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THE OUT-OF-SAMPLE FAILURE OF EMPIRICAL EXCHANGE RATE MODELS: SAMPLING ERROR OR MISSPECIFICATION?

by

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I. Introduction[1/]

A companion study [Meese-Rogoff (1981)] compares the out-of-sample fit of various structural and time series exchange rate models, and finds that the random walk model[2/] performs as well as any estimated model at one to twelve month horizons for 1970's dollar/mark, dollar/pound, dollar/yen and trade-weighted dollar exchange rates.[3/] The structural models perform poorly ever though their forecasts are purged of all uncertainty concerning the future paths of their explanatory variables by using actual realized values.

The present study demonstrates that the dismal performance of the structural models is not attributable to the sample distribution of the coefficient estimates. We rule out that explanation by showing that the models (with autoregressive error terms) perform poorly at one to twelve month forecast horizons over a wide range of coefficient values. These values are based on the theoretical and empirical literature on money demand and purchasing power parity. Since the coefficent-constrained models only require estimation of the intercept terms, it is possible to look at longer forecast horizons here than in our other study. There the relative superiority of the random walk model over the structural models diminishes as the forecast horizon approaches twelve months. The present study explores the possibility that the structural models may improve on the random walk model forecasts at horizons of twelve to thirty-six months.

The main part of the paper is contained in section 3, which discusses the coefficient-constrained experiments. In section 2, vector autoregressions (VAR) are used to identify the factors that influence the exchange rate over short versus long horizons. The results from the VAR experiments also highlight the difficulties in finding legitimate instruments with which to estimate the structural models, thus motivating the constrained-coefficient approach of section 3. Section 4 asks which of the common building blocks of the structural models is most likely to have failed. It appears that the breakdown of empirical exchange rate equations is the international counterpart of the breakdown of money demand specifications.

2. Decomposing the Forecast Error Variance of the Exchange Rate at Long and Short Horizons

Before proceeding to tests of the representative structural exchange rate models, we first examine a vector autoregression consisting of the exchange rate and the explanatory variables of these models: relative money supplies, relative outputs,[4/] relative short-term and long-term interest rates, and trade balances. The VAR is a tool for analyzing the relative importance of the explanatory variables in exchange rate model forecasts at both short and long horizons. As a by-product, the VAR also provides information on whether the conventional exogeneity assumptions used in estimation of the structural models are appropriate.

A convenient normalization for estimation of the VAR is one in which the contemporaneous value of each variable is regressed against lagged values of all the variables; e.g., the exchange rate equation is given by

$$(1) \quad s_t = a_{i1}s_{t-1} + a_{i2}s_{t-2} + \ldots a_{in}s_{t-n} + B'_{i1}X_{t-1} + B'_{i2}X_{t-2} + \ldots B'_{in}X_{t-n} + u_{it},$$

where s_t is the (logarithm of the) exchange rate at time t and X_{t-j} is a vector of lagged values of the other included variables (listed above). Expressing the VAR system in the form of equation (1) facilitates estimation, as ordinary least squares equation by equation is an efficient estimation strategy. This normalization does not, however, preclude contemporaneous interactions between the variables, as these effects are captured in the covariance matrix of the disturbance terms u_{it}. The uniform lag length n across all (seven) equations is estimated using Parzen's (1975) lag length selection criterion.[5]

We estimate the VAR model for the dollar/mark, dollar/pound, and dollar/yen exchange rates over the floating rate period; the data consist of monthly observations for March 1973 through June 1981 (our seasonal adjustment procedures are described in the data appendix). Once having obtained the coefficient estimates, the dynamic interactions among the variables are most easily studied with the use of the moving average (MA) representation, which is derived by inverting the autoregressive (AR) representation to express each of the endogenous variables in terms of the disturbances or innovations [the u_{it} in (1) for example]. Studying the MA representation is complicated by the fact that the disturbance terms in the MA (or AR) representation are in general contemporaneously correlated; see Sims (1980) or Fischer (1981). So in order to simulate a "typical" shock to a given variable it is necessary to recognize that the expectation of other disturbances in the system, conditional on the particular shock of interest, are usually nonzero. Unfortunately, for two correlated disturbances $z_1(t)$ and $z_2(t)$, if the $E[z_1(t)| z_2(t) = 1] = \alpha$, α an arbitrary constant, it is not in general true that $E[z_2(t)| z_1(t) = \alpha] = 1$. Because of this fact, there is no unique way to simulate "typical" shocks to these systems of endogenous variables when contemporaneous variable interactions are present. (In other words, when the covariance matrix of the disturbance terms is nondiagonal.) In order to identify a typical shock to the VAR system with a particular variable, we will follow the Sims' (1980) procedure of specifying a variable ordering a priori. The variable ordering essentially specifies that the first variable is predetermined with respect to all other variables, that the second variable is predetermined with respect to all but the first, etc. The identification of the VAR systems is pursued in greater detail in the technical appendix.

The multi-horizon forecast error variance decompositions listed in tables 1-3 are based on a variable ordering with the logarithm of U.S. to foreign relative money supplies $m-\overset{*}{m}$ first, followed by the logarithm of relative outputs $y-\overset{*}{y}$, the short-term interest differential $r_S - \overset{*}{r}_S$, the long-term interest differential $r_L - \overset{*}{r}_L$, the U.S. and foreign trade balances TB and $T\overset{*}{B}$, and the logarithm of the dollar price of foreign currency s. In tables 4-6 the variable ordering is reversed. In the U.S.-German system, the largest estimated contemporaneous correlation is 44% between the short and long-term interest differential equations. The other estimated contemporaneous correlations range from 5 to 20%. These results suggest that the variable ordering is potentially important in the U.S.-German VAR system. And indeed there are some differences between the U.S.-German VAR systems, regular versus reverse order, at both short and long forecast horizons. Note that in the regular (reverse) order system, exchange rate and long-term interest rate innovations account for 78.6% (93.7%) and 12.8% (4.9%) of the one-month-ahead forecast error variance of the exchange rate, and 48.1% (60.1%) and 15.4% (10.5%) of the 36-month-ahead forecast error variance of the exchange rate. However, the significance of these differences cannot be ascertained from tables 1 and 4. We have not yet performed the requisite (expensive) stochastic simulations to obtain estimates of the dispersion of these forecast error variance decompositions. Of course, the data necessarily contain less information about long-run variable interactions than short-run. Similar observations apply to the U.S.-U.K. and U.S.-Japanese VAR systems.

A second important observation to be made from tables 1-6 is that no variable appears to be exogenous to the VAR system. Abstracting from coefficient uncertainty, an exogenous variable would manifest itself as follows: at all horizons a variable's own innovations would account for all of its forecast error variance, so there would be a one in the column corresponding to a variable's own innovation and zeros elsewhere. (Block exogeneity is the obvious multivariate generalization.)[6] In the U.S.-German VAR the exchange rate, relative incomes, the long-term interest differential, and the German and U.S. trade balances all appear to have large exogenous components, since for both variable orderings and all horizons (1-36 months) own innovations in these variables explain at least 48%, 55%, 50%, 59%, and 65% of their respective forecast error variances. For the U.S.-U.K. VAR, own innovations in the exchange rate, the U.K. and U.S. trade balances, and the long-term interest differential account for most of the forecast error variance of these variables. In the U.S.-Japanese system, it is the exchange rate, the Japanese and U.S. trade balances, and relative incomes that have this property.

The last feature of tables 1-6 that we wish to emphasize concerns the difference between those factors which appear to explain the forecast error variance of the three bilateral exchange rates at short horizons (1-3 months) as opposed to longer horizons (1-3 years). Based on the numbers reported in these tables it is clear that own innovations in exchange rates explain a large fraction of the exchange rate forecast error variance at one and three month forecast horizons, while innovations in the other variables become relatively more important at horizons of one and thre years. This result is not atypical of VARs estimated on macro-economic data; see Fischer (1981).

All of the features of tables 1-6 noted above suggest both (1) the difficulty in specifying the menu of variables to include in a structural exchange rate equation, and (2) the problems associated with finding legitimate instruments with which to consistently estimate the parameters of these models. The latter difficulty has led to the constrained-coefficient methodology of the next section.

