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WAS IT REAL? THE EXCHANGE RATE-INTEREST DIFFERENTIAL RELATION, 1973-1984

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ABSTRACT

The main result of Meese and Rogoff [1983 a,b] is that small structural exchange rate models forecast major dollar exchange rates no better than a naive random walk model. This result obtains even when the model forecasts are based on actual realized values of the explanatory variables. Here we improve our methodology by implementing a new test of out-of-sample fit; the test is valid even for overlapping long-horizon forecasts. We find that the dollar exchange rate models perform somewhat less badly over the recent Reagan regime period than over the episodes studied previously. The methodology is also applied to the mark/yen and mark/pound exchange rates, and to real exchange rates. Finally, we test to see if real exchange rates and real interest differentials can be represented as a cointegrated process. The evidence suggests that there is no single common influence inducing nonstationarity in both real exchange rates and real interest differentials.

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I. Introduction

What forces cause exchange rates to be so volatile? Unfortunately, it has not proven easy to identify the key determinants of exchange rate movements. Even the theoretically elegant asset models of exchange rate determination have fared poorly. Using a rolling regression methodology, Meese and Rogoff (1983a, b) find that both sticky- and flexible-price monetary models of exchange rate determination fail to outforecast a naive random walk model, even when the forecasts are based on actual realized values ("news") of the explanatory variables. While it would be unrealistic to expect these simple models to explain a large percentage of the variance of exchange rate changes, it is still surprising that exchange rate movements seem to have so little correlation with changes in relative monies, incomes, interest rates, inflation rates, and current accounts. Some may find it tempting to conclude that it is necessary to use more sophisticated models; e.g., models with a wider menu of domestic and foreign assets held by risk averse agents. However, current evidence suggests that these portfolio balance models do not yield better results.[1]

The purpose of the present paper is fourfold. First, we wish to update our earlier results on the out-of-sample fit of small-scale monetary exchange rate models. Our two earlier studies covered the period March 1973-June 1981. Here we add three years of data and investigate whether the models perform any better over the more recent forecasting period.

1 See Frankel (1982), Langenborg (1985), Rogoff (1984) or the Federal Reserve staff studies paper #135 by Danker et al. (1985).

By extending the data set, we are also able to meaningfully examine longer forecasting horizons than was previously possible. Second, and perhaps foremost, we improve on the methology of our earlier studies by *formally* comparing the out-of-sample fit of each model with that of the random walk model; the test we develop is valid even for multiple forecast horizons. In our earlier studies, we were only able to present formal comparisons for one-month (non-overlapping) forecasts. Other, relatively minor, econometric improvements here are that we implement some different estimation techniques, including one where the rolling regression coefficient estimates depend more heavily on recent observations. We also make a crude attempt to incorporate fiscal deficits into the empirical models. Third, we report detailed results for two non-dollar exchange rates in this study: the mark/pound and the mark/yen rates. (As in our previous work, we also examine the dollar/mark, dollar/yen and dollar/pound rates.) One reason for considering non-dollar rates is to see whether money demand instability in the U.S. is an important explanation of the models' poor performances.[2] For the same reason, we devote considerable attention to examining real versions of our selected structural exchange rate models. Fourth, we further explore the relationship between real exchange rates and real interest differentials by testing to see if the two series can be represented as a cointegrated process [see Granger (1983)]. The evidence here suggests that there is no common influence inducing nonstationarity in both real exchange rates and real interest rate differentials.

Section II contains the results for nominal exchange rate models over the forecasting period November 1980-June 1984. In contrast to our previous results based on earlier time periods, we now do find instances where the structural models have lower root-mean-square forecast errors (RMSE) [or mean absolute forecast errors (MAE)] than the random walk model. However, only in the case of the dollar/yen rate do these differences approach statistical significance (using the formal test discussed above and in Appendix B). In Section III, we derive *real* exchange rate versions of the models and present results for the forecasting periods

2 Simpson and Porter (1980), among others, have documented the instability of U.S. money demand for conventionally measured aggregates such as M1 and M2.

January 1977-June 1984, and November 1980-June 1984. In the real exchange rate models, the real interest rate differential is the most important explanatory variable. We find that it usually enters with the theoretically anticipated sign; i.e., a high real interest rate differential implies a high real exchange rate. However, only in *one* case (the mark/pound rate at one month horizons for the November 1980-June 1984 forecasting period) do any of the models ever *significantly* outforecast the random walk model. Still, the results are somewhat more favorable than those obtained in our previous work. In Section IV, we consider a number of modifications to the real exchange rate models of Section II, and present representative results for the dollar/yen rate. We consider longer-term interest rates, cumulative government deficits, lagged adjustment mechanisms, and alternative estimation techniques. None of these modifications yields any substantial improvement. Section V contains a test of the cointegration of real exchange rates and real interest rate differentials. Section VI concludes.

II. The Out-of-Sample Fit of Nominal Exchange Rate Models

This section updates the results of Meese and Rogoff [1983a, b] on the out-of-sample fit of simple monetary (nominal) exchange rate models. We improve on our earlier methodology by implementing a test which allows us to formally compare the RMSE of alternative models at multiple forecast horizons.

As the main focus of the present paper is on the real exchange rate, our description of the nominal exchange rate models will be quite brief.[3] The models considered here are all variants of the monetary model of exchange rate determination; they differ only in their assumptions about the price adjustment mechanism and the behavior of the "long-run" (flexible-price equilibrium) real exchange rate. (An example of a sticky-price model is Dornbusch's [1976]

3 The reader is referred to Meese and Rogoff (1983b) for a more complete discussion of these models.

"overshooting" model.)[4] The quasi-reduced form specification of all the models we test is subsumed in the general specification below:

(1) $$s = a_0 + a_1(m - m^*) + a_2(y - y^*) + a_3(r_s - r_s^*)$$

$$+a_4(\pi^e - \pi^{*e}) + a_5(\overline{TB} - \overline{TB}^*) + u,$$

where star (*) superscripts denote the foreign country, s is the logarithm of the nominal exchange rate (the domestic price of foreign currency), m - m* is the logarithm of the ratio of the home to foreign money supply, y - y* is the logarithm of the ratio of home to foreign output, $r_s - r_s^*$ is the short-term interest differential, and $\pi^e - \pi^{*e}$ is the long-term expected inflation differential. Finally, $\overline{TB}$ and $\overline{TB}^*$ are the cumulated home and foreign trade balances, and u is a disturbance term.

It is important to recognize that although equation (1) contains only contemporaneous explanatory variables on the right-hand side, the exchange rate does depend on market expectations about future fundamentals. These expectations are embodied in the interest differential and the expected inflation differential, which are both endogenous variables. Expected future fundamentals (say, money supplies and real incomes) would appear on the right-hand side if one were to solve out for a true reduced-form instead of using the quasi-reduced form of equation (1). Of course, one would then have to find a good method for measuring market expectations about the fundamental exogenous variables.

