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Characterizing Real Exchange Rate Behaviour of Selected East Asian Economies*

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This paper attempts to evaluate aspects of the real effective exchange rate behaviour of five Asian countries by analysing (i) the extent to which the real exchange rates revert toward, in the long run, the purchasing power parity level; (ii) the persistent effect of a nominal exchange rate adjustment on the real exchange rate; and (iii) the equilibrium relationship between the real effective exchange rates and the relative prices of tradable versus non-tradable goods. The five countries under study are South Korea, Malaysia, the Philippines, Singapore and Thailand.

I. Introduction

Real exchange rates are increasingly being recognized as one of the key relative prices in determining the successful outcome of any stabilization policy and in achieving efficient allocation of resources in developing economies (Sachs and Collins, 1989). Since movements in the real exchange rates represent some combination of nominal exchange rate and relative price levels adjustments, the actual level of the real exchange rates will depend on the nominal exchange rate regimes, the wage price mechanisms and the stances of macro economic policy. In this paper, we analyse several aspects of the real exchange rate behaviour of five East Asian countries since the abandonment of the Bretton Woods System. The five countries are South Korea, Malaysia, the Philippines, Singapore and Thailand. In Section II, we present a descriptive analysis of the sources of the short-run real exchange rate variability of these countries. In Section III, we analyse how far the real

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exchange rates, in the long run, tend to revert toward the constant purchasing power parity level. In Section IV, we evaluate an important policy issue, i.e. how persistent is the effect of nominal exchange rate adjustment on the real exchange rate. In Section V, we look into the equilibrium relationship between the real exchange rates and the relative prices of tradable versus non-tradable goods. Section VI summarizes the overall findings of the paper.

II. Short Run Real Exchange Rate Variability

In this Section, we provide a brief discussion of the evolution of the nominal exchange rate regimes of the five East Asian countries following the termination of the Bretton Woods arrangement. We then present summary measures of the monthly changes in the logarithms of the real effective exchange rates, the logarithms of the nominal effective exchange rates and the logarithms of the relative price levels.

With the advent of generalized floating, both Malaysia and Singapore allowed their currencies to float until 1975 when both countries decided to peg their currencies to baskets of trading partners' currencies. Korea and Thailand continued initially to peg their currencies to the US Dollar but switch to pegging their currencies to basket composites after a major devaluation of the nominal exchange rates. Korea switched from the dollar peg to the composite peg in 1980 following a 20 percent devaluation while Thailand adopted the basket pegging arrangement after a devaluation in 1985. The Philippines officially adopted the US Dollar peg system, while making several adjustments to the nominal parity through a series of devaluations.

Table 1 shows the means and variances of the monthly changes in the logarithms of the real effective rates, together with the movements of the monthly changes of the logarithms of its components, viz, the nominal effective exchange rates and the relative price ratios. The sample period for Korea, the Philippines, Singapore and Thailand is from January 1974 to June 1988, while the sample period for Malaysia is from January 1975 to October 1987.

The nominal effective exchange rate index of each country is calculated as a geometric trade-weighted index, with the weights determined from the multilateral merchandise trade flows (imports plus exports) between the country concerned and its fifteen leading trading partners.[1] The component bilateral nominal exchange rate is defined as

[1] The multilateral weighting scheme assigns the weights to each trading partner accord-

Table 1

MEANS AND VARIANCES OF MONTHLY PERCENTAGE CHANGES IN EXCHANGE RATES AND INFLATION DIFFERENTIALS

Countries	Mean Δq	Mean Δe	Mean $\Delta(p^f\text{-}p)$	Var. Δq	Var. Δe	Var. $\Delta(p^f\text{-}p)$	Cov. $[\Delta e, \Delta(p^f - p)]$
South Korea	0.33	0.554	−0.521	4.612	4.451	2.353	−1.096
Malaysia	0.148	0.1477	−0.0001	2.806	1.424	1.189	0.193
Philippines	0.023	0.851	−0.828	7.360	9.123	4.345	−3.054
Singapore	0.0288	0.052	−0.235	2.263	1.264	1.115	−0.058
Thailand	0.232	0.363	−0.130	3.835	3.636	0.964	−0.383

Note: Δq = percentage change in real effective exchange rate, Δe = percentage change in nominal effective exchange rate, $\Delta(p^f\text{-}p)$ = percentage in relative prices.

the domestic currency price of a unit of foreign currency. The real effective exchange rate is computed by deflating the nominal rates by the relative wholesale price indices.[2]

Table 1 presents the decomposition of the sources of real exchange rate variability into the variation in the nominal effective exchange rate, the variation in the inflation differential and the covariation between the nominal exchange rate and inflation differential.[3] In Table 1, q represents the logarithm of the effective real exchange rate, e the logarithm of the nominal effective exchange rate, p^f the logarithm of the weighted wholesale price indices of the trading partners and p the logarithm of the domestic wholesale price index.

The variances of monthly changes in the real exchange rates in Korea, the Philippines and Thailand are dominated by variances in the nominal exchange rates. For these countries, the average ratio of the

ing to the partners' merchandise trade share in the total trade flows of the fifteen countries. The trade shares are averaged over the period 1979 to 1981. The base period for the indices of the real effective exchange rate was correspondingly set at 1979-1981.

2 Monthly observations of the nominal exchange rates and the wholesale price indices are taken from the IMF, *International Financial Statistics,* with the exception of the producer price index for Malaysia. The latter is obtained from the Department of Statistics, *Monthly Producer Price Indices.* The availability of the producer price index only for the period January 1975 to October 1987 at the time when the empirical research was carried out was the reason why the sample period for Malaysia was different from other countries.

3 See Mussa (1986).

variance of Δe to the variance of $\Delta(p^f - p)$ is 2.6. In Malaysia and Singapore, the sources of the short-run variability of the real exchange rates changes are evenly shared by the variability in the nominal exchange rate and the variability of the relative price changes. For these two countries, the average ratio of the variance of Δe to the variance of $\Delta(p^f - p)$ is 1.

There is little in the way of a systematic offsetting relationship between the fluctuations in the nominal exchange rate and the movements in the relative prices that would help keep the real exchange rates constant, as implied by the relative purchasing power parity relationship.[4]

III. Cointegration and the Purchasing Power Parity in the Long Run

In this Section, we investigate the long run validity of the purchsing power parity relationship for the five countries exchange rates based on the cointegration between domestic and international tradable goods prices. In the long run, when the effect of goods arbitrage is allowed to work itself out, the domestic price of tradable goods tend to converge to the international price level

(1) $\quad p_t^{*f} - \alpha p_t = X_t$

where $p_t^{*f} = e + p^f$ and X_t is the deviation at time t between the domestic tradable goods price level and the international price level. Purchasing power parity relationship holds if the estimated $\alpha = 1$ and $\{X_t\}$ is a stationary process.

A natural way to test for the long run PPP relationship based on (1) is to use the cointegration test proposed by Engle and Granger (1987). The essential notion behind the cointegration test is that while the sequences of p_t^{*f} and p_t may by themselves be non stationary and tend to drift apart at any one time, cross border commodity arbitrate tend to force these prices together into a stable equilibrium relationship. In such a situation, the domestic price level and the exchange rate adjusted international price level are said to be cointegrated, thereby restricting the real exchange rate to follow a stationary process.

The Engle and Granger tests of cointegration are based on the residuals of the cointegration regression estimated by the ordinary least

4 This observation is similar to those that have been widely documented in the short-run exchange rate behaviour of industrial countries. See, for example, Mussa (1986).

squares. The OLS estimates of the cointegration parameter are "super consistent" in the sense that the order of convergence of the OLS estimate to its true parameter value is much faster than in the stationary case. The cointegration regression based on (1) can be specified as

$$\text{(2)} \qquad p_t^{*f} = \alpha_0 + \alpha_1 p_t + X_t$$

The null hypothesis is that there is no cointegration between p_t^{*f} and p_t against the alternative of cointegration. Under the null hypothesis, the sequence of $\{X_t\}$ is a non stationary process with unit root. The presence of unit root in $\{X_t\}$ is tested using the cointegration Durbin-Watson (CRDW) test. Alternatively, the presence of unit root in $\{X_t\}$ can be evaluated using the augmented Dickey-Fuller (ADF) statistic from the following regression

$$\text{(3)} \qquad \Delta X_t = -\phi X_{t-1} + \sum_{i=1}^{p} \delta_i \Delta X_{t-i} + \epsilon_t$$

The Engle-Granger test procedure is based on the assumption that X_t is an iid normal variate with constant variance. However recent studies have shown that the exchange rates tend to exhibit conditional heteroskedasticity.[5] In addition, the existence of the "peso problem" is likely to induce serial dependence in the real exchange rate. Phillips (1987) has shown that a non parametric correction to the Engle-Granger test will allow the cointegration test to be robust to the error process of the cointegration regression. In the following cointegration tests of the purchasing power parity relationship based on (1), we report both the Engle-Granger tests results as well as the results based on the Phillips Z_t test.

