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ABSTRACT

Uncovered interest parity (UIP) has been almost universally rejected in studies of exchange rate movements, although there is little consensus on why it fails. In contrast to previous studies, which have used relatively 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 the coefficients on interest differentials are of the correct sign, and *almost all* are closer to the UIP value of unity than to the zero coefficient implied by the random walk hypothesis. We then use a small macroeconomic model to explain the differences between the short- and long-horizon results. Regressions run on data generated by stochastic simulations 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 risk premium shocks in the face of endogenous monetary policy. 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.

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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 (1990b) reports few cases where even 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 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. The evidence in favor of UIP over longer horizons is not conclusive, however. Most of the estimated parameters are less than the theoretical value of

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[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 (1993)). Terminological issues are discussed further in Section II.

unity; the standard errors on some of the parameters are sufficiently large to allow a wide range of possible alternatives; and interest rate differentials still explain less than one half of the total observed variance in exchange rates.

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 outlined by McCallum (1994a). 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. Long-horizon regressions, in contrast, yield coefficients close to unity, consistent with estimation results using actual data. 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.

Most explanations for the failure of UIP posit the existence of either time-varying risk premia, expectational errors, or "peso" problems. Our results support the view that the failure of short-horizon UIP reflects the endogeneity of interest rate movements in the face of time-varying risk premia. As macroeconomic "fundamentals" play little role in tying down exchange rates at short horizons, risk premia play a greater role in contaminating short-horizon than long-horizon tests of UIP. In contrast, the model's intrinsic dynamics, particularly the response of output and inflation to movements in exchange rates and interest rates, tend to dominate over longer horizons. As the role of shocks to risk premia fades over these horizons, the empirical support for UIP rises.

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:

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

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

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

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

Equation (2) is a risk-free arbitrage condition that holds regardless of the preferences of investors. 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:

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

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

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

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 change in the expected exchange rate 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:

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[3]Indirect tests of UIP have been performed using surveys of published forecasts of exchange rates, but the latter may be suspect as measures of market expectations for the reasons noted in Bryant (1995).

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

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]

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

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:

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

Under the assumption that the composite error term $\varepsilon_{t,t+k}$, which consists of both risk premia and expectational errors, 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 exchange 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

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[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.

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

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 (1990b), 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).[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 was performed using the Generalized Methods of Moments (GMM) estimator of Hansen (1992). Estimation using the 6- and 12-month horizon data at a quarterly frequency led to overlapping observations, inducing moving average (MA) terms in the residuals. To correct for this, the standard errors of the parameter estimates were corrected for MA terms of an order equal to the number of overlapping quarters at each horizon (i.e., MA(1) in the case of 6-month data and MA(3) in the case of 12-month data).

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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.

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 (1990b). 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 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

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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 exhibit systematic expectational errors over short samples. Krasker (1980) provided the original empirical analysis for Mexico; the implications for the unbiasedness hypothesis are lucidly discussed in Obstfeld (1989).

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.

III. Long-horizon Estimates

As noted in the introduction, short-horizon tests of UIP have been facilitated by the availability of interest rate series that correspond closely to the requirements for covered interest parity. 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.[10] 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, we have used domestic yields on government bonds of maturities that correspond roughly, but not necessarily exactly, to assumed constant maturities. As well as possible deviations in actual maturity from assumed values, these on-shore instruments may be subject to differences in tax regime, capital controls, etc. Because these data will 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 from its hypothesized value of unity under UIP.

The first data set we have employed to test long-horizon UIP consists of the benchmark government bond yields used by Edison and Pauls (1993).[11] These are end-of-month yields on outstanding government bonds for the G-7 countries, generally 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 was then regressed on the 10-year lagged differential in the associated bond yield. Given that floating rates were generally introduced in 1973, after allowing for the 10-year lag on the interest differential, the available estimation period consisted of 1983Q1–1998Q1 (given limitations on the availability of bond yield data for Italy, the sample period for the lira began in 1985Q1).