The assumptions that $a_1 = 1$ and that $a_2 < 0$ are common to all the models; an increase in the supply of money or a decrease in the transaction demand for money causes a depreciation of the exchange rate. The flexible-price monetary model posits that $a_3 > 0$, under the assumption that the predominant shocks are monetary and that a rise in interest rates signals

4 For other monetary models, see Bilson (1978), Frenkel (1976), Frankel (1979), and Hooper and Morton (1982).

a rise in expected inflation. The sticky-price monetary model (or Dornbusch-Frankel [D-F] model) similarly posits that a rise in expected inflation causes a depreciation. However, a rise in the *real* interest differential causes an appreciation (when disturbances are primarily monetary), so that $a_3 < 0$, $a_4 > 0$. Hooper and Morton's (H-M) model contains the further assumption that a rise in the cumulated current account (relative to trend) signals an appreciation of the long-run flexible-price real exchange rate, so that $a_5 < 0$.

The main result of Meese and Rogoff [1983 a,b] was that the models subsumed in equation (1) forecast major dollar exchange rates no better than a naive random walk model, even when their forecasts are based on *realized* values of the explanatory variables. Thus "news" about market fundamentals would appear to be of little value in predicting nominal exchange rates. It is true that some types of news (omitted variables) are difficult to quantify: political events, financial crises, and central bankers' views of equilibrium exchange rates. These types of news may have considerable explanatory power. The effects of "unquantifiable news" could not be studied in the current experimental design, but we do provide some indirect evidence in Section V.[5]

Aside from the omitted variable caveat, the main result of our prevous work was robust to a variety of estimation techniques, specifications of the underlying money demand functions, alternative serial correlation or lagged adjustment corrections, and measures of forecast accuracy [including both RMSE and MAE]; it was also robust to allowing for different coefficients on home and foreign variables. The two earlier studies were based on data from March 1973 through June 1981; the two forecasting periods considered were January 1977-June 1981 and December 1978-June 1981.[6]

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5 Our quasi-reduced form specifications will still be valid in the presence of certain types of speculative bubbles (since interest rates enter as explanatory variables). Burmeister, Flood, and Garber (1983) note the observational equivalence of bubbles and certain types of omitted variables. See also Hamilton and Whiteman (1985).

6 We readily admit that it is possible to find shorter sample periods where one of the models strongly outperforms the random walk model for a particular exchange rate; see, for example, Woo (1985). However, our results highlight the fact that the correlations of exchange rates and market fundamentals over these subperiods are not structural.

In Table 1 below, we present representative rolling regression results for the current Reagan regime, November 1980-June 1984.[7] The coefficients of the models are first estimated for the period March 1973 through October 1980. These estimates are used to produce forecasts at one- to twelve-month horizons, employing actual realized values of the explanatory variables. Then we add the data for November 1980, re-estimate the models, and repeat the procedure. As the RMSE statistics listed in Table 1 indicate, the models do not perform particularly well over the recent forty-four month period. However, the D-F model does do better than the random walk (RW) model for the dollar/yen rate and the mark/pound rate. The improvement for the dollar/yen, while not dramatic, falls just short of being statistically significant. (The entries in parentheses in Table 1 are for the asymptotically N(0,1) statistic discussed in Appendix B. A positive entry indicates improvement over the RW model.) The H-M model improves on the RW model for the dollar/pound and mark/pound exchange rates at all horizons, though again the improvement falls short of statistical significance. For the dollar/yen rate, there is a very slight improvement at six- and twelve-month horizons. Results based on MAE are qualitatively similar. Finally, we note that the RMSE for the forward rate (not listed) are slightly higher than those for the spot rate across all horizons and currencies for the November 1980-June 1984 sample period.

In part for illustrative purposes, and in part because "market timing" tests have recently become popular in the finance literature, Table 1 also lists the percent of the time that each model correctly forecasts the direction of change in the exchange rate. If the null hypothesis is that a structural model predicts the one-month direction of change of the exchange rate no better than a coin toss, then a model which successfully predicts the direction of change 61 percent of the time (27 out of 44 months) would constitute significant evidence against the null at the ninety-five percent significance level. (The longer horizon forecasts are serially correlated, so the statistics there must be regarded as purely descriptive.) While the impor-

7 The results presented in Table 1 are based on Cochrane-Orcutt estimation. In our previous papers, we also considered both instrumental variable (IV) techniques and constrained coefficient estimation. The GLS forecasting results were typically no worse than IV. See Appendix C for a listing of the full-sample GLS coefficient estimates.

tance of the direction of change metric is debatable,[8] it is interesting to note that the models seem to perform somewhat better by this criterion than by either RMSE or MAE.

We have previously explored a number of alternative explanations for the poor out-of-sample fit of monetary exchange rate models, including simultaneous equations bias, sampling error, risk premia, measurement of inflationary expectations, failure of the constraint that home and foreign country variables enter with equal but opposite sign, nonlinearities, and shifting parameters. Other explanations of our results are that real disturbances are importance (and are not captured by H-M's model), and that there are speculative bubbles (omitted variables). The real disturbance explanation is appealing because it is consistent with the low volatility of forward premia [see Flood (1981)]. As a practical matter, however, it may not be as easy to identify and measure real disturbances as it is monetary disturbances (though this remains an open question). The next two sections of this paper will be devoted to exploring whether the models perform poorly because of shifts in the underlying money demand specifications.

III. Real Exchange Rate Models

It has often been asserted that the most robust relationship in empirical exchange rate models is between the real exchange rate and the real interest rate differential.[9] In the next two sections, we confirm that the sign of the correlation between these two variables is indeed typically consistent with the predictions of sticky-price monetary models. However, models featuring the real interest differential (short or long) do not appear to fit out-of-sample significantly better than the random walk model. In this section, we present both in-sample

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8 See, for example, Henriksson and Merton (1983) or Levich (1980). While the direction of change statistics listed in Table 1 is of descriptive interest, a model which forecasts small changes accurately but misses big market turns is not necessarily of any great use. The direction of change statistics is only at all meaningful when the distribution of forecast errors is symmetric.

9 See, for example, the 1984 *Economic Report of the President*. Shafer and Loopesko (1983) present evidence on the relationship between real interest rates and real exchange rates.

and out-of-sample results for real versions of the Dornbusch-Frankel (D-F) and Hooper-Morton (H-M) models. (In the flexible-price monetary model, of course, the real exchange rate is exogenous.) By implementing real versions of the models, we also hope to abstract from misspecification of the money demand functions, which are a crucial building block of the monetary models.

There is more than one way to derive the real versions of the D-F and H-M models. One can, for example, substitute out for (m-m*) in equation (1) by using the underlying money demand equations. After the imposing coefficient constraints, the resulting equation simplifies to one involving the real exchange rate, the real interest differential, and cumulated trade balances. It is instructive, however, to adopt an alternative approach. This approach uses the fact that in the models subsumed in equation (1), all endogenous variables adjust monotonically at the same constant rate to their flexible-price values (along the saddlepath).[10] In particular, this implies that the real exchange rate $q_t \equiv s_t + p_t^* - p_t$ adjusts towards its flexible-price value $\bar{q}_t$ according to

$$(2) \qquad E_t(q_{t+k} - \bar{q}_{t+k}) = \theta^k(q_t - \bar{q}_t), \; 0 < \theta < 1 \;,$$

where $\bar{q}_t$ is the real exchange rate which would prevail at time t if all prices were fully flexible, p is the home-currency price of the domestically-produced good, p* is the foreign-currency price of the (different) good produced abroad, and θ is a speed of adjustment parameter which depends on all the parameters in the model. Note that $\bar{q}_t$ is not in general equal to $E_t\bar{q}_{t+k}$ in (2). However, both the D-F and H-M models impose the assumption that $\bar{q}$ follows a random walk; thus

$$(3) \qquad E_t\bar{q}_{t+k} = \bar{q}_t \;.$$

10 See Obstfeld and Rogoff (1984). There are, of course, more general versions of the Dornbusch model in which variables can adjust at different rates and in which adjustment needn't be monotonic.

