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# ABSTRACT

## On the Japanese Yen-US Dollar Exchange Rate: A Structural Econometric Model Based on Real Interest Differentials*

In this paper the short- and long-run movements of the Japanese yen-US dollar exchange rate are modelled for the recent floating period. The modern general-to-specific approach is used as our econometric framework. In contrast to some other exchange rate studies, we interpret multiple cointegrating vectors using economic theory. Among the findings are sensible and significant long-run relationships, and dynamic equations which describe the movements of the exchange rate and satisfy a battery of diagnostic tests. The models are shown to produce good in-sample forecasting performance and also out-of-sample forecasting performance which dominates a random walk.

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Keywords: cointegration, exchange rates, econometric modelling

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# NON-TECHNICAL SUMMARY

In this paper the medium- (short-) and long-run movements of the Japanese yen-US dollar exchange rate are modelled for the recent floating period. During this time, this exchange rate has drawn considerable attention from policy-makers and has been used as an instrument aimed at reducing US deficits after the Plaza meeting in 1985. Despite the intellectual consensus about the appropriate determinants of exchange rates over the long-run time horizon, however, most previous studies have failed to outperform a simple statistical model in the out-of-sample context. As a result, some researchers seem to have lost interest in so-called fundamental determinants, such as prices and interest rates, for understanding exchange rate movements. The main objective of this paper, however, is to construct an exchange rate model using fundamentals. The modelling approach relies on recent techniques and on a particular specification of the exchange rate relationship which has not been considered previously.

On the technical side, the modern general-to-specific approach is used as the econometric framework. This method copes with several problems which may have produced poor empirical results in previous studies. These potential problems include simultaneity bias, *a priori* restrictions on parameters, the static nature of the specification, and inappropriate treatment of the time-series properties of data. Although the classical version of this approach is frequently used in exchange rate studies, few attempts have been made to use its modern version which exploits multivariate cointegration methods. In addition, we propose one method for identifying the unique cointegrating relationships. This is an important step in constructing a model which makes both economic and statistical sense.

The theoretical model attempts to explain exchange rate movements using fundamental determinants and as a consequence our model is better suited to explaining the medium- to long-run behaviour of this exchange rate. The choice of our specification partly reflects the increasing body of literature which supports such determinants having explanatory power over exchange rate movements at longer-term time horizons. The model is derived from the real interest differential model which was originally developed by Frankel (1979). The model differs from the original specification in that it abstracts from money market equilibrium conditions, however, thereby avoiding any uncertainty regarding the money demand function. The simplification of the original model also has the advantage of assisting in the identification of the number of cointegrating vectors. Since the modern general-to-specific approach is based

on the system, it is essential to find out which long-run relationships correspond to which equations.

Among the findings are sensible and significant long-run relationships, and dynamic equations which describe the motion of the two exchange rates and satisfy a battery of diagnostic tests. The models are shown to produce good in-sample forecasting performances and also an out-of-sample forecasting performance which dominates a simple random walk. Indeed, the model is able to beat the random walk in all forecasting time horizons. The study shows that this is attributable to three factors: an appropriate long-run specification; abstraction of money markets in our specification; and finally, the introduction of dynamics in the short-run (medium-run) model. This strong evidence that the yen exchange rate can be explained by the fundamental determinants offers encouragement to policy-makers who often have to assess how far the yen is from underlying fundamentals. This paper therefore reasserts the importance of fundamental variables in understanding short-run exchange rate movements. Current exchange rate research, often named 'market microstructure', points to the importance of factors such as the bid-ask spread in understanding the short-run exchange rate movements, and often regards fundamentals as irrelevant in this time horizon. Although consideration of non-fundamentals would seem interesting and plausible, as far as we are aware this line of study is still some way from reaching a consensus and is not strongly supported by empirical evidence. We think that the introduction of fundamentals may well give some further scope for the micro-structure model.

## 1. Introduction

In this paper we use a simplified version of the real interest differential (RID) model, first proposed in Frankel (1979), to model the Japanese yen exchange rate against the US dollar, over the period 1975, quarter 3 to 1994, quarter 3. The motivation for our study arises from the recently-noted success for variants of the monetary model. Thus, for example, MacDonald and Taylor (1992,1993) have shown that the monetary model produces sensible long-run equilibrium relationships and outperforms a random walk in out-of-sample forecasting exercises. One issue relating to such studies is that there are often multiple cointegrating vectors and it is hard to interpret all of them in the context of the monetary model. Using a simplified version of the RID, we offer a way of placing some structure on the cointegrating relationships. Our modelling strategy involves using the so-called structural econometric modelling of Hendry and Mizon (1993). This approach is especially useful in the current application since it can handle in a natural way problems associated with previous estimates of monetary models, such as a lack of model dynamics, *a priori* restrictions on parameters and simultaneous equation bias.[1] The modelling approach has the further advantage that it contains a rigorous set of criteria for assessing a models in-sample as well as out-of sample fit. As an additional way of assessing our exchange rate models validity, we use what has become the acid test of an exchange rate model, namely the Meese and Rogoff (1983) out-of-sample root mean squared error (RMSE) forecasting criterion.

The outline of the remainder of this paper is as follows. Section 2 briefly explains the modern general-to-specific modelling methodology and, in Section 3, we derive the simplified RID model. A distinguishing feature of our version of the RID model is that, in contrast to the original version of Frankel (1979), it does not rely upon money demand functions for its derivation. Section 4 describes our data set and the time-series properties of the individual series. In Section 5 we propose a way of identifying unique cointegrating vectors in a multivariate cointegration test, and then go on to discuss the modern general-to-specific approach before presenting our empirical results. A state of the art test, comparing our models' performance with that of the AR(1) model, in an out-sample context, is conducted in section 6. Finally, in section 7 the overall performance of our models is summarised.

## 2. Modern General-to-specific Econometric Modelling Approach

The concept of a general-to-specific (GS) approach may be traced back to Sargan (1964), but Mizon (1976), Hendry and Mizon (1993) and Hendry and Richard (1982, 1983) and Hendry (1995) are probably the most relevant studies here. We have chosen this strategy because it incorporates recent developments in the time-series literature in a natural way (in particular, the multivariate cointegration methods of Johansen (1988) and offers an econometric modelling strategy combining data consistency with economic intuition.

More specifically, the GS approach consists of two steps. In the first step the long-run behaviour of the time-series in the system are examined using the multivariate cointegration methods of Johansen (1988). Therefore, an unrestricted VAR which is congruent with the data must be constructed. Assuming the relevant variables are I(1) then the following unrestricted VAR which has a p-dimensional vector autoregressive process, of order k, may be expressed as:

$$z_t = \sum_{i=1}^{k} A_i z_{t-i} + \mu + \Psi D_t + \varepsilon_t \qquad [1]$$

where z is a vector of stochastic endogenous variables, $\mu$ is the constant term, and D is a dummy vector. This dummy vector may include one point and/or seasonal dummies as well as stationary variables. The error, $\varepsilon$, is a Gaussian error vector (i.e. $\varepsilon \sim NIID(0, \Sigma)$). Johansen shows that equation [1] can be transformed into the vector error correction representation (VECM).

$$\Delta z_t = \Gamma_1 \Delta z_{t-1} + \cdots + \Gamma_{k-1} \Delta z_{t-k+1} + \Pi z_{t-1} + \mu + \Psi D_t + \varepsilon \qquad [2]$$

where $\Pi = -I + \Pi_1 + \ldots + \Pi_i$ and $\Gamma$ here is defined as $\Gamma = I - \Gamma_1 -, \ldots, \Gamma_k$. For cointegration to exist amongst the variables, $\Pi$ must be of reduced rank r<p, where the rank, r, determines the number of linearly independent stationary relationships between the levels of variables (and thus it implies that there are p-r nonstationary relationships in the model). The decomposition of $\Pi$ into the loadings ($\alpha$) and cointegrating parameters ($\beta$) must be carried out such that $\beta z$ is stationary, where $\beta z$ has the interpretation of an error correction mechanism (ECM).

Often in a cointegration study no structure is placed on the cointegrating vectors. However, in the presence of multiple cointegrating vectors it is not clear what interpretation may be placed on them. There is therefore an increasing trend towards imposing some restrictions on the cointegrating parameters which are consonant with economic theory (see, for example, Johansen and Juselius (1992)). Equation [2] with such restrictions imposed is called the constrained VAR (CVAR). Exclusion tests are then conducted on the dynamic components of the CVAR and a parsimonious VAR (PVAR) is derived. The PVAR is the system to which our final model will be compared. The successful final model must contain as much explanatory power as the PVAR, which can be tested using an encompassing test. In addition, the long-run weak exogeneity of variables can be tested by imposing restrictions on the loadings. Therefore, one of the long-standing criticisms against the reduced form equation on which most previous exchange rate studies rely can be statistically tested.

The second stage of the modelling process involves a successful reduction and re-parameterisation of the PVAR to the structural econometric model (SEM). The SEM can be evaluated by several criteria (Hendry and Richard 1983), some of which are already discussed in the process of obtaining the structural econometric model from the unrestricted VAR. Hendry and Richard discuss six criteria, the first three relating to the information sets available. The first condition is that the errors are innovation process and therefore white noise. This means that the current errors cannot be predicted using their own past history. The second condition is related to the present information set and is that of weak exogeneity. This is an essential condition for a model to be simplified without any loss of information. The third condition relates to future data and requires that the parameters of interest be time-invariant. This condition addresses the Lucas critic and is therefore connected with the concept of super-exogeneity. The fourth condition is that the SEM must be consistent with economic theory since if this is not the case then the model is unlikely to be useful from an economic perspective.

