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**What Determines Real Exchange Rates?
The Long and Short of It**

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Abstract

This paper presents a reduced-form model of the real exchange rate. Using multilateral cointegration methods, the model is implemented for the real effective exchange rates of the dollar, the mark, and the yen, over the period 1974-1993. In contrast to much other research using real exchange rates, there is evidence of significant and sensible long-run relationships for a simplified version as well as for the full version of the model. The estimated long-run relationships are used to produce dynamic equations, which outperform a random walk and produce sensible dynamic patterns in the context of an impulse response analysis.

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SUMMARY

Recently, there has been a revival of interest in modeling the long-run behavior of nominal bilateral exchange rates using "fundamentals" such as relative prices. In general, this approach has established that, for the recent floating period, although some form of purchasing power parity (PPP) seems to hold on a single-currency basis for most countries, it does not conform to a strict interpretation of PPP. This is because the standard homogeneity restrictions often do not hold and also because adjustment to equilibrium is very slow.

When researchers have used long runs of historical time-series data (about 100 years of annual data) or panel data sets for the recent float covering at least 30 countries, results have conformed more to strict PPP. The key to resolving the failure of strict PPP for the recent floating period lies in understanding the forces that keep a nominal exchange rate away from an equilibrium based solely on relative prices. This is relates partly to the rigidity of prices in the face of nominal shocks and partly to the impact of real disturbances.

This paper presents a reduced-form model of the real exchange rate consisting of two components: a real interest differential and a set of fundamentals, including net foreign asset accumulation, productivity bias, and fiscal balances. Using multivariate cointegration methods, the model is implemented for the real effective exchange rates of the U.S. dollar, the Deutsche mark, and the Japanese yen from 1974 Q1 to 1993 Q2. In contrast to other similar research there is evidence of significant and sensible long-run relationships for this model and also for a simplified version of the model. The estimated long-run relationships are used to produce dynamic equations, which outperform a random walk and produce sensible dynamic patterns in the context of an impulse response analysis.

I. INTRODUCTION

Recently there has been a revival of interest in modeling the long-run behavior of nominal bilateral exchange rates using 'fundamentals' such as relative prices.[2] In general, this line of research has established that for the recent floating period weak-form purchasing power parity (PPP) would seem to hold on a single currency basis, but strong-form PPP does not.[3] Additionally, the adjustment to equilibrium in PPP-based equations is painfully slow. In order to obtain strong-form PPP results, and relatively rapid adjustment, researchers have used long runs of historical time series data (see, for example, Abuaf and Jorion (1990) and Diebold, Husted and Rush (1991)), or panel data sets defined for the recent float (see, for example, Frankel and Rose (1996), MacDonald (1988) and Wei and Parsely (1995)).

However, it is still of interest, from both an academic and policy perspective, to establish sensible long-run relationships for a single currency using data only from the period of recent floating. The key to resolving the failure of strong-form PPP lies in understanding the forces that keep a nominal exchange rate away from a PPP equilibrium. Undoubtedly an element of this is related to the rigidity of prices in the face of nominal shocks (as in the seminal Dornbusch (1976) model) while the remainder reflects the impact of real disturbances. MacDonald and Marsh (1996) have demonstrated that proxying such real and nominal disturbances using interest rates produces sensible PPP-based equilibrium exchange rates and also impressive out-of-sample forecasts. The objective in the current paper may be seen as an attempt to specify the real factors proxied in the MacDonald and Marsh paper and also to empirically model their influence on the equilibrium effective exchange rates of the U.S. dollar, German mark, and Japanese yen over the period of recent floating.

There have been a number of previous attempts at modeling equilibrium real exchange rates for the recent floating period and such work has not proved particularly fruitful. For example, modeling exercises which use single currency data fail to establish a significant long-run link between real exchange rates and fundamentals, such as real interest differentials (see Meese and Rogoff (1988), Edison and Pauls (1993) and Coughlin and Koedijk (1990)).[4] Given the rather negative conclusions from this 'behavioral' literature, can another examination of this

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[2]See the surveys of Breuer (1995), Froot and Rogoff (1995) and MacDonald (1995).

[3]In terms of an equation which conditions an exchange rate on relative prices, weak form PPP requires that the generated residual series be stationary, while strong form PPP requires the former condition plus homogeneity. See MacDonald (1993) for a further discussion.

[4]As in nominal exchange rate studies, one way of addressing such failures when using data for the recent float has been to estimate equilibrium real exchange rate relationships using panel data sets—Chinn (1996), Chinn and Johnson (1996), MacDonald (1996), MacDonald, Marsh and Nagayasu (1996).

kind of modeling be justified?[5] We believe it can. One key message to come from the literature on modeling nominal exchange rates is that the econometric methods used, and also the model specification, can have a crucial bearing on the findings of significant and sensible long-run relationships for single currencies.[6] In this paper we use the method of Johansen (1988, 1991) and report evidence of sensible and significant long-run relationships. An attractive feature of this econometric method is that it also facilitates computing the short-run dynamic behavior of our chosen exchange rates. Although this is a secondary objective of our work it is, nevertheless, of interest to examine how an exchange rate returns to its equilibrium value after a disturbance and to pitch our dynamic models against the random walk paradigm.

In terms of the policy debate regarding equilibrium real exchange rates, much of discussion has focussed on the concept of a fundamental equilibrium exchange rate (FEER), an explicitly normative approach which offers an appealing way of thinking about the evolution of actual and equilibrium real exchange rates.[7] However, although the FEER's concept has a number of attractive features, the main difficulty associated with it is one of tractability in terms of the need to have a fully specified multilateral structural model and, further, it does not provide an empirical link between a real exchange rate and its determinants. A key attraction of behavioral time series methods for analyzing single currency real exchange rates is the relative ease with which they may be computed and the fact that they do spell out the links with the underlying fundamental determinants.

The outline of the remainder of this paper is as follows. In the next section we use a decomposition of the real exchange rate which facilitates a discussion of the factors introducing systematic trends into the behavior of the equilibrium real exchange rate. These factors are defined in Section III and are labeled the fundamentals exclusive of real interest rates (FERID); they include variables such as net foreign asset accumulation, productivity bias and fiscal balances. Using the real uncovered interest rate parity condition we go on, in Section IV, to define a static relationship for the current equilibrium exchange rate in terms of the FERID variables and the real interest differential (RID). We propose operationalizing this model using a vector error correction framework. In Section V our data sources are given and the construction of the proxies for our variables, introduced in Section IV, are defined. Our estimates of the long-run exchange rate relationships and short-run dynamic results are presented in Section VI. The paper closes with a concluding section.

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[5]Recent studies in the spirit of that implemented here are Faruqee (1995) and Stein (1995).

[6]See, for example, Cheung and Lai (1993a), Kugler and Lenz (1993), MacDonald (1993) and MacDonald and Taylor (1993).

[7]This concept was originally proposed by Williamson (1985) and is based on an internal-external balance framework. See Wren-Lewis (1990) for a useful overview of the concept; see also the papers contained in Williamson (1995). An alternative internal-external balance approach to modeling equilibrium exchange rates has been derived by Stein (1995) and labeled the NATREX.

II. A REAL EXCHANGE RATE DECOMPOSITION

In this section we briefly discuss some real exchange rate decompositions which are useful in motivating our empirical tests. The real exchange rate, defined with respect to a general or overall price level, such as the CPI, is given by

$$q_t \equiv s_t - p_t^* + p_t \tag{1}$$

where q_t denotes a real exchange rate, s_t denotes the nominal spot exchange rate, defined as the foreign currency price of a unit of home currency (this is the most convenient definition since in our empirical application we use effective exchange rates), p_t denotes a price level and an asterisk denotes a foreign magnitude. Lower-case letters denote logarithms of the variables. In this context, therefore, a rise (fall) in q_t denotes an appreciation (depreciation) of the general real exchange rate. A similar relationship may be defined for the price of traded goods as

$$q_t^T \equiv s_t - p_t^{T*} + p_t^T, \tag{2}$$

where a T superscript indicates that the variable is defined for traded goods. If the prices in (2) are composite terms then, as we shall emphasize below, for q_t^T to be constant we have to assume that each of the goods prices which enters p_t^T *has* an equivalent counterpart in p_t^{T*}, and the weights used to produce each of these composite price levels are the same.[8]

We assume that the general prices entering (1) may be decomposed into traded and non-raded components as:

$$p_t = (1 - \alpha_t) p_t^T + \alpha_t p_t^{NT}, \tag{3}$$

$$p_t^* = (1 - \alpha_t^*) p_t^{T*} + \alpha_t^* p_t^{NT*}, \tag{3'}$$

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[8]We could, of course, define (2) simply for a single traded good and it would therefore capture the law of one price. However, it is more appropriate for our purposes to think in terms of overall traded-goods price levels. The construction of such is presumably less controversial than the use of identical weights to construct general price levels.

where the α's denote the shares of nontradeable goods sectors in the economy, and are assumed to be time-varying, and NT denotes a non-traded good. By substituting (3), (3') and (2) in (1) a general expression for the long-run equilibrium real exchange rate, $\bar{q}_t$, may be obtained as

$$\bar{q}_t \equiv q_t^{T} + \alpha_t^* (p_t^{NT*} - \alpha_t (p_t^{T} - p_t^{NT}). \tag{4}$$

Equation (4) is illuminating since it highlights three potentially important sources of long-run real exchange rate variability: nonconstancy of the real exchange rate for traded goods, which will arise if the kinds of goods entering international trade are imperfect substitutes and there are factors (discussed below) which introduce systematic variability into q_t^T; movements in the relative prices of traded to non-traded goods between the home and foreign country, due to, say, productivity differentials in the traded goods sectors; differing time-variability of the weights used to construct the overall prices in the home and foreign country. Let us consider each of these sources of variability in greater detail.

III. SOURCES OF TRENDS IN THE LONG-RUN REAL EXCHANGE RATE.

A. The Traded-Non-traded Price 'Ratio'

The first group of factors we consider relate to the relative price of traded to non-traded goods across countries, captured in equation (4) by the term $(p_t^{T*} - p_t^{NT*}) - (p_t^{T} - p_t^{NT})$. One way of interpreting this term is to think of it capturing factors which impinge on the relative price of non-traded goods, without necessarily affecting the relative price of traded goods

Balassa-Samuelson

Perhaps the best known source of systematic changes in the relative price of traded to non-traded goods is the Balassa-Samuelson effect.[9] This presupposes that the nominal exchange rate moves to ensure the relative price of traded goods is constant over time; that is, $q_t^T = c$. Productivity differences in the production of traded goods across countries can introduce a bias into the overall real exchange rate because productivity advances tend to be concentrated in the traded goods sector. The possibility of such advances in the non-traded sector is limited or non-existent; the productivity of haircuts, for example, has probably been constant since at least Byzantian times! If the prices of traded and non-traded products are linked to wages, wages linked to productivity, and wages linked across non-traded and traded industries, then the relative price of traded goods will rise less rapidly over time for a country with relatively high productivity in the tradeable sector: the real exchange rate, defined using overall price

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[9]The Balassa-Samuelson effect has been empirically investigated by, *inter alia*, Hsieh (1982), Marston (1990) and Miles-Ferretti (1994).

indices, appreciates for fast growing countries, even when the law of one price holds for traded goods; in terms of (4), if the home country is the relatively fast growing country it will have a positive $(p_t^{T*} - p_t^{NT*}) - (p_t^T - p_t^{NT})$ term thereby pushing $\bar{q}_t$ above q_t^T [10] (remember the currency is here defined as the foreign currency price of a unit of home currency).[10]

The demand side and non-traded goods

The existence of non-traded goods may allow a demand side bias which pushes an exchange rate away from its PPP level defined using traded goods prices. Assuming unbiased productivity growth, Genberg (1978) has demonstrated that if the income elasticity of demand for non-traded goods is greater than unity, the relative price of non-traded goods will rise as income rises (that is, as income rises households will spend a disproportionate amount of their income on services). This relative price change will be reinforced if, as seems likely, the share of government expenditure devoted to non-traded goods is greater than the share of private expenditure, and if income is redistributed to the government over time.[11]

We may therefore think of the second term in (4) as having the following general functional form:

$$(p_t^{T*} - p_t^{NT*}) - (p_t^{NT} - p_t^T) = g(\overset{+}{PROD}, \overset{+}{DEM}), \tag{5}$$

where *PROD* is a measure of productivity bias and *DEM* represents demand side bias. For the reasons noted above, a rise in the domestic value of either of these variables will, *ceteris paribus*, generate an appreciation of the overall real exchange rate.

B. Imperfect Substitutability of Traded Goods Prices

The factors in the last section can affect the real exchange rate even if traded goods are perfect substitutes across countries and q_t^T is constant. The constancy of the real exchange rate defined with respect to traded goods prices is not, however, uncontroversial. For example, there is now considerable evidence to suggest that the kinds of goods produced by industrial countries are not perfect substitutes and therefore the idea that price differences are quickly arbitraged away is completely unrealistic.[12] We now turn to some of the factors which may introduce systematic variability into q_t^T.

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[10]See Hallwood and MacDonald (1994) for further details.

[11]See Hallwood and MacDonald (1994) for further details.

[12]See Isard (1977), Kravis and Lipsey (1978) and Rogers and Jenkins (1995).

National savings and investment and the real exchange rate

The relative price of traded goods, q_t^T, is a major determinant of the goods and nonfactor services component of the current account. The current account, in turn, is driven by the determinants of national savings and investment and since one key component of national savings is the fiscal balance it follows that the fiscal balance is a determinant of the q_t^T component of the REER. Initial interest in the relationship between the government fiscal deficit and the real exchange rate was stimulated by the Reagen experiment in the 1980s (see, for example, Evans (1985)) and, more recently, with the desire on the part of Clinton administration for fiscal consolidation (see the references in Clark and Laxton (1995)). The effect of fiscal policy on the real exchange rate may be discussed by asking the question: will fiscal consolidation strengthen or weaken the external value of a currency?

