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ABSTRACT

Five East Asian currencies -- the Indonesian rupiah, Korean won, Singapore dollar, Taiwanese dollar, and the Thai baht -- are modeled in the framework of a monetary specification augmented by the relative price of nontradables. This relative price variable proxies for the Balassa-Samuelson effect in East Asian real exchange rates identified in Chinn (1997b). All of the currencies fit the long run implications of various types of monetary models, according to Johansen (1988) multivariate cointegration tests. Exchange rates do the bulk of adjustment toward equilibrium, except in the cases of the Thai baht and the New Taiwan dollar. For these currencies, interest rates and money supplies move to restore equilibrium. In ex post simulations, the out-of-sample fit of the estimated models is relatively good for the won, Singapore and New Taiwan dollars, and for the baht, although in no case is the exact magnitude and timing of the currency clashes predicted. The estimated model completely fails to track the rupiah out-of-sample.

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1 INTRODUCTION

In the wake of the turmoil in East Asian currency markets, several East Asian countries -- among them Korea, Thailand and Indonesia -- allowed their currencies to float freely during the summer of 1997. These events marked a substantial departure from the previous practice of tightly managing, or pegging, currencies against either the US dollar, or a basket of currencies. Now that these exchange rates are more determined by market forces, it is increasingly important to discern what, if any, macroeconomic factors systematically move nominal exchange rates.[1]

Using cointegration techniques I am able to detect and identify the long run relationship between exchange rates and money supplies, interest and inflation rates, and the relative price of nontraded goods for the Korean won, Singapore dollar, New Taiwan dollar, Indonesian rupiah and the Thai baht over the 1981-1996 period. The econometric results also indicate that over this sample period, exchange rates for the won, Singapore dollar, and rupiah responded to long run deviations from the estimated equilibrium relationships. However, for the baht and the New Taiwan dollar, most of the adjustment toward equilibrium took place through the adjustment of domestic money supplies or interest rates.

While these results pertain to a period prior to the currency crises of 1997, they may still have relevance for the post-crises behavior of these exchange rates. It is likely, for instance, that the long run relationship between the relative price of nontraded goods and the nominal exchange rate is invariant to the regime. Hence, whether the long run relationships identified through the

[1] In posing this question, I assume that the likelihood of a return to pegged exchange rates is nil. There are many reasons for this view. As documented in a number of papers, the degree of integration in both goods and financial capital is increasing over time. Furthermore, it is not clear that currency board arrangements could circumvent the impediments to fixing exchange rates, given the large potential liabilities represented by ailing banking systems in Indonesia and Thailand.

statistical tests remain stable is an empirical issue. Furthermore, the channel by which adjustment to restore equilibrium takes place depends only in part upon the policy regime in place; that is the theoretical issue of the Lucas critique may not necessarily be empirically valid.[2] The out-of-sample performance of these econometric specifications supports the relevance of these for certain currencies, including the won and the New Taiwan dollar.

One is tempted to question the feasibility of estimating such structural models, given the well-known difficulties in predicting this asset price. However, the conventional view of exchange rate determination has changed substantially since the early 1980s when the papers of Meese and Rogoff (1983a,b) illustrated the deficiencies of the monetary models of the time. During the early 1990s, several studies of G-7 currencies provided robust evidence that macroeconomic fundamentals do affect nominal exchange rates, although the relationships are only apparent at much lower frequencies than previously considered.

This paper proceeds in the following manner. In Section 2, the monetary model of the exchange rate is presented. In Section 3, the econometric methodology is discussed. The empirical estimates and out-of-sample performance of the models are examined in Section 4. Section 5 draws out the policy implications for these countries, and sets forth some concluding points.

2. THE MONETARY MODEL

2.1 The Standard Model

The asset-based approach to monetary model of the exchange rate is generally represented

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[2] For a discussion of the empirical relevance of the Lucas Critique, see Ericsson and Irons (1995).

as:

$$s_t = \beta_0 + \beta_2(m_t - m_t^*) + \beta_3(y_t - y_t^*) + \beta_4(i_t - i_t^*) + \beta_5(\pi_t - \pi_t^*) + \beta_6\omega \qquad \textbf{(1)}$$
$$\omega_t \equiv [(p_t^T - p_t^N) - (p_t^{T*} - p_t^{N*})]$$

where m_t is the (log) nominal money stock, p_t is the (log) price level, y_t is (log) income, i_t and π_t are the interest and expected inflation rates, respectively. The last term ω is the price of nontradable goods relative to tradable goods.

In the standard monetary model, as discussed by Frankel (1983), the coefficients have structural interpretations which may vary with the assumptions in effect. In monetary models, β_2 equals unity, while $\beta_3 < 0$, and represents the income elasticity of money demand. If prices are assumed to be flexible (Frenkel, 1976), then the interest rate and inflation differential are the same, and the condition $\beta_4 > 0$ holds. β_4 is equal to the absolute value of the interest semi-elasticity of money demand. On the other hand, if prices are sticky and there is secular inflation (Frankel, 1979), then $\beta_4 < 0$ and the magnitude of this parameter is positively related to price stickiness; the more rapid price level adjustment is, the smaller this coefficient is, in absolute value terms. In the Frankel (1979) model $\beta_5 > 0$, and is positively related to the interest semi-elasticity of money demand and negatively related to the degree of price stickiness. In most treatments of the monetary approach, long run PPP is assumed to hold economy-wide, and thus $\beta_6 = 0$. In this study we allow for deviations from economy-wide PPP; that is, nontradable goods prices are assumed not to be equalized across countries. The interpretation of β_6 will be discussed at further length below.

At this juncture, it is useful to review the empirical evidence in favor of such models, insofar as the developed country currencies are concerned. Some skepticism is warranted, given the findings of Meese and Rogoff (1983a,b) that univariate structural models did not usually outperform a random walk in out of sample simulations. However, starting with the work of MacDonald and Taylor (1994), robust evidence of long run relationships was obtained so that over long horizons, exchange rates do appear to be related to monetary fundamentals. Mark (1995) finds that monetary factors affect exchange rates, out-predicting a random walk at horizons of three years and more. Chinn and Meese (1995) extend this finding to sticky-price monetary models.

2.2 Developing Country Issues

There are three key issues that must be addressed when applying the monetary model to less developed countries (LDCs): capital mobility and substitutability; money demand stability, and; the presence of nontradables.

Capital mobility and substitutability: Because the monetary approach is built on perfect capital mobility and substitutability, it is not clear that these models would hold well for East Asian countries. As is well documented, some of these countries are only now removing restrictions on the capital account, and indeed Korea is still in the process of liberalizing its external accounts (Chinn and Maloney, 1998). Furthermore, perfect capital substitutability, defined by Frankel (1983) as the condition where bonds denominated in different currencies are viewed as perfect substitutes, is also not likely to hold. Hence, the model's predictions are unlikely to be borne out in the short run because the assumption of uncovered interest rate parity is violated. On the other hand, the covered interest differentials and the exchange risk premium are likely to be stationary, so the model's long run predictions are still relevant .[3]

Money demand stability: An empirically stable counterpart to equation (1) requires, in part, the presence of stable money demand functions. In LDCs, where economies are subject to monetization, increasing financial intermediation, or financial repression, *a priori* assertion of money demand stability not plausible. On the basis of Engle-Granger cointegration tests, Tseng and Corker (1991) assert that a stable cointegrating relationship holds for Indonesia, Korea, Singapore and Thailand. Dekle and Pradhan (1997) update these results for several Southeast Asian countries, using the more powerful Johansen (1988) methodology, and find that cointegrating relationships hold for nominal money supplies in Singapore and Thailand. Further, in the cases of narrow and broad Thai money, the restriction of homogeneity in price levels cannot be rejected. Indonesia constitutes the exception. In this last case, they identify a cointegrating relationship in money demand, allowing for structural shifts. The econometric analysis will take into account the possibility of such functional shifts.

