Real Innovations and Real U.S. Dollar Exchange Rates

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**Abstract**

This paper examines the relative importance of real shocks in explaining real U.S. dollar exchange rates. Using the equilibrium relationship among the real exchange rate and domestic and foreign interest rates, the number of a priori or untestable restrictions needed to just identify the structural model from the reduced form estimates is reduced from three to just one. The empirical model suggests that real shocks have dominated the behavior of the U.S. dollar real exchange rate versus the Japanese yen and the German mark but that nominal shocks play a significant role in explaining the behavior of the real U.S. - Canadian dollar exchange rate and the real dollar - pound exchange rate.
INTRODUCTION

Since the advent of the modern floating system of exchange rates, theoretical models of exchange rate behavior have been a spectacular failure. The seminal study by Meese and Rogoff (1983) demonstrated that a simple random walk model outperformed several important and popular theoretical models in forecasting nominal exchange rates. Baillie and Pecchenino (1991) conclude that the fundamental flaw in these theoretical models is the assumption that purchasing power parity (PPP) holds at some point in time. Numerous studies over the last two decades document the failure of PPP for bilateral U.S. Dollar exchange rates over the floating regime period (see Dornbusch, 1987).\(^1\) This failure of PPP is particularly troubling for asset-approach models of exchange rate determination since PPP is a fundamental building block of these monetary and portfolio balance models. Many economists, beginning with Stockman (1980), have concluded that the non-stationary behavior of real exchange rates is due to the predominance of real shocks over those that originate in the monetary sector. Real shocks, such as taste, technology and government spending shocks, alter the relative price of international goods and have permanent effects on real exchange rates. Nominal shocks, due primarily to monetary disturbances, should only change the money price of international goods but leave the real price unchanged, at least in the long-run.

The importance of permanent versus transitory shocks has been inferred by some by analyzing the univariate behavior of real exchange rates (see Abuaf and Jorion, 1990, Diebold, et al., (1991), and Liu and He, 1991). This approach can be misleading. Quah (1992) demonstrates that all univariate measures of the size, in Cochrane’s (1988) sense, of the random walk component of a time series are equivalent and equal

\(^1\)There has been some recent evidence that supports PPP by relying on cross-sectional data sets. See Lothian (1995) and
to the spectral density of the changes in the series at frequency zero. The problem is one of identification. In the univariate setting the various innovations impinging on real exchange rates are not statistically separated out from one another. Quah (1992) suggests that the remedy to this underidentification problem is to expand the information set by including other variables that are correlated with the variable of interest.

To shed more light on the various disturbances that affect real exchange rates, a literature has developed in which real and monetary shocks are empirically identified and their relative contribution to exchange rates is examined in a multivariate setting. The evidence is not conclusive. Clarida and Gali (1994) report significant nominal shock effects on the real exchange rate. The abstract of their paper reads, "We find that demand shocks,..., explain the majority of the variance in real exchange rate fluctuations, while supply shocks explain very little." Lastrapes (1992) finds that the supply or real shocks dominate the real U.S. dollar exchange rate. Evans and Lothian (1993) come down somewhere in the middle acknowledging that while nominal shocks play a significant role they are dominated by real innovations to the real exchange rate.

This study contributes to the current debate by re-examining U.S. dollar real exchange rates versus the U.K. pound, Canadian dollar, German mark, and Japanese yen. The identification of the structural innovations from the reduced form errors differs substantially from the prior works. A simple equilibrium model of the balance of payments is used to derive implications about the cointegrability of the real exchange rate and the nominal interest rates in the two countries that comprise each bilateral relationship versus the U.S. dollar. This model has been analyzed by Johansen and Juselius (1992) and Crowder (1996) for the U.S. - U.K. relationship. The evidence suggests that the system in each bilateral relationship is driven by two permanent innovations or common (stochastic) trends and one transitory innovation
or equilibrium error. The implication is that the (stochastic) trend in each of the three variables in each system has at most two sources. The identification of the real permanent shock is achieved by resorting to a long-run neutrality restriction operationalized by restricting the nominal permanent shock to have no long-run impact on the real exchange rate. This identification is similar to that used by Blanchard and Quah (1989), Lastrapes (1992), and Clarida and Gali (1994), but differs in the treatment of the long run equilibrium relationship among the three variables in each system.\footnote{The existence of cointegration actually implies certain restrictions that aid in the structural identification of the model. See King, et al. (1991), Warne (1993) and Crowder, et al. (1994).}