3. Predicting and Explaining the Exchange Rate Out of Sample Using Structural Models with Constrained Coefficients

Elsewhere [Meese-Rogoff (1981)] we employ rolling regressions to construct out-of-sample forecasts of the exchange rate using three structural models: a flexible-price monetary model (Frenkel-Bilson), a sticky-price monetary model (Dornbusch-Frankel), and a sticky-price asset model which incorporates the trade balance (Hooper-Morton).[7/] The fact that these structural models do not outperform the random walk model at horizons of one to twelve months cannot be attributed to the inherent unpredictability of the explanatory variables; this uncertainty is purged from the forecasts by using realized explanatory variable values. Still, there remains the possibility that our small-sample results can be attributed to poor parameter estimates rather than specification error. This possibility is especially worrisome in light of the estimated VAR models presented in the previous section. They indicate that it is difficult to find legitimate exogenous variables in the three structural exchange rate models. If this is the case, then consistent coefficient estimation becomes problematic and requires a priori knowledge of the serial correlation process of the error terms. These possible estimation problems may explain why the instrumental variables techniques implemented in our other study do not yield better results than ordinary least squares.

Here we explore a range of constrained-coefficient models and present evidence that our other results concerning one to twelve month forecast horizons cannot be explained by coefficient uncertainty. In addition, since the coefficient-constrained models do not require a significant portion of the limited floating-rate data set for estimation, we are able to look at longer forecast horizons.

3a. The representative structural models

All three of the structural exchange rate models we consider are based on a common money demand specification, thereby allowing us to impose coefficient constraints on a consistent basis across models. The quasi-reduced form specification of each of the models is subsumed in the general specification below:

$$(2) \qquad s = a_0 + a_1(m - \overset{*}{m}) + a_2(y - \overset{*}{y}) + a_3(r_s - \overset{*}{r}_s) + a_4(\pi^e - \overset{*}{\pi}{}^e) + a_5(\overline{TB} - \overset{*}{\overline{TB}}) + u,$$

where $(\pi^e - \overset{*}{\pi}{}^e)$ is the expected long-term inflation differential, $\overline{TB}$ and $\overset{*}{\overline{TB}}$ are the cumulated U.S. and foreign trade balances, u is a disturbance term, and the other variables are as defined in section 2 above. In (2) we have imposed the usual constraint that domestic and foreign variables affect the exchange rate with coefficients of equal but opposite sign; this constraint is relaxed in a limited number of experiments both here and in our earlier study.[8/] We choose not to specify an ad-hoc lagged adjustment mechanism in (2), preferring to model the dynamics using an autoregressive error term as described below.

All three models hypothesize first-degree homogeneity of the exchange rate with respect to relative money supplies, or $a_1 = 1$. The Frenkel-Bilson, or flexible-price monetary model, formed by differencing two identical money demand specifications while imposing purchasing power parity (PPP), posits the additional coefficient restrictions $a_2 < 0$, $a_3 > 0$, $a_4 = a_5 = 0$.

The Dornbusch-Frankel, or sticky-price monetary model also hypothesizes that the coefficient on relative incomes $a_2 > 0$, but in contrast to the Frenkel-Bilson model hypothesizes that the coefficient on the short-term interest differential $a_3 < 0$, and that the coefficient on the long-term expected inflation differential $a_4 > 0$. The derivation of these coefficient restrictions is exposited in Frankel (1979). The principal theoretical difference between the Frenkel-Bilson model and the Dornbusch-Frankel model is that the latter allows for short-run deviations from purchasing power parity due to sticky domestic prices. Prices adjust only gradually, in response to both excess demand, which depends on the terms of trade, and to secular inflation differentials [$\pi^e - \overset{*}{\pi}{}^e$ in equation (2)]. The long-run or flexible price exchange rate $\bar{s}$ is derived in the same manner as s in the Frenkel-Bilson model except that it depends on $\pi^e - \overset{*}{\pi}{}^e$, which is equal to the long-run short-term interest differential:

$$(3) \qquad \bar{s} = -\alpha + (m - \overset{*}{m}) - \emptyset(y - \overset{*}{y}) + \lambda(\pi^e - \overset{*}{\pi}{}^e).$$

Using money demand functions of the form

$$(4a) \qquad m - p = \alpha - \lambda r_s + \emptyset y$$

$$(4b) \qquad \overset{*}{m} - \overset{*}{p} = \alpha - \lambda \overset{*}{r}_s - \emptyset \overset{*}{y},$$

and a price adjustment equation of the form

$$(5) \qquad (p - \overset{*}{p})_{t+1} - (p - \overset{*}{p})_t = \theta(s - p + \overset{*}{p})_t + (\pi^e - \overset{*}{\pi}{}^e)_t,$$

Frankel demonstrates that augmented regressive expectations are rational:[9/]

$$(6) \qquad s^e_{t+1} - s_t = \theta(\bar{s} - s)_t + (\pi^e - \overset{*}{\pi}{}^e)_t,$$

where s^e_{t+1} is the exchange rate expected to prevail at time t+1 based on period t information. Substituting (3) into (6) for $\bar{s}$, and also imposing uncovered interest parity by substituting $r_s - \overset{*}{r}_s$ for $s^e_{t+1} - s_t$, one arrives at the quasi-reduced form of the Dornbusch-Frankel model:

$$(7) \qquad s = -\alpha + (m - \overset{*}{m}) - \emptyset(y - \overset{*}{y}) - \frac{1}{\theta}(r_s - \overset{*}{r}_s) + (\frac{1}{\theta} + \lambda)(\pi^e - \overset{*}{\pi}{}^e).$$ [10/]

So in the Dornbusch-Frankel model a_3, the coefficient on the short-term interest differential $r_s - \overset{*}{r}_s$, does not depend on the nominal interest rate semi-elasticity of the demand for real balances λ. Rather it depends on the negative of the inverse of θ, the coefficient on excess demand in the price adjustment equation. The coefficient on the expected long-run inflation differential, a_4, is the sum of $1/\theta$ and λ.

The Hooper-Morton trade-weighted dollar model imposes the same constraints as the Dornbusch-Frankel model, except that it allows unanticipated shocks to the U.S. trade balance to affect the PPP or long-run real level of the exchange rate. In our bilateral version of their model, incipient trend U.S. trade balance surpluses require an appreciation of the long-run real exchange rate, while incipient trend foreign surpluses require a depreciation. Thus, $a_5 < 0$. It should be noted that the random walk model is also subsumed in the general specification (2). That model is given by $a_1 = a_2 = a_3 = a_4 = a_5 = 0$, and $u_t = u_{t-1} + e_t$, where e_t is a white noise process.

3b. A description of the coefficient constraints

The least controversial constraint we impose is that a_1, the coefficient on the logarithm of relative money supplies, is unity. While we shall not consider other values for a_1, we do experiment with different definitions of the money supply; the reserve adjusted base, M1-B, and M2 (in conjunction with their respective foreign counterparts).[11/]

Widespread agreement is lacking on the values of the other parameters. For example, there are a range of theoretical and empirical estimates of the interest and income elasticities of money demand. The quantity theory puts the income elasticity at one, and the interest elasticity at zero. Alternatively, the Baumol (1952) - Tobin (1956) inventory theoretic approach, in its simplest form, can be used to derive an income elasticity of .5 and an interest elasticity of -.5. Taking into account integer constraints raises the income elasticity towards one and the interest elasticity towards zero; see Barro (1976). The Miller-Orr (1966) model of a firm's optimal cash-management procedures yields an interest elasticity of -.33. The income elasticity suggested by that model ranges from .33 to .67, depending on whether a rise in income brings a rise in the number of transactions or in the average size of transactions. The Whalen (1966) model of the precautionary demand for money also suggests an interest elasticity of -.33. In addition it yields an income elasticity which depends on how the size versus frequency of transactions changes as income rises, ranging from .33 to 1. Finally, we consider empirical estimates of the demand for money, for which Goldfeld's (1973) paper is a standard reference. He estimates the income elasticity of money demand to be .19 in the short run, and .68 in the long run; his short-run and long-run interest elasticities are -.064 and -.23. Since the present study takes the approach of modeling the serial correlation properties of the error term rather than specifying an ad-hoc lagged adjustment mechanism, it follows that the higher long-run elasticities are more relevant for our purposes.

