Review of International Economics, 8(1), 20–43, 2000

The Usual Suspects? Productivity and Demand Shocks and Asia–Pacific Real Exchange Rates

*Menzie D. Chinn**

Abstract

A productivity-based model of East Asian relative prices and real exchange rates is tested using calculated productivity levels for China, Indonesia, Japan, Korea, Malaysia, Philippines, Singapore, Taiwan, and Thailand. Time-series regressions of the exchange rate on relative productivity ratios indicate such a relationship for Japan, Malaysia, and the Philippines (and Indonesia and Korea when oil prices are included). Panel regression provides slightly more encouraging results when the panel encompasses a subset of countries (Indonesia, Japan, Korea, Malaysia, and the Philippines). Neither government spending nor the terms of trade appear to be important factors.

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).

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. However, with the exception of Japan, one would not expect the Harrod–Balassa–Samuelson (hereafter HBS) effect to be very apparent.[1]

In this paper, I examine the relevance of the HBS effect in East Asian countries and compare it against the importance of other effects by rounding up "the usual suspects": sectoral productivity levels, government spending ratios, the real price of oil, and the terms of trade. 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. However, past research on these economies has been hampered by the unavailability of sectoral productivity data. Thus, previous researchers have been able to examine only one portion of the HBS hypothesis—that relative prices of traded to nontraded goods should determine in large part real exchange rates (Isard and Symansky, 1996). Yet relative prices could be driven by either the mechanism outlined by HBS, or by long-run demand-side factors such as an income-elastic preference for the consumption of nontradables. Examination of the correlation between relative prices and the real exchange rate cannot differentiate between the two explanations.[2]

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* Chinn: University of California, Santa Cruz, CA 95064, USA. Tel: (831) 459-2079; Fax: (831) 459-5900; E-mail: chinn@cats.ucsc.edu. I thank Peter Isard and the UCSC Brown Bag group for discussions that stimulated this research, Yin-Wong Cheung, Rob Feenstra, Jaewoo Lee, Eric Ramstetter, Nirvikar Singh, seminar participants at U.C. Davis, U.C. Irvine and Mount Holyoke College, an anonymous referee and the Associate Editor for valuable comments. This paper was also presented at the conference on "International Economic Links and Policy Formation" (Kobe University, June 1997), and at the CEPR/Austrian National Bank conference on "Real Exchange Rates: Recent Theories and Evidence" (Vienna, April 1998). It was completed while the author was a visiting scholar at the Center for Pacific Basin Monetary and Economic Studies of the Federal Reserve Bank of San Francisco. Financial support of faculty research funds of the University of California, Santa Cruz, is gratefully acknowledged.

Another set of papers incorporates the assumption that overall productivity growth is highly correlated with the tradable/nontradable productivity differential. I will argue below that this is an inappropriate assumption which may lead to misleading inferences. Hence a contribution of this paper is to use data more consistent with testing the validity of 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, while productivity ratios generally *do* appear to be related to the real exchange rate. Furthermore, the results do not appear to be much stronger when appeal is made to panel regression techniques.

2. Model Derivation and 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, p^*, is constructed, with the same weight ascribed to nontraded goods. Using the definition of the real exchange rate as:

$$q_t \equiv (s_t + p_t^* - p_t), \tag{2}$$

(where s is the log of the domestic currency price of foreign currency), the following holds:

$$\begin{aligned} q_t &= (s_t - p_t^T + p_t^{T*}) - \alpha[(p_t^N - p_t^T) - (p_t^{N*} - p_t^{T*})] \\ q_t &= q_t^T - \alpha\tilde{p}_t. \end{aligned} \tag{3}$$

Although there are many alternative decompositions that can be undertaken, equation (3) is the most relevant since many economic models point to the second term as the determinant of the real exchange rate, while the first term—the common currency relative price of home and foreign tradables (q^T)—is assumed to be zero by purchasing power parity (PPP) as applied to traded goods. Examining the Asia–Pacific currencies, Isard and Symansky (1996) undertake a slightly different decomposition which allows the nontraded shares (α) to change over time. They obtain the result that the first term accounts for almost all of the movements in the real rate for the Chinese, Indonesian, Japanese, Philippine, and Thai currencies. In part this is an unsurprising conclusion, given the relatively high volatility of s, which enters in the q^T term.