The results are not consistent across all four bilateral real exchange rates. The real shock dominates the nominal shocks at all horizons and for the entire floating exchange rate regime in the dollar-yen and dollar-mark systems, but nominal shocks play an important role in explaining the short-to-medium-term movements in the U.S. - Canadian dollar and dollar-pound real exchange rates.\footnote{Although there are three structural innovations in each system, only two leave permanent imprints on the data. The third is a purely transitory shock. Since the focus here is on the source of the stochastic trend in real exchange rates, the discussion centers around the two permanent innovations that I have dichotomized into real and nominal. It is possible that the transitory shock is also real in nature, but since it doesn’t have a permanent effect on any of the variables, including the real exchange rate, the source of this shock need not be economically identified.} This result is robust to the alternative (and counterintuitive) long run identification restriction, i.e., that real shocks have no long run impact on the real exchange rate.

The next section briefly outlines the simple model of balance of payments equilibrium that serves as a basis for the choice variables to be used in the empirical analysis. Section 3 discusses the econometric procedures and the specific identification issues. Section 4 presents the empirical results and section 5 concludes.
A SIMPLE MODEL OF THE REAL EXCHANGE RATE

This section presents a very simple model of bilateral balance of payments equilibrium which implies a long run equilibrium relationship between the real exchange rate and nominal interest rates on domestic and foreign bonds. This model is not intended to be the focus of the empirical analysis. It is presented only to give theoretical justification to the possibility that the real exchange rate and nominal interest rates are cointegrated.4

The balance of payments (BOP) under a floating exchange rate regime is given as \( C_t + K_t = 0 \), where \( C_t \) is the current account balance and \( K_t \) is the capital account balance. In this model the current account is predominantly a function of the real exchange rate as in equation (1).

\[
C_t = a(s_t - p_t + p_t^*)
\]  

(1)

The variable \( s_t \) is the log of the spot exchange rate and \( p_t \) and \( p_t^* \) are the logs of the domestic and foreign price levels, respectively. The capital account, \( K_t \), is modelled as a function of deviations from uncovered interest parity as in equation (2).

\[
K_t = \gamma(i_t - i_t^* - \Delta s_{t+k}^e)
\]  

(2)

The domestic interest rate on a k-period bond is denoted by \( i_t \) and the foreign interest rate is \( i_t^* \). The term \( \Delta s_{t+k}^e \) represents expected nominal exchange rate depreciation. Using the BOP equilibrium to combine equations (1) and (2) yields,

4The more common hypothesized relationship between the real exchange rate and real interest rates is somewhat problematical in the context of a cointegrated system since real interest rates appear to be stationary across the industrialized countries. Evidence supporting this can be found in Crowder and Hoffman (1996), Mishkin (1992) and Crowder (1996). Furthermore, Meese and Rogoff (1988) and Edison and Pauls (1993) demonstrate that there is very little evidence of a systematic relationship between real exchange rates and real interest differentials.
\[ \rho_t = -\left(\gamma/a\right)(\hat{t} - \hat{t}^* - \Delta \hat{s}_{t+k}^e) \]  
where \( \rho_t = s_t - p_t + \hat{p}_t^* \).

Since the expected spot rate depreciation is stationary, equation (3) represents a long run equilibrium relationship between the real exchange rate and the domestic and foreign interest rates. The next section discusses how this long run equilibrium can be exploited to achieve structural identification of the shocks or innovations driving the three variables.