Preliminary tests for the existence of unit roots, using the Dickey-Fuller (1979) t-test and the Phillips-Perron (1988) test, in the p_f^{*f} and p_t series are reported in Table 2. The results indicate that the null hypothesis of unit root cannot be rejected in the p_t^{*f} and p_t series for all the five countries. A similar test for the presence of unit roots in the Δp_t^{*f} and Δp_t is rejected. Hence, both the p_t^{*f} and p_t series are integrated of order 1 for the five countries.

The cointegration regressions for each country, with p_t^{*f} and p_t being alternatively normalized as dependent variable, are estimated and the results are reported in Table 3. The point estimates of the cointegration parameter, α_1, are all positive and in most cases approximate the

[5] See, for example, Cumby and Obstfeld (1984).

PPP value of unity. The R^2 of the equations are generally high, with the exception of the Singapore regressions. However, such estimates are

Table 2

TESTS OF UNIT ROOTS IN p^{*f} AND p

$$y_f = \alpha_0 + \beta_0 t + \beta_1 y_{t-1} + \sum_{i=1}^{p} \beta_{2i}\, \Delta y_{t-i} + w_t$$

	p^{*f}			p		
Countries	$\hat{\beta}_1$	$\tau\beta_1$	$Z(\tilde{\beta}_1)$	$\hat{\beta}_1$	$\tau\beta_1$	$Z(\tilde{\beta}_1)$
South Korea	0.941	−2.861	−2.912	0.935	−2.335	−2.593
Malaysia	0.973	−1.363	−1.501	0.982	−2.163	−2.188
Philippines	0.989	−1.213	−1.441	0.983	−0.682	1.071
Singapore	0.972	−2.833	−2.851	0.987	−0.612	−0.988
Thailand	0.937	−2.644	−2.667	0.953	−2.367	−2.476

Notes: $\tau\beta_1$ is the Dickey-Fuller t-test statistic of the null hypothesis that $\beta_1 = 1$ and its critical values are from Fuller (1976). The critical values for $\tau\beta_1$ at 0.05 level are −3.45 and −3.43 for sample sizes 100 and 250 respectively. $Z(\hat{\beta}_1)$ is the Phillips-Perron t-test statistic and it has similar critical value as $\tau\beta_1$ statistic. 12 lags ΔY_{t-1} are used. In the Phillips-Perron test, the lag window e is also set at 12.

Table 3

COINTEGRATION REGRESSIONS

Countries		**CRDW**	**R^2**
South Korea	$p_t = -3.745 + 0.896\, p_t^{*f}$	0.031	0.95
	$p_t^{*f} = 4.449 + 1.055\, p_t$	0.032	0.95
Malaysia	$p_t = 4.509 + 0.709\, p_t^{*f}$	0.021	0.69
	$p_t^{*f} = -4.445 + 0.985\, p_t$	0.025	0.68
Philippines	$p_t = -4.677 + 1.006\, p_t^{*f}$	0.114	0.98
	$p_t^{*f} = 4.722 + 0.978\, p_t$	0.115	0.98
Singapore	$p_t = -0.718 + 0.562\, p_t^{*f}$	0.012	0.475
	$p_t^{*f} = 5.384 + 0.850\, p_t$	0.009	0.475
Thailand	$p_t = -1.767 + 0.682\, p_t^{*f}$	0.031	0.88
	$p_t^{*f} = 3.393 + 1.287\, p_t$	0.032	0.88

Notes: CRDW is the cointegration Durbin-Watson test statistic and its critical value at 5 percent is 0.386 (Engle and Granger, 1987).

essentially spurious, given that the CRDW statistics are well below the critical value of 0.39 at the 5 percent level.[6]

The failure of the CRDW test to reject the null hypothesis of no cointegration is further confirmed by the augmented Dickey-Fuller test reported in Table 4, where the ADF statistics are all not significant. The computed Phillips's Z_t statistics also failed to reject the null hypothesis of no cointegration at the 5 percent level.[7]

Table 4

AUGMENTED DICKEY-FULLER AND PHILLIPS Z_t TESTS OF THE RESIDUALS OF THE COINTEGRATION REGRESSIONS

	Normalised on P_t		**Normalised on p^{*f}**	
Countries	τ_ϕ	$\mathbf{Z}_t$	τ_ϕ	$\mathbf{Z}_t$
South Korea	−1.161	−1.077	−0.937	−1.113
Malaysia	−1.767	−1.894	−1.583	−1.312
Philippines	−2.405	−2.105	−2.376	−2.276
Singapore	−1.167	−1.221	−1.492	−1.656
Thailand	−1.239	−1.484	−0.860	−1.572

Notes: τ_ϕ is the Augmented Dickey-Fuller test statistic and its critical value at 5 percent is −3.17 (Engle and Granger, 1987). Z_t is the Phillips Z_t test statistic and its critical value at 5 percent level is −2.7619 (Phillips and Ouliaris, 1990). In the ADF tests, the order of the autoregressive lags is set at 12 and in the Z_t test the lag window is also set at 12.

6 As Phillips (1986) has shown, in regressions involving variables which are integrated of order 1, the t-ratio does not have a limiting distribution but actually dverge as the sample size increases and thereby biasing the test towards rejection of the null hypothesis of no relationship between the variables. The covariance between the nominal effective exchange rate and relative prices in Malaysia, while positive, is rather small. In general, the small covariance observed in Malaysia and Singapore suggests that the nominal exchange rates in these two countries had not been adjusted at different intervals according to inflation rate differential. Rather, the exchange rate policy seems to be aimed at maintaining certain degree of nominal exchange rate stability as indicated by the relatively small variance of the nominal effective exchange rates.

7 We have also, following the suggestions of Banerjee *et al.* (1986) and others, conducted tests of cointegration based on whether $\{p_t{}^{*f}\}$ and $\{p_t\}$ obey an "error-correction" process. These authors have pointed out that the unrestricted error-correction formulation of the cointegration test is subject to lesser small sample bias since it does not incorporate the constraint imposed by the estimated equilibrium regression while at the same time it allows for the dynamics of adjustment through the inclusion of the lagged terms. The error-correction VARs are estimated for the five countries and the Engle-Granger F-tests also failed to reject the null hypothesis of non-cointegration. The results are available from the author on request.

The findings that the domestic and international price levels do not cointegrate implies that the real effective exchange rates for the five countries folow non-stationary processes characterized by stochastic trends. As confirmed in Table 5, the Dickey-Fuller and the Phillips-Perron test failed to reject the null hypothesis of unit root in all the five real effective exchange rate series.

Table 5

TESTS OF UNIT ROOTS IN THE REAL EXCHANGE RATES

Countries	$\hat{\beta}_1$	$\tau\beta_1$	$\mathbf{Z}(\tilde{\beta}_1)$
South Korea	0.986	−0.959	−0.962
Malaysia	0.993	−0.562	−0.712
Philippines	0.923	−2.673	−2.708
Singapore	0.959	−1.066	−1.088
Thailand	0.967	−0.606	−0.702

Note: 12 lags in the change in the real exchange rate were used.

IV. The Relationship between the Nominal and Real Effective Exchange Rates in the Long Run

The evidence presented in Section II indicates the short-run variability of the real exchange rates is substantially affected by the variability of the nominal exchange rates. In this Section, we investigate the extent to which the short-run impact of the nominal exchange rate exchanges on the real exchange rate tend to persist in the long run. The answer to this inquiry would be of considerable importance to the exercise of exchange rate policy where the monetary authority may try to influence the path of the real exchange rate, in the long run, through adjustment in the nominal exchange rate.

