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An Empirical Investigation of the Japanese /U.S. Exchange Rate. 1979-1994

By

© Xiangping Tang

A thesis submitted to the Faculty of Graduate Studies and Research in partial fulfilment of the requirements for the degree of *Master of Arts*.

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ABSTRACT

Monetary models of exchange rate determination have performed poorly in the 1980s. We briefly investigate performance of the conventional monetary models. The results show that they cannot beat the random walk model. One possible reason for the poor performance may be that short-run dynamics has not been properly modeled. In this thesis, an error-correction model of the monetary approach was constructed and estimated with Japanese yen/US dollar over the first quarter of 1979 to the first quarter of 1994. The estimated model in capable of beating the random walk model and other basic structural models in post-sample forecasting.

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1. INTRODUCTION

Major exchange rates have been floating against each other after the breakdown of the Bretton Woods Agreement in 1973. An extensive literature has emerged on modelling the behaviour of exchange rates since then. The main structural exchange rate model is the monetary model, original proposed by Mundell (1968). According to this theory, the exchange rate is determined by the demand and supply for the stock of money supplies, which are viewed as financial assets.

In this paper we will briefly survey the conventional structural models of exchange rates, ranging from the flexible price monetary model to the sticky price monetary model, the real interest differential monetary model, the current account monetary model, the portfolio balance monetary model, the dynamic stock-flow monetary model, and the relative price monetary model. We will also investigate the performance of various models by estimating them using the data for the Japanese yen and US dollar exchange rate over the period of the first quarter of 1979 to first quarter of 1994. Our initial results indicate that flexible price model, real interest differential model, and portfolio balance model have the expected coefficient estimates. However their performance is less favourable than the random walk model in out-of-sample forecasting. The estimation procedures used in previous models suffer from several statistical problems, many of which have been noted by previous authors. A main statistical issue is that many time-series have stochastic trends in the form of unit roots. Failure to take account of cointegration which has been discussed by Engle and Granger (1987) is an important factor for why the simple models of exchange rate determination have broken down and failed to generate reliable forecasting. While the functional relationship between the exchange rate and other variables may hold in the long run, it may not hold in the short run, such as the behaviour in a quarterly times series. The techniques of cointegration and error-correction are useful tools for extracting a long-run relationship from noise short-run dynamics. Therefore, estimation of the exchange rate equation must take proper account of the cointegration relationship and bring out short-run dynamics by way of an error-correction model. In this paper, an error-correction model is constructed and estimated for the Japanese yen/US dollar.

The monetary model is based on two important behavioral assumptions. purchasing power parity (PPP) holds, 2) uncovered interest rate parity. To investigate the empirical support for the monetary model, we also test whether PPP holds as a long equilibrium relation. Our results show the PPP may not hold as a long run relation in a quarterly data series. However, the deviation from PPP is useful in explaining the short-run behaviour of the exchange rate. Here, we also discuss the literature on foreign exchange market efficiency, which is the same as uncovered interest parity tests, which is another building block of the monetary model of the exchange rate.

Our preferred version of the monetary model is similar to MacDonald and Taylor (1993). The sample period is from the first quarter of 1979 to the first the quarter of 1994. We have found that this monetary model has strong explanatory power and reliable forecasting capability. It can outperform the random walk model.

In this paper, we also estimate other structural models by conventional procedures.

2. LITERATURE REVIEW

2.1 Theoretical Models of Exchange Rate Behaviour

During the past two decades, there has been an enormous growth in the literature on exchange rate economics. But what are the variables that determine exchange rates? In this section, we survey some of the basic structural monetary models of exchange rate determination. All theoretical models considered here are based on the asset market approach.

1. The Flexible-Price Monetary Model

The monetary approach with flexible prices views the nominal exchange rate as the relative price of national monies, therefore its value is determined by relative money demand and supply. This model assumes (i) all goods prices are completely flexible, (ii) domestic and foreign assets are perfect substitutes, (iii) there is perfect capital mobility, (iv) the money supply and real income are determined exogenously, and (v) domestic money is demanded only by domestic residents, and foreign money only by foreign residents. These assumptions imply that purchasing power parity and uncovered interest parity both hold. When the variables are written in logarithmic form, PPP can be expressed as:

$$S = p - p^{*} \tag{1}$$

where S is the spot exchange rate, the domestic currency price of foreign exchange, p and

p^* are the domestic and foreign price. Domestic and foreign prices are determined by demand and supply in domestic and foreign money markets. Following conventional money demand functions employed by most empirical studies, we can infer the money market equilibrium:

$$p = m - \beta_2 y + \beta_3 i \tag{2}$$

$$p^* = m^* - \beta_2 y^* + \beta_3 i^* \tag{3}$$

where m and y are the nominal money supply and real output, respectively, expressed in logarithms, and i is the nominal interest rate (not in logarithms); β_2 and β_3 are coefficients which are assumed to be identical across countries and an asterisk (*) denotes the variables for the foreign country.

From the above three equations, appending the error term e and a constant β_0, we can obtain an estimable exchange rate equation:

$$s=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+e \tag{4}$$

Based on uncovered interest parity, $i-i^* = d = \pi - \pi^*$ is the differential of the anticipated inflation rate or d is the expected rate of depreciation forward premium. Then, the above equation becomes:

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(\pi-\pi^*)+e \qquad (5)$$

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3 E_t(S_{t+1}-S_t)+e \qquad (6)$$

with the following expectation: $\beta_1 = 1$, $\beta_2 < 0$ and $\beta_3 > 0$. In the flexible price monetary model, the domestic interest rate can be regarded as being independent of the exchange rate. The usual explanation is that under perfect capital flow, the domestic interest rates is basically determined by the foreign interest rate. The domestic money supply and real output are determined exogenously too.

From equations (5) and (6), we can see that if the domestic money supply grows faster than the foreign money stock, the exchange rate will depreciate. The intuition here is very obvious. The proponents of this model also justify it by saying that a relative rise in domestic real income creates an excess demand for the domestic money stock. As agents try to increase their real money balances, they reduce expenditure and prices fall until money market equilibrium is achieved. As prices fall, the force of PPP ensures an exchange rate appreciation. Conversely, an increase in the domestic interest rate reduces the demand for money and so leads to an depreciation of the exchange rate. An exactly converse analysis can apply to the foreign variables.

Hodrick's (1978) tests of the above equation for the US dollar/German mark and UK pound/US dollar over the 1970s are supportive of the flexible price model. For example, the dollar/mark equation has the domestic and foreign money supply terms close to plus and minus one, the income coefficients are both correctly signed, but only the US term is significant. Additionally, the interest rate coefficients are both significant, but the German interest rate has the wrong sign.

Bilson (1978) tests this equation for the German mark/UK pound exchange rate over the period January 1972 to April 1976. Although all the coefficients are correctly signed in this equation, only one coefficient differs significantly from zero. The insignificance of the coefficients combined with the high value of R^2 lead Bilson to conclude that multicollinearity is a problem, and the presence of autocorrelation is probably the problem of model misspecification.

Based on quarterly data for Germany over 1973 to 1979, Dorbusch (1983) provides another test of the above equation where he adds the long term interest differential as a proxy for the anticipated inflation differential. The results showed that the estimated coefficients are generally insignificant. The simple monetary model offers a poor description for this period.

In the latter half of the 1970s, the flexible price monetary approach ceased to be an accurate description of the behaviour of exchange rates, especially for a number of small open economies. For example, in the UK over the period 1979 to 1981 the sterling nominal effective exchange rate appreciated substantially even though the UK money supply grew rapidly relative to growth in the world money supply. The real exchange rate appreciated by about 40% over this period and this was followed by an equally sharp fall over the 1981 to 1984 period.

There are several reasons to explain the poor performance of this model. Haynes and Stone (1981) suggest that the reason may be traced to the constraints imposed on relative monies, incomes and interest rates. Subtractive constraints used in this model are particularly restrictive because they can lead to biased estimates and also to sign reversals.

Driskell and Sheffrin (1981) argue that the poor performance of the monetary model may be caused by the failure to account for the simultaneity bias which is caused by having the expected change in the exchange rate on the right hand side of monetary equation.

A further explanation for the failure of the monetary approach equations may be due to the instability of the money demand function. In order to capture the instability of the demand function, some economists incorporate shifts in the money demand function by introducing a relative velocity shift term. However, the presence of significant first-order serial correlation remains a problem in all the reported equations.

The other popular reason for its failure may lie in an assumption underlying all the monetary models, i.e. assets are perfect substitutes. Relaxing this assumption, we will introduce a broader model, the sticky price monetary model.

2. The Sticky-Price Monetary Model

The flexible price model contains a number of strong assumptions, particularly the assumption that all markets are in approximate equilibrium and all goods prices are perfectly flexible. The sticky price monetary model allows that prices are sticky in the short run. This implies an initial fall in the real money supply and consequent rise in interest rates in order to clear the money market. The rise in domestic interest rate then leads to capital inflow and an immediate appreciation of the nominal exchange rate, which can be greater than the long-run equilibrium value. Later a slow depreciation is expected. During the adjustment process, prices and exchange rates may move in opposite directions.

One of the most influential papers that deals with this model, is Dornbusch (1976a,b). His analysis allows different speeds of adjustment for the goods and money markets in his model. The equation under this model states that:

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+e \tag{7}$$

with $\beta_1 > 1$, $\beta_2 < 0$ and $\beta_3 < 0$ expected.

In this model, the domestic money and real outputs are still assumed to be exogenous. The major difference from the flexible price approach is that domestic and foreign goods are no longer considered perfect substitutes, and adjustment of goods prices to new equilibrium is not instantaneous. The lagged response is due to the costs of adjustment, or lack of complete information. However, he does not conduct empirical tests on this model.

The typical test of the Dornbusch model of overshooting is done by Driskill (1981). He analyzes the quarterly dollar/swiss franc rate over 1973-1979 period and reports that the elasticity of the exchange rate in response to an unanticipated monetary disturbance exceeds unity, indicating that overshooting exists.

Other tests of the sticky price model have been conducted by Hacche and Townend (1981) and Backus (1984). Backus (1984) tests the above equation for the

Canadian dollar/US dollar for 1971 to 1980, but finds no evidence of overshooting. The estimated results of a more dynamic version sticky price model by Hacche and Townend (1981) suggest overshooting of the exchange rate, but in other respects the estimated equation is unsatisfactory. Many coefficients are insignificant and wrongly signed and the equation does not exhibit sensible long run properties.

The poor performance of the sticky price model can be attributed to the same reasons that we have mentioned in the flexible price model.

3. The real interest differential monetary model

The shortcoming of the sticky price model is that it does not allow a role for differences in secular rates of inflation. Frankel (1979a) includes the real interest differential as an additional explanatory variable, and proposes the real interest differential model, which exhibits the features of both the flexible and the sticky price model. The flexible price monetary model may apply to the situation where variation in the inflation differential is large, such as in the German hyperinflation of the 1920's, while the sticky price monetary theory may be applicable when variation in the inflation differential is small. But neither of the flexible-price or sticky-price models is appropriate for explaining moderate inflation differential as it has been among the major industrialized countries in the 1970's.

Frankel (1979a) develops a real interest differential model, which combines the sticky-price model in the sense of slow adjustment in the goods market and the flexible-price theory in the sense of a secular rate of inflation. The new feature of this model is that the exchange rate is negatively related to the nominal interest differential, but positively related to the expected long run inflation differential.

The model assumes that PPP only holds in the long run, so the equation becomes:

$$s=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4(\pi-\pi^*)+e \tag{8}$$

with $\beta_1 = 1$, $\beta_2 < 0$, $\beta_3 < 0$ and $\beta_4 > 0$ expected.

Frankel(1979a) tests this hypothesis using monthly data for the mark/dollar rate over the period from July 1974 to February 1978. The regression results from both OLS and Cochrane-Orcutt techniques are robust with regard to the correct signs of all coefficients as hypothesized. Moreover, when the Cochrane-Orcutt technique is used to correct strong first-order autocorrlation, the significance level is even higher.

