The Importance of Real Innovations in Explaining Real U.S. Dollar Exchange Rates: Evidence from the Yen-Dollar Exchange Rate

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1Very preliminary. Quote at your own risk.
Abstract

This paper examines the relative importance of real shocks in explaining real U.S. dollar exchange rates. The apparent unit root in real exchange rates over the modern floating regime is quite puzzling. If nominal perturbations are the norm then these shocks should leave no permanent imprint on the real exchange rate. If real shocks are the dominant source of economic fluctuations, then real exchange rates may be permanently altered to the extent that relative prices change across countries. This would be true even if the law of one price held for all traded goods. I use an equilibrium relationship among the real exchange rate, real domestic income, real foreign income and domestic and foreign interest rates derived from the open economy version of the IS-LM model to examine the importance of real shocks in the determination of the real yen-dollar exchange rate. By imposing three estimated cointegrating relations on the system, the number of a priori or untestable restrictions needed to just identify the structural model from the reduced form estimates is reduced from ten to four. The empirical model suggests that real shocks have dominated the behavior of the U.S. dollar real exchange rate versus the Japanese yen.
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) is a valid description of the joint behavior of prices and nominal exchange rates. 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 and Rogoff , 1996 for surveys of this literature).

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.

Asset models of exchange rate determination aside, economists are loathe to part with such a simple and intuitively appealing theory as PPP as it embodies many ideas that economists hold near and dear, i.e., rational expectations, long-run money neutrality, efficient markets, etc. This may explain the continued effort by economists to resuscitate PPP. The effort has yielded some successes.
The evidence is much more favorable to PPP when very long spans of data are employed. For example, Crowder (1996), Lothian and Taylor (1997) and Diebold, et al. (1991) all present evidence that supports PPP when a century or more of data are used in the analysis. This is thought to be the result of nominal disturbances dominating real shocks over long time periods. There has also been some recent evidence that supports PPP over the modern float by relying on panel data sets, i.e., Abuaf and Jorion (1990) and Lothian (1994). Panel tests have greater power asymptotically than conventional unit root tests.

The studies that find support for PPP are, however, not without problems. For instance, when using long spans of data several different exchange rate regimes are embodied in the one sample period. How changes in the regimes affect tests for PPP has not been studied in detail. One possibility is that changes in regimes create additive outliers in the real exchange rate series that bias the tests in favor of PPP. Panel studies that find support for PPP have only just become subjected to the scrutiny that standard unit root tests have undergone for the last decade. Very little is known about the small sample properties of the panel estimators in the context of unit roots. Issues of model specification and size distortion in the tests seem to be good starting points for the sort of critical inquiry necessary to validate the panel estimators promise.

The underlying empirical issue is a simple one. Do real exchange rates have constant first moments? Are real exchange rates mean stationary processes?\(^1\) If we are to answer this question in the negative, a natural follow-up would be, "How important is the non-stationarity in real exchange rates?" Another might be, "What are the sources of the time dependency in the mean of real exchange rates?" These two

\(^1\)The distinction between mean stationary or first moment stationary is in contrast to the pure definition of a stationary series as one that has all moments that are time invariant. PPP has nothing to say about moments of real exchange rates beyond the first.
questions relate to the issue of permanent versus transitory components of economic
time series. Such a dichotomy arises in the context of non-stationary time series since
for such series some innovations have permanent impacts on the level of the series.

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
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 ex-
change rates versus the Japanese yen. The identification of the structural innovations
from the reduced form errors differs substantially from the prior work. A simple open economy IS-LM model is used to motivate the empirical model employed. The empirical model employs a system which includes the real exchange rate, real income and nominal interest rates in the two countries. The evidence suggests that the system in each bilateral relationship is driven by two permanent innovations or common (stochastic) trends and three transitory innovations. The implication is that the (stochastic) trend in each of the five variables 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 variables in the system.\(^2\)

The results suggest that the real permanent shock dominates the nominal shocks at most forecast horizons for the entire floating exchange rate regime in the yen-dollar system.\(^3\) The next section briefly outlines the simple open economy IS-LM model 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.


