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ABSTRACT

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We review Irving Fisher's seminal work on UIP and on the closely related equation linking interest rates and inflation. Like Fisher, we find that the failures of UIP are connected to individual episodes in which errors surrounding exchange rate expectations are persistent, but eventually transitory. We find considerable commonality in deviations from UIP and PPP, suggesting that both of these deviations are driven by a common factor. Using a dynamic latent factor model, we find that deviations from UIP are almost entirely due to forecasting errors in exchange rates, a result consistent with those reported by Fisher a century ago.

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Irving Fisher, Expectational Errors, and the UIP Puzzle

Of the three major international parity relations, uncovered interest rate parity (UIP) has proven to be the most troublesome empirically. According to UIP, the difference between interest rates in two different currencies will equal the rate of change of the exchange rate between those currencies. However, most studies fail to find this positive one-to-one relation and, indeed, many find a negative relation.[1]

To Irving Fisher, who may have been the first economist to formulate the UIP condition, these anomalous results probably would not have come as a much of a surprise (Dimand, 1999). Fisher viewed UIP as the dual of the interest rate compared to inflation relation or what has come to be called "the Fisher Equation."[2] He saw both as examples of a general relation that links interest rates in different standards.

In discussing this relation, he did so under the heading of "appreciation and interest," first in a monograph of that title published in 1896 by the American Economic Association, and later in two books on the subject of interest-rate determination, the *Rate of Interest* (1907) and his later and more often cited *Theory of Interest* (1930).

In the Fisher equation, the interest rates in question are the nominal and real rates of interest and the link between them, the expected rate of inflation, i.e., the rate at which money depreciates (or appreciates) in terms of goods. For UIP, the interest rates are the nominal interest rates of the two countries in question and the link between them is the

[1] 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, Foresi, and Telmer (2001), Chinn and Meredith (2001), Bekaert, Wei, and Xing (2003), and Brennan and Xia (2005).

expected rate of change of the exchange rate, the rate at which the one currency is expected to depreciate (or appreciate) in terms of the other.

In Fisher's discussions of the empirical evidence surrounding the relation between interest and inflation, he saw the relation as very often subject to violation in the real world. The reason, he claimed, was that people generally did not "adjust at all accurately and promptly" to changes in the behavior of prices but did so only with a long lag (1930). He made much the same argument for UIP, presenting evidence of incomplete and delayed adjustment of nominal interest rate differentials to exchange rate movements and also of episodes of what now fall under the heading of "peso problems."

In this paper, we first briefly review Fisher's work on this subject. When we re-examine the performance of UIP since the advent of floating exchange rates in the 1970s, we find the evidence is consistent with Fisher's conjectures.

Like Fisher, we find that the failures of UIP are related to individual episodes in which errors surrounding exchange rate expectations have been persistent, but which in the end are transitory. The first evidence that supports this inference is the improving performance of UIP that we find as we average the data over progressively longer periods. Our second piece of evidence comes from analysis of UIP in conjunction with the other two key international parity conditions, purchasing power parity (PPP) and real interest rate equality (RIE). The short-tern deviations away *from* UIP and PPP are both substantial and highly correlated. This evidence points to exchange rate forecast errors, as opposed to risk premia, as the major force driving the UIP deviations.

We see evidence of this in another way when we examine the deviations away from RIE, which in principle are independent of exchange rate forecast errors. These deviations are relatively small in comparison to the deviations away from the other two parity conditions, and are less correlated with them.

A third piece of evidence derives from a dynamic latent factor model that we use to estimate the magnitudes of the effects of risk premia and exchange rate forecast errors on the UIP relation. For all of the currencies for which we estimate this model, exchange rate forecast errors again appear to be the principal force behind deviations from UIP. Moreover, these results are robust to alternative model specifications and across countries and time periods.

Fisher's claim, made a century ago, that "unforeseen monetary changes" are the major cause of departures from UIP and the appreciation-interest relation more generally appear confirmed.

The paper is organized as follows. In section I we review Fisher's work on UIP. In Section II we provide empirical results on the short and long run behaviour of UIP. In Section III we analyse the sources of deviations away from UIP. In Section IV we show that it is predominately expectational errors which account for the deviations away from UIP in the short run.

I. Fisher on UIP and the relation between appreciation and interest

Fisher's investigation of UIP centered on two bodies of data: yields on U.S. bonds over the period 1870 to 1896, one bond payable in gold and the other in paper, or "greenback," currency; and yields on Indian bonds traded in London between 1865 and 1894, one bond payable in sterling and the other in silver rupees. Fisher discusses these results first in his monograph *Appreciation and Interest* (1896), and then in his two books on the subject (1907, 1930).

In his analysis of the U.S. data, Fisher discusses two important episodes, the 1879 resumption of specie payments and the decades surrounding that episode, and the 1896 presidential election and three years preceding it. In both events he found evidence of behavior consistent with theory. Prior to resumption, yields on currency bonds exceeded yields on gold bonds as they should have, given the expectations of an appreciation in the value of the paper currency relative to gold. At its peak in 1870, the spread between the two stood at 100 basis points. As time passed and the prospects of resumption increased, the spread narrowed, and by mid-1878 had reversed sign. Over the next 15 years the spread between the yields on currency and gold bonds averaged only -37 basis points, and in the earlier part of that period generally stood at -20 basis points or less.

Fisher went on to compare the expected rates of appreciation of the greenback implicit in the yield differentials prior to resumption. In his comparisons he used realized rates over progressively shorter periods, beginning in January 1870 and ending in each instance in January 1879, the actual date of resumption. The expected rate at the start of this sample was 0.8 percent per annum compared to a realized rate of 2.1 percent per annum, a ratio of a bit less than two fifths. Such underestimation was not at all unusual. Not until 1877 did the ratio finally break out of that general range. For a time in 1874 it actually went negative, implying expectations of depreciation rather than appreciation.

If adjustment was incomplete for most of the period prior to resumption, it was certainly not the case in the years leading up to the 1896 presidential election. During that episode, the first of the two peso-problems uncovered by Fisher, which we noted above, developed. Yields on currency bonds and gold bonds both increased, and the spread between the two progressively widened from 30 basis points in 1893 to a peak of 110 basis points in 1896. Fisher's explanation, which subsequent research substantiates, attributed these developments to the free-silver agitation and the fears of impending inflation and dollar depreciation that it engendered.[3] "Both the increases and the wedging apart of the two rates are explainable as effects of the free-silver proposal and its incorporation (July 1896) in the platform of the democratic [sic] party," Fisher wrote (1930).

Fisher conducted a similar analysis using the yield data for India. In the period 1865-1874 when the exchange rate was stable, the yields on gold and silver rupee bonds were almost identical, differing on average by roughly 20 basis points. Then, in 1875, as the rupee began to depreciate, the spreads gradually widened, from an average of close to 40 basis points in the period between 1875 and 1878, to 64 basis points during the period 1879-1887, to over 100 basis points from 1888 through the first half of 1890. After further depreciation in the half decade that followed, the exchange rate stabilized at the par value of 16d/rupee.

Fisher pointed out that market reactions, both to the initial decline and to the eventual stabilization of the rupee, although basically in line with theory, came with substantial lags. In the latter instance, market participants apparently anticipated a further depreciation in the exchange rate, but this depreciation never actually materialized. This incident is the second of the two peso problems highlighted by Fisher.

[3] Hallwood, et al. (2000) provide econometric evidence supporting this interpretation. For historical discussions of this episode see Friedman and Schwartz (1963, Chapter 3; 1982, Chapter 7).