Nontradable goods and PPP: Most studies of nominal exchange rate determination rely upon purchasing power parity holding in the long run; in other words, the long run real exchange rate is constant. Because this assumption is so grossly violated empirically in this region (Isard and Symansky, 1996; Chinn, 1997b), it is necessary to allow the long run real exchange rate to vary over time. In order to accomplish this, one needs to specify the determinants of this variable.

To modify the monetary model, let the log aggregate price index be given as a weighted

[3] A related issue pertains to existence of a "market determined" interest rate. Many East Asian interest rates are regulated, or at least subject to informal intervention by the monetary authorities. Chinn and Dooley (1997) discuss some of the macroeconomic implications of regulated banking systems.

average of log price indices of traded (T) and nontraded (N) goods:

$$p_t = (1-\alpha)p_t^T + \alpha p_t^N \tag{2}$$

where α is the share of nontraded goods in the price index. Suppose further that the foreign country's aggregate price index is similarly constructed. Furthermore, assume the long run nominal exchange rate is given by long run purchasing power parity *only in tradable goods*,[4]

$$s_t = p_t^T - p_t^{T*} \tag{3}$$

One then obtains the expression in (1), where the ω term is the relative price of nontradables.[5] The coefficient $\beta_6 = \alpha > 0$.

The relative price variable may be determined by any number of factors. In the Balassa (1964) and Samuelson (1964) model, relative prices are driven by relative differentials in productivity in the tradable and nontradable sectors.[6] With respect to the East Asian countries, there is a widespread belief that such factors are of central importance.

Relative prices may also be affected by demand side factors. In the long run, the rising

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[4] In Chinn (1998) I show that this assumption is consistent with the data for Indonesia, Korea, Singapore and Thailand, when the reference country is the United States. For Taiwan, purchasing power parity appears to hold for broad price indices (CPIs) over the 1975M01-1996M12 period, so making the assumption that PPP holds for a component of the aggregate price index should not pose any problems in estimation.

[5] The tradable and nontradable goods inflation rates are assumed equal ($\Pi = \Pi^T$) for simplicity.

[6] This view is adopted in DeGregorio and Wolf (1994), Canzoneri, Cumby and Diba (1996), Chinn (1997a,b) among others. The first two studies examine annual total factor productivity data for 14 OECD countries in a panel context, while Chinn (1997a) undertakes a higher frequency analysis. He uses quarterly time series regressions where labor productivity in manufacturing is used as a proxy for relative sectoral productivity, for the US, Canada, Germany, Japan and the UK.

preference for services, which are largely nontradable, may induce a trend rise in the relative price of nontradables. Over shorter horizons, government spending on public services may also induce changes in relative prices (DeGregorio and Wolf (1994) and Chinn (1997a)).

If one wished to differentiate between the two channels, it would be useful to substitute out for the individual determinants of the relative price variable in the square brackets. However, this is not the main objective of this study. In any event, sectoral productivity data is not readily available for many of the countries being investigated in this study. Hence, I proxy these combined Balassa-Samuelson and demand side effects with a relative price variable.[7]

3. ECONOMETRIC METHODOLOGY

The series being analyzed -- nominal exchange rates, relative money supplies, incomes, and so forth -- are integrated of order one [$I(1)$], so that individually they do not tend to revert back to their means. However, this does not mean that the series taken together are unrelated. Rather they may (or may not) exhibit a long run relationship. The existence of a long run relationship in levels of $I(1)$ variables is termed cointegration. As discussed in Phillips and Loretan (1991), one can proceed along a number of avenues. The main choice is between single equation methods and multiple equation methods. In this study, we opt to use the latter, for two reasons. First, single equation methods are ill suited to testing for situations in which there is more than one cointegrating vector. Second, one is interested not only in how exchange rates adjust, but also how other macroeconomic variables in the system adjust, especially since several of these

[7] While the ω variable is likely to be somewhat endogenous, this will not prove to be a problem for inference, as the econometric methodology allows for endogenous variables.

countries have actively managed their exchange rates by manipulating other policy instruments such as money supplies.[8]

The current standard in testing for cointegration in time series is the full-system maximum likelihood estimation technique of Johansen (1988) and Johansen and Juselius (1990). In the context of this model, this procedure involves estimating the system:

$$
\begin{aligned}
\Delta s_t &= \gamma_{10} + \Phi_1 ECT_{t-1} + \sum_{i=1}^{k} \gamma_{1i}\Delta s_{t-i} + \sum_{i=1}^{k} \zeta_{1i}\Delta\hat{m}_{t-i} + \sum_{1=i}^{k} \nu_{1i}\Delta\hat{y}_{t-i} \\
&\quad + \sum_{i=1}^{k} \xi_{1i}\Delta\hat{i}_{t-i} + \sum_{i=1}^{k} \mu_{1i}\Delta\hat{\pi}_{t-i} + \sum_{i=1}^{k} \chi_{1i}\omega_{t-i} + \varepsilon_{1t} \\
\Delta\hat{m}_t &= \gamma_{20} + \Phi_2 ECT_{t-1} + \sum_{i=1}^{k} \gamma_{1i}\Delta s_{t-i} + \sum_{i=1}^{k} \zeta_{2i}\Delta\hat{m}_{t-i} + \sum_{1=i}^{k} \nu_{2i}\Delta\hat{y}_{t-i} \\
&\quad + \sum_{i=1}^{k} \xi_{2i}\Delta\hat{i}_{t-i} + \sum_{i=1}^{k} \mu_{2i}\Delta\hat{\pi}_{t-i} + \sum_{i=1}^{k} \chi_{2i}\omega_{t-i} + \varepsilon_{2t} \\
\Delta\hat{y}_t &= \gamma_{30} + \Phi_3 ECT_{t-1} + \sum_{i=1}^{k} \gamma_{3i}\Delta s_{t-i} + \sum_{i=1}^{k} \zeta_{3i}\Delta\hat{m}_{t-i} + \sum_{1=i}^{k} \nu_{3i}\Delta\hat{y}_{t-i} \\
&\quad + \sum_{i=1}^{k} \xi_{3i}\Delta\hat{i}_{t-i} + \sum_{i=1}^{k} \mu_{3i}\Delta\hat{\pi}_{t-i} + \sum_{i=1}^{k} \chi_{3i}\omega_{t-i} + \varepsilon_{3t} \\
\Delta\hat{i}_t &= \gamma_{40} + \Phi_4 ECT_{t-1} + \sum_{i=1}^{k} \gamma_{4i}\Delta s_{t-i} + \sum_{i=1}^{k} \zeta_{4i}\Delta\hat{m}_{t-i} + \sum_{1=i}^{k} \nu_{4i}\Delta\hat{y}_{t-i} \\
&\quad + \sum_{i=1}^{k} \xi_{4i}\Delta\hat{i}_{t-i} + \sum_{i=1}^{k} \mu_{4i}\Delta\hat{\pi}_{t-i} + \sum_{i=1}^{k} \chi_{4i}\omega_{t-i} + \varepsilon_{4t} \\
\Delta\hat{\pi}_t &= \gamma_{50} + \Phi_5 ECT_{t-1} + \sum_{i=1}^{k} \gamma_{5i}\Delta s_{t-i} + \sum_{i=1}^{k} \zeta_{5i}\Delta\hat{m}_{t-i} + \sum_{1=i}^{k} \nu_{5i}\Delta\hat{y}_{t-i} \\
&\quad + \sum_{i=1}^{k} \xi_{5i}\Delta\hat{i}_{t-i} + \sum_{i=1}^{k} \mu_{5i}\Delta\hat{\pi}_{t-i} + \sum_{i=1}^{k} \chi_{5i}\omega_{t-i} + \varepsilon_{5t} \\
\Delta\omega_t &= \gamma_{60} + \Phi_6 ECT_{t-1} + \sum_{i=1}^{k} \gamma_{6i}\Delta s_{t-i} + \sum_{i=1}^{k} \zeta_{6i}\Delta\hat{m}_{t-i} + \sum_{1=i}^{k} \nu_{6i}\Delta\hat{y}_{t-i} \\
&\quad + \sum_{i=1}^{k} \xi_{6i}\Delta\hat{i}_{t-i} + \sum_{i=1}^{k} \mu_{6i}\Delta\hat{\pi}_{t-i} + \sum_{i=1}^{k} \chi_{6i}\omega_{t-i} + \varepsilon_{6t} \\
ECT &\equiv [\beta_1 s + \beta_2\hat{m} + \beta_3\hat{y} + \beta_4\hat{i} + \beta_5\hat{\pi} + \beta_6\omega]
\end{aligned}
\tag{4}
$$

