due to carry trade related capital flows and reversals of this appreciation because of reversion to fundamentals (e.g., Froot and Ramadorai (2005)). Note that the deviation of the exchange rate from PPP is unobservable, but we can construct a variable that captures the same information, up to a constant. Specifically, consider the log real exchange rate

$$q_t = s_t + (z_t^* - z_t)\,, \tag{12}$$

where q and s are the log real and nominal exchange rates, respectively, and z and z^* denote the log price levels in the domestic and foreign country, respectively. Under PPP, the real exchange rate is constant; thus, the observed real exchange rate equals the deviation of the exchange rate from this PPP implied level, up to an unknown constant. In the context of a regression analysis, this unknown constant will appear in the intercept.

The empirical model in Jorda and Taylor (2012) combines the standard forward premium regression in equation (2) with the real exchange rate in equation (12) above:

$$\Delta s_{t,t+1} = \alpha + \psi_1 (i_{t,1} - i_{t,1}^*) + \psi_2 q_t + \varepsilon_{t,t+1}\,. \tag{13}$$

They motivate the real exchange rate variable as the deviation from the fundamental equilibrium exchange rate, although they do not provide a motivating theoretical model since they are primarily interested in forecasting and the associated trading strategies. They estimate various specifications employing the two variables in equation (13) using monthly data across multiple exchange rates for the period 1986-2008 and report results consistent with ours.

A. Exchange Rate Determination: Evidence

Table 4, Panel A presents summary statistics for the log real exchange rate series for the nine currency pairs of the G10 countries. The means are essentially meaningless in that they reflect the normalization of the price level series in the two countries. It is not surprising that the series are very persistent given the persistence of the exchange rate series, and the relatively strong positive correlation between the series is also expected.

For the G10 countries, we run the bivariate version of the forward premium regression in equation (13) using the deviation of the exchange rate from PPP. We estimate regressions of annual exchange rate changes (overlapping monthly) on the log real exchange rate and the interest rate differential at the beginning of the year (and special cases thereof). The results are reported in Table 4, Panel B. For ease of comparison, the top line for each exchange rate reports the standard UIP regressions, which are also reported in Tables 1 and 2. The second line reports the regression with the log real exchange rate, and the final line reports the results from the full specification.

The first notable result in Table 4, Panel B is that, both alone and in the full specification, the log real exchange rate appears with a negative and statistically significant coefficient for all nine currency pairs. This negative coefficient is consistent with the intuition from the explanation provided above. When the real exchange rate is high, i.e., the dollar has appreciated less or depreciated more than would be suggested by the relative inflation rates in the two countries, this effect is expected to reverse in the coming year. Moreover, this reversion to PPP, or expected currency "crash", explains a significant fraction of the variation in exchange rate changes on its own, with R^2s averaging 15.2% for the G10 currencies.

The second notable result is that including both the interest rate differential and the deviation from PPP variables substantially increases the explanatory power of the regression. For example, for the USD/DEM exchange rate, the R^2 increases to 27.1% (from 1.8% in the UIP regression and 18.2% in the real exchange rate regression). This pattern is not unusual and holds for all the other currency pairs (except USD/NOK for which the increase is small). In fact, the R^2 increases on average to 24.9% versus 4.0% in the UIP and 15.2% in the real exchange rate regressions, respectively. Clearly, controlling for both the PPP reversion effect and the carry trade effect together enhances our ability to identify both effects and increases the explanatory power for exchange rates.

Consistent with the existing literature described above, the results presented here help explain why interest rate differentials on their own do not explain exchange rate movements. Including the real exchange rate variable that directly measures the magnitude of the deviation from PPP helps better isolate two offsetting effects—the carry trade effect and the eventual reversion to PPP, both of which are proxied for by the interest rate differential. When the real exchange rate is added to the standard UIP regression, the magnitude of the coefficient on interest rate differentials increases, i.e., the coefficient becomes more negative, for all nine of the exchange rates. Specifically, the average coefficient, ψ_1, in equation (13) is -1.80 compared to an average for the analogous coefficient, β in equation (2) of -0.88. In other words, partially fixing the omitted variable problem in the standard forward premium regression in equation (2) more than doubles the magnitude of the average coefficient on the interest rate differential. In other words, when the regression controls for the reversion to PPP, the interest rate differential is left to pick up only the carry trade effect, which is then seen to be much larger.