In the H-M model, $\bar{q}_t$ is actually posited to be a function of the cumulated current account (which itself is posited to follow a random walk).

Substituting equation (3) into equation (2), one can obtain

$$(4) \qquad q_t = \alpha(E_t q_{t+k} - q_t) + \bar{q}_t \ ,$$

where $\alpha = 1 / (\theta^k - 1) < 0$. To obtain an equation relating real exchange rates to real interest rates, we make use of the uncovered interest parity relation

$$(5) \qquad E_t s_{t+k} - s_t = {}_k r_t - {}_k r_t^* \ ,$$

where ${}_k r_t$ is the k-period nominal interest rate. Equation (5) implies that

$$(6) \qquad E_t(q_{t+k} - q_t) = {}_k R_t - {}_k R_t^* \ ,$$

where the k-period real interest rate ${}_k R_t \equiv {}_k r_t - (E_t p_{t+k} - p_t)$.

Substituting equation (6) into equation (4) yields

$$(7) \qquad q_t = \alpha({}_k R_t - {}_k R_t^*) + \bar{q}_t \ .$$

Adding an exogenous risk premium to equation (5) will add an exogenous forcing term to equation (7).

Equation (7) relates the real exchange rate to the real interest differential and to the flexible-price equilibrium real exchange rate; this is the equation which will be used in our forecasting experiments. [In the H-M model, $\bar{q} = f(\overline{TB}, \overline{TB}^*)$.] Before turning to our empirical results, we wish to contrast our derivation of equation (7) with the popular "identities" approach.[11] This approach begins with the relation between the real interest rate differential and the expected rate of change of the real exchange rate which obtains when uncovered interest rate parity holds, equation (6). Rearranging terms in (6), we have

$$(8) \qquad q_t = -({}_kR_t - {}_kR_t^*) + E_t q_{t+k} \ .$$

In equation (8), the coefficient of the real interest rate differential is one (or, if we annualize the real interest differential, then the coefficient is approximately k). To obtain an "estimating" equation using the identities approach, it is typically assumed that $E_t q_{t+k} \cong E_t \bar{q}_{t+k}$, and then some proxy is used for $E_t \bar{q}_{t+k}$. Because the coefficient on $R - R^*$ is given under the null, it is not clear how to best interpret estimates of (8).

In Table 2, we present in-sample instrumental variable estimates of equation (7) in first-difference form for the dollar/mark, dollar/yen, and dollar/pound rates only. The estimates are based on monthly data for March 1974-June 1984. We have chosen k=3, so that the variable $R - R^*$ represents a three-month real interest differential, which is formed by using the ex post realized three-month inflation differential as a proxy for its expected value (that is, by imposing rational expectations). The resulting equation is then estimated by McCallum's technique.[12] This method would be impractical, of course, if we were to employ long-term real interest rates. (Indeed, it is probably more difficult to measure long-term expected inflation differentials than short-term expected inflation differentials.) The Lagrange multiplier test statistics reported in Table 2 indicate that no further serial correlation is required (with the exception of the H-M model on the dollar/pound rate).

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11 See, for example, Isard (1983).

12 See McCallum (1976). Because the Lagrange Multipler tests of serial correlation do not indicate any residual serial correlation for all but one model, it does not appear likely that we would obtain substantially different results by using a technique such as two-step, two-stage least squares, which would correct for a residual MA error term. See Cumby, Huizinga and Obstfeld (1983).

The good news in Table 2 is that for the dollar/mark and dollar/yen rates, both the real interest differential and the cumulated trade balance variable have the theoretically expected signs. A rise in the real interest differential between dollar and mark assets leads to an appreciation of the dollar; so too does a trade balance surplus. (In the H-M model, a trade balance surplus signals an appreciation of the long-run real exchange rate.) In fact, many of the regressions obtained in this study yield the right sign on the real interest differential (short or long), with the exception of some regressions involving the dollar/mark and dollar/pound rates. This finding provides weak confirmation for the widely-held view that amidst all the volatility of exchange rates, there has been a consistent relationship between real exchange rates and real interest rates. The bad news in Table 2 is that taken individually, the coefficients are not statistically significant. Also, the real interest rate coefficients should always be greater than .25 in absolute value [see equations (4) and (7); the constraint is .25 instead of 1 because the three-month real interest differentials have been annualized.] Last, an F-test for a structural break in November 1980 rejects the null hypothesis that the coefficients have been stable across the two subperiods for only the dollar/pound rate.[13]

The in-sample results reported in Table 2 are reinforced by the rolling regression (forecasting) results presented in Tables 3 and 4. (A tabulation of MAE in place of RMSE again yields a similar picture.) Table 3 lists results for the forty-four month forecasting period November 1980 - June 1984. As with the in-sample results in Table 2, the results in Table 3 are based on three-month real interest differentials; however, *current* three-month inflation is used as a proxy for expected inflation. In contrast to Meese and Rogoff (1983a,b), we do find cases where the structural models outperform the RW model at horizons twelve months and under. Indeed, each model outperforms the RW model for at least one forecast horizon for every exchange rate (with the exception of the D-F model for the dollar/mark and H-M model for

13 The acceptance of the stability hypothesis for all but one equation would appear to be good news for the real exchange rate models. However, the acceptance is more a consequence of the poor fit of the unconstrained model, as it is frequently the case that the signs of the real interest differential and cumulated trade balance variables are incorrect over the two subperiods, March 1973 - October 1980 and November 1980 - June 1984.

the dollar/yen). However, the difference is statistically significant only in the case of the mark/pound rate at one-month horizons.

Table 4 lists RMSE for the forecasting period December 1977 - June 1984 (ninety months). Because we have three years more data than in our previous studies, it is possible to meaningfully examine forecast horizons up to thirty-six months. While one or both of the structural models outperform(s) the RW model for all the exchange rates for at least one forecast horizon, the improvement is never statistically significant.

The results in Tables 3 and 4 are based on a generalized least squares (GLS) estimation procedure with a correction for a first-order autoregressive process (denote the AR coefficient as ρ). Similar results obtain when the models are estimated in first-difference form or by instrumental variables techniques. In Table 5, we present "average" coefficients for the regressions used to construct Tables 3 and 4. These coefficients are formed as follows: for the November 1980 - June 1984 forecasting period, the initial rolling regressing estimates are through October 1980; denote this vector of coefficient estimates by $<\alpha_1>$. Then we add data for November 1980 and form a new set of coefficient estimates, $<\alpha_2>$, etc. The average coefficients are then the sample mean of these estimates:

$$(9) \qquad \frac{1}{44}\sum_{s=1}^{44} <\alpha_s> = <\bar{\alpha}>$$

The coefficients in parentheses in Table 5 are the standard deviations of these estimates; they measure the stability of the coefficient estimates across the forty-four different rolling regression periods. As reported above, the coefficient on the real interest differential typically has the theoretically expected (negative sign), except for the dollar/mark and dollar/pound rates (for both models). The theoretically expected (negative sign) for the cumulated trade balance differential obtains only for the dollar/yen rate (forty-four month horizon) and the mark/yen (both horizons). Not surprisingly, the most stable coefficient is ρ, the coefficient for the first-order autoregressive disturbance term. The coefficient estimates are less than unity but as is well-known, we cannot use conventional significance levels to test this hypothesis (see Section V below).

