Monetary Policy and Long-Horizon Uncovered Interest Parity

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Uncovered interest parity (UIP) has been almost universally rejected in studies of exchange rate movements. In contrast to previous studies, which have used short-horizon data, we test UIP using interest rates on longer-maturity bonds for the Group of Seven countries. These long-horizon regressions yield much more support for UIP—all of the coefficients on interest differentials are of the correct sign, and almost all are closer to the UIP value of unity than to zero. We then use a macroeconomic model to explain the differences between the short- and long-horizon results. Regressions run on model-generated data replicate the important regularities in the actual data, including the sharp differences between short- and long-horizon parameters. In the short run, the failure of UIP results from the inter-action of stochastic exchange market shocks with endogenous monetary policy reactions. In the long run, in contrast, exchange rate movements are driven by the "fundamental," leading to a relationship between interest rates and exchange rates that is more consistent with UIP. [JEL F21, F31, F41]

Few propositions are more widely accepted in international economics than that uncovered interest parity (UIP) is at best useless—or at worst perverse—

as a predictor of future exchange rate movements. This finding has been replicated in an extensive literature, including the initial studies by Bilson (1981), Longworth (1981), and Meese and Rogoff (1983). 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 under UIP, and not a single case where it exceeds the theoretical value of unity.[1] This resounding unanimity on the failure of the predictive power of UIP must be virtually unique in the empirical literature in economics.

A notable aspect of almost all published studies, however, is that UIP has been tested using financial instruments with relatively short maturities, generally 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 longer-term, fixed-maturity interest rate data were difficult to obtain. The third reason is that some pioneering studies were also concerned with testing the hypothesis of *covered* interest parity, which required observations on forward exchange rates of the same maturity as the associated financial asset. In any event, relatively thick forward exchange markets exist only to a maximum horizon of 12 months.

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. Accordingly, this paper tests the UIP hypothesis using instruments of considerably longer maturity than those employed in past studies. Our results for the exchange rates of the major industrial countries differ strikingly from those obtained using shorter horizons. For instruments with maturities ranging from 5 to 10 years, *all* of the coefficients on interest rate differentials in the UIP regressions are of the correct sign. Furthermore, *almost all* of the coefficients on interest rates are closer to the UIP 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.

To explain the apparently anomalous differences in tests of UIP using short- and long-horizon data, we develop a small macroeconomic model that extends the framework of McCallum (1994). In particular, we incorporate a more general monetary reaction function that causes interest rates to respond to innovations in output and inflation, as opposed to the exchange rate–targeting framework used by McCallum. Stochastic simulations of the model are used to generate a synthetic database for replicating UIP tests. Standard short-horizon regressions using

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[1] In our terminology, tests of UIP are interchangeable with tests of the unbiasedness hypothesis—that is, that the coefficient on lagged interest differentials in regressions of exchange rate changes is unity. This usage is somewhat loose, however. As discussed in Section I, the combined hypotheses of UIP and rational expectations are sufficient (but not necessary) for unbiasedness to hold. UIP, in itself, is neither necessary nor sufficient for unbiasedness.

these synthetic data yield negative coefficients on short-term interest rates of roughly the same magnitude as found in most short-horizon studies; thus the model can explain the "forward discount bias" found in such studies. Long-horizon regressions, in contrast, yield coefficients close to unity, consistent with our estimation results using actual data.[2] The long-horizon results differ sharply from the short-horizon results because the model's "fundamentals" play a more important role in tying down exchange rate movements over longer horizons. More generally, the data generated by the simulations for the endogenous variables mimic closely the key properties of actual data for the Group of Seven (G-7) countries.[3]

I. 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.[4] Algebraically, CIP can be expressed as

$$F_{t,t+k} \; / \; S_t \; = \; I_{t,k} \; / \; I_{t,k}^{*}, \tag{1}$$

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 lowercase letters), equation (1) becomes

$$f_{t,t+k} \; - \; s_t \; = \; \left(i_{t,k} \; - \; i_{t,k}^{*} \right). \tag{2}$$

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 accordingly:

$$f_{t,t+k} \; = \; s_{t,t+k}^{e} \; - \; rp_{t,t+k}. \tag{3}$$

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[2]As pointed out by the referee, there are two uses of the term "long-horizon regressions." The first involves regressions using long-maturity instruments, and the second pertains to regressions relating a change in a variable between t and $t + k$ to a right-hand side variable dated at time t (as in Chinn and Meese, 1995). In the case we are examining here, the two definitions apply.

[3]We do not mean to exclude other possible explanations for the forward premium puzzle, including the presence of noise traders (e.g., Taylor and Allen, 1992). Indeed, one could interpret the shock to the UIP relationship as arising from the actions of noise traders.

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

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

$$\Delta s_{t,t+k}^{e} = \left(i_{t,k} - i_{t,k}^{*}\right) - rp_{t,t+k}. \tag{4}$$

Narrowly defined, UIP refers to the proposition embodied in equation (4) when the risk premium is zero—consistent, for instance, with the assumption of risk-neutral investors. 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.[5] 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 $>_{t, t+k}$ that is uncorrelated with all information known at t, including the interest differential and the spot exchange rate:

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

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

$$\Delta s_{t,t+k} = \left(i_{t,k} - i_{t,k}^{*}\right) - rp_{t,t+k} + \xi_{t,t+k}, \tag{6}$$

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

The regression equation generally used to test the unbiasedness hypothesis is

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

Under the combined assumptions of risk neutrality and rational expectations, the disturbance in equation (7) reduces to the deviation from the outturn for the exchange rate from its rational expectation, that is, $\xi_{t, t+k}$ in equation (6). By definition, this is orthogonal to all information known at time t, rational expectations, the disturbance in equation (7) reduces to the deviation between the outturn for inflation and its rational expectation, i.e. $\xi_{t, t+k}$ in equation (6). By definition, this term is orthogonal to the interest differential and all other information available at time t, so the probability limit of the slope parameter in equation (7) will be one. It should be noted that these combined assumptions—referred to as the risk-neutral efficient-markets hypothesis, or RNEMH—are sufficient, but not necessary for unbiasedness. The only necessary condition is that the deviations from risk neutrality and rational expectations be uncorrelated with the interest differential. Thus,

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[5]Indirect tests of UIP have been performed using surveys of published forecasts of exchange rates, with mixed results (Chinn and Frankel, 1994 and 2002). See Bryant (1995) for a discussion.

the failure of unbiasedness must be consistent with two phenomena: (i) deviations from risk neutrality and/or rational expectations, and (ii) an economic mechanism that results in a correlation between these deviations and the interest differential. including the interest differential. Thus the regression equation is equivalent to equation (6) absent the risk premium, and the probability limit of the slope parameter will equal one, along with the condition that no other regressor known at time t will enter the regression. It should be noted that these combined assumptions—referred to as the risk-neutral efficient-markets hypothesis, or RNEMH (Clarida and Taylor, 1997)—are sufficient, but not necessary, for unbiasedness. The only necessary condition is that any deviations from risk neutrality or rational expectations be uncorrelated with the interest differential. Failure of unbiasedness, then, must reflect two phenomena: (i) deviations from risk neutrality and/or rational expectations, and (ii) economic channels that generate a correlation between the interest differential and these deviations.

Regarding the constant term, nonzero values may still be consistent with UIP. Jensen's inequality, for instance, implies that the expectation of a ratio (such as the exchange rate between two currencies) is not the same as the ratio of the expectations (see Meese, 1989).[6] 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 II may not share the same default attributes, so 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, which is similar in magnitude to the null under the UIP hypothesis, *but of the opposite sign.* In another survey of the literature, MacDonald and Taylor (1992) observe that ". . . [various researchers] all report a result suggesting a sound rejection of the unbiasedness hypothesis: a significantly negative point estimate of β" (page 31).[7] Thus, the common perception that the failure of UIP indicates that short-run exchange rate movements are best characterized as a random walk is not strictly true: over short horizons, most studies find that exchange rates move *inversely* with interest differentials.[8]

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[6]As noted in Engel (1996), however, a constant term due to Jensen's inequality is likely to be small in practice.

[7]Other surveys that report similar results include Isard (1995) and Lewis (1995). A qualified exception is the study by Flood and Rose (1996), which finds that the coefficient on the interest differential is closer to its UIP value during periods when exchange rate realignments within Europe's Exchange Rate Mechanism (ERM) were expected (and observed).

[8]The perception that exchange rates are random walks probably reflects the interpretation of studies that have tested the random walk hypothesis against specific, but limited, alternatives. The influential study by Meese and Rogoff (1983), for instance, found that the random walk outperformed covered interest rate parity, as well as structural exchange rate models, during the late 1970s and early 1980s. But they did not test the random walk against more general alternatives to UIP with an unconstrained coefficient on the interest differential.

Table 1. Short-Horizon Estimates of b

$$\Delta s_{t,t+k} = \alpha + \beta\left(i_{t,k} - i_{t,k}^*\right) + \varepsilon_{t,t+k}$$

Currency	Maturity		
	3 months	6 months	12 months
Deutsche mark	−0.809* (1.134)	−0.893*** (0.802)	−0.587*** (0.661)
Japanese yen	−2.887*** (0.997)	−2.926*** (0.800)	−2.627*** (0.700)
U.K. pound	−2.202*** (1.086)	−2.046*** (1.032)	−1.418*** (0.986)
French franc	−0.179 (0.904)	−0.154 (0.787)	−0.009 (0.773)
Italian lira	0.518 (0.606)	0.635 (0.670)	0.681 (0.684)
Canadian dollar	−0.477*** (0.513)	−0.572*** (0.390)	−0.610*** (0.490)
Constrained panel[1]	−0.757*** (0.374)	−0.761*** (0.345)	−0.536*** (0.369)

Notes: Point estimates from the regression in equation (7) (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample is 1980: Q1–2000: Q4. *, **, *** indicate different from null of unity at, respectively, the 10 percent, 5 percent, and 1 percent marginal significance level.

[1]Fixed-effects regression. Standard errors adjusted for serial correlation (see text).

To illustrate the performance of short-horizon UIP for the exchange rates of the G-7 countries, Table 1 presents estimates of equation (7) from 1980 to 2000. 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.[9] 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 Methods of Moments (GMM) estimator of Hansen (1982) to correct the standard errors of the parameter estimates for MA 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).

The results confirm the failure of UIP over short horizons, similar to other studies. At each horizon, five 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 −0.8 at the 6-month horizon to −0.5 at the 12-month horizon.[10] 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 sufficiently large that it would be difficult to reject almost any plausible hypothesis. The adjusted R^2 statistics of these regressions (not shown) are very low, and occasionally negative.

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[9]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.

[10]Since by construction there is clearly cross-equation correlation of the error terms, it might appear that seemingly unrelated regression estimation (SURE) would be appropriate. However, the right-hand side variables are not strictly exogenous, so SURE would not yield consistent estimates.

The range of slope coefficients is somewhat larger than reported in most previous studies, with estimates for the lira yielding positive estimated values for β at all of the short horizons. This result is consistent with the findings of Chinn and Frankel (1994), who estimate highly positive values of β for some of the currencies—including the lira—that depreciated in the aftermath of the 1992 ERM crisis. They interpret this as evidence that the "peso problem" may be relevant in explaining earlier results that were unfavorable to UIP.[11] Interestingly, however, reestimation of the equation for the lira excluding the post-1991 period leaves a positive estimate for β, suggesting that the ERM crisis is not the main explanation for the anomalous value found over the full sample. Rather, it seems that the stochastic process driving short-term movements in the lira has systematically differed from other major currencies.

II. Long-Horizon Estimates

Some Basic Results

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 onshore 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 (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 (7).

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 of the G-7

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[11]The "peso problem" refers to the possibility that market expectations reflect the risk of "large" events that do not actually occur over the sample period. This can lead to biased estimates of slope parameters in samples that are too short to accurately reflect the small probability of large events. In other words, rational investors may appear (misleadingly) to exhibit systematic expectational errors over short samples. The implications for the unbiasedness hypothesis are lucidly discussed in Obstfeld (1989).

countries with 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.[12] Given that generalized floating began in 1973, after allowing for the 10-year lag on the interest differential, the available estimation period consisted of 1983: Q1–2000: Q4. (With limitations on the availability of bond yield data for Italy, the sample period for the lira begins in 1987: Q1.)

The results of these regressions are reported in the first panel of Table 2 (Panel 2a). They represent a surprising and stark contrast to the short-horizon results reported in Section I. 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 and the deutsche mark, the point estimates are very close to unity, and the franc also evidences a high coefficient. The yen, pound, and lira are the three cases in which UIP is statistically rejected. The adjusted R^2 statistics are also higher than in a typical short-horizon regression, with the proportion of the explained variance in the deutsche mark and the pound approaching one-half.

Because 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, Panel 2a. Its value of 0.616 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 United Kingdom, Canada, and the United States, 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, Panel 2b. The estimated slope parameters are about as close to unity as in the corresponding regressions using benchmark yields, although the pattern of coefficients is slightly different. Moreover, the constrained panel point estimate of 0.682 is closer to the posited value.[13]

We repeat the exercise with constant-maturity five-year yields for Germany, the United Kingdom, Canada, and the United States over the 1980: Q1–2000: Q4 period, to match the sample to that for our short-horizon results. The results reported in Table 2, Panel 2c are again quite favorable to the UIP hypothesis: for all three of these currencies, the slope coefficients are statistically indistinguishable from the implied value of unity. The estimate for the Deutsche mark is particularly

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[12]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 nonoverlapping 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 and 4.5 times the number of observations available otherwise.

[13]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 (2001) 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.

Table 2. Long-Horizon Tests of Uncovered Interest Parity

$$\Delta s_{t,t+k} = \alpha + \beta\left(i_{t,k} - i_{t,k}^*\right) + \varepsilon_{t,t+k}$$

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

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	R^2	N
Deutsche mark	0.003 (0.004)	0.924 (0.232)		0.44	72
Japanese yen	0.037 (0.005)	0.399 (0.144)	***	0.10	72
U.K. pound	−0.003 (0.004)	0.563 (0.104)	***	0.44	72
French franc	0.005 (0.011)	0.837 (0.442)		0.04	72
Italian lira[1]	−0.013 (0.007)	0.197 (0.151)	***	0.00	56
Canadian dollar	−0.001 (0.002)	1.120 (0.335)		0.21	72
Constrained panel[2]	. . .	0.616 (0.148)	***	0.53	360

Notes: Point estimates from the regression in equation (7) (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1983: Q1–2000: Q4. *, **, *** indicate different from null of unity at, respectively, the 10 percent, 5 percent, and 1 percent marginal significance level.

[1]Sample period: 1987: Q1–2000: Q4.
[2]Fixed-effects regression, excluding the lira. Sample period: 1983: Q1–2000: Q4.

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

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	R^2	N
Deutsche mark	0.004 (0.004)	0.918 (0.214)		0.45	72
Japanese yen	0.036 (0.006)	0.431 (0.170)	***	0.10	72
U.K. pound	0.003 (0.003)	0.716 (0.102)	***	0.45	72
Canadian dollar	−0.005 (0.003)	0.603 (0.254)		0.08	72
Constrained panel[1]	. . .	0.682 (0.143)	***	0.65	288

Notes: Point estimates from the regression in equation (7) (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1983: Q1–2000: Q4. *, **, *** indicate different from null of unity at, respectively, the 10 percent, 5 percent, and 1 percent marginal significance level.

[1]Pooled regression, with fixed effects. Sample period: 1983: Q1–2000: Q4.

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

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	R^2	N
Deutsche mark	−0.000 (0.012)	0.870 (0.694)		0.08	84
U.K. pound	−0.000 (0.015)	0.455 (0.385)		0.03	84
Canadian dollar	−0.009 (0.009)	0.373 (0.464)		0.02	84
Constrained panel[1]	. . .	0.674 (0.412)		0.10	252

Notes: Point estimates from the regression in equation (7) (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1980: Q1–2000: Q4. *, **, and *** indicate different from null hypothesis at, respectively, the 10 percent, 5 percent, and 1 percent marginal significance level.

[1]Fixed-effects regression. Standard errors adjusted for serial correlation (see text).

close to unity at 0.870, while those for the pound and Canadian dollar are closer to zero. However, in no case can one reject either the null of zero or unit slope.

There are only two other studies that we are aware of that test the unbiasedness hypothesis over similar horizons. Flood and Taylor (1997) calculate three-year exchange rate changes and collect 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. Flood and Taylor obtain a coefficient on the interest differential of 0.596 with a standard error of 0.195. Thus both hypotheses, that β equals either zero or unity, can be rejected. Alexius (2001) examines 14 long-term bond rates of varying maturities for the 1957–97 period, also drawn from *IFS*.[14] Her study also finds evidence in favor of the unbiasedness hypothesis at long horizons, although it is difficult to interpret these statistical results as being consistent with UIP, as the sample encompasses periods of fixed exchange rates and extensive capital controls.

Nonetheless, 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.

Robustness Tests

We also examined whether the results were sensitive to the data frequency, sample period, and data types. First, robustness to data frequency was assessed by resorting to annual data. Using the five-year interest rates, we selected the last observation from each year and implemented the corresponding long-horizon regressions. To the extent that the number of overlapping horizons is reduced considerably (truncation lags of only 4, instead of 19, are now required), one might expect the small-sample bias of the Hansen-Hodrick standard errors to be mitigated. The results are reported in Table 3. Once again, all point estimates are positive and insignificantly different from one. The fixed-effects estimate of the slope coefficient is 0.514. Interestingly, none of the substantive conclusions change as one moves from the results of Table 2, Panel 2c to those of Table 3.

Second, we check whether expanding the sample (so that it no longer corresponds to that of the short-horizon results) has a substantial impact upon our results. Bekaert, Wei, and Xing (2002) have argued that the finding of long-horizon UIP is specific to the post-1980s sample. Hence, we use a sample of 1977: Q1–2000: Q4 (using five-year interest rates beginning in 1972). In this case, the coefficients drop somewhat, but remain positive. Moreover, in no case can the null hypothesis of a unit slope coefficient be rejected.[15]

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[14]The *IFS* data are somewhat problematic in that the definitions of the long-term bonds is not homogeneous across countries and over time.

[15]Indeed, only by restricting our sample to approximately correspond to Bekaert, Wei, and Xing's sample (equivalent to 1977: Q1–1996: Q3 at a quarterly frequency), thereby dropping the latest observations, can we obtain rejection of the unit slope coefficient. Note that this earlier period (with five-year interest rates corresponding to those in 1972: Q1) will better capture the effects of capital controls on onshore interest rates. For instance, Frankel (1984) observes that capital controls on short-term rates were only removed in Japan in the early 1980s.

Table 3. Long-Horizon Tests of Uncovered Interest Parity, Annual Data

$$\Delta s_{t,t+k} = \alpha + \beta\left(i_{t,k} - i_{t,k}^*\right) + \varepsilon_{t,t+k}$$

5-Year Government Bond Yields
(MA(4)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	R^2	N
Deutsche mark	0.001 (0.013)	0.608 (0.902)		0.03	21
U.K. pound	0.001 (0.018)	0.402 (0.529)		0.02	21
Canadian dollar	−0.006 (0.009)	0.608 (0.534)		0.04	21
Constrained panel[1]	. . .	0.514 (0.473)		0.06	63

Notes: Point estimates from the regression in equation (7) (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1980: Q1–2000: Q4. *, **, and *** indicate different from null hypothesis at, respectively, the 10 percent, 5 percent, and 1 percent marginal significance level.

[1]Fixed-effects regression. Standard errors adjusted for serial correlation (see text).

An alternative question is whether it is truly the *maturity* of the interest rate, rather than the long horizon, that matters. Estimating the same long-horizon regression in equation (7), but substituting the three-month interest differential for the five-year interest differential, fails to produce the long-horizon results obtained earlier. Half of the point estimates are negative (although not statistically significantly different from zero); and in all cases, the null of $\beta = 1$ can be rejected. This is not surprising, as Chen and Mark (1995) found that long-horizon regressions using as determinants *short-term* interest rates yielded negative coefficients.

Finally, we address one complication that arises with the use of long-term bond data—that the reported yields are not zero-coupon rates. We checked to see if the results were sensitive to our use of yield-to-maturity rates for constant maturities instead of zero-coupon yields.[16] The estimates did not differ substantially. In the case of the deutsche mark, the zero-coupon data result in slightly higher point estimates (0.367 vs. 0.305) and larger standard errors (0.821 vs. 0.768). The pound provides a slight contrast, with a slightly lower estimate (0.413 vs. 0.477) and slightly larger standard error (0.401 vs. 0.345).

III. Explaining the Results

The stark differences between the results of tests of UIP using short- versus long-horizon data are a puzzling anomaly. None of the standard explanations for the UIP puzzle—risk premiums, expectational errors, or peso problems—appear at first glance to offer an explanation for why the results should be so different using essentially the same sample periods for the tests. Here, we propose a solution to the UIP puzzle based on the properties of a small macroeconomic model that incorporates feedback mechanisms between exchange rates, inflation, output, and

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[16]We thank Geert Bekaert for graciously allowing us to use his zero-coupon yield series.

interest rates. In particular, the model generates simulated data that are fully consistent with the stylized facts: that regressions using short-horizon data yield negative slope coefficients and explain little if any of the variance in exchange rates, while long-horizon regressions yield coefficients close to unity and explain a much higher proportion of exchange rate movements.

The model is in the spirit of the framework outlined in McCallum (1994), but allows for a richer interaction between interest rates and exchange rates. Stochastic simulations of the model are performed to generate a synthetic database, which is then used to replicate standard short- and long-horizon tests of UIP. The regressions using the synthetic data are similar to those obtained using actual data for the G-7 countries, with a pronounced difference between the short- and long-horizon parameter estimates. In the short run, shocks in exchange markets lead to monetary policy responses that result in a negative correlation between exchange rates and interest rates, contrary to the unbiasedness hypothesis under UIP. Over the longer term, in contrast, exchange rates and interest rates are determined by the macroeconomic "fundamentals" of the model, and thus behave in a manner more consistent with the conventional UIP relationship.

McCallum's framework is based on a two-equation system consisting of a UIP relationship augmented by a monetary reaction function that causes interest rates to move in response to exchange rate changes:

$$\Delta s^e_{t,t+1} = \left(i_t - i^*_t\right) - \eta_t$$

$$\left(i_t - i^*_t\right) = \lambda \Delta s_t + \sigma\left(i_{t-1} - i^*_{t-1}\right) + \omega_t,$$

where $i_t - i^*_t$ represents the interest differential, η_t is a stochastic shock to the UIP condition, and ω_t is an interest rate shock. McCallum is agnostic about the nature of the factors that underlie η. We follow the same convention, simply calling it for the time being an "exchange market" shock. McCallum solves this model to show that the parameter on the interest rate in the reduced-form expression for the change in the next-period exchange rate is $-\sigma/\lambda$, which will be negative, given conventional parameter values.[17]

The applicability of McCallum's interest rate reaction function has been criticized by Mark and Wu (1996), who find a value of λ that is small and insignificant for Germany, Japan, and the United Kingdom. More generally, his reaction function does not incorporate variables that are usually believed to be of concern to policymakers, such as inflation and output. In this sense, McCallum's model does not provide a complete characterization of macroeconomic interactions, but it serves the narrower purpose of illustrating how a negative correlation between interest rates and exchange rate movements might be generated in a consistent framework.

To generalize McCallum's model and allow a richer characterization of the feedback process between interest rates and exchange rates, we extend it by including equations for output and inflation. The monetary reaction function is then spec-

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[17]He also allows for first-order autocorrelation in η. In this case, the parameter on the interest rate becomes $(\rho - \sigma)/\lambda$, which McCallum argues will also be negative for plausible parameter values.

Table 4. Simulation Model

Uncovered interest parity:

$$\Delta s^e_{t,t+1} = \hat{i}_t - \eta_t$$

Monetary reaction function:

$$i_t - \hat{\pi}_t = 0.5(\hat{\pi}_t + \hat{y}_t)$$

Inflation (π) equation:

$$\hat{\pi}_t = 0.6\hat{\pi}_{t-1} + (1 - 0.6)\hat{\pi}^e_{t,t+1} + 0.25\hat{y}_t + 0.1\Delta(s_t - \hat{p}_t) + v_t$$

Output (y) equation:

$$\hat{y}_t = 0.1(s_t - \hat{p}_t) - 0.5(\hat{i}^{l,e}_t - \pi^{l,e}_i) + 0.5\hat{y}_{t-1} + \varepsilon_t$$

Price level (p) identity:

$$\hat{p}_t = \hat{p}_{t-1} + \hat{\pi}_t$$

Exchange rate (s) identity:

$$s_t = s_{t-1} + \Delta s_t$$

Long-term expected interest rate:

$$\hat{i}^{l,e}_t = (1/5)(\hat{i}_t + \hat{i}^e_{t,t+1} + \hat{i}^e_{t,t+2} + \hat{i}^e_{t,t+3} + \hat{i}^e_{t,t+4}) + \mu_t$$

Long-term expected inflation rate:

$$\hat{\pi}^{l,e}_t = (1/5)(\hat{\pi}_t + \hat{\pi}^e_{t,t+1} + \hat{\pi}^e_{t,t+2} + \hat{\pi}^e_{t,t+3} + \hat{\pi}^e_{t,t+4})$$

Note: ^ denotes a variable expressed relative to the same U.S. variable.

ified so that interest rates adjust in response to movements in output and inflation, using the rule proposed by Taylor (1993). To the extent that output and inflation are affected by the exchange rate, interest rates will still respond to innovations in the disturbance in the UIP relationship, but through a less direct channel than originally posited by McCallum. The model is described in Table 4, where the variables are interpreted as being measured relative to those in the partner country against which the exchange rate is defined—in this case, the United States. The periodicity is assumed to be annual, and all variables are expressed at annual rates.

The inflation equation is an expectations-augmented Phillips curve: current period inflation adjusts in response to past inflation, expected future inflation, the current output gap, and the current change in the real exchange rate.[18] The theoretical

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[18]Equivalently, the equation can be rewritten in terms of the change in the nominal exchange rate by bringing the inflation term (Δp_t) to the left-hand side and dividing through the other parameters by $(1 + 0.1)$.

justification for this type of equation is discussed in Chadha, Masson, and Meredith (1992). Parameter values have been chosen to be broadly consistent with the empirical evidence, using panel data for the G-7 countries. The output equation is a standard open-economy investment-saving (IS) curve, where output responds to the real exchange rate, the expected long-term real interest rate, and the lagged output gap. The parameters have been chosen such that a 10 percent appreciation in the real exchange rate reduces output by 1 percent in the first year and by 2 percent in the long run; a 1 percentage point rise in the real interest rate lowers output by 0.5 percent in the first year and 1 percent in the long run.[19] The long-term interest rate is determined as the average of the current short-term interest rate and its expected value over the four subsequent periods—thus, the long-term rate can be thought of as a five-year bond yield that is determined by the expectations theory of the term structure. Expected long-term inflation is defined similarly in constructing the real long-term interest rate.

Stochastic elements are introduced via four processes, all of which are assumed to be white noise: exchange market shocks (η_t), inflation shocks (ν_t), output shocks (ε_t), and term structure shocks (μ_t). We characterize the solution using numerical simulations based on the stacked-time algorithm for solving forward-looking models described in Armstrong and others (1998).[20] An important feature of the solution path is that expectations are consistent with the model's prediction for future values of the endogenous variables, based on available information about the stochastic processes. As the innovation terms η_t, ν_t, ε_t, and μ_t are assumed to be independent and uncorrelated, the information set consists of the contemporaneous innovations as well as the lagged values of the endogenous variables.[21] In this sense, expectations are fully rational, given the model structure. Nevertheless, agents lack perfect foresight, because they cannot anticipate the sequence of future innovations that determine the realizations of the endogenous variables. As the innovations are white noise, so are the associated expectational errors.

The only other information needed to perform the simulations is the variance of the stochastic processes. These were chosen to yield simulated variances of exchange rates, inflation, output, and long-term interest rates that are consistent with the stylized facts for the G-7 countries, as discussed below. Specifically, the standard deviation of the year-to-year movement in the log of the exchange rate, averaged across the G-7 countries (excluding the United States, the numeraire currency) is about 12.0 percentage points for the 1975–97 period. The standard deviations in the year-to-year movements in inflation and the log of output (relative to the United States) are much lower, at about 2.0 percentage points and 1.9 percentage points, respectively; the standard deviation of the long-term interest rate

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[19]These responses are broadly consistent with the average values across the G-7 countries embodied in MULTIMOD, the IMF's macroeconomic simulation model (Masson, Symansky, and Meredith, 1990).

[20]The performance of this algorithm is compared with that of other forward-looking solution methods in Juillard and others (1998). The simulations were performed using Portable Troll version 1.031. Data and programs are available on request from the authors.

[21]At any point in time, the conditional expectation of the future values of the innovations is zero, given the assumption that they are white noise.

differential is 1.1 percentage points. Simulations indicated that these values were consistent with a standard deviation for the exchange market innovation of 9.5 percentage points, for the inflation innovation of 1.2 percentage points, for the output innovation of 1.8 percentage points, and the term structure shock of 0.9 percentage points.

Figure 1. Impulse Response Functions to Standardized Shocks

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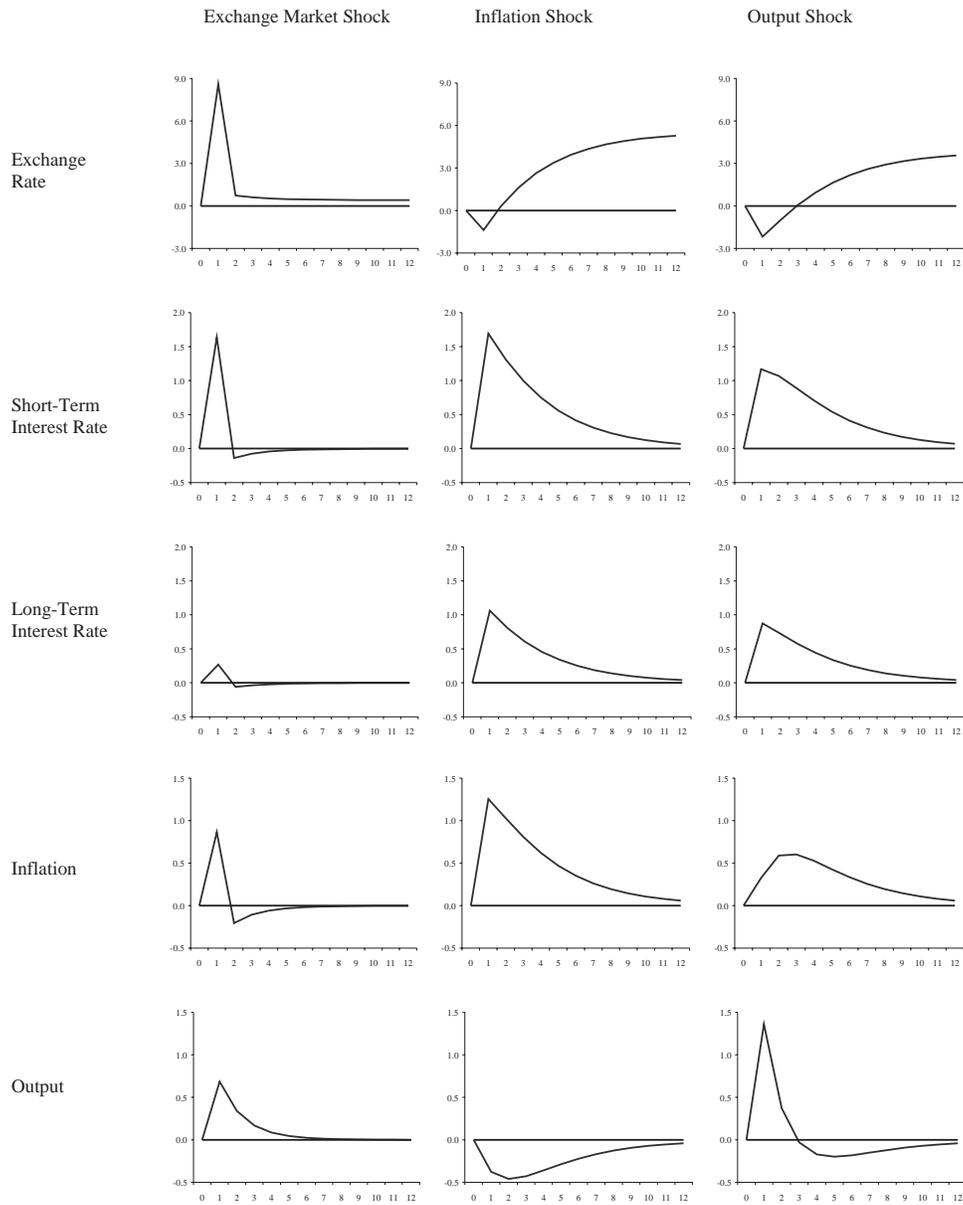
To illustrate the model's properties, Figure 1 shows impulse responses for standardized innovations in the disturbances for the exchange rate, inflation, and output. In the face of a temporary exchange market shock, the log of the exchange rate rises (in other words, depreciates) by 9 percentage points in the first period. This raises inflation by almost 1 percentage point, and the log of output by 0.7 percentage point. Under the Taylor Rule, these movements in inflation and output cause the short-term interest rate to rise by slightly more than 1.5 percentage points. In the second period, the shock dissipates and the log of the exchange rate *appreciates* by about 8 percentage points, reversing the initial increase in inflation and the short-term interest rate, while output declines toward its baseline level. The exchange rate appreciation in the second period occurs in spite of a higher lagged short-term interest rate, implying the opposite response to that predicted by UIP. This reflects the rise in the lagged exchange market shock, which generates a perverse short-run correlation between the lagged interest rate and the next-period change in the exchange rate. From a low-frequency perspective, though, the effects of the exchange market shock show little persistence. This is reflected in the muted response of the long-term interest rate (defined here as the five-year bond yield), which increases by only 0.25 percentage point in the first period before returning close to baseline in the second.

An inflation shock causes the short-term interest rate to rise by roughly the same amount in the first period as an exchange market shock. The exchange rate initially appreciates in response to higher interest rates, followed by depreciation in subsequent periods, as implied by the well-known "overshooting" model of Dornbusch (1976).[22] In all periods after the first period (when the shock hits), the relationship between the change in the exchange rate and the lagged interest rate is consistent with UIP, in contrast to the situation with an exchange market shock. The long-term interest rate also initially rises by much more, indicating that the inflation shock has greater persistence in its effects on short-term interest rates. This difference is important, because it implies a greater covariance between the long-term interest rate and the future change in the exchange rate under an inflation shock than under an exchange market shock. This, in turn, puts greater weight on comovements in interest rates and exchange rates that are UIP-consistent at longer horizons.

Similarly, an output shock causes short- and long-term interest rates to rise on impact, and the exchange rate initially appreciates, followed by subsequent depreciation. Although the changes in interest rates are not as large as under an inflation shock, the results are qualitatively similar—long-term interest rates rise by much more than with an exchange market shock, which again result in more weight being placed on UIP-consistent movements in the data at longer horizons.

To confirm the intuition provided by the impulse response functions, stochastic simulations were performed on the model. Each simulation was performed over an 85-year horizon, with the first 30 and last 30 years being discarded to avoid contamination from beginning- and end-point distortions. This left a synthetic

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[22]The long-run depreciation of the nominal exchange rate under an inflation shock reflects an increase in the domestic price level, which is not tied down under the Taylor Rule. The *real* exchange rate returns to its initial level in the face of a temporary inflation shock.

Table 5. Actual and Simulated Moments of Model Variables		
	Actual	Simulated
Standard deviation of:		
Δs_t	12.0	12.1
$\Delta \hat{\pi}_t$	2.0	1.9
$\Delta \hat{y}_t$	1.9	1.9
$\Delta \hat{\imath}_t$	3.4	3.2
$\Delta \hat{\imath}_t^1$	1.1	1.2
Correlation of:		
$\hat{\imath}_t, \hat{\imath}_{t-1}$	0.52	0.50

Note: "^" denotes a variable expressed relative to its U.S. equivalent.

sample of 25 observations (interpreted as years) for estimation, roughly corresponding to the length of the period available for performing UIP regressions using actual data. This process was repeated 5,000 times to generate a hypothetical population of 5,000 such draws.

Table 5 compares some important moments of the simulated data with the actual values for the G-7 countries during the 1975–98 period.[23] It is apparent that the model replicates closely the observed volatility in the actual data. Interestingly, volatility in short-term interest rates is similar in the simulated and actual data, even though no explicit rate shock is incorporated in the interest rate reaction function in the model. In addition, the serial correlation of short-term interest rates is very similar in the simulations to that in the actual data, in spite of the fact that an "interest rate smoothing" term is not included in the reaction function and the model's innovations are serially uncorrelated. This reflects the propagation over time of shocks via the lagged dependent variables in the inflation and output equations.[24]

For each draw, standard UIP regressions were run, using horizons varying from one to five years. The average parameters as well as their standard errors are shown in Figure 2. The most striking result is the difference in the slope parameters at the one-year horizon versus those at longer horizons. For the one-year horizon, the average slope parameter of –0.50 is similar to those obtained in Section I using data for the G-7 countries. Given the average standard error of 0.78, it would be possible to reject the hypothesis that β equals unity at conventional levels of significance in the typical sample, but not the hypothesis that it is zero. In the five-year regressions, the average estimated value of β is 0.74, with a standard error of 0.31. Thus, one could reject the hypothesis that β equals zero in the typical sample, but not that it equals unity. These results are broadly consistent with the pooled regressions using long-horizon data reported in Section II.

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[23]The G-7 data are measured relative to U.S. values for the other six countries.

[24]This contrasts with McCallum's model, which requires the assumption of serially correlated exchange market shocks to generate serial correlation in interest rates, even though the model incorporates an interest rate smoothing mechanism.

Figure 2. UIP Slope Parameter from Model Simulations

Regression horizon
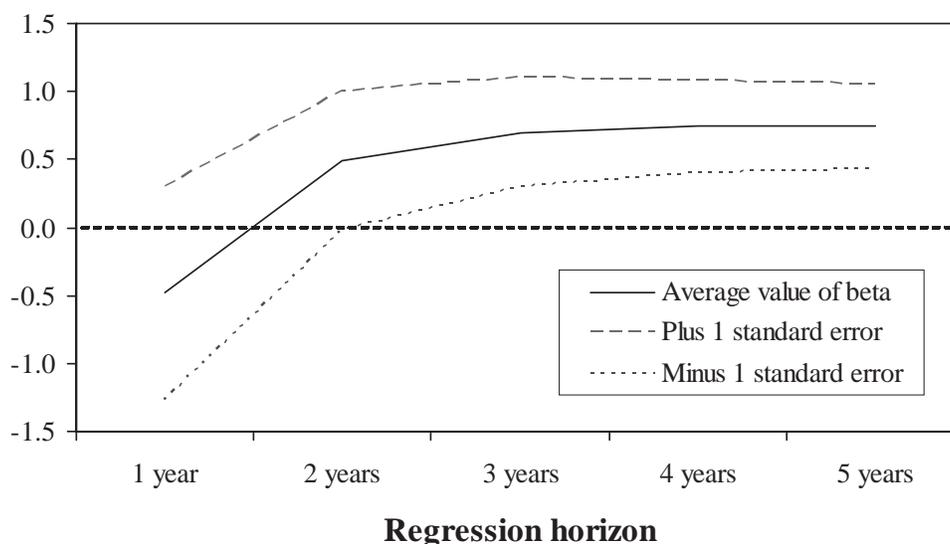
Another interesting comparison between the regressions involves the $\overline{R}^2$ statistics. The average value in the one-year regressions is only 0.001, indicating virtually no explanatory power, similar to the regressions using actual data. For the five-year regressions, in contrast, the average value rises to 0.21. Again, this is consistent with the stylized facts from the long-horizon regressions. In any case, even in the context of a model whose structure is unchanging over time and where agents are endowed with fully rational expectations, interest differentials explain only a small component of the variance in longer-term exchange rate movements.

So it appears that a small, forward-looking macroeconomic model with a conventional structure is capable of explaining the important stylized facts relating to tests of UIP at short and long horizons. The failure of UIP over short horizons is consistent with the endogeneity of interest rates in the face of disturbances in exchange markets. Over the longer term, in contrast, the model's underlying dynamics dominate and UIP performs better.[25] We interpret this as evidence that the bias in interest rates as a predictor of exchange rate movements arises from the behavior of monetary authorities in "leaning against the wind" in the face of exchange rate shocks via their effect on output and inflation.

This framework is designed to explain the source of the correlation between deviations from RNEMH and the interest differential, and the implications for short- and long-horizon UIP tests. The framework does not, however, provide a motivation for the underlying shocks in exchange markets of the size needed to

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[25]This is consistent with the finding in Mark (1995) and Chinn and Meese (1995) that short-horizon movements in exchange rate are dominated by noise, while longer-term movements can be related to economic fundamentals.

generate the observed volatility in exchange rates. Conceptually, these shocks must reflect either risk premiums or deviations from rational expectations. Regarding risk premiums, it is well known that conventional consumption-based asset pricing models are unable to generate shocks of the magnitude required to explain observed price fluctuations, not only in exchange markets, but in financial markets more generally. More recent analyses based on "first-order" risk aversion, such as Bekaert, Hodrick, and Marshall (1997), also generate risk premiums that are far smaller than the shocks needed to explain the actual data. Beyond this, it has also proved difficult to relate ex post exchange risk premiums to macroeconomic factors.

Regarding deviations from rational expectations, it has been found that surveys of exchange rate forecasts generally fail to predict future exchange rate movements. The model used here cannot explain this regularity, as the rational expectation of agents regarding the future change in the exchange rate (i.e., the solution value of $i_t - \eta_t$) will be an unbiased predictor of the actual change. In the absence of an explanation for this puzzle, the possibility cannot be ruled out that expectational errors explain the differences in results at short versus long horizons. As documented by Froot and Ito (1989), short-term expectations tend to "overreact" relative to long-term expectations. Furthermore, Chinn and Frankel (2002) find that there is some evidence of reversion to purchasing power parity (PPP) at longer (five-year) horizons, while such evidence is more difficult to find at shorter horizons. These observations could support the argument that expectations are less "biased" (for whatever reasons) at long horizons, and hence may be more conducive to finding UIP.

IV. 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 standard test of UIP yields 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 UIP may work better at longer horizons.

The difference in the results is shown to be consistent with the properties of a conventional macroeconomic model. In particular, a temporary disturbance to the UIP 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. The model, though, cannot explain why such shocks are as large as needed to explain observed exchange rate volatility. Neither can it explain why tests using survey data on exchange rate expectations fail to uncover an unbiased relationship

between expected and actual exchange rate movements. So there are puzzles that remain to be explored.

Regardless of the reasons for the failure of UIP at short horizons, from an unconditional forecasting perspective, the conclusion remains that the simple form of the UIP condition is essentially useless as a predictor of short-term movements in exchange rates, even if there might be some information imbedded in the term structure of forward rates.[26] Over longer horizons, however, our results suggest that UIP may significantly outperform naive alternatives such as the random walk hypothesis, although it is still likely to explain only a relatively small proportion of the observed variance in exchange rates.

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[26]See, for instance, Clarida and Taylor (1997) and Clarida and others (2003).

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