

Before the fall: were East Asian currencies overvalued?

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Abstract

Two major approaches to identifying the equilibrium exchange rate are implemented. First, the concept of purchasing power parity (PPP) is tested and used to define the equilibrium real exchange rate for the Hong Kong dollar, Indonesian rupiah, Korean won, Malaysian ringgit, Philippine peso, Singapore dollar, New Taiwanese dollar and the Thai baht. The calculated PPP rates are then used to evaluate whether these seven East Asian currencies were overvalued. A variety of econometric techniques and price deflators are used. As of May 1997, the HK\$, baht, ringgit and peso were overvalued according to this criterion. The evidence is mixed regarding the Indonesian rupiah and NT\$. Second, a monetary model of exchange rates, augmented by a proxy variable for productivity trends, is estimated for five currencies. An overvaluation for the rupiah and baht is indicated, although only in the latter case is the overvaluation substantial (17%). The won, Singapore dollar and especially the NT\$ appear undervalued according to these models. © 2000 Elsevier Science B.V. All rights reserved.

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1. Introduction

One of the key questions surrounding the 1997 East Asian crises is whether the precipitous currency declines could have been predicted. At first glance, there would seem to be a substantial body of empirical work answering this question in the affirmative. Goldfajn and Valdes (1999) have conducted an exhaustive cross country examination which demonstrates that overvaluation is a precursor of a currency crash. In addition, a number of papers have pointed to exchange rate overvaluation as a robust empirical determinant of currency crises (Frankel and Rose, 1996a; Sachs et al., 1996; Kaminsky and Reinhart, 1999). Hence, the presence of overvaluation is potentially important for policy purposes because of its role as a component of an early warning system (see e.g. Berg et al., 2000).

In the East Asian case, all the regional currencies, save Hong Kong's, lost value; hence a natural conclusion is that these currencies were overvalued on the eve of the crisis. In fact, however, currencies not previously thought to be overvalued such as the Korean won, Singapore dollar and Taiwanese dollar, *also* depreciated suggesting that the characterization of overvaluation is often applied tautologically.

In this paper, the issue of 'overvaluation' is taken seriously from an econometric perspective. However, before this endeavor can be accomplished, one must take a stand upon the theoretical framework that will guide the statistical analysis. There are at least three broad definitions in use (see Williamson, 1994; Milesi-Ferretti and Razin, 1996):

1. Price based criteria, such as purchasing power parity and its variants.
2. Model based criteria, based on a formal model of nominal exchange rates.
3. Solvency and sustainability based criteria, which make reference to trends in the current account and the external debt to GDP ratio.

It turns out that the relevance of each criterion is inversely related to the difficulty of implementing it. Price based criteria are relatively easy to implement, but do not address the economically interesting question of whether a particular exchange rate is at an optimal level, besides that defined by a no-arbitrage condition. On the other hand, the sustainability measures can make reference to an optimal level, but are very difficult to calculate as they require a fully-fleshed out macroeconomic model. Moreover, in order to make a statement about optimality, they need to take a stand on representative agent behavior.¹ Consequently, in this paper a more modest goal of implementing the first two criteria is set forth.

The paper proceeds in the following manner. In Section 2, the price based measures are described, the tests for purchasing power parity undertaken, and the

¹See e.g. Bayoumi et al. (1994) and Driver and Wren-Lewis (1999) for developed countries, and Hinkle and Montiel (1999) for less developed countries. Furthermore, overvaluation may not be very relevant in models involving moral hazard, or in insurance-based explanations (Chinn et al., 1999).

calculations of equilibrium rates reported. In Section 3, a monetary model of nominal exchange rates, augmented by a relative price variable that proxies for productivity differentials, is estimated, and then used to calculate the equilibrium exchange rates. Section 4 concludes.

To anticipate the results, I find that there is evidence that real exchange rates for most East Asian currencies are mean reverting over the 1975–1996 period, although this finding is specific to particular base currencies (US dollar or Japanese yen) and deflators (CPI and PPI). The Hong Kong dollar, Malaysian ringgit, Philippine peso, and Thai baht were overvalued on the eve of the 1997 crisis. Surprisingly, the won appears undervalued (as does the Singapore dollar). There is mixed evidence regarding the Indonesian rupiah and NT\$. The former is between 6% undervalued and 30% overvalued, while the latter is between 9% undervalued and 7% overvalued.²

Evidence of a long-run (cointegrating) relationship between exchange rates, monetary fundamentals, and the relative price of tradables and non-tradables is also obtained. The implied dis-equilibria are consonant with those obtained using the price-based measures for Singapore, Thailand and Taiwan, but not for Indonesia and Korea. For the former, an 8% overvaluation is detected, while the exchange rate of the latter is approximately at equilibrium.

It appears that it is possible to detect evidence of overvaluation prior to the East Asian crises, but the evidence is not resounding. In particular, the estimated misalignments are not usually very large, and are not well correlated with the severity of the subsequent currency crashes. This leads to the conclusion that overvaluations are neither a necessary nor sufficient condition for a currency crisis.

2. Price based measures of equilibrium real exchange rates

2.1. Theoretical background to purchasing power parity

The equilibrium exchange rate is often associated with an international version of the Law of One Price (LOP): abstracting from transport costs, identical goods in different countries have the same price, when expressed in common currency terms. An arbitrage argument is usually offered to explain why this condition should hold.

At this juncture, data limitations intrude. Typically, one does not have prices for identical goods; rather one observes (log) price indices, p , for bundles of goods. These indices do not usually ascribe the same weights to each good, nor are the quality attributes of these goods identical, so that direct testing of LOP is not

²Note that *changes* in the real exchange rate are not interpreted as measures of overvaluation, as for instance Corsetti et al. (1998) do. Real appreciation may very well be a statistically significant precursor of a currency crisis, but its role is not examined here.

possible. What can be tested is how well purchasing power parity (PPP) holds up to a constant κ which depends upon the base year of the price indices,

$$s_t + p_t^* = p_t + \kappa \quad \text{implies} \quad q_t \equiv s_t - p_t + p_t^* = \kappa \quad (1)$$

where s and q are the log nominal and real exchange rates, respectively.

The consensus in the profession is that PPP clearly does not hold continuously, and perhaps does not hold even over long periods when one interprets the price index as one pertaining to a broad set of goods and services (see Froot and Rogoff, 1995).³ Since some of the items in a typical consumption or production bundle are not tradable and subject to international price pressures from international trade, this result is not completely unexpected. On the other hand, since consumer bundles might be more similar across countries than producer or wholesale bundles, consumer price indices (CPIs) may provide a more consistent measure of price levels and thus of real exchange rates.

Adopting an agnostic view on the issue, calculations using a measure using as a proxy the wholesale or producer price index (WPI or PPI), which covers goods considered to be highly tradable, are also presented. Finally, if the countries of interest are primarily exporting to third country markets, then the export price index may in principle be the more appropriate deflator. In practice, export unit value indices are notoriously subject to measurement error; moreover, the composition of the bundles of exports are likely to vary even more widely across countries than the corresponding PPI or CPI bundles.

2.2. Methodology

The standard approach to testing for an equilibrium real exchange rate based on prices is to implement a unit root test, such as the following Augmented Dickey–Fuller (ADF). As is well known, such tests possess low power against local alternatives. Hence previous attempts to find mean reversion in the post-Bretton Woods period, using univariate techniques, have usually failed.