A similar, but smaller, effect operates on the real exchange rate. While this variable primarily proxies for deviations from PPP, it also picks up the carry trade effect to a lesser degree. Consequently, adding the interest rate differential to a regression with the real exchange rate increases the magnitude of the coefficient on the latter variable as well. Specifically, the average coefficient decreases from -0.31 in the univariate regression to -0.39 in the bivariate regression.

B. Reconciling the Evidence

Section II.C of this paper provides a new way to look at the forward premium puzzle, using past forward interest rate differentials rather than current interest rate differentials. While the theoretical implications of these two approaches are similar, the empirical results show stark differences.

In this subsection, motivated by the empirical results in Section III.A, we offer an explanation and new evidence to reconcile these results, namely that past forward interest rate differentials at different horizons pick up the two opposing effects – a carry effect and reversion to PPP – to different degrees, yielding horizon-dependent coefficients and R^2s.

Specifically, while it is clear that the empirical exchange rate model in (13) can fit the stylized facts of the forward premium puzzle, it is not obvious that this model is consistent with the forward interest rate differential results of Section II.C. First, there is the need to reconcile the forward premium anomaly (i.e., a negative β in equation (2) and negative ψ_1 in equation (13)) with the forward interest rate differential results (i.e., positive βs in equation (6)). Note that the carry trade component of exchange rates explains why PPP does not hold and β is negative in equation (2) and (13). If there is a positive probability that exchange rates will revert back to PPP, however, then the effect of the crash component on the regression coefficient in equation (6) partially (or even fully) reverses the effect of the carry trade. For short horizons, the carry trade effect dominates and the coefficient is negative. For longer horizons, the role of the forward interest rate differential as a proxy for the magnitude of the PPP violation and hence the size of a crash, should it occur, can become the more important factor, and the coefficient becomes positive.

Second, the coefficients in the forward interest rate differential regressions appear to increase in maturity and the explanatory power of the forward interest rate differential also increases in the horizon over which the regressions are estimated, i.e., with information about countries' future interest rates that becomes increasingly stale. The intuition is that the current interest rate differential has information not only about future real rate differentials, which, due to the carry trade, lead to exchange rate movements, but also about past interest rate differentials. Depending on the probability of a reversion, these past differentials tell us something about the future magnitude of the reversion in the exchange rate. Of course, replacing the forward interest rate differentials of the forward premium regressions with crash-specific variables (e.g., such as deviations in PPP) would improve the fit of the exchange rate model in equation (6). In other words, stale forward interest rate differentials are just proxies for potential reversion states.

As a first pass, we decompose interest rate differentials into forward interest rate differentials (set j years ago) and the difference between the two. Note that if the expectations hypothesis of interest rates were approximately true, then this decomposition would be equivalent to breaking interest rate differentials into their expected value (set j years ago) and unexpected shocks over these j years. Specifically, Table 5, Panel A presents results for the regression

$$\Delta s_{t,t+1} = \alpha_j + \phi_{0,j}[(i_{t,1} - i^*_{t,1}) - (if^{\,j,j+1}_{t-j} - if^{\,j,j+1^*}_{t-j})] + \phi_j(if^{\,j,j+1}_{t-j} - if^{\,j,j+1^*}_{t-j}) + \varepsilon_{t,t+1}, \quad (14)$$

for the three exchange rates (USD/DEM, USD/GBP, USD/CHF) and each horizon. For all three currencies, the coefficient ϕ_j is generally positive and increasing (albeit noisily) in the horizon. In contrast, the $\phi_{0,j}$ coefficients are all negative and declining in magnitude as the horizon increases. The R^2s are quite impressive.