For this purpose, we estimated an unrestricted vector autoregression (VAR) consisting of the first difference in the logarithms of the nominal effective exchange rate, domestic wholesale price and the weighted foreign wholesale price.[8] From the estimated coefficients of the VAR we

[8] The unit root tests on the logarithms of the nominal exchange rate, domestic price level and weighted foreign price level indicate that these variables are integrated of order 1. Engle and Granger (1987) have indicated that the first-difference VAR specification is

derived the implied impulse response functions of the real exchange rates to a given shock in the nominal exchange rate.[9]

Table 6 shows the cumulative impulse response functions of Δe, Δp^f and Δp, together with the implied impulse response function for Δq, to a one-standard deviation shock in Δe. The impulse response functions are cumulated over sixty months to capture the long-run response of these variables. The cumulative impulse response functions indicate the level of these variables five years after a given nominal exchange rate shock.

A unit shock in the nominal effective exchange rate generally leads to an increase of at least one percent in the nominal rate in the long run in all the countries. A rather large response is found in the case of the Philippines, where a 1 percent shock in the nominal exchange rate results in a 4.4 percent depreciation in the long-run. Since the moving average representations are derived from the reduced form estimates, the large depreciation of the Philippines Peso possibly reflects the vicious circle effect arising from the dynamic interaction between the nominal exchange rate and price level. A vicious circle effect is especial-

not appropriate if the levels of the variables are cointegrated. A series of cointegration test on the three variable system using the Engle and Yoo (1987) augmented Dickey-Fuller procedure failed to reject the hypothesis of no cointegration among the three variables.

9 The impulse response functions are computed from the orthogonalized moving average representations of the estimated VAR in the following way. Let $y_t = (\Delta e, \Delta p^f, \Delta p)$. The VAR representation can be written as $A(L)y_t = u_t$, where $A(L) = \sum_{j=0}^{k} A_j L^j$, $A_0 = I$, and $E\ (u, u_t') = \Omega$. Inverting the $A(L)$ we get the moving average representation: $Y = A(L)^{-1} u_t$. To evaluate the dynamic response of the variables in y to an innovation in $\{\Delta e_t\}$, the contemporaneous innovations of these variables are orthogonalized by means of a Choleski factorization of Ω. Let $\epsilon_t = Gu_t$ and G is chosen to be a lower triangular matrix such that $G\Omega G' = \Gamma$. Γ is a diagonal matrix. Letting $C(L) = A(L)^{-1} G(L)^{-1}$, we can written $y_t = C(L)\ \epsilon_t$ where

$$C(L) = \begin{bmatrix} C_{11}(L) & C_{12}(L) & C_{13}(L) \\ C_{21}(L) & C_{22}(L) & C_{23}(L) \\ C_{31}(L) & C_{32}(L) & C_{33}(L) \end{bmatrix}$$

The long run change in e, p^f, and p to a unit shock in e is given by $C_{11}(1)$, $C_{21}(1)$ and $C_{31}(1)$ respectively. Therefore the implied long run change in q following a given shock in Δe is $C_{11}(1) + C_{21}(1) + C_{31}(1)$.

To implement the VAR consisting of the above three variables, two decisions have to be made. First is the ordering of the variables. The computed moving average presentations reported in Table 2 are based on the following ordering: $\Delta p^f / \Delta e$, Δp. The results are not sensitive to alternative orderings since the correlation among each pair of contemporaneous innovations are small.

The second decision concerns the choice of lag length for the variables in the VAR. Here, we impose a uniform lag length of twelve lags across equations and variables.

Table 6

CUMULATIVE IMPULSE RESPONSE FUNCTIONS (60 months) TO A ONE-STANDARD DEVIATION SHOCK IN THE NOMINAL EFFECTIVE EXCHANGE RATE

Country	Response to shock in			
	Δe	Δpf	Δp	Δq
South Korea	1.607	−0.007	0.784	0.816
Malaysia	1.907	−0.291	0.321	1.295
Philippines	4.382	−0.185	3.844	0.353
Singapore	1.747	−0.111	0.234	1.402
Thailand	1.269	0.156	0.913	0.202

ly likely in an inflation-prone country like the Philippines.

A marked difference exists in the response of the domestic wholesale price level to a unit shock in the nominal exchange rate between Korea, the Philippines and Thailand on the one hand and Malaysia together with Singapore on the other. In the first group of countries, a 1 percent shock in the nominal exchange rate leads to an increse in the domestic price level that varies from 0.8 percent (in the case of Korea) to 3.8 percent (as in the case of the Philippines). In Malaysia and Singapore, there appears to be a greater degree of price stickiness in response to exchange rate depreciation, where a percent innovation in the nominal exchange rate leads only to an increase of 0.3 percent in the domestic price level.

The small responsiveness of the domestic price level to nominal exchange rate changes in Malaysia and Singapore implies that changes in nominal rate exert a considerable persistence effect on the real exchange rate in these two countries. Typically, in these two countries, the real exchange rate depreciates in the long run by 1.3 percent following a unit shock in the nominal exchange rate. In contrast, in the Philippines and Thailand, a nominal exchange rate shock has no permanent impact on the real exchange rate.

V. The Relative Price of Tradable to Non-Tradable Goods and the Equilibrium Real Exchange Rates

The earlier analysis suggests that there is no systematic tendency for the real exchange rate, following a given shock, to revert, constant PPP equilibrium level. However, where real shocks cause permanent shifts in the relative price of tradable to non-tradable goods, the underlying equilibrium exchange rate itself is subject to change. The underlying shocks that cause equilibrium shifts in the relative price of tradable goods, can possibly originate from, among others, productivity growth differential between the tradable and non-tradable sector (Balassa, 1964; Marston, 1987), shifts in the external term-of-trade (Edwards, 1988), changes in the composition of government expenditure (Frenkel and Razin, 1987; Durlauf and Staiger, 1990).

We wish to analyse, therefore, the extent to which the real exchange rate reverts to a shifting equilibrium that is determined by the relative price of tradable goods.[10]

To obtain a measure of the real exchange rate that allows for long-run shifts in the relative price of tradable goods, we estimate the following cointegration equation

$$(4) \qquad q_t = a_0 + a_1 \; ptng_t + \tilde{q}_t$$

where ptng is the proxy for the relative price of tradable to non-tradable goods differential and it is defined as $(p_w - p_c) - (p_w^f - p_c^f)$. p_w and p_c are the logarithms of the domestic wholesale and consumer price indices respectively.[11] p_w^f and p_c^f are the logarithms of the weighted trading partners' wholesale and consumer price indices respectively. The residual of the regression $\tilde{q}_t$, is taken to be the real exchange rate adjusted for shifts in the relative price of tradable goods differential.

Applying the Engle and Granger cointegration tests on $\{\tilde{q}_t\}$ will indicate whether the series is stationary. In Table 7, we present the CRDW and ADF tests of the null hypothesis that $\{\tilde{q}_t\}$ is a non stationary process with unit root. The estimated cointegration parameters, $\hat{a}_1$, are of the right signs, indicating that an increase in the relative price of tradable goods leads to, in the long run, a depreciation of the real effective exchange rates in the five countries. However, the CRDW and the ADF test

10 For the use of movements in the differential relative tradable goods prices as an empirical proxy for shifts in the equilibrium real exchange rate, see Clements and Frenkel (1980), Koedijk and Schotman (1990). Genberg (1978) and Lothian (1990), on the other hand, proxied the movement in the equilibrium real exchange rate by a smooth deterministic trend.

11 The ratio of the domestic consumer price to domestic wholesale price was used as a proxy for the real exchange rate largely because implicit price deflators or price indices for broad classes of tradable and non-tradable sectors are not available for the sample of countries. The former represents an "indirect" approach to proxying the real exchange rate and has been widely used in the literature. See Dwyer (1992).

