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INTEGRATION, COINTEGRATION AND THE FORECAST CONSISTENCY OF STRUCTURAL EXCHANGE RATE MODELS

by

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ABSTRACT

INTEGRATION, COINTEGRATION AND THE FORECAST CONSISTENCY OF STRUCTURAL EXCHANGE RATE MODELS

Exchange rate forecasts are generated using some popular monetary models of exchange rates, in conjunction with several estimation techniques. We propose an alternative set of criteria for evaluating forecast rationality, which entails the following requirements: the forecast and the actual series i) have the same order of integration, ii) are cointegrated, and iii) have a cointegrating vector consistent with long run unitary elasticity of expectations. When these conditions hold, we consider the forecasts to be "consistent". These criteria appear to be more appropriate for forecasts generated by structural models than typical measures of forecast rationality, since such models rely upon serially correlated measures of the fundamentals.

We find that it is fairly easy for the generated forecasts to pass the first requirement of consistency that the series be of the same order of integration. However, cointegration fails to hold the farther out the forecasts extend. At the one year ahead horizon, most series and their respective forecasts do not appear cointegrated. Finally, of the cointegrated pairs, the restriction of unitary elasticity of forecasts with respect to actual appears not to be rejected in general. The exception to this pattern is in the case of the error correction models in the longer subsample.

1 Introduction

Numerous studies have compared the forecasting performance of various exchange rate models, structural and non-structural, against that of the random walk model. Some recent attempts are Cheung (1993, fractional integration models), Diebold and Nason (1990, nonparametric methods), Chinn (1991, nonlinear models), Meese and Rose (1991, nonlinear models), and Chinn and Meese (1994, structural models and long horizons). Results from these studies tend to corroborate the results reported by Meese and Rogoff in their original papers (1983a,b); that is, it is extremely difficult to out-predict a random walk model of the exchange rate using structural or other time series models. This result has held up for a wide variety of forecast metrics, structural and time series models, estimation techniques, and sample periods.

This study attempts to evaluate forecasts from structural models based on the time series properties of these forecasts. Instead of examining the commonly used measures of forecast accuracy, such as the mean squared error, mean absolute deviation, and the serial correlation of the forecast errors, we explore some basic time series properties of forecasts.[1] In particular, we examine whether forecasts from structural models and the spot exchange rate series i) have the same order of integration, ii) are cointegrated, and iii) have a cointegrating vector consistent with long run unitary elasticity of expectations.

[1] For a recent example of this methodology, see Zarnowitz and Braun (1992). Frankel and Froot (1987) examine the attributes of exchange rate forecasts.

The first property relates to the persistence of forecasts and spot exchange rates, as measured by the order of integration. The other two properties are related to how exchange rates and their respective forecasts are related in the long-run. While exchange rate forecasts may deviate from the observed exchange rates in the short-run, we expect a forecast of any practical relevance should have the above properties. We label the condition where these three properties hold as the "consistency" of a forecast.[2] That is, a forecast is consistent if it has a one-to-one relationship with the spot exchange rate in the long-run.[3] This notion of consistency focuses on the long-run property of forecasts, and hence is weaker than the one conventionally used in evaluations of forecast rationality. It does not, for example, impose any further restrictions on the forecast errors, above and beyond the requirement that they be weakly covariance stationary.[4] That means a forecast can meet the requirement of consistency and, at the same time, it does not satisfy the usual notion of a "rational" forecast. This can happen, for example, when the correct model is

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[2] The usage of "consistency" here is different from that in econometrics, where it denotes convergence in probability. It also differs from a recent definition attributable to Froot and Ito (1989).

[3] Fischer (1989) and Liu and Maddala (1992) apply the concepts of integration and cointegration to testing for relationships between the survey-based forecasts and the actual series. Fischer does so in the context of the US money stock, while Liu and Maddala address exchange rates.

[4] In the literature, a forecast is said to be "rational" if the forecast errors have a zero mean <u>and</u> zero serial correlation.

used but the data on the fundamentals are contaminated by stationary measurement errors. Such a situation is <u>very</u> likely to occur in the case of typical asset-based models which incorporate information on industrial production, money stocks and price indices. Thus, the consistency requirement represents a more realistic way to evaluate exchange rate forecasts from structural models.

The consistency property of forecasts from three structural exchange rate determination models are examined. It can be verified that forecasts from the random walk model are consistent if the spot exchange rate data follow an I(1) process. Thus, even though it is not explicitly considered, the random walk model can serve as a benchmark for comparison.

To anticipate our results, we find that it is fairly easy for the generated forecasts to pass the first requirement of consistency that the series be of the same order of integration. However, cointegration fails to hold the farther out the forecasts extend. Finally, of the cointegrated pairs, the restriction of unitary elasticity of forecasts with respect to actual appears not to be rejected in general, with the exception of the error correction model forecasts in the longer subsample.

The remainder of the paper is organized as follows. In section 2, we briefly review the literature on exchange rate forecast evaluation. Section 3 presents the structural models. Procedures used to estimate these models and generate forecasts are also discussed in this section. The tests for the order of integration, and for cointegration are described in Section 4. Section 5 first describes the data and then reports the empirical results. Section 6 concludes.

