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**The Long-Run Nominal Exchange Rate:  
Specification and Estimation Issues**

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**Abstract<sup>1</sup>**

We use monthly data from May 1973 to December 1991 to estimate a textbook version of the monetary model of the nominal exchange rate determination. We use a modified version of the Phillips and Loretan (1991) *Two-Sided Dynamic Least Squares*. This method accounts for the serial correlation in the residuals, the simultaneity, and *cointegration* among the *regressors*. The latter condition is consistent with the restriction that the system is homogeneous of degree zero in the money supply differential and the exchange rate. We show that most of the empirical problems known to be associated with monetary models can be ameliorated.

The views expressed in this paper are those of the authors and do not necessarily reflect the views of the Reserve Bank of New Zealand.

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## 1. Introduction and motivation

Nominal Exchange rate determination models have been controversial in the literature. In a series of papers, Meese and Rogoff (eg 1983) and Meese (1990) criticised the exchange rate models of the 1970s, the flexible-price monetary model (FPMM) of Frenkel (1976) and Bilson (1978), the sticky-price monetary model (SPMM) of Dornbusch (1976), and the real interest rate differential model (RIDM) of Frankel (1979) in the sense that their out-of-sample performance is no better than that obtained from 'naive models' such as a random walk model. This paper deals with some of the problems identified by Meese (1990).

Meese (1990) identifies empirical problems of exchange rate models as (a) estimation problems, and (b) specification problems. Potential estimation problems include simultaneity, the reliance on limited information estimation techniques, imposition of inappropriate constraints or misspecified dynamics, and small sample bias. Specification problems include non-linearity in the data generating mechanism for exchange rates, omitted variables, and the inappropriate modelling of expectation formation.

With regard to estimation, problems such as spurious regressions (Granger and Newbold, 1974) were recognised. Other problems are poor fit, statistically insignificant parameters, sign-reversal (see for example, Dornbusch (1978) and Frankel (1979)), serial correlations in the residuals (for example see, Bilson (1978), Hodrick (1978), Putnam and Woodbury (1979), Dornbusch (1978, 1980), Frankel (1984), Haynes and Stone (1981), Backus (1984), and Hooper and Morton (1982)). Simultaneity and multicollinearity have been identified as common problems in this class of reduced-form exchange rate models (for example, see Frenkel and Mussa (1985), p. 724 and MacDonald, 1988, P. 145).

Recent developments in econometrics include new methods to estimate regression equations with integrated series, namely the cointegration analysis (for example Engle and Granger (1987), Johansen (1988) and Johansen and Juselius (1990)). Using these methods, the evidence seems to indicate that the monetary models can explain large portions of the long-run movements in the exchange rate. Baillie and Selover (1987), and McNown and Wallace (1989) use the Engle and Granger (1987) approach. MacDonald and Taylor (1992, 1994), McNown and Wallace (1994)<sup>2</sup>, Chinn and Meese (1995) and Kim and Mo (1995) use the Johansen technique and provide evidence that the monetary model of the exchange rate determination outperforms the random walk model in the long-run.

Unfortunately, one cannot interpret the estimated cointegration vectors from unrestricted VARs as elasticities. In general, cointegration vectors are obtained from a reduced form system where all of the variables are assumed to be jointly endogenous. Consequently, they cannot be interpreted as representing structural equations because, in general it is impossible to go from the reduced form back to the structure. To obtain estimates of long-run elasticities, one must impose some identifying restrictions on the system. A discussion is found in Dickey, Jansen, and Thornton (1991), and recently in Wickens (1996).

An important implication of the *empirical* failure of exchange rate models of the 1970s is that critics of the post-Bretton Woods floating experience have interpreted periods of increased turbulence as representing exchange rate movements that are not justified by the underlying economic fundamentals. The position of the critics can be represented by rejecting the null hypothesis that the variance of the exchange rate is equal to the variance of the fundamentals

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<sup>2</sup> They report incorrect signs of the cointegration parameters.

in favour of the alternative hypothesis that the variance of the exchange rate exceeds that of the fundamentals. Formal representations and tests of ‘excess volatility’ include models of speculative bubbles (Flood and Garber (1987), West (1987), and Frankel and Rose (1994)), Flood and Rose (1995), and noise trading (DeLong et al, 1990). It is well known that testing procedures for bubbles and excess volatility are subject to a ‘joint hypotheses’ criticism. Wu (1995) uses an elegant Kalman filter procedure and provides empirical evidence that strongly argues against speculative bubbles. However, long before Wu’s paper, Mussa (1990) argued that rational bubbles are empirically irrelevant.

Recently, Bartolini and Bodnar (1995), Goldberg and Frydman (1996), Lothian and Taylor (1996), and Mark (1995) have provided evidence supporting the monetary model of the exchange rate determination.

This paper extends recent attempts to re-examine the monetary model.<sup>3</sup> We estimate a textbook version of the monetary model of the exchange rate determination via a modified version of the Phillips and Loretan (1991) *Two-Sided Dynamic Least Squares*. We show that if some of the regressors are cointegrated then the Phillips-Loretan regression equation can be re-written in such a way that is consistent with the restriction that the system is homogeneous of degree zero in the money supply differential and the exchange rate. In this case, the restriction that the money supply differential has a coefficient of unity is arrived at from modifying the Phillips-Loretan method.

We also show that fundamentals implied by the model (money supply differential, income differential, short-term interest rate differential, and the long-term interest rate differential) can reasonably explain movements of the trade-weighted US dollar. The signs of the estimated coefficients are consistent with what the model predicts.<sup>4</sup> The residuals are white-noise, serially uncorrelated, and homoscedastic. And finally, we reject the null hypothesis that the variance of the fundamentals is equal to the variance of the nominal exchange rate in favour of the alternative hypothesis that the variance of the fundamentals is greater than that of the exchange rate, which is consistent with Friedman (1953), and Johnson (1969).