The low power of such unit root tests may be due to the imposition of inappropriate common factor restrictions implicit in the ADF specification (Kremers et al., 1992). In estimating an ADF on the real exchange rate, one forces the short-run dynamics for the exchange rate and both price levels to be the same. In principle, there is no reason to believe that this condition should hold. A more general specification implied by cointegration is:

$$\Delta s_t = \gamma_{10} + \Phi_1 \text{ECT}_{t-1} + \sum_{i=1}^k \gamma_{1i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{1i} \Delta p_{t-i} + \sum_{i=1}^k \nu_{1i} \Delta p_{t-i}^* + \varepsilon_{1t}$$

³For a contrasting view, see the recent panel work by Frankel and Rose (1996b) and MacDonald (1996).

$$\Delta p_t = \gamma_{20} + \Phi_2 \text{ECT}_{t-1} + \sum_{i=1}^k \gamma_{2i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{2i} \Delta p_{t-i} + \sum_{i=1}^k v_{2i} \Delta p_{t-i}^* + \varepsilon_{2t} \quad (2)$$

$$\Delta p_t^* = \gamma_{30} + \Phi_3 \text{ECT}_{t-1} + \sum_{i=1}^k \gamma_{3i} \Delta s_{t-i} + \sum_{i=1}^k \zeta_{3i} \Delta p_{t-i} + \sum_{i=1}^k v_{3i} \Delta p_{t-i}^* + \varepsilon_{3t}$$

$$\text{ECT} \equiv [\beta_1 s + \beta_2 p + \beta_3 p^*]$$

Johansen (1988) and Johansen and Juselius (1990) describe the maximum likelihood method of estimating this vector error correction model (VECM) and for testing cointegration. A likelihood ratio test can be applied to the restriction that $(\beta_1 \beta_2 \beta_3)$ takes on the value $(1 \ -1 \ 1)$. Cheung and Lai (1993b) are among the first to apply this approach; they find evidence for cointegration, but reject the unitary coefficient restriction implied by strict PPP.

Since one has prior information on the form of the cointegrating vector, a more powerful test of the null of no cointegration against the alternative of cointegration with a pre-specified cointegrating vector can be applied. Horvath and Watson (1995) tabulate the critical values for a Wald test on the Φ coefficients equaling zero. Rejection of this null hypothesis implies cointegration because the variables, either singly or jointly, revert back to the conditional mean defined by the cointegrating vector. Edison et al. (1997) apply this test and find mixed evidence for PPP for the G-7 currencies during the post-Bretton Woods.

2.3. Data

The countries under study include Hong Kong–PRC, Indonesia, Korea, Malaysia, the Philippines, Singapore, Taiwan, and Thailand. Bilateral real exchange rates against the US dollar and the Japanese yen are generated. Most series are from the IMF's *International Financial Statistics*, and span the 1970.01–1997.09 period. The Taiwanese data are from Bank of China, *Financial Statistics*, various issues, as recorded in Federal Reserve Bank of San Francisco electronic database. The exchange rates are end-of-month data, expressed in US\$/local currency unit [inverse of *IFS* line *ae*]. Exchange rates against the yen are calculated by dividing by the US\$/yen rate.

For the broad deflator, the CPI [*IFS* line 64] is used. The 'tradable' price deflator is proxied by the PPI or WPI data reported in *IFS* line 63. The Indonesian PPI data exclude petroleum products [*IFS* line 63a], while the Hong Kong PPI data are quarterly, from the Hong Kong Department of Census and Statistics. The export price data is the export unit value index [*IFS* line 74].

In principle, one might like to use a trade weighted measures of the real exchange rate. The problem that one encounters is that the pattern of trade flows change substantially over the sample period and hence so too do the appropriate trade-weights. Nonetheless, the trade-weighted CPI deflated real exchange rates

calculated by the IMF, and the PPI deflated real exchange rates reported by Morgan-Guaranty⁴ are also analyzed.

The time series patterns of the multilateral exchange rates do not differ greatly between those of the two bilateral exchange rates, except in a couple instances, most prominently the Taiwanese dollar.⁵ It turns out that the results using only the bilateral exchange rates will in general be sufficient to make inferences regarding stationarity of the real exchange rate. This outcome makes sense as the US and Japan accounted for a large portion of these countries' imports and exports in 1996.

Overall, there is no obvious difference between trends in CPI deflated dollar or yen rates. Trends in PPI adjusted exchange rates are typically smaller (in absolute value) than their CPI-deflated counterparts, suggesting that the PPI deflators may yield greater evidence of stationarity. Interestingly, the real exchange rates defined using export price indices exhibit substantial trend depreciations, with the exception of the Singapore dollar. The presence of substantial, but imprecisely estimated, drift terms suggests that such price indices are subject to greater measurement error. In particular, the export bundles of these newly industrializing countries have probably changed substantially over time, introducing drift in the price indices (the Japanese yen is an exception; it is likely that the composition of the Japanese export bundle has changed less drastically over the sample period).

2.4. Empirical results

The ADF test was applied to all the real exchange rate series. Only approximately four cases appeared to be stationary, a proportion about consistent with what would be expected to occur by chance.

The results of applying the Johansen procedure to bilateral rates against the US\$ over the 1975.01–1996.12 period are reported in Appendix Table A1.⁶ There is substantial evidence of cointegration, even using the finite sample critical values of Cheung and Lai (1993a) for almost all cases. However, with few exceptions (such as the Philippine peso), the estimates do not conform to the PPP hypothesis when either the CPI or export price index are used as a deflator. Using the PPI, one finds that the point estimates are closer to their hypothesized values, and in the case of the Singapore dollar, conform very closely to posited values. [Greater discussion of these estimates are contained in Chinn (1999a).]

Table 1 reports the results of applying the Horvath–Watson procedure for the

⁴The IMF series are described in Zanetto and Desruelle (1997). The Morgan Guaranty series are the 'broad' effective exchange rate indices, based on 1990 trade weights for the 1987–1997 period. Prior to that, the 1980 trade weights are used. See *World Financial Markets* (1993). Note that the HK series is calculated using a Hong Kong retail price series, rather than a PPI.

⁵Note that while there are large divergences in the early period, the analysis will be conducted on the floating rate period data, starting from 1975. Hence, the large divergences evident in the early 1970s do not influence the subsequent econometric analysis.

⁶Using a longer sample encompassing the post-crisis period would only increase the ability to detect mean reversion (Fujii, 2000).

Table 1
Horvath–Watson test results for US\$, Japanese yen and trade-weighted exchange rates^a

	HK	IN	JP	KO	MA	PH	SI	TH	TI
Panel 1.1: CPI									
US\$ <i>k</i>	12	1	12	12	11	11	12	1	1
<i>W</i>	11.549**	17.913***	4.484	14.323***	5.121	1.456	2.277	18.523***	10.544*
yen <i>k</i>	1	12	na	12	12	12	12	12	12
<i>W</i>	8.773	9.309°		13.005**	4.104	8.674	6.202	4.936	16.914***
TWXR <i>k</i>	1	1	1	1	1	1	1	1	1
<i>W</i>	7.414	4.569	8.372	4.059	1.322	2.356	2.061	1.408	9.198°
Panel 1.2: PPI									
US\$ <i>k</i>	na	1	12	12	2	12	12	5	4
<i>W</i>		10.544*	3.497	13.742**	2.649	4.036	7.095	12.050**	4.368
yen <i>k</i>	na	12	na	11	1	2	3	2	6
<i>W</i>		4.640		3.465	12.413**	9.867*	2.919	7.346	2.267
TWXR <i>k</i>	1	1	1	1	1	1	1	1	1
<i>W</i>	10.098*	2.990	7.066	5.918	10.318*	8.674	0.111	1.403	1.835
Panel 1.3: Export price indices									
US\$ <i>k</i>	12	11	12	1	4	12	1	1	na
<i>W</i>	9.630°	5.366	3.926	3.926	7.690	8.374	3.337	22.270***	
yen <i>k</i>	1	11	na	1	1	2	2	1	na
<i>W</i>	6.198	3.838		5.110	0.762	4.705	3.323	2.644	

^a Notes: Asterisks indicate significance at the *10%, **5%, or ***1% MSL; ° indicates borderline significance. *k* is the number of first difference lags included in the VECM (selected by Schwartz Information Criterion, for lags up to 12). *W* is the Wald statistic. Critical values are *9.72, **11.62, and ***15.41, from Horvath and Watson (1995). TWXR denotes trade weighted exchange rate (PPI deflated series from Morgan Guaranty, CPI deflated series from IMF).

