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THE USUAL SUSPECTS? PRODUCTIVITY AND DEMAND SHOCKS AND ASIA-PACIFIC REAL EXCHANGE RATES

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ABSTRACT

The evidence for a productivity-based explanation for real exchange rate behavior of East Asian currencies is examined. Using sectoral output and employment data, relative prices and relative productivities are calculated for China, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan and Thailand. Time series regressions of the real exchange rate on relative prices indicate a role for relative prices for Indonesia, Japan and Korea. When examining real exchange rates and relative productivity ratios, one finds a relationship for Japan, Malaysia, and the Philippines. Only when augmenting the regressions with real oil prices are significant relationships obtained for Indonesia and Korea. Panel regression results are slightly more supportive of a relative price view of real exchange rates. However, the panel regressions incorporating productivity variables, as well as other demand side factors, are less encouraging, except for a small subset of countries (Indonesia, Japan, Korea, Malaysia and the Philippines). Surprisingly, government spending does not appear to be a determinant of real exchange rates in the region.

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1 Introduction

The study of long run real exchange rate determination has enjoyed a renaissance in recent years. With the development of detailed sectoral databases, economists have been able to analyze the empirical foundations for productivity based models of the real exchange rate, such as those of Harrod (1933), Balassa (1964) and Samuelson (1964). These models imply that real exchange rate appreciation should be correlated with differentials in tradable/nontradable-sector productivity growth.[1]

For the developed countries, the resurgence has been sparked by the publication of the OECD's *International Sectoral Data Base* (*ISDB*) which provides sectoral total factor productivity data.[2] However, with the exception of Japan, one would not expect the Harrod-Balassa-Samuelson (hereafter HBS) effect to be very apparent.[3]

In this paper, I examine whether one can verify the relevance of the HBS effect in East Asian countries by rounding up "the usual suspects": sectoral productivity levels, government spending ratios, and real oil prices. The East Asian economies are exactly the type for which Balassa posited the relevance of the HBS effect: economies characterized by rapid growth, presumably due to rapid manufacturing (and hence traded) sector productivity growth. In the past, researchers have been constrained by the unavailability of sectoral productivity data, and either restricted themselves to examining only the first portion of the of the HBS hypothesis -- that relative prices of traded to nontraded goods should determine in large part real exchange rates (*e.g.*, Isard and Symansky, 1996 for the APEC countries) -- or have assumed that overall productivity growth is highly correlated with the

[1] See Asea and Corden (1994) for a review.

[2] See for instance DeGregorio and Wolf (1994), Strauss (1995, 1996), and Canzoneri, Cumby and Diba (1996).

[3] In fact, it is difficult to detect evidence in support of purely supply-side effects on the real exchange rate in time series data. Rather, one finds that demand side shocks, such as government spending, must be included if one is to find a role for traded/nontraded relative productivity levels (see DeGregorio and Wolf (1994) and Chinn and Johnston (1996) in a panel context).

tradable/nontradable productivity differential (*e.g.*, Bahmani-Oskooee and Rhee, 1996 for Korea). Hence a modest contribution of this paper is to use data more consistent with the underlying economic hypotheses.

This paper proceeds in the following fashion. In Section 2, the literature on real exchange rates is reviewed. The data and the construction of the key variables are described in Section 3. In section 4 the relationship between the real exchange rate and the relative price of tradables and nontradables is re-examined. Section 5 investigates the role of sectoral productivity differentials in the determination of real exchange rates. Section 6 continues with an application of panel cointegration methodologies to the two relationships. Section 7 concludes. To anticipate the results, I find that the relative price of tradables to nontradables explains some, but not all, of the variation in real exchange rates. Somewhat surprisingly, productivity ratios *do* appear to be related to the real exchange rate, but only in certain cases. Furthermore, the results do not appear to be much stronger when appeal is made to panel regression techniques.

2 Literature Review

The starting point for most investigations of the linkage between the relative price of nontradables and the real exchange rate relies upon the following construction. Let the log aggregate price index be given as a weighted average of log price indices of traded (T) and nontraded (N) goods:

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

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:

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

Then the real exchange rate is given by:

$$q_t \equiv (s_t + p_t^* - p_t) \tag{3}$$

where s is the log of the domestic currency price of foreign currency. For $\alpha = \alpha^*$, the following holds:

$$q_t = (s_t + p_t^T - p_t^{T*}) - \alpha[(p_t^N - p_t^T) - (p_t^{N*} - p_t^{T*})] \tag{4}$$

Although there are many alternative decompositions that can be undertaken, equation (4) is the most relevant since most economic models make reference to the second term as the determinant of the real exchange rate, while the first is assumed to be zero by purchasing power parity (PPP) as applied to traded goods. Engel (1995) examines the relative contributions of both of these components to the variability of the US real exchange rate against other G-7 currencies, and finds that the first component accounts for almost all of it, over many horizons, and for many price indices. Isard and Symansky (1996) undertake a slightly different decomposition, allowing the nontraded shares (α's) to change over time. This procedure yields a breakdown of the real exchange rate changes into three components: (i) changes in the relative price of traded goods; (ii) changes in the relative price of traded to nontraded goods; and changes in share weights (α's). They obtain the surprising result, consonant with Engel's, that the first term accounts for almost all of the movements in the real rate for China, Indonesia, Japan, Philippines, and Thailand. Chinn (1996) examines East Asian exchange rates in this accounting sense, and using cointegration techniques finds that there is some evidence that in the long run, relative prices of tradables and nontradables do explain real exchange rates, at least for certain currencies.

In order to move away from accounting identities one requires a model, such as the HBS framework. The relative prices of nontradables and tradables will be determined solely by productivity differentials, under the stringent conditions that capital is perfectly mobile internationally, and factors of production are free to move between sectors. Substituting out for relative prices yields:

$$q_t = (s_t - p_t^T + p_t^{T*}) - \alpha[(\frac{\theta^N}{\theta^T})a_t^T - a_t^N] + \alpha[(\frac{\theta^{N*}}{\theta^{T*}})a_t^{T*} - a_t^{N*}] \tag{5}$$

where a^i is total factor productivity in sector i, and the θs are coefficients from the sectoral production functions.

Most researchers have proceeded under the assumption that the first term is I(0). This implies cointegrating relationships of the form:

$$q_t = -\alpha(p_t^N - p_t^T) + \alpha(p_t^{N*} - p_t^{T*}) \tag{6}$$

and

$$q_t = -\alpha[a_t^T - a_t^N] + \alpha[a_t^{T*} - a_t^{N*}] \tag{7}$$

respectively (where the production functions in the tradable and nontradable sectors are assumed to be the same, so that the θs cancel out in equation (7)). Equations (6) and (7) will motivate the empirical work. [4]

3 Data and Variable Construction

Since the central contribution of this paper is the application of econometric techniques to disaggregate data, the details of the data and the construction of the empirical counterparts to the theoretical variables merit some discussion (additional details are in the Data Appendix). The countries examined include the China, Indonesia, Japan, Korea, Malaysia, the Philippines, Singapore, Thailand, Taiwan and the US, for the period 1970-1992. The data are drawn from a variety of sources. The nominal exchange rate is the bilateral period average, expressed against the US$ (in $/f.c.u.). Price levels are GDP deflators drawn from the World Bank's *World Tables*. The real exchange rates are then expressed in terms of GDP deflators, rather than the typical CPIs. However, the difference in the behavior of the variables is minor.

[4] Both equations have been exploited extensively. Equation (6) has been examined most recently by Kakkar and Ogaki (1994) for several exchange rates. Equation (7) has been estimated by Hsieh (1982), Marston (1990), and the references in footnote 2.

The prices of tradables and nontradables are proxied by the price deflator for manufacturing, and for "other", which includes mostly services, construction, and transportation, respectively. These data are calculated as the ratio of nominal to real (in 1987 domestic currency units) sectoral output, as reported in *World Tables*, except in certain cases, such as Malaysia and Taiwan, where the Asian Development Bank's *Key Indicators of Developing Asian and Pacific Countries* are used. Note that this ratio omits the agricultural sector. Hence, unlike the approach adopted in Engel (1996) and Chinn (1996), the overall price deflator is *not* a weighted average of the two sectoral deflators.

Since several of these countries have at certain times exported a large amount of agricultural products, one might wonder if the results reported below are sensitive to the equating of manufactures with tradables. I have also examined an alternative measure of the relative price of tradables and nontradables, where a weighted average of agricultural and manufacturing prices is used in place of the manufacturing deflator. None of the statistical results for cointegration involving relative prices hinge upon this alternative definition, although the time series pattern for Thailand, and the trends for China, Indonesia and Korea, do change.[5]

The exchange rate series and the respective relative price series are displayed in Figures 1 through 9; the average annual rate of change for these series are also reported Panel 1 of Table 1. To make ocular regressions easier to conduct, the exchange rates are expressed in *inverse* terms (foreign currency unit/ US$) and normalized to 1974 = 0 in these figures; hence the country pairs should covary positively according to the theory described above. With the exception of China and Taiwan, there does appear to be some comovement of the posited nature.