**IDENTIFICATION OF A STRUCTURAL MODEL**

The identification of a structural model from the dynamic reduced form has been extensively discussed since Sims (1980) first proposed the use of vector autoregressions (VARs) in analyzing the dynamic relationship among time series variables. The issues surrounding structural identification can be understood by specifying a reduced form VAR in (4),

\[ \Phi(L)X_t = \mu + \varepsilon_t \]  
where \( X_t \) is a \( p \times 1 \) vector of stationary time series variables, \( \Phi(L) \) is a matrix polynomial in the lag operator where \( \Phi_0 = I \), \( \mu \) is the constant and \( \varepsilon_t \) is the reduced form error with covariance matrix \( \Omega \). A corresponding structural model is assumed to exist as in (5),

\[ A(L)X_t = \delta + \nu_t \]  

\(^5\)The relevant citations include, but are not limited to, Bernanke (1986), Sims (1986), Shapiro and Watson (1988), Blanchard and Quah (1989), and Warn (1993).
where \( E[\nu_1\nu_2] = I \). The relationship between the structural and reduced form parameters is \( \Phi(L) = A_0^{-1}A(L) \), \( \mu = A_0^{-1}\delta \), and \( A_0\Omega A_0' = I \). Identification, in general, would require \( p^2 \) independent restrictions on (4). Arbitrary normalization of each VAR equation provides \( p \) restrictions and the assumption of structural error independence provides another \( p(p - 1)/2 \) restrictions. From this it is clear that exact identification of the structural model depends upon \( p(p - 1)/2 \) additional restrictions to be imposed by the econometrician. These restrictions usually take the form of restrictions on the \( A_0 \) matrix.

Sims (1980) original specification made \( A_0 \) a lower triangular matrix implying a contemporaneous recursive structure to the model. Bernanke (1986) and Sims (1986) independently suggested applying non-recursive exclusion and/or general restrictions on \( A_0 \) that were more justifiable based upon economic reasoning. Shapiro and Watson (1988) and Blanchard and Quah (1989) used long-run restrictions to identify the structural model from the reduced form estimates. This is achieved by noting that the VARs in (4) and (5) have invertible Wold moving average representations (MAR). The reduced form MAR is given by,

\[
X_t = \zeta + C(L)\epsilon_t
\]  

where \( C(L) = \Phi(L)^{-1} \) and \( \zeta = \Phi(L)^{-1}\mu \). The corresponding structural MAR is related to the reduced form MAR by noting the \( D(L) \), the structural MA matrix polynomial, is equal to \( C(L)A_0^{-1} \). Furthermore, the long-run structural impact matrix is related to the reduced form by \( D(1) = C(1)A_0^{-1} \). General restrictions on the \( D(1) \) matrix can be imposed by a suitable choice for \( A_0 \). These are then interpreted as long-run as opposed to contemporaneous restrictions on the interactions of the variables in \( X_t \). Blanchard and Quah (1989) argue that economic theory provides stronger implications for the long-run interaction among macro aggregates than it does about the contemporaneous relationship between these variables. Such "neu-
trality” restrictions are more justifiable on theoretical grounds.

Note that for the long-run identification restrictions to be binding, some of the variables in $X_t$ must be the first differences of non-stationary variables. That is the $C(1)$ matrix must not vanish for these type of restrictions to be relevant. This issue is important since many important examples exist in macroeconomics and finance where the $C(1)$ matrix is not of full rank. This implies that only a subset of the innovations have a permanent impact on the variables. The other innovations have only transitory effects and cannot be identified from restrictions on $C(1)$. A relevant situation is when the variables of interest are cointegrated. The theoretical model presented in section 2 implies that $p = 3$ and that the number of additional a priori restrictions needed to achieve exact identification is three. Assuming that the variables are I(1), the model also implies one cointegrating relation among the three variables. This means that the rank of $C(1)$ is two or that only two of the three innovations leaves permanent imprints on the data. Furthermore, $C(1)$ must satisfy the restrictions implied by cointegration$^6$, namely that $\beta'C(1) = 0$ and $C(1)\alpha = 0$, where $\alpha$ and $\beta$ are $p \times 1$ full rank matrices such that $\alpha\beta' = \Phi(1)$ from (4). These imply two restrictions on $C(1)$ that help aid in the identification of $D(1)$.