\(^3\)Although there are five structural innovations in each system, only two leave permanent imprints on the data. The other three are purely transitory. 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 some of the transitory shocks are also real in nature, but since they don’t have a permanent effect on any of the variables, including the real exchange rate, the source of this shock is not of direct interest.
A MODEL OF THE REAL EXCHANGE RATE

I will attempt to motivate the empirical model specification by appealing to a very simple model of the real exchnage rate. This model is not intended to represent a model of testable restrictions or hypotheses but simply a way to introduce the variables used in the empirical analysis.

The open economy version of the IS-LM model, or Mundell-Fleming model, is a standard in international finance textbooks. The model is based upon specifications for goods market equilibrium, money market equilibrium and balance of payments equilibrium. The IS curve describes the locus of real income and nominal interest rate combinations that represent equilibrium in the goods market. A general functional form is given in (1).

\[ Y^{IS} = f(i, \ldots) + \nu_g \]  

(1)

This specifies goods market equilibrium as a function of the nominal interest rate. One can think of interest rates, exogenous to the goods market, as affecting investment, consumption, etc. The term \( \nu_g \) represents a productivity shock. The money market equilibrium specification is given by (2).

\[ i^{LM} = f(Y, \ldots) + \nu_\pi \]  

(2)

The money market specification treats real income as exogenous and the innovation \( \nu_\pi \) represents a nominal shock such as a money growth or inflation expectations shock.

The final component to this model is the balance of payments (BOP) specification.

\[ CA(Y, \rho, \ldots) + KA(i, \rho, \ldots) = 0 \]  

(3)

Here the current account (CA) balance is a function of real income and and a the real exchange rate. The capital account (KA) is a function of the nominal interest
rate and the real exchange rate. Similar equations are assumed describe the foreign economy equilibrium so that a system comprising foreign and domestic real income, foreign and domestic nominal interest rates and the real exchange rate will capture the salient features of the open economy relationship. The focus in this study is the real exchange rate and the contribution of the innovations impinging on it to its behavior, especially the permanent shocks affecting it. A VAR will be an appropriate way to measure the dynamic interactions of the variables without specifying a structure ahead of time.

IDENTIFICATION OF A STRUCTURAL VAR 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.\(^4\) The issues surrounding structural identification can be understood by specifying a reduced form VAR in (4),

\[ \Phi(L)X_t = \mu + \varepsilon_t \]  

(4)

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)

\(^4\)The relevant citations include, but are not limited to, Bernanke (1986), Sims (1986), Shapiro and Watson (1988), Blanchard and Quah (1989), King, et al. (1991) and Warne (1993).
where $E[\nu\nu'] = I$. 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,

$$X_t = \zeta + C(L)\varepsilon_t$$  \hspace{1cm} (6)

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
"neutrality" 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 empirical model has five variables so that $p = 5$. Therefore, the number of additional a priori restrictions needed to achieve exact identification is ten. Assuming that the variables are I(1), the existence of cointegrating relations among the five variables means that the rank of $C(1)$ is less than five implying that some of the structural innovations leave permanent imprints on the data while others have only temporary effects. Furthermore, $C(1)$ must satisfy the restrictions implied by cointegration, 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 restrictions on $C(1)$ that help aid in the identification of $D(1)$.

There are two implications from the foregoing discussion that I would like to highlight. 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 five variable case to be analyzed in the next section, there exists three cointegrating relations Therefore, no long-run neutrality restrictions can be imposed on the third, fourth and fifth structural innovation since they are already neutral.

\footnote{See Engle and Granger (1987) and Johansen (1991).}
with respect to the variables in the system by definition. That is, such a restriction is not binding nor independent of the cointegrating restrictions.\(^6\) In general then, the econometrician needs to combine both long-run and contemporaneous restrictions to achieve exact identification of the structural model.\(^7\) 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 \textit{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**

**Data**

The data used in this study represent monthly observations beginning in February 1971 and ending in July 1996 on the U.S. dollar spot exchange rate vis à vis the Japanese yen and the consumer price index from each of these countries. These data

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\(^6\)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 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.