The fifth condition asserts the necessity for data accuracy and data admissibility. The former is concerned with consistency in the choice of aggregated economic data with the information in the economic mechanism. The aggregated data used in the study must be consistent with the data generating process (DGP) of the true model. The model must also be consistent with the data constraints. Finally, the theory of reduction implies that all SEMs constructed from the sample DGP (or the parsimonious model) are comparable. In a single equation context, rival models can be produced using a different choice of endogenous variables but different models can be constructed in the system context, also by a different choice of conditioning variables. However, all successful SEMs must be nested in the parsimonious model although they need not nest each other. An encompassing test is used to determine if the SEM has the same informational content as the PVAR.

## 3. Theoretical Model Description

The version of the RID model utilised in this paper posits, as in the original RID model, that the long-run exchange rate is determined by PPP. However, this concept only pertains in the very long-run, a period much longer than that captured by the data sample used in this paper (for example, recent work has shown that PPP holds reasonably well when using approximately one hundred years of annual data - see MacDonald (1995)). Our preferred long-run, or equilibrium, relationship is one which allows deviations from PPP to be governed by a real interest differential. This seems appealing since the latter can capture the underlying real variables which keep an exchange rate away from its PPP determined level, and also the effect of sticky prices (see MacDonald and Marsh (1997) for a further discussion). In contrast to the original RID model (see Frankel (1979)), however, our simplified version does not feature money supplies or income levels. This may have at least two advantages in econometric modelling. First, it allows removal of uncertainty concerning the stability of the demand for money function and, second, for practical purposes, it is likely that a more meaningful model can be constructed because unique cointegrating vectors are more easily identifiable. Clearly one disadvantage of such a simplification is that the effect of money supplies on the exchange rates cannot be examined directly, although recent studies (Clarida and Gali 1994, and Eichenbaum and Evans, 1995) tend to indicate that the link between money and exchange rates is rather weak.

More specifically, there are three assumptions underpinning our model. The first is that in the 'long-run' PPP holds:

$$\bar{s}_t = p_t - p_t^* \qquad [3]$$

where $\bar{s}_t$ denotes the logarithm of the long-run nominal exchange rates, $p_t$ is the logarithm of the price index and an asterisk refers to a foreign variable. Our assumption of long-run PPP is much less restrictive than the treatment of PPP in other types of monetary models (e.g., the flexible price monetary model) in which PPP must hold continuously. For a sample period such as the recent float, an exchange rate is unlikely to be continuously at its PPP-determined level. In recognition of this we allow the current 'equilibrium' rate to diverge from its PPP level by a real interest rate differential. The rationale for this may be seen in the following way. Assume that the expected change in currency depreciation is a function of the difference in the current and long-run exchange rates, and of expected inflation differentials ($E_t\Delta p_{t+1}$ - $E_t\Delta p^*_{t+1}$ ):

$$E_t s_{t+1} - s_t = -\theta(s_t - \bar{s}_t) + E_t\Delta p_{t+1} - E_t\Delta p^*_{t+1} \qquad [4]$$

where $E_t(\cdot_{t+1})$ is the expected value of $(\cdot)$ at time t+1 based on the available information set at time t. This equation says that, *ceteris paribus,* investors will anticipate exchange rates at a lower level in the future (at t +1) when the current exchange rate is higher than its long-run level. The same effect on expected future exchange rates can be obtained when the expected inflation differential between countries declines. The coefficient $\theta$ is an approximate measure of the adjustment speed: a relatively low value of $\theta$ indicates that the current deviation of an exchange rate from PPP is relatively persistent, while an infinite value of $\theta$ means that the current exchange rate is always at the PPP level. Nominal interest rates are assumed to be linked across countries by the UIP condition:

$$E_t s_{t+1} - s_t = i_t - i_t^* \qquad [5]$$

where $i$ and $i*$ are the domestic and foreign nominal short-term interest rates, respectively. On using equations [3]-[5], we can derive the following simplified version of the RID model:

$$s_t = p_t - p_t^* - \theta^{-1}\left[\left(i_t - E_t \Delta p_{t+1}\right) - \left(i_t^* - E_t \Delta p_{t+1}^*\right)\right] \qquad [6]$$

which simply summarises the fact that an exchange rate will be above or below its PPP determined level by a proportion due to the real interest differential. The signs in equation [6] are consistent with theoretical predictions, and prices satisfy the homogeneity and symmetrical parameter restrictions.

Equation [6] is the relationship we use to define our cointegrating relationship. It is clear from the above discussion that this is not a 'true' equilibrium relationship since it contains an adjustment component pertaining to the real interest differential. However, the rationale for using [6] as our long-run relationship is that in a period such as the recent float it is evident that real interest differentials have not been zero, or equal to a constant, on average throughout the period (the massive observed net capital flows are in themselves a testament to this). Ignoring the importance of interest differentials in our long-run relationships would produce a misspecified relationship (and this is something we discuss further below). If we had a longer time span of data (say 100 years) we would expect [3] to hold more closely since it is to be expected that real interest differentials, and the underlying real shocks which they capture, would average out over such a long time span (and this indeed is borne out by the empirical evidence - see MacDonald (1995))

## 4. Data Description

All data are from the *OECD Main Economic Indicators*. We employ Japanese quarterly seasonally adjusted data over the period 1975:Q3-1994:Q3, with the eight final observations kept for the purpose of out-of-sample forecasts. Our quarterly data are transformed from a monthly basis by averaging all the observations within a period. Wholesale price indices (WPI), which are seasonally unadjusted, are used as our price measures. The interest rate variables are short-term yields, with a 3 month maturity, and all variables except interest rates are expressed in logarithmic form. In constructing real interest rates we subtract a measure of the expected inflation rate from the nominal interest rate[2]. The expected inflation rates are calculated using an autoregressive distributed lag (ADL) model.[3] Initially, eight lags are taken to ensure that the model does not suffer from serial correlation, and then insignificant lagged variables are removed one by one. This process is continued until the residuals do not violate the non-autocorrelation condition. This method implies that investors form expectations about future prices using all present and past information on prices and any discrepancy between the actual and expected value of prices is attributed solely to 'news' factors which arise after the formation of expectations. The following specifications, seemed appropriate for our purposes:

| | | |
|---|---|---|
| Japan 1973:Q3-1994:Q4 | $p_t = 1.75p_{t-1} - 0.75p_{t-2} + 0.01p_{t-3} - 0.04p_{t-5} + 0.02p_{t-6}$ | |
| | $R^2 = 0.99$ | AR 1-5 (5,73) = 1.25 [0.30] |
| US 1974:Q1-1994:Q4 | $p_t = 1.55p_{t-1} - 0.75p_{t-2} + 0.54p_{t-3} - 0.32p_{t-4} - 0.02p_{t-5} + 0.03p_{t-6} - 0.03p_{t-7} + 0.01p_{t-8}$ | |
| | $R^2 = 0.99$ | AR 1-5 (5, 71) = 0.88 [0.50] |

The AR1-5 statistic is a test of autocorrelation based on the Lagrange Multiplier test, expressed in the F-form (Harvey 1981).

Since the econometric methods outlined in Section 2 rely on the variables entering the $z$ vector being I(1), we have conducted a standard set of Augmented-Dickey Fuller (ADF) unit root tests. These are summarised in Table 1. Using the critical values of MacKinnon (1991), our ADF results show that most variables seem to be integrated of degree one (i.e., I(1)) since the null hypothesis can be rejected when they are differenced, but not when they are in levels. The Japanese real interest rate appears to be I(0). Henceforth, we shall regard all variables as I(1), except the Japanese real rate which is treated as an exogenous variable in our study.

## 5. Application of Modern General-to-Specific Approach

### 5.1. Long-run Study

In our empirical implementation of the RID model, our initial focus is on the long-run relationship, this being the first stage of the general-to-specific approach. Our model contains four endogenous variables (an exchange rate, domestic and US prices and the US real interest rate); to be consistent with our unit root test results the Japanese real interest rate enters the system as a deterministic element. Centred seasonal dummies are included in the system.

To determine the other deterministic components of our models, and identify the unique cointegrating vectors, we propose the following method.[4] Our method utilises the technique proposed by Hendry and von Ungern-Sternberg (1981) [HUS] in the context of a single equation consumption function study. Their method calculates the static steady-state condition using parameters obtained from an autoregressive distributed lag (ADL) version of the model. It has, at least, two useful purposes in the system modelling context. First of all it provides useful information regarding the value of cointegrating parameters, in terms of their magnitude and consistency with *a priori* signs. Secondly, the HUS method allows us to examine, given the specification of the model, whether the equilibrium relationships require constant and time trend terms. These are two vital pieces of information in attempting to identify the unique cointegrating vectors. In summary form, our method involves the following three steps[5]:

> *Step 1:* Use the Johansen test to examine the number of significant cointegrating vectors in the system. In defining the cointegrating vectors there are three potential deterministic specifications. The first contains a constant restricted to the cointegrating relationship (model 2); the second has an unconstrained constant (model 3); while the third has a time trend in the cointegrating vector (model 4).
>
> *Step 2*: Estimate the cointegrating parameters, corresponding to all of the endogenous variables, or variables which economic theory predicts will converge on the steady state (e.g., exchange rates in our study), using an ADL. The choice of Models 2, 3 or 4 is determined by the final specification of the static steady state conditions calculated by the ADL.
>
> *Step 3*: Impose the estimated values of parameters from the ADL on the cointegrating vectors in as many cases as possible. Here it is important that at least the signs of these imposed parameters be consistent with economic theory. The acceptance of this joint linear restriction can be tested using a likelihood ratio test.