In the *Theory of Interest*, regarding this incident Fisher wrote:

> "[T]he legal par was reached in 1898 and was maintained thereafter, subject only to the slight variations of exchange due to the cost of shipping specie. *But until the par was proved actually stable by two or three years' experience, the public refused to have confidence that gold and the rupee were once more to run parallel.* Their lack of confidence was shown in the difference in the rates of interest in gold and rupee securities during the transition period, 1893-1898, and the two or three succeeding years." (Emphasis is ours)

The rest of Fisher's empirical evidence concerned the behavior of nominal interest rates within countries, in particular Britain and the United States, but also France, Germany, India, and Japan, and also, in the *Rate of Interest* (but not the *Theory of Interest*), China. This evidence ranged from brief historical descriptions of important episodes, to comparisons of the direction of change in nominal interest rates and rates of price change between subperiods of varying lengths, chosen according to whether prices were rising or declining, to comparisons of the standard deviations of nominal interest rates and ex *post* real interest rates derived from the subperiod data, and later, in *The Theory of Interest*, to his computation of simple correlations between contemporaneous values of nominal interest rates and inflation rates and the estimation of distributed lag relations between those two variables.

The standard deviations of *ex post* real interest rates were many multiples of the standard deviations of the nominal interest rates in every instance, ratios four to eight times greater in the data analyzed in *The Rate of Interest* and seven to 13 times greater in the data analyzed in *The Theory of Interest*.

Fisher's comparisons of the changes in inflation rates and *ex post* real interest rates in *The Rate of Interest* tell a similar story. Increases in inflation went hand in glove with decreases in *ex post* real rates, again implying much less than complete adjustment in nominal rates.

Fisher's summation of this evidence is highly illuminating (1907):

> There are two possible explanations for [this inverse relation]. … One is that when prices are rising the cause may not be monetary but may lie in a progressive scarcity of commodities produced and exchanged ... The second reason is that these [price] movements are only imperfectly foreseen"

He went on to argue:

> Doubtless both of these causes play a part in the explanation in particular cases. *Nevertheless there is internal evidence to show that in general the latter factor – unforeseen monetary changes – is the more important.* This evidence consists in the fact that commodity interest fluctuates so widely in some instances becoming negative. (Emphasis is ours)

However, he concluded that "When long periods of price movements are taken, the influence of appreciation on interest is more certain, because in averages covering so many years we may be sure that accidental causes are almost wholly eliminated." And he presented evidence for Britain and the United States for which he used averages spanning a decade or more that was entirely consistent with this statement (1907). The direction of the movements in nominal interest rates and inflation in seven of the eight cases is the same and the variability of the *ex post* real rates is much closer to that of the nominal rates in both countries than in the data for the shorter subperiods.

II. **UIP regressions**

There is substantial evidence that UIP does not in fact hold, or at least not in the short term. Engel (1996) provides a comprehensive survey of this literature. We also show that empirically, UIP does not hold in the short run. However, we are able to explain why deviations from UIP occur and what is driving these results.

We use monthly data for the period January 1970 to December 2005 for 18 countries relative to the United States: Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, and the United Kingdom. We obtain most of these data from the International Monetary Fund's *International Financial Statistics*. Exchange rates are denominated in units of foreign currency per U.S. dollar; interest rates are short-term domestic Treasury bill or money market rates.

II.A. *UIP Regressions*

We begin by running standard UIP regressions of the following form for each country individually over the full sample period:

$$s_{t+1} - s_t = \alpha + \beta(i_t - i_t *) + e_{t+1} \qquad (1)$$

where s_{t+1}- s_t is the one-period change in the log spot exchange rate and i_t- i_t* is the corresponding foreign interest differential compared to that of the U.S.

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, we would expect that on average, the foreign currency would appreciate by one percent over the next *n* periods. Therefore, a test of the hypothesis $\alpha = 0$ and $\beta = 1$ in (1) provides a test of uncovered interest parity.

However, most studies reject this hypothesis. Indeed, one of the most puzzling features 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 by its other name as "the forward premium puzzle," is now so well documented that it has taken on the aura of a stylized fact. As a result it has spawned an extensive second-generation body of literature that attempts to explain it.

Insert Table 1

The regression results we report in Table 1 are very much in line with other results reported in other studies. In 13 of the 18 countries, the estimates of β are zero or negative, and in 14 are significantly different from the theoretical value of unity at the five per cent level or below. In all instances the coefficients of determination in these regressions are extremely low, and in most instances zero. In only one case, Spain, is the estimate of β both positive and significantly different from zero. In short, there is little if any evidence that UIP holds in the monthly data for these countries.

We run similar regressions using Fisher's data for the United States and India. We obtain these data from Tables 11 and 12 in Chapter 19 of *The Theory of Interest* (1930). Our results are reported in Table 2 and are similar to those reported in Table 1.[4] In the U.S. case, the estimate of the slope coefficient β is positive, and in the Indian case, negative. In both instances, these estimates are both nonsignificantly different from zero and nonsignificantly different from unity. We are interested to see that although Fisher finds subperiods in which UIP has some validity, the relation still does not hold well in general.

Insert Table 2

But there is more to the story. As discussed above, Fisher's explanation of the failures of UIP and the appreciation-interest relation centre more generally on small-sample problems and other "accidental" factors affecting that relation. To investigate the possible effects of such influences, we run rolling regressions and regressions, using pooled data averaged over progressively longer time periods.

II. B. ***Rolling regressions***

In Figure 1, the solid line plots cross-country averages of the slope coefficients of five-year rolling regressions based on equation (2) and estimated for the G7 countries plus the Netherlands. The dotted lines show a plus or minus one-standard-deviation range for those averages. The averages are for the period 1975:2 and 2005:12. We report the full results in Table 3.

Insert Table 3

We plot the coefficients at the starting points of the sample periods over which the regressions are run. The standard deviations are deviations of the coefficients in the individual-country regression about the average.

What stands out in the figure are the often sizable variations in the slope coefficients over time. We see periods, such as the early 1970s and early 1990s, during which most of the individual-country coefficients are positive and for a time close to unity. But we see these periods followed by long periods of systematic movements away from this UIP value.

Insert Figure 1

The first such departure occurred in the 1980s. At the start of the decade, we see a gradual decrease in the magnitude of the regression slope coefficients and then large negative values by the middle of the decade. This decade was the period of Reagan-Volcker disinflation, when the Federal Reserve contained and then reversed the process of rising inflation. However, expectations for the inflation decline changed more slowly. During the following five year period, we see a gradual reversion towards unit slope coefficients in the rolling UIP regressions and hence a return to UIP.

A second major shock was the 1992 ERM crisis, when the United Kingdom, followed later by Italy and Spain, pulled out of the European Monetary System. Here we see a sharp rise in the average slope coefficient.

A third major event, like the one in the1980s in which we see falling and eventually negative slope estimates, began in the mid to late 1990s prior to the introduction of the euro. This event appears attributable to the uncertainty that accompanied that episode.

II.C. *UIP regressions using temporally averaged data*

If the current problems surrounding UIP are in fact episodic phenomena that are due, as Fisher put it, to "accidental causes," then his solution of averaging the data is appropriate.

We do this averaging in Figure 2 and in the regressions in Table 3. In the three panels of Figure 2 we show the plots of the UIP relation based on five-year, 15-year, and full-period averages of the data for our 18 countries. To provide a theoretical frame of reference, we also show a 45 degree line drawn through the origin. In Table 3 we list the corresponding regression results.

In the five-year averaged data there is a positive, but nevertheless weak, relation between the exchange rate change and the interest differential. However, the picture changes markedly as the period over which we average the data lengthens. We see this relation clearly improving in the bottom two panels of Figure 1. When we look at the 15-year averages, we find a strong positive relation, and an even stronger relation for the full-period averages.

Insert Figure 2

The regression results confirm these observations. As the period over which we compute the averages lengthens, the slope coefficients in the regressions increase from less than 0.02 to 0.75, and the standard errors of those regressions decrease from close to six percentage points to 1.2 percentage points. Although we can always reject the hypothesis of a unit slope, it is clear from these results that as a long-run first approximation, UIP contains a substantial kernel of truth.

