JOURNAL OF FINANCIAL AND QUANTITATIVE ANALYSIS Vol. 51, No. 3, June 2016, pp. 875–897

doi:10.1017/S0022109016000302

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 that 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.

I. Introduction

Well over 100 papers document, in some form or another, the forward premium anomaly, namely, that future exchange rate changes do not move on a one-for-one basis 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.

This paper looks at the forward premium anomaly 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

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a subset of four of the G10 currencies, for which we have sufficient interest rate data, over the time period 1980–2010. In stark contrast to current research on UIP, 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.

These results need to be interpreted with caution because both the coefficient estimates and R^2s are noisy, and inference is difficult due to small-sample concerns that we quantify in an extensive Monte Carlo exercise. Therefore, we extend the analysis by providing a plausible explanation for these findings that can be examined in a larger cross-section of currencies. Specifically, we explain why UIP 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 although interest rate differentials lead to capital flows from the carry trade, associated currency movements, purchasing power parity (PPP) violations, and the rejection of UIP; the buildup 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 that argue exchange rate changes are a function of two key state variables: the interest rate 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, and the deviation from PPP captures the reversal of this effect in the longer term as exchange rates revert to fundamentals.

To test this intuition, 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. Finally, we reconcile these empirical results with those from 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 5 years going back to 1976. Data for the term structure of zero-coupon interest rates are derived from London Interbank Offered Rate (LIBOR) data (with maturities of 6 and 12 months) and swap rates (2-, 3-, 4-, and 5-year semiannual swap rates).[1] Because swap data become available only 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 1 month to 5 years for this subset of countries.[2] Swap and LIBOR data are 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, 1-year spot interest rates and 1-year forward interest rates from years 1–2, 2–3, 3–4, and 4–5. For the remainder of the countries we compute continuously compounded, 1-year spot interest rates starting in Jan. 1980 (or later as dictated by data availability).

Using the exchange rate data, we compute annual changes in the natural log of the exchange rates with the U.S. dollar (USD) as the base currency, starting in Jan. 1980 (or later as dictated by the availability of interest rate data) and ending in Dec. 2010 (i.e., we examine changes in the USD/foreign currency 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 $t+1$.[3] Thus, the 4- to 5-year forward interest rate in Jan. 1976, the first observation, is matched with the annual exchange rate change from Jan. through Dec. 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 are sufficient interest rate data. Finally, we also combine this exchange rate data with Consumer Price Index data to construct real exchange rates for all country pairs. Further discussion of these series is postponed until Section III.

[1] 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- and 12-month maturities and semiannual swap rates for maturities of 24, 36, 48, and 60 months. Therefore, the only maturities we need to spline are 18, 30, 42, and 54 months. We maximize the smoothness of the spline function over these unknowns by minimizing the sum of squared deviations.

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

[3] 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 previously, these annual quantities are calculated on a monthly overlapping basis to maximize the information content of the empirical analysis.

To summarize, the final data set consists of annual exchange rate changes, with the first observation corresponding to calendar year 1980 and the last to calendar year 2010 (361 observations sampled monthly) for 5 of the 9 exchange rates, with start dates ranging from Feb. 1986 to Jan. 1993 for the other 4. For all countries we also have matched 1-year, spot interest rates covering a sample whose dates correspond 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 $j = 1, \dots, 4$, respectively (all with 361 observations). Panels A and B of Table 1 contain descriptive statistics for these variables.

B. Existing Evidence

The expectations hypothesis for exchange rates (also known as forward parity or the unbiasedness hypothesis) is commonly written as

$$\mathrm{E}_t s_{t+j} = f_t^j, \tag{1}$$

TABLE 1

Preliminaries

Panels A and B of Table 1 report summary statistics (mean, standard deviation (SD), 1st-order autocorrelation, 12th-order autocorrelation, and cross correlations) for annual changes in the natural log of the 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–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 $j = 0, \dots, 4$, 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. The U.S. Dollar (USD), British Pound (GBP), German Mark (DEM), Swiss Franc (CHF), Australian Dollar (AUD), Canadian Dollar (CAD), Japanese Yen (JPY), New Zealand Dollar (NZD), Norwegian Krone (NOK), and Swedish Krone (SEK) are the G10 currencies.