We are now ready to specify a complete grid of constraints for the Frenkel-Bilson model. The income elasticity constraints considered are .5, .65, .75, .85, and 1. This grid excludes the lowest ranges of income elasticities obtained in the theoretical models; we implicitly assume that income growth is accompanied by some growth in the size of transactions. The interest rate semi-elasticity constraints include -3, -4.5, -6, -7.5, and -10. The latter grid encompasses interest rate elasticity priors ranging from somewhere between -.18 and -.21 to -.60 and -.70, depending on the bilateral exchange rate. The semi-elasticity priors are obtained by dividing the interest elasticity priors by the average prevailing level of short-term interest rates during the sample.

The grids of constraints for the Dornbusch-Frankel and Hooper-Morton models incorporate the same range of income elasticity and interest-rate semi-elasticity constraints as the Frenkel-Bilson model grid. The two sticky-price models also require the specification of a grid for θ, the speed of adjustment parameter in the goods market. We choose a range of constraints for θ using the fact that it also represents the speed at which deviations from the long-run real exchange rate are damped. The grid for θ

is based on the assumption that between 33% and 100% of today's deviation from PPP is expected to be eliminated one year hence. This range encompasses Genberg's (1978) estimates as well as those of Frankel (1979), both of which are based on data for Germany. Since in the Dornbusch-Frankel model the coefficient on relative short-term interest rates, a_3, is equal to $-1/\theta$, our grid of priors for a_3 in that model includes -1, -2, and -3 . (The values for θ and therefore $-1/\theta$ are conceived on an annual basis, since the short-term interest differentials and expected long-term inflation differentials in the data set are annualized.) The coefficient on the expected long-run inflation differential a_4 is equal to $\lambda + 1/\theta$, where $-\lambda$ is the interest semi-elasticity of money demand. The grid of constraints for a_4 is 4, 7, 9, 11, 13 , which includes the minimum and maximum possible values of $\lambda + 1/\theta$ given the individual grids of constraints for λ and $1/\theta$. For consistency, we exclude from our overall grid for the Dornbusch-Frankel model combinations of a_4 and a_3 such that $a_4 - a_3$ is less than 3 or greater than 10, the bounds on the grid for λ.

The coefficients on the cumulative monthly trade balances (taken as deviations from trend) in the Hooper-Morton model are based primarily on Hooper and Morton's work. We assume that a billion dollar U.S. trade balance surplus above trend level leads, ceterus paribus, to an offsetting .3 to .5 percent appreciation of the dollar; that is, a_5 is .003 or .005. The results reported below are robust to using values of a_5 of .01 or .02. For simplicity, and in order to limit the size of the large grid of coefficient constraints for the Hooper-Morton model, we assume that a foreign trade balance surplus has an effect on the exchange rate of equal magnitude but opposite sign.

The final variable for which it is necessary to specify a grid of constraints is one which we logically know nothing about-- the error term u_t. We assume that u_t follows a first order autoregressive process:

(8) $$u_t = \rho u_{t-1} + e_t = e_t/(1 - \rho L),$$

where e_t is white noise and L is the lag operator. The grid for the autoregressive parameter ρ is 0, .2, .4, .6, .8, and 1.0 , so both the no serial correlation case and the first-difference case are covered.[12/] The decision to analyze only a first-order autoregressive process is made in part to limit the size of the parameter grids, but it is also in part because of the results of our other study. There, optimal linear combinations of structural models and very general autoregressive time series models are analyzed. A wide variety of optimal lag length selection criteria are used in developing the time series components of the forecasts; these criteria generally select a lag length of one for the univariate models.

Given the range of constraints we have selected, the grid for the Frenkel-Bilson model contains 150 different combinations of parameters, the grid for the Dornbusch-Frankel model has 330 elements, and the Hooper-Morton model grid has 660 elements.[13/]

3c. Results

The grid of parameter values developed above is now used to perform two basic experiments, designed to compare the structural models to the random walk model at forecast horizons of one, three, six, twelve, eighteen, twenty-four, thirty, and thiry-six months. The "ex-post" and "ex-ante" forecasting experiments differ mainly in whether forecasts are generated using realized values of the explanatory variables (ex-post), or using predictions of the explanatory variables based on information available at the time of the forecast (ex-ante).[14/] The other difference is that ex-ante forecasting begins in June 1975 while ex-post forecasting covers the entire sample period. The ex-ante experiment requires enough observations for first-round estimation of the VAR that generates predictions of the explanatory variables. (Because the ex-ante experiment is quite expensive to conduct, it is performed only for the Dornbusch-Frankel model.) Otherwise, the experiments are conducted in identical fashion. Constant terms corresponding to each constellation of parameter values are estimated using rolling regressions. The autoregressive component of forecasts made at time t are based on the period t error term.

The results of the ex-post forecasting experiment are broadly characterized in table 7, where the structural model "forecasts" are compared with the random walk model forecasts on the basis of RMSE and MAE.[15/][16/] For each model and exchange rate, table 7 reports the shortest forecast horizon, in months, at which 0.1%, 10%, 25% and 50% of each model's parameter grid outpredicts the random walk model when realized values of the explanatory variables are used. Table 7 demonstrates that the results of Meese-Rogoff (1981) cannot be explained by parameter uncertainty. For the entire parameter grid and for all three exchange rates, the structural models never improve at all, much less significantly, on the random walk model in MAE or RMSE at forecast horizons less than twelve months. However, at horizons of twelve months or more--longer than we could examine in our study based on estimated coefficients--the RMSE and MAE of the models do sometimes improve on the random walk model. This result is tempered by the fact that the minimum RMSE or MAE coefficient configurations bounce around at different forecasting horizons. Still the percentage of the parameter grids which improve on the random walk model does increase with forecast horizon. Overall these essentially in-sample results--in-sample, because not all coefficient configurations improve on the random walk model--must be interpreted with caution.

Table 8 presents best representative parameter values for each of the models, together with their corresponding RMSE and MAE.[17/] These two statistics are also given for the random walk model. At 36 months, the best representative coefficient values for the Dornbusch-Frankel and Hooper-Morton models do about 50% better than the random walk model in RMSE and MAE; the Frenkel-Bilson model only does about 30% better.

Since the models do not forecast well at short horizons in the ex-post experiment, it is not surprising that the one model considered in the ex-ante experiment does poorly at short horizons as well.[18/] Tables 9 and 10 present results for ex-ante forecasting experiment with the Dornbusch-Frankel model. No parameterization of that model ever improves on the random walk model in MAE for horizons under 12 months; the threshold horizon is even longer when RMSE is the metric. Furthermore, for the dollar/pound and dollar/yen exchange rates, over 90% of the parameter grids fail to beat the random walk model in MAE or RMSE at any horizon. It is true, however, that at 36 months the best representative Dornbusch-Frankel model performs almost as well in the ex-ante experiment as the best representative Dornbusch-Frankel model in the ex-post experiment; compare tables 8 and 10. Again, we should emphasize that the evidence presented here on the possible forecasting superiority of the structural models is essentially in-sample, since not all configurations of the parameter constraints improve on the random walk model.

Also reported in Table 10 are the forecasting properties of the seven-variable VAR system of section 2. This model, estimated by rolling regressions, is a true ex-ante forecaster. The VAR outforecasts the random walk model at three-year horizons for the dollar/DM rate. It does worse at one-year horizons for that exchange rate, though, and worse at all horizons for the dollar/pound and dollar/yen exchange rates. It is possible that these results can improved by imposing probabilistic priors on the VAR; see Litterman (1979). (An identified structural model such as the Dornbusch-Frankel model can be thought of as a VAR with a priori restrictions.)

4. The Poor Performances of the Structural Models: Possible Causes

The constrained-coefficient experiments of section 3 reinforce the results of our earlier study. The selected structural models (with autoregressive error terms) fail to forecast or even explain out of sample as well as the random walk model at horizons of up to twelve months. The models do sometimes produce better forecasts than the random walk model at longer horizons, but in an unstable fashion. As noted in section 2, the limited floating rate data set necessarily contains more information about short than about long forecast horizons.