III. Short-run behavior and the sources of UIP deviations

In theory, uncovered interest parity is an ex ante concept, positing equality of expected nominal returns across countries:

$$i_t = i_t* + \mathrm{E}_t\left[\left(s_{t+1} - s_t\right)\right] \tag{2}$$

However, empirical investigations of UIP generally use actual, ex post changes in exchange rates as a proxy for their unobservable ex ante counterparts. Therefore, deviations from UIP empirically can arise because of differences between actual and expected exchange rate changes and because of differences in the riskiness of the two assets:

$$i_t - i_t* - (s_{t+1} - s_t) = \rho_t - \varepsilon_{st}\,, \tag{3}$$

where ε_{st} is the exchange rate forecast error and ρ_t is the ex ante risk premia.

The risk premia will be positive (or negative) if investors require an expected excess return on a currency to compensate for the risk of holding it. Under the usual assumptions of rational expectations, exchange rate forecast errors will be random.

However, as Fisher pointed out, there are conditions under which these errors might in fact be systematic over time. One situation in which errors are systematic will occur is if investors anticipate changes in the underlying process generating the return distribution. In modern terminology we call this phenomenon the "Peso problem." A second situation is that of a monetary shock, in the form of a sudden shift in monetary regime. Before investors learn about the true process that generates the returns, there may be a period in which forecast errors again are systematic over time, rather than random. Fisher discussed the first of these two cases in the context of the 1896 U.S. presidential election and the second in the context of the stabilization of the rupee.

III. A. ***The three-parity framework***

To begin to disentangle the effects of risk premia and systematic exchange rate forecast errors on UIP, we use the framework developed in Marston (1997). We examine the deviations from UIP together with those from PPP and RIE.

We consider PPP, written here in terms of expected rates of change of the variables:

$$E_t\,[\pi_{t+1} - \pi_{t+1}^*] = E_t\,[s_{t+1} - s_t] \tag{4}$$

where π and π^* are the rates of inflation in the home and foreign countries, respectively.

Deviations from PPP can arise either as a result of exchange rate forecast errors, ε_{st}, inflation forecast errors, ε_{pt}, or expected changes in the real exchange rate θ_t:

$$E_t\,[\pi_{t+1} - \pi_{t+1}^*] - E_t\,[s_{t+1} - s_t] = \varepsilon_{st} + \varepsilon_{pt} + \theta_t \tag{5}$$

When we compare (5) and (3), we see that risk premia do not affect PPP deviations, but exchange rate errors affect both UIP and PPP deviations.

However, UIP, PPP, and RIE are not independent. The deviations from any one of these relations is equal to the algebraic sum of the deviations from the other two. Thus, by subtracting (5) from (3), we get an equation for the real-interest differential, $r - r^*$, in the form:

$$r_t - r_t^* = \rho_t - \theta_t - \varepsilon_{pt}\ , \tag{6}$$

where ε_{pt} is an inflation forecast error.

We note that when we compare (6) with (3), UIP deviations and RIE deviations have only risk premia as a common source. Exchange rate forecast errors do not matter for RIE.

When we compare the time paths of deviations from PPP and RIE with those of from UIP we can infer the causes of the UIP deviations observed in the data. We make this comparison first in Figure 3, where we use dollar-pound as a representative currency and plot the deviations from the three parity conditions, and in Table 4, where we present the correlations between the deviations from the three parity conditions for the Euro (prior to 1999 we use the Deutschmark), British pound sterling, and Japanese yen against the dollar.

Insert Figure 3

What immediately strikes us about Figure 3 are, on the one hand, the high correlation between the UIP and PPP deviations and the similar and substantial magnitudes of both, and on the other hand, the low correlation between these deviations and the RIE deviations, which are very much smaller in magnitude. In Table 4 we illustrate this for all three major currencies using correlation statistics for the full sample period.

Insert Table 4

Exchange rate forecast errors are common to equations (3) and (5), and explain UIP and PPP deviations. However, they do not appear in equation (6) to explain RIE deviations. In contrast, equation (6) and equation (3) have risk premia as a common factor. Therefore, we infer that exchange rate forecast errors rather than risk premia are the major driving force between UIP deviations.

III. B. A dynamic factor approach to decomposing the UIP relation

To investigate the process driving the UIP deviations, we adopt a dynamic latent factor model as in Harvey (1991). Although this type of model has been extensively used in other fields, univariate models generally predominate in exchange rate studies (e.g., Wolff, 1991; and Nijman, Palm, and Wolff, 1993).

In equations (3), (5), and (6) we wrote the deviations from UIP, PPP, and RIE, respectively, in terms of the risk premium, ρ_t, exchange rate forecast errors, ε_{st}, inflation forecast errors, ε_{pt}, and expected changes in the real exchange rate, θ_t. Since these last two variables never appear separately in any of the equations, we cannot disentangle their effects. Thus, we have a three-equation system with three common factors, risk premia, exchange rate forecast errors, and a factor that combines inflation forecast errors and expected changes in the real exchange rate. Each of the three parity conditions is affected by just two of these factors, which makes it possible for us to distinguish between the effects of risk premia and exchange rate forecast errors in deviations away from the UIP relation.

Risk premia affect only nominal and real interest differentials, not inflation differentials. Systematic errors in forecasting exchange rates affect only nominal, not real interest differentials. Thus, we have a system of three parity condition equations with three unknown factors. So, by estimating any combination of two parity conditions, we are able to observe the ex post effects of risk premia and exchange rate forecast errors on deviations from UIP, and the ex post effects of risk premia and the combination of expected changes on the real exchange rate and inflation forecast errors on deviations from RIE.

We model this set of joint parity conditions by estimating a dynamic latent factor model for the unobservable components on the right hand sides of equations (3) and (6).

We estimate and identify the model on the UIP and the RIE parities using the following set of equations in (7). In (7), we again denote the common latent factor, the risk premia, by ρ_t:

$$\begin{pmatrix} i_t - i_t{*} - (s_{t+1} - s_t) \\ r_t - r_t{*} \end{pmatrix} = \begin{pmatrix} c_{UIP} \\ 0 \end{pmatrix} + \begin{pmatrix} 1 \\ 1 \end{pmatrix} \rho_t + \begin{pmatrix} \nu_t^{UIP} \\ \nu_t^{RIE} \end{pmatrix}. \tag{7}$$

We assume the measurement errors ν_t^{UIP} and ν_t^{RIE} to be iid with a zero correlation and variances σ_{UIP}^2 and σ_{RIE}^2, respectively. We consider the common factor for the risk premium, ρ_t, to be a latent factor, modeled by an AR(1) process

$$\rho_t = c_\rho + \phi_\rho \rho_{t-1} + \eta_{\rho t}, \qquad \eta_{\rho t} \sim N(0, \sigma_\rho^2) \tag{8}$$

The model described by equations (7) and (8) is a state space model. For identification reasons we do not include a constant term in the measurement equation for RIE; the constant term in the UIP equation is denoted by c_{UIP}. We estimate the model parameters and latent factors by maximum likelihood and compute the likelihood function recursively, using the Kalman filter. Once we have determined the common factor, the risk premia, we can identify the exchange rate forecast error ε_{st} and the joint component composed of the inflation forecast error and expected real exchange rate change, $\theta_t + \varepsilon_{pt}$, as

$$\hat{\varepsilon}_{st} = (s_{t+1} - s_t) - (i_t - i_t^*) + \hat{\rho}_t \tag{9}$$

and

$$\hat{\theta}_t + \hat{\varepsilon}_{pt} = \hat{\rho}_t - (r - r_t^*), \tag{10}$$

where a hat denotes an estimated value.

In Table 5 we present estimation results for the main currencies in our sample, the euro, the British pound sterling, and the Japanese yen. Since we obtain similar results for all other countries in our sample, we focus on only the major currency combinations.