Our methodology is to examine the time series properties of the data first. We report the results in section two. The model is presented in section three. In section four, we fit the model to the data, test standard restrictions, and evaluate the model’s performance. Finally, we test the null hypothesis that the variance of the change in the fundamentals is equal to the variance of the change in the nominal exchange rate. A summary of the results is presented in section five.

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<sup>3</sup> Meese and Rose (1991) investigate one of the potential problems identified by Meese (1990), namely the non-linearity of the data generating process of the exchange rate and reports that the poor explanatory of exchange rate models cannot be attributed to non-linearity.

<sup>4</sup> Frankel (1979) uses long-term interest rate in his model as proxy for expected inflation. Kim and Mo (1995) use both short and long-term interest rates in their cointegration framework. Unfortunately, they do not report the estimated coefficient. Here are some examples. Bartolini and Bodnar (1995) use short-term interest rates. MacDonald and Taylor (1994, 1991) use long-interest rate. McNown and Wallace (1994, 1989) use short-term interest rate. Johansen and Juselius (1992) use short-term interest rate. Baillie and Pecchenino (1991) use short-term interest rate.

## 2. The data

### 2.1 Source of the data

We use the following notation in the paper:  $e_t$  is the trade-weighted US dollar,  $m_t$  is the money supply,  $Y_t$  is real output,  $r_t$  is the nominal 30-day interest rate, and  $r_t^B$  is the long-term nominal interest rate on bonds. Money supply data are seasonally adjusted. All variables are measured in natural logarithms except for the interest rates. An asterisk is used to denote the foreign magnitudes.

The data are from various volumes of the International Monetary Fund's *International Financial Statistics*. The exchange rate is defined as the price in US dollars of a unit of foreign currency (ie, an increase in the exchange rate represents a depreciation of the US dollar) at the end of each month. We use trade-weights (1990) for the United States with the United Kingdom, Germany, Japan, Canada, France, and Italy (Group of Seven) to compute the trade-weighted exchange rate. We choose to work with the trade-weighted dollar rather than working with six bilateral exchange rates because both modelling and estimation are more tractable.

The data are monthly from May 1973 to December 1991.<sup>5</sup>

The money supply is M1. It is defined as currency and demand deposits outside the banks (line 34). For the United Kingdom, money is M0, which is defined as coins and notes outside the banks plus bankers operational deposits with the Bank of England. The money supply is an index with the base equal to a 100 in September 1989. Income is measured by the industrial production index (line 66..c) with the base equal to a 100 in September 1989. The nominal short-term interest rates are money market rates (line 60b), except for the United Kingdom and Canada, where treasury bill rates are used instead (line 60c).<sup>6</sup> The CPI is the consumer price index with the base equal to a 100 in September 1985 (line 64). The long-term interest rates are long-term (10 year) yields on government bonds (line 61). We use the same weights to compute a trade-weighted money supply, the inflation rates, and interest rates. The variables  $m'$ ,  $Y'$ ,  $r'$  and  $r'^B$  are plotted in figures 1 to 4, where  $m'$  is  $(m - m^*)$ ,  $r'$  is  $(r - r^*)$ ,  $Y'$  is  $(Y - Y^*)$ , and  $r'^B$  is  $(r^B - r^{B*})$ . The foreign magnitudes represent the G7 countries without the United States (eg, G6).

### 2.2 Unit roots

We test for unit roots using the Augmented Dickey-Fuller test (Dickey and Fuller, 1979, 1981, and Said and Dickey, 1984), and the Z test (Phillips, 1987). The results are reported in

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<sup>5</sup> Currently, I am testing a longer data set and bilateral exchange rate.

<sup>6</sup> IFS publications do not report money market rates for the UK and Canada.

table 1. The results indicate that we cannot reject the null hypothesis that the variables  $e$ ,  $m'$ ,  $Y'$ ,  $r'$ , and  $r'^B$  are unit root processes. Although there is a disagreement on the power of the tests and there seems to be a general consensus that these variables may contain unit roots.

We *only* test the null hypothesis that there is no cointegration between the variables using the Johansen and Juselius (1990) method.<sup>7</sup> Results are reported in table 2. At the 1 percent level, we reject the null hypothesis that there is no cointegration relationship between  $e$ ,  $m'$ ,  $Y'$ ,  $r'$ , and  $r'^B$ . Correction for the critical value using the procedure in Chueng and Lia (1993) still indicates that the null hypothesis can be rejected, except perhaps for the case of  $e$  and  $r'^B$   $\lambda$  max statistic (the trace statistic still rejects the null). Note that we report a significant cointegration relationship between the regressors  $m'$  and  $r'$ . This relationship will play an important role in estimating the model using the modified Phillips-Loretan method.<sup>8</sup>

### 3. The model

We use the monetary approach to model the exchange rate in the long run. The core equation of the model is the demand for real balances. The demand for money is specified as a function of real income, short-term domestic nominal interest rate, long-term domestic interest rate, short-term foreign interest rate, and the long-term nominal foreign interest rate. Craig and Fisher (1996, p. 163) note that short-term interest rates are probably poor proxies for the opportunity cost of holding money, being overly smoothed or even largely based on relatively inflexible official rates. Thus two rates might simply pick up different characteristics of the opportunity cost of holding money. Also, capital markets are sufficiently well-integrated so foreign interest rates can be justified as explanatory variables in the demand for money equations.<sup>9</sup>

$$m_t^d = P_t + \phi Y_t - \lambda_1 r_t - \lambda_2 r_t^B - \lambda_3 r_t^* - \lambda_4 r_t^{B*} + \xi_t \quad (1)$$

Similarly, the rest of the G7 countries have a similar demand for real balances function.

$$m_t^{d*} = P_t^* + \phi Y_t^* - \lambda_1 r_t^* - \lambda_2 r_t^{B*} - \lambda_3 r_t - \lambda_4 r_t^B + \xi_t^*. \quad (2)$$

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7 We do not impose or test any restriction about the parameters or the cointegration space.