In this section we try to trace the instability or misspecification of these empirical exchange rate equations to their building blocks, such as uncovered interest parity[19/], the particular money demand specification, the proxies for inflationary expectations, and the goods markets specifications. These building blocks are not, of course, strictly independent.

The assumption of uncovered interest parity has been strongly challenged by recent work on exchange rate risk premia.[20/] However, while some authors find evidence of risk premia, the weight of the evidence is that the magnitudes involved are not large. Nevertheless, volatile time-varying risk premia remain a possible explanation of the results.

The goods market specifications of the three representative structural models are relatively simple. The flexible-price Frenkel-Bilson monetary model imposes purchasing power parity, even in the short run. The sticky-price Dornbusch-Frankel monetary model allows for short-run deviations from PPP. The Hooper-Morton model is similar except that it attempts to incorporate movements in the long-run PPP level of the exchange rate by assuming that these movements take place in response to unanticipated trade balance (current account) deficits or surpluses. While short-run PPP does not provide an accurate characterization of the 1970's,[21/] there is no strong evidence that the long-run PPP level of the exchange rate changed significantly. The results in Meese and Rogoff (1982) suggest also that although the deviations from PPP damp quite slowly, the rate at which they damp is relatively stable.

The performance of the Dornbusch-Frankel and Hooper-Morton models are potentially quite sensitive to the use of a variable other than the long-term interest differential as a proxy for the long-run expected inflation differential. Although we did not find a proxy which yielded better results [see footnote (15)], this issue merits further attention.

However, the major problem with the structural models considered here may be the instability of the underlying money demand specifications. The recent breakdown of U.S. money demand relationships was first noted by Goldfeld (1976) and is documented extensively by Simpson and Porter (1980). Conventional empirical money demand specifications such as equations (4) of section 3 have consistently underpredicted U.S. M1 velocity since mid-1974. For this reason, the present study uses M1-B, for which the systematic bias over the sample period is much smaller, and the new definition of M2, for which the bias is negligible. But equations (4) still fail to predict these aggregates or the reserve-adjusted base with any notable degree of precision. As reported above, our exchange rate experiment results are not sensitive to which of these aggregates (together with their respective foreign counterparts) we employ.

Whether or not money demand instability and/or misspecification is responsible for the exchange rate results, it is certainly true that the conventional money demand equation does not work well when expressed in terms of U.S. minus foreign variables. That equation [(4(a) minus 4(b)] fails Chow (1960) tests for the stability of the intercept term at four different breaks in the sample. It also fails Goldfeld-Quandt (1965) tests of homoscedastic disturbance terms over the same sample breaks.[22/]

To investigate the possibility that our results are generated solely by money demand instability in the U.S., we performed ex-post forecasting experiments using the Dornbusch-Frankel model on the pound/mark, pound/yen and yen/mark cross exchange rates. For the case of the yen/mark rate, we found coefficient values for which the model pulled even with the random walk model as early as six months. But the subsequent improvement at longer horizons never exceeded 30%. (The pound/mark and pound/yen cross-rate results are no better than the results for the various dollar exchange rates.)

In sum, money demand instability is an important potential explanation of our results, but further work is needed to demonstrate that time-varying risk premia, volatile long-run real exchange rates, or poor measurement of inflationary expectations are not the dominant problems.

5. Conclusions

The unimpressive out-of-sample performance of the Frenkel-Bilson, Dornbusch-Frankel and Hooper-Morton empirical exchange rate models cannot be attributed to inconsistent or inefficient parameter estimates. These models fail to yield any improvement over the random walk model in mean absolute or root mean squared error one to twelve months out of sample for a broad range of theoretically plausible coefficient values even when autoregressive error terms are introduced. Thus it is unlikely that more efficient estimation techniques, such as imposing all the cross-equation rational expectations restrictions, will yield parameter estimates which do better.[23/] The coefficient-constrained models do prevail at longer horizons but in an unstable fashion; the best coefficient values bounce around depending on the forecast horizon.[24/]

While the breakdown of empirical exchange rate models may be due to volatile time-varying risk premia, volatile long-run real exchange rates, or poor measurement of inflationary expectations, the main problems appear to lie in their money demand specifications. If this is the case, then the same improvements which resuscitate domestic empirical money demand equations should lead to great improvements in empirical exchange rate equations as well.

Footnotes

[1/]The authors have benefited from the comments and suggestions of Jeffrey Frankel, Robert Flood, Robert Hodrick, Peter Hooper, Peter Isard, and Julio Rotemberg. We are indebted to Julie Withers and Tamara McKann for excellent research assistance. This paper represents the views of the authors and should not be interpreted as reflecting the views of the Board of Governors of the Federal Reserve System or other members of its staff.

[2/]The structural models are described here in section 3 below. It should be noted that all the models considered are derivatives of the monetary or asset approach in that they specify real money demand at home and abroad as a function of real income, short-term interest rates and possibly wealth. The random walk model predicts that today's exchange rate will obtain at all future dates.

[3/]In a study of the dollar/pound rate, Hacche and Townend (1981) use different methods to arrive at a similar conclusion; that the models do a very poor job of explaining the dollar/pound rate. The present study examines the three bilateral dollar rates and also cross-rates.

[4/]The assumption that U.S. and foreign variables enter exchange rate equation systems with equal but opposite signs is relaxed later in a limited number of experiments on the structural models. Economizing on variables in the otherwise highly parameterized VAR systems is quite important, so we only estimate the VAR models with the relative variables.

5/ Parzen's (1975) criterion selects an order ℓ^* which minimizes

$$CAT(\ell) = \mathrm{trace}\left(\frac{N}{T} \sum_{j=1}^{\ell} V_j^{-1} - V_\ell^{-1}\right), \ \ell = 1, 2, \ldots L,$$

where N is the number of variables in the VAR, T is sample size, L is the maximal order considered, and V_j is an estimate (adjusted for degrees of freedom) of the covariance matrix of disturbances for the model with ℓ lags of each variable. Asymptotically, the order selected is never less than the true order, assuming the true order is finite.

6/ In Meese-Rogoff (1981), the block exogeneity assumptions of the Dornbusch-Frankel model are formally tested.

7/ See Bilson (1978, 1979), Frenkel (1976), Dornbusch (1976b), Frankel (1979, 1981), and Hooper and Morton (1982). The identification of particular empirical models with authors who contributed significantly to their development follows one conventional nomenclature. It is relevant to note, however, that several of these same authors have analyzed more than one of the three models. For example, Frenkel (1981b) discusses a sticky-price model, and emphasizes that the flexible-price model is a limiting approximation which is applicable in a highly inflationary environment. Dornbusch (1976a) examines a flexible-price monetary model with traded and nontraded goods.

8/ Haynes and Stone (1981) suggest that a problem with the representative structural models we consider is the restriction of equal but opposite coefficients on domestic and foreign variables. In Meese-Rogoff (1981), relaxing this restriction by letting certain domestic and foreign variables--incomes, money supplies and cumulated trade balances--enter equation (2) separately yielded no forecasting improvement. Here we tried separating the incomes and trade balances, but again found no forecasting improvement.

9/ It is also straightforward to show that deviations from purchasing power parity caused by monetary shocks are expected to damp at rate θ.

10/ Frankel uses both long-term interest differentials and past inflation differentials as proxies for $\pi^e - \overset{*}{\pi}^e$, the flexible price or long-run expected inflation differential.

11/ For the dollar/pound rate we use M3's, since there is no data on M2 for the U.K. The results presented later in this section are based on M1 (M1-B) data. However, we obtain very similar results with the different monetary aggregates.

12/ For the Dornbusch-Frankel model, we also experimented with a range of constraints on ρ concentrated between .8 and 1. The lowest end of this range produced the best results.

13/ Recall that the Dornbusch-Frankel and Hooper-Morton model grids exclude combinations of a_3 and a_4 incompatible with the range of constraints specified for the interest rate semi-elasticity of real money demand λ.