Haynes and Stone (1981) estimate an unconstrained version of the real interest differential model. They find that coefficient values are broadly supportive of the real interest differential model, in particular, the sign on the relative money term is as predicted by the theory. However, they find the problem of multicollinearity: high R^2 combined with few statistically significant variables. Driskill and Sheffrin (1981) find no support for the real interest differential model and suggest that the reason for its failure may lie in an assumption underlying all the monetary models, i.e., assets are perfect substitutes.

4. The current account monetary model

Some economists have incorporated the current account into asset market models. The proposal is simply to include wealth, in addition to income, as a transactions variable in the monetary demand function. A foreign current account surplus will redistribute wealth from domestic residents to foreign residents, simultaneously raising foreign money demand, lowering domestic money demand, and raising the exchange rate. Unlike in the portfolio balance model, it does not matter if the current account is financed or even more than financed by foreign exchange intervention. Foreign exchange intervention does not alter the level of private sector wealth, only its currency composition, and thus will not affect the demand for domestic or foreign money. Nor, if the intervention is sterilized, will it affect the supply of domestic or foreign money. Frankel (1982) modified his interest differential monetary model by adding wealth (Federal government debt + cumulation of past current account surpluses). The modified equation becomes

$$s=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4(\pi-\pi^*)+\beta_5(w-w^*)+e \tag{9}$$

with $\beta_1 = 1$, $\beta_2 < 0$, $\beta_3 < 0$, $\beta_4 > 0$, and $\beta_5 > 0$ expected.

When this equation is applied to the Greman mark/US dollar exchange rate from 1974-1980, Frankel (1982) finds that with the exception of real income, the coefficients of all variables are significant and have correct signs. Giancarlo et. al (1993), pursue the time varying coefficients approach in estimating this current account model using the lira/dollar exchange rate. Their results are rather dismal, because, with exception of the constant term, no coefficient is significant at 5% level. Most of the other coefficients show wrong signs and the goodness of fit is generally poor.

5. The Portfolio Balance Model

Similar to the flexible-price and sticky-price monetary models, the level of the exchange rate in the portfolio balance model is determined, at least in the short run, by supply and demand in the markets for financial assets. The exchange rate, however, is a principal determinant of the current account of the balance of payments. A surplus (deficit) on the current account represents a rise (fall) in net domestic holdings of foreign assets which in turn affects the level of wealth. Wealth is a major determinant of the level of asset demand and asset demands influence the exchange rate. Thus the portfolio balance model is inherently a dynamic model of exchange rate adjustment that incorporates the current account, the price level, and the rate of asset accumulation.

Like the sticky price model, the portfolio balance model allows one to distinguish between short-run equilibrium which supply and demand equate in asset markets, and the dynamic adjustment to long run equilibrium which is a static level of wealth and the system has no tendency to move over time. Unlike the sticky price model, it also allows for the full interaction between the exchange rate, the balance of payments, the level of wealth, and the stock equilibrium.

The demand for domestic bonds is an increasing function of wealth. Domestic bonds and foreign bonds are not perfect substitutes for each other. There are many reasons why domestic and foreign assets might not be perfect substitutes. Differences in liquidity, government tax laws, default risks, political risks and exchange risks may differentiate domestic bonds from foreign bonds. The portfolio balance approach presumes that the exchange risk makes domestic and foreign bonds imperfect substitutes. In order to diversify the exchange rate risk investors want to balance their bond portfolio between domestic and foreign bonds in proportions that depend on the expected relative rates of return (that is risk premium, $\rho = i - i^* - E\Delta S$, here $E\Delta S$ is the expectation of the change of exchange rate).

A very simple portfolio-balance model assumes static expectation, that is, $E\Delta S = 0$. As a consequence, the exchange rate is simply determined by the relative supply of bonds and the interest differential, that is

$$S=\beta_0+\beta_1(i-i^*)+\beta_2 B-\beta_3 B^*+e \tag{10}$$

where B is domestic bond supply while B^* is foreign bond supply.

A much more general portfolio balance model combined with the interest differential model was derived by Frankel (1983). The exchange rate equation becomes:

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4(\pi-\pi^*)+\beta_5(B-B^*)+e \tag{11}$$

with $\beta_1 = 1$, $\beta_2 < 0$, $\beta_3 < 0$ $\beta_4 > 0$, $\beta_5 > 0$ expected.

In contrast to the monetary approach to the exchange rate, relatively less empirical work has been done on the portfolio balance approach since good disaggregated data on non-monetary assets are difficult to obtain.

By using the above equation and monthly data from January 1974 to October 1978 on the mark/dollar rate, Frankel finds that the coefficient of the relative bond supply appears significant but with a sign that is the opposite from what is hypothesized by the simple portfolio balance model. He attributes this to German government intervention to increase its holdings of dollar assets in order to keep the deutsche mark from appreciating against the US dollar.

Branson, Halttunen, and Masson (1977) estimate a log-linear version of the portfolio balance model similar to this for the deutsche mark/US dollar exchange rate over the period August 1971-December 1976. However, they drop the terms relating to domestic and foreign bond holding because of the ambiguous effect on the exchange rate. However, as Bisignano and Hoover (1982) point out, this rather arbitrary exclusion will generally result in biased regression coefficients. Although the estimates reported by Branson, Halttunen, and Masson (1977) are supportive of the portfolio balance model, once account is taken of acute first-order autocorrelation residuals, only one coefficient, that on the U.S. money supply, remains statistically significant. After specifying a simple reaction function that is purported to capture the simultaneity between the exchange rate and the money supply, they find other estimates remain statistically significant. Branson, Halttunen, and Masson re-estimate their equation using two stage least squares and report more satisfactory estimates of the empirical portfolio balance model.

6. The dynamic stock-flow model

Much of the literature on floating exchange rates addresses the question of the long run relationships between money, exchange rates, and price levels. The exception is the Dornbusch model, that has emphasized the role of slowly adjusting commodity price levels in the short run and intermediate run exchange rate dynamics. The most striking implication of the Dornbusch model is that the exchange rate may overshoot in the short run. Following this initial overshoot, the exchange rate then monotonically approaches its long run equilibrium value. An alternative to Dornbusch's view of exchange rate dynamics has developed. This allows short run undershooting and nonmonotonic exchange rate and price level adjustment instead of short run overshooting and monotonic exchange rate and price level adjustment to long run equilibrium.

Based on the uncovered interest arbitrage assumption, that is i-i^* equals the expected change of the exchange rate from t to $t+1$, money market equilibrium and Dornbusch's sluggish price adjustment equation, Driskill (1981) derives a reduced form which generalizes the Dornbusch model to permit imperfect capital mobility:

$$\begin{aligned} S_t = {} & \beta_0+\beta_1 S_{t-1}+\beta_2 (m-m^*)_t+\beta_3 (p-p^*)_{t-1}+\beta_4 (y-y^*)_t \\ & +\beta_5 (y-y^*)_{t-1}+z_t \end{aligned} \tag{12}$$

with constraints $\Sigma\beta_i = 1$, $\beta_1 < 0$, $\beta_2 > 1$, $\beta_3 > 0$, $\beta_4 < 0$ and $\beta_5 < 0$, here z is a first-order serially correlated random variable.

He applies this equation to the quarterly average data of Swiss franc/U.S. dollar rate for the period 1973-1977. The overall explanatory power of the equation is quite good. All coefficients except the coefficient of first-order autoregression are significantly different from zero at the 0.05 level for the one-tailed test. The Durbin h statistic indicates no serial correlation in the disturbance.

7. The relative price monetary model

The equilibrium exchange rate is influenced by both real and monetary factors which operate through their influence on the relative demands and supplies of monies. Previous work shows that one of the important channels through which real factors affect the exchange rate is the relative price of traded to nontraded goods. In the paper of Clements and Frenkel (1980), they adopt the analytical framework developed by Dornbusch (1976a) by adding the relative price of traded to non-traded goods to the conventional monetary model. Their equation is as follows:

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4\left(\frac{p^T}{p^N}-\frac{p^{T*}}{p^{N*}}\right)+e \qquad (13)$$

where p^T and p^N denote the prices of traded goods and nontraded goods. β_4 is the elasticity of the exchange rate with respect to the relative price which is expected to be equal to the relative share of spending on non-traded goods with value between 0 and 1.

The model is applied to the monthly dollar/pound exchange rate from 1921 to 1925 during which exchange rates were flexible which corresponds with their previous work on exchange rates. They use the wholesale price indices as proxies for the prices of traded goods wages as proxies for prices of nontraded goods. Their results show that the estimated model has expected coefficient signs except the sign of the income ratio is positive.

On the assumption that long run PPP applies only to traded goods, Wolff (1987) suggests adding a relative price of traded goods to nontraded goods to the interest differential model. However, instead of wages, he proxies the price of non-traded goods by consumer price indices. The equation becomes

$$s=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4(\pi-\pi^*)+\beta_5(q-q^*)+e \tag{14}$$

where $q\text{-}q^*=log((p^T/p^N)/(p^{T*}/p^{N*}))$.

He estimates this model and compares it with the model that doesn't include a relative price model. His conclusion is that on average, the forecasting results for the models with the real exchange rate index do not differ very much from the results without the relative price index for the mark/dollar and pound/dollar exchange rates. The index variable makes quite a difference, however, for the yen/dollar exchange rate.

Wolff attributes this improvement of the yen/dollar exchange rate to the fact that Japan has concentrated much more than Germany or the United Kingdom on growth through productivity gains in the tradable sector.

2.2 Unit Root and Cointegration Test

The literature on empirical exchange rate models uses time series data, but time series data have many econometric problems. A common property of the time series used in the exchange rate literature is that most of the variables are integrated of order 1 or *I(1)*, i.e. the time series data have unit roots. The conventional standard tests on the model are inappropriate, when the disturbance term in the regression equation has a unit root. However, if the model with *I(1)* variables has a stationary disturbance term, then we can still regress this model and obtain valid statistical inference. However, if the disturbance is nonstationary, the usual asymptotic results cannot be expected to apply, since spurious regression between totally unrelated but nonstationary variable may occur. Thus, it is important to test for the presence of a unit root. But what is a unit root? Why is it important to economics?

Many economic time series trend upward over time. This is especially true of exchange rates. There are two ways to model trending time series:

$$y_t=\gamma_0+\gamma_1 t+u_t \tag{15}$$

$$y_t=\gamma_1+y_{t-1}+u_t \tag{16}$$

where the error terms u_t will in general not be independently and identically distributed. The first of the model is trend-stationary, that is stationary around a linear trend. In contrast, the second model is a random walk with drift. The drift parameter γ in the second equation is the same as the trend parameter γ in the first equation. It causes y_t to trend upward over time. But the behaviour of y_t is very different in the two cases, because in the first case detrending it will produce a variable that is stationary, while in the second case, if we wrongly try to remove trend by including t as regressor, we will get spurious regression, which is a serious practical problem.

The distinguishing factor for the second model is the presence of a unit root. If all roots are outside the unit circle, the process is stationary. If any root is equal to or less than 1 in absolute value, the process is not stationary. A root that is equal to 1 in absolute value is called a unit root. When a process has a unit root, it is said to be integrated of order one. A series that is $I(1)$ must be differenced once in order to make it stationary. The obvious way to choose between these two equations is to nest them both within a more general model,

$$y_t=\beta_0+\beta_1 t+\alpha y_{t-1}+u_t \tag{17}$$

If y_{t-1} is subtracted from both sides, the above equation becomes,

$$\Delta y_t=\beta_0+\beta_1 t+(\alpha-1)\, y_{t-1}+u_t \tag{18}$$

where Δ is the first difference operator. If $\alpha < 1$, equation (18) is equivalent to equation (16), whereas, if $\alpha = 1$, it is equivalent to equation (15). Thus the unit root test is to test the null hypothesis that $\alpha = 1$ against the one-sided alternative that $\alpha < 1$.

If the variables in a regression equation contain unit roots, in general, the residual will also have a unit root. But there exists a special case in which variable with stochastic trends move together, so that the residual is actually stationary. That is a case of cointegration. If variables are co-integrated, then groups of economic variables are linked by a long run equilibrium relationship. Although the variables may drift away from equilibrium in the short-run, economic forces may be expected to restore equilibrium in the long-run.