Step 1 is a conventional one. In step 2, the signs of the adjustment coefficients can also provide some information on which variable is forcing the equation to the equilibrium path, and thus equilibrium. In order to ensure that the expected steady state specification is indeed correct, a unit root test can also be implemented. As mentioned, the estimates from the ADL suffer much less from small sample bias than those from the static model such as the Engle-Granger procedure (Banerjee *et al* 1993), and furthermore, whether variables are statistically significant or not can be tested and insignificant values dropped from the specification. Ideally, the imposition of all estimated values on cointegrating parameters is accepted by the likelihood-ratio test. But since these estimates are sensitive to the existence of the deterministic variables such as one-point dummies and simultaneous biases, in practice it is often difficult to impose all estimates from the ADL on the cointegrating vectors in the Johansen test. Success in finding the unique cointegrating vectors (i.e., ECMs) will enable us to construct the parsimonious VAR which incorporates the ECMs.

Using the notation of Hansen and Juselius (1995), when the lag length is two in the pre-differenced model (i.e. the unrestricted VAR), then expressing the deterministic terms explicitly, we can rewrite equation [2] as:

$$\Delta z_t = \Gamma_t \Delta z_{t-1} + \alpha \begin{pmatrix} \beta \\ \mu_1 \\ \delta_1 \end{pmatrix}' \tilde{z}_{t-1} + \alpha_{\perp}\mu_2 + \alpha_{\perp}\delta_2 t + \Psi D_t + \varepsilon_t \qquad [8]$$

where $\tilde{z}_{t-1}$ is the error correction vector as a result of a successful reduction of ranks. The Greek letter, $\beta$, consists of parameters corresponding to the long-run specification, and $\mu$ and $\delta$ correspond to the intercepts and time trends, respectively. The subscripts on the deterministic terms indicate if they relate to the long-run equilibrium relationship or the short-run model. The subscript 1, for instance, denotes intercepts or linear trends in the cointegration space. The existence of the deterministic components in the system is determined in our procedure by the estimates of the ADLs. Other notation is the same as in equation [2]. For the model to be balanced, the combination of $\beta\tilde{z}_{t-1}$ must be stationary.

At this stage it is not important to interpret each parameter since the unrestricted VAR is purely a statistical model, so it cannot be expected that its parameters have an economically meaningful interpretation. However, whether or not the model is capable of representing the DGP is crucial for the subsequent modelling process and this can be judged by the diagnostic tests of the residuals which we would expect to be white noise. Table 2 summaries the diagnostic tests of the residuals from the model and on the basis of these, we can conclude that the model satisfies the condition of residual whiteness. The US price equation is the only one which exhibits some residual non-normality (a significant ARCH effect), otherwise the rest satisfy the necessary conditions. Figure 1 shows the in-sample fit of the model: the actual and fitted values of differenced endogenous time-series, standardised residuals, histogram of standardised residuals with normal distribution, and correlogram of residuals.

The remainder of this section follows the above-noted three steps in order to derive the parsimonious VAR. First, the multivariate cointegration test of Johansen is used to examine the number of significant cointegration vectors in the cointegrating space (r). Our models constrain the constant term to lie in the cointegrating vectors (this is based on our ADL estimates, discussed below). The 90, 95 and 99 percent critical values are employed, and the rejection of the null is indicated by asterisk marks in table 3. Tests are based on the trace statistic (Trace) and the maximum eigenvalue statistic ($\lambda$-max) (Johansen 1988) and critical values are from Osterwald-Lenum (1992). The null hypothesis for the existence of at most r cointegrating vectors and the alternative for the existence of greater than r can be determined using the trace statistic (Trace) which is calculated using the maximum likelihood function as follows:

$$Trace = -2\ln(Q) = -T\sum_{i=r+1}^{p}\ln\left(1-\hat{\lambda}_i\right)$$

where r = 0, ..., p-1, T the number of observations, $\hat{\lambda}$ are eigenvalues and Q is the ratio of restricted maximum likelihood to the unrestricted one. Another statistic is the maximum eigenvalue statistic (λ-max) and this is calculated as:

$$\lambda - \max = -T\ln\left(1-\hat{\lambda}_{r+1}\right)$$

Note that the test hypothesis of the maximum eigenvalue statistic is different from that of the trace statistic, and its null hypothesis is that there are r cointegrating vectors and the alternative is that r + 1 cointegrating vectors exist.

Since the distributions of the above statistics are sensitive to the sample size, we also calculate the corresponding small sample corrected values of Reimers (1992) (reported as $Trace^S$ and $\lambda\text{-max}^S$).[6] We interpret these tests as suggesting that there are two cointegrating vectors and we therefore present the static long-run equation solved from the ADL which corresponds to the exchange rate and domestic price equations.

Japan

Exchange Rate: $s = 2.787p - 1.727p^* + 0.610(i^* - E\Delta p^*)$

Sample: 1976:Q1-1994:Q3, Diagnostic Tests of the ADL: AR F(5, 56) = 1.052 [0.397]
ARCH F(4, 53) = 1.476 [0.223] Normality $Chi^2$ (2) = 5.526 [0.063] $Xi^2$ F(28, 32) = 1.628 [0.092] RESET F(1, 60) = 0.005 [ 0.944]

Price: $p = 0.836 + 0.267s + 0.519p^* + 0.959(i^* - E\Delta p^*)$

Sample: 1976:Q2-1994:Q3, Diagnostic Tests of the ADL: AR F(5, 45) = 1.078 [0.386]
ARCH F(4, 42) = 2.112 [0.096] Normality $Chi^2$ (2) = 4.286 [0.117] $Xi^2$ F(46, 3) = 0.223 [0.992]
RESET F(1, 49) = 0.469 [ 0.497]

Both equations pass the diagnostic tests and the signs of parameters in the exchange rate equations are all correct. In addition to residual tests used in table 2, the regression specification test (RESET) developed by Ramsey (1969), which examines the null hypothesis of the correct specification, is included in this table. The PPP condition holds: although all price coefficients exceed unity, their signs are all correct and the value of coefficients are not too far from their expected values. In addition, the real interest rate terms are correctly signed. As for the price equations, domestic and US prices are positively related, as we would expect from PPP, and the positive sign on the exchange rate is also consistent with PPP. We also implemented a unit root test on the residual series for each equation and both series proved to be stationary.

The ADL estimates provide us with some information about how to identify the two significant cointegrating vectors in an economically meaningful way. Further they also indicate that the constant term must enter in the cointegrating vector for the price equation. The non-existence of the constant and time trend term in the exchange rate equation is consistent with economic theory and empirical results from the Engle-Granger 2-step procedure.

The next step is related to the theory of reduction (e.g. Hendry 1995) which is also related to the issue of the identification of unique cointegrating vectors. Here, one of the recognised reduction methods, namely the exclusion of variables from the cointegrating vectors, is used. The restricted cointegrating vectors are reported in table 4[7] and it is worth noting that these restrictions satisfy the generic identification condition suggested by Johansen and Juselius (1994). For both models, the first cointegrating vectors are normalised in such a way that they correspond to the domestic price equations, and the second vectors to the exchange rate equations. These joint restrictions on the cointegrating vectors are supported by the data using a likelihood-ratio test ($Chi^2$ (1) = 0.04 [0.85]). The shaded values in table 4 are based on estimates from the ADL, and this table shows that it is difficult to impose all of the estimates for the ADL on the multivariate model. Nonetheless, most variables enter the relationships with the correct signs and their values are very close to these of the ADL (in instances where a variable appears in both relationships). Figure 2 show that equilibrium conditions for these two cases. The upper graphs in these figures present $\beta'\tilde{z}$ in equation [8], and the lower graphs $\beta' R_k$. Although the former figures show the actual disequilibirums, the latter ones are more relevant since the short-run distortion is removed in this graph (Hansen and Juselius 1995). All $\beta' R_k$ graphs show that these linear combinations are stationary.

A preliminary feel for whether our structural econometric models are likely to be successful is obtained from the signs of the loadings. All loading signs corresponding to the domestic price and exchange rate equations are negative, indicating the correct specification of our steady state conditions. Finally, in order to confirm the legitimacy of constructing a system, as opposed to a reduced form, long-run weak exogeneity tests are conducted for these two models. The results from the joint restrictions on both cointegrating vectors and loadings show that the model must be constructed as a system since the relevant $Chi^2$ (7) statistic is 38.54, with a probability value of 0.00. This suggests that previous studies which rely upon a reduced form equation are likely to suffer from problems of exogeneity bias.