US\$ and the yen bilateral rates and the trade weighted exchange rate (TWXR). The export price deflated rates can be dispensed with immediately, as only the Thai baht shows up as stationary. Turning to the PPI based results, one finds that the Wald test statistic rejects the no-cointegration constraint for the trade weighted Hong Kong dollar, the rupiah, won and baht (against the US\$) and the ringgit and Philippine peso (against the yen).⁷

Some other interesting results are obtained. First, the PPI-deflated won/yen rate does not mean-revert, which is surprising given the apparently close link between the Korean and Japanese economies. Second, the trade weighted indices do not typically evidence mean reversion, with the exception of the PPI deflated Hong Kong dollar and the Malaysian ringgit. This finding may obtain because of the shifts in the trade weights used in calculating the Morgan-Guaranty series.

Table 1 also indicates that the US\$ based HK\$, rupiah, baht, won and NT\$ are stationary. The last two are also stationary against the yen. For the NT\$, the rejection of the no-cointegration null is at the 1% MSL for the yen, and only at the 10% for the US dollar. This finding of CPI cointegration outcome mirrors the finding that a productivity based tradables/non-tradables model does not explain the New Taiwan dollar (Chinn, 2000). For the won, the stationarity appears to be greater for the US\$ based rate.

Considering all the CPI- and PPI-deflated rates (against the dollar, yen and multilateral), evidence of mean reversion is found for all the region's currencies. Therefore, the results reported above are more favorable to the PPP hypothesis than those obtained in previous studies of the East Asian currencies. Phylaktis and Kassimatis (1994) and Fukuda and Kano (1997) find mean stationarity in PPI deflated bilateral exchange rates of the won and peso. Lee (1999) finds mean reversion for the PPI deflated rupiah, won, ringgit, peso and Singapore dollar (expressed against the US dollar) over longer samples spanning both the pre- and post-Bretton Woods periods. Note that Bahmani-Oskooee (1993), Tang and Butiong (1994), Baharumshah and Ariff (1997) and Chou and Shih (1995) also find evidence of cointegration for several currencies, but reject the symmetry and proportionality conditions that are required for mean reversion in real exchange rates.

2.5. *Estimated equilibrium rates*

Based on the Johansen test result for Singapore and the Horvath–Watson test results, mean stationarity of the real rate (and not merely a cointegrating relationship) is found in for all currencies, with respect to at least one reference currency (dollar or yen) or deflator (CPI or PPI). The mean real exchange rates are estimated using a regression of the real rate on a constant over the 1975–1996 period. In omitting the post-crisis observations, the procedure biases against

⁷Since Morgan-Guaranty does not report the nominal trade-weighted rates corresponding to the real exchange rates, I use the IMF's nominal trade weighted series as the nominal exchange rate and to infer the rest-of-world PPIs.

detecting an overvaluation. Clearly, including post-crisis data would increase the chances of finding an overvaluation.

While the cointegration results are specific to the numeraire (dollar, yen or multilateral) or deflator, for the sake of completeness all the deviations from mean as of 1997.05 are reported in Table 2, with the entries in **bold face** denoting cases where the calculations are appropriate, in light of the cointegration tests.

The results indicate a May 1997 overvaluation of the Malaysian ringgit (14%), Philippine peso (10%) and Thai baht (7%).⁸ On the other hand, the Indonesian rupiah, Korean won and Singapore dollar appear to be undervalued. Of these currencies, the overvaluation measure yields the most counter-intuitive results for the rupiah and won, two currencies that suffered precipitous declines in values. Calculating the deviations over the 2 years preceding the crisis does not change the overall pattern of results very much.

An alternative estimate of the equilibrium rate, validated by the stationarity findings of Section 2.4, allows for a trend in the CPI-deflated exchange rate is calculated by estimating over the same 1975–1996 period the regression,

$$q_t^{\text{CPI}} = \mu_0 + \mu_1\tau + v_t \quad (3)$$

where τ is a time trend. The estimated misalignments τ are reported in Table 3, with the valid entries indicated in **bold face**.

According to the cointegration tests, the CPI based dollar measures are valid for the HK\$, rupiah, baht, won and NT\$ (both the yen and the dollar in the last two cases). For the last currency, the dollar measure is the most appropriate (both the Johansen and Horvath–Watson results agree), and here the evidence is for a slight undervaluation. The won's undervaluation against the dollar is only 2%, as compared to the 9% in the PPI calculations. Overall, the won, HK\$, and baht calculations are in concurrence with those obtained using the PPI measures. Only in the case of the Indonesian rupiah does it appear that the use of the CPI yields a substantially different view: using the CPI, a 30% overvaluation is implied. It should be noted, however, that the PPI deflated real exchange rate exhibits stronger evidence of mean reversion than does the CPI deflated rate.

The equilibrium rates and actual levels from 1990 onward are plotted in Figs. 1–8. The implied over- and undervaluations derived from the PPI-measures are broadly consistent with historical accounts.⁹

⁸While the Hong Kong dollar is also overvalued by some 20%, according the calculations, the interpretation of the Hong Kong calculation is problematic, since Morgan Guaranty uses a retail price index as a proxy for the PPI. Over the 1990–1998 period, the retail price index has moved more in tandem with the CPI than the PPI, prompting concerns about the robustness of this particular conclusion.

⁹For instance, this measure implies that the Singapore dollar was overvalued in 1979–1982 period, while Moreno (1988: 192) asserts that Singapore's industry lost competitiveness during this period. By contrast, Hong Kong did not experience a substantial deterioration in competitiveness during this period, a view confirmed by the Hong Kong export-price deflated measure. Perhaps the strongest confirmation of this approach's utility comes from the overvaluation indicated at the end of the 1970s, an overvaluation that coincides with the extreme deterioration in the Korean external accounts.

Table 2

Deviations from PPP as predicted by PPI-deflated real rates (misalignment = $q_t^{\text{PPI}} - \hat{c}$)^a

	Hong Kong	Indonesia	Japan	Korea	Malaysia	Phil.	Singapore	Thailand	Taiwan
US\$	-0.046	-0.056	0.089	-0.093	0.078	0.189	-0.058	0.070	-0.029
yen	-0.225	-0.145	na	-0.182	0.136	0.100	-0.150	-0.019	-0.079
TWXR	0.204	-0.252	0.006	-0.184	-0.041 ^{b/}	0.097	0.044	-0.034	-0.061

^aNotes: $q = s - p + p^*$, where S is measured in US\$/local currency unit, yen/local currency unit or an index of the trade-weighted value of the local currency, p and p^* are log CPIs. A positive (negative) number indicates an overvaluation (undervaluation) of the local currency. Figures in **bold face** indicate valid estimates of misalignment, according to the cointegration tests.