Panel A. Summary Statistics: Exchange Rates

					Autocorrelation	
Exchange Rate	Start Date	No. of Obs.	Mean (%)	SD (%)	1st Order	12th Order
USD/GBP	1/1980	361	−1.35	11.46	0.93	−0.02
USD/DEM	1/1980	361	0.67	12.71	0.93	0.10
USD/CHF	1/1980	361	1.59	12.53	0.92	0.00
USD/AUD	10/1988	256	0.63	12.26	0.93	−0.16
USD/CAD	1/1980	361	0.40	6.93	0.93	−0.05
USD/JPY	1/1980	361	3.20	11.98	0.93	0.11
USD/NZD	10/1988	256	0.77	13.54	0.94	−0.05
USD/NOK	2/1986	288	0.81	10.87	0.91	−0.26
USD/SEK	1/1993	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

(continued on next page)

TABLE 1 (continued)

Preliminaries

Panel B. Summary Statistics: Forward Interest Rate Differentials

$if^{j,j+1} - if^{*j,j+1}$	j	Mean (%)	SD (%)	Autocorrelation 1st Order	Autocorrelation 12th Order
United States–United Kingdom	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
United States–Germany	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
United States–Switzerland	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
United States–Australia	0	−2.09	1.97	0.98	0.67
United States–Canada	0	−0.62	1.32	0.95	0.60
United States–Japan	0	3.23	2.05	0.97	0.55
United States–New Zealand	0	−2.63	1.46	0.96	0.45
United States–Norway	0	−1.94	2.65	0.98	0.66
United States–Sweden	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

where s_{t+j} is the natural log of the spot price of foreign currency at time $t+j$, and f_t^j is the natural 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 (natural 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

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

where $\Delta s_{t,t+1} \equiv s_{t+1} - s_t$. Under 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.[4]

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 by Fama (1984). Although 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 (e.g., Bekaert and Hodrick (1993), Bekaert (1995), (1996), Bekaert, Hodrick, and Marshall (1997), Mark and Wu (1998), and Graveline (2006)).

As a first look at equation (2), Panel C of Table 1 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, confirming the well-known negative relation between exchange rates and interest rate differentials. Although the estimates are fairly noisy, tests of the null hypothesis that the coefficients equal 1 can be resoundingly rejected for 7 of the 9 U.S.–G10 currency pairs.

The low R^2s in most of the regressions are also notable, and this feature is both disappointing and puzzling.[5] 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. Because inflation is fairly predictable (see, e.g., Fama and Gibbons (1984)), and inflation differentials are a fundamental driver of exchange rates under PPP, 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), forward parity, is almost always cast in terms of interest rate differentials and then tested using equation (2), UIP. 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 use equation (1) to define expected changes in future exchange rates as the difference between two forward exchange rates. That is,

(3) $$\mathrm{E}_t[\Delta s_{t+j,t+k}] \quad = \quad f_t^k - f_t^j,$$

[4]See, for example, 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)).

[5]This low R^2 result parallels Meese and Rogoff (1983), who 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. For a theoretical analysis of this issue, see Engel and West (2005).

where $k > j$. Under the unbiasedness hypothesis, 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; that is,

$$\text{(4)} \qquad \mathrm{E}_t[\Delta s_{t+j,t+k}] \;=\; k\,(i_{t,k} - i^*_{t,k}) - j\,(i_{t,j} - i^*_{t,j}).$$

Rearranging the interest rate differential terms in equation (4), and using the definition of forward interest rates,[6] we get

$$\text{(5)} \qquad \begin{aligned} \mathrm{E}_t[\Delta s_{t+j,t+k}] &= (k\,i_{t,k} - j\,i_{t,j}) - (k\,i^*_{t,k} - j\,i^*_{t,j}) \\ &= (k-j)\left(if_t^{j,k} - if_t^{j,k*}\right), \end{aligned}$$