Both outcomes are in fact potentially correct—it just depends on which particular view of the world is adopted. In the traditional Mundell-Fleming two-country model, a tightening of fiscal policy, which increases a country's national savings,[13] would lower the domestic real interest rate and generate a (permanent) real currency depreciation which, in turn, would produce a permanent current account surplus.[14] The real currency depreciation would also occur in flexible price models (see Clark and Laxton (1995)). What we are picking up in all these models is the 'crowding in' effect of the exchange rate depreciation; the necessity for aggregate demand to equal aggregate supply forces this result irrespective of the class of model.

The basic Mundell-Fleming model, however, ignores the effects of the stock-flow implications of the initial current account imbalance. Models which account for the stock implications of the initial fiscal tightening are portfolio balance models (see, for example, Branson (1977), Allen and Kenen (1980), Blanchard and Dornbusch (1984)) and the asset market/ balance of payments synthesis model of Frenkel and Mussa (1988) and Mussa (1984)). In the context of this class of model, the long-run is defined as a point at which the current account is balanced or, to put it slightly differently, any interest earnings on net foreign assets are offset by a corresponding trade imbalance. Hence if the fiscal consolidation is permanent it will imply a permanent increase in net foreign assets and an appreciation of the long-run real exchange rate.[15] Other expenditure effects can be analyzed in a similar fashion.

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[13]The analysis of national savings and investment and their effects on the real exchange rate is central to the IMF's analysis of real exchange rates; see Clark et at (1994). See also Clark and Laxton (1995) for a discussion of the short- and long-run effects of a fiscal contraction.

[14]See Hallwood and MacDonald (1994) for a discussion of the two-country Mundell-Fleming model.

[15] This result, of course, presupposes the absence of Ricardian equivalence.

In terms of national savings and investment the other key determinant of the q_t^T component of the real effective exchange rate is private sector net savings. It is often assumed that such savings are relatively constant over time. However, the independent effect of secular determinants of savings on the net foreign asset position should not be discounted. This, however, is unlikely to be a good working assumption for a country like Japan and even for the United States; where as early evidence suggested a relatively constant U.S. savings rate (see David and Scadding (1974), recent work finds a significant trend in U.S. savings (see Gokhale et al (1996)). More generally, Masson et al (1993) note: "demographic variables that reflect the age structure of the population seem to be important determinants of the cross country variations of saving rates .. and hence should affect net foreign asset positions."

The real price of oil

Changes in the real price of oil can also have an effect on the relative price of traded goods, usually through their effect on the terms of trade. The importance of this variable was highlighted by the dramatic increases in the real price of oil in the 1970s (for example, in the early 1970s the real price of oil rose by approximately 65 percent) and the equally dramatic fall in the mid-1980s (by approximately 50 percent). In comparing a country that is self-sufficient in oil with one which requires to import oil, the former, *ceteris paribus*, would appreciate in terms of the other currency as the price of oil rose. More generally, countries which have at least some oil resources could find their currencies appreciating relative to countries which do not have oil resources.

The effect of the various variables discussed in this section on the real exchange rate expressed in terms of traded goods may be summarized using the following relationship:

$$q_t^T = f(\overset{+}{FISC}, \overset{+}{PS}, \overset{?}{ROIL}), \tag{6}$$

where *FISC* captures the effect of relative fiscal balances on the equilibrium real exchange rate, *PS* represents private sector savings and *ROIL* is the real price of oil. The signs above the variables summarize the long-run effects of these variables on the real exchange rate.

C. Systematic Trends in the α Weights

It is widely accepted that the weights used to construct overall price series differ across countries (see, for example, Dornbsuch (1987)). Often in PPP calculations such differences are assumed constant across countries and therefore in a relative PPP calculation, or indeed when looking at the time series properties of the real exchange rate, they do not matter. Since the evidence on the relative importance of this effect is unclear, we therefore do not explicitly model the time variability of the α weights. However, since we explicitly check the in-sample stability of our estimated coefficients, we do not believe that this is a serious omission.

Combining (5) and (6) we obtain the following general relationship for the equilibrium real exchange rate, where the signs above the variables should be obvious from the above discussion.

$$\bar{q}_t \overset{-}{=} h(\overset{+}{PROD}, \overset{+}{DEM}, \overset{+}{FISC}, \overset{+}{PS}, \overset{+/-}{ROIL}). \tag{7}$$

In the next section we detail how (7) may be operationalized.

IV. THE ADJUSTMENT OF THE REAL EXCHANGE RATE TO STATIC EQUILIBRIUM AND ECONOMETRIC METHODS

In the last section we discussed the key determinants of the long-run equilibrium real rate. In this section we address the issue of how the actual exchange rate adjusts to the long-run rate.[16] To tie up the short-run with the longer run perspective, we start by introducing the familiar uncovered interest parity (UIP) condition:

$$E_t(\Delta s_{t+k}) = (i_t - i_t^*), \tag{8}$$

where an i_t denotes a nominal interest rate, Δ is the first difference operator, E_t is the conditional expectations operator, t+k defines the maturity horizon of the bonds and other symbols have the same interpretation as before.[17] Equation (8) may be converted into a real relationship by subtracting the expected inflation differential — $E_t(\Delta p_{t+k} - \Delta p_{t+k}^*)$ — from both sides of the equation. After rearrangement this gives:

$$q_t = E_t(q_{t+k}) - (r_t - r_t^*), \tag{9}$$

where $r_t = i_t - E_t(\Delta p_{t+k})$ is the *ex ante* real interest rate. Expression (9) describes the current equilibrium exchange rate as being determined by two components, the expectation of the real exchange rate in period t+k and the real interest differential with maturity *t+k*. We assume that

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[16]Our operationalization of the short-run real exchange rate equation follows Isard (1983), Meese and Rogoff (1988), Edison and Pauls (1993), Koedijk and Schotman (1990) and Baxter (1994).

[17]We have not explicitly included a risk premium in (8); however, the dynamic structure of our estimated short-run exchange rate models implies that they are entirely consistent with the existence of a time-varying risk premium.

the unobservable expectation of the exchange rate, $E_t(q_{t+k})$, is the equilibrium exchange rate defined in the previous section, namely $\bar{q}_t$:[18]

$$q_t = \bar{q}_t - (r_t - r_t^*),$$ (9')

In our model, therefore, the actual equilibrium exchange rate given by (9') comprises two components: the first component, $\bar{q}_t$, driven by the fundamentals exclusive of the real interest differential (FERID) discussed in the previous section, and the real interest differential (RID). The equilibrium condition represented by (9) is static and is unlikely to hold continuously. How then does the actual rate adjust to the rate given by (9)?

Since the variables in the FERID and RID terms and q_t are potentially I(1) processes (this issue is considered in the next section), and since there may exist cointegrating relationships amongst these variables, we propose using a cointegration framework to calculate the static relationship given by (9'). Specifically, we define the (nx1) vector of variables, consisting of the variables contained in the vector FERID and RID and q_t as x_t and assume that it has a vector autoregressive representation of the form:

$$x_t = \eta + \sum_{i=1}^{p} \Pi x_t + \varepsilon_t,$$ (10)

$$\Delta x_t = \eta + \sum_{i=1}^{p-1} \Phi_i \Delta x_{t-i} - \Pi x_{t-1} + \varepsilon_t,$$ (11)

where η is a (nx1) vector of deterministic variables, and ε is a (nx1) vector of white noise disturbances, with mean zero and covariance matrix Ξ. Expression (10) may be reparameterized into the vector error correction mechanism (VECM) as:
where Δ denotes the first difference operator, Φ_i is a (nxn) coefficient matrix
(equal to $-\sum_{j=i+1}^{p} \Pi_j$), Π is a (nxn) matrix (equal to $\sum_{i=1}^{p} \Pi_i - I$) whose rank determines the

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[18]This assumption has been invoked by, for example, Meese and Rogoff (1988).

number of cointegrating vectors.[19] If Π is of either full rank, n, or zero rank, Π=0, there will be no cointegration amongst the elements in the long-run relationship (in these instances it will be appropriate to estimate the model in, respectively, levels or first differences). If, however, Π is of reduced rank, r (where r<n), then there will exist (nxr) matrices α and β such that Π=$\alpha\beta'$ where β is the matrix whose columns are the linearly independent cointegrating vectors and the α matrix is interpreted as the adjustment matrix, indicating the speed with which the system responds to last period's deviation from the equilibrium level of the exchange rate. Hence the existence of the VECM model, relative to say a VAR in first differences, depends upon the existence of cointegration.[20] As we have noted, for our model to be valid, cointegration must exist amongst the variables in (10).

We test for the existence of cointegration amongst the variables contained in x_t using two tests proposed by Johansen. The likelihood ratio, or Trace, test statistic for the hypothesis that there are at most r distinct cointegrating vectors is

$$TR = T \sum_{i=r+1}^{N} \ln(1 - \hat{\lambda}_i), \tag{12}$$

where $\hat{\lambda}_{r+1}, ... , \hat{\lambda}_N$ are the N-r smallest squared canonical correlations between x_{t-k} and Δx_t series (where all of the variables entering x_t are assumed I(1)), corrected for the effect of the lagged differences of the x_t process. (For details of how to extract the λ's, see Johansen (1988), and Johansen and Juselius, (1991). Additionally, the likelihood ratio statistic for testing at most r cointegrating vectors against the alternative of r+1 cointegrating vectors—the maximum eigenvalue statistic—is given by (13):

$$LR = T \ln(1 - \lambda_{r+1}) \tag{13}$$

Johansen (1988) shows that (12) and (13) have a non-standard distribution under the null hypothesis. He does, however, provide approximate critical values for the statistic, generated by Monte Carlo methods (see also Osterwald-Lenum (1993)). It has been pointed out by, for example, Cheung and Lai (1993b) that these statistics may be subject to size distortions depending on the chosen DGP and sample size. To correct for the possibility of such in this paper we follow Reimers (1992) and report, in addition to (12) and (13), the small sample

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[19]It is straightforward to demonstrate that equation (11) is simply a reparameterization of a VAR in levels.

[20]The so-called Granger representation theorem (see Engle and Granger (1987)) implies that if there exists cointegration amongst a group of variables there must also exist an error correction representation.

corrected formulas:

$$TR = (T - np) \sum_{i=r+1}^{N} \ln(1 - \lambda_i),$$ (12')

and

$$LR = (T - np)\ln(1 - \lambda_{r+1}).$$ (13')

Although an examination of long-run exchange rate relationships is instructive, it can nevertheless be problematic since an interpretation of the coefficients in the long-run relationship as, say, elasticities is based on the (often implicit) ceteris paribus assumption that a unit shock does not have an effect on the other variables as well. For example, a fiscal shock will likely affect the real interest differential and perhaps also net foreign assets if it alters national savings. Since such interrelationships are summarized in our VAR model, we may use this to get a feel for these relationships. To do this, we employ an impulse response representation of the VAR. Such an approach has the more general benefit of illustrating the short-run dynamic responses of our group of three exchange rates with respect to the fundamentals. Additionally, the impulse response framework should give a feel for how long it takes a real exchange rate to adjust back to equilibrium following a real disturbance. This is potentially useful since in a systems-based modeling framework, such as that adopted here, the single-equation adjustment speeds (the α's) are difficult, if not impossible, to interpret.

Although impulse response methods have been used in a number of applications elsewhere, and therefore the method is well-known, practically all previous applications ignore the implications of potential cointegrating relationships in the calculation of the impulse responses. In this paper we calculate the impulse responses with the long-run relationships imposed. The standard impulse response approach involves calculating the moving average (MA) representation of the VAR system (10) and examining the response of the exchange rate change to orthogonal impulses. More specifically, the approach involves the following. On the assumption that all of the variables in the vector x_t are stationary (we return to this assumption below), then Wold's decomposition theorem implies the following canonical MA representation for x_t:

$$x_t = \eta + \sum_{i=0}^{\infty} \Psi_i \varepsilon_{t-i}$$ (14)

where of terms not previously defined, $\Psi_0 = I_n$ and the infinite sum is defined as the limit in mean square. This relationship may then be used to examine the effect of shocks, as represented by the white noise disturbances, , on the elements of the x_t vector. However, a common problem with this is that since the covariance matrix Σ_ε, is unlikely to be diagonal it

is difficult to interpret the effects of a particular shock on, say, the exchange rate. This is because the shock will in all likelihood have a contemporaneous effect on other shocks which, in turn, will have an impact on the exchange rate making it impossible to unravel the sole influence of the initial shock. A standard way of dealing with this problem is to use the MA representation with orthogonalised innovations. That is,

$$x_t = \sum_{i=0}^{\infty} \Theta_i \omega_{t-i} \tag{15}$$

where the components of ω are uncorrelated and a matrix P is chosen so that Σ_ω has unit variance (that is, $\Sigma_\omega = P^{-1} \Sigma_s (P^{-1})' = I_k$). The matrix P can be any solution of $PP^{-1} = \Sigma_\varepsilon$ and perhaps the most popular assumption is that P is chosen, using a Choleski factorization, as a lower triangular nonsingular matrix with positive diagonal elements; other decompositions, such as the 'structural' decompositions of Bernanke (1986) and Blanchard and Quah (1989) also exist. In the (stable) case the Ψ_i converge to zero as $i \to \infty$ and $\Sigma_x(h)$ converges to the covariance matrix of x_t as $h \to \infty$; however, this does not necessarily occur in the case of unstable, integrated or cointegrated VAR processes. Nevertheless, even for such processes it is still possible, as demonstrated by Lutkepohl (1993), to construct Ψ_i and θ_i. This is what we do here, subject to the restriction that there is one cointegrating vector for each of the systems (i.e. r = 1).