where the carats (^) denote relative differences. For simplicity, I have written (4) assuming only

[8] For Korea and Taiwan, see Moreno (1996), for Singapore, Moreno and Spiegel (1997). See also Glick and Moreno (1994).

one cointegrating vector, although in principle nothing prevents the existence of more.

The procedure yields a trace statistic on which a likelihood ratio (LR) test can be conducted for the null of *r* cointegrating vectors against the alternative of *m* cointegrating vectors. The asymptotic critical values for this test are reported in Osterwald-Lenum (1992). Cheung and Lai (1993), among others, have shown that finite sample critical values may be more appropriate given the relatively small samples which are generally under study. In this study, I will refer to inferences made using both sets of critical values.

The multivariate approach also enables one to examine what variables do the adjustment to restore equilibrium. The typical focus on a single equation, normalized on the exchange rate, may miss such adjustment mechanisms. An example of this would be if money supplies are manipulated by the central bank to target a given nominal exchange rate. If this is the case, the exchange rate might not exhibit any response to the disequilibrium represented by the error correction term.

The lag length vor the VECMs is selected such that the residuals are serially uncorrelated, using the Box-Ljung Q-statistic for fourth order serial correlation, and a 5% marginal significance level.

4. EMPIRICAL ANALYSIS

4.1 Data

The analysis is conducted on quarterly data over the period 19781Q1 to 1998Q1, mostly drawn from IMF's *International Financial Statistics*, and from the Bank of China's *Financial Statistics*. Exchange rates are end-of-period, in US$/local currency unit. Money is either narrow money (*IFS* line 34) or broad money (*IFS* lines 34 plus 35). Income is GDP in 1990 currency units. Interest rates are interbank rates (*IFS* lines 60b). Inflation rates are calculated as the annualized first difference of the log- CPI.

Exchange rates are plotted (in solid lines) in Appendix Figures A1-A5. Both the interest and inflation rates are plotted in Appendix Figures A6-A10. (Further details regarding the data are reported in the Data Appendix).

The relative price variable is not directly observed; hence I use a proxy variable, which is calculated as:

$$\tilde{\omega} \equiv \log\left[\frac{(PPI^{US}/CPI^{US})}{(PPI^{*}/CPI^{*})}\right] \tag{5}$$

Assuming $\alpha = 0.5$, and the share of the CPI accounted for by nontradables is 50%, the posited value for β'_6 in equation (6),

$$s_t = \beta_0 + \beta_2(m_t - m_t^*) + \beta_3(y_t - y_t^*) + \beta_4(i_t - i_t^*) + \beta_5(\pi_t - \pi_t^*) + \beta'_6\tilde{\omega} \tag{6}$$

is approximately unity. To the extent that actual data for these countries deviates from these assumptions, this coefficient will differ from a value of one, although it should be positive in sign.

The $\tilde{\omega}$ term is plotted for each currency in Appendix Figures A1-A5 (as the dashed lines). A similar specification incorporating a relative price variable is used in Dornbusch (1976b), Wolff (1987) and Chinn and Meese (1995).

4.2 Empirical Results

The cointegration results[9] are reported in Table 1. First consider the currencies of the newly industrializing countries (NICs) – Korea, Singapore, and Taiwan. In the first row are the likelihood ratio (LR) statistics for the test of the null of zero cointegrating vectors against the alternative of one. (Since this is the most relevant choice given the results, I report only this LR test statistic.) The second row shows the 5% asymptotic critical values for this test; finite sample critical values adjusted using the method suggested by Cheung and Lai (1993) are shown in brackets. The implied number of cointegrating vectors using the asymptotic critical values and, in brackets, the number using the finite sample critical values, are reported in the third row. In three cases, there is evidence of multiple cointegrating vectors according to the asymptotic critical values. Using the finite sample adjusted critical values, one finds evidence of borderline cointegration for Korea and Taiwan. In the latter case, the LR statistic of 217.6 exceeds the adjusted critical value of 191.5. In the former, the LR statistic of 142.6 is very close to the adjusted 5% critical value of 150.6, so that if a 10% MSL were used, one would find evidence of one cointegrating vector. The decision is not so clear-cut in the case of Singapore, where the LR statistic of 135.2 is far below the adjusted critical value of 251.2. Given this ambiguity, I will discuss the case of the Singapore dollar, keeping in mind the weak evidence for cointegration.

The long run relationship for the dollar-won exchange rate (Column 2) fits the augmented

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[9] Preliminary unit root tests were conducted on the series included in the cointegration tests. Only the Indonesian interest rate, the Indonesian, Korean, and Singapore quarterly inflation rate differentials appear to be stationary, using a 5% significance level. However, the rejections of the unit root null are sensitive to sample period, and, for the latter three, to definition of the inflation rate (annual differences vs. quarterly differences).

monetary model well. The coefficient on narrow money is 1.352, within one standard error of the implied value of unity. The coefficient on relative income is -1.056, which implies unit-elastic money demand. The interest differential enters in with a negative sign, indicating that an increase in Korean interest rates relative to US rates induces an appreciation of the won. This set of results is consistent with a sticky price model of the exchange rate (although the interest rate coefficient is not significant at conventional levels). Inflation enters in with a positive sign, as expected, so that an increase in Korean inflation induces a depreciation of the won. Finally, the relative price variable enters with the appropriate sign, and significantly so. The in-sample fit for the won is illustrated in Figure 1.

For Taiwan, it is not possible to fit a model using narrow money, or broad money. Rather, the only specification that fits, with the expected signs, is one where US broad money, and Taiwanese quasi-money, enter in separately. I also include a dummy variable to account for the shift in money demand in 1984Q4 identified by Kuo (1990), as well as a dummy variable to account for a shift in capital account openness in 1989Q1 (see Chinn and Maloney, 1998). The results of estimating this specification are reported in Column 5. The US money coefficient has the expected positive value of 0.719, and the Taiwanese quasi-money coefficient, of -1.023. Both estimates are significantly different from zero. Relative income, interest rates and the relative nontradables price coefficients are also all correctly signed and statistically significant. The overall in-sample performance of this equation is shown in Figure 2.