For example, for the USD/DEM exchange rate, at forward rate horizons of one to four years, the ϕ_j s are 0.55, 0.32, 1.33, and 2.40, respectively, while the $\phi_{0,j}$ s are –2.32, –1.03, –0.96, and –0.88. From the standpoint of our explanation of exchange rate determination, the positive and increasing coefficients on the forward interest rate differentials are capturing the probability and magnitude of a reversion of the currency to PPP, while the negative coefficients on the forecast error in the exchange rate regression are capturing the carry trade effect. Thus, the negative $\phi_{0,j}$ explains why the forward premium anomaly exists from a statistical viewpoint, that is, why we get negative coefficients and low R^2s in Table 1, Panel C. Breaking up current interest rates into the two components separates information about the magnitude and probability of future currency reversions to PPP contained in the forward curve from current interest rates. By not breaking them up, the two information sources offset each other, leading to a low R^2.

One interesting stylized fact from the regression results in Table 2 is that using dated (i.e., old) information in forward interest rate differentials increases explanatory power for future exchange rates. The explanation provided in Section III.A argues that this information is important because these differentials predict the reversion component of future changes in exchange rates. That is, past forward interest rate differentials predict not only the interest rate differential (i.e., carry effect) but also the reversion to PPP. As the horizon increases, the latter term dominates.

To better understand these results, we estimate an analogous regression to equation (13), namely annual exchange rate changes (overlapping monthly) on our PPP deviation measure and the past forward interest rate differential (instead of the interest rate differential):

$$\Delta s_{t,t+1} = \alpha_j + \beta_j (if_{t-j}^{j,j+1} - if_{t-j}^{j,j+1^*}) + \gamma_j q_t + \varepsilon_{t,t+1}. \tag{14}$$

These results are reported in Table 5, Panel B for the three available currencies (DEM, GBP, CHF), relative to the USD, over the horizons j = 0,…,4. The horizon j = 0 is equivalent to the regression specification (23) with results also provided in Table 4, Panel B.

Table 5, Panel B provides two pieces of evidence in support our explanation for exchange rate determination described above. First, for horizons j = 1,...,4, six of twelve coefficients on the forward interest rate differential are now negative (three significantly so). For the regressions in Table 2 that did not include the PPP deviation variable, eleven of twelve coefficients on the forward interest rate differential were positive. Recall that the forward interest rate differential has information about two components – the carry effect and the reversion to PPP. Therefore, the reason that the coefficients flip sign is that in regression specification (14) the inclusion of q_t proxies for the reversion to PPP component of the forward interest rate differential, leaving just the carry effect. As documented in Table 1, Panel C, the carry effect has a negative sign.

Second, and equally important, in contrast to Table 2, Table 5, Panel B shows that the R^2s now generally decrease with the horizon (with the exception of the final horizon for GBP). The reason is that past forward interest rate differentials (due to their staleness) are a poorer measure of the carry effect than the current interest rate differential. Of course, the magnitude of the R^2s are higher for regression specification (14) with j = 0 than not only the j = 1,…,4 horizons but also the alternative forward interest rate differential specifications given by either equations (6) or (14).

IV. Concluding Remarks

The forward premium puzzle is one of the more robust and widely studied phenomena in financial economics. Our paper makes two important contributions to this large literature.

First, we document that recasting the UIP regression in terms of lagged forward interest rate differentials, rather than spot interest rate differentials, deepens the puzzle. Specifically, the coefficients in these regressions are positive in contrast to the negative coefficients in the standard UIP specification, and the R^2s and coefficients are generally increasing in the horizon.

Second, motivated by the existing literature on the forward premium anomaly, we provide an explanation that both fits the existing evidence and reconciles it with our new evidence. The key insight is that exchange rate changes reflect two distinct but related phenomena. A carry trade effect associated with interest rate differentials pushes exchange rates in the opposite direction to that predicted by a standard model of PPP. However, exchange rates revert back to their fundamental levels. Forward interest rate differentials at different horizons pick up both of these conflicting effects to different degrees, yielding horizon-dependent coefficients and R^2s. We show that it is possible to decompose these two effects using either forward interest rate differentials and shocks to these differentials, or spot interest rate differentials and real exchange rates. The data are consistent with these decompositions and provided further support for the proposed explanation for exchange rate movements.