V. DATA SOURCES AND DEFINITIONS [21]

The sample period is 1975:Q1 to 1993:Q2, with data from 1974:Q1-Q4 used to construct lags. We use two measures of the real effective exchange rate. The first, LREER, is the multilateral CPI-based real effective exchange rate for the domestic country relative to its G7 partner countries, expressed in logarithms. The second, LREER1, is the equivalent ULC-based real effective exchange rate.[22] We use a number of FERID variables to capture the influence of the fundamentals noted in Section 3. We experiment with two variables to proxy for PROD. The first, LTNT, is the ratio of the domestic consumer price index to the wholesale price index relative to the equivalent foreign (trade weighted) ratio, expressed in logarithms.[23] The second is LPROD constructed from rates of growth in real output in manufacturing at home relative to the trade weighted foreign equivalent. We capture the effect

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[21]I am indebted to Susanna Mursula for constructing the data set used in this paper.

[22]Since the results for LREER1 are qualitatively very similar to those with LREER, we do not discuss them further in this paper. They are available from the author on request.

[23]This ratio is designed to proxy the ratio of traded to non-traded prices and was recommended by Kakkar and Ogaki (1993).

of fiscal deficits using the term FBAL which is the domestic fiscal balance as a proportion of GDP relative to the weighted sum of the partner countries (where the weights are those used to construct the effective exchange rates).[24] NFA is the ratio of the domestic country's net foreign asset position to GDP; it also captures the effect of fiscal policy on the real exchange rate as well as other factors more closely associated with private sector savings, such as demographics.

Two variables are used to capture the effect of commodity shocks. The terms of trade, LTOT, is constructed as the ratio of domestic export unit value to import unit value as a proportion of the equivalent effective foreign ratio, expressed in logarithms. ROIL is the real price of oil defined as the ratio of the nominal price of oil to the domestic country's wholesale price index, again expressed in logarithms. Finally, we use two relative real interest rate terms: RRL which is the long-term real interest differential constructed using the domestic 10-year nominal bond yield minus a centered 12-quarter moving average of the inflation rate minus the equivalent foreign effective; RRS is the equivalent short term differential, where three-month treasury bill rates and a 4-quarter centered been moving average were used.[25] The effect of all of these variables on the static equilibrium exchange rate (9) is summarized in (16).

$$q_t = h(\overset{+}{LRPROD}/\overset{}{LTNT}, \overset{+}{FBAL}, \overset{+}{NFA}, \overset{+/-}{LTOT}, \overset{+/-}{ROIL}, \overset{+}{RRL}/RRS) \tag{16}$$

VI. RESULTS

A. Unit Root Testing

In Table 1 we present a set of univariate tests for the null hypothesis that each of the series defined in the last section contains a unit root. These statistics are standard Dickey-Fuller tests for a unit root in the autoregressive representation for each of the series. Overall, the results are supportive of the hypothesis that each of the series, across each of the countries, does indeed contain a unit root. The tenor of these results finds support in Table 2 where we report multivariate tests for unit roots in each of the series. These tests are based on the two sets of VAR systems constructed for each currency (with long and, alternatively, short rates) and

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[24]The fiscal data come from the OECD Analytical Database, where the fiscal balance is defined as current government receipts minus current disbursements.

[25]Edison and Pauls (1993) demonstrate that the actual proxy used for expected inflation makes little difference to the outcome in this context.

have the null hypothesis of stationarity, given the cointegration space.[26] All of the reported test statistics have a high p-value, indicating the null of stationarity can be rejected for each variable in each of the systems. Before considering our implementation of the general model (16), we consider, first, a simplified version of (9') in which the long-run real exchange rate is assumed constant.

B. Long-Run Relationships

The lock between real exchange rates and real interest rates

Testing the relationship between a real exchange rate and a real interest differential, conditional on a constant equilibrium rate, has proven to be a relatively popular, although unsuccessful, way of modeling real exchange rates. We re-examine the model here because it should serve as a useful benchmark with which to compare our more general model and also

to establish if using more powerful econometric methods than those used by others produces satisfactory results. The model we estimate has the following form:

$$q_t = \beta_0 + \beta_1 r_t + \beta_2 r_t^* + \varphi_t \tag{17}$$

Equation (17) may be derived from (9') by assuming $\bar{q}_t = \beta_0$ a constant; we let the real interest rates be unconstrained in our estimation and test the restriction that they enter the relationship with equal and opposite signs. Meese and Rogoff (1988), Edison and Pauls (1993), Throop (1994), Coughlin and Koedijk (1990) use the Engle-Granger two-step cointegration method to estimate (17), for a variety of currencies and time periods, and find no evidence of a long-run relationship.[27] One reason for this may simply lie in the econometric technique used to estimate (17). Thus, Banerjee and others (1986) have noted that the small sample properties of the Engle-Granger method are poor. Additionally, if the regressors in (17) are endogenous and (or) the errors exhibit serial correlation then the asymptotic distribution of the coefficients will depend on nuisance parameters. Researchers have

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[26]These statistics are calculated on the basis of the number of significant cointegrating vectors, discussed in the next section (assumed to be 1), and have a χ^2 distribution. The 5 per cent critical value is 14.07.

[27]Throop (1994), using an error correction relationship for the real exchange rate/real interest rate relationship reports some evidence for cointegration on the basis of the estimated t-ratio on the error correction term; however, this is not significant on the basis of a small sample correction.

demonstrated that in testing equilibrium relationships for the nominal exchange rate, econometric methods robust to simultaneity bias and potential endogeneity can make a significant difference to the outcome.[28] Is the same true in the current application? We estimate (17) using the methods discussed in Section 4. These methods should produce asymptotically optimal estimates because they incorporate a parametric correction for serial correlation (which comes from the underlying VAR structure) and the systems nature of the estimator means the estimates should be robust to simultaneity bias.

Our results from estimating (17), using, alternatively, short and long rates, are reported in Tables 3 and 4, respectively. These tables contain Trace and LMax statistics, along with the normalized cointegrating relationships, adjustment coefficients and residual diagnostics. For all three currencies, we note that there is evidence of one significant cointegrating vector on the basis of both the Trace and LMax statistics. [29] [30] Normalising the significant cointegrating vector reveals that both interest rate terms are correctly signed. For example, a one percentage point increase in the German short rate produces an appreciation of the effective mark rate of around 2 percent. Tests of the hypothesis that the coefficients on the home and foreign interest rates are equal and opposite (LL(1)) cannot be rejected for any of the currencies at the 5 percent level.

In Table 3, the LB, LM and NM statistics are multivariate residual diagnostic tests: LB is Hoskings multivariate Ljung-Box statistic, LM(1 and 4) are multivariate Godfrey (1988) LM-type statistics for first and fourth order autocorrelation and NM(6) is a Doornik and Hansen (1994) multivariate normality test. Reported numbers are p-values and indicate, in general, an absence of serial correlation, although there is some evidence of non-normality in the Japanese yen and US dollar systems. In terms of the coefficients of determination, the explanatory power ranges from 0.16 for the dollar to 0.36 for the mark.

In Table 4 a similar set of results to those portrayed in Table 3 is presented for real exchange rates and real long-term interest rates. The picture here is broadly similar to that reported in Table 3. There is again evidence of significant long-run relationships for all three currencies, interest rate coefficients are generally correctly signed (apart from the coeffcients in the Japanese equation). The multivariate residual diagnostics again indicate an absence of serial correlation across the three systems, but there is some evidence of non-normality. For the

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[28]See, *inter alia*, Cheung and Lai (1993a), Kugler and Lenz (1993), MacDonald (1993) and MacDonald and Marsh (1994).

[29]Given the relatively parsimonious nature of this exchange rate relationship we do not adjust the critical values using a small sample correction.

[30]For each of the U.S. relationships plots of the residuals revealed two large outliers: one in 1980:Q2 and the other in 1982:Q4. We use two intervention dummy variables to model these outliers.

mark and dollar, explanatory power increases relative to the short rate systems, but the explanatory power of the yen stays unchanged.[31]

In sum, then, what is perhaps the simplest real exchange rate model does not do too badly relative to the metric set by other researchers, and also in terms of producing statistically significant long-run relationships which, in turn, produce dynamic equations that explain a reasonable percentage of the in-sample performance of an exchange rate change.

A non-constant real equilibrium exchange rate

In Tables 6 through 8 we present our cointegration results and associated statistics, for our general exchange rate model in which the equilibrium real exchange rate, $\bar{q}_t$, is time dependent and assumed to be a function of the variables contained in the vector FERID. As in the case of the simplest model, we experimented with both short and long interest rates. The LMax and Trace statistics are contained in Tables 5 and 6 for the systems containing short and long interest rates, respectively. On the basis of the standard set of significance values (that is, the values unadjusted for small sample bias), there is very strong evidence of cointegration for all three currencies, regardless of the interest rate measure. For both the dollar and and yen systems we have let the interest rate terms be unconstrained. This is based partly on the pretesting noted in Tables 3 and 4 and also on the fact that the relationships with the unconstrained interest rates produced the more appealing cointegrating vectors (in terms of having correctly signed coefficients). However, as is evident, the systems reported in Tables 5 and 6 are rather large and heavily parameterised. We have therefore adjusted the Trace an Lmax statistics using Reimer's (1992) small sample correction (see equations (12′) and (13′) above), reported in the columns labeled T-np. With these adjusted statistics the picture changes—there is now one statistically significant vector for each currency. We therefore proceed on the basis of one significant cointegrating vector for each of the currencies.

Estimates of the cointegrating vectors associated with the largest eigenvalues for each system are presented in Table 7 for both short and long rates. Each of the vectors has been normalised on the exchange rate (that is, the LREER has a coefficient of -1, so the coefficients are written in equation format). Across all of the exchange rate combinations there is a very good strike record in terms of correctly signed coefficients. Thus, all but 6, out of 46, are correctly signed and of plausible magnitude. The magnitude of coefficients is roughly comparable across the two sets of systems although there are some sign changes: real rates for the Japanese yen are correctly signed in the short rate system but incorrectly signed in the long

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[31]We also computed cointegrating relationships for both the short and long interest rate specifications, where the constant term entered unconstrained in the VAR. The results, both in terms of the number of significant cointegrating relationships and the coefficient estimates, were very similar to those reported with a constrained constant term.

rate system. Notice that the coefficient on FBAL has a sign consistent with a stock-flow model for Japan and the US but a 'traditional' Mundell-Fleming sign for the mark.[32] In absolute terms the fiscal balance is also more important for the mark. The Balassa-Samuleson effect, proxied here by the LTNT variables, enters all of the equations with relatively large coefficients, and indicates a more than proportional response of the real exchange rate in four out of the six cases.

In Figures 1 to 3 we plot the three real exchange rates against the appropriate relative prices. It is clear from these figures that there is more to nominal effective exchange rates than simple PPP (see MacDonald (1995) for a more detailed discussion). In Figures 4 to 9 we plot the long-run equilibrium values of the three currencies, derived from the equations reported in Table 7 against the actual outcomes.[33] It is evident from these figures that all three exchange rates seem to track the fundamental's-based equilibrium rates quite closely, certainly in terms of the broad trends, and this is so irrespective of the interest rate measure adopted. Interestingly, the plots for the U.S. dollar suggest that the fundamentals we have used track the steep appreciation and subsequent depreciation of the dollar, although the peak of the appreciation seems to be a non-fundamental, or purely speculative, phenomenon. Of course, this kind of discussion begs the question of whether the actual data fundamentals, used here to define the long-run equilibrium, were calibrated at sustainable levels throughout the sample. For example, it may be that the fiscal stance of the U.S. in the early 1980s was not the most appropriate and therefore one should recalibrate the equilibrium exchange rate using values of the relative fiscal position which more closely mirror sustainable values; this, however, is the topic of a separate paper.

In table 8 we present some diagnostics for the dynamic equations constituting the three VAR systems. We present the same array of diagnostics as reported in tables 3 and 4 for the simple RID model. The R^2's in Table are ordered to correspond to the variable listings in Table 7. Hence the first R^2 corresponds to the exchange rate equation, the second to the domestic real interest rate equation, and so on. In sum, the R^2's for the exchange rate equation are around double the value for the corresponding model in which the equilibrium rate is assumed constant; for the three short rate equations the average value is 0.55, while for the long rate equations the average value is 0.61. The portmanteau diagnostics are also respectable as all of the models pass the serial correlation tests and four pass the normality test.

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[32]The short- and long-run effects of fiscal policy on the real exchange rate are discussed in Clark and Laxton (1995).

[33]It is perhaps worth stressing that these are the fitted values from the significant cointegrating vector and not the fitted values from the exchange rate equation in the VAR system. The fitted values from the latter show an almost exact correspondence with the actual values throughout the sample period.