Summarizing the results of this section, we do find a consistent correlation between the real interest differential and the real exchange rate (though of the wrong sign for the dollar/mark and dollar/pound). Generally speaking, models emphasizing real interest differentials do *not* improve significantly on the random walk model. (Though the D-F model does yield a significant improvement for the mark/pound rate at one-month horizons.)

IV. Alternative Real Exchange Rate Model Specifications

In this section we show that the dollar/yen results of Section III are robust to using long-term real interest differentials in place of short-term differentials, to using a lagged adjustment mechanism in place of an autoregressive disturbance, to using cumulated government deficits in constructing the proxy for the long-term real exchange rate, and to using an estimator which places a greater weight on recent observations. In the interest of brevity, we only present results for the dollar/yen rate.

Models (2) and (6) in Table 6 are the same as the D-F and H-M models of the previous section, except that the long-term government bond differential is used in place of the three-month treasury bill differential. (Note that equation (7) holds for real interest differentials of any maturity.) Inflationary expectations are proxied by the current twelve-month inflation rate. A comparison of Table 6 with Tables 3 and 4 reveals that the D-F model performance is unaffected by the use of long-term rates. (The H-M model performs slightly better at the thirty-six month horizon with the long rates.) A lagged adjustment specification [models (3) and (7)] never yields improvement over the models reported in Tables 3 and 4. (The lagged adjustment specification is obtained from equation (7) by adding q_{t-1} as an additional explanatory variable.) The lagged adjustment model performs quite poorly over the longer forecasting periods. In models (4) and (8) we attempt to allow for coefficient drift by weighting observations by $.95^{t_0-t}$, where t_0 is the final observation period. This modification improves the performance of the D-F model over the forty-four month period, and improves the H-M model at thirty-six month horizons over the ninty-month period.

There has been considerable recent discussion of whether federal government deficits affect the real exchange rate.[14] The models we are considering already incorporate one major channel through which deficits affect exchange rates, namely the real interest rate differential. Federal deficits can also cause changes in the exchange rate by affecting the long-run real exchange rate required to generate an equilibrium current account. (If transitory government deficits cause transitory current account deficits then, in some models, a long-run real exchange rate depreciation may be required to produce a trade balance surplus to offset the service account deficit.) To adequately capture the effects of government deficits, it would be necessary to estimate a fully dynamic model of private saving and portfolio behavior, a task beyond the scope of this paper. However, for descriptive purposes, we chose to include the difference between cumulated home and foreign federal government deficits as an additional explanatory variable in both the D-F and H-M models. As can be observed from the results for models (5) and (9) in Table 6, inclusion of the cumulated deficit variable as a proxy for the long-run real exchange rate yielded little forecasting improvement over the random walk model.[15]

Finally, we considered three purely statistical models as alternatives to the random walk model: (1) a random walk with drift, (2) an ARIMA (0,1,1) or simple exponential smoothing, and (3) a local trend predictor.[16] The last two procedures were employed to test the possibility that there exists exploitable (positive) serial correation in monthly exchange rate

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14 The relationship between deficits and exchange rates has been a major issue in recent editions of the *Economic Report of the President*.

15 Similar results obtain when the Japanese cumulated deficit was converted to dollars using the mean exchange rate over the period. Entering the deficit variables separately yielded similar forecasting results. (The same comments hold for cumulated trade balanaces.) See Appendix A for data sources.

changes. This empirical regularity has been uncovered using *daily* exchange rate data; Dooley and Shafer (1976, 1983) document the profitability of simple filter rules on daily exchange rate data over the modern floating rate period. At *monthly* horizons, however, our alternative statistical models forecast no better than the random walk model without drift, so the results of these experiments are not reported.

It is also known that exchange markets are characterized by periods of relative calm and turbulance (Mussa (1979)), so that exchange rate regression disturbances typically exhibit conditional heteroskedasticity. (See Cumby and Obstfeld (1982), among others.) We did not experiment with estimation techniques robust to this possibility, as it seems unlikely to improve the accuracy of point forecasts. In terms of forecasting applications, the major contribution of the literature on conditional heteroskedasticity has been to refine our estimates of the k-step ahead forecast error variances. Nevertheless, the forecasting performance of models allowing for conditional heteroskedasticity merits further attention.

V. Tests of the Cointegration of Real Exchange Rates and Real Interest Differentials

It is clear from the evidence presented in the preceding sections that real exchange rates exhibit behavior which at least borders on nonstationarity. In this section, we use nonstructural methods to investigate whether the same factor which introduces nonstationarity in real exchange rates introduces a similar order of nonstationarity in real interest differentials. Our tests are based on the concept of cointegration, first introduced by Granger (1983), and

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16 The sample autocorrelations and partial autocorrelations of monthly exchange rate changes do not typically suggest models other than the ARIMA (0,1,1) or a random walk. The local trend predictor is based on a linear trend estimated from all observations that have occurred since the last "turning point." A turning point is defined by the occurrence of at least an α% change of direction from the previous change of the exchange rate. A value of $\alpha = 5\%$ typically produced the lowest RMSE exchange rate forecasts for the local trend model.

expanded upon in Granger and Engle (1984). A vector time series is cointegrated of order (d,b) if each element needs to be differenced d times to achieve stationarity, but yet there exists a (not necessarily unique) linear combination of the two vectors which only needs to be differenced (d-b) times to achieve stationarity. It is not unusual to find evidence of unit roots in the autoregressive representation of asset price data. In our context, we are interested in exploring the possibility that the nonstationarity we find in real exchange rates can be accounted for by the nonstationarity of real interest differentials.

In Table 7, we report results of a test for a unit root in autoregressive representations of the real exchange rate, and short- and long-term real interest differentials, for the U.S.-German, Japanese, and U.K. data sets. The results in Table 7 are based on regressions of the following general form.

$$(10) \qquad (q_t - q_{t-1}) = b_0 + b_1 q_{t-1} + b_2(q_{t-1} - q_{t-2})$$

$$+ b_3(q_{t-2} - q_{t-3}) + \varepsilon_t \ .$$

The b_i are constant parameters and ε_t is a white noise disturbance. If the autoregressive representation of q_t contains a unit root (is integrated of order one), the t-ratio for b_1 should be consistent with the hypothesis $b_1 = 0$. Conventional t-tables are inappropriate for this hypothesis test, so we use the results of Dickey and Fuller (1979) and the tabulated distribution in Fuller (1976, p. 373) to interpret the t-ratio. The approximate critical values for this ratio, using a five percent significance level and T = 100, are -2.89 and -.05. From inspection of Table 7, we can therefore accept the null hypothesis of a unit root for the real exchange rate and real long-term interest differential equations. But we reject it for real short-term interest differential equations for the dollar rates vis-a-vis the pound, yen and mark.

