Testing Uncovered Interest Parity at Short and Long Horizons *

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Abstract

The hypothesis that interest rate differentials are unbiased predictors of future exchange rate movements has been almost universally rejected in empirical studies. In contrast to previous studies, which have used short-horizon data, we test this hypothesis using interest rates on longer-maturity bonds for the G-7 countries. The results of these long-horizon regressions are much more positive — the coefficients on interest differentials are of the correct sign, and almost all are closer to the predicted value of unity than to zero. These results are robust changes in data type and to base currency (i.e., Deutschemark versus US dollar). We appeal to an econometric interpretation of the results, which focuses on the presence of simultaneity in a cointegration framework.

JEL Classification: F21, F31, F41

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1. INTRODUCTION

Few propositions are more widely accepted in international economics than that the "unbiasedness hypothesis" -- that interest rate differentials are unbiased predictors of future exchange rate movements -- performs poorly. Indeed, a common finding is that exchange rate move in the opposite direction to that predicted by the hypothesis. In a survey of 75 published estimates, Froot and Thaler (1990) report few cases where the sign of the coefficient on interest rate differentials in exchange rate prediction equations is consistent with the unbiasedness hypothesis, and not a single case where it exceeds the theoretical value of unity. This resounding unanimity on the failure of the predictive power of interest differentials is virtually unique in the empirical literature in economics.

A notable aspect of almost all published studies, however, is that the unbiasedness hypothesis has been tested using financial instruments with relatively short maturities, generally of 12 months or less. There appear to be at least three reasons for this practice. The first is constraints on sample size, given that generalized exchange rate floating began only in the early 1970s. This was particularly problematic in the early 1980s, when the floating-rate period was shorter than the maturity of longer-dated financial instruments. The second is that appropriate longer-term, fixed-maturity interest rate data were difficult to obtain. The third is that such longer term rate data were, and remain, based on onshore assets; hence the effects of incipient and extant capital controls could not be easily accounted for.

Fortunately, the length of the floating-rate period is now much longer than when the initial studies were performed, and the availability of data on yields of comparable longer-dated instruments across countries has increased. Furthermore, the effects of formal and informal

impediments to capital flows are now much attenuated relative to the 1970s and early 1980s. Accordingly, this paper tests the unbiasedness hypothesis using instruments of considerably longer maturity than those employed in past studies. Our results for the dollar-based exchange rates of the major industrial countries differ strikingly from those obtained using shorter horizons. For instruments with constant maturities of 5 or 10 years, *all* of the coefficients on interest rate differentials in the unbiasedness regressions are of the correct sign. Furthermore, almost all of these coefficients on interest rates are closer to the predicted value of unity than to the zero coefficient implied by the random walk hypothesis. Finally, as the "quality" of the bond yield data in terms of their consistency with the requirements underlying UIP increases, the estimated parameters typically become closer to those implied by the unbiasedness hypothesis.

The paper is structured as follows. Section 2 reviews the unbiasedness hypothesis, summarizes the existing evidence over short horizons, and provides updated results from 1980 through early 2000. Section 3 presents estimates of the unbiasedness hypothesis using data on government bond yields for the G-7 countries. Section 4 provides an econometric rationalization for the results that are obtained. Section 5 examines the question of whether the results are specific to the US dollar. Section 6 provides concluding remarks.

2. A Review of the UIP Hypothesis and Short-Horizon Evidence

It is convenient to introduce notation and concepts by starting with the covered interest parity (CIP) condition, which follows from the assumption of arbitrage between spot and forward foreign exchange markets. If the conditions for risk-free arbitrage exist, the ratio of the forward

to the spot exchange rate will equal the interest differential between assets with otherwise similar characteristics measured in local currencies.[1] Algebraically, CIP can be expressed as:

$$F_{t,t+k} \ / \ S_t \ = \ I_{t,k} \ / \ I^*_{t,k} \ , \qquad \textbf{(1)}$$

where S_t is the price of foreign currency in units of domestic currency at time t, $F_{t,t+k}$ is the forward value of S for a contract expiring k periods in the future, $I_{t,k}$ is one plus the k-period yield on the domestic instrument, and $I^*_{t,k}$ is the corresponding yield on the foreign instrument. Taking logarithms of both sides (indicated by lower-case letters), equation (1) becomes:

$$f_{t,t+k} - s_t \ = \ (i_{t,k} - i^*_{t,k}) \ . \qquad \textbf{(2)}$$

Equation (2) is a risk-free arbitrage condition that holds regardless of investor preferences. To the extent that investors are risk averse, however, the forward rate can differ from the expected future spot rate by a premium that compensates for the perceived riskiness of holding domestic versus foreign assets. We define the risk premium, η, accordingly:

$$f_{t,t+k} \ = \ s^e_{t,t+k} \ + \ \eta_{t,t+k} \ . \qquad \textbf{(3)}$$

Substituting equation (3) into (2) then allows the expected change in the exchange rate from period t to period $t{+}k$ be expressed as a function of the interest differential and the risk premium:

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[1] These conditions include identical default risk and tax treatment, the absence of restrictions on foreign ownership, and negligible transactions costs.