2 A Brief Review

It is widely recognized that current exchange rate models fit poorly on post Bretton Woods data. Meese (1990) and Frankel and Rose (1994) provide recent surveys and references. The problem is not a paucity of possible explanations, but rather an embarrassing over-abundance. These include simultaneity problems, improper modeling of expectations formation, the presence of nonlinearities in the data generation mechanism (DGM) of exchange rates, and over-reliance on the representative agent paradigm. This stylized fact has in turn spawned an enormous empirical literature attempting to overturn this stylized fact.

Simultaneity issues were addressed in the original Meese and Rogoff (1983b) paper by using a grid search over the parameter space. Most of the models incorporate the rational expectations assumption, or impose uncovered interest parity; relaxing the first condition, by use of survey measures of exchange rate expectations, has not been shown to improve forecast accuracy. In fact, such forecasts appear to be very biased (Frankel and Froot, 1987). Attempts to account for a time varying risk premium have also been unsatisfactory (Frankel, 1983). Accounting for nonlinearities in the function form has also not been particularly successful in improving out of sample forecasting (Meese and Rose, 1991; Chinn, 1991). Finally, attempting to introduce heterogeneity into a formal macro model of exchange rate determination was undertaken by Chinn (1994), with some limited success. It would be fair to conclude that the general record of structural exchange rate modeling has been fairly dismal, with the following caveat: in almost all these papers, the usual metrics have been used -- mean forecast error, root mean squared error, and mean absolute error. The use of the proposed consistency criterion will offer a different perspective on evaluating exchange rate forecasts.

3 Exchange Rate Models: Estimation and Forecasting

3.1 Exchange Rate Models

This study examines the consistency property of forecasts from three monetary models: the Frenkel (1976) and Mussa (1976) flexible price model; the Dornbusch (1976a) and Frankel (1979) sticky price model; and the Dornbusch (1976b) tradables-nontradables model. All these models start with conventional money demand functions for both the domestic and foreign economies, and impose the condition that expected depreciation equal the nominal interest differential plus an exogenous risk premium on domestic assets that may or may not be zero. These models can be written, respectively, as:

$$\textit{Model 1: } s = (m-m^*) - \phi(y-y^*) + \mu(i-i^*) \tag{1}$$

$$\textit{Model 2: } s = (m-m^*) - \phi(y-y^*) + (\mu+1/\theta)(\pi-\pi^*) - (1/\theta)(i-i^*) \tag{2}$$

Model 3: $$s = (m - m^*) - \phi(y - y^*) + (\mu + 1/\theta)(\pi - \pi^*) - (1/\theta)(i - i^*) + \beta q \qquad \textbf{(3)}$$
$$q \equiv ((p^T - p^N) - (p^{T*} - p^{N*}))$$

where s, m, y and q are the logarithms of the exchange rate (domestic currency per unit of foreign currency), money supply, real income and the relative price of tradables to nontradables, and i and π are the levels of the nominal interest and inflation rates, respectively. An asterisk denotes a foreign variable.

Model 1 contains only the terms in monies, incomes and nominal interest rates, and relies on the further assumption that purchasing power parity (PPP) holds. This "flexible price" monetary model subsumes the Lucas (1982) model since the latter model contains monies and real incomes but no interest rate term.

Model 2, a "sticky price" monetary model does not assume PPP holds at all times. Instead it assumes slow adjustment of goods prices relative to asset prices, thus yielding the well-known overshooting characteristic.

Our third model is motivated by the failure of purchasing power parity to hold for broad price indices, such as the consumer price index and GNP deflators. One approach is to make an explicit recognition of nontraded goods, and to posit that PPP only holds for tradable goods (Dornbusch, 1976b). If the aggregate price level index can be represented by a Cobb-Douglas function of the individual nontraded and traded price indices (with weight β on nontradables) then model 3 is obtained.

3.2 Estimation

Since it is generally accepted that exchange rates and their fundamentals are well approximated by unit root processes, we will estimate all three of these models in first difference form, using OLS and 2SLS procedures. An instrumental variable approach such as 2SLS is appropriate because the right hand side variables -- such as interest rates and money stocks -- can plausibly be interpreted as being jointly determined with the exchange rate.[5]

In addition to the first-difference specification, we also implement the error correction version of these models. The error correction model (ECM) variants include the error correction term (to be discussed below) lagged once, and the first difference of fundamentals lagged once. Thus all regressors in the ECM models are predetermined, and one month ahead forecasts are true ex-ante forecasts.

The Chinn and Meese (1995) methodology is used to construct the error correction term that captures the long-run relationship between exchange rates and their fundamentals. We assume that log linear versions of equations 1-3 are appropriate in the long run, and impose a set of coefficient restrictions for each of the models. These values are given in Table 1. For all models, the money supply and income elasticities are the same (unity and .75, respectively). The coefficients on interest rates, inflation rates and relative prices vary by model, although the coefficients on the first two variables are functions of the interest rate semi-elasticity, which we assume is 4.5. The goods market speed of

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[5] Assuming rational expectations, appropriate instrumental variables include elements in the information set such as lagged variables. We use lags 2 - 4 of the right hand side variables, since there is evidence of MA1 serial correlation in the first difference specifications.

adjustment parameter is taken to be .5 on an annual basis; this corresponds to deviations from PPP damping at rates .94 for monthly data. The final parameter of interest is the share of nontradables in the aggregate price index, ß, which we take to be 0.5.[6]

3.3 The Forecasting Exercise

We evaluate the out-of-sample explanatory power of our representative models over two forecast periods. Our choice of forecast periods is arbitrary; the first starts with the end of the recession in the U.S. in 1982, and the second corresponds to the period after the Louvre Accord in April 1987.

