Sources of Real Exchange-Rate Fluctuations: Empirical Evidence from Four Pacific Basin Countries*

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I. Introduction

In international finance, the competing hypothesis that the logarithms of real exchange rates appear to be well-described as random walks during a period of floating rate, has found some support. Studies by Adler and Lehmann [2], Hakkio [12] and Mark [18] are examples of work supporting this view. If a real exchange rate process has been approximated to be non-stationary it implies there is little tendency for the real exchange rate to be mean-reverting, and also that deviations from the purchasing power parity (PPP) are permanent.

Modern models of exchange rate determination suggest that real shocks can induce permanent changes in the real exchange rate. Stockman [21] pointed out that according to the equilibrium approach, the behavior of real exchange rates since the collapse of Bretton Woods could reflect, not the importance of sluggish price-level adjustment, but rather the influence of real shocks with significant permanent components. Moreover, the evidence in Campbell and Clarida [5] and Huizinga [13] suggests that real exchange rates contain both permanent and transitory components, with most of their variation explained by the permanent components.

However recent studies employing long-run data, or more powerful tests, tend to confirm the PPP hypothesis. Studies by Abuaf and Jorion [1], Kim [16], Whitt [22] and Diebold, Husted, and Rush [7] are four such examples. Their evidence shows that real exchange rates are stationary, which implies that the behavior of real exchange rates are influenced by transitory nominal disturbances. Hence, the validity of PPP remains at issue.

Identifying the sources of exchange rate fluctuations is important not only for establishing the validity of PPP but also for achieving successful exchange rate stabilization. Attempts to stabilize exchange rate changes that are due to economic fundamentals could be futile and even harmful to the economy. In addition, measuring the relative importance of permanent and transitory shocks on exchange rates is essential for exchange rate modeling. If exchange rates are dominated by real shocks then the "equilibrium approach" offered by Stockman [21] is appropriate in analyzing the behavior of exchange rates. Conversely, if the evidence suggests the contrary, then the "disequilibrium approach" of Dornbusch [8] should be considered as an alternative. As a consequence,

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it is important to develop an empirical analysis to measure the relative importance of permanent and transitory shocks on exchange rates.

Recently, the approach of long-run structural VAR proposed by Blanchard and Quah [4] has been used to examine the sources of the macroeconomic fluctuations.[1] Lastrapes [17] and Evans and Lothian [10] extend this approach in examining the sources of real exchange rate fluctuations. In general, their findings support the importance of real shocks. To date, the majority of studies concerning the sources of real exchange rate fluctuations have been conducted for developed countries; most of these studies indicate that real shocks dominate in the variability of real exchange rates.

However, it is interesting to consider the behavior of the real exchange rate for the Pacific Basin countries. Over the past decade, countries in the Pacific Basin have instituted a continuing policy of financial market liberalization and as a result have experienced rapid growth, which has lead to increasingly strong trading ties to the United States.

The purpose of this paper is to apply the method of long-run structural VAR to analyze the influence of real shocks on the fluctuations of real exchange rates for four Pacific Basin countries. Our paper attempts to distinguish between empirically real, as opposed to nominal sources of exchange rate fluctuations. We find that real shocks dominate the variability of real exchange rates. This is consistent with the finding of Lastrapes [17].

Our paper is organized as follows: Section II provides the empirical model, discusses the econometric method, and examines the model's identification. Section III presents empirical results, and our conclusions are summarized in section IV.

II. The Structural VAR Model

In this section, we apply the long-run structural VAR approach to examine the influence of real shocks on the fluctuations of real exchange rates. We utilize Blanchard and Quah's [4] method to show how our assumptions help in characterizing the exchange rate process, and how this process can be recovered from the data.

Suppose that the economic system is driven by two structural shocks: real disturbances, and monetary disturbances. Furthermore, let us assume the monetary shock has no long-term effect on the real exchange rate; at the same time these two disturbances are uncorrelated and have unit variance.

