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# Uncovered interest-rate parity over the past two centuries

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A B S T R A C T

We study the validity of uncovered interest-rate parity by constructing ultra-long time series that span two centuries. The forward-premium regressions yield positive slope estimates over the whole sample. The estimates become negative only when the sample is dominated by the period of 1980s. We also find that large interest-rate differentials have significantly stronger forecasting powers for currency movements than small interest-rate differentials. Furthermore, when we regress domestic currency returns on foreign bonds against returns on domestic bonds as an alternative test of the parity condition, the null hypotheses of zero intercept and unit slope cannot be rejected in most cases. These results are consistent with a world in which expectations formation is highly imperfect and characterized on the one hand by slow adjustment of expectations to actual regime changes and on the other by anticipations for extended periods of regime changes or other big events that never materialize. An historical account of expected and realized regime changes adds credence to this explanation and illustrates how uncovered interest-rate parity holds over the very long haul but nevertheless can be deviated from over long periods of time due to ex post-expectation errors.


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## 1. Introduction

Uncovered interest-rate parity (*UIP*) is one of three key international financial relations that are used repeatedly in the fields of international finance and open-economy macroeconomics in both model construction and other analytical work. The other two, purchasing power parity (*PPP*) and real interest-rate equality, which a decade ago appeared to be of questionable empirical applicability, have now been at least somewhat rehabilitated. Uncovered interest-rate parity, however, has not been nearly as fortunate.

Indeed, one of the most puzzling feature of exchange rate behavior since the advent of floating exchange rates in the early 1970s is the tendency for countries with high interest rates to see their currencies appreciate rather than depreciate as *UIP* would suggest. This *UIP* puzzle, known in its other guise as "the forward-premium puzzle," is now so well documented that it has taken on the aura of a stylized fact and as a result has spawned a second generation of papers attempting to account for its existence (see, for instance, Fama, 1984; Hodrick, 1987; Bekaert and Hodrick, 1993; Bekaert, 1995; Dumas and Solnik, 1995; Engel, 1996; Flood and Rose, 1996; Bansal, 1997; Bakshi and Naka, 1997; Backus et al., 2001; Chinn and Meredith, 2005; Bekaert et al., 2007; Brennan and Xia, 2006).

Most of the empirical investigation focuses on the sample period of the late 1970s and the 1980s, a period dominated by the persistent appreciation of the dollar. We conjecture that the negative results are in part driven by the unique features of this sample period. We further conjecture that while the market tolerates small deviations from *UIP* for a relatively long period of time due to market frictions such as transaction costs, *UIP* – like *PPP* – will hold much better over the long run and reversions to the parity condition will become stronger and more obvious when the deviations are large.

We test these hypotheses with a careful selection of both data and methodology. First, we construct ultra-long time series on two currency pairs, the French franc versus the pound sterling and the US dollar versus sterling. The time series span two centuries so that our tests will be free from any local features of a short sample period. Equipped with these long data series, we run the forward-premium regression for both the whole sample and some interesting subsample periods. We also run the regression using rolling windows to see how the estimates vary with the sample periods. Second, to test the hypothesis that small deviations from *UIP* are tolerated while large deviations are likely to be followed by reversions to parity, we apply the principle of extreme sampling and run regressions conditional on large and small *UIP* deviations to detect differences in the regression slopes. We investigate these phenomena further using a non-linear specification that allows smooth transition from small to large deviations.

The results confirm our hypotheses. First, we run forward-premium regressions of depreciation rates on nominal interest-rate differentials. *UIP* implies that the regression slope should be one while traditional evidence often generates negative estimates. Our regressions over the long time series generate results much more in accord with the expectation hypothesis: The regressions slopes are positive for both currency pairs and the slope estimate is not significantly different from one for franc-sterling. Hence, the uncovered interest-rate parity holds much better over the long run than suggested by traditional evidence over short samples.

Second, our rolling window regressions confirm our conjecture that the negative slope estimates are mainly a special feature of the late 1970s and the 1980s. Indeed, once the start of the sample period switches to the early 1970s, the regression slope estimates become negative for both currency pairs. In all three countries, this was a period of substantial inflation, the culmination of an historically unprecedented three-decade period in which inflation trended up continually and reached historic peacetime highs. The first break in the process came in 1979 in the UK following the Thatcher election and the ensuing move to much tighter monetary policies. A similar shift began in the US shortly after Paul Volcker became Federal Reserve Chairman and gained momentum following the 1980 Reagan election. The shift in France came several years later. Nevertheless, in all three countries, the public remained skeptical for several years after the actual policy changes. Expected inflation rates remained high several years after the actual inflation rates had decreased significantly. Therefore, we argue that the negative regression slopes during this sample period are mainly a result of a failure of expectations to adjust over an extended period of time to the regime switch.

Third, when we run regressions conditional on large deviations, the regression slopes for both pairs of currencies are not significantly different from unity, thus confirming our conjecture that although the market tolerates small deviations from *UIP* for a relatively long period of time, *UIP* holds over the long run and the reversion to parity becomes stronger when the deviation is large.

The forward-premium regression is formulated based on the martingale hypothesis on the forward exchange rates. Linking that hypothesis with the covered interest-rate parity leads to the test of *UIP*. Another perspective from which to view *UIP* is the hypothesis that investing in foreign and domestic bonds should generate the same expected returns when they are computed under the same currency. Based on this perspective, we propose an alternative test of *UIP* by comparing the returns on investing in the foreign and domestic bonds. We regress the return on the foreign bond in domestic currency, which is the sum of the foreign interest rate and the currency depreciation rate, against the return on the domestic bond, which is simply the domestic interest rate. Under the null hypothesis of *UIP*, the intercept estimate should be zero and the slope estimate should be one. The regression results cannot reject either hypothesis in most cases. Furthermore, when we plot and compare the two return series, we observe that their differences are insignificantly different from zero on average, although large deviations can occur and persist during some sample periods.

Finally, to better understand the behavior of *UIP* during the past two centuries, we decompose the *UIP* deviation into two components: (1) the deviation from real interest-rate equality and (2) the deviation from purchasing power parity. An historical account of the major *UIP* deviations indicates that during the nineteenth century these deviations are mainly due to deviations from real interest-rate equality, but the *UIP* deviations during the more modern period are mainly driven by deviations from purchasing power parity. Furthermore, no matter whether they are from nominal or real factors, most of the deviations can be attributed to one of the two following problems: (1) A peso problem, where the investors are anticipating a large event that only materializes after the specific sample period and hence can only be captured by an ultra-long sample, or (2) a missed expectation problem, e.g., a regime or policy switch that investors fail to realize is happening for an extended period of time. Both problems become severe when the data set only covers a relatively short sample period. The most efficient way to deal with these problems is to construct an ultra-long sample, which we do in this paper.

Taken together, our results suggest that *UIP* works better (1) at long investment horizons (e.g., yearly versus monthly), (2) over long time periods (e.g., centuries versus decades), (3) in the presence of large deviations, and (4) in terms of holding-period returns. All four cases reduce the potential impact of random noises, expectation errors, and/or missed expectations. The first three cases reduce the errors via averaging. For the last case, if there is a temporary expectation error producing the currency movement, its impact on the forward-premium regression is larger than its impact on the holding-period-return regression.[2]

In an influential paper, Fama (1984) attributes the behavior of forward and spot exchange rates to a time-varying risk premium. Fama shows that a negative slope estimate from the *UIP* regression implies that the risk premium on a currency must (1) be negatively correlated with its expected rate of depreciation and (2) have greater variance. Modern currency pricing models, e.g., Backus et al. (2001) and Leippold and Wu (2007), can accommodate flexible enough risk-premium specifications to generate a negative regression slope, but the implied market price of risk often varies too much to seem plausible. Our empirical investigation indicates that it is not the market price of risk that varies wildly over time, but the regression estimates based on short samples that we cannot rely on.

Nevertheless, even with the ultra-long sample and with non-linear regressions or extreme sampling techniques, we find that the overall predictive performance of *UIP* is rather poor, especially over shorter periods and for small interest-rate differentials. The universally low *R*-squares from these predictive regressions are consistent with market efficiency, but they also show that if there is a stylized fact about *UIP*, it is not the anomalous negative relation between interest-rate differentials and exchange rate depreciation rates observed in the 1980s, but the fact that there is very often little relationship one way or the other. Especially over short periods, exchange rates move very much independently of the interest-rate differentials between the two countries.

[2] For a review of corroborative evidence, see Chinn (2006).

In related literature, Baillie and Bollerslev (2000), Bekaert and Hodrick (2001), and Bekaert et al. (2007) show how small sample can bias the regression tests to overwhelmingly reject the expectation hypothesis even when it holds. Lewis (1989), Gourinchas and Tornell (2004), and Bacchetta and van Wincoop (2009), among others, discuss the effect of changing beliefs and expectation errors on exchange rates. The peso problem has also been repeatedly raised in the literature with regard to how it can dramatically alter the results on expectation-hypothesis regressions. Early reference can be traced to Fisher (1907). Recent references include Lewis (1988), Engel and Hamilton (1990), Kaminsky (1993), Evans and Lewis (1995), Hallwood et al. (2000), Farhi and Gabaix (2008), Burnside et al. (2008), and Campbell et al. (2009). The literature has proposed statistical methods to correct for small-sample bias Bekaert et al. (1997), but the ultimate remedy for peso problem is to use an ultra-long sample.

The remainder of the paper is organized as follows. The next section describes the construction of the long data series. Section 3 presents the results from forward-premium regressions and regressions on holding-period returns. Section 4 considers the extreme sampling and smooth transition non-linear regressions. Section 5 performs an historical analysis over the past two centuries. Section 6 concludes.

## 2. Data construction

The data set consists of annual observations of dollar-sterling and franc-sterling exchange rates, as well as long-term and short-term interest rates for France, the United Kingdom (UK), and the United States (US). The data are compiled from several sources. Refer to Appendix A and Lothian and Taylor (1996) for details on the exchange rate series and Lothian (2000) for details on the interest-rate series.

The dollar-sterling exchange rate data span more than two centuries (209 years) from 1791 to 1999. The franc-sterling exchange rate starts in 1802 with 198 years of data. The long-term interest rates for all three countries and the short-term series for the UK and France start in 1800. The short-term interest rate for the US starts in 1831. The short-term interest-rate series for France have missing observations from 1914 through 1924. The ending point of all data series is 1999.