Insert Table 5

For the GBP, we find that the common factor ρ_t is nonstationary. The regression coefficients on the lagged risk premium coefficient, φ_t range from 0.941 for the Japanese yen and 0.961 for the pound sterling. (This nonstationarity is not a problem in this model, since the state space model does not require stationarity.) Both of the specific factors are stationary.

In Figure 4 we plot the time series of the two latent-factor estimates. In comparing the two visually, we note that in all three cases, the scales for the estimates of the risk premia are much smaller than those for the exchange rate forecast errors. The variance of the risk premia is less than 0.07 in all cases, but the variance of the exchange rate and the associated forecast errors are roughly ten times larger.

Insert Figure 4

Our results are extremely robust to using the alternative combinations of UIP and PPP, and PPP and RIE to derive parameter estimates.

IV. Expectational Errors and the UIP puzzle

Exchange rate forecast errors appear to play a much more important role in terms of variability than do risk premia. The question that we now wish to consider is the relative size of the impacts of the two on UIP deviations. To answer this question, we decompose the slope coefficient in the UIP regression (2) in terms of the moments derived in Table 5. Here, we also show that it is the forecasting errors, and the subsequent large variance in forecasting exchange rates, that account for the large negative estimated slope coefficients normally found in UIP regressions

To see the effect of exchange rate expectations errors on results of UIP regression results, we first write the estimated slope coefficient for the UIP regression in terms of the standard OLS formula:

$$\hat{\beta} = \frac{\operatorname{cov}\left(i_t - i_t^*, s_{t+1} - s_t\right)}{\operatorname{var}\left(i_t - i_t^*\right)}. \tag{11}$$

It is clear from (11) that a negative slope coefficient can only occur if $\mathrm{cov}(i_t - i_t^*, s_{t+1} - s_t)$, the covariance between the interest differential and the exchange rate change, is negative.

To determine the specific effects of risk premia and exchange rate expectation errors on the regression coefficient, we use the expression for the deviations from UIP given in equation (3) to get:

$$\mathrm{cov}(i_t - i_t^* - (s_{t+1} - s_t), s_{t+1} - s_t) = \mathrm{cov}(\rho_t, s_{t+1} - s_t) - \mathrm{cov}(\varepsilon_{st}, s_{t+1} - s_t). \qquad (12)$$

We can rewrite this equation in terms of the covariance of the interest differential and the change in the exchange rate as:

$$\mathrm{cov}(i_t - i_t^*, s_{t+1} - s_t) = \mathrm{cov}(\rho_t, s_{t+1} - s_t) + \mathrm{var}(s_{t+1} - s_t) - \mathrm{cov}(\varepsilon_{st}, s_{t+1} - s_t). \qquad (13)$$

The result is an expression for the slope coefficient from the UIP regression in terms of both the risk premia and the exchange rate forecast errors. Since a negative slope in the UIP regression can only occur if $cov(i_t - i_{t^*}, s_{t+1} - s_t)$ is negative, then either the covariance of exchange rate changes with the risk premia has to be negative, or the covariance of exchange rate changes with the exchange rate forecast error has to be positive, and in either instance the value must be of sufficient magnitude to outweigh the effects of the other right-hand-side terms in (13).

Using the results from the dynamic latent factor model, we can estimate the variances and covariances of the various factors and decompose the significance of these two variables on the beta estimates. Table 6 shows these moments.

Insert Table 6

We find very large positive estimates for the variances and covariances of the variables surrounding exchange rate expectations errors. The covariance between these errors and exchange rate changes ranges from 8.9 for the GBP to 10.3 for the other two currencies. The variance of the exchange rate has a similar magnitude. Although the covariance of the risk premia with the exchange rate is negative, the empirical estimates for the all three currencies are all small in value with the highest value being -0.1 for the sterling.

In the final two rows of Table 6 we report the ratios

$$\frac{\operatorname{cov}(\rho_t, s_{t+1} - s_t)}{\operatorname{var}(i_t - i_t*)}$$

and

$$\frac{\operatorname{var}(s_{t+1} - s_t) - \operatorname{cov}(\varepsilon_{st}, s_{t+1} - s_t)}{\operatorname{var}(i_t - i_t*)}$$

as a percentage of the alternative estimate for the UIP slope coefficient in equation (14)

$$\frac{\operatorname{cov}(\rho_t, s_{t+1} - s_t) + \operatorname{var}(s_{t+1} - s_t) - \operatorname{cov}(\varepsilon_{st}, s_{t+1} - s_t)}{\operatorname{var}(i_t - i_t*)} \,. \tag{14}$$

For both the GBP and the yen, the first component in the numerator of (14), which measures the relative effect of risk premia, on average accounts for 30% of the size of the estimated betas from the five-year rolling regressions. In contrast, errors in exchange rate forecasts account on average for 70%. For the euro, the contrast between the two effects is even greater, showing 5% and 95% for the risk premia and the exchange rate forecast error, respectively.

Interestingly, the results support the work of Bacchetta et al. (2007) on the predictability of excess returns on foreign exchange markets due to the predictability of expectational errors, and the results from the survey data on the dispersion of beliefs by Jongen et al. (2006) who find long periods of large dispersion in market participants' forecasts of exchange rates. Bacchetta and Wincoop (2005) attribute this predictability to 'rational inattention,' a situation in which investors are rational, but due to significant information costs, are slow in responding to new information.

These empirical results provide additional support for Fisher's proposition that the negative slope coefficients that are often obtained in UIP regressions depend mainly on errors in exchange rate expectations and not on time variation in risk premia.

Conclusion

Our results on UIP are consistent with those reported a century or more ago by Irving Fisher in his studies of the relation between appreciation and interest, both in its UIP and interest compared to inflation versions. Consistent with Fisher's view, we find evidence of the important role played by episodic phenomena in disturbing that relation. Like Fisher, we too find that the influence of such phenomena dissipates over time.

We conclude that there are long-run deviations from parity conditions that appear to be caused by large, but infrequent, shocks to the monetary environment. These shocks systematically affect the error in forecasting the change in exchange rates. Over the long term, these errors are less important and we find empirical support for UIP.

Using Marston's (1997) analysis, we investigate the possibility of a common factor driving deviations from parity conditions. We find extremely high correlation coefficients between UIP and PPP deviations that we identify with exchange rate forecasting errors.

Using a dynamic latent variable model, we estimate the risk-premia and exchange rate-forecast-error parameters that drive changes in deviations from UIP. We find evidence of large and persistent forecasting errors. Although we can hypothesize about what may be driving the persistent errors in forecasting exchange rates, we do not attempt to model predictability in excess returns. Instead, we provide strong empirical support that it is indeed expectational errors rather than risk premia that underlie the short-run deviations from UIP.

Appendix: Robustness analysis

In this appendix we present evidence on the robustness of the estimated factors from the dynamic latent factor models (7) and (8).

Instead of using the risk premium as our latent variable, we could model the exchange rate forecast error as a latent variable. (We note that as a final alternative to specification, we could model the joint risk premium and inflationary forecast term, $\theta_t + \varepsilon_{pt}$, as a latent variable. However, we do not show these results here.) By modeling the exchange rate forecast error as a latent variable, we combine the deviations from UIP and PPP in the measurement equation to arrive at an alternative model that takes the form:

$$\begin{pmatrix} i_t - i_t{*} - (s_{t+1} - s_t) \\ (\pi_{t+1} - \pi_t) - (\pi_{t+1}{*} - \pi_t{*}) - (s_{t+1} - s_t) \end{pmatrix} = \begin{pmatrix} c'_{UIP} \\ 0 \end{pmatrix} + \begin{pmatrix} -1 \\ -1 \end{pmatrix} \varepsilon_{st} + \begin{pmatrix} \nu_t^{UIP} \\ \nu_t^{PPP} \end{pmatrix}, \tag{15}$$

and

$$\varepsilon_{st} = c_s + \phi_s \varepsilon_{s,t-1} + \eta_{st}, \qquad \eta_{st} \sim N(0, \sigma_s^2), \tag{16}$$

with errors ν_t^{UIP} and ν_t^{PPP} assumed to be independently and identically normally distributed. Again, we estimate the parameters and the latent forecast errors ε_{st} through the Kalman filter.