C. Short-run Dynamics

The random walk

Ever since the seminal paper by Meese and Rogoff (1983) the benchmark by which a fundamentals-based exchange rate model is assessed is by comparison to a simple random walk. As we noted earlier, such comparisons have not favoured real exchange rate models (see MacDonald and Taylor (1992) and Frankel and Rose (1995). Although we do not believe that beating a random walk should be the last word on the performance of an exchange rate model, especially when the primary objective of that model is to discover something about the longer run trends in exchange rates, we nevertheless thought it worthwhile to subject our models to a random walk horse race. This seems worthwhile because there is evidence that when the kinds of dynamic error correction models utilised in this paper are used to estimate nominal exchange rate models, they are able to beat a random walk.[34]

Our out-of-sample forecasts are constructed using the approach adopted by Meese and Rogoff (1983).[35] In particular, we re-estimated all of the models reported above for the period 1975:Q1 – 1987:Q2 and then calculated out-of-sample forecasts for horizons of one to eight quarters ahead. Each model, and the associated set of out-of-sample forecasts, was then re-estimated period by period through to 1993 Q2. Root Mean Square Error (RMSE) statistics were then computed for the forecasts and the ratio of these to the RMSE for a simple random walk model was calculated. The ratios are reported in Table 9. An asterisk indicates that the model was unable to beat a random walk.

For Germany it is evident that the simple RID models, in which the equilibrium real rate is assumed constant, produces forecasts which are marginally better than a random walk for all horizons apart from horizon 8 in the model with short interest rates. However, the models in which we incorporate the FERID variables into our equilibrium relationship clearly dominate both a random walk and the simple models. Model 3, the model with short rates, performs best, but both models turn in a dramatic improvement over a random walk by quarter 8. The pattern for the other two currencies is rather different from that of the mark. For the US dollar, there is no evidence of the simple models (1 and 2) outperforming a random walk, while for the yen there are three occurences of the simple models beating a random walk. However, for both these currencies there is considerable evidence of our general models beating a random walk. For the yen we can beat the random walk at all horizons and the system containing long rates dominates the comparable system with short rates. Similarly for the US dollar, the long rate system outperforms a random walk at all horizons, although the general system with short rates only outperforms the random walk at three horizons.

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[34]See MacDonald and Taylor (1993, 1994) and the model of MacDonald and Marsh (1996), which contains a nominal analogue to the simple real interest differential model considered here.

[35]That is to say, we construct 'perfect foresight' forecasts.

Impulse response functions

In Figures 10 to 12 the impulse response of the logarithmic change of the real effective exchange rates of our three long interest rate systems (the qualitative picture from the short rate systems is similar) are analysed with respect to orthogonalised shocks in each of the underlying fundamental variables. In each of the figures the impulse responses are bounded by two standard error bands, calculated using bootstrap methods. In particular, these bands were constructed using the sample standard deviation of the empirical distribution from a bootstrap simulation on the reduced form errors with 2000 replications. The variable ordering in these systems is: FBAL, ROIL, NFA, LTNT, LTOT, RRS/L, LREER. The ordering is intended to reflect the relative exogeneity of the series (FBAL most exogenous, LREER, least exogenous). The general tenor of the results contained in these figures is that the short-run exchange rate dynamics in response to a shock are rich, and the impact of a shock is often relatively long-lived and statistically significant.

For example, in the case of the United States, a 1 percent rise in the home real interest rate produces a one percent exchange rate appreciation by quarter 2, the exchange rate then depreciating to quarter 7. Both the productivity and terms of trade shocks produce (positive) exchange rate overshoots in the first quarter. The net foreign asset shock results in a less than proportionate appreciation of the real rate and the appreciation is long-lived. The fiscal balance shock initially produces an appreciation of the exchange rate although this is fairly rapidly reversed and there are a preponderance of negative changes after quarter 2.

For Germany, there is no evidence of overshooting with respect to any of the variables and, apart from the real interest rate shock and net foreign asset shock, the time profiles for the exchange rate are similar to their U.S. counterparts. Perhaps the major difference between the United States and German systems is that in the latter all of the exchange rate changes are insignificantly different from zero by about quarter 16. The response of the Japanese yen rate to the set of shocks is broadly similar to the German case.

One interesting feature of the impulse response results is that they give a feel for how long it takes a real exchange rate to adjust to a real disturbance, or shock. As we noted in the introduction, researchers have only been able to obtain half-lives of around 4 years when using data samples with extended spans (sucha s long runs of historical data or panel data sets)); single-equation, single-currency estimates for the recent floating period suggest half-lives of around 20 years (see MacDonald (1995)) and therefore very slow adjustment of exchange rates to shocks. Our single equation estimates for the recent float suggest relatively rapid adjustment after a shock. Thus, across the three exchange rate systems we observe the extinction of exchange rate changes between quarters 16 and 20, suggesting helf-lives of between two and two and one-half years.

VII. SUMMARY AND CONCLUSIONS

In this paper we have reexamined the determinants of real exchange rates in a 'long-run' setting. We presented a model of the equilibrium exchange rate which featured productivity

and terms of trade effects, in addition to fiscal balances, net foreign assets and real interest rates, as key fundamental determinants. Our model was shown to produce significant and sensible long-run relationships for the real effective exchange rates of the mark, dollar and yen, and it seemed much better suited than relative prices to explain the long-run trends in effective exchange rates. We also reported evidence of significant long-run relationships for a simplified version of our model and we noted that such significance contrasted with practically all of the extant research on this relationship.

Although our main focus in this paper was the long-run determinants of real exchange rates, it has become the acid test of a fundamentals-based exchange rate model that it should outperform a random walk model in terms of having a lower root mean square error. We found that our general real exchange rate model passed this test for each of the currencies. In general, systems which included long maturity interest rates did better than systems with short rates. The base line real exchange rate model — the model with a constant equilibrium real exchange rate — did not do so well in terms of the forecasting criterion. The short-run behavior of our model was further examined by calculating impulse response functions for real exchange rates with respect to orthogonalised shocks in our fundamental variables. The impulse response analysis provided a set of results which were intuitively plausible and statistically significant. Additionally, this analysis suggests that the impact of shocks on the real exchange rate is rapidly offset and this contrasts markedly with the single equation half-lives that others have reported using only data from the recent float.

We believe that our modelling exercises can be interpreted as indicating that fundamentals do have an important, and significant, bearing on the determination of both long- and short-run exchange rates. One way in which our work could be extended would be to utilise the methods of this paper to decompose real exchange rate behavior into both nominal and real components.

Table 1. Univariate Unit Root Tests

| | L | | ΔL | |
	μ	τ	μ	τ
German Data				
LREER	-1.70	-1.92	-3.72	-3.82
FBAL	0.80	-1.41	-3.19	-3.78
LTOT	-2.97	-2.79	-4.69	-4.66
LTNT	-1.69	-1.38	-2.78	-2.97
RRS	-2.65	-2.84	-4.48	-4.44
RRL	-2.29	-3.24	-3.82	-3.92
NFA	-0.66	-2.02	-3.72	-3.79
LROIL	-1.79	-2.19	-4.74	-4.81
Japanese Data				
LREER	-1.28	-2.65	-4.52	-4.45
FBAL	0.49	-1.41	-1.85	-4.09
LTOT	-1.39	-2.31	-4.25	-4.28
LTNT	-1.22	-2.63	-4.03	-3.99
RRS	-3.17	-3.16	-4.41	-4.35
RRL	-2.64	-2.52	-5.28	-5.47
NFA	0.27	-1.94	-3.32	-3.49
LROIL	-2.10	-2.23	-5.28	-5.34
U.S. Data				
LREER	-1.91	-2.06	-3.29	-3.26
FBAL	-2.63	-2.75	-1.78	-1.82
LTOT	-1.74	-2.32	-4.06	-4.02
LTNT	-1.87	-2.39	-4.07	-4.04
RRS	-2.24	-1.37	-4.01	-4.41
RRL	-2.09	-3.37	-3.70	-3.96
NFA	-0.15	-1.96	-1.99	-1.82
LROIL	-1.53	-2.26	-5.43	-5.46

Notes. The numbers denote augmented Dickey Fuller (ADF) t-ratios, where the lag length used in the underlying autoregression was chosen using the Schwarz selection criterion. The column headings μ and τ indicate, respectively, that only a constant and a constant plus a time trend are included in the underlying autoregression. The L and ΔL denote, respectively, that the unit root test relates to the level and first difference of the appropriate variable. The variables listed in the first column are as defined in the text. The five per cent critical values for the ADF statistics are approximately -2.89, without a time trend, and -3.43 with a time trend.

Table 2. Multivariate Unit Root Tests

	LREER	RRS/L	FBAL	LTOT	LTNT	NFA	LROIL
German Mark							
RRS	45.75	45.99	53.21	40.23	51.03	45.33	47.90
RRL	31.44	35.16	42.05	30.13	33.56	34.08	32.69
Japanese Yen							
RRS	59.82	26.27	61.40	57.15	53.22	56.52	60.72
RRL	62.30	45.54	66.67	59.00	53.24	60.56	63.34
U.S. Dollar							
RRS	66.12	52.19	65.81	64.84	64.57	64.28	65.75
RRL	52.09	35.15	55.32	50.33	53.63	52.46	51.89

Notes. The numbers are chi-squared statistics (with 7 degrees of freedom) and are tests of the null hypothesis of stationarity for each of the variables defined in the column heading. The statistics are calculated for the VAR systems with short and long interest rates entering as the alternate interest rate measures. The statistics are computed under the assumption that $r=1$; the 5 per cent critical value is 14.07.

Table 3. Real Exchange Rates and Real Short Term Interest Rates

	Germany	Japan	US
LR1	20.83	37.40*	21.49
LR2	8.91	12.22	12.26
LR3	4.67	3.75	3.82
TR1	33.69*	53.36*	37.57*
TR2	12.86	15.97	16.08
TR3	4.67	3.75	3.82
β_0	4.670	0.369	4.767
β_1	0.116	0.286	0.406
β_2	-0.189	-0.185	-0.328
LL(1)	3.17(0.07)	5.72(0.06)	2.42(0.12)
LB(18)	0.03	0.04	0.67
LM(1)	0.63	0.19	0.78
LB(4)	0.13	0.21	0.88
NM(6)	0.53	0.01	0.02
$R^2 q$	0.36	0.28	0.16
$R^2 r$	0.41	0.44	0.36
$R^2 r*$	0.37	0.39	0.34

Notes. The entries in the rows labeled LR1 to LR3 and TR1 to TR3 are, respectively, the estimates of the λMax (equation 13) and TRACE (equation 12) statistics. β_0, β_1, and β_2 are estimates of the constant, the coefficient on the domestic interest rate and the coefficient on the foreign interest rate. The R^2 's represent coefficients of determination for the variable subscripted. The numbers in the columns LB(18), LB(4), LM(4) and NM(6) are marginal significance levels; these statistics are discussed in the text. An asterisk denotes significance at the five per cent level.

Table 4. Real Exchange Rates and Real Long Term Interest Rates

	Germany	Japan	U.S.
LR1	23.72*	19.54	31.94*
LR2	9.35	9.46	9.86
LR3	6.96	4.60	4.71
TR1	40.03*	33.60*	46.51*
TR2	16.32	14.06	14.57
TR3	6.96	4.60	4.71
β_0	4.698	0.313	-4.664
β_1	0.067	-0.194	0.471
β_2	-0.162	0.285	-0.434
LL(1)	3.59(0.06)	0.87(0.35)	0.97(0.32)
LB(18)	0.01	0.03	0.02
LB(4)	0.01	0.05	0.51
LB(4)	0.79	0.08	0.63
NM(6)	0.45	0.10	0.03
R^2q	0.46	0.28	0.22
R^2r	0.25	0.42	0.53
R^2r*	0.35	0.37	0.39

Notes: See Table 3.

Table 5. Number of Cointegrating Vectors in the Complete Model (Short Rates)

	Germany		Japan		U.S.	
	T	T-np	T	T-np	T	T-np
LR1	80.80*	50.22*	71.01*	44.14*	85.19*	52.96*
LR2	48.84*	30.36	53.98*	33.56	67.23*	41.79
LR3	35.78*	22.24	39.48*	24.54	42.15*	26.20
LR4	20.36	12.65	34.43	21.46	35.70	22.19
LR5	17.44	10.84	27.99	17.40	28.39	17.64
LR6	10.36	6.44	18.12	11.26	21.59	13.42
LR7	5.17	3.21	12.90	8.01	15.93	9.92
LR8	-	-	10.28	6.30	6.78	4.21
TR1	218.76*	135.98*	267.60*	166.34*	302.96*	188.33*
TR2	137.96*	85.76	196.60*	122.21	217.77*	135.37
TR3	89.12	55.59	143.20*	89.02	150.54*	93.58
TR4	53.33	33.15	103.72*	64.47	108.39*	67.38
TR5	32.97	20.49	69.29	43.07	72.69	45.19
TR6	15.53	9.65	41.30	25.67	44.29	27.53
TR7	5.17	3.21	23.18	14.41	22.70	14.11
TR8	-	-	10.28	6.39	6.78	4.21
E1	0.53		0.64		0.76	
E2	0.46		0.43		0.46	
E3	0.27		0.37		0.39	
E4	0.17		0.34		0.27	
E5	0.16		0.21		0.24	
E6	0.09		0.17		0.19	
E7	0.04		0.12		0.15	
E8	-		0.10		0.11	

Notes: The entries in the rows labeled LR1 to LR8 and TR1 to TR8 are, respectively, the estimates of the λMax (equation 13) and TRACE (equation 12) statistics. The entries in the rows labeled E1 to E8 are the eigenvalues for each cointegrating vector.