Firstly, we will consider a vector of stationary variables $\mathbf{X}_t$ and a vector of structural shocks $\mathbf{U}_t$. In this paper, $\mathbf{X}_t$ includes the change in real and nominal exchange rate and $\mathbf{U}_t$ consists of the real and monetary innovations. Then $\mathbf{X}_t$ can be represented by:

$$\mathbf{X}_t = \sum_{k=0}^{\infty} \mathbf{A}_k \mathbf{U}_{t-k} = \mathbf{A}(L)\mathbf{U}_t, \quad \mathrm{Var}(\mathbf{U}_t) = \mathbf{I} \tag{1}$$

where $\mathbf{X}_t = \{\Delta s_t, \Delta q_t\}'$, Δs_t and Δq_t are the change in nominal and real exchange rates respectively, and $\mathbf{U}_t = \{u_t^m, u_t^r\}'$, where u_t^m and u_t^r denote for monetary innovations (e.g., money supply) and real innovations (e.g., resources endowment, technology, preference) respectively. $\mathbf{A}_i$'s are matrices of unknown structural coefficients and the disturbances are unobservable.

1. An advantage of the long-run structural VAR model is that the model does not impose contemporaneous restrictions, but it allows the data to determine short-run dynamics based conditionally on particular long-run restrictions.

Secondly, we will show how to recover this representation from the data. Because $\mathbf{X}_t$ is stationary, it has a unique Wold-moving average representation:

$$\mathbf{X}_t = \mathbf{C}(L)\mathbf{V}_t, \quad \mathrm{Var}(\mathbf{V}_t) = \boldsymbol{\Omega}, \;\; \mathbf{C}_0 = \mathbf{I}. \tag{2}$$

After estimating the vector autoregressive representation of $\mathbf{X}_t$, one can then invert the estimated coefficients to obtain $\mathbf{C}_i$'s.[2]

Next, comparing equations (1) and (2), we can obtain:

$$\mathbf{V}_t = \mathbf{A}_0\mathbf{U}_t \tag{3}$$

$$\mathbf{A}_i = \mathbf{C}_i\mathbf{A}_0, \quad i = 1, 2, \ldots p. \tag{4}$$

Since $\mathbf{C}_i$ is obtained by inverting VAR's estimated coefficients, $\mathbf{A}_i$ is solved when $\mathbf{A}_0$ is known. Identifying $\mathbf{A}_0$ to solve the system is essential, because Zellner [23] proved that OLS estimates of such a VAR system are consistent and efficient if each equation has precisely the same set of explanatory variables.[3] The unique determination of the four elements of $\mathbf{A}_0$ requires four independent restrictions. The variance-covariance matrix in the equation (3), $\mathbf{A}_0\mathbf{A}_0' = \boldsymbol{\Omega}$, provides three restrictions. One more restriction is needed to solve the $\mathbf{A}_0$ matrix. Following Lastrapes [17], we assume that monetary shocks have no long-run effect on the real exchange rate. This assumption implies that the element in the second row and the first column of the $\mathbf{A}(1)$ matrix is zero. This restriction, along with those restrictions provided by the variance-covariance matrix, can be used to determine the $\mathbf{A}_0$ matrix and therefore get all $\mathbf{A}_i$ matrices. Having the matrices of $\mathbf{U}_t$ and $\mathbf{A}_i$'s, one can figure out the process of $\mathbf{X}_t$ from equation (1). Finally, we use the impulse-response functions and variance decompositions to illustrate the important dynamic character of the empirical model.

There are two potential problems with the interpretation of these shocks as both nominal and real. First, it may be that nominal shocks do have permanent effects on the real exchange rate (as for large nominal shocks under Baldwin's [3] hysteresis model).[4] If this impact is small relative to that of real shocks, however, the identification scheme we use is approximately correct as shown by Blanchard and Quah [4, 659, 668–69]. Second, the assumption that exchange rates are subject to only two structural disturbances may be misleading. In fact, it is likely that there are many sources for the disturbances which effect the variability of exchange rates. Blanchard and Quah [4, 669–672] derive reasonable conditions under which the existence of multiple shocks does not vitiate our identification of nominal and real shocks. It is not apparent that these conditions are testable, so the results of this paper are based on the assumption that they approximately hold.

2. Assume that the VAR model is $\mathbf{X}_t = \mathbf{B}(L)\mathbf{X}_t + \mathbf{V}_t$, then the moving average representation of $\mathbf{X}_t$ is $\mathbf{X}_t = (\mathbf{I} - \mathbf{B}(L))^{-1}\mathbf{V}_t = \mathbf{C}(L)\mathbf{V}_t$. As $(\mathbf{I} - \mathbf{B}(L))\mathbf{C}(L) = \mathbf{I}$, the $\mathbf{C}_i$ will be known when $\mathbf{B}_i$ is estimated by the VAR model in the usual way.