After estimation we can extract the risk premium ρ_t and the inflationary components $\theta_t + \varepsilon_{pt}$ as follows:

$$\hat{\rho}_t = i_t - i_t * - (s_{t+1} - s_t) + \hat{\varepsilon}_{st} ,$$

$$\hat{\theta}_t + \hat{\varepsilon}_{pt} = (\pi_{t+1} - \pi_t) - (\pi_{t+1} * - \pi_t *) - (s_{t+1} - s_t) + \hat{\varepsilon}_{st} \ .$$

In Table 7 we present the parameter estimates for the bivariate model. The model consists of both the deviations of UIP and PPP. We note that the latent forecast error has a slightly positive, albeit nonsignificant, autoregressive coefficient.

To compare the estimated series for the forecast errors in the two models, we compute correlations between the risk premia and forecasts errors that we compute from the dynamic factor models comprising equations (7) and (8) and equations (15) and (16), respectively.

In Table 8 we report these results. The table shows that the time series for risk premia ρ_t and the forecast errors ε_{st} that we compute from the two models are very, almost extremely, highly correlated, thus providing extremely robust results to the model specification.

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Table 1: Summary of results of individual-country UIP regressions

We run the regressions summarized below by using monthly data from 1970:2-2005:12 obtained from *International Financial Statistics*. These regressions take the form

$$s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + e_{t+1}, \qquad \text{(T.1)}$$

where $s_{t+1} - s_t$ is the one-period change in the log of the foreign compared to the U.S. spot exchange rate and $i_t - i_t^*$ is the corresponding foreign compared to the U.S. interest differential. We note that for some countries interest-rate data are only available at a later starting date.

Full Sample Regression	**Intercept**	***Standard Error***	***t-Stat α=0***	**Beta**	***Standard Error***	***t- Stat B=1***	**R Square**	***SEE***	**Nobs**
Austria	-3.327	1.931	-1.723	-1.135	0.705	-3.029	0.006	37.449	431
Belgium	0.663	2.065	0.321	-1.488	0.854	-2.915	0.007	37.828	431
Canada	0.846	1.065	0.794	-0.463	0.480	-3.048	0.002	17.616	431
Denmark	0.837	2.317	0.361	-0.430	0.511	-2.800	0.002	37.572	408
Finland	-1.440	0.176	-8.173	0.002	0.005	-196.260	0.000	3.657	431
France	1.654	2.166	0.764	-0.982	0.733	-2.704	0.004	37.075	431
Germany	-2.678	2.214	-1.209	-1.431	0.854	-2.847	0.008	38.528	366
Greece	-3.343	4.106	-0.814	0.217	0.381	-2.055	0.001	39.054	250
Ireland	-0.929	2.576	-0.361	0.559	0.501	-0.881	0.004	38.459	331
Italy	-0.369	3.032	-0.122	0.609	0.559	-0.699	0.003	36.398	346
Japan	-2.450	0.121	-20.176	-0.008	0.003	-316.039	0.016	2.512	431
Netherlands	-1.699	0.158	-10.752	0	0.004	-241.009	0	3.277	431
Norway	-0.508	2.136	-0.238	0.139	0.493	-1.747	0	34.511	413
Portugal	4.946	1.774	2.788	-0.180	0.232	-5.087	0.001	36.828	431
Spain	3.024	0.395	7.657	0.262	0.093	-7.939	0.048	6.325	324
Sweden	0.646	1.959	0.330	0.314	0.532	-1.289	0.001	35.761	431
Switzerland	-6.287	2.955	-2.128	-1.206	0.604	-3.654	0.011	42.088	364
UK	5.557	2.541	2.187	-1.930	0.783	-3.741	0.014	35.270	431

Table 2: Results of UIP regressions based on Irving Fisher’s (1930) data for U.S. gold and greenback bonds and Indian sterling and rupee bonds

In the regressions summarized below we use the data reported in Tables 11 and 12 of Fisher's *The Theory of Interest* (1930).

These regressions take the form

$$s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + e_{t+1}, \quad \text{(T.1)}$$

where $s_{t+1} - s_t$ is the one-period change in the log of the foreign compared to the U.S. spot exchange rate and $i_t - i_t^*$ is the corresponding foreign compared to the U.S. interest differential.

Average	**Intercept**	***Standard Error***	***t Stat α=0***	**Beta**	***Standard Error***	***t Stat β=1***	**R Square**	***SEE***
U.S. Bonds	-1.037	0.724	-1.433	2.608	1.434	1.122	0.091	4.283
Indian Bonds	-0.020	1.369	-0.014	-2.012	2.435	-1.237	0.019	4.312

Table 3: Results of UIP regressions for nonoverlapping averages of the data

The regressions we summarize below are pooled regressions that we run using the averaged data. These regressions take the form

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

where $s_{t+1} - s_t$ is the one-period change in the log of the foreign compared to the U.S. spot exchange rate and $i_t - i_t^*$ is the corresponding foreign compared to the U.S. interest differential.

Observations are missing for some countries, since some series have later starting dates. The countries we analyze are Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, the Netherlands, Norway, Portugal, Spain, Sweden, Switzerland, and the UK, all compared to the US.

Average	**Intercept**	***Standard Error***	***t-Stat α=0***	**Beta**	***Standard Error***	***t-Stat β=1***	**R Square**	***SEE***	**Nobs**
5-year	-0.088	0.631	-0.14	0.259	0.168	-4.413	0.023	5.864	102
15-year	-0.457	0.314	-1.452	0.667	0.097	-3.428	0.612	1.576	32
35-year	-0.729	0.349	-2.089	0.758	0.110	-2.207	0.749	1.245	18

Table 4: Correlation of deviations from ex-post UIP, PPP, and RIE

The currencies we use are the euro (EUR), British pound sterling (GBP), and Japanese yen (JPY), all compared to the U.S. dollar. The estimation period is from January 1976 to December 2005

Full sample	UIP & PPP	UIP & RIE	PPP & RIE
Euro	0.994	0.034	0.11
Japan	0.981	0.030	0.171
UK	0.981	-0.020	0.153

Table 5: Estimation results for the dynamic factor model (7) and (8)

This table presents parameter estimates for the dynamic factor model consisting of equations:

$$\begin{pmatrix} i_t - i_t^* - (s_{t+1} - s_t) \\ r_t - r_t^* \end{pmatrix} = \begin{pmatrix} c_{UIP} \\ 0 \end{pmatrix} + \begin{pmatrix} 1 \\ 1 \end{pmatrix} \rho_t + \begin{pmatrix} \nu_t^{UIP} \\ \nu_t^{RIE} \end{pmatrix} \qquad \text{(T.2)}$$

Where the risk premium, ρ_t, is modeled as:

$$\rho_t = c_\rho + \phi_\rho \rho_{t-1} + \eta_{\rho t},\ \eta_{\rho t} \sim N(0, \sigma_\rho^2) \qquad \text{(T.3)}$$

The currencies we use are the euro (EUR), British pound sterling (GBP), and Japanese yen (JPY), all compared to the U.S. dollar. The estimation period is from January 1976 to December 2005.