$$\Delta s^{e}_{t,t+k} = (i_{t,k} - i^{*}_{t,k}) - \eta_{t,t+k} , \tag{4}$$

Narrowly defined, UIP refers to the proposition embodied in equation (4) when the risk premium is zero; this outcome would be consistent, for instance, with the assumption of risk-neutral investors.[2] In this case, the expected exchange rate change equals the current interest differential. Equation (4) is not directly testable, however, in the absence of observations on market expectations of future exchange rate movements.[3] To operationalize the concept, UIP is generally tested jointly with the assumption of rational expectations in exchange markets. In this case, future realizations of s_{t+k} will equal the value expected at time t plus a white-noise error term $\xi_{t,t+k}$ that is uncorrelated with all information known at t, including the interest differential and the spot exchange rate:

$$s_{t+k} = s^{re}_{t,t+k} + \xi_{t,t+k} , \tag{5}$$

where $s^{re}_{t,t+k}$ is the rational expectation of the exchange rate at time $t+k$ formed in time t. Substituting equation (5) into (4) yields the following relationship:

$$\Delta s_{t,t+k} = (i_{t,k} - i^{*}_{t,k}) - \eta_{t,t+k} + \xi_{t,t+k} , \tag{6}$$

where the left-hand side of equation (6) is the realized change in the exchange rate from t to $t+k$.

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[2] Note that some approximations and simplifying assumptions have been made in order to arrive at this expression. See Engel (1996).

[3] Indirect tests of UIP have been performed using surveys of published forecasts of exchange rates. Chinn and Frankel (1994, forthcoming) find mostly positive correlations between the forward discount and the expected depreciation, which is consistent with UIP.

Under the "unbiasedness" hypothesis, the last two terms in equation (6) are assumed to be orthogonal to the interest differential. Thus, in a regression context, the estimated parameter on the interest differential will have a probability limit of unity in the following regression:

$$\Delta s_{t,t+k} = \alpha + \beta\,(i_{t,k} - i^*_{t,k}) + \varepsilon_{t,t+k}\,. \tag{7}$$

A sufficient condition for this to be observed is that there be no risk premium in equation (4) (i.e. that UIP hold), and that expectations be rational -- jointly referred to as the risk-neutral efficient-markets hypothesis (RNEMH). In this case, the disturbance in equation (7) becomes simply the rational expectations forecast error $\xi_{t,t+k}$, which by definition is orthogonal to all information known at time *t*, including the interest differential. RNEMH is not necessary, however, for the unbiasedness hypothesis to hold. All that is required is that any risk premium and/or non-rational expectations error be uncorrelated with the interest differential. RNEMH, however, does imply the somewhat stronger restriction that no other regressors known at time *t* should have explanatory power, as the disturbance in equation (7) will be white noise.

Regarding the constant term, non-zero values may be explained by Jensen's inequality, which implies that the expectation of a ratio is not the same as the ratio of the expectations (although this term is likely to be small in practice). Alternatively, relaxing the assumption of risk-neutral investors, the constant term may reflect a constant risk premium demanded by investors on foreign versus domestic assets. Default risk could play a similar role, although the latter possibility is less familiar because tests of UIP (as well as CIP) generally use returns on assets issued in offshore markets by borrowers with comparable credit ratings. In contrast, the

long-term government bonds used for estimation in Section 3 may not share the same default attributes, so that a pure default risk premium might exist.

As noted above, estimates of equation (7) using values for k that range up to one year resoundingly reject the unbiasedness restriction on the slope parameter. The survey by Froot and Thaler (1990), for instance, finds an average estimate for β of -0.88. Similar results are cited in surveys by MacDonald and Taylor (1992) and Isard (1995), among others.

To update this characterization of the dismal performance of short-horizon interest rates as predictors for movements in the exchange rates, Table 1 presents estimates of equation (7) for the period 1980Q1 to 2000Q1. The exchange rates of the other six countries were expressed in terms of U.S. dollars, and the 3-, 6-, and 12-month movements in exchange rates were regressed against differentials in eurocurrency yields of the corresponding maturity.[4] Estimation using the 6- and 12-month horizon data at a quarterly frequency led to overlapping observations, inducing (under the rational expectations null hypothesis) moving average (MA) terms in the residuals. Following Hansen and Hodrick (1980), we used the Generalized Method of Moments (GMM) estimator of Hansen (1992) to correct the standard errors of the parameter estimates for moving average serial correlation of order k-1 (i.e., MA(1) in the case of 6-month data and MA(3) in the case of 12-month data).[5]

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[4] Yields and exchange rates were both constructed as the average of bid and offer rates on the last trading day of each quarter. Exchange rate movements and interest differentials are expressed at annual rates.

[5] Under the null, the a rectangular window should be used. A Bartlett window is used instead, to guarantee positive semi-definiteness of the variance-covariance matrix.

The results confirm the failure of UIP over short horizons, similar to other studies. At each horizon, four of the six estimated coefficients have the "wrong" sign relative to the unbiasedness hypothesis. The average coefficient is around -0.8, similar to the value in the survey by Froot and Thaler (1990). Panel estimation with slope coefficients constrained to be identical across countries yields estimates ranging from about -0.6 at the 6-month horizon to -0.4 at the 12-month horizon.[6] In most cases it is possible to reject the hypothesis that β equals unity; in cases where UIP cannot be rejected, the standard errors of the estimated parameters are quite large.[7] All of the adjusted R^2 statistics (not reported) are very low, and occasionally negative.