For Singapore (Column 3), the results are somewhat less definitive. The broad money supply and income enter in with posited sign. However, only the latter is statistically significant (money is borderline significant). The relative price variable is completely insignificant (and wrong signed). This result mirrors that in Chinn (1997b) which indicated the irrelevance of the tradable/nontradable productivity model for that country's currency. What does seem to matter are nominal interest rates and inflation rates; both evidence statistically significant relationships with the nominal exchange rate. This pattern of results is in accord with a sticky price model of the exchange rate. The in-sample fit is shown in Figure 3.

As for the two LDCs of Thailand and Indonesia, there is evidence of cointegration for the latter, but mixed evidence for the former. For the Thai baht (Column 4), the LR test statistic for the null of zero against the alternative of one cointegrating vector is 135.2, which exceeds the unadjusted critical value of 68.5. According to the asymptotic critical values, there are two cointegrating vectors. In contrast, the adjusted critical value is 154.2 which is larger than the test statistic. *Assuming* one cointegrating vector, one obtains plausible coefficients. The coefficient on relative broad money is 1.654, is statistically significant, but within one standard error of the expected value of unity. The income and relative price coefficients are also correctly signed, and statistically significant. Only the interest differential is insignificant in the long run. The predicted values of the baht are presented in Figure 4.

In the case of Indonesia (Column 1, Figure 5), some revision to the basic model is required. While the relative price variable seems to have some relationship to the exchange rate, the real price of oil has an even larger effect. One can think of this variable as proxying for terms of trade effects. Hence, I substitute the oil price variable for the relative price variable.

In order to account for the money demand shifts identified by Dekle and Pradhan (1997), the regressions are augmented by two dummies, one for 1983Q2 and 1988Q3.[10] The LR statistic

[10] I also include a dummy for 1986Q1 to account for a spike in interest rates.

is 264.0. Using the adusted critical values, one finds evidence of a single cointegrating vector. The long run coefficient on money is 0.535, while that on income is -0.546. Both are correctly signed and statistically significant. The coefficient on interest rates is 0.343 which is very small, implying a rapid rate of price level adjustment. Only the inflation differential is not statistically significant.

The coefficient on the price of oil is 0.653, is highly significant, and implies that a one percentage point increase in the real (US$) price of petroleum induces a 0.653 percentage point appreciation of the rupiah against the dollar. This result is consistent with the findings in Chinn (1997b) regarding the effect of the real price of oil on the real exchange rate.

Table 2 reports the estimated response of each of the individual variable to the error correction terms (the Φ's in equation (4)), for each of the currencies. In the cases of three currencies – the rupiah, the won and the Singapore dollar – the exchange rate responds to the error correction term by moving to reduce the disequilibrium. The rate of response is very rapid in the case of the rupiah: 0.645, and in fact this seems to be sole adjusting variable. For the won, it is somewhat slower: 0.027.

For the baht and New Taiwan dollar, the response of the exchange rate is perverse (significantly so for the latter). In both cases, adjustment to the conditional mean appears to be effected by changes in monetary policy; in both cases money moves to close the gap. In the Taiwanese dollar's case, this is seen explicitly by way of the -0.375 coefficient (statistically significant at the 1% MSL) on Taiwanese quasi-money. For the Thai baht, this can be inferred from the positive and significant estimate on relative US-Thai money.

Interest rates also bear a portion of adjustment in the cases of the Singapore and New Taiwan dollars, and for the Thai baht. For the Singapore dollar and the Thai baht, interest rates fall in response to weakened domestic currency; that is, when the Singapore dollar is weak, Singapore interest rates fall to restore equilibrium. In the case of the New Taiwan dollar, the reverse occurs.

The preceding discussion may appear to suggest that the model does not fit with real world behavior in at least one dimension. That is, if the exchange rate is weaker than the long run value implied by the model at time t, the interest rate *falls* between time periods t and t+1.The common response of policymakers to a weak currency, including in the recent currency crises, is to *raise* interest rates.

A closer look at the results for the Singapore dollar and the Thai baht is instructive. When the error correction term is negative at time t -- so that the Singapore dollar is weak -- then between time period t and t+1 the Singapore dollar strengthens *and* Singapore interest rates rise relative to US interest rates. In other words, for the these two currencies, an increasing interest rate is associated with an strengthening currency.

4.3 Out-of-Sample Forecasting

The results presented thus far speak to the in-sample fit of the models. In order to assess the robustness, and relevance, of these empirical estimates, I conduct a series of out-of-sample forecasting exercises for the 1997Q1-1998Q1 period. The first is to use the estimated VECMs from the sample period extending up to 1996Q4 to forecast out-of-sample using the actual lagged values of the right hand side variables. This procedure – also termed historical simulation -- yields a static forecast of the exchange rate series, and can be thought of as a means to purge the forecasts of any uncertainty regarding the future values of the right hand side variables in the exchange rate equation. If the forecasts match the actual fairly well, this suggests that the estimated error correction model is stable. For each currency I also conduct a dynamic simulation, which uses the recursively generated values for each of the right hand side variables. The resulting predictions are true *ex ante* forecasts. Since the consensus view is that the 1997 currency crises were unexpected – at least with respect to timing – these forecast errors can be thought of as the news that occurred over this period.

In Figure 6, the static forecast for the won follows the actual with a one quarter lag. The precipitous drop in the won's value in 1997Q4 was partly picked by the model. The dynamic forecast implies only a continuation of the trend in won depreciation, confirming that the currency crash was largely unanticipated by the macroeconomic variables included in the regressions. The static forecast for the New Taiwan dollar (Figure 7) is slightly more successful, which is unsurprising given the absence of a crash for this currency. In fact, the model implies weaker currency for the middle of 1997 than actually obtained.

The static forecast for the Singapore dollar equation matches the movements in the actual, but does not indicate as severe a depreciation as actually occurred. By the end of 1997, the model predicts a 9% stronger Singapore dollar than was actually observed. The results are even striking for the Thai baht and the Indonesian rupiah. In these two cases the static forecast error for 1997Q4 is about 47% and 80% respectively. However, by 1998Q1, the static forecast for the Thai baht is about on the mark. The Indonesian rupiah, on the other hand, is mis-predicted by 130%. Hence, it is unlikely that the relationship estimated for the 1982Q1-1996Q4 period will be of much use for prediction in the foreseeable future.

5. CONCLUSIONS

The results in this paper suggest that the monetary approach is applicable to several East

Asian currencies. However, models that assume PPP for broad price indices are unlikely to be altogether successful, except in the case of the Singapore dollar, where the sticky-price monetary model appears to work. The augmented monetary model appears to fit the data well for a number of currencies. The fact that real factors, as represented by the relative price of nontradables, matters so much in East Asia mirrors results obtained by Hoffmaister and Roldos (1997), using panel structural VARs.

One policy implication of the ability of the monetary model to fit the behavior of these currencies is that high capital mobility is probably a good description of external regime facing these countries. Attempts to insulate these economies from external shocks will therefore be quite difficult to undertake.