Cointegration is a relatively new statistical concept, pioneered by Granger and Weiss(1983), and Engle and Granger (1987). Engle and Granger (1987) define a series x_t to be integrated of order d, i.e., $x_t \sim I(d)$ if the series becomes covariance stationary after being differenced d times to be integrated of order 0. If two variables x_t and y_t are both $I(d)$, then it will generally be true that a linear combination

$$z_t = x_t - \alpha y_t \tag{19}$$

will be also be $I(d)$. However, it may happen that $z_t \sim I(d\text{-}b)$ where $b > 0$ and x_t and y_t are then said to be cointegrated of order (d, b). A particularly important case is where x_t and y_t are both $I(1)$, but z_t is $I(0)$.

If x_t represents a g dimensional vector of random variables and all the components are $I(d)$, then if there exists a vector $\alpha \neq 0$ such that

$$z_t = \alpha' x_t \sim I(d-b) \tag{20}$$

then α is known as a cointegrating vector. The basic monetary model may be useful in describing a long-run relationship characterised by $z_t \sim I(0)$. It implies that z_t may depart from zero in the short-run, but in the long-run, the average value of z_t is zero.

Engle and Granger (1987) formulate and analyze seven testing methods that can be used to test cointegration. In all seven tests, the null hypothesis is noncointegration against the alternative of cointegration. These tests are based on the equation:

$$x_t = \mu + b y_t + e_t \tag{21}$$

1. Durbin-Watson statistic from the cointegration regression. If DW statistic is sufficiently large, e_t is stationary and so x_t and y_t are cointegrated.
2. Modified Dickey-Fuller type regression to test if the estimated time series of the residual e_t from the cointegration regression has a unit root. If there is a unit root, x_t and y_t are not cointegrated.
3. Augmented Dickey-Fuller test, similar to 2, but additional lags of Δe_t are used to be sure that the residuals from the DF regression are serially uncorrelated.
4. Restricted VAR test based on the error correction model,

$$\Delta x_t = c_1 + d_1 e_{t-1} + \epsilon_t \tag{22}$$

$$\Delta y_t = C_2 + d_2 e_{t-1} + \delta x_t + \eta_t \tag{23}$$

If d_1 and d_2 are significantly different from zero in joint form, x_t and y_t satisfy an error correction model and are thereby cointegrated.

5. Augmented Restricted VAR test which is similar to 4, but with additional lags of Δx_t and Δy_t.

6. Unrestricted VAR test to examine the joint significance of the coefficients, β's, from the following two equations to determine if x_t and y_t satisfy a VAR in their levels.

$$\Delta x_t = \alpha_1 + \beta_1 y_{t-1} + \beta_2 x_{t-1} + \epsilon_t \tag{24}$$

$$\Delta y_t = \alpha_2 + \beta_3 y_{t-1} + \beta_4 x_{t-1} + \sigma \Delta x_t + \eta_t \tag{25}$$

If they are significantly different from zero, Δx_t and Δy_t depend on their levels and so may follow an error-correction equation, thereby cointegrated.

7. Augmented UVAR, similar to 6, but with additional lags of Δx_t and Δy_t.

Based on the critical values, they examine the power of the tests and find that first test has the best performance in a first-order system. For higher-order systems, Augmented Dickey-Fuller test is recommended.

Baillie and Selover (1987) estimate monetary models of exchange rate determination. These models include the pure flexible price monetary model of Frenkel

(1976) and Bilson (1978b), the sticky price monetary model of Dornbusch (1976a) and the real exchange rate differential model of Frankel (1979a).

The general equation is:

$$S_t=\beta_1(m-m^*)_t+\beta_2(y-y^*)_t+\beta_3(r-r^*)_t+\beta_4E_t(P-P^*)_{t+1}+e_t \qquad (26)$$

where

- S: end of month exchange rates in terms of US dollars per unit of foreign currency
- m: money supply, M1 equivalent in billions of currency units
- y: index of industrial production as a proxy for real income
- r: short run interest rate as monthly average
- E_tP_{t+1} long run interest rate, or bond yield as a proxy for the expected rate of inflation
- e: error term

Applying an OLS procedure to the monthly data from March 1973 to December 1983 for United States, United Kingdom, France, West Germany, Japan and Canada, the above equation is estimated. Although many of the coefficient estimates are significant, the sign and magnitude are substantially different from what is expected. Then, they estimate the same equation by assuming a first order autoregressive (AR(1)) error process on e_t, $e_t = \rho e_{t-1} + e_t$. The coefficients from the estimation are insignificant and exhibit an

AR(1) error process with a root close to unit, that is $\rho = 1$.

Before testing for cointegration, they check the degree of integration of the variables in equation (26) by employing the Augmented Dickey-Fuller (ADF) statistics. They find that the variables possess different orders of cointegration and conclude that equation (26) exit as long run equilibrium. Even after differencing some variables, the estimates of the equation remain disappointing.

ADF test on the residuals from equation (26), except the output differential and short run interest differential which are integrated of zero degree, shows that the unit root restrictions cannot be rejected. This implies that no cointegrating relationship can be found between the nominal exchange rate and the two other *I(1)* variables.

They also do a cointegration test on the nominal exchange rate and relative price by using the Dickey-Fuller statistic. They find that, except France, there is no cointegrating relationship for the other four currencies. The finding implies that departures from PPP tend to be quite long.

From the above, they conclude that there is no long-run relationship as suggested by the pure monetary model. Therefore, attempts at using these models to forecast will generally produce poor results.

Baillie and Bollerslev (1987) try to examine the cointegration relation between seven spot and forward rates, and to detect the existence of a common stochastic trend in a multivariate time series model.

Before doing that, they first test the unit root in the univariate time series representations of the daily spot and one-month forward exchange rates from March

1,1980 to January 28, 1985 for the currencies of UK, West Germany, France, Italy, Switzerland, Japan and Canada vs US dollar.

Suspecting the presence of serial correlation and time-dependent heteroscedasticity, they apply the Phillips-Perron testing method.

The test involves the following three equations and relevant test statistics.

1. $y_t = \alpha y_{t-1} + e_t$, test statistic $Z(t_\alpha)$ for H_0: $\alpha = 1$ against the stationary alternative H_1: $\alpha < 1$.

2. $y_t = u^* + \alpha^* y_{t-1} + e^*_t$, test statistics $Z(t_\alpha)$, $Z(\phi_1)$ for with and without a drift, that is H_0: $\alpha^* = 1$ and H_0: $u^* = 0, \alpha^* = 1$.

3. $y_t = u^* + \beta(t-n/2) + \alpha y_{t-1} + e_t$, test statistics $Z(t_\alpha)$, $Z(\phi_3)$ for with and without time trend, that is H_0: $\beta = 0$, $\alpha = 1$ and H_0: $u = 0$, $\beta = 0$, $\alpha = 1$ and $Z(\phi_2)$ for time trend and a drift, that is H_0: $\alpha = 1$.

The results from the six test statistics give strong evidence for the presence of a unit root for all seven currencies. All the series appear to be stationary in their first difference by means of Lagrange Multiplier test.

Based on the equation $S_{t+1} = a + bF_t + e_{t+1}$, where e_{t+1} denotes the OLS residual, they test the cointegration for the seven pairs of spot and forward exchange rates for a unit root in e_t. However, the Dickey-Fuller critical values are not adopted as most researchers do. They are concerned that the Dickey-Fuller critical value will be numerically too small, leading to the rejection of a unit root in e_t too often. Instead, they adopt the values reported in Engle and Yoo (1987). The results give strong evidence of cointegration between daily spot and one-month forward exchange rates, i.e. reject the null hypothesis of a unit root of the error term.

2.3 Purchasing Power Parity Test:

There are many explanations for the poor performance of monetary models. Misspecfication of reduced form equations has been linked to the building blocks used to construct these models. Possible misspecifications range from the money demand specifications, uncovered interest parity to purchasing power parity. Other explanations attribute the failure of monetary models to simultaneous equation bias owing the endogeneity of interest rates, the imposition of static and dynamic restrictions, and possible dominance of real over monetary factors in the determination of exchange rates. These explanations may all play a role in the failure of monetary models.

The assumption of uncovered interest parity (UIP) has been questioned. We will discuss the UIP in section 2.4 The Efficient Market Hypothesis. The failure of PPP to hold is another important cause underlying the failure of the monetary models.

Purchasing Power Parity (PPP) is a theory of exchange rate determination. It asserts that the exchange rate change between two currencies over time is determined by the change in the two countries' relative price levels. The absolute form of the purchasing power parity (PPP) theory is as follows:

$$S = P/P^* \tag{27}$$

where S: exchange rate, i.e. number of units of domestic currency per unit of foreign currency.

P: price index in domestic country

P^*: price index in foreign country

The absolute version of PPP relates an exchange rate to the absolute price level of all goods in the two countries on the assumption of the "law of one price". If there are no trade obstacles between these two countries, the price of a given good in one country will be the same as that in the other country when it is quoted in the same currency. If goods prices, denominated in the same currency, are equalized across countries, and if the same goods enter each country's basket of consumption with the same weights, then absolute PPP prevails.

However, absolute PPP does not exist in the real world, because of transport costs, tariffs and quotas among countries. The presence of these obstacles limits the applicability of absolute PPP.

The relative version of PPP therefore restates the theory in terms of changes in relative price levels and the exchange rate: $e = kP/P^*$, where k is a constant reflecting the given obstacles to trade. Given these obstacles, an increase in the domestic price level relative to that abroad will result an equi-proportionate depreciation of the domestic currency:

$$\hat{e} = \hat{p} - \hat{p}^* \tag{28}$$

where a ^ denotes a percentage change.

How well does PPP explain the actual data of exchange rates and national price levels? Actually, almost all versions of the PPP theory do badly in explaining the movement of the exchange rate.

Relative PPP is sometimes a reasonable approximation for low-frequency data. In other words, PPP may hold in the long run, especially if the frequency is longer than a year. For example, a dramatic violation of relative PPP occurs in the years after 1979 for the U.S. dollar/German mark exchange rate. The dollar first sustained a massive appreciation against the mark even though the U.S. price level continued to rise faster than that of Germany. Next the dollar depreciated by far more than PPP would have predicted. Relative PPP did hold over the period of 1964-1983 taken as a whole: over those two decades, the percentage rise in the dollar/mark exchange rate was very close to the percentage increase in the U.S. price level relative to the German price level. Studies of other currencies also confirm the pattern exhibited by the data of the German mark/US dollar.

In our paper, we follow Enders' (1988) methods to test PPP. Enders employs two different tests to estimate the Purchasing Power Parity (PPP) relationship under fixed and flexible exchange rate regimes. His model of PPP is as follows:

$$S_t P_t^* - \alpha P_t = d_t \tag{29}$$

where

S_t = US dollar price of foreign exchange in period t

P_t^* = index of the foreign price level

P_t = index of the US price level

d_t = a stochastic disturbance representing a deviation from PPP

α = constant term

For PPP to hold, α should be equal to 1 and d_t must be integrated of order zero. The author employs two alternative tests, ARIMA and Cointegration.

I. ARIMA Model

The author rewrites the Purchasing Power Parity equation as follows:

$$\frac{S_t \cdot P_t^*}{P_t} = \alpha + d1_t \tag{30}$$

where $S_t \cdot P^*_t/P_t$ is real exchange rate in time t period, $d1_t$ is a stochastic disturbance. For notational simplicity, we define the real exchange rate as R_t.

He assumes d_t is ARIMA(n,0,0), thus the ARIMA model becomes:

$$R_t = \alpha_0 + \sum \alpha_i R_{t-i} + e_t \tag{31}$$

In this case, PPP requires that $\alpha_0/(1-\Sigma\alpha_i) = 1$ and for all characteristic roots to lie within the unit circle.