Table 7

TESTS OF COINTEGRATION BETWEEN REAL EFFECTIVE EXCHANGES AND THE RELATIVE PRICE OF TRADABLE GOODS DIFFERENTIAL

Countries		
South Korea	$q_t = 4.701 - 0.544$ ptng$_t$ CRDW $= 0.046$ ADF $= -0.776$	$R^2 = 0.036$
Malaysia	$q_t = 4.462 - 0.242$ ptng$_t$ CRDW $= 0.042$ ADF $= -0.445$	$R^2 = 0.076$
Philippines	$q_t = 4.618 - 0.027$ ptng$_t$ CRDW $= 0.113$ ADF $= -2.705$	$R^2 = 0.071$
Singapore	$q_t = 4.661 - 0.429$ ptng$_t$ CRDW $= 0.009$ ADF $= -0.3198$	$R^2 = 0.066$
Thailand	$q_t = 4.697 - 1.176$ ptng$_t$ CRDW $= 0.054$ ADF $= -0.753$	$R^2 = 0.172$

Notes: q = real effective exchange rate,
ptng = relative price of tradable to non-tradable goods differential.

statistics are well below the critical values at the 5 percent level. Consequently, the null hypothesis of the presence of unit root in $\{\tilde{q}_t\}$ cannot be rejected, suggesting that the real exchange rates of these countries, following a given shock, do not revert back to the equilibrium path determined by the differential relative price of tradable goods.

However, as Cochrane (1988), Poterba and Summers (1988) and others have shown, the unit root tests have low power in distinguishing between a pure random walk process and an alternative non-stationary process which contains considerable transitory, but persistent, mean reverting component. To adequately capture the slow decaying transitory component in the process, we need to examine the behaviour of the autocorrelations of a series over a sufficiently long time horizon.[12]

12 The slow decaying components could arise from persistence of nominal shocks (which do not change the equilibrium real exchange rate). In addition, Cutler, Poterba and Summer (1990) have shown that speculative trading between rational speculators and feedback traders can cause successive deviation of asset prices away from the fundamental values over the short horizon but with eventual reversion back to the fundamentals in the long run. In the foreign exchange market, a major negative feedback trading activity is carried

To evaluate whether the real exchange rates adjusted for relative price of tradable goods contain transitory components which cause reversion towards the equilibrium level at a longer time horizon, we employ the Fama and French (1988) long-horizon regression approach. As Fama and French have pointed out, in the context of a model of stock prices which contain the sum of random walk and transitory components, tests of slow mean-reverting component can be made by examining the autocorrelations of the stock returns for increasing holding periods. Let $r_1(k)$ be the first-order autocorrelations of the real exchange rate change over a time interval of length k. The Fama-French test of mean reversion of transitory component rests on the evidence of negative $r_1(k)$ at different horizons, $r_1(k)$ can be shown to be the slope, β_k, of the following regression

$$(5) \qquad \tilde{q}_{t+k} - \tilde{q}_t = \beta_0 + \beta_k(\tilde{q}_t - \tilde{q}_{t-k}) + u_t$$

Assuming that the random walk and transitory components are uncorrelated, β_k can be shown to approximate the fraction of the variance of the k-period change that are explained by the transitory component. If $\tilde{q}$ is a pure random walk, then $\beta_k = 0$ for all k. If $\tilde{q}$ is entirely made up of the purely stationary component, the β_k approaches $-1/2$ for large k. Where $\tilde{q}$ consists of both the random walk and the stationary components, the evidence of mean reversion will lead to values of β_k being negative.

Estimating equation (5) for various k's involves the use of overlapping data. Given the relatively small sample size, we confine the estimates to the horizons from $k = 1$ to 48. The results are presented in Table 8. Because of the use of overlapping data, the residuals of the regressions for $k > 1$ follow a moving average process of order $k-1$. To obtain consistent estimates of the variance-covariance matrix of β_k, we employ the Newey-West (1987) version of the generalized methods of moments. The Newey-West procedure is robust to both the presence of serial correlation and conditional heteroskedasticity in the regression residuals. We have not corrected for the potential downward bias in the estimated β_k (Fama and French, 1988; Kim, Nelson and Startz, 1989).

out by the Central Bank in its attempt to smoothern short-run exchange rate fluctuation through a policy of "leaning against the wind."

Frenkel and Froot (1990) survey evidence indicate that in the short horizon, traders in the foreign exchange market tend to trade on rules based on technical analysis. Trading on the basis of the chartists rules serves to drive the exchange rate away from the existing fundamental values. Only in the longer run are expectations more firmly grounded on fundamentals.

Table 8

ESTIMATED SLOPES OF THE k-period REGRESSIONS

$$\Delta\tilde{q}_{t+k}=\beta_0+\beta_k\,\Delta\tilde{q}_{t-k}+u_t$$

k	South Korea	Malaysia	Philippines	Singapore	Thailand
1	0.058	0.106	0.082	0.112	0.082
	(0.807)	(1.762)	(1.074)	(0.934)	(1.074)
3	0.072	0.284	0.045	0.273	0.029
	(0.957)	(1.371)	(0.449)	(1.244)	(0.369)
6	0.081	0.325	0.036	0.241	0.138
	(1.048)	(1.081)	(0.512)	(1.116)	(1.565)
12	0.115	0.127	−0.358	0.046	−0.119
	(1.407)	(1.461)	(−4.346)	(0.696)	(−1.462)
24	−0.534	−0.612	−0.530	−0.318	−0.459
	(−5.858)	(−7.465)	(−4.328)	(3.699)	(−2.366)
36	−0.894	−0.839	−0.203	0.592	−0.363
	(6.924)	(−7.009)	(−1.706)	(2.179)	(−2.606)
38	−0.249	0.212	−0.901	−0.935	1.997
	(1.886)	(1.691)	(−10.246)	(−4.865)	(10.04)

Note: Figures in parenthesis are t-statistics calculated from the Newey-West consistent standard errors.

The t-statistics (which is computed as the ratio of the OLS slope to its Newey-West standard errors) indicate that, for short horizons from 1 to 12 months, the null hypothesis that $\beta_k=0$ cannot be rejected. This suggests that at relatively short time horizons, the movements of the real exchange rates are dominated by the random walk component which tends to restrict reversion back to the equilibrium path determined by shift in the relative price of tradable goods. From 24 months onwards the $\hat{\beta}_k$'s are negative and significantly different from zero, suggesting that these exchange rates can be expected to revert back to their equilibrium path at least two years following a given shock. Even allowing for the possible downward bias in β_k, the estimates indicate the presence of substantial equilibrium reverting transitory components in the real exchange rates. With the exception of Singapore, around 50 percent of the deviation of the real exchange rate from its equilibrium relative price path tend to be reversed at the two year horizon. For Singapore, the proportion of transitory component at the two year horizon is 30 percent. At the 48 month horizon, the estimated transitory components reach its maximum of 90 percent in the Philippines Peso and the Singapore Dollar. The sample size is not large enough however, for us to estimate β_k at considerably longer horizons in order to detect its movement back to zero.

We next investigate the typical dynamic adjustment of the real exchange rate in response to a shock in the relative price of tradable goods different. For this purpose, we estimated a bivariate VAR consisting of 12 lags of Δq and Δptng from which we computed the impulse response functions.