Fig. 1 plots the two exchange-rate series for the past two centuries. Prior to 1914, franc-sterling shows almost no movement. This stability is a product of the specie standards to which both countries adhered – a bimetallic standard pre-1875 and gold thereafter for France and gold from 1821 on for the UK. Dollar-sterling over this period is more variable, but this is mainly due to behavior in two wartime episodes – the lower US than UK inflation during the Napoleonic Wars and the higher US than UK inflation during the Civil War. Over most of the rest of the period, dollar-sterling appears as stable

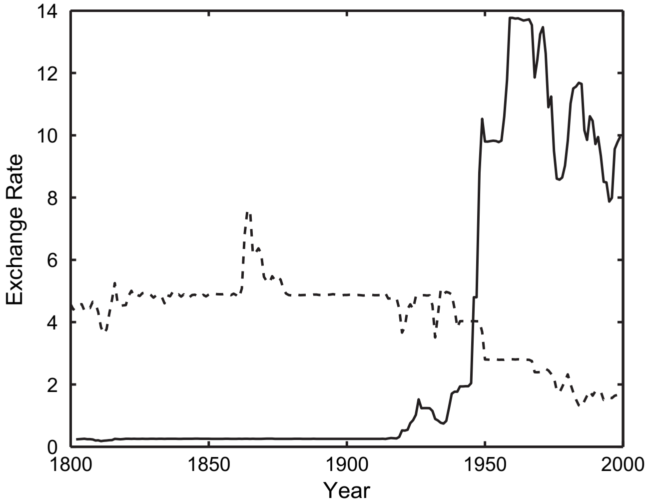


**Fig. 1.** Exchange rates over the past two centuries. The solid line denotes the franc-sterling exchange rate, and the dashed line denotes the dollar-sterling rate, from 1800 to 1999. Refer to Appendix A for the sources of the exchange rates.

as franc-sterling. As the twentieth century wears on, the picture changes dramatically. During the two World Wars and their aftermaths, the inter-war years, and under the post-Bretton-Wood's float, exchange-rate variability is markedly greater.

Fig. 2 plots the time series for both the short-term (solid lines) and the long-term (dashed lines) interest rates. Interest rates for all countries were abnormally high during late 1970s and early 1980s, but came down in the 1990s. Overall, the interest rates, especially the long-term rates, follow similar patterns of movement among the three countries. The short-term rates are more volatile than the long-term rates, but they follow each other closely. The one major exception is the US short-term rate during the nineteenth century. Between 1831 and 1873, US short-term interest rates are much higher and much more volatile than the long-term counterparts. These rates also deviate from interest rates in the other two countries. This is a period when a number of severe banking panics took place in the US. The anomalous behavior of the US short-term rates is probably a result of this fact. The US short-term rates are for commercial paper and hence may include a portion of credit premium, which can become significant during crisis-laden periods.

Table 1 reports the summary statistics of the exchange rates and the interest rates. Since we have missing data for the French short-term interest rates during the First World War and the years immediately thereafter, we compute the summary statistics excluding those missing data points. Due to the telescoping property of the log exchange rates, we measure the mean depreciation rates through a simple regression of log exchanges rates over time. The standard deviation measures the standard error of this regression slope estimate. Over the past two centuries, sterling appreciated about 2.46 percent per year against the French franc, and depreciated about 0.45 percent per year against the US dollar, but as is obvious from Fig. 1, this trend is mainly due to the realignments of both the sterling and the franc relative to the dollar in 1949. While the magnitudes do not exactly match, the differentials in long-term rates are in line with this trend. For example, on average, the French long-term rate is 0.38 percentage point higher than the UK long rate, partially compensating for the currency depreciation of the franc. Correspondingly, the US dollar long-term rate is slightly lower than the UK rate, in line with the slight appreciation of the dollar. The direction of the short-term interest-rate differentials, however, is counter to intuition. On average, the appreciating currency also has a higher short-term interest rates. There are at least two potential reasons for this. First, the short-term rate does not forecast currency movements as well as the long-term interest rate. Second, in the US case, due to the data that are used, the short-term interest-rate series may also contain a significant portion of credit premium that contaminates any relation between the interest rates and currency movements.

Both currency depreciation rates exhibit moderate mean reversion. The annual first-order autocorrelation is 0.21 for the franc-sterling depreciation rate and 0.22 for the dollar-sterling depreciation rate. Assuming a first-order autoregressive process, we also compute the half life, the length of time by which the autocorrelation declines by half of its first-order autocorrelation value. The half lives for both depreciation rates are less than six months. In contrast, the interest rates are much more persistent, with annual autocorrelations between 0.78 and 0.91, and half lives ranging from three to thirty years. The most persistent series is the UK long-term interest rate, the least is the US short-term rate. Overall, the long-term rates are more persistent than the short-term rates and the interest-rate differentials are less persistent than the interest rates themselves.

## 3. Uncovered interest-rate parity regressions

### *3.1. Forward-premium regression over the past two centuries*

Based on the martingale hypothesis, the forward exchange rate should be an unbiased forecast of the future exchange rate. A popular forecasting relation is formulated as:

$$s_{t+1} - s_t \;=\; \alpha + \beta(f_t - s_t) + e_{t+1}, \tag{1}$$

where $s_t$ denotes the logarithm of the exchange rate and $f_t$ denotes the one-period-ahead forward exchange rate. Under the martingale hypothesis that the forward risk premium is zero, $\alpha = 0$ and $\beta = 1$. Furthermore, by covered interest-rate parity,

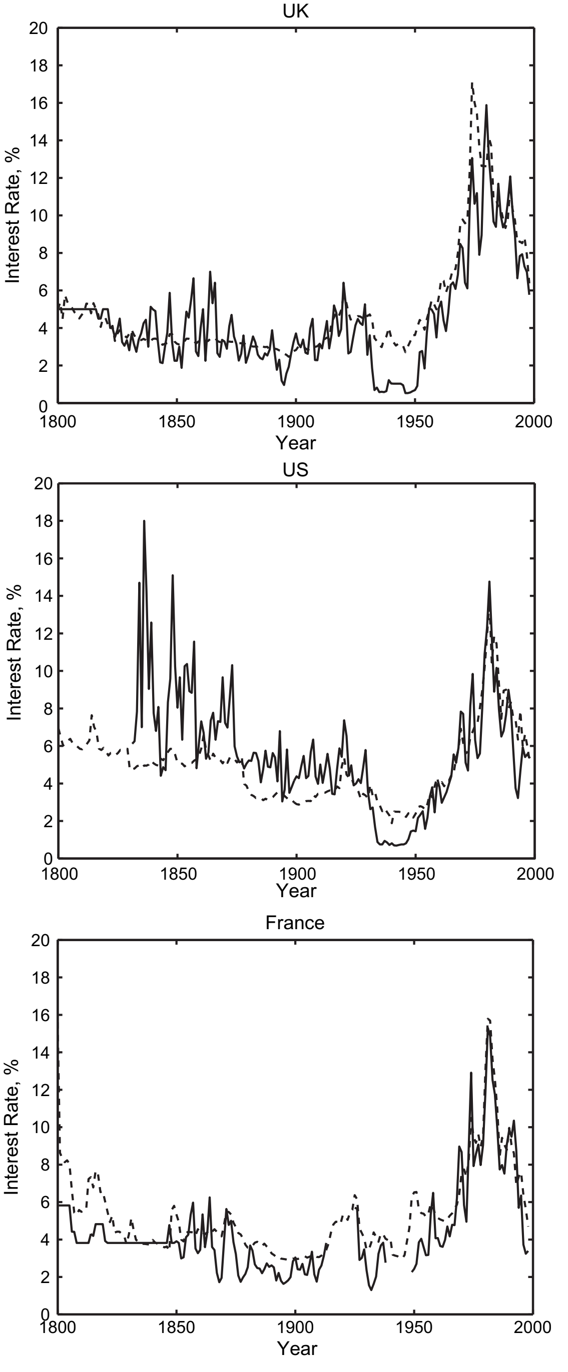


**Fig. 2.** Short-term and long-term interest rates over the past two centuries. The solid lines denote the short-term interest rates. The dashed lines denote the long-term interest rates. Refer to Appendix A for the sources and definitions of these interest rates.

**Table 1**
Summary statistics of exchange rates and interest rates.

| | *ds* | Long-term rates | | | Short-term rates | | |
|---|---|---|---|---|---|---|---|
| | | *r* | *r** | *dr* | *r* | *r** | *dr* |
| Home = France; Foreign = UK | | | | | | | |
| Mean | 2.46 | 5.33 | 4.94 | 0.38 | 4.45 | 4.40 | −0.18 |
| Std Dev | 0.10 | 2.40 | 2.90 | 1.71 | 2.52 | 2.78 | 1.34 |
| Auto | 0.21 | 0.89 | 0.98 | 0.79 | 0.90 | 0.91 | 0.60 |
| Half life | 0.44 | 6.20 | 31.67 | 2.92 | 6.65 | 6.95 | 1.34 |
| Skewness | 3.84 | 2.03 | 2.05 | −0.16 | 1.90 | 1.40 | −0.08 |
| Kurtosis | 22.65 | 5.15 | 3.86 | 9.40 | 4.17 | 2.38 | 0.70 |
| Home = US; Foreign = UK | | | | | | | |
| Mean | −0.45 | 4.94 | 4.94 | −0.01 | 5.66 | 4.40 | 1.28 |
| Std Dev | 0.03 | 1.96 | 2.90 | 2.04 | 3.12 | 2.78 | 3.01 |
| Auto | 0.22 | 0.95 | 0.98 | 0.93 | 0.78 | 0.91 | 0.79 |
| Half life | 0.46 | 14.92 | 31.67 | 9.01 | 2.77 | 6.95 | 2.88 |
| Skewness | −0.44 | 1.34 | 2.05 | −1.68 | 0.95 | 1.40 | 0.90 |
| Kurtosis | 5.34 | 3.09 | 3.86 | 5.16 | 1.73 | 2.38 | 2.53 |

Note: Entries report the summary statistics of exchange-rate depreciation rates, interest rates, and interest-rate differentials. We use *ds* to denote the exchange rate annual depreciation rates in percentages and $r$ and $r*$ to denote domestic (France or US) and foreign (UK) interest rates, also in annualized percentages. The column under *dr* denotes the interest-rate differential $dr = r - r*$. Data are annual, starting in 1800 for US dollar and UK sterling and 1803 for French franc.

$$f_t - s_t = r_t - r_t^*, \tag{2}$$

where $r_t$ and $r_t^*$ denote the domestic and foreign interest rate on a one-period zero coupon bond, respectively. Replacing the forward premium in equation (1) with the interest-rate differential in equation (2), we have

$$s_{t+1} - s_t = \alpha + \beta(r_t - r_t^*) + e_{t+1}. \tag{3}$$

A test of the hypothesis $\alpha = 0$, and $\beta = 1$ is a test of both the absence of a forward risk premium and uncovered interest-rate parity. Under the *UIP* hypothesis, if the return on a domestic $n$-period zero coupon bond is one percentage point per annum higher than that on a foreign bond, one would expect, on average, the foreign currency to appreciate by one percent over the next $n$ periods. In practice, one often finds the hypothesis grossly violated. Most puzzling of all, however, is that the estimate for $\beta$ is often negative.