Most researchers have proceeded under the assumption that q^T is $I(0)$ and q is $I(1)$. This implies a cointegrating relationship of the form:

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

Chinn (1996) examines East Asian exchange rates in the accounting sense where the aggregate price index is by construction the weighted sum of tradables and nontradables price indices, and using cointegration techniques determines that there is evidence that relative prices of tradables and nontradables do explain real exchange rates except for the Chinese, Indonesian, and Thai currencies.

In order to move beyond 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 another cointegrating relation:

$$q_t = -\alpha[(a_t^T - a_t^N) - (a_t^{T*} - a_t^{N*})] \equiv -\alpha\,\tilde{a}_t, \tag{5}$$

where a^i is total factor productivity (TFP) in sector i, and the sectoral production functions are assumed identical. Equations (4) and (5) will motivate the empirical work.

While the variables of primary interest are the productivity indicators, it is obvious that other factors can influence relative prices, including demand-side influences. One such variable is government spending. In the conventional interpretation, where government consumption is thought to fall mainly on nontradable goods (such as services), government spending should raise the relative price of nontradable goods, and hence appreciate the domestic currency. Balvers and Bergstrand (1997, 1998) have observed that this conclusion arises if government and private consumption are substitutes in the representative agent's utility function. If they prove instead to be *complements*, this "consumption tilting" effect could mitigate, or even reverse, the conventional expected effect.

The real exchange rate can also be affected by changes in wealth which may arise from movements in the terms of trade. As a check, I also investigate whether changes in the real price of oil may induce similar effects.

The assumption that the real exchange rate is integrated of order one merits some discussion. Time-series tests of the real exchange rate spanning the floating-rate period typically fail to reject the unit-root null, and this is true in the dataset used in this paper too. Admittedly, recent studies employing panel unit-root tests have rejected the unit-root null (Frankel and Rose, 1996). However, these results have proven sensitive to the currencies included in the regression (Papell, 1997). Given the ambiguity of the results, and *a priori* expectations regarding the nonstationarity of exchange rates in the region, the assumption of integrated real exchange rate series appears appropriate (the results of the panel unit-root tests reported in section 3 are consistent with this assumption).

3. Data and Variable Construction

Exchange Rates and Prices

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 China, Indonesia, Japan, Korea, Malaysia, the Philippines, Singapore, Thailand, Taiwan, and the US, for the period 1970–92. The annual data are drawn from a variety of sources. The nominal exchange rate is the bilateral period average, expressed against the US$ (in $/foreign currency unit). 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.

The price of tradables is proxied by the price deflator for manufacturing, and for the price of nontradables, the "other" price deflator, which includes services, construction, mining, and transportation. 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* is used.[3]

The exchange rate series and the respective relative price series are displayed in Figures 1–9. 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.

Sectoral Productivity Levels and Proxies

As remarked above, previous 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. 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:

$$q_t = \kappa_0 + \kappa_1(a_t - a_t^*) + u_t, \tag{6}$$

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

$$a_t = \omega a_t^T - \omega a_t^N. \tag{7}$$

The α coefficient in (5) is not identified; rather only the product $\kappa = \omega\alpha$ can be estimated. A more complicated situation arises if the coefficients on the sectoral a are not the same. Then an omitted variable situation arises, and cointegration is unlikely to be found, even if the model in (5) holds true.[4] 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:

$$A_i^L \equiv Y_i / N_i, \tag{8}$$

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 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. For instance, Union Bank of Switzerland (1997) reports weekly working hours for industrial workers ranging from 38 hours (in Singapore) to 48 hours (Manila). 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,