It is puzzling that real long term interest differentials appear to be nonstationary, given the current system of highly integrated capital markets. This anomaly is also characteristic of the nominal long-term government bond rate differential, at least for the data used in this study. As such, we cannot attribute the apparent nonstationarity in the long-term interest differential to our expected inflation proxy. Both the nominal and real *short-term* interest differentials do appear to be stationary in levels.[17] The stationarity of short-term rates is consistent with Mussa's (1979) observation that forward premia exhibit much less volatility than any of the individual variables related by covered interest parity. (Nominal k-period interest rate differentials equal the forward premium for regularly traded forward rates, k=1,3,6 or 12 months.) The puzzling nonstationarity of long-term interest differentials might be a consequence of the lack of homogeneity of our long-term bond yields, the use of on-shore instead of Euromarket rates, or the lack of organized forward markets for long maturities.

The most probable explanation of the nonstationarity of long-term real interest rate differentials is the low power of the unit root tests to detect borderline stationary alternatives. We know for example that it cannot be literally true that real exchange rates have a unit root, but a coefficient of .95 in a regression of the real exchange rate on a single lag of itself is quite plausible. It suggests that deviations from purchasing power parity take five years to damp down. Given our sample size, the probability of a type II error (accept a coefficient of 1.0 when it is really .95) is roughly 80%; see Evans and Savin (1981, p. 771).

The nice feature of the cointegration tests reported below is that they can be meaningful even if there is high probability of a type II error. Effectively they test whether some linear combination of the "large variance components" of real exchange rates and real interest differentials effectively cancel one another, leaving an "equilibrium error" with small variance. If there does not exist a linear combination of real exchange rates and real interest differentials that is itself a stationary process, it suggests the relationship between the two variables is at best tenuous, or that a highly variable factor has been omitted from the real exchange rate - real interest differential relation.

17 Test statistics for the hypothesis $b_1 = 0$ in (10) for the U.S.-German, U.S.- Japanese, and U.S.-U.K. three month nominal interest rate differential are -3.08, -1.82, and -3.54, respectively. The failure to reject the unit root for the U.S.-Japanese interest differential might be a consequence of Japanese capital controls in effect over much of the sample period.

Since real short-term interest differentials appear to be stationary in levels, real exchange rates and real short-term interest differentials cannot be cointegrated. To test whether real exchange rates and real long-term interest rates are cointegrated we employ the preferred tests of Granger and Engle (1984). An "equilibrium regression" of q_t on the real interest differential is run (or the reverse regression) and the residuals are examined for nonstationary behavior.[18] The results in Table 8 suggest that real exchange rates and long-term real interest rate differentials are *not* cointegrated. Given this evidence, we should not necessarily expect to find damped (stationary) forecast errors when predicting real exchange rates with real interest rate differentials. Our earlier estimates of the D-F and H-M models confirm this fact, as the first order serial correlation parameter is always close to unity.

Real interest differentials and real exchange rates are linked by international parity conditions, so our findings of no cointegration suggest that a variable omitted from relation, possibly the expected value of some future real exchange rate,[19] must have large variance as well. Alternatively, the set of shocks inducing near nonstationarity in real exchange rates can not be the same as the set of shocks impinging on real interest rate differentials.

VI. Conclusions

The results we have presented are slightly more favorable than the results of our earlier studies. We do find that the real exchange rate and the real interest differential have the theoretically anticipated sign (except for some regressions involving the dollar/mark and dollar/pound rates). However, the relationship is not statistically significant, and real interest

18 The forward regression (real spot rate on the real interest differential) or the reverse regression can be used to test for cointegration. In the limit the data matrix becomes colinear so the coefficient of the reverse regression is equal to the reciprocal of the coefficient of the forward regression. Simultaneity does not pose a problem for the test procedure either; see Granger and Engle (1984).

19 Isard (1983, pp. 22-23) discusses the identity that links real exchange rates, real interest differentials, the expected future real exchange rate, and the risk premium.

differentials do not provide significant improvement over a random walk model when forecasting real exchange rates (except in a few isolated cases). Thus, the results of our forecasting experiments corroborate the in-sample findings of both Alder and Lehmann (1983) and Hakkio (1984). These authors also provide evidence for the martingale behavior of real exchange rates.

One popular explanation of why monetary models perform so poorly is that the disturbances impinging on exchange markets are predominantly real.[20] Thus models which focus primarily on monetary disturbances should not be expected to explain very much. While this hypothesis deserves further attention, it is not yet certain whether it will be helpful in building better empirical exchange rate models. It has proven extremely difficult to identify which real factors (such as technology shocks or changes in preferences) affected exchange rates over what periods. Another popular current explanation of the failure of monetary exchange rate models is the existence of self-fulfilling expectations or exchange market bubbles. This is an active area of theoretical research.[21] Empirical evidence for the existence of bubbles or extraneous variables in exchange markets is mixed, and this explanation of the failure of monetary exchange rate models also merits further research.[22]

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20 See, for example, Flood (1981) or Barro (1983).

21 Obstfeld and Rogoff (1983, 1985) present the theoretical case for ruling out divergent (explosive or implosive) bubbles.

22 See Meese (1986), West (1985), and Woo (1984).

Appendix A

The data is sampled monthly from March 1973 - June 1984. All the asset market data are end-of-month point sample from Federal Reserve Board (FRB) data base. [In Meese and Rogoff (1983a,b) asset market data were drawn from the same day as foreign money supply figures. The results do not seem to depend much on this issue, and end-of-month data are much more conveniently obtained.] The exchange rates are N.Y. noon bids. The short-term interest rates are three-month interbank rates, and long-term rates are five- to ten-year government bond yields. The trade balance and industrial production data are from the OECD. CPI data is from the FRB. Japanese fiscal deficit data are taken from the Bank of Japan's *Economic Statistics Monthly*, table 80, section 3, item "balance of ordinary receipts and payments with the public." U.S. fiscal deficit data are taken from *International Financial Statistics*, line 80. All raw series are seasonally unadjusted.

Appendix B

Let $e(1,t)$ and $e(2,t)$ denote the period t forecast errors from models (1) and (2) respectively. Let $A = \{a_{ij}\}, i,j = 1,2$ denote the covariance matrix of forecast errors. Define $x(t) = e(1,t) - e(2,t)$ and $y(t) = e(1,t) + e(2,t)$. A test of the null hypothesis $a_{11} = a_{22}$ is easily conducted using $x(t)$ and $y(t)$, as $a_{11} = a_{22}$ only when $cov(x(t),y(t))$ is zero. Assuming the vector process $(e(1,t), e(2,t))$ is independent and identically distributed as a $N(0,A)$, then a uniformly most powerful unbiased test of $a_{11} = a_{22}$ can be based on the sample correlation coefficient of $x(t)$ and $y(t)$, $\hat{\rho}(x,y)$. Specifically, a test of $a_{11} = a_{22}$ can be based on a statistic with a t-distribution with T-2 degrees of freedom:

$$t_{(T-2)} = \hat{\rho}(x,y)(T-2)^{\frac{1}{2}} / (1 - \hat{\rho}(x,y)^2)^{\frac{1}{2}} \ ,$$

for $T > 2$. (See Hogg and Craig (1970), pp. 339-342, or Granger and Newbold (1977), pp. 281-282.) T is the number of known forecast errors.