14/ In the absence of greater knowledge about the true underlying structure than is inherent in equation (2), it is not possible to take advantage of any correlation between the error term and the explanatory variables in generating the ex-post forecasts. Such correlation is likely, though, given the endogeneity of the explanatory variables indicated by the VAR's in section 2. In fact, if the variance of the error term is large and its (unknown) covariance with the relevant linear combination of the explanatory variables (the "fundamentals") negative, our ex-post forecaster need not dominate optimal ex-ante forecasters. In this perverse case, knowing that the realized fundamentals suggest a higher exchange rate means that you should guess a lower exchange rate.

15/ The results for the Dornbusch-Frankel and Hooper-Morton models reported in Tables 7-10 are obtained using long-term interest differentials as a proxy for expected long-run inflation differentials. It is important to recognize that these models are potentially quite sensitive to this variable. However, using instead current-period inflation differentials, a moving average of past inflation differentials, or future inflation differentials, yields qualitatively similar results in the ex-post forecasting experiments (Tables 7 and 8). We did not try these other proxies in the expensive ex-ante experiments (Tables 9 and 10).

16/ Let k = 1, ..., 36 denote the forecast step, N_k the total number of forecasts in the projection period for which the actual value A(t) is known, F(t) denote the forecast value, and let forecasting begin in period (t+1). Define

$$\text{Mean absolute error} = \sum_{s=0}^{N_k - 1} \left| F(t + s + k) - A(t + s + k) \right| / N_k$$

$$\text{Root mean square error} = \left\{ \sum_{s=0}^{N_k - 1} \left[F(t + s + k) - A(t + s + k) \right]^2 / N_k \right\}^{1/2}$$

Our use of mean absolute error covers problems that might arise if, as suggested by Westerfield (1977), exchange rate changes are drawn from a stable Paretian distribution with infinite variance. The mean errors of the models (not reported) are small relative to mean absolute errors in almost all cases where $\rho > .2$, indicating that the structural models are not simply systematically over or underpredicting.

17/ The "best" representative set of parameter constraints for each model in Table 8 is chosen in an ad hoc fashion as the one which comes in first (also ahead of the random walk model) at the greatest number of horizons. The maximum improvements over the random walk model in MAE or RMSE at 36 month horizons exhibited by these representative models are as large as those exhibited by any other parameter configurations.

18/ While only one model is considered in the ex-ante experiment, note that all three models yield qualitatively equivalent results for the ex-post experiment. Also, since the Dornbusch-Frankel model predicts that the exchange rate will return in the long run to its flexible-price or Frenkel-Bilson model value, we should a priori expect the performance of both models at long forecast horizons to be quite similar.

19/ In later versions of the Hooper-Morton model this assumption is relaxed.

20/ See for example Hansen and Hodrick (1980a,b), Cumby and Obstfeld (1981), Hakkio (1981), Tryon (1979), Bilson (1981), Meese and Singleton (1980), or Geweke and Feige (1979). Hansen and Hodrick study this issue in their paper contained in this volume.

21/ Isard (1976), Genberg (1978), and Frenkel (1981a,b) provide evidence on this point. In this context, it would be useful to remind the reader that our identification of particular models with particular authors oversimplifies the history, development, and application of these models. [See footnote (7)].

22/ The breaks in the sample at which these stability tests are conducted are chosen arbitrarily and correspond to (1) June 1974 - the start of the mature float, (2) November 1976 - the approximate sample midpoint, (3) November 1978 - the dollar support program, and (4) October 1979 - the change in Federal Reserve operating procedures. The tests were conducted using all parameter configurations with the grids for (λ, $\emptyset$) reported in section 3b.

23/ See Driskell and Sheffrin (1981) or Glaessner (1982). These more sophisticated statistical techiques may provide superior expectations proxies, however.

24/ If the true structural model were known, and combined with an accurate representation of the serial correlation process of the error term, then such a model would produce minimum MSE forecasts at all horizons.

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Technical Appendix

In this appendix we describe the triangularization of the VAR system used in section 2 to analyze the dynamic effects of an innovation to a particular variable. First suppose the t-th observation of the VAR is represented by

$$\text{(A1)} \quad [I_N - A(L)]\underline{y}_t = \underline{u}_t,$$

where $[I_N - A(L)]$ is a matrix polynomial in the lag operator L, $\underline{y}_t$ is the N x 1 vector of variables in the system, $E(\underline{u}_t) = \underline{0}$, and $Var(\underline{u}_t) = V$, positive definite. Using the Cholesky factorization $V = WW'$, where W is lower triangular, we can transform (A1) to the system

$$\text{(A2)} \quad W^{-1}[I_N - A(L)]\underline{y}_t = W^{-1}\underline{u}_t = \underline{e}_t,$$

where $E(\underline{e}_t) = \underline{0}$ and $VAR(\underline{e}_t) = I_N$, the order N identity matrix. Since W^{-1} is also lower triangular, the system (A2) is recursive as described in the text. The moving average representation of (A2) is

$$\text{(A3)} \quad \underline{y}_t = [I_N - A(L)]^{-1}W\,\underline{e}_t,$$

and in this expression the contemporaneous value of the first component of $\underline{e}$ enters all N equations, the contemporaneous value of the second component of $\underline{e}$ enters the last N-1 equations, etc. Because the decomposition of V is not unique, studying the effect of the uncorrelated innovations $\underline{e}_t$ on $\underline{y}_t$ will depend on the variable ordering unless V is diagonal, i.e. unless the system (2) has no contemporaneous interactions among variables.

Expression (A3) is also used to construct the variance decompositions of tables 1-6. Since all components of $\underline{e}_t$ have unit variance, the variance of y_{it} (the i-th element of the vector $\underline{y}_t$) is the sum of squares of the elements in the i-th row of $[I_N - A(L)]^{-1}W$. The percentage of the forecast error variance of y_{it} explained by the j-th innovation e_{it} (the j-th element in the vector $\underline{e}_t$) is calculated as the ratio of the sum of squares of the (i, j) element of $[I_N - A(L)]^{-1}W$ to the variance of y_{it}.

Data Appendix

The data set consists of seasonally unadjusted monthly observations over the period March 1973 to June 1981. All the raw data are seasonally adjusted using dummy variables (the results reported in the text are insensitive to the use of more sophisticated seasonal adjustment procedures described in Meese-Rogoff (1981)). In the U.K. data set, the spot and forward exchange rates, short-term interest rate, and long-term bond rate are always drawn from the same date. Because daily bond series are not readily available for Japan and Germany, only the exchange and interest rates correspond in these data sets. All other series are monthly data, and all data are taken from publicly available sources.

The bilateral data sets draw exchange rate data from identical sources, as follows:

One, Six, and Twelve-Month Forward Exchange Rates

Data Source: Data Resources, Inc. data base.

Series: One, six, and twelve-month forward bid rates in U.S. dollars per local currency unit.

Description: Daily data based on 10:00 a.m. opening New York market rates.

Three-Month Forward and Spot Exchange Rates

Data Source: Federal Reserve Board data base.

Series: Three-month forward and spot bid rates in U.S. dollars per local currency unit.

Description: Daily data based on 12:00 noon New York market rates.

Sources of the other data series are discussed below by country.

Germany

Bond Yields

Data Source: Deutsche Bundesbank, Statistical Supplement to the Monthly Reports of the Deutsche Bundesbank, Series 2, Securities Statistics, Table 7b.

Series: Yields in percent per annum on fully taxed outstanding bonds of the Federal Republic of Germany.

Description: Monthly data. Data are calculated as averages of four bank-week return dates including the end-of-month yield of the preceding month.

Consumer Prices

Data Source: Deutsche Bundesbank, Monthly Report of the Deustche Bundesbank, Table VIII-7.

Series: Total cost of living index for all households.

Description: Monthly index.

Industrial Production

Data Source: O.E.C.D., Main Economic Indicators.

Series: Total industrial production.

Description: Monthly index.

Interest Rates (Three-Month)

Data Source: Frankfurter Allegemeine Zeitung.

Series: "Geldmarkt Vierteljahresgeld" in percent per annum. (3-month interbank rate).

Description: Daily data.

Monetary Base (Reserve-Adjusted)

Data Source: Deutsche Bundesbank, Monthly Report of the Deutsche Bundesbank, Table II-1 (components of the unadjusted monetary base) and Table IV (average reserve ratio).