Enders employs monthly data for three major US trading partners: Germany, Canada and Japan. In order to compare PPP in the 1960s versus the 1970s and 1980s, he divides the data series into two periods: January 1960-April 1971 (the period of fixed exchange rates) and January 1973-November 1986 (the period of flexible exchange rates).

Maximum likelihood estimates were obtained based on AR(1) and AR(2) models.

The results show that for each country in each period, the predicted steady state value ($\alpha_0/(1-\Sigma\alpha_i)$) of the real exchange rate does not significantly differ from unity as expected. However, the results also show that, for all countries, point estimates of the characteristic root for the AR(1) model are not statistically different from unity at conventional 1%, 5% and 10% significance levels, which implies that the random walk hypothesis be rejected.

Using a standard F test for convergence of an AR(2) model, the results show that it is not possible to reject the hypothesis that the real exchange rate follows a random walk. He also found that the real exchange rate was far more volatile in the 1973-1986 period than in the 1960-1971 period, which is consistent with the fact that real supply shocks and lack of monetary coordination were characteristic of the latter period. Overall, the results of the ARIMA test provide mixed support for the PPP hypothesis, i.e. it is hard to claim that PPP held in either period.

II. Cointegration and the Error Correction Model

Another way of testing the hypothesis that the real exchange rate is stationary is to test for cointegration between the series P_t and $S_tP^*_t$. Enders (1988) used three steps to estimate α and to test for the stationarity of the residual.

a. Regress S_tP^* on P_t to get the equilibrium relationship or long run PPP relationship.

b. Check the residuals of the equilibrium regression for stationarity using the Dickey-Fuller (1981) test for a unit root,

$$(1-L)\partial_t=-\pi_0\partial_{t-1}+\sum_{i=1}^{n}(1-L)\pi_i\partial_{t-i}+v_t \qquad (32)$$

Here d_t is stationary, the coefficient for π_0 in the following regression should be statistically different from zero, where v_t is an *i.i.d.* disturbance with zero mean and L is the lag operator.

c. If the null of no cointegration is rejected, the residuals of the equilibrium regression can be used as instruments to estimate an error correction model.

The test for cointegration in Ender's model provided mixed evidence for PPP. Point estimates of long run rate α are far from unity. However, there is strong evidence for cointegration of the US and Japanese price level during the Bretton Woods period and a weak support for cointegration of the US and Canada price levels after 1973. The error correction model shows that the foreign price (Japan or Canada) adjusts to any deviation from PPP but not the US price, which is consistant with the fact that United States occupies a unique position in world trade.

Corbae and Quiliaris (1988) also employ the theory of cointegrated processes to test whether PPP holds as a long run relation. If PPP is true, intercountry trade arbitrage ensures that deviations from a linear combination of spot exchange rates and domestic and foreign price levels should be stationary.

PPP can be expressed as $S{\cdot}P^* = P$ where S, P, P^* denote the spot exchange rate, domestic and foreign price levels respectively. The condition can be rewritten as $r'X_t = S_t$ where $r' = (1, -1, -1)$ and $X'_t = (lnP_t, lnS_t, lnP_t^*)$. If $r'X$ possesses a unit root, that is $r'X$

is nonstationary, then there exists permanent divergence from PPP.

Since a mixture of variables of $I(1)$ and $I(2)$ is trivially cointegrated, they first detect a unit root for all data series by two methods, namely, the Augmented Dickey-Fuller (ADF) test (Said and Dickey, 1984) and the Phillips and Perron (1988) Z statistics. The data used are monthly averages of daily Canadian dollar, French franc, Deutsche mark, Italian lira, Japanese yen and British pound-US dollar spot exchange rates and monthly consumer price indices for each of the above countries from the beginning of July 1973 to the end of September 1986. The results from Z statistics suggest that except for the Japanese price level, the logarithm of spot exchange rates and consumption price indices are nonstationary, i.e. have a unit root. The ADF statistic confirms these finding.

Applying the ADF and Z statistics, Corbae and Oviliavis can not reject the null hypothesis that the real exchange rate has a unit root for all five countries. In other words, the deviations from PPP have no tendency to converge to a long run equilibrium path on the basis of monthly data. From the results, the authors conclude that the monetary model of the exchange rate understates the role of real disturbance in the world economy.

2.4 The Efficient Market Hypothesis:

Uncovered interest parity (UIP) is another important assumption for monetary models. According to uncovered interest parity, the interest differential should be equal to the expected rate of depreciation of the exchange rate:

$$i_t - i_t^* = \Delta S_{t+k}^e \tag{33}$$

Under the condition of covered interest parity (CIP) and rational expectations, UIP implies that the forward rate should serve as an optimal predictor of the future spot rate. This means that on a maintained hypothesis of CIP, indirect tests of UIP can be viewed as a test of efficiency of the forward exchange market. In this litereture, the interest of researchers has been to determine if the foreign exchange market is an efficient market in which exchange rates fully reflect all available information rapidly. In the following section, we will briefly survey efficient market tests.

Following Fama (1970), an efficient market can be defined as a market which "fully reflects" all relevant information instantly. Thus it should not be possible for a market operator to earn abnormal profits. As Levich (1979) has emphasized, in order to implement the hypothesis empirically and to make sense of the term "fully reflect", some views of equilibrium expected returns or equilibrium prices is required. Using equilibrium expected returns, the excess market return on asset i is given by:

$$z_{it} = x_{it} - E(\overline{x_{it}} \mid I_t - 1) \tag{34}$$

where x_{it} is the one period percentage return, I_t is the information set, a bar denotes an equilibrium value and z_{it} represents the excess market return. If the market for asset i is efficient then the sequence z_{it} should be orthogonal to the information set $(E(z_{it}/I_{t-1})=0)$ and serially uncorrelated.

The above formulation of the market efficiency hypothesis (MEH) is assumed under the following conditions:

(i) the market is competitive

(ii) no transactions costs

(iii) information is costless to acquire and is used rationally

From the above formulation and the assumptions, it becomes clear that testing the MEH is conditional on the fulfilment of the assumptions stated above and the additional assumption about the presence of the risk premium. Therefore, the MEH is a joint hypothesis of efficiency and the validity of all of the assumptions, specifically the assumptions relating to the risk premium. By its nature, it is impossible to test the components of the MEH, separately.

There are several cases to be considered here for testing the market efficiency hypothesis.

Type I.

$$s_t = a_0 + a_1 f_{t-1} + u_t \tag{35}$$

If speculators are risk neutral, market efficiency implies that $\alpha_0 = 0$, $\alpha_1 = 1$ (the joint hypothesis of unbiasedness) and the forecast error u_t should be serially uncorrelated and orthogonal to the information set i.e. $E(u_t/I_{t-1}) = 0$.

If speculators are risk averse we would expect α_0 to be significantly different from zero, the error term to be correlated with f_{t-1} resulting in biased and inconsistent estimates of α_1 and the error term to be non-white.

Edwards (1982) argues that the error term u_t contains exogenous shocks or 'news' which have not yet been incorporated into the forward rate since significant events may occur after a forward contract is signed. And these significant events may still have an impact on future spot rates. He further argues that the error structures for all countries may contain 'news' effects, and thus there may be additional information gained from the cross country error structure correlations. So, he uses seemingly unrelated regression estimations and finds that unbiased forward exchange rate hypothesis holds. The four exchange markets studies by Frenkel (1980) all pass the test of unbiasedness: the computed F statistic, which tests the joint hypothesis that $\alpha_0 = 0$ and $\alpha_1 = 1$, cannot be rejected at the 5 percent significance level. Furthermore, the Durbin-Watson statistics do not indicate the presence of first-order autocorrelation.

Although the above equation has become a popular way of testing the efficiency of the forward exchange market, other studies have tested this relationship in rates of change. This follows from the findings of a number of researchers (Hansen and Hodrick, 1980; Meese and Singleton, 1982; Meese and Rogoff, 1984; MacDonald and Taylor, 1987). Since the stochastic processes generating s_t and f_{t-1} may be non-stationary and in fact contain a unit root, thus, on subtracting s_{t-1} from s_t and f_{t-1} in equation (35), we obtain this equation,

Type II

$$s_t - s_{t-1} = a_0 + a_1(f_{t-1} - s_{t-1}) + u_t \tag{36}$$

It is expected that $\alpha_0 = 0$, $\alpha_1 = 1$ and u_t is a white noise process orthogonal to the information set on which agents form their expectations.

Frankel (1980) also reports estimates of the above equation, but cannot reject the null hypothesis that $\alpha_0 = 0$, and $\alpha_1 = 1$. The forward premium explains only a small fraction of the actual variation of the spot rate, suggesting the main part of exchange rate changes are due to the arrival of new information.

Most researchers who tested the above equation found that the joint efficiency hypothesis was rejected by the data. They have attributed such rejection either to the existence of a time-varying risk premium or to the irrationality of expectations.

An alternative test of the optimality of the forward rate as a predictor of the exchange rate change is to conduct orthogonality tests of forecast errors. The estimated equation is of the form:

$$s_{t+k}-f_t=\Gamma X_t+w_{t+k} \tag{37}$$

where X_t is a vector of variable known at time t, which is the econometrician's observed portion of the true information set I_t, which is available to agents. Γ is a vector of parameters, and w_{t+k} is an error term. A test of market efficiency is equivalent to showing that Γ is a null vector. In this case, the error in forecasting the exchange rate using the current forward rate cannot be forecast using current information, i.e. it should be orthogonal to elements of the information set available at time t. Orthogonality tests of efficiency may be split into those that include only lagged forecast errors in the conditioning information set (weak form) and those that include information additional to lagged forecast errors in the information set (semistrong form).

a. Weak form error orthogonality tests take the form:

$$s_{t+1}-f_t^{t+1}=a+b_i\sum_{i=1}^{n}(s_{t+1}-f_t^{t+1})_{t-i}+u_{t+1} \tag{38}$$

where EMH implies that the constant and all other coefficients should equal zero and u_{t+1} should be a white noise process.

Hansen and Hodrick (1980) estimate this equation with i arbitrarily equal to 0 and 1 using weekly data on three month forward rates for three currencies, the Swiss franc, Italian lira and German mark (all against US dollar) and find statistically significant coefficients. Frankel (1979b), who includes a single lagged value of the forecasting error in his study of six currencies for the period January 1973-8, also find statistically significant lagged forecast errors for the German mark/US dollar, UK pound/US dollar and Italian lira/US dollar. Gweke and Feige (1979), who also set $i = 1$ in the above equation, reject weak form efficiency only for one currency (the Canadian dollar/US dollar) out of seven currencies tested for the period 1972.9 to 1977.1.

b. Semi-strong form tests take the form:

The definition of semi-strong form efficiency given earlier refers to the test of the error orthogonality property which utilizes more information than simply the past history of forecasting errors. Hodrick (1980) defines a semi-strong form test as one in which the forecast error is regressed on the own lagged forecast error and lagged forecast errors from other exchange markets. Thus in a regression of the forecast error for market i:

$$s^{j}_{t+1}-f^{j}_{t}=a+\sum_{j=1}^{n} b_{ij}(s^{j}_{t+1}-f^{j}_{t})_{t-1}+e_{t+1} \tag{39}$$

on its own lagged forecast error and the lagged forecasting error from j other markets, semi-strong form efficiency implies that the constant and b_{ij} terms should be statistically insignificant. A semi-strong from test may also be captured by the equation:

$$s_{t}-f_{t-1}=a_{0}+A_{1}X_{t-1}+u_{t} \tag{40}$$

where X represents an $nx1$ vector of publicly available information, such as money supplies, α_0 and vector A_1 are estimated parameters and it is expected that they equal zero.

For the recent floating experience, Hansen and Hodrick find that lagged forecasting errors do have significant explanatory power in the cases of the Canadian-US dollar, the German mark-US dollar and the Swiss franc-US dollar exchange markets and therefore that the EMH must be rejected for these markets. Geweke and Feige (1978) find that the hypothesis that the coefficients are equal to zero can only be rejected for the Canadian dollar from a selection of seven currencies; however, estimating equation 2.4.8 for the seven currencies jointly, using ZSURE, results in the hypothesis that all the coefficients are insignificant being rejected at the one percent level.

