

Long Run Real Exchange Rate Overvaluation: Evidence from Asia

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Abstract:

This paper studies the bilateral real exchange rates of 5 East Asian economies vis-à-vis the U.S. dollar and tests whether permanent changes in these real exchange rates can be explained by permanent changes in the relative prices of nontraded- to traded-goods. We find that, with the exception of the South Korean-U.S. real exchange rate, the other 4 real exchange rates are cointegrated with the relative prices of nontraded- to traded-goods. We propose a new measure of real exchange rate misalignment based on this long run relationship, and calculate estimates of misalignment prior to the Asian currency crisis. For all 4 countries, the proposed measure yields protracted and economically significant overvaluations prior to the crisis. By contrast, a PPP-based measure yields the counterintuitive result that the real exchange rates were significantly *undervalued* prior to the Asian currency crisis.

JEL Classification System: F31 (Foreign Exchange), C22 (Time Series Models)

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1. Introduction

The real exchange rate, defined as the relative price of domestic goods to foreign goods, is a key relative price for any open economy.¹ Economists and policy makers have long held the view that a misaligned real exchange rate significantly reduces the welfare of a country (Edwards 1988, Willet 1986). Overvalued real exchange rates, for example, are often blamed for deteriorating the trade balance of a country and making it more vulnerable to economic instability. Recent work (Kaminsky, Lizondo and Reinhart 1998) suggests that a rapid appreciation of the real exchange rate, presumably a reflection of misalignment, is among the most successful predictors of currency crises.

Identifying real exchange rate misalignments is therefore one of the most important and challenging tasks confronting students of international economics. The theory of Purchasing Power Parity (PPP) has traditionally been viewed as determining the long run equilibrium exchange rate, and is often used to assess exchange rate misalignments (Artus, 1978). However, an important drawback of the PPP-based misalignment measures is that deviations of exchange rates from their PPPs tend to be near permanent. Although the recent literature on PPP has documented some evidence of mean reversion, the unit-root nonstationarity of real exchange rates is virtually a stylized fact in international finance.²

¹ There are several different definitions of the real exchange rate currently in use. This paper uses the “PPP-definition” of the real exchange rate. The reader is referred to Edwards (1988) for a discussion of the various definitions and their relationships.

² For instance, Baillie and McMahon (1989), Corbae and Ouliaris (1988), Engel (2000), Mark (1990), Meese and Rogoff (1988), O'Connell (1998), and several other researchers document evidence suggesting that real exchange rates are unit-root nonstationary,

One possible explanation for the persistent deviations from PPP that has received attention in the recent literature is the presence of nontradable goods in the general price indices used in constructing real exchange rates.³ If PPP holds only for tradables, then any factor that permanently changes the relative price of nontradables (such as a permanent productivity shock) will cause a permanent shift in the real exchange rate. Repeated permanent shocks will therefore create long run comovements between relative prices of nontradable goods and the real exchange rate.

The main objectives of this paper are to test whether changes in the relative price of nontradables can account for the long run changes in the real exchange rates of Asian economies, and if so, to utilize these long run comovements to construct an alternative measure of real exchange rate misalignment. Although measurement of exchange rate misalignments is (justifiably) a very contentious issue, there are at least four reasons why this measure may be considered reasonable and useful.

First, it is compatible with a wide variety of economic models that use the nontradables approach to real exchange rate determination, such as the celebrated Balassa-Samuelson [Balassa (1964) and Samuelson (1964)] model. It is therefore grounded in economic theory.

thereby implying that PPP fails to hold even in the long-run. While several recent studies (e.g. Frankel and Rose 1995, Lothian and Taylor 1996) have found evidence of mean reversion in real exchange rates, the slow rate of convergence to PPP suggests that persistent real disturbances could plausibly be playing an important role. See the surveys by Froot and Rogoff (1995) and Rogoff (1996) for further details.

³ See, for example, Chinn (1997), De Gregorio, Giovannini and Wolf (1994), DeLoach (1997), Engel (2000) and Kakkar and Ogaki (1999).