While the study has focused on the pre-crisis period, for the Korean won, the New Taiwan dollar, and to a lesser extent the Singapore dollar, the model's long and short run predictions appear to hold outside the sample period. That means that the change in the exchange rate regime since July 1997 has not rendered irrelevant the estimates reported in this study. The Thai baht equation exhibits less stability out-of-sample; however, the model still tracks the baht fairly well in the sense that the forecast error as of March 1998 is virtually zero. The results for these four currencies indicate that the estimated models -- like virtually all other empirical models -- are not well suited to explaining behavior during speculative crises.

The Indonesian rupiah equation seems least able to cope with the post-sample data. Given the extreme political uncertainty surrounding the Indonesian economy during 1997-1998, one should perhaps not be overly surprised that *any* econometric model fails to capture asset behavior.

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Data Appendix

The data are from IMF, *International Financial Statistics*, November 1997 CD-ROM, except for data for Taiwan: Bank of China, *Financial Statistics*, various issues, as recorded in Federal Reserve Bank of San Francisco electronic database. IFS data updated using economic data sharing system at the IMF, April 1998.

□ Exchange rates, IFS line ae, in US$/national currency unit, end of period.

□ Narrow money, *IFS* line 34, in national currency units. Broad money is narrow money plus quasi- money *IFS* line 35, in national currency units.

□ Income is real GDP, *IFS* line 99b.r, in 1990 national currency units. The Korean and Singapore GDP series are seasonally adjusted over the 1970Q1-98Q1 and 1985Q1-98Q1 periods, respectively, using the X-11 seasonal adjustment additive procedure. Taiwanese GDP is originally in 1991 New Taiwan $, but is rebased to 1990 units. Indonesian data is from the IMF's Indonesia country desk (provided by Ilan Goldfajn). Thai GDP is estimated using the annual relationship between GDP, exports, imports, the real exchange rate and time, and quarterly data on these variables to generate a quarterly GDP series. For 1997 data, a constant 0.4% annualized growth rate is assumed.

□ Interest rates are short term, interbank interest rates, *IFS* line 60b, in decimal form. Indonesian missing values interpolated using an estimated AR(1).

□ Consumer price index, *IFS* line 64, 1990 = 100.

□ Producer price index, *IFS* line 63, 1990 = 100. Indonesian data excludes petroleum prices.

□ Inflation is the first difference of log(CPI), annualized.

□ Relative price variable:

$$\tilde{\omega} \equiv \log(PPI^{US}/CPI^{US}) - \log(PPI^{*}/CPI^{*})$$

Table 1
Johansen Cointegration Results
Long Run Parameters

Coeff	Pred	IN	KO	SI[1/]	TH[1/]	TI[2/]
LR		264.0	142.6	198.5	135.2	217.6
c.v.		94.2[188.3]	94.2[150.6]	94.2[251.2]	68.5[154.2]	94.2[191.5]
CR's		3[1]	4[0]	3[0]	3[0]	4[1]
s	(-1)	-1	-1	-1	-1	-1
m	(1)	0.535***	1.352***	0.908	1.654**	0.729***
		(0.066)	(0.480)	(0.574)	(0.730)	(0.094)
m*	(-1)	-0.535***	-1.352***	-0.908	-1.654**	-1.023***
		(0.066)	(0.480)	(0.574)	(0.730)	(0.068)
y-y*	(-)	-0.546***	-1.056*	-2.212**	-3.008***	-2.171**
		(0.120)	(0.593)	(0.928)	(1.303)	(0.869)
i-i*	(-)	-0.343*	-2.121	-11.921**	0.010	-1.417***
		(0.186)	(1.412)	(5.005)	(0.192)	(0.405)
π-π*	(+)	0.170	5.562*	11.368***		
		(0.160)	(2.938)	(4.292)		
$\tilde{\omega}$			3.655***	-0.133	2.038**	1.361***
			(1.225)	(0.198)	(1.015)	(0.420)
p^{oil}		0.653***				
		(0.026)				
k+1		5	4	5	5	5
N		60	64	48	45	59
Smpl		82Q1-96Q4	81Q1-96Q4	85Q1-96Q4	85Q4-96Q4	82Q2-96Q4
Dummies		1983Q2 1988Q3 (1986Q1 only)				1984Q4 1989Q1

Notes: "Coeff" is the β coefficient from equation (4). "Pred" indicates predicted sign. LR is the likelihood ratio test statistic for the null of zero cointegrating vector against the alternative of one. c.v. is the asymptotic critical value for the test of zero cointegrating vectors against the alternative of one [finite sample critical values in brackets]. CR's is the number of cointegrating relations implied by the asymptotic critical values [finite sample critical values]. Coefficients are long run parameter estimates from the Johansen procedure described in the text. k+1 is the number of lags in the VAR specification of the system. N is the effective number of observations included in the regression. Smpl is the sample period. Dummies are indicator variables taking on a value of one at the indicated date onward (except for the 1986Q1 dummy which takes on a value of 1 only in that quarter).

1/ Broad money.

2/ US broad money, Taiwanese quasi-money.

Table 2
Vector Error Correction Model Results:
Estimates of Reversion Rates

Coeff	Pred	IN	KO	SI[1/]	TH[1/]	TI[2/]
Δs	(-)	-0.687*** (0.062)	-0.027* (0.016)	-0.167** (0.070)	0.141 (0.106)	0.483*** (0.174)
Δm	(+)	0.141 (0.182)	0.066 (0.071)	-0.043 (0.083)	0.493*** (0.164)	-0.078 (0.052)
Δm^*	(-)	-0.141 (0.182)	-0.066 (0.071)	0.043 (0.083)	-0.493*** (0.164)	-0.374*** (0.073)
$\Delta(y-y^*)$	(-)	-0.025 (0.052)	-0.054** (0.026)	0.032 (0.062)	0.015 (0.110)	-0.039 (0.058)
$\Delta(i-i^*)$	(±)	0.200 (0.127)	-0.021 (0.016)	-0.062*** (0.018)	-0.562*** (0.180)	0.139** (0.067)
$\Delta(\pi-\pi^*)$	(+)	-0.051 (0.174)	0.109*** (0.030)	0.119** (0.050)		
$\Delta\tilde{\omega}$	(+)		-0.008 (0.007)	0.012 (0.144)	0.075 (0.117)	-0.089 (0.060)
Δp^{oil}		0.420 (0.380)				
adj-R^2 SER		.92 0.018	.38 0.015	.42 0.016	-0.07 0.011	.46 0.035

Notes: Coefficients are the Φ coefficients from equation (4). "Pred" is predicted sign. SER is the standard error of the regression.
1/ Broad money.
2/ US broad money, Taiwanese quasi-money.