The results of these regressions are reported in the first panel of Table 3. They represent a surprising and stark contrast to the short-horizon results reported in Section II. In all cases, the estimated slope coefficient is positive, with four of the six values lying closer to

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[10]The advent of long-dated swap markets in the 1980s has allowed the testing of long-horizon *covered* interest parity, albeit over relatively short sample periods (see Popper, 1993, and Fletcher and Taylor, 1996). Unfortunately, the availability of swap data is still too limited to allow meaningful tests of uncovered interest parity.

[11]We are grateful to Hali Edison for making these data available to us from the Federal Reserve Board data bank.

unity than to zero. For the Canadian dollar, the point estimate (1.104) is very close to unity, while the deutschemark and the franc also evidence high coefficients. The yen, pound and lira are the three cases in which UIP is statistically rejected. But for all currencies except the lira, the hypothesis that β equals zero can also be strongly rejected. The adjusted R^2 statistics are also typically higher than in the 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.[12] The resulting point estimate is reported under the entry "constrained panel" at the bottom of Table 3a. Its value of 0.645 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 horizons up to one year reported in Table 1.

For Japan, Germany, the U.K., and the U.S., it was also possible to obtain synthetic "constant maturity" 10-year yields from interpolations of the yield curve of outstanding government securities. The regressions using measures of long-horizon interest differentials based on these data are reported in Table 3b. For each of the three associated exchange rates, the estimated slope parameter is closer to unity than in the regressions using benchmark yields (although the difference is slight for the deutschemark). Moreover, the panel point estimate of 0.708 is also closer to the posited value. The improvement in the results, although modest, suggests that part of the reason why UIP fails using benchmark yields relates to discrepancies between the assumed and actual maturities of the outstanding securities. In other words, improvements in the quality of the data appear to systematically shift the results toward supporting the UIP hypothesis.

Similar constant-maturity 5-year yields were obtained for Germany, the U.K., Canada, and the U.S. Regressions of 5-year changes in exchange rates on the interest differentials implied by these data are reported in Table 3c. The results are even more favorable for the UIP hypothesis: for both the deutschemark and the pound, the slope coefficients are economically and statistically indistinguishable from the implied value of unity, while the null of zero under the random walk hypothesis is strongly rejected. In the case of the Canadian dollar, the point estimate is 1.34, but the 50 percent confidence bound for the point estimate easily encompasses the value of unity. Similarly, the panel estimate is essentially equal to the theoretically implied value.

The only other study that we are aware of that tests UIP over horizons of longer than 12 months is Flood and Taylor (1997). They regress 3-year changes in exchange rates on

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

annual average data on medium-term government bonds from the IMF's *International Financial Statistics*. The data over the 1973–92 period are then pooled for a sample of 21 countries. They find a coefficient on the interest differential of 0.596 with a standard error of 0.195, thus the hypotheses that β equals either zero or unity can both be rejected. These results are broadly in line with our 10-year results, although our 5-year results using constant maturity data are more supportive of UIP. This difference may reflect the fact that our end-period, constant-maturity data are more closely aligned with the requirements underlying the UIP hypothesis, although differences in country coverage and sample periods may also play a role. In any event, it is reassuring that the Flood and Taylor 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 and sample periods.

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 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, risk premium shocks dominate the relationship between interest rates and exchange rates, generating a negative correlation—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 equation (4) augmented by a monetary reaction function that causes interest rates to move in response to exchange rate changes:

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

$$i_t = \lambda \Delta s_t + \sigma i_{t-1} + \omega_t ,$$

where i_t represents the interest differential, η_t is a risk premium shock, and ω_t is an interest rate 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 values for the parameters.[13]

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, the 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 exchange rate risk premium, 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.[14] The theoretical justification for this type of equation is discussed in Chadha, Masson, Meredith (1993). Parameter values have

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

[14]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).