### 5.2. Medium-run Study

5.2.1. Parsimonious VAR

Using the long-run equilibrium conditions derived in the last section, the next step in our estimation strategy involves estimating the PVAR. The parsimonious representation of the system is expressed as equations [2] and [8], and therefore, it consists of differenced current and lagged endogenous variables and dummies as well as error correction mechanisms (ECMs). Our PVARs are based on two-lagged endogenous variables, which is consistent with our unrestricted VAR, two ECM terms, which are calculated from the equilibrium conditions analysed in the long-run study, and lagged Japanese real interest rates which are assumed to be exogenous variables. The ECMs are defined as:

$$ECM1_t = -0.267\Delta s_t + \Delta p_t - 0.519\Delta p^*_t - 1.231\Delta (i^* - E\Delta p^*)_t + ECM1_{t-1}$$

$$ECM2_t = \Delta s_t - 2.999\Delta p_t + 1.916\Delta p^*_t + 1.820\Delta(i^* - E\Delta p^*)_t + ECM2_{t-1}$$

A PVAR is somewhere between a statistical model and an economic model in the sense that some long-run structure, consistent with economic theory is imposed, but it is still a VAR relationship. In that sense, the whiteness of the residuals of the PVAR should not be violated. Table 5 contains the details of the PVAR and the corresponding diagnostic tests. The statistics indicate a good fit and only the US price equation exhibits evidence of ARCH effects. We therefore regard both PVAR models as adequate representations of the DGPs.

5.2.2. Structural Econometric Model

In constructing our SEM systems we have sequentially deleted all statistically insignificant variables from the PVAR system, and our final models are shown in table 6. The fitted and actual values of the differenced exchange rates both in-sample and out-of-sample are shown in figure 3. These models are estimated by FIML and satisfy both order and rank conditions for identification.

5.2.3. Evaluation of Models

In order to evaluate our models we use the criteria suggested by Hendry and Richard (1983) and noted in section 2. These are: data coherency; consistency with theory; data admissibility; parameter constancy; encompassing; and weakly exogenous repressors. Since the exogeneity status of variables has already been checked, we do not discuss this criterion further here.

• Data coherency

The condition of data coherency requires the residuals from our models to be white noise processes, so that their past history will not improve forecastability of the variables. A summary of our residual diagnostic tests is presented in table 6. Using the same criteria applied previously, neither any individual equation nor the system as a whole suffers from autocorrelation, non-normality or residual heterogeneity. Most notably, the system now does not exhibit any ARCH effects. Therefore, we regard these structural econometric models as data coherent.

• data admissibility/theory consistency

Although the consistency of our long-run parameters with economic theory has already been discussed, we need to verify the consistency of the coefficient signs from our estimated structural econometric models. Since some variables remain in the model to ensure residual whiteness we only discuss those parameters which enter the models with a 5 per cent significance level.

The overall performance of our model is quite impressive since the vast majority of coefficients enter with the correct a priori sign. More specifically, in the exchange rate equations the majority of inflation and interest rate variables enter with the correct sign. For example, in the Japanese exchange rate equation, the effect of Japanese inflation on the change in the exchange rate is positive, which is consistent with relative PPP, although US inflation exhibits an ambiguous effect on the exchange rate with one lag positive and one lag negative. In addition, the lagged Japanese interest rate is, in sum, negatively correlated with the change in the exchange rate. The lagged difference in the exchange rate enters the exchange rate equations significantly. Finally, the corresponding ECMs are significantly negative (using a 10 percent significance level) a feature necessary for model stability.

Reasonably successful results are also obtained for the inflation and differenced interest rate equations. For the inflation equations, there appears to be important persistence of 'own' inflation in the two equations. Additionally, US inflation enters the Japanese inflation equation significantly, but Japanese inflation does not enter the US inflation equation significantly. Relative PPP, which links the exchange rate change to inflation, does not appear to hold in the Japanese inflation equations since the Japanese rate of inflation is a negatively related to the exchange rate, suggesting that an appreciation of the Japanese yen does not have an impact in attenuating inflation. Judging from the plot of the Japanese WPI, this trend was strong in the 1970s. Although an increase in WPI appears to slow down when the yen appreciation increased dramatically since the middle of the 1980s, the overall effect of the exchange rate on domestic inflation is very small. This may be due to a price stickiness. In the real interest equation, we note an ambiguous relationship for the US interest rate. Finally, the ECM1 terms in the inflation equations are significantly negative (at the 1 percent level) and the ECM2 terms, which corresponds to the disequilibrium conditions of the exchange rate, also enter significantly into the inflation equations. However, the latter violates the weak exogeneity condition although as Hendry, Neale and Ericsson (1990) demonstrate, using a Monte Carlo study, this only imparts a small bias.

• parameter constancy

Another criterion for a successful model is that the parameters be invariant over time and is seen as essential if the model is to be used for forecasting or for policy simulation. In order to test parameter constancy we have calculated out-of-sample forecasts using the fixed coefficients from our final models[8] and figure 4 gives a graphical portrayal of these results for the period 1992:Q4-1994:Q3 period. We note that these out-of-sample forecasts are within the 95 percent confidence level. Statistically, parameter constancy is tested using a $Chi^2$ statistic and the results confirm our figures - F(32, 59) = 0.463 [0.99]).

• encompassing

Finally, we have tested whether the final models encompass parsimonious VARs using a likelihood-ratio test. This guarantees that no information is lost in the process of reduction and that the structural econometric models have as much explanatory power as the parsimonious VARs. The test results for the model is $Chi^2$ (25) = 14.459 [0.95]. This high acceptance ratio may be attributed to the fact that the reduction process ceased just before any violation of residual whiteness occurred and therefore the residuals of our structural econometric models are as white as those of the parsimonious VARs.

## 6. Can Our Structural Econometric Models Beat AR Models?

The poor performance of the asset class of models in an out-of-sample context is highlighted in the seminal paper of Meese and Rogoff (1983) in which they show that a random walk model outperforms these asset class models in terms of a 1-step-ahead forecast over a one-to-twelve month period for the US dollar/DM and US dollar/yen. Although comparing one's models with an AR(1) has become a rather fashionable way to finish a papers, it is important to bear in mind that it is not an absolute criterion to compare the performance of econometric models.

First of all, the use of a 1-step-ahead forecast, as opposed to the multi-step-ahead forecast, is criticised by Schinasi and Swamy (1987) on the grounds that the 1-step-ahead forecast favours the random walk model since a random walk model using the lagged dependent variables has access to more information than the monetary models which have no lagged dependent variables in Meese and Rogoff's specification. More importantly, Ericsson (1992) argues that the minimum mean square forecast error is not a sufficient condition for the model to satisfy parameter constancy or to encompass a rival model. As a result, he concludes that although the model may well satisfy the condition of both parameter constancy and minimum RMSE, it may not represent an appropriate econometric forecasting model. Therefore, strictly speaking, the 1-step ahead forecast alone cannot be a criterion to assess the forecastability of the system as in our structural econometric models. However, since the forecast accuracy of the model not only checks the performance of the model in the out-of-sample context, but also examines the models' performance in the in-sample context (Schinasi and Swamy 1987) and since lagged endogenous variables are allowed in our models, we shall use it in order to evaluate our models as one of the general criteria.

An out-of-sample forecasting test is conducted using the same criterion which Meese and Rogoff used. For consistency, the parameters are re-estimated each time in calculating the forecast values, but the full information maximum likelihood (FIML) method is used in our study. There are three commonly used measures to compare forecasts of the model: the mean error (ME), mean absolute error (MAE) and root mean square error (RMSE). Meese and Rogoff (1983) use all three of them, while mainly utilising the last criterion. Although use of ME and MAE enables researchers to find out whether the model consistently under- or over-predicts, the RMSE alone is used in our analysis. The RMSE is calculated using the following formula:

$$\mathrm{RMSE} = \sqrt{\frac{\sum_{i=1}^{n}\left(s^{f}_{T+i} - s_{T+i}\right)^{2}}{N}}$$

where $s^f$ and s denote the forecast and actual values of and exchange rate, and N is the number of ex-post forecasts conducted. In our case, N = 1, ..., 8, (1992:Q3-1994:Q3). Table 7 presents Theil's U statistics which provide an instant comparison of the RMSE between a random walk and the model of interest. A value of Theil's U statistic below unity indicates that our model outperforms the random walk model while a statistic above unity indicates that our model is inferior to the random walk counterpart. Only the forecasting performance of our exchange rate equations are considered in here; therefore, the price and interest rate equations are ignored. remainder of this section, and the exchange rate equations are treated here as if they are single equations. The significance of this exercise is that our model uses error correction terms which are explicitly identified as those from the exchange rate equations. Table 7 summarises the performance of our model. This table shows some supportive empirical evidence favourable to our exchange rate model in the out-of-sample forecasting context during the sample 1992:3-1994:3. In particular, the ratio is well below unity over all 8 time horizons, suggesting that our model convincingly outperforms the simple random walk model. Furthermore, this ratio falls as the forecasting time horizon increases, reflecting the fact that our model is a more appropriate one in the medium- to long-run.