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Table 1: Preliminaries

Panel A: Summary Statistics – Exchange Rates

Exchange Rate	Start Date	No. of Obs.	Mean (%)	SD (%)	1st Order Autocorr.	12th Order Autocorr.
USD/GBP	1/80	361	-1.35	11.46	0.93	-0.02
USD/DEM	1/80	361	0.67	12.71	0.93	0.10
USD/CHF	1/80	361	1.59	12.53	0.92	0.00
USD/AUD	10/88	256	0.63	12.26	0.93	-0.16
USD/CAD	1/80	361	0.40	6.93	0.93	-0.05
USD/JPY	1/80	361	3.20	11.98	0.93	0.11
USD/NZD	10/88	256	0.77	13.54	0.94	-0.05
USD/NOK	2/86	288	0.81	10.87	0.91	-0.26
USD/SEK	1/93	205	0.44	11.92	0.92	-0.06

Correlations

Exchange Rate	USD /DEM	USD /CHF	USD /AUD	USD /CAD	USD /JPY	USD /NZD	USD /NOK	USD /SEK
USD/GBP	0.71	0.64	0.61	0.47	0.29	0.60	0.76	0.74
USD/DEM		0.94	0.53	0.34	0.47	0.58	0.86	0.84
USD/CHF			0.45	0.25	0.53	0.52	0.80	0.75
USD/AUD				0.78	0.20	0.90	0.67	0.79
USD/CAD					0.06	0.66	0.55	0.66
USD/JPY						0.22	0.14	0.20
USD/NZD							0.69	0.78
USD/NOK								0.86

Panel B: Summary Statistics – Forward Interest Rate Differentials

$if^{j,j+1} - if^{*j,j+1}$	j	Mean (%)	SD (%)	1st Order Autocorr.	12th Order Autocorr.
US-UK	0	-1.71	1.99	0.95	0.51
	1	-1.05	1.20	0.89	0.40
	2	-1.08	1.28	0.91	0.51
	3	-0.92	1.35	0.88	0.52
	4	-0.96	1.52	0.91	0.60
US-Germ.	0	1.13	2.37	0.98	0.74
	1	1.38	1.83	0.97	0.73
	2	1.56	1.52	0.96	0.71
	3	1.57	1.46	0.96	0.71
	4	1.48	1.51	0.97	0.72
US-Switz.	0	2.57	2.52	0.98	0.75
	1	3.00	2.29	0.95	0.77
	2	3.22	1.82	0.96	0.77
	3	3.23	1.76	0.94	0.74
	4	3.18	1.77	0.96	0.76
US-Aus.	0	-2.09	1.97	0.98	0.67
US-Can.	0	-0.62	1.32	0.95	0.60
US-Jap.	0	3.23	2.05	0.97	0.55
US-NZ	0	-2.63	1.46	0.96	0.45
US-Nor.	0	-1.94	2.65	0.98	0.66
US-Swe.	0	-0.35	2.02	0.99	0.57

Panel C: The Forward Premium Puzzle – 1-Year Horizon

Exchange Rate	α	Std. Err.	β	Std. Err.	R^2
USD/GBP	-2.78	0.02	-0.84	0.88	2.11
USD/DEM	1.47	0.02	-0.71	0.71	1.77
USD/CHF	4.83	0.02	-1.26	0.60	6.41
USD/AUD	-0.43	0.03	-0.50	0.97	0.66
USD/CAD	0.28	0.01	-0.20	0.65	0.14
USD/JPY	11.63	0.02	-2.61	0.53	20.01
USD/NZD	-0.25	0.06	-0.39	1.98	0.17
USD/NOK	0.55	0.02	-0.14	0.66	0.11
USD/SEK	0.00	0.02	-1.24	1.05	4.44