Table 3

Deviations from PPP as predicted by trends in CPI-deflated real rates (misalignment = $q_t^{\text{CPI}} - \hat{q}_t^{\text{CPI}}$)^a

	HK	Indonesia	Japan	Korea	Malaysia	Phil.	Singapore	Thailand	Taiwan
US\$	0.159	0.303	-0.160	-0.024	0.167	0.237	0.126	0.133	-0.087
yen	0.371	0.463	na	0.136	0.327	0.397	0.286	0.293	0.074
TWXR	0.052	0.313	-0.302	-0.082	0.222	0.105	0.110	0.154	-0.087

^aNotes: $q = s - p + p^*$, where S is measured in US\$/local currency unit, p and p^* are log CPIs. A positive (negative) number indicates an overvaluation (undervaluation) of the local currency. Figures in **bold face** indicate valid estimates of misalignment, according to the cointegration tests. Predictions from Eq. (3) (see text).

2.6. Comparisons and robustness checks

The most commonly used numeraire is a trade weighted exchange rate (e.g. Sachs et al., 1996; Goldfajn and Valdes, 1999). The resulting deviations are reported in the third row of Table 3. If one were anticipating pervasive exchange rate overvaluation on the eve of the 1997 crises, then this methodology would ratify such expectations. The rupiah is overvalued by 31%, the ringgit by 22% and the baht by 15%. Such findings buttress the argument that dollar pegs, combined with the dollar's appreciation against the yen (undervalued by 16% in these calculations), were a major impetus for the currency crises (Ito et al., 1998). Yen based calculations yield even greater estimates of overvaluation. Of course, there is no statistical evidence to justify the use of either of these measures.

Given our uncertainty regarding all types of PPP calculations, it makes sense to undertake some robustness checks against the use of different sample periods [these detailed results are reported in Chinn (1999a)]. First, the equilibrium values

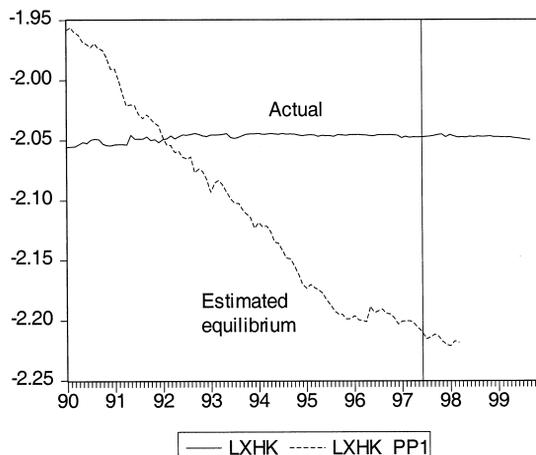


Fig. 1. Hong Kong dollar/US dollar exchange rate and CPI equilibrium rate.

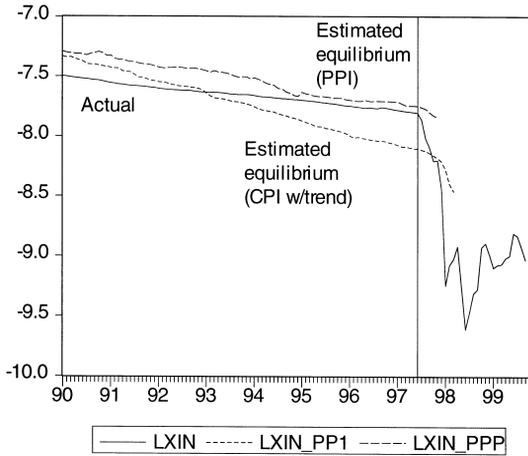


Fig. 2. Indonesian rupiah/US dollar exchange rate, CPI and PPI equilibrium rates.

were recalculated using the 1986.01–1996.12 period instead of the 1975.01–1996.12 span. If the real exchange rate series were truly mean stationary, changing the sample period should not matter very much, and in fact, the estimates do not change significantly, with the exception of the Indonesian rupiah. In this case, the rupiah is estimated to be approximately 9% overvalued as of 1997.05, as well as for the 2-year period preceding that.¹⁰

3. A model-based measure of overvaluation

3.1. The monetary model of nominal exchange rates

The Section 2 has provided a framework for estimating the long-run equilibrium real exchange rate. In order to obtain a short-run model, one may wish to relax some of the assumptions embodied in either the PPP or productivity-based formulations. Solving for the nominal exchange rate in Eq. (1), and substituting out the price levels with inverted money demand functions yields the following expression for the monetary model of the exchange rate:

$$s_t = \beta_0 + \beta_2(m_t - m_t^*) + \beta_3(y_t - y_t^*) + \beta_4(i_t - i_t^*) + \beta_5(\pi_t - \pi_t^*) + \beta_6\omega \quad (4)$$

$$\omega_t \equiv \left[(p_t^T - p_t^N) - (p_t^{T*} - p_t^{N*}) \right]$$

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

¹⁰See also the PPP based estimates presented in Furman and Stiglitz (1998).

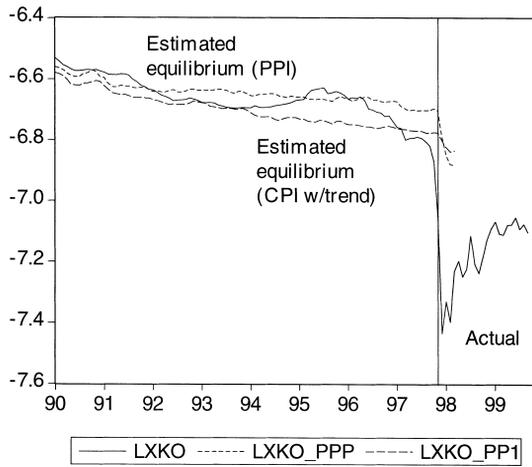


Fig. 3. Korean won/US dollar exchange rate, and PPI and CPI equilibrium rates.

income, i_t and π_t are the interest and expected inflation rates, respectively. The last term ω is the inter-country price of non-tradable goods relative to tradable goods. As in the previous formulations, Eq. (4) can also be construed as a long-run relationship.

In the standard monetary model, 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 sticky (Dornbusch, 1976) and there is secular inflation (Frankel, 1979),

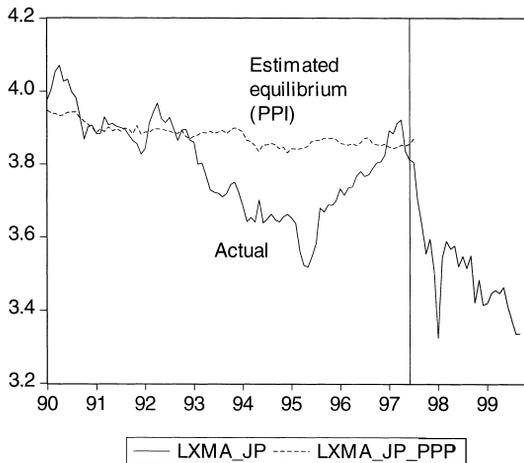


Fig. 4. Malaysian rupiah/Japanese yen exchange rate and PPI equilibrium rate.

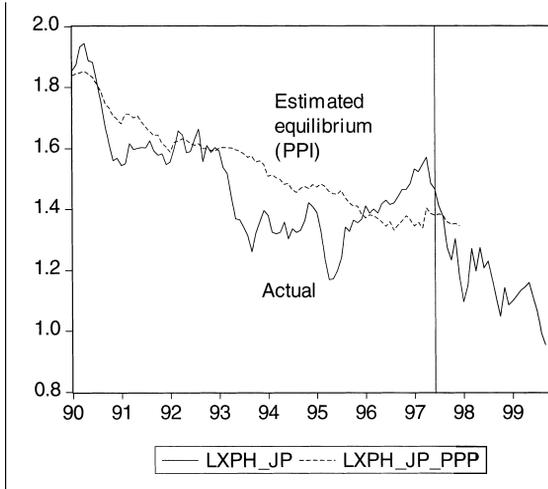


Fig. 5. Philippine peso/Japanese yen exchange rate and PPI equilibrium rate.

then $\beta_4 < 0$ and $\beta_5 > 0$, while $\beta_4 > 0$ and $\beta_5 = 0$ if prices are perfectly flexible (Frenkel, 1976). In most treatments of the monetary approach, long-run PPP is assumed to hold economy-wide, and thus $\beta_6 = 0$.