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 to 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,

$$\text{(6)} \qquad \Delta s_{t,t+1} \;=\; \alpha_j + \beta_j\left(if_{t-j}^{j,j+1} - if_{t-j}^{j,j+1*}\right) + \varepsilon_{t-j,t+1}.$$

Under the expectations hypothesis of exchange rates, the annual exchange rate change from t to $t+1$ should move on a one-for-one basis 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), Panel A of Table 2 provides estimates over different horizons and across a subset of the G10 currencies for tests of the expectations hypothesis of exchange rates.[7] This analysis requires a history of long-term forward interest rates, and, as described in Section II.A, we have such data for the United States, the United Kingdom, Switzerland, and Germany. In contrast to Panel C of Table 1, 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 1 to 4 years yield coefficients of 0.68, 0.76, 2.02, and 3.17 for the USD/DEM exchange rate.

[6]The annualized forward interest rate is defined as $if_t^{j,k} \equiv (k\,i_{t,k} - j\,i_{t,j})/(k-j)$.

[7]Using different specifications, Chinn and Meredith (2005), Bekaert, Min, and Ying (2007), and Chinn and Quayyum (2012) also analyze implications for UIP at short and long horizons.

TABLE 2

The Expectations Hypothesis of Exchange Rates

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

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

using annual data sampled monthly. All regressions are run using exchange rate data over the period 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 2-sided p-values, respectively, for the tests $\beta = 1$ and $\beta = 0$ under the Monte Carlo scheme described in the Appendix. Panel B reports tests of the hypotheses that $\beta = 1$ and that the βs are equal for various horizons. The LM 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.

Panel A. Regression Results

						$\beta_j = 1$		$\beta_j = 0$		
Exchange Rate	j	α_j	Std. Err.	β_j	Std. Err.	SD	p-Value (%)	SD	p-Value (%)	R^2
USD/GBP	0	−2.78	1.99	−0.84	0.88	0.91	4.73	0.91	33.00	2.11
	1	−0.38	2.35	0.92	1.30	1.01	93.20	1.00	33.47	0.94
	2	2.34	2.13	3.41	1.01	1.14	4.06	1.11	0.45	14.61
	3	0.44	1.77	1.94	1.02	1.35	47.54	1.30	13.27	5.22
	4	1.09	1.90	2.54	0.87	1.71	36.34	1.65	12.04	11.28
USD/DEM	0	1.47	1.87	−0.71	0.71	0.97	7.87	0.97	43.69	1.77
	1	−0.27	1.95	0.68	1.20	1.08	75.19	1.07	49.87	0.96
	2	−0.52	2.17	0.76	1.29	1.22	83.68	1.20	50.75	0.83
	3	−2.50	2.37	2.02	1.38	1.44	46.50	1.39	14.22	5.33
	4	−4.02	2.37	3.17	1.32	1.83	23.29	1.77	7.34	14.16
USD/CHF	0	4.83	2.32	−1.26	0.60	1.04	3.41	1.03	20.75	6.41
	1	2.06	2.52	−0.16	0.97	1.15	29.28	1.14	87.96	0.08
	2	0.23	3.51	0.42	1.22	1.30	64.00	1.27	72.72	0.38
	3	−2.92	3.97	1.40	1.32	1.53	78.27	1.49	33.11	3.85
	4	−4.99	3.61	2.07	1.21	1.94	57.24	1.89	26.28	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