\(^7\)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.
are from OECD main economic indicators. The interest rate for Japan is the average three-month CD rate for all banks collected by the Bank of Japan. The U.S. interest rate is the 3-month CD rate from the St. Louis Federal Reserve Bank. Japanese real income is measured by total household income divided by the CPI. The U.S. real income measure is real personal income from the National Income and Products Accounts. The data are plotted in figure 1.

**Unit Root and Cointegration Analysis**

Each of the series in figure 1 appear to be non-stationary. The real ¥/$ exchange rate clearly has a negative drift over the sample. This could be due to a deterministic trend instead of a stochastic one, but that would be a violation of PPP as well. A battery of univariate unit root tests indicate that each of the series are integrated of order one. These tests have been shown to be biased in favor of the null when multiplicative outliers are the cause of structural breaks in the series as demonstrated 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 such 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 endogenized. These univariate results will be shown to be confirmed by the multivariate evidence to be presented subsequently.

The estimation and testing of the long-run equilibrium (cointegrating) relationship in the system is accomplished by using the Johansen (1988) procedure. This procedure is based upon a transformation of the reduced form VAR into its error correction mechanism (ECM) form as in (7).

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*Theses are available upon request.*
\( \Delta X_t = \mu + \Pi X_{t-1} + \sum_{j=1}^{j-1} \Gamma_j \Delta X_{t-j} + \varepsilon_t \) \hspace{1cm} (7)

The tests for the number of cointegrating relations, \( r \), among the \( p \) variables in \( X_t \) are tests of the rank of the long-run parameter matrix \( \Pi = -\Phi(1) = \alpha'\beta' \). There are three possibilities that are of interest. When \( \Pi \) has full rank \( p \), the variables in \( X_t \) are all stationary in levels. If \( \Pi \) has rank equal to zero then there are no stationary variables in \( X_t \) nor are there any cointegrating relationships among \( X_t \). Finally, when \( \Pi \) has reduced rank of \( r \) then there are \( r \) independent linear combinations of \( X_t \) that are stationary, including the possibility that some of the variables in \( X_t \) are stationary alone, forming what I call a trivial cointegrating relation. When \( \Pi \) has reduced rank of \( r \), it can be decomposed into two \( p \times r \) matrices \( \alpha \) and \( \beta \) which represent the error correction space and cointegration space, respectively. Johansen has developed two tests for the rank of \( \Pi \), the maximum eigenvalue (\( \lambda \)-max) test and the trace test. As the name of the first test suggests, these tests are based upon the number of non-zero eigenvalues in a Cholesky decomposition of the matrix of cannonical correlations of \( X_t \).

Table 1 presents the summary statistics from the cointegration analysis.\(^9\) The maximum eigenvalue and trace test results are quite clear that there are at least two cointegrating relations. The p-values of the test statistics are generated using the response surface procedure of MacKinnon et al. (1996). The evidence regarding the significance of the third cointegrating vector is somewhat marginal. Although the

\(^9\)The lag length of the VECM was set at thirteen 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 equations probably due to ARCH effects. Gonzalo (1994) demonstrates via Monte Carlo analysis that the Johansen estimator is quite robust to such departures from normality. However, any inference on the VECM parameters must be done using a heteroscedasticity-consistent covariance matrix. This is accomplished in the present study by using the Newey and West (1987) estimates for the covariance matrix.
\(\lambda\)-max test rejects the null of two cointegration vectors in favor of three at the 7% level of significance, the trace test only rejects at the 15% level of significance. The evidence of three cointegrating vectors in the system is mixed. Figure 1 displays 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.\(^{10}\) Further evidence of the the number of cointegrating vectors is provided in figure 3 where the trace statistics are calculated recursively using a procedure advocated by Hansen and Johansen (1993).\(^{11}\) Figures 4 through 7 display the recursive estimates of the four largest eigenvalues along with asymptotic two-standard error bands. The largest two eigenvalues are always significantly different from zero while the third is only significant in the later part of the sample. The fourth largest eigenvalue is never significant. Given these results I will proceed under the maintained hypothesis that \(r = 3\).