Second, given that the real exchange rate is cointegrated with the relative price of nontradables, any gap between the actual real exchange rate and its estimated equilibrium value is not “sustainable” and will eventually be reversed. This is a natural requirement for any estimate of a “long run equilibrium value”, but is not satisfied by PPP-based measures of overvaluation.

Third, this framework allows for measurement errors in the prices of tradables and nontradables. This is important because final goods can seldom be characterized as purely traded or purely nontraded. Most goods that are considered tradable, for example, have a nontraded component in the form of retailing services.

Fourth, the proposed measure is easy to compute and only utilizes data that are available over a long time period for several countries.

Our results suggest that, for four of the five Asian economies considered here, their bilateral real exchange rates vis-à-vis the U.S. dollar are indeed cointegrated with the relative prices of nontradable to tradable goods. We apply the nontradables based measure of real exchange rate misalignment to these four economies prior to the Asian currency crisis of 1997, and compare the results to a PPP-based measure of misalignment. Somewhat surprisingly, the PPP-based misalignment measure indicates that the Indonesian, Thai and Philippine currencies were *undervalued* prior to the crisis! By contrast, the proposed nontradables-based measure generally indicates persistent and economically significant overvaluations over a three-year period prior to the crisis.

In related work, Chinn (1998) has also examined the evidence for exchange rate misalignments for Asian economies prior to the crisis. In addition to the PPP-based measures, he computes a measure of misalignment based on a monetary model of

exchange rate determination. He finds that the estimated misalignments do not match the prior expectations of an overvaluation very well.⁴ An advantage of our framework over the monetary model is its parsimony. Since the model only requires data for three variables (the bilateral real exchange rate, the domestic relative price of nontradables and the foreign relative price of nontradables), this also enables the use of a longer span of data which is important for a reasonable performance of cointegration-based models (Pierse and Snell, 1995).⁵

2. Model

Consider a world economy with two countries: country H is the home country and country F is the foreign country. In each country, there are two goods: good T is tradable and good N is nontradable. Assume that the general price index of a country, measured by the GDP-deflator, can be expressed as a geometric average of prices of the traded and nontraded goods. That is:

$$P_j^{GDP} = c_j^{GDP} [P_j^N]^{\alpha_j} [P_j^T]^{(1-\alpha_j)}, \quad j = H, F, \quad (1)$$

⁴ Chinn (1998) also augments his monetary model with the relative price of nontradables, but he does not find this variable to be significant in the case of Indonesia, Malaysia, Singapore and Thailand. Section 4 provides a comparison with his results.

⁵ Chinn's sample ranges from 7 years of quarterly data for Indonesia, to 20 years of quarterly data for Singapore, with an average of about 13 years. By contrast, our sample ranges from 23 years of annual data for Singapore to 49 years of annual data for the Philippines, with an average of 35 years.

where α_j is the share of nontradables in the price index. c_j^{GDP} is any factor that causes the geometric average of traded and nontraded goods prices to deviate from the price level, such as a measurement error. It is assumed that c_j^{GDP} is stationary.

Let E be the nominal exchange rate: E units of the domestic currency purchase one unit of the foreign currency. The relative price of nontradable goods in terms of tradable goods is denoted by

$$Q_j = \frac{P_j^N}{P_j^T}, j = H, F. \quad (2)$$

The real exchange rate, E_r , is defined by the general price index:

$$E_r = \left(\frac{P_H^{GDP}}{E P_F^{GDP}} \right). \quad (3)$$

An increase in the real exchange rate increases the price of home goods relative to foreign goods, and thus corresponds to a real appreciation. Since traded goods across countries may not be identical, PPP may not hold even for the tradables in the short run. It is assumed that PPP holds for the tradables in the long run:

$$\ln(P_H^T) = \ln(E) + \ln(P_F^T) + u, \quad (4)$$

where u is a stationary random variable with zero mean.⁶ Combining equations (1) through (4), the real exchange rate may be written as

$$\ln(E_r) = \theta + \alpha_H \ln(Q_H) - \alpha_F \ln(Q_F) + \varepsilon, \quad (5)$$

⁶ The stationarity of u implies that a deviation from PPP for tradables is transitory, and vanishes in the long run.