The results for the USD/GBP and USD/CHF exhibit similar patterns. These results are quite different from the significant negative coefficients that plague Panel C of Table 1 (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. Panel B of Table 2 reports tests that the coefficients are equal and that the coefficients are equal to 1. 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 1 (labeled "= 1"). In the former case, columns 4 and 5 provide the restricted coefficient estimate under the Lagrange multiplier (LM) test and the associated standard error. Turning to the results, the LM tests yield only two rejections at the 10% level, both for the hypothesis that the coefficients equal 1 at all horizons. In contrast, the Wald tests yield rejections in all but four cases.[8] Thus, there is definitely evidence, although 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, for monthly overlapping data. Therefore, one of the difficulties in studying multistep-ahead forecast regressions like those specified in equation (6) is the availability of data. Although sophisticated econometrics have somewhat alleviated the problem (Hansen and Hodrick (1980), 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), 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 on 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. Panel A of Table 1 shows that exchange rates are much more variable than interest rate differentials, and they are also relatively unpredictable (see Panel A of Table 2). 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 in Section II.D that also provides a comparison with alternative long-horizon methodologies for forward premium regressions (e.g., see Chinn and Meredith (2005), Chinn and Quayyum (2012)).

For now, Panel A of Table 2 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 tests that $\beta = 1$ and $\beta = 0$ (in the column "p-value"), that is, the percentage of the replications in which the absolute magnitude of deviation of

[8] We employ both the LM 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 overrejects while the LM test underrejects, results that are consistent with our simulation evidence discussed later.

the estimated coefficient from 1 equals or exceeds the deviation for the estimated coefficient from the actual data. For these calculations, we simulate under the null hypothesis in each case (i.e., $\beta = 1$ or $\beta = 0$) 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 simulated p-values are broadly consistent with those from standard hypothesis tests of UIP of the individual coefficients using the estimated standard errors. Specifically, the short-horizon ($j = 0$) coefficients are statistically significantly different from 1, as is the coefficient for $j = 2$ for the USD/GBP. However, the hypothesis that the coefficients are different from 0 can be rejected only for $j = 2$ for the USD/GBP using the simulated p-values, whereas this hypothesis is rejected for a number of other long-horizon coefficients at the 10% level using the estimated standard error.

As a final comment on the evidence, note that in Table 2 the regression R^2s have a tendency to increase with the horizon. Whereas the dependent variable, that is, 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 later 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 restricted vector autoregressive specification 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; that is, we assume UIP holds. In the other experiment, we assume exchange rates follow a random walk; that is, 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. A detailed description of the simulation procedure is given in the Appendix.

Columns 6–9 of Panel A of Table 2 discussed in Section II.C and Table 3 report the key results. Panel A of Table 3 reports results under the expectations hypothesis of exchange rates ($\beta_j = 1$). It compares the coefficients and 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), and it also examines the properties of both single-horizon and cross-horizon test statistics in the former regression. The statistics in this panel, and in the remainder of Table 3, are calculated from

TABLE 3

Monte Carlo Results

Table 3 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, $\beta_j = 1$ (Panel A), or a random walk in exchange rates, $\beta_j = 0$ (Panel B). These observations are then aggregated to construct samples of 361 annual, monthly overlapping observations. (See the Appendix for a detailed description and Richardson and Smith (1992) for an analysis of the benefits of using overlapping observations.) We report statistics on the coefficient estimates and R^2s from the forward premium regressions (see Section II.B),

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

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

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

"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, respectively. For the test statistics, we report the percentage of the simulations that reject the null hypothesis at the 10%, 5%, and 1% levels.

Panel A. $\beta_j = 1$

	Forward Interest Rates					Long Horizon				
j	Mean β_j	SD β_j	True R^2	Mean R^2	SD R^2	Mean β_j	SD β_j	True R^2	Mean R^2	SD R^2
0	0.97	0.91	4.12	6.97	7.46	0.97	0.91	4.12	6.97	7.46
1	0.91	1.01	3.24	5.99	6.83	0.93	0.93	6.79	11.43	11.76
2	0.87	1.14	2.42	5.13	6.15	0.90	0.95	8.32	14.13	14.27
3	0.82	1.35	1.64	4.36	5.47	0.86	0.96	8.96	15.44	15.55
4	0.76	1.71	0.91	3.54	4.63	0.83	0.98	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