3.2. Modifications to account for 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

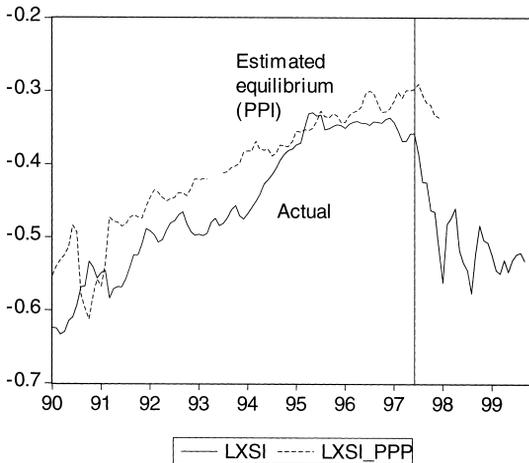


Fig. 6. Singapore dollar/US dollar exchange rate and PPI equilibrium.

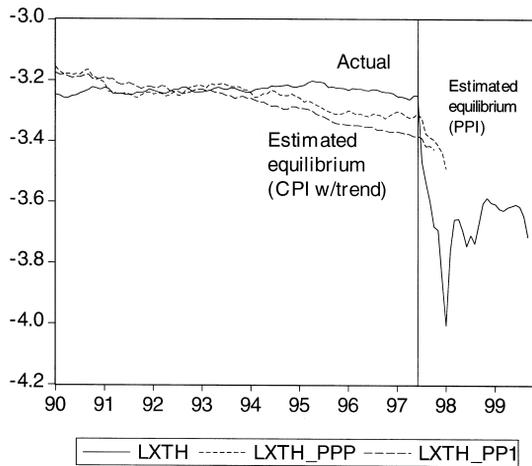


Fig. 7. Thai baht/US dollar exchange rate and PPI and CPI equilibrium rates.

East Asian newly industrializing countries. As is well documented, some of these countries are only now removing restrictions on the capital account, and indeed Korea and Taiwan are still in the process of liberalizing its external accounts (Chinn and Maloney, 1998). Hence, neither covered nor uncovered interest parity is likely to hold. On the other hand, even if there are parity deviations, they may not be sustainable in the long-run, so the model's predictions are still of some interest.¹¹

Another issue pertains to the stability of the money demand function imbedded in Eq. (4). In emerging market economies subject to monetization, increasing financial intermediation, or financial repression, money demand functions maybe time-varying. Tseng and Corker (1991) find stable cointegrating relationships hold for Indonesia, Korea, Malaysia, Singapore and Thailand. Using the more powerful Johansen (1988) methodology, Dekle and Pradhan (1999) update these results for several Southeast Asian countries and conclude that, with the exception of Indonesia, there is no evidence of real money demand cointegration.¹² In the Indonesian case, they identify a cointegrating relationship in money demand only after allowing for structural shifts.

Perhaps the most important issue pertains to the relevance of PPP for *broad* price indices. The results from Section 2 should suggest the questionable value of this assumption. Because this assumption is so grossly violated empirically for certain East Asian currencies (Isard and Symansky, 1996; Chinn, 2000), it is necessary to allow the long-run real exchange rate to vary over time.

¹¹Time invariant risk premia will be subsumed into the constant of the cointegrating vector.

¹²Dekle and Pradhan (1999) do find that cointegrating relationships hold for nominal money supplies. Furthermore, in the cases of narrow Malaysian, and narrow and broad Thai money, the restriction of homogeneity in price levels cannot be rejected.

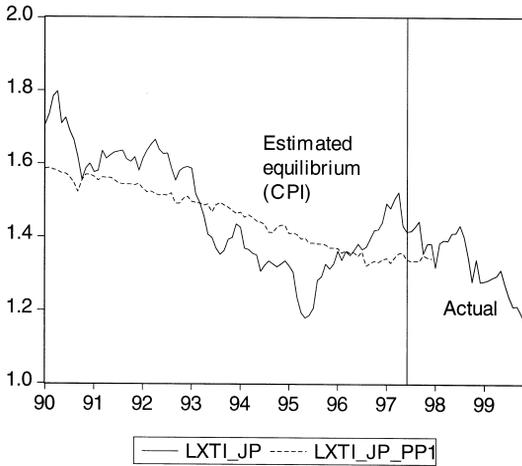


Fig. 8. New Taiwan dollar/Japanese yen exchange rate and CPI equilibrium rate.

Let the log aggregate price index be given as a weighted average of log price indices of traded (T) and on-traded (N) goods:

$$p_t = (1 - \alpha)p_t^T + \alpha p_t^N$$

where α is the share of non-traded goods in the price index. Suppose further that the foreign country's aggregate price index is similarly constructed. Rearranging, and allowing for sticky prices and long-run PPP only for tradables prices yields Eq. (4) where $\beta_6 > 0$.

The relative price variable ω 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 non-tradable sectors. Relative prices may also be affected by demand side factors. In the long-run, the rising preference for services, which are largely non-tradable, may induce a secular trend in the relative price of non-tradables.

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 with respect to the exchange rate. Unfortunately, sectoral productivity data is not available at a quarterly frequency for many of the countries being investigated. Hence, these Balassa-Samuelson and demand side effects are proxied with a relative price variable

3.3. Methodology, data and empirical results¹³

The statistical analysis is conducted on quarterly data over the 1982–1996 period

¹³This section is based on Chinn (1999b).

for Indonesia, Korea, Singapore, Thailand and Taiwan (the 1997–1998 period is reserved for out of sample calculations.) The other East Asian countries are omitted because of data limitations. Exchange rates are end-of-period, in US\$/local currency unit. Money is either narrow money or broad money in the case of Taiwan. Income is GDP in real currency units. Interest rates are interbank rates. Inflation rates are calculated as the annual change in the log of the price level, as measured by the CPI. The relative price variable is calculated as the log ratio of PPI to CPI.

The relative price variable is calculated as follows:

$$\alpha \left[(p^T - p^N)(p^{T^*} - p^{N^*}) \right] \approx \omega \equiv \log \left[\frac{\text{PPI}^{\text{US}} / \text{CPI}^{\text{US}}}{\text{PPI}^* / \text{CPI}^*} \right] \quad (5)$$

The implied long-run cointegrating relationship, in terms of observable variables, is Eq. (4) with $\beta_6 = 1$ for $\alpha = 0.50$. A similar specification incorporating a relative price variable is used in Chinn and Meese (1995). The Indonesian equation replaces the relative price variable with the real price of oil, which serves as proxy for the terms of trade.

The Johansen (1988) methodology cited earlier is used to test for the presence of cointegrating relationships between exchange rates, money stocks, incomes, interest and inflation rates and relative prices.

3.4. Model fit and estimated misalignments

The cointegration results are reported in Table 4. First consider the currencies of 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. 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 (1993a) 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 the cases of Korea and Taiwan, there is evidence of at least one cointegrating vector, while for Singapore, the evidence is much weaker.