where $\varepsilon = \{ \ln(c_H^{GDP}) - \ln(c_F^{GDP}) \} - E\{ \ln(c_H^{GDP}) - \ln(c_F^{GDP}) \} + u$ is also a zero-mean stationary random variable, and $\theta = E\{ \ln(c_H^{GDP}) - \ln(c_F^{GDP}) \}$.⁷

Equation (5) implies that the real exchange rate will move together with the domestic and foreign relative prices of nontradables in the long run. In order to estimate equation (5), one needs measures for the relative price of nontradables in the home and foreign countries. However, it is difficult to measure the prices of *purely* traded and nontraded goods, as most final goods are likely to have *both* traded and nontraded components. To alleviate this measurement problem, two price indices that assign *different* weights to the traded and nontraded components are used.

The *Consumer Price Index (CPI)* is based on a fixed basket of goods and services consumed by the average household, and is likely to have a large share of nontraded goods in the form of retailing services, housing, transportation and other services. By contrast, the *Wholesale Price Index (WPI)* is generally limited to agricultural and manufacturing sector goods that are largely tradable. It also uses prices that exclude (nontraded) retailing services. Thus the *CPI* and *WPI* may be written as:

$$P_{CPI}^j = c_{CPI}^j [P_N^j]^{\beta_j} [P_T^j]^{1-\beta_j}, j = H, F \quad (6a)$$

$$P_{WPI}^j = c_{WPI}^j [P_N^j]^{\delta_j} [P_T^j]^{1-\delta_j}, j = H, F. \quad (6b)$$

Here c_{CPI}^j (c_{WPI}^j) is any factor that causes a deviation between the geometric average of unobserved traded and nontraded goods prices and the *CPI* (*WPI*), such as a measurement

⁷ The stationarity of ε follows from the stationarity of c_j^{GDP} and u . It is zero-mean by construction.

error. It is assumed that $1 > \beta_j > \delta_j > 0$ so that the *CPI* has a larger share of nontradables relative to the *WPI*.

As long as these shares are stable over time, the ratio of the *CPI* to the *WPI* will provide a reasonable proxy for the relative price of nontradables for the purpose of estimating a cointegrating regression. To see this, divide equation (6a) by (6b) and express the result in logs:

$$\ln(P_{CPI}^j / P_{WPI}^j) = \ln(c_{CPI}^j / c_{WPI}^j) + (\beta_j - \delta_j) \ln(Q_j), \quad j = H, F. \quad (7)$$

Solving for the unobservable $\ln(Q_j)$ from equation (7) and substituting in equation (5) gives

$$\ln(E_r) = \lambda + \left(\frac{\alpha_H}{\beta_H - \delta_H} \right) \ln(P_{CPI}^H / P_{WPI}^H) - \left(\frac{\alpha_F}{\beta_F - \delta_F} \right) \ln(P_{CPI}^F / P_{WPI}^F) + \eta, \quad (8)$$

where $\lambda = \theta + E\{\ln(c_{CPI}^H / c_{WPI}^H) - \ln(c_{CPI}^F / c_{WPI}^F)\}$ is a constant, and

$\eta = \varepsilon + \{\ln(c_{CPI}^H / c_{WPI}^H) - \ln(c_{CPI}^F / c_{WPI}^F)\} - E\{\ln(c_{CPI}^H / c_{WPI}^H) - \ln(c_{CPI}^F / c_{WPI}^F)\}$ is a zero-mean stationary random variable.

Equation (8) implies that the ratio of *CPI* to *WPI* can be used as a proxy for the relative price of nontradables. We can test the model by testing for cointegration, and checking for the signs and statistical significance of the estimated parameters.⁸ This equation forms the basis of the empirical work.