3. Long-horizon Estimates

As noted in the introduction, short-horizon tests of the unbiasedness hypothesis have been facilitated by the availability of interest rate series that correspond closely to the requirements for CIP. Data of comparable quality for longer-horizon instruments generally are much less readily available. In particular, it is difficult to obtain longer-term rates in offshore markets on thickly-traded instruments of a known fixed maturity. For the purposes of this study, then, we have used data that are inherently somewhat less pure from the point of view of the UIP hypothesis. Specifically, these on-shore instruments may be subject to differences in tax regime, capital controls, etc., such that CIP might be violated. Nonetheless, based on the findings by Popper

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[6] These are fixed effects regressions which allow for a different constant across currencies. The standard errors are constructed to allow for cross-currency correlations, as well as serial correlation due to overlapping horizons. See Frankel and Froot (1987) for details.

[7] Except for the 3 month horizon regressions, one cannot formally test the null of a zero coefficient since the standard errors are constructed under the null hypothesis that $\beta=1$.

(1993) that covered interest differentials at long maturities are not appreciably greater than those for short (up to one year) maturities, we do not expect that rejections of long-horizon UIP will be driven by deviations from CIP. Another problem is that some of our interest rate series are for debt instruments with maturities that only approximate the posited horizons, and are not the zero-coupon yields that would be exactly consistent with equation (1).

Even if these data tend to exhibit more "noise" than those used for short-horizon tests of UIP, for conventional errors-in-variables reasons we would expect the coefficient on the interest differential in these long-horizon regressions to be biased *toward* zero, and away from its hypothesized value of unity. Hence, the results we obtain should be conservative in nature.

The first data set we employ to test long-horizon unbiasedness consists of updated data on the benchmark government bond yields used by Edison and Pauls (1993). These are end-of-month yields on outstanding government bonds for the G-7 countries of 10-year maturity at the date of issuance. The 10-year change in the exchange rate versus the dollar for the other six currencies is then regressed on the 10-year lagged differential in the associated bond yield.[8] Given that generalized floating began in 1973, after allowing for the 10-year lag on the interest differential, the available estimation period consisted of 1983Q1–2000Q1 (given limitations on the availability of bond yield data for Italy, the sample period for the lira begins in1985Q1).

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[8] The serial correlation problem becomes a potentially serious issue as the number of overlapping observations increases rapidly with the instrument maturity. One way to overcome the problem is to use only non-overlapping data; however, this procedure amounts to throwing away information. Boudoukh and Richardson (1994) argue that, depending upon the degree of serial correlation of the regressor and the extent of the overlap, using overlapping data is equivalent to using between 3 to 4.5 times the number of observations available otherwise.

The results of these regressions are reported in the first panel of Table 2. They represent a surprising and stark contrast to the short-horizon results reported in Section 2. In all cases, the estimated slope coefficient is positive, with four of the six values lying closer to unity than to zero. For the Canadian dollar, the point estimate (1.100) is very close to unity, while the Deutschemark and the franc also evidence high coefficients. The yen, pound and lira are the three cases in which UIP is statistically rejected. The adjusted R^2 statistics are also typically higher than in a typical short-horizon regressions, with the proportion of the explained variance in the Deutschemark and the pound approaching one half.

Since there are relatively few independent observations in the single-currency regressions, additional power can be obtained by pooling the data and constraining the slope coefficient to be the same across currencies. The resulting point estimate is reported under the entry "constrained panel" at the bottom of Table 2.a. Its value of 0.592 is well below unity; on the other hand, it is closer to unity than to zero, a substantial difference from the panel estimates obtained for short horizons reported in Table 1.

For Japan, Germany, the U.K., and the U.S., it was also possible to obtain synthetic "constant maturity" 10-year yields from interpolations of the yield curve of outstanding government securities. The regressions using measures of long-horizon interest differentials based on these data are reported in Table 2.b. The estimated slope parameters are as close -- or closer -- to unity than in the corresponding regressions using benchmark yields. Moreover, the panel point estimate of 0.726 is substantially closer to the posited value. The improvement in the results, although modest, suggests that part of the reason why unbiasedness is still rejected when using benchmark yields relates to discrepancies between the assumed and actual maturities of the

outstanding securities. In other words, improvements in the quality of the data appear to systematically shift the results toward supporting the UIP hypothesis.[9] Figure 1 illustrates the difference in the two interest rate differentials, and compares these series to the exchange rate depreciation.

Similar constant-maturity 5-year yields were obtained for Germany, the U.K., Canada, and the U.S. Results of regressions of 5-year changes in exchange rates on the corresponding interest differentials are reported in Table 2.c. The results are equally favorable to the UIP hypothesis: for all three of these currencies, the slope coefficients are statistically indistinguishable from the implied value of unity, as is the panel estimate.