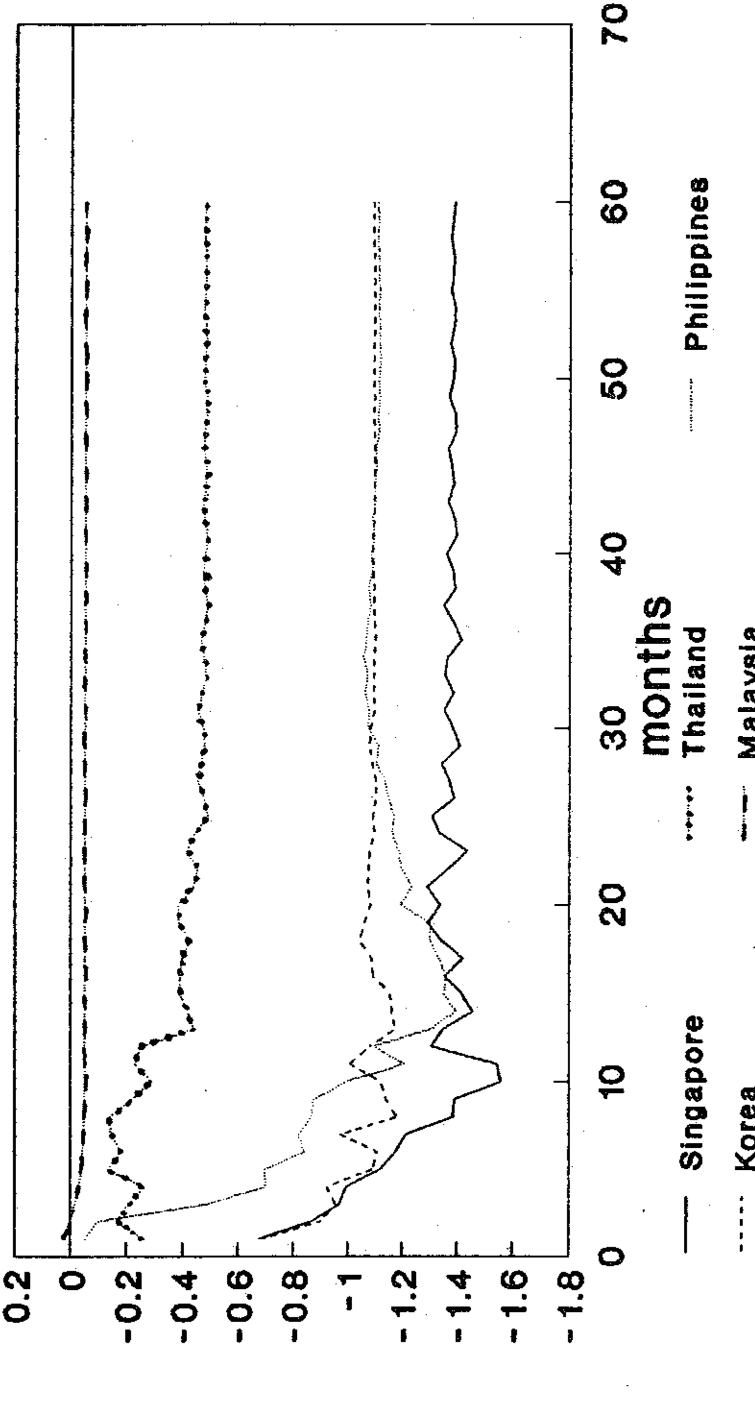
Figure 1 shows the plots of the accumulated response of the real effective exchange rates, over a 60 month period, to a one-standard deviation shock in the relative price of tradeable good differential. For all the five currencies, the real exchange rates appreciate in response to an innovation in the relative price tradable goods. However, the strength of the response varies considerably. In the case of Korea, the Philippines and Singapore, a one-standard deviation shock in the relative price of tradable goods causes a long-run appreciation of the real exchange rate by slightly over 1 percent. In the case of the Korean Won and the Singapore Dollar, the real appreciation is fairly rapid, reaching its peak around eight to ten months following the shock. The Philippines Peso's real appreciation is more sluggish, reaching its maximum only around fourteen months after the shock. The Thai Baht real exchange rate appears to be moderately responsive to the relative price shock. At the extreme of the response pattern is the Malaysian Ringgit, which shows hardly any appreciation in response to the relative price of tradable goods shock.

Finally, we try to obtain some measure of the long-run equilibrium real exchange rate and to determine the extent to which the actual movements of the real effective exchange rates are explained by shifts in the equilibrium rate. For this purpose we adopt the Beveridge-Nelson (1981) approach, in which the permanent or the long-run equilibrium component is viewed as the infinite future period forecast conditioned on the current and past-values of the corresponding variable itself.[13] Here the notion of the long-run is not tied to a specific time frame but rather that the forecast should be far enough into the future to eliminate the influence of all types of transitory disturbances. The permanent level of the real exchange rate can be specified as

13 Huizinga (1987) and Cumby and Huizinga (1990) have employed the Beveridge-Nelson approach to determine the long-run equilibrium exchange rate.

Figure 1

Response of Real Exchange Rate to a One-Standard Deviation Shock in the Relative Price of Tradable Goods



(6) $$q_t^P = \lim_{k \to \infty} E(q_{t+k} - k\Delta\bar{q})$$

where q_t^P is the permanent component of the real exchange rate at time t and $\Delta\bar{q}$ is the unconditional mean rate of change in q_t. Equation (6) can be rewritten as

(7) $$q_t^P = q_t + E_t(\Delta q_{t+1} - \Delta\bar{q}) + E_t(\Delta q_{t+2} - \Delta\bar{q}) + \ldots.$$

We use the bivariate VAR consisting of twelve monthly lags of the real effective exchange rate and relative tradable price differential to generate the conditional forecast. To obtain the sum of the conditional infinite forecast, we rearrange the VAR into its first-order companion matrix form

(8) $$\begin{bmatrix} \Delta q_t \\ \Delta q_{t-1} \\ \cdot \\ \cdot \\ \cdot \\ \Delta q_{t-11} \\ \Delta ptng_t \\ \Delta ptng_{t-1} \\ \cdot \\ \cdot \\ \cdot \\ \Delta ptng_{t-11} \end{bmatrix} = \begin{bmatrix} \alpha_{1,1} & \alpha_{1,2} & \ldots\ldots\ldots & \alpha_{1,24} \\ 1 & 0 & \ldots\ldots\ldots & 0 \\ \cdot & & & \cdot \\ \cdot & & & \cdot \\ \cdot & & & \cdot \\ \cdot & & & \cdot \\ \cdot & & & \cdot \\ \beta_{1,1} & \beta_{1,2} & \ldots\ldots\ldots & \beta_{1,24} \\ 0 & \ldots & 0\ 1\ 0 \ldots & 0 \\ \cdot & & & \cdot \\ \cdot & & & \cdot \\ \cdot & & & \cdot \\ 0 & \ldots\ldots\ldots & & 0\ 1\ 0 \end{bmatrix} \begin{bmatrix} \Delta q_{t-1} \\ \Delta q_{t-2} \\ \cdot \\ \cdot \\ \cdot \\ \Delta q_{t-12} \\ \Delta ptng_{t-1} \\ \cdot \\ \cdot \\ \cdot \\ \Delta ptng_{t-12} \end{bmatrix} + \begin{bmatrix} v_{1t} \\ v_{2t} \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ \cdot \\ v_{24t} \end{bmatrix}$$

where all the variables are expressed as deviations from the sample means. (8) can be written more compactly as

(9) $$Z_t = AZ_{t-1} + V_t$$

The compansion matrix form has the useful property that the infinite period forecast of Δq can be obtained from[14]

(10) $$e' \sum_{j=1}^{\infty} A^j Z_t = e' A(I - A)^{-1} Z_t$$

where e' is a selection vector (1,0 ....0).

14 See Campbell and Shiller (1989) and Cochrane (1990).

The permanent component of the real exchange rate at t can then be obtained as

(11) $$q_t^p = q_t + e'A(I-A)^{-1}Z_t$$

The plots of the permanent/equilibrium real effective exchange rates together with the actual real effective exchange rates for the five currencies are given in Figures 2 to 6.

For all the currencies, the movements of the actual effective exchange rate generally track the movements of its permanent component well. While the deviation of the actual real exchange rate from the permanent level tend to be serially correlated, such discrepancies tend to be eliminated over time. This observation is consistent with the long horizon mean reverting behaviour that was found earlier on.

While there are long-run parallel movements between the actual and the permanent real exchange rates for the five currencies, there exists considerable differences in the way in which the monthly changes in the actual exchange rate reflect the month-to-month stochastic shifts in the equilibrium exchange rate. Table 9 shows the R^2 of a regression of percentage change in the real exchange rate against the percentage change in the estimated permanent component. Around 80 percent of the monthly variation in the Korean and the Thai real exchange rates are accounted by movements in the permanent components. On the other hand, only 30 percent of the monthly fluctuations in the Malaysian Ringgit real exchange rate are explained by changes in the permanent component, suggesting considerable "over shooting" and "undershooting" episodes in the broad swing of the Ringgit real exchange rate.