As remarked above, attempts to test the Harrod-Balassa-Samuelson model's predictions regarding East Asian productivity levels and currencies have been hampered by the lack of sectoral productivity data. Consequently, previous researchers have resorted to proxies, such as GDP *per capita*, which is clearly inadequate. Showing a correlation between the real

[5] Attempting an analogous procedure for calculating sectoral labor productivity is a much more difficult enterprise, as agriculture employs a large number of part-time workers.

exchange rate and income levels merely validates what Samuelson called "The Penn Effect," and cannot in itself demonstrate that productivity matters directly. Relative prices may depend upon demand side factors, such as a greater preference for nontraded goods as income rises. Even use of aggregate productivity does not circumvent the difficulty in isolating the HBS effect. For instance, Bahmani-Oskooee and Rhee (1996) estimate the following equation for Korea:

$$q_t = \kappa_0 + \kappa_1(a_t - a_t^*) + u_t \tag{8}$$

which implicitly assumes that economy-wide a and the difference of the sectoral a's are cointegrated. Taking the simplest case,

$$a_t = \omega a_t^T - \omega a_t^N \tag{9}$$

The α coefficients are not identified; rather only the product $\kappa = \omega\alpha$ can be estimated. A more complicated situation arises if the coefficients on the sectoral a's are not the same. Then an omitted variable situation arises, and cointegration is unlikely to be found, even if the model in (7) holds true.[6] It is therefore critical to find adequate proxies for the *sectoral* productivity levels.

I calculate average labor productivity as the proxy for sectoral total factor productivity; average labor productivity is obtained by dividing real output in sector i by labor employment in sector i:

$$A_i^L \equiv Y_i / N_i \tag{10}$$

where i takes on the same categories as before. The output series are drawn from the same sources used for the relative price deflators. The employment figures are drawn from the *World Tables*, the ADB *Key Indicators*, and for the developed countries, the International

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[6] In fact, if the aggregate price index is a Cobb-Douglas function of tradable and nontradable prices, then the equation using aggregate productivity can only represent a cointegrating relation if the cointegrating coefficients are allowed to time-vary nonlinearly with a^T/a^N.

Labour Office's *Yearbook of Labour Statistics*.

Two limitations of the data should be stressed. First, since these labor employment statistics are not adjusted for part time workers, one may very well have qualms about the reliability of these proxies for labor productivity. To cross- check the results, I have compared the calculations for manufacturing productivity against those reported by the *World Tables* for several countries. These figures match quite well. Furthermore, for the US and Japan, we have data on labor productivity in the traded (industrial output, transportation) and nontradable (services, construction) sectors from the OECD's *International Sectoral Data Base*.[7] These series also match quite well for manufacturing vs. tradables, and "other" versus nontradables. These results serve to improve one's confidence that the proxies we use are not wholly inadequate.

Second, the proxy variable is labor productivity, rather than total factor productivity as suggested by the model. Canzoneri, Cumby and Diba (1996) have argued that use of labor productivity is favored because it is less likely to be tainted by mis-estimates of the capital stock. In any event, there is little possibility of circumventing this problem. To my knowledge, almost all calculations of East Asian total factor productivity over long spans of time have been conducted on an economy-wide basis (see for a recent example Bosworth and Collins, 1997), with the exception of Alwyn Young's (1995) recent analysis.

The (inverse) exchange rate and relative productivity variables are displayed in Figures 10-18 (rescaled to a common base year equal zero); the average annual rates of change are reported in Panel 2 of Table 1. Once again, China and Taiwan fail to exhibit the expected correlation. In the other cases, the covariation is not very pronounced, although in a eclectic model of relative price determination, demand side factors such as government spending comes into play. Hence, it appears worthwhile to proceed with the statistical analysis.

The other variables are government consumption spending, in nominal terms. The variable used in the regressions is the log of the ratio of nominal government consumption to nominal GDP. It would be preferable to have real government consumption divided by real

[7] These categorizations are the same as those used in DeGregorio and Wolf (1994) and Chinn and Johnston (1996).

GDP as the government spending measure. In the absence of these data (and the likelihood that such government spending deflator data would be even more flawed than developed country equivalents), I use the nominal measure. Real oil prices are measured as the $ price per barrel of Dubai Fateh, deflated by the US CPI (in log terms).

Almost all the series appear to be integrated of order 1, according to ADF tests (with constant and trend) with lags of order 1 and 2. The exceptions are the US-China productivity differential, the US-Malaysia relative price ratio (although only 12 observations are available for this series), the Singapore real exchange rate (all at the 5% MSL) and the US-Taiwan relative productivity differential (at the 1% MSL). I will proceed under the presumption that all series are I(1), although the econometric results will be interpreted in light of these unit root test results.

4 The Real Exchange Rate and Relative Prices

4.1 Econometric Specification

If the relative price of traded goods is stationary, then the series q, $p^{N*}-p^{T*}$, $p^{N}-p^{T}$, are cointegrated with the cointegrating vector implied by theory of $(-1\ \alpha^{*}\ -\alpha)$. Best current practice is to use the multivariate maximum likelihood technique of Johansen (1988) and Johansen and Juselius (1990). Unfortunately, this procedure yields very dispersed estimates (for reasons suggested by Stock and Watson (1993)). Moreover, since there are very few observations, after adjusting the critical values for the loss of degrees of freedom, only a few instances of cointegration are found.[8]

Given the implausible estimates derived from the Johansen procedure, I implement a single equation error correction model, as suggested by Phillips and Loretan (1991). Typically, one has to worry about endogeneity in any single equation regression. Fortunately, in this model, it is plausible to assume a recursive structure such that $(p^{N}-p^{T})$ and $(p^{N*}-p^{T*})$ are weakly exogenous for q. Hence, in this paper I adopt a single-equation error correction

[8] There are at most 22 observations. If one adjusts the critical values for the loss of degrees of freedom due to estimating the parameters, then one cannot in general reject the null hypothesis of no cointegrating vectors. See Cheung and Lai (1993).

modeling approach, *viz*:

$$\Delta q_t = \mu + \sum_{i=1}^{k} \gamma_i \Delta q_{t-i} + \sum_{j=1}^{m} \delta_j \Delta \tilde{p}_{t-j} + \phi[q_{t-1} + \alpha \tilde{p}_{t-1}] + \nu_t \tag{11}$$
$$\tilde{p} = (p^N - p^T) - (p^{N*} - p^{T*})$$

This equation is estimated using nonlinear least squares (NLS), the asymptotic distribution of the *t*-statistic associated with the ϕ parameter is approximately Normal under the above assumptions.

4.2 Empirical Results

Regressions of the form of (11) were estimated for all real exchange rates. In general k=j=1, except where noted. The results in Table 2 indicate that the real exchange does exhibit some mean reversion. In all cases except China, Taiwan and Thailand, the coefficient on the lagged error correction term appears to be statistically significant. In these cases the estimated rate of reversion toward equilibrium is above 25%. For the exception of China, the statistical results are unsurprising, given the behavior of the real exchange rate over the sample period, and the clear violations of the assumptions of the Harrod-Balassa-Samuelson model -- full capital mobility, and free intersectoral factor mobility. For Taiwan and Thailand, the failure to find reversion is more puzzling.

The coefficient on relative nontradable/tradable prices is less precisely estimated. The three significant estimates appear in the plausible range of 0.28 to 0.76 (recall this coefficient has the interpretation of the share of total expenditures on nontradables). Even the nonsignificant estimates are plausible, with the exception of Singapore and Taiwan. In light of Young's (1995) findings regarding the close rates of productivity growth in manufacturing and service sectors in both these countries, it would be unsurprising to find a lack of a role for relative prices.[9]

Note that these specifications have omitted time trends. Standard econometric practice is to include deterministic time trends in such regressions; inclusion of such time trends in these regressions usually drastically reduces the statistical significance of the point estimates on the relative price variable. Note, however, that allowing for deterministic time trends implies the presence of a deterministic time trend in the cointegrating vector, such as[10]:

$$q_t = d_t - \alpha(p_t^N - p_t^T) + \alpha(p_t^{N*} - p_t^{T*}) + \epsilon_t \qquad (12)$$
$$d_t = \gamma + \xi t$$

Hence, any deterministic time trend must be in the relative price of tradables (when expressed in a common currency). Breuer (1994) points out that PPP for tradables implies not only trend stationarity, but also mean stationarity, so I rule out time trends.

In this interpretation, Korea evidences the most rapid rate of convergence to PPP for traded goods. The implied half life of a deviation is less than a year. On the other end of the spectrum, the implied half life for a Indonesian deviation is 2.3 years. Whether these are rapid rates depends upon one's perspective. Phylaktis and Kassimatis (1994) examined WPI deflated real exchange rates, which roughly corresponds to the manufacturing sector price deflated exchange rates used in this study. They found that the half life of deviations (estimated across eight currencies, allowing for a time trend) is about one year. If the deviation from equilibrium I obtain is attributable to deviations from PPP for tradable goods, then 2.3 years is a relatively long half life.

One must be forthright about the limited results obtained here. For China, Taiwan and Thailand, there is no statistically substantive evidence for reversion of any sort. For Malaysia, the Philippines and Singapore, there is some evidence that there is reversion, but the effect of relative prices is too imprecisely estimated to make any definite conclusions. Only for

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[9] An alternative interpretation is that the Taiwanese deflators are very mis-measured. Young (1995: 662-63) notes the unusual manner in which Taiwan accounts for public sector employee productivity. On statistical grounds, the failure to find cointegration is also not surprising since US-Taiwan relative productivity appears to be I(1) according to ADF tests.

[10] See Zivot (1996) for the conditions where this restriction is inappropriate.

Indonesia, Japan and Korea are the posited effects identified; and for Japan, one is already aware of the evidence in favor of the relative price effect (Strauss, 1996; Chinn and Johnston, 1996; Chinn, 1996).