8 For evidence on cointegration see MacDonald and Taylor (1994), McNown and Wallace (1994), Kim and Mo (1995). Johansen and Juselius (1992) find a cointegration relationship between the exchange rate and interest rate. Dickey, Jansen, and Thornton (1991) find cointegration relationships between money, income, and prices. Baillie and Pecchenino (1991) find cointegration between money, income, and prices but not the exchange rate. Sephton and Larsen (1991) find no cointegration relationships in the exchange rate models.

9 Friedman (The Quantity Theory of Money - A Restatement) has many more interest rates in the demand for money equation.

Subtracting equation (2) from (1)-depending on how terms are collected- gives:

$$m_t^d - m_t^{d*} = P_t - P_t^* + \phi(Y - Y^*)_t - (\lambda_1 - \lambda_3)(r - r^*)_t - (\lambda_2 - \lambda_4)(r^B - r^{B*})_t + u_t, \quad (3)$$

where  $u_t$  is  $\zeta_t$  and  $\zeta_t^*$ .

Assuming that  $m^d = m^s$  in the long-run, and that PPP holds as a good approximation (see Lothian and Taylor, 1996), the model reduces to:

$$e_t = m'_t - \phi Y'_t + (\lambda_1 - \lambda_3)r'_t + (\lambda_2 - \lambda_4)r'^B_t + v_t, \quad (4)$$

where  $m'$ ,  $Y'$ ,  $r'$ , and  $r'^B$  are defined in section 2.1, and  $v_t$  is an error term with a mean zero and a variance  $\sigma_v^2$ , and includes  $\zeta_t$ ,  $\zeta_t^*$  and  $u_t$ .

The money supply differential has a one-to-one effect on the exchange rate. Thus, an increase in the money supply at home relative to that abroad last month depreciates the dollar. There are few good reasons for this restriction (see Frankel, 1979). Among the justifications for this restriction (see Frankel, 1979) is that the system should be homogenous of degree zero in the money supply differential and the exchange rate. The parameter  $\phi$  has a negative sign so an increase in real output at home relative to real output abroad appreciates the dollar.

The signs on the interest rates depend on the magnitudes of  $\lambda_1$ ,  $\lambda_2$ ,  $\lambda_3$  and  $\lambda_4$ . In the SPMM (Dornbusch, 1976) and RIDM (Frankel, 1979), the coefficient on short-term interest rate differential is hypothesised to be negative. Its interpretation implies that an increase in the short-term interest rate induces capital inflow and hence, appreciation of the dollar. However, in the FPMM, the sign is positive, which implies a depreciation of the currency because the increase in short-term interest rate reflects inflationary pressures. The long-term interest rate on bonds is used in the RIDM (Frankel, 1979) as a proxy for inflation expectations so the coefficient on the long-term interest rate differential is hypothesised to be positive. Obviously, these are testable hypotheses.<sup>10</sup> Therefore, we will leave the signs of the coefficients on the interest rates to be determined empirically. The *long-run* reduced form model is given by:

$$e_t = m'_t + \beta_2 Y'_t + \beta_3 r'_t + \beta_4 r'^B_t + v_t \quad (5)$$

$$\beta_2 = \phi, \quad \beta_3 = \lambda_1 - \lambda_3, \quad \beta_4 = \lambda_2 - \lambda_4.$$

#### 4. Estimation of the model, and the results

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<sup>10</sup> Figure (4) plots the exchange rate and the long-term interest rate differential.

For estimation we use the *Two-Sided Dynamic Least Squares* method of Phillips and Loretan (1991) to estimate the reduced-form model. To illustrate, we use a two-variable system

$$y_t = \alpha'x_t + u_{0t} \quad (6)$$

$$X_t = x_{t-1} + u_{xt} \quad (7)$$

one can derive the general format given by:<sup>11</sup>

$$y_t = \beta'x_t + \sum_{i=-k}^k \delta_i \Delta x_{t-i} + \sum_{s=1}^T \rho_s (e_{t-s} - \beta'x_{t-s}) + v_t, \quad (8)$$

Phillips and Loretan (1991) show that this single-equation technique is asymptotically equivalent to a maximum likelihood on a full system of equations under Gaussian assumptions. This technique provides modified OLS estimators that are efficient statistically and whose t-ratios can be used for inference in the usual way. The method takes into account both serial correlation in the errors and endogeneity in the regressors that results from the existence of a cointegration relationship between the LHS and the RHS variables.

The modification of the Phillips-Loretan (1991) method when the *regressors* are cointegrated can be derived as follows:

Let

$$x_{1t} = \beta_2 x_{2t} + u_{1t}. \quad (9)$$

$$x_{2t} = x_{2t-1} + u_{2t}, \quad (10)$$

where the errors are assumed to be  $u_{xt} = (u_{1t}, u_{2t})'$ , stationary, Gaussian with zero mean and spectral density  $f_{uu}(\lambda)$  with  $f_{uu}(0) > 0$ .