3. If the underlying structural model provides a set of over-identifying restrictions on the reduced form, OLS is no longer optimal.

4. Baldwin [3] pointed out that if market-entry costs are sunk, nominal shocks can alter domestic market structure and thereby have persistent real effects.

Table I. Unit-Root Test

	s		Δs		q		Δq	
	$Z(\tau_\mu)$	$Z(\tau_\tau)$	$Z(\tau_\mu)$	$Z(\tau_\tau)$	$Z(\tau_\mu)$	$Z(\tau_\tau)$	$Z(\tau_m u)$	$Z(\tau_\tau)$
Japan	−0.203	−2.319	−7.950*	−7.983*	−0.947	−1.993	−7.949*	−7.968*
Korea	−1.879	−1.848	−3.483*	−3.558*	−0.984	−1.796	−4.268*	−4.502*
Taiwan	−0.538	−1.923	−4.589*	−4.550*	−0.858	−2.645	−6.137*	−6.003*
Philippines	−2.533	−0.991	−7.465*	−8.174*	−2.181	−1.682	−10.218*	−11.083*

Notes:
1. $Z(\tau_\mu)$ and $Z(\tau_\tau)$ denote the Phillips-Perron [20] unit-root test without trend and with trend. The critical values of $Z(\tau_\mu)$ and $Z(\tau_\tau)$ are obtained from Dickey and Fuller [6].
2. *denotes for rejection of the unit root null at the 5% level.

III. Empirical Results

The Data

In our paper we investigate the time series properties of the real exchange rate of four Pacific Basin countries: Japan, Korea, Taiwan and the Philippines. The data includes quarterly exchange rates which are all relative to the U.S. dollar, and the consumer price index (CPI). As of 1973 Japan started using the flexible exchange rate system, so we use the data from Japan spanning the period from 1974Q1 to 1994Q4. As Korea, Taiwan and the Philippines began their financial market liberalization regimes in 1981, we employ the data of these three countries over the period of 1981Q1–1994Q4. The data have been obtained from International Financial Statistics except those for Taiwan, which were obtained from Financial Statistics Monthly Taiwan District, the Republic of China. All variables are in logarithms.

Integration and Co-integration Properties of the Data

To justify the appropriateness of the structural VAR, we need to show that each individual series is integrated of order one, and that both nominal and real exchange rates are not co-integrated. If real and nominal exchange rates are non-stationary but co-integrated with each other, then the VAR model should be replaced by an error correction representation.

We employ the Philips-Perron [20] unit-root test to examine the integration properties of the real and nominal exchange rates. Both statistics of $Z(\tau_\mu)$ and $Z(\tau_\tau)$ show that exchange rates are non-stationary in level but are stationary in first difference. Therefore, the results from Table I indicate that both real and nominal exchange rates can be characterized as integrated processes with order 1 in all countries.

Having established that individual time series are integrated of order 1, we then proceed to test for co-integration between nominal and real exchange rates. The Engle and Granger (EG) [9] two-step regression is the simplest method. Table II reports the result of the co-integration test with EG's methods. Both the Durbin-Watson (CRDW) statistics and Augmented Dickey-Fuller (ADF) statistics fail to support the existence of co-integration relationships between real and nominal exchange rates in all currencies.

As well as the methods of EG, we also applied the multivariate cointegration techniques provided by Johansen [14] and Johansen and Juselius [15] to test for co-integration between real and nominal exchange rates. The optimality of Johansen's estimation technique has been shown by

Table II. Engle-Granger's Cointegration Test (Model: $x = \alpha + \beta y$)

	Japan	Korea	Taiwan	Philippines
$x = q, y = s$				
CRDW	0.067	0.039	0.199	0.039
ADF(4)	−2.098	−1.603	−2.455	−0.036
$x = s, y = q$				
CRDW	0.112	0.026	0.235	0.286
ADF(4)	−2.533	0.219	−2.338	−0.921

Note:
1. q and s are the real and nominal exchange rate respectively.
2. The critical values of CRDW and ADF(4) can be obtained from Engle and Granger [9].

Table III. Johansen's Cointegration Test

	λ_{Max}		Trace	
	$H_0 : r = 0$	$H_0 : r \leq 1$	$H_0 : r = 0$	$H_0 : r \leq 1$
Japan	9.449	4.324	13.773	4.324
Korea	11.393	1.244	12.637	1.244
Taiwan	12.264	5.182	17.446	5.182
Philippines	12.074	3.730	15.804	3.730

Note:
1. λ_{Max} denotes the maximum-eigenvalue statistics.
2. The critical values of the maximum-eigenvalue and trace statistics are obtained from Johansen and Juselius [15].