We run the above forward-premium regression on both pairs of currencies. Since neither the short-term interest rates nor the long-term interest rates match the required maturity of one year, we perform alternate sets of regressions using first short rates and then long rates as the right-hand-side variable. Table 2 reports the results of these regressions. In contrast to most findings in the literature, the regression slope estimates for $\beta$ over the past two centuries are positive for all four regressions. The slope estimates of $\beta = 0.38$ (long rate) and 0.14 (short rate) for dollar-sterling are significantly different from the null hypothesis of $\beta = 1$, but the estimates for the franc-sterling regression $\beta = 0.73$ (long rates) and 0.97 (short rates) are not statistically different from the null hypothesis of one. Furthermore, none of the intercept estimates for $\alpha$ are significantly different from zero. Therefore, we conclude that the *UIP*, or the forward-premium puzzle disappears over our ultra-long sample period. *UIP* may be violated during a particular short period, but it holds much better over the long haul.

Comparing the regression results based on long-term interest rates and short-term interest rates, we find that the regressions with the long-term rates generate results closer to the expectations hypothesis. In the case of the France–UK pair, the slope estimate with the short-term rates is closer to one in magnitude, but the standard error of the estimate is also significantly larger. As a result, the estimate of 0.97 is neither significantly different from zero nor from one. In contrast, the regression slope of 0.73 based on the long-term interest rates is significantly different from zero, but not significantly different from the null value of one. In the case of the US–UK pair, the slope estimate becomes closer to zero when using the short-term rates instead of the long-term rates. Therefore, the long-term rates predict the currency movement better than the short-term rates do. Both the instruments used and the term

**Table 2**
Forward-premium regressions.

| | France/UK | | | | US/UK | | | |
|---|---|---|---|---|---|---|---|---|
| | Long-term rates | | Short-term rates | | Long-term rates | | Short-term rates | |
| | $\alpha$ | $\beta$ | $\alpha$ | $\beta$ | $\alpha$ | $\beta$ | $\alpha$ | $\beta$ |
| Estimates | 0.02 | 0.73 | 0.00 | 0.97 | −0.01 | 0.39 | −0.01 | 0.14 |
| Std Error | 0.01 | 0.43 | 0.01 | 0.86 | 0.00 | 0.28 | 0.01 | 0.16 |
| *t*-statistics | 1.65 | −0.63 | 0.73 | −0.03 | −1.08 | −2.21 | −1.27 | −5.27 |
| *p*-value | 0.10 | 0.53 | 0.46 | 0.97 | 0.28 | 0.03 | 0.20 | 0.00 |
| $R^2$, $N$ | 0.01 | 197 | 0.03 | 176 | 0.01 | 199 | 0.00 | 168 |

Note: Entries report summary properties of the forward-premium regression: $s_{t+1} - s_t = \alpha + \beta(r_t - r_t^*) + e_{t+1}$. Standard errors are constructed based on Newey and West (1987) with three lags. The number of lags is optimally chosen according to Andrews (1991) with a VAR(1) specification on the residuals. The *t*-statistics and *p*-value are constructed based on the hypothesis: $\alpha = 0$, $\beta = 1$. Data are annual from 1800 to 1999. In the last row, we report the *R*-squares on the left cell and the number of actual observations on the right cell, for each regression.

mismatch are likely to play a role in this result, particularly for the US given our use of a commercial paper rate. The commercial paper rate doubtlessly contains some credit risk component, which can be both sizable and highly volatile during years of financial crisis. Furthermore, the slope and curvature of a yield curve is mainly generated with rates within two years of maturity Backus et al. (1998). Hence, on average, the difference between the one-year rate and three-month rate can be larger than that between one year and ten-year rates. Both data issues generate measurement errors, which can bias the regression slope toward zero Bekaert and Hodrick (1993). After correction for these biases, the regression slope estimates should be even closer to one, the null value.

### 3.2. *Subperiod analysis of forward-premium regressions*

To analyze the robustness of the results, we divide the sample into three subperiods and perform an analysis within each subperiod. The three subperiods roughly correspond to three different broad regimes: (1) 1800–1913, the varied regimes of the early nineteenth century, the US Civil War period, and the heyday of the classical gold standard; (2) 1914–1949, the War and inter-war years, periods of substantial inflation punctuated by the economic dislocations of the 1920s and 1930s; and (3) 1950–1999, the post-WWII period characterized by the quasi-fixed exchange rates of Bretton Woods and the managed float and, in the case of France, the recent move to the euro.

Table 3 reports the sample mean estimates (and standard errors for the estimates in parentheses) for the depreciation rates and interest-rate differentials, as well as the estimates of the forward-premium regression for the three subperiods. For comparison, we also report the corresponding statistics for the full sample in the last two rows of each panel. The results based on the long-term interest rates again conform to the hypothesis better than those based on the short-term series. The regression slope estimates are positive in all subperiods when using the long-term rates, but become negative in some cases when the short-term interest rates are used.

The slope estimates for the subperiod regressions exhibit large fluctuations from period to period. Thus, although the expectations hypothesis holds approximately over the past two centuries, the regression slope estimates can vary dramatically from the null value in any given subperiod, illustrating the potential danger of drawing conclusions based on a short sample.

Unconditionally, the mean values of the deprecation rates and the interest-rate differentials move mostly in the same direction. For example, during the nineteenth century, sterling appreciates against both the dollar and the franc, but the UK interest rates are on average lower than those of the other two countries. We also find that the mean interest-rate differentials are much larger than the exchange-rate changes in this period. This, however, is expected, given that the exchange rates are mostly fixed during this period, with only occasional realignments.

During the World Wars and inter-war period, sterling on average depreciated against the dollar, but appreciated (dramatically) against the franc. Correspondingly, the mean interest-rate differential is negative between the US and UK, but positive between France and the UK, again consistent with

**Table 3**
Subperiod analysis on the forward-premium regression.

| Periods | $ds$ | Long-term rates | | | | Short-term rates | | | |
|---|---|---|---|---|---|---|---|---|---|
| | | $dr$ | $\alpha$ | $\beta$ | $(N, R^2)$ | $dr$ | $\alpha$ | $\beta$ | $(N, R^2)$ |
| A. Home = France; Foreign = UK | | | | | | | | | |
| 1800–1913 | 0.08 (0.04) | 0.93 (0.13) | −0.00 (0.00) | 0.35 (0.26) | 111 0.01 | −0.13 (0.12) | 0.00 (0.00) | −0.05 (0.35) | 111, 0.00 |
| 1914–1949 | 7.95 (1.02) | 0.30 (0.18) | 0.08 (0.04) | 10.05 (3.63) | 35 0.10 | 0.63 (0.78) | −0.01 (0.03) | 7.21 (1.77) | 14, 0.42 |
| 1950–1999 | −0.50 (0.27) | −1.09 (0.56) | 0.01 (0.01) | 0.70 (0.49) | 49 0.05 | −0.59 (0.35) | −0.00 (0.01) | −0.18 (0.56) | 49, 0.00 |
| Whole | 2.46 (0.18) | 0.32 (0.20) | 0.02 (0.01) | 0.73 (0.43) | 197 0.01 | −0.19 (0.15) | 0.00 (0.01) | 0.97 (0.86) | 176, 0.03 |
| B. Home = US; Foreign = UK | | | | | | | | | |
| 1800–1913 | 0.10 (0.04) | 1.32 (0.14) | −0.00 (0.01) | 0.42 (0.73) | 113 0.00 | 3.44 (0.44) | 0.00 (0.01) | −0.02 (0.15) | 82, 0.00 |
| 1914–1949 | −0.44 (0.17) | −0.75 (0.10) | 0.00 (0.02) | 1.36 (2.67) | 35 0.01 | 0.38 (0.19) | −0.01 (0.01) | 1.18 (2.20) | 35, 0.01 |
| 1950–1999 | −1.58 (0.15) | −2.50 (0.53) | −0.00 (0.02) | 0.32 (0.64) | 49 0.01 | −1.68 (0.34) | −0.01 (0.02) | −0.08 (0.94) | 49, 0.00 |
| Whole | −0.45 (0.07) | −0.01 (0.28) | −0.01 (0.00) | 0.39 (0.28) | 199 0.01 | 1.28 (0.42) | −0.01 (0.01) | 0.14 (0.16) | 168, 0.00 |

Note: Entries report the estimates and standard errors (in parentheses) of the mean depreciation rates ($ds = s_{t+1} - s_t$, in percentage per annum), mean interest-rate differentials ($dr = r - r^*$, in annual percentage), and the estimates of the following uncovered interest-rate parity regression: $s_{t+1} - s_t = \alpha + \beta(r_t - r_t^*) + e_{t+1}$, under each subperiod. Under the column labeled "($N$, $R^2$)", we report the number of available observations for the regression under each subperiod under the first row and the R-square of the regression under the second row. Standard errors for the sample mean and regression estimates are constructed based on Newey and West (1987) with three lags.

expectations. During this period, the average depreciation rate of sterling against dollar was close to the average of the interest-rate differential between the two countries in magnitude. The franc, in contrast, experienced a dramatic depreciation against UK, much bigger than the mean interest-rate differential.

Finally, during the last period of floating exchange rates, sterling depreciated against both the dollar and the franc. At the same time, UK interest rates were higher than those of the other two countries.

In summary, the violation of the uncovered interest-rate parity is much smaller than generally portrayed in the literature, especially when we take an ultra-long perspective and use a more stable interest-rate series. Nevertheless, during any particular subperiod, the regression slope can deviate dramatically from the null value of one.

### 3.3. *Rolling forward-premium regressions*

To investigate further the influence of the sampling period on the slope estimate, we re-run the forward-premium regression with a rolling window. We fix the ending period of the regression at the last observation (1999), but move the starting period progressively forward from 1802 to 1989. Fig. 3 plots the regression slope estimates (solid lines) of the rolling regression, as well as the 95 percent confidence intervals (dashed lines), as a function of the starting period. In the case of franc-sterling (left panel), the regression slopes are not significantly different from one (the null value) until the starting period moves into the mid 1970s. Similar findings apply to the case of dollar-sterling (the right panel). The slope estimates are positive when the regression is run on the whole sample period but begin to become negative when the start of the sample period is after the early 1970s. Interestingly, the slope estimate becomes positive again when the regression focuses on the 1990s, a finding also documented in Flood and Rose (2001). Indeed, most of the traditional evidence on negative regression slopes is based on samples dominated by late 1970s and early 1980s, with the US dollar as the numeraire. Nevertheless, the negative slopes on franc-sterling during this sample period indicate that the violations of the expectation hypothesis on the forward premium during this sample period not only applies to currencies benchmarked to the US dollar, but also to other exchange rates without the involvement of dollar. Bekaert (1995) presents similar findings.

In all three countries, this was a period of substantial inflation, the culmination of an historically unprecedented three-decade period of inflation. By the end of the 1970s, inflation, which on a long-term average basis had trended up steadily since the 1950s, reached historic peacetime highs in all three countries. The first break in the process came in 1979 in the UK following the Thatcher election and the ensuing move to much tighter monetary policies. A similar shift began in the US shortly after Paul Volcker became Federal Reserve Chairman and gained momentum following the 1980 Reagan election. In France, the shift came several years later.