Table 2 displays statistics from the ECM estimation using the three cointegration vectors and a lag length of \(\ell = 13\), which was chosen as the minimum lag necessary to ensure residuals from all VAR equations were free from autocorrelation. The Q-statistic is the LM test for residual serial correlation and is distributed as a \(\chi^2(51)\) statistic.\(^{12}\) There is clearly no significant residual autocorrelation as evidenced by the large p-values associated with this test.

The exclusion tests are likelihood ratio (LR) tests of the null hypothesis that the

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\(^{10}\)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).

\(^{11}\)In this procedure all of the short-run dynamics are concentrated out and the recursive estimation is conducted on the long-run model only. As argued by Hansen and Johansen (1993), when the parameters of interest are the long-run parameters, i.e. \(\beta\), setting the short-run parameters to their full sample values reduces the possible variation in test results due to instability in the short-run dynamics of the system.

\(^{12}\)The consistency of this test requires that the number of autocorrelations tested increases with the sample size.
variable of interest may be excluded from the all of the cointegration vectors. This statistic has a $\chi^2(3)$ distribution. The evidence is very clear cut. All of the variables enter at least one of the three cointegrating relations.

The stationarity statistic tests the hypothesis that one of the cointegration vectors contains only the variable of interest. This implies that, conditional upon three cointegrating relations, the variable of interest is stationary by itself or that it forms a trivial cointegrating relation. This test can be thought of as complementary to the univariate unit root tests. This statistic is distributed as a $\chi^2(2)$ variate. There is no evidence that any of the five variables are stationary.

Finally, the weak exogeneity statistic tests the hypothesis that the variable of interest is weakly exogenous to the parameters of interest, in this case the cointegrating vector estimates. A weakly exogenous variable is one whose conditional distribution does not depend upon the error correction terms and thus need not be modelled explicitly to gain efficient estimates of the parameters of interest. More importantly, a weakly exogenous variable can be economically identified as a source of one of the common trends. There is some evidence that the Japanese nominal interest rate is weakly exogenous with respect to $\beta$, the cointegration vectors, but its marginal significance level is 11%.

**Stability Tests**

If the dynamic analysis of the system is to be valid the system and its parameter estimates must be relatively stable over the sample. Structural breaks and regime shifts may lead to spurious correlations and dynamics. Figure 8 displays recursive LR tests of the hypothesis that the cointegration space is equal to its full-sample value normalized by the appropriate critical value. There is some evidence of instability in the relationship in the late 1979 and disappears in late 1981. This is the period in which the U.S. Federal Reserve Board changed in monetary policy operat-
ing procedures to targeting non-borrowed reserves. Figure 9 displays the recursive
tests of the stability of the cointegration space when the model is augmented with an
intervention dummy variable that is defined as one from October 1979 until October
1981 and zero in all other periods. This modelling strategy appears to eliminate the
instability in the cointegration vector estimates. Figure 10 plots the recursive trace
statistics obtained from the model with the intervention dummy variable. Figure 11
plots the permanent and transitory components from each series implied by the fitted
common trends representation. Estimation of the common trends representation and
structural identification are presented next.

Structural Identification

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
$\beta_1 \beta_1 = 0, \alpha' \alpha_1 = 0, \Phi^*(1) = \Phi_1^* + \ldots + \Phi_{n-1}^*$ 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 re-
strictions 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 $\beta_1^0(\alpha'_1 \Phi^*(1)\beta_1^0)^{-1}\alpha'_1$ where
$\beta_1^0 = A(0)\beta_1$. Thus the identifying restrictions can be imposed directly on $\beta_1$.\footnote{The choice of $\beta_1^0$ must also satisfy the condition that the structural errors are independent. This
is ensured when $(\beta_1^T \beta_1^T)^{-1} \beta_1^T C(1)\Omega C(1)'\beta_1^T (\beta_1^T \beta_1^T)^{-1}$ is diagonal, where $\beta_1^T = \beta_1^0(\alpha'_1 \Phi^*(1)\beta_1^0)^{-1}$. Taking the Cholesky factor of this matrix and multiplying it by $\beta_1^0$ will achieve this independence
requirement.} Since there are two common trends or $k = 2$, one a priori restriction must be im-

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posed to achieve identification of the long-run components. But, since \( r = 3 \), i.e., there are three cointegrating vectors, requiring three contemporaneous restrictions to identify the short-run components of the model.