There are two implications from the foregoing discussion. The first and most obvious is that the knowledge of cointegration rank and space, in general, will yield restrictions that reduce the number of a priori restrictions needed to achieve exact identification of the structural model. The second, and more subtle implication is that existence of cointegration among the variables reduces the set of possible restrictions that the econometrician can legitimately impose. For example, in the three variable case to be analyzed in the next section, there exists one cointegrating relation. Therefore, no long-run neutrality restrictions can be imposed on the third structural innovation since it is already neutral, with respect to the variables in the

system, by definition. That is, such a restriction is not binding nor independent of the cointegrating restrictions. In general then, the econometrician needs to combine both long-run and contemporaneous restrictions to achieve exact identification of the structural model. Specifically, when there exists \( r \) cointegrating relations among \( p \) variables, there will be \( k = p - r \) permanent innovations that can be identified via long-run restrictions and \( r \) transitory innovations that can only be identified using short run or contemporaneous restrictions. The restrictions on \( C(1) \) implied by the cointegrating relations yields \( kr \) independent restrictions leaving just \( p(p - 1)/2 - kr \) independent \( a \) priori restrictions needed to just identify the model. Of these \( k(k-1)/2 \) should be long-run restrictions and \( r(r - 1)/2 \) should be contemporaneous or short run restrictions.

**EMPIRICAL RESULTS**

The data used in this study represent monthly observations beginning in January 1975 and ending in October 1993 on the U.S. dollar spot exchange rate vis-à-vis The British pound, the Canadian dollar, the German mark and the Japanese yen and the consumer price index from each of these four countries. For the U.S. and Canada,

\footnote{Crowder (1995a) demonstrates that Blanchard and Quah’s original model can be represented as a cointegrated system and that the long run neutrality restriction they impose is actually over-identifying. The test of the over-identifying restriction is not rejected in the Blanchard and Quah data but is rejected when an updated sample is used. This example demonstrates the special case where some of the variables in a system are I(1) and others are I(0). This system can (and probably should) be represented as a cointegrated system, albeit a trivial form of cointegration. Specifically, the result that the system is driven by a reduced number of (stochastic) trends must be accounted for in the identification stage.}

\footnote{Gali (1992) does this but ignores the cointegrating relationship between real M1, income and nominal interest rates and is thus still subject to the problem discussed. See Crowder, et al. (1995) for an analysis that properly accounts for the Fisher and equilibrium money demand relations to identify a structural model.}
the interest rates are three month yields on certificates of deposit. The U.K. interest rate is the three-month IBOR. For Germany the interest rate used is the three month FIBOR. There is no measure of the interest rate for Japan that is consistent with the interest rates used for the other countries over the entire sample.\(^9\) From 1975:1 to 1979:4 the three month call rate was used. From 1979:5 to 1993:10 the three month commercial deposit rate is used. All of the data are taken from the OECD Main Economic Indicators. The data are plotted in figure 1.

A battery of univariate unit root tests indicate that each of the series are integrated of order one.\(^10\) These tests have been shown to be biased in the presence of structural breaks by Perron (1989). Zivot and Andrews (1992) propose a recursive testing procedure that endogenously chooses a break point, if one exists, and tests the unit root hypothesis. The results for these tests are presented in figure 2, where the test statistics have been normalized by the appropriate critical value meaning that values greater than one imply statistical significance at the 5% level. There is no evidence that any of the series are stationary even when possible structural breaks are accounted for.

The estimation and testing of the long-run equilibrium (cointegrating) relationship in each BOP system is accomplished by using the Johansen (1988) procedure.\(^11\) Table 1 presents the summary statistics from the cointegration analysis.\(^12\) There is evidence

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\(^9\)The call money rate is available over the entire sample, but is not entirely consistent with the other countries interest rates. It turns out that the Japanese call rate and the commercial deposit rate are cointegrated with vector \([1, -1]'\) over the period which they overlap. Furthermore, the results do not change when the call rate is used over the entire period.

\(^10\)Theses are available upon request.

\(^11\)The methodology is well known, so I will forgo a discussion here. The uninitiated reader is referred to the citations listed in the references.