Series: The unadjusted base is calculated in millions of DM as total Bundesbank assets less the reserve adjustment balancing asset, foreign and domestic public authority deposits, SDR allocations, EMCF gold contributions, liquidity paper liabilities, and "other" liabilities. The reserve adjustment is made by multiplying the unadjusted base statistic by [.631 + 3.2/total average reserve ratio] where .631 = the currency percentage of the unadjusted base in the base period (January 1980) and 3.2 = [1-.631] [the total average reserve ratio in the base period].

Description: Monthly data. Data for components of the unadjusted base refer to the last banking day of the month. The average reserve ratio is a monthly average statistic.

Money Supplies

Data Source: Deutsche Bundesbank, Monthly Report of the Deustche Bundesbank, Table I-2.

Series: Money stock M1 and money stock M2 in millions of DM.

Description: Monthly data. Data refer to the last banking day of the month.

Adjustment: A break in the series, caused by the introduction of a new method of computation, occurs in December 1973. The 1973 statistics are adjusted using the ratio of the new to the old statistic for December 1973.

Trade Balance

Data Source: O.E.C.D., Main Economic Indicators.

Series: Trade balance (f.o.b. - c.i.f.) in billions of DM.

Description: Monthly data.

Japan

Bond Yields

Data Source: Data prior to 1981 are taken from Bank of Japan, Economic Statistics Monthly, Table 71(2). 1981 data are taken from Planning and Research Department, Tokyo Stock Exchange, Monthly Statistics Report, Table 8-1.

Series: Yields in percent per annum on listed government bonds (Tokyo Stock Exchange).

Description: Monthly data. Data refer to the last banking day of the month.

Consumer Prices

Data Source: Bank of Japan, Economic Statistics Monthly, Table 119(1).

Series: General consumer price index for all Japan.

Description: Monthly index.

Industrial Production

Data Source: O.E.C.D., Main Economic Indicators.

Series: Total industrial production.

Description: Monthly index.

Interest Rates (Three-Month)

Data Source: Federal Reserve Board data base.

Series: "Over two-month ends" bill discount rate (Tokyo Stock Exchange) in percent per annum.

Description: Daily data based on Reuters quotes.

Money Supplies

Data Source: Bank of Japan, Economic Statistics Monthly, Table 4.

Series: M1 and M2+CD in 100 million yen.

Description: Monthly data. Data refer to the last banking day of the month.

Trade Balance

Data Source: O.E.C.D., Main Economic Indicators.

Series: Trade Balance (f.o.b. - c.i.f.) in billions of yen.

Description: Monthly data.

United Kingdom

Bond Yields

Data Source: Financial Times

Series: "British funds, Undated, War loans 3½" in percent per annum.

Description: Daily data.

Consumer Prices

Data Source: Department of Employment, Employment Gazette, Table 6.4.

Series: General index of retail prices, all items.

Description: Monthly index.

Industrial Production

Data Source: O.E.C.D., <u>Main Economic Indicators</u>.

Series: Total industrial production.

Description: Monthly data.

Interest Rates (Three-Month)

Data Source: <u>Financial Times</u>.

Series: Three-month local authority deposits (London money rates) in percent per annum.

Description: Daily data.

Monetary Base (Reserve-Adjusted) and Money Supplies

Data Source: Bank of England, <u>Quarterly Bulletin</u>, Table 1 (monetary base components) and Table II (money supplies).

Series: Money stock M1 and money stock sterling M3 in millions of pounds. The reserve-adjusted monetary base is calculated in millions of pounds as total currency in circulation plus bankers' deposits.

Description: Monthly data. Data refer to the third Wednesday of the month (second in December).

Trade Balance

Data Source: O.E.C.D., <u>Main Economic Indicators</u>.

Series: Trade balance (f.o.b. - c.i.f.) in millions of pounds.

Description: Monthly data.

<u>United States</u>

With the exception of the trade balance statistics, all data are taken from the Federal Reserve Board data base. Many of these series are published in the <u>Federal Reserve Bulletin</u>, and all are available to the public.

Bond Yields

Series: Government bonds with at least 10 years to maturity.

Description: Daily data.

Consumer Prices

Series: Consumer Price Index.

Description: Monthly index.

Industrial Production

Series: Total industrial production.

Description: Monthly index.

Interest Rates (Three-month)

Series: Treasury bill rates.

Description: Daily data.

Monetary Base (Reserve-Adjusted) and Money Supplies

Series: Reserve-adjusted monetary base, M1-B, M2, and M3.

Description: Weekly Wednesday data.

Trade Balance

Data Through 1978:

Data Source: Department of Commerce, Highlights of U.S. Export and Import Trade, Exports Table E-1; Imports Table I-1.

Series: Domestic and foreign exports, excluding Department of Defense shipments, in millions of $ on a F.A.S. value basis; General imports in millions of $ on a Customs Valuation basis changing to a F.A.S. basis in 1974.

Description: Monthly data.

Adjustment: 1973 statistics are adjusted to a F.A.S. value basis using the 1974 average ratio of Customs Valuation to F.A.S. value.

1979-1981 Data:

Data Source: Department of Commerce, Summary of U.S. Export and Import Merchandise Trade, December 1980 (advance statistics for Highlights of U.S. Export and Import Trade), Exports Table 3; Imports Table 5.

Series: Total domestic exports, excluding Department of Defense grant-aid, in millions of $ on a F.A.S. value basis; General imports in millions of $ on a F.A.S. value basis.

Description: Monthly data.

Table 1

Unconstrained U.S.-German VAR, March 1973-June 1981, regular variable order

Proportions of forecast error variance k months ahead attributable to each innovation[a/]

Forecast error variance in	k	Innovation in: $m-m^*$	$y-y^*$	$r_s-r_s^*$	$r_L-r_L^*$	TB	TB^*	s
$m-m^*$	1	.843	.005	.132	.004	.001	.004	.011
	3	.629	.003	.259	.029	.019	.040	.020
	12	.293	.007	.295	.100	.035	.249	.016
	36	.252	.007	.268	.122	.037	.254	.061
$y-y^*$	1	.006	.942	.000	.035	.011	.005	.001
	3	.006	.856	.002	.010	.008	.027	.001
	12	.014	.646	.004	.223	.029	.070	.014
	36	.015	.622	.004	.224	.031	.072	.032
$r_s-r_s^*$	1	.007	.024	.838	.073	.030	.007	.021
	3	.011	.017	.641	.080	.097	.124	.029
	12	.008	.016	.491	.081	.091	.282	.030
	36	.011	.017	.469	.088	.088	.276	.051
$r_L-r_L^*$	1	.024	.000	.161	.781	.002	.002	.030
	3	.056	.054	.126	.683	.007	.028	.044
	12	.152	.099	.101	.524	.036	.048	.040
	36	.147	.096	.108	.510	.039	.052	.048
TB	1	.080	.111	.009	.005	.769	.001	.025
	3	.078	.144	.011	.005	.721	.014	.026
	12	.068	.167	.010	.063	.617	.024	.050
	36	.068	.163	.010	.071	.596	.028	.063
TB^*	1	.021	.034	.063	.008	.042	.832	.001
	3	.024	.061	.047	.012	.036	.813	.006
	12	.023	.101	.037	.061	.032	.676	.069
	36	.028	.096	.040	.069	.030	.649	.086
s	1	.041	.011	.022	.128	.005	.006	.786
	3	.077	.008	.022	.163	.032	.010	.687
	12	.155	.034	.018	.125	.064	.050	.553
	36	.155	.033	.030	.154	.058	.090	.481

[a/]Notes for tables 1-6: Columns of the table correspond to innovations in a particular variable for the specified forecast horizon k = 1, 3, 12, and 36. The rows add to one because the total forecast error variance attributable to each variable on the left of the table is allocated across the seven innovations. Abstracting from coefficient uncertainty, an exogenous variable would manifest itself as follows: At all horizons a variable's own innovations would account for all of its forecast error variance, so there would be a one in the column corresponding to a variable's own innovation and zeros elsewhere.