The only other studies that we are aware of that test the unbiasedness hypothesis over horizons of longer than 12 months are Flood and Taylor (1997), and Alexius (1999). Flood and Taylor regress 3-year changes in exchange rates on annual average data on medium-term government bonds from the IMF's *International Financial Statistics* (*IFS*). The data over the 1973–92 period are then pooled for a sample of 21 countries. They find a coefficient on the interest differential of 0.596 with a standard error of 0.195. Thus the hypotheses that β equals either zero or unity can both be rejected. These results are broadly in line with our 10-year results, although our 5-year results using constant maturity data are more supportive of the unbiasedness hypothesis than theirs. This difference may reflect the fact that our end-period,

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[9] A more appropriate data set would include zero coupon constant maturity interest rate series. Unfortunately these data are not readily available on a cross country basis. Alexius (1999) applies a correction to account for the absence of zero coupon yields, and obtains improved results relative to those based on unadjusted data. Presumably using adjusted data in our context would have a similar effect.

constant-maturity data better fulfill the requirements underlying the UIP hypothesis, although differences in country coverage and sample periods may also play a role.

In the study by Alexius, 14 long term bond rates (of uncertain maturities) for the 1957Q1-1997Q4 period are drawn from *International Financial Statistics* (*IFS*). She attempts to control for the measurement error arising from uncertain maturities, and the role of coupon payments.[10] Her study also finds substantial evidence in favor of the unbiasedness hypothesis at long horizons. For the Deutschemark, the OLS point estimate for the duration- and coupon-adjusted series is 0.820, which is remarkably close to our estimate of 0.851 for the 10 year constant maturity yields. On the other hand, her estimates for the yen and the pound (0.209 and 0.278, respectively) are somewhat lower than the estimates we report in Table 2.b of 0.418 and 0.713. Some of this difference may be due to the longer sample she uses, which encompasses a period of substantial capital controls.

In any event, it is reassuring that despite data and methodological differences, these results are similar to those obtained in our regressions, suggesting that the difference between short- and long-horizon tests of UIP may be robust across countries, sample periods and estimation procedures.

4. Explaining the Results Econometrically

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[10] The *IFS* data are somewhat problematic in that the definitions of the long term bonds is not homogeneous across countries and over time. Moreover, her data sample spans periods of both fixed and flexible rate regimes, as well as an era when capital controls were pervasive.

The rather strikingly different results obtained at different horizons should be placed in the context of recent findings that, when the unbiasedness proposition is couched in terms of cointegrating relationships, one finds that it is much more difficult to reject the null hypothesis of unbiasedness (e.g., Evans and Lewis, 1995). Here, we are not so much concerned with the specific finding regarding cointegration with the posited values, but rather the econometric implications of estimating equation (7). If the expected spot and forward rate are cointegrated, then it must be true that the current spot and forward rate are also cointegrated. It turns out that it is more convenient to work with this representation (Zivot, 2000). According to the Engle-Granger Representation Theorem, one can write this latter cointegrated system as:

$$\begin{aligned} \Delta s_t &= \gamma_{10} + \Phi_1[s_{t-1} - \delta_1 f_{t-1} - \delta_0] + \sum_{i=1}^{j} \gamma_{1i}\Delta s_{t-i-1} + \sum_{i=1}^{j} \zeta_{1i}\Delta f_{t-i-1} + \varepsilon_{1t} \\ \Delta f_t &= \gamma_{20} + \Phi_2[s_{t-1} - \delta_1 f_{t-1} - \delta_0] + \sum_{i=1}^{j} \gamma_{2i}\Delta s_{t-i-1} + \sum_{i=1}^{j} \zeta_{2i}\Delta f_{t-i-1} + \varepsilon_{2t} \end{aligned} \quad \textbf{(8)}$$

where the horizon has been set to one ($k = 1$) for simplicity of exposition. As pointed out by Phillips (1991), single-equation estimation of (8.a) is plagued by asymptotic bias as long as the forward rate is not weakly exogenous. This assertion can be verified by enumerating the steps necessary to convert equation (8) to (7). First, one must assume weak exogeneity of f (implying that $\Phi_2 = 0$, so that we can ignore the second equation). Subsuming the constant into the cointegrating vector, one obtains

$$\begin{aligned} \Delta s_t &= b_0 \Delta f_t + \Phi_1[s_{t-1} - \delta_1 f_{t-1} - \delta_0] \\ &\quad + \sum_{i=1}^{j-1} b_i \Delta s_{t-i} + \sum_{i=1}^{j-1} c_i \Delta f_{t-i} + u_t \end{aligned} \quad \textbf{(9)}$$

where b_i and c_i are functions of the variances and covariances of ε_1 and ε_2, and u is a function of ε_1 and ε_2, and their variances and covariances. In particular, $b_0 = \sigma_{12}/\sigma_{22}$, which equals zero only when the correlation between the ε's is zero. Imposing the restrictions $\delta_0 = 0$ and $\delta_1 = 1$,[11] equation (9) can be rewritten as:

$$\begin{aligned} \Delta s_t = & - \Phi_1[f_{t-1} - s_{t-1}] + b_0 \Delta f_t \\ & + \sum_{i=1}^{j-1} b_i \Delta s_{t-i} + \sum_{i=1}^{j-1} c_i \Delta f_{t-i} + u_t \end{aligned} \tag{10}$$

Notice that equation (10) degenerates to equation (7) if and only if $b_0 = 0$, $b_i = c_i = 0$ for all i, *as well as* $\delta_0 = 0$, $\delta_1 = 1$ (Moore, 1994; Villanueva, 1999).