This procedure can only be appiled to unbiased, serially uncorrelated, normally distributed forecast errors. For multiple (k-step-ahead) forecast horizons, a sequence of forecast errors will in general follow a moving average (MA) process of order $(k-1)$. In this case, we can still use $x(t)$ and $y(t)$ to construct a test of $a_{11} = a_{22}$, but asymptotic distribution theory for time series is required. Given our assumptions, $x(t)$ and $y(t)$ are themselves $MA(k-1)$ stationary processes. Using results in Hannan (1970), chapter 4, we know that the sample covariance of $x(t)$ and $y(t)$, $\hat{cov}(x(t),y(t))$, is a consistent estimator of the population covariance.

In addition,

(B1) $$\sqrt{T}\left[c\hat{o}v(x(t),y(t)) - cov(x(t),y(t))\right] \xrightarrow{L} N(0,B),$$

where B is a complicated function of the autocovariances and fourth cumulants of the vector process (x(t), y(t)); see Hannan (1970), p. 209. Assuming normality of forecast errors, we can ignore fourth cumulants. Exploiting the $MA(k-1)$ behavior of optimal k-step-ahead forecasts, B can then be consistently estimated by

(B2) $$\hat{B} = \sum_{s=-k+1}^{k-1} (1 - |s| / T)\Big\{c\hat{o}v(x(t),x(t-s))c\hat{o}v(y(t),y(t-s)) +$$

$$c\hat{o}v(x(s),y(t-s))c\hat{o}v(y(s),x(t-s))\Big\}.$$

The test statistics reported in the paper are thus

(B3) $$c\hat{o}v(x(t),y(t)) / (\hat{B} / T)^{\frac{1}{2}},$$

which is approximately N(0,1) for large T.

An alternative estimator of B that does not require normality of forecast errors can be calculated using the generalized method of moments (GMM) methodology described in Hansen and Singleton (1982) or White and Domowitz (1984). The estimate of the covariance of $x(t)$ and $y(t)$ is chosen to satisfy the orthogonality condition

(B4) $$0 = \frac{1}{T}\Sigma\left[c\hat{o}v(x(t),y(t)) - (x(t)-\bar{x})(y(t)-\bar{y})\right]$$

A consistent method of moments estimator of B is then

$$\text{(B5)} \qquad \hat{B} = \sum_{s=-k+1}^{k+1} \frac{1}{T} \sum_{t} (1 - |s| / T) \left[x(t)y(t)x(t-s)y(t-s) - c\hat{o}v\ (x(t), y(t))^2 \right] .$$

The test statistic (B3) is again appropriate. The use of (B2) or (B5) to estimate B made no qualitative difference in the results reported in Tables 1, 3, 4, and 6.

Several caveats are in order. First, multiple step-ahead forecasts (large k) are typically biased; the mean error of these forecasts (not reported in the text) generally increases in absolute value with k. Second, the assumption of normality of forecast errors is also suspect for large k (the normality assumption is not needed for the GMM methodology). Third, our test statistic (B3) has not been subjected to any simulation experiments. Last, the rolling regression estimates of $e(1,t)$ and $e(2,t)$ are based on different sample sizes. To correct for small sample bias in (B3), the sample size can be kept constant across all regressions, or forecast errors can be weighted by $\sqrt{(T-t)/T}$, where T is the last (largest) sample, and $(T-t)$ is the number of observations in the regressions that generated $e(1,t)$ and $e(2,t)$.

Appendix C

Average Coefficient Estimates for Rolling Regressions[a]

Exchange rate	Model	Constant	Money supplies $m-m^*$	Incomes $y-y^*$	Short-term rates $r_S-r_S^*$	Long-term rates $r_L-r_L^*$	Cumulated trade balances $TB-TB^*$	A.R. coefficient ρ
	D-F	.25	-3.5	5.7	-.66	-.11		.96
$/mark		(.13)	(5.4)	(2.3)	(.06)	(.01)		(.002)
	H-M	.26	-3.1	-5.6	-.7	-.1	.12	.96
		(.11)	(5.7)	(2.8)	(.1)	(.05)	(1.2)	(.002)
	D-F	.12	8.8	17.9	-.53	.04		.98
$/yen		(.04)	(7.4)	(3.5)	(.13)	(.00)		(.002)
	H-M	.12	5.0	22.3	-.6	.04	-.17	.97
		(.08)	(9.5)	(5.0)	(.3)	(.00)	(.07)	(.01)
	D-F	-.04	31.6	-6.1	-.04	-.05		.96
$/pound		(.16)	(3.3)	(1.4)	(.01)	(.00)		(.006)
	H-M	-.05	22.0	-5.1	-.02	-.05	11.0	.97
		(.11)	(2.8)	(1.6)	(.01)	(.00)	(2.5)	(.005)
	D-F	.25	14.6	-16.5	-.27	.01		.93
mark/yen		(.16)	(4.6)	(3.3)	(.12)	(.00)		(.004)
	H-M	-.21	12.6	-11.7	-.35	.01	-.17	.92
		(.10)	(4.9)	(2.6)	(.13)	(.00)	(.02)	(.006)
	D-F	-.30	29.4	2.9	-.12	.01		.95
mark/pound		(.03)	(2.5)	(1.6)	(.06)	(.01)		(.001)
	H-M	-.18	12.1	2.0	-.14	.02	-9.9	.96
		(.06)	(4.3)	(1.8)	(.05)	(.00)	(.6)	(.004)

Coefficient[b]

[a] Each initial rolling regression is estimated (GLS) over March 1973 - October 1980. Then a month of data is added, and the coefficients are re-estimated. The final estimation is over March 1973 - October 1984. Denoting $\alpha(i,j)$ as the estimate for $\alpha(i)$ for regression period j, the table entries are

$$\bar{\alpha}(i) = \frac{1}{44} \sum_j \alpha(i,j)$$

that is, the average coefficient estimates. Entries in parentheses are the standard deviation of these estimates, i.e.,

$$\hat{\sigma}_\alpha = \left[\frac{1}{43} \sum_j (\alpha(i,j) - \bar{\alpha}(i))^2 \right]^{\frac{1}{2}} .$$

b All table entries are multiplied by 10^2, except those in the cumulated trade balance column (which are multiplied by 10^4), and those in the "ρ" column (which have not been adjusted).