Phillips [19] in terms of symmetry, unbiasedness, and efficiency properties. A Monte-Carlo study by Gonzalo [11] supports the superior properties of the Johansen estimation technique relative to several other techniques.

In order to implement Johansen's procedure, one needs to determine the optimal lag length in the VAR system. The lag length of the chosen VAR was 2 in all countries except for Japan, where the lag length was set to 3. Our procedure for choosing the optimal lag length was to test up from a general VAR(1) system until increasing the order of the VAR by 1 lag could not be rejected by using a likelihood-ratio statistic. The residuals from the chosen VAR were then checked for whiteness. If the residuals in any equation proved to be non-white, we sequentially chose a higher lag structure until they were whitened.

Table III lists the results of the co-integration test which were derived using the Johansen method. Both the trace and the maximum eigenvalue statistics demonstrate the existence of nonco-integration among variables. These findings are also consistent with those from Mark [18]. The finding of no co-integration implies that there is no linear long-run equilibrium relationship between nominal, and real exchange rates, over the empirical period for these countries.

To sum up, each individual series is integrated of order 1, and empirical evidence fails to support the existence of co-integrating relationships between real and nominal exchange rates in all currencies. Thus, the structural VAR can be applied to examine the sources of real exchange-rate fluctuations.

Impulse-Response Functions and Variance Decompositions

Impulse-response functions describe the dynamic characteristics of the empirical model. The