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 lagged output.[15] 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.[16] 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: risk premium shocks (η_t), inflation shocks (v_t), and output shocks (ε_t).[17] We characterize the solution using numerical simulations based on the stacked-time algorithm for solving forward-looking models developed by Juillard, Laxton, McAdam, and Pioro (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, v_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.

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[15]Potential output is assumed to be exogenous in this model, thus output can equivalently be thought of as the deviation of output from its potential level.

[16]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)).

[17]Simulations were also performed with a stochastic disturbance in the term-structure equation, as discussed below.

[18]The simulations were performed using Portable Troll version 1.033. 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.

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 are much lower, at about 2.0 percent and 1.9 percent respectively.[20] Experimental simulations indicated that these values were broadly consistent with a standard deviation for the risk premium innovation of 9.7 percent, for the inflation innovation of 1.3 percent, 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 risk premium 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 risk premium, 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 risk premium 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 a risk premium 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).[21] 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 a risk premium shock. The long-term interest rate initially rises by much more than under a risk premium shock, 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

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[20]Consistent with the interpretation of the model variables, inflation and output are measured as deviations from U.S. levels.

[21]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.

the long-term interest rate and the future change in the exchange rate under an inflation shock than under a risk premium 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 under the risk premium 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 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. The populations 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 risk premium 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:[22]

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

	Actual	Simulated
Standard deviation of:		
Δs_t	12.0	12.5
$\Delta \pi_t$	2.0	2.0
Δy_t	1.9	1.4
Δi_t	3.4	3.3
Δi^l_t	1.1	0.9
Correlation of:		
i_t, i_{t-1}	0.52	0.50

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 uncorrelated. This reflects the propagation over time of uncorrelated innovations via the lagged dependent variables in the inflation and output equations.[23]

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.[24] To check the sensitivity of the results to this assumption, simulations were preformed with additional stochastic disturbances added to the term structure relationships.[25] 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

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

[24]See McCallum (1994b) for a discussion of this literature.

[25]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.

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 many of the stylized facts relating to tests of UIP. The failure of UIP over short horizons reflects the endogeneity of interest rates in the face of shocks to the exchange risk premium, while the model's "fundamentals" dominate over the longer term.[26] Nevertheless, it does not explain all of the puzzles. One issue the model cannot shed light on is the variance in the risk premium shocks that is needed to generate the observed volatility in exchange rates. It is well known that conventional consumption-based asset pricing models are unable to generate risk premiums of the order of magnitude required to explain fluctuations in asset prices, not only in exchange markets, but in almost all financial markets. It has also proved difficult to relate ex post exchange risk premia to macroeconomic factors. A central question, however, is whether the models underlying these approaches are a useful starting point for characterizing the objective function and actual behavior of market participants.

The 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 - η_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 (1994a) find that there is some evidence of reversion to PPP 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

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

hypothesized value of unity than to zero. These results confirm the earlier conjectures of Mussa (1979) and Froot (1990b) 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 increase in the risk premium 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. Over longer horizons, the temporary effects of risk premium shocks fade and the model results are dominated by other dynamics that are consistent with the UIP hypothesis.

The results suggest, then, that risk premium shocks are capable of explaining the main stylized facts regarding UIP, although the model 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.

Nevertheless, to the extent that risk premium shocks are the main source of deviations from UIP, this has important implications for economic modeling at longer horizons. In particular, it would validate the use of open-interest parity conditions to characterize expected exchange rate movements in structural macroeconomic models. One would need, however, to take a stand on the macroeconomic determinants of the exchange risk premium, and carefully consider the nature of the shocks likely to perturb the economy. 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 Tests of Uncovered Interest Parity

Regression equation: $\Delta s_{t+k} = \alpha + \beta\,(i - i^*)_t + \varepsilon_{t+k}$

Sample period: 1980Q1 – 1998Q1 (71 degrees of freedom)