5 The Real Exchange Rate, Productivity and Demand Shocks

5.1 Econometric Specification

Since the productivity data span even shorter periods than the relative price data, conservation of degrees of freedom is at a premium. In principle, the NLS procedure adopted in the previous section is feasible. However, estimation of the implied error correction models yielded highly implausible estimates. Kremers, Ericsson and Dolado (1992, hereafter KED) propose a test where the cointegrating vector is imposed *a priori*. The test for cointegration can then be applied quite simply by evaluating the t-statistic on the error correction term. As Zivot (1996) has pointed out, the distribution for this t-statistic depends upon a number of assumptions, most importantly on the validity of the imposed a priori cointegrating vector. Since in the absence for a role for demand side shocks,

$$p_t^N - p_t^T = a_t^T - a_t^N \tag{13}$$

I substitute relative productivity terms for relative prices, and impose the constraint $\alpha = 0.5$. This implies that the share of nontradables in the aggregate price index is one-half. Then I estimate the equation, allowing for the level of relative productivity to enter in separately. Zivot shows that if this variable enters in significantly, then the imposed cointegrating vector is invalid. For the instances in which this is true, I estimate the equation unconstraining the cointegrating vector, once again using NLS.

To summarize, the unconstrained specification is:

$$\begin{aligned} \Delta q_t &= \mu + \sum_{i=1}^{k} \gamma_i \Delta q_{t-i} + \sum_{j=1}^{m} \delta_j \Delta \tilde{a}_{t-j} \\ &\quad + \phi[q_{t-1} + \alpha \tilde{a}_{t-1}] + \nu_t \\ \tilde{a} &\equiv (a^T - a^N) - (a^{T*} - a^{N*}) \end{aligned} \tag{14}$$

and the constrained specification is:

$$\begin{aligned} \Delta q_t &= \mu + \sum_{i=1}^{k} \gamma_i \Delta q_{t-i} + \sum_{j=1}^{m} \delta_j \Delta \tilde{a}_{t-j} \\ &\quad + \phi[q_{t-1} + 0.5\tilde{a}_{t-1}] + \nu_t \end{aligned} \tag{15}$$

5.2 Empirical Results

Preliminary regressions indicated that the $\alpha = 0.5$ restriction was violated by the data in only two cases -- Singapore and Taiwan. Hence, regressions of the form of (14) were estimated for all real exchange rates except those two currencies, in which case equation (15) was implemented. In general k=j=1 (except where noted). The results in Table 3 indicate that the real exchange does appear to be cointegrated with relative traded/nontraded productivity levels for Japan, Malaysia and the Philippines.[11] Exchange rates for Singapore and Taiwan appear to be unexplained by relative productivity differentials, as indicated by the zero reversion and wildly implausible coefficient estimate on productivity in the former, and the positive estimate on productivity in the latter. These results for Taiwan contrast strongly with those obtained by Wu (1996) who finds that the real exchange rate is cointegrated with relative productivities *and* relative unit labor costs, expressed in a common currency. However, the point estimates on the relative productivity variables imply that over 100% of the output in the US and Taiwan are nontradables.

[11] In the case of the Philippines, I report the estimates using the World Bank measure of manufacturing productivity because of uncertainty about the accuracy of the calculated productivity measure (see the Appendix). Using the measure used in the other regressions *raises* the estimated rate of reversion and the goodness of fit statistic. Hence I have been conservative in my approach.

Notice that the formulation embodied in these regressions assumes that supply shocks, in the form of productivity growth, are the only determinants of the relative price of tradables to nontradables. Dropping that assumption, I allow for government spending to affect the contemporaneous real exchange rate. If government spending falls mostly on nontradables, then a change in government spending should induce a current appreciation in the real exchange rate. The results of this re-specification are reported in Table 4. Inclusion of this variable typically reduces the significance level for the coefficient on the error correction term. Only Japan's coefficient remains statistically significant. At the same time, the government spending variable often enters in the regression with unexpected sign. For Malaysia and Taiwan, changes in the local government spending to GDP ratio cause depreciation of the local currency. This is consistent with most government spending falling on tradable, rather than nontradable, goods. On the other hand, US government spending, when it is significant, does appreciate the US$ (this is the Korean case).

In previous productivity-based studies of the real exchange rate, real oil prices have entered significantly (Chinn and Johnston, 1996; DeGregorio and Wolf, 1994). In the Japanese case where all oil is imported, the channel could either work its way through the wealth effect, or via shifts in the production function. to investigate this effect, I augmented the regression specifications in Tables 3 and 4 with oil prices (either in levels or first differences). Real oil prices prove to be a significant factor in three cases -- Indonesia (a substantial oil exporter) and Japan and Korea (two oil importers). The results of reestimating the error correction models with oil prices included are reported in Table 5. Now all error correction terms are statistically significant, although sometimes only at the 20% MSL.

Log real oil prices enter into the cointegrating vectors for Indonesia and Japan, and into the short run dynamics for Korea. The direction of effects is consistent with priors: Increases in oil prices appreciates the Indonesian Rupiah against the US$, and depreciates the Yen (in both short and long run) and Won (in the short run).

It appears that the evidence, after accounting for the role of oil, is fairly strong for a productivity-based model of real exchange rates in the East Asian region. Indonesia, Japan, Korea, Malaysia and the Philippines appear to fit the general pattern. However, China, Singapore, Taiwan and Thailand appear recalcitrant.

6 Panel Evidence

Since the time series data are only partly informative regarding cointegration, a reasonable manner in which to proceed is to exploit the cross-section information. At first glance a cross-section approach has some appeal. In Figures 19 and 20, the data in Table 1 are plotted (here the real exchange rate is expressed in US$/foreign currency unit, so a negative correlation is implied). China has been excluded in both figures. In both cases, the negative relationship is apparent. For the exchange rate - relative price relationship, a simple OLS regression yields:

$$\begin{aligned} \Delta q &= \underset{(0.004)}{0.014} - \underset{(0.095)}{0.175}\times\Delta\tilde{p} + u_t \\ \overline{R}^2 &= .26 \quad N=8 \end{aligned}$$

For the exchange rate - relative productivity relationship, one obtains:

$$\begin{aligned} \Delta q &= \underset{(0.007)}{0.016} - \underset{(0.209)}{0.772}\times\Delta\tilde{a} + u_t \\ \overline{R}^2 &= .64 \quad N=8 \end{aligned}$$

Hence I estimate a panel error correction models corresponding to those estimated in the previous sections. The econometric basis for undertaking such regressions is discussed in Chinn and Johnston (1996).

6.1 The Relative Price of Tradables/Nontradables

In Table 6, the relationship between the real exchange rate and the relative price of tradables versus nontradables is investigated. Rather surprisingly, In a specification that includes time and currency dummies, but constrains the slope coefficients to be the same over currencies, the estimated rate of reversion to equilibrium is -0.09 (statistically significant at the 1% MSL). This is somewhat slower than the rates estimated using time series data, and implies a rather long deviation half life of about 7 years; it is not clear if this result is due to misspecification. A Wald test on a common rate of reversion coefficient is not rejected;

however, this may be due to the poor fit of the model. The point estimate of -0.679 on the coefficient on the relative price variable is statistically significant at the 3% MSL.

Since the Chinese coefficient appears to be individually significant at the 20% MSL in an unconstrained specification, it makes sense to examine the results excluding the Chinese series. In column 2, the estimated rate of reversion is much more rapid -- 0.243 -- implying a deviation half life of about 2 1/2 years. The coefficient on the relative price variable is again plausible in magnitude, and.both key coefficients are highly significant.

6.2 Relative Productivity Levels

In the remaining columns of Table 6, I examine the role of productivity and other factors in a sample *excluding* China. If relative prices do not explain the evolution of this real exchange rate, it is unlikely that relative productivity levels will perform any better. Column 3 reports results for the basic specification. Although the estimated rate of reversion is somewhat slower than that indicated in the relative price specification, the point estimate of -0.153 is statistically significant. Unfortunately, the point estimate for the effect of sectoral productivity is not significant.

The influence of productivity may be obscured by the omission of relevant variables. In the next two columns, I augment the specification with changes in government spending, and with oil prices. Government spending does not enter with statistical significance in either specification. Real oil prices do not enter into the long run relation, although they certainly do in the short run, via contemporaneous changes. The absence of a role for oil prices is not surprising given the heterogeneity in the countries. Indonesia is an oil exporter, while Japan and Korea are oil importers. Still a restriction on a common slope coefficient on long run oil prices is not rejected by a Wald test; only the restriction on short run effects is rejected (at a high level of significance).

Since no time series support was obtained for the productivity based model for Singapore, Taiwan and Thailand, it seems inappropriate to pool these countries in with the others. In Table 7, I report the results restricting the sample even further. Results for the basic specification, that augmented by government spending changes, and by oil prices are reported.

In all cases, the rate of reversion is estimated quite precisely at about 17% per annum. Furthermore, the point estimates for the productivity coefficient are plausible, although only significant at the 20% MSL. In order to insure that the results are not being driven by inappropriate cross-currency restrictions, I have conducted a series of Wald tests. The only restriction clearly rejected by the data is that on the short-run oil price coefficient. Unconstraining this coefficient does not change the estimates on the other variables much, so one can conclude that there is some panel evidence in support of the augmented productivity based model.

7 Conclusions

In this paper, I have examined the evidence for a productivity-based explanation for long run movements in East Asian real exchange rates. I find, somewhat in contrast to Isard and Symansky (1996), that there is some evidence in favor of such explanations. However, the evidence is by no means conclusive, especially as it relates to some of the countries now growing the most rapidly -- China, and Thailand.