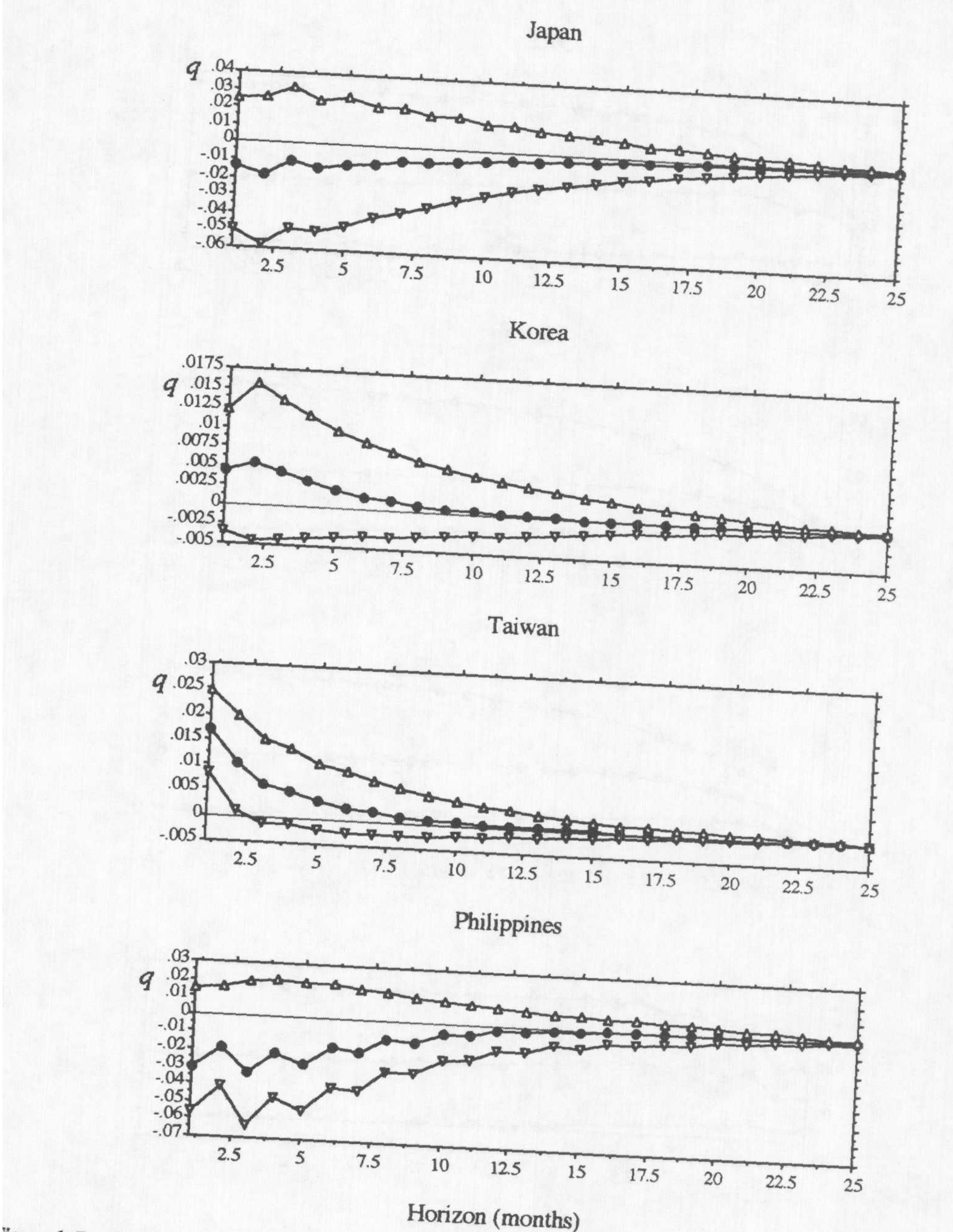

Figure 1. Reponse of Real Exchange Rate to Nominal Shock

responses of the individual levels are generated by accumulating the $\mathbf{A}_i$ coefficients.[5] Figure 1 illustrates the responses of the level of real exchange rates to a monetary disturbance. A

5. For example, with equation (1) we can get $x_1 = x_0 + a_0u_1$, and $x_2 = x_1 + a_0u_2 + a_1u_1$. Inserting the expression of into that of x_2 yields $x_2 = x_0 + a_0u_2 + (a_0 + a_1)u_1$. Repeating this operation, one can derive the following equation:

$$x_t = x_0 + a_0u_t + (a_0 + a_1)u_{t-1} + \ldots + \sum_{i=0}^{t-1} a_i u_1.$$

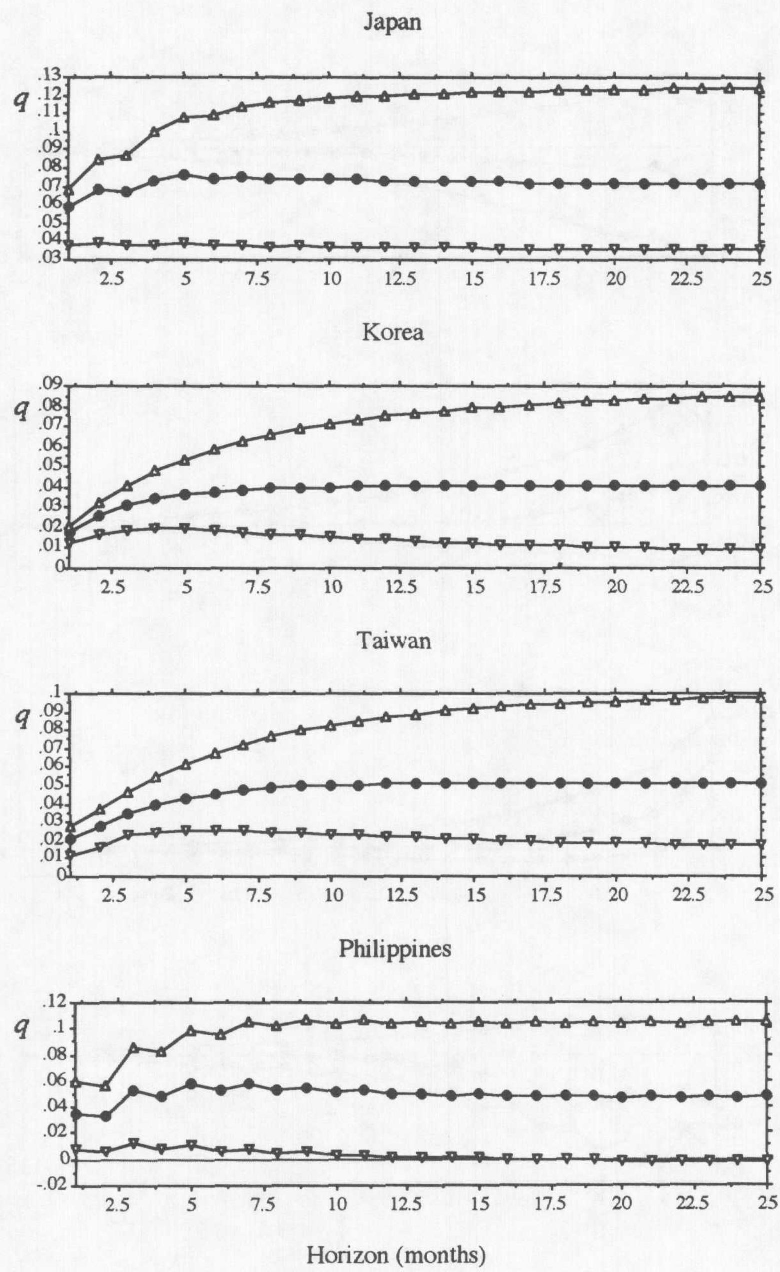

Figure 2. Response of Real Exchange Rate to Real Shock

two-standard error confidence interval, indicates the precision of the impulse-response function estimates. These bounds are constructed by using the sample standard deviation for empirical distribution from a Monte-Carlo simulation on the reduced form errors with 100 replications. Due to nominal shock, the responses in Figure 1 decompose gradually. This result is consistent with our assumption that a nominal shock has no effect on the real exchange rate in the long run. Figure

Table IV. Variance Decompositions of Real and Nominal Exchange Rates

	Japan		Korea	
	Relative Contribution of Real Shock (in %) to		Relative Contribution of Real Shock (in %) to	
k	q	s	q	s
1	95.1 (17.6)	89.2 (21.6)	94.7 (12.6)	89.3 (15.6)
4	93.1 (16.7)	88.4 (20.5)	95.0 (10.5)	89.6 (11.4)
8	92.9 (16.6)	88.3 (20.2)	94.9 (10.1)	89.4 (11.1)
12	92.9 (16.7)	88.3 (20.2)	94.9 (9.9)	89.4 (11.0)
24	92.9 (16.7)	88.3 (20.1)	94.9 (9.9)	89.4 (11.0)

	Taiwan		The Philippines	
	Relative Contribution of Real Shock (in %) to		Relative Contribution of Real Shock (in %) to	
k	q	s	q	s
1	60.7 (18.5)	99.6 (7.4)	58.6 (30.6)	22.6 (28.7)
4	62.0 (14.5)	94.6 (7.4)	55.7 (28.6)	27.0 (26.8)
8	62.3 (14.0)	94.4 (7.5)	53.4 (28.4)	28.0 (26.5)
12	62.3 (13.9)	94.4 (7.5)	53.3 (28.4)	28.0 (26.4)
24	62.3 (13.9)	94.4 (7.5)	53.3 (28.4)	28.0 (26.4)

Note:
1. q and s are the real and nominal exchange rate respectively, and k is the forecast horizon in months.
2. Value in parentheses is the empirical standard error from a Monte-Carlo simulation with 100 replications.

2 demonstrates the responses of the level of real exchange rates to a real disturbance, and if caused by a real shock, the responses of real exchange rates keep a constant level in the long run.[6]

Figure 1 illustrates that the initial impact of a monetary shock on real exchange rates varies across countries. Facing a nominal shock, the real exchange rates of won and NT$ rise immediately, but the real exchange rates of the yen and of the peso fall at once. However, a real shock raises the real exchange rates of the four currencies immediately. The response of the real exchange rate to monetary shocks goes to zero in about 15 quarters or even less in some cases (see Figure 1). On the contrary, the response of the real exchange rate to a real shock is persistent. In addition, the long-run effect of the real shock is greater than the initial impact of the shock.

Variance decompositions measure the relative contribution of forecast error variance of each shock as a function of forecast horizon. While the impulse-response function reveals the dynamic effects of a one-time shock, the variance decomposition is a convenient measure of the relative importance of such shocks to the system. In this paper, we have only discussed the effect of the real shock on real and nominal exchange rates.

Table IV reports the results of variance decompositions in the real and nominal exchange rates due to real shocks of various steps. Estimated standard errors from Monte-Carlo simulation are shown in parentheses. The role of real shocks in driving the real and nominal exchange rates

6. The empirical results from variance decompositions and impulse response analysis are not significantly affected when the lag length of the VAR model is set to 4 in all countries or when the quarterly data is replaced by the monthly data.