VI. Conclusions

In this paper, we set about analysing the behaviour of the real effec-

Table 9

R^2 OBTAINED FROM REGRESSING PERCENTAGE CHANGE IN REAL EXCHANGE RATE AGAINST THE PERCENTAGE IN THE PERMANENT COMPONENT

South Korea	Malaysia	Philippines	Singapore	Thailand
0.80	0.31	0.66	0.57	0.82

Figure 2

Actual and Permanent Component of the Real Effective Exchange Rate — Korea

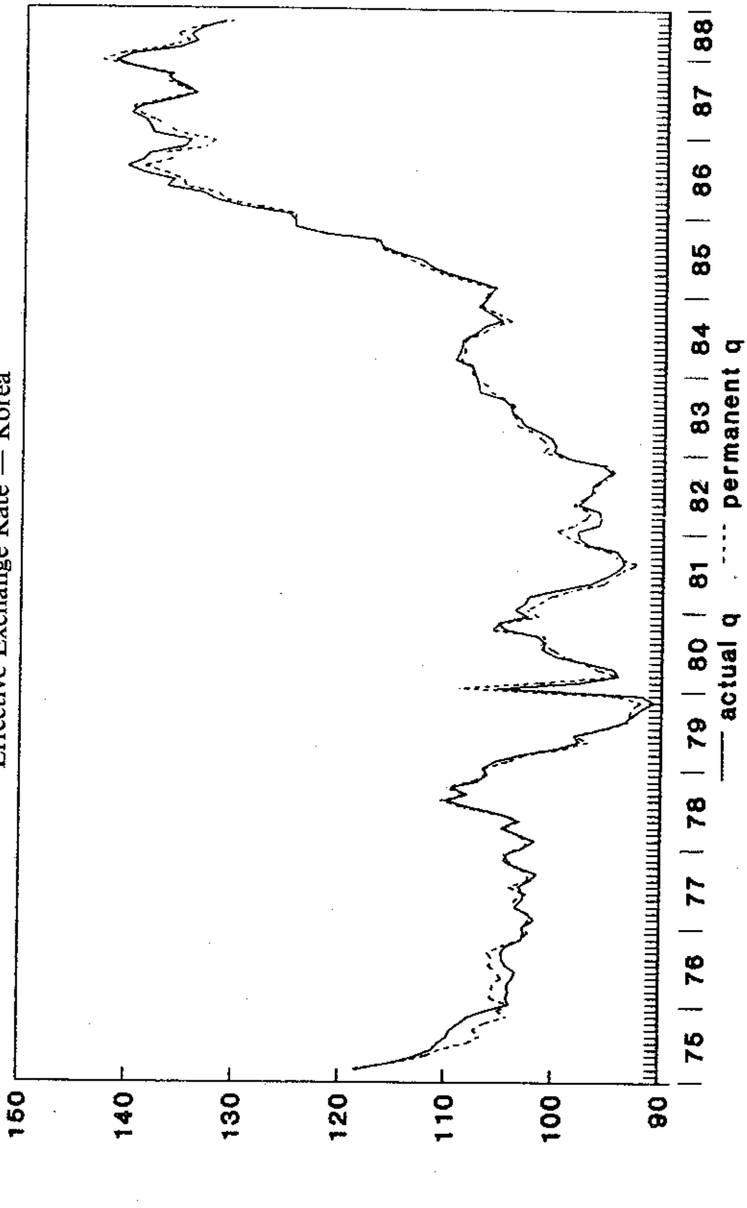
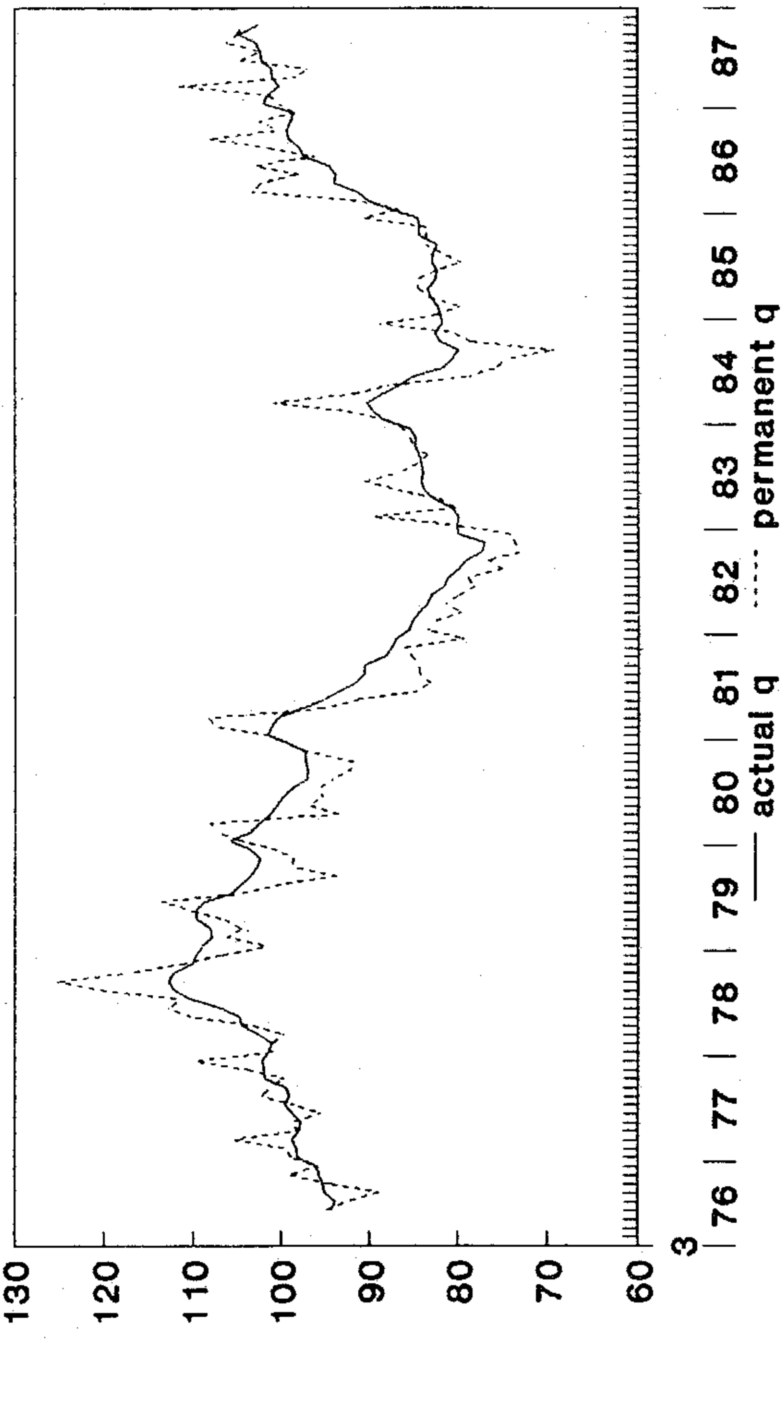


Figure 3
Actual and Permanent Component of the Real
Effective Exchange Rate — Malaysia