⁸ The plausibility of the magnitude of the estimated coefficients may be established as follows. Suppose that the share of nontradables in the GDP-deflator is about 0.5, about 0.7 in the *CPI*, and about 0.3 in the *WPI*. Then the coefficients of the relative price of nontradables in equation (8) should be greater than 1 in absolute value. Estimates of

3. Data and its Trend Properties

We use annual data for five Asian countries and the United States from the *IFS* CD-ROM produced by the *IMF*. The Asian countries are Indonesia, Korea, Singapore, Philippines, and Thailand.⁹ Indonesia, Korea, Philippines and Thailand essentially maintained a currency peg against the U.S. dollar. During the crisis, all four countries were forced to abandon the peg, and their currencies suffered substantial depreciations. Asian countries are treated as the home country and the United States as the foreign country.

The statistical testing begins by examining the evidence for the existence of unit roots in the real exchange rate and the proxy for the relative prices of nontradables. Table 1 reports the results of the Phillips-Perron (1988) Z_t test and Park's $J(1,5)$ test for the null hypothesis of a unit root against the alternative of trend stationarity. Neither test rejects the null hypothesis at the 10% significance level for either variable for any of the countries. Thus the assumptions that, these Asian countries' bilateral real exchange rates vis-à-vis the U.S. dollar as well as the U.S. and Asian relative prices of nontradables, are unit-root nonstationary and possess stochastic trends, are supported by these results.

Prior to estimating equation (8), it is also necessary to ensure that the coefficients of home and foreign relative prices of nontradables are identified. If the relative prices of nontradables in the home (Asian) country and the U.S. share a common stochastic trend, then their coefficients cannot be identified by any econometric method. Table 2 tests whether this is the case using the Phillips-Ouliaris (1990) t -ratio test and Park's (1990)

these coefficients significantly smaller than 1 in absolute value may be construed as evidence against the model.

⁹ Malaysia was not included because its available sample size was only about 10 years.

$I(1,5)$ test for the null hypothesis of no stochastic cointegration. The Phillips-Ouliaris test is not significant for any of the five countries at conventional significance levels, whereas the $I(1,5)$ test is significant at the 10% level only for Thailand. Thus, with the exception of Thailand, for which the evidence is mixed, for all other countries the relative price of nontradables does not appear to be cointegrated with the U.S. relative price of nontradables. Figure 1 shows a plot of the Thai relative price of nontradables and its estimated value based on the U.S. relative price of nontradables. Most of the long run movements in the Thai relative price of nontradables can be explained by corresponding movements in the U.S. relative price, further supporting the statistical evidence in favor of cointegration.

4. Empirical Results

Having verified that the conditions necessary for testing the existence of a long run relationship between the real exchange rate and the relative prices of nontradables are empirically supported, the next step is to estimate equation (8) via a cointegrating regression. Since the economic model implies cointegration, it is desirable to test the null hypothesis of cointegration to control the probability of rejecting a valid economic model. Park's (1992) Canonical Cointegrating Regressions (CCR) procedure is used to test the null hypothesis of stochastic cointegration and the deterministic cointegration restriction.¹⁰ The CCR estimators are asymptotically efficient and have asymptotic

¹⁰ See Campbell and Perron (1991) and Ogaki and Park (1997) for definitions of stochastic and the deterministic cointegration restriction. The results reported in this paper use Ogaki's (1993) Gauss CCR-package.

distributions that can essentially be considered as normal distributions, so that their standard errors can be interpreted in the usual way.

Table 3 reports the results of the estimation of equation (8) by the CCR procedure.¹¹ With the exception of South Korea, the $H(1,2)$ test does not reject the null hypothesis of stochastic cointegration for any of the countries at the 1% significance level. The $H(1,3)$ test, which also maintains stochastic cointegration as the null hypothesis, is significant for Indonesia and Thailand at the 5% significance level, but not for the other countries. The $H(0,1)$ statistic tests the deterministic cointegration restriction implied by the model. With the exception of South Korea, the $H(0,1)$ statistic does not reject the deterministic cointegration restriction at conventional significance levels for any of the countries. Thus, with the exception of South Korea, both the stochastic and deterministic cointegration restrictions implied by the model are supported empirically.

The coefficients of the relative prices of nontradables have the signs predicted by the model for all the countries. For the Indonesia-U.S. regression, the coefficient of the Indonesian relative price of nontradables is not statistically significant, and for the Thailand-U.S. regression (Case 1) both coefficients are insignificant. For all other regressions, the coefficients are statistically significant at conventional levels. Their magnitudes are also plausible after taking into account the standard errors. This is further evidence in favor of the model.