The long-run relationship for the won exchange rate (Column 2) fits the augmented monetary model well. The coefficients on narrow money and relative income are not significantly different from that implied by theory. The interest differential enters in with a negative (although insignificant) sign, which is consistent with a sticky price model of the exchange. Inflation enters in with a positive sign. Finally, the relative price variable enters with the appropriate sign, and significantly so.

For Taiwan, it is not possible to fit a model using either narrow or broad money. Rather, the only specification that fits, with the expected signs, is one where US

Table 4
Long-run parameters of the monetary model^a

Coeff	IN	KO	SI ^b	TH ^b	TI ^c
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]
CRs	3[1]	4[0]	3[0]	3[0]	4[1]
m	0.535*** (0.066)	1.352*** (0.480)	0.908 (0.574)	1.654** (0.730)	0.729*** (0.094)
m^*	-0.535*** (0.066)	-1.352*** (0.480)	-0.908 (0.574)	-1.654** (0.730)	-1.023*** (0.068)
$y - y^*$	-0.546*** (0.120)	-1.056* (0.593)	-2.212** (0.928)	-3.008*** (1.303)	-2.171** (0.869)
$i - i^*$	-0.343* (0.186)	-2.121 (1.412)	-11.921** (5.005)	0.010 (0.192)	-1.417*** (0.405)
$\pi - \pi^*$	0.170 (0.160)	5.562* (2.938)	11.368*** (4.292)		
ω		3.655*** (1.225)	-0.133 (0.198)	2.038** (1.015)	1.361*** (0.420)
p^{oil}	0.653*** (0.026)				
$k + 1$	5	4	5	5	5
N	60	64	48	45	59
Smpl	82.1–96.4	81.1–96.4	85.1–96.4	85.4–96.4	82.2–96.4
Dummies	1983.2 1988.3 (1986.1 only)			1984.4 1989.1	

^aNotes: 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]; CRs 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 1986.1 dummy which takes on a value of 1 only in that quarter). Source: Chinn (1999b)

^bBroad money.

^cUS broad money, Taiwanese quasi-money.

broad money, and Taiwanese quasi-money, enter in separately. The results of estimating this specification are reported in Column 5.¹⁴ The US money coefficient

has the expected positive value of 0.719, and the Taiwanese quasi-money coefficient, of -1.023 . Relative income, interest rates and the relative non-tradables price coefficients are also all correctly signed and statistically significant.

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), while nominal interest rates and inflation rates exhibit statistical significance.

As for Thailand and Indonesia, there is evidence of cointegration for the latter, but mixed evidence for the former. Assuming one cointegrating vector for the baht/dollar exchange rate relation, one obtains plausible coefficients. The coefficient on relative broad money is 1.654, is statistically significant and 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 case of Indonesia (Column 1) one finds evidence of a single cointegrating vector.¹⁵ 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. Note that 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 (2000) regarding the effect of the real price of oil on the real exchange rate.

The models reported in Table 4 are used to generate equilibrium exchange rates. The base year effects are estimated using the sample up to 1996.4 (consistent with the estimation sample). Table 5 reports the implied deviations from equilibrium for all of 1997. As of 1997.2, the rupiah and baht are overvalued, with the baht substantially overvalued by 17%. However, both currencies are apparently becoming more overvalued in the quarters leading up to the crisis. The won, and Singapore and New Taiwan dollars are all undervalued (in ascending order of absolute magnitudes). The 1997.3 deviation is actually more appropriate for examination of the won's behavior, as the Bank of Korea did not give up on its defense until October. At this juncture, the undervaluation has increased to 13%; this is because interest rates had risen, which implies a stronger currency in the long-run. That the won continued to weaken attests to the deviation from the long-run relationship in the quarters preceding the crisis.¹⁶

¹⁴A dummy variable to account for the shift in money demand in 1984.4 (Kuo, 1990), as well as a dummy variable to account for a shift in capital account openness in 1989.1 (Chinn and Maloney, 1998) is included.

¹⁵In order to account for the money demand shifts identified by Dekle and Pradhan (1999), the regressions are augmented by two dummies, one for 1983.2 and 1988.3, and a dummy for 1986.1 to account for a spike in interest rates.

Table 5

Deviations from equilibrium as predicted by monetary model (misalignment = $s_t - \hat{s}_t$)^a

	IN	KO	SI	TH	TI
1997:1	0.013	-0.068	-0.168	0.093	-0.098
1997:2	0.077	-0.033	-0.084	0.172	-0.112
1997:3	-0.441	-0.130	-0.118	-0.039	-0.142
1997:4	-0.661	-0.901	-0.872	-0.191	-0.231

^aNotes: $s = \log(S)$, where S is measured in US\$/local currency unit; misalignment is the prediction error long-run relation. A positive (negative) value indicates an overvaluation (undervaluation) of the local currency. Out of sample prediction error from long-run relationship estimated over the 1974.1–1996.4 period. Source: Author's calculations based on results reported in Chinn (1999b).

4. Conclusions

This paper has documented the findings of mean reversion for several exchange rates over the 1975–1996 period; these results can be interpreted as detection of purchasing power parity. It is important to observe that these findings are often currency- and deflator-specific. As long as one is willing to entertain the PPP criterion as a measure of equilibrium exchange rates, one finds that there is some evidence of overvaluation on the eve of the 1997 currency and financial crises. There is little disagreement between valid indicators (where validity is judged on the basis of the cointegration tests), excepting the rupiah and NT\$.

Evidence is also presented that monetary fundamentals affect exchange rates over the long-run, for specific currencies. For certain countries, monetary models were either inappropriate (Hong Kong) or not estimated because of missing data (Malaysia) or short sample periods (Philippines). In the cases for which data were available, only Singapore presented uncertain evidence regarding cointegration.

Taken together, the various models yielded the estimated misalignments at mid-1997 summarized in Fig. 9 (PPIDEV, CPIDEV, and MONDEV are the PPI, CPI and monetary model implied deviations, respectively). The different approaches agree that the Singapore and Taiwan dollars and the won were undervalued, while the HK\$, peso and ringgit were overvalued. As for the rupiah, there is no agreement as to the degree of misalignment.

Clearly, the concept of overvaluation has some empirical content. The Singapore and New Taiwan dollars were both undervalued, and both suffered only modest declines in value. The peso, ringgit and baht were overvalued, and did experience crashes. Unfortunately for the overvaluation-cum-crisis hypothesis, the undervalued won also crashed, while the overvalued Hong Kong dollar did not.

Thus, overvaluation is an important factor in economic crises only in certain instances. In some events, overvaluation may be dominated by other factors such as large government liabilities in the form of implicit guarantees to bail out insolvent banking systems. Hence, to the degree that these episodes constitute financial —

¹⁶See Husted and MacDonald (2000) for a panel perspective on this question.

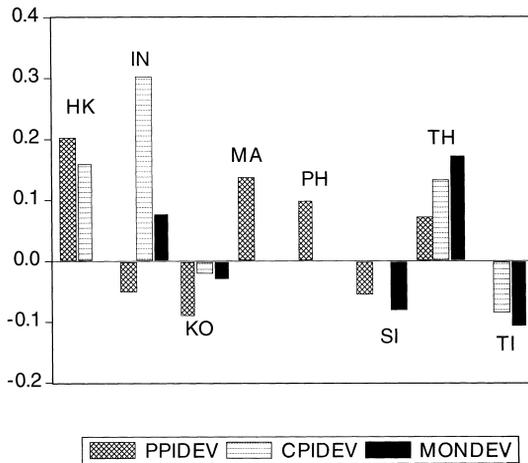


Fig. 9. PPP misalignment (PPI and CPI) and monetary model misalignment measures.

rather than currency — crises, overvaluation may not consistently presage difficult times.¹⁷

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Appendix A

IFS denotes IMF, *International Financial Statistics*, November 1997 and November 1999 CD-ROMs, updated using the IMF's Economic Data Sharing System (EDSS) in April 1998. *FS* denotes Bank of China, *Financial Statistics*, various issues, as recorded in Federal Reserve Bank of San Francisco electronic database.