\(^12\)The lag length of the VECM was set at seven in each case as this was the minimum lag length necessary to eliminate serial correlation in the errors as measured by the LR test based upon 24 autocorrelations. The residuals did display strong evidence of non-normality in the interest rate.
of one cointegration vector in each system, although the evidence is weak for all but the U.S.-Japanese relationship. The normalized estimates of the cointegration vectors are given in table 2. Figure 3 plots each of the estimated equilibrium relationships. Univariate unit root tests on these estimated equilibria reject the null hypothesis at the 10% level of significance or higher.

The existence of cointegration implies that the $\Phi(L)$ matrix has a singularity which complicates the inversion to yield the MAR of equation (6). The MAR for each system is computed by noting that $C(1) = \beta_1 (\alpha_1 \Phi^*(1) \beta_1)^{-1} \alpha_1$ (see Johansen, 1991), where $\beta_1 \beta_1 = 0$, $\alpha_1 \alpha_1 = 0$, $\Phi^*(1) = \Phi^*_1 + \ldots + \Phi^*_k$ and $\Phi^*(L) = (\Phi(L) - \Phi(1))(1 - L)^{-1}$ from equation (4). Johansen (1991) has given $\alpha_1$ the interpretation of the common trends, as in Stock and Watson (1988), and $\beta_1 (\alpha_1 \Phi^*(1) \beta_1)^{-1}$ is interpreted as the factor loadings or how the common trends are passed on to the variables in the system. From this form of $C(1)$ it can be seen that restrictions on $C(1)$ can be imposed either through restrictions on the common trends, $\alpha_1$, or the factor loadings, $\beta_1 (\alpha_1 \Phi^*(1) \beta_1)^{-1}$. The structural total impact matrix $D(1)$ can be written as $\beta_1^0 (\alpha_1 \Phi^*(1) \beta_1)^{-1} \alpha_1$ where $\beta_1^0 = A_0 \beta_1$. Thus the identifying restrictions can be imposed directly on $\beta_1$. Since there are two common trends or $k = 2$, one a priori restriction must be imposed to achieve identification of the long-run components. But, since $r = 1$, i.e., there is only

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13 The usual Dickey-Fuller critical values are no longer valid as demonstrated by Engle and Granger (1987). The appropriate critical values can be found in McKinnon (1991).

14 The choice of $\beta_1^0$ must also satisfy the condition that the structural errors are independent. This is ensured when $(\beta_1^0 \beta_1^*)^{-1} \beta_1^* \Omega C(1) \beta_1 (\beta_1^0 \beta_1^*)^{-1}$ is diagonal, where $\beta_1^* = \beta_1^0 (\alpha_1 \Phi^*(1) \beta_1^0)^{-1}$. Taking the Cholesky factor of this matrix and multiplying it by $\beta_1^0$ will achieve this independence requirement.
one cointegrating vector, the assumption of structural error independence is enough to identify the short run components of the model. No extra restrictions are necessary. The existence of cointegration has supplied two of the three extra restrictions needed to exactly identify the structural model.

The variables in each system are ordered as $X_t = [\rho_t, i_t, i_t^*]'$. The single extra identification restriction imposed is that the second permanent innovation, i.e., the structural innovation to the U.S. nominal interest rate, has no long-run effect on the real exchange rate. This restriction suggests interpreting structural shocks to the real exchange rate as real permanent innovations and structural shocks to the U.S. nominal interest rate as nominal permanent innovations. This interpretation further suggests that a common nominal trend among these interest rates may be traced back to the U.S. Crowder (1996) presents evidence that a common trend links the G-7 inflation rates over the post-war period. Extending the results among G-7 inflation rates to nominal interest rates via a Fisher (1930) relation provides some support for this interpretation of the nominal permanent shock in each system in the current study. Given this economic interpretation, the restriction is sensible in that the real exchange rate is a relative price and should not be influenced permanently by purely nominal changes. This restriction does not prevent real exchange rates from a rich dynamic response to nominal shocks in the short to intermediate term. Figure 4 plots each series and their fitted common trend. Note that the trend terms more closely follow the real exchange rate and the U.S. nominal interest rate suggesting that the economic interpretation given the two common trends in each system is consistent with the data.