3. Bivariate Autoregression Approach

A number of researchers notably Hakkio (1981) test the EMH as a set of non-linear cross-equation restrictions on the parameters of a vector autoregression of spot and forward rates. Suppose the current rate of depreciation and the forward premium together form a jointly covariance stationary process. Then Wold's decomposition (Hannan, 1970) implies the existence of a unique, infinite-order moving average representation. In finite samples, this can generally be approximated by a finite-order vector autogression. In the context of the spot-forward relationship Hakkio (1981) and Baillie et al. (1983) have shown that the spot-forward relationship is modelled as a bivariate vector autoregression:

$$s_{t+1}=\sum_{i=0}^{n} a_i^1 f_{t-i}+\sum_{i=0}^{n} a_i^2 s_{t-i}+u_{1,t+1} \tag{41}$$

$$f_{t+1}=\sum_{i=0}^{n} a_i^3 f_{t-i}+\sum_{i=0}^{n} a_i^4 s_{t-i}+u_{2,t+1} \tag{42}$$

The EMH generates a set of complex non-linear restrictions between a_i^1, a_i^2, a_i^3, a_i^4. In order to test these restrictions the model is first estimated unconstrained then reestimated with the constraints implied by the EMH imposed and a likelihood ratio constructed.

Hakkio (1981) uses this methodology to test the EMH restrictions on a vector autoregression of the rate of depreciation and change in the forward rate using weekly data and one-month forward rates for the period April 1973 to May 1977 for five currencies against the US dollar/Dutch guilder, German mark, Canadian dollar, Swiss franc and UK sterling. He estimates the vector autoregressions both with and without the restrictions imposed and computes likelihood ratio statistics. The results generally reject the EMH for all currencies.

Levy and Nobay (1986) extend this methodology to develop a test of the EMH based on vector autoregression moving average processes for the rate of depreciation and forward premium. Again using weekly data and one-month forward rates for five currencies against the US dollar/UK sterling, German mark, Swiss franc, France franc and Canadian dollar for the period January 1976 to December 1981, the EMH is easily rejected in all cases.

The rejection of the EMH may imply that foreign exchange markets are not efficient, or rational expectations do not hold. However, many researchers also attribute the failure of the EMH to the presence of a time varying risk premium, the existence of the peso problem, the rational bubbles phenomenon, inefficient information processing, or parameter instability.

3. EMPIRICAL STUDIES

The data in this study are taken from the International Monetary Fund's *International Financial Statistics* and *Main Economic Indicators*, and run from the first quarter of 1979 through the first quarter of 1994. A full description of the data and data sources is contained in Appendix 1.

In this section, we plan to evaluate some popular exchange rate models on the basis of a simple data set. We want to investigate how well they perform in forecasting. Suspecting that failure to take account of short-run dynamics may be important, we perform tests for a unit root and cointegration. From the performance of these models, we choose one particular monetary model as the preferred forecasting tool. To investigate the empirical validity for the monetary models, PPP test will be done. From the performance of some basic exchange rate models, we choose one particular monetary model as basis to obtain our particular error-correction model. Forecasting based on our model outperform other monetary models.

3.1 Conventional regression models:

As a concise sample of the existing empirical literature on the exchange rate, we choose the following seven structural models for evaluation. They are:

1. The Flexible Price Monetary Model:

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(\pi-\pi^*)+e \tag{43}$$

with expectation: $\beta_1 = 1$, $\beta_2 < 0$ and $\beta_3 > 0$.

2. The Sticky Price Monetary Model

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+e \tag{44}$$

with $\beta_1 > 1$, $\beta_2 < 0$ and $\beta_3 < 0$ expected.

3. The Real Interest Differential Monetary Model

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4(\pi-\pi^*)+e \tag{45}$$

with $\beta_1 = 1$, $\beta_2 < 0$, $\beta_3 < 0$ and $\beta_4 > 0$ expected.

4. The Current Account Monetary Model

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4(r-r^*)+\beta_5(TB-TB^*)+e \tag{46}$$

with $\beta_1 = 1$, $\beta_2 < 0$, $\beta_3 < 0$, $\beta_4 > 0$, and $\beta_5 > 0$ expected.

5. The Portfolio Balance Model

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4(\pi-\pi^*)+\beta_5(B-B^*)+e \tag{47}$$

with $\beta_1 = 1$, $\beta_2 < 0$, $\beta_3 < 0$ $\beta_4 > 0$, $\beta_5 > 0$ expected.

6. The Dynamic Stock-Flow Model

$$S_t=\beta_0+\beta_1 S_{t-1}+\beta_2(m-m^*)_t+\beta_3(p-p^*)_{t-1}+\beta_4(y-y^*)_t+\beta_5(y-y^*)_{t-1}+Z_t \tag{48}$$

with constraints $\Sigma\beta_i = 1$, $\beta_1 < 0$, $\beta_2 > 1$, $\beta_3 > 0$, $\beta_4 < 0$ and $\beta_5 < 0$,

7. The Relative Price Monetary Model

$$S=\beta_0+\beta_1(m-m^*)+\beta_2(y-y^*)+\beta_3(i-i^*)+\beta_4(r-r^*)+\beta_5(q-q^*)+e \tag{49}$$

where $q\text{-}q^*=log((p^T/p^N)/(p^{T^*}/p^{N^*}))$, $\beta_1 = 1$, $\beta_2 < 0$, $\beta_3 < 0$, $\beta_4 > 0$ and $0 < \beta_5 < 1$.

where:

S: logarithm of exchange rate

m: logarithm of money supply (Japan)

y: logarithm of real income (Japan)

i: short run interest rate (Japan)

r: long run interest rate (Japan)

TB: logarithm of cumulated trade balance (Japan)

B: logarithm of government bond (Japan)

P^T: logarithm of consumer price index (CPI) used as proxies for expected price levels (Japan)

P^N: logarithm of producer price index used as prices of traded goods (Japan)

q: price ratio of traded-nontraded goods

That is $q = \log (p^T/p^N)$

*: The same variables that corresponding to USA

The OLS regression results are reported in Table 1. Some of the models have expected properties while others seem to be inconsistent with the theoretical predictions. For the flexible price model, real interest differential model and portfolio balance model, the estimated coefficients have the expected signs, however the restriction that the estimated coefficients of the money differential be unity is rejected at 5% significant level.

The coefficients of other models do not appear to be consistent with what is expected. In the case of the sticky price monetary model, coefficients have the expected signs, except for the positive estimated coefficient for the short-run interest rate, other coefficients have expected signs. The results of the current-account monetary model are not favourable because most coefficient estimates have wrong signs. In the case of the dynamic stock-flow model, only the estimated parameters on real income and the price level have the correct signs. In addition, the restriction that the sum of estimated coefficients is unity is not confirmed. The relative price monetary model has the wrong signs on the terms of money supply and price ratio of traded and nontraded goods. None of these models can pass the restriction that the sum of monetary supply parameters be unity, including three of the better-performing models.

Overall, half of the coefficient estimates are not consistent with what theories predict, especially the signs, although most of the coefficients are significant. The R^2s are close to one, and Durban Watson values are high; The first order autoregression coefficients are also relatively high.

The results of our regression are consistent with those of previous studies. Meese, R. and Rogoff (1983) have suggested the following factors for the observed poor performance.

First, the empirical work on efficiency of foreign exchange markets and exchange rate risk premium has thrown doubt on the assumption of uncovered interest parity, which is commonly assumed. In the presence of a risk premium, interest parity should be:

$$r_t = r_t^* + \Delta s_{t+1}^e + \rho_t \tag{50}$$

Only if the risk premium, ρ_t, is zero, does uncovered interest parity hold. A number of authors, including Bilson (1981), Cumby and Obstfeld (1981), Geweke and

Feige (1979), Hakkio (1981) have found evidence of risk premia. But the risk premium may not be very large. Therefore, deviations from uncovered interest parity fully explain the poor forecasting performance of the structural models.

Second, the specification of the goods market in the models may be suspect. There is strong empirical evidence that purchasing power parity, on which the monetary approach is based, does not hold in the short run.

A third explanation points to the instability of the money demand functions. As documented by Goldfeld (1976), U.S. money demand shifted downward starting from 1973. Other studies also found that Japanese money demand was unstable. Booth and Poloz (1988) concluded that the performance of real interest differential model improved with the use of shift adjusted money, a free dynamic structure and nonimposition of the static restrictions. Broadly speaking, the underlying parameters in the models shifted over time due to changes in global trade patterns, or changes in policy regimes. And these simple structural models with constant coefficients are not well suited to describe the volatile structural change in the real world.

Fourth, measuring inflationary expectations presents another major problem. The sticky price model, real interest differential model, current account model portfolio balance model and relative price model here are potentially quite sensitive to the proxy used for the expected long run inflation differential. It is possible that long term interest rates and past inflation rates are poor proxies for the expectations variable.

Fifth, imposing the constraint that domestic and foreign variables enter these equations in a differential form implicitly assumes that the parameters of the domestic and foreign money demand and price adjustment equations are equal. While this assumption is conventional in empirical applications, it may be a potential source of misspecification.

The above list of explanations may account for the poor performance of the seven structural models. Among the seven models, the flexible price model, real interest differential model, and portfolio balance model perform better, because all the estimated coefficients of these models have correct signs, except the coefficient of $m - m^*$ is not close to unity. In view of such results, we will later compare the forecasting performance of these three models with that of a random walk model and our preferred version of a monetary model.

In the exchange rate literature, the random walk model (RWM) often is commonly used as a benchmark for comparison of forecasting performance. If the movement of an exchange rate follows a random walk, the best predictor of the future spot rate is, then, the current spot rate. The random walk model is a useful reference. If an exchange rate model cannot beat the RWM , then it is useless as a tool for forecasting. The RWM assumes that the change in the spot rate is unpredictable. The spot rate has already incorporated all relevant information to date and is therefore is the best forecast of the future spot rate. It's form is:

$$S^{e}_{t+1} = S_t \tag{51}$$

where $S_{t+1}{}^{e}$, and S_t are both in logarithm forms.

The accuracy of out-of-sample forecasting used in this paper for comparing structural models with the RWM is measured by three statistics: mean error (ME), mean absolute error (MAE) and root mean square error (RMSE). These are defined as follows:

$$ME = \sum_{s=0}^{n_k-1} \frac{[X_{t+s+k} - Y_{t+s+k}]}{n_k} \tag{52}$$

$$MAE = \sum_{s=0}^{n_k-1} \frac{|X_{t+s+k} - Y_{t+s+k}|}{n_k} \tag{53}$$

$$RMSE = [\sum_{s=0}^{n_k-1} \frac{[X_{t+s+k} - Y_{t+s+k}]^2}{n_k}]^{\frac{1}{2}} \tag{54}$$

where k : forecast step,

n_k : total number of forecasts,

Y_t : actual exchange rate,

X_t : the forecast exchange rate,

t : forecasting beginning period.

Root mean square error is the principal criterion for comparing forecasts. The other two measures are also useful. For example, RMSE is inappropriate under some conditions such as when exchange rates are governed by a non-normal stable Paretian process with infinite variance. Therefore, we also include mean absolute error, which is especially useful when the exchange rate distribution has fat tails. Mean error is another measurement similar to MAE. Both of them can help to tell whether a model systematically over- or under-predicts.

In this paper the initial sample period for estimation is 1979.I-1990.II. The forecasts are a one-quarter-ahead forecast generated for one horizon over the post sample period of 1990.III-1994.I. The RMSE, MAE and ME for each model are listed in Table 5.1. Each RMSE value is larger than corresponding MAE value, while each MAE value is larger than MA value, indicating that the models do not systematically overpredict or underpredict. We also note that the random walk model, whose RMSE is 0.02308, MAE, 0.01024, ME, 0.00629, outperforms in RMSE, MAE and ME over other models. Our results presented above obviously show that the flexible price, real interest differential and portfolio balance monetary approach are not significantly better than the random walk model, a result consistent with previous studies.