In each instance, a series of announcements accompanied these moves to less inflationary policy, but these announcements did not do much initially to alter market expectations on the inflation rates. In the UK, many otherwise perspicacious observers even several years after the fact argued that no policy change had occurred and that inflation could be expected to increase rather than decrease. A major cause of this skepticism was the behavior of the broad monetary aggregate that the Bank of England had chosen as its target. This aggregate had accelerated for reasons that had nothing to do with policy and had no implications for price and spending behavior. Nevertheless, this acceleration created a false perception among the public that the expected inflation rate remained high for several years after the actual inflation came down, thus creating an extended period of missed targets between the forecasts and the realization in inflation rates. In this connection, see the discussion in Darby and Lothian (1983) and Sargent (1983).

A similarly slow adjustment of expectations also took place in the US. An interesting bit of evidence in this regard is provided by the ten-year inflation forecasts collected by the Federal Reserve Bank of Philadelphia.[3] Fig. 4 plots the inflation forecasts (dashed line) and the realized inflation rates (solid

[3] Federal Reserve Bank of Philadelphia, 2003, "Long-term inflation forecasts: Expected inflation over the next 10 years" http://www.phil.frb.org/files/spf/cpie10.txt.

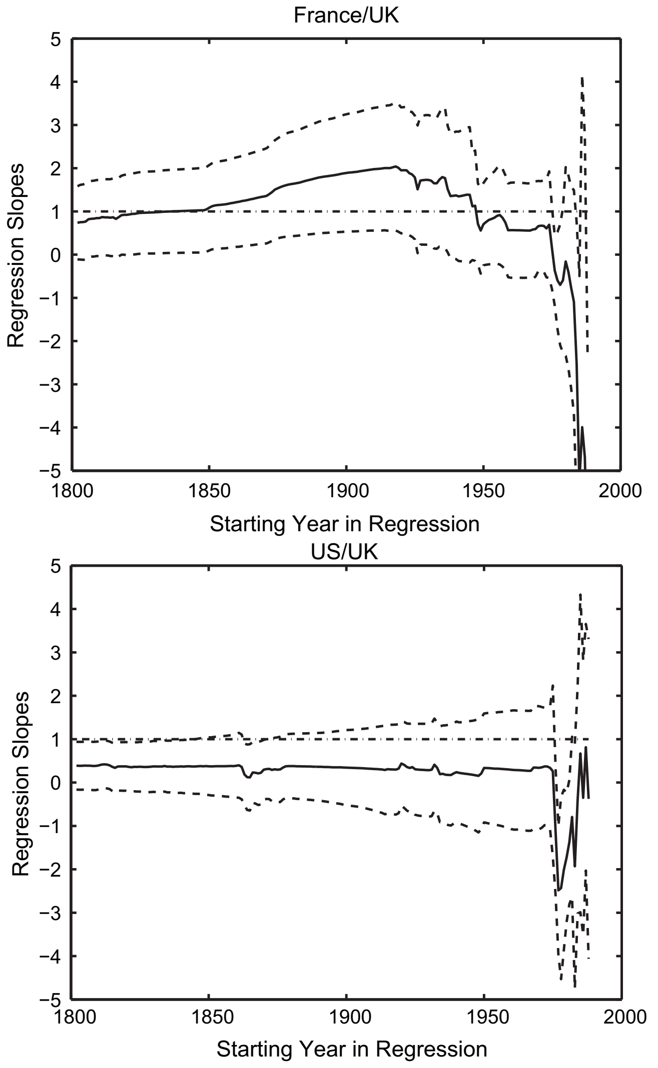


**Fig. 3.** Rolling forward-premium regression slope estimates over the past two centuries. The solid lines are the slope estimates for the forward-premium regression on franc-sterling on the left panel and dollar-sterling on the right panel. The dashed lines are the 95 percent confidence intervals, constructed according to Newey and West (1987) with three lags. The dash-dotted lines represent the null value of one. Data are annual. The regressions are rolling forward from 1802 to 1989. The *x*-axis shows the starting period of each regression, with the ending period fixed at 1999.

line), with the difference given by the dash-dotted line. Inflation forecasts made throughout the 1980s are systematically higher than the rates actually realized. Such evidence points to failing expectations with respect to the regime switch, rather than time-varying risk premia as suggested by Fama (1984), as the key reason for the negative slope estimates of the *UIP* regression during this sample period.

### *3.4. Analyzing uncovered interest-rate parity via holding-period returns*

The forward-premium regression is formulated based on the martingale hypothesis on the forward exchange rate. Combining this hypothesis with covered interest-rate parity, we can also interpret the

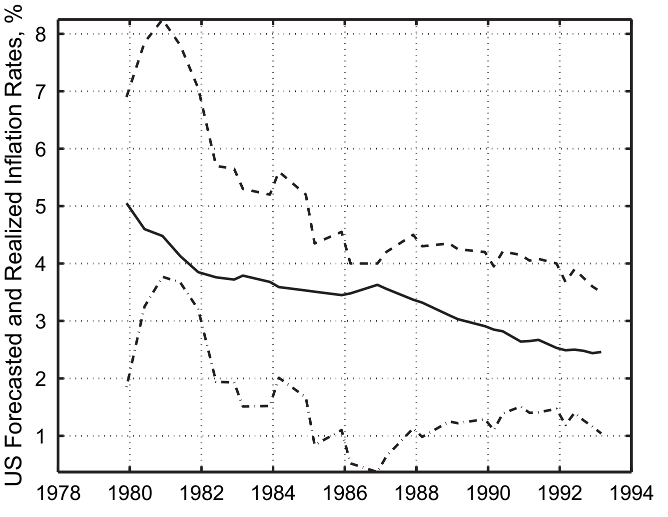


**Fig. 4.** Forecasted and realized inflation in the US. The dashed line denotes the ten-year inflation forecasts, the solid line denotes the realized inflation rate, and the dash-dotted line denotes the forecasting error, the difference between forecasted and realized inflation rates.

regression results as a test of uncovered interest-rate parity. In this section, we propose an alternative regression test of uncovered interest-rate parity based on the holding-period returns on domestic and foreign bonds. Under the hypothesis of uncovered interest-rate parity, investing in domestic and foreign bonds should on average generate the same return when denominated in the same currency. Investing in a domestic bond and holding it to maturity generates a certain return of $r_t$. Investing in a foreign bond, holding it to maturity, and converting it back to the domestic currency, generates a return of $s_{t+1} - s_t + r_t^*$, which includes both the foreign interest rate, which is certain at the time of investment $t$, and the return due to currency appreciation or depreciation, which is uncertain at time $t$ and becomes known at time $t+1$, the maturity date of the bond.

We can investigate the unconditional validity of uncovered interest-rate parity by comparing the average returns from these two types of investments. We can also investigate the conditional validity of uncovered interest-rate parity by regressing the return on the foreign bond investment against the return on the domestic bond investment,

$$s_{t+1} - s_t + r_t^* = \alpha + \beta r_t + e_{t+1}. \tag{4}$$

The null hypothesis of uncovered interest-rate parity implies $\alpha = 0$ and $\beta = 1$ for this alternative regression.

Table 4 reports the sample estimates (and the standard errors of the estimates in parentheses) of the mean return on holding foreign bonds, the mean return on holding domestic bonds, their mean difference, and the estimates of the alternative *UIP* regression in (4), both for different subperiods and the whole sample period. This holding-period-return regression generates results that are even closer to the prediction of uncovered interest-rate parity, especially for the whole sample. Unconditionally over the whole sample, the mean returns on investing in foreign and domestic bonds are not statistically different from each other for the France–UK pair, regardless of which interest-rate series we use. For the US–UK pair, the mean return difference is also insignificant when using the long-term interest rates. When using the short-term interest rates, we find that investing in the US bond generates significantly higher returns on average than investing in the UK bond. However, as we discussed above, this mean difference is more likely due to the credit premium on the commercial paper used for the US short-term rate than due to violation of uncovered interest-rate parity.

Conditionally for the whole sample period, the regression slope estimates are not statistically different from their null values of one in three of the four cases tested, with the only exception being

**Table 4**
Analyzing uncovered interest-rate parity via holding-period returns.

| Periods | Long-term rates | | | | | | Short-term rates | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|---|
| | $ds + r*$ | $r$ | $dUIP$ | $\alpha$ | $\beta$ | $(N, R^2)$ | $ds + r*$ | $r$ | $dUIP$ | $\alpha$ | $\beta$ | $(N, R^2)$ |
| A. Home = France; Foreign = UK | | | | | | | | | | | | |
| 1800–1913 | 3.58 (0.33) | 4.44 (0.23) | −0.86 (0.29) | 0.00 (0.01) | 0.71 (0.29) | 111, 0.09 | 3.58 (0.33) | 3.51 (0.18) | 0.07 (0.33) | 0.02 (0.01) | 0.40 (0.22) | 111, 0.02 |
| 1914–1949 | 14.68 (4.30) | 4.31 (0.26) | 10.37 (4.28) | 0.08 (0.22) | 1.49 (4.46) | 35, 0.00 | 7.93 (6.20) | 3.08 (0.51) | 4.85 (5.91) | −0.14 (0.09) | 7.13 (3.65) | 14, 0.23 |
| 1950–1999 | 8.93 (1.28) | 7.81 (0.74) | 1.13 (1.15) | 0.03 (0.03) | 0.75 (0.36) | 49, 0.09 | 8.93 (1.28) | 6.96 (0.84) | 1.97 (1.22) | 0.05 (0.02) | 0.63 (0.33) | 49, 0.09 |
| Whole | 6.84 (1.03) | 5.26 (0.31) | 1.58 (1.00) | 0.03 (0.02) | 0.73 (0.35) | 197, 0.02 | 5.37 (0.73) | 4.43 (0.36) | 0.93 (0.63) | 0.01 (0.01) | 1.03 (0.24) | 176, 0.12 |
| B. Home = US; Foreign = UK | | | | | | | | | | | | |
| 1800–1913 | 3.60 (0.53) | 4.86 (0.21) | −1.26 (0.50) | −0.00 (0.02) | 0.79 (0.45) | 113, 0.03 | 3.33 (0.60) | 6.77 (0.51) | −3.44 (0.73) | 0.03 (0.01) | 0.11 (0.14) | 82, 0.00 |
| 1914–1949 | 3.18 (1.23) | 3.26 (0.26) | −0.08 (1.15) | −0.01 (0.04) | 1.26 (1.46) | 35, 0.01 | 1.67 (1.31) | 2.88 (0.65) | −1.21 (1.19) | −0.01 (0.02) | 1.07 (0.55) | 35, 0.06 |
| 1950–1999 | 7.77 (1.26) | 6.40 (0.74) | 1.38 (1.21) | 0.05 (0.03) | 0.42 (0.57) | 49, 0.03 | 6.42 (1.28) | 5.87 (0.76) | 0.55 (1.29) | 0.04 (0.02) | 0.46 (0.52) | 49, 0.03 |
| Whole | 4.42 (0.58) | 4.94 (0.27) | −0.52 (0.51) | −0.00 (0.02) | 0.93 (0.43) | 199, 0.07 | 3.72 (0.67) | 5.66 (0.43) | −1.94 (0.66) | 0.01 (0.01) | 0.46 (0.20) | 168, 0.04 |

Note: Entries report the estimates and standard errors (in parentheses) of the mean return on investing in foreign bond ($ds + r^* = s_{t+1} - s_t + r_t^*$, in percentage per annum), the mean return on investing in domestic bond ($r$, in percentage per annum), and their mean difference ($dUIP = ds + r* - r$, in percentage per annum), as well as the intercept and slope estimates of the following holding-period-return regression: $s_{t+1} - s_t + r_t^* = \alpha + \beta r_t + e_{t+1}$, under each subperiod. Under the column labeled "$(N, R^2)$", we report the number of available observations for the regression under each subperiod under the first row and the R-square of the regression under the second row. Standard errors for the sample mean and regression estimates are constructed based on Newey and West (1987) with three lags.

the US–UK pair with the short-term interest rates derived from commercial paper. Even for the subperiod regressions, the slope estimates are not significantly different from one for both currency pairs over each of the three subperiods when we use the long-term interest rates. Only when we use short-term interest rates can we reject the null hypothesis of $\beta = 1$ in three instances. For the France–UK pair, the slope estimate is significantly lower than one (0.40 with a standard error of 0.22) during the subperiod 1800–1913. For the US–UK pair, the slope estimate is also significantly lower than one during the same subperiod (0.11 with a standard error of 0.14). The slope estimate is significantly lower than one for the whole sample at 0.46, with a standard error of 0.20.