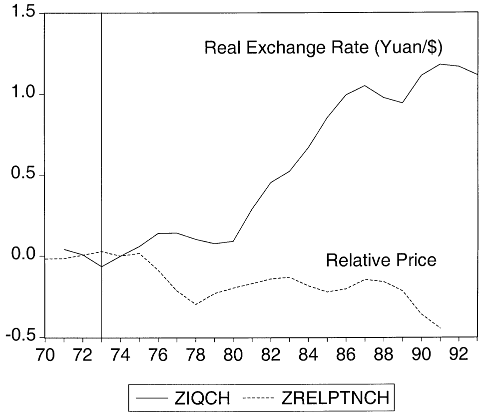

Figure 1. Relative Price: China

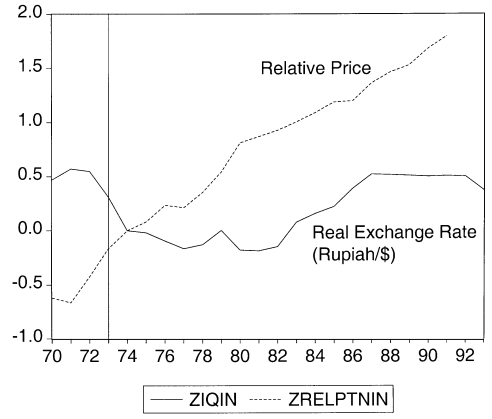

Figure 2. Relative Price: Indonesia

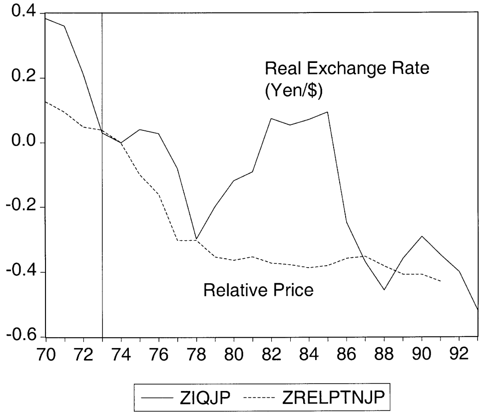

Figure 3. Relative Price: Japan

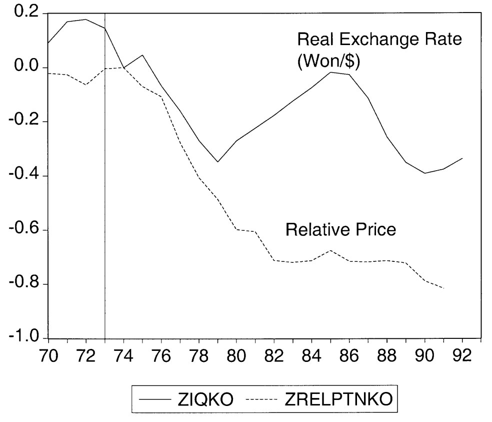

Figure 4. Relative Price: Korea

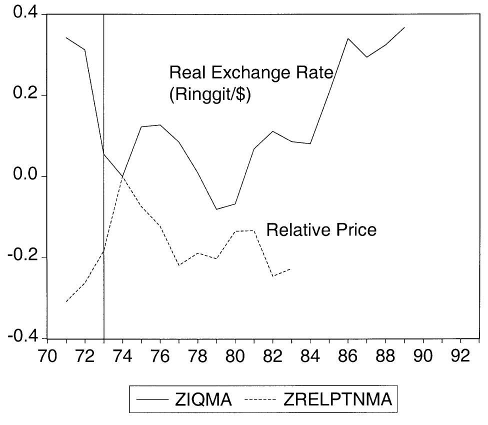

Figure 5. Relative Price: Malaysia

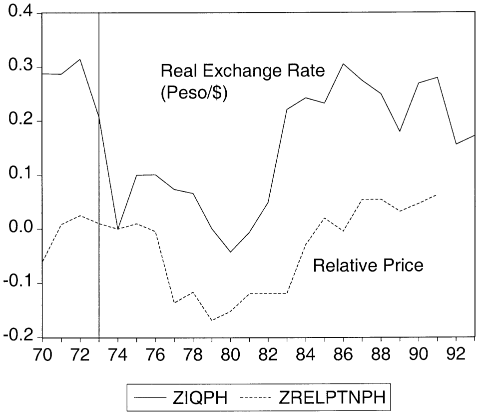

Figure 6. Relative Price: Philippines

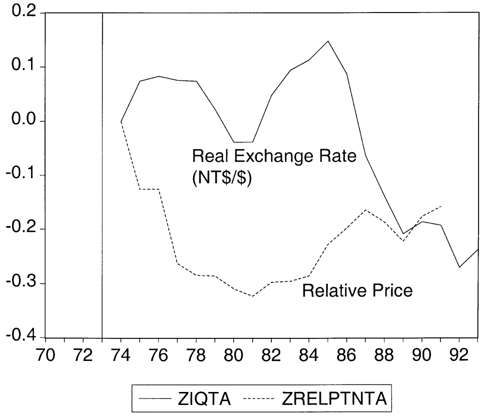

Figure 7. Relative Price: Taiwan

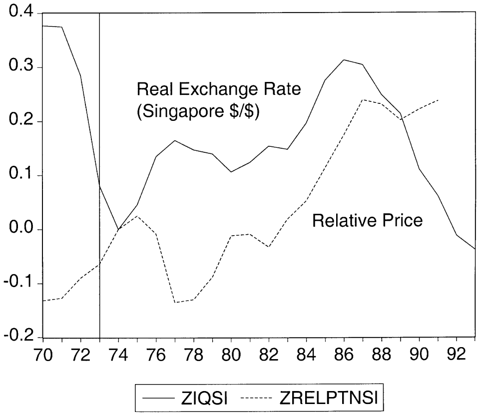

Figure 8. Relative Price: Singapore

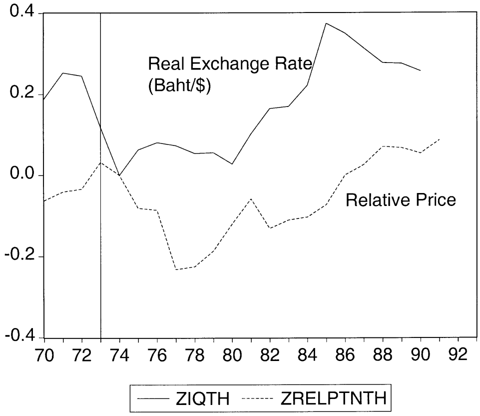

Figure 9. Relative Price: Thailand

transportation) and nontradable (services, construction) sectors from the OECD's *International Sectoral Data Base*.[5] These series also match quite well for manufacturing versus tradables, and "other" versus nontradables. These outcomes serve to improve one's confidence that the proxies used are not implausible.

Second, the proxy variable is labor productivity, rather than TFP as suggested by the model. Canzoneri, Cumby and Diba (1996) have argued that use of labor productivity is to be preferred 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 (one exception is Alwyn Young's (1995) analysis).[6]

The (inverse) exchange rate and relative productivity variables are displayed in Figures 10–18 (rescaled to a common base year equal zero). Once again, China and Taiwan fail to exhibit the expected correlation. In the other cases, the covariation is not very pronounced, although in an eclectic model of relative price determination, demand-side factors such as government spending come into play.