Writing the errors in terms of the Hilbert projection onto the space spanned by  $\{\Delta x_{2s}\}_{s=-\infty}^{\infty}$  and  $\{u_{1s}\}_{s=-\infty}^{\infty}$  as follows:

$$u_{0t} = \sum_{j=-\infty}^{\infty} P_{2j} \Delta x_{2t+j} + \sum_{j=0}^{\infty} P_{1j} u_{1t-j} + u_{0,xt} \quad (11)$$

Note that in this decomposition we need only be concerned with leads of variables that are I(1). The errors  $u_{0,xt}$  are orthogonal to the full history of the I(1) process  $x_{2t}$  and orthogonal to the present and past of the regressor equilibrium errors  $u_{1t}$  so that  $E(\Delta x_{2t+k}, u_{0,xt}) = 0$ , where

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<sup>11</sup> The errors must be orthogonal, therefore, the presence of the leads in the regression is to eliminate any feedback from  $v_t$  to the RHS variables, and to ensure valid conditioning.

( $k = 0, \pm 1, \pm 2, \dots$ ) and  $E(u_{1t-j}, u_{0,xt}) = 0$  where ( $j = 0, 1, 2, \dots$ ). The spectral density of the process  $u_{0,xt}$  is  $f_{00,x}(\lambda)$

In this model with cointegrated regressors that combine equation (1) with (9), and (10), we obtain the following formulation by substituting (11) into (1).

$$y_t = \beta' \alpha x_t + \sum_{j=-\infty}^{\infty} P_{2j} \Delta x_{2t+j} + \sum_{j=0}^{\infty} P_{1j} u_{1t-j} + u_{0,xt}, \quad (12)$$

or

$$y_t = \beta'_2 x_{2t} + \sum_{j=-\infty}^{\infty} P_{2j} \Delta x_{2t+j} + \sum_{j=0}^{\infty} Q_{1j} u_{1t-j} + u_{0,xt}, \quad (13)$$

where  $Q_{10} = P_{10} + \beta_1$  and  $Q_{1j} = P_{1j}$ ,  $j > 0$ . The corresponding empirical regression equation for our system with  $y_t = e_t$ ,  $x_{1t} = m'_t$ ,  $x_{2t} = r'_t$ ,  $x_{3t} = Y'_t$ , and  $x_{4t} = r'^B_t$  is:

$$e_t = \beta_0 + \beta_3 x_{3t} + \beta_4 x_{4t} + \sum_{i=-k}^k \delta_{1i} \Delta x_{2t-i} + \sum_{i=-k}^k \delta_{2i} \Delta x_{3t-i} + \sum_{i=-k}^k \delta_{3i} \Delta x_{4t-i} + \sum_{j=0}^m \gamma_j (x_{1t-j} - \beta_2 x_{2t-j}) + \sum_{s=1}^T \rho_s (e_{t-s} - \beta_0 - \beta_3 x_{3t-s} - \beta_4 x_{4t-s}) + v_t. \quad (14)$$

Note that the term  $(x_{1t} - \beta_2 x_{2t})$  imposes a unit coefficient on  $x_{1t}$  (the money supply differential) because of the cointegration relationship in equation (9). The theoretical restriction that the money supply differential has a unit coefficient comes as a natural result of equations (9) and (10). This is consistent with the assumption that in the *long run*, the system is homogeneous of degree zero in the money supply differential and the nominal exchange rate (see Phillips, 1997).

The researcher determines the number of lags and leads in equation (14). However, size distortion can easily be introduced by over-fitting the model. We start by fitting a general lag structure of six lags and six leads. We test the significance of these lags and leads by testing backward using the Wald test. Unnecessary lags and leads are eliminated. Each time, we test the residuals for whiteness. We choose a system with three lags and two leads.<sup>12</sup> The variable  $j$  is set equal to zero.

In what follows, we review and interpret our results first. Second, we test some standard hypotheses about the parameter estimates, check the out-of-sample performance of the model, and finally we test the null hypothesis that the variance of the changes in fundamentals is equal to the variance of the changes in the exchange rate.

<sup>12</sup> We tried a system with one lag and one lead only. The fit is as good as that of the higher order system ( $R^2$  is 0.97), the errors are serially uncorrelated but the coefficient on the real output differential is not statistically significant.

The effective sample used in the estimation spans the period November 1973 to October 1991.<sup>13</sup> Estimates of  $\beta_0$ ,  $\beta_2$ ,  $\beta_3$  and  $\beta_4$  are reported in table 3. All coefficient estimates are significant.<sup>14</sup> We report t-ratios and  $\chi^2$  statistics because the t-ratios are only asymptotically normal in non-linear regressions (Galant and Jorgenson, 1979). The coefficient on the real output differential is negative, which means that an increase in the domestic real output relative to foreign output appreciates the dollar. These signs are consistent with all versions of the monetary model of exchange rate determination.

We find the effect of the 30-day interest rate differential on the spot rate to be positive. This implies that an increase in the domestic 30-day interest rate relative to the foreign rate raises (depreciates) the exchange rate. This result is consistent with the FPMM, Bilson (1978) and Frenkel (1976) monetary approach of the exchange rate determination. An increase in the 30-day bill rate at home reduces the demand for domestic currency relative to foreign currency. This is a rise in the exchange rate defined as the price of foreign currency. Although goods prices can be sticky in the short-run, the result is consistent with well integrated capital and money markets. The result, however, is inconsistent with Dornbusch (1976) SPMM and Frankel (1979) RIDM, where the increase in the short-term bill rate at home attracts foreign capital, which lead to an appreciation of the domestic currency.

The long-term interest rate on bonds has a negative effect on the spot rate, which implies an appreciation of the dollar. This implies that the increase in the long-term interest rate at home relative to that abroad appreciates the US dollar. The increase in long-term interest rate on US bonds makes these bonds attractive to residents in the rest of the G7, which induces capital inflows into the US economy in the long run. However, the sign of  $\beta_4$  is inconsistent with Frankel (1979) because he uses  $r_t^B$  as a proxy for inflation expectations. This result is consistent with empirical observations of the data (see figure 4).