Panels A and B report summary statistics (mean, standard deviation, first-order autocorrelation, twelfth-order autocorrelation, and cross correlations) for annual changes in log exchange rates and 1-year forward interest rate differentials at various horizons, sampled monthly (horizon j = 0 corresponds to spot interest rates). Panel C reports coefficient estimates, corresponding standard errors (heteroskedasticity and autocorrelation adjusted using the Newey and West (1987) method), and R^2s from the forward premium regression at the 1-year horizon

$$\Delta s_{t,t+1} = \alpha + \beta(i_{t,1} - i^*_{t,1}) + \varepsilon_{t,t+1}.$$

Exchange rate data cover 1/1980–12/2010 and interest rate data cover 1/1980–1/2010, 1/1979–1/2009, 1/1978–1/2008, 1/1977–1/2007, and 1/1976–1/2006 for horizons j = 0,…,4, respectively, for a maximum total of 361 monthly observations (with later start dates and fewer observations as dictated by data availability and noted in Panel A). See Section II.A for a detailed description of the data.

Table 2: The Expectations Hypothesis of Exchange Rates

Panel A: Regression Results

Exchange Rate	j	α_j	Std. Err.	β_j	Std. Err.	SD	P-value (%)	R^2
USD/GBP	0	-2.78	1.99	-0.84	0.88	0.91	4.73	2.11
	1	-0.38	2.35	0.92	1.30	1.01	93.20	0.94
	2	2.34	2.13	3.41	1.01	1.14	4.06	14.61
	3	0.44	1.77	1.94	1.02	1.35	47.54	5.22
	4	1.09	1.90	2.54	0.87	1.71	36.34	11.28
USD/DEM	0	1.47	1.87	-0.71	0.71	0.97	7.87	1.77
	1	-0.27	1.95	0.68	1.20	1.08	75.19	0.96
	2	-0.52	2.17	0.76	1.29	1.22	83.68	0.83
	3	-2.50	2.37	2.02	1.38	1.44	46.50	5.33
	4	-4.02	2.37	3.17	1.32	1.83	23.29	14.16
USD/CHF	0	4.83	2.32	-1.26	0.60	1.04	3.41	6.41
	1	2.06	2.52	-0.16	0.97	1.15	29.28	0.08
	2	0.23	3.51	0.42	1.22	1.30	64.00	0.38
	3	-2.92	3.97	1.40	1.32	1.53	78.27	3.85
	4	-4.99	3.61	2.07	1.21	1.94	57.24	8.59

Panel B: Hypothesis Tests

Exchange Rate	j	Test	β	Std. Err.	Deg. Of Freedom	LM Stat.	P-value	Wald Stat.	P-value
USD/GBP	1-4	=	1.33	0.70	3	4.81	0.19	7.22	0.07
	0-4	=	0.69	0.49	4	4.97	0.29	10.16	0.04
	1-4	=1			4	4.75	0.31	8.87	0.06
	0-4	=1			5	5.20	0.39	10.17	0.07
USD/DEM	1-4	=	1.27	0.63	3	3.83	0.28	4.33	0.23
	0-4	=	-0.64	0.56	4	5.26	0.26	8.87	0.06
	1-4	=1			4	3.93	0.42	4.72	0.32
	0-4	=1			5	10.62	0.06	12.65	0.03
USD/CHF	1-4	=	-0.20	0.53	3	4.39	0.22	6.00	0.11
	0-4	=	-1.07	0.52	4	4.90	0.30	8.17	0.09
	1-4	=1			4	4.40	0.35	6.80	0.15
	0-4	=1			5	12.60	0.03	18.06	0.00

Panel A reports coefficient estimates, corresponding standard errors (heteroskedasticity and autocorrelation adjusted using the Newey and West (1987) method) and R^2s from the forward premium regression (see Section II.B)

$$\Delta s_{t,t+1} = \alpha_j + \beta_j (if_{t-j}^{j,j+1} - if_{t-j}^{j,j+1^*}) + \varepsilon_{t-j,t+1},$$

using annual data sampled monthly. All regressions are run using exchange rate data over 1980–2010 (see Section II.A for a detailed description of the data). The columns labeled "SD" and "P-value" report simulated cross-sectional standard deviations of the estimated coefficient and two-sided P-values for the test β =1, respectively, under the Monte Carlo scheme described in Appendix A. Panel B reports tests of the hypotheses that β =1 and that the βs are equal for various horizons. The Lagrange Multiplier test statistics (LM Stat.) impose the relevant restrictions and the Wald test statistics (Wald Stat.) are based on the unrestricted parameter estimates. We report the restricted parameter estimate and associated standard error where relevant.

