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A RECONSIDERATION OF THE UNCOVERED INTEREST PARITY RELATIONSHIP

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ABSTRACT

The paper first presents reasons for viewing the uncovered interest-parity (UIP) relationship as more important, in terms of economic analysis, than the unbiasedness of forward rates as predictors of future spot exchange rates. The two hypotheses are closely related, so that test rejections of the latter tend to cast doubt on the former, but are not identical--so unbiasedness rejections are not conclusive for UIP.

Next, some representative evidence is presented that pertains to alternative versions of the unbiasedness test. Although $s_t = \alpha + \beta f_{t-1} + \varepsilon_t$ and $s_t - s_{t-1} = \alpha + \beta(f_{t-1} - s_{t-1}) + \varepsilon_t$ are equivalent under the null hypothesis of $\beta = 1.0$, they represent different classes of alternative hypotheses. Empirically, they give rise to extremely different outcomes, estimates of β being very close to 1.0 in the former equation but in the vicinity of -3.0 in the latter. In a generalized specification that includes both as special cases, the results strongly favor the second specification--thereby rejecting unbiasedness.

Finally, three possible explanations for the $\beta = -3$ result are considered and related to the UIP condition. Of these three, the latter two--one involving systematically irrational expectations and the other an additional relationship reflecting monetary policy behavior--are consistent with UIP. The policy-response hypothesis, that monetary authorities manage interest-rate differentials so as to resist rapid changes in exchange rates and in these differentials, is attractive conceptually and is capable of explaining not only the $\beta = -3$ finding, but also several other notable features of the data.

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I. Introduction

One of the most extensively studied topics in any area of economics over the past decade has been the efficiency--or, more precisely, the apparent inefficiency--of the forward market for foreign exchange. In particular, an enormous number of studies and surveys have considered the putative failure of forward exchange rate premia to serve, during the floating rate era, as unbiased predictors of future changes in spot rates.[1] With the condition of covered interest parity obtaining, as it must to avoid riskless arbitrage possibilities, such failure would be very closely related to the violation of uncovered interest parity, but the latter has received much less explicit attention in the literature.[2] It is, however, the contention of the present paper that this relative emphasis is unwarranted. Indeed, it would appear that the uncovered interest parity (UIP) relation is both more enlightening and more important from the perspective of economic analysis than the question of foreign exchange market unbiasedness. Consequently, an attempt is made in the present paper to reexamine outstanding issues and evidence from the perspective of the UIP relation. The reexamination indicates that the evidence is indeed inconsistent with forward-rate unbiasedness, despite some inaccurate arguments to that effect in the literature,[3] but that this evidence does not entail rejection of UIP. Indeed, the most convincing explanation for the failure of unbiasedness to hold is consistent with UIP.

Organizationally, our treatment begins in Section II with a brief review of the main concepts at hand and a discussion of the role of UIP in exchange rate analysis. Next, Section III presents a number of empirical regularities bearing on the apparent failure of UIP and on the extent of that failure. Then in Section IV three alternative explanations for the anomalies in the data are outlined and their relative merits explored. The most promising of these explanations relies strongly on the hypothesis that the monetary policy authorities systematically manage interest rate differentials so as to resist rapid changes in exchange rates but also to smooth interest-rate movements. This form of behavior implies the presence of a second relationship, prevailing together with UIP as parts of a simultaneous system, that has been critically neglected in most of the relevant empirical analysis. Recognition of the postulated policy equation yields a model that is capable of rationalizing not only the unbiasedness evidence but also several other notable features of the data pertaining to spot exchange rates and forward discounts. The development of this model is, accordingly, a major feature of the paper. A brief conclusion to the latter is provided, finally, in Section V.

II. Concepts and Preliminaries

As stated above, numerous writers have documented and emphasized the apparent failure, over the floating-rate years since 1973, of forward exchange rates to serve as "unbiased" predictors of future spot rates. The most extensive survey of existing evidence has been provided in a monograph by Hodrick (1987), who finds "a major conclusion... is that very strong evidence exits against the hypothesis that forward exchange rates... are unbiased predictors of future spot rates" (1987, p.4). In a similar vein, Meese (1989, p.165) reports that tests by a variety of authors "have established beyond all reasonable doubt that the composite unbiased forward rate hypothesis is inconsistent with the data" while MacDonald (1988, p.197) concludes that "the broad thrust of the research reported above would seem to suggest an overwhelming rejection of the [unbiasedness hypothesis] as applied to the forward market for foreign exchange."

For the discussion of these concepts and others related to UIP, we will need some notation. Accordingly, let s_t denote the log of the spot exchange rate expressed as the price, in "home-country" monetary units, of foreign exchange--either a single foreign currency or some weighted-average bundle. Similarly, let f_t denote the log of the one-period forward rate, i.e., the home currency price in period t of a unit of foreign exchange to be paid for and delivered in period t + 1.[4] Then f_t might be said to be an unbiased predictor of s_{t+1} if $\alpha = 0.0$ and $\beta = 1.0$ in the relation

(1) $\alpha + \beta f_t = E_t s_{t+1}$,

where $E_t s_{t+1} \equiv E(s_{t+1}|\Omega_t)$ is the conditional expectation of s_{t+1} formed on the basis of the information set Ω_t available at time t. For reasons explained by Hodrick (1987, p.28) and Meese (1989, p.162), however, it is reasonable to permit a non-zero value of α in (1) and express the unbiasedness hypothesis as $\beta = 1.0$.

With rational expectations, it is of course true that the expectational error $\varepsilon_{t+1} = s_{t+1} - E_t s_{t+1}$ will be uncorrelated in the population with all elements of Ω_t, in which case equation (1) can be rewritten as

(2) $s_{t+1} = \alpha + \beta f_t + \varepsilon_{t+1}$,

a form that is suitable for consistent estimation by means of OLS (ordinary least squares) if ε_t is white noise or by other standard techniques if ε_t features heteroschedasticity.[5] Such a form was, indeed, utilized in early research on the subject by Frenkel (1976)(1981), Levich (1978), and others.

More recently, however, most analysts have not utilized (2) in conducting unbiasedness tests but have instead relied on the relation

(3) $s_{t+1} - s_t = \alpha + \beta(f_t - s_t) + \varepsilon_{t+1}$,

which is equivalent to (2) under the tested hypothesis $\beta = 1.0$.[6] As it happens, empirical tests with post-1973 data tend to reject $\beta = 1.0$ rather decidedly in formulation (3) but to support $\beta = 1.0$ in (2), a point that will be taken up in detail below.

The issue of whether $\beta = 1.0$ in (2) and (3) is admittedly of considerable interest. But to the present writer that issue seems to have been somewhat overemphasized, since it is not the case that rejections of $\beta = 1.0$ strictly imply that the foreign exchange market is functioning inefficiently in the sense of Pareto optimality. If market participants are averse to risk, then a time-varying risk-premium term would appear in equations (1) - (3), and could be correlated with ε_{t+1} so as to eliminate the presumption that $\beta = 1.0$ should be found in regressions of the form (2) and (3). Furthermore, even if some departure from Pareto optimality were established, it would not follow automatically that government intervention in the exchange market would be appropriate; governmental behavior often features notable inefficiencies of its own.

Arguably of greater importance from the perspective of economic analysis of exchange rates is the closely related but distinct condition of uncovered interest parity (UIP), which may be expressed as

(4) $$R_t - R_t^* = s_{t+1}^e - s_t + \xi_t.$$

Here the idea is that home and foreign interest rates (R_t and R_t^*) on similar one-period loans will differ systematically only to the extent of expected depreciation in the relative value of the home currency, $s_{t+1}^e - s_t$, where s_{t+1}^e is the value of s_{t+1} expected as of period t. Unsystematic--i.e., random--sources of discrepancy between R_t and R_t^* are represented by the disturbance term ξ_t, which might represent time-varying aggregation or other effects, as well as risk premia.[7] Some writers have adopted a terminology according to which UIP is said not to hold if a term like ξ_t is included in the relation. That convention seems undesirable, however, since it would make rejections of UIP rather uninteresting and since relation (4) with ξ_t included is the version that appears in the leading econometric models. Consequently, ξ_t will be included in the analysis that follows.