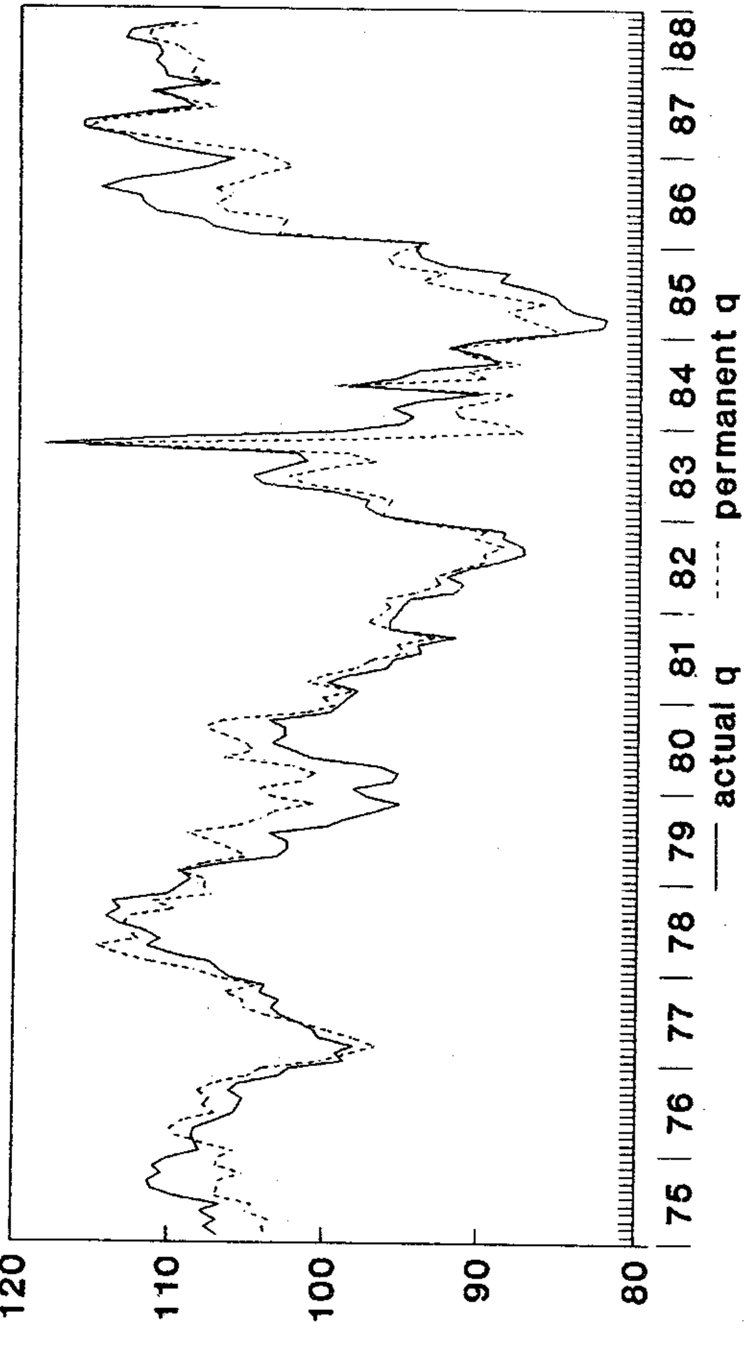


Figure 4
Actual and Permanent Component of the Real
Effective Exchange Rate — Philippines

Figure 5

Actual and Permanent Component of the Real Effective Exchange Rate — Singapore

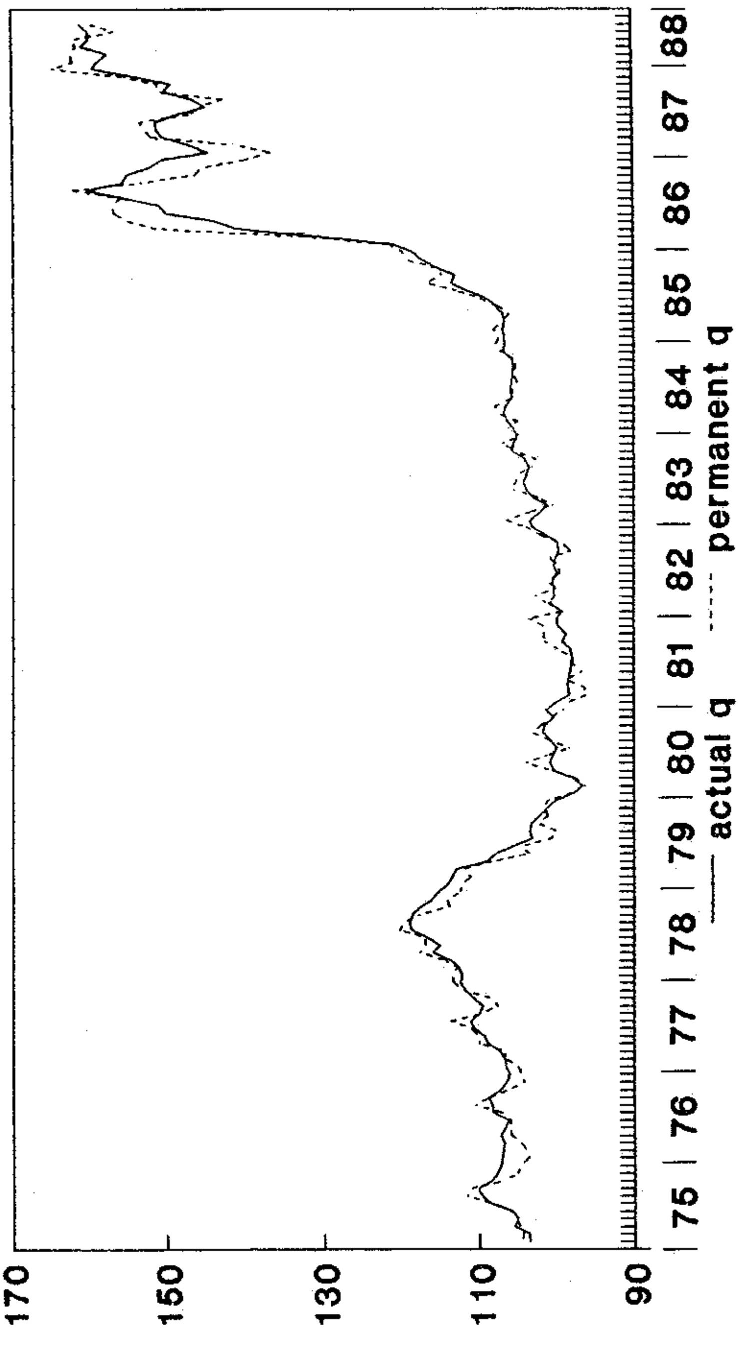
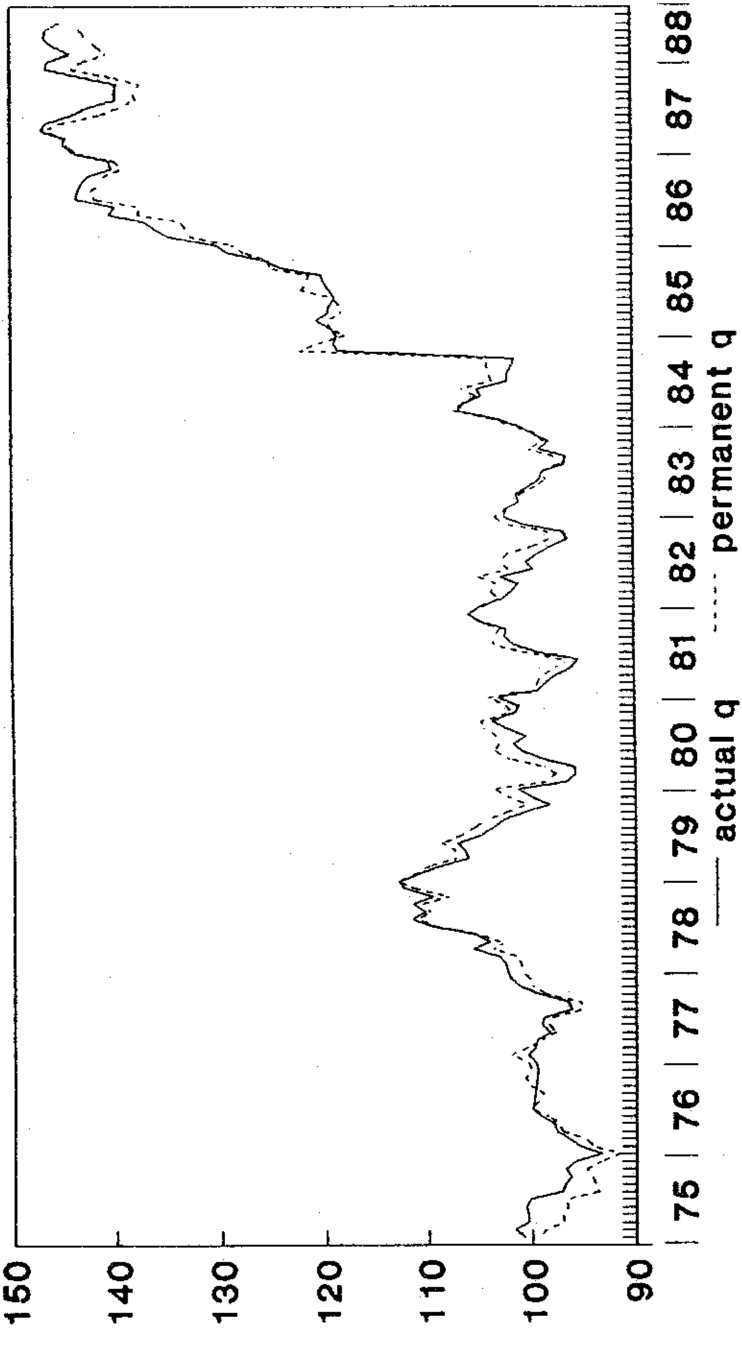