In the experiments reported below, the original estimation period for the first sample is 1973.06 through 1982.12 (115 observations). We then "roll" through our sample ending in 1993.08 to produce 128, 123, and 117 one-, six-, and twelve-month ahead forecasts, respectively. Whenever necessary, forecasts use actual realized values of the RHS variables. As we "roll" through each forecast period, parameter estimates are updated with the addition of each new data point. The original estimation period for the second sample is 1973.06 to 1987.06 (169 observations). We then perform an analogous "rolling regression" procedure, to produce 74, 69, and 63 one-, six-, and twelve-month ahead forecasts.

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[6] For explanation for parameter selections, see the discussion in Chinn and Meese (1995).

4 Unit Root Test and Cointegration Analysis[7]

4.1 Unit Root Test

For a time series $\{y_t\}$, t=1,...,t', the ADF unit root test is based on the regression

$$\Delta y_t = c + \mu t + \pi y_{t-1} + \sum_{j=1}^{k} \gamma_j \Delta y_{t-j} + u_t \qquad \textbf{(4)}$$

Δ is the differencing operator defined by $\Delta y_t \equiv y_t - y_{t-1}$. The following procedure is used to determine the lag order parameter k. First, the Akaike Information Criterion and the Schwartz Bayesian Information Criterion (AIC and SBC respectively) are used to select the lag order among specifications k = 1, ..., 13. This is in accord with Hall's (1994) finding that such a lag selection process can improve both the size and power of the ADF test. Then, residuals from the selected specification are tested for serial correlation. If significant serial correlation is detected, the lag length is increased until the model passes the residual test. (In most cases the two criteria yield similar inferences and so in order to conserve space, we only report the results based on the AIC.)

The unit root null hypothesis is rejected if π estimate is significantly less than zero. Since the usual t-statistic for π does not have a standard t-distribution, finite sample critical values that adjusted for both sample size and lag order effects are

[7] Readers familiar with these econometric techniques may skip the section and move directly to Section 5.

used to determine the significance of the ADF statistic (Cheung and Lai, forthcoming).

4.2 Testing for Cointegration

Consider in general an m x 1 vector $\mathbf{x}_t$ of I(1) variables.

$$A(L)x_t = \mu + u_t \tag{5}$$

Where L is the lag operator, $A(L) = I - A_1L - \ldots - A_pL^p$, μ is a vector of constants, and u_t is a vector of white noise Gaussian disturbances, with mean zero and variance Ω. By writing $A(L) = A(1)L^p + (1-L)A^*(L)$, where $A^*(0) = I$, one obtains an equivalent VAR(p) representation for equation (5):

$$= \mu + \Gamma_1\Delta x_{t-1} + \Gamma_2\Delta x_{t-2} + \ldots \Gamma_{p-1}\Delta x_{t-p+1} + \Pi x_{t-p} \tag{6}$$

where Γ_1, Γ_2, ... Γ_{p-1}, Π are m x m matrices of unknown parameters. The Johansen (1991) cointegration test can be conducted as follows.

Consider the m x 1 residual vectors R_{0t} and R_{pt}, which are obtained by implementing the OLS regressions:

$$x_t = c_1 + \gamma_{1,1}(1-L)x_{t-1} + \ldots + \gamma_{1,p-1}(1-L)x_{t-p+1} \tag{7}$$

$$= c_2 + \gamma_{2,1}(1-L)x_{t-1} + \ldots + \gamma_{2,p-1}(1-L)x_{t-p+1} + \tag{8}$$

Then define the product moment matrices of the residuals as:

$$S_{ij} = n^{-1}\sum_{t=1}^{n}\hat{w}_{it}\hat{w}_{jt}' \quad for\ i,j = 0,p \tag{9}$$

Solving the eigenvalue problem:

$$|\lambda S_{pp} - S_{p0}S_{00}^{-1}S_{0p}| = 0 \qquad \textbf{(10)}$$

yields a set of eigenvalues and corresponding eigenvectors. Ranking these from largest to smallest:

$$\begin{gathered}\hat{\lambda}_1 > \hat{\lambda}_2 > ... > \hat{\lambda}_m \\ \hat{V} = (\hat{v}_1, \hat{v}_2, ..., \hat{v}_m) \\ \textit{where} \;\; \hat{V}' S_{pp} \hat{V} = I_m \end{gathered} \qquad \textbf{(11)}$$

Johansen proposes two tests for inferring the number of cointegrating vectors. The trace statistic for testing the hypothesis of at most r cointegrating vectors:

$$J_T = -n \sum_{i=r+1}^{m} \log(1-\hat{\lambda}_i) \qquad \textbf{(12)}$$

Under the null hypothesis of at most r cointegrating vectors, the J_T statistic given by (12) is asymptotically distributed as:

$$tr\{ \int_0^1 (dW) F' \, [\int_0^1 FF' dt]^{-1} \int_0^1 F (dW)' \} \qquad \textbf{(13)}$$

where W is an (m-r) dimensional standard Brownian motion on the unit interval [0,1], and F = $\{F_1(n), ..., F_{m-r}(n)\}$ is defined by $F_j(n) = W_j(n) - \int_0^1 W_j(s)ds$, j = 1,..., m-r. The quantiles of the distribution in (13) are tabulated in Johansen and Juselius (1990) using simulations.