In addition to measures of forecast errors, we also compare models on the basis of forecasting the direction of change. This yardstick of comparison is less stringent, but very useful for practical purposes. As indicated by Table 5.2 the three structural models correctly predicted the direction of change 9 out of 15 forecasting quarters (Table 5.2). But in general, structural models are less favourable than the random walk model, which predicted the direction of change 7 over 15 forecasting quarters.

In the previous section, we state the shortcomings of these models. Failure to beat the random walk model may due to these reasons. However, there may be another reason to explain this failure. It is that these models fail to account for the long run and short run relationships among the variables. They only consider the long run equilibrium in the model. As reviewed earlier, the exchange rate and variables in exchange rate equations all appear to be nonstationary. The low DW statistics and strong presence of first order-autocorrelation are evidence that simple regression in levels with nonstationary variables is not appropriate. In view of this consideration, in the following part, we obtain our model by using an error-correction model which takes account of these factors and may improve upon the conventional monetary models. Our model is derived from the flexible price model, the best performer in predicting the future exchange rate among the seven models that we have evaluated.

3.2 Unit root and cointegration tests

In order to determine the degree of integration of the variables in the above seven models, we employ the conventional Augmented Dickey Fuller (ADF) statistic. The hypothesis of a unit root in an autoregression is tested by means of estimating the following two different models.

$$\text{(1)} \qquad \Delta y_t=\alpha_0+\alpha_1 y_{t-1}+\sum_{j=1}^{k}\gamma_j \Delta y_{t-j}+e_t \tag{55}$$

I. Test hypothesis: H_0: $\alpha_1=0$

H_1: $\alpha_1 \neq 0$

Test statistics: $n\alpha_1$ and τ

II. Test hypothesis: H_0: $\alpha_0=\alpha_1=0$

H_1: $\alpha_0 \neq 0$ or $\alpha_1 \neq 0$

Test statistics: F-test Φ_1

$$(2) \qquad \Delta y_t=\alpha_0+\alpha_1 y_{t-1}+\alpha_2 t+\sum_{j=1}^{h}\gamma_j \Delta y_{t-j}+e_t \qquad (56)$$

I. Test hypothesis: H_0: $\alpha_1=0$

H_1: $\alpha_1 \neq 0$

Test statistics: $n\alpha_1$ and τ

II. Test hypothesis: H_0: $\alpha_0=\alpha_1=\alpha_2=0$ (zero drift)

H_1: $\alpha_0 \neq 0$ or $\alpha_1 \neq 0$ or $\alpha_2 \neq 0$

Test statistics: F-test Φ_2

III. Test hypothesis: H_0: $\alpha_1=\alpha_2=0$ (non-zero drift)

H_1: $\alpha_1 \neq 0$ or $\alpha_2 \neq 0$

Test statistics: F-test Φ_3

It should be noted that the critical values of the above test statistic based upon the standard t-statistic crucially depends on the sample size and whether or not a constant term is included. The value of h is chosen to be sufficiently large so that the e_t is a close approximation to white noise. However, an unnecessarily large value of h and the consequent inclusion of insignificant lagged Δy_t variables can reduce the power of the test. But when we run the model by SHAZAM, the program will automatically chose the suitable lagged value.

The results of applying the ADF test are reported in Table 2.1 and Table 2.2. They show that the null hypothesis of a unit root cannot be rejected at the conventional significant level. The tests of Augmented Dickey and Fuller on all variables, m - m^*, y - y^*, i - i^*, r - r^*, P - P^*, q - q^*, B - B^* are integrated of degree one except TB - TB^* which is integrated of degree two.

In the case of the exchange rate, we find that ΔS_t is stationary. Therefore, the S_t series is integrated of degree one.

While in general a linear combination of variables integrated of degree one is also integrated of one, there is a special case in which it is integrated of degree zero, or, the variables are cointegrated. Based on the seven structural regression models, we check if the linear combinations of all the variables in each models are cointegrated. The results (Table 2.3) for cointegration tests are disappointing with few parameter estimates approaching significance. The null hypothesis of no cointegration be rejected for all the seven structural models, that is to say the residual series for these models may all possess unit roots. S_t are not cointegrated with other variables in each model. So from our seven conventional regression models, even though the flexible price, real interest differential, and portfolio monetary models appear to have the estimated coefficients with the expected signs, there is a risk that these structural models may capture only spurious relationships among variables that contain stochastic trends.

3.3 Model: Monetary Approach to the Exchange Rate

The presence of unit roots and noncointegration may be the reason why conventional models fail. What should one do is a situation in which all the variables are not stationary? The classical approach to dealing with a model with integrated variables, especially in the time series literature, has been to difference them as many times as needed to make them stationary. This approach has the merit of simplicity. Once all series have been transformed to stationarity, dynamic regression models may be specified in the usual way, and standard asymptotic results apply. The problem with this approach is that differencing eliminates the opportunity to estimate any relationships between the levels of the dependent and independent variables. In the case of cointegration, such relationships exist, and they are often of considerable economic interest. Thus simply using differenced data is often not an appropriate strategy.

A second approach is to estimate some sort of an error correction model. The error correction model is a useful way to estimate a regression model when variables have unit roots and variables are cointegrated, furthermore, it combines the long run and short run factors together.

But what is the error-correction model? The Engle-Granger Representation Theorem implies that any system of cointegrated variables that has an ARIMA representation may be written in the form of an error-correction model. Our model is a special case of an ARIMA model because there is no autocorrelation evidence in our error term. We will verify this later. In particular, if we have a system of variables X that are cointegrated with cointegration vectors given by a matrix β, then we write the dynamics of the system as:

$$\Delta X_t=\alpha(\beta X_{t-1})+\sum_{j=1}^{n}\gamma_j\Delta X_{t-j}+e_t \tag{57}$$

Here, the vector βX gives the deviation of the system from its long run equilibrium relationships, while α measures the speeds of adjustment towards those relationships for each variable and the lagged $\gamma_i \Delta X$ captures other short-run dynamics. If β were known, there would be no problem in estimating the error correction model by least squares, otherwise, we have to estimate the long run equilibrium relationship first to get the estimated value of β, and then substitute it in the estimating ECM model.

We will estimate our exchange rate equation using the non-linear least squares methodology described by Phillips and Loretan (1991). Unlike the Engle and Granger (1987) two-step procedure, with this method we can estimate simultaneously both the long- and short-run relationships in the form:

$$\Delta y_t = \rho(L)(y_{t-1} - x_{t-1}\beta) + z_t(x_t, y_t)\gamma(L) + u_t \tag{58}$$

where y is the dependent variable i.e., spot exchange rate, x and z are vectors of long-run and stationary short-run explanatory variables, respectively. $\rho(L)$ and $\gamma(L)$ are polynomials in the lag operator, β is the cointegration vector between y and x, u_t is an i.i.d mean zero error term. Phillips and Loretan (1991) find that, given the presence of cointegration, standard hypothesis-testing statistics, such as t-ratios and F-statistics, can be used in the usual way to test hypotheses about β and γ (not α). In general, the above equation may include any number of lags of the variables.

In this paper, we will use such an error-correction model to forecast the exchange rate, the model is derived from the flexible-price monetary formulation (57). The reason that we choose the flexible-price model is that it has a better performance based on previous evaluation. The basic model is:

$$S_t=m_t-m_t^*-\eta y_t+\eta^* y_t^*+\varepsilon(i_t-i_t^*), \tag{59}$$

where the definition of the variables here is the same as before, but we relax the constraint that the parameters of domestic and foreign variables are identical except for the interest rate differential. The restriction for identical coefficients may not be valid, and is not essential for the theory. Haynes and Stone (1981) found that such restrictions on the monetary models were rejected by their data. A further assumption underlying the flexible price monetary model is that the risk-premium ϱ_t is identically zero and therefore UIP holds:

$$i_t=i_t^*+E_tS_{t+1}-S_t, \tag{60}$$

The combination of the above equations will generate the following spot exchange rate equation:

$$s_t=\frac{1}{1+\varepsilon}m_t-\frac{1}{1+\varepsilon}m_t^*-\frac{\eta}{1+\varepsilon}y_t+\frac{\eta^*}{1+\varepsilon}y_t^*+\frac{\varepsilon}{1+\varepsilon}E_tS_{t+1}, \tag{61}$$

The above equation includes the unobserved variable E_tS_{t+1}. By applying a mathematical expectation to the above equation, we generate an expression which specifies the expected value of exchange rates in any future period. The expression may be repeatedly substituted into the above equation to obtain a "news model":

$$s_t=\frac{1}{1+\varepsilon}\sum_{j=0}^{\infty}\left(\frac{\varepsilon}{1+\varepsilon}\right)^j E_t[m_{t+j}-m^*_{t+j}-\eta y_{t+j}+\eta^* y^*_{t+j}] \qquad (62)$$

This is a monetary model with rational expectations. It differs from the simple flexible price monetary model. The exchange rate in this rational expression monetary model not only depends on current values of money supplies and real incomes, but also depends on their infinite expected future values. The solution for the spot exchange rate requires expectations of the future paths of all the exogenous variables. So it is necessary to obtain an observable expression for the expectation terms it contains.

Following Campbell and Shiller (1987), S_{t-1} and x_t are subtracted from either side of the above equation and rearranging it:

$$s_t-s_{t-1}=\sum_{i=1}^{\infty}\left[\frac{\varepsilon}{1+\varepsilon}\right]^i E(\Delta x^*_{t+i}|I_t) \qquad (63)$$

where $x_t = [m_t - m^*_t, -\eta y_t + \eta^* y^*_t]$

Since m_t, m_t^*, y_t, and y_t^* are first-difference stationary $I(1)$ variables, then the above equation is $I(0)$. Thus the right hand side of the above equation is also stationary, i.e. the exchange rate is cointegrated with m_t, m_t^*, y_t, and y_t^*.

The right side of above equation is an infinite stationary series. According to the Wold Decomposition Theorem, there exists a finite-order auto regressive representation when the infinite series is inverted. Following this model, our error-correction model is derived from the above infinite stationary series equation.

Also we need to establish a cointegration equation in order to capture the long-run relationship among these variables. The cointegration equation is:

$$s_t=\beta_0+\beta_1(m_t-m_t^*)+\beta_2 y_t-\beta_2^* y_t^*+\beta_3(i_t-i_t^*)+z_t \tag{64}$$

The explanatory variables in the above equation are in levels. We also need to test whether the variables are cointegrated. If they are cointegrated, then we can specify an easily interpreted form for an exchange rate equation, like an error-correction model.

There are several tests for cointegration. The first approach is that advocated by Engle and Granger (1987). As we have mentioned before, their two-step procedure requires us first to estimate a static ordinary least squares regression of the above equation, and then test the resulting residuals for stationarity.

An alternative single-equation cointegration test with possibly more power has been proposed by Hansen (1990). The potential increase in power is due to the fact Hansen's approach yields a limiting distribution that is invariant to the number of regressors. Briefly, the test requires us to estimate the same static regression as above using the Cochrane-Orcutt procedure.

Another approach to testing for cointegration is the systems approach developed by Johansen (1988) and Johansen and Juselius (1990). Recent Monte Corlo work by Gonzalo (1989) suggests that the Johansen systems approach to estimating and testing for cointegration performs better than both the methods advocated by Engle and Granger (1987) and by Hansen (1990). This approach involves using maximum likelihood methods to estimate a full vector-autoregressive system of equations that includes both levels and first differences. The rank of the coefficient matrix corresponding to the level terms is then used to test the number of cointegrating vectors.

In this paper, we only apply the approach advocated by Engle and Granger (1987) since it is much easier to carry out. Our test statistic is τ = -4.01 whose absolute value is larger than its critical value, -3.81. Given that the estimate statistic is close to the critical value and the cointegration test is known to not be powerful, we proceed under the maintained hypothesis that cointegration exists. We use estimated z in estimating our ECM. This implies that our model captures the variables that have a significant influence on the long-run exchange rate over this period. Our next step is to formulate and estimate an error-correction model for the exchange rate.