Before considering the analytical importance of (4), let us note its relationship to (3) by recalling that the absence of riskless arbitrage possibilities implies that, when transaction costs are neglected,

(5) $$R_t - R_t^* = f_t - s_t.$$

There exists a very strong theoretical presumption that this covered interest-parity relationship must hold, for if (say) the left-hand side exceeded the right, then a market participant could trade foreign for home-country currency on the spot market and sell the proceeds forward, thereby earning additional interest over the subsequent period without incurring any additional risk. Thus condition (5) should be expected to prevail and, in fact, existing studies have tended to verify that observed discrepancies are bounded by transaction costs.[8]

Substitution of (5) into (4) and rearrangement yields, however,

(6) $$s^e_{t+1} - s_t = f_t - s_t - \xi_t$$

which is almost the same as (3) with $\alpha = 0.0$ and $\beta = 1.0$. If expectations are rational, so that $s^e_{t+1} = E_t s_{t+1}$, then (6) and (3) will coincide if $\alpha = 0.0$, $\beta = 1.0$, and $\xi_t = -\varepsilon_{t+1}$. Consequently, rejection of the unbiasedness hypothesis serves, to a substantial extent, to reflect discredit onto the UIP condition. That the hypotheses are not identical is, nevertheless, demonstrated by consideration of the possibility that expectations are not rational.[9]

Turning then to the analytical importance of the UIP condition, the main fact to be kept in mind is that it appears as a key behavioral relationship in virtually all of the prominent current-day models of exchange rate determination. These include not only small models used in theoretical analysis, but also a number of the more ambitious and carefully specified of today's array of multicountry econometric models--those used by international organizations as well as individual open-economy policy analysts. Among recently constructed systems that incorporate both UIP and rational expectations are the prominent models of Taylor (1989)(1990), McKibbon and Sachs (1989), and the IMF's MULTIMOD (Masson, Symansky, and Meredith, 1990). With these systems an analyst could in principle substitute some other expectational hypothesis for full rationality; indeed simulation experiments with static or adaptive expectations have occasionally been conducted by the models' maintainers. But without UIP (or some portfolio-theory modification[10]), the systems are simply incomplete and therefore unusable.

It is of course logically possible that formal statistical rejections might be found for the highly specific hypothesis that the crucial response coefficient β equals 1.0 exactly, but with estimated values nevertheless falling reasonably close to 1.0 in terms of economically meaningful magnitudes. In such a case one could easily reconcile the divergent points of view implicit in existing evidence and current modeling practice. But that is not the situation that actually prevails. Instead, according to Froot and Thaler (1990, p. 182), "the average coefficient across some 75 published estimates is -0.88.... A few are positive but not one is equal to or greater than the null hypothesis of $\beta = 1$." Thus the bulk of the evidence indicates not just that exchange rate changes fail to move one-for-one with interest differentials (and forward premia), but rather that these changes are substantial and in the opposite direction to that implied by UIP. Taken at face value, that contradiction would seem to suggest that models incorporating UIP--i.e., most current models--are misspecified to a truly drastic extent.

The problem under discussion can be highlighted by temporarily supposing that the monetary authorities at home and abroad use the short-term interest rates R_t and R_t^* as their policy instruments--a supposition that has a basis in actual practice--and also that they conduct policy in a manner that makes the interest differential, $x_t = R_t - R_t^*$, exogenous. (The latter supposition is not realistic.) Then a two-equation system consisting of the policy rule plus a generalized version of (4) with rational expectations, i.e.,

(7) $$s_t = E_t s_{t+1} - \beta x_t + \xi_t,$$

will ostensibly constitute a complete model of spot rate determination.[11] And if the true value of β is negative, rather than 1.0, spot rates will move in response to x_t shocks in the opposite direction from that normally presumed. An unexpected, policy-induced increase in the home country's interest rate will, that is, reduce rather than enhance the exchange value of the home-country currency. Such a response would conflict not only with the beliefs of model builders, but also practical men including actual policymakers.

The term "ostensibly" must be included in the statement of the preceding paragraph, incidentally, because with an exogenous policy process for x_t, the two-equation system under discussion would feature indeterminacy (not multiplicity) of nominal variables. As in the case of a single closed economy, extensively analyzed in McCallum (1981)(1986), nominal determinacy requires that the monetary policy authority utilize an interest rate rule that reflects some concern for the path of some nominal variable such as the price level, money stock, or exchange rate.

A few additional insights can be gained by examination of a slightly extended, and more orthodox, model of spot rate determination for a small economy. Such a model would consist of (7) plus the following four relationships:

(8) $$m_t - p_t = c_o + c_1 y_t + c_2 R_t + v_{8t}$$

(9) $$y_t = b_o + b_1 r_t + b_2 q_t + b_3 g_t^* + v_{9t}$$

(10) $$q_t = s_t - p_t + p_t^*$$

(11) $$r_t = R_t - (E_t p_{t+1} - p_t)$$

Here (8) is a home-country money demand function, that could without major effect be revised to make a weighted average of p_t and $p_t^* + s_t$ the relevant price level index, while (9) is an open-economy IS (or saving-investment) relation in which the real interest rate and real exchange rate are critical determinants. These last two variables are defined by (10) and (11). Then with a policy rule specifying behavior of m_t and with R_t^*, p_t^*, y_t^* given from abroad, the relations (7) - (11) plus $x_t = R_t - R_t^*$ serve to determine the ongoing behavior of p_t, R_t, s_t, q_t, r_t, and x_t if price level flexibility is assumed so that y_t is continually equal to its full-employment value $\bar{y}_t$. With sticky prices assumed instead, some specification of price adjustment behavior would need to be appended. Models with structures of these two types are, of course, representative of classical and Keynesian orthodoxy, respectively.

The three points implied by such structures that are germane to our current discussion are these. First, the forward rate f_t appears nowhere in the system of relations just outlined. That fact serves to support the notion advanced above that the UIP relation is of much greater importance for open-economy macroeconomics than the forward-rate unbiasedness issue itself. Second, it remains true that the exchange-rate response to a monetary policy shock would be very different if the coefficient β were negative rather than equal to the UIP value of 1.0. Third, these considerations would continue to be of relevance and importance in much the same way if the model's assumption of rational expectations were replaced with some other specific hypothesis about expectation formation.

It should be added that our emphasis on the conceptual distinction between the UIP and unbiasedness hypotheses does not imply that issues concerning the former must be empirically implemented with interest-rate data rather than forward premia. If covered interest parity holds, as is widely believed to be the case, then the variable $R_t - R_t^*$ and $f_t - s_t$ should be so highly correlated that they could be used interchangeably, with the choice made on the basis of data availability and convenience. In fact, for those reasons the empirical work reported below will actually be based on measurements on $f_t - s_t$ even though it is the relationship between spot-rate movements and $R_t - R_t^*$ that is of primary interest.

III. Empirical Regularites

In this section the object will be to outline several empirical regularities, a few of which may be unfamiliar even to specialists in exchange-rate behavior, that serve to delineate the issues and puzzles referred to above. The measurements that will be utilized throughout our tables and discussion are \$/DM, \$/£, and \$/Yen rates taken from the data base of the Bank for International Settlements (BIS). Logarithms of spot rates and 30-day forward rates, both measured on the final day of each month, are denoted s_t and f_t, respectively.[12] In most cases the sample period extends from January 1978 through July 1990 (denoted 1978.01 - 1990.07). The raw data series are reported in Appendix A.

The first and perhaps most striking (although familiar) of the empirical regularities relating to the issues at hand is provided by the sharp contrast in estimates of the forward-rate "unbiasedness" coefficient β that result from the two specifications (2) and (3), i.e., from the equations $s_t = \alpha + \beta f_{t-1} + \varepsilon_t$ and $s_t - s_{t-1} = \alpha + \beta(f_{t-1} - s_{t-1}) + \varepsilon_t$. For the first of these, OLS estimates of β are very close to the value of 1.0 that represents the unbiasedness hypothesis. Only one of the three estimates, reported in the first panel of Table 1, departs from 1.0 by more than one (estimated) standard error. These standard errors, moreover, are all smaller than 0.02. When, however, one uses instead $s_t - s_{t-1}$ and $f_{t-1} - s_{t-1}$ as the dependent and regressor variables, which would amount to the same specification as before if $\beta = 1.0$ held exactly, the results are entirely different. In particular, the estimated β values range from -3.3 to -4.7, figures that are significantly negative even with the much larger standard errors that obtain; see the second panel of Table 1. This discrepancy, between estimates close to 1.0 and others around -3.0, was first noted by Tryon (1979) and Longworth (1981) and has since been mentioned by many writers. What conclusions are appropriate?

Table 1

OLS Regression Estimates, 1978.01 - 1990.07

Exchange Rate	Variables	Estimates (std. errors) Const.	Slope	R^2	Statistics SE	DW
$/DM	s_t on f_{t-1}	-0.0092 (0.012)	0.9896 (0.016)	0.963	0.0362	2.05
$/£	" "	-0.0137 (0.009)	0.9770 (0.016)	0.960	0.0359	1.82
$/Yen	" "	-0.0464 (0.068)	0.9913 (0.013)	0.975	0.0380	1.84
$/DM	s_t-s_{t-1} on $f_{t-1}-s_{t-1}$	-0.0161 (0.006)	-4.3030 (1.70)	0.041	0.0351	2.19
$/£	" "	-0.0078 (0.0032)	-4.7403 (1.095)	0.111	0.0332	2.21
$/Yen	" "	0.0153 (0.0052)	-3.3265 (1.173)	0.051	0.0364	2.02
$/DM	s_t-s_{t-2} on $f_{t-1}-s_{t-2}$	-0.0012 (0.003)	0.9398 (0.083)	0.461	0.0362	1.93
$/£	" "	0.0015 (0.003)	1.0231 (0.086)	0.486	0.0362	1.89
$/Yen	" "	-0.0007 (0.003)	1.0372 (0.085)	0.502	0.0380	1.92
$/DM	s_t-s_{t-3} on $f_{t-1}-s_{t-3}$	-0.0019 (0.003)	1.0539 (0.061)	0.670	0.0362	2.18
$/£	" "	-0.0015 (0.003)	1.0450 (0.060)	0.667	0.0361	1.93
$/Yen	" "	-0.0009 (0.003)	1.0513 (0.058)	0.684	0.0380	1.96

An answer that appears with some frequency in the literature (e.g., MacDonald (1988, p.181), Meese and Singleton (1982, p.1034)) is that since s_t and f_t give evidence of being generated by nonstationary time-series processes--see Meese and Singleton (1982)--specification (2) is inappropriate for OLS analysis so results based on (3) should be the more reliable.