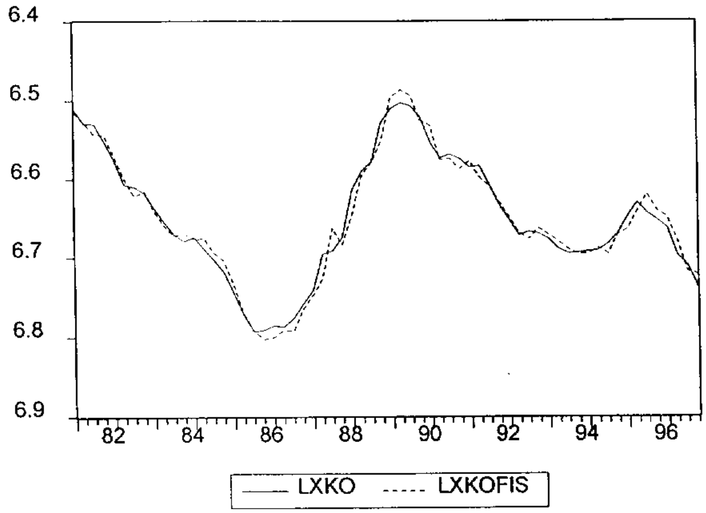

Figure 1: Korean won and in-sample fit

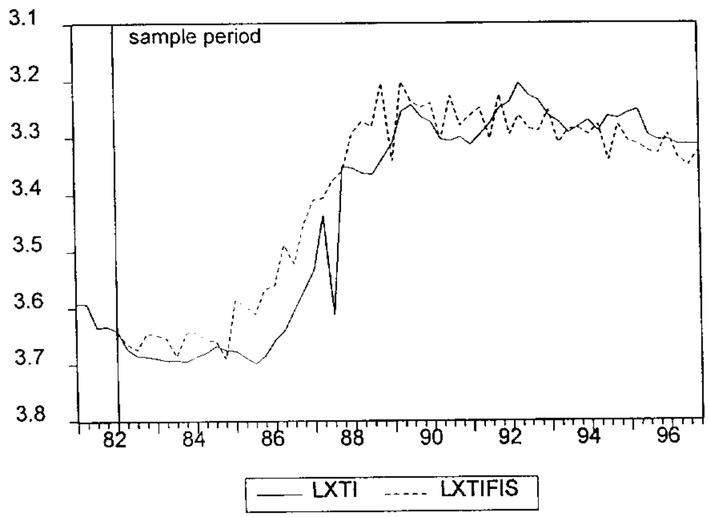

Figure 2: New Taiwan dollar and in-sample fit

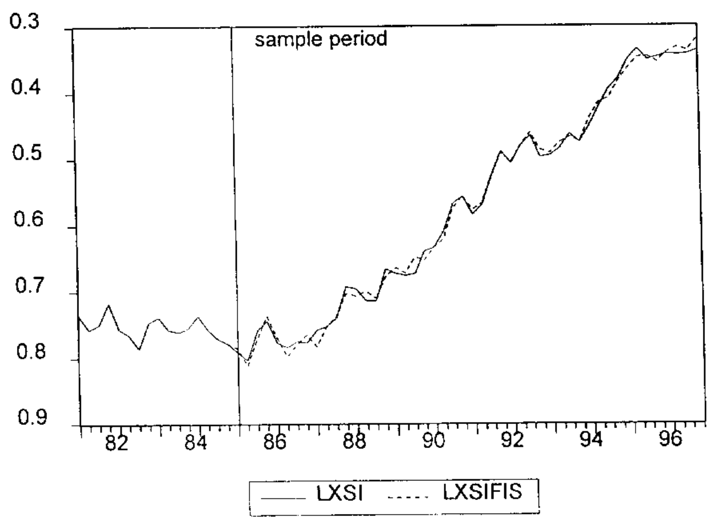

Figure 3: Singapore dollar and in-sample fit

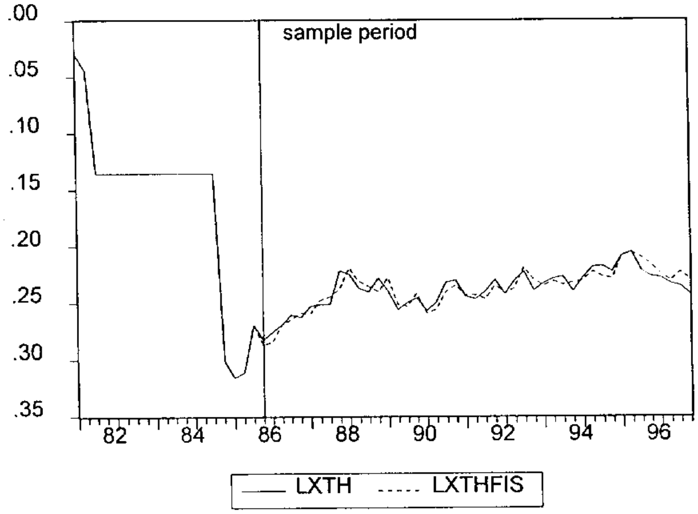

Figure 4:Thai baht and in-sample fit

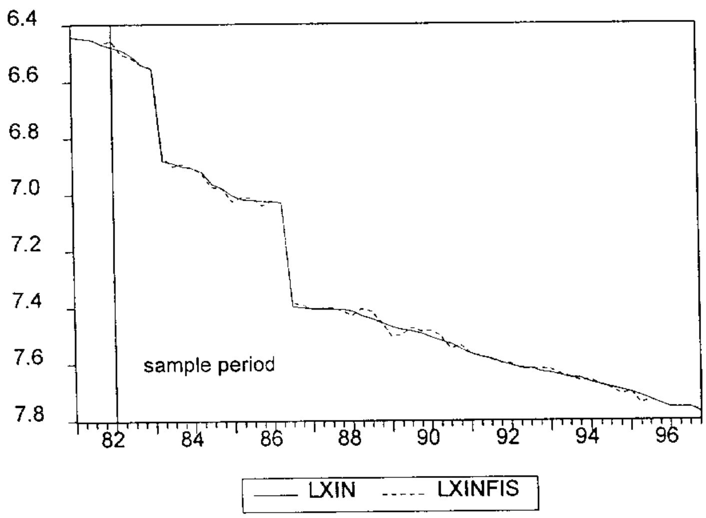

Figure 5: Indonesian rupiah and in-sample fit

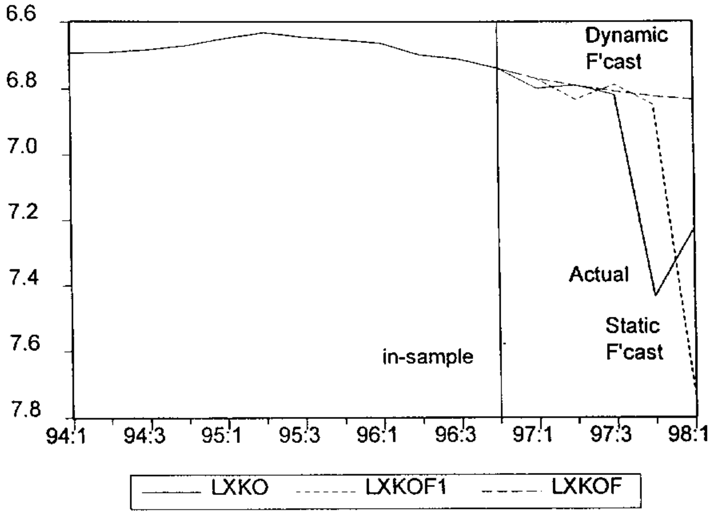

Figure 6: Korean won, static and dynamic forecasts

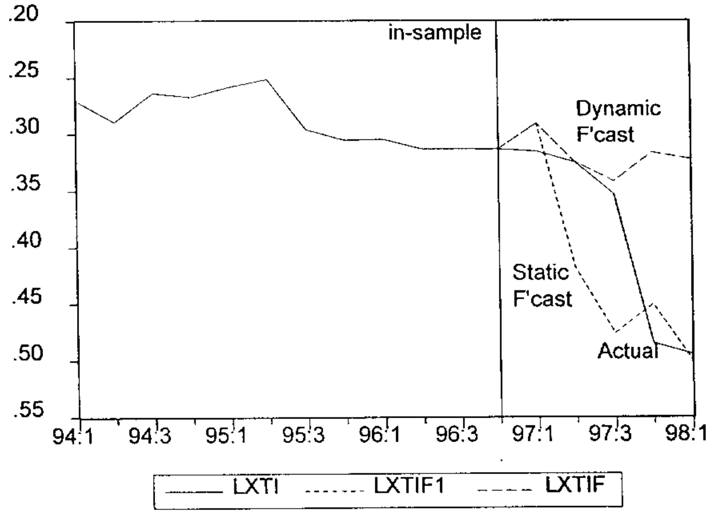

Figure 7: New Taiwan dollar, static and dynamic forecasts

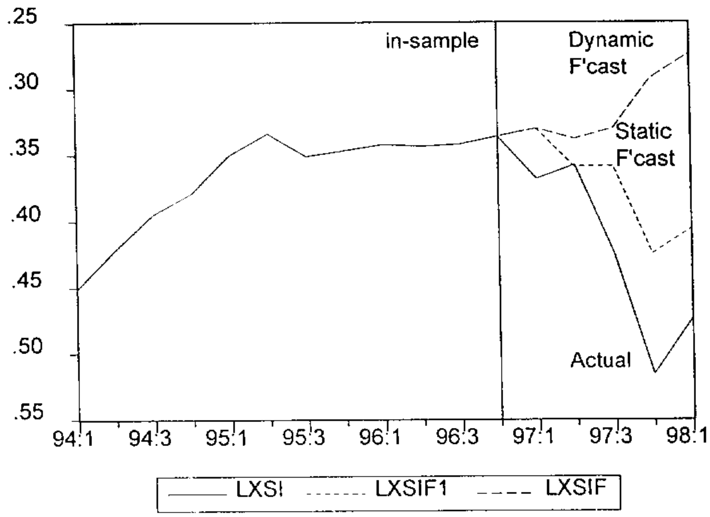

Figure 8: Singapore dollar, static and dynamic forecasts

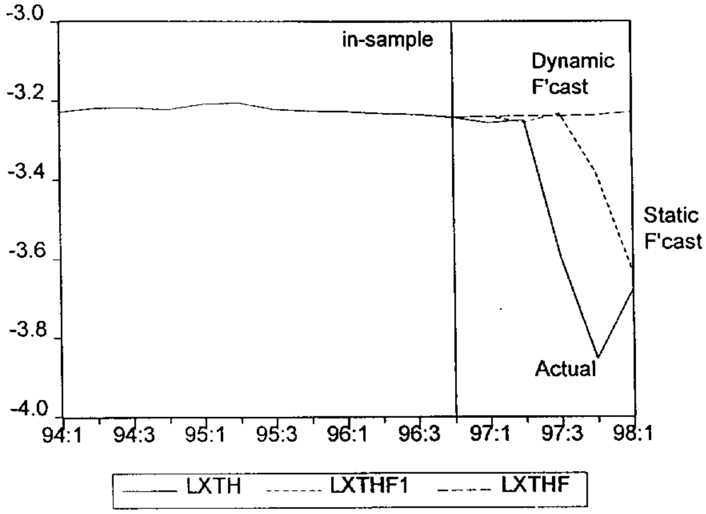

Figure 9: Thai baht, static and dynamic forecasts

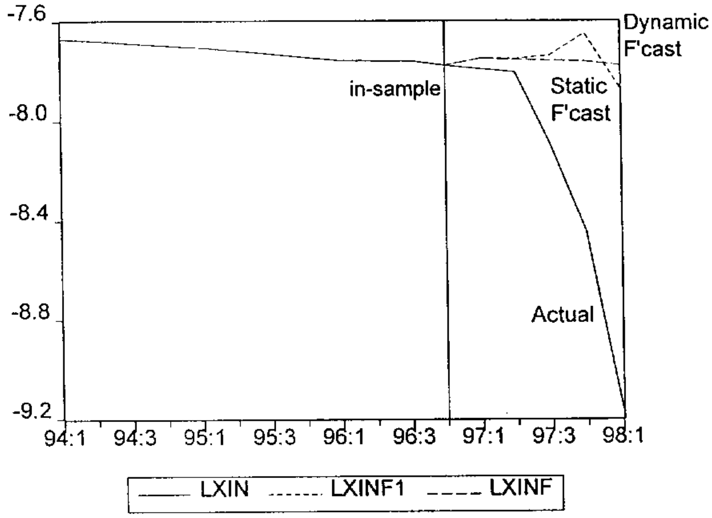

Figure 10: Indonesian rupiah, static and dynamic forecasts

APPENDIX

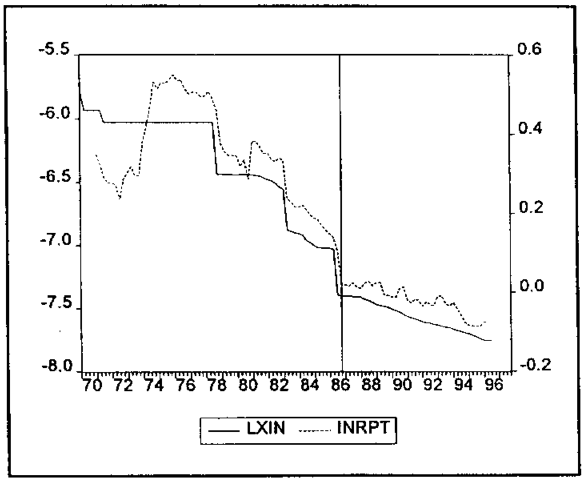

Figure A1: Indonesian rupiah and relative prices

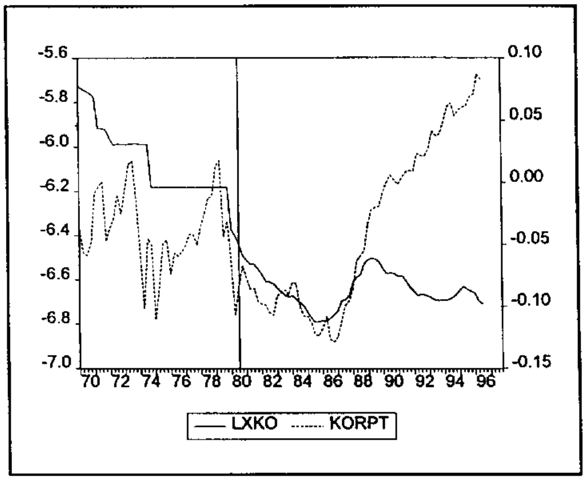

Figure A2: Korean won and relative prices

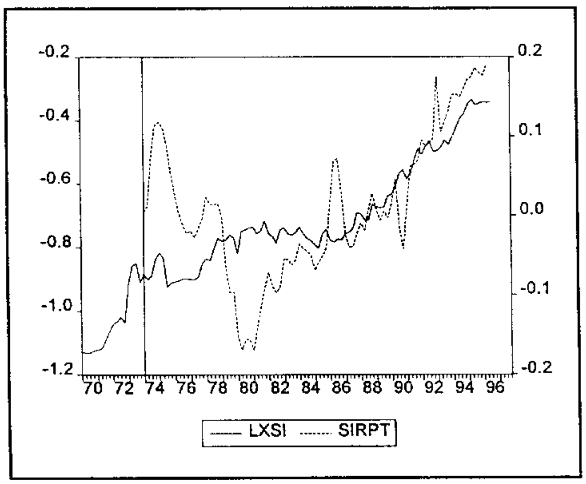

Figure A3: Singapore dollar and relative prices

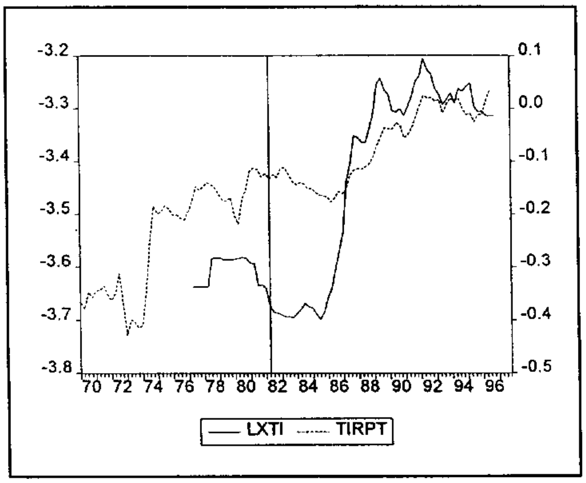

Figure A4: New Taiwan dollar and relative prices

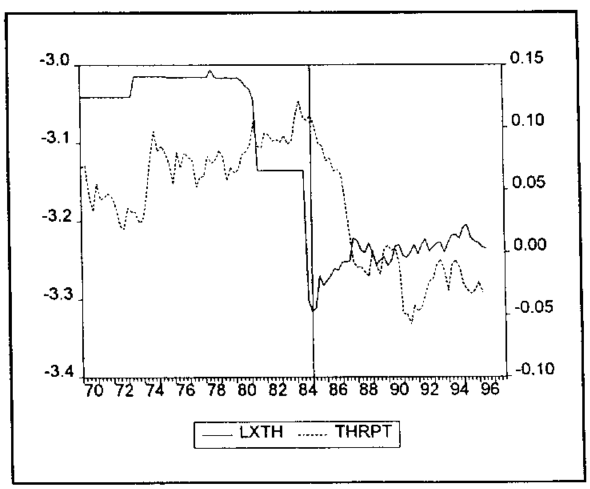

Figure A5: Thai baht and relative prices

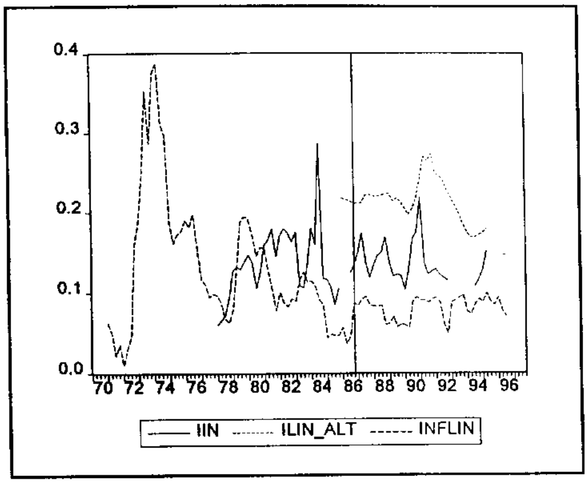

Figure A6: Indonesian money market rate, lending rate and inflation rate

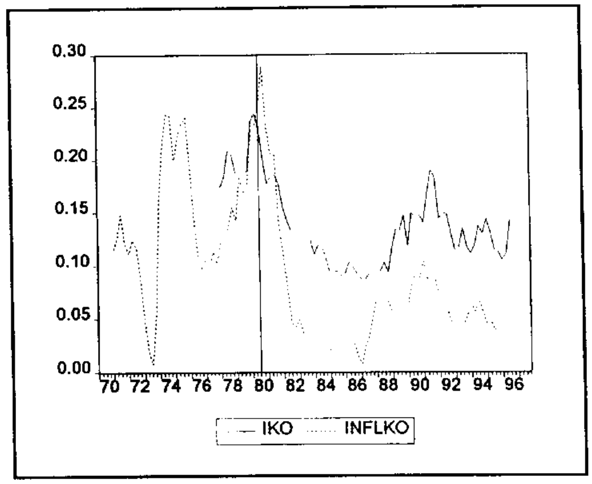

Figure A7: Korean money market rate and inflation rate

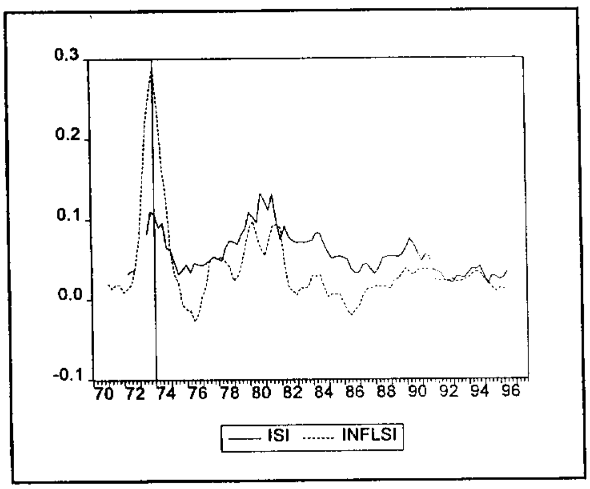

Figure A8: Singapore money market rate and inflation rate

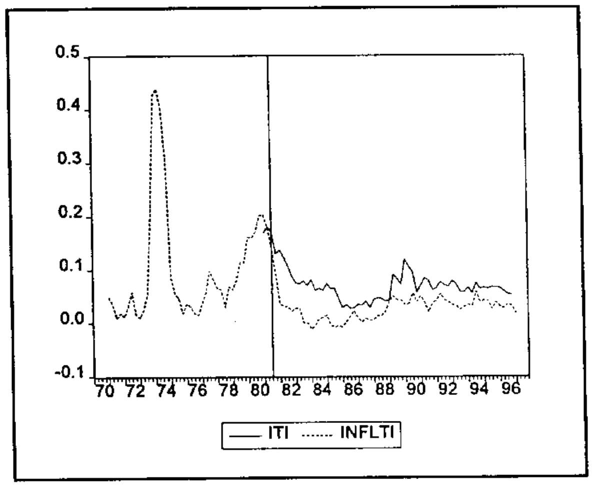

Figure A9: Taiwanese money market rate and inflation rate

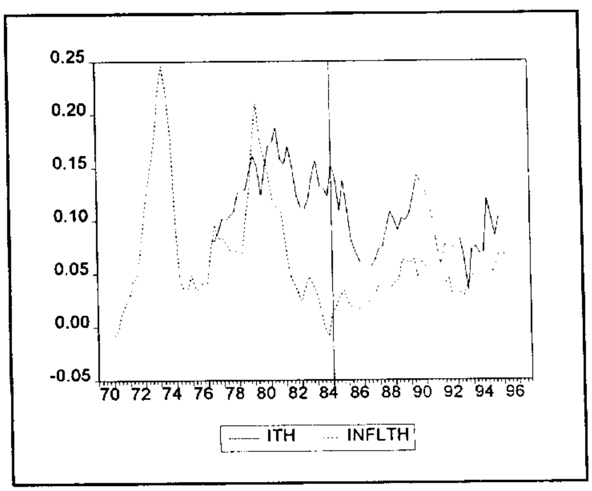

Figure A10: Thai money market rate and inflation rate