¹¹ Since the Thai relative price of nontradables may be cointegrated with its U.S. counterpart, we estimate two cointegrating regressions for Thailand, one with both relative prices, and another with only the Thai relative price.

The evidence in favor of cointegration between bilateral real exchange rates and relative prices of nontradables suggests that real exchange rate must eventually revert back to its “long run equilibrium value” implied by the model. Thus, given sufficient time, any disequilibrium will vanish. However, the speed with which such departures from equilibrium are eliminated may well vary across countries. It is possible to estimate the speed of reversion to equilibrium by estimating an error correction model.

Table 4 reports the results of regressing the first difference of the real exchange rate on the lagged error correction term and lags of the first differences of the relative prices of nontradables.¹² The first column of Table 4 reports the estimated coefficient of the error correction term. The coefficient has the expected negative sign for countries and is also statistically significant with the exception of Thailand in Case 2. The second column calculates the implied half-life of a deviation from the long run equilibrium. The half-life is less than 2.5 years for Indonesia, Singapore and Thailand (Case 1), but is nearly 5 years for the Philippines. These estimates are quite reasonable given that most studies find a half-life of nearly 5 years for convergence to Purchasing Power Parity. Thus, relative to the constant equilibrium value of the real exchange rate implied by PPP, convergence to the equilibrium value implied by the nontradables-based model is significantly faster.

As mentioned in the Introduction, the long run relationship between the real exchange rate and its equilibrium value implied by the model provides a natural estimate of the extent of “misalignment”. We exploit this relationship to calculate the implied

¹² We do not estimate an error correction model for South Korea given that the evidence does not support a long run relationship between the real exchange rate and relative prices of nontradables for this economy.

degree of misalignment prior to the Asian currency crises of 1997. The first column of Table 5 reports the average overvaluation during the three years preceding the currency crisis (i.e., 1994 through 1996). For all four countries, and in both cases for Thailand, the nontradables-based model suggests an overvaluation, varying from a low of about 7% (Thailand under Case 1) to a high of nearly 20% (Philippines). The fourth column of Table 5 reports the estimated overvaluation at the end of 1996, and these numbers are slightly larger than the corresponding average reported in the first column. This suggests that not only was overvaluation persistent over the 3 years preceding the crisis, it was also increasing in magnitude.

Since PPP is a widely used benchmark for calculating exchange rate overvaluation, it is useful to compare these numbers to those derived from the traditional PPP model. The second and fifth columns of Table 5 show the estimated overvaluation under the assumption that PPP holds in the long run. The equilibrium real exchange rate is assumed to be a constant and its value is estimated by the sample mean of the real exchange rate. The PPP based estimates suggest that the Indonesian and Philippine real exchange rates were significantly *undervalued* prior to the crisis (-12 % and -16%, respectively)! The Singaporean real exchange rate appears to be significantly overvalued (25%), whereas the Thai real exchange rate is close to its equilibrium value. These results are in stark contrast to those based on the nontraded-goods model, and are also contrary to the prior expectations of most economists and financial analysts.

We also compare our results to those based on the monetary model, as reported by Chinn (1998). Chinn's estimates of overvaluation are based on the monetary model of exchange rates and are reported in the third (and sixth) column of Table 5. For the three

years prior to the crisis, Chinn finds a modest overvaluation of less than 5% for Indonesia and Thailand, a significant undervaluation for the Philippines (-26%) and a huge overvaluation for Singapore (45%). On the eve of the crisis, Indonesian and Thai real exchange rates appear to be in close to equilibrium, whereas the Philippine and Singapore real exchange rates are significantly undervalued and overvalued, respectively. These estimates also do not appear to be intuitively plausible.

These results are interesting because they imply that, of the three overvaluation measures considered here, only the one based on the nontraded-goods model yields estimates that are consistent with prior expectations of an overvaluation before the Asian currency crisis.