The maximal eigenvalue statistic for testing H(r-1) against r cointegrating vectors

$$\lambda_{max} = -n \log(1-\hat{\lambda}_r) \quad \textbf{(14)}$$

According to our definition of "consistency", forecasts should be cointegrated with the actual series. Failing this, forecasts could drift infinitely far away from the actual series.

4.3 The Cointegrating Vector

A stronger requirement for the consistency of a forecast is that the coefficients in the cointegrating vector are (1 -1). A likelihood ratio can be employed to test the restrictions on the cointegrating vectors.

The maximum likelihood estimators for ß and α are:

$$\begin{aligned} \hat{\beta} &= (\hat{v}_1, \hat{v}_2, ..., \hat{v}_r) \\ \hat{\alpha} &= S_{0p}\hat{\beta} \end{aligned} \quad \textbf{(15)}$$

In the particular case of one forecast, and one actual, series, α and ß are both 2 x 2 matrices.

Johansen and Juselius describe how linear constraints on the cointegrating vector can be tested. Following Johansen (1991) and Johansen and Juselius (1990), the hypothesis of a linear constraint on the cointegrating vector can be expressed as:

$$H_G\!: \ \beta = GB \quad \textbf{(16)}$$

where G is a know m x r_0 matrix of full rank r_0, and B is a r_0 x r matrix of unknown parameters ($m \geq r_0 \geq r$). If $r_0 = r$, the cointegrating space is fully specified. If $r_0 = m$, then no restriction is imposed on ß. Note that G is the matrix that defines the coefficient restriction. In terms of (16), the unitary elasticity restriction is described by (1 -1)', so $r_0 = 1$ in this case.

Johansen (1991) demonstrates that the likelihood ratio test statistic for H_G is given by:

$$J_G = -T\sum_{j=1}^{r} \ln\{(1-\lambda_j)/(1-\lambda_j^*)\}$$

where $\lambda_1^*,\ldots,\lambda_r^*$ are the largest eigenvalues of $G'S_{p0}S_{00}S_{0p}G$ with respect to $G'S_{pp}G$. In other words, the r largest roots of:

$$|\lambda G'S_{pp}G - G'S_{p0}S_{00}^{-1}S_{0p}G| = 0$$

Under the null hypothesis that the restriction H_G is valid, Johansen shows that the limiting distribution of J_G is given by a χ^2 distribution with $r(m-r_0)$ degrees of freedom.

5 Estimation Results

Monthly data from OECD's Main Economic Indicators are used. The exchange rate is the end-of-period spot rate, in US$/foreign currency unit. The narrow measure of money, as defined by OECD, is used for money. Income is proxied by industrial production. Interest rates are either 3 month CD rates, or a daily call money rate, in the case of Japan. Inflation rates are measured as annual log-differences. Finally, tradables and nontradables prices are proxied by producer and consumer price indices, respectively.

Details are provided in the Data Appendix.

5.1 Unit Root Test Results

The unit root test results are presented in Table 1. In accord with previous research, we find that we cannot reject the null of unit roots in all the actual nominal exchange rate series (using the 5% marginal significance level). Similarly, when testing the forecast series, we find that all the forecast series in the longer post-1982 sample also appear integrated. These results therefore fulfill the first condition of consistent forecasts -- that is that the series share the same order of integration. However, for the shorter post-Louvre sample, several $/Yen forecasts reject the unit root null. Since the outcome is a rejection of the null hypothesis, this result cannot be attributed to the low power of the unit root tests. Nor can the source of this result be located in the specific estimation technique -- both OLS and two stage least squares specifications appear to be trend stationary at one-month (six-month for OLS) or all horizons (2SLS), across a variety of models. Hence, it appears that the peculiarity is specific to the forecasts of the Yen/Dollar for this shorter forecasting period.

5.2 Cointegration Test Results

The results of applying the Johansen cointegration test to spot exchange rates and forecasts are reported in Table 2. We applied the cointegration test only to those series that shared the same order of integration.

For the post-1982 sample at the one-month ahead horizon, all forecasts are cointegrated with the actual series, except the Canadian error correction specification for model 2. At the six-month ahead horizon, all but two pairs are cointegrated -- OLS Model 3 for Germany and the error correction specification of Model 1 for Canada. For the one year ahead forecasts, a majority of the pairs fail to reject the null of no cointegration. Interestingly, all the one year ahead Canadian dollar forecasts are cointegrated with the actual exchange rate.

For the shorter post-Louvre sample involving one-month ahead forecasts, we find all the pairs (for which both series appear I(1)) appear cointegrated. For six-month ahead forecasts, however, the null of no cointegration is not rejected for one Canadian dollar exchange rate forecast. Moving to the one year horizon, a large number of series do not reject the no cointegration null -- 19 out of 24 cases for which both series of the pair are I(1). The five series which appear to be cointegrated are once again highly currency specific -- in this case, to the Canadian dollar.