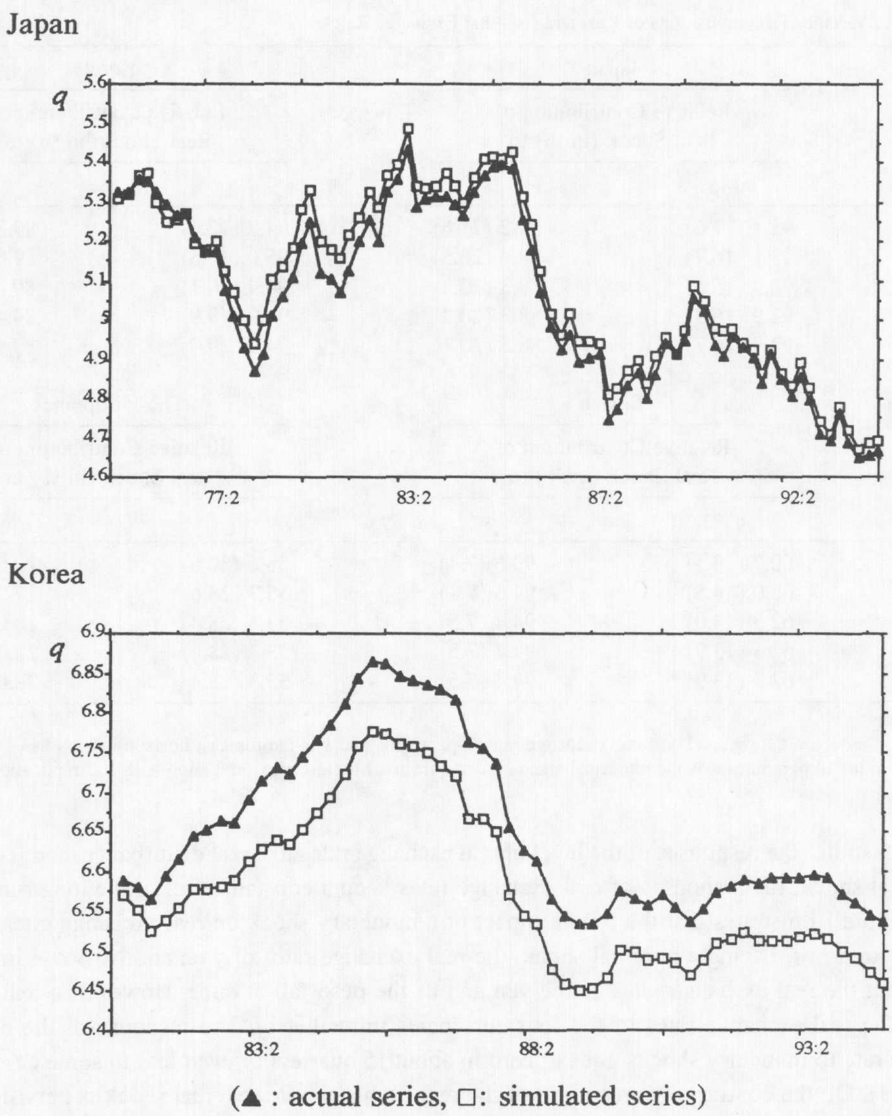

Figure 3. Real Exchange Rate Due to Real Shock

is apparent. For the yen, the real innovation causes a 95.1 percent of variation in the real exchange rate at one quarter, and the percent of variation falls to 92.9 percent at eight quarters. Furthermore, for the yen and the won, more than 90 percent of variation in real exchange rates are due to real shocks at all time horizons. The contribution of real disturbances to the real exchange rate of the peso and NT$, is about 50 and 60 percent respectively, at all time horizons. Hence, the role of real shocks varies across the currencies. Real shocks are more important for the yen and won and less important for the peso and the NT$. On the other hand, there is an essential difference in the degree of real shocks to all four currencies. In general, real shocks dominate in the variations of the real exchange rate. This result accords with the finding in Lastrapes [17].

For nominal exchange rates, the role of real shocks seems also to be more important in Japan, Korea, and Taiwan than in the Philippines. Except for Taiwan, real shocks are less important in