The variables in each system are ordered as \( X_t = [y_t, \hat{i}_t, \rho_t, y_t^*, \hat{i}_t^*]' \). The restriction imposed to identify the permanent real shock from the permanent nominal shock 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 U.S. real income. This restriction suggests interpreting structural shocks to the U.S. real income 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 the interest rates may be traced back to the U.S. through, say, monetary policy. 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 U.S. real income should not be influenced permanently by purely nominal changes. This restriction does not prevent U.S. real income from a rich dynamic response to nominal shocks in the short to intermediate term.

The restriction is most easily imposed by computing a general form for \( \beta_\perp \) (using, for instance, the generalized inverse of \( \beta \)) and setting the first two columns of \( A(0) = [(I_k*\beta_\perp)^{-1}:0] \), where \( I_k \) is the \( k \times k \) identity matrix.\(^{14}\) The estimate of the cointegration

\(^{14}\text{The third column of } A(0) \text{ is restricted by the assumption that the permanent and transitory innovations are orthogonal. Specifically, the covariance matrix between the permanent and transitory innovations is } \alpha_\perp' \Omega \alpha_\perp, \text{ where } A(0)_r \text{ is the } r \times p \text{ submatrix of } A(0) \text{ such that } A(0) = [(I_k* \beta_\perp)^{-1}:A(0)_r]\). The orthogonality restriction implies that this covariance be equal to zero. Therefore, \( A(0)_r \) must lie in the null space of \( \alpha_\perp \). This obviously implies that \( A(0)_r \) must lie in the space spanned by \( \alpha \).
space is \( \hat{\beta} = \begin{bmatrix} 1.00 & 1.00 & 1.00 \\
1.41 & 0.30 & 0.53 \\
-0.37 & 0.21 & 0.04 \\
-1.68 & -1.18 & -0.85 \\
1.91 & -0.48 & 3.69 \end{bmatrix} \), with a corresponding restricted orthogonal complement given by \( \beta^0_\perp = \begin{bmatrix} 1.00 & 0.00 \\
-1.01 & 1.31 \\
0.71 & 0.50 \\
-0.10 & -0.04 \end{bmatrix} \). 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 U.S. real output, is enough to identify the separate permanent innovations.

The dynamic analysis of the structural models yields impulse response functions (IRFs), forecast error variance decompositions (VDCs) and historical residual decompositions (HDCs).

Figure 12 displays the impulse responses due to a permanent real innovation along with simulated 90% confidence bounds.\textsuperscript{15} If we interpret this shock as a productivity innovation originating in the U.S. then we should observe a dollar appreciation following the shock. Since the real exchange rate is defined as yen-per-dollar, the IRF of the real exchange rate conforms to economic intuition. Two years after the shock,

\textsuperscript{15}It should be noted that the simulated confidence bounds assume that the underlying data processes are normally distributed. When this assumption is violated the interpretation of the confidence intervals must be viewed as only an approximation to the true confidence boundaries.
however, it is seen that the real value of the dollar versus the yen has declined. Figure 13 displays the IRFs due to the permanent nominal innovation. Here both U.S. and Japanese real income increase temporarily but eventually return to their original equilibrium paths. This shock should result in a domestic depreciation assuming its origin is the U.S. and that is precisely what is observed in figure 13.

The relative importance of these two shocks in explaining the real exchange rate can be gleaned from table 3 which presents the variance decompositions of the real exchange rate. Note that almost half of the variance of the one-step ahead forecast error in the real ¥/$ exchange rate is due to the second transitory innovation. The importance of this transitory innovation in explaining the real exchange rate does not die out for many years suggesting a significant mean reverting component in the real ¥/$ exchange rate.