The restriction is most easily imposed by computing a general form for $\beta_1$ (using, for instance, the generalized inverse of $\beta$) and setting the first two columns of $A_0 = [(I_k * \beta_1)^{-1:0}]$, where $I_k$ is the $k \times k$ identity matrix.\footnote{The third column of $A_0$ is restricted by the assumption that the permanent and transitory}
system, the estimate of the cointegration vector is \( \hat{\beta} = [1.00 \ 0.08 \ -0.09]' \), with a corresponding restricted orthogonal complement given by \( \beta^0_\perp = \begin{bmatrix} 1.00 & 0.00 \\ 0.00 & 1.00 \\ 11.11 & 0.89 \end{bmatrix} \)\(^{16}\).

The zero in the \( \beta^0_\perp (1, 2) \) element is identifying, while that in the \( \beta^0_\perp (2, 1) \) element is not. When the orthogonality of the errors restriction is imposed (see note 11), the zero in the \( \beta^0_\perp (1, 2) \) element is preserved while that in the \( \beta^0_\perp (2, 1) \) element is not. This lone zero restriction on common factor loading matrix, which implies that innovations in the second common trend have no long-run impact on the real exchange rate, is enough to give the model a structural interpretation.

The dynamic analysis of the structural models yields impulse response functions (IRFs), forecast error variance decompositions (VDCs) and historical residual decompositions (HDCs). Figure 6 displays the impulse responses from the U.S. - Canada structural system along with simulated 90% confidence bounds.\(^{17}\) Note that the real shock has no statistically significant effect on the U.S. interest rate and only a very short-lived effect on the Canadian rate occurring about six months after the initial impulse. The real exchange rate has very small but significant responses to both the permanent nominal shock and the transitory shock. These results suggest that although the long-run equilibrium among the variables is weak, there is information contained in the two interest rates that is relevant for describing the real exchange

\(^{16}\)Although the two interest rate parameters are small they are statistically significant. Exclusion tests on these two parameters yields \( \chi^2(2) \) statistics of 18.06 and 15.86, respectively.

\(^{17}\)It should be noted that the simulated confidence bounds assume that the underlying data processes are normally distributed. When this assumption is being violated the interpretation of the confidence intervals must be viewed as only an approximation to the true confidence boundaries.
rate process. Furthermore, it implies that there is a significant role for nominal shocks in explaining the behavior of the U.S. - Canadian dollar real exchange rate. Figures 7 and 8 display the impulse responses for the U.S. - Germany and U.S. - Japan systems, respectively. These are similar to the patterns exhibited in the U.S. - Canada system, but differ in one very important respect. Neither the permanent nominal innovation nor the transitory innovation has any statistically significant effect on the real exchange rate in either system. For these two real exchange rates the evidence suggests that real shocks dominate the effects of nominal shocks almost completely. On the other hand figure 9, which displays the results for the U.S. - U.K. system, suggests a significant role for nominal shocks for this real exchange rate, very similar to the U.S. - Canada system.

Table 3 presents the h-step ahead forecast error variance decomposition along with simulated 90% confidence intervals. The forecast error variance at increasing horizons is decomposed into the variance attributable to each of the structural innovations. In the Canada and U.K. systems, non-real shocks, i.e., the permanent nominal and transitory innovations, explain almost 40% of the forecast error variance at the one-month horizon. In stark contrast, the Japan and German systems yield less than 10% of the variance to these two non-real innovations. This dichotomy is revealed in examining the confidence intervals of the four systems. At the one-month horizon in the Canada and U.K. systems the real innovation may explain as little as 8% of the variance in the forecast error. While in the Japan and German systems the minimum explained by real shocks is 71% and 43%, respectively.

Figures 10 through 13 display the historical decomposition of the one-step ahead forecast error into the component attributable to the real shock (upper panel) and the component attributable to the permanent nominal shock and transitory shock (lower panel) for the Canadian dollar, German mark, Japanese yen and U.K. pound real exchange rates versus the U.S. dollar, respectively. The scales on each of the
two panels is kept constant to facilitate the comparison. Clearly, the real shock has accounted for the overwhelming majority of the deviations of the real exchange rates from their respective equilibrium paths for the dollar-yen and dollar-mark exchange rates. The HDCs in figure 13 for the dollar-pound real exchange rate clearly suggest that nominal and/or transitory innovations have played a very important role in explaining this real asset price. This appears to be especially true in the late 1970s to early 1980s and again in the early 1990s.