Table 6. Number of Cointegrating Vectors in the Complete Model (Long Rates)

	Germany		Japan		U.S.	
	T	T-np	T	T-np	T	T-np
LR1	82.59*	51.33*	85.80*	53.34*	105.74*	65.73*
LR2	44.86*	27.88	63.45*	39.44	46.41*	28.84
LR3	32.41	20.15	36.35	22.59	37.77	23.47
LR4	23.10	14.36	30.36	18.87	23.27	14.47
LR5	16.82	10.46	28.51	17.76	20.80	12.93
LR6	11.64	7.34	20.06	12.47	15.84	9.85
LR7	4.12	2.56	11.96	7.43	12.82	7.97
LR8	-	-	5.55	3.45	8.67	5.38
TR1	215.55*	133.99*	282.04*	175.32*	271.07*	168.50*
TR2	132.96*	82.65	196.24*	121.98	165.33*	102.77
TR3	88.09	54.76	132.79*	82.55	119.17*	74.08
TR4	55.68	34.61	96.44*	59.95	81.40*	50.60
TR5	32.58	20.25	66.08	41.08	58.13*	36.13
TR6	15.76	9.79	37.58	23.36	37.33*	23.21
TR7	4.12	2.56	17.51	10.88	21.49*	13.36
TR8	-	-	5.55	3.45	8.67	5.38
E1	0.51		0.63		0.76	
E2	0.38		0.46		0.46	
E3	0.32		0.39		0.39	
E4	0.19		0.29		0.27	
E5	0.11		0.22		0.24	
E6	0.10		0.18		0.19	
E7	0.03		0.13		0.15	
E8	-		0.10		0.11	

Notes. See Table 5.

Table 7. Estimates of Long-run, or Equilibrium, Exchange Rates

Short Rates

	R	R*	FBAL	LTOT	LTNT	NFA	LROIL	CON
DM	0.030	-	-0.377	1.138	0.998	0.002	-0.130	4.628
JY	0.027	-0.089	0.018	-0.136	1.592	-0.045	0.498	0.517
USD	-0.094	-0.009	0.025	0.873	1.804	0.021	0.016	4.769

Long Rates

	R	R*	FBAL	LTOT	LTNT	NFA	LROIL	CON
DM	0.002	-	-0.809	1.146	1.329	0.008	0.237	4.538
JY	-0.012	0.014	0.004	0.022	0.821	-0.002	0.338	0.469
USD	0.084	-0.173	0.010	0.441	1.305	0.015	0.016	4.890

Notes. The numbers in this tables are normalised (on the exchange rate) cointegrating coefficients for the two systems (with short and long rates, respectively) discussed in the text. The relevant exchange rate is defined in the first column and the fundamental variables are defined in the text.

Table 8. Equation Diagnostics from VECM Models

Short Rates

	R^2	R^2	R^2	R^2	R^2	R^2	R^2	R^2	LB	LM	NM
DM	0.66	0.67	-	0.72	0.56	0.71	0.47	0.68	0.29	0.66	0.71
JY	0.65	0.67	0.73	0.57	0.78	0.82	0.47	0.63	0.70	0.20	0.25
USD	0.34	0.56	0.42	0.57	0.46	0.63	0.62	0.43	0.01	0.00	0.00

Long Rates

	R^2	R^2	R^2	R^2	R^2	R^2	R^2	R^2	LB	LM	NM
DM	0.48	0.58	-	0.49	0.45	0.68	0.27	0.52	0.46	0.75	0.08
JY	0.67	0.65	0.68	0.58	0.78	0.81	0.55	0.63	0.08	0.12	0.03
USD	0.54	0.68	0.63	0.65	0.58	0.79	0.72	0.64	0.08	0.13	0.00

Notes. The numbers reported in this table represent residual diagnostics corresponding to the equation orderings in table 8. The R^2s denote coefficients of determination and their ordering is consistent with the variable ordering in table 7. LB, LM and NM are as defined in the text.

Table 9. Out-of-Sample Forecast Results: RMSE Ratios.

German Mark

Forecast Horizon	Model 1	Model 2	Model 3	Model 4
1	0.938	0.928	0.737	0.871
2	0.921	0.940	0.636	0.873
3	0.962	0.976	0.627	0.832
4	0.988	0.973	0.522	0.713
5	0.994	0.922	0.389	0.559
6	0.943	0.858	0.273	0.453
7	0.937	0.840	0.249	0.366
8	*	0.907	0.255	0.346

Japanese Yen

Forecast Horizon	Model 1	Model 2	Model 3	Model 4
1	0.888	0.919	0.841	0.766
2	0.978	*	0.887	0.808
3	*	*	0.898	0.802
4	*	*	0.922	0.800
5	*	*	0.961	0.774
6	*	*	0.891	0.625
7	*	*	0.791	0.511
8	*	*	0.752	0.452

U.S. Dollar

Forecast Horizon	Model 1	Model 2	Model 3	Model 4
1	*	*	0.997	0.992
2	*	*	0.964	0.930
3	*	*	0.965	0.838
4	*	*	*	0.754
5	*	*	*	0.728
6	*	*	*	0.785
7	*	*	*	0.902
8	*	*	*	0.818

Notes. Model 1: Short Interest Rates/ Constant Equilibrium; Model 2: Long Interest Rates/ Constant Equilibrium; Model 3: Short Interest Rates + FERID; Model 4: Long Interest Rates+FERID.

Figure 1

lreer=_____   lrcpi=_____

German mark effective and
Relative Consumer Prices

Figure 2

lreer=_____   lrcpi=_____

US Dollar Real Effective and
Relative Consumer Prices

Figure 3

Ireer=_____   Ircpi=_____

Japanese Yen Real Effective and
Consumer Prices

Figure 4

US Dollar: Short Rates, FERID

Figure 5

Lreer=_____ Fitted=.......

US Dollar: Long Rates,
FERID

Figure 6

German Mark: Short Rates
FERID

Figure 7

German Mark: Long Rates
FERID

Figure 8

Japanese yen: Short Rates
FERID

Figure 9

Lreer=_____   Fitted=.......