Putting the estimated z in the finite-order auto regressive equation as we have mentioned before and following the idea of MacDonald and Taylor (1993), we lag each variable five times to capture the autocorrelation phenomena of the data.

Our error-correction model becomes:

$$\Delta s_t = \alpha + \sum_{i=0}^{n} \beta_i \Delta s_{t-i-1} + \sum_{i=0}^{n} \gamma_i \Delta m_{t-i} + \sum_{i=0}^{n} \delta_i \Delta m^*_{t-i} + \sum_{i=0}^{n} \theta_i \Delta y_{t-i}$$
$$+ \sum_{i=0}^{n} \kappa_i \Delta y^*_{t-i} + \sum_{i=0}^{n} \lambda_i \Delta i_{t-i} + \sum_{i=0}^{n} \mu_i \Delta i^*_{t-i} + \rho \hat{z}_{t-1} + u_t \tag{65}$$

where u_t, z_t denote disturbance terms and z_t also denotes an error correction term which is the estimated residual z_t of equation (62), and ρ is the speed of adjustment expected to be negative so that the path of the exchange rate is convergent. Thus a positive value of z_t implies that s_t is above its long run equilibrium level, and the exchange rate tends to depreciate in the next period.

Before estimating this model, first we check whether the variables in the model are $I(0)$. From Table 2.1, 2.2 we see that all the variables except z are $I(1)$ but their first difference is integrated of degree of zero, which satisfies the condition for the error correction model.

Here we regress our model and compare it with the other monetary models discussed before and the random walk model (RWM). The regression is run by setting n = 1 or 2, 3, 4, 5 respectively. In order to find the optimal value of the lag, Akaike Information Criteria (AIC) should be calculated. Akaike's AIC criterion is defined as follows:

$$AIC(n) = \ln \hat{\sigma}_n^2 + \frac{2n}{T} \tag{66}$$

where σ_n^2 is the ML estimator of the residual variance obtained from a model with lag length n, that is $\sigma_n^2 = SSE/T$. For this criteria the lag length estimate n is chosen so as to minimize the criterion used. AIC values for the lags from 1 to 5 in this model are 0.145E-3, 0.151E-3, 0.152E-3, 0.130E-3, 0.83E-4. We therefore conclude that $n = 5$ is the optimal lag number. We regress this model with five lag terms for each explanatory variable.

The estimated results show not all of the coefficients are statistically significant. In order to improve efficiency, eliminating insignificant coefficients is necessary. By successively omitting the variables with the lowest t-statistics and reestimating the model, it eventually reduces to the specification reported in Table 4.

In the next part, we subject the model to a variety diagnostic tests to determine its adequacy. The aim is to ensure that the inferences presented in Table 4 are valid, and to see whether the model is useful for forecasting exchange rates out-of-sample.

Our first step to ensure the validity of the inferences is to test for serial correlation in the equation's residuals. The Durbin-Watson value is not valid when lagged exchange rates are added as explanatory variables. So only Durbins's H, M tests, and a Bruesch-Godfrey (BG) test are carried out here. The test results are reported in Table 3.1. Both Durbins H and M tests show that there is no autocorrelation in the error term, since Durbins H, M statistic values are -0.4041, -0.2578 and BG value are 0.1603, 1.2348, 3.1942 and 3.9506 which are all less than their corresponding 5% critical value, and also the null hypotheses of no AR(1), AR(2), AR(3), AR(4) are rejected at 5% level by the BG tests.

Next, we consider whether the data show evidence of autoregressive heteroscedasticity. Since the foreign exchange rates are generally estimated using time series data where the residuals are quite small for a number of successive periods of time, then much larger for a while, then smaller again, and so on. Thus the error term maybe show the evidence of autoregressive heteroscedasticity. There has been a great deal of literature on ways to model this phenomenon. Engle (1982) first proposed the concept of autoregressive conditional heteroscedasticity (ARCH) which is a quite general phenomenon in exchange rate data. The basic idea of ARCH is that the variance of the error term at time t depends on the size of the squared error terms in previous time periods. Engle's (1983) ARCH model is designed to analyze the stochastic process that is serially autocorrelated with zero mean and conditional variance which is a function of past errors. The ARCH model can be stated as follows:

$$
\begin{gathered}
y_t|I_{t-1} \sim N(x_t\beta, h_t) \\
h_t = \alpha_0 + \alpha_1\epsilon_{t-1}^2 + \alpha_2\epsilon_{t-1}^2 + \ldots + \alpha_p\epsilon_{t-p}^2 \\
\epsilon_t = y_t - x_t\beta \\
\alpha_0 > 0, \alpha_i \geq 0, i = 1, 2, \ldots, p.
\end{gathered}
\tag{67}
$$

where $x_t\beta$ is a linear combination of lagged endogenous and exogenous variables included in the information set I_t ,β is a vector of unknown parameters, h_t can be generalized to include current and lagged x's to provide a wider class of possible parameterizations of heteroscedasticity.

There are several reasons for choosing this particular representation. Cumby and

Obstfeld (1984) and Hodrick and Srivastava (1984) provide evidence that the forecast error is heteroscedastic. The ARCH model is a convenient specification for incorporating heteroscedasticity into the estimation procedure.

Since ARCH implies that the variances are not homogeneous, we test for heteroscedasticity to confirm this. BP and LM tests are conducted here. (Table 3.2) We also use LM tests to try to identify the lag number in the ARCH model (Table 3.2). Our results are less than the respective critical value, so we cannot reject the null hypothesis of homogeneity. Therefore, the model is not an ARCH model, and the standard tests can be finished by simple OLS techniques.

On the basis of the above tests, we find no evidence that would cast serious doubt on the inferences presented in Table 4. Therefore, we now turn to consider other tests of the model's adequacy.

One criterion of a model's adequacy suggested by Hendry and Richard (1982) is parameter constancy over the sample period, since stability of the parameter estimates is a key factor in the model's ability to forecast accurately out-of-sample. As Hendry (1979) has shown, dynamic specification can be critical to the constancy of equations, so parameter inconstancy may suggest that the specified dynamic structure is inadequate. To test this, we apply CUSUM and CUSUMSQ tests via recursive residuals, which show no evidence of parameters instability.

We also examine the specification for possible evidence of non-linearities. We use the Ramsey (1969) Reset test. The Reset test requires adding powers of the predicted dependent variable to the original set of explanatory variables and determining whether the coefficients of these variables are significantly different from zero. The Reset test is F-distributed. Test results (Table 3.4) show evidence of non-linearity. Such non-linearity may due to relavent variables omitted which can lead to biased and inconsistent estimates.

Finally, we test for normality. For access to this literature we can see White and MacDonald (1980) and references listed in Judge, et al (1985). We will perform two tests. One is a GF test, the other is an LM test. Under the null hypothesis that the errors are normally distributed, our results are ambiguous for these two tests. The value of the GF test is 33.301 which exceeds the critical value $\chi^2(2) = 5.991$, while the LM value is 3.7711 which does not exceed the critical value (see Table 3.3). So we cannot draw any conclusion from these two tests. Non-normality is common in financial time series and given our large number of observations and the absence of serial correlation, this non-normality is unlikely to affect our inferences.

In sum, fairly rigorous diagnostic tests find no evidence that would suggest that our inferences in the preceding sections are suspect. Thus our post-sample forecasting based on this error-correction model is valid.

So far, we have investigated only the properties of the estimated monetary model. A stronger criterion to evaluate the usefulness of the model would be to determine how well they perform out of sample. In order to carry out this function, we reserve the last 15 observations from 1990.III to 1994.I for out-of-sample forecasting evaluations. The system is estimated beginning with the period 1990.III and one quarter ahead forecasts of the exchange rate are generated. The methods used here are the same as that when we forecast the structural regression models. Ultimately, 15 quarter ahead forecasts were obtained. The realized values of the estimated coefficients are used to generate all forecasts. The statistics used to gauge the out-of-sample properties of the models are the root mean square error (RMSE) and mean absolute error (MAE), and mean error (ME). These are 0.01564 (RMSE), 0.01156 (MAE), -0.0003 (ME) (Table 5.1). Compared with the conventional flexible price model, the real interest differential model, the portfolio balance model and the random walk model, our model outperforms all of them at a one quarter horizon (Table 5.1). As indicated by Table 5.2, the error-correction model also correctly predicted the direction of change for 11 out of 15 forecasting quarters, so is a more accurate predictor in direction.

However we notice that the speed of adjustment ρ in our model is -0.0016, and it is insignificant at the 5% significance level. This insignificant value provides ambiguous evidence for using this error-correction model. In order to prove that our error-correction model is still a good model despite this, we compare this model with a similar model without the error correction term z_t. The forecast results for both these models are reported in Tables 5.1 and 5.2. They show that our error-correction model beats the similar model without the error correction term z_t in the RMSE, MAE, ME and the direction of the forecast errors.

The results of here suggest that the monetary model of the exchange rate may still be useful in forecasting the exchange rate.

4. CONCLUSIONS

After a brief survey of the conventional monetary models of the exchange rate, we investigate the performance of these models by estimating them using data for the Japanese yen/US dollar exchange rate over the period from the first quarter of 1979 to the first quarter of 1994. We find these models fail to beat the random walk model. One of the reasons why these conventional models fail is that most relevant variables have unit roots. In this paper, an error correction model is constructed and estimated to improve forecasting performance.

Although there are problems in our error-correction model, it still has strong explanatory power and reliable forecasting capability. It outperforms the random walk model and other conventional structural models.

Due to time constraint, the exchange rate study in this paper is a preliminary work. If time permits, we will extend our study to other exchange rates.

Table 1. The OLS estimates of the Conventional Models

vars	mdl 1	mdl 2	mdl 3	mdl 4	mdl 5	mdl 6	mdl 7
S_{t-1}						0.003* 0.001	
$m-m^*$	0.33* 0.09	0.66* 0.13	0.26* 0.12	-0.08 0.11	0.23* 0.10	0.19* 0.11	0.30* 0.12
$m-m^*_{t-1}$							
$y-y^*$	-0.14* 0.07	-1.76* 0.12	-1.37* 0.11	-1.10* 0.09	-1.37* 0.10	-1.49* 0.12	-0.41* 0.11
$y-y^*_{t-1}$						0.71* 0.13	
$i-i^*$		0.01* 0.003	-0.003 0.003	0.0006 0.002	-0.001 0.003		0.001 0.003
$i-i^*_{t-1}$							
$r-r^*$	0.17* 0.22		0.19* 0.03	0.10* 0.03	0.14* 0.03		0.16* 0.03
$r-r^*_{t-1}$							
$TB-TB^*$				-0.01* 0.001			
$B-B^*$					0.07* 0.02		
$p-p^*_{t-1}$						0.06 0.12	
$q-q^*$							-0.19* 0.09
const	4.18* 0.04	3.95* 0.063	4.22* 0.065	4.40* 0.058	4.12* 0.064	4.40* 0.570	4.18* 0.067
R^2	0.989	0.982	0.989	0.993	0.991	0.985	0.990
D.W.	0.613	0.496	0.652	0.639	0.659	0.829	0.670
ρ	0.602	0.678	0.589	0.560	0.588	0.544	0.599

Note: 1. Value below the estimated coefficients are standard error.
2. ρ is the coefficient of AR(1) errors.
3. * denotes significant at the 5% level.