To summarise, our version of the RID model convincingly beats the random walk paradigm in an out-of-sample forecasting context and this finding is very different to the perceived wisdom on exchange rate modelling (see Frankel and Rose (1995)). This distinctive outcome can be attributed to (at least) three factors. First, our model does not suffer from the uncertainty relating to the money demand function which often plagues other empirical exchange rate studies. Second, we were able to successfully identify the steady-state of the exchange rate (i.e., an error correction term) which forces the exchange rate back onto its equilibrium path and, third, the short-run model contained complex dynamic interactions between the exchange rate change and fundamentals.

## 7. Summary and Conclusion

In this paper, we have used a simplified version of the RID model to explain the Japanese yen-US dollar bilateral exchange rate over the recent floating period. The modern general-to-specific approach is employed and empirical results show that our models, particularly the exchange rate equations, prove to be particularly successful, judged by standard criteria.

We firstly examined the long-run equilibrium relationships using multivariate cointegration methods and find evidence of two cointegrating vectors. In order to identify the unique cointegrating vectors which also make economic sense, estimates of the cointegrating parameters from an autoregressive distributed lag (ADL) are used as a guide. Some restrictions are partly imposed to satisfy the generic ranking condition suggested by Johansen and Juselius (1994). The acceptance of these linear restrictions on the cointegrating vectors is tested using a likelihood-ratio test. As a result, our long-run cointegrating parameters used in the subsequent modelling are very close to those from an ADL.

In the short-run study, criteria suggested by Hendry and Richard (1983) are used to evaluate our models. These criteria include data coherency, theory consistency, data admissibility, parameter constancy, encompassing and weakly exogenous repressors. Our models satisfy all such conditions, and it should be underlined that all parameters in these two exchange rate equations are consistent with theory. Particularly, it is worthwhile mentioning that the Japanese exchange rate is a negative function of its real interest rates and all ECMs enter with significantly negative coefficients, which implies that the corresponding equations are converging to their long-run path.

In conclusion, we believe that the evidence reported in this paper clearly provides a justification for using fundamental determinants to understand the medium- to long-run movements of exchange rates. It would seem that exchange rate models which eschew fundamentals have rather thrown the baby out with the bathwater!

[1] See MacDonald and Taylor (1992) for a concise summary of potential reasons for the poor performance of the monetary class exchange rate models.

[2] Our model of constructing the real interest rate differs from Frankel (1979), who used long bond yields for proxy expected inflation.

[3] There are many ways to calculate the proxy for the expected inflation rates. Here we follow the method often used in demand for money studies (e.g., Muscatelli 1989).

[4] Alternatively, Johansen (1992) who advocates the use of the Pantula principle, by which the joint hypothesis of the rank order and deterministic components can be tested jointly.

[5] As a result, our method is similar to Bagliano *et al* (1991) in a way that parameter restrictions, which are calculated outside the model, are imposed for the identification purpose. However, our method is more flexible in that we rely more heavily upon a statistical properties of each equation.

[6] However, Doornik and Hendry (1994) argue that it is inconclusive if this type of correction is preferred.

[7] All restrictions on cointegrating vectors and loadings are conducted using the econometric packages, CATS in RATES developed by Hansen and Juselius (1995).

[8] Therefore, the method used here differs from the rolling estimation method used in Meese and Rogoff (1983).

Table 1 Unit Root Test by ADF

| | Level | | First Difference | |
|---|---|---|---|---|
| | Con | Con + Trend | Con | Con + Trend |
| $s^Y$ | -0.193 | -2.719 | -6.434 ** | -6.475 ** |
| $p^Y$ | -2.192 | -2.047 | -3.438 ** | -3.482 ** |
| $p^{US}$ | -2.688 | -1.476 | -4.923 ** | -5.849 ** |
| $(i - E\Delta p)^Y$ | -3.232 * | -3.744 * | -4.294 ** | -4.349 ** |
| $(i - E\Delta p)^{US}$ | -0.933 | -1.484 | -9.799 ** | -9.803 ** |

Note: One asterisk (*) indicates that statistics are significant at the 5 percent level, and two asterisks (**) at the 1 percent level. Seasonal dummies are included in the calculation for price indices and productivity differentials. Critical values in Osterwald-Lenum (1992) are used here. The sample covers 1973:Q3-1994:Q3 for testing the unit root of exchange rates and prices in levels, and it covers 1974:Q4-1994:Q3 for exchange rates and prices in differences. The sample period for the real interest rates in levels is 1975:Q1-1994:Q3, and that for these in differences is 1975:Q2-1994:Q3.

Table 2. Diagnostic Tests of Unrestricted VAR, 1975:Q3-1992:Q3

| | s | p | p* | $i - E\Delta p$ | $i^*-E\Delta p^*$ |
|---|---|---|---|---|---|
| Japanese RID | | | | | |
| AR 1-5 F(5, 44) | 0.904 [0.487] | 1.289 [0.286] | 0.665 [0.652] | -- | 1.452 [0.225] |
| Normality $Chi^2$ (2) | 1.701 [0.427] | 0.081 [0.960] | 2.382 [0.304] | -- | 2.460 [0.292] |
| ARCH F(4, 41) | 0.872 [0.489] | 0.086 [0.986] | 3.533 [0.014] * | -- | 1.224 [0.315] |
| $Xi^2$ F(24, 24) | 0.672 [0.832] | 0.519 [0.943] | 1.113 [0.398] | -- | 1.085 [0.422] |
| Vector AR 1-5 (80, 104) | 1.114 [0.301] | | | | |
| Vector Normality $Chi^2$ (10) | 11.37 [0.182] | | | | |
| Vector F(240, 163) | 0.486 [1.000] | | | | |

Note: see table 1. The Lagrange-Multiplier (LM) test is used for detecting serial correlation and the normality test is based upon Doornik and Hansen (1994). The residual is also tested for ARCH effects and heteroscedasticity ($Xi^2$). A detailed description of these tests is made in Engle (1982) and White (1980) respectively.

Table 3. Johansen Tests Applied to the Japanese Simplified RID Model, 1975:Q4-1992:Q3

| | λ-max | λ-max $^S$ | Trace | Trace $^S$ |
|---|---|---|---|---|
| $H_0$: rank = r | | | | |
| r == 0 | 33.71 ** | 27.76 $^{\otimes}$ | 77.60 ** | 63.91 ** |
| r <= 1 | 23.67 * | 19.49 $^{\otimes}$ | 44.89 ** | 36.15 * |
| r <= 2 | 16.45 * | 13.55 | 20.23 * | 16.66 |
| r <= 3 | 3.778 | 3.112 | 3.778 | 3.112 |

Note: The marks, $^{\otimes}$, * and **, attached to these statistics indicate that these values are statistically significant at the 10, 5 and 1 percent level.

Table 4. Restricted Normalised Cointegrating Vectors and Loadings

| Japanese RID | | | | | | |
|---|---|---|---|---|---|---|
| Cointegrating | -0.267 | 1.000 | -0.519 | -- | -1.231 | -0.863 |
| Vectors | 1.000 | -2.999 | 1.916 | -- | 1.820 | 0.000 |
| Loadings | -1.751 | -0.310 | | | | |
| | -0.496 | -0.072 | | | | |
| | -0.303 | -0.138 | | | | |
| | 0.448 | 0.208 | | | | |

Table 5. Japanese Parsimonious VAR

| Explanatory Variables | Endogenous Variables | | | | |
|---|---|---|---|---|---|
| | $\Delta s_t$ | $\Delta p_t$ | $\Delta p^*_t$ | $\Delta(i-E\Delta p)_t$ | $\Delta(i^*-E\Delta p^*)_t$ |
| $\Delta s_{t-1}$ | 0.176 | -0.001 | 0.067 | | -0.051 |
| | (0.205) | (0.037) | (0.041) | | (0.062) |
| $\Delta s_{t-2}$ | -0.306 | -0.104 | -0.003 | | -0.110 |
| | (0.203) | (0.037) | (0.041) | | (0.061) |
| $\Delta p_{t-1}$ | 0.660 | 0.663 | -0.241 | | -0.320 |
| | (1.067) | (0.194) | (0.215) | | (0.323) |
| $\Delta p_{t-2}$ | 1.185 | 0.548 | 0.330 | | 0.372 |
| | (1.050) | (0.190) | (0.211) | | (0.317) |
| $\Delta p^*_{t-1}$ | 1.547 | 0.927 | 0.639 | | 1.378 |
| | (0.979) | (0.177) | (0.196) | | (0.295) |
| $\Delta p^*_{t-2}$ | -1.129 | -0.678 | -0.306 | | 0.017 |
| | (1.353) | (0.245) | (0.272) | | (0.408) |
| $\Delta(i^*-E\Delta p^*)_{t-1}$ | -0.455 | -0.022 | 0.058 | | -0.361 |
| | (0.695) | (0.126) | (0.139) | | (0.209) |
| $\Delta(i^*-E\Delta p^*)_{t-2}$ | -0.045 | 0.091 | 0.070 | | 0.015 |
| | (0.420) | (0.076) | (0.084) | | (0.127) |
| ECM1 | -1.822 | -0.475 | -0.250 | | 0.470 |
| | (0.680) | (0.123) | (0.137) | | (0.205) |
| ECM2 | -0.331 | -0.066 | -0.121 | | 0.216 |
| | (0.205) | (0.037) | (0.041) | | (0.062) |
| Constant | -0.008 | 0.003 | 0.008 | | 0.003 |
| | (0.036) | (0.007) | (0.007) | | (0.011) |
| $(i-E\Delta p)_t$ | -3.007 | -0.661 | -0.095 | | 0.041 |
| | (1.346) | (0.243) | (0.270) | | (0.406) |
| $(i-E\Delta p)_{t-1}$ | 2.412 | 0.198 | 0.161 | | -0.888 |
| | (1.781) | (0.322) | (0.357) | | (0.537) |
| $(i-E\Delta p)_{t-2}$ | -0.429 | 0.142 | -0.203 | | 0.741 |
| | (1.096) | (0.198) | (0.220) | | (0.331) |
| Diagnostic Tests | s | p | $p^*$ | $i-E\Delta p$ | $i^*-E\Delta p^*$ |
| AR (5, 46) | 0.510 | 1.339 | 0.611 | | 0.947 |
| | [0.768] | [0.265] | [0.692] | | [0.460] |
| Normality $Chi^2$ (2) | 1.083 | 0.135 | 2.663 | | 1.694 |
| | [0.582] | [0.935] | [0.264] | | [0.429] |
| ARCH (4, 43) | 1.532 | 0.115 | 3.862 | | 1.654 |
| | [0.210] | [0.976] | [0.009] ** | | [0.178] |
| $Xi^2$ (29, 21) | 0.536 | 1.052 | 1.010 | | 1.102 |
| | [0.940] | [0.459] | [0.499] | | [0.415] |
| Vector AR (80, 112) | 1.253 | | | | |
| | [0.135] | | | | |
| Vector Normality $Chi^2$ (8) | 10.282 | | | | |
| | [0.246] | | | | |
| Vector $Xi^2$ (290, 140) | 0.734 | | | | |
| | [0.985] | | | | |