The residuals from the restricted model are tested for the absence of serial correlation, whiteness, and homoscedasticity. Results are reported in table 4. The Bartlett's-Kolmogorov-Smirnov test indicates that we cannot reject the null hypothesis that the residuals are white-noise.<sup>15</sup> We also report the DW statistic, Q statistics, and the LM(12) statistic to test for the absence of serial correlations.<sup>16</sup> Results indicate that we cannot reject the null hypothesis that the residuals of the model are serially uncorrelated. The Test for ARCH and the Bruesch-Pagan test also indicate that the residuals are homoscedastic. We conclude that the model is not misspecified.

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<sup>13</sup> Note that we lost observations from the beginning and from the end of the sample because of the lags and leads.

<sup>14</sup> We also estimated the equation with a dummy variable that takes a value of one from October 1979 to October 1982 to account for the Fed's changes in its operating procedure. We found the dummy variable to be insignificant. Engsted and Tanggaard (1994) find that changes in the monetary regimes have no effect on the long-run.

<sup>15</sup> These tests are non-parametric, and conducted using frequency domain analysis.

<sup>16</sup> LM(12) includes LM(1).

Phillips and Loretan recommend that diagnosis of the residuals and out-of-sample forecasting may be necessary because of the possibility of over-fitting. Pagan (1989) shows that because of relationships existing between the simulation residuals and the estimation residuals, no new information is forthcoming from a simulation. The same information can be extracted by a thorough analysis of the estimation residuals. A one-step ahead forecast is sufficient for diagnostic purposes. We have already shown that residuals from the model pass a range of diagnostic tests. Recently, Berkowitz and Giorgianni (1997) show that little can be gained from estimating and forecasting error correction models for the exchange rate for horizons greater than one time period.

In this paper, we look at both the one-step-ahead forecast and the 48-step-ahead forecast of the model. We interpret the forecast as *an additional diagnostic test*. Cautiously, our results may indicate that our model outperforms the random walk model. The one-step forecast is, however, consistent with our analysis of the residuals.

To forecast, we use the restricted model. We fit the model using a sample from November 1973 to September 1987 using the Phillips-Loretan method. Thus, we save 48 observation (June 1987 to October 1991) for  $e_t$ , the contemporaneous RHS variables, their contemporaneous growth rates, and their lags, and also the leads. Then we compute the forecasts for 48-months. At each horizon, we computed root mean squared error (RMSE). To compare with the random walk model, we compute RMSE from a random walk model. Results are reported in table 5. The RMSE is smaller for all seven horizons except one month.

To test the restrictions implied by the model we re-estimate the model without restrictions.<sup>17</sup> Then we check the residuals for serial correlations, homoscedasticity, and ARCH effects. Restrictions on interest rates holds at the 5 percent level, but that on income only holds at the 10 percent level. We test the hypotheses that  $Y + Y^* = 0$ ,  $r + r^* = 0$ ,  $r^B + r^{B^*} = 0$  individually. We also test whether all restrictions hold simultaneously. The Wald statistics reported in table 6 indicate that we cannot reject the restrictions implied by the model for the interest rates, but for output, the statistic is a boarder line.

Figure (5) is a plot of the logarithm of the exchange rate and the *long-run* fundamentals,  $f_t$  computed using the parameter estimates reported in table (3) i.e.  $f_t = 0.60 + m_t' - 1.09y_t' + 0.05r_t' - 0.09r_t'^B$ . The correlation coefficient between  $e_t$  and  $f_t$  is 0.75.<sup>18</sup> Figure (5) seems to show a good correlation between the fundamentals and the trade-weighted dollar, and particularly during the mid 1980s where the US dollar appreciated against all currencies.

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<sup>17</sup> We do not report the results to save a space. Results are available upon requests.

<sup>18</sup> Note that this plot is not for the actual and the fitted values from the regression.

Friedman (1953) and Johnson (1969) argue that the volatility of the spot rate during the floating regime should be explained by the volatility of the underlying macroeconomic fundamentals. To test this proposition, we test the null hypothesis that the variance of the first difference of the fundamentals  $\sigma_{\Delta f}^2$  is equal to the variance of the change in the logarithm of the trade-weighted dollar against the alternative hypothesis that  $\sigma_{\Delta f}^2 > \sigma_{\Delta e}^2$ . The variance  $\sigma_{\Delta f}^2$  is 0.000968 and  $\sigma_{\Delta e}^2$  is 0.000275.<sup>19</sup> The F value is 3.54 which is significant at the 1 percent level. Our results are consistent with Friedman (1953) and Johnson (1969) who argue that volatility of the floating exchange rate is a reflection of the volatility of the underlying market fundamentals.

## 5. Summary

Meese (1990) identifies specification and estimation problems with the exchange rate models of the 1970s, namely, the flexible-price monetary model, FPMM (Frenkel, 1976), the sticky-price monetary model, SPMM (Dornbusch, 1979), and the real interest rate differential model, RIDM (Frankel, 1979). Potential estimation problems include simultaneity, the reliance on limited information estimation techniques, imposition of inappropriate constraints or misspecified dynamics, and small sample bias. Specification problems are non-linearity in the data generating mechanism for exchange rates, omitted variables, or the inappropriate modelling of expectation formation. These problems may have contributed to the empirical failure of the exchange rate models (for example see, Bilson (1978), Hodrick (1978), Putnam and Woodbury (1979), Dornbusch (1978, 1980), Frankel (1984), Haynes and Stone (1981), Backus (1984), and Hooper and Morton (1982)).