The real permanent innovation only begins to dominate the real exchange rate after two years. By the end of the first year the two permanent innovations are each responsible for roughly one-third of the variance in the real exchange rate. This implies that while the real permanent innovation is important in explaining the unit root in the real ¥/$ exchange rate, the nominal permanent innovation is almost as important. One cannot attribute the failure of PPP to real innovations only. This may suggest some foreign exchange market anomaly where purely nominal shocks are not neutral in the long run.

Figure 17 displays the decomposition of the real exchange rate residuals into its structural components. The real permanent innovation appears to be much more important in the 1980s and the nominal more important in the 1970s and 1990s. The transitory innovations contribute significantly to the historical forecast error.
CONCLUSION

This paper uses structural VAR methodology to examine the importance of various shocks in explaining the real ¥/$ exchange rate over the modern floating regime with particular emphasis on the resul

REFERENCES


[8] William J. Crowder. The International Convergence of Inflation Rates During Fixed and Floating Exchange Rate Regimes. *Journal of International Money and Fi-


Table 1. Tests for the Number of Cointegrating Relations

<table>
<thead>
<tr>
<th>H₀:</th>
<th>Eigenvalues</th>
<th>λ–max</th>
<th>P-value</th>
<th>Trace</th>
<th>P-value</th>
</tr>
</thead>
<tbody>
<tr>
<td>r = 0</td>
<td>0.1744</td>
<td>56.15</td>
<td>&lt; 0.001</td>
<td>112.43</td>
<td>&lt; 0.001</td>
</tr>
<tr>
<td>r ≤ 1</td>
<td>0.1001</td>
<td>30.90</td>
<td>0.018</td>
<td>56.28</td>
<td>0.007</td>
</tr>
<tr>
<td>r ≤ 2</td>
<td>0.0663</td>
<td>20.09</td>
<td>0.069</td>
<td>25.37</td>
<td>0.149</td>
</tr>
<tr>
<td>r ≤ 3</td>
<td>0.0169</td>
<td>5.00</td>
<td>0.742</td>
<td>5.28</td>
<td>0.778</td>
</tr>
<tr>
<td>r ≤ 4</td>
<td>0.0010</td>
<td>0.28</td>
<td>0.597</td>
<td>0.28</td>
<td>0.597</td>
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</table>

Table 2. Various Statistical Tests from Cointegrated VAR

<table>
<thead>
<tr>
<th>Dep. var.</th>
<th>Q-Stat</th>
<th>P-value</th>
<th>Exclusion</th>
<th>P-value</th>
<th>Stationarity</th>
<th>P-value</th>
<th>Weak Exog.</th>
<th>P-value</th>
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</thead>
<tbody>
<tr>
<td>Yₜₛ</td>
<td>19.98</td>
<td>0.99</td>
<td>21.46</td>
<td>0.00</td>
<td>16.13</td>
<td>0.00</td>
<td>12.49</td>
<td>0.01</td>
</tr>
<tr>
<td>Yₗₚ</td>
<td>45.04</td>
<td>0.71</td>
<td>28.40</td>
<td>0.00</td>
<td>13.57</td>
<td>0.00</td>
<td>26.26</td>
<td>0.00</td>
</tr>
<tr>
<td>Rₜₛ</td>
<td>40.98</td>
<td>0.84</td>
<td>26.24</td>
<td>0.00</td>
<td>14.94</td>
<td>0.00</td>
<td>23.55</td>
<td>0.00</td>
</tr>
<tr>
<td>Rₗₚ</td>
<td>33.63</td>
<td>0.97</td>
<td>14.06</td>
<td>0.00</td>
<td>18.95</td>
<td>0.00</td>
<td>6.10</td>
<td>0.11</td>
</tr>
<tr>
<td>X</td>
<td>36.56</td>
<td>0.94</td>
<td>40.01</td>
<td>0.00</td>
<td>13.85</td>
<td>0.00</td>
<td>20.78</td>
<td>0.00</td>
</tr>
</tbody>
</table>

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Table 3. Variance Decomposition of Real Exchange Rate