	EUR	**GBP**	**JPY**
c_{UIP}	-0.019 (0.145)	0.096 (0.143)	0.148 (0.158)
c_ρ	0.003 (0.003)	0.002 (0.003)	-0.006 (0.003)
ϕ_ρ	0.943 (0.018)	0.961 (0.002)	0.941 (0.016)
σ_ρ	0.050 (0.002)	0.073 (0.002)	0.069 (0.002)
σ_{UIP}	3.209 (0)	3.008 (0)	3.303 (0)
σ_{RIE}	0.024 (0)	0.018 (0)	0.012 (0)

Table 6: Moments

The currencies we use are the euro (EUR), British pound sterling (GBP), and Japanese yen (JPY), all compared to the U.S. dollar. The estimation period is from January 1976 to December 2005. We calculate *OLS beta* coefficient from the UIP regression

$$s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + e_{t+1}, \qquad \text{(T.1)}$$

where $s_{t+1} - s_t$ is the one-period change in the log of the foreign compared to the U.S. spot exchange rate and $i_t - i_t^*$ is the corresponding foreign compared to the U.S. interest differential over the sample period January 1976 to December 2005. We calculate the *Implied beta* from equation

$$\frac{\operatorname{cov}(\rho_t, s_{t+1} - s_t) + \operatorname{var}(s_{t+1} - s_t) - \operatorname{cov}(\varepsilon_{st}, s_{t+1} - s_t)}{\operatorname{var}(i_t - i_t^*)}. \qquad \text{(T.4)}$$

	EUR	**GBP**	**JPY**
$\operatorname{var}(\rho_t)$	0.0221	0.0669	0.0417
$\operatorname{var}(s_{t+1} - s_t)$	10.348	8.9698	10.272
$\operatorname{var}(i_t - i_t^*)$	0.0395	0.0328	0.0448
$\operatorname{cov}(i_t - i_t^*, s_{t+1} - s_t)$	-0.0477	-0.0651	-0.0882
$\operatorname{var}(\varepsilon_{st})$	10.387	8.9736	10.382
$\operatorname{cov}(\rho_t, s_{t+1} - s_t)$	-0.0438	-0.1005	-0.071
$\operatorname{cov}(\varepsilon_{st}, s_{t+1} - s_t)$	10.323	8.9130	10.290
OLS beta	-1.204	-1.943	-1.805
Implied beta	-2.097	-3.164	-3.435
Variaton due to $\operatorname{cov}(\rho_t, s_{t+1} - s_t)$	4.76%	29.48%	29.86%
Variaton due to $\operatorname{var}(s_{t+1} - s_t)$- $\operatorname{cov}(\varepsilon_{st}, s_{t+1} - s_t)$	95.24%	70.52%	70.14%

Table 7: Estimation results for the dynamic factor model using deviations from UIP and PPP

The table presents parameter estimates for the dynamic factor model consisting of equations by the exchange rate forecast error,

$$\begin{pmatrix} i_t - i_t * - (s_{t+1} - s_t) \\ (\pi_{t+1} - \pi_t) - (\pi_{t+1} * - \pi_t *) - (s_{t+1} - s_t) \end{pmatrix} = \begin{pmatrix} c'_{UIP} \\ 0 \end{pmatrix} + \begin{pmatrix} -1 \\ -1 \end{pmatrix} \varepsilon_{st} + \begin{pmatrix} \nu_t^{UIP} \\ \nu_t^{PPP} \end{pmatrix} \qquad \text{(T.5)}$$

and $\varepsilon_{st} = c_s + \phi_s \varepsilon_{s,t-1} + \eta_{st}$, $\qquad \eta_{st} \sim N(0, \sigma_s^2)$ (T.6)

The currencies we use are the euro (EUR), British pound sterling (GBP), and Japanese yen (JPY). All exchange rates are against the U.S. dollar. The estimation period is from January 1976 to December 2005.

	EUR	GBP	JPY
c_{UIP}	0.053 (0.008)	0.034 (0.013)	-0.089 (0.010)
c_s	0.024 (0.163)	-0.081 (0.147)	-0.134 (0.111)
ϕ_s	0.041 (0.034)	0.081 (0.053)	0.090 (0.047)
σ_s	3.228 (0.105)	2.983 (0.103)	3.214 (0.102)
σ_{UIP}	0.022 (0.014)	0.258 (0.004)	0.207 (0.001)
σ_{PPP}	0.150 (0.007)	0.002 (0.012)	0.012 (0.002)

Table 8: Correlations between estimated risk premia and forecast errors

We compute correlations between time series for risk premia ρ_t and forecast errors estimates ε_{st}, for the two dynamic factor models 1) for the risk premia, by ρ_t:

$$\begin{pmatrix} i_t - i_t^* - (s_{t+1} - s_t) \\ r_t - r_t^* \end{pmatrix} = \begin{pmatrix} c_{UIP} \\ 0 \end{pmatrix} + \begin{pmatrix} 1 \\ 1 \end{pmatrix} \rho_t + \begin{pmatrix} \nu_t^{UIP} \\ \nu_t^{RIE} \end{pmatrix} \tag{T.2}$$

Where the risk premium, ρ_t, is modeled as:

$$\rho_t = c_\rho + \phi_\rho \rho_{t-1} + \eta_{\rho t}\,,\ \eta_{\rho t} \sim N(0, \sigma_\rho^2) \tag{T.3}$$

And 2) by the exchange rate forecast error,

$$\begin{pmatrix} i_t - i_t^* - (s_{t+1} - s_t) \\ (\pi_{t+1} - \pi_t) - (\pi_{t+1}^* - \pi_t^*) - (s_{t+1} - s_t) \end{pmatrix} = \begin{pmatrix} c'_{UIP} \\ 0 \end{pmatrix} + \begin{pmatrix} -1 \\ -1 \end{pmatrix} \varepsilon_{st} + \begin{pmatrix} \nu_t^{UIP} \\ \nu_t^{PPP} \end{pmatrix} \tag{T.5}$$

and $\varepsilon_{st} = c_s + \phi_s \varepsilon_{s,t-1} + \eta_{st}$, $\quad \eta_{st} \sim N(0, \sigma_s^2)$ (T.6)

The currencies we use are the euro (EUR), British pound sterling (GBP), and Japanese yen (JPY). All exchange rates are against the U.S. dollar. The estimation period is from January 1976 to December 2005.

	EUR	**GBP**	**JPY**
Risk premia ρ_t	0.958	0.975	0.969
Forecast errors ε_{st}	0.999	0.999	0.999

Figure 1: Averages of coefficients from five-year rolling regressions for the G7 countries and one-standard-deviation bounds

We report the beta estimates from the 5 year rolling regression summarized below by using monthly data from 1970:2-2005:12 obtained from *International Financial Statistics*. These regressions take the form

$$s_{t+1} - s_t = \alpha + \beta(i_t - i_t^*) + e_{t+1}, \qquad (T.1)$$

where $s_{t+1} - s_t$ is the one-period change in the log of the foreign compared to the U.S. spot exchange rate and $i_t - i_t^*$ is the corresponding foreign compared to the U.S. interest differential. We note that for some countries interest-rate data are only available at a later starting date. We take the average for the G7 countries, and also report the one-standard-deviation bounds.