¹⁷See Corsetti et al. (1998), Chinn et al. (1999) and Chinn and Kletzer (forthcoming).

A.1. Monthly data

- Exchange rates, *IFS* line *ae*, in US\$/local currency unit, end of period.
- Consumer price index, *IFS* line 64, 1990 = 100. Hong Kong CPI data is seasonally adjusted, and obtained from the EDSS.
- Producer price index, *IFS* line 63, 1990 = 100. Indonesian data excludes petroleum prices. Hong Kong data is quarterly, starting from 1991.1 Source: Hong Kong Department of Census and Statistics, personal communication from Winnie Tam.
- Export price index, *IFS* line 74, 1990 = 100.
- Trade-weighted real exchange rates (CPI-deflated). 1990 = 100, 1988–1990 trade weights. Source: IMF Information Notice System.
- Broad trade-weighted real exchange rates (PPI-deflated). 1990 = 100, 1990 trade weights for 1987–1997; 1980 trade weights for 1970–1986. Hong Kong series adjusted by Hong Kong retail price index. Source: Morgan Guaranty, <http://www.jpmorgan.com>.

A.2. Quarterly data

- Exchange rates, *IFS* line *ae*, in US\$/local currency unit, end of period.
- Narrow money, *IFS* line 34, in national currency unit.
- Broad money is narrow money plus quasi-money *IFS* line 35, in national currency units.
- Income is real GDP, *IFS* line 99*b.r*, in 1990 national currency units for the US and Korea. Malaysian income is proxied by industrial production. Taiwanese GDP is from *FS*, in 1991 New Taiwan dollars. Singapore income data is proxied by industrial production *IFS* line 66*ey* until 1983.4, and real GDP thereafter. Indonesian GDP data is unpublished data obtained from the IMF. Thai data is interpolated using an annual relationship between output, imports, exports, and the real exchange rate, and quarterly data up to 1991; thereafter is actual quarterly GDP data, obtained from the Bank of Thailand website.
- Interest rates are short term, interbank interest rates, *IFS* line 60*b*, in decimal form.
- CPI, *IFS* line 64, 1990 = 100.
- PPI, *IFS* line 63, 1990 = 100. Indonesian PPI data excludes petroleum prices (*IFS* line 63*a*).
- Inflation is 4-quarter difference of $\log(\text{CPI})$.
- Relative price variable: calculated as

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

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

A.3. Appendix Table 1. Johansen cointegration results

	HK	IN	JP	KO	MA	PH	SI	TH	TI
Panel A1.1: CPI									
k	1	1	1	1	1	1	1	1	1
#[#]	1[1]	0[0]	2[2]	2[2]	1[1]	1[1]	1[1]	1[1]	2[2]
c	-0.292	-5.942	-29.885	5.135	3.405	2.940	13.148	2.337	8.033
β_1	1	1	1	1	1	1	1	1	1
β_2	-0.828 (0.408)	1.535 (4.038)	-7.930 (3.296)	0.940 (1.337)	0.829*** (0.347)	-0.572 (0.762)	0.257 (1.236)	1.130*** (0.305)	-1.886* (0.531)
β_3	0.455*** (0.178)	1.482 (2.154)	15.505** (7.278)	-0.633* (0.902)	-1.363*** (0.539)	0.560 (0.377)	-3.019 (2.645)	-0.937*** (0.287)	0.847 (0.707)
LnLik	959.8	1538.2	1676.7	1900.9	1957.5	1640.0	1928.4	1916.1	1636.8
Smpl	80.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12
N	203	264	264	264	264	264	264	264	264
Panel A1.2: PPI									
k	–	1	1	1	1	1	1	1	1
#[#]	–	0[0]	0[0]	1[1]	0[0]	2[2]	1[1]	1[1]	1[0]
c	–	21.862	3.335	6.998	-0.269	4.669	1.519	2.124	6.116
β_1	–	1	1	1	1	1	1	1	1
β_2	–	-5.131 (5.335)	-1.924*** (0.130)	-0.996 (0.444)	1.684*** (0.838)	-1.299*** (0.112)	1.071 (0.126)	0.382*** (0.293)	-1.651* (0.184)
β_3	–	1.908 (1.070)	2.253*** (0.276)	0.870* (0.268)	-1.405*** (0.645)	0.964 (0.031)	0.856 (0.163)	-0.148*** (0.233)	1.029 (0.250)
LnLik	–	1154.2	1597.8	1664.2	956.9	1256.2	1499.3	1632.0	1585.5
Smpl	–	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12
N	–	264	264	264	264	264	264	264	264
Panel A1.3: Export price indices									
k	1	1	1	7	1	1	3	1	–
#[#]	2[2]	1[1]	1[1]	0[0]	1[1]	1[1]	1[1]	1[1]	–
c	7.231	-2.029	-2.188	-8.770	1.321	1.761	12.104	2.634	–
β_1	1	1	1	1	1	1	1	1	–
β_2	-8.153 (14.198)	2.866*** (0.104)	-1.468*** (0.130)	2.562** (1.630)	0.132*** (0.048)	-0.835 (0.259)	-1.998*** (0.279)	0.387** (0.349)	–
β_3	7.138 (11.560)	-0.804 (0.078)	2.042 (0.228)	0.749 (0.684)	-0.214 (0.040)	1.127 (0.097)	-0.525 (0.238)	0.507 (0.227)	–
LnLik	1470.1	501.2	1072.2	683.6	868.0	595.1	1128.7	1132.2	–
Smpl	76.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	75.01– 96.12	–
N	256	237	264	76	207	204	212	249	–

Notes: k is lag in VECM specification. #[#] is the number of cointegrating vectors according to a likelihood ratio test on the maximal eigenvalue statistic,

using asymptotic [finite sample] critical values. Finite sample critical values from Cheung and Lai (1993a,b). β_i are cointegrating vector coefficients. Asterisks denotes significance at the 10%, **5% and ***1% MSL for the null hypothesis of $\beta_2 = -1$ or $\beta_3 = 1$. LnLik is the log likelihood statistic, Smpl is sample, N is number of observations.