As thus stated, however, this answer is unsatisfactory. For, as Hodrick (1987, pp.28-29) and Meese and Singleton (1982, p.1030) have recognized, the presumed nonstationarity of s_t and f_t does not destroy consistency of the OLS estimator of β provided by relation (2). Inapplicability of the standard sampling theory is a consequence of nonstationarity, of course, so tests of hypotheses such as $\beta = 1.0$ that are based on the OLS standard errors cannot be considered reliable. But this in itself does not imply that the point estimates from (2) are misleading, nor does it rationalize the difference between estimates obtained from (2) and (3).

To emphasize this last point, notice that if the object of subtracting s_{t-1} from the variables on both sides of equation (2) is merely to generate stationary variables, it should serve just as well to subtract s_{t-2} or s_{t-j} for any $j = 1,2,\ldots$. But when estimates are obtained for the resulting specifications with $j = 2,3,\ldots$, the results revert to ones qualitatively similar to those for specification (2) and quite unlike those for $j = 1$. This phenomenon is documented for $j = 2$ and 3 in the third and fourth panels of Table 1, where slope estimates attached to $f_{t-1} - s_{t-j}$ (for $j = 2$ and 3) are all within one standard error of 1.0.[13]

It might also be mentioned that (3) does not represent, as is occasionally suggested in the literature, a differenced version of (2). Results for such a version, with Δs_t regressed on Δf_{t-1}, are shown in the first panel of Table 2. As will readily be seen, the slope coefficients are

Table 2

Additional OLS Regression Estimates, 1978.01 - 1990.07

Exchange Rate	Variables	Estimates (std. errors) Const.	Slope	R^2	Statistics SE	DW
\$/DM	Δs_t on Δf_{t-1}	0.0019 (0.0029)	-0.0632 (0.0816)	0.0040	0.0358	1.96
\$/£	" "	-0.0002 (0.0029)	0.0244 (0.0823)	0.0006	0.0352	1.99
\$/Yen	" "	-0.0031 (0.0030)	0.0381 (0.0821)	0.0014	0.0374	1.99
\$/DM	Δs_t on Δf_t	0.0000 (0.000)	1.0017 (1.0024)	0.9992	0.00104	1.60
\$/£	" "	0.0000 (0.0001)	1.00143 (0.0027)	0.9988	0.00118	1.42
\$/Yen	" "	0.0000 (0.0001)	1.0006 (0.0025)	0.9991	0.0011	1.65
\$/DM	s_t on f_t	-0.0027 (0.0006)	1.0008 (0.0007)	0.9999	0.00170	0.37
\$/£	" "	0.0029 (0.0006)	0.9977 (0.0112)	0.9998	0.00247	0.23
\$/Yen	" "	0.0070 (0.0045)	1.0020 (0.0008)	0.9999	0.00251	0.20

insignificantly different from zero for all three currencies.

The difference between estimates obtained from (2) and (3) will be discussed at greater length below. First, however, it is appropriate to emphasize that, although the coefficients on the lagged forward discount $f_{t-1} - s_{t-1}$ in (3) are "statistically significant" at conventional levels, the $f_{t-1} - s_{t-1}$ variable nevertheless contributes extremely little in terms of predictive or explanatory power. One direct indication of this is provided by the R^2 values given in Table 1, but it is arguably more revealing to examine simple scatter diagrams relating Δs_t to $f_{t-1} - s_{t-1}$. Such diagrams are included, consequently, as three parts of Figure 1. In each of these diagrams, the general impression is clearly of the weakness, rather than the strength, of any underlying relationship.

If the $f_{t-1} - s_{t-1}$ terms in (3) were not "significant," it would then trivially be true that the reported process for s_t was insignificantly different from a random walk--possibly one with drift. That suggests that one consider whether low-order autoregressive or moving-average terms are useful in modeling the univariate s_t processes. That such terms are not useful--that the univariate s_t processes are indeed close to random walks for all three of the currencies examined--is shown by the values reported in Table 3. There it will readily be seen that the residual standard-error magnitudes for the random-walk specification (located in row 1) are little different from those in panel 2 of Table 1. In addition, Table 3 indicates rather clearly that one cannot formally reject the random-walk specification in favor of other ARMA models with one or two parameters.

Next let us briefly consider the univariate time-series properties of $f_t - s_t$, the forward discount. From estimates of ARMA models with the seven specifications listed in Table 4, one gains the impression that an ARMA (1,0,1) specification accommodates the data reasonably well. It is true that

Figure 1 -- Scatter Diagrams for Equation (3)

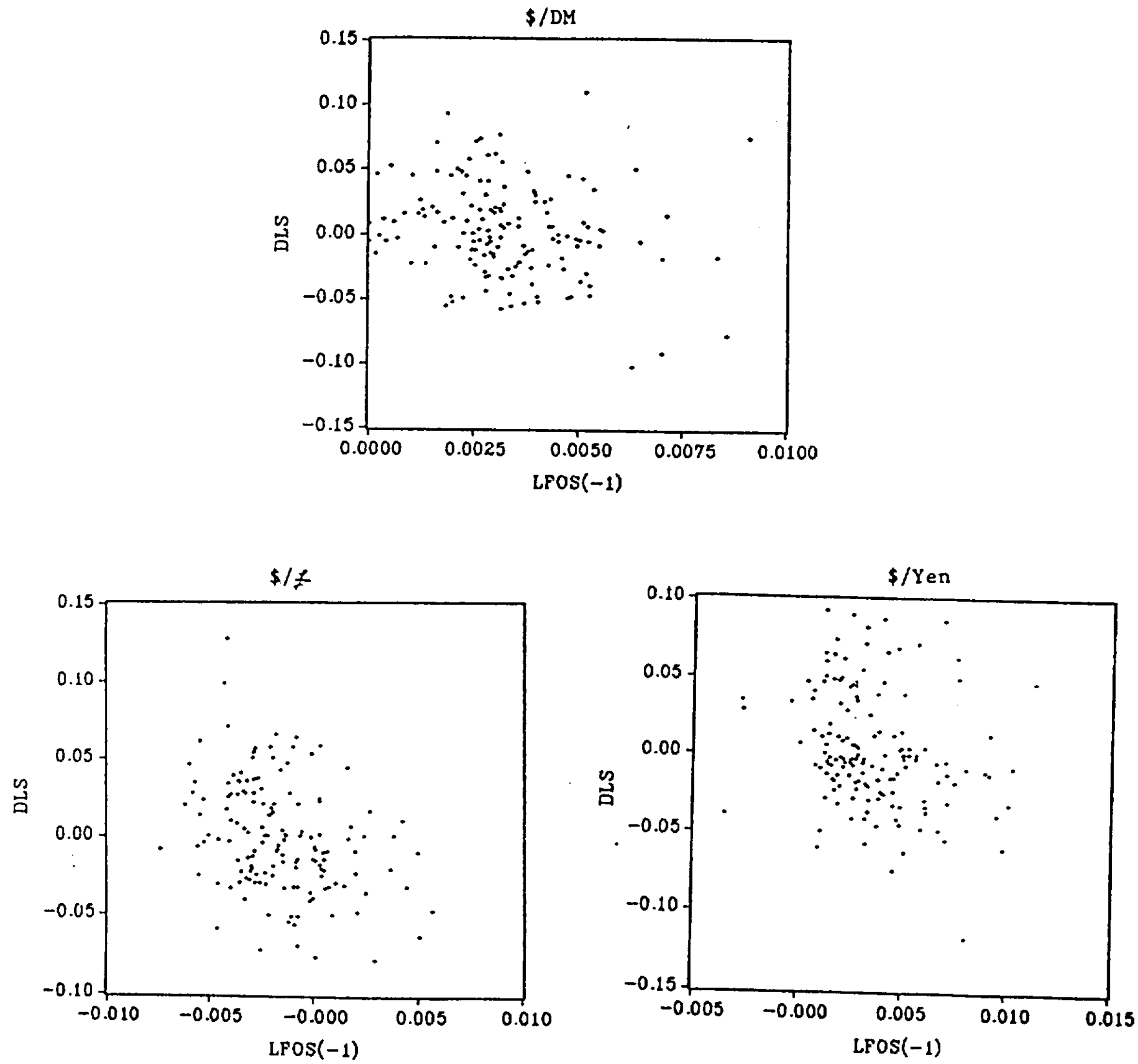

Table 3

Summary Statistics for Various ARMA Models of Spot-Rate Logarithms, 1978.01 - 1990.07