Japanese Yen: Long Rates,
FERID
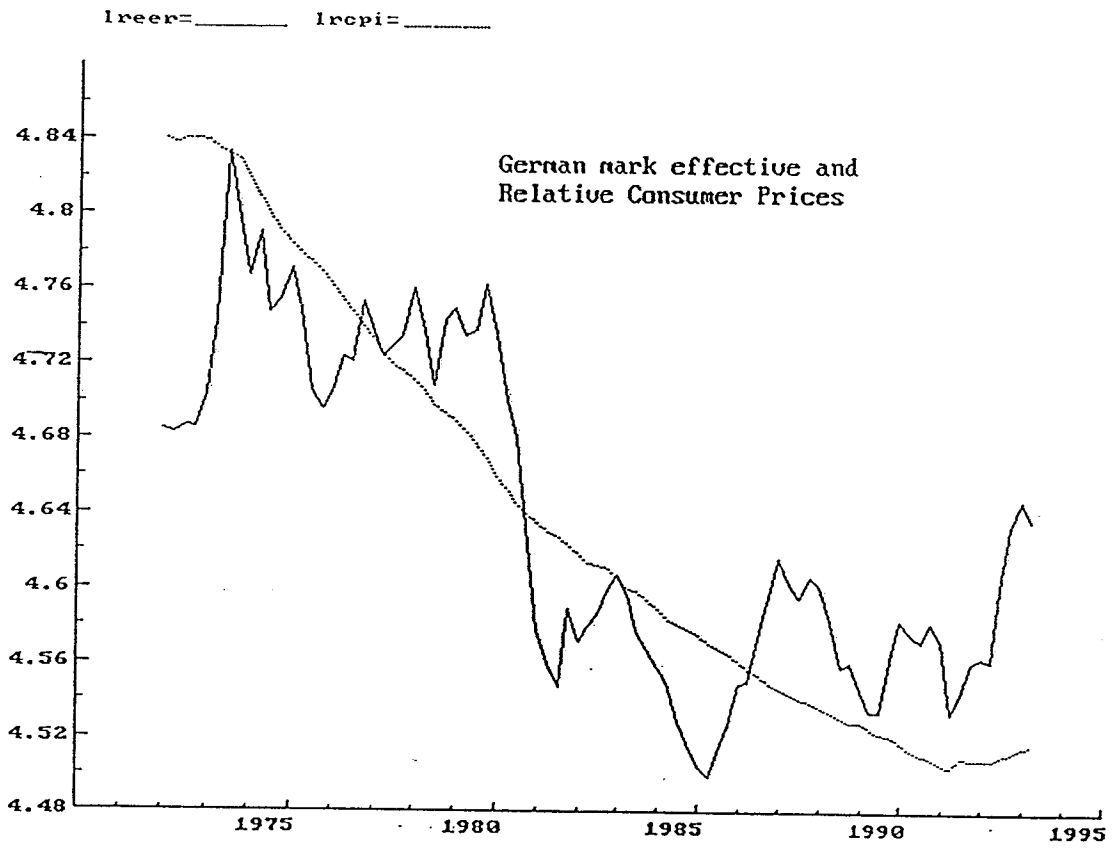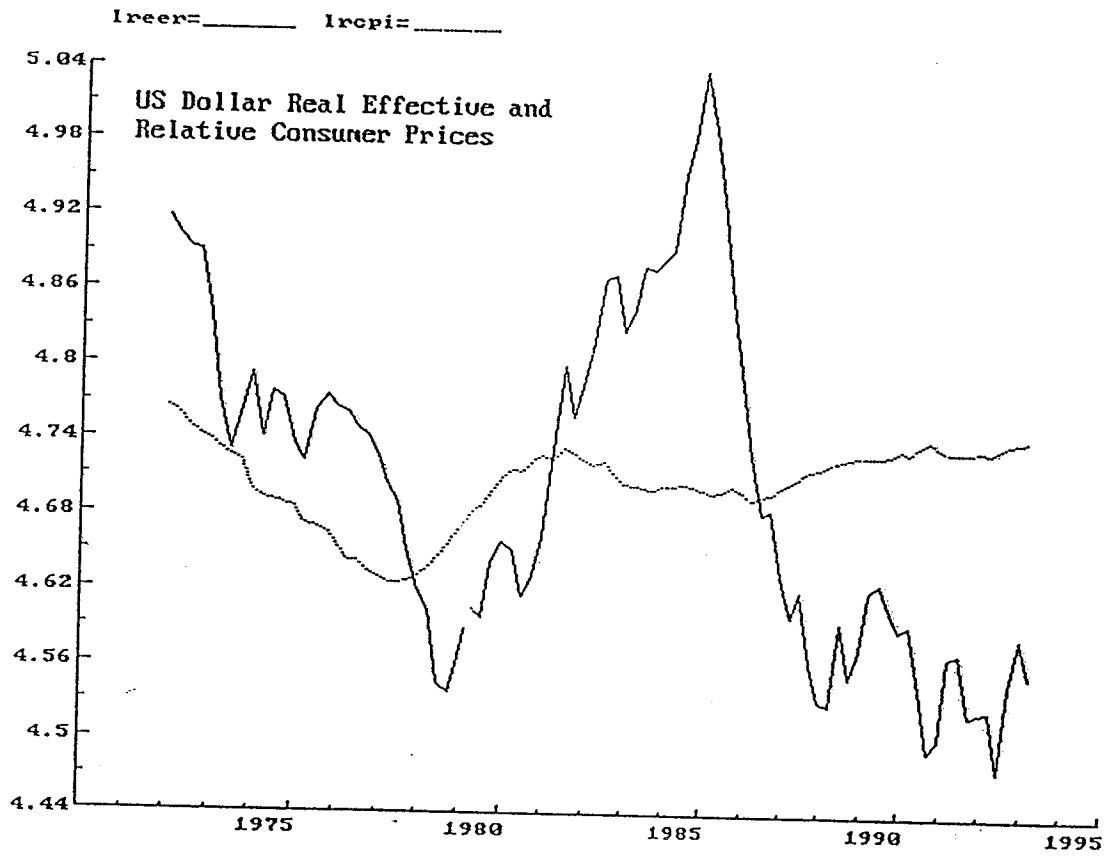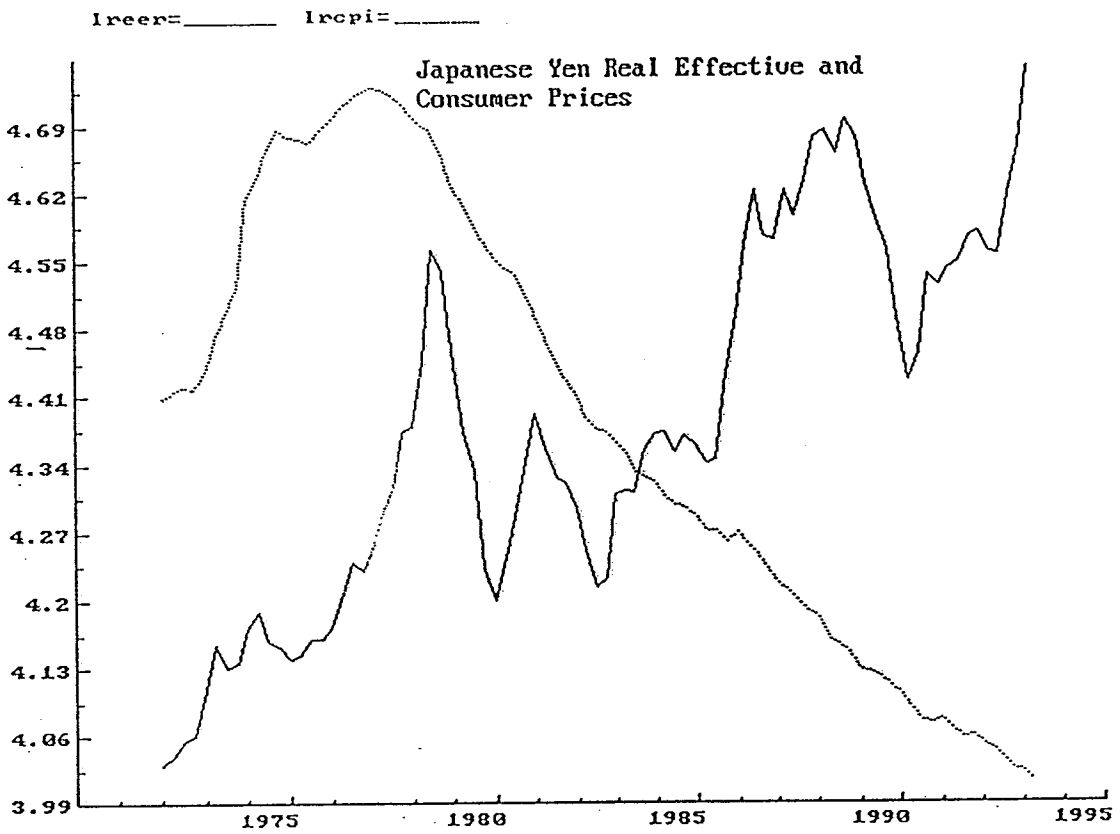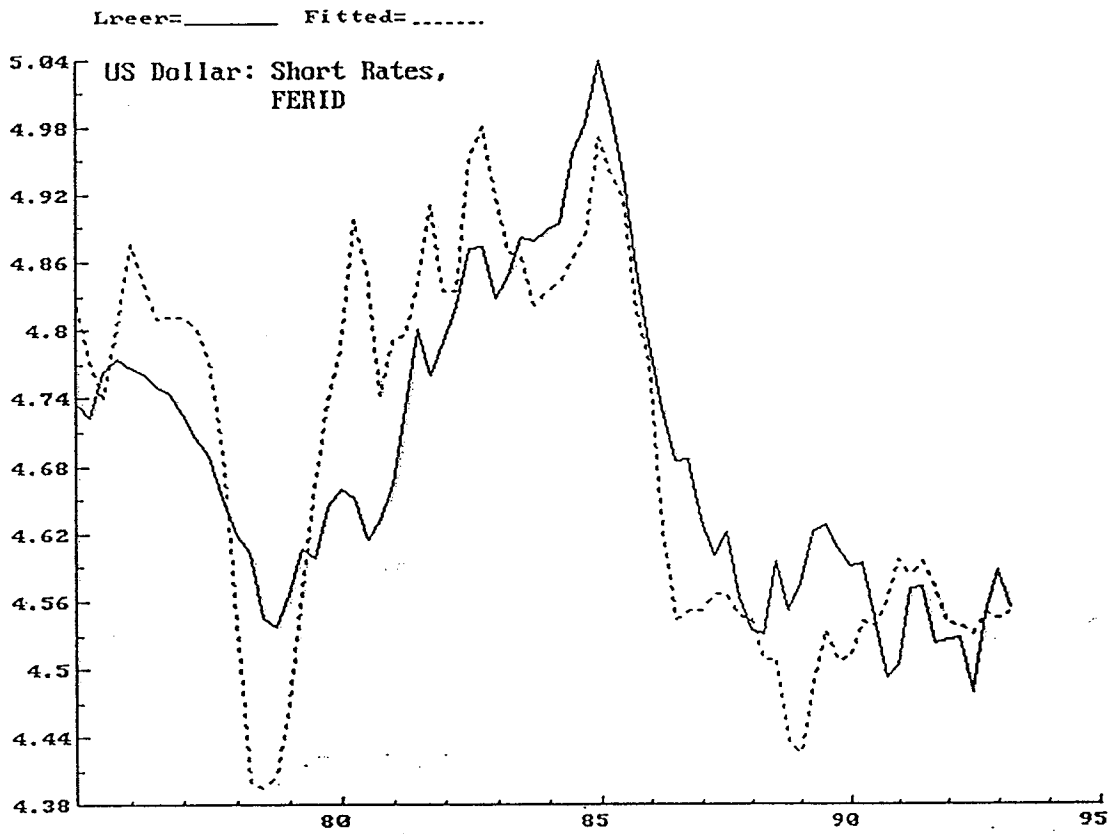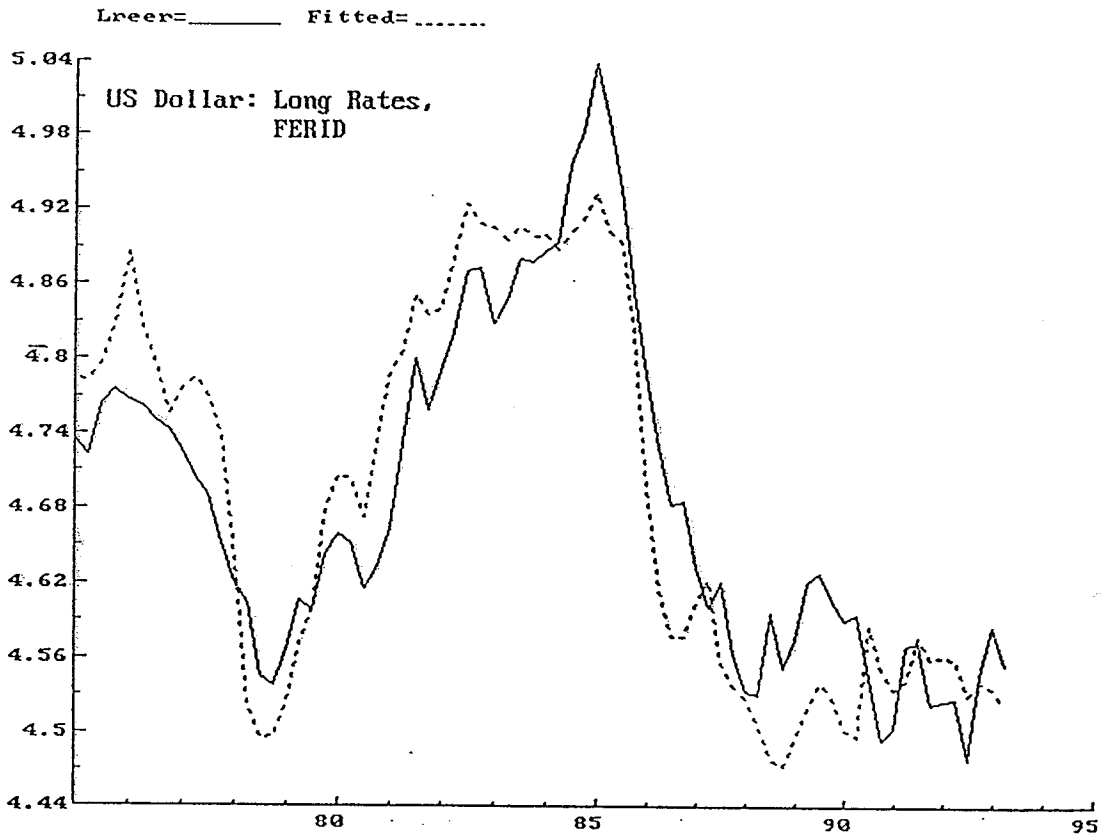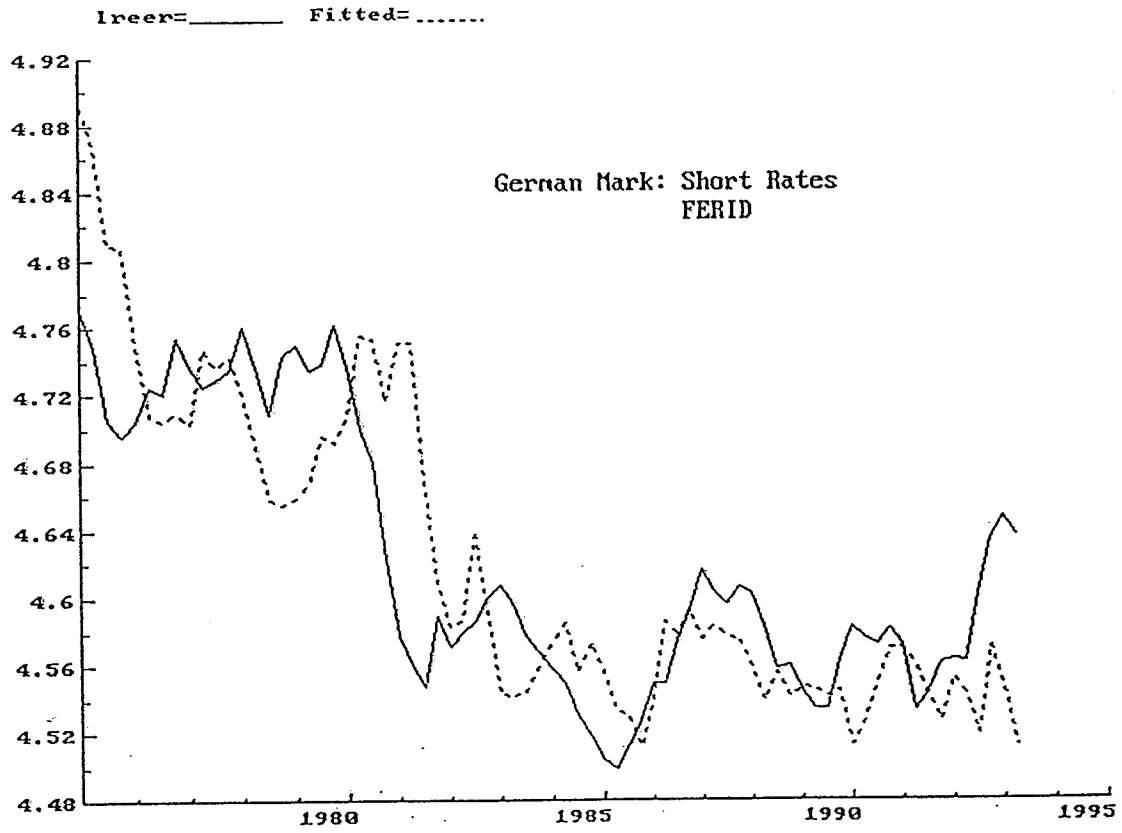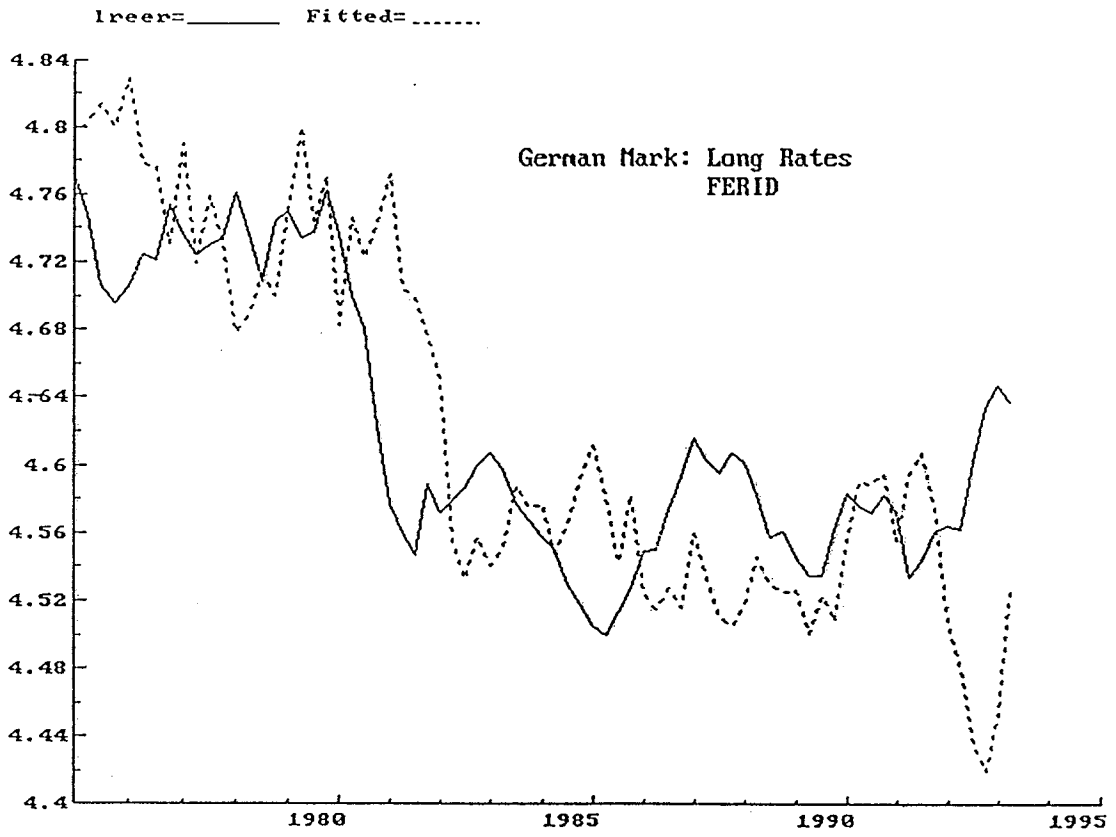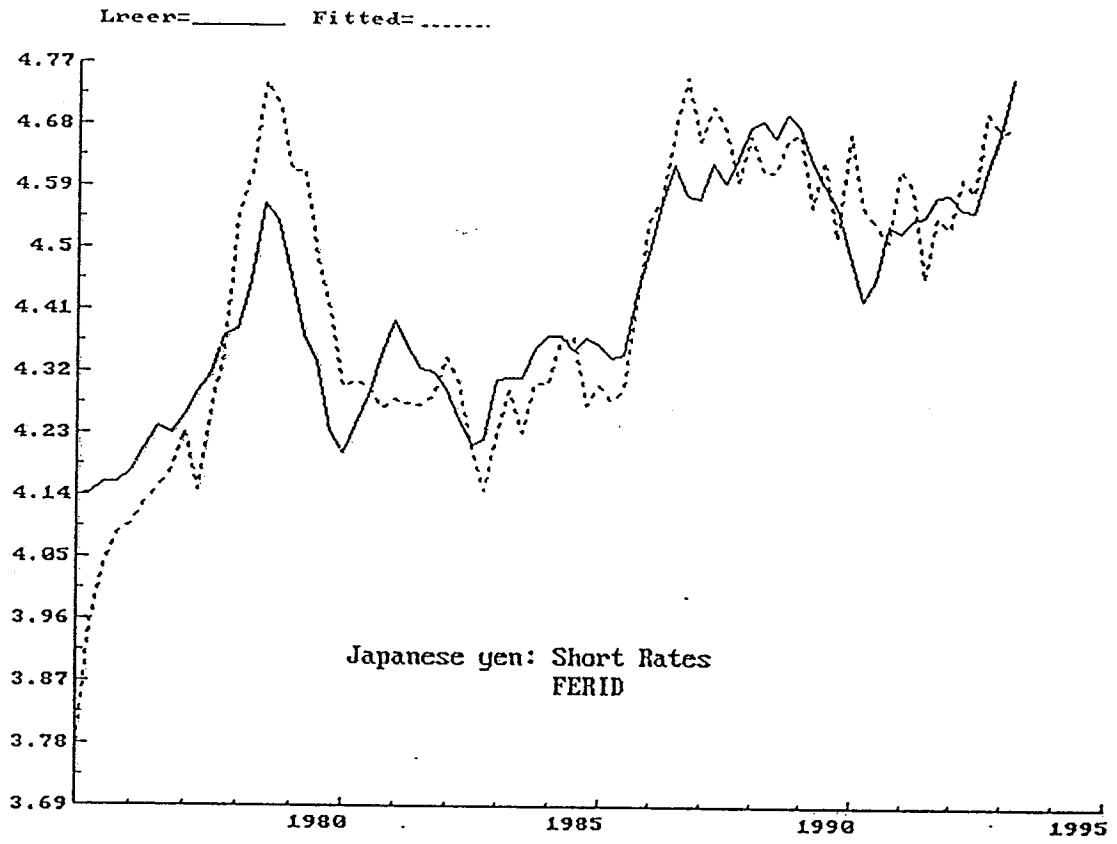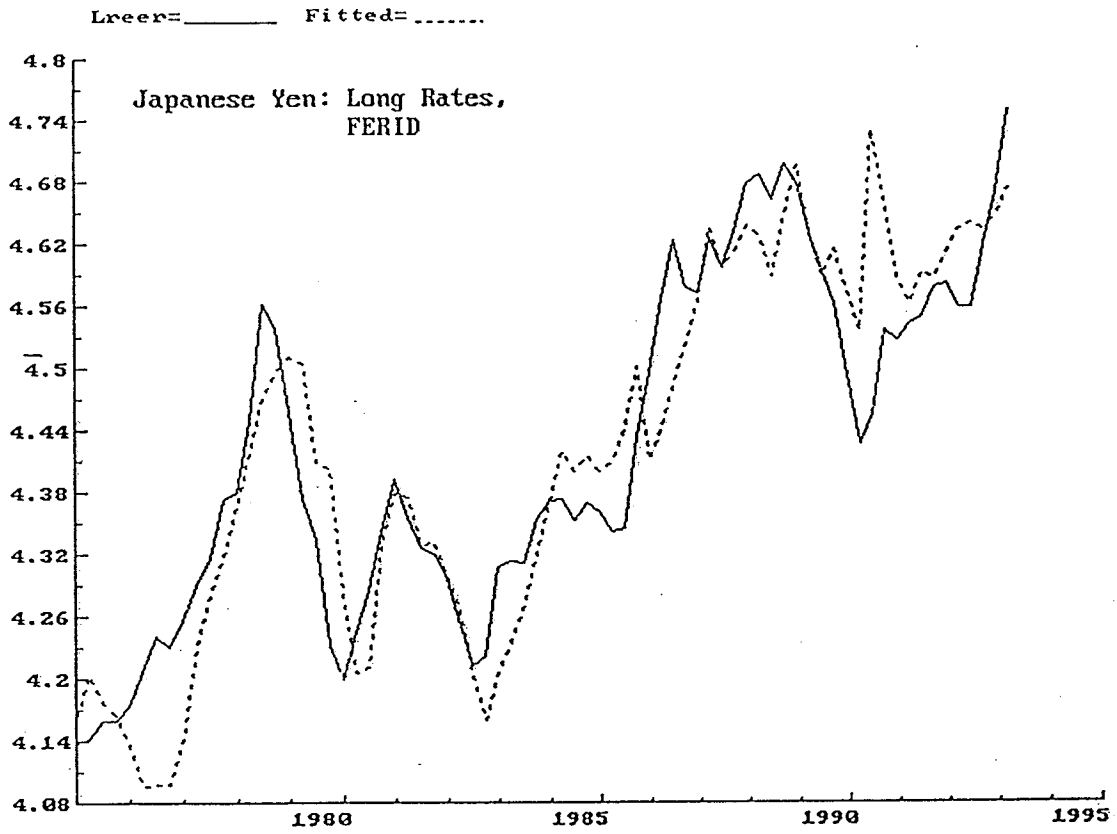
Figure 10