Panel 1.a: 3-month Eurocurrency Yields
(standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject null of: $\beta = 0$	$\beta = 1$	$\bar{R}^2$	D.W.
Japanese yen	-0.133***	-3.440	***	***	0.101	2.11
	(0.040)	(0.998)				
German deutschemark	-0.007	-0.646		*	-0.008	1.83
	(0.037)	(1.160)				
U.K. pound	0.057*	-2.036	*	***	0.034	1.87
	(0.032)	(1.211)				
French franc	0.018	0.273			-0.013	1.68
	(0.036)	(1.010)				
Italian lira	-0.027	1.462	*		0.022	1.53
	(0.051)	(0.858)				
Canadian dollar	0.016	-0.516		***	-0.005	2.12
	(0.013)	(0.589)				
Constrained panel[1]	...	-0.507		***	-0.000	...
		(0.405)				

1 Sample period: 1980Q1-1997Q4.
* Different from null at 10 percent significance level.
** Different from null at 5 percent significance level.
*** Different from null at 1 percent significance level.

Table 1. (continued)

Panel 1.b: 6-month Eurocurrency Yields
(MA(1)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject null of: $\beta = 0$	$\beta = 1$	$\bar{R}^2$	D.W.
Japanese yen	-0.134*** (0.032)	-3.387 (0.783)	***	***	0.168	1.32
German deutschemark	-0.012 (0.030)	-0.861 (0.839)		**	0.005	1.03
U.K. pound	0.053 (0.030)	-1.963 (1.176)	*	***	0.058	1.15
French franc	0.017 (0.034)	0.269 (0.920)			-0.012	0.91
Italian lira	-0.046 (0.046)	1.847 (0.795)	***		0.022	0.92
Canadian dollar	0.017* (0.008)	-0.535 (0.391)		***	0.005	1.15
Constrained panel[1]	...	-0.568 (0.390)		***	0.011	...

[1] Sample period: 1980Q1-1997Q4.
* Different from null at 10 percent significance level.
** Different from null at 5 percent significance level.
*** Different from null at 1 percent significance level.

Table 1. (continued)

Panel 1.c: 12-month Eurocurrency Yields
(MA(3)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject null of: $\beta = 0$	$\beta = 1$	$\bar{R}^2$	D.W.
Japanese yen	-0.126*** (0.030)	-2.996 (0.706)	***	***	0.230	0.56
German deutschemark	-0.011 (0.027)	-0.581 (0.682)		***	0.001	0.38
U.K. pound	0.037 (0.026)	-1.268 (1.052)		**	0.041	0.53
French franc	0.011 (0.035)	0.491 (0.889)			-0.006	0.30
Italian lira	-0.055* (0.038)	1.994 (0.647)	***		0.061	0.32
Canadian dollar	0.015* (0.008)	-0.464 (0.477)		***	0.006	0.47
Constrained panel	...	-0.321 (0.423)		***	0.022	...

[1] Sample period: 1980Q1-1997Q4.
* Different from null at 10 percent significance level.
** Different from null at 5 percent significance level.
*** Different from null at 1 percent significance level.

Table 2. Short-Horizon Tests of UIP Over Different Sample Periods
(Estimated value of β in equation (7))

	1980Q1–1988Q4			1989Q1–1998Q1		
	3-month	6-month	12-month	3-month	6-month	12-month
Japanese yen	-4.52	-4.45	-4.15	-3.86	-3.97	-3.15
German deutschemark	-3.15	-4.82	-4.01	-0.54	-0.49	-0.47
U.K. pound	-5.10	-5.33	-4.07	1.06	1.02	1.15
French franc	0.21	0.42	0.76	0.20	-0.17	-0.11
Italian lira	0.89	1.50	1.98	2.60	2.36	2.08
Canadian dollar	-0.45	-1.06	-2.12	-0.69	-0.52	-0.01
Constrained panel	-1.28	-1.39	-0.82	-0.18	-0.30	-0.33

Table 3. Long-Horizon Tests of Uncovered Interest Parity

Regression equation: $\Delta s_{t+k} = \alpha + \beta\,(i - i^{*})_t + \varepsilon_{t+k}$

Sample period: 1983Q1 – 1998Q1 (59 degrees of freedom)