Fig. 5 plots the time series of holding-period returns for investments in the foreign (dashed lines) and domestic (solid lines) bonds. In all cases, the dashed lines vary much more than the solid lines, showing the additional exchange-rate risk in investing in the foreign bonds. Nevertheless, the two lines show a substantial degree of co-movements over time, evidence in support of uncovered interest-rate parity. One obvious deviation is seen in the panel with US–UK short-term interest rates. During the early 1800s, the solid line, which is the short-term US dollar commercial paper rate, is visibly higher than the dashed line, which is the return on the UK bond. Again, we attribute this deviation to the credit component of the commercial paper rate.

Fig. 6 plots the time series of return differences between the two types of investments. Despite the large and sometimes persistent deviations, all four lines hover around zero, showing the overall validity of *UIP*.

These time series plots suggest that although *UIP* holds in the long run, random shocks to the exchange rates can generate large short-term deviations from *UIP*. The strong co-movements between

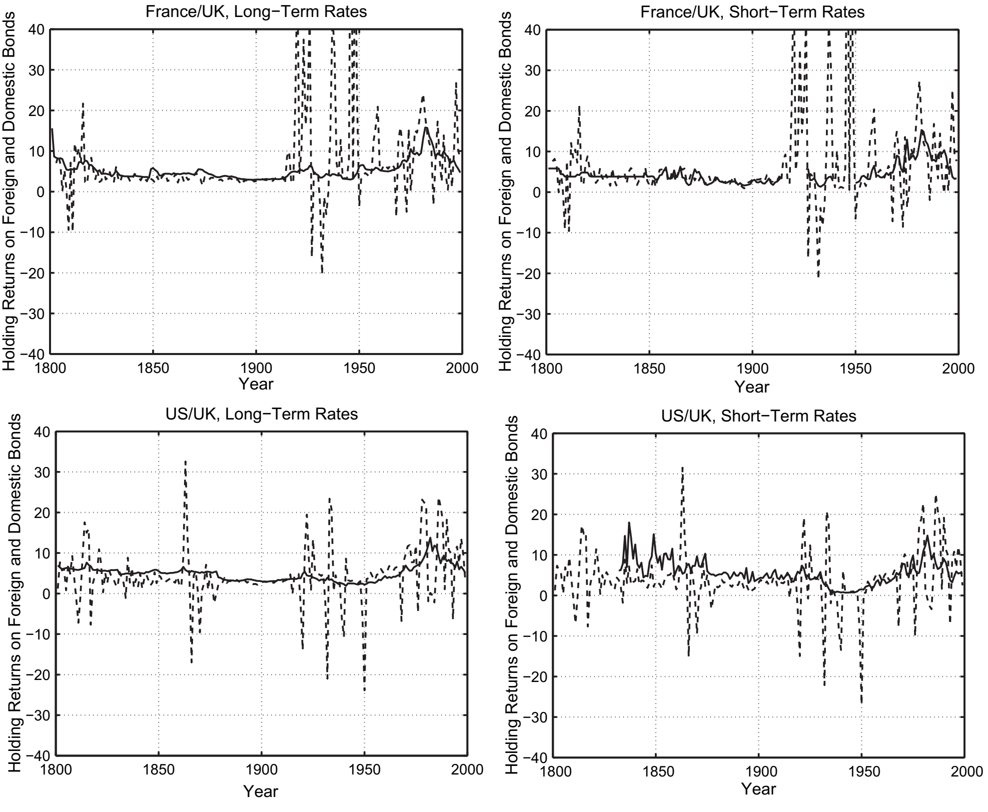


**Fig. 5.** Returns on investing in foreign and domestic bonds over two centuries. The solid lines are the holding-to-maturity returns on domestic bonds, and the dashed lines are the returns on investing in the corresponding foreign bonds and converting the proceeds to the domestic currency. In the four panels, the UK is the foreign country and France and the US are the domestic countries.

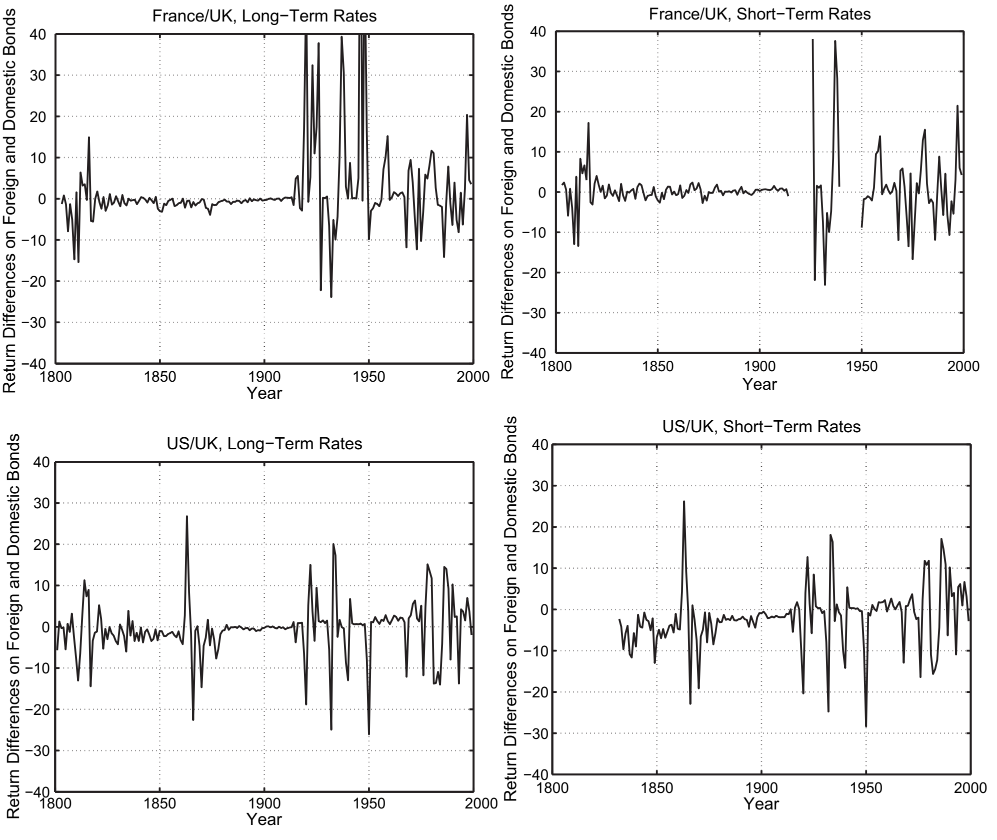


**Fig. 6.** Return differentials on investing in foreign and domestic bonds over two centuries. Solid lines are the return differences between investing in a foreign bond and a domestic bond by holding both to maturity and converting the foreign proceeds to domestic currency. In the four panels, the UK is the foreign country and France and the US are the domestic countries.

the two return series further suggest that these random shocks affect the forward-premium regressions more than they affect the holding-period-return regressions. This point can be illustrated through an extreme but simple example where the two interest rates $r$ and $r^*$ move perfectly together and with the same magnitude, then *UIP* dictates that the currency returns will vary randomly around zero. A holding-period-return regression will recover the *UIP* condition of zero intercept and unit slope, but a forward-premium regression will degenerate as the interest-rate differentials become a series of zeros.

## 4. Extreme sampling and non-linear forecasting relations

Due to inherent market frictions such as transaction costs, exchange rate adjustment may not follow the interest-rate differentials (forward premiums) instantaneously. In particular, relatively small magnitudes of interest-rate differentials may be tolerated in the market without inducing any directional movement on the exchange rate. Large interest-rate differentials, however, are less likely to persist without inducing corresponding movements in the exchange rate. A second, but not mutually exclusive, explanation for such behavior revolves around measurement error Bekaert and Hodrick (1993). In the presence of such errors, large and persistent interest-rate differentials have a much higher signal to noise ratio and hence are much more likely to contain the market's view on how the exchange rate will move in the future. As a result, uncovered interest-rate parity should hold better during periods of large interest-rate differentials.

To test this hypothesis, we propose two specifications. The first applies the principles of extreme sampling and the second relies on a smooth transition between different regimes.

### 4.1. Extreme sampling

The idea of extreme sampling is to run regressions conditional on the absolute magnitude of a signal being large. For our application, we use the absolute value of the interest-rate differential as the criterion for conditioning and run the following regression,

$$s_{+1} - s_t = \alpha + \beta^S (r_t - r_t^*) I_{t \in S} + \beta^L (r_t - r_t^*) I_{t \in L} + e_{t+1}, \tag{5}$$

where the superscripts $S$ and $L$ refer to small and large absolute realizations of the interest-rate differential, respectively. The term $I_{t \in S}$ is an indicator function that equals one if period $t$ has a small interest-rate differential and zero otherwise. $I_{t \in L}$ is analogously defined for large interest-rate differential periods. A similar extreme sampling technique has been applied in Huisman et al. (1998).

We run the regression in (5) based on different extreme sampling criteria. We define the criteria based on percentiles of the data. Specifically, we sort the absolute value of the interest-rate differential and then identify the cut-off value for different percentiles from 90 percentile to 99 percentile. We then define interest-rate differentials with absolute magnitude higher than this cut-off value as extreme periods ($L$) and those smaller as normal observations ($S$). Table 5 reports the regression estimates based on the different percentile criteria. Since we have found that the long-term interest-rate differential forecasts the currency movement better than the short-term interest-rate differential, we only report extreme sampling analysis results based on the long-term interest-rate data. For each coefficient, we report its estimate in the first column and the standard error in the second column. Under $\|dr\|$, we report the cut-off values on the absolute value of the interest-rate differential that corresponds to each percentile.