5. Conclusions

This paper studies whether permanent changes in relative prices of nontradables can account for permanent changes in the real exchange rates of 5 East Asian economies. It is found that, with the exception of the South Korean case, all other real exchange rates are cointegrated with the relative prices of nontradable. A new measure of real exchange rate misalignment, based on the long run relationship between the real exchange rate and the relative prices of nontradables, is also considered. The proposed measure is easy to compute, and when applied to the 4 economies for which cointegration is found, yields intuitive plausible estimates of persistent overvaluation prior to the Asian currency crisis of 1997. By contrast, PPP-based measures of misalignment indicate an undervalued real exchange rate prior to the crisis.

In using the nontradables-based measure of misalignment, we implicitly allow for both supply shocks (e.g., differential productivity shocks in the tradable and nontradable

goods' sectors) as well as demand shocks (e.g., changes in preferences) to affect the equilibrium real exchange rate by altering the relative price of nontraded-goods. By contrast, the PPP approach, although very popular, assumes that all shocks are necessarily temporary. This is unlikely to be true especially with regard to shocks to technology.¹³

Our approach is also consistent with recent work that views the real exchange rate as comprising both permanent and transitory components, as opposed to the conventional dichotomy that views the real exchange rate as either a purely stationary process or a random walk. Mark and Choi (1997) show that models in which the long run real equilibrium exchange rate is identified as the permanent component of the real exchange rate outperform models in which long run PPP holds in terms of out-of-sample forecasting power. Generating an artificial century long time-series of the U.S.-UK real exchange rate, Engel (2000) shows through extensive Monte Carlo simulations that the real exchange rate contains an economically significant permanent component associated with the relative price of nontraded-goods that may go undetected by standard unit root tests.

Given the importance placed on the role of real exchange rate overvaluation in currency and balance-of-payments crises, and the limited empirical support for long run

¹³ For instance, Kakkar (2002) documents evidence suggesting that sectoral total factor productivities are unit-root nonstationary, and are cointegrated with the relative prices of nontradables in fourteen OECD economies.

PPP, it is hoped that the proposed measure will be found to be a useful and practical alternative to the existing PPP-based measures.

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Table 1**Unit Root Tests for Real Exchange Rates and Relative Prices of Nontraded-Goods**

Country/ <u>Sample</u>	$\ln(E_r)$		$\ln(Q_N)$	
	Z_t^a	$J(1,5)^b$	Z_t^a	$J(1,5)^b$
Indonesia (1971-96)	-1.812	10.800	-1.533	1.967
Korea (1966-96)	-2.198	0.896	-1.108	7.971
Philippines (1949-96)	-1.466	1.883	-1.384	3.018
Singapore (1974-96)	-0.352	13.632	-1.915	3.357
Thailand (1953-96)	-2.288	0.837	-1.937	1.365
USA (1949-96)	-----	-----	-1.491	2.897

^a Z_t denotes the Phillips-Perron (1988) t-ratio test for the null hypothesis of a unit root against the alternative of trend stationarity. Critical values used are from MacKinnon (1991)

^b $J(1,5)$ denotes Park's (1990) test for the null hypothesis of a unit root against the alternative of trend stationarity. The 1%, 5% and 10% critical values are 0.1228, 0.2950 and 0.4520, respectively. These are taken from Ogaki (1993).

Table 2

**Tests for the Null Hypothesis of No Stochastic Cointegration between Domestic and
US Relative Prices of Nontraded-Goods**

<i>Country/Sample</i>	Z_t^a	$I(1,5)^b$
Indonesia (1971-1996)	-2.669	2.793
Korea (1967-1996)	-1.936	4.028
Philippines (1949-1996)	-2.411	1.011
Singapore (1974-1996)	-1.811	3.193
Thailand (1953-1996)	-3.135	0.304 [*]

^a Z_t denotes the Phillips-Ouliaris (1990) t-ratio test for the null hypothesis of no stochastic cointegration. Critical values used are from MacKinnon (1991)

^b $I(1,5)$ denotes Park's (1990) test for the null hypothesis of no stochastic cointegration. The 1%, 5% and 10% critical values are 0.1027, 0.2506 and 0.4984, respectively. These are taken from Ogaki (1993).