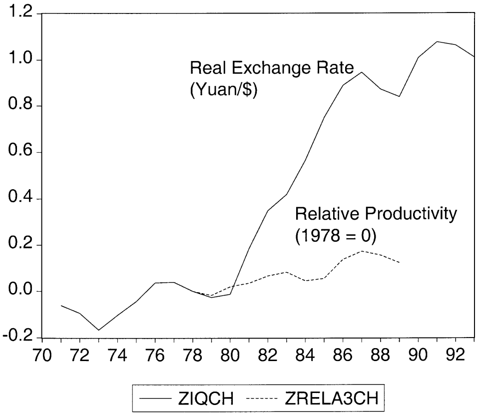

Figure 10. Relative Productivity: China

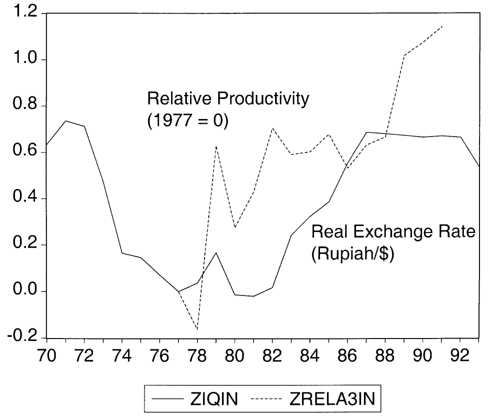

Figure 11. Relative Productivity: Indonesia

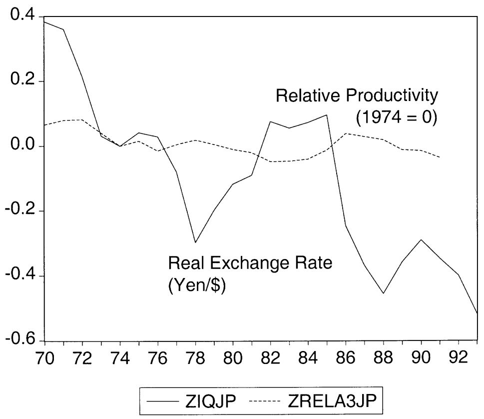

Figure 12. Relative Productivity: Japan

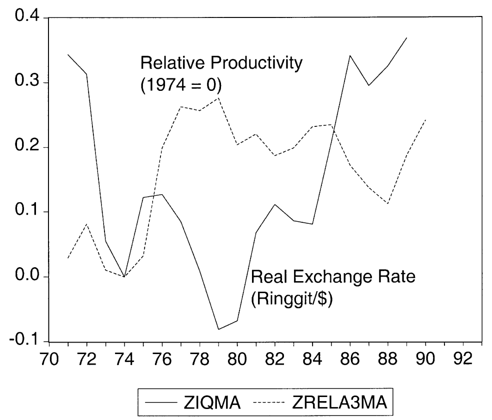

Figure 13. Relative Productivity: Malaysia

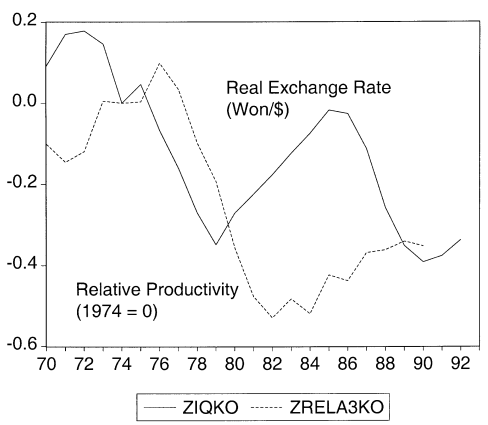

Figure 14. Relative Productivity: Korea

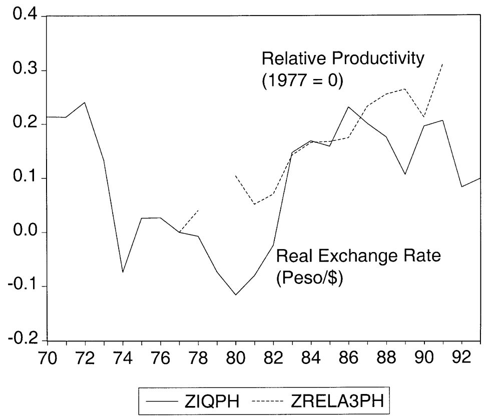

Figure 15. Relative Productivity: Philippines

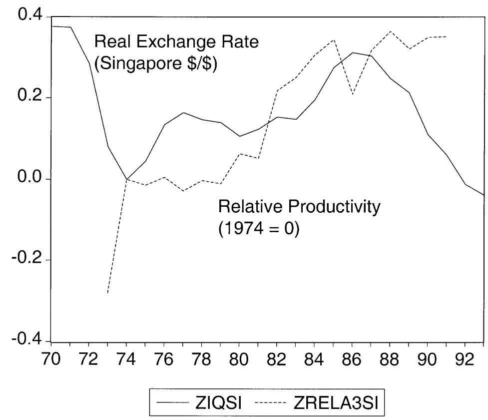

Figure 16. Relative Productivity: Singapore

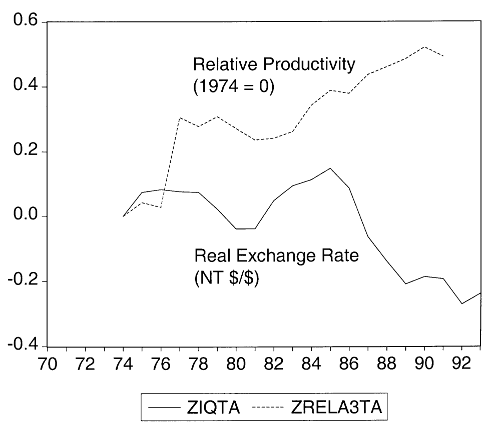

Figure 17. Relative Productivity: Taiwan

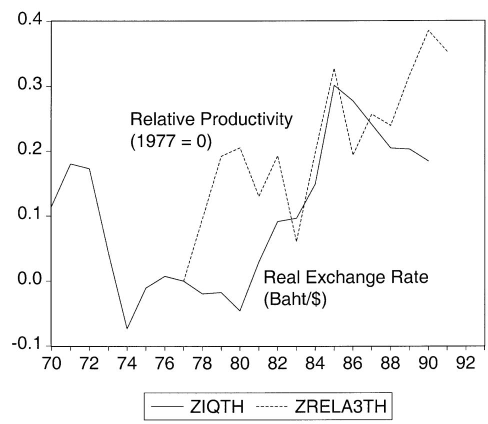

Figure 18. Relative Productivity: Thailand

A number of other studies have used GDP per capita as a measure of the sectoral productivity differential. This is clearly inadequate. Showing a correlation between the real 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 (Linder, 1961). *This* effect is often proxied by GDP per capita. To investigate this competing hypothesis, I gather data on this variable, measured in constant 1985 "international" dollars, expressed in relative log terms. The data span the period 1970–90, and are drawn from the Penn World Tables, Mark 5.5, described in Summers and Heston (1991).

Other Variables

The other variables include government spending, the terms of trade, and the real price of oil. The government-spending variable used in the regressions is the log of the ratio of nominal government consumption to nominal GDP. I also check whether the results are sensitive to use of the log ratio of real government consumption spending (deflated using the nontradables deflator) to real GDP. The terms of trade are measured as the log ratio of the export to import price, both measured in US$ terms. 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, the Singapore real exchange rate (all at the 5% marginal significance level), and the US–Taiwan relative productivity differential (at the 1% marginal significance level). Since this proportion of rejections could arise merely by chance, 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.