Figure 6

Actual and Permanent Component of the Real
Effective Exchange Rate — Thailand



tive exchange rate of five East Asian economies. We found that the purchasing power parity relationship does not hold both in the short and in the long run. In the short run, the variability of the real effective exchange rates are dominated by variability of the nominal effective exchange rates and there is no offsetting movements between the nominal effective exchange rates and the relative prices to stabilizes the real exchange rates. In the long run, the domestic tradable price level do not cointegrate with the international price level.

Nominal exchange rate shocks appear to have persistent effect on the real exchange rate, at least in the low inflation economies like Malaysia and Singapore. This suggests that the nominal exchange rate in these countries can be used to influence the adjustment of the real exchange rate towards its equilibrium path. Another issue that is important for the conduct of the exchange rate policy is the extent to which the swing the real exchange rate predictable. Our analysis indicates that real exchange rates of the five countries tend to revert, in the longer run, to the equilibrium path determined by shifts in the relative price of tradable to non-tradable goods. This implies some degree of predictability of the real exchange rates in the longer run.

References

Balassa, B., "The Purchasing Power-Parity Doctrine: A Reappraisal," *Journal of Political Economy,* 72, 1964, 584-596.

Banerjee, A., D.F. Hendry and G.W. Smith, "Exploring Equilibrium Relationship in Econometrics through Static Models: Some Monte Carlo Evidence," *Oxford Bulletin of Economics and Statistics,* 48, 1986, 253-277.

Beveridge, S. and C.R. Nelson, "A New Approach to Decomposition of Economic Time Series into Permanent and Transitory Components with Particular Attention to the Measurement of the 'Business Cycle'," *Journal of Monetary Economics,* 7, 1981, 151-174.

Campbell, J.Y. and R.J. Shiller, "Dividend-Price Ratio and Expectations of Future Dividends and Discount Factors," *The Review of Financial Studies,* 1, 1989, 195-228.

Clements, K.W. and J.A. Frenkel, "Exchange Rates, Money and Relative Prices: The Dollar-Pound in the 1920's," *Journal of International Economics,* 10, 1980, 249-262.

Cochrane, J.H., "How Big Is the Random Walk in GNP," *Journal of Political Economy,* 96, 1988, 893-919.

Cochrane, J.H., "Univariate vs Multivariate Forecasts of GNP Growth and Stock Returns: Evidence and

Implications for Persistence of Shocks, Detrending Methods, and Tests of the Permanent Income Hypothesis," mimeo, University of Chicago, Department of Economics, 1991.

Cumby, R.E. and J. Huizinga, "Predictability of Real Exchange Rate Changes in the Short and Long Run," mimeo, University of Chicago, Graduate School of Business, 1990.

——— and M. Obstfeld, "International Interest Rate and Price Level Linkages Under Flexible Exchange Rates: A Review of Evidence," in J.O. Bilson and R.C. Marston (eds.), *Exchange Rate Theory and Practice,* Chicago, University of Chicago Press, 1984.

Dickey, D.A. and W.A. Fuller, "Likelihood Ratio Statistics for Autoregressive Time Series with a Unit Root," *Journal of American Statistical Association,* 74, 1979, 355-367.

Durlauf, S.N. and R.W. Staiger, "Compositional Effects of Government Spending in a Two-Country, Two-Sector Production Model," *Journal of International Economics,* 29, 1990, 333-347.

Dwyer, J., "The Tradeable Non-Tradable Dichotomy: A Practical Approach," Australian Economic Papers, 31, 1992, 443-459.

Edwards, S., "Temporary Term-of-Trade Disturbances, the Real Exchange Rate and the Current Account," *Economica,* 56, 1989, 343-357.

Engle, R.F. and C.W.J. Granger, "Cointegration and Error Correction: Representation, Estimation and Testing," *Econometrica,* 55, 1987, 251-276.

Engle, R.F. and B.S. Yoo, "Forecasting and Testing in Cointegrated System," *Journal of Econometrics,* 35, 1987, 43-59.

Fama, E.F. and K.R. French, "Permanent and Temporary Components of Stock Prices," *Journal of Political Economy,* 96, 1988, 246-273.

Frenkel, J.A. and A. Razin, *Fiscal Policies and the World Economy: An Intertemporal Approach,* Cambridge, MA, MIT Press, 1987.

——— and K.A. Proof, "Chartists, Fundamentalists and Trading in the Foreign Exchange Market," *American Economic Review,* 80, 1990, 151-185.

Fuller, W.A., *International to Statistical Time Series,* New York, John Wiley & Sons, 1976.

Genberg, H., "Purchasing Power Parity Under Fixed and Flexible Exchange Rates," *Journal of International Economics,* 8, 1978, 247-276.

Huizinya, J., An Empirical Investigation of the Long Run Behavior of Real Exchange Rates," in K. Brunner and A. Meltzer (eds.), *Carnegie-Rochester Series on Public Policy,* Amsterdam, North Holland, 1987, 149-214.

Kim, M.J., C.R. Nelson and R. Startz, "Mean Reversion in Stock

Prices] A Reappraisal of the Empirical Evidence," mimeo, 1989.

Koedijk, K.G. and P. Schotman, "How to Beat the Random Walk: An Empirical Model of Real Exchange Rates," *Journal of International Economics,* 29, 1990, 311-325.

Krugman, P.R., "Equilibrium Exchange Rates," in W.H. Branson (eds.), *International Policy Coordination and Exchange Rate Fluctuations,* Chicago, University of Chicago Press, 1990, 149-187.

Lothian, J.R., "A Century Plus of Yen Exchange Rate Behavior," *Japan and the World Economy,* 2, 1990, 47-70.

Marston, R.C., "Real Exchange Rates and Productivity Growth in the United States and Japan," in S.W. Arndt and J.D. Richardson (eds.), *Real-Financial Linkages Among Open Economies,* Cambridge, MA, MIT Press, 1987.

Mussa, M., "Nominal Exchange Rate Regime and the Behavior of the Real Exchange Rates: Evidence and Implications," in K. Brunner and A.H. Meltzer (eds.), *Carnegie-Rochester Series on Public Policy,* Amsterdam, North Holland, 1986, 117-214.

Newey, W.K. and K.D. West, "A Simple, Positive Semi-Definite, Heteroskedasticity and Autocorrelation Consistent Covariance Matrix," *Econometrica,* 55, 1987 703-708.

Perron, P., "The Great Crash, The Oil Price Shock, and the Unit Root Hypothesis," *Econometrica,* 57, 1989, 1361-1401.

Phillips, P.C.B., "Understanding Spurious Regressions in Econometrics," *Journal of Econometrics,* 33, 1986, 311-340.

______, "Time Series Regressions with Unit Root," *Econometrica,* 55, 1987, 277-301.

______and S. Ouliaris, "Asymptotic Properties of Residual Based Tests for Cointegration," *Econometrica,* 58, 1990, 165-193.

Poterba, J.M. and L.H. Summers, "Mean Reversion in Stock Prices: Evidence and Implications," *Journal of Financial Economics,* 22, 1988, 27-59.

Sachs, J.D. and S.M. Collins, (eds.) *Developing Country Debt and Economic Performance,* Vol. 3, Chicago, University of Chicago Press, 1989.

Stock, J., "Asymptotic Properties of Least Squares Estimators of Cointegrating Vectors," *Econometrica,* 55, 1987, 1035-1056.