Table 3: Monte Carlo Results

Panel A: β_j = 1

	Forward Interest Rates			Long Horizon		
j	True R^2	Mean R^2	SD R^2	True R^2	Mean R^2	SD R^2
0	4.12	6.97	7.46	4.12	6.97	7.46
1	3.24	5.99	6.83	6.79	11.43	11.76
2	2.42	5.13	6.15	8.32	14.13	14.27
3	1.64	4.36	5.47	8.96	15.44	15.55
4	0.91	3.54	4.63	8.88	15.71	15.99

	Correlation of β_j							
	Forward Interest Rates				Long Horizon			
j	1	2	3	4	1	2	3	4
0	0.86	0.69	0.53	0.37	0.95	0.86	0.78	0.69
1		0.85	0.65	0.46		0.96	0.89	0.80
2			0.83	0.58			0.97	0.90
3				0.77				0.97

Panel B: β_j = 0

	Forward Interest Rates			Long Horizon		
j	True R^2	Mean R^2	SD R^2	True R^2	Mean R^2	SD R^2
0	0.00	3.07	4.09	0.00	3.07	4.09
1	0.00	3.05	4.05	0.00	5.63	7.14
2	0.00	3.01	3.99	0.00	7.62	9.35
3	0.00	2.95	3.91	0.00	9.00	10.78
4	0.00	2.78	3.72	0.00	9.87	11.63

	Correlation of β_j							
	Forward Interest Rates				Long Horizon			
j	1	2	3	4	1	2	3	4
0	0.87	0.69	0.54	0.37	0.95	0.87	0.79	0.70
1		0.85	0.65	0.45		0.96	0.89	0.81
2			0.82	0.56			0.97	0.90
3				0.76				0.97

Panel C: Hypothesis Tests

Hypothesis		LM Test			Wald Test		
$\beta_j = 1$	Level (%)	10	5	1	10	5	1
	Rejection (%)	14.28	5.31	0.24	38.72	28.70	14.71
β_j equal	Level (%)	10	5	1	10	5	1
	Rejection (%)	12.99	4.91	0.27	26.92	17.49	6.82

Table A.1 reports the results from a Monte Carlo simulation in which we generate 100,000 replications of 432 monthly observations from a model that imposes the expectations hypothesis of interest rates and either the expectations hypothesis of exchange rates, $\beta_j = 1$, or a random walk in exchange rates, $\beta_j = 0$. These observations are then aggregated to construct samples of 361 annual, monthly overlapping observations. (See Appendix A for a detailed description and Richardson and Smith (1992) for an analysis of the benefits of using overlapping observations.) Panels A and B report statistics on the coefficient estimates and R^2s from the forward premium regressions (see Section II.B)

$$\Delta s_{t,t+1} = \alpha_j + \beta_j (if_{t-j}^{j,j+1} - if_{t-j}^{j,j+1^*}) + \varepsilon_{t-j,t+1},$$

and the long-horizon regressions, after Chinn and Meredith (2005),

$$\Delta s_{t,t+j} = \alpha_j + \beta_j (i_{t,j} - i_{t,j}^*) + \varepsilon_{t,t+j}.$$

"True" refers to the analytical (infinite sample) value, and "Mean" and "SD" refer to the mean and standard deviation of the values across the simulations. For the test statistics, Panel C reports the percent of the simulations that reject the null hypothesis at the 10%, 5%, and 1% levels under $\beta_j = 1$.