References

- Baharumshah, A.Z., Ariff, M., 1997. Purchasing power parity in South East Asian countries economies: a cointegration approach. *Asian Econ. J.* 11, 141–154.
- Bahmani-Oskooee, M., 1993. Purchasing power parity based on effective exchange rates and cointegration: 25 LDCs experience with its absolute formulation. *World Dev.* 21, 1023–1031.
- Balassa, B., 1964. The purchasing power parity doctrine: a reappraisal. *J. Polit. Econ.* 72, 584–596.
- Bayoumi, T., Clark, P., Symansky, S., Taylor, M., 1994. The robustness of equilibrium exchange rate calculations to alternative assumptions and methodologies. In: Williamson, J. (Ed.), *Estimating Equilibrium Exchange Rates*. Institute for International Economics, Washington, DC, pp. 19–60.
- Berg, A., Borensztein, E., Milesi-Ferretti, G.M., Pattillo, C., 2000. Anticipating balance of payments crises — the role of early warning systems. IMF Occasional Paper 186. IMF, Washington, DC.
- Cheung, Y.-W., Lai, K.S., 1993a. Finite-sample sizes of Johansen's likelihood ratio tests for cointegration. *Oxf. Bull. Econ. Stat.* 55, 313–328.
- Cheung, Y.-W., Lai, K.S., 1993b. Long-run purchasing power parity during the recent float. *J. Int. Econ.* 34, 181–192.
- Chinn, M.D., 1999a. Measuring misalignment: PPP and East Asian currencies in the 1990s. IMF Working Paper WP/99/120. International Monetary Fund, Washington, DC.
- Chinn, M.D., 1999b. On the Won and other East Asian currencies. *Int. J. Finance Econ.* 4, 113–127.
- Chinn, M.D., 2000. The usual suspects: productivity and demand shocks and Asia-Pacific real exchange rates. *Rev. Int. Econ.* 8, 20–43.
- Chinn, M.D., Dooley, M.P., Shrestha, S., 1999. Latin America and East Asia in the context of an insurance model of currency crises. *J. Int. Money Finance* 18, 659–681.
- Chinn, M.D., Kletzer, K. Imperfect information, domestic regulation and financial crises. In: Glick, R., Moreno, R., Spiegel, M., (Eds.), *Financial Crises in Emerging Markets*. Cambridge University Press, Cambridge (forthcoming).
- Chinn, M.D., Maloney, W.F., 1998. Financial and capital account liberalization in the Pacific Basin: Korea and Taiwan. *Int. Econ. J.* 12, 1–22.
- Chinn, M.D., Meese, R.A., 1995. Banking on currency forecasts: is change in money predictable? *J. Int. Econ.* 38, 161–178.
- Chou, W.L., Shih, Y.C., 1995. Long-run real exchange rates in the Four Little Dragons. *J. Int. Trade Dev.* 4, 184–202.
- Corsetti, G., Pesenti, P., Roubini, N., 1998. Paper Tigers? A preliminary assessment of the Asian crisis. NBER Working Paper #6783. NBER, Cambridge.
- Dekle, R., Pradhan, M., 1999. Financial liberalization and money demand in the ASEAN countries. *Int. J. Finance Econ.* 4, 205–215.
- Dornbusch, R., 1976. Expectations and exchange rate dynamics. *J. Polit. Econ.* 84, 1161–1176.
- Driver, R., Wren-Lewis, S., 1999. FEERs: a sensitivity analysis. In: MacDonald, R.R., Stein, J. (Eds.), *Equilibrium Exchange Rates*. Kluwer Academic Publishers, Boston.
- Edison, H., Gagnon, J., Melick, W., 1997. Understanding the empirical literature on purchasing power parity: the post-Bretton Woods era. *J. Int. Money Finance* 16, 1–17.
- Frankel, J.A., 1979. On the Mark: a theory of floating exchange rates based on real interest differentials. *Am. Econ. Rev.* 69, 610–622.

- Frankel, J.A., Rose, A.K., 1996a. Currency crashes in emerging markets: an empirical treatment. *J. Int. Econ.* 41, 351–368.
- Frankel, J.A., Rose, A.K., 1996b. A panel project on purchasing power parity: mean reversion within and between countries. *J. Int. Econ.* 40, 209–224.
- Frenkel, J.A., 1976. A monetary approach to the exchange rate: doctrinal aspects and empirical evidence. *Scand. J. Econ.* 78, 200–224.
- Froot, K.A., Rogoff, K., 1995. Perspectives on PPP and long-run real exchange rates. In: Grossman, G., Rogoff, K. (Eds.), *Handbook of International Economics*, vol. 3. Elsevier Press, Amsterdam, pp. 1648–1684.
- Fujii, E., 2000. Exchange rate and price adjustment in the aftermath of the Asian crisis, mimeo. Otaru University, Hokkaido, Japan.
- Fukuda, S., Kano, T., 1997. International price linkage within a region: the case of East Asia. *J. Jpn. Int. Econ.* 11, 643–666.
- Furman, J., Stiglitz, J., 1998. Economic crises: evidence and insights from East Asia. *Brookings Pap. Econ. Activity* 1988 (2), 1–136.
- Goldfajn, I., Valdes, R., 1999. The aftermath of appreciations. *Q. J. Econ.* 114, 229–262.
- Hinkle, L.E., Montiel, P.J., 1999. *Exchange Rate Misalignment*. Oxford University Press/World Bank, New York.
- Horvath, M.T.K., Watson, M.W., 1995. Testing for cointegration when some of the cointegrating vectors are prespecified. *Econometric Theory* 11, 984–1014.
- Husted, S., MacDonald, R.R., 2000. The Asian currency crash: were badly driven fundamentals to blame? *J. Asian Econ.* (forthcoming).
- Isard, P., Symansky, S., 1996. Long run movements in real exchange rates. In: Ito, T., Isard, P., Symansky, S., Bayoumi, T. (Eds.), *Exchange rate movements and their impact on trade and investment in the APEC region*. Occasional Paper 145. International Monetary Fund, Washington, DC.
- Ito, T., Ogawa, E., Sasaki, Y.N., 1998. How did the dollar peg fail in Asia? NBER Working Paper No. 6729. National Bureau of Economic Research, Cambridge.
- Johansen, S., 1988. Statistical analysis of cointegrating vectors. *J. Econ. Dyn. Control* 12, 231–254.
- Johansen, S., Juselius, K., 1990. Maximum likelihood estimation and inference on cointegration — with applications to the demand for money. *Oxf. Bull. Econ. Stat.* 52, 169–210.
- Kaminsky, G.L., Reinhart, C.M., 1999. The twin crises: the causes of banking and balance-of-payments problems. *Am. Econ. Rev.* 89, 473–500.
- Kremers, J., Ericsson, N., Dolado, J., 1992. The power of cointegration tests. *Oxf. Bull. Econ. Stat.* 54, 325–348.
- Kuo, S.W.Y., 1990. Liberalization of the financial market in Taiwan in the 1980s. *Pacific-Basin Capital Markets Res.* 1, 7–26.
- Lee, D.Y., 1999. Purchasing power parity and dynamic error correction: Evidence from Asia Pacific economies. *Int. Rev. Econ. Finance* 8.
- MacDonald, R.R., 1996. Panel unit root tests and real exchange rates. *Econ. Lett.* 50, 7–11.
- Milesi-Ferretti, G.M., Razin, A., 1996. Current account sustainability: selected East Asian and Latin American experiences. NBER Working Paper #5791. National Bureau of Economic Research, Cambridge.
- Moreno, R., 1988. Exchange rates and monetary policy in Singapore and Hong Kong. In: Cheng, H.-S. (Ed.), *Monetary Policy in Pacific Basin Countries*. Kluwer, Boston, pp. 173–200.
- Phylaktis, K., Kassimatis, Y., 1994. Does the real exchange rate follow a random walk? The Pacific Basin perspective. *J. Int. Money Finance* 13, 476–495.
- Sachs, J., Tornell, A., Velasco, A., 1996. Financial crises in emerging markets: the lessons from 1995. *Brookings Pap. Econ. Activity* 1996 (1), 147–215.
- Samuelson, P., 1964. Theoretical notes on trade problems. *Rev. Econ. Stat.* 46, 145–154.
- Tang, M., Butiong, R.Q., 1994. Purchasing power parity in Asian developing countries: a cointegration test. *Statistical Report Series #17*. Asian Development Bank, Manila.

- Tseng, W., Corker, R., 1991. Financial liberalization, money demand, and monetary policy in Asian countries. Occasional Paper 84. International Monetary Fund, Washington, DC.
- Williamson, J. (Ed.), 1994. Estimating equilibrium exchange rates. Institute for International Economics, Washington, DC.
- World Financial Markets, 1993. Morgan Guaranty, New York, 19 November.
- Zanello, A., Desruelle, D., 1997. A primer on the IMF's information notices system. Working Paper WP97/71. International Monetary Fund, Washington, DC.