<table>
<thead>
<tr>
<th>Forecast Horizon</th>
<th>$\varepsilon_1$</th>
<th>$\varepsilon_2$</th>
<th>$\varepsilon_3$</th>
<th>$\varepsilon_4$</th>
<th>$\varepsilon_5$</th>
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</thead>
<tbody>
<tr>
<td>1</td>
<td>14.63</td>
<td>34.45</td>
<td>0.00</td>
<td>44.47</td>
<td>1.57</td>
</tr>
<tr>
<td></td>
<td>{54.39, 0.09}</td>
<td>{63.90, 5.46}</td>
<td>{0.00, 0.00}</td>
<td>{72.37, 22.21}</td>
<td>{13.64, 0.01}</td>
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<tr>
<td>4</td>
<td>22.56</td>
<td>30.24</td>
<td>0.85</td>
<td>39.32</td>
<td>1.07</td>
</tr>
<tr>
<td></td>
<td>{62.62, 1.00}</td>
<td>{63.78, 3.06}</td>
<td>{2.09, 0.19}</td>
<td>{67.05, 17.67}</td>
<td>{9.06, 0.17}</td>
</tr>
<tr>
<td>8</td>
<td>23.37</td>
<td>31.76</td>
<td>0.58</td>
<td>37.66</td>
<td>1.14</td>
</tr>
<tr>
<td></td>
<td>{63.36, 1.04}</td>
<td>{66.50, 3.37}</td>
<td>{1.69, 0.22}</td>
<td>{65.60, 16.39}</td>
<td>{9.09, 0.27}</td>
</tr>
<tr>
<td>12</td>
<td>28.70</td>
<td>28.68</td>
<td>0.94</td>
<td>34.24</td>
<td>1.65</td>
</tr>
<tr>
<td></td>
<td>{66.74, 2.51}</td>
<td>{65.60, 2.76}</td>
<td>{3.68, 0.21}</td>
<td>{61.83, 14.35}</td>
<td>{9.72, 0.42}</td>
</tr>
<tr>
<td>24</td>
<td>32.21</td>
<td>24.03</td>
<td>6.29</td>
<td>28.41</td>
<td>3.52</td>
</tr>
<tr>
<td></td>
<td>{66.59, 4.15}</td>
<td>{59.48, 3.60}</td>
<td>{13.23, 1.87}</td>
<td>{53.28, 10.90}</td>
<td>{13.28, 1.08}</td>
</tr>
<tr>
<td>48</td>
<td>33.55</td>
<td>25.68</td>
<td>10.06</td>
<td>22.79</td>
<td>3.26</td>
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<tr>
<td></td>
<td>{65.53, 5.88}</td>
<td>{59.91, 3.91}</td>
<td>{15.10, 4.68}</td>
<td>{44.23, 9.08}</td>
<td>{12.18, 1.07}</td>
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<td>120</td>
<td>41.47</td>
<td>29.47</td>
<td>8.11</td>
<td>14.06</td>
<td>2.55</td>
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<tr>
<td></td>
<td>{74.10, 6.44}</td>
<td>{67.52, 3.32}</td>
<td>{15.43, 6.22}</td>
<td>{31.94, 6.22}</td>
<td>{12.43, 0.91}</td>
</tr>
</tbody>
</table>
Fig. 1.
FIG. 2. Zivot-Andrews Unit Root Tests
Fig. 3. No Intervention Dummies
Fig. 4.
Second Largest Eigenvalue Plus/Minus Two SE

Fig. 5.
Fourth Largest Eigenvalue Plus/Minus Two SE

Fig. 7.
Recursive LR Tests of Constant Cointegration Space

Fig. 8. No Intervention Dummies
Recursive LR Tests of Constant Cointegration Space

Fig. 9. With Intervention Dummies
Fig. 10. With Intervention Dummies
FIG. 12. Impulse Responses from Shock to First Permanent Innovation
FIG. 13. Impulse Responses from Shock to Second Permanent Innovation
FIG. 14. Impulse Responses from Shock to First Transitory Innovation
Fig. 15. Impulse Responses from Shock to Second Transitory Innovation
FIG. 16. Impulse Responses from Shock to Third Transitory Innovation
Fig. 17. Historical Decomposition