Panel 3.a: Benchmark Government Bond Yields, 10-Year Maturity
(MA(39)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject null of: $\beta = 0$	$\beta = 1$	$\bar{R}^2$	D.W.
Japanese yen	-0.039*** (0.004)	0.487 (0.101)	***	***	0.197	0.18
German deutschemark	-0.008*** (0.002)	0.829 (0.147)	***		0.424	0.23
U.K. pound	0.005 (0.004)	0.567 (0.104)	***	***	0.447	0.21
French franc	-0.007 (0.020)	0.885 (0.508)	*		0.027	0.07
Italian lira[1]	0.008 (0.007)	0.214 (0.155)	*	***	0.009	0.11
Canadian dollar	0.000 (0.007)	1.104 (0.657)	*		0.145	0.17
Constrained panel[2]	...	0.635 (0.111)	***	***	0.654	...

1 Sample period for the Italian lira limited to 1987Q1–1998Q1 (43 degrees of freedom).
2 Excluding the lira; sample period 1985Q1 - 1997Q4.
* Different from null at 10 percent significance level.
** Different from null at 5 percent significance level.
*** Different from null at 1 percent significance level.

Table 3. continued

Sample period: 1983Q1 – 1998Q1 (59 degrees of freedom)

Panel 3.b: 10-Year Government Bond Yields
(MA(39)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject null of: $\beta = 0$	$\beta = 1$	$\bar{R}^2$	D.W.
Japanese yen	-0.038*** (0.005)	0.564 (0.146)	***	**	0.213	0.18
German deutschemark	-0.009*** (0.002)	0.836 (0.128)	***	*	0.468	0.23
U.K. pound	-0.003 (0.004)	0.719 (0.109)	***	***	0.446	0.32
Constrained panel [1]	...	0.708 (0.090)			0.695	...

[1] Sample period: 1985Q1-1997Q4.
* Different from null at 10 percent significance level.
** Different from null at 5 percent significance level.
*** Different from null at 1 percent significance level.

Table 3. (continued)

Sample period: 1983Q1 – 1998Q1 (59 degrees of freedom)

Panel 3.c: 5-Year Government Bond Yields
(MA(19)-adjusted standard errors in parentheses)

	$\hat{\alpha}$	$\hat{\beta}$	Reject null of: $\beta = 0$	Reject null of: $\beta = 1$	$\bar{R}^2$	D.W.
German deutschemark	0.002 (0.016)	0.914 (0.360)	***		0.066	0.12
U.K. pound	0.002 (0.010)	1.084 (0.415)	***		0.011	0.23
Canadian dollar[1]	0.008 (0.007)	1.337 (0.410)	***		0.157	0.18
Constrained panel[2]	...	1.010 (0.344)	***		0.396	...

1 Sample period for Canada limited to 1985Q4–1998Q1 (44 degrees of freedom).
2 Sample period 1986Q1 - 1997Q4.
* Different from null at 10 percent significance level.
** Different from null at 5 percent significance level.
*** Different from null at 1 percent significance level.

Table 4. Simulation Model

Uncovered interest parity:

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

Monetary reaction function:

$$i_t - \pi_t = 0.5\,(\pi_t + y_t)$$

Inflation (π) equation:

$$\pi_t = 0.6\pi_{t-1} + (1-0.6)\pi^{e}_{t,t+1} + 0.25y_t + 0.1\Delta(s_t - p_t) + \nu_t$$

Output (y) equation:

$$y_t = 0.1(s_t - p_t) - 0.5(i^{l,e}_{t} - \pi^{l,e}_{t}) + 0.5\,y_{t-1} + \varepsilon_t$$

Price level (p) identity:

$$p_t = p_{t-1} + \pi_t$$

Exchange rate (s) identity:

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

Long-term expected interest rate:

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

Long-term expected inflation rate:

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

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

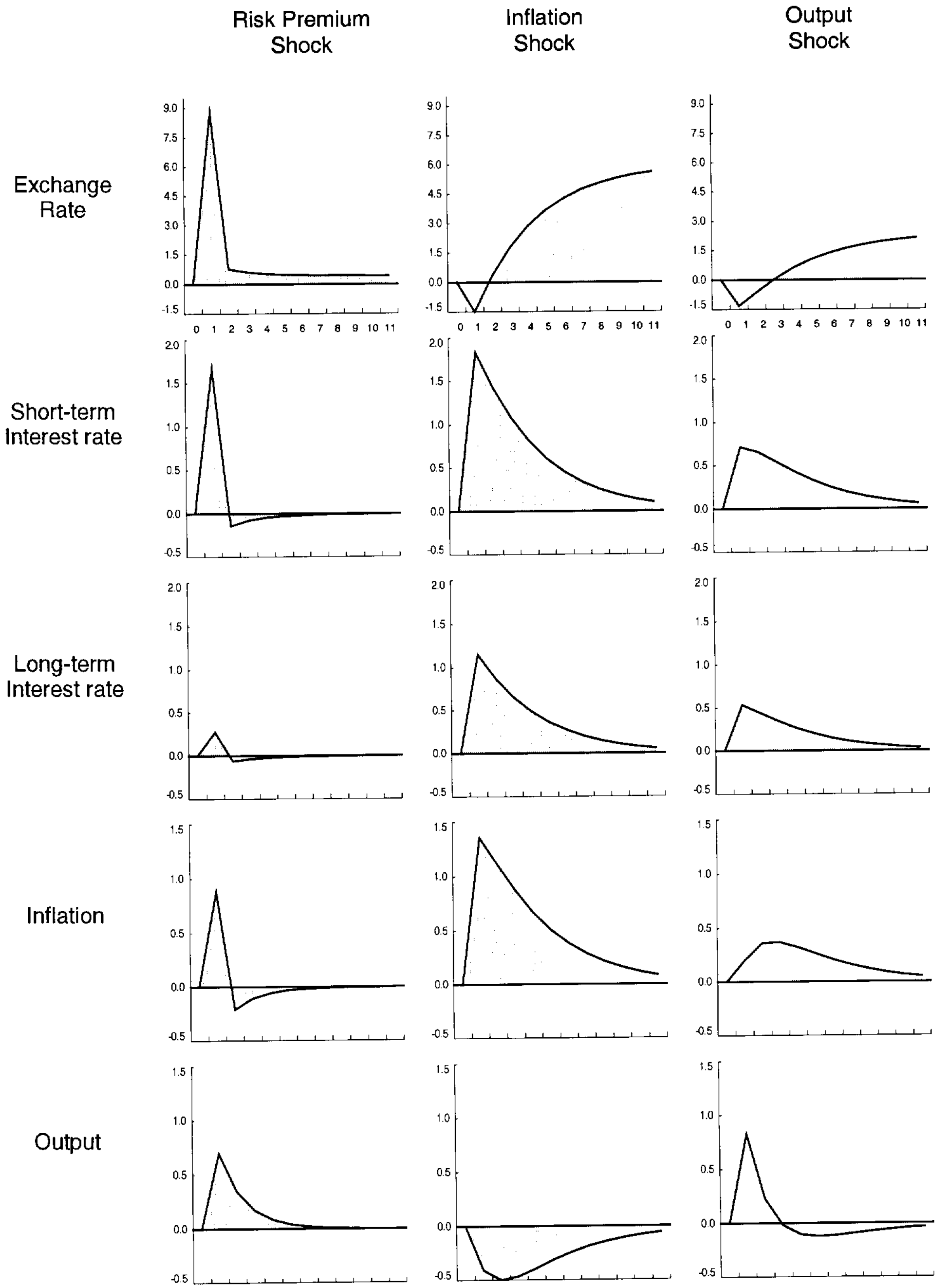

Figure 2. Estimated Slope Parameters from UIP Regressions at Different Horizons Using Data From Stochastic Simulations