Table 2

Unconstrained U.S.-Japan VAR, March 1973-June 1981, regular variable order

Proportions of forecast error variance k months ahead attributable to each innovation

Forecast error variance in	k	Innovation in: $m-m^*$	$y-y^*$	$r_s-r_s^*$	$r_L-r_L^*$	TB	TB^*	s
$m-m^*$	1	.943	.003	.001	.002	.012	.024	.014
	3	.866	.045	.004	.010	.024	.021	.029
	12	.551	.298	.021	.009	.045	.039	.038
	36	.424	.319	.070	.013	.048	.045	.080
$y-y^*$	1	.044	.929	.009	.001	.009	.000	.008
	3	.038	.915	.009	.008	.005	.020	.004
	12	.111	.637	.086	.019	.042	.090	.014
	36	.120	.531	.111	.019	.048	.072	.099
$r_s-r_s^*$	1	.142	.076	.770	.006	.002	.001	.003
	3	.183	.092	.581	.022	.030	.058	.034
	12	.110	.058	.284	.058	.069	.348	.073
	36	.115	.054	.238	.076	.120	.320	.076
$r_L-r_L^*$	1	.128	.014	.294	.520	.000	.003	.041
	3	.252	.039	.207	.342	.024	.022	.113
	12	.219	.046	.108	.184	.079	.092	.172
	36	.224	.071	.104	.172	.186	.085	.157
TB	1	.011	.016	.032	.000	.904	.005	.031
	3	.010	.022	.059	.011	.858	.008	.031
	12	.011	.032	.078	.017	.726	.097	.040
	36	.012	.036	.079	.024	.671	.111	.068
TB^*	1	.011	.036	.031	.064	.030	.823	.005
	3	.010	.045	.023	.095	.040	.780	.007
	12	.015	.051	.035	.105	.103	.626	.066
	36	.019	.041	.039	.093	.130	.523	.154
s	1	.011	.007	.007	.020	.071	.055	.828
	3	.016	.004	.004	.014	.139	.080	.741
	12	.016	.004	.003	.057	.285	.127	.507
	36	.026	.006	.003	.060	.313	.138	.455

Table 3

Unconstrained U.S.-U.K. VAR, March 1973-June 1981, regular variable order

Proportions of forecast error variance k months ahead attributable to each innovation

Forecast error variance in	k	Innovation in: $m-m^*$	$y-y^*$	$r_s-r_s^*$	$r_L-r_L^*$	TB	TB^*	s
$m-m^*$	1	.911	.023	.059	.001	.006	.001	.000
	3	.789	.041	.142	.004	.004	.019	.001
	12	.398	.058	.292	.052	.004	.140	.057
	36	.270	.032	.200	.076	.028	.148	.245
$y-y^*$	1	.009	.820	.039	.012	.026	.027	.066
	3	.034	.600	.029	.087	.059	.073	.123
	12	.058	.431	.019	.138	.066	.183	.104
	36	.058	.407	.021	.145	.068	.177	.124
r_s-r_s	1	.047	.018	.880	.034	.001	.019	.000
	3	.034	.012	.830	.031	.001	.089	.003
	12	.136	.018	.641	.050	.000	.144	.009
	36	.159	.017	.600	.060	.002	.139	.024
$r_L-r_L^*$	1	.027	.007	.209	.725	.003	.018	.011
	3	.026	.032	.190	.637	.005	.066	.043
	12	.018	.091	.149	.528	.025	.050	.138
	36	.022	.086	.160	.496	.036	.051	.148
TB	1	.013	.041	.113	.016	.755	.018	.044
	3	.014	.050	.126	.069	.634	.057	.049
	12	.018	.052	.117	.078	.511	.082	.143
	36	.036	.047	.105	.074	.420	.115	.203
TB^*	1	.028	.069	.052	.002	.008	.830	.010
	3	.054	.106	.049	.038	.034	.710	.008
	12	.118	.128	.044	.068	.030	.596	.015
	36	.119	.116	.052	.079	.033	.555	.045
s	1	.000	.021	.014	.026	.007	.005	.926
	3	.001	.018	.008	.038	.008	.035	.891
	12	.012	.006	.002	.088	.048	.168	.675
	36	.066	.007	.058	.099	.059	.150	.560

Table 4

Unconstrained U.S.-German VAR, March 1973-June 1981,
reverse variable order

Proportions of forecast error variance k months
ahead attributable to each innovation

Forecast error variance in	k	Innovation in: $m-m^*$	$y-y^*$	$r_s-r_s^*$	$r_L-r_L^*$	TB	TB^*	s
$m-m^*$	1	.703	.017	.111	.047	.042	.017	.062
	3	.522	.011	.138	.191	.025	.058	.054
	12	.281	.024	.081	.298	.020	.264	.032
	36	.243	.022	.067	.287	.026	.264	.089
$y-y^*$	1	.011	.868	.007	.018	.068	.020	.009
	3	.009	.779	.014	.068	.057	.059	.013
	12	.014	.580	.049	.048	.075	.094	.096
	36	.015	.558	.050	.146	.076	.095	.059
$r_s-r_s^*$	1	.011	.016	.522	.350	.013	.085	.002
	3	.008	.013	.343	.333	.083	.115	.004
	12	.014	.018	.232	.279	.085	.353	.019
	36	.015	.018	.222	.272	.084	.343	.047
$r_L-r_L^*$	1	.000	.000	.002	.963	.008	.002	.024
	3	.013	.068	.013	.848	.014	.017	.026
	12	.080	.145	.014	.663	.048	.023	.027
	36	.081	.140	.015	.647	.051	.028	.038
TB	1	.011	.036	.000	.006	.918	.005	.023
	3	.014	.053	.008	.006	.873	.022	.023
	12	.017	.070	.024	.032	.757	.032	.068
	36	.018	.069	.025	.036	.733	.036	.083
TB^*	1	.001	.014	.001	.019	.028	.932	.004
	3	.003	.025	.009	.016	.023	.906	.017
	12	.003	.050	.034	.023	.024	.750	.117
	36	.008	.047	.035	.028	.024	.721	.137
s	1	.001	.005	.001	.049	.004	.003	.937
	3	.013	.007	.001	.068	.045	.007	.859
	12	.056	.013	.001	.060	.024	.054	.690
	36	.073	.013	.002	.105	.116	.091	.601

Table 5

Unconstrained U.S.-Japan VAR, March 1973-June 1981
reverse variable order

Proportions of forecast error variance k months
ahead attributable to each innovation

Forecast error variance in	k	Innovation in: $m-m^*$	$y-y^*$	$r_s-r_s^*$	$r_L-r_L^*$	TB	TB^*	s
$m-m^*$	1	.749	.106	.028	.060	.013	.036	.010
	3	.686	.178	.020	.039	.034	.024	.019
	12	.394	.396	.023	.045	.064	.049	.024
	36	.271	.469	.045	.066	.051	.051	.045
$y-y^*$	1	.000	.898	.058	.002	.015	.019	.009
	3	.003	.857	.050	.016	.014	.044	.015
	12	.014	.693	.073	.088	.016	.092	.022
	36	.014	.658	.068	.090	.028	.071	.071
$r_s-r_s^*$	1	.011	.001	.505	.417	.016	.049	.001
	3	.056	.001	.503	.274	.022	.124	.020
	12	.046	.017	.308	.134	.052	.320	.124
	36	.052	.020	.273	.117	.107	.302	.130
$r_L-r_L^*$	1	.009	.003	.000	.826	.012	.122	.026
	3	.134	.002	.017	.599	.026	.118	.109
	12	.151	.029	.032	.304	.133	.105	.246
	36	.193	.068	.035	.283	.148	.097	.226
TB	1	.006	.016	.005	.007	.842	.009	.115
	3	.006	.017	.041	.006	.800	.010	.119
	12	.009	.017	.072	.010	.672	.114	.105
	36	.010	.029	.075	.012	.622	.131	.120
TB^*	1	.012	.008	.000	.001	.033	.928	.018
	3	.008	.013	.008	.003	.046	.906	.016
	12	.017	.035	.058	.006	.119	.739	.026
	36	.019	.038	.053	.012	.131	.601	.146
s	1	.001	.000	.002	.020	.000	.001	.976
	3	.004	.000	.004	.014	.024	.012	.942
	12	.009	.001	.024	.017	.137	.042	.739
	36	.016	.002	.027	.016	.175	.088	.676