Note: figures in ( ) are standard errors, and these in [ ] are t-probability.

Table 6. Japanese Structural Econometric Model

$\Delta s_t = 0.325\Delta s_{t-1}$ - $0.277\Delta s_{t-2}$ + $1.524\Delta p_{t-2}$ + $1.894\Delta p^*_{t-1}$ - $1.673\Delta p^*_{t-2}$ - $2.227(i\text{-}E\Delta p)_t$
[0.005] [0.125] [0.034] [0.010] [0.061] [0.003]

+ $1.277\Delta(i\text{-}E\Delta p)_{t-1}$ -$1.510ECM1_{t-1}$ - $0.296ECM2_{t-1}$
[0.065] [0.006] [0.087]

AR (5, 47) = 1.200 [0.324] Normality $Chi^2$ (2) = 0.888 [0.641] ARCH (4, 44) = 0.803 [0.530]
$Xi^2$ (28, 23) = 0.576 [0.918]

$\Delta p_t$ = - $0.098\Delta s_{t-2}$ + $0.472\Delta p_{t-1}$ + $0.635\Delta p_{t-2}$ + $1.019\Delta p^*_{t-1}$ - $0.692\Delta p^*_{t-2}$ - $0.493(i\text{-}E\Delta p)_t$
[0.001] [0.000] [0.000] [0.000] [0.000] [0.000]

+ $0.217(i\text{-}E\Delta p)_{t-2}$ + $0.064\Delta(i^*\text{-}E\Delta p^*)_{t-2}$ - $0.481ECM1_{t-1}$ - $0.067ECM2_{t-2}$
[0.008] [0.034] [0.000] [0.035]

AR (5, 47) = 2.405 [0.051] Normality $Chi^2$ (2) = 0.243 [0.885] ARCH (4, 44) = 0.818 [0.947]
$Xi^2$ (28, 23) = 1.043 [0.463]

$\Delta p^*_t$ = 0.006 + $0.053\Delta s_{t-1}$ - $0.426\Delta p_{t-1}$ + $0.442\Delta p_{t-2}$ + $0.691\Delta p^*_{t-1}$ - $0.191\Delta p^*_{t-2}$ - $0.143(i\text{-}E\Delta p)_{t-2}$
[0.080] [0.016] [0.000] [0.000] [0.000] [0.132] [0.010]

- $0.289ECM1_{t-1}$ - $0.121ECM2_{t-1}$
[0.003] [0.000]

AR (5, 47) = 1.499 [0.208] Normality $Chi^2$ (2) = 2.843 [0.243] ARCH (4, 44) = 2.105 [0.096]
$Xi^2$ (28, 23) = 1.179 [0.346]

$\Delta(i^*\text{-}E\Delta p^*)_t$ = - $0.099\Delta s_{t-2}$ + $1.125\Delta p^*_{t-1}$ -$0.381\Delta(i^*\text{-}E\Delta p^*)_{t-1}$ - $0.465(i\text{-}E\Delta p)_{t-1}$ + $0.4669(i\text{-}E\Delta p)_{t-2}$
[0.001] [0.000] [0.000] [0.012] [0.012]

+ $0.486ECM1_{t-1}$ + $0.182ECM2_{t-1}$
[0.000] [0.000]

AR (5, 47) = 1.572 [0.186] Normality $Chi^2$ (2) = 0.181 [0.913] ARCH (4, 44) = 0.409 [0.801]
$Xi^2$ (28, 23) = 1.435 [0.190]

Diagnostic Tests for the Structural Model

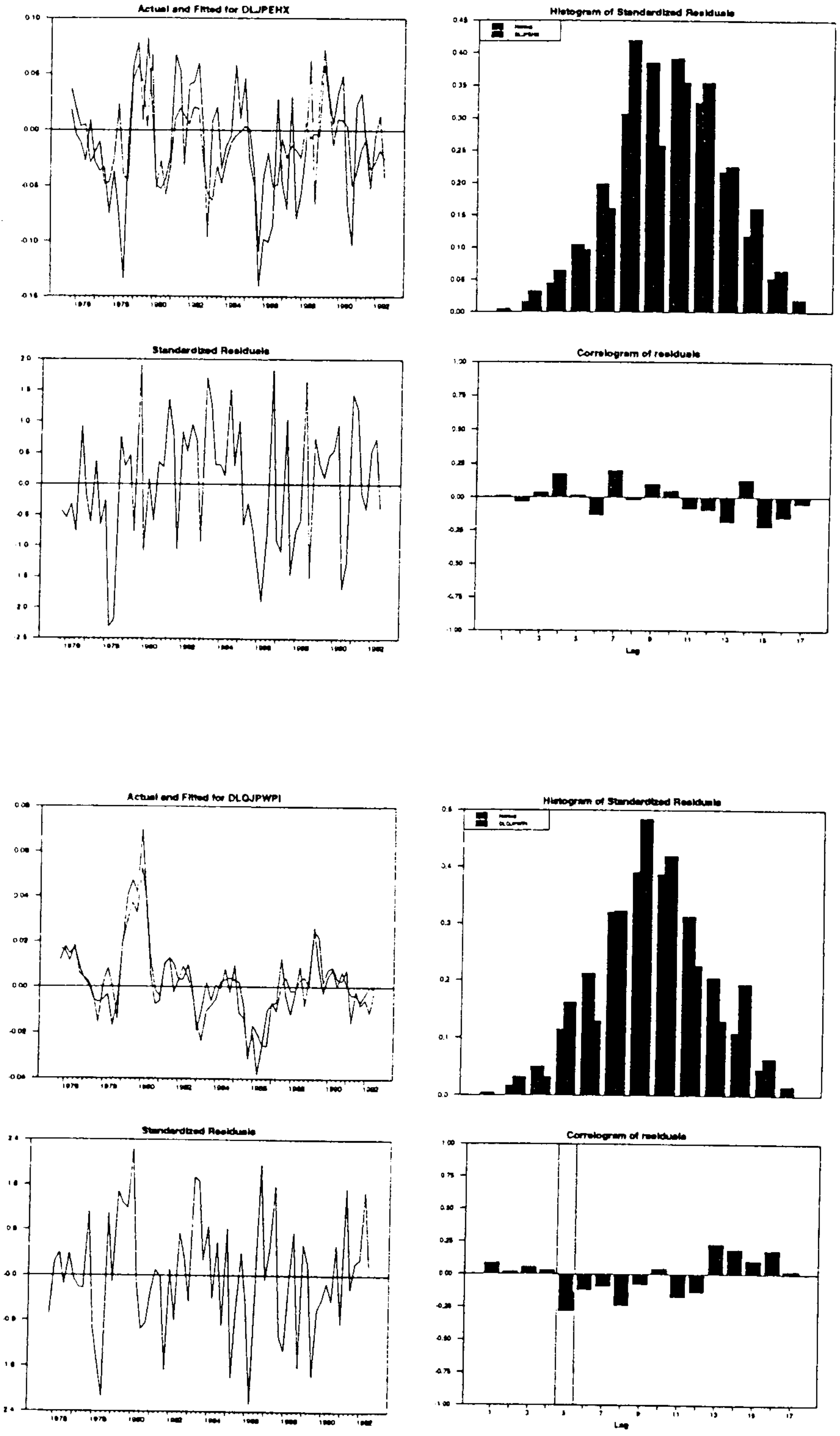
Vector AR (80, 144) = 1.296 [0.089]
Vector Normality $Chi^2$ (8) = 8.120 [0.414]
Vector $Xi^2$ F(280, 224) = 1.229 [0.054]