Single-Horizon Hypothesis Tests: Forward Interest Rates

			Level (%)		
j	Mean Z	SD Z	10	5	1
0	−0.04	1.42	23.35	15.97	7.04
1	−0.11	1.42	23.53	16.20	7.22
2	−0.15	1.43	23.84	16.40	7.34
3	−0.17	1.44	24.31	16.77	7.50
4	−0.19	1.43	23.95	16.53	7.37

Cross-Horizon Hypothesis Tests: Forward Interest Rates

Hypothesis		LM Test			Wald Test		
$\beta_j = 1$	Level (%)	10	5	1	10	5	1
	Rejection (%)	13.23	4.64	0.16	37.69	27.73	14.03
β_j equal	Level (%)	10	5	1	10	5	1
	Rejection (%)	12.23	4.46	0.20	26.05	16.74	6.36

(continued on next page)

TABLE 3 (continued)

Monte Carlo Results

Panel B. $\beta_j = 0$

	Forward Interest Rates					Long Horizon				
j	Mean β_j	SD β_j	True R^2	Mean R^2	SD R^2	Mean β_j	SD β_j	True R^2	Mean R^2	SD R^2
0	0.00	0.91	0.00	3.07	4.09	0.00	0.91	0.00	3.07	4.09
1	0.00	1.00	0.00	3.05	4.05	0.00	0.92	0.00	5.63	7.14
2	0.00	1.11	0.00	3.01	3.99	0.00	0.93	0.00	7.62	9.35
3	0.00	1.30	0.00	2.95	3.91	0.00	0.94	0.00	9.00	10.78
4	0.00	1.65	0.00	2.78	3.72	0.00	0.95	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

Single-Horizon Hypothesis Tests: Forward Interest Rates

			Level (%)		
j	Mean Z	SD Z	10	5	1
0	0.00	1.42	23.30	15.91	7.00
1	−0.01	1.42	23.39	15.92	7.00
2	−0.01	1.41	23.23	15.86	6.95
3	0.00	1.41	23.13	15.77	6.84
4	0.00	1.40	22.82	15.34	6.71

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.

The coefficient estimates in both regressions exhibit a small downward bias that increases in the horizon. However, the magnitudes of this bias are close to an order of magnitude smaller than the deviations of the estimated coefficients from 1. More important, 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 1 to 4 years, the bias, that is, the difference between the mean R^2 from the simulations and the true R^2, ranges from 2.75% (5.99% simulated vs. 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 vs. 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.[9]

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

In terms of the single-horizon test statistics in equation (6), the more detailed results confirm those presented in Table 2. We report the means and standard deviations of the Z-statistic for testing the null hypothesis, and the percentage of simulations in which the null is rejected at significance levels ranging from 1% to 10%.[10] Estimated standard errors are somewhat understated; thus, the test statistic tends to be overstated in magnitude, and the null is rejected somewhat too frequently.

Inferences from the simulation results for the cross-horizon Wald and LM tests are more nuanced. Again we report the percentage of the simulations in which the null hypothesis is rejected. Consistent with Berndt and Savin (1997) and Bekaert and Hodrick (2001), the Wald test substantially overrejects the null hypothesis, whereas the LM test tends to underreject the null hypothesis, especially for high significance levels. For example, for the hypothesis $\beta_j = 1$ across all five horizons, the LM test rejects 4.6% and 0.2% of the time at the 5% and 1% levels, respectively, whereas the Wald test rejects the null hypothesis in 27.7% and 14.0% of the simulations. Moreover, whereas the LM test performs similarly for both the $\beta_j = 1$ and $\beta_j = 0$ hypotheses, the small-sample properties of the Wald test are much worse for the hypothesis $\beta_j = 1$.

Panel B of Table 3 reports the results under the assumption that the exchange rate follows a random walk ($\beta_j = 0$). The biases in the coefficient estimates are negligible in either regression specification, but 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 an R^2 bias that ranges from 2.78% to 3.05%, whereas the biases in the long-horizon regressions increase with the horizon up to 9.87%. Finally, the properties of the single-horizon test statistics are similar to those reported in Panel A.