Meese and Rose (1991) tested the DGP of major exchange rates and found no evidence of non-linearities. In this paper, we investigate some of Meese's (1990) ideas about the reasons that might be behind the failure of the asset-price monetary models. We estimate the reduced-form monetary model for the post-Bretton Woods trade-weighted dollar with the rest of the G7 countries via a modified version of the *Two-Sided Dynamic Least Squares* method of Phillips and Loretan (1991) and Phillips (1997). The procedure is proved to be asymptotically equivalent to a full information maximum likelihood on a full system of equations. It accounts for serial correlation and the endogeneity of the regressors. We show that:

- The estimated reduced form coefficients to have the signs implied by the model.
- The estimated elasticities have plausible magnitudes. For example, the income elasticity is not different from unity, and the interest rate elasticities are about 0.05 and 0.09 for the 30-day and the 10-year bond interest rates respectively.

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<sup>19</sup> The rejection region  $F > F_{\alpha}$ , where  $F_{\alpha}$  is chosen so that  $P(F > F_{\alpha}) = \alpha$  when the numerator has a  $\nu_1 = (T_1 - 1)$  degrees of freedom, and the denominator has a  $\nu_2 = (T_2 - 1)$  degrees of freedom.

- Restrictions implied by the model hold.
- The residuals of the model are serially uncorrelated, white-noise, and homoscedastic.
- The out-of-sample forecast seems to support our finding that the residuals are indeed white-noise and homoscedastic.
- We also show that the variance of fundamentals as defined in our model to be greater than that of the nominal exchange rate, which is consistent with Friedman (1953) and Johnson (1969).

Our future research considers using bilateral exchange rates instead on the trade-weighted dollar and extending the data to 1998. We will be examining whether the relationship between the exchange rate and macroeconomic variables during the 1990s is different from that in 1980s.

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**Table 1**  
**Testing for Unit Root**

$X_t$	Lags	ADF		ADF		Phillips (Z)			
		<u>Constant, No Trend</u>	<u>Constant &amp; Trend</u>	<u>Constant, No Trend</u>	<u>Constant &amp; Trend</u>				
$e$	11	(a <sub>1</sub> )	-2.1289	(a <sub>2</sub> )	-1.4966	(a <sub>3</sub> )	-3.5380	(a <sub>4</sub> )	-1.9416
		(b <sub>1</sub> )	2.3611	(b <sub>2</sub> )	1.5898	(b <sub>3</sub> )	1.6121	(b <sub>4</sub> )	1.4146
				(c <sub>2</sub> )	2.2902			(c <sub>4</sub> )	1.9447
$m'$		(a <sub>1</sub> )	-2.2945	(a <sub>2</sub> )	-2.6962	(a <sub>3</sub> )	-6.3287	(a <sub>4</sub> )	-11.048
		(b <sub>1</sub> )	3.1516	(b <sub>2</sub> )	2.9442	(b <sub>3</sub> )	4.3037	(b <sub>4</sub> )	3.5977
				(c <sub>2</sub> )	3.8931			(c <sub>4</sub> )	4.1796
	3	(a <sub>1</sub> )	-3.1443	(a <sub>2</sub> )	-3.3856	(b <sub>3</sub> )	13.935	(a <sub>4</sub> )	16.115
		(b <sub>1</sub> )	4.9706	(b <sub>2</sub> )	3.8408	(b <sub>3</sub> )	3.6385	(b <sub>4</sub> )	2.7737
				(c <sub>2</sub> )	5.7338			(c <sub>4</sub> )	4.1286
$r'$	13	(a <sub>1</sub> )	-1.8633	(a <sub>2</sub> )	-1.8440	(a <sub>3</sub> )	-12.723*	(a <sub>4</sub> )	-12.999
		(b <sub>1</sub> )	1.8935	(b <sub>2</sub> )	1.2891	(b <sub>3</sub> )	3.1203	(b <sub>4</sub> )	2.1653
				(c <sub>2</sub> )	1.7768			(c <sub>4</sub> )	3.1783
$r'^B$	0	(a <sub>1</sub> )	-1.8346	(a <sub>2</sub> )	-7.4855*	(a <sub>3</sub> )	-6.5541	(a <sub>4</sub> )	-7.1411
		(b <sub>1</sub> )	1.7380	(b <sub>2</sub> )	1.1883	(b <sub>3</sub> )	1.6672	(b <sub>4</sub> )	1.1367
				(c <sub>2</sub> )	1.7820			(c <sub>4</sub> )	1.7046

**a1:** t-statistic to test ( $D=0$ ), the 5% level C.V.= -2.86. **b1:** F-statistic to test ( $D=0, \delta_0=0$ ), the 5% level C.V.= 4.59.

**a2:** t-statistic to test ( $D=0$ ), the 5% level C.V.= -3.41.

**b2:** F-statistic to test ( $D=0, \delta_0=0$ ), the 5% level C.V. = 4.68.

**c2:** t-statistic to test ( $D=0, \delta_0=0, \delta_1=0$ ), the 5% level C.V. =6.25.

**a3:** Z ( $D=0$ ) without trend, the 5% level C.V.= -14.1.

**b3:** Z ( $\delta_1=0, D=0$ ) without trend, the 5% level C.V. =4.59.

**a4:** Z with a constant and a trend ( $D=0$ ), the 5% level C.V. = -21.7.

**b4:** Z test with a constant and a trend ( $\delta_0=0, \delta_1=0, D=0$ ), the 5% level C.V.=4.68.

**c4:** Z test with a constant and a trend ( $\delta^1=0, D=0$ ), the 5% C.V.= 6.25.

Lags are determined by AIC and SC Information Criteria.

Asterisk means significant at the 5% level.