Table 6

Unconstrained U.S.-U.K. VAR, March 1973-June 1981, reverse variable order

Proportions of forecast error variance k months ahead attributable to each innovation

Forecast error variance in	k	Innovation in: $m-m^*$	$y-y^*$	$r_s-r_s^*$	$r_L-r_L^*$	TB	TB^*	s
$m-m^*$	1	.749	.106	.028	.060	.013	.036	.010
	3	.686	.178	.020	.039	.034	.024	.019
	12	.394	.396	.023	.049	.064	.049	.024
	36	.271	.470	.045	.066	.051	.051	.045
$y-y^*$	1	.000	.898	.058	.002	.015	.019	.009
	3	.003	.857	.050	.016	.014	.044	.015
	12	.014	.693	.073	.088	.016	.092	.022
	36	.014	.658	.068	.090	.028	.071	.072
$r_s-r_s^*$	1	.011	.001	.505	.417	.016	.049	.001
	3	.056	.001	.503	.274	.022	.024	.020
	12	.046	.017	.308	.134	.052	.320	.124
	36	.052	.020	.273	.117	.107	.302	.130
$r_L-r_L^*$	1	.009	.003	.000	.826	.012	.122	.026
	3	.134	.002	.017	.594	.026	.118	.109
	12	.151	.029	.032	.304	.133	.105	.246
	36	.143	.068	.035	.283	.148	.097	.226
TB	1	.006	.016	.005	.007	.842	.009	.115
	3	.006	.017	.041	.006	.800	.010	.119
	12	.009	.014	.072	.010	.672	.114	.105
	36	.010	.029	.075	.012	.622	.131	.120
TB^*	1	.012	.008	.000	.001	.033	.928	.018
	3	.008	.013	.008	.003	.046	.906	.016
	12	.017	.035	.058	.006	.119	.739	.026
	36	.019	.038	.053	.012	.131	.601	.146
s	1	.001	.000	.002	.020	.000	.001	.976
	3	.004	.001	.004	.014	.024	.012	.942
	12	.009	.001	.024	.017	.137	.072	.739
	36	.016	.002	.027	.016	.175	.088	.676

Table 7

Shortest forecast horizon (in months) for which at least x% of each model's parameter grid improves on the random walk model in MAE/RMSE when realized values of the explanatory variables are used.

	Exchange rate:	$/DM		$/£		$/Yen	
	Metric:	MAE	RMSE	MAE	RMSE	MAE	RMSE
Model	Threshold	Months ahead					
Frenkel-Bilson (Grid size = 150)	0-1%	24	30	18	24	12	12
	10%	30	30	18	24	18	18
	25%	30	30	24	30	24	24
	50%	36	36	30	36	36	30
Dornbusch-Frankel model (Grid size = 330)	0-1%	12	18	18	18	12	12
	10%	18	18	24	24	12	12
	25%	30	30	30	36	12	12
	50%	-	-	-	-	24	18
Hooper-Morton model (Grid size = 660)	0-1%	12	18	18	18	12	12
	10%	18	18	24	24	12	12
	25%	30	30	30	36	12	18
	50%	-	-	-	-	24	18

MAE is mean absolute error, RMSE is root mean square error. The text contains a description of the parameter grids. To read table 7: Consider the entry in the second row, first column (30). The earliest forecast horizon at which at least 10% of the monetary model parameter grid can improve on the random walk model in MAE for the $/DM rate is 30 months out.

Table 8

Comparing the random walk model and the structural models

(with their best representative parameter configurations)

when realized values of the explanatory variables are used.

Model (with best parameter configurations)	Horizon	\$/DM MAE	\$/DM RMSE	\$/£ MAE	\$/£ RMSE	\$/Yen MAE	\$/Yen RMSE
Random Walk model	1	2.4	3.2	2.0	2.5	2.1	3.0
	3	4.8	6.2	3.2	5.1	4.2	5.7
	12	9.4	10.9	9.8	11.5	10.6	13.8
	36	18.1	21.0	23.4	25.4	19.4	23.3
(a_2, a_3, ρ)		(-.5, 4.5, .4)		(-1, 3, .8)		(-.5, 4.5, .8)	
Frenkel-Bilson	1	9.1	11.4	4.2	6.1	4.5	6.4
	3	11.5	14.2	8.7	11.1	7.9	11.5
	12	12.2	15.2	13.5	16.6	9.7	13.3
	36	12.6	17.0	15.5	18.8	10.2	14.5
(a_2, a_3, a_4, ρ)		(-.85, -1, 6, .4)		(-.5, -1, 4, 0)		(-.5, -1, 9, .8)	
Dornbusch-Frankel model	1	5.5	6.9	8.1	10.0	4.4	8.4
	3	8.1	9.7	8.6	10.5	7.4	9.4
	12	8.8	10.8	10.4	12.3	7.0	8.5
	36	8.2	10.5	8.3	10.0	8.8	10.2
$(a_2, a_3, a_4, a_5, \rho)$		(-.5, -1, 7, .005, 0)		(-.5, -1, 4, .005, 0)		(-1, -1, 9, .003, 8)	
Hooper-Morton model	1	8.3	10.4	4.0	10.0	4.9	6.7
	3	8.8	11.0	8.5	10.5	8.3	10.9
	12	9.2	11.6	10.0	12.0	8.8	11.7
	36	9.3	11.6	10.5	12.5	9.3	12.2

MAE (mean absolute error) and RMSE (root mean squared error) are approximately in percentage terms, since "forecasts" are for the logarithm of the exchange rate. Forecasts are compared over the period March 1973-June 1981.

Table 9

Shortest forecast horizon (in months) for which at least x% of the Dornbusch-Frankel model's parameter grid improves on the random walk model in MAE/RMSE when predicted values of the explanatory variables are used.[a/]

		$/DM		$/£		$/Yen	
		MAE	RMSE	MAE	RMSE	MAE	RMSE
	Threshold	Months ahead					
Dornbusch-Frankel model (Grid size = 330)	0-1%	12	24	30	36	12	18
	10%	12	24	-	-	-	-
	25%	36	36	-	-	-	-
	50%	-	-	-	-	-	-

[a/]The description of table 9 is essentially the same as given for Table 8. However, in contrast to the experiment of Table 8, in which realized values of the explanatory variables are used to generate forecasts of the exchange rate, table 9 is based on an experiment in which the explanatory variables are forecast with a VAR.

Table 10

Comparing the random walk model, a VAR model estimated by rolling regressions, and the best representative Dornbusch-Frankel model when predicted values of the explanatory variables are used.

Model	Horizon	$/DM MAE	$/DM RMSE	$/£ MAE	$/£ RMSE	$/Yen MAE	$/Yen RMSE
Random Walk model	1	2.2	3.0	2.0	2.6	2.3	3.2
	3	4.0	5.2	4.1	5.2	4.7	6.2
	12	10.1	11.7	10.1	11.5	13.2	16.1
	36	24.2	26.2	18.8	21.2	24.8	28.1
(a_2, a_3, a_4, ρ)		(-5, -1, 4, 1.0)		(-.5, -1, 4, 0)		(-5., -3, 11, .8)	
Dornbusch-Frankel model (with best parameter configuration)	1	2.4	3.2	14.0	17.3	2.9	4.2
	3	4.6	5.7	14.7	18.0	6.6	8.6
	12	7.3	10.9	17.5	20.1	12.4	16.0
	36	12.4	19.1	9.0	10.8	17.0	18.7
Unconstrained VAR model	1	5.2	6.3	5.4	6.3	4.9	6.4
	3	7.9	9.5	8.0	9.6	7.4	9.5
	12	11.1	13.2	17.3	19.3	15.9	19.6
	36	16.9	18.5	39.5	44.8	37.5	40.6

MAE (mean absolute error) and RMSE (root mean squared error) are approximately in percentage terms, since forecasts are for the logarithm of the exchange rate. Forecasts are compared over the period June 1975-June 1981.