As the percentile increases and hence the criterion for large observations becomes more stringent, the estimate for $\beta^S$ declines and becomes closer to zero while the estimate for $\beta^L$ increases and becomes more positive. For the France–UK case, at a 90 percentile criterion, $\beta^S = 1.44$ is actually bigger than $\beta^L = 0.30$ and hence runs against the extreme sampling hypothesis. But once the criterion increases to 99 percentile, $\beta^S = 0.34$ becomes insignificantly different from zero while $\beta^L = 2.33$, which is even significantly greater than one. The same pattern applies to the US–UK case. As the percentile increases, $\beta^S$ declines and becomes insignificantly different from zero while $\beta^L$ increases and becomes more positive.

This phenomenon is vividly captured by the graphics in Fig. 7, in which we plot the estimates of $\beta^L$ (solid lines) and $\beta^S$ (dashed lines) as a function of the extreme sampling criteria in terms of percentiles. As the percentile increases and hence the criterion becomes more stringent for large interest-rate differentials, the slope estimate for the large realization ($\beta^L$) increases while the slope estimate for the small realization ($\beta^S$) decreases. Thus, as we have conjectured, larger interest-rate differentials have a bigger impact, or more significant forecasting power, on the currency movement.

At the extreme case of 99 percentile, we have only two sample points where the absolute magnitude of the interest-rate differentials are "large" for both pairs of currencies and these two sample points refer to the large interest-rate differentials in 1974 and 1975 for both pairs. For France–UK, the interest-rate differentials are $-0.92$ and $-1.17$ in 1974 and 1975, respectively. The next year's exchange-rate depreciation rates are respectively $-16.87$ and $-9.86$, respectively. For US–UK, the interest-rate differentials are $-2.23$ and $-2.63$ for these two years, and the following years exchange-rate depreciation rates are $-5.20$ and $-20.70$. For both pairs, the realized exchange-rate depreciation rates not only have the same sign as the interest-rate differentials, but also exhibit a larger magnitude of change than the interest-rate differentials, thus generating slope estimates of $\beta^L$ greater than one.

As the percentile increases, the $R$-squares of the regressions also improve. At 99 percentile extreme sampling, the $R$-square is 2.1 percent for France–UK and 4.0 percent for US–UK, higher than the $R$-squares from the linear regressions, about one percent for both exchange rates. Nevertheless, the overall forecasting power remains extremely small, even with the help of extreme sampling.

**Table 5**
Uncovered interest-rate parity under extreme sampling.

| Percentile | $\alpha$ | $\beta^S$ | $\beta^L$ | $\|\|dr\|\|$ | $R^2$ |
|---|---|---|---|---|---|
| A. Home = France; Foreign = UK | | | | | |
| 90 | 1.42 (1.01) | 1.44 (0.81) | 0.30 (0.57) | 2.44 | 0.016 |
| 91 | 1.42 (1.02) | 1.35 (0.77) | 0.31 (0.59) | 2.49 | 0.015 |
| 92 | 1.43 (1.03) | 1.26 (0.74) | 0.34 (0.61) | 2.55 | 0.014 |
| 93 | 1.53 (1.05) | 1.00 (0.68) | 0.51 (0.57) | 2.69 | 0.011 |
| 94 | 1.59 (1.07) | 0.86 (0.66) | 0.60 (0.58) | 2.76 | 0.010 |
| 95 | 1.68 (1.09) | 0.71 (0.65) | 0.75 (0.56) | 2.86 | 0.010 |
| 96 | 1.77 (1.09) | 0.57 (0.60) | 0.93 (0.54) | 3.03 | 0.011 |
| 97 | 1.87 (1.06) | 0.39 (0.55) | 1.24 (0.36) | 3.67 | 0.013 |
| 98 | 1.91 (1.06) | 0.37 (0.50) | 1.47 (0.30) | 4.06 | 0.015 |
| 99 | 1.92 (1.02) | 0.34 (0.40) | 2.33 (0.21) | 6.27 | 0.021 |
| B. Home = US; Foreign = UK | | | | | |
| 90 | −0.43 (0.51) | 0.23 (0.39) | 0.50 (0.33) | 2.52 | 0.021 |
| 91 | −0.43 (0.50) | 0.22 (0.37) | 0.52 (0.34) | 2.57 | 0.021 |
| 92 | −0.42 (0.50) | 0.21 (0.35) | 0.54 (0.34) | 2.63 | 0.022 |
| 93 | −0.42 (0.49) | 0.20 (0.33) | 0.56 (0.35) | 2.81 | 0.022 |
| 94 | −0.41 (0.50) | 0.19 (0.36) | 0.59 (0.34) | 2.91 | 0.023 |
| 95 | −0.43 (0.49) | 0.23 (0.36) | 0.56 (0.36) | 3.23 | 0.022 |
| 96 | −0.47 (0.51) | 0.30 (0.36) | 0.49 (0.39) | 3.62 | 0.020 |
| 97 | −0.41 (0.51) | 0.18 (0.35) | 0.69 (0.26) | 3.98 | 0.025 |
| 98 | −0.37 (0.51) | 0.07 (0.36) | 0.98 (0.14) | 5.63 | 0.037 |
| 99 | −0.42 (0.49) | 0.14 (0.30) | 1.24 (0.27) | 8.20 | 0.040 |

Note: Entries report estimates of the following extreme sampling regression: $s_{+1} - s_t = \alpha + \beta^S (r_t - r_t^*) I_{t \in S} + \beta^L (r_t - r_t^*) I_{t \in L} + e_{t+1}$, where $S$ and $L$ denote small and large realizations on the absolute value of the interest-rate differentials, respectively. The differentiation between small and large is based on the percentiles of the data, as shown in the first column of the table. The column under $||dr||$ reports the critical value of the interest-rate differential to make this differentiation for each regression. The last column reports the R-square of each regression. For each estimate ($\alpha, \beta^S, \beta^L$), the left column reports the regression estimate while the right column reports its standard error in parentheses, which is constructed based on Newey and West (1987) with three lags. The regressions are based on long-term interest rates.

### *4.2. A smooth transition non-linear regressive model*

The extreme sampling analysis shows that the relation between the rate of exchange-rate depreciation and the interest-rate differential is inherently non-linear. One particular statistical characterization of non-linear adjustment that appears to work well for exchange rates is the smooth transition non-linear regression model Granger and Terasvirta (1993). In these models, adjustment takes place in every period but the speed of adjustment varies with the magnitude of the interest-rate differential. A particularly simple formulation that is applicable to our case can be specified as follows,

$$s_{+1} - s_t = \alpha + \beta\left(r_t - r_t^* - \mu\right) + \gamma\left(1 - e^{-\lambda\left(r_t - r_t^* - \mu\right)^2}\right)\left(r_t - r_t^* - \mu\right) + e_{t+1}, \tag{6}$$

where $\mu$ denotes a long run mean of the interest-rate differential, the transition function $\phi = 1 - e^{-\lambda(r_t - r_t^* - \mu)^2}, \lambda > 0$, is between zero and one as the deviation of the interest-rate differential increases from zero to infinity. The transition function is centered around a mean interest-rate differential level $\mu$. The transition parameter $\lambda$ determines the speed of transition between the two extreme regimes, with lower values of $\lambda$ implying slower transition. The inner regime corresponds to $r_t - r_t^* = \mu$ so that $\phi = 0$ and equation (6) becomes a linear forecasting relation,

$$s_{+1} - s_t = \alpha + \beta\left(r_t - r_t^* - \mu\right) + e_{t+1}. \tag{7}$$

The outer regime corresponds to the case when $r_t - r_t^* - \mu \to \infty$ and $\phi = 1$ so that equation (6) becomes a different linear forecasting relation,

$$s_{+1} - s_t = \alpha + (\beta + \gamma)\left(r_t - r_t^* - \mu\right) + e_{t+1}, \tag{8}$$

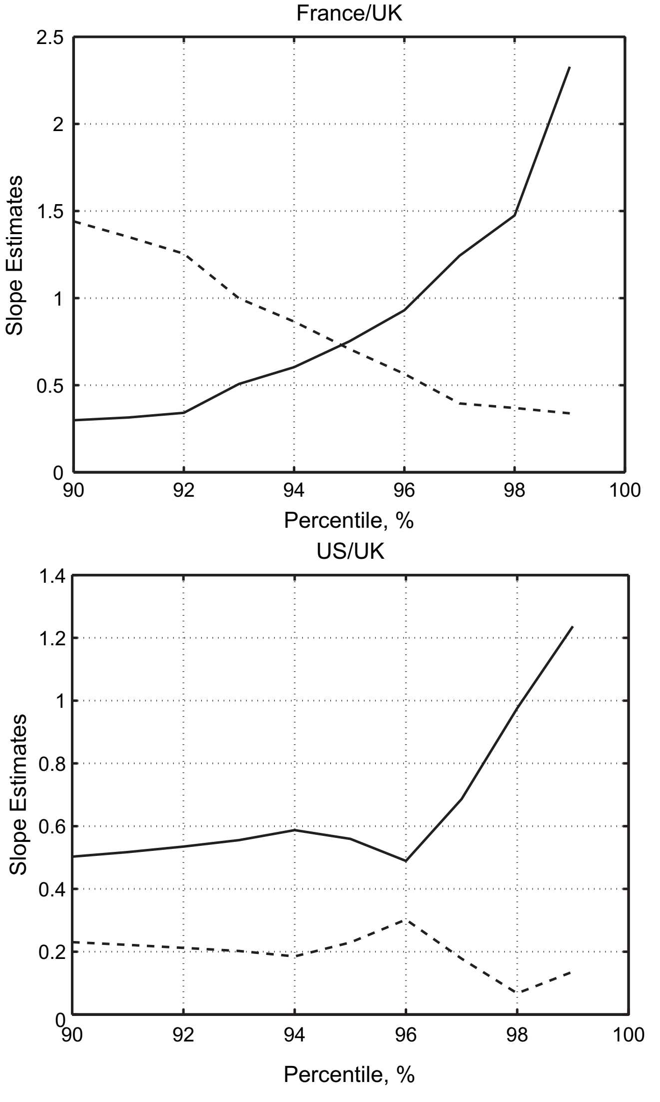


**Fig. 7.** Regression slopes under different extreme sampling criteria. Lines represent slope estimates of the following extreme sampling regression under different sampling criteria: $s_{+1} - s_t = \alpha + \beta^S(r_t - r_t^*)I_{t\in S} + \beta^L(r_t - r_t^*)I_{t\in L} + e_{t+1}$, where $S$ denotes the sample periods with the absolute value of the interest-rate differential ($\|dr\|$) is within a critical value and $L$ denote the sample periods when the interest-rate differential is outside this critical value. The critical value on the absolute interest-rate differential is based on the percentile of the data. The lines denote the slope estimates under different percentiles. In particular, the solid lines represent the estimates for $\beta^L$ and the dashed lines represent the estimates for $\beta^S$.

with a different regression slope. Thus, equation (6) provides a smooth transition between these two limiting cases (regimes) and can be regarded as a smoother version of extreme sampling. The slope coefficient $\beta$ represents the response of the exchange rate to small interest-rate differentials and the coefficient $(\beta + \gamma)$ corresponds to the response to large interest-rate differentials.