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.

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. Although this story is just one potential explanation for the observed behavior of UIP using spot and forward interest rate differentials, we provide additional supporting empirical evidence. Specifically, we show how the recent empirical exchange rate determination models of Jorda and Taylor (2012), which depend on

[10] By definition, the Z-statistic should have a mean of 0 and a standard deviation of 1 under the null hypothesis.

two key state variables (i.e., the interest rate differential and the deviation of the current exchange rate from that implied by PPP), are consistent with the standard forward premium anomaly and our contrasting results using forward interest rate differentials.

There have been a number of recent papers in the area of exchange rate determination that 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, cause 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 (2014), Berge, Jorda, and Taylor (2010), and 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 (2008), Plantin and Shin (2010)). Moreover, theories based on speculative dynamics (e.g., Plantin and Shin (2010)) and existing empirical work (e.g., Brunnermeier et al. (2008), Jorda and Taylor (2012)) imply that this carry risk should be increasing in the deviation from 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 (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. 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 natural log real exchange rate

$$q_t = s_t + \left(z_t^* - z_t\right), \tag{7}$$

where q and s are the natural log real and nominal exchange rates, respectively, and z and $z*$ denote the natural 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 of Jorda and Taylor (2012) combines the standard forward premium regression in equation (2) with the real exchange rate in equation (7):

$$(8) \qquad \Delta s_{t,t+1} \quad = \quad \alpha + \psi_1 \left(i_{t,1} - i^*_{t,1} \right) + \psi_2 q_t + \varepsilon_{t,t+1}.$$

Jorda and Taylor motivate the real exchange rate variable as the deviation from the fundamental equilibrium exchange rate, although they do not provide a motivating theoretical model because they are primarily interested in forecasting and the associated trading strategies. They estimate various specifications employing the two variables in equation (8) using monthly data across multiple exchange rates for the period 1986–2008 and report results consistent with ours.

A. Exchange Rate Determination: Evidence

Panel A of Table 4 presents summary statistics for the natural log real exchange rate series for the 9 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 (8) using the deviation of the exchange rate from PPP. We estimate regressions of annual exchange rate changes (overlapping monthly) on the natural log real exchange rate and the interest rate differential at the beginning of the year (and special cases thereof). The results are reported in Panel B of Table 4. 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 natural log real exchange rate, and the final line reports the results from the full specification.

The first notable result in Panel B of Table 4 is that, both alone and in the full specification, the natural log real exchange rate appears with a negative and statistically significant coefficient for all 9 currency pairs. This negative coefficient is consistent with the intuition from the explanation provided earlier. When the real exchange rate is high, that is, 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

TABLE 4

Real Exchange Rates and the Expectations Hypothesis of Exchange Rates

Panel A of Table 4 reports summary statistics for natural log real exchange rates over the period Jan. 1980 to Jan. 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–West (1987) method), and R^2s from the estimation of the augmented forward premium regression (see Section III.A for details),

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

using annual data sampled monthly.

Panel A. Summary Statistics

Exchange Rate	Start Date	No. of Obs.	Mean	SD	Autocorrelation: 1st Order	Autocorrelation: 12th Order
USD/GBP	1/1980	361	0.55	0.12	0.97	0.54
USD/DEM	1/1980	361	0.18	0.16	0.98	0.69
USD/CHF	1/1980	361	−0.27	0.16	0.98	0.69
USD/AUD	10/1988	256	−0.34	0.15	0.98	0.70
USD/CAD	1/1980	361	−0.21	0.11	0.98	0.83
USD/JPY	1/1980	361	6.85	0.19	0.98	0.79
USD/NZD	10/1988	256	−0.50	0.16	0.98	0.68
USD/NOK	2/1986	288	−1.89	0.12	0.97	0.61
USD/SEK	1/1993	205	−1.99	0.14	0.97	0.64

Exchange Rate	Correlations: 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

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 earlier, 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 the 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, that is, the coefficient becomes more negative, for all 9 of the exchange rates. Specifically, the average coefficient, ψ_1, in equation (8) 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. That is, 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.