Impulse Responses Of Real Effective Exchange Rate (US)
Long Interest Rate System.
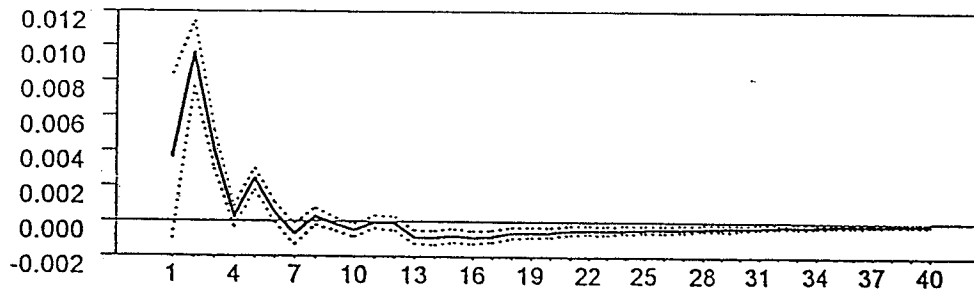
(a.) Orthogonal Shock To Long Run Real Domestic Interest Rate.

(b.) Orthogonal Shock To Long Run Real Partner Interest Rate.
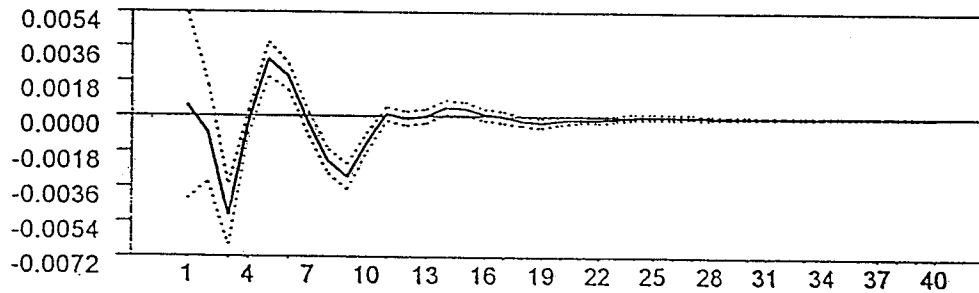
(c.) Orthogonal Shock To Relative Productivity.
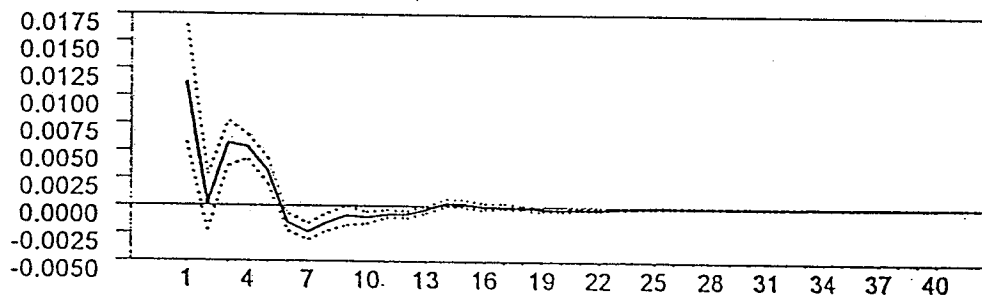
(d.) Orthogonal Shock To Terms Of Trade.
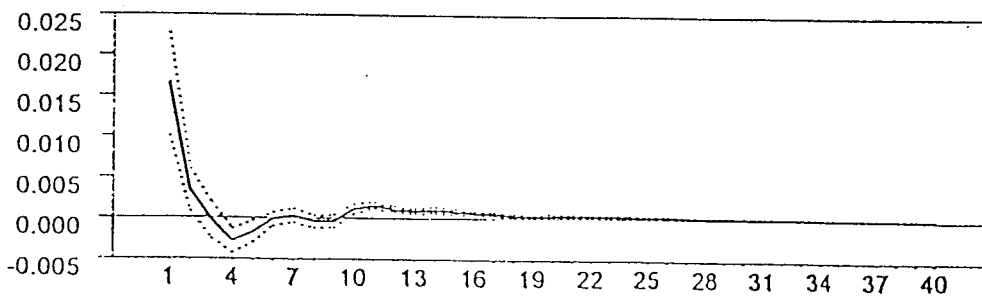
Figure 10 (Concluded)
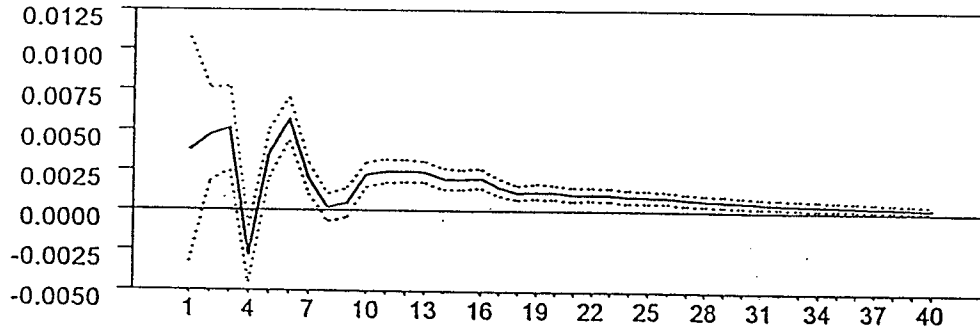
(e.) Orthogonal Shock To Net Foreign Assets.

(f.) Orthogonal Shock To Real Oil Price.
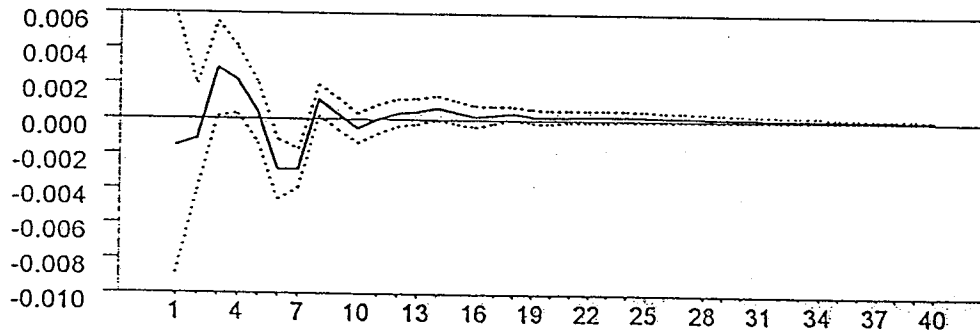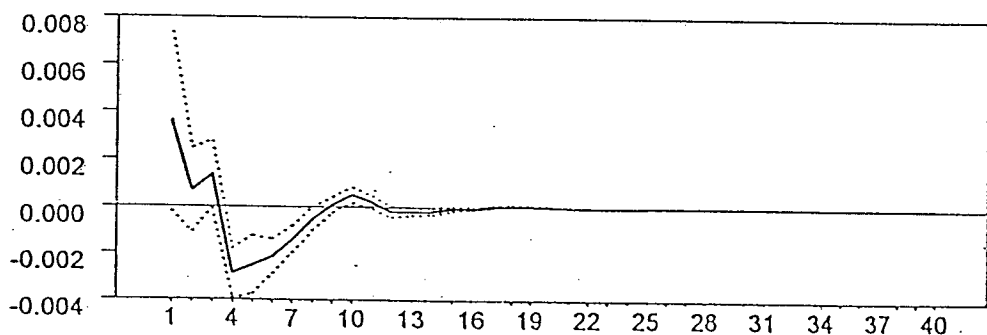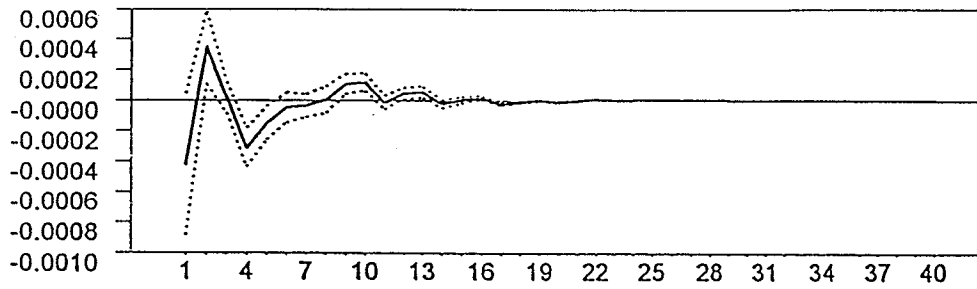

(g.) Orthogonal Shock To Relative Fiscal Balance
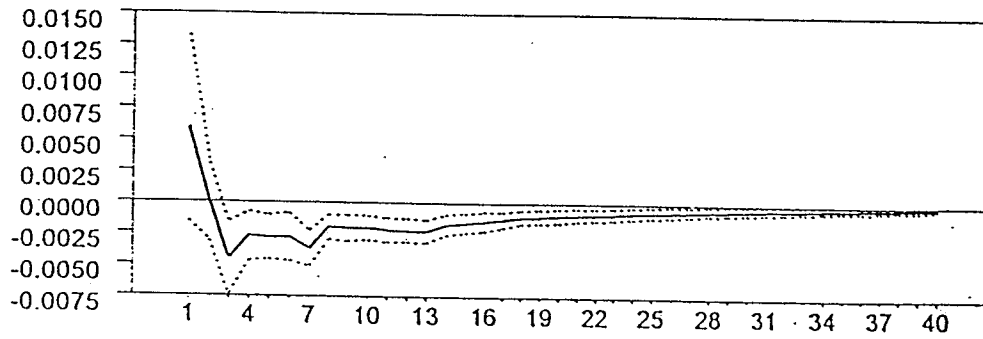
Figure 11

Impulse Responses Of Real Effective Exchange Rate (Germany) Long Interest
Rate System.