Overall, as the forecast horizon extends out to 12 months ahead, the proportion of cointegrated pairs usually drops drastically: 10 out of 27 in the post-1982 sample. This pattern holds with even greater force for the post-Louvre sample, with only five out of 24 fulfilling the requirement of cointegration. The observed cointegration pattern seems not to be totally explained by the decrease in sample sizes and the consequence drop in the power. For the post-1982 sample, the sample size decreases from 123 to 117 (for the six-month ahead and twelve-month ahead forecasts, respectively). On the other hand the rejection rate of the no-cointegration null drops from 25/27 to 10/27. In the case of post-Louvre sample, the observed rejection frequency declines to 5/24 from 22/23, as the number of observations shrinks to 63 from 69.

These 12-month ahead results seem to be specific to currencies. Japan/US and German/US pairs seldom appear cointegrated. In fact, most of the cointegrated pairs are Canadian. A somewhat disappointing result is that error correction models do not appear to be distinguishable from other specifications, in terms of their cointegration characteristics. However, the one-year ahead horizon is considerably shorter than the three year horizons for which Chinn and Meese (1995) found positive results. Indeed, for the shorter horizons the ECMs did not systematically outperform other estimation methods, in their study.

5.3 Elasticity of Expectations

A requirement of forecast consistency is that not only do the forecast and actual series share the same stochastic trend, but also that the cointegrating vector be (1 -1). The results of implementing this test are reported in Table 3. Using the likelihood ratio test on the data from the post-1982 sample, at the one-month horizon, most of the rejections of unitary elasticity come from forecasts derived from error correction models -- 7 out of the 8 cases reject. The other 6 are distributed evenly over the

OLS and 2SLS specifications. At the six-month horizon, this pattern is repeated, with 6 out of 8 error correction specifications rejecting unitary elasticities. The other 2 rejections are for 2SLS specifications. At the one year horizon, only 1 out of the 10 cases rejects -- a 2SLS specification of Model 1 for the Canadian dollar.

Thus, at the one-month ahead horizon, this restriction is rejected in one half of the cases (at the 5% level). At six-month ahead, only one-third reject. At the one year horizon, only one out of 10 series rejects. However, it is important to note that the number of cointegrated pairs at this horizon is substantially smaller than before. Hence, as the forecast horizon extends forward, the number of cointegrated pairs declines, but of those that are cointegrated, more pass the test of coefficient restrictions.

In the post-Louvre sample, the restriction on the cointegrating vector is only rejected three times, at the 1-month-ahead horizon. This outcome seems to reflect the lower power of the tests given the shorter span of data.

5.4 Discussion

Our results show that it is fairly easy for the generated forecasts to pass the requirement of same order of integration. The failure of the forecast and the exchange rate to have the same order of integration only accounts for 6% of the rejections. Most of the rejections are attributed to the absence of cointegrating relationship and the non-unitary elasticity of forecasts. About

26% of the I(1) pairs of forecasts and exchange rates are found to be not cointegrated.

It is found that cointegration fails to hold the farther out the forecasts extend. At the 12-month ahead horizon, most exchange rate series and their respective forecasts do not appear cointegrated. Among the cointegrated cases, 22% of them fail the unitary elasticity of forecasts condition. Specifically, the non-unitary elasticity results are found mostly among the one-month ahead forecasts and those from the error correction specification in the post-1982 samples. Table 4 summarizes these results. In sum, 87 out of the total 162 cases satisfy the consistency requirement.

The pattern of consistency results appears to be currency specific. The Canadian dollar forecasts exhibit the strongest evidence of forecast consistency. 36 of 87 consistent forecast series are from Canadian Dollar exchange rate models. Compared with the Japanese Yen and German Mark, it may be easier to explain Canadian Dollar exchange rate movements because of the close linkages, both economic and geographic, between the U.S. and Canada.

Regarding the estimation methodology, the error correction approach generates the least number of consistent forecast series. It accounts for 25% of the consistent cases. This seems to be at variance with results reported in Chinn and Meese (1995). However, it is noted that the horizon considered by Chinn and Meese is 3 years while the longest horizon considered in the current study is one year.

The choice of model specifications show no distinguishable effect on the forecast consistency. Of the 87 consistent forecast series, 26 are generated from the flexible price monetary model, 29 from the sticky price model, and the remaining are from the model that incorporates the relative price of tradables and nontradables. This pattern indicates that the inclusion of additional fundamental variables in the exchange rate equation does not detectably improve forecasting performance at these horizons, a result that corroborates the existing consensus regarding the difficulty in forecasting exchange rates.

6 Concluding Remarks

In this study, we have applied a test of rationality looser than that imposed by the typical rational expectations methodology. Specifically, our definition of consistency requires only that the forecast and the actual series be cointegrated (and hence necessarily of the same order of integration), with cointegrating vector (-1 1). These criteria are more appropriate for evaluating forecasts generated from structural models which incorporate macroeconomic data. Such macro data usually impart serial correlation to the forecast series, which invalidates at least one of the standard criteria for rationality.

Forecasts evaluated are one-month, six-month, and twelve-month ahead forecasts for Canadian Dollar, German Mark, and Japanese Yen. These exchange rate forecasts are generated from three commonly used structural exchange rate models. Three different estimating methods and two forecasting periods are considered.

We find that it is fairly easy for the generated forecasts to pass the first requirement of consistency that the series be of the same order of integration. However, cointegration fails to hold the farther out the forecasts extend. At the 12 month ahead horizon, most series and their respective forecasts do not appear cointegrated. Finally, of the cointegrated pairs, the one-month ahead forecasts and those from the error correction estimating method tend to reject the restriction of unitary elasticity of forecasts with respect to actual.