^{*} Significant at the 10% level.

Table 3

**Canonical Cointegrating Regressions between Real Exchange Rates and Relative
Prices of Nontraded-Goods**

<u>Sample</u>	β_1^a	β_2^a	$H(0,1)^b$	$H(1,2)^b$	$H(1,3)^b$
Indonesia-US					
1971-96	0.3261 (0.2972)	-1.3625 (0.5158)	2.3912 (0.1220)	5.0091 (0.0252)	12.9424 (0.0015)
Philippines-US					
1949-96	2.1331 (0.4805)	-1.5582 (0.6440)	0.6945 (0.4046)	0.9603 (0.3271)	5.4244 (0.0664)
Singapore-US					
1974-96	1.7261 (0.4218)	-2.5146 (0.6891)	4.6652 (0.0308)	0.1604 (0.6888)	2.0088 (0.3663)
Thailand-US					
1953-96 (Case 1)	0.0931 (0.6847)	-0.4543 (0.4480)	0.6362 (0.4251)	0.3059 (0.5802)	0.8627 (0.6496)
1953-96 (Case 2)	-2.2177 (0.7301)	-----	0.0015 (0.9691)	2.2010 (0.1379)	3.7247 (0.1553)
Korea-US					
1966-96	3.7152 (1.3626)	-3.7294 (1.9722)	8.3310 (0.0039)	36.4713 (0.0000)	82.1209 (0.0000)

^a Standard errors are in parenthesis.

^b $H(0,1)$ tests the null hypothesis of the deterministic cointegration restriction. $H(1,2)$ and $H(1,3)$ test the null hypothesis of stochastic cointegration. P -values are in parenthesis.

Table 4**Estimated Speed of Reversion to the Long Run Equilibrium Real Exchange Rate**

$$\Delta \ln(E_r)_t = \alpha_0 + \alpha_1(ECT)_{t-1} + \sum_{i=1} \lambda_{1i} \Delta \ln(E_r)_{t-i} + \sum_{i=1} \lambda_{2i} \Delta \ln(Q_H)_{t-i} + \sum_{i=1} \lambda_{3i} \Delta \ln(Q_F)_{t-i} + \varepsilon_t$$

Country	α_1^a	Implied Half-Life Estimate ^b (Years)
Indonesia	-0.807** (0.180)	0.86
Philippines	-0.143* (0.080)	4.85
Singapore	-0.425* (0.201)	1.63
Thailand (Case 1)	-0.283** (0.129)	2.45
Thailand (Case 2)	-0.042 (0.080)	16.5

^a α_1 is the coefficient of the error correction term, *ECT*, and measures the speed of adjustment to the long-run equilibrium. For e.g., $\alpha_1 = -0.283$ means that, holding all else constant, 28.3% of the overvaluation in the real exchange rate relative to its long-run equilibrium level, is eliminated in one year. Standard errors are in parenthesis.

^b The half-life refers to the duration of time required for 50% of the real exchange rate overvaluation to dissipate. It is calculated by solving the following exponential decay equation for *T*: $e^{\alpha_1 T} = 0.5$.

* and ** indicate statistical significance at the 10% and 5% levels of significance, respectively.

Table 5**Estimated Percentage Overvaluation of the Real Exchange Rate Prior to Crisis**

Country	Average Overvaluation (1994-96)			Overvaluation (1996)		
	Nontraded Model	PPP Model	Monetary Model [*]	Nontraded Model	PPP Model	Monetary Model [*]
Indonesia	10.72	-12.12	4.67	12.68	-9.70	0.93
Philippines	19.98	-16.16	-26.45	20.09	-10.58	-24.63
Singapore	12.06	24.32	45.39	11.65	27.26	35.41
Thailand (Case 1)	6.54	-0.0505	0.0371	10.01	0.00	0.01
Thailand (Case 2)	18.10	-0.0505	0.0371	23.01	0.00	0.01

* These figures are based on Chinn's (1998, Table 7) estimation of a monetary model using quarterly data. The Average Overvaluation is measured for the period 1995Q2 to 1997Q1, and the Overvaluation is measured for 1997Q1.