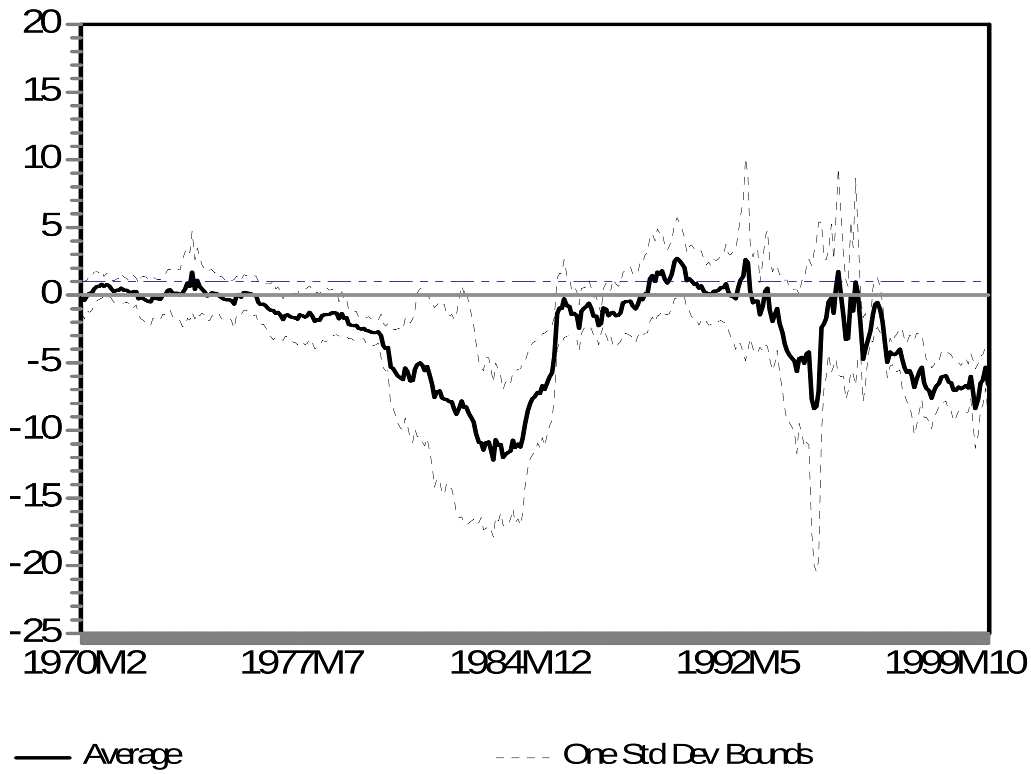
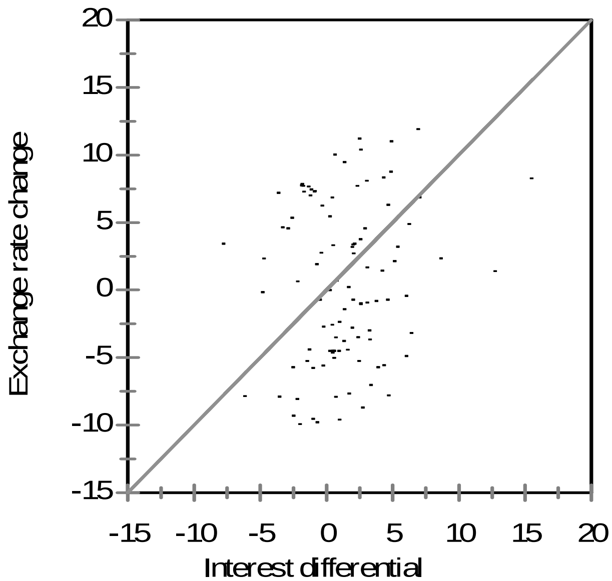


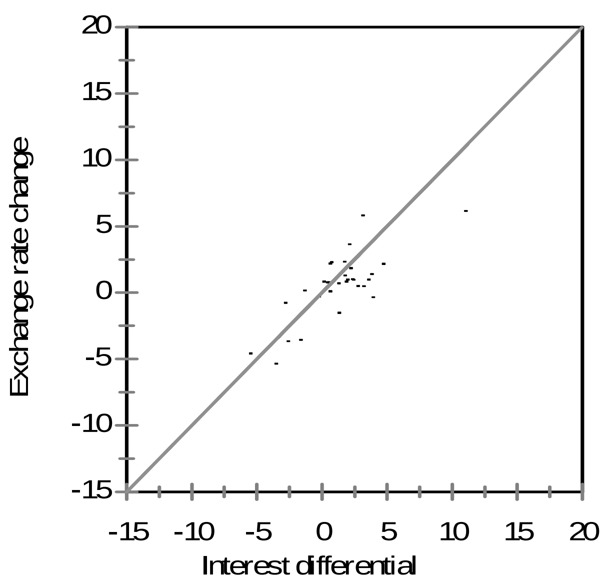
Figure 2: UIP relations based on five-year, 15-year and full-period averages

In the charts below we plot period averages of the exchange rate change against the interest differential for all 18 countries. For Spain and Portugal these data begin in 1985. The 5-year sample periods were 1970-1975, 1975-1980, 1985-1990, 1990-1995, 1995-2000, 2000-2005; the 15 year sample periods were 1975-1990 and 1990-2005.

Five-year Averages



Fifteen-year Averages



Full-period Averages

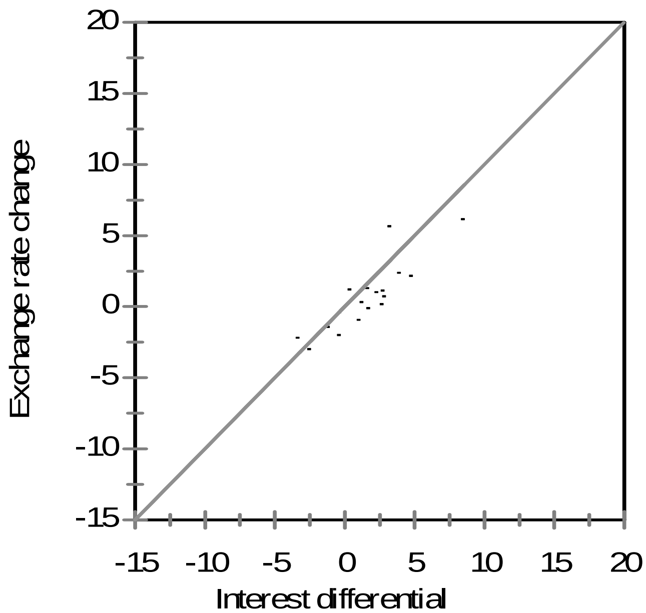


Figure 3: Ex post deviations from UIP, PPP, and RIE for the US and the UK from January 1970 – December 2005.

We plot ex post deviations from the three equations:

$$i_t - i_t{}^* - (s_{t+1} - s_t) = \rho_t - \varepsilon_{st}, \qquad \text{(T.3)}$$

$$E_t[\pi_{t+1} - \pi_{t+1}{}^*] - E_t[s_{t+1} - s_t] = \varepsilon_{st} + \varepsilon_{pt} + \theta_t \qquad \text{(T.5)}$$

$$r_t - r_t{}^* = \rho_t - \theta_t - \varepsilon_{pt}, \qquad \text{(T.6)}$$

Where s_{t+1}- s_t is the one-period change in the log spot exchange rate, and i_t- i_t* is the corresponding foreign interest differential compared to that of the U.S. π and π* are the rates of inflation in the home and foreign countries. $r - r^*$ is the real-interest differential. ρ_t, is the risk premia, ε_{st}, is the exchange rate forecast error, ε_{pt}, is the inflation forecast error, *and* θ_t: is the expected changes in the real exchange rate .

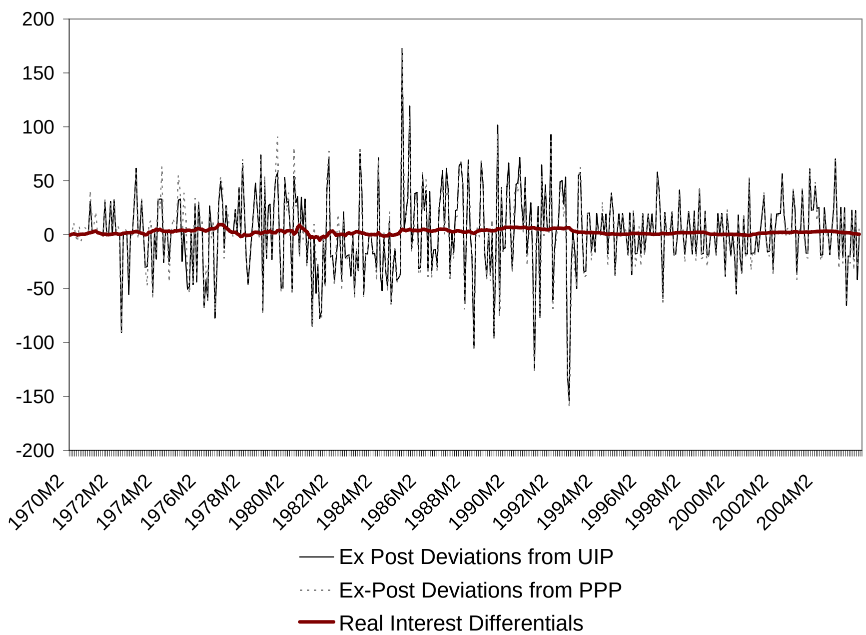


Figure 4: Estimated factors from the dynamic factor model

The table plots the parameter estimates from the dynamic factor model consisting of equations:

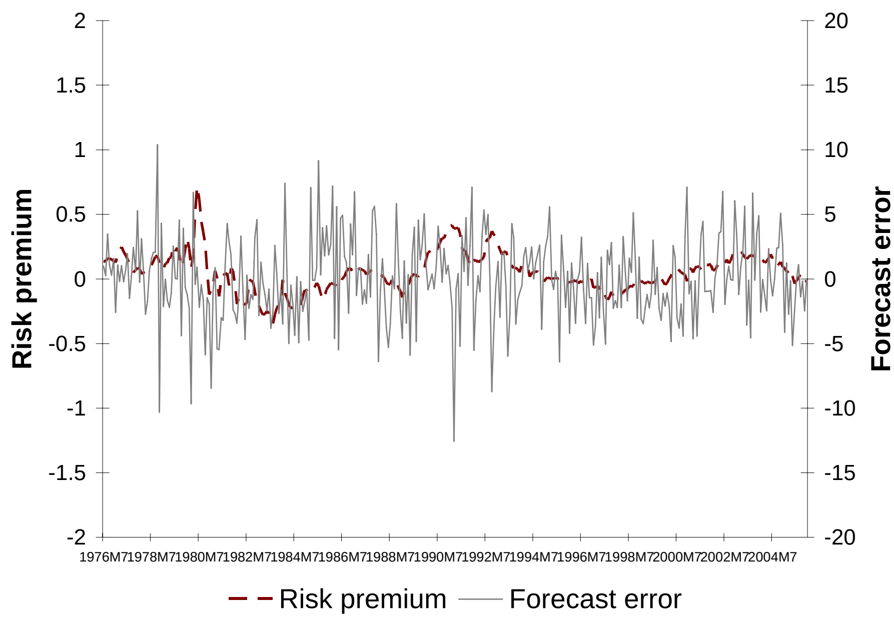
$$\begin{pmatrix} i_t - i_t * - (s_{t+1} - s_t) \\ r_t - r_t * \end{pmatrix} = \begin{pmatrix} c_{UIP} \\ 0 \end{pmatrix} + \begin{pmatrix} 1 \\ 1 \end{pmatrix} \rho_t + \begin{pmatrix} \nu_t^{UIP} \\ \nu_t^{RIE} \end{pmatrix} \quad \text{(T.2)}$$

Where the risk premium, ρ_t, is modeled as:

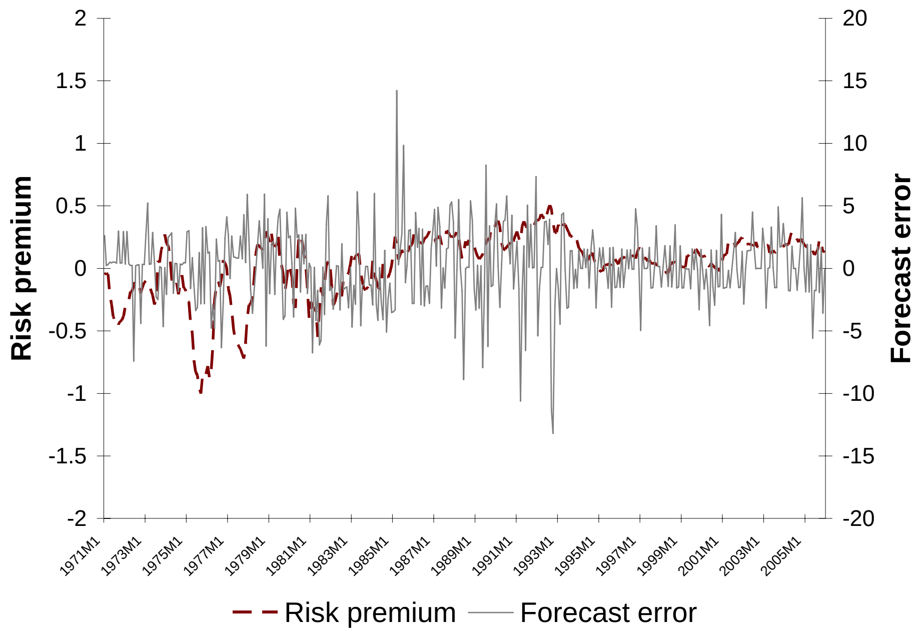
$$\rho_t = c_\rho + \phi_\rho \rho_{t-1} + \eta_{\rho t}\,,\ \eta_{\rho t} \sim N(0, \sigma_\rho^2) \quad \text{(T.3)}$$

The currencies we use are the euro (EUR), British pound sterling (GBP), and Japanese yen (JPY), all compared to the U.S. dollar. The estimation period is from January 1976 to December 2005.

Euro

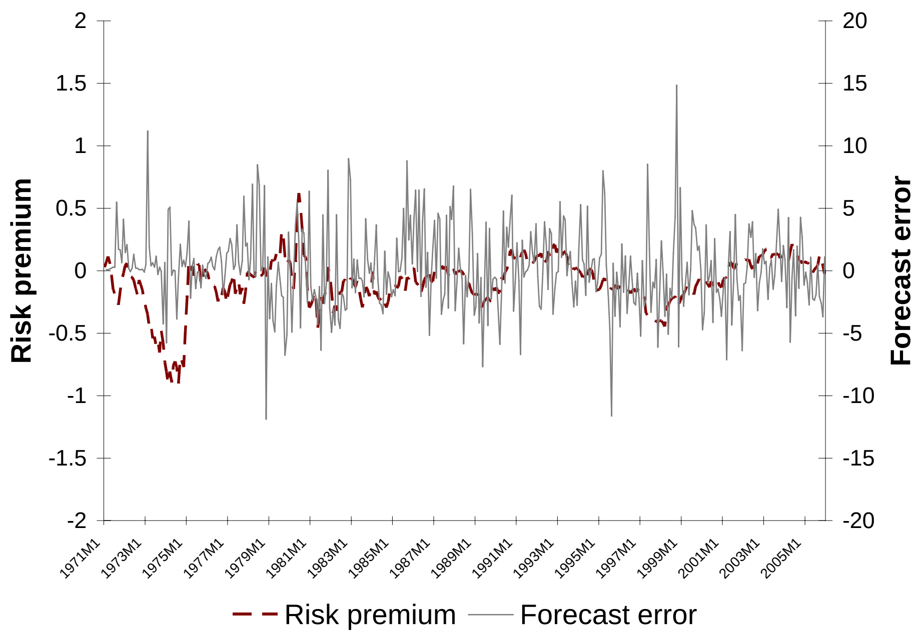


GBP:



Risk premium
Forecast error
2
1.5
1
0.5
0
-0.5
-1
-1.5
-2
20
15
10
5
0
-5
-10
-15
-20
1971M1
1973M1
1975M1
1977M1
1979M1
1981M1
1983M1
1985M1
1987M1
1989M1
1991M1
1993M1
1995M1
1997M1
1999M1
2001M1
2003M1
2005M1
Risk premium
Forecast error

JPY:



Risk premium
Forecast error
2
1.5
1
0.5
0
-0.5
-1
-1.5
-2
20
15
10
5
0
-5
-10
-15
-20
1971M1
1973M1
1975M1
1977M1
1979M1
1981M1
1983M1
1985M1
1987M1
1989M1
1991M1
1993M1
1995M1
1997M1
1999M1
2001M1
2003M1
2005M1
Risk premium
Forecast error