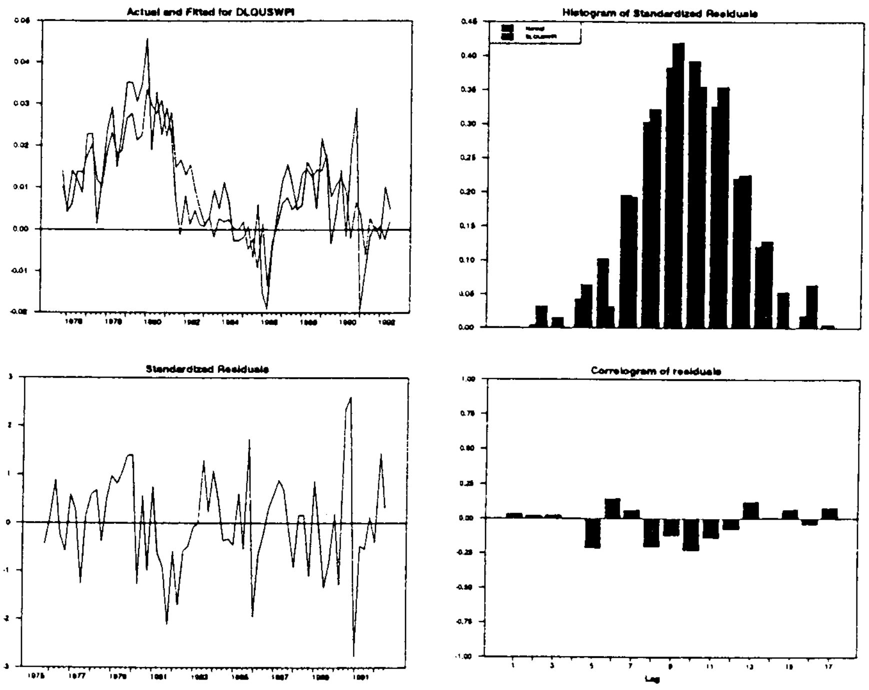
Note: figures in [ ] are t probability.

Table 7. Theil's U Statistics

| | |
|---|---|
| 1992:Q4 | 0.572 |
| 1993:Q1 | 0.491 |
| 1993:Q2 | 0.379 |
| 1993:Q3 | 0.291 |
| 1993:Q4 | 0.263 |
| 1994:Q1 | 0.175 |
| 1994:Q2 | 0.133 |
| 1994:Q3 | 0.149 |

Figure 1. Actual and Fitted Values and Residuals




Actual and Fitted for DLQUSWPI
Histogram of Standardized Residuals
Standardized Residuals
Correlogram of residuals
Lag

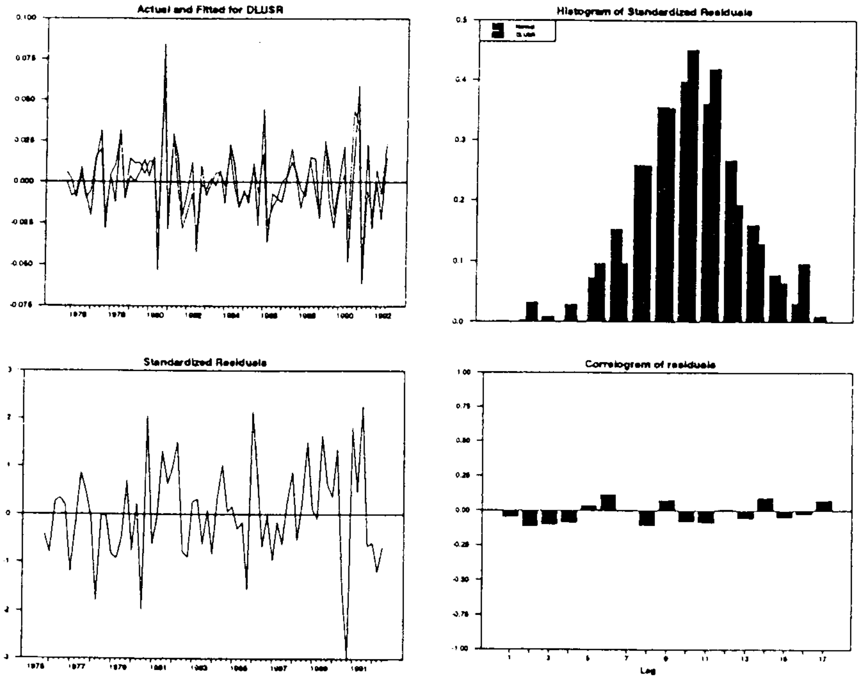

Actual and Fitted for DLUSR
Histogram of Standardized Residuals
Standardized Residuals
Correlogram of residuals
Lag