Table 2.1 Unit Root Tests for Variables of Japan (a)

$$\Delta y_t = \alpha_0 + \alpha_1 y_{t-1} + \sum_{j=1}^{p} \gamma_j \Delta y_{t-j} + \epsilon_t$$

Variable	z-test	τ-test	F-test Φ_1	I(d)
S	-0.0507	-0.0278	0.9291	I(1)
ΔS	-14.154	-3.6370	6.7709	I(0)
$m-m^*$	-2.0301	-0.7573	0.4335	I(1)
$\Delta(m-m^*)$	-------	-3.3162	5.7305	I(0)
$y-y^*$	-0.6286	-0.2877	0.4267	I(1)
$\Delta(y-y^*)$	-52.698	-13.117	87.545	I(0)
$i-i^*$	-------	-1.3034	0.6260	I(1)
$\Delta(i-i^*)$	-------	-4.2810	9.1670	I(0)
$r-r^*$	-6.4991	-1.7842	1.5919	I(1)
$\Delta(r-r^*)$	-------	-4.5761	10.472	I(0)
$TB-TB^*$	-2.9150	-1.4667	1.3262	I(2)
$\Delta(TB-TB^*)$	-------	-2.2828	2.6063	I(1)
$\Delta\Delta(TB-TB^*)$	-------	-5.7654	16.624	I(0)
$B-B^*$	-3.9025	-1.4136	0.9994	I(1)
$\Delta(B-B^*)$	-58.41	-18.651	176.28	I(0)
$P-P^*$	-1.9585	-0.7322	0.4281	I(1)
$\Delta(P-P^*)$	-16.632	-3.8652	7.7304	I(0)
$q-q^*$	-14.465	-2.8610	4.4753	I(1)
$\Delta(q-q^*)$	-------	-6.0882	18.551	I(0)

Table 2.2 Unit Root Tests for Variables of Japan (b)

$$\Delta y_t=\alpha_0+\alpha_1 y_{t-1}+\alpha_2 t+\sum_{j=1}^{p}\gamma_j \Delta y_{t-j}+e_t$$

Variable	z-test	τ-test	F-test Φ_1	F-test Φ_2	I(d)
S	-7.4989	-2.1791	2.7770	3.1474	I(1)
ΔS	-31.452	-9.4661	33.340	49.658	I(0)
m-m*	-8.7094	-2.4604	2.7654	3.9844	I(1)
Δ(m-m*)	-------	-3.3081	3.8955	5.6141	I(0)
y-y*	-8.1999	-2.4228	2.9686	4.0189	I(1)
Δ(y-y*)	-54.028	-12.960	59.153	87.199	I(0)
i-i*	-------	-3.4132	4.3162	6.3609	I(1)
Δ(i-i*)	-------	-4.2249	5.9917	8.9839	I(0)
r-r*	-6.1518	-1.7558	1.0955	1.6430	I(1)
Δ(r-r*)	-------	-4.6317	7.1760	10.763	I(0)
TB-TB*	-2.3617	-1.0658	0.98527	1.2303	I(2)
Δ(TB-TB*)	-------	-2.4861	2.0713	3.1063	I(1)
ΔΔ	-------	-5.6906	10.869	16.298	I(0)
B-B*	-2.8638	-1.0149	1.4013	2.1017	I(1)
Δ(B-B*)	57.782	-18.648	122.72	181.65	I(0)
P-P*	-8.3140	-2.3707	2.6222	3.7560	I(1)
Δ(P-P*)	-25.141	-5.8467	12.426	18.302	I(0)
q-q*	-14.523	-2.8396	2.9390	4.0328	I(1)
Δ(q-q*)	-------	-6.0265	12.147	18.203	I(0)

Table 2.3 Cointegration Tests

The test for no cointegration is given by two types of the cointegrating equations and the test for a unit root in the estimated residuals u_t :

$$y_{t1}=\beta_0+\sum_{j=2}^{m}\beta_j y_{tj}+u_t$$

$$y_{t1}=\beta_0+\beta_1 t+\sum_{j=2}^{m}\beta_j y_{tj}+u_t$$

$$\Delta\hat{u}_t=\alpha_0\hat{u}_{t-1}+\sum_{j=1}^{p}\gamma_j\Delta\hat{u}_{t-j}+v_t$$

equation	c/n z-test	c/n τ-test	c/t z-test	c/t τ-test	coint
1	-22.051	-4.3454	-27.560	-4.4409	not
2	-17.716	-3.8041	-------	-1.7596	not
3	-22.687	-4.3386	-29.297	-4.7776	not
4	-28.647	-5.4691	-------	-4.1853	not
5	-23.852	-4.3148	-------	-4.7237	not
6	-------	-2.6638	-------	-3.0021	not
7	-22.467	-4.1765	-29.296	-4.7773	not

Note: c/n denotes constant and no trend
c/t denotes constant and trend

Table 3.1 Test for Autocorrelation

H_0: no autocorrelation
H_1: autocorrelation

Test Stat	Results	D.F	5% Critical value	Reject/not
Durbins H	-0.4041	--	Z =-1.96	not reject
Durbins M	-0.2578	19	t_{19} =-2.09	not reject
BG				
--AR(1)	0.1603	1	$\chi^2(1)=3.84$	not reject
--AR(2)	1.2348	2	$\chi^2(2)=5.99$	not reject
--AR(3)	3.1942	3	$\chi^2(3)=7.82$	not reject
--AR(4)	3.9506	4	$\chi^2(4)=9.49$	not reject

Table 3.2 Test for Heteroscedasticity and ARCH effects

H_0: homoskedasticity
H_1: heteroscedasticity

Test Stat	value	D.F.	$\chi_{0.05}^2$	Reject or Not Reject
BP	7.6724	4	9.4877	not reject
White	12.7680	20	31.4104	not reject
LM;				
ARCH(1)	0.9001	1	3.8415	not reject
ARCH(2)	0.9032	2	5.9915	not reject
ARCH(3)	0.9033	3	6.2514	not reject
ARCH(4)	1.5410	4	7.7794	not reject

Table 3.3 Test for Normality

H_0: error term is normal distributed
H_1: error term is not normal distributed

Test Stat	Results	D.F.	$\chi_{0.05}^2(2)$	Reject or Not Reject
GF	33.301	2	5.991	reject
LM	3.7711	2	5.991	not reject

Table 3.4 Test for Non-linearity

H_0: Specification for Non-linearity
H_1: Specification for linearity

Test Stat	Results	D.F.	Reject or Not
RESET(2)	1.0012	1,19	not reject
RESET(3)	0.9407	2,18	not reject
RESET(4)	0.6269	3,17	not reject

Table 3.5 Results of CUSUM and CUSUMSQ tests

RECURSIVE RESIDUALS - SIGNIFICANCE LEVEL = 5%

OBS	REC-RES	CUSUM	BOUND	LOWER	CUSUMSQ	UPPER
27	-0.70196E-02	0.[illegible]2844	4.6635	-0.30277	0.08641	0.40277
28	-0.19650E-02	0.64403	5.0878	-0.25277	0.09319	0.45277
29	0.95219E-02	0.0[illegible]31	5.5115	-0.20277	0.25219	0.50277
30	0.18579E-02	0.4[illegible]28	5.9354	-0.15277	0.25824	0.55277
31	-0.29766E-02	0.10639	6.3594	-0.10277	0.27378	0.60277
32	-0.41696E-02	0.86935	6.783	-0.05277	0.30427	0.65277
33	0.27382E-02	0.36831	7.2073	-0.00277	0.31742	0.70277
34	-0.18110E-02	0.69970	7.6313	0.04723	0.32317	0.75277
35	-0.35554E-03	0.76476	8.0552	0.09723	0.32339	0.80277
36	0.77640E-03	0.62269	8.4792	0.14723	0.32445	0.85277
37	0.10403E-01	1.28095	8.9031	0.19723	0.51425	0.90277
38	-0.86942E-02	0.30994	9.3271	0.24723	0.64682	0.95277
39	-0.24199E-02	0.75275	9.7510	0.29723	0.65709	1.00277
40	-0.97834E-02	2.54294	10.1750	0.34723	0.82494	1.05277
41	0.23604E-02	2.11102	10.5990	0.39723	0.83471	1.10277
42	0.45646E-02	1.27578	11.0229	0.44723	0.87125	1.15277
43	0.66208E-02	0.06429	11.4469	0.49723	0.94813	1.20277
44	0.35850E-02	0.59171	11.8708	0.54723	0.97067	1.25277
45	0.40884E-02	1.33982	12.2948	0.59723	0.99998	1.30277
46	0.11005E-03	1.35995	12.7188	0.64723	1.00000	1.35277

HARVEY(1990,EQUATION 2.10 RECURSIVE T-TEST=0.3041 WITH 19 D.F.
HARVEY(1990,EQUATION 2.12 HETEROSKEDASTICITY TEST=0.5753 WITH M = 6

Table 4. Results of an Error-Correction Model

variable	coefficient	st. error	t-ratio
$S_{t-1}-S_{t-2}$	-0.0067	0.0052	-1.269
$S_{t-2}-S_{t-3}$	0.1897	0.1430	1.327
$S_{t-4}-S_{t-5}$	0.1942	0.1305	1.488
Δm_{t-2}	-0.0664	0.0663	-1.001
Δm^*_t	0.1499	0.0693	2.164*
Δm^*_{t-2}	-0.0559	0.0862	-0.648
Δy_t	-0.5651	0.0288	-19.61*
Δy_{t-2}	-0.1552	0.1085	1.431
Δy_{t-4}	-0.0762	0.0790	0.964
Δy^*_t	0.8091	0.0766	10.57*
Δy^*_{t-2}	-0.1214	0.1655	-0.734
Δy^*_{t-4}	-0.1536	0.1253	-1.226
Δi_t	0.0041	0.0024	1.727*
Δi_{t-3}	-0.0066	0.0024	-2.738
Δi_{t-4}	0.0053	0.0030	1.752*
Δi_{t-5}	-0.0044	0.0020	-2.227*
Δi^*_{t-1}	0.0014	0.0021	0.676
Δi^*_{t-2}	0.0015	0.0024	0.616
Δi^*_{t-3}	0.0032	0.0016	1.959*
Δi^*_{t-5}	0.0024	0.0012	1.999*
zj1	-0.0016	0.0114	-0.136
constant	-0.0047	0.0022	-2.100*

Note: * denotes significance at 5% level.

Table 5.1 Results of RMSE, MAE and ME measurement.

Model	Flex	Rea Int	Portfo	our mdl	our mdl no z_t	R.W.
RMSE	0.02697	0.03354	0.05653	0.01564	0.02435	0.02308
MAE	0.01847	0.02086	0.04560	0.01156	0.02019	0.01524
ME	0.00964	0.01051	0.02409	-0.0003	0.01015	0.00629

Table 5.2 Direction of forecast errors (forecast number:15)

Model	Flex	Rea Int	Portfo	our mdl	our mdl no z_t	R.W.
correct direct	9/15	9/15	9/15	11/15	8/15	7/15

Appendix: Data Sources

The data for this study related to the Japanese Yen-U.S. dollar exchange rate are taken from the International Monetary Funds's *International Financial Statistics* database and *Main Economic Indicators*

S_t : logarithm of the yen price of the US dollar, taken from *International Financial Statistics* (line ae), published by the International Monetary Fund.

$B\text{-}B^*$: logarithm of the ratio of the Japanese bond to the US bond, taken from *International Financial Statistics* (line 36ab). The US bond has been converted to yen using the corresponding exchange rate.

$m\text{-}m_*$: logarithm of the ratio of the Japanese money supply to the US money supply. Money supplies are seasonally adjusted M1 figures, taken from *International Financial Statistics* (line 34b)

$y\text{-}y^*$: logarithm of the ratio of the Japanese real income to the US real income. Seasonally adjusted figures for industrial production are taken as real income from *Main Economic Indicators*, published by the Organization for Economic Cooperation and Development.

i-i^*: short term interest rate differential, taken from *International Financial Statistics* (line 60b 60c).

r-r^*: long term interest rate differential, which are government bond yield rates, taken from *International Financial Statistics* (line 61).

p-p^*: expected inflation rate differential. The logarithm of consumer price indexes (CPI) are used as proxies for expected price levels, which are taken from *International Financial Statistics* (line 64).

q-q^*: logarithm of the relative price differential. Relative price is the price ratio of traded goods to non-traded goods. Consumer price indexes are used as proxies for the prices of nontraded goods while producer price indexes are for the prices of traded goods. Both two price indexes are taken from *International Financial Statistics* (line 64 63)

TB-TB^*: the trade balance differential. Accumulated current account are taken from the *Main Economic Indicators* in billion of yen for apan and in millions of dollar for the United States which has been converted to billion of yen using the corresponding exchange rate. Both are deflated by the respective consumer price indexes.

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