Overall, 87 out of 162 cases satisfy the requirement of consistency. In term of the model performance, our results show that about half of the forecasts generated by each of the three structural models are consistent; that is they have a one-to-one relationship with the actual exchange rates in the long-run. Obviously, this is not an ideal performance. However, the results indicate these structural exchange rate models are capable of generating forecasts that are related to the actual series in the long-run.

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TABLE 1
Augmented Dickey-Fuller Test Results

Model		Sample 1 1983:01-1993:08 1-month ahead	Sample 1 1983:01-1993:08 6-month ahead	Sample 1 1983:01-1993:08 12-month ahead	Sample 2 1987:06-1993:08 1-month ahead	Sample 2 1987:06-1993:08 6-month ahead	Sample 2 1987:06-1993:08 12-month ahead
<u>Differences</u>							
Canada	1	-1.286	-1.390	-1.706	-1.282	-1.219	-1.380
	2	-1.029	-1.409	-1.609	-1.305	-1.279	-1.451
	3	-1.036	-1.394	-1.580	-1.297	-1.292	-1.473
Germany	1	-1.731	-1.870	-2.148	-2.732	-2.466	-2.067
	2	-1.754	-1.888	-2.214	-2.277	-2.156	-2.391
	3	-1.929	-1.963	-2.433	-2.261	-2.140	-2.449
Japan	1	-1.707	-1.828	-1.615	-2.510	-3.486*	-3.414
	2	-1.696	-1.842	-1.619	-2.482	-3.512*	-3.463
	3	-1.704	-1.836	-1.594	-2.475	-2.661	-3.359
<u>Error Correction</u>							
Canada	1	-1.108	-1.426	-1.525	-1.184	-1.402	-1.824
	2	-1.147	-1.473	-1.488	-1.192	-1.475	-1.755
	3	-1.154	-1.516	-1.450	-1.218	-1.530	-1.726
Germany	1	-1.789	-2.211	-2.283	-2.082	-2.831	-1.519
	2	-1.691	-2.200	-2.321	-2.813	-2.711	-1.586

	3	-1.695	-2.156	-2.361	-2.809	-2.693	-1.764
Japan	1	-1.628	-1.762	-1.368	-1.802	-2.933	-2.747
	2	-1.589	-1.801	-1.492	-1.825	-3.043	-2.414
	3	-1.598	-1.819	-1.463	-1.867	-3.063	-2.556
2SLS							
Canada	1	-1.206	-1.249	-1.627	-1.198	-1.238	-1.643
	2	-1.195	-1.665	-1.866	-1.120	-1.386	-1.748
	3	-1.225	-1.978	-2.073	-1.249	-1.943	-2.598
Germany	1	-2.010	-2.164	-3.261	-2.856	-2.669	-2.697
	2	-1.648	-2.197	-2.743	-2.843	-3.164	-2.556
	3	-1.547	-1.792	-3.031	-2.914	-2.971	-2.285
Japan	1	-2.248	-2.005	-2.313	-3.905*	-5.771*	-5.428*
	2	-2.453	-2.240	-2.145	-3.485*	-3.416	-4.930*
	3	-2.332	-2.303	-2.341	-3.492*	-3.517*	-5.027*

Notes: ADF statistics for regressions selected by AIC. * indicates significance at 5% MSL using Cheung and Lai (forthcoming) finite sample critical values. See Appendix 1 for details of the unit root tests.

TABLE 2.1
Cointegration Test Results
Sample 1: 1983:01-1993.08

Model		Forecasting Horizons					
		1-month ahead		6-month ahead		12-month ahead	
Differences							
Canada	1	88.938*	0.905	16.692*	9.178	32.142*	3.919
	2	76.369*	0.856	66.034*	4.582	27.915*	4.134
	3	105.561*	0.991	64.554*	4.324	26.554*	3.734
Germany	1	64.490*	1.865	55.108*	1.702	8.310	3.972
	2	38.244*	1.407	55.135*	1.736	8.216	4.128
	3	36.957*	1.393	15.128	2.272	8.672	3.956
Japan	1	82.985*	0.433	52.992*	0.684	11.901	1.297
	2	81.983*	0.439	51.685*	0.681	11.635	1.314
	3	80.412*	0.432	51.133*	0.693	12.009	1.273
Error Correction							
Canada	1	19.764*	1.096	14.050	4.563	28.289*	3.106
	2	17.090	1.132	66.998*	3.290	25.560*	2.938
	3	66.001*	0.893	69.886*	3.317	24.584*	3.159
Germany	1	52.433*	1.999	55.370*	1.804	5.885	4.297
	2	45.954*	1.720	55.713*	1.768	6.960	4.786
	3	44.564*	1.822	54.638*	1.815	6.774	4.840
Japan	1	104.395*	0.450	59.806*	0.781	13.010	1.071
	2	98.589*	0.464	67.235*	0.798	13.743	

							1.244
	3	100.130*	0.452	62.678*	0.800	12.901	1.218
<u>2SLS</u>							
Canada	1	26.411*	1.066	59.000*	2.549	33.468*	2.465
	2	55.370*	0.797	55.642*	3.045	24.487*	2.761
	3	41.411*	0.888	42.388*	2.923	20.036*	3.095
Germany	1	82.962*	1.084	26.080*	1.778	11.237	3.426
	2	78.926*	1.072	28.664*	1.834	9.117	3.490
	3	78.399*	0.941	42.835*	1.576	11.539	3.114
Japan	1	73.561*	0.370	47.703*	0.817	15.064	1.235
	2	77.417*	0.382	44.767*	0.801	15.888	1.140
	3	74.157*	0.395	37.538*	0.830	17.701*	1.249

Notes: Maximal Eigenvalue statistics for Johansen regressions (lag lengths selected by AIC). * indicates significance at 5% MSL using Cheung and Lai (1993) finite sample critical values. See Appendix 2 for detailed regression results.