Figure 2. Japanese Disequilibrium Conditions

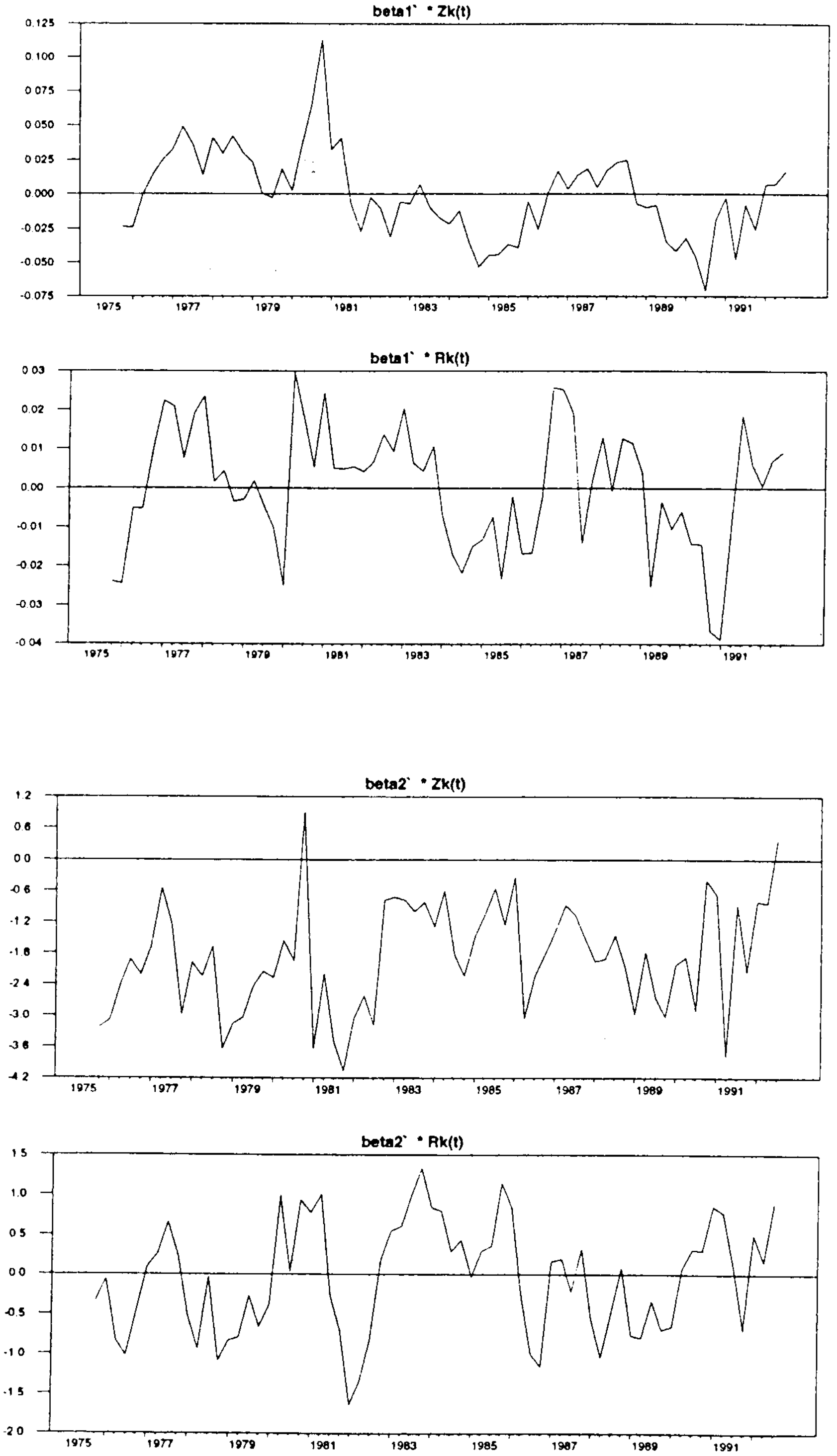


Figure 3. Actual and Fitted Differenced Time-Series of Japan

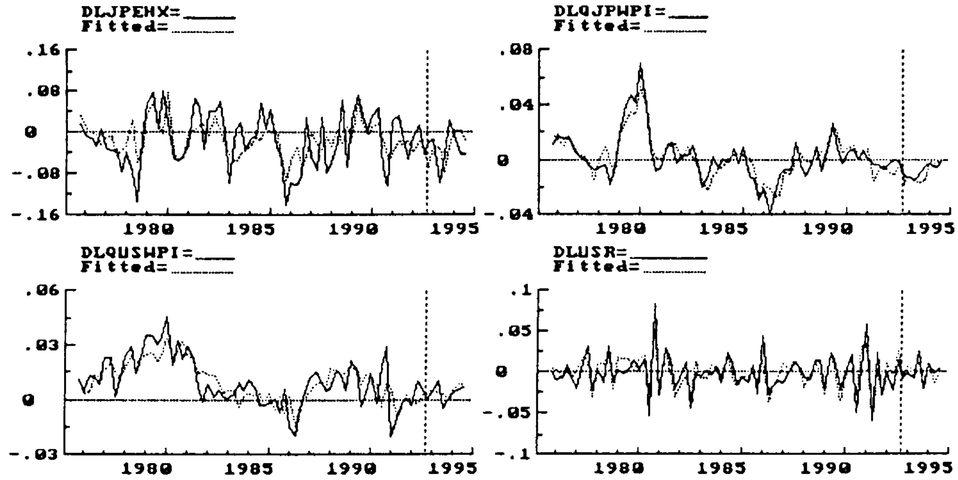
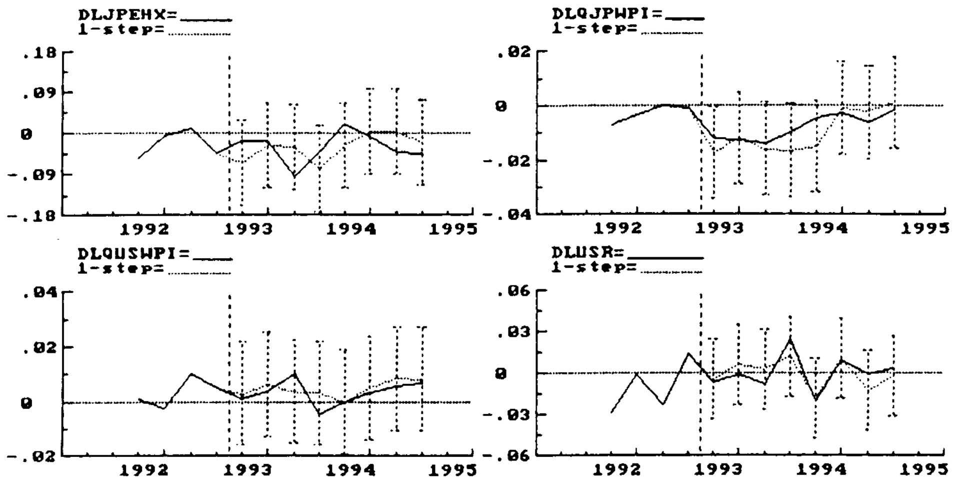


Figure 4. 1-step Forecast of Each Equation of Japan, 1992:Q3-1994:Q3



## REFERENCES


Banerjee, A., J. Dolado, J. W. Galbraith and D. F. Hendry, (1993), *Co-integration, Error Correction and the Econometric Analysis of Non-stationary Data,* Oxford: Oxford University Press.

Bagliano, F. C., C. A. Favero and V. A. Muscatelli, (1991), 'Cointegration and Simultaneous Models: An Application to the Italian Money Demand', *Working Paper,* No. 9108, University of Glasgow.

Clarida, R. and J. Gali, (1994), 'Sources of Real Exchange Rate Fluctuations: How Important Are Nominal Shocks?', *Carnegie-Rochester Conference Series on Public Policy,* 41, 1-56.

Doornik, J. A. and H. Hansen , (1994), 'A practical Test of Multivariate Normality', *mimeo,* University of Oxford.

Doornik, J. A. and D. F. Hendry, (1994), *PcFiml 8.0: Interactive Econometric Modelling of Dynamic System,* London: International Thomson Publishing.

Eichenbaum, M. and C. L. Evans, (1995), 'Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rates', *Quarterly Journal of Economics,* 110, 957-1009.

Engle, R. F., (1982), 'Autoregressive Conditional heteroscedasticity, with Estimates of the Variance of United Kingdom Inflations', *Econometrica,* 50, 987-1007.

Ericsson, N. R., (1992), 'Parameter Constancy, Mean Square Errors and Measuring Forecast Performance: An Exposition, Extensions and Illustration', *Journal of Policy Modelling,* 14, 4, 465-495.

Frankel, J. A., (1979), 'On the Mark: A Theory of Floating Exchange Rates Based on Real Interest Differentials', *American Economic Review,* 69, 4, 610-621.

Frankel, J. and A. Rose (1995), 'A Survey of Empirical Research on Nominal Exchange Rates', in Grossman and Rogoff (eds.) *The Handbook of International Economics*, Vol. 3 North Hollan.

Hansen, H. and K. Juselius, (1995), 'Cats in Rats: Cointegration Analysis of Time Series', Evanston: Estima.

Harvey, A. C., (1981), *The Econometric Analysis of Time Series,* Deddington: Philip Allen.

Hendry, D. F., (1995), *Dynamic Econometrics,* Oxford: Oxford University Press.

Hendry, D. F. and J-F. Richard, (1982), 'On the Formation of Empirical Models in Dynamic Econometrics', *Journal of Econometrics,* 20, 3-33.

Hendry, David F. and J-F. Richard, (1983), 'The Econometric Analysis of Economic Time Series', *International Statistical Review,* 51, 111-163.

Hendry, D. F. and G. E. Mizon, (1993), 'Evaluating Dynamic Econometric Models by Encompassing the VAR', in P. C. P. Phillps (ed.), *Models, Methods and Applications of Econometrics,* Oxford: Basil Blackwell.

Hendry, D. F., A. J. Neale , N. R. Ericsson, (1990), *PC-NAIVE: An Interactive Program for Monte Carlo Experimentation in Econometrics.* Institute of Economics and Statistics, University of Oxford.

Hendry, D. F. and T. von Ungern-Sternberg, (1981), 'Liquidity and Inflation Effects on Consumers' Expenditure', in A. S. Deaton (ed.) *Essays in the Theory and Measurement of Consumers' Expenditure,* Cambridge: Cambridge University Press.

Johansen, S., (1988), 'Statistical Analysis of Cointegration Vectors', *Journal of Economic Dynamics and Control,* 12, 231-254.

Johansen, S., (1992), 'Determination of Cointegration Rank in the Presence of a Linear Trend', *Oxford Bulletin of Economics and Statistics,* 54, 383-397.

Johansen, S. and K. Juselius, (1992), 'Testing Structural Hypotheses in a Multivariate Cointegration Analysis of the PPP and the UIP for UK', *Journal of Econometrics,* 53, 211-244.

Johansen, S. and K. Juselius, (1994), 'Identification of the Long-run and the Short-run Structure: An Application to the ISLM Model', *Journal of Econometrics,* 63, 7-36.

MacDonald, R., (1995), 'Long-Run Exchange Rate Modelling', *IMF Staff Papers,* 42, 437-489.

MacDonald, R. and I. W. Marsh (1997), 'On Fundamentals and Exchange Rates: A Casselian Approach', forthcoming, *Review of Economics and Statistics.*

MacDonald, R. and M. P. Taylor, (1992), 'Exchange Rate Economics: A Survey', *IMF Staff Papers,* 39, 1-57.

MacDonald, R. and M. P. Taylor, (1993), 'The Monetary Approach to the Exchange Rae: Rational Expectations, Long-run Equilibrium and Forecasting', *IMF Staff Papers,* 40, 89-107.

MacKinnon, J. G., (1991), 'Critical Values for Cointegration Tests', in R. F. Engle and C. W. Granger (eds.), *Long-run Economic Relationships,* Oxford: Oxford University Press.

Meese, R. and K. Rogoff (1983), 'Empirical Exchange Rate Models of the Seventies: Do They Fit out of Sample?', *Journal of International Economics,* 14, 3-24.

Mizon, G. E., (1976), 'Model Selection Procedures', in M. J. Artis and A. R. Nobay (eds.), *Studies in Modern Economic Analysis: The Proceedings of the Association of University Teachers of Economics,* Oxford: Basil Blackwell.

Muscatelli, V. A., (1989), 'A Comparison of the 'Rational Expectations' and 'General-to-specific' Approaches to Modelling the Demand for M1', *Oxford Bulletin of Economics and Statistics*, 51, 4, 353-375.

Osterwald-Lenum, M., (1992), 'A Note with Quantiles of the Asymptotic Distribution of the Maximum Likelihood Cointegration Rank Test Statistics', *Oxford Bulletin of Economics and Statistics*, 54, 3, 461-472.

Ramsey, J. B., (1969), 'Tests of Specification Errors in Classical Linear Least Squares Regression Analysis', *Journal of the Royal Statistical Society B*, 31, 350-371.

Reimers, H. E., (1992), 'Comparisons of Tests for Multivariate Cointegration', *Statistical Papers*, 33, 335-359.

Sargan, J D (1964) 'Wages and Prices in the United Kingdom: A Study in Econometric Methodology', in P. E. Hart, G. Mills and J. K. Whitaker (eds.), *Econometric Analysis for National Economic Planning*, London: Butterworths, 25-63.

Schinasi, G. J. And P. A. V. B. Swamy, (1987), 'The Out-of-sample Forecasting Performance of Exchange Rate Models When Coefficients are Allowed to Change', *International Finance Discussion papers*, No. 301, Board of Governors of Federal Reserve System.

White, H., (1980), 'A heteroscedastic-consistent Covariance Matrix Estimator and a Direct Test for Heteroscedasticity', *Econometrica*, 48, 817-838.