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, which generates novel and striking empirical results. Motivated by these results, in Section III.A we offer an explanation, namely, that past forward interest rate differentials at different horizons pick up the two opposing effects (i.e., the carry effect and reversion to PPP) to different degrees, and new empirical evidence consistent with this explanation. What remains is to reconcile these two sets of empirical results.

First, note that the carry-trade component of exchange rates explains why PPP does not hold and why the coefficient on interest rate differentials is negative in equations (2) and (8). However, if there is a positive probability that exchange rates will revert back to PPP, 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.

Of course, replacing the forward interest rate differentials of the forward premium regressions with crash-specific variables (e.g., deviations in PPP) improves 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.

To verify this intuition, 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, Panel A of Table 5 presents results for the regression

$$(9) \quad \Delta s_{t,t+1} = \alpha_j + \varphi_{0,j}\left[\left(i_{t,1} - i^*_{t,1}\right) - \left(if^{j,j+1}_{t-j} - if^{j,j+1*}_{t-j}\right)\right] + \varphi_j\left(if^{j,j+1}_{t-j} - if^{j,j+1*}_{t-j}\right) + \varepsilon_{t,t+1},$$

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 $\varphi_{0,j}$ coefficients are all negative and declining in magnitude as the horizon increases. The R^2s are quite impressive.

The positive and increasing coefficients on the forward interest rate differentials are capturing the probability and magnitude of a reversion of the currency to

TABLE 5

Decomposing Interest Rate Differentials

Panel A of Table 5 reports coefficient estimates, corresponding standard errors (heteroskedasticity and autocorrelation adjusted using the Newey–West (1987) method), and R^2s from the estimation of the bivariate regression of interest rate and forward interest rate differentials (see Section III.B for details),

$$\Delta s_{t,t+1} = \alpha_j + \varphi_j\left(if^{j,j+1}_{t-j} - if^{j,j+1^*}_{t-j}\right) + \varphi_{0,j}\left[\left(i_{t,1} - i^*_{t,1}\right) - \left(if^{j,j+1}_{t-j} - if^{j,j+1^*}_{t-j}\right)\right] + \varepsilon_{t,t+1},$$

using annual data sampled monthly. Panel B reports coefficient estimates, corresponding standard errors (heteroskedasticity and autocorrelation adjusted using the Newey–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):

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

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

Panel A. Augmented Forward Interest Rate Regression I

Exchange Rate	j	α_j	Std. Err.	φ_j	Std. Err.	$\varphi_{0,j}$	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

Exchange Rate	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

PPP, and the negative coefficients on the forecast error in the exchange rate regression are capturing the carry-trade effect. Thus, the negative $\varphi_{0,j}$ explains why the forward premium anomaly exists from a statistical viewpoint, that is, why we get negative coefficients and low R^2s in Panel C of Table 1. 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.

To better understand these results, we estimate an analogous regression to equation (8), 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):

$$\Delta s_{t,t+1} = \alpha_j + \beta_j \left(if_{t-j}^{j,j+1} - if_{t-j}^{j,j+1*} \right) + \gamma_j q_t + \varepsilon_{t,t+1}. \tag{10}$$

These results are reported in Panel B of Table 5 for the three available currencies (DEM, GBP, CHF), relative to the USD, over the horizons $j = 0, \ldots, 4$. The horizon $j = 0$ is equivalent to the regression specification in equation (8), with results also provided in Panel B of Table 4.

Panel B of Table 5 provides two pieces of evidence in support of our explanation for exchange rate determination. First, 6 of 12 coefficients on the forward interest rate differential for horizons $j = 1, \ldots, 4$ are now negative (3 significantly so). For the regressions in Table 2 that did not include the PPP deviation variable, 11 of 12 coefficients on the forward interest rate differential were positive. Recall that the forward interest rate differential has information about both components, the carry effect and the reversion to PPP, although the latter dominates at longer horizons. Therefore, the reason that the coefficients flip signs is that in the regression specification in equation (10), q_t proxies for the reversion-to-PPP component, leaving just the carry effect. As documented in Panel C of Table 1, the carry effect has a negative sign.