TABLE 2.2
Cointegration Test Results
Sample 2: 1987:06-1993.08

Model		Forecasting Horizons					
		1-month ahead		6-month ahead		12-month ahead	
<u>Differences</u>							
Canada	1	51.436*	1.558	46.821*	1.714	29.688*	0.150
	2	58.117*	1.654	44.921*	1.544	28.258*	0.151
	3	58.464*	1.669	45.231*	1.422	26.982*	0.105
Germany	1	22.357*	3.560	30.453*	2.615	7.411	4.148
	2	28.868*	3.758	33.023*	2.707	8.912	4.382
	3	22.151*	2.074	33.856*	2.740	8.518	4.505
Japan	1	51.020*	0.755	--	--	11.538	0.774
	2	51.728*	0.750	--	--	11.755	0.717
	3	51.534*	0.739	20.552*	0.094	11.109	0.814
<u>Error Correction</u>							
Canada	1	64.098*	1.307	16.278	2.282	27.103*	0.220
	2	56.766*	1.311	42.917*	1.291	18.678	0.105
	3	50.561*	1.330	41.420*	1.331	19.224*	0.202
Germany	1	30.048*	4.242	35.002*	3.172	8.846	2.727
	2	22.756*	7.607	33.280*	3.352	10.938	3.094
	3	23.395*	7.756	35.319*	3.081	9.647	3.306
Japan	1	72.128*	0.596	28.950*	0.005	5.054	0.782

	2	72.884*	0.606	30.592*	0.000	6.216	0.981
	3	72.151*	0.595	30.111*	0.002	6.982	1.055
<u>2SLS</u>							
Canada	1	37.035*	1.332	33.724*	0.209	18.697	0.001
	2	48.980*	1.715	28.921*	0.178	12.697	0.018
	3	19.485*	1.611	27.006*	0.112	10.614	0.418
Germany	1	48.015*	6.073	17.852*	3.512	9.603	5.669
	2	43.252*	6.187	18.546*	5.578	10.705	3.137
	3	39.643*	5.870	15.368*	5.636	8.529	2.790
Japan	1	--	--	--	--	--	--
	2	--	--	18.870*	0.248	--	--
	3	--	--	--	--	--	--

Notes: Maximal Eigenvalue statistics for Johansen regressions (lag lengths selected by AIC). * indicates significance at 5% MSL using Cheung and Lai (1993) finite sample critical values. "--" indicates failure to find the same degree of integration between forecast and actual series. See Appendix 2 for detailed regression results.

TABLE 3
Cointegration Vector Restrictions Test

Model		Sample 1 1983:01-1993:08 1-month ahead	6-month ahead	12-month ahead	Sample 2 1987:06-1993:08 1-month ahead	6-month ahead	12-month ahead
<u>Differences</u>							
Canada	1	7.61*	1.19	1.21	0.92	0.11	0.60
	2	3.91*	3.22	1.07	0.00	0.19	0.64
	3	2.46	3.57	1.13	0.05	0.16	0.82
Germany	1	4.64*	0.52	--	1.60	0.11	--
	2	1.02	0.17	--	1.43	0.34	--
	3	1.72	--	--	0.55	0.10	--
Japan	1	0.04	0.06	--	2.68	--	--
	2	0.06	0.09	--	2.54	--	--
	3	0.09	0.32	--	2.70	1.36	--
<u>Error Correction</u>							
Canada	1	8.27*	--	3.24	0.95	--	1.55
	2	--	12.36*	2.69	0.56	0.72	--
	3	11.54*	11.43*	2.46	0.74	1.18	1.19
Germany	1	9.44*	3.64	--	2.18	0.27	--
	2	21.16*	14.27*	--	0.02	0.00	--
	3	14.71*	10.29*	--	0.43	0.40	--
Japan	1	7.81*	7.19*	--	3.92*	0.14	--
	2	2.30	2.99	--	4.51*	0.00	--
	3	4.56*	4.58*	--	3.95*	0.01	--
<u>2SLS</u>							
Canada	1	7.05*	14.89*	7.00*	1.51	0.07	--
	2	6.42*	7.82*	2.76	0.36	0.00	--
	3	0.02	0.01	0.09	0.00	0.54	--

Germany	1	2.45	1.14	--	1.41	0.09	--
	2	5.18*	1.44	--	0.76	1.11	--
	3	1.99	1.33	--	0.05	0.15	--
Japan	1	0.09	0.35	--	--	--	--
	2	0.45	0.98	--	--	0.00	--
	3	0.20	0.51	0.66	--	--	--

Notes: The entries are the test statistics for the restriction on the cointegrating vector of (-1 1), which is distributed chi-squared. A * indicates rejection at the 5% level. "--" indicates failure to find the same degree of integration between forecast and actual series, or a failure to find cointegration using the 5% MSL. See Appendix 3 for detailed results.