In order to assess whether a more powerful test for unit roots, exploiting the cross-sectional information, might detect evidence of mean stationary real exchange rates, I also applied a Levin and Lin (1992) regression to the exchange rate series. The t-statistic on the lagged real exchange rate (allowing for country fixed-effects) was –1.502, well above the Levin–Lin critical value for a marginal significance level of 10%.

4. The Real Exchange Rate and Relative Prices

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). Unfortunately, this procedure yields very dispersed estimates (for reasons suggested by Stock and Watson (1993)). Moreover, after adjusting the critical values for the loss of degrees of freedom, only a few instances of cointegration are found.[7]

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 I adopt a single-equation error-correction modeling approach, viz:

$$\Delta q_t = \mu + \sum_{i=1}^{k} \gamma_i \Delta q_{t-i} + \sum_{j=1}^{m} \delta_j \Delta \tilde{p}_{t-j} + \varphi[q_{t-1} + \alpha \tilde{p}_{t-1}] + v_t,$$
$$\tilde{p} = (p^N - p^T) - (p^{N*} - p^{T*}). \tag{9}$$

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.

Empirical Results

Regressions of the form of (9) were estimated for all real exchange rates. In general $k = j = 1$, except where noted. The results in Table 1 indicate that the real exchange does exhibit some mean reversion. In all cases except China, Taiwan, and Thailand, the coefficient on the error-correction term appears to be statistically significant. In these

Table 1. NLS Error Correction Regressions: Real Exchange Rates and Relative Prices

Coefficient	*(Sgn)*	*CH*	*IN*	*JP*	*KO*	*MA*	*PH*	*SI*	*TA*	*TH*
Method		NLS	NLS	NLS	NLS	NLS	NLS	NLS	NLS	NLS
ECT	(–)	–0.024	–0.259***	–0.