**Table 2**  
**Testing for Cointegration<sup>a</sup>**

	Lag <sup>b</sup>	r=0		r=1		r=2		r=3		r=4	
		<u>8max</u>	<u>Trace</u>	<u>8max</u>	<u>Trace</u>	<u>8max</u>	<u>Trace</u>	<u>8max</u>	<u>Trace</u>	<u>8max</u>	<u>Trace</u>
(1) $e, m'$	3	12.19*	13.01*	0.83	0.83	na	na	na	na	na	na
(2) $e, Y'$	6	11.52*	14.43*	2.90	2.90	na	na	na	na	na	na
(3) $e, r'$	3	11.65*	13.97*	2.32	2.32	1.76	1.76	na	na	na	na
(4) $e, r'^B$	12	9.56	14.66*	4.66	4.66	1.28	1.28	na	na	na	na
(5) $e, m', Y', r', r'^B$	3	38.41*	83.66*	21.96*	45.25*	14.43*	23.29*	7.90	8.86	0.96	0.96
(6) $m', Y', r', r'^B$	3	22.95*	49.32*	16.60*	26.37	6.55	9.77	3.22	3.22	na	na
(7) $m', r'$	3	11.78*	18.75*	6.97	6.97	na	na	na	na	na	na
(8) $m', Y'$	3	9.68	16.24*	6.55	6.55	na	na	na	na	na	na
(9) $m', r'^B$	3	7.92	10.32	2.40	2.40	na	na	na	na	na	na

a: Using the Johansen procedure, we test the null hypothesis that  $r=0$ , where  $r$  is the number of cointegration vectors.

b: To determine the number of lag differences we start by fitting a general lag structure of 18 lags. Unnecessary lags are eliminated by testing backward using SCI criterion. The residuals are tested for whiteness each time using LM(1) and LM(4) tests.

Asterisks means statistically significant at the 1 percent level.

Chung and Lai (1993) correction for critical value is  $T/T-nk$  where  $T=212$ ,  $n$  is the number of variables, and  $k$  is the number of lags.

5: The estimated cointegration parameters normalised on  $e$ , are  $-0.026$  (constant),  $-1.3$  ( $m'$ ),  $-2.04$  ( $Y'_t$ ),  $0.05$  ( $r'_t$ ), and  $-0.094$  ( $r'^B_t$ ).

**Table 3**  
**Estimated Long-Run Coefficients For Restricted Model**  
**Sample November 1973 to June 1991**  
**Dependent Variable  $e_t$**

Regressors	Coefficients	t-statistics	$\chi_1^2$ <sup>a</sup>
Constant	-0.60	-47.00* (0.0001)	(0.0001)*
$Y_t'$	-1.09	-1.71# (0.0869)	(0.0010)*
$r_t'$	0.05	3.01* (0.0001)	(0.0001)*
$r_t'^B$	-0.09	-7.03* (0.0001)	(0.0001)*
$DW$	2.09		
$\sigma$	0.015		
$\lambda_0$	0.08	2.92* (0.0001)	(0.0001)*
$\rho$	-0.89	-41.66* (0.0001)	(0.0001)*

a: We test the null hypothesis that all parameters = 0 except for output where we test whether the parameter=1.

We estimated the model with a dummy variable that takes a value of 1 from October 1973 to October 1982 the period of the Fed's changes of its operating procedure. We found it to be insignificant.

The P values are for the  $\chi^2$  statistic suggested by Gallant and Jargenson (1979) for non-linear models.

Asterisks denote significant at the 5% level.

**Table 4**  
**Diagnostics Tests for the Residuals of Restricted model**

Test	Null Hypothesis	Restricted Model	Inference
(1) Bartlett's Kolmogrov-Smirnov	White-noise	0.0718	Cannot Reject
(2) Breusch-Pagan	Homoscedasticity	32.11 (0.18)	Cannot Reject
(3) ARCH(1)	Homoscedasticity	1.37 (0.24)	Cannot Reject
(4) ARCH(12)	Homoscedasticity	13.98 (0.30)	Cannot Reject
(5) LM(12)	No higher-order serial correlation	11.89 (0.45)	Cannot Reject
(6) Q(6)	No serial correlation	2.81 (0.24)	Cannot Reject
(7) Q(12)	No higher-order serial correlation	7.63 (0.47)	Cannot Reject

(1) The Bartlett's Kolmogrov-Smirnov critical values are 0.0981 and 0.1176 at the 5% and 1% levels respectively.

(2) The test is  $\sim \chi^2_{(26)}$ .

P-values are in parentheses.

**Table 5**  
**Forecasts of the Restricted Model and the RW model**

Month ahead	RMSE	RMSE(RW)
1	0.0174 <sup>a</sup>	0.0170
3	0.0299	0.0306
6	0.0367	0.0437
9	0.0347	0.0489
12	0.0355	0.0505
24	0.0339	0.0530
48	0.0362	0.1000

a: One-step-a head forecast.

**Table 6**  
**Wald Statistics for Testing Restrictions**

Hypotheses	Unrestricted Model	DF
$H_0^1: Y + Y^* = 0$	5.98(0.01)*	1
$H_0^2: r + r^* = 0$	2.61(0.10)	1
$H_0^3: r^B + r^{B*} = 0$	0.68(0.40)	1
$H_0^4: H_0^1 \cap H_0^2 \cap H_0^3$	9.80 (0.02)*	4

Wald Statistics are  $\chi^2$  distributed.

Asterisks denote significant at the 5% level.