CONCLUSION

In this study I have used a simple model of the balance of payments to provide an equilibrium relationship between the real exchange rate and domestic and foreign interest rates. The interest rates provide information about the nominal shocks impinging on the real exchange rate. But, theoretically only real shocks should have important long-run effects on the real exchange rate. Using this idea and the equilibrium relations among the variables, complete structural identification of the statistical model is achieved. The existence of the cointegrating relationship among the variables allows the incorporation of additional information via the increased dimension of the VAR without the need to impose even more identification restrictions than has been done in the previous literature.

The results imply that real U.S. dollar exchange rates versus the Japanese yen and German mark have been dominated by real shocks over the modern floating exchange rate regime. In contrast, the real U.S. dollar exchange rates relative to the Canadian dollar and U.K. pound show evidence of significant effects from nominal innovations. Therefore one may reject the random walk hypothesis for these two real asset prices. Even so, in the long-run all four real exchange rates appear to be dominated by real shocks which may help to explain the empirical failure of monetary and other asset approaches in the determination of exchange rates.
An important consideration is the extent to which the simple model used in this analysis is dominated by other models at capturing the nominal shocks that are so crucial to the underlying asset approach theories. Certainly more complex models exist and they may prove to be more adept at explaining the behavior of real exchange rates. The one drawback to using more variables in an expanded model is that the restrictions that must be imposed to achieve identification become more difficult to justify on theoretical grounds. Recently MacDonald and Taylor (1994) have found that the simple monetary model outperforms various important measures of forecasting accuracy when the cointegrating relations implied by the model are imposed. This may be a fruitful avenue to explore in future research.
REFERENCES


Table 1. Trace Tests for the Number of Cointegrating Relations

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Conventional critical values

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<td>90 %</td>
<td>26.79 13.33 2.69</td>
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Note: Critical values are taken from Osterwald–Lenum (1992).
Table 2. Estimates of the Normalized Cointegration Space

\[ x_t = \begin{bmatrix} \rho_t & i_t & i_t^* \end{bmatrix} \]

U.S. - Canada

\[ \hat{\beta}_t = \begin{bmatrix} 1.000 & 0.080 & -0.089 \end{bmatrix} \]

U.S. - Japan

\[ \hat{\beta}_t = \begin{bmatrix} 1.000 & -0.055 & 0.373 \end{bmatrix} \]

U.S. - Germany

\[ \hat{\beta}_t = \begin{bmatrix} 1.000 & 0.001 & -0.150 \end{bmatrix} \]

U.S. - U.K.

\[ \hat{\beta}_t = \begin{bmatrix} 1.000 & 0.088 & -0.158 \end{bmatrix} \]
Table 3. Forecast Error Variance Decompositions

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<td>94.78</td>
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<td>(95.34, 46.84)</td>
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Note: The restriction used to separate the effects of the two orthogonal permanent innovations is that Prm2 has no long-run effect on the real exchange rate. Monte Carlo simulated standard errors are in parentheses below each estimate.
FIG. 1. Data Used in the Analysis

FIG. 2. Zivot-Andrews Unit Root Tests
**Fig. 3. Estimated Cointegrating Vectors**

**Fig. 4. Canadian and Japanese Data with Permanent Components**
Fig. 5. German and UK Data with Permanent Components

Fig. 6. Impulse Responses from the US-Canada System
Fig. 7. Impulse Responses from US-German System

Fig. 8. Impulse Responses from US-Japan System
**Fig. 9.** Impulse Responses from US-UK System

**Fig. 10.** Historical Decomposition of Real US-Canada Exchange Rate
**Fig. 11. Historical Decomposition of Real US-Japan Exchange Rate**

**Fig. 12. Historical Decomposition of Real US-German Exchange Rate**
FIG. 13. Historical Decomposition of Real US-UK Exchange Rate