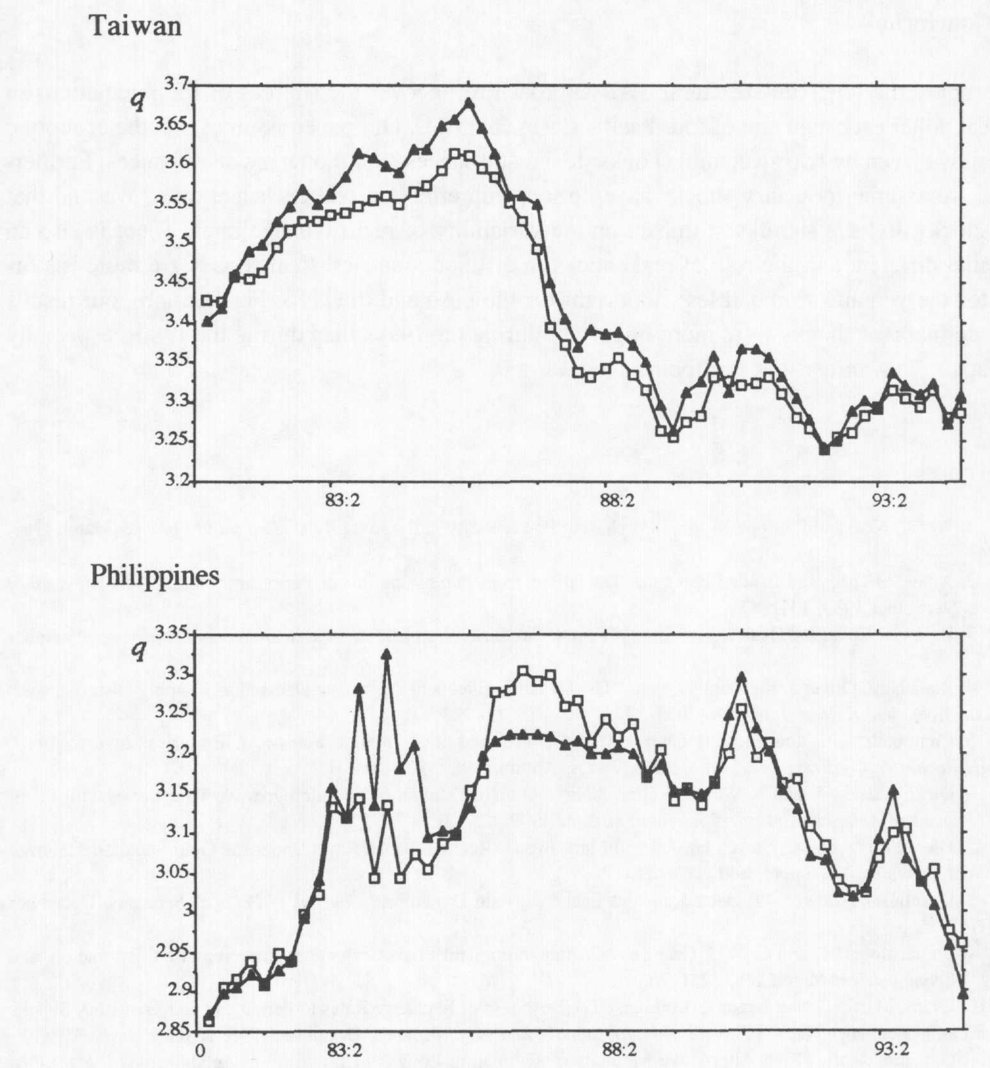

Figure 3. (continued) Real Exchange Rate Due to Real Shock

explaining the nominal rate than in explaining the real rate. Since the real exchange rate is the nominal exchange rate, weighted by the relative price levels, the previous findings are probably a result of the differences of relative price levels. In general, real shocks still play a dominant role in nominal exchange rates.

By removing the estimated nominal shock, we show, in Figure 3, the predicted value, by simulating the real shock only. The fit between the predicted and actual series is surprisingly good in Japan. For Korea, the movement of the predicted series is almost the same as that of the actual series. For Japan, Taiwan and the Philippines, the simulated series fits better during the 1990s, as opposed to the 1980s. Therefore, the patterns of Japan, Taiwan and the Philippines show that the real shock was more important during the 1990s than during the 1980s.

IV. Conclusion

We employ the long-run structural VAR approach to uncover the sources of the fluctuations in the real dollar exchange rate of four Pacific Basin countries. Our paper assumes that the economic system is driven by two structural shocks: real disturbances, and monetary disturbances. Furthermore, we assume monetary shocks have no long-run effect on real exchange rates. We find that real shocks do have significant impact on the variability of real exchange rates. There is also an essential difference in the role of real shocks in distinct countries. Real shocks are more important for the yen and won and less important for the peso and the NT$. Furthermore, our results indicate that real shocks were more important during the 1990s than during the 1980s, especially for Japan, Taiwan and the Philippines.

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