

Pacific Basin Working Paper Series

On the Won: And Other East Asian Currencies

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Working Paper No. PB97-07

Center for Pacific Basin Monetary and Economic Studies
Economic Research Department
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October 1, 1997

Abstract: A sticky price monetary model (Frankel, 1979) of exchange rates is applied to quarterly data on seven currencies: the Indonesian rupiah, Korean won, Malaysian ringgit, Philippine peso, Singapore dollar, Taiwanese dollar and the Thai baht. The model proves empirically unsuccessful, except in the case of the baht, and to a lesser extent, the Singapore dollar. A monetary model, augmented by the relative price of nontradables, is developed. This relative price variable proxies for the Balassa-Samuelson effect in East Asian real exchange rates identified in Chinn (1997b). The Korean won is best described by this modified model, while the Indonesian rupiah, Philippine peso and Taiwanese dollar also fit the model's long run predictions. On the other hand, the Malaysian ringgit proves difficult to econometrically model. This inability is disappointing because these two currencies have recently been allowed to float more freely.

JEL: F31, F41, F47

Acknowledgments: I am indebted to Eswar Prasad for comments on an earlier draft of the paper; I also thank Jaewoo Lee and the UCSC brown bag for very useful comments. Sona Shrestha provided assistance in collecting data. This paper was written while the author was a visiting scholar at the Federal Reserve Bank of San Francisco and as a resident fellow at the Rockefeller Foundation's Study and Conference Center in Bellagio, Italy. Financial support of faculty research funds of the University of California is gratefully acknowledged.

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1 INTRODUCTION

In the wake of the turmoil in East Asian currency markets, several East Asian countries -- Thailand, Indonesia, the Philippines and Malaysia -- allowed their currencies to float relatively freely during the summer of 1997. These moves represented 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

One is tempted to question the feasibility of such an endeavor, 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. Another strand of literature attacked the long run determinants of *real* exchange rates, where the determinants included government spending and productivity differentials. The conclusion of this approach is that productivity differentials matter, as do other macroeconomic variables.

These results for developed country currencies have a number of policy implications for the newly industrializing countries (NICs) of East Asia. As the NICs continue to liberalize their domestic financial markets and capital accounts, the behavior of their exchange rates will begin to resemble those of the developed countries. Central banks will no longer be able to simultaneously control exchange

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, Malaysia and Thailand.

rates and conduct independent monetary policies. Rather, exchange rate variability will generally increase, as long as central banks are unwilling to subordinate domestic monetary policy to exchange rate targets. This outcome is likely since the requisite policies will be increasingly painful, in the face of increased openness to financial capital inflows.

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 evidence for the monetary approach is examined in Section 4. Section 5 augments the monetary model with real variables primarily motivated by Balassa (1964) and Samuelson conjecture (1964). Section 6 draws out the policy implications for these countries.

To anticipate the empirical results, I find that the monetary model describes the short and long run behavior of the Thai baht, and to a lesser extent the Singapore dollar, fairly well. The Korean won, Indonesian rupiah, Taiwanese dollar and Philippine peso fit the long run predictions of the augmented model, although only the won and peso evidence an economically and statistically significant response of the currency to a disequilibrium. Finally, the Malaysian ringgit is not well explained by either model.

2. THE CONVENTIONAL MONETARY MODEL

2.1 The Model As Applied to Developed Country Currencies

The asset-based approach to monetary models of the exchange rate, as examined for instance by Mark (1995), relies upon a money demand equation of the form:

$$m_t^d - P_t = \phi Y_t - \lambda i_t \quad (1)$$

where m_t is the (log) nominal money stock, p_t is the (log) price level, y_t is (log) income and i_t is the

interest rate. Hence, the demand for real balances is an increasing function of income, and a decreasing function of the interest rate. The parameters have structural interpretations. ϕ is the income elasticity, and λ the interest semi-elasticity, of money demand.

Building upon this money demand function, the condition that expected depreciation equals the nominal interest differential (*i.e.*, perfect capital mobility *and* substitutability), and *continuous* purchasing power parity one obtains the flexible price monetary model:

$$s_t = (m_t - m_t^*) - \phi(Y_t - Y_t^*) + \lambda(i_t - i_t^*) \quad (2)$$

where s_t is the logarithm of the exchange rate (US\$ per unit of foreign currency) and $*$ denotes the local country variable.

The intuition for this type of model is as follows: the larger the money supply, m , relative to the determinants of the money demand, such as income and interest rates, the weaker (the higher) the exchange rate. When income rises, more money is demanded for transactions purposes, and the exchange rate falls (or appreciates). Interest rates decrease money demand, so that the higher the domestic interest rate, the lower the demand for money and hence the weaker the domestic currency.

The model described above imposes an assumption that prices are perfectly flexible; for many purposes this is an undesirable restriction. One way to relax this assumption is to allow prices to adjust slowly, and monotonically, so as to close the gap between the current and the long run price level.

Allowing for secular inflation, the exchange rate behaves as:

$$s_{t+1} - s_t = -\Theta(s_t - \overline{s_t}) + (\Pi_t - \Pi_t^*) \quad (3)$$

where Π_t is the expected CPI inflation rate from t to $t+1$, and Θ is the rate of reversion of the price level

to its long run value, \bar{s} .² Equation (3) states that the exchange rate depreciates whenever the current exchange rate is stronger than its long run value (denoted by the overbar). Imposing rational expectations, one then obtains the following model due to Frankel (1979) following Dornbusch (1976a):

$$= (m_t - m_t^*) - \phi (y_t - y_t^*) - \left(\frac{1}{\Theta}\right) (i_t - i_t^*) + \left(\lambda + \frac{1}{\Theta}\right) (\Pi_t - \bar{\Pi}) \quad (4)$$

Now interest rate differentials and inflation rate differentials enter in separately whereas in the previous formulation, the inflation differential and the interest differential were one and the same.

When the money supply is increased in this model, domestic real money balances increase too, since prices are sticky. To equilibrate the money market, domestic interest rates must fall. However, since uncovered interest parity must hold, the exchange rate must be appreciating over time, even though in the long run the exchange rate must be depreciated relative to its initial value. Hence, the response of the exchange rate to an interest rate decrease is an instantaneous depreciation of the currency. The more rigid prices are, the larger this “overshooting” effect. This model is more appropriate than the flexible price version if only because it accords better with casual empiricism -- higher interest rates, *ceteris paribus*, are associated with stronger currencies.

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

² The long run nominal exchange rate moves with $\Pi - \bar{\Pi}$.

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. At shorter horizons, such as at the weekly and monthly frequencies, currencies will continue to experience short term fluctuations that are largely inexplicable.

2.2 Developing Country Issues

Because the monetary approach is built on perfect capital mobility and substitutability, it is unreasonable to expect that these models would hold very well for East Asian NICs. 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, 1996).

Even in countries with relatively open capital accounts, such as Malaysia and Indonesia, may evidence covered interest differentials. Woo and Hirayama (1996) argue that the central banks' ability to *drastically* change monetary policies and financial regulations grants them an ability to punish speculators. Monetary authorities in the region have some short run latitude in manipulating interest rates while targeting the exchange rates. Hence, covered interest parity (even if forward rate markets exist) is unlikely to hold. Perfect capital substitutability, defined by Frankel (1983) as the condition where bonds denominated in different currencies are viewed as perfect substitutes, is also ruled out. Summing up, 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, such deviations will be more difficult to sustain in the long run, so the model's predictions are still of some interest.

A related issue pertains to existence of a "market determined" interest rate. While interest rate series are reported for all of these NICs, many of them are regulated, or at least subject to informal intervention by the monetary authorities. When the money market rate is not freely determined, it is

unclear whether the market clearing conditions assumed in the monetary approach hold.³

Perhaps a more fundamental issue pertains to the stability of the money demand functions. Recall that equation (4) is built upon a money demand function in equation (1). Even in developed countries, there is substantial debate over the stability of such equations in light of financial innovation. In LDCs, where economies are subject to monetization, increasing financial intermediation, or financial repression, *a priori* assertion of money demand stability is even less sustainable. On the basis of Engle-Granger cointegration tests, Tseng and Corker (1991) assert that a stable cointegrating relationship holds for Indonesia, Korea, Malaysia, Singapore and Thailand. Dekle and Pradhan (1996) update these results for some Southeast Asian countries, using the more powerful Johansen (1988) methodology, and conclude that there is no evidence of real money demand cointegration, with the exception of Indonesia.⁴ In this last case, they identify a cointegrating relationship in money demand, allowing for structural shifts. However, the identified breakpoints do not fall within the rupiah's floating period examined here.

2.3 Cross-Section Evidence for the Monetary Approach

Before embarking upon a statistical analysis of the exchange rate equations, it might be of interest to see how well the relationship holds in a cross-section, over long periods of time. First I calculate the fundamentals (Z),⁵ as defined according to the relationship:

³ See Chinn and Dooley (1997) for a closer examination of this issue of financial sector regulation.

⁴ Dekle and Pradhan (1996) do find that cointegrating relationships hold for nominal money supplies. Further, in the cases of narrow Malaysian, and narrow and broad Thai money, the restriction of homogeneity in price levels cannot be rejected.

⁵ The calculations of the fundamentals require that some assumption be made regarding the income elasticity of narrow money demand. I assume that this is unity for the US. For Indonesia, Korea, Malaysia, the Philippines, Singapore and Thailand, this parameter takes on the values of 1.2, 0.8, 1.1, 0.7, 0.9 and 1.0, respectively (see Tseng and Corker, 1991, Table 1). For Taiwan, I used Dynamic OLS estimates over the 1981-96 period, which indicate an income elasticity of 1.5).

$$Z_t = (m_t - m_t^*) - (\phi^{US} Y_t - \phi^* Y_t^*)$$

The fundamentals equal the exchange rate if the fundamentals follow a random walk, and price levels are perfectly flexible. This last assumption is appropriate when considering long run relationships.

In Figure 1, the average quarterly change in nominal exchange rates (in US\$/currency unit) is plotted against the change in the fundamentals for the 1975.1-84.4 period. Figure 2 presents the same variables for the 1985.1-94.4 period.

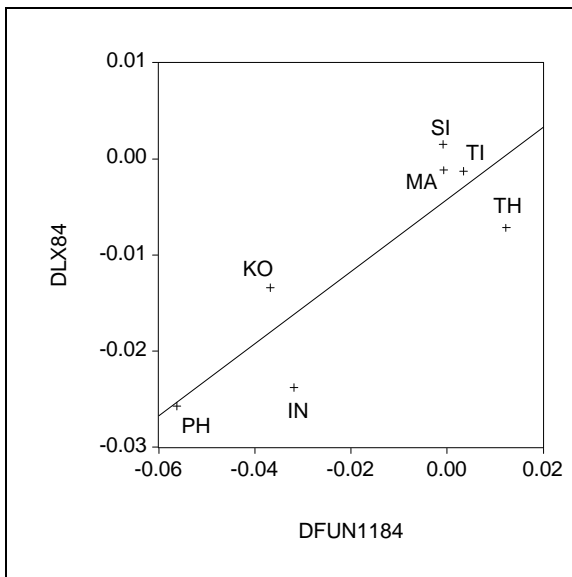


Figure 1: Average quarterly changes in exchange rates and fundamentals, 1975-84

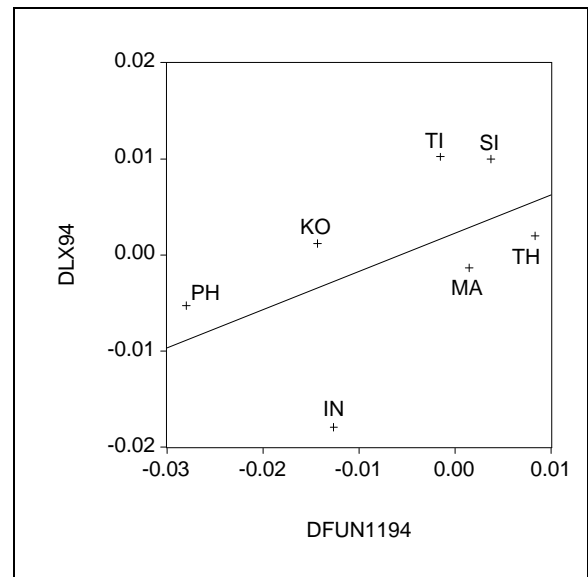


Figure 2: Average quarterly changes in exchange rates and fundamentals, 1985-94

The data conform to the posited relationship, in the sense that the relationship is positive. The slope coefficient is 0.38 (t-statistic of 3.86) for the earlier period; it is 0.40 (t-statistic of 1.37) for the later. While the fit is quite good for the first period, it *deteriorates* substantially in the second period (adjusted $R^2 = 0.70$ to 0.13). Indonesia appears to be something of an outlier; however, omitting this

observation still produces the same deterioration.⁶ This pattern could arise for any number of reasons, including those mentioned earlier. For the moment, I defer the discussion of these issues until after the econometric investigation.

3. ECONOMETRIC APPROACH

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). 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. Given the large number of variables (typically nine, or five in relative differences form) and lags of at least one, such finite sample critical values are typically so large as to rule out any opportunity of rejecting the null of no-cointegration. Moreover, such maximum likelihood techniques require that the error correction model for each endogenous variable be adequately modeled by the selected specification; this seems a particularly onerous requirement for countries marked by extreme instability in their functional relationships.

I rely on estimates derived from nonlinear least squares (NLS) regression. Phillips and Loretan (1991) argue that the following NLS estimator is optimal among single-equation estimators:

$$\Delta s_t = \gamma_0 + \gamma_1 (s_{t-1} - BX_{t-1}) + \sum_{i=+1}^{-1} \Psi_i \Delta X_{t+i} + u_t \quad (6)$$

where X is a vector of explanatory variables; the specifications include quarterly dummies to account for deterministic seasonality. The leads of the differences of the right hand side variables serve to orthogonalize the error term. If s reacts to the disequilibrium then γ_1 should be negative, and statistically

⁶ Using the truncated sample, the adjusted R^2 falls from 0.73 in the early period to 0.19 in the later. The slope coefficient falls from 0.34 to 0.27, with no statistical significance in the later period.

significantly different from zero.

This approach, which allows for the possibility of endogenous right hand side variables, is appropriate since many of these East Asian currencies are managed by their central banks.⁷ In particular, one can think of the domestic monetary authorities "reacting" to exchange rate developments; in this case, the assumption of money supply exogeneity would be invalid. While cointegration ensures super-consistency of OLS estimates of the cointegrating vector parameters, regardless of endogeneity, the small sample characteristics of such estimates is often quite poor.

To check that the residuals are approximately white noise, lags of first differences are added such that the null of no serial correlation cannot be rejected at the 10% marginal significance level. The F-test version of the Godfrey-Breusch LM test for fourth order serial correlation is used for this purpose.

4. EMPIRICAL ANALYSIS

4.1 Data

The data are drawn from the IMF's *International Financial Statistics* March 1997 CD-ROM, for all series, except for those of Taiwan. The Taiwanese data are drawn from the Bank of China's *Financial Statistics*, as recorded in the Federal Reserve Bank of San Francisco's electronic database. They are quarterly data and span the period 1970.1 to 1996.4

Exchange rates are end-of-period, in US\$/foreign currency unit. They are also plotted (in solid lines) in Appendix Figures A1-A7. Money is either narrow money (*IFS* line 34) or broad money (*IFS* lines 34 plus 35). Income is usually GDP in 1990 currency units. The exceptions are Taiwan (in 1991

⁷ For Korea and Taiwan, see Moreno (1996), for Singapore, Moreno and Spiegel (1997). See also Glick and Moreno (1994).

New Taiwan \$) and Malaysia, where industrial production is used since the quarterly Malaysian GDP series spans only a few years. Interest rates are usually 3 month interbank rates. For the Indonesian series, a medium term interest rate is used because the short term rate exhibits several large breaks which cause a substantial loss of degrees of freedom if used. Inflation rates are calculated as the annual change in the log of the price level, as measured by the CPI. Both the interest and inflation rates are plotted in Appendix Figures A8-A14. (Further details regarding the data are reported in the Data Appendix).

4.2. Empirical Results (I)

The results of implementing the regressions, using a narrow measure of money, are reported in Table 1. Since the focus of interest is on exchange rate determination under floating rates, I omit fixed/pegged period data from each regression (the specific sample periods are indicated in the tables, and also in the Appendix Figures as the periods after the vertical lines). In certain cases, a time trend is included if it is statistically significant.

Turning attention to the original East Asian NICs of Korea, Singapore, and Taiwan, one finds that the simple sticky-price monetary model does not perform very well in two of the three cases. While all three currencies exhibit exchange rate reversion to equilibrium, the Singapore dollar and Taiwanese dollar have statistically significant and negative coefficients on money. Income variable coefficients are also perversely signed (although not significantly so, at the 10% level). Interest rates enter with correct sign only in the case of the Singapore dollar. It is noteworthy that the interest rate variable is the key factor in the sticky-price monetary model, so its statistical significance lends some credence to monetary model. The perverse sign on the monetary aggregate could be due to financial innovation common in all international financial centers (as alluded to in Dekle and Pradhan, 1996: 11-12). The Korean won exhibits some reversion to mean, although only at the 20% significance level. Money and income have

the correct signs, while interest rates and inflation rate differentials do not.

Of the Southeast Asian currencies, neither the Indonesia nor Malaysian appear to fit the model very well either. The Indonesian rupiah equation does not exhibit cointegration with the monetary fundamentals, while the Malaysian ringgit displays perversely signed (and statistically significant) coefficients on money and income. On the other hand, the Thai baht has money and income coefficients of the correct sign, so some positive results are obtained. The Philippine peso also fails to exhibit cointegration.

The results of using a broader measure of money -- approximately M2 -- are not appreciably better. The only improvement is that the money measure now has a correct, but insignificant, sign for the NT\$ (this result using broad money is reported in Table 1).

The fact that the interest differentials are often incorrectly signed suggests that the model may be inappropriate. I investigate the possibility that prices are flexible in these less developed countries, so that one should search for a cointegration vector consistent with the Bilson (1978) model in equation (2). The Bilson model is the correct one if the interest differential equals the inflation differential, which in turn equals the expected rate of depreciation. However, this assumption is unlikely to hold unless interest rates are market determined, a condition that is fairly implausible for several of these countries. Consequently, I estimate equation (2) using inflation differentials in the place of interest differentials, as in Frenkel's (1976) formulation:

$$s_t = (m_t - m_t^*) - \phi(Y_t - Y_t^*) + \lambda(\Pi_t - \Pi_t^*) \quad (7)$$

The results (not reported) indicate no improvement, except in one case: the Thai baht shows up in the broad money specification with correctly signed and significant coefficients. Other than that, it seems apparent that monetary models, assuming either instantaneous or long-run purchasing power

parity, are inadequate for the task of explaining East Asian exchange rates.

5. REAL DETERMINANTS OF REAL EXCHANGE RATES

5.1 A Theoretical Modification

Most investigations of nominal exchange rate determination, including the foregoing, 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.

$$P_t = (1-\alpha)P_t^T + \alpha P_t^N \quad (8)$$

To modify the monetary model, let the log aggregate price index be given as a weighted average of log price indices of traded (T) and nontraded (N) goods:

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:

$$P_t^* = (1-\alpha^*)P_t^{T*} + \alpha^* P_t^{N*} \quad (9)$$

One can then write the relative price level as:

$$\begin{aligned}
P_t - P_t^* &= \alpha (P_t^N - P_t^{N*}) + (1-\alpha) (P_t^T - P_t^{T*}) \\
&= \alpha [(P_t^N - P_t^T) - (P_t^{N*} - P_t^{T*})] + (P_t^T - P_t^{T*})
\end{aligned} \tag{10}$$

Now let the nominal exchange rate behave as follows:

$$s_{t+1} - s_t = -\Theta (s_t - \overline{s_t}) + (\Pi_t^T - \Pi_t^{T*}) \tag{11}$$

In other words, *tradable* goods inflation, Π^T , is what is relevant. Further, the long run nominal exchange rate is given by purchasing power parity *only in tradable goods*,

$$\overline{s_t} = \overline{P_t^T} - \overline{P_t^{T*}} \tag{12}$$

so that one then obtains the following expression:

$$\begin{aligned}
s_t &= (m_t - m_t^*) - \phi (Y_t - Y_t^*) - \left(\frac{1}{\Theta} \right) (i_t - i_t^*) + \lambda (\Pi_t - \Pi_t^*) \\
&+ \frac{1}{\Theta} (\Pi_t^T - \Pi_t^{T*}) - \alpha [(P_t^N - P_t^T) - (P_t^{N*} - P_t^{T*})]
\end{aligned} \tag{13}$$

where the price of nontradables (tradables) is proxied by the CPI (PPI). In the empirical section, $\Pi = \Pi^T$ for simplicity.

The relative price variable in square brackets [.] 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. 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.

Relative prices may also be affected by demand side factors. In the long run, the rising 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. Both DeGregorio and Wolf (1994) and Chinn (1997a) incorporate such effects.

In principle, one would like to substitute out for the determinants of the relative price variable in the square brackets, especially since the price of tradables is likely to be endogenous to the exchange rate. Unfortunately, sectoral productivity data is not readily available for many of the countries being investigated in this study. Hence, I proxy these Balassa-Samuelson and demand side effects with a relative price variable, which is calculated as:

$$-\alpha [(p^N - p^T) - (p^{N*} - p^{T*})] \approx \omega \equiv \log \left[\frac{(PPI^{US} / CPI^{US})}{(PPI^* / CPI^*)} \right] \quad (14)$$

The ω term is plotted for each currency in Appendix Figures A1-A7 (as the dashed lines). The long run cointegrating relationship, in terms of observable variables, is then:

$$m_t - m_t^* = \phi (y_t - y_t^*) - \left(\frac{1}{\Theta} \right) (i_t - i_t^*) + \left(\lambda + \frac{1}{\Theta} \right) (\pi_{t+1} - \pi_{t+1}^*) \quad (15)$$

A similar specification incorporating a relative price variable is used in Dornbusch (1976b), Clements and Frenkel (1980), Wolff (1987) and Chinn and Meese (1995).

5.2 Empirical Results (II)

I first redo the cross-section analysis by examining a scatterplot of the average change in the nominal exchange rate change against the average change in fundamentals, revised to include real factors. This yields Figures 3 and 4. In both, there is now a definite positive relationship, with a slope

coefficient of 0.30 (t-statistic of 3.43) in the early period,⁸ and 0.59 (t-statistic of 2.87) in the later. More importantly, the adjusted R^2 is 0.55 in the later period, which is substantially above the 0.13 obtained using only monetary factors.⁹

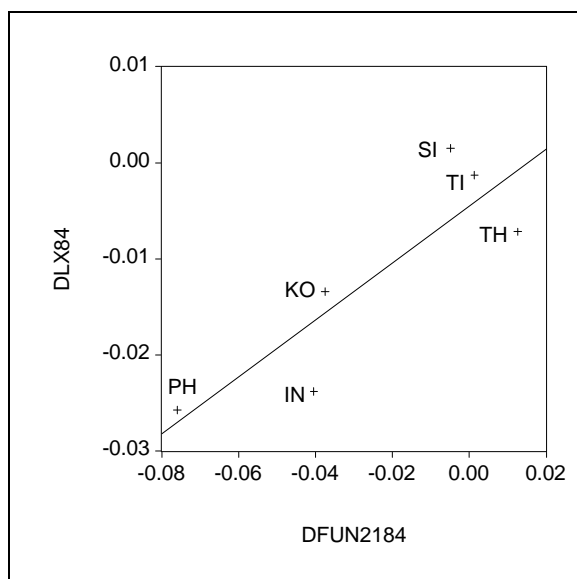


Figure 3: Average change in nominal exchange rates against average change in fundamentals, 1975.1-1984.4

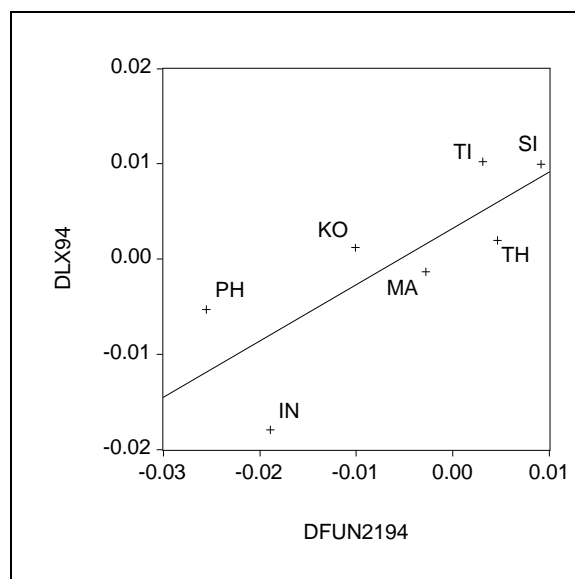


Figure 4: Average change in nominal exchange rates against average change in fundamentals, 1985.1-1994.4

The results of running the regression on equation (15) with narrow money are reported in Table 2. The Korean, Malaysian, Taiwanese, Thai and possibly Singapore currencies evidence reversion to equilibrium. In all cases money is correctly signed, except for Singapore and Taiwan. Unfortunately, only in the case of Korea is the money coefficient correctly signed *and* statistically significant. Income

⁸ No PPI data is available for Malaysia during the early subsample, so the regressions are not strictly comparable. However, omitting Malaysia in the later subsample yields a similar pattern, so no conclusions hinge upon the Malaysian observation.

⁹ In this case, I set the coefficient on ω to unity, which it should be if the share of nontradables in the CPI is 0.5, and that in the PPI is zero.

coefficients typically show up with statistical insignificance, as do the interest and inflation coefficients.

As for the relative price variable, it is correctly signed with the exception of the Singapore dollar and Thai baht (both non-significant); the only instance for which the point estimate is statistically different from zero is the Korean case. Repeating the exercise using a broader measure of money yields similar results, although the Thai baht's relative price variable now has a negative and statistically significant coefficient.

5.3 Interpretation

It is important to place these statistical results in perspective. In the empirical literature on developed country exchange rate determination, it is commonplace to obtain sign reversals and apparent parameter instability, even after accounting for issues of nonstationarity. Typically, these results are based on quarterly data extending over 23 years of post-Bretton Woods data. In contrast, the results in Tables 1 and 2 apply to periods as short as seven years (for the Indonesian rupiah) and economies undergoing extremely rapid structural change. The fact that the estimates are not often statistically significant should not be surprising in this context.

The Korean won is the one currency that best fits the augmented monetary model. The rate of reversion is fairly rapid, implying a half life of a deviation of about 5 quarters. The long run point estimate for money is 0.636; one cannot reject the null hypothesis of a unit coefficient implied by theory. As for the most important variable, relative prices also enter in positively and significantly; furthermore one cannot reject a unit coefficient. This outcome lends credence to the Bahmani-Oskooee and Rhee (1996) and Chinn (1997b) results showing a role for productivity in the long run real value of the won.

There is an interesting parallel between the behavior of the Korean won and the Japanese yen. In both cases, a monetary model, augmented by a relative price variable works quite well (see Chinn and Meese, 1995, for results for the Yen). It appears that similarities in the industrialization process and the

mode of financial regulation spill over into the behavior of these two countries' currencies.

While the relative price variable is not usually significant in the other cases, inclusion of this variable causes the money coefficient to become positive in all cointegrated cases except the Thai baht. In this latter case, the monetary approach appears more appropriate since the relative price variable is not statistically significant, and the null hypothesis of a unitary coefficient can be statistically rejected. Furthermore, the money variable enters in with near significance *only* if this relative price variable is omitted. This result is consistent with the finding in Chinn (1997b) that relative prices and relative productivity differentials do not play a role in movements in the baht.

The Taiwan dollar results, like those of the Singapore dollar, consistently show a negative coefficient on money. In the case of Singapore, Dekle and Pradhan (1996) do not identify a cointegrating vector corresponding to money demand, so one might not be surprised by the perverse sign on money stocks. The Taiwanese dollar is a bit more difficult to interpret. Kuo (1990) does identify two periods of liberalization which may have induced some money demand instability. The first predates the beginning of the sample period, and is in any case adjudged ineffectual. The second is dated at November 1983.

I incorporate an indicator variable which takes on a value of unity at 1983.4 onwards, as well as a variable constructed as the interaction between the indicator variable and the relative money stock variable directly into the regression specification in (15). Since this procedure yielded very large standard errors, I implemented the following two step procedure: first estimate the long run relationship in levels, and second estimate an error correction model (ECM) using the identified cointegrating vector as the error correction term (ECT). The first step is accomplished using the Stock and Watson (1993) dynamic OLS procedure.

$$s_t = \xi_0 + \sum_{i=-2}^{-1} \beta_i \Delta X_{t+i} + v_t \quad (16)$$

where the leads and lags serve to account for endogeneity of the right hand side variables. The results of implementing this procedure are found in columns 1 and 2 of Table 3 (using narrow and broad money, respectively). The long run coefficients are between the solid horizontal lines.

Although the money variable still has the incorrect sign in column 1, it is no longer statistically significant at conventional levels. In fact, the only significant coefficient is that on relative prices, with a point estimate of 0.893. The reversion coefficient is very large, at -0.355, implying that a deviation half life of about 1.7 quarters. Still, this hardly constitutes a resounding affirmation of the augmented monetary approach given the sign on money.

Estimation with broad money yields a positive and significant long run coefficient. Moreover, income, inflation and relative prices now have the correct sign ascribed to them. However, the exchange rate does not exhibit any mean reversion. Further investigation indicates that the money stock responds to the disequilibrium (although the estimate is not statistically significant).¹⁰ This outcome is consistent with active targeting of the exchange rate via monetary policy.

Two currencies evidence no long run relationship: The Philippine peso and the Indonesian rupiah.¹¹ To check that the no-cointegration finding is not a function of the estimation technique, a DOLS estimation is also implemented. Neither a sticky-price monetary model, nor a simple Frenkel

¹⁰ The point estimate of -0.06 implies that a half life of a deviation closed by monetary policy is about 4 years (the 2 standard error bounds encompass no reversion, and a one year half life).

¹¹ In the former, the assumptions of perfect capital mobility and substitutability may be so strongly violated, given the extreme political unrest during the mid-1980s, that the posited long run relationship does not hold. I am unaware of any studies regarding covered interest differentials for the Philippines. However, the average interest differential is on the order of 10-11% per annum, while the annual rate of depreciation is on the order of 8%, suggesting an average political risk plus exchange risk premium on the order of 3%. See Chinn and Dooley (1997).

specification, yields cointegration with the expected signs. However, a Frenkel specification, augmented by the relative price variable and a time trend does yield a plausible set of estimates, as reported in column 2 of Table 3. Money and relative prices enter significantly. Repeating the two step procedure outlined for the Taiwanese dollar, I use the implied cointegrating vector as an error correction term. The resulting estimated rate of reversion is 0.560, suggesting that the Philippine peso does respond strongly to the fundamentals. The only caveat is that the trend term carries a large proportion of weight in the cointegrating vector; extrapolation of this trend effect beyond the sample period is problematic.

In the case of the Indonesian rupiah, identification of the long run parameters is hampered by the brevity of the floating rate period combined with a lack of recent money data. I repeat the two step procedure with the following modifications. The first is that I make the assumption that a long run relationship holds over the entire 1970-96 period, after accounting for structural breaks; the second, that inflation rates proxy for interest rates as in a Frenkel specification in equation (7) (since no interest rate series continuously spans the entire period; see Appendix Figure A8).

The two known structural breaks in the Indonesian money demand relationship that Dekle and Pradhan (1996) identify in 1983.2 and 1988.3 are econometrically accounted for by indicator variables.¹² These breaks are associated with liberalization efforts (decontrol of interest rates and opening of the banking sector to new entrants, respectively). The results are reported in Table 3 in columns 4 and 5 for the monetary and augmented monetary approaches. The first of these specifications indicates a negative long run coefficient on money, while in the second, money appears with the correct, albeit insignificant, coefficient. The relative price variable enters the second equation with correct sign and statistical significance. In Chinn (1997b) I find that oil prices, along with relative productivity

¹² Dekle and Pradhan (1996) using annual data identify statistical breaks in the cointegrating relationship for the years. The specific quarters correspond to the official implementation of new regulations.

differentials, are an important determinant of Indonesian real exchange rates; one interpretation of this correlation is that the price of oil works through wealth effects. Including the price of oil in the monetary specification, but *omitting relative prices*, yields the estimates in column 6. The long run coefficients on money and inflation are now statistically significant in the negative direction. Including both oil prices and relative prices yields the estimates in column 7. While the money coefficient is again in the correct direction, the only variable with statistical significance is the relative price variable. Given the results, the preferred specification is in column 5. This specification indicates that the modified model explains some portion of the long run behavior of the rupiah.

To assess the short run dynamics, I estimate the implied error correction model over the 1986.4-93.2 period, *using the cointegrating vector estimated from the entire sample*. The error correction term coefficient is the relevant one, if exchange rates respond to the fundamentals. In no case is the coefficient negative, except for that in column 4, the augmented monetary specification. Unfortunately, the coefficient is economically small (-0.006) and statistically insignificant; however a cursory glance at the time series for the rupiah (Appendix Figure A1) is sufficient to convince one that the monetary authorities were seeking to stabilize the currency around a trend.

In order to investigate how the adjustment to equilibrium takes place, I estimate a series of error correction models. The most obvious candidate for investigation is the Indonesian narrow money stock. The response of this variable to a disequilibrium is 0.07. In words, a 1% deviation induces the central bank to "lean against the wind" by increasing the money stock by .07% more than it otherwise would have. Unfortunately, this point estimate is not statistically significant. Interestingly, it appears that most of the adjustment is borne by the real sector -- in fact the relative price variable ω responds *very* strongly to a disequilibrium.¹³ This is probably due to the fact that importables play a large role in the

¹³ The relative price variable responds at a rate of 0.28 per quarter (p-value of .018).

price of Indonesian tradables. Consequently, due to the nature of the exchange rate regime it is not possible to isolate the short run dynamics of the exchange rate under free floating. However, one can be relatively confident that accelerated inflation will cause more rapid exchange rate depreciation.

6. CONCLUSIONS

The results in this paper suggest that the monetary approach has some applicability in East Asia. However, models that assume PPP for broad price indices are unlikely to be altogether successful, except in the cases of the Thai baht and perhaps the Singapore dollar, where the 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's greatest confidence is in the estimated relationship for the Korean won, in which case there is a statistically significant finding of cointegration, with money stocks and relative prices entering in the expected manner and rapid reversion to equilibrium. Allowing for some structural shifts, the long run relationships between the Indonesian rupiah and the Taiwan dollar and their respective fundamentals can also be explained. However, until these two central banks allow their currencies to float more freely, it will be difficult to identify *the short run dynamics* of the exchange rate (as opposed to those of money or prices).

The Singapore dollar results are simultaneously discouraging and encouraging. Money stocks enter in with incorrect (although marginally significant) sign. On the other hand, according to the results in Table 1, the interest differential, the key variable in the real interest differential model, shows up very statistically significant. In light of the substantial financial innovation that has taken place over

the sample period, perhaps one should not be surprised that the monetary aggregates appear with ambiguous effects.

Finally, the Malaysian ringgit and, to a lesser extent, the Philippine peso prove to be resistant to econometric investigation. Since these are two of the four currencies newly floating, this ignorance is particularly troubling.

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Data Appendix

The data are from IMF, *International Financial Statistics*, March 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.

□ 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. Malaysia income is proxied by industrial production. Taiwanese GDP, in 1991 New Taiwan \$. For Indonesia, and Singapore, GDP data is a centered 4 quarter moving average of annual data. Interpolation is via the formula:

$$\begin{aligned}Y_{\tau, Q1} &= .375 \times Y_{\tau-1} + .625 \times Y_{\tau} \\Y_{\tau, Q2} &= .125 \times Y_{\tau-1} + .875 \times Y_{\tau} \\Y_{\tau, Q3} &= .875 \times Y_{\tau} + .125 \times Y_{\tau+1} \\Y_{\tau, Q4} &= .625 \times Y_{\tau} + .375 \times Y_{\tau+1} \\for \tau &= 1971, \dots, 1991\end{aligned}$$

□ Interest rates are short term, interbank interest rates, IFS line 60b, in decimal form. Indonesian interest rates are long term, from Morgan-Guaranty database. 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 4-quarter difference of log(CPI).

□ Relative price variable:

$$-\alpha [(p^N - p^T) - (p^{N*} - p^{T*})] \approx \omega \equiv \log(PPI^{US}/CPI^{US}) - \log(PPI^*/CPI^*)$$

[which is appropriate if $\alpha = 0.5$, and CPI contains one half nontradables].

Table 1
ECM Regressions: Monetary Model of the Exchange Rate

Coeff	Pred	IN	KO	MA	PH	SI	TI	TI ^a	TH
ECT	(-)	-0.077 (0.199)	-0.047† (0.035)	-0.345** (0.084)	0.028 (0.130)	-0.202*** (0.076)	-0.216** (0.095)	-0.026 (0.057)	-0.805*** (0.206)
m-m*	(1)	0.339 (0.874)	0.816 (0.880)	-0.338** (0.078)	0.751 (1.340)	-1.012** (0.452)	-0.666*** (0.153)	0.466 (4.288)	0.124† (0.086)
Y-Y*	(-)	-3.653 (10.728)	-1.751 (1.619)	0.314 (0.066)	-13.064 (47.581)	1.011† (0.784)	0.219 (0.306)	-3.499 (8.799)	-0.112** (0.043)
i-i*	(-)	1.652 (4.095)	3.602 (3.603)	0.339 (0.322)	-23.093 (114.416)	-2.826*** (1.058)	-0.972 (1.395)	27.090 (57.330)	0.078 (0.349)
Π-Π*	(+)	1.218 (3.801)	-0.889 (2.166)	0.038 (0.359)	33.362 (153.898)	-0.880 (0.754)	0.189 (0.936)	-9.541 (27.021)	-0.058 (0.457)
T	(?)	-0.0031** (0.0012)				0.0024** (0.0010)			
\bar{R}^2		.998	.980	.858	.993	.981	.989	.989	.775
N		27	62	84	45	85	58	55	41
Smp1		86.4-93.2	80.3-95.4	73.4-94.3	81.2-95.4	74.1-95.1	81.3-95.4	82.2-95.4	85.1-95.1
LM(4)		0.769 [.575]	1.749 [.161]	1.191 [.324]	0.751 [.569]	1.870 [.128]	2.350 [.074]	2.811 [.046]	0.371 [.582]
Iter.		113	15	70	43	47	150	150	59

Notes: Regression coefficients from Nonlinear Least Squares. †{*}(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. LM(4) is the Breusch-Godfrey F- test for serial correlation of order 4 [p-values in brackets]. N is the effective number of observations included in the regression. Iter. is the number of iterations necessary for convergence. Sample is the sample period.
 \bar{a} / Broad money.

Table 2
ECM Regressions: Augmented Monetary Model of Exchange Rates

Coeff	IN	KO	MA	PH	SI	TI	TH	
ECT	(-)	-0.071 (0.255)	-0.132** (0.057)	-0.419** (0.151)	-0.139 (0.140)	-0.148† (0.091)	0.320*** (0.100)	-0.777*** (0.250)
m-m*	(1)	0.293 (0.987)	0.636* (0.364)	0.025 (0.215)	0.343 (0.465)	-1.222† (0.782)	-0.437*** (0.140)	0.069 (0.094)
Y-Y*	(-)	-3.712 (14.232)	-0.391† (0.304)	-0.092 (0.467)	-0.839 (3.172)	1.830 (1.658)	-0.013 (0.281)	-0.081* (0.044)
i-i*	(-)	1.980 (6.934)	1.220 (1.396)	0.206 (1.045)	9.488 (8.367)	-4.939† (3.582)	-0.032 (1.359)	0.007 (0.366)
Π-Π*	(+)	0.916 (4.187)	0.331† (1.308)	-1.324 (1.495)	-8.544 (8.161)	-2.200 (2.318)	-0.930 (0.861)	0.160 (0.457)
ω	(1)	1.115 (6.127)	2.803* (1.685)	0.623 (1.037)	-0.975 (2.899)	-0.365 (0.634)	0.552 (0.777)	-0.047 (0.200)
τ	(?)	-0.0027† (0.0017)				0.0031*** (0.00121)		
\bar{R}^2	.998	.976	.885	.993	.981	.992	.804	
N	27	62	41	45	83	58	41	
Smpl	86.4-93.2	80.3-95.4	84.3-94.3	81.2-95.4	74.3-95.4	81.3-95.1	85.1-95.1	
LM(4)	0.952	2.127	2.116	0.525	1.842	2.018	0.861	
	[.495]	[.100]	[.147]	[.719]	[.029]	[.119]	[.517]	
Iter.	123	52	16	20	54	124	82	

Notes: Regression coefficients from Nonlinear Least Squares. †{*(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. LM(4) is the Breusch-Godfrey F test for serial correlation of order 4 [p-values in brackets]. N is the effective number of observations included in the regression. Iter. is the number of iterations necessary for convergence. Sample is the sample period.

Table 3
Two-Step Regressions: Taiwan, Philippines and Indonesia

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Coeff	TI ^a	TI ^b	PH ^c	IN ^d	IN ^d	IN ^d	IN ^d
ECT	-0.355*** (0.109)	0.118† (0.085)	-0.560*** (0.154)	0.007 (0.009)	-0.006 (0.025)	0.008 (0.011)	0.007 (0.025)
m-m*	-0.209† (0.154)	0.816** (0.360)	0.088*** (0.020)	-0.547*** (0.176)	0.127 (0.171)	-0.232 (0.199)	0.099 (0.210)
y-y*	0.246* (0.138)	-1.473** (0.634)	-0.284 (0.456)	3.558*** (0.804)	0.587 (0.743)	3.060*** (0.743)	1.032 (0.977)
i-i*	-0.025 (0.333)	1.364** (0.573)					
π-π*	0.270 (0.616)	2.361† (1.669)	0.219 (0.241)	-0.692*** (0.199)	0.878** (0.398)	-0.847** (0.335)	0.515 (0.415)
ω	0.893*** (0.307)	1.362** (0.643)	1.471*** (0.325)	1.842*** (0.405)		1.440*** (0.470)	
p ^{oil}					0.225*** (0.082)	0.080 (0.079)	
time			-0.015*** (0.001)				
D84q4	6.189*** (2.224)	1.362† (6.218)					
D84q4 x(m̂)	-0.298*** (0.108)	-0.460† (0.298)					
D83q2				-0.633*** (0.082)	-0.132† (0.103)	-0.483*** (0.083)	-0.190* (0.107)
D88q3				0.057† (0.104)	-0.036 (0.042)	0.245*** (0.093)	0.048 (0.099)
R ²	.407	.353	.345	.142	.156	.079	.107
N	57	58	45	27	27	27	27
Smp1	82.2-96.2	82.1-96.2	81.3-96.2	86.4-93.2	86.4-93.2	86.4-93.2	86.4-93.2
LM(4)	1.665 [.179]	1.675 [.176]	0.999 [.429]	2.368 [.116]	1.538 [.264]	1.943 [.180]	1.024 [.445]

Notes: ECT is coefficient on error correction term in second stage regression. Long run coefficients from Dynamic OLS (Stock and Watson, 1993, DOLS1 estimator), with lags from +2 to -2 of first differences. †{*}[**][***] indicates significance at the 20%{10%}(5%)[1%] MSL. DOLS standard errors are heteroscedasticity consistent, calculated using Bartlett window with k=5. Summary statistics refer to second stage regression. LM(4) is the Breusch-Godfrey F test for serial correlation of order 4 [p-values in brackets]. N is the effective number of observations included in the second stage regression. Sample is the sample period.

a/ 4 lags of income and relative prices. Narrow money.

b/ 4 lags of income and 3 lags of relative prices. Broad money used.

c/ 2 lags of exchange rates, income and relative prices, and 4 lags of inflation.

d/ 5 lags of income.

APPENDIX

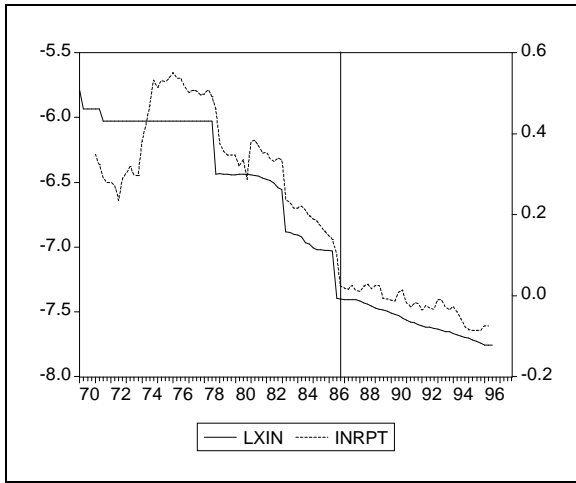


Figure A1: Indonesian rupiah and relative prices

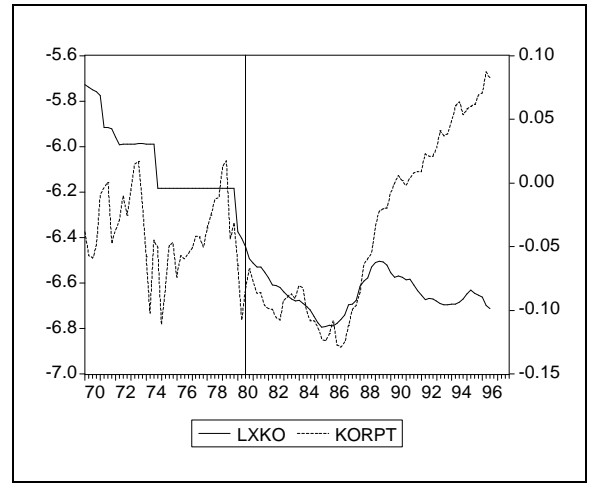


Figure A2: Korean won and relative prices

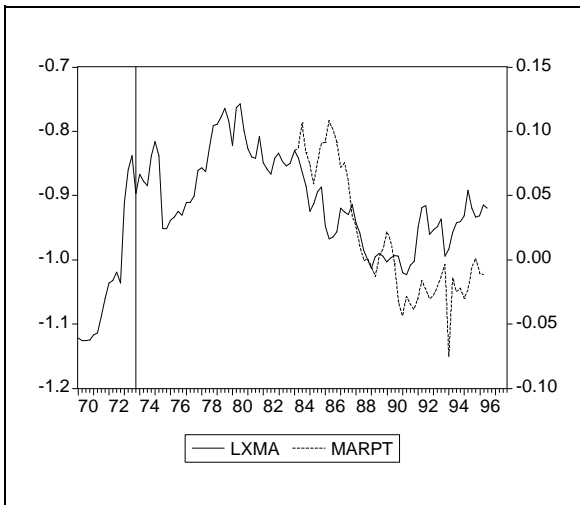


Figure A3: Malaysian ringgit and relative prices

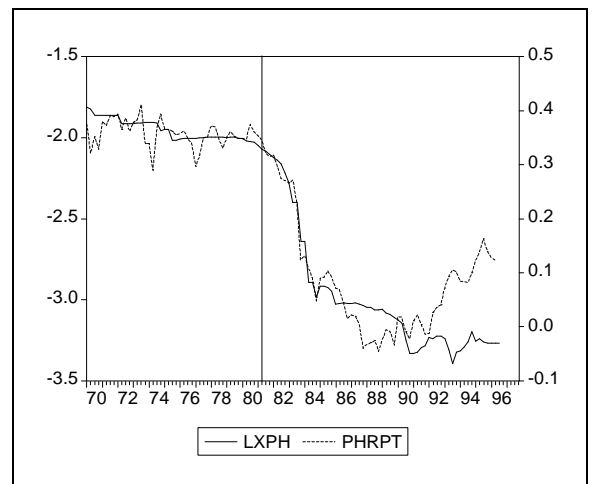


Figure A4: Philippine peso and relative prices

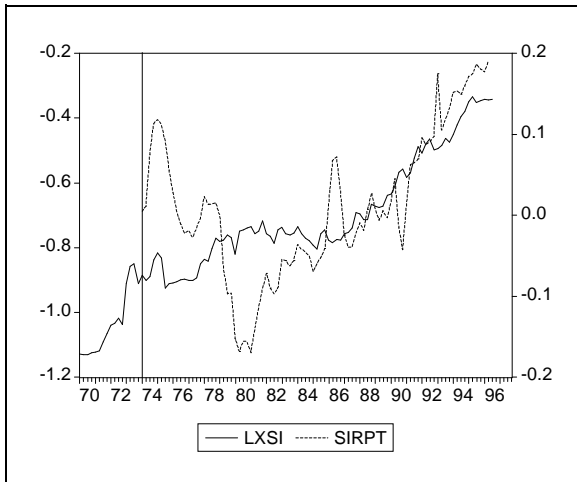


Figure A5: Singapore dollar and relative prices

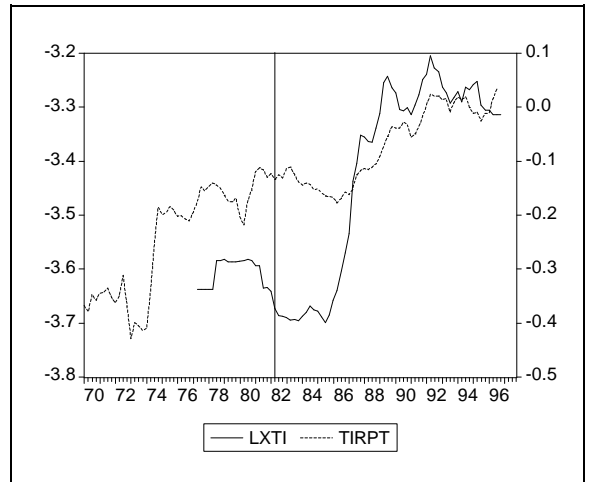


Figure A6: New Taiwan dollar and relative prices

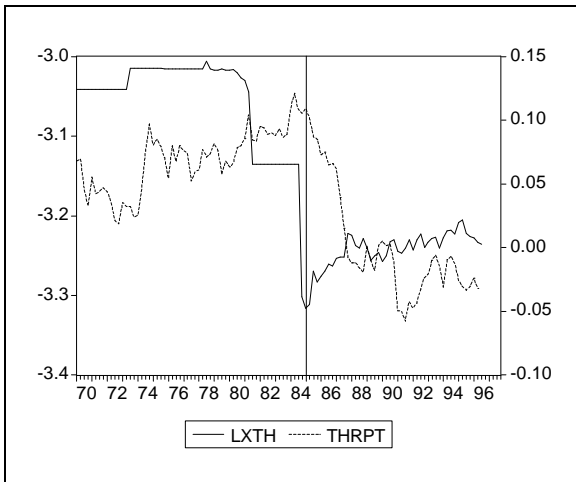


Figure A7: Thai baht and relative prices

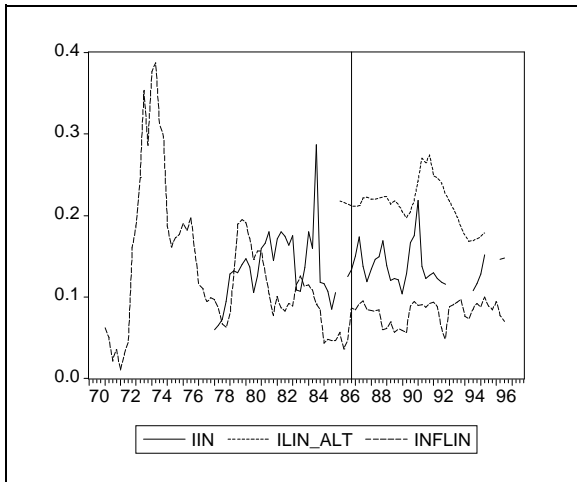


Figure A8: Indonesian money market rate, lending rate and inflation rate

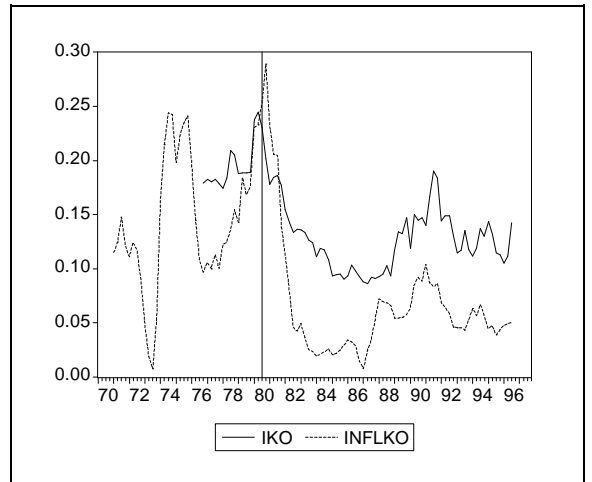


Figure A9: Korean money market rate and inflation rate

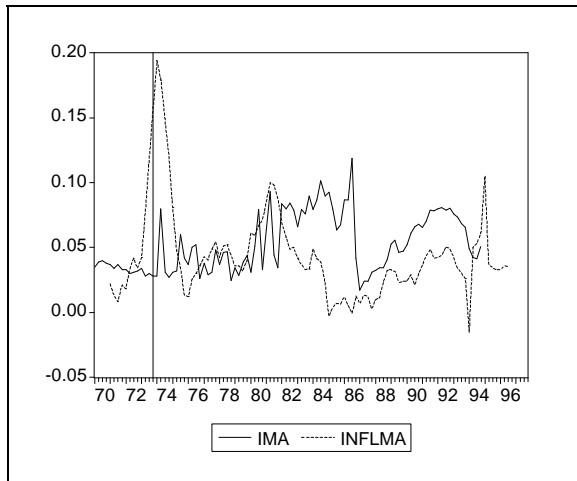


Figure A10: Malaysian money market rate and inflation rate

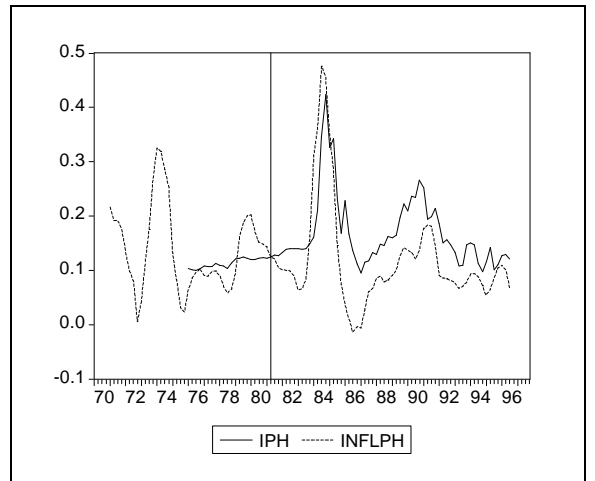


Figure A11: Philippine 3 month T-bill rate and inflation rate

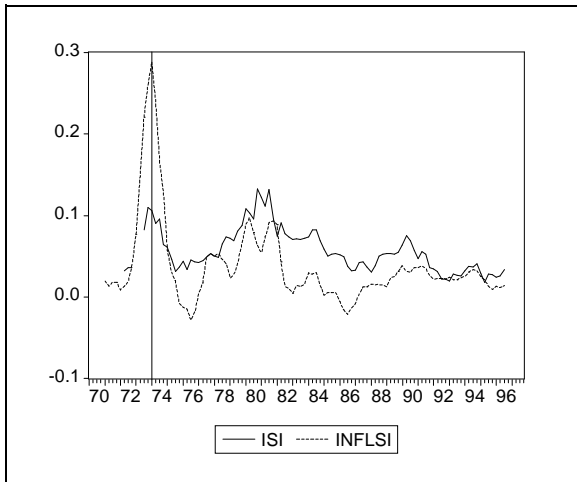


Figure A12: Singapore money market rate and inflation rate

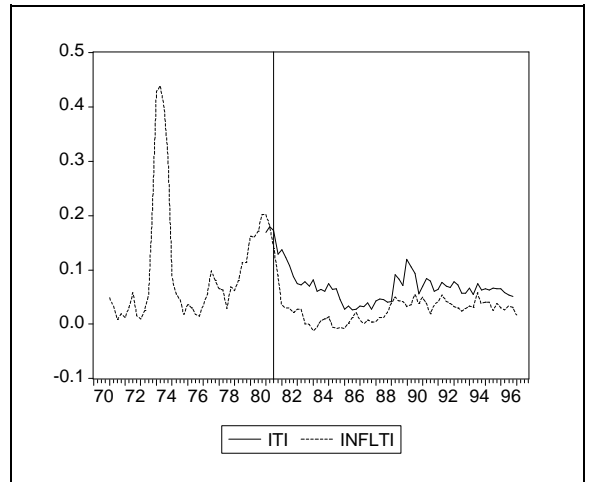


Figure A13: Taiwanese money market rate and inflation rate

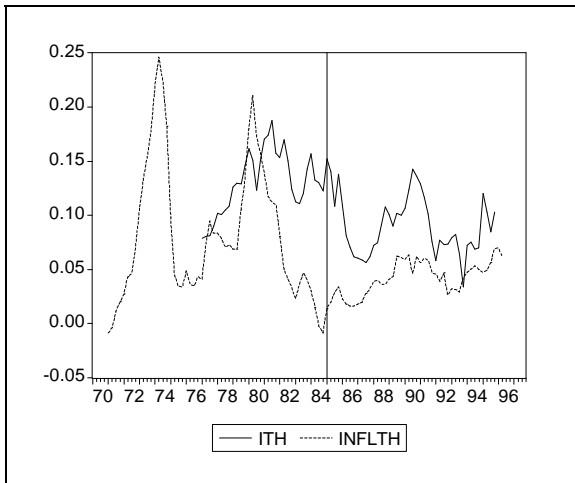


Figure A14: Thai money market rate and inflation rate