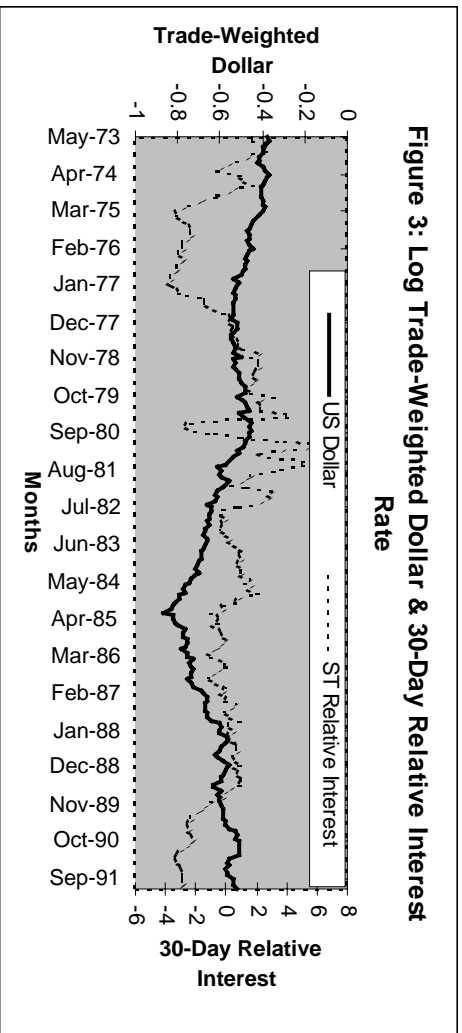
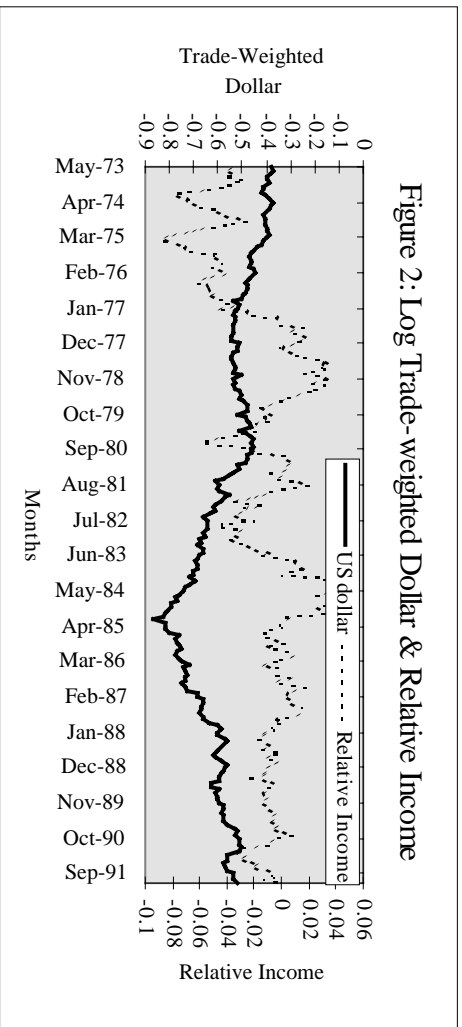
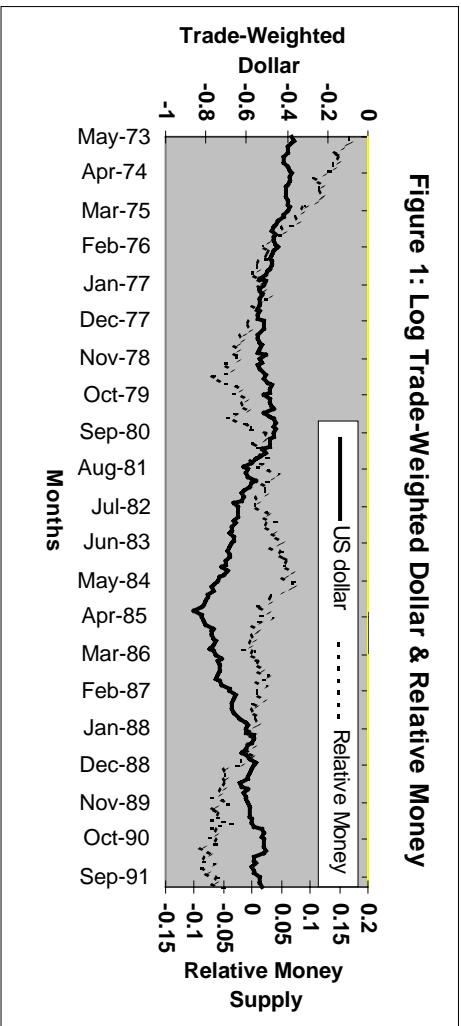


Figure 4: Log trade-Weighted Dollar & 10-Year Bond Relative Interest Rate

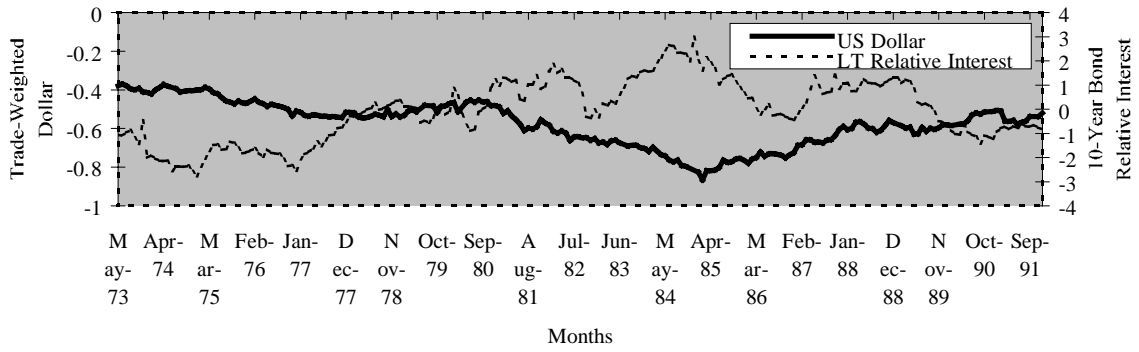


Figure 5: Log Trade-weighted Dollar & Fundamentals