Table 4: Real Exchange Rates and the Expectations Hypothesis of Exchange Rates

Panel A: Summary Statistics

Exchange Rate	Start Date	No. of Obs.	Mean	SD	1st Order Autocorr.	12th Order Autocorr.
USD/GBP	1/80	361	0.55	0.12	0.97	0.54
USD/DEM	1/80	361	0.18	0.16	0.98	0.69
USD/CHF	1/80	361	-0.27	0.16	0.98	0.69
USD/AUD	10/88	256	-0.34	0.15	0.98	0.70
USD/CAD	1/80	361	-0.21	0.11	0.98	0.83
USD/JPY	1/80	361	6.85	0.19	0.98	0.79
USD/NZD	10/88	256	-0.50	0.16	0.98	0.68
USD/NOK	2/86	288	-1.89	0.12	0.97	0.61
USD/SEK	1/93	205	-1.99	0.14	0.97	0.64

Correlations

Exchange Rate	USD /DEM	USD /CHF	USD /AUD	USD /CAD	USD /JPY	USD /NZD	USD /NOK	USD /SEK
USD/GBP	0.73	0.63	0.65	0.39	0.27	0.63	0.71	0.63
USD/DEM		0.96	0.79	0.39	0.60	0.80	0.86	0.91
USD/CHF			0.69	0.26	0.72	0.74	0.81	0.87
USD/AUD				0.86	0.03	0.93	0.89	0.74
USD/CAD					-0.09	0.68	0.82	0.55
USD/JPY						0.06	0.15	0.45
USD/NZD							0.83	0.81
USD/NOK								0.80

Panel B: Regression Results

Exchange Rate	α	Std. err.	ψ_1	Std. err.	ψ_2	Std. err.	R^2
USD/GBP	-2.78	1.99	-0.84	0.88			2.11
	25.57	6.86			-0.49	0.12	27.01
	25.68	6.12	-1.49	0.68	-0.54	0.10	33.44
USD/DEM	1.47	1.87	-0.71	0.71			1.77
	6.60	3.31			-0.33	0.12	18.24
	10.00	2.94	-1.69	0.53	-0.42	0.11	27.05
USD/CHF	4.83	2.32	-1.26	0.60			6.41
	-6.28	3.25			-0.29	0.13	13.55
	-4.38	3.21	-2.45	0.59	-0.45	0.11	33.56
USD/AUD	-0.43	3.02	-0.50	0.97			0.66
	-7.59	4.77			-0.24	0.14	8.99
	-15.52	6.56	-1.87	1.00	-0.36	0.14	15.82
USD/CAD	0.28	1.12	-0.20	0.65			0.14
	-2.98	2.29			-0.16	0.10	6.65
	-4.15	2.88	-0.78	0.80	-0.19	0.10	8.57
USD/JPY	11.63	2.10	-2.61	0.53			20.01
	145.85	74.68			-0.21	0.11	11.06
	191.28	60.37	-3.00	0.52	-0.26	0.09	36.85
USD/NZD	-0.25	5.57	-0.39	1.98			0.17
	-15.24	7.38			-0.32	0.15	15.13
	-26.82	8.94	-2.51	1.85	-0.42	0.12	20.96
USD/NOK	0.55	1.89	-0.14	0.66			0.11
	-72.26	28.11			-0.39	0.15	18.61
	-74.79	26.99	-0.36	0.55	-0.40	0.15	19.37
USD/SEK	0.00	2.18	-1.24	1.05			4.44
	-69.28	31.63			-0.35	0.16	17.40
	-85.06	24.15	-2.03	0.94	-0.43	0.12	28.41

Panel A reports summary statistics for log real exchange rates over the period January 1980 to January 2010 (with later start dates as dictated by data availability). Panel B reports coefficient estimates, corresponding standard errors (heteroskedasticity and autocorrelation adjusted using the Newey and West (1987) method) and R^2s from the estimation of the augmented forward premium regression (see Section IV.A for details):

$$\Delta s_{t,t+1} = \alpha + \psi_1 (i_{t,1} - i^*_{t,1}) + \psi_2 \, q_t + \varepsilon_{t,t+1},$$

using annual data sampled monthly.