The non-linear regression in (6) is estimated by minimizing the mean squared error of the regression residuals using numerical non-linear least square packages. The estimation results are reported in Table 6. For both exchange rates, the estimates for $\beta$ are negative, but the estimates for $\gamma$ are large and positive. These results are consistent with our findings from the extreme sampling analysis. Exchange rate movements only respond to large interest-rate differentials, but not to small ones.

Compared to the linear regression, the percentage of explained variance ($R$-square) increases to 2.2 percent for France–UK and 3.9 percent for US–UK, similar to the performance of the extreme sampling regression. Thus again, although the forecasting performance can increase slightly via a non-linear, and hence more flexible specification, the overall forecasting power of interest-rate differentials for currency movement is still very low.

## 5. An historical account of *UIP* deviations

The upshot of the findings reported so far is that over the long term, the *UIP* puzzle largely disappears. In that sense, *UIP* "works." The problem, however, is that it does not work all that well as a forecasting relation. The standard deviations of the *UIP* regression residuals are large relative to the standard deviations of exchange-rate changes. In this section, we try to uncover the reasons why this is the case. To do so, we divide the sample into subperiods that correspond to historical regimes and analyze the properties of exchange rates and interest rates under each regime.

To understand the source of *UIP* deviations under different historical periods, we decompose the deviation from *UIP* into a real component and a nominal component. The real component captures the real interest-rate differential, while the nominal component captures the deviation from the purchasing power parity. To understand this decomposition, consider the following open-economy Fisher equation:

$$dr = d\rho + \overline{d\pi}, \tag{9}$$

where $d\rho$ denotes the difference between real interest rates in the two countries and $\overline{d\pi}$ denotes the difference between the relevant anticipated rates of inflation in the two countries.[4] Adding the anticipated depreciation rate $\overline{ds}$ to both sides of the Fisher equation results in,

$$\overline{ds} - dr = \overline{ds} - \overline{d\pi} - d\rho. \tag{10}$$

The left hand side is the anticipated deviation from *UIP*, which we label as $\overline{dUIP}$. The first two terms on the right hand side constitutes the anticipated deviation from *PPP*, which we denote as $\overline{dPPP}$. Thus, we can decompose the anticipated deviation from *UIP* into two components: (1) the anticipated deviation from purchasing power parity and (2) the difference between the real interest rates of the two countries,

$$\overline{dUIP} = \overline{dPPP} - d\rho. \tag{11}$$

Similar decomposition can be found in Gokey (1994) and Hollifield and Yaron (2001).

Since *PPP* is being expressed in growth rate form, the on-going debate concerning the existence and size of a permanent component in the real exchange rate is beside the point. No economist to our knowledge has argued that the real exchange rate is $I(2)$. We would expect, therefore, that the anticipated exchange-rate depreciation rates $\overline{ds}$ and the anticipated inflation rates $\overline{d\pi}$ converge even if the levels of nominal exchange rates and relative prices did not. Deviations from growth rate *PPP* would require the existence of recurrent real shocks, or factors producing recurrent shifts in exchange-rate expectations. An example of the former is continually faster productivity growth in one country over the other. An example of the latter is continually increasing or decreasing fears about future inflation in one of the two countries. Evidence presented in Lothian and Simaan (1998) suggests that such factors have little empirical relevance over the longer term.

The differential between the real interest rates in the two countries can be further decomposed into two components: (1) the differential between the real returns on real assets internationally and (2) the differential between the differentials in the real returns on real assets and on nominal assets (bonds) in

[4] We use the overline to denote anticipated rates, in contrast to the ex post-realized rates.

**Table 6**
Smooth transition non-linear uncovered interest parity regression model.

| Parameters | France/UK | | US/UK | |
|---|---|---|---|---|
| | Estimates | Std err | Estimates | Std err |
| $\alpha$ | 0.014 | (0.030) | −0.003 | (0.008) |
| $\beta$ | −2.544 | (2.775) | −0.715 | (0.697) |
| $\mu$ | 0.022 | (0.016) | −0.020 | (0.009) |
| $\gamma$ | 3.183 | (2.555) | 2.355 | (1.924) |
| ln($\lambda$) | 7.828 | (2.136) | 5.731 | (1.814) |
| *R*-square | 0.022 | | 0.039 | |

Note: Entries report the estimates of the following non-linear forecasting relation: $s_{+1} - s_t = \alpha + \beta(r_t - r_t^* - \mu) + \gamma(1 - e^{-\lambda(r_t - r_t^* - \mu)^2})(r_t - r_t^* - \mu) + e_{t+1}$. The estimates are based on long-term interest rates.

the two countries domestically (Friedman and Schwartz, 1982, page 513 forward). The first component reflects factors affecting the degree of arbitrage among countries such as capital controls. The second component reflects incomplete financial intermediation within countries, differences in the quality of the two financial assets (risk premia), or measurement or expectation error.

Table 7 reports subperiod averages of exchange rate depreciation rates ($ds$), interest-rate differentials ($dr$), the deviations from *UIP* ($dUIP = ds - dr$), as well as four additional variables that are related to the above decompositions: inflation differentials ($d\pi$), real interest-rate differentials ($d\rho = dr - d\pi$), deviations from *PPP* in growth rate form ($dPPP = ds - d\pi$), and real GDP grow rate differentials ($dy$). Equipped with these sample averages of *UIP* deviations and the relevant real and nominal components and variables, we proceed to analyze the fundamental source of *UIP* deviation under each sample period. We divide the past two centuries into two broad periods, with 1914 the dividing line. We show that *UIP* deviations before and after 1914 are generated from different sources.

**Table 7**
Historical decomposition of uncovered interest parity deviations.

| Subperiods | *dr* | *ds* | *dUIP* | $d\pi$ | $d\rho$ | *dPPP* | *dy* |
|---|---|---|---|---|---|---|---|
| A. Home = France; Foreign = UK | | | | | | | |
| **1803–1913** | **0.89** | **0.07** | **−0.83** | **0.09** | **0.81** | **−0.02** | **−0.25** |
| 1803–1815 | 1.71 | 0.42 | −1.29 | 0.51 | 1.19 | −0.10 | 0.63 |
| 1816–1874 | 0.82 | −0.03 | −0.85 | 0.15 | 0.67 | −0.18 | −0.48 |
| 1875–1913 | 0.40 | 0.02 | −0.39 | 0.14 | 0.26 | −0.12 | −0.38 |
| **1914–1999** | **−0.47** | **4.26** | **4.73** | **3.82** | **−4.29** | **0.44** | **0.52** |
| 1914–1925 | 0.38 | 11.58 | 11.20 | 10.11 | −9.74 | 1.46 | 1.05 |
| 1926–1939 | 0.20 | 3.89 | 3.68 | 4.07 | −3.87 | −0.19 | −0.63 |
| 1940–1949 | 0.57 | 17.85 | 17.28 | 20.49 | −19.92 | −2.64 | −1.31 |
| 1950–1999 | −1.07 | −0.11 | 0.96 | −1.10 | 0.03 | 0.99 | 1.08 |
| 1950–1973 | −0.60 | 0.14 | 0.74 | 0.88 | −1.48 | −0.74 | 2.07 |
| 1974–1999 | −1.51 | −0.35 | 1.16 | −2.93 | 1.42 | 2.58 | 0.17 |
| B. Home = US; Foreign = UK | | | | | | | |
| **1800–1913** | **1.28** | **0.14** | **−1.14** | **0.19** | **1.09** | **−0.05** | **2.19** |
| 1800–1815 | 1.29 | 0.80 | −0.49 | 1.62 | −0.32 | −0.82 | 1.65 |
| 1816–1860 | 1.79 | 0.07 | −1.72 | −0.63 | 2.42 | 0.70 | 2.21 |
| 1861–1874 | 2.12 | 0.67 | −1.45 | 1.96 | 0.16 | −1.28 | 2.99 |
| 1875–1913 | 0.39 | −0.24 | −0.63 | −0.08 | 0.47 | −0.16 | 2.10 |
| **1914–1999** | **−1.75** | **−1.28** | **0.46** | **−1.53** | **−0.21** | **0.25** | **1.13** |
| 1914–1925 | −0.54 | −0.07 | 0.47 | −0.63 | 0.09 | 0.56 | 2.61 |
| 1926–1939 | −0.91 | −0.61 | 0.30 | 0.66 | −1.58 | −1.27 | −0.66 |
| 1940–1949 | −0.84 | −1.85 | −1.01 | −1.26 | 0.42 | −0.59 | 2.87 |
| 1950–1999 | −2.45 | −1.65 | 0.80 | −2.42 | −0.03 | 0.76 | 0.92 |
| 1950–1973 | −2.25 | −1.70 | 0.55 | −1.63 | −0.62 | −0.07 | 1.10 |
| 1974–1999 | −2.63 | −1.61 | 1.02 | −3.14 | 0.51 | 1.53 | 0.76 |

Note: The symbols $dr$, $ds$, $dUIP$, $d\pi$, $d\rho$, $dPPP$ and $dy$ denote, respectively, the nominal long-term interest-rate differential, the percentage change in the nominal exchange rate, the deviation from *UIP* ($dUIP = ds - dr$), the inflation rate differential, the real interest-rate differential ($d\rho = dr - d\pi$), the deviation from *PPP* ($dPPP = ds - d\pi$), and the real GDP growth rate differential. All figures are expressed in percentage per annum terms. The UK is the numeraire in all instances.

### 5.1. The nineteenth and early twentieth centuries

In the case of France versus the UK, the behavior of the interest-rate differential (*dr*) and the currency depreciation rate (*ds*) was fairly homogeneous over the course of the century following the end of the Napoleonic Wars. The gap between *dr* and *ds* was positive but averaged only a bit over 80 basis points. This gap was traceable arithmetically to the gap between the two countries' real interest rates. Purchasing power parity for its part held almost perfectly over the period, with the average deviation merely two basis points per annum. Underlying the high French real interest rates, according to Homer and Sylla (1996), were risk premia resulting from the uncertainties surrounding the French political situation during this period, the occasional major changes in regime that took place in the first half century or so, and the subsequent tensions with Germany that resulted in the Franco-Prussian War in 1870.

Viewed in a purely technical sense, however, *UIP* was violated. A British investor and his heirs who bought French rentes at the end of the Napoleonic Wars and held them until the start of World War I would have averaged 80 basis points per year more than on an equivalent investment in British consols. He would, however, have borne the political risk alluded to above and had he held the rentes until 1920 would have seen the additional return over the previous century more than fully eroded by the depreciation of the franc relative to sterling during WWI and its immediate aftermath.

The experience of the US versus the UK over the same long period was similar on average to that of France versus the UK, but differed in the details. Again the interest-rate differential *dr* exceeded the currency depreciation rate *ds* and in this instance by more than one percentage point per annum. Again this was the arithmetic result of a difference in real interest rates between the two countries rather than a deviation from *PPP*.