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Table 1

Root mean square forecast errors for the logarithm of the nominal exchange rate, 11/80 - 6/84.[a]

Exchange rate	Model: Horizon	Random walk	Dornbusch-Frankel		Hooper-Morton		Correctly predicts direction of change[b] (percent) D-F	H-M
	1 month	3.1	3.1	(-.1)	3.2	(-.2)	48	50
$/mark	6 months	7.9	8.4	(-.5)	8.5	(-.1)	56	56
	12 months	8.7	11.1	(-1.0)	11.4	(-1.0)	10	03
	1 month	3.5	3.3	(1.4)	3.5	(-.1)	55	54
$/yen	6 months	7.8	7.0	(1.6)	7.7	(.1)	67	50
	12 months	9.0	7.5	(1.6)	8.7	(.1)	81	64
	1 month	3.0	3.1	(-.4)	2.8	(1.5)	54	75*
$/pound	6 months	9.1	9.2	(-.8)	7.5	(1.4)	49	92
	12 months	12.8	13.6	(-.8)	9.4	(1.2)	52	100
	1 month	3.2	3.3	(-.2)	3.2	(.0)	57	66*
mark/yen	6 months	6.7	9.3	(-1.2)	7.4	(-.6)	38	64
	12 months	10.1	13.2	(-.8)	8.1	(.8)	36	80
	1 month	3.1	3.0	(.2)	2.9	(.8)	61 *	57
mark/pound	6 months	6.6	5.9	(1.2)	6.4	(.5)	69	69
	12 months	6.6	5.4	(1.1)	7.0	(.8)	91	82

[a]The 44-month forecasting period runs from 11/80 - 6/84. Initial rolling regression runs 3/73 - 10/80. The entries in parentheses are for the asymptotically N(0,1) statistic discussed in appendix B; they compare the RMSE of the D-F or H-M model with that of the RW.

[b]The direction of change statistic indicates the number of times each model correctly predicted the direction of change of the exchange rate. Using a binomial test for one-month horizons only, the 95% critical value is 61%. An asterisk denotes significance at the five percent level.

Table 2

In-sample coefficient estimates and stability tests for real versions of Dornbusch-Frankel and Hooper-Morton models.[a]

Coefficient:		Constant ($\times 10^2$)	Real Interest diff. R - R*	Cumulated trade balance TB - TB* ($\times 10^4$)	LM-test $x^2(12)$[b]	Standard Error of Regression	Stability test[c] F(13,105) F(14,104)
Real exchange rate (q)	Model						
$/mark	D - F	1.4 (1.3)	-.25 (-.24)		13.1	.031	.93
	H - M	0.9 (0.8)	-.40 (-.40)	-13. (-.64)	14.2	.032	.88
$/yen	D - F	2.1 (1.7)	-.34 (-.98)		8.1	.033	1.60
	H - M	1.9 (1.4)	-.33 (-.94)	-8.0 (-.58)	8.7	.034	1.57
$/pound	D - F	1.6 (1.6)	.01 (.04)		20.1	.032	4.32*
	H - M	1.7 (1.7)	.04 (.14)	17. (.88)	23.3*	.032	1.63

[a] Monthly data, March 1973 - June 1984. All regressions are reported in first-difference form. Estimates are based on McCallum's technique, since the realized three-month inflation rate is used as proxy for expected inflation in forming the three-month real interest differentials. Instruments for the H-M model are lagged four R - R*, TB - TB* and q, and current value and four lags of relative incomes, y - y*. Instruments for the D - F model are the same except for lag four of TB - TB*. t-statistics are in parentheses. Each regression also includes monthly seasonal intercept dummies, the coefficients of which are not reported.

[b] The LM-test for serial correlation is used since the regressors in the real exchange equation cannot be regarded as being independent of the disturbance. A typical LM-statistic is calculated as sample size multiplied by R^2 from the regression of the residual on the regressors of the appropriate structural model plus 12 lags of the residuals. See Harvey (1981).

[c] Tests for stability compare slope coefficients, and seasonal dummies pre- and post-November 1980. The null hypothesis of no change is rejected at the 95% significance level for only the D - F model of the $/pound rate. Tests are based on a "Lagrange multiplier" F statistic, as the denominator is an estimate of the residual variance of the constrained (no structural break) model. An asterisk denotes significance at a five percent level.

Table 3

Root mean square forecast errors
for real exchange rate models, 11/80 - 6/84.[a]

Exchange rate	Horizon	Random Walk	Dornbusch-Frankel	Hooper-Morton	Correctly predicts direction of change[b] (percent) D-F	H-M
$/mark	1 month	3.2	3.4 (-0.1)	3.1 (0.1)	45	55
	6 months	8.5	11.2 (-0.7)	7.8 (0.5)	31	69
	12 months	9.6	17.0 (-2.3)*	8.9 (0.2)	0	42
$/yen	1 month	3.6	3.5 (0.1)	3.6 (-0.0)	55	50
	6 months	8.8	8.5 (0.1)	8.9 (-0.0)	56	54
	12 months	10.8	11.1 (-0.1)	12.2 (-0.2)	45	36
$/pound	1 month	3.2	3.0 (0.2)	3.0 (0.2)	57	68*
	6 months	9.3	8.6 (0.1)	7.9 (0.3)	49	74
	12 months	13.4	12.5 (0.1)	10.4 (0.5)	52	100
mark/yen	1 month	3.1	3.0 (0.6)	3.0 (0.1)	66*	68*
	6 months	6.4	7.6 (-0.8)	6.0 (0.1)	46	74
	12 months	9.8	11.5 (-0.6)	6.7 (1.2)	42	91
mark/pound	1 month	3.1	2.7 (2.0)*	2.8 (1.8)	68*	57
	6 months	7.0	6.4 (0.8)	6.5 (0.8)	62	62
	12 months	7.3	5.9 (0.8)	6.2 (0.8)	76	88

[a]See footnote a to Table 1.

[b]See footnote b to Table 1.

* denotes significance at a 5% level.

Table 4

Root mean square forecast errors
for real exchange rate models, 1/77 - 6/84.[a]

Exchange rate	Model Horizon	Random Walk	Dornbusch-Frankel	Hooper-Morton	Correctly predicts direction of change (percent)[b] D-F	H-M
	1 month	3.4	3.4 (-0.2)	3.3 (0.3)	57*	53
$/mark	6 months	8.2	9.7 (-1.1)	8.7 (-0.4)	45	59
	12 months	13.0	16.1 (-1.0)	13.7 (-0.3)	29	47
	36 months	31.5	30.4 (0.5)	24.9 (0.8)	62	82
	1 month	3.7	3.7 (0.2)	3.6 (0.6)	49	51
$/yen	6 months	10.6	10.6 (0.1)	10.7 (-0.0)	53	52
	12 months	15.4	14.5 (0.6)	15.2 (0.1)	52	42
	36 months	20.1	16.2 (0.9)	20.2 (-0.0)	51	44
	1 month	3.2	3.0 (1.1)	3.0 (1.5)	56*	61*
$/pound	6 months	8.2	8.9 (-0.7)	8.5 (-0.3)	39	49
	12 months	13.4	14.0 (-0.3)	13.5 (-0.1)	34	53
	36 months	29.5	23.9 (1.0)	27.2 (0.5)	64	56
	1 month	3.6	3.5 (0.3)	3.5 (0.6)	63*	62*
mark/yen	6 months	10.7	10.6 (0.1)	10.2 (0.7)	59	65
	12 months	16.4	15.5 (0.6)	13.9 (1.6)	53	87
	36 months	15.5	17.5 (-0.6)	12.2 (1.0)	45	58
	1 month	3.0	3.0 (0.6)	3.0 (0.5)	60*	53
mark/	6 months	8.1	8.7 (-0.8)	8.8 (-1.0)	47	49
pound	12 months	12.4	13.5 (-0.6)	13.5 (-0.8)	44	48
	36 months	28.6	30.6 (-0.4)	29.0 (-0.3)	24	45

[a]Initial rolling regressions run 3/73 - 12/76. See footnote a to Table 1.