Second, and equally important, in contrast to Table 2, Panel B of Table 5 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 is higher for the regression specification in equation (10) with $j = 0$ than not only the $j = 1, \ldots, 4$ horizons but also the alternative forward interest rate differential specifications given by either equation (6) or equation (9).

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 these results and the existing literature, we provide an intuitive model of exchange rate determination, further empirical evidence consistent with this intuition, and a reconciliation of the two sets of empirical evidence. The key insight is that exchange rate changes reflect two distinct but related phenomena: a carry-trade effect associated with interest rate differentials, which pushes exchange rates in the opposite direction to that predicted by UIP, and reversion to fundamentals at longer horizons. Forward interest rate differentials at different horizons pick up these opposing effects to different degrees, yielding horizon-dependent coefficients and R^2s. We decompose these two effects using both forward interest rate differentials and shocks to these differentials, along with spot interest rate differentials and real exchange rates.

Appendix. Monte Carlo Analysis

The purpose of the Monte Carlo analysis is to assess the small-sample properties of our specification and also compare them to those of the alternative long-horizon regressions.

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:[11]

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

where

(A-2)
$$\varepsilon_{t,t+1} \equiv \begin{bmatrix} \varepsilon^s_{t,t+1}\\ \varepsilon^1_{t,t+1}\\ \varepsilon^2_{t,t+1}\\ \vdots\\ \varepsilon^{59}_{t,t+1}\\ \varepsilon^{60}_{t,t+1} \end{bmatrix} \sim \text{MVN}(0, \Sigma).$$

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

(A-3)
$$\Sigma = \left[\begin{array}{c|ccccc}
\sigma^2_s & 0 & \cdots & 0 & \cdots & 0\\
\hline
 & \sigma^2_i & \cdots & \upsilon^{t-1}_{ij}\rho_{ij}\sqrt{\upsilon^{t-1}_i}\sigma^2_i & \cdots & \upsilon^{59}_{ij}\rho_{ij}\sqrt{\upsilon^{59}_i}\sigma^2_i\\
 & & \ddots & & & \vdots\\
 & & & \upsilon^{t-1}_i\sigma^2_i & & \upsilon^{59-t}_{ij}\rho_{ij}\sqrt{\upsilon^{t+58}_i}\sigma^2_i\\
 & & & & \ddots & \vdots\\
 & & & & & \upsilon^{59}_i\sigma^2_i
\end{array}\right].$$

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

The variances of the shocks to forward interest rate differentials decline in maturity and the correlations between the shocks to forward interest rate 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 0 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 results we report are for the values $\rho = 0.95$, $\sigma_s = 3.4641\%$, $\sigma_i = 0.0289\%$, $\upsilon_i = 0.995$, $\rho_{ij} = 0.995$, and $\upsilon_{ij} = 0.99$, but the inferences drawn from the Monte Carlo analysis are not sensitive to the precise choice of the parameters.

Define the state vector as

$$y_{t+1} = \begin{bmatrix} \Delta s_{t,t+1} \\ i_{t+1,1} - i^*_{t+1,1} \\ if^{1,2}_{t+1} - if^{1,2*}_{t+1} \\ \vdots \\ if^{58,59}_{t+1} - if^{58,59*}_{t+1} \\ if^{59,60}_{t+1} - if^{59,60*}_{t+1} \end{bmatrix}. \tag{A-4}$$

Equations (A-1)–(A-2) imply that $y_{t+1} \sim \text{MVN}(0, \Omega)$, where Ω is a function of ρ and Σ. The simulation procedure is as follows:

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

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: $\Delta s_{t,t+1} = \varepsilon^s_{t,t+1}$. Finally, we repeat the previous analyses, 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.

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