As in the France–UK case, *PPP* again held almost perfectly over the period as a whole. Unlike the French case, in which political risk appeared to have been responsible for the real interest-rate gap, the higher US than UK real interest rates appear due to higher real returns to investment in real assets in the US. Real GDP growth (*dy*), our proxy for such returns, was roughly two percentage points per annum faster in the US than in the UK over the period, while for France versus the UK the real growth differential was actually reversed.

The other noticeable difference between the two cases was the much greater variability of both *dr* and *ds* and of the difference between the two, the deviation from *UIP*, in the US–UK case. This greater variability was largely the result of greater variability in two episodes – the US Civil War period and the early part of the Greenback period that followed it and the decade or so surrounding the Napoleonic Wars and the War of 1812. Interestingly, however, in both of these episodes *ds* and *dr* do track one another reasonably well in terms of broad movements, even though the magnitude of the differential between the two at times widens dramatically.

### 5.2. The twentieth century from 1914 to 1999

In 1914 the world changed. This certainly was true for monetary behavior, and not surprisingly, it carried over into both exchange rate and price level behavior. In the nineteenth century, suspension of specie payments during wartime and exchange-rate depreciation were followed by resumption and appreciation. Wartime price level increases were followed by post-war decreases. The gold standard, as Bordo and Kydland (1995) have argued, was a commitment mechanism and investors for the most part caught on. Late nineteenth century US experience was the exception that proved the rule. Fears that the US would leave gold caused the spread between US and UK interest rates to widen. When these fears proved unfounded, US interest rates fell and, as it turned out, proved too low given the somewhat higher US than UK inflation over the next decade and a half. See the discussions of this episode in Friedman and Schwartz (1963, 1982) and the subsequent econometric analysis of Hallwood et al. (2000).

Only in the UK following World War I did this nineteenth-century pattern of inflation followed by deflation continue. In 1925, the UK returned to gold at the pre-war parity as a conscious policy decision. In France in the early inter-war years, in contrast, the severe inflation of World War I actually was followed by further inflation. As a result, the price level and the nominal franc-sterling exchange rate wandered even further from their nineteenth century levels. In the US, the price level eventually did come back to earlier levels, but mostly as a result of the Great Depression.

Over the full period from 1914 to 1999, the average annual deviations from *UIP* for franc-sterling were markedly greater than the average deviation from 1803 to 1913 (4.73 versus −0.83 percentage point), but were actually quite a bit less for dollar-sterling (0.46 versus −1.14 percentage points). The franc-sterling result, as data for the subperiods indicate, was heavily influenced by behavior in the wartime and inter-war periods. In all three of these episodes, expectations of inflation clearly were out of sync with what eventually transpired. French nominal interest rates averaged higher than their UK counterparts but the differentials were small in comparison to the realized differentials in inflation. Ex post-real interest rates, therefore, were highly negative. Exchange rates did, however, change roughly in line with the inflation differentials so that deviations from *PPP* for both currencies continued to be small over the full period on average.

The disparity between differentials in nominal interest rates and movements in exchange rates in the two war periods may in part reflect the controls that governments at various times placed on interest rates. A more important source of divergences, we suspect, were problems of expectations formation. If the war-related inflations were unanticipated, or expected to have only transient effects on price levels, nominal interest rates would generally have been too low before the fact and nominal interest-rate differentials, therefore, a poor predictor of subsequent changes in exchange rates.

Similar problems of expectations formation appear to have surrounded the moves to lower inflation regimes in the early 1980s in the UK and US. As we argued in a previous section, such peso-like problems very likely account for much of the seemingly anomalous behavior of nominal interest rates and exchange rates relative to one another and to realized inflation rates in the late 1970s and early 1980s in both the US and the UK.

During this period, we see an average deviation from *UIP* for dollar-sterling considerably above its twentieth century non-war average and an average deviation from *UIP* for franc-sterling slightly greater than its non-war average. In both cases, these are accompanied by deviations from *PPP* that are large both in comparison to twentieth century experience as a whole and to the *UIP* deviations. This positive association between *UIP* and *PPP* deviations has been documented previously by Gokey (1994) and Marston (1997). A plausible explanation for the association between the two, as Lewis (1988), Marston (1997), and Peruga (1996) have argued, revolves around learning. As inflation fell and learning set in, nominal exchange rates and short-term interest rates adjusted with a lag and long-term interest rates with greater lags still. Sizable deviations from both *UIP* and *PPP* were the end result.

Our long-term evidence as well as the results of several studies utilizing cross-country data for recent decades, e.g., Flood and Taylor (1996) and Lothian and Simaan (1998), suggests that in the end such problems disappear.

## 6. Conclusion

Uncovered interest-rate parity is one of three theoretical relations that are used repeatedly in analytical work in international finance and international monetary economics. Stated in its simplest form, the conclusion to which *UIP* gives rise is that countries with high nominal interest rates relative to interest rates abroad are countries with depreciating currencies. The problem, however, is that over the past several decades we have often seen the exact opposite taking place.

In this paper, we attribute these widely documented *UIP* failures to the coincidence of two empirical artifacts: (1) the unique features of the late 1970s and the 1980s and (2) the noise surrounding small *UIP* deviations. Each in its own way gives rise to expectation problems. The first because of what appear to be imperfectly foreseen regime changes; the second because of signal-extraction problems of an errors-in-variable variety.

We control for both by constructing an ultra-long time series spanning two centuries and by running regressions conditional on large deviations from *UIP*. We find that traditional forward-premium regressions yield positive slope estimates over the whole sample period and that these estimates only become negative when the 1980s make up a major portion of the sample period. When we estimate an alternative regression based on holding-period returns on foreign versus domestic bonds, the null hypothesis of *UIP* can no longer be rejected over the whole sample period. We also find that large interest-rate differentials have stronger forecasting powers for currency movements than small interest-rate differentials. Finally, an historical account of expected and realized regime changes illustrates how the expectation hypothesis underlying *UIP* holds over the very long haul but can be deviated from for long periods of time, due either to failures of expectations to adjust quickly enough to regime and other broad-based policy changes or to anticipations over extended periods of large events that in the end never actually materialize.

These are the positive findings, the parts of the glass that are full. There is also a truly major part that is empty, which is the overall poor predictive performance of *UIP* over shorter periods and for small interest-rate differentials. If there is a *UIP* puzzle, it is not as commonly believed the anomalous negative relationship between the interest-rate differential and the rate of exchange-rate depreciation observed in the 1980s, but the fact that there is very often little relationship one way or the other.

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## Appendix A. Data sources

The data are constructed from a variety of sources, including:

1. Board of Governors of the Federal Reserve System, Federal Reserve Bulletin, various issues.
2. Michael D. Bordo, "The Bretton Woods International Monetary System: An Historical Overview," in Michael D. Bordo and Barry Eichengreen, eds., A Retrospective on the Bretton Woods System, Chicago: University of Chicago Press for the NBER, 1993 [and associated data diskettes].
3. Michael D. Bordo and Lars Jonung, The Long-Run Behaviour of the Velocity of Circulation, The International Evidence, New York: Cambridge University Press, 1987 [and associated data diskettes].
4. Milton Friedman and Anna J. Schwartz, Monetary Trends in the United States and the United Kingdom, Chicago: University of Chicago Press for the NBER, 1982.
5. Sydney Homer, A History of Interest Rates, 2nd ed. New Brunswick, NJ: Rutgers University Press, 1977.
6. International Monetary Fund, International Financial Statistics (IFS), various issues, and companion CD ROM.

*United Kingdom*

*Short-term interest-rate*

1831–1844, Overend–Guerney average annual rate for first-class 3-month bills from Mitchell (1988), table entitled "Financial Institutions 15. The Market Rate of Discount – 1824–1980."
1845–1869, average annual rate for 3-month bank bills from the same source.
1870–1986, average annual rate for 3-month bank bills from Bordo and Jonung (1987) and Bordo (1993).
1987–1999, average annual money market rate from International Financial Statistics.

*Long-term bond yields*

1791–1869, average annual yields on three percent consols for from Mitchell, table entitled "Financial Institutions 13. Yield on Consols."
1870–1975, average annual yields on three percent consols from Bordo and Jonung (1987).
1976–1994, average annual government bond yield from IFS.

*United States*

*Short-term interest rates*

1831–1899, average annual commercial paper rate from Homer (1977), Table 44.
1900–1975, average annual prime 60–90 day commercial paper rate from Homer (1977), Table 51.

1976–1999, average annual prime 60–90 day commercial paper rate from various issues of the Federal Reserve Bulletin.

*Long-term bonds yields*

These are US treasury bonds for the most part. They are a mixture of maturities. A rough guess as to the average is 15–20 years prior to 1977 after which it is 30 years. The data were taken from Global Financial Data.

*Exchange rate*

1791–1796, annual averages of the White exchange rate series in the form of percent deviations of sterling from parity (in dollars per pound) from Table Appendix Table 1, pp. 610–612 in Officer (1983) adjusted by the parity values in his Table 5.
1797–1820, annual averages of the White exchange rate series in the form of percent deviations of sterling from parity inclusive of U.K. paper currency depreciation from worksheets provided by Lawrence Officer, adjusted by parity values in Officer (1983), Table 5.
1821, annual average of the White series inclusive of paper currency depreciation (first quarter) and the Appendix Table 1 (remaining three quarters) adjusted by the parity values in Officer (1983), Table 5.
1822–1829, same construction as for 1791–1796.
1830–1899, annual averages of percent deviations of sterling from parity from Officer (1985), pp. 563–565, adjusted by the parity values in Officer (1983), Table 5 (which are variable until 1837 and fixed at 4.8666 thereafter) and further adjusted in the years 1837–1843, 1857 and 1862–1878 for US currency depreciation on the basis of the estimates reported in Warren and Pearson, (1935), Table 2, p. 154; 1900–1985, Friedman and Schwartz (1982), Table 4.9, pp.130–137; 1976–1990, IFS.

*France*

*Short-term interest rates*

1863–1899, average annual open-market discount rate from Homer (1977), Table 27.
1900–1913, average annual open-market discount rate from Homer (1977) Table 61.
1925–1948, average annual private discount rate from Homer, Table 61.
1949–1994, average annual money market rate from IFS.

*Long-term bond yields*

1800–1825, average annual yield on five percent French government rentes minus 67 basis points, the difference between the yield on the five percent rentes and the yield on three percent rentes in 1826), from Table 25.
1826–1899, average annual yield on three percent rentes, from Table 25;
1900–1948, average annual yield on three percent perpetual rentes, Table 60.
1949–1994, average annual government bond yield from IFS.

*Exchange rate*

1803–1940 and 1945–80, Paris franc/sterling exchange rate from British Historical Statistics (BHS), table entitled "Financial Institutions 22. Foreign Exchange Rates-1609–1980," pp. 702–703, adjusted for a break in 1931 by taking a weighted average of the 124.06 rate prevailing for the first three quarters of that year and the 94.02 rate in the last quarter.

1941–1944, derived as a cross rate using New York dollar/sterling rates and Swiss quotations of franc/dollar rates graciously provided by Phillipe Jorion.
1981 on, derived as a cross rate from yearly average dollar/sterling and franc/dollar rates from the IFS.

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