Table 5: Decomposing Interest Rate Differentials

Panel A: Augmented Forward Interest Rate Regression I

Exchange Rate	j	α_j	Std. err.	ϕ_j	Std. err.	$\phi_{0,j}$	Std. err.	R^2
USD/GBP	1	-1.69	2.37	0.37	1.37	-1.10	0.84	4.27
	2	0.97	2.30	2.60	1.21	-0.78	0.68	16.44
	3	-0.84	2.16	1.14	1.45	-0.68	0.87	6.59
	4	0.32	1.97	2.03	1.22	-0.38	0.80	11.69
USD/DEM	1	-0.68	1.95	0.55	1.15	-2.32	0.93	9.79
	2	-0.28	2.17	0.32	1.37	-1.03	0.67	4.05
	3	-1.84	2.40	1.33	1.57	-0.96	0.71	8.42
	4	-3.20	2.45	2.40	1.55	-0.88	0.80	16.84
USD/CHF	1	2.53	2.67	-0.66	0.87	-2.40	1.03	11.32
	2	1.47	3.40	-0.30	1.13	-1.66	0.70	9.68
	3	-0.36	4.00	0.30	1.37	-1.49	0.65	12.50
	4	-1.84	4.04	0.83	1.42	-1.32	0.66	15.65

Panel B: Augmented Forward Interest Rate Regression II

	j	α_j	Std. Err.	β_j	Std. Err.	γ_j	Std. Err.	R^2
USD/GBP	0	25.68	6.12	-1.49	0.68	-0.54	0.10	33.44
	1	29.37	6.86	-2.18	1.11	-0.60	0.13	30.90
	2	22.93	8.15	1.31	1.30	-0.42	0.16	28.55
	3	25.32	6.79	1.16	0.99	-0.47	0.12	28.82
	4	26.29	5.65	2.20	0.88	-0.47	0.11	35.39
USD/DEM	0	10.00	2.94	-1.69	0.53	-0.42	0.11	27.05
	1	10.21	3.24	-1.45	1.00	-0.42	0.12	21.19
	2	9.73	3.84	-1.32	1.46	-0.39	0.12	20.13
	3	5.24	3.53	0.62	1.08	-0.31	0.11	18.66
	4	2.00	3.43	2.46	1.07	-0.28	0.11	26.36
USD/CHF	0	-4.38	3.21	-2.45	0.59	-0.45	0.11	33.56
	1	-5.18	2.99	-2.64	0.61	-0.54	0.11	26.62
	2	-3.70	3.55	-2.19	1.41	-0.45	0.13	19.30
	3	-6.21	4.28	-0.05	1.12	-0.29	0.11	13.55
	4	-9.01	4.18	1.31	1.03	-0.24	0.12	16.52

Panel A reports coefficient estimates, corresponding standard errors (heteroskedasticity and autocorrelation adjusted using the Newey and West (1987) method) and R^2s from the estimation of the bivariate regression of interest rate and forward interest rate differentials (see Section IV.B for details):

$$\Delta s_{t,t+1} = \alpha_j + \phi_j (if_{t-j}^{j,j+1} - if_{t-j}^{j,j+1^*}) + \phi_{0,j}[(i_{t,1} - i_{t,1}^*) - (if_{t-j}^{j,j+1} - if_{t-j}^{j,j+1^*})] + \varepsilon_{t,t+1},$$

using annual data sampled monthly. Panel B reports coefficient estimates, corresponding standard errors (heteroskedasticity and autocorrelation adjusted using the Newey and West (1987) method) and R^2s from the estimation of the bivariate regression of deviations from PPP and forward interest rate differentials (see Section IV.B for details):

$$\Delta s_{t,t+1} = \alpha_j + \beta_j (if_{t-j}^{j,j+1} - if_{t-j}^{j,j+1^*}) + \gamma_j q_t + \varepsilon_{t,t+1}.$$

All regressions are run using exchange rate data over 1980–2010 (see Section II.A for a detailed description of the data).