To examine whether the standard assumption of weak exogeneity of the forward rate is justified at either the short or long horizons, we generate implicit forward rates using the exact relationship in equation (1), for both the 3 month and 5 year horizons. We then test for cointegration between the forward rate and the future spot rate[12] using the Johansen (1988) maximum likelihood procedure. The results are reported in Table 3; in Panel 3.a are the cointegration results for the 3 month forward rates and the future spot rates, and in Panel 3.b are the corresponding results for the 5 year implicit forward rates.

The first column displays the likelihood ratio for the Maximal Eigenvalue statistic. The 5% critical value for rejecting the null hypothesis of no cointegrating vectors in favor of the

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[11] See Brenner and Kroner (1995) complications involved in imposing the $\delta_0 = 0$ restriction in the cointegrating vector.

[12] In principle, either specification is valid asymptotically. Zivot (2000) argues for testing the cointegrating vector involving the contemporaneous forward and spot rate, while Villanueva (1999) reports results demonstrating that lagged forecast errors yield more unambiguous results.

alternative of one is 15.41. All the currencies evidence cointegration. Furthermore, if the long run unbiasedness hypothesis is imposed, then in all cases save one, the forward -- and not spot -- rate responds to the disequilibrium. The sole exception is the yen, in which case the spot rate responds as well (although in a perverse fashion).

For the 5 year implicit forwards and the corresponding future spot rates, cointegration is detected for the pound, while less evidence of cointegration is detected for the Deutschemark and Canadian dollar. For the pound one obtains the result that at horizons of 5 years, the spot rate responds to the lagged cointegrating vector Φ_1 with high statistical significance, while the forward rate does not. That is, long term interest rate differentials are weakly exogenous in this system.

Unfortunately, the cointegration evidence for the other two currencies is weaker. If one uses the more powerful Horvath-Watson (1995) test imposing the unbiasedness hypothesis, one finds that test statistic for the Deutschemark of 3.50 is somewhat less than the 10% critical value of 4.73 for the case with a zero mean in the variables (although it is much less than the corresponding critical value of 8.30 for the possibly more relevant nonzero-mean case). If one were willing to impose the prior of cointegration (see Kremers, Ericsson and Dolado, 1992), then the t-statistic on Φ_1 is 1.589, while that on Φ_2 is only 0.987. The data thus seem to suggest that the 5 year Deutschemark forward rate -- corresponding to the interest differential -- is less endogenous than the spot rate.

For the Canadian dollar, slight evidence of cointegration can be detected. The results in the Panel 3.b are for the Horvath-Watson regressions where the null of long run unbiasedness is imposed. The forward rate appears to be more responsive to the forward forecast error than the

spot rate, contradicting the argument posed above. However, these latter results are merely suggestive because we are not able to detect cointegration in the sample we have.[13]

For two of the three currencies for which we have data, it appears that the forward rate is weakly exogenous at long horizons, while at short horizons the spot rate is more likely to be weakly exogenous. From a purely statistical standpoint, this explains some of the differences in the results obtained at short and long horizons.

5. Extending the Results

The results we have reported up to this point have been based on data using the US dollar as the reference currency. However, it may be the case that the dollar is an exceptional currency, in terms of its adherence, or lack thereof, to the uncovered interest parity relationship. In order to investigate this question, depreciations and interest differentials are re-expressed against the Deutschemark, and the regressions described in Sections 2 and 3 are estimated. The results are reported in Tables 4 and 5.

In Table 4, for all the currencies save the franc and lira, the coefficient estimates on the short term interest differential are negative. The results at the long horizons are mixed. Using the benchmark bond yields, one finds that the yen and pound regressions exhibit negative coefficients, while the lira estimate is statistically significantly different from unity. As mentioned earlier, these benchmark bond yields are imprecise measures. Fortunately, for both of

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[13] An interesting aspect of the Canadian dollar is the large Canada-US interest differential which appeared in 1990 with the collapse of the Meech Lake accords, and disappears in 1997 (see Clinton, 1998).

these currencies, constant maturity rates are available. Using this data, one finds the UK coefficient is now positive, while the yen remains negative. It appears to be the case that accounting for the maturity mismeasurement is important, although not sufficient to overturn the rejection of UIP. At the five year horizon, the results are less ambiguous; both coefficients are positive, and not statistically different from the value of unity.

Why might these results based upon the Deutschemark be less clear about the relevance of UIP? Once one controls for the horizon length, the sole exception to the finding of a positive coefficient is the yen-Deutschemark rate. Japan and Germany are two countries which implemented decontrol of the government bond market later than did the US. It is suggestive that if one restricts the sample to 1990Q1-2000Q1 (corresponding to post-1980 interest rate data), the estimated coefficients rise in value.[14]

6. Conclusions

We find strong evidence for the G-7 countries that the perverse relationship between interest rates and exchange rates is a feature of the short-horizon data that have been used in almost all previous studies. Using longer horizon data, the results of standard test of UIP yield strikingly different results, with slope parameters that are positive and closer to the hypothesized value of unity than to zero. These results confirm the earlier conjectures of Mussa (1979) and Froot and Thaler (1990) that the unbiasedness proposition may better apply at longer horizons.