[b]The direction of change statistic indicates the number of times each model correctly predicted the direction of change of the exchange rate. Using a binomial test for one-month horizons only, the 95% critical value is .56.

* denotes significance at a 5% level.

Table 5

Average coefficient estimates for real exchange rate models[a]

Exchange rate	Model	Constant ($\times 10^3$)	$R - R^*$	$\overline{TB}-\overline{TB}^*$ ($\times 10^4$)	ρ
3/73 - 11/80 - 6/84.					
$/mark	D-F	0.8 (1.9)	.11 (.02)		.96 (.004)
	H-M	2.1 (1.0)	.13 (.02)	6.6 (.7)	.96 (.002)
$/yen	D-F	1.4 (1.0)	-.04 (.01)		.97 (.004)
	H-M	1.6 (1.3)	-.04 (.01)	-.01 (.1)	.97 (.002)
$/pound	D-F	1.4 (1.4)	.23 (.01)		.97 (.01)
	H-M	1.8 (1.1)	.23 (.01)	7.0 (2.4)	.98 (.01)
mark/yen	D-F	0.6 (1.1)	-.06 (.01)		.95 (.004)
	H-M	0.7 (0.6)	-.06 (.01)	-.18 (.02)	.94 (.01)
mark/pound	D-F	1.2 (0.6)	-.02 (.01)		.97 (.01)
	H-M	-0.2 (0.5)	-.01 (.01)	6.8 (1.6)	.97 (.01)
3/73 - 1/77 - 6/84					
$/mark	D-F	4.5 (3.9)	.17 (.06)		.94 (.03)
	H-M	4.0 (4.8)	.17 (.05)	4.6 (2.6)	.92 (.05)
$/yen	D-F	2.0 (1.7)	-.03 (.02)		.96 (.02)
	H-M	2.7 (2.8)	-.03 (.01)	.06 (.1)	.96 (.02)
$/pound	D-F	0.3 (2.3)	.22 (.02)		.96 (.01)
	H-M	0.6 (2.6)	.21 (.03)	6.4 (3.8)	.97 (.01)
mark/yen	D-F	-0.6 (1.8)	-.05 (.02)		.96 (.01)
	H-M	-0.4 (1.7)	-.06 (.02)	-.22 (.1)	.94 (.01)
mark/pound	D-F	-1.7 (3.3)	-.08 (.09)		.97 (.01)
	H-M	-2.8 (2.8)	.07 (.09)	5.0 (3.1)	.97 (.01)

[a]Each initial rolling regression is estimated (GLS) over 3/73 - 10/80 (12/76). Then a month of data is added, and coefficients are updated, etc. Forty-four (ninety) sets of coefficient estimates are thus obtained; table entries are the average of these estimates. Entries in parantheses are the standard deviations of these forty-four (ninety) element vectors.

Table 6

Root-mean square errors for the real $/yen exchange rate[a]

		November 1980 - June 1984			January 1977 - June 1984			
	Horizon:	1 month	6 months	12 months	1 month	6 months	12 months	36 months
Model								
(1)	Random walk	3.6	8.8	10.8	3.7	10.6	15.4	20.1
(2)	D-F with long rate differential	3.5 (0.4)	8.5 (0.4)	11.1 (-0.2)	3.7 (0.1)	10.6 (0.1)	14.5 (0.6)	16.2 (0.9)
(3)	D-F with lagged adjustment	3.6 (0.0)	9.1 (-0.5)	12.4 (-0.8)	3.7 (-0.1)	13.4 (-1.6)	28.1 (-2.1)*	b
(4)	D-F with geometric weights	3.6 (-0.2)	8.1 (1.5)	9.2 (1.2)	3.8 (-0.6)	10.9 (-0.4)	15.5 (-0.0)	19.4 (0.2)
(5)	D-F with cumulated deficits	3.6 (-0.1)	9.2 (-0.5)	12.5 (-0.7)	3.7 (0.0)	10.8 (-0.2)	15.3 (0.1)	24.9 (-1.0)
(6)	H-M with long rate differential	3.6 (-0.1)	8.9 (-0.2)	12.2 (-0.6)	3.6 (0.6)	10.7 (-0.1)	15.2 (-0.1)	20.1 (-0.0)
(7)	H-M with lagged adjustment	3.8 (-1.1)	11.8 (-1.6)	20.2 (-1.4)	3.8 (-0.7)	14.6 (-2.1)*	30.6 (-2.5)*	b
(8)	H-M with geometric weights	3.6 (-0.0)	9.8 (-0.4)	17.4 (-0.9)	3.7 (0.2)	11.5 (-0.5)	19.5 (-1.1)	12.3 (1.0)
(9)	H-M with cumulated deficits	3.6 (-0.3)	9.2 (-0.6)	12.7 (-0.8)	3.7 (0.0)	11.3 (-0.9)	16.7 (-0.7)	23.1 (-0.8)

[a] The entries in parentheses are for the asymptotically N(0,1) statistic discusssed in Appendix B. A negative entry indicates that the forecast error variance of the random walk is smaller than that of the structural model.

[b] The long run stock adjustment forecasts were nonsensical because the coefficient on the lagged real exchange rate was not constrained to be less than one.

Table 7

Tests for unit roots using equation (10).[a]

Dependent variable		$\hat{b}_1$	"DF t-ratio" for b_1	$\hat{b}_2$	$\hat{b}_3$
\$/mark	q_t	-.012	-.613	.105	.016
"	R_t (short)	-.169	-3.336*	.185	-.079
"	R_t (long)	-.016	-.836	-.027	.082
\$/pound	q_t	-.024	-1.161	.096	.060
"	R_t (short)	-.301	-4.931*	.300	-.076
"	R_t (long)	-.077	-2.500	.139	.221
\$/yen	q_t	-.035	-1.583	.147	-.017
"	R_t (short)	-.347	-4.970*	.194	.149
"	R_t (long)	-.060	-2.200	.005	-.087

[a] Regressions with q_t and R_t (short) as dependent variable are estimated over 9/73 - 6/84 (130 observations), while regressions with R_t (long) as regressand are estimated over 6/74 - 6/84 (121 observations). In all cases, no further lags of the dependent variable were necessary to whiten the residuals. An asterisk denotes rejection of the unit root hypothesis using a least a 5% significance level. All regressions included a constant and seasonal dummies.

Table 8

Tests for cointegration of real exchange rates

and real long-term interest differentials[a]

Tabulated results are based on the following regression:

$$\text{(T1)}\quad q_t = \text{constants} + c.R_t \text{ (long)} + \text{disturbance.}$$

Test 1: Reject the hypothesis of no cointegration at approximately a 5% significance level if the Durbin-Watson (DW) statistic exceeds .28; see Granger and Engle (1984) tables II and III.

Test 2: Reject the hypothesis of no cointegration at approximately a 5% significance level if the Dickey-Fuller (DF) t-ratio of equation (8) applied to the residuals of (T1) is less than -3.2; see Granger and Engle (1984) tables II and III.

	Dependent variable q_t	$\hat{c}$	Test 1 DW statistic	Test 2 DF statistic
	$/mark	-4.345	.163	-2.251
real	$/pound	.018	.048	-1.015
	$/yen	-2.058	.118	-2.023

[a] All regressions are estimated over 6/74 - 6/84 (121 observations) and include a constant and 11 seasonal dummies.