Model specification for s_t (p,d,q)	SE, DW Statistics for: $/DM	$/£	$/Yen
(0,1,0)	0.0357, 2.10	0.0351, 1.94	0.0373, 1.91
(1,1,0)	0.0358, 1.97	0.0352, 1.99	0.0374, 2.00
(2,1,0)	0.0355, 2.00	0.0353, 1.99	0.0375, 2.00
(0,1,1)	0.0358, 2.00	0.0352, 1.97	0.0374, 1.99
(0,1,2)	0.0356, 2.00	0.0353, 1.98	0.0375, 1.99
(1,1,1)	0.0358, 1.96	0.0353, 1.99	0.0375, 2.00
(1,0,0)	0.0358, 2.09	0.0350, 1.92	0.0373, 1.90
(2,0,0)	0.0359, 1.97	0.0374, 2.00	0.0351, 1.99

Note: In the 24 models there are no parameter estimates that are as large as twice their standard error. (In the final two specifications, this statement applies to the AR1 parameter relative to 1.0, not 0.0, and does not apply to the constant term.) In the first column, p and q denote the number of autoregressive and moving-average terms while d is the number of times that s_t has been differenced.

the AR parameters are large and the MA parameters small, so that the (1,0,1) processes are not drastically far from random walks.[14] But the overall impression is definitely that of a stationary process. The other noteworthy feature of the estimated models is that the residual variability is much smaller than for the Δs_t process. Estimated disturbance variances for the (1,0,1) models are, in fact, only about 1/1000 as large as for Δs_t (a ratio that is about 1/32 in terms of standard deviations).

That the variability of $f_t - s_t$ is itself small relative to that of Δs_t is well-known from time plots of these two variables presented by Frenkel and Mussa (1981), Copeland (1989, p. 290), and many others. Also well-known from frequently-presented graphs is that s_t and f_t track each other very closely--their time plots can scarcely be distinguished on a dot-matrix printout. It is perhaps less well-known, however, that Δs_t and Δf_t are also very highly correlated. In fact, the residual variance is actually smaller for a Δs_t on Δf_t regression than for one with s_t and f_t as variables, as is documented in the second and third panels of Table 2. Precisely what to make of this fact is unclear, except for the apparent suggestion that s_t and f_t typically move in response to the <u>same</u> shocks, whatever those might be.

Let us now return to the topic that began this section, namely, the discrepancy in β estimates from equations (2) and (3). Since one of our main objectives is to determine the general magnitude of the slope coefficient in the UIP relationship, and since covered interest parity is being assumed to hold, it is of importance to know which of the two indicated magnitudes, roughly +1.0 or -3.0, represents the more reliable estimate of the unbiasedness parameter. And it was mentioned above that nonstationarity of s_t and f_t would not necessarily rule out the former.

Table 4

Summary Statistics for Various ARMA Models of Forward Discount
1978.01 - 1990.07

Model specification for $f_t - s_t$ (p,d,q,)	SE, DW Statistics for: $/DM	$/£	$/Yen
(1,0,0)	0.00099, 1.46	0.00115, 1.83	0.00110, 1.58
(0,0,1)	0.00112, 1.17	0.00183, 1.93	0.00168, 0.88
(1,0,1)	0.00093, 2.03	0.00115, 1.97	0.00108, 1.94
(2,0,0)	0.00095, 1.87	0.00115, 2.01	0.00108, 1.99
(0,1,0)	0.00104, 1.59	0.00117, 1.93	0.00112, 1.65
(1,1,0)	0.00102, 1.85	0.00118, 2.00	0.00111, 1.95
(0,1,1)	0.00100, 2.14	0.00117, 1.99	0.00111, 2.01
(1,1,1)	0.00099, 2.00	0.00115, 2.01	0.00111, 1.96

Note: In the first-column, p and q denote the number of autoregressive and moving-average terms while d is the number of times that $f_t - s_t$ has been differenced.

To understand the choice between estimates based on (2) and (3), it is useful to recognize that, although the null hypotheses with $\beta = 1.0$ are equivalent, the classes of alternative hypotheses admitted by the two formulations are different. For example, the unconditional growth-rate implications of (2) and (3) are quite distinct when $\beta \neq 1.0$. In particular, equation (2) implies $E\Delta s_t = \beta E\Delta f_t$ whereas (3) implies $E\Delta s_t = E\Delta f_t$ regardless of the value of β. For many analysts, the latter implication will be more attractive as a maintained hypothesis so (3) will, to them, be the preferred specification. And in this case, it transpires that formulation (2) has distinctly unattractive properties. In particular, (3) can be written, as Hodrick (1987, p.30) has noted, as

(3′) $$s_t = \beta f_{t-1} + (1-\beta)s_{t-1} + \varepsilon_t.$$

But that arrangement indicates that if s_t and f_t are both nonstationary and integrated of order 1, then s_t and f_t will be cointegrated, assuming ε_t is stationary, with the cointegrating vector equal to 1.0. And in such a circumstance, a regression of s_t on f_{t-1} (or indeed on f_{t-j} more generally) will tend to result in a slope coefficient of 1.0, regardless of the value of β in (3). So regression (2) will provide no information regarding the magnitude of β in (3), which makes (2) highly inappropriate as an empirical vehicle if (3) provides the more attractive class of alternative hypotheses. (For related arguments, see Meese (1989) and Barnhart and Szakmary (1991).)

It is also the case, however, that from the other point of view a somewhat similar situation obtains. Specifically, if $s_t = \alpha + \beta f_{t-1} + \varepsilon_t$ is deemed the more interesting alternative, formulation (3) will be inappropriate for estimation of β. For with s_t and f_t close to random-walk processes, $E_t s_{t+1}$ will be very closely approximated by s_t. Thus s_{t-1} will by (1) approximately equal $\alpha + \beta f_{t-1}$ and

(12) $$f_{t-1} - s_{t-1} = f_{t-1} - \alpha - \beta f_{t-1} \doteq (1-\beta)f_{t-1} - \alpha,$$

so the regressor in expression (3) will be approximately equal to a nonstationary, random-walk variable. Under the $\beta = 1.0$ hypothesis of principal interest, furthermore, the $f_{t-1} - s_{t-1}$ regressor degenerates to a constant, making nonsense of (3) as a regression specification. From the descriptive statistics reported above we know that $f_{t-1} - s_{t-1}$ is not strictly speaking a random-walk process--but also that it is not terribly far removed from one and that its variability is quite small in comparison with that of Δs_t. Consequently, unreliable estimation of β in (2) should be anticipated from formulation (3).

Thus the choice of a maintained hypothesis, within which the null hypothesis $\beta = 1.0$ is to be embedded, is a matter of considerable importance. This conclusion, which is frequently overlooked, carries more implications than might be apparent initially. It implies, for example, that the interesting line of argument recently put forth by Goodhart, McMahon, and Ngama (1990) cannot be accepted as entirely convincing. For the contention of that paper is basically that there will be a tendency for the slope coefficient in (3) to be estimated as a negative magnitude when the true situation is that (2) obtains with a slope close to 1.0 but smaller than the coefficient relating s_t to f_t in another relationship.[15] That correct arithmetic point is irrelevant, however, if (2) is not the appropriate maintained hypothesis. And, as mentioned above, the choice of (3) over (2) as a maintained hypothesis is for many analysts a compelling one on _a priori_ grounds, as a consequence of the implied unconditional growth-rate properties of the system.

Because of the importance of this choice between formulations (2) and (3), it will be worthwhile to develop an alternative argument that may be still more attractive and convincing. This argument, which is empirical rather than _a priori_, is based on a version of (8) that leaves the coefficients unconstrained. Let us write that relation as

(13) $$s_t = \beta f_{t-1} + \gamma s_{t-1} + \varepsilon_t.$$

Then from the perspective of either (2) <u>or</u> (3) the unbiasedness hypothesis is that $\beta = 1.0$ and $\gamma = 0$. Thus if $\hat{\beta}$ and $\hat{\gamma}$ are the OLS estimators of β and γ, two estimates of β will be provided by each regression, namely, realizations of $\hat{\beta}$ and $\tilde{\beta} = 1-\hat{\gamma}$. For the three BIS exchange rates under discussion, the estimates over 1978.01 - 1990.07 are reported in Table 5. Both estimates have the property of consistency, but standard errors are highly suspect. As it happens, however, the Table 5 results are strikingly supportive of specification (3): in all three cases, the $\hat{\beta}$ and $\tilde{\beta}$ estimates are virtually identical and quite close in value to those of panel 2 in Table 1. We conclude that the β parameter in (3) is significantly different from the 1.0 value implied by the unbiasedness hypothesis. The estimates in specification (2) are (spuriously) close to 1.0 nevertheless because s_t and f_t are cointegrated; the slope coefficient in (2) provides a consistent estimate of the "long run" effect--the sum of the parameter values in a distributed-lag relationship.[16]

<u>IV. Competing Explanations</u>

We turn now to the critical task of attempting to give some plausible interpretation to the empirical regularities outlined above. To a large extent, the job is to understand why β estimates in the vicinity of -3 or -4 result from application of OLS to formulation (3) and what implications this fact carries for the validity of the UIP hypothesis. Of special concern is the usefulness, in light of these β estimates, of exchange-rate models that incorporate the UIP condition as a basic structural component.