TABLE 4
Summary: Consistent Forecasts

	Model	Sample 1 1983:01-			Sample 2 1987:06-		
		1-month ahead	6-month ahead	12-month ahead	1-month ahead	6-month ahead	12-month ahead
<u>Differences</u>							
Canada	1	--	C	C	C	C	C
	2	--	C	C	C	C	C
	3	C	C	C	C	C	C
Germany	1	--	C	--	C	C	--
	2	C	C	--	C	C	--
	3	C	--	--	C	C	--
Japan	1	C	C	--	C	- -	--
	2	C	C	--	C	--	--
	3	C	C	--	C	C	--
<u>Error Correction</u>							
Canada	1	--	--	C	C	- -	C
	2	--	--	C	C	C	--
	3	--	--	C	C	C	C
Germany	1	--	C	--	C	C	--
	2	--	--	--	C	C	--
	3	--	--	--	C	C	--
Japan	1	- -	- -	- -	- -	C	--
	2	C	C	--	--	C	--
	3	--	--	--	--	C	--

2SLS

Canada	1	--	--	--	C	C	--
	2	--	--	C	C	C	--
	3	C	C	C	C	C	--
Germany	1	C	C	--	C	C	--
	2	--	C	--	C	C	--
	3	C	C	--	C	C	--
Japan	1	C	C	C	--	- -	--
	2	C	C	--	--	C	--
	3	C	C	C	--	--	--

Notes: "C" indicates forecasts that pass all three requirements for consistency.

Data Appendix

OVERVIEW

In general, the data are seasonally unadjusted monthly data, derived from OECD Main Economic Indicators (MEI). The data covers the period 1973.06 to 1993.08.

Exchange Rates

- Series: Spot exchange rates.
- Description: End of period spot rates, in US$ per foreign currency unit (US$/C$, US$/DM, US$/¥.

Money Stocks

- Series: M1
- Description: OECD definition narrow money, billions of local currency units, end of period.

Income Proxy

- Series: Industrial production
- Description: Total manufacturing, 1985=100.

Interest Rate

- Series: 3 month interest rate.
- Description: CD rate for US and Canada, Frankfurt rate for Germany, and call money rate for Japan.

Consumer Price Index

- Series: Consumer Price Index
- Description: CPI-All items, 1985=100.

Producer Price Index

- Series: Producer Price Index
- Description: PPI for manufacturing, 1985=100

Inflation Rate

- Series: Inflation rate
- Description: Annual log-difference in the CPI inflation rate.

Real Exchange Rate Indicator Variable

- Series: Ratio of Tradables/Nontradables ratio.

■ Description: log((PPI/CPI)/(PPI*/CPI*)) where * denotes the foreign country.

Appendix 1
Unit Root Test Results

These tables show the results of the ADF tests for:

M1F1	1 month ahead forecasts, 1983:01 onwards
M1F6	6 "
M1F12	12 "
M2F1	1 month ahead forecasts, 1987:06 onwards
M2F6	6 "
M2F12	12 "

Country denotes:

CN	Canadian
GY	Germany
JP	Japan

A suffix __ denotes:

none	OLS
E	ECM
T	TSLS

Series:

XR	actual
FR	forecast from Frenkel-Bilson
DR	forecast from Dornbusch-Frankel
BR	forecast from Balassa

Lag indicates number of lags of first difference used in ADF regression.

ASTAT indicates the ADF statistic

UROOT: "R" indicates rejection at the 5% MSL, "A" indicates failure to reject.

Appendix 2
Cointegration Test Results

These tables show the cointegration test results for:

TM1F1A.XLS 1 month ahead forecasts, 1983:01 onwards
TM1F6A.XLS 6 "
TM1F12A.XLS 12 "
TM2F1A.XLS 1 month ahead forecasts, 1987:06 onwards
TM2F6A.XLS 6 "
TM2F12A.XLS 12 "

Each model is denoted by a code of up to 8 digits. The first 2 denote whether it is for the longer (m1) or shorter (m2) subsample. the third digit indicates the model, either Frenkel-Bilson (f), Dornbusch-Frankel (d) or Balassa (b). The fifth and possibly sixth digits indicate the forecast horizon (either 1, 6, or 12). The final digit indicates the estimation method, either OLS (blank), error correction (e) or two stage least squares (t).

EISTAT are the maximal eigenvalues, in ascending order; TRSTAT are the trace statistics, in ascending order. The EISTAT and TRSTAT under the 10% and 5% CV brackets are the simulated 10% MSL and 5% MSL critical values.

The entries under DECISION indicate the results using either the 10% or 5% critical values for the null hypothesis of 0 cointegrating vectors (r = 0), or of greater than or equal to one cointegrating vectors (r ≥ 1). "A" indicates failure to reject; "R" indicates rejection.

Appendix 3
Test of Restriction on Cointegrating Vector

These tables show the test results of the restriction on the cointegrating vector:

M1F1BST	1 month ahead forecasts, 1983:01 onwards
M1F6BST	6 "
M1F12BST	12 "
M2F1BST	1 month ahead forecasts, 1987:06 onwards
M2F6BST	6 "
M2F12BST	12 "

LAG is the lag used in the Johansen procedure

BSTAT is the test statistic.

C_V is the normalized cointegrating vector.