[14] Frankel (1984) argues that the Japanese market in short term instruments was decontrolled only in 1980. If the sample is truncated, then the benchmark rate coefficient rises from -1.023 to -0.658; the constant maturity rate coefficient rises from -0.851 to -0.561.

These findings are generally replicated for exchange rates and interest differentials based upon the Deutschemark, rather than the more commonly used dollar numeraire. They also extend to other sample periods, and other measures of interest rates. Hence, one can be reasonably certain that our findings are not a statistical fluke.

From an econometric perspective, the differential results can be explained in the context of endogeneity of the right hand side variable. Deciding what type of economic model induces such an endogeneity is a more contentious issue. In a related paper (Meredith and Chinn, 1998), we suggest the difference in the results is consistent with the properties of a conventional macroeconomic model. In particular, a temporary disturbance to the uncovered interest parity relationship causes the spot exchange rate to depreciate relative to the expected future rate, leading to higher output, inflation, and interest rates. Higher interest rates are then typically associated with an ex post future appreciation of the exchange rate at short horizons, consistent with the forward discount bias typically found in empirical studies. Over longer horizons, the temporary effects of exchange market shocks fade and the model results are dominated by more fundamental dynamics that are consistent with the UIP hypothesis.

An alternative explanation for these results has been forwarded by Ogaki (1999), who relies upon exogenously determined segmentation between short and long term bond markets.

Regardless of the reasons for the failure of the unbiasedness hypothesis at short horizons, from an unconditional forecasting perspective, the conclusion remains that interest differentials are essentially useless as predictors of short-term movements in exchange rates. Over longer horizons, however, our results suggest that interest differentials may significantly outperform

naive alternatives such as the random-walk hypothesis, although they are still likely to explain only a relatively small proportion of the observed variance in exchange rates.

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Table 1. Estimates of β

$$\Delta s_{t,t+k} = \alpha + \beta (i_{t,k} - i^*_{t,k}) + \varepsilon_{t+k} . \qquad (7)$$

	Maturity		
Currency	3 mo.	6 mo.	12 mo.
Deutschemark [1]	-0.740*	-0.834**	-0.524***
	(1.134)	(0.811)	(0.669)
Japanese yen	-2.740***	-2.820***	-2.665***
	(1.039)	(0.837)	(0.720)
U.K. pound	-2.166***	-1.979***	-1.367**
	(1.111)	(1.056)	(1.007)
French franc [1]	0.208	0.160	0.244
	(0.954)	(0.822)	(0.772)
Italian lira	0.991	1.006	1.127
	(0.697)	(0.701)	(0.645)
Canadian dollar	-0.423***	-0.497***	-0.599**
	(0.525)	(0.395)	(0.501)
Panel[2]	-0.529***	-0.592***	-0.415***
	(0.403)	(0.366)	(0.400)

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample is 1980Q1-2000Q1. * (**)[***] Different from null of unity at 10%(5%)[1%] marginal significance level.

1 Sample period: 1980Q1-1999Q3.

2 Fixed effects regression. Sample period: 1980Q1-1999Q4.

Table 2. Long-Horizon Tests of Uncovered Interest Parity

$$\Delta s_{t,t+k} = \alpha + \beta\,(i_{t,k} - i^*_{t,k}) + \varepsilon_{t+k}\,. \qquad (7)$$

Panel 2.a: Benchmark Government Bond Yields, 10-Year Maturity
(MA(39)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	$\bar{R}^2$
Deutschemark	0.006 (0.003)	0.851 (0.180)		0.40
Japanese yen	0.038 (0.005)	0.388 (0.144)	***	0.10
U.K. pound	-0.003 (0.004)	0.562 (0.106)	***	0.43
French franc	0.006 (0.012)	0.837 (0.439)		0.04
Italian lira[1]	-0.010 (0.006)	0.212 (0.149)	***	0.01
Canadian dollar	-0.001 (0.003)	1.100 (0.486)		0.16
Constrained panel[2]	...	0.592 (0.134)	***	0.52

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1983Q1-2000Q1. * (**)[***] Different from null of unity at 10%(5%)[1%] marginal significance level.

1 Sample period: 1987Q1–2000Q1.

2 Fixed effects regression, excluding the lira. Sample period: 1983Q1 - 1999Q4.

Table 2. continued

Panel 2.b: 10-Year Government Bond Yields
(MA(39)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	$\bar{R}^2$
Deutschemark	0.006 (0.003)	0.851 (0.165)		0.42
Japanese yen	0.037 (0.006)	0.418 (0.168)	***	0.08
U.K. pound	0.003 (0.003)	0.713 (0.104)	***	0.44
Constrained panel [1]	...	0.726 (0.068)	***	0.69

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1983Q1-2000Q1. * (**)[***] Different from null of unity at 10%(5%)[1%] marginal significance level.

[1] Pooled regression, with fixed effect for the yen. Sample period: 1983Q1-1999Q4.

Table 2. (continued)

Panel 2.c: 5-Year Government Bond Yields
(MA(19)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	$\bar{R}^2$
Deutschemark	0.003 (0.014)	0.759 (0.581)		0.06
U.K. pound	0.000 (0.016)	0.679 (0.321)		0.06
Canadian dollar	-0.001 (0.010)	0.742 (0.474)		0.06
Constrained panel[1]	...	0.715 (0.402)		0.11

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1983Q1-2000Q1. * (**)[***] Different from null hypothesis at 10%(5%)[1%] marginal significance level.

1 Fixed effects regression. Sample period:1983Q1 - 1999Q4.

Table 3. Johansen Cointegration Test Results

$$\begin{aligned}\Delta s_t &= \gamma_{10} + \Phi_1[s_{t-1} - f_{t-2}] + \sum_{i=1}^{j}\gamma_{1i}\Delta s_{t-i-1} + \sum_{i=1}^{j}\zeta_{1i}\Delta f_{t-i-2} + \varepsilon_{1t} \\ \Delta f_{t-1} &= \gamma_{20} + \Phi_2[s_{t-1} - f_{t-2}] + \sum_{i=1}^{j}\gamma_{2i}\Delta s_{t-i-1} + \sum_{i=1}^{j}\zeta_{2i}\Delta f_{t-i-2} + \varepsilon_{2t}\end{aligned} \tag{7}$$

Panel 3.a: 3 Month Horizon

	LR	Φ_1	Φ_2	j	N
Deutschemark [1]	13.66*	1.145 (0.957) [1.196]	0.140** (0.056) [2.488]	2	80
Japanese yen	18.91**	3.644*** (1.233) [2.957]	0.163** (0.067) [2.453]	2	82
U.K. pound	20.79***	1.368 (1.154) [1.186]	0.177** (0.078) [2.269]	3	82
French franc [1]	21.47***	0.039 (0.871) [0.044]	0.340*** (0.028) [3.957]	2	80
Italian lira [2]	28.84***	-0.723 (0.752) [0.962]	0.371*** (0.093) [3.992]	2	81
Canadian dollar	13.62*	0.348 (0.583) [0.597]	0.284*** (0.079) [3.607]	2	82

Notes: LR is the likelihood ratio for the Maximal Eigenvalue test of the H_0 of zero cointegrating vectors against H_A of one cointegrating vector. 15.41 and 20.04 are the 5% and 10% critical values (Osterwald and Lenum, 1992). Point estimates from OLS regression OLS regression (standard errors in parentheses) [absolute values of the *t*-statistics in brackets]. *j* is the number of lags in the VAR representation of the cointegrated system. *N* is the number of observations. Sample period: 1980Q1-2000Q1. * (**)[***] Different from null hypothesis at 10%(5%)[1%] marginal significance level.

1 Sample period: 1980Q1-1999Q3.

2 Sample period: 1980Q1-1999Q4.

Panel 3.b: 5 Year Horizon					
	LR	Φ_1	Φ_2	j	N
Deutschemark	4.87	-0.041† (0.026) [1.589]	0.026 (0.026) [0.987]	3	82
U.K. pound	16.76**	-0.052* (0.027) [1.939]	0.060 (0.040) [1.484]	2	82
Canadian dollar	6.44	-0.015 (0.022) [0.691]	0.073** (0.036) [2.034]	2	82

Notes: LR is the likelihood ratio for the Maximal Eigenvalue test of the H_0 of zero cointegrating vectors against H_A of one cointegrating vector. 15.41 and 20.04 are the 5% and 1% critical values (Osterwald and Lenum, 1992). Point estimates from OLS regression (standard errors in parentheses) [absolute values of the *t*-statistics in brackets]. *j* is the number of lags in the VAR representation of the cointegrated system. *N* is the number of observations. Sample period: 1980Q1-2000Q1.
{†}* (**)[***] Different from null hypothesis at {20}10%(5%)[1%] marginal significance level.

[1] Sample period: 1980Q1-2000Q1.

Table 4. Estimates of β
Using the Deutschemark as the Base Currency

$$\Delta s_{t,t+k} = \alpha + \beta\,(i_{t,k} - i^{*}_{t,k}) + \varepsilon_{t+k}\,. \qquad (7)$$

	Maturity		
Currency	3 mo.	6 mo.	12 mo.
Japanese yen	-1.187*	-1.307**	-0.589
	(1.338)	(1.250)	(1.400)
U.K. pound	-1.068***	-0.759***	-0.376*
	(0.901)	(0.812)	(0.857)
French franc	0.951	0.799	0.678
	(0.264)	(0.185)	(0.239)
Italian lira	0.237***	0.232***	0.125***
	(0.278)	(0.231)	(0.265)
Canadian dollar	-0.889	-0.782**	-0.593**
	(1.257)	(0.911)	(0.789)

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1980Q1-1999Q3. * (**)[***] Different from null of unity at 10%(5%)[1%] marginal significance level.

Table 5: Long-Horizon Tests of Uncovered Interest Parity
Using the Deutschemark as the Base Currency

$$\Delta s_{t,t+k} = \alpha + \beta\,(i_{t,k} - i^*_{t,k}) + \varepsilon_{t+k}\,. \qquad (7)$$

Panel 5a: Benchmark Government Bond Yields, 10-Year Maturity
(MA(39)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject $H_0: \beta = 1$	$\bar{R}^2$
Japanese yen	0.032 (0.001)	-1.023 (0.093)	***	0.38
U.K. pound	-0.056 (0.005)	-1.037 (0.275)	***	0.17
French franc	0.000 (0.012)	0.446 (0.200)	***	0.17
Italian lira[1]	-0.034 (0.008)	0.112 (0.113)	***	0.03
Canadian dollar	1.010 (0.007)	1.185 (0.195)		0.22

Panel 5b: 10-Year Government Bond Yields
(MA(39)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject $H_0: \beta = 1$	$\bar{R}^2$
Japanese yen	0.032 (0.002)	-0.851 (0.133)	***	0.29
U.K. pound	-0.018 (0.020)	0.251 (0.409)	*	0.00

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1983Q1-2000Q1. * (**)[***] Different from null of unity at 10%(5%)[1%] marginal significance level.

1 Sample period: 1987Q1–2000Q1.

2 Fixed effects regression, excluding the lira. Sample period: 1983Q1 - 1999Q4.

Table 5. (continued)

Panel 5.c: 5-Year Government Bond Yields
(MA(19)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	$\bar{R}^2$
U.K. pound	0.006 (0.024)	0.963 (0.657)		0.15
Canadian dollar	-0.007 (0.015)	0.515 (0.411)		0.01

Notes: Point estimates from the regression in equation 1 (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1983Q1-2000Q1. * (**)[***] Different from null hypothesis at 10%(5%)[1%] marginal significance level.

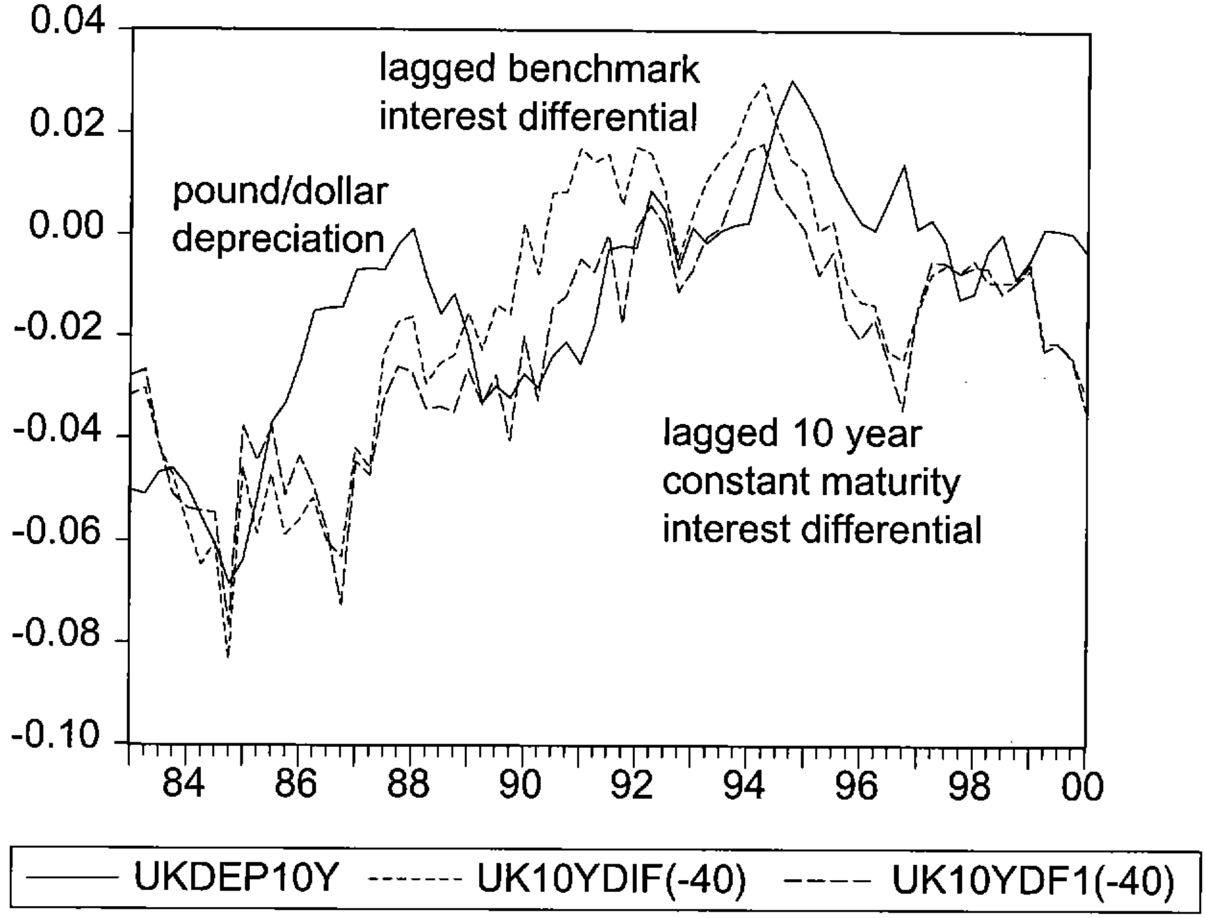

Figure 1: Ex post pound/dollar 10 year depreciation (annualized), and benchmark and constant maturity interest differentials (lagged).