(a.) Orthogonal Shock To Long Run Real Interest Rate Differential.
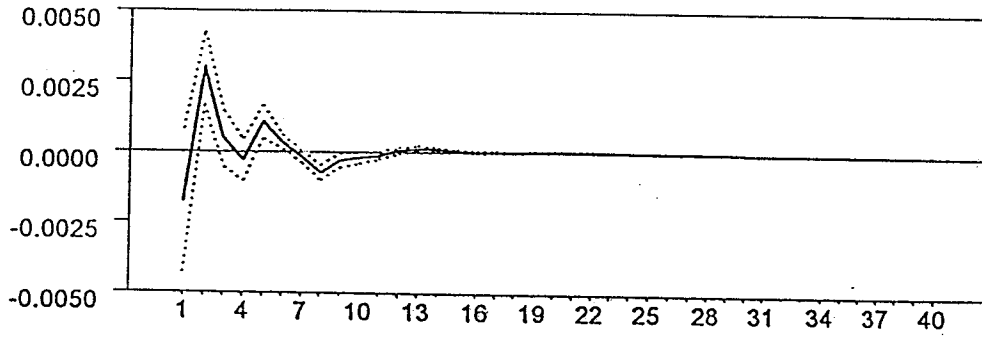
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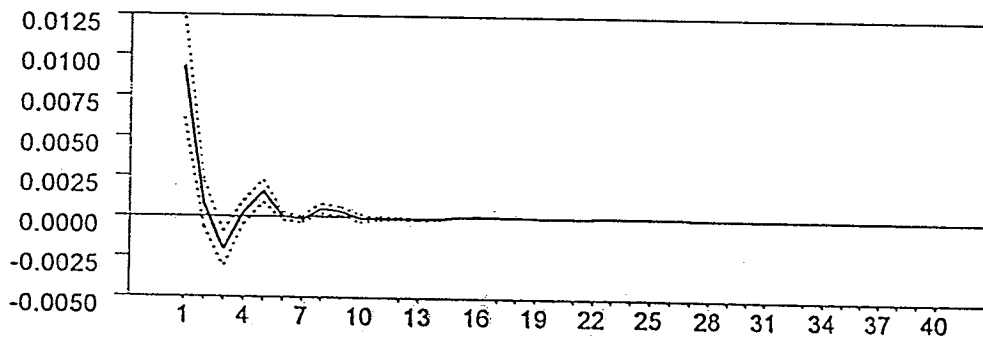
(c.) Orthogonal Shock To Terms Of Trade.
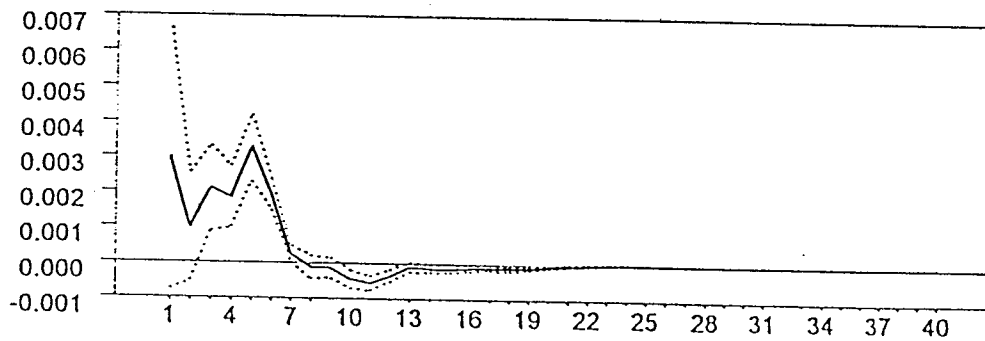
Figure 11 (Concluded)

(d.) Orthogonal Shock To Net Foreign Assets.
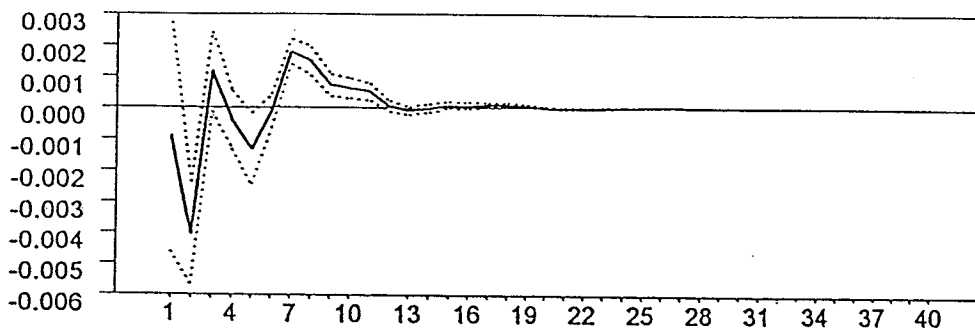
(e.) Orthogonal Shock To Real Oil Price.
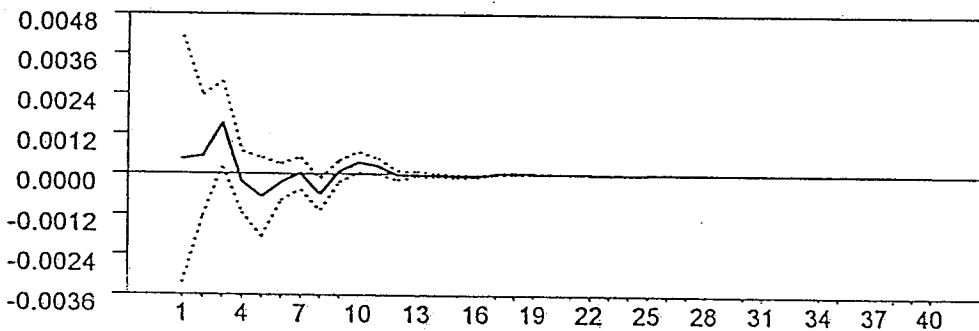
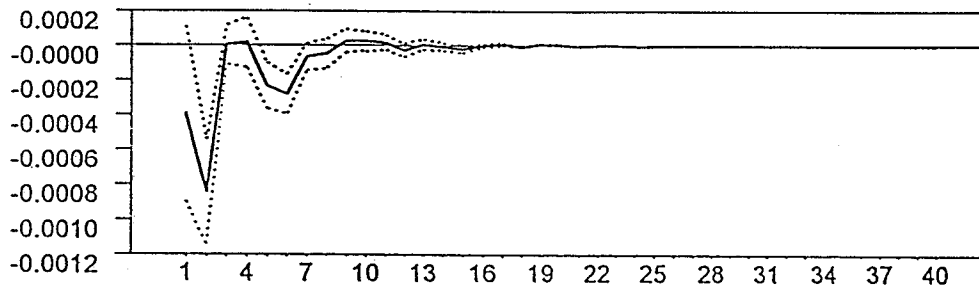(f.) Orthogonal Shock To Relative Fiscal Balance

Figure 12

Impulse Responses Of Real Effective Exchange Rate (Japan)
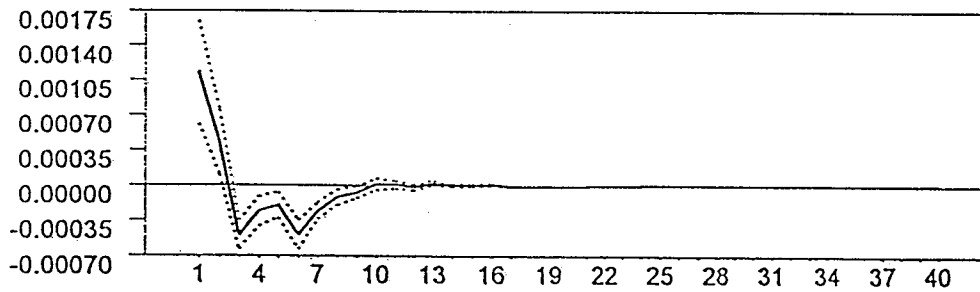Long Interest Rate System.

(a.) Orthogonal Shock To Long Run Real Domestic Interest Rate.

(b.) Orthogonal Shock To Long Run Real Partner Interest Rate.

(c.) Orthogonal Shock To Relative Productivity.
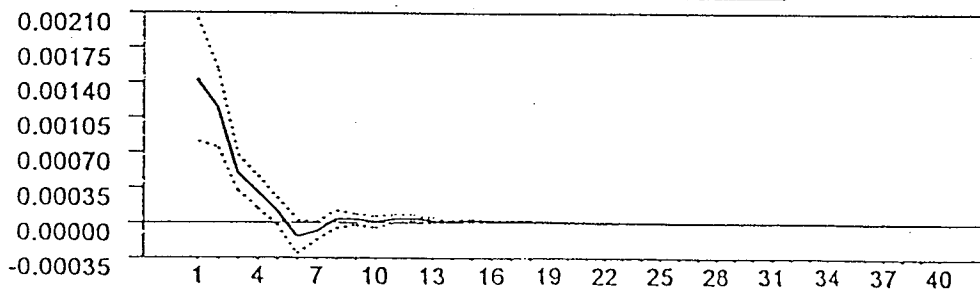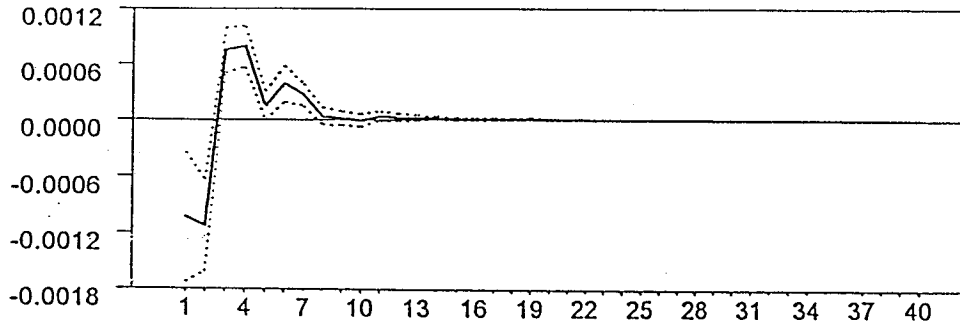

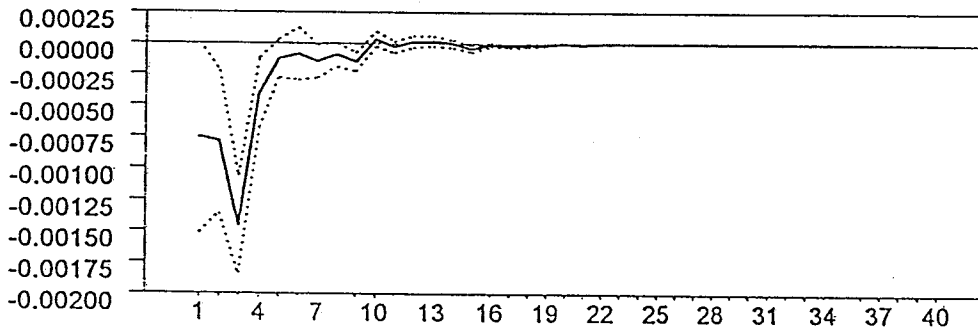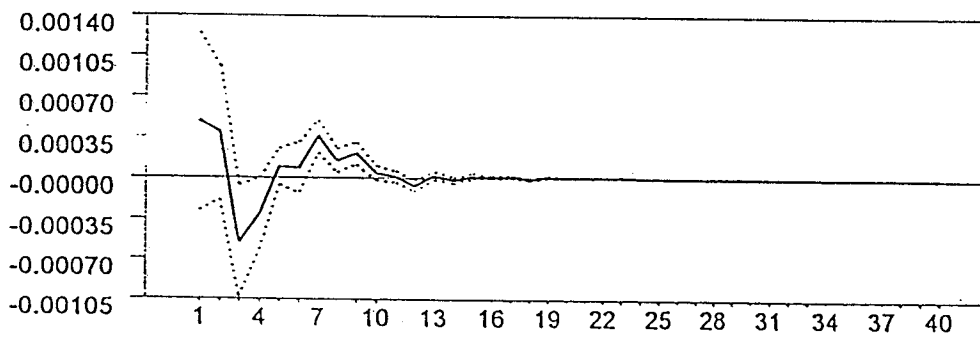(d.) Orthogonal Shock To Terms Of Trade.

Figure 12 (Concluded)

(e.) Orthogonal Shock To Net Foreign Assets.

(f.) Orthogonal Shock To Real Oil Price.

(g.) Orthogonal Shock To Fiscal Balance.

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