There are at least three quite distinct explanations that we shall consider. The first and most straightforward of these is that the UIP relation (4) does not hold; that instead expected exchange-rate depreciation is related to the interest differential as in

(14) $$s^e_{t+1} - s_t = \alpha + \beta\,(R_t - R^*_t) - \xi_t$$

with a value for β in the vicinity of -3 or -4 (henceforth, "-3" for

Table 5

OLS Regression Estimates of Eq. (13),

1978.01 - 1990.07

Exchange Rate	Estimates (std. errors) Const.	β	$1-\gamma$	Statistics R^2	SE	DW
$/DM	0.006 (.013)	-4.428 (1.71)	-4.414 (1.71)	0.965	0.0351	2.17
$/£	-0.003 (.009)	-4.630 (1.11)	-4.621 (1.11)	0.966	0.0333	2.17
$/Yen	-0.077 (.065)	-3.661 (1.19)	-3.644 (1.19)	0.978	0.0363	2.02

brevity). The second possibility to be considered is that $\beta = 1.0$ in expression (14) but that expectations are not rational--$s^e_{t+1} \neq E_t s_{t+1}$--with the prevailing mode of irrationality being such as to generate negative estimates of β in equation (14) or (3). Finally, there is the possibility that policy is conducted via manipulation of $R_t - R^*_t$, with adjustments to exchange rate movements implying a second simultaneous relationship whose form leads to the estimates that are typically observed. Let us then discuss these three contending explanations in turn.[17]

The possibility that β equals -3 in relation (14) is the one that implies the most drastic departure from prevailing views on exchange rate determination. It amounts to a direct denial of uncovered interest parity, which, as mentioned above, is a crucial cornerstone of most empirical and theoretical models. A different sort of denial, one that augmented relation (14) with terms involving relative supplies of home and foreign-country debt (as in the "portfolio approach" literature), would merely constitute a modification or refinement of UIP that would be non-drastic and fully intelligible. But, unfortunately, the evidence does <u>not</u> support this sort of denial.[18] And the form of denial under discussion--one that implies large negative values of β in (14)--would seem to be virtually unintelligible in terms of economic analysis. It would, for instance, imply that a policy-induced increase in home country interest rates would, with foreign conditions and expectations of future conditions unchanged, induce a depreciation--a reduction in the current foreign value of the home currency. As mentioned above, such an implication is inconsistent not only with existing models, but also with views that have been held by actual policymakers for many decades--indeed, for over a century.

As it happens, however, there exists a body of evidence that tends to discredit this first explanation. The evidence in question is that of Frankel and Froot (1990) and Froot and Frankel (1989), which relies on <u>survey</u>

data regarding expectations.[19] For most applications, that type of data is justifiably viewed as unreliable by many analysts, but the details of the current application are relatively favorable. Specifically, equation (14) is intended to represent the behavior of the same group of market participants as those who take part in the surveys. Furthermore, the measures of the variables f_t and s_t, used to represent $R_t - R_t^*$, are ones that these participants have readily available and of which they are almost certainly aware. And when survey-data values, rather than ex post Δs_{t+1} values, are used for $s_{t+1}^e - s_t$ in equation (14) with $f_t - s_t$ for $R_t - R_t^*$, then β is actually estimated to lie in the vicinity of 1.0--See Frankel and Froot (1990, Table 2).

We now turn to the second of the competing explanations. To see that systematic expectational errors could account for strongly negative estimates of β when it in fact equals 1.0, suppose (following Frankel and Froot (1990)) that expected depreciation $s_{t+1}^e - s_t$ systematically exceeds that which transpires when the forward discount $f_t - s_t$ is positive. In other words, suppose that $\gamma > 0$ in the relation

(15) $$\Delta s_{t+1}^e - \Delta s_{t+1} = \gamma(f_t - s_t) + \text{random noise}$$

where $\Delta s_{t+1}^e \equiv s_{t+1}^e - s_t$. Then combining the latter with (14) after using $f_t - s_t = R_t - R_t^*$ gives

(16) $$\Delta s_{t+1} = \alpha + (\beta-\gamma)(f_t - s_t) + \text{noise}.$$

Clearly, then, if $\gamma > 1.0$ a negative estimate will be found for the slope coefficient if in fact $\beta = 1.0$.

This possibility would be of little interest if there were no particular reason to believe that expectations might be systematically erroneous in the manner postulated by (15). As it happens, however, Froot and Frankel (1989) and Frankel and Froot (1990) have presented some evidence suggesting that such could be the case. In particular they utilize--as mentioned above--survey data pertaining to the expectations held by market participants. Combining this data with standard observations on actual Δs_{t+1} and $f_t - s_t$ values, their estimates of the γ parameter in expression (15) are positive and in the vicinity of 3.0 for most of the data sets studied.[20] Such a value would clearly correspond to a slope estimate of -2.0 in (5) even if the true β were 1.0, as implied by UIP. The work of these researchers serves, therefore, to lend some support to the second explanation in comparison with the first.

There is, however, a serious difficulty associated with this second possibility. In part the difficulty is one of understanding how to motivate the particular form of irrationality expressed in (15) for any given value of γ. But the trouble goes even deeper, in the following sense: it is hard even to comprehend what form of behavior is expressed by equation (15). A formula that represented Δs^e_{t+1} in terms of variables known at time t would at least be intelligible, if perhaps hard to justify, but condition (15) involves values not known at the time of expectation formation. It is, therefore, non-operational--it fails to specify how one could simulate a model in which expectations are thusly formed. In that regard it compares unfavorably with the non-rational expectations scheme proposed by DeGrauwe (1988) or the "extrapolative" and "adaptive" formulas studied by Frankel and Froot (1987). It is unclear, however, that any of these schemes would be adequate to explain the "$\beta = -3$" results; substitution into (3) yields expressions relating $s_t - s_{t-1}$ to $f_t - s_t$, not $f_t - s_{t-1}$.[21]

It would seem worthwhile, then, to consider the third possible line of explanation, involving policy behavior. The basic idea is that monetary policymakers in both home and foreign countries have some tendency to resist rapid changes in exchange rates. When a nation's currency is tending to rise in value, for example, monetary policymakers will tend to be a bit more expansionary than otherwise. The main <u>instrument</u> of monetary policy in most nations, moreover, is a short term interest rate--the federal funds rate in the United States, for example. Thus the expansionary tendency will manifest itself as a fall in R_t and in $R_t - R_t^*$ relative to their values in the absence of this postulated policy response. Actual monetary authorities also tend to smooth interest rate movements, however, in the sense of keeping R_t from departing too far from its value in the recent past.[22]

Together, these policy tendencies might be represented in a highly stylized form in the following relation, in which x_t denotes $R_t - R_t^*$ and ζ_t represents random policy influences:

(17) $$x_t = \lambda(s_t - s_{t-1}) + \gamma x_{t-1} + \zeta_t.$$

Here $\lambda > 0$, since s_t is measured as the home-currency price of foreign exchange, while $0 < \gamma \leq 1$ to reflect R_t smoothing as described. To illustrate the implications of the postulated form of policy behavior let us combine (17) with the UIP relation expressed as

(18) $$s_t = s_{t+1}^e - x_t + \xi_t.$$

Here the relation is written in a manner designed to emphasize that it is basically concerned, under the regime at hand, with the determination of the current exchange rate s_t in response to expectations and interest rates available to market participants at home and abroad. The stochastic disturbance term ξ_t reflects not pure expectational error, but rather the myriad of minor influences that keep $s_t = s_{t+1}^e - x_t$ from holding exactly. In addition, and to highlight the contrast with the second competing explanation (irrational expectations) without necessarily denying its potential usefulness, let us initially assume that expectations are fully rational (i.e., that $s_{t+1}^e = E_t s_{t+1}$). Finally, we assume that ζ_t and ξ_t are generated by purely random white-noise processes.

Under these assumptions it is possible to obtain an analytical solution to the system (17)(18) for Δs_t.[23] We start by substituting (17) into (18) and writing $E_t s_{t+1} - s_t$ as $E_t \Delta s_{t+1}$ to obtain

(19) $$E_t \Delta s_{t+1} = \lambda \Delta s_t + \gamma x_{t-1} + \zeta_t - \xi_t.$$

The relevant state variables for Δs_t are then x_{t-1}, ζ_t, and ξ_t so to obtain the linear solution that is free of bubble or bootstrap effects, we conjecture that it is of the form

(20) $$\Delta s_t = \phi_1 x_{t-1} + \phi_2 \zeta_t + \phi_3 \xi_t.$$

The undetermined-coefficients procedure as outlined in McCallum (1983) then gives rise to the following implied identities:

(21) $$\phi_1^2 \lambda - \lambda\phi_1 - \gamma + \phi_1\gamma = 0$$

$$\phi_1\phi_2\lambda - \phi_2\lambda + \phi_1 - 1 = 0$$

$$\phi_1\phi_3\lambda - \phi_3\lambda + 1 = 0$$

The first of these has two solutions for ϕ_1, namely 1 and $-\gamma/\lambda$. Application of the procedure discussed on pp. 146-147 of McCallum (1983) indicates that it is the latter that represents the bubble-free solution, however, so we conclude that $\phi_1 = -\gamma/\lambda$, $\phi_2 = -1/\lambda$, and $\phi_3 = 1/(\lambda+\gamma)$. Thus the solution for Δs_t is

(22) $$\Delta s_t = -\gamma/\lambda\ x_{t-1} - (1/\lambda)\zeta_t + (1/(\lambda+\gamma))\xi_t.$$

We see, then, that with γ and λ both positive, the coefficient attached to x_{t-1} will be negative in a regression with dependent variable Δs_t. In fact, with values of γ fairly close to 1.0, representing considerable persistence in interest-rate differentials, "reasonable" values for λ will imply coefficient magnitudes in excess of 1.0 in absolute value. With, for example, $\gamma = 0.80$ and $\lambda = 0.20$ we get a coefficient of -4 in the relation (22), which is of exactly the same form as the regression equation (3) that produces the anomalous estimates of β. But those estimates would be generated by the system under discussion despite its specification that the slope coefficient (corresponding to β) in the UIP relation is 1.0.

The foregoing result is quite encouraging and suggestive of the possibility that it is a muddling together of two distinct behavioral relationships--policymakers' as well as market participants'--that is responsible for slope estimates like -3 in regressions of form (3). But, unfortunately, there is another dimension along which the model at hand fails to match the data. That failure concerns the univariate behavior of x_t. If, that is, the solution expression (22) is substituted into (17), the resulting expression is

$$(23) \qquad x_t = (\lambda/(\lambda+\gamma))\ \xi_t,$$

in which the x_{t-1} values have canceled out. Thus the implication is that, with ξ_t and ζ_t white noise variates, the process for x_t should be close to white noise. In fact, as we have seen in Section III, the time series properties of $x_t = f_t - s_t$ are more complex, being well represented by ARMA (1,0,1) specifications with sizeable AR and small MA parameters.

There is, however, a simple modification of the system that is capable of remedying the foregoing difficulty. Suppose, in particular, that the disturbance term ξ_t is generated by a first-order autoregressive process rather than one that is white noise. Then we have

$$(24) \qquad \xi_t = \rho\xi_{t-1} + u_t \qquad\qquad |\rho| < 1.0$$

where the presumption is that $\rho > 0$. With this specification, it is straightforward to infer that the bubble-free solution for Δs_t of the form (20) differs from (22) only by having a coefficient of $1/(\lambda + \gamma - \rho)$ attached to ξ_t. But that form is not the one that is relevant for comparison with equation (3), for ξ_t will be correlated with x_{t-1} when ξ_t is generated by (24). Substitution of the latter yields, however, the appropriate expression, namely,

$$(25) \qquad \Delta s_t = [(\rho - \gamma)/\lambda]\ x_{t-1} - (1/\lambda)\zeta_t + [1/(\lambda + \gamma - \rho)]u_t.$$

In this expression, the composite disturbance is white noise and uncorrelated with x_{t-1}. A negative coefficient on x_{t-1} is not inevitable, but with γ close to 1.0 that sign is extremely likely. And it is entirely plausible that λ could be small enough to make the composite coefficient's absolute value equal to 3, 4, or an even greater magnitude. Now, moreover, substitution into (17) results in the following expression for x_t:

(26) $$x_t = [\lambda/(\lambda + \gamma - \rho)]\ \xi_t.$$

Thus with ξ_t autoregressive, that property will also pertain to x_t -- thereby approximating the time series properties of $f_t - s_t$ in Table 4, where we identified an ARIMA (1,0,1) specification with a small MA parameter value.

In fact, there are a number of additional ways in which the time-series properties of the theoretical system under discussion match those of the actual data series summarized in Tables 1-5. First, and most basically, it is clear from (25) and (26) that the variance of the one-period forecast error will be much greater for s_t than for x_t. With λ sufficiently small, the ratio could easily be as great as indicated in Tables 3 and 4. Next, that case would imply that the process for s_t would be <u>close to</u> that of a random walk, since the first term on the right-hand side of (25) would provide little variability compared with the third.

In addition, Δs_t and Δx_t can in this case be represented as

(27) $$\Delta s_t = [1/(\lambda + \gamma - \rho)]\ [(\rho - \gamma)\ \xi_{t-1} + u_t] - (1/\lambda)\ \zeta_t$$

and

(28) $$\Delta X_t = [\lambda/(\lambda + \gamma - \rho)]\ [(\rho - 1)\ \xi_{t-1} + u_t],$$

so with γ close to 1.0 s_t and x_t will reflect the same shocks if the variance of ζ_t is relatively small. And with a small value of λ, as we are assuming, $\Delta f_t = \Delta s_t + \Delta x_t$ will tend to move almost one-for-one with Δs_t, thereby mimicking the near-unity regression coefficient obtained in the second panel of Table 2. Furthermore, with Δs_t being close to white noise--s_t close to a random walk --it then also follows that Δs_t will be almost unrelated to Δf_{t-1}, as in the first panel of Table 2.

All in all, then, the specification provided by (17), (18), and (24) yields the implication that reasonable parameter values will provide an impressive match to the time series data on spot rates s_t, forward discounts x_t, and the relationship of Δs_t to x_{t-1}. That this result can be obtained in such an extremely simple setup--and with parameter values that accord with one's intuition--suggests rather strongly that policy response may indeed be the main source of the puzzling findings highlighted in Tables 1 through 5. If so, one conclusion would be that, since (18) is one relation of the model, uncovered interest parity holds despite the failure of the unbiasedness hypothesis.

V. Conclusions

The arguments put forth in this paper can be summarized briefly as follows. First, there are reasons for viewing the uncovered interest parity (UIP) relationship as more important, in terms of economic analysis, than the unbiasedness of forward rates as predictors of future spot exchange rates. The two hypotheses are closely related, so that test rejections of the latter tend to cast doubt on the former, but are not identical--so unbiasedness rejections are not conclusive for UIP.

Next, some representative evidence is presented that pertains to alternative versions of the unbiasedness test. Although $s_t = \alpha + \beta f_{t-1} + \varepsilon_t$ and $s_t - s_{t-1} = \alpha + \beta(f_{t-1} - s_{t-1}) + \varepsilon_t$ are equivalent under the null hypothesis of $\beta = 1.0$, they represent different classes of alternative hypotheses. Empirically, they give rise to extremely different outcomes, estimates of β being very close to 1.0 in the former equation but in the vicinity of -3.0 in the latter. In a generalized specification that includes both as special cases, the results strongly favor the second specification--thereby rejecting unbiasedness.

Three possible explanations for the $\beta = -3$ result are considered and related to the UIP condition. Of these three, the latter two--one involving systematically irrational expectations and the other an additional relationship reflecting monetary policy behavior--are consistent with UIP. The policy-response hypothesis, that monetary authorities manage interest-rate differentials so as to resist rapid changes in exchange rates and in these differentials, is attractive conceptually and is capable of explaining the β = -3 finding. Furthermore, when combined with the assumption of first order AR disturbances, this policy-response hypothesis also has several notable implications consistent with the univariate time series properties of the spot rate and forward-discount series, the latter serving as a proxy for the interest differential. The most attractive of the three possible hypotheses is, then, one that does not contradict uncovered interest parity.

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Footnotes

[1]Useful surveys have been provided by Levich (1985), Hodrick (1987), MacDonald and Taylor (1989), Meese (1989), and others. Individual studies that have been particularly influential include Frankel (1976), Bilson (1981), Hansen and Hodrick (1980), Fama (1984), and Froot and Frankel (1989).

[2]Some discussion has been provided, of course, including useful contributions by Cumby and Obstfeld (1984) and MacDonald (1988, pp.208-213).

[3]Since first drafting the paper, I have learned of a recent contribution by Barnhart and Szakmary (1991) that reaches conclusions regarding unbiasedness that are similar to those developed here. That contribution does not discuss UIP, however, and does not put forth any model comparable to that of Section IV below.

[4]Forward rates also exist for two or more periods into the future, of course, but to avoid inessential complications we shall restrict our discussion to one-period forward rates and interest rates.

[5]With s_t and f_t included in Ω_t, ε_t should be serially uncorrelated given that we are focusing on one-period rates.

[6]Early tests using specification (3) were conducted by Tryon (1979) and Longworth (1981). More recent examples are too numberous to list.

[7] Some writers would argue that time-varying premia represent the <u>only</u> reason for including a random disturbance term in (4) or other similar relations. Indeed, such a position represents a point of view that is at present rather widely held. It is this author's opinion that such a point of view is misguided; that it is proper (though unfashionable) to include disturbance terms in all behavioral relations intended for empirical application. The reason, briefly, is that all useful models are simplifications of reality that fail to account for behavioral influences of many kinds. Consequently, relations that hold exactly in theoretical models will not in empirical applications, even for a single agent. And at the market level of aggregation the "error" will be greater and more complex because agents are not, in fact, all alike. In addition, there is in practice usually some amount of measurement error, even in studies of exchange rate data.

[8] See Fratianni and Wakeman (1982) for one study and Levich (1985) or MacDonald (1988) for surveys.

[9] This statement should not be understood as a claim that unbiasedness and UIP can be different <u>only</u> with irrational expectations. Indeed, another source of differences is emphasized below in Section IV.

[10] Accumulated evidence has not, however, been supportive of portfolio-approach modifications of the UIP condition. On this, see Isard (1988) and MacDonald (1988).

[11] The qualification "ostensibly" will be discussed momentarily.

[12] By simply using the final-day rates we are failing to make the alignment correction explained by Hodrick (1987, pp. 36-37) and utilized by most recent researchers. It is my impression that this misalignment has little effect on the reported results. Some evidence in support of this impression is provided by Bekaert and Hodrick (1991), Table 1.

[13] According to Goodhart, McMahon, and Ngama (1990, p. 3), some related results have been obtained in an unpublished study by Pope and Peel (1989).

[14] The values of the estimated AR parameters are 0.64, 0.87 and 0.87, respectively, with 0.53, 0.08, and 0.19 for the MA parameters.

[15] There is, incidentally, a second flaw in the Goodhart, McMahon, and Ngama (1990) argument. Specifically, the argument presumes that a simple bivariate relationship connecting s_t to f_t is structural; but no theory is presented in support of that crucial hypothesis. Indeed, the nature of the relationship is not even mentioned.

[16] Similar conclusions have very recently been reached by Barnhart and Szakmary (1991).

[17] Our list of possible explanations differs from most of those that have been advanced by researchers, which typically include time-varying risk premia and the "peso problem" as prominent contenders (see, e.g., Hodrick (1987) and Meese (1989)). Recent discussions by Frankel and Froot (1990) and Froot and Thaler (1990) seem to convincingly eliminate these possibilities, despite the huge volume of high-quality research that they have stimulated.

[18] See again the reviews by MacDonald (1988), Isard (1988), and Levich (1985).

[19] The survey data used by Frankel and Froot (1987) (1990) is collected initially by the Economists' Financial Report of London and Money Market Surveys (MMS) of Redwood City, California. The Economist's surveys have been conducted every six weeks since 1981 whereas the MMS surveys have been conducted each two weeks since late 1982. For more details, see Frankel and Froot (1987).

[20]See Frankel and Froot (1990), Table 3, and Froot and Frankel (1989), Table VI. The MMS case with one-month rates contains erroneous data, according to (1990, p.8), and should be ignored.

[21]Furthermore, at short horizons, the extrapolative formula yields a positive extrapolative coefficient; i.e., δ is positive in $s^e_{t+1} - s_t = \delta(s_t - s_{t-1})$. Substitution in (3) then yields $\delta(s_t - s_{t-1}) = f_t - s_t + \xi_t$, implying a positive slope parameter.

[22]revious discussions that mention policy response as a likely reason for the $\beta = -3$ finding include Isard (1988, p.186) and Boyer and Adams (1988). The former includes no explicit model, however, while the policy equation in the latter is quite different from ours.

[23]With rational expectations, the system (17) (18) will not determine the level of s_t. In what follows that matter will be ignored, as it is independent of the issues of special concern in this paper. It is my conjecture, however, that nominal determinacy would require that (17) be regarded as a limiting case of some policy rule in which some nominal variable appears, as analysed in the closed-economy context in McCallum (1986)

Appendix A

obs	SDM	FDM	SPND	FPND	SYEN	FYEN
1977.09	2.307400	2.301300	1.746800	1.748300	264.5000	263.8300
1977.10	2.252800	2.246600	1.834700	1.838200	250.3500	249.5300
1977.11	2.227800	2.223100	1.815000	1.815000	244.2000	243.4800
1977.12	2.105000	2.095800	1.906000	1.906600	240.0000	238.7500
1978.01	2.111800	2.105000	1.949200	1.949400	241.7400	240.8600
1978.02	2.036000	2.029600	1.934200	1.933900	238.6500	237.6300
1978.03	2.023000	2.015700	1.856100	1.856300	223.4000	222.1300
1978.04	2.067800	2.060100	1.828700	1.824200	223.9000	222.8000
1978.05	2.100800	2.093300	1.822300	1.818900	223.1500	222.2500
1978.06	2.075300	2.066500	1.860200	1.855300	204.5000	203.5200
1978.07	2.041300	2.032500	1.931300	1.926600	190.8000	189.6300
1978.08	1.986500	1.978200	1.942100	1.938400	190.0000	188.9800
1978.09	1.938600	1.928600	1.972100	1.966600	189.1500	188.0600
1978.10	1.736700	1.725800	2.089000	2.087500	176.0500	174.6300
1978.11	1.923400	1.911200	1.949400	1.946400	197.8000	195.9800
1978.12	1.828000	1.815200	2.034700	2.033000	195.1000	193.1300
1979.01	1.861600	1.851900	1.995400	1.991200	201.4000	200.0500
1979.02	1.851500	1.842300	2.023500	2.017700	202.3500	201.1000
1979.03	1.867600	1.859000	2.068500	2.065000	209.3000	208.5000
1979.04	1.901900	1.892500	2.058100	2.055100	219.1500	218.0300
1979.05	1.909100	1.901600	2.067500	2.065000	219.7000	218.6800
1979.06	1.848200	1.840300	2.167900	2.161600	217.0000	215.9000
1979.07	1.837700	1.829700	2.282300	2.275100	216.9000	215.9000
1979.08	1.827800	1.820900	2.251000	2.246600	220.0500	219.2000
1979.09	1.742500	1.733700	2.196500	2.194500	223.4500	222.3000
1979.10	1.806600	1.797400	2.077500	2.077400	237.8000	236.1800
1979.11	1.730000	1.722200	2.193000	2.188500	249.5000	248.2300
1979.12	1.731500	1.722800	2.224500	2.220100	239.9000	238.7000
1980.01	1.739400	1.731400	2.265900	2.258400	238.8000	237.6300
1980.02	1.772300	1.759900	2.278300	2.273400	249.8000	248.3700
1980.03	1.941900	1.924300	2.167400	2.170900	249.7000	248.6800
1980.04	1.801500	1.792300	2.266000	2.258600	238.3000	237.9800
1980.05	1.786000	1.784500	2.330000	2.317300	224.4000	225.0000
1980.06	1.758200	1.757900	2.361800	2.344300	218.1500	218.9200
1980.07	1.785100	1.783900	2.343700	2.329200	226.8500	227.4700
1980.08	1.792300	1.788500	2.392400	2.381500	219.2000	219.2800
1980.09	1.811300	1.804000	2.388700	2.383000	212.0000	211.7300
1980.10	1.909200	1.898700	2.439000	2.435600	211.7500	210.7800
1980.11	1.925700	1.909700	2.360000	2.370000	216.7500	215.1000
1980.12	1.959000	1.942300	2.384500	2.393700	203.6000	201.5000
1981.01	2.116700	2.103000	2.386500	2.393500	205.2000	203.6500
1981.02	2.129500	2.126700	2.208500	2.214500	208.8500	207.3500
1981.03	2.101800	2.097600	2.244900	2.249600	211.4000	209.9700
1981.04	2.214500	2.204000	2.139100	2.148600	215.0000	212.9500
1981.05	2.327400	2.316600	2.072300	2.082800	223.5000	221.5000
1981.06	2.390900	2.378500	1.946200	1.957300	225.7500	223.5300
1981.07	2.464500	2.447000	1.856700	1.866000	239.7500	237.0300
1981.08	2.429000	2.417500	1.837800	1.844600	228.7500	226.6500
1981.09	2.322500	2.313300	1.800700	1.801300	231.5500	229.6800
1981.10	2.254200	2.247000	1.845500	1.844100	233.5000	231.8600
1981.11	2.203500	2.201300	1.968500	1.962500	214.1500	213.3200
1981.12	2.254800	2.249100	1.915000	1.912300	220.2500	218.8800

obs	SDM	FDM	SPND	FPND	SYEN	FYEN
1982.01	2.308500	2.301200	1.884500	1.884300	228.4500	227.0500
1982.02	2.386000	2.376700	1.818300	1.819300	235.2000	233.5200
1982.03	2.414200	2.401300	1.780200	1.783400	248.3000	246.3900
1982.04	2.332700	2.320500	1.792500	1.795500	236.3000	234.5900
1982.05	2.345200	2.333900	1.791000	1.792900	243.7000	242.2500
1982.06	2.459800	2.446100	1.739600	1.743800	255.5500	253.7300
1982.07	2.454500	2.447200	1.741000	1.741900	256.6500	255.4200
1982.08	2.497200	2.491100	1.716500	1.716500	259.6000	258.7300
1982.09	2.527600	2.520100	1.692700	1.693500	269.4000	268.4800
1982.10	2.566800	2.561100	1.676000	1.676000	277.4000	276.6900
1982.11	2.487200	2.482400	1.612300	1.611300	253.4500	253.0000
1982.12	2.376500	2.369900	1.613500	1.611800	235.3000	234.7200
1983.01	2.447500	2.440900	1.533900	1.531300	238.4000	237.9400
1983.02	2.421200	2.413600	1.522000	1.518500	235.5500	235.1500
1983.03	2.426500	2.417300	1.476300	1.474900	239.3000	238.6700
1983.04	2.458100	2.449500	1.564000	1.562300	237.7000	237.1000
1983.05	2.519000	2.509700	1.608600	1.607400	238.6000	238.0000
1983.06	2.541900	2.532000	1.529500	1.529600	239.8000	239.1700
1983.07	2.643500	2.632200	1.519300	1.519800	241.5000	240.8200
1983.08	2.706800	2.696100	1.492500	1.493000	246.7500	246.0700
1983.09	2.639100	2.630600	1.494800	1.494700	236.1000	235.5500
1983.10	2.626400	2.617700	1.494100	1.494700	233.6500	233.0000
1983.11	2.697000	2.688700	1.463200	1.464200	234.2000	233.7000
1983.12	2.723800	2.714500	1.450500	1.451300	232.0000	231.2800
1984.01	2.813900	2.805200	1.403200	1.403700	234.7400	234.0900
1984.02	2.605800	2.596500	1.489100	1.490200	233.2800	232.6300
1984.03	2.590000	2.579600	1.442500	1.444600	224.7500	223.9500
1984.04	2.717400	2.705100	1.398500	1.401300	226.3000	225.4000
1984.05	2.733300	2.720700	1.385500	1.388300	231.6300	230.7000
1984.06	2.784200	2.769500	1.353500	1.356900	237.4500	236.2800
1984.07	2.896400	2.880400	1.306500	1.305800	245.5000	244.2800
1984.08	2.887000	2.871800	1.311200	1.312400	241.7000	240.6200
1984.09	3.025300	3.011000	1.247500	1.248000	245.4000	244.3800
1984.10	3.029600	3.018800	1.217000	1.216100	245.3000	244.5400
1984.11	3.096300	3.087800	1.199800	1.198900	246.5000	246.0000
1984.12	3.148000	3.140200	1.162200	1.161100	251.5800	251.1000
1985.01	3.167700	3.161500	1.125400	1.121400	254.7800	254.3400
1985.02	3.322500	3.314100	1.093000	1.088500	259.0000	258.4800
1985.03	3.093000	3.085400	1.242500	1.237500	250.7000	250.1500
1985.04	3.090200	3.083300	1.239300	1.234800	251.4000	250.8800
1985.05	3.089200	3.083700	1.273000	1.267700	251.7800	251.4900
1985.06	3.060700	3.055100	1.295000	1.289500	248.9500	248.6200
1985.07	2.788400	2.780500	1.429000	1.424700	236.6500	236.2900
1985.08	2.781800	2.773900	1.397500	1.393100	237.1000	236.8000
1985.09	2.669900	2.661600	1.401500	1.396800	216.0000	215.6700
1985.10	2.616800	2.609900	1.443000	1.438900	211.8000	211.7100
1985.11	2.512000	2.504500	1.484500	1.480300	202.0500	202.0300
1985.12	2.461300	2.454500	1.441000	1.436700	200.6000	200.4500
1986.01	2.389200	2.382900	1.414500	1.409500	192.6500	192.4000
1986.02	2.218500	2.212300	1.465500	1.459700	180.4000	180.1500
1986.03	2.317500	2.312000	1.481000	1.475900	179.3000	179.0000
1986.04	2.186500	2.182500	1.543500	1.538400	168.1000	167.8000

obs	SDM	FDM	SPND	FPND	SYEN	FYEN
1986.05	2.312700	2.307900	1.483000	1.479300	172.0500	171.7100
1986.06	2.198600	2.193800	1.529000	1.524800	163.9500	163.5900
1986.07	2.094000	2.090900	1.492500	1.488300	154.1500	153.9300
1986.08	2.052000	2.049600	1.478500	1.473400	156.0500	155.9300
1986.09	2.020700	2.018000	1.446500	1.440700	153.6300	153.4600
1986.10	2.067600	2.065500	1.400000	1.394300	161.4500	161.2200
1986.11	1.977300	1.974800	1.436000	1.430300	162.2000	161.9600
1986.12	1.940800	1.937700	1.475500	1.470200	160.1000	159.7700
1987.01	1.808500	1.805700	1.529500	1.523900	152.3000	152.0800
1987.02	1.826800	1.823200	1.543000	1.537100	153.1500	152.8900
1987.03	1.805100	1.800900	1.605000	1.599900	145.8500	145.4900
1987.04	1.786400	1.782100	1.664500	1.660600	139.6500	139.3800
1987.05	1.821500	1.816700	1.626500	1.624100	144.1500	143.7500
1987.06	1.829900	1.824600	1.608000	1.605200	146.7500	146.3200
1987.07	1.855400	1.850900	1.593500	1.590500	149.2500	148.8400
1987.08	1.815200	1.810600	1.628000	1.624200	142.3500	141.9700
1987.09	1.838300	1.832500	1.629000	1.625900	146.3500	145.9000
1987.10	1.739300	1.734100	1.715000	1.712000	138.5500	138.1900
1987.11	1.635400	1.629000	1.832500	1.831200	132.4500	132.0200
1987.12	1.581500	1.576500	1.871000	1.868800	122.0000	121.6800
1988.01	1.675900	1.671200	1.773000	1.770700	127.1800	126.9000
1988.02	1.688400	1.683100	1.773500	1.769900	128.1200	127.8200
1988.03	1.659300	1.654500	1.880000	1.877400	124.5000	124.1800
1988.04	1.668300	1.663000	1.876000	1.874500	124.8200	124.4600
1988.05	1.726700	1.720300	1.845800	1.846000	124.8000	124.3900
1988.06	1.821100	1.815900	1.711000	1.708700	132.2000	131.7800
1988.07	1.881000	1.876100	1.713500	1.710200	132.5300	132.1000
1988.08	1.874800	1.869400	1.681500	1.676200	134.9700	134.5800
1988.09	1.879800	1.874500	1.685500	1.680700	134.3000	133.8700
1988.10	1.768400	1.763300	1.780000	1.774900	125.0000	124.5700
1988.11	1.735400	1.728700	1.847100	1.841700	121.8500	121.2900
1988.12	1.780300	1.774300	1.809500	1.803700	125.9000	125.3800
1989.01	1.864600	1.859700	1.761300	1.756200	129.1300	128.6300
1989.02	1.829600	1.824500	1.739800	1.735300	127.1500	126.5400
1989.03	1.892700	1.886400	1.688800	1.685100	132.5500	131.8400
1989.04	1.878300	1.871900	1.690000	1.685700	132.4900	131.8800
1989.05	1.985800	1.979900	1.573300	1.567600	142.7000	142.1200
1989.06	1.952500	1.948000	1.550200	1.543900	143.9500	143.4000
1989.07	1.866000	1.861800	1.664200	1.656500	138.4000	137.9500
1989.08	1.960400	1.957300	1.570000	1.563600	144.2800	143.8700
1989.09	1.868300	1.865300	1.625200	1.617700	139.3500	138.8800
1989.10	1.837500	1.835300	1.577200	1.568400	142.1500	141.8300
1989.11	1.789500	1.788600	1.567800	1.559500	142.9000	142.6400
1989.12	1.697800	1.696800	1.605500	1.595900	143.4000	143.1700
1990.01	1.682600	1.682600	1.682800	1.674300	144.4000	144.2100
1990.02	1.691800	1.691400	1.684700	1.675400	148.5200	148.3700
1990.03	1.694400	1.694400	1.642800	1.634200	157.6500	157.4800
1990.04	1.680300	1.679600	1.636000	1.626500	159.0800	158.8900
1990.05	1.691000	1.690400	1.682000	1.672400	151.7500	151.6200
1990.06	1.671500	1.671200	1.741800	1.732300	152.8500	152.7500
1990.07	1.596000	1.595800	1.852700	1.841300	147.5000	147.4400

Note: Here S and F denote spot and one-month forward rates, respectively, with DM, PND, and YEN designating the currency in question. For DM and YEN the values are currency units per U.S. dollar while for PND the value are dollars per PND. All reported calculations are based on rates in terms of dollars per currency unit.