Furthermore, it must be admitted that the real exchange rate exhibits large swings away from the equilibrium rate as predicted by either the relative price of tradables to nontradables, or relative productivities. The most plausible explanation for this finding -- that relative prices of *traded* goods can exhibit substantial persistence -- suggests that closer examination of the changing composition of East Asian exports and imports over time is warranted.[12] Hence here, as elsewhere, the primary culprits may not be among the "usual suspects."

[12] I have in mind models of imperfect competition, and pricing to market. Ito, Isard and Symansky (1996) have linked machine export ratios to real exchange rate appreciation. Presumably the motivation is that machine goods are differentiated and less likely to be ruled by arbitrage.

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Data Appendix

Exchange rates: period average: line *rf*, IMF, *International Financial Statistics*, November 1996 CD-ROM; and for Taiwan, Bank of China, *Financial Statistics*, various issues.

Consumer price indices: line 64, IMF, *International Financial Statistics*, November 1996 CD-ROM; and for Taiwan, Bank of China, *Financial Statistics*, various issues.

Nominal GDP and real GDP: World Bank, *World Tables, 1995* CD-ROM; and Asian Development Bank, *Key Indicators of Developing Asian and Pacific Countries*, various issues; and Bank of China, *Financial Statistics*, various issues.

Sectoral output, at factor cost, in nominal and real terms: World Bank, *World Tables, 1995* CD-ROM; and for Malaysia, and Taiwan, Asian Development Bank, *Key Indicators of Developing Asian and Pacific Countries*, various issues.

Sectoral employment: Asian Development Bank, *Key Indicators of Developing Asian and Pacific Countries*, various issues; and International Labour Office, *Yearbook of Labour Statistics*, various issues.

Labor productivity: For all Asian LDCs, sectoral output/sectoral employment: "Tradable" = manufacturing; "Nontradable" = other. For US and Japan, indices from OECD *International Sectoral Data Base* (*ISDB*). "Tradable" = industry, mining, transportation, agriculture. "Nontradable" = services, construction, government.

Alternate manufacturing labor productivity measure: World Bank, *World Tables, 1995* CD-ROM.

Government spending. Government consumption, from World Bank, *World Tables 1995* CD-ROM. (G## is government consumption spending, divided by GDP).

Oil prices: IMF, *International Financial Statistics*, CD-ROM. (LRPOIL is log price of oil, *IFS* line 76aad, deflated by US CPI).

Real exchange rate: (Q##) Nominal exchange rate adjusted by aggregate GDP deflators. (Q1##) Nominal exchange rate adjusted by CPIs.

"Relative prices": (other sector deflator/manufacturing sector deflator) in log terms.

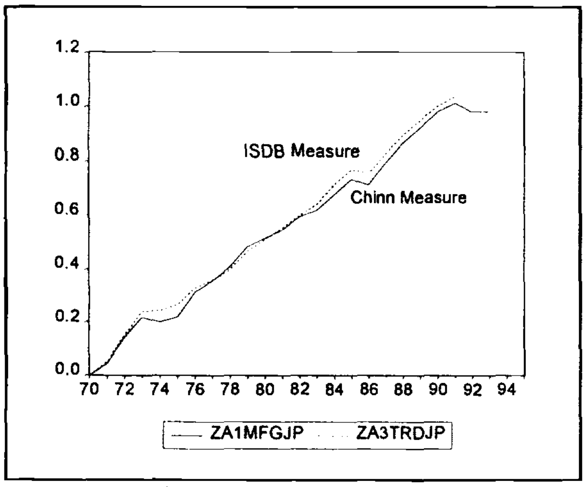
"Relative productivity": (manufacturing sector productivity/other sector productivity) in log terms.

Appendix: Comparison of Productivity Measures

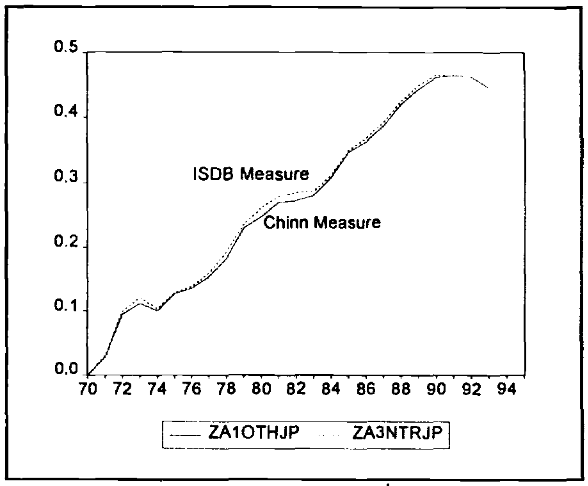
In this appendix, I discuss the characteristics of the data used in the paper. The first aspect is the degree of correspondence between the calculated series with alternative measures of tradable and nontradable sector productivity. The second is match of the World Bank's measures of manufacturing productivity versus those I calculate from World Bank and ADB data.

A1. Tradable and Nontradable Sector Productivity Measures

Of the East Asian countries, Japan is the only one for which both World Bank *and* OECD data are available. The two measures are depicted below, for both the tradable and nontradable sectors. The series are rescaled to a common base of 1970 = 0.



Appendix Figure 1: Japanese Tradable Productivity Measures



Appendix Figure 2: Japanese Nontradable Productivity Measures

The two series match up in each figure quite well, despite differing coverage. In Appendix Figure 1, the "Chinn Measure" is World Bank real manufacturing output divided by manufacturing employment, obtained in this case from ILO (in log terms). The "ISDB Measure" is labor productivity from the OECD's *International Sectoral Data Base*, and covers total industrial output and agriculture. In Appendix Figure 2, the Chinn Measure is World Bank output in services and construction, divided by employment in that sector.

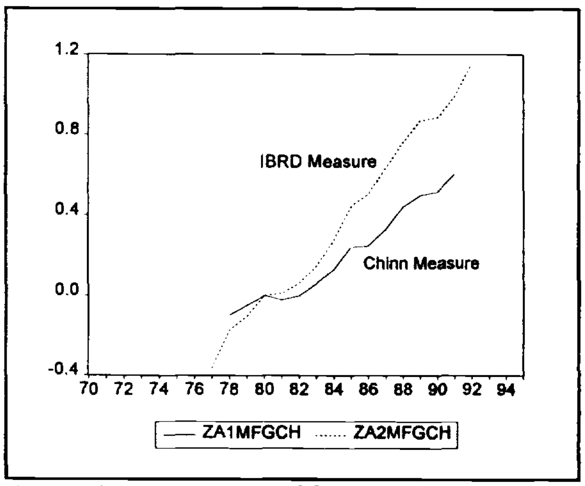
For other East Asian countries, the Chinn Measure is World Bank or ADB output divided by either World Bank or ADB employment figures; these employment figures are usually adjusted by adding together the appropriate categories to match the sectoral breakdown of output.

A2. World Bank and Calculated Measures of Manufacturing Productivity

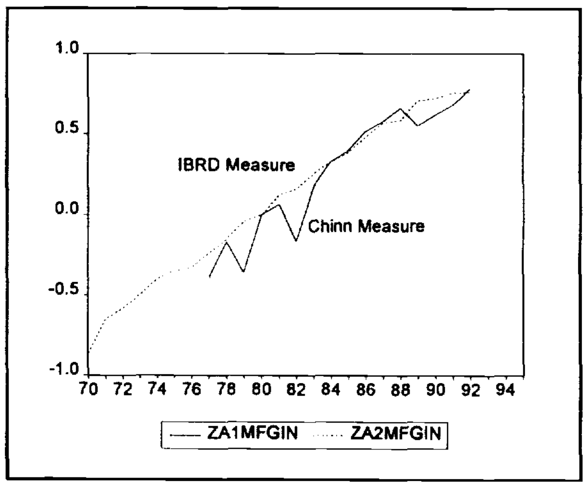
For the other East Asian countries, there is no way in which to determine directly the validity of the nontradables proxy for productivity used. However, a partial assessment can be made by inspecting the reported World Bank manufacturing productivity estimates (labeled "IBRD Measure" in the figures below) and those I calculate. These pairs of variables are depicted in Appendix Figures 3 through 7 for China, Indonesia, Korea, Philippines and Singapore. Each pair of series are rescaled to 1980 = 0.

As one can see, the two series that show the greatest deviation are China and the Philippines. In the other cases for which World Bank measures, the trends are similar. In the Chinese case, the IBRD measure shows much more rapid growth, so that there is a substantial underprediction of manufacturing productivity by 1991. I do not report the results using the World Bank measure because they fail to provide any plausible estimates (the wrong sign is obtained on relative productivity). However, because of this discrepancy, the Chinese regression results involving productivity are probably uninformative.

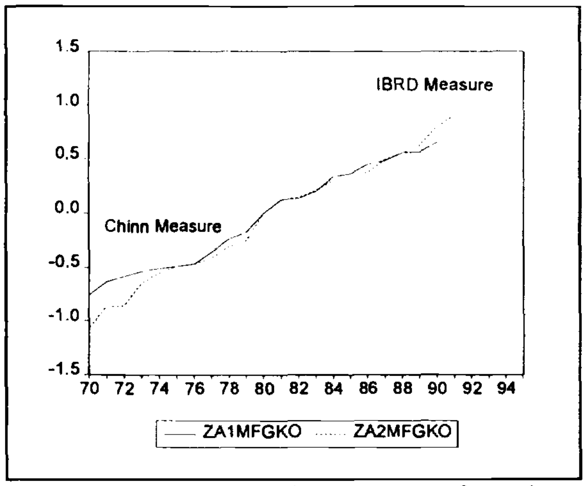
In the Philippines case, it is unclear why the series are so divergent; the IBRD measure shows a much less pronounced fall in the wake of Marcos' fall, and a more rapid rise in 1990 to 1991. Given the availability of the IBRD measure, I have used this measure in the reported regression results.



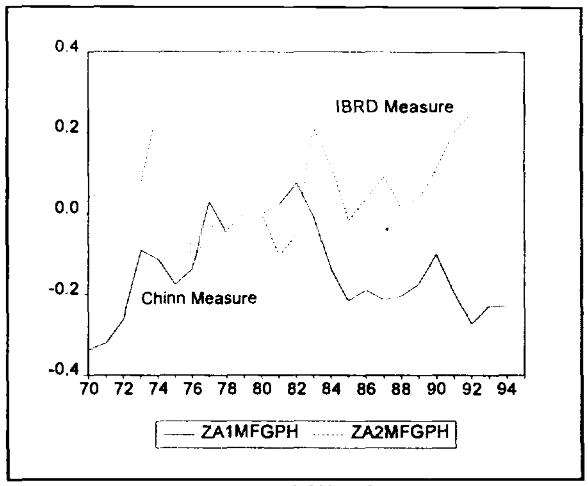
Appendix Figure 3: Chinese Manufacturing Productivity Measures



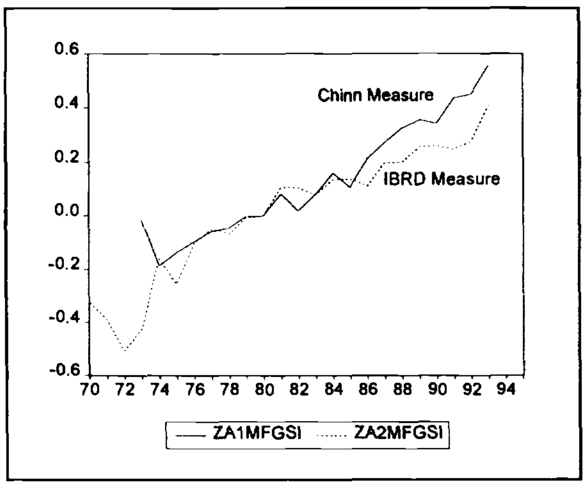
Appendix Figure 4: Indonesian Manufacturing Productivity Measures



Appendix Figure 5: Korean Manufacturing Productivity Measures



Appendix Figure 6: Philippines Manufacturing Productivity Measures



Appendix Figure 7: Singapore Manufacturing Productivity Measures

Table 1
Changes in Real Exchange Rates, Relative Prices, and Relative Productivity

Panel 1.1: Real Exchange Rates and Relative Prices

Country	Real Exchange Rate (-q)	Relative Prices	Obs.	Sample
China	0.0569	-0.0216	20	1971-91
Indonesia	0.0018	0.11517	20	1970-90
Japan	-.03484	-0.0265	21	1970-91
Korea	-.02231	-0.0378	21	1970-91
Malaysia	-.02144	0.00680	12	1971-83
Philippines	-.00039	0.00584	21	1970-91
Singapore	-.01505	0.01756	21	1970-91
Taiwan	-.01135	-0.0093	17	1974-91
Thailand	0.0035	0.00583	20	1970-90

Panel 1.2: Real Exchange Rates and Relative Productivities

Country	Real Exchange Rate (-q)	Relative Produc-ivity	Obs.	Sample
China	0.0763	0.01118	11	1978-89
Indonesia	0.0479	0.08165	14	1977-90
Japan	-.03484	-0.0049	21	1970-91
Korea	-.02419	-0.0125	20	1970-90
Malaysia	0.0014	0.00877	18	1971-90
Philippines	0.0262	0.02058	12	1977-78, 80-91
Singapore	-.00108	0.03513	18	1973-91
Taiwan	-.01135	0.02898	17	1974-91
Thailand	0.0142	0.02962	13	1977-90

Notes: Annual change, expressed in log differences. Real exchange rate expressed as $ln(1/Q)$ or as $-q$.

Table 2
NLS Error Correction Regressions
Real Exchange Rates and Relative Nontradable/Tradable Prices

$$\Delta q_t = \mu + \sum_{i=1}^{k} \gamma_i \Delta q_{t-i} + \sum_{j=1}^{m} \delta_j \Delta \tilde{p}_{t-i} + \phi(q_{t-1} + \alpha \tilde{p}_{t-1}) + \nu_t$$
$$\tilde{p} \equiv (p^N - p^T) - (p^{N*} - p^{T*})$$

Coeff	CH	IN	JP	KO	MA	PH	SI	TA	TH
Est. Meth.	NLS	NLS	NLS	NLS	NLS	NLS	NLS	NLS	NLS
ECT	-0.024 (0.058)	-0.259*** (0.070)	-0.404** (0.166)	-0.523*** (0.171)	-0.699*** (0.153)	-0.488* (0.273)	-0.433*** (0.100)	-0.184† (0.110)	-0.286 (0.191)
p^N-p^T	-0.575 (9.183)	-0.464*** (0.120)	-0.761* (0.391)	-0.280** (0.104)	-0.703 (0.387)	-0.407 (0.663)	-0.010 (0.171)	2.203 (1.423)	-0.281 (0.773)
Δq_{t-1}	0.467* (0.236)		0.469* (0.237)	0.556** (0.226)	0.282 (0.149)	0.286 (0.284)	0.776 (0.144)	0.545** (0.206)	0.452 (0.299)
Δq_{t-2}				0.476* (0.251)					
Δp_{t-1}		0.632** (0.233)	-0.798 (0.690)	-0.207 (0.288)		-0.372 (0.433)	-0.405 (0.212)	0.292 (0.215)	-0.130 (0.301)
Δp_{t-2}		0.486 (0.194)							
$\bar{R}^2$	.97	.94	.70	.85	.55	.51	.80	.88	.70
N	20	20	21	20	12	21	21	17	19
LM(2)	3.167	0.824	0.986	5.150	2.571	5.296	2.348	2.556	0.735
p-val	[.205]	[.662]	[.611]	[.080]	[.277]	[.071]	[.309]	[.279]	[.692]
Iter.	1	1	4	1	1	4	4	4	3
Smpl	1973-92	1973-92	1972-92	1973-92	1973-84	1972-92	1972-92	1976-92	1972-90

Notes: Regression coefficients from Nonlinear Least Squares. †{*}(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. "LM(2)" is the Breusch-Godfrey TR^2 test for serial correlation of order 2. "N" is the effective number of observations included in the regression. "Iter." is the number of iterations necessary for convergence. "Smpl" is the sample period.

Table 3
Kremer-Ericsson-Dolado Error Correction Regressions
Real Exchange Rates and Relative Tradable/Nontradable Productivity Levels

$$\Delta q_t = \mu + \sum_{i=1}^{k} \gamma_i \Delta q_{t-i} + \sum_{j=1}^{m} \delta_j \Delta \tilde{a}_{t-i} + \phi(q_{t-1} + \alpha \tilde{a}_{t-1}) + \nu_t$$

$$\tilde{a} \equiv (a^T - a^N) - (a^{T*} - a^{N*})$$

Coeff	CH	IN	JP	KO	MA	PH[1/]	SI	TA	TH
Est Meth.	OLS	OLS	OLS	OLS	OLS	OLS	NLS	NLS	OLS
ECT	-0.059 (0.080)	-0.071 (0.135)	-0.260** (0.120)	-0.050 (0.146)	-0.302* (0.157)	-0.258* (0.123)	0.000 (0.167)	-0.244† (0.172)	-0.183 (0.171)
a^T-a^N	-0.5	-0.5	-0.5	-0.5	-0.5	-0.5	-142.109 (57848.94)	1.309† (0.796)	-0.5
Δq_{t-1}	0.443† (0.290)	0.150 (0.279)	0.310† (0.209)	0.407† (0.289)	0.412* (0.235)	0.243 (0.307)	0.528*** (0.230)	0.643** (0.248)	0.239 (0.300)
Δq_{t-2}								-0.351 (0.295)	
Δa_{t-1}				0.348 (0.249)	0.605† (0.391)		-0.261† (0.161)	-0.218 (0.177)	
Δa_{t-2}							-0.289** (0.124)		
$\bar{R}^2$	.94	.88	.69	.85	.57	.16	.82	.87	.76
N	12	15	21	20	17	14	17	16	13
LM(1)	0.177 [.674]	0.195 [.659]	0.046 [.830]	1.746 [.186]	0.428 [.513]	0.177 [.674]	2.045 [.153]	0.095 [.758]	0.150 [.700]
Iter.	--	--	--	--	--	--	3	5	--
Smpl	1979-90	1978-92	1972-92	1972-91	1973-89	1978-92	1976-92	1977-92	1978-90

Notes: Regression coefficients from Nonlinear Least Squares. †{*}(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. LM(1) is the Breusch-Godfrey TR^2 test for serial correlation of order 1 [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.

1/ Philippines regression uses World Bank measure of manufacturing productivity. See text.

Table 4
Kremer-Ericsson-Dolado Error Correction Regressions
Real Exchange Rates, Relative Tradable/Nontradable Productivity Levels and Government Spending Changes

$$\Delta q_t = \mu + \sum_{i=1}^{k} \gamma_i \Delta q_{t-i} + \sum_{j=1}^{m} \delta_j \Delta \tilde{a}_{t-i} + \phi(q_{t-1} + \alpha \tilde{a}_{t-1}) + \kappa_1 \Delta g_t - \kappa_2 \Delta g_t^* + \nu_t$$

$$\tilde{a} \equiv (a^T - a^N) - (a^{T*} - a^{N*})$$

Coeff	CH	IN	JP	KO	MA	PH1/	SI	TA	TH
Method		OLS	OLS	OLS	OLS	OLS	NLS	OLS	OLS
ECT		-0.169† (0.103)	-0.260* (0.126)	-0.023 (0.137)	-0.077 (0.122)	-0.173 (0.133)	-0.128 (0.179)	0.032 (0.071)	0.087 (0.372)
a^T-a^N		-0.5	-0.5	-0.5	-0.5	-0.5	0.310 (1.004)	-0.5	-0.5
Δq_{t-1}		0.193 (0.229)	0.247 (0.232)	0.329† (0.282)	0.458*** (0.157)		0.535* (0.257)	0.899*** (0.227)	0.095 (0.457)
Δq_{t-2}									
Δa_{t-1}		0.190* (0.095)		0.321 (0.232)	0.818** (0.298)		-0.168† (0.123)		
Δa_{t-2}							-0.289** (0.124)		
Δg_t		-0.585 (1.082)	-1.089 (1.131)	-1.338* (0.715)	0.409 (0.728)	-1.287† (0.919)	0.168 (0.422)	0.084 (0.429)	-0.217 (0.966)
Δg^*_t		0.741** (0.256)	-0.307 (0.894)	0.127 (0.241)	-0.767*** (0.195)	0.116 (0.293)	-0.191† (0.131)	-1.074*** (0.322)	-0.512 (0.561)
$\bar{R}^2$		.94	.69	.87	.81	.20	.79	.91	.73
N		14	21	20	17	14	18	17	13
LM(1)		0.360 [.549]	0.026 [.871]	3.030 [.082]	0.001 [.972]	0.879 [.348]	1.143 [.285]	0.349 [.554]	0.012 [.913]
Iter.		--	--	--	--	--	3	--	--
Smpl		1979-92	1972-92	1972-91	1973-89	1978-92	1975-92	1976-92	1978-90

Notes: Regression coefficients from Nonlinear Least Squares. †{*}(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. LM(1) is the Breusch-Godfrey TR^2 test for serial correlation of order 1 [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.

1/ Philippines regression uses World Bank measure of manufacturing productivity. See text.

Table 5
Quasi-KED Error Correction Regressions
Real Exchange Rates, Relative Tradable/Nontradable Productivity Levels, Government Spending Changes and Oil Prices

$$\Delta q_t = \mu + \gamma_1 \Delta q_{t-1} + \delta_1 \Delta \tilde{a}_{t-1} + \phi(q_{t-1} + \alpha \tilde{a}_{t-1} - \xi p^{oil}_{t-1}) + \kappa_1 \Delta g_t - \kappa_2 \Delta g^*_t + \zeta p^{oil}_t + \nu_t$$
$$\tilde{a} \equiv (a^T - a^N) - (a^{T*} - a^{N*})$$

Coeff	IN	IN	JP	JP	KO	KO
Estimation Method	NLS	NLS	NLS	NLS	OLS	OLS
ECT	-0.601** (0.260)	-0.361† (0.240)	-0.216** (0.010)	-0.209* (0.105)	-0.150† (0.099)	-0.130† (0.098)
a^T-a^N	-0.347** (0.152)	-0.529* (0.272)	-0.5	-0.5	-0.5	-0.5
p^{oil}	0.386*** (0.096)	0.253 (0.191)	0.381† (0.275)	-0.369 (0.298)	--	--
Δq_{t-1}	0.393† (0.255)	0.263 (0.218)	0.401* (0.229)	0.371† (0.206)	0.792 (0.218)	0.665 (0.229)
Δa_{t-1}	0.208* (0.105)	0.222** (0.087)				
Δp^{oil}_t			-0.210*** (0.060)	-0.205*** (0.071)	-0.106*** (0.037)	-0.092**1/ (0.040)
Δg_t		-0.657* (0.299)		-0.461 (0.972)		-1.071† (0.660)
Δg^*_t		0.657* (0.299)		0.148 (0.812)		0.216 (0.221)
$\bar{R}^2$	.91	.94	.79	.81	.89	.89
N	14	14	21	21	20	20
LM(1)	0.242 [.623]	0.002 [.967]	1.248 [.264]	1.302 [.254]	0.716 [.398]	2.576 [.108]
Iter.	6	6	6	6	--	--
Smpl	1979-92	1979-92	1972-92	1972-92	1972-91	1972-91

Notes: Regression coefficients from Nonlinear Least Squares. †{*}(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. LM(1) is the Breusch-Godfrey TR^2 test for serial correlation of order 1 [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.
1/ Oil price change lagged once.

Table 6
Panel NLS Error Correction Model Regressions

	Full Sample	No China	No China	No China	No China
ECT	-0.091*** (0.030)	-0.243*** (0.047)	-0.153** (0.057)	-0.146** (0.057)	-0.149 (0.057)
p^N-p^T	-0.679** (0.300)	-0.411*** (0.093)			
a^T-a^N			-0.238 (0.309)	-0.231 (0.329)	-0.219 (0.325)
p^{oil}					0.111 (0.712)
Δq_{t-1}	0.181** (0.084)	0.241*** (0.084)	0.257*** (0.100)	0.275*** (0.101)	0.258** (0.104)
Δp_{t-1}	0.169† (0.127)	0.197† (0.131)			
Δa_{t-1}			0.192*** (0.067)	0.190*** (0.068)	0.188*** (0.068)
$\Delta p^{oil}{}_t$					0.095** (0.043)
Δg_t				-1.010 (0.880)	-2.924 (2.206)
$\Delta g^*{}_t$				-0.014 (0.115)	-0.000 (0.117)
Wald(ECT)	0.206 [.650]	4.583 [.032]			
Wald(p^N-p^T)	0.632 [.427]	3.588 [.058]			
Wald(a^T-a^N)			0.138 [.710]	0.011 [.918]	
Wald(Δg^*)				0.073 [.786]	
Wald(p^{oil})					2.389 [.122]
Wald(Δp^{oil})					102.04 [.000]
$\bar{R}^2$	.99	.99	.99	.99	.99
N	166	146	131	131	131

Notes: Regression coefficients from Nonlinear Least Squares; regressions include time and currency effects. †{*}(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. N is the effective number of observations included in the regression. Wald(X) indicates a Wald test on the restriction of common slope coefficients on the X variable.

1/ Oil price change lagged once.

Table 7
Panel NLS Error Correction Model Regressions on Restricted Sample

	No China, Singapore, Taiwan and Thailand		
ECT	-0.166** (0.081)	-0.164** (0.081)	-0.167** (0.081)
a^T-a^N	-0.620† (0.394)	-0.631† (0.407)	-0.614† (0.400)
p^{oil}			0.129 (0.769)
Δq_{t-1}	0.266** (0.133)	0.260* (0.135)	0.237* (0.139)
Δa_{t-1}	0.280*** (0.095)	0.279*** (0.095)	0.275*** (0.096)
$\Delta p^{oil}{}_t$			0.100* (0.056)
Δg_t		-1.077 (1.119)	-3.198 (2.756)
$\Delta g^*{}_t$		0.040 (0.195)	0.072 (0.200)
Wald(a^T-a^N)	0.001 [.977]		
Wald(Δg^*)		0.004 [.952]	
Wald(p^{oil})			2.204 [.138]
Wald(Δp^{oil})			76.460 [.000]
Wald(Δq)	3.136 [.077]	2.732 [.098]	0.056 [.098]
$\bar{R}^2$	.99	.99	.99
N	84	84	84

Notes: Regression coefficients from Nonlinear Least Squares; regressions include time and currency effects. †{*}(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. N is the effective number of observations included in the regression.

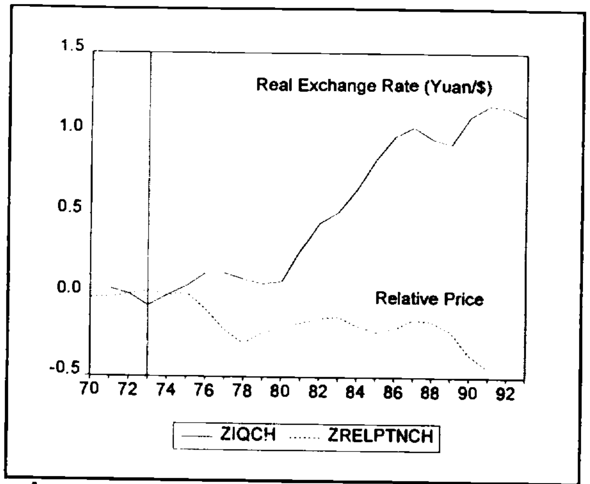


Figure 1: China

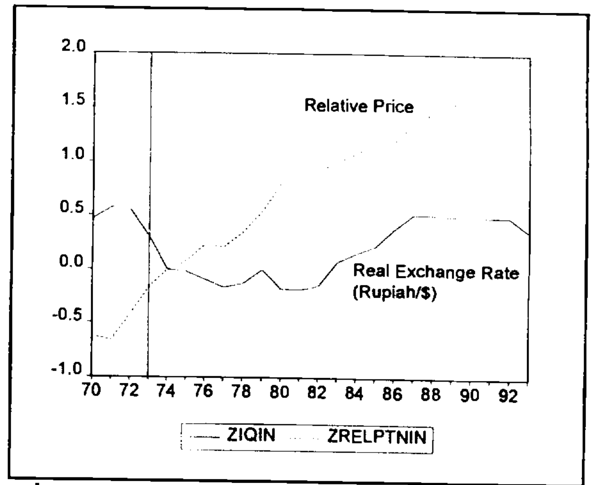


Figure 2: Indonesia

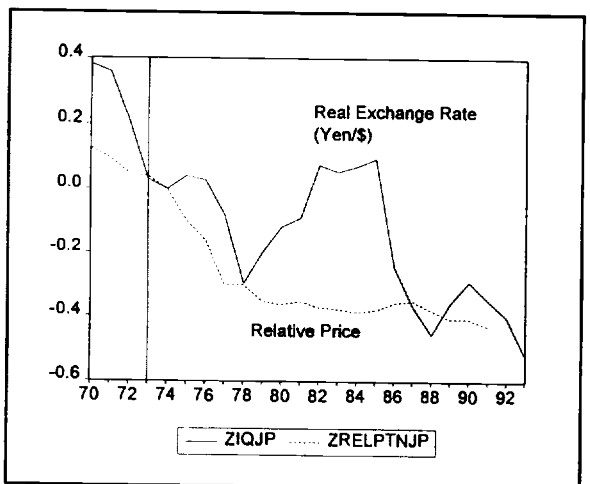


Figure 3: Japan

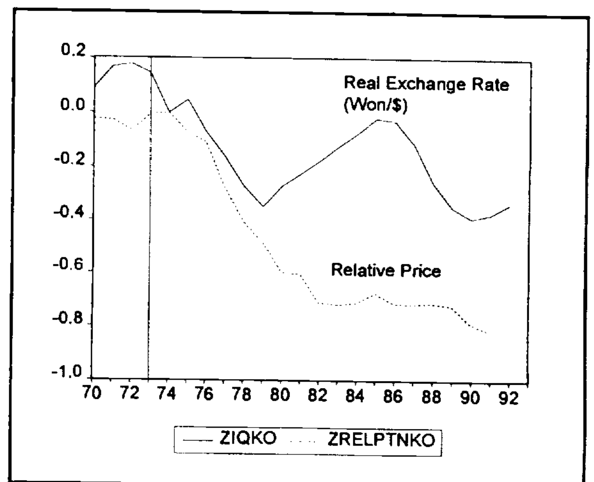


Figure 4: Korea

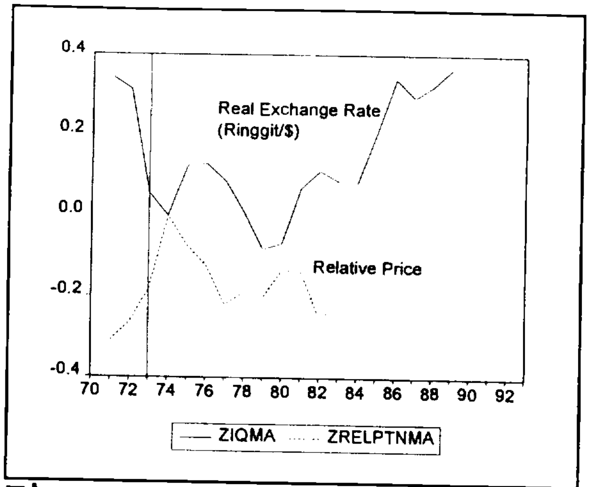


Figure 5: Malaysia

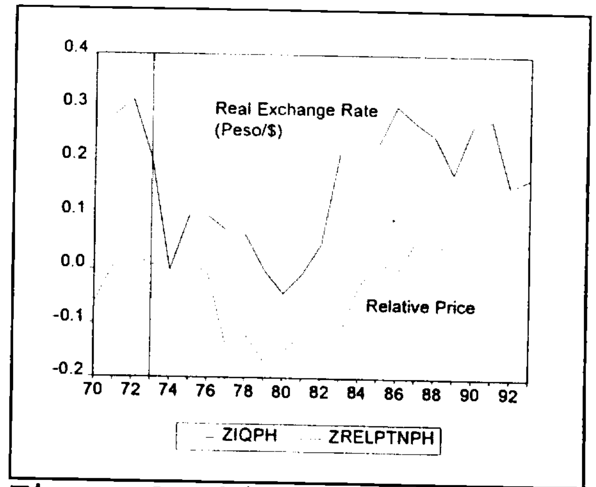


Figure 6: Philippines

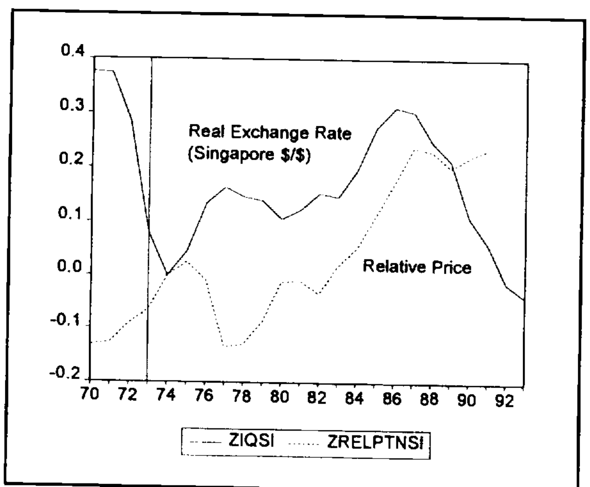


Figure 7: Singapore

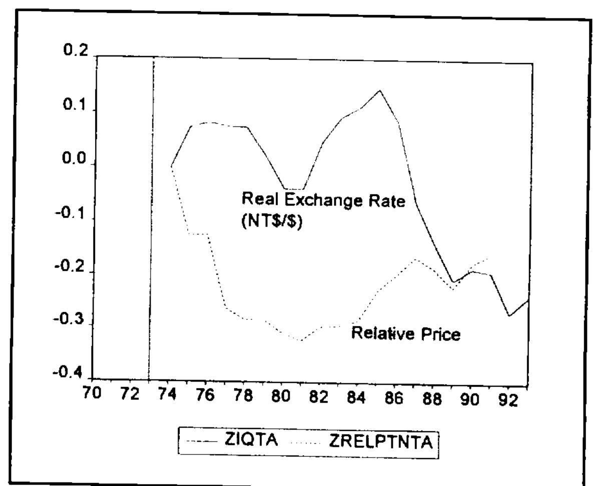


Figure 8: Taiwan

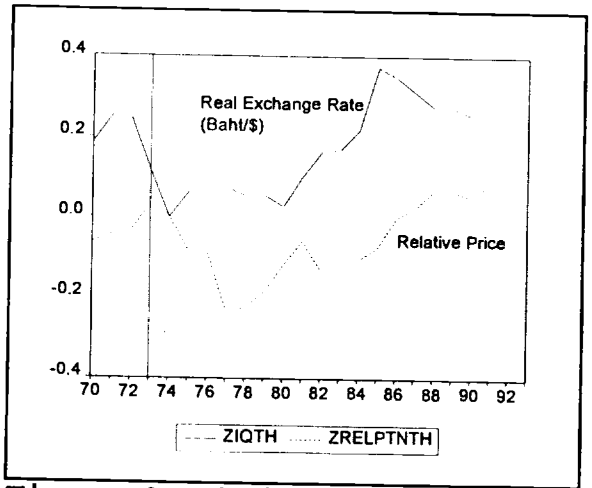


Figure 9: Thailand

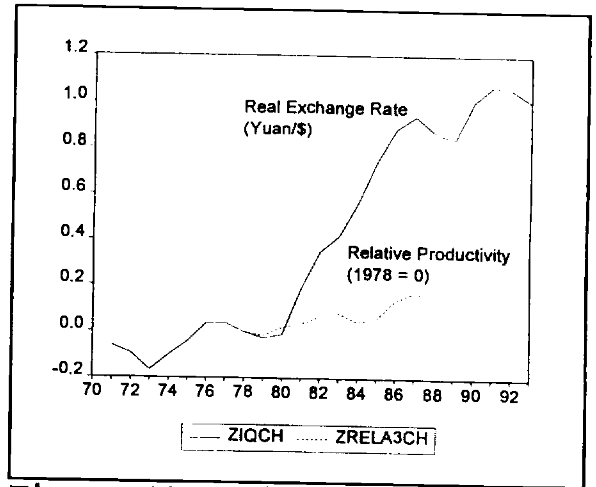


Figure 10: China

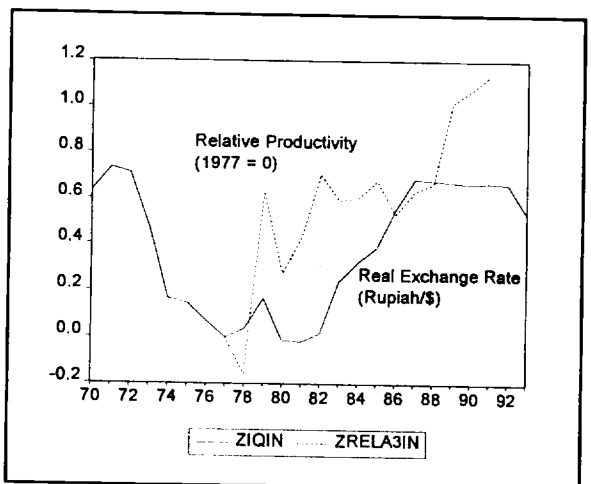


Figure 11: Indonesia

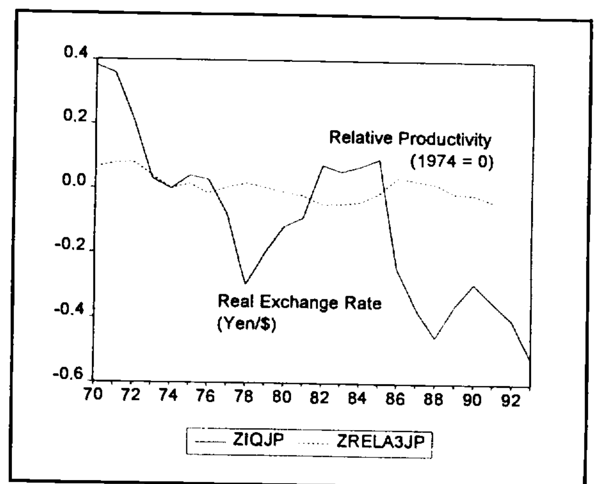


Figure 12: Japan

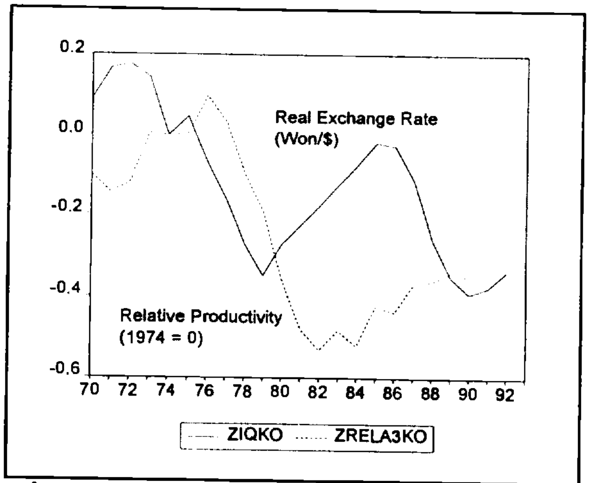


Figure 13: Korea

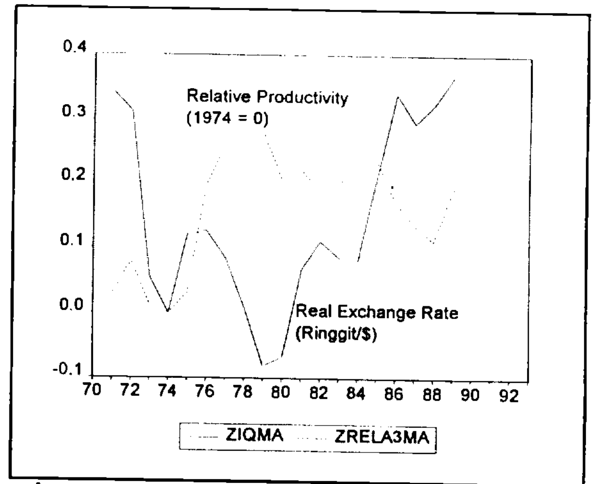


Figure 14: Malaysia

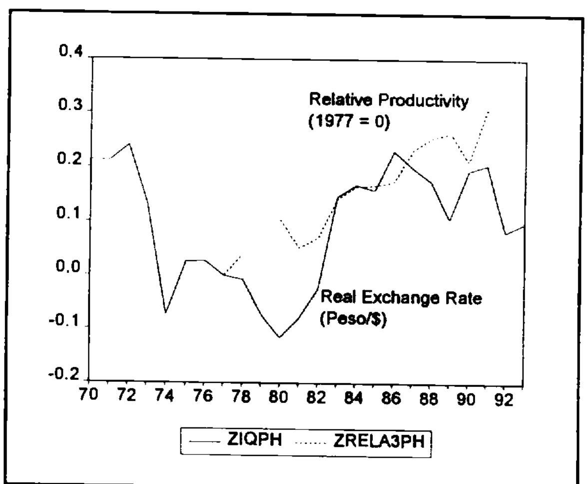


Figure 15: Philippines

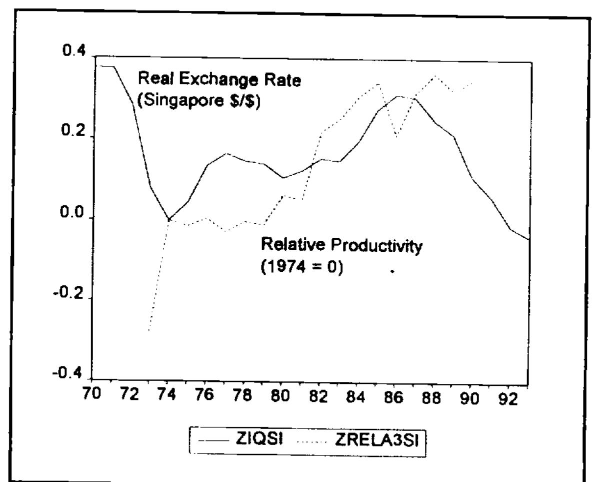


Figure 16: Singapore

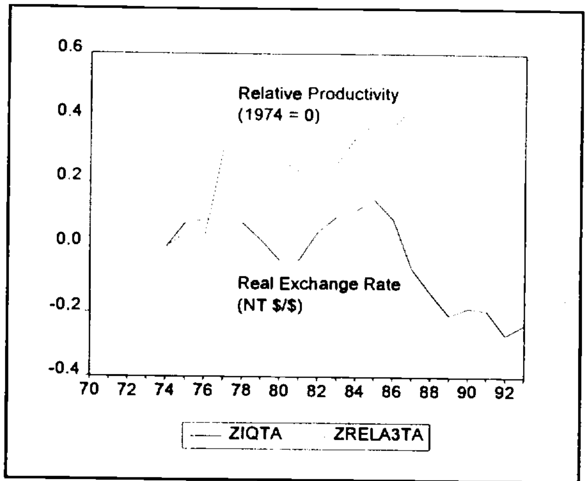


Figure 17: Taiwan

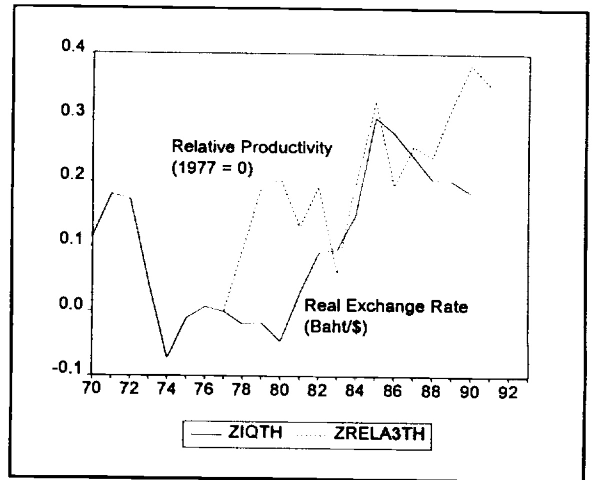


Figure 18: Thailand

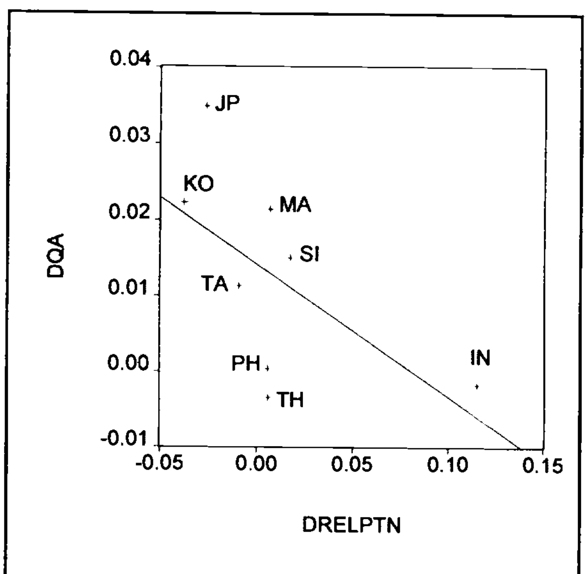


Figure 19: Cross Section of Real Exchange Rates and Relative Prices (excl. China)

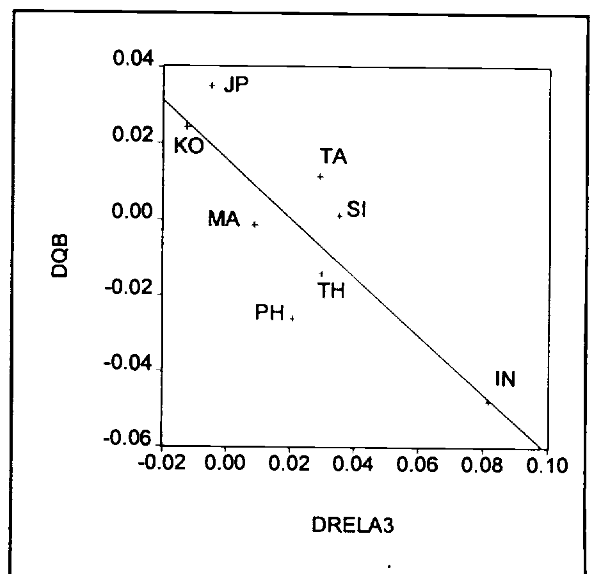


Figure 20: Cross Section of Real Exchange Rates and Relative Productivities (excl. China)