Monetary Policy and Long-Horizon Uncovered Interest Parity *

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Abstract

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 G-7 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 interaction of stochastic exchange market shocks with endogenous monetary policy reactions. In the long run, in contrast, exchange rate movements are driven by the "fundamentals," leading to a relationship between interest rates and exchange rates that is more consistent with UIP.

JEL Classification: F21, F31, F41

Keywords: uncovered interest parity, unbiasedness hypothesis, long-horizon, monetary policy, Taylor Rule

I. INTRODUCTION

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 (1990) reports 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 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 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

[1] We define UIP here in terms of the unbiasedness hypothesis that the coefficient on interest differentials in exchange rate regressions is unity. Other authors sometimes term this the "risk-neutral efficient markets hypothesis" (see, e.g., Clarida and Taylor (1997)). Terminological issues are discussed further in Section II.

forward exchange rates of the same maturity as the associated financial asset. In the event, relatively thick forward exchange markets only exist 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 (1994a). In particular, a more general monetary reaction function is incorporated 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 based are used to generate a synthetic database for replicating UIP tests. Standard regressions using 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. 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 remarkably closely the key properties of actual data for the G-7 countries.

The paper is structured as follows. Section II reviews the UIP hypothesis, summarizes the existing evidence over short horizons, and provides updated results from 1980 through early 1998. Section III presents estimates of the UIP hypothesis using data on government bond yields for the G-7 countries. Section IV develops a model that is consistent with the key features of the observed data, while Section V provides concluding remarks.

II. 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.[2] Algebraically, CIP can be expressed as:

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

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

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 lower-case letters), equation (1) becomes:

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

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:

$$f_{t,t+k} \;=\; s^e_{t,t+k} \;+\; rp_{t,t+k} \, . \tag{3}$$

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:

$$\Delta s^e_{t,t+k} \;=\; (i - i^*)_{t,k} - rp_{t,t+k} \, , \tag{4}$$

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.[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:

$$s_{t+k} = s^{re}_{t,t+k} + \xi_{t,t+k} , \qquad \textbf{(5)}$$

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) gives the following relationship:[4]

$$\Delta s_{t,t+k} = (i - i^*)_{t,k} - rp_{t,t+k} + \xi_{t,t+k} , \qquad \textbf{(6)}$$

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

It is natural, then, to test the combined hypothesis of UIP and rational expectations via the regression equation:

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

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

[4]This condition is also sometimes referred to as the "risk-neutral efficient markets hypothesis." In the absence of risk neutrality, market efficiency does not require that the forward exchange rate equals its expected future level. Tests of this more general version of market efficiency are not possible, however, in the absence of direct measures of risk premia in exchange markets.

Under the assumption that the composite error term $\epsilon_{t,t+k}$ is orthogonal to the interest differential, the estimated slope parameter in equation (7) should be unity. This is generally referred to as the "unbiasedness hypothesis" in tests of UIP. In addition, no other regressors known at time *t* should have explanatory power, as all available information should be captured in the rational expectation of $\Delta s_{t,t+k}$ as reflected in the period-*t* interest differential. Regarding the constant term, non-zero 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)).[5] 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 III 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, 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

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

estimate of β” (page 31).[6] 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.[7]

To illustrate the performance of short-horizon UIP for the exchange rates of the G-7 countries, Table 1 presents estimates of equation (7) for the period from the first quarter of 1980 to the first quarter of 1998. 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.[8] 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 (1992) to correct the standard errors of the parameter estimates for moving

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[6]Other recent 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 the ERM were expected (and observed).

[7]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.

[8]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.

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).

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 -0.3 at the 12-month horizon. 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. Only for the lira is it possible to reject the random-walk model while not also rejecting UIP. All of the adjusted R^2 statistics are very low, and occasionally negative.

The range of slope coefficients is somewhat larger than reported in most previous studies, with estimates for the lira yielding an estimated value for β of about 1½ at the 3-month horizon and almost 2 at the 12-month horizon. These anomalies are consistent with the results of Chinn and Frankel (1994), who find 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.[9] Interestingly, however, reestimation of the equation for the lira excluding

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[9]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 (continued...)

the post-1991 period leaves a estimate for β well above unity, 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.

More generally, the parameters for some countries proved robust to changes in the sample, while those for other countries experienced sharp shifts. As shown in Table 2, splitting the observations into the 1980-88 and 1989-98 sub-periods yields estimates for the slope parameters that are qualitatively similar across the two periods for the yen and the lira. For the pound sterling, in contrast, the results are dramatically different, with the more recent period yielding slope parameters that are close to the null under UIP. The constrained panel estimates give coefficients that are less negative at all three horizons during the more recent period. This finding could be interpreted as providing some support for one explanation for the failure of UIP—that markets had not had enough experience with floating exchange rates in the 1970s and early 1980s to understand the process driving them, resulting in large forecast errors. But the evidence for "market learning" is weak, given that the signs of the pooled coefficients in the more recent sample—which begins 16 years after the introduction of generalized floating—are all still perverse. So if markets have learned from the experience under floating rates, they haven't learned much, and the result found in earlier studies that the sign on the interest differential in UIP regressions is perverse still holds.

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[9](...continued)
exhibit systematic expectational errors over short samples. The implications for the unbiasedness hypothesis are lucidly discussed in Obstfeld (1989).

III. LONG-HORIZON ESTIMATES

III.A. 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 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 (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 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.[10] Given that generalized floating began in 1973, after allowing for the 10-year lag on the interest differential, the available estimation period consisted of 1983Q1–2000Q4 (given limitations on the availability of bond yield data for Italy, the sample period for the lira begins in1987Q1).

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 and Deutsche mark, the point estimates are very close to unity, while 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 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 0.616 is well below unity; on

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[10] 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 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., Canada 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 about as close to unity as in the corresponding regressions using benchmark yields, although the pattern of coefficients is slightly different. Moreover, the panel point estimate of 0.682 is closer to the posited value.[11]

We repeat the exercise with constant-maturity 5-year yields for Germany, the U.K., Canada, and the U.S. over the 1980Q1-2000Q4 period, to match the sample to that for our short horizon results. The results reported in Table 2.c 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 Deutschemark is particularly 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 calculate 3-year exchange rate changes and collect average data on medium-term government bonds from the IMF's *International Financial*

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[11] 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.

Statistics (*IFS*). The data over the 1973–92 period are then pooled for a sample of 21 countries. They obtain 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. Alexius (2001) examines 14 long term bond rates of varying maturities for the 1957-1997 period, also drawn from *IFS*.[12] 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 uncovered interest parity, 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.

III.B. 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, the last observation from each year was selected, and the corresponding long horizon regressions implemented. 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

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

insignificantly different from zero. 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.c 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 et al. (2002) have argued that the findings of long horizon UIP is specific to the post-1980's sample. Hence, we use a sample of 1977Q1-2000Q4 (using 5 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 be rejected.[13]

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 (7), but substituting the 3 month interest differential for the 5 year interest differential, does not produce the same set of results as obtained before. 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 (1996) found that long horizon regressions using as fundamentals *short term* interest rates yielded negative coefficients.

Finally, we address one complication that arises with the use of long term bond data. One drawback of using bond data – as opposed to offshore deposit rates – is that one has to account

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[13] Indeed, only by restricting our sample to approximately correspond to Bekaert et al.'s sample (equivalent to, 1977Q1-1996Q3 at a quarterly frequency), thereby dropping the latest observations, can we obtain rejections of the unit slope coefficient. Note that this earlier period (with 5 year interest rates corresponding to those in 1972Q1) will more capture the effects of capital controls on onshore interest rates. For instance, Frankel (1984) concludes that capital controls on short term rates were only removed in Japan in the early 1980's.

for the fact 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.[14] The estimates did not differ substantially depending upon the series. In the case of the Deutschemark, 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).

IV. 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 premia, expectational errors, or peso problems—appears 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 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 (1994a), but allows for a richer interaction between interest rates and exchange rates. Stochastic simulations of the

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

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 manner more consistent with the conventional UIP relationship.

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

$$\Delta s^{e}_{t,t+1} = (i_t - i_t^*) - \eta_t$$

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

where $i_t - i_t^*$ represents the interest differential, η_t is a stochastic shock to the uncovered interest parity condition, and ω_t is an interest rate shock. McCallum is agnostic about the nature of the factors that underly 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 -σ/λ, which will be negative given conventional parameter values.[15]

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 U.K. 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 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 specified 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.

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[15]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.

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.[16] The theoretical justification for this type of equation is discussed in Chadha, Masson, Meredith (1993). 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 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 ½ percent in the first year and 1 percent in the long run.[17] 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 three processes, all of which are assumed to be white noise: exchange market shocks (η_t), inflation shocks (ν_t), and output shocks (ϵ_t). We characterize the solution using numerical simulations based on the stacked-time algorithm for

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[16]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).

[17]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)).

solving forward-looking models described in Armstrong, Black, Laxton, and Rose (1998).[18] 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, and ϵ_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.[19] 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 relative variance of the three stochastic processes. These were chosen to yield simulated variances of exchange rates, inflation, and output that are consistent with the stylized facts for the G-7 countries. Specifically, the standard deviation of the year-to-year movement in the exchange rate, averaged across the G-7 countries (excluding the U.S., the numeraire currency) is about 12.0 percent for the 1975-97 period. The standard deviations in the year-to-year movements in inflation and output (relative to the US) are much lower, at about 2.0 percent and 1.9 percent respectively. Experimental simulations indicated that these values were broadly consistent with a standard deviation for the

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

[19]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.

exchange market innovation of 9.7 percentage points, for the inflation innovation of 1.3 percentage points, and for the output innovation of 1.1 percent.

To illustrate the model's properties, Figure 1 shows impulse responses for standardized innovations in each disturbance. In the face of a temporary exchange market shock, the exchange rate depreciates by 9 percent in the first period. This raises inflation by almost 1 percent, and output by ¾ percent. Under the Taylor Rule, these movements in inflation and output cause the short-term interest rate to rise by slightly over 1½ percentage points. In the second period the shock dissipates and the exchange rate *appreciates* by about 8 percent, 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 5-year bond yield), which increases by only ¼ 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 Dornbush (1976).[20] 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, while 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 results 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 a 140-year horizon, with the first 30 and last 30 years being discarded to avoid contamination from beginning- and end-point considerations. This yielded a "sample" of 80 observations for each simulation. This process was repeated 50 times to generate a hypothetical population of 50 such

[20]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.

samples. For each sample, standard UIP regressions were run using horizons varying from 1 to 10 years. The results of the 1-year and 5-year regressions for a representative population of 50 simulations are shown in Table 5.

The most prominent feature of the results is the difference in the slope parameters between the regressions at the 1-year horizon versus those at longer horizons. For the 1-year regressions, the average slope parameter of -0.50 is the same order of magnitude as those obtained in Section II using data for the G-7 countries. Given the average standard error of 0.42, it would easily be possible to reject the hypothesis that β equals unity with a high level of confidence in the typical sample. In both the 5-year and 10-year regressions, the average estimated β is 0.82, with a standard error of only 0.18. Thus, one could reject the hypothesis that β equals zero at conventional confidence levels, but not generally reject the hypothesis of unity. These results are generally consistent with the pooled regressions using long-horizon data reported in Section III. There are large outliers in some of the samples, however. The short-horizon coefficient ranges from -1.37 to 0.37 across samples, indicating the variability in estimation results that could be obtained using samples even as long as 80 periods. The 5-year results are somewhat more tightly clustered, ranging from 0.53 to 1.32. For both the short- and long-horizon regressions, the average standard error of β found in individual samples is similar to the standard deviation calculated across the 50 samples, suggesting that the calculated standard errors in the regressions are indeed good estimates of the sample variability of the coefficients.

Another interesting comparison between the regressions involves the adjusted R^2 statistics. The average value in the 1-year regressions is only 0.01, indicating a virtual complete lack of explanatory power, similar to the regressions using actual data. For the 5-year regressions, in contrast, the average value rises to 0.21, with some draws as high as 0.46. Again, this is consistent with the stylized facts from the actual long-horizon regressions reported above. Thus, even in the context of a model whose structure is unchanging over time and where agents are assumed to have fully rational expectations, interest differentials do not explain the bulk of the variance in longer-term exchange rate movements. This reflects the influence of future innovations that are inherently unpredictable in affecting the future exchange rate path.

Figure 2 illustrates the pattern of the slope coefficients at alternative horizons for three different populations of simulations. They yield very similar results, with the average slope coefficients ranging from -0.4 to -0.5 at the 1-year horizon, but becoming significantly positive at horizons of 2 years and more. Indeed, the 3-year horizon coefficients are quite close to the values of 0.7–0.8 reached at 5- and 10-year horizons. It is also interesting to note that the coefficients do not asymptote toward unity at these longer horizons, but rather stabilize at levels somewhat below the value implied by the unbiasedness hypothesis. This reflects the diminishing—but nonnegligible—role that exchange market shocks continue to play at longer horizons. The implication is that it may be unrealistic to expect to find coefficients centered on unity in UIP tests at any horizon, even in the absence of measurement errors in the data. These results are consistent with those obtained using actual long-horizon data. Flood and Taylor estimate a coefficient at a 3-year horizon of about 0.6, only slightly below that implied by the synthetic regressions. Our 5-year results are actually somewhat more favorable to UIP than implied by the synthetic regressions, but the coefficients differ by less than one standard deviation of the value estimated in the synthetic regressions. Our 10-year results are only slightly below the synthetic value.

In terms of the consistency of the model-generated volatility of the main variables with actual data, the following tabulation compares the mean results across the simulations with average values across the G-7 countries for 1975-97:[21]

	Actual	Simulated
Standard deviation of:		
Δs_t	12.0	12.5
$\Delta\hat{\pi}_t$	2.0	2.0
$\Delta\hat{y}_t$	1.9	1.4
$\Delta\hat{i}_t$	3.4	3.3
$\Delta\hat{i}^l_t$	1.1	0.9
Correlation of:		
$\hat{i}_t, \hat{i}_{t-1}$	0.52	0.50

Note: "^" denotes a variable expressed relative to the US.

The model replicates closely the observed volatility in the actual data, although the simulated standard deviations of changes in output and long-term interest rates are somewhat lower than the averages observed for the G-7 countries. It is interesting to note that changes in short-term interest rates exhibit similar volatility in the simulations as in the observed data, even though no explicit interest rate shock is incorporated in the model. In addition, the correlation of short-term interest rates with their lagged values 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

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

uncorrelated. This reflects the propagation over time of uncorrelated innovations via the lagged dependent variables in the inflation and output equations.[22]

The somewhat lower simulated variance of the long-term interest rate compared with the actual data may reflect the absence of error terms (i.e. risk premia) in the term structure equations, contrary to empirical evidence (see McCallum (1994b)). To check the sensitivity of the results to this assumption, simulations were preformed with additional stochastic disturbances added to the term structure relationships.[23] The short-horizon estimation results were virtually unaffected by this addition, while the slope parameters in the long-horizon regressions declined modestly—for instance, the average value of β in the 5-year regressions fell from 0.82 to 0.78, while that at the 10-year horizon fell from 0.78 to 0.68. This is not surprising, as the term premium introduces what amounts to an error in the regressor in the long-horizon regressions. For conventional reasons, this source of noise would bias the estimated coefficient toward zero. But the magnitude of the effect is modest in the simulations and the estimated parameter remains similar to those obtained with actual data.

So it appears that a small, forward-looking macroeconomic model with a conventional structure is capable of explaining important stylized facts relating to tests of UIP. The failure of UIP over short horizons is consistent with the endogeneity of interest

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[22]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.

[23]The standard deviations of the disturbances were calibrated to raise the standard deviation of the year-to-year movements in the long-term (10-year) bond yield to match that in the observed data, i.e. 1.1 percentage points.

rates in the face of disturbances in exchange markets. Over the longer term, in contrast, the model's "fundamentals" dominate and UIP performs better.[24]

Nevertheless, this framework cannot explain another major puzzle—the source of the disturbances in exchange markets required to generate the observed volatility in exchange rates. It is well known that conventional consumption-based asset pricing models are unable to generate risk premia of the magnitude required to explain observed price fluctuations, not only in exchange markets, but in almost all financial markets. More recent analyses based on "first-order" risk aversion, such as Bekaert, Hodrick, and Marshall (1997), also generate risk premia that are far smaller than the shocks observed in the data. Beyond this, it has also proved difficult to relate ex post exchange risk premia to macroeconomic factors. It appears then, that alternatives to approaches based solely on agents' aversion to consumption risk are needed to explain the stylized facts of predictable excess returns in asset markets.

A second puzzle is why 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 "over react" relative to long-term expectations. Furthermore, Chinn and Frankel (2002) find that there is some evidence of reversion to PPP

[24]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.

at longer (5-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.

V. 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 UIP may work better at longer horizons.

The difference in the results is shown to be fully 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. 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 UIP is essentially useless as a predictor of short-term movements in exchange rates. 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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Table 1. Short-Horizon Estimates of β

$$\Delta s_{t,t+k} \quad = \quad \alpha \ + \ \beta\,(i_{t,k} \ - \ i^{*}_{t,k}) \ + \ \epsilon_{t+k} \ . \qquad \textbf{(6)}$$

Currency	Maturity 3 mo.	6 mo.	12 mo.
Deutschemark	-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 1980Q1-2000Q4. * (**)[***] Different from null of unity at 10%(5%)[1%] marginal significance level.

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

Table 2. Long-Horizon Tests of Uncovered Interest Parity

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

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

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	R^2	N
Deutschemark	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: 1983Q1-2000Q4. * (**)[***] Different from null of unity at 10%(5%)[1%] marginal significance level.

1 Sample period: 1987Q1–2000Q4.

2 Fixed effects regression, excluding the lira. Sample period: 1983Q1 - 2000Q4.

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

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: 1983Q1-2000Q4. * (**)[***] Different from null of unity at 10%(5%)[1%] marginal significance level.

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

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: $\beta = 1$	R^2	N
Deutschemark	-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: 1980Q1-2000Q4. * (**)[***] Different from null hypothesis at 10%(5%)[1%] marginal significance level.

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

Table 3. Long-Horizon Tests of Uncovered Interest Parity:

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

	$\hat{\alpha}$	$\hat{\beta}$	Reject H_0: $\beta = 1$	R^2	N
Deutschemark	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 (6) (serial correlation robust standard errors in parentheses, calculated assuming k-1 moving average serial correlation). Sample period: 1980Q1-2000Q4. * (**)[***] Different from null hypothesis at 10%(5%)[1%] marginal significance level.

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

Table 4. Simulation Model

Uncovered interest parity:

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

Monetary reaction function:

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

Inflation (π) equation:

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

Output (y) equation:

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

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:

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

Long-term expected inflation rate:

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

Note: ^ denotes a variable expressed relative to the US.

Table 5. Regression Results from Stochastic Simulations

	Regression horizon: 1 year	5 years	10 years
Estimated slope coefficient (β)			
Average	-0.50	0.82	0.78
Maximum	0.23	1.32	1.21
Minimum	-1.37	0.53	0.30
Standard deviation	0.37	0.17	0.21
Standard error of β			
Average	0.42	0.18	0.18
Maximum	0.53	0.22	0.22
Minimum	0.31	0.14	0.14
Standard deviation	0.04	0.02	0.02
Adjusted R^2			
Average	0.01	0.21	0.19
Maximum	0.08	0.46	0.44
Minimum	-0.01	0.06	0.03
Standard deviation	0.03	0.09	0.10

Figure 1. Impulse Response Functions to Standardized Shocks

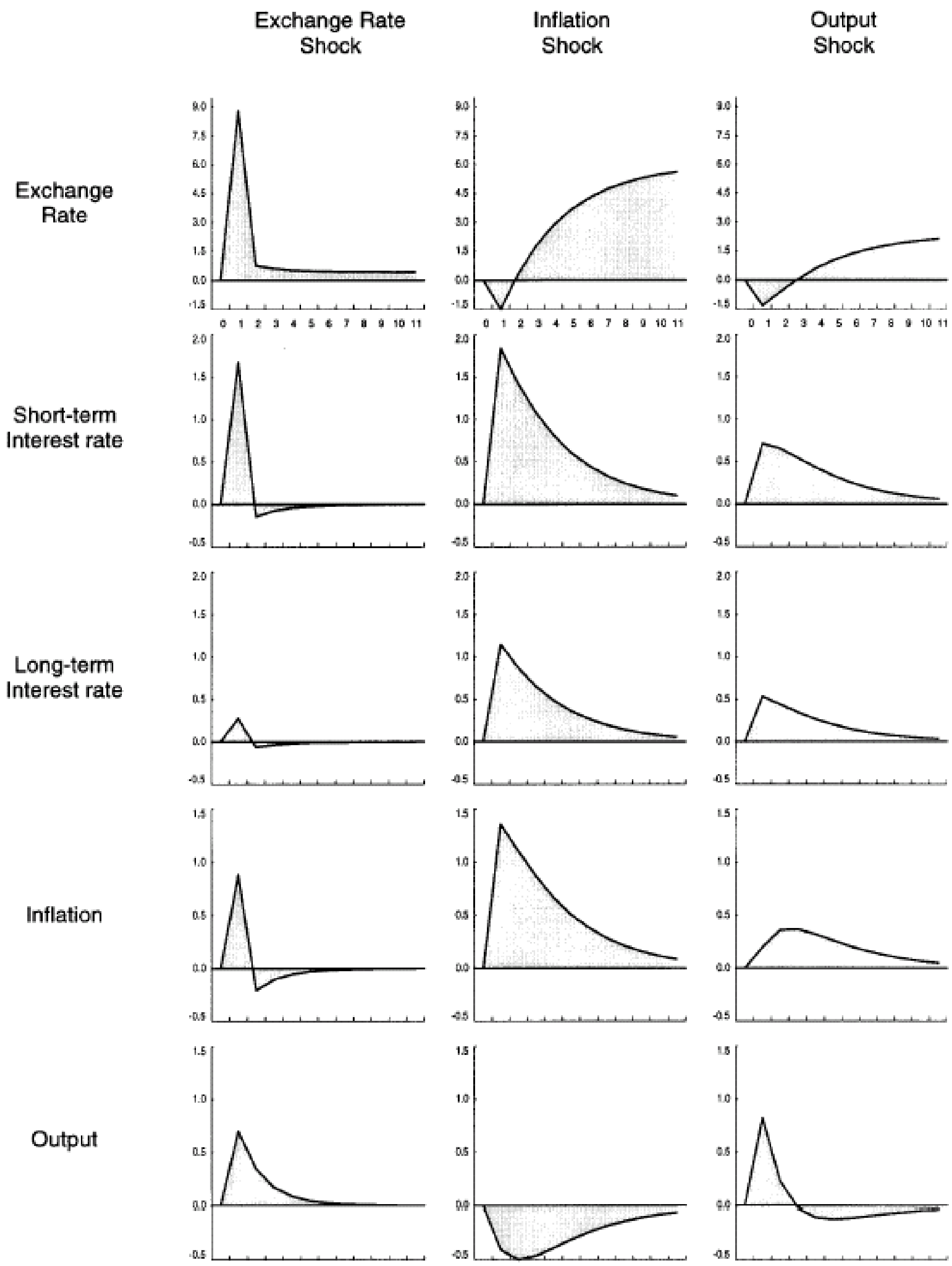

Figure 2. Estimated Slope Parameters from UIP Regressions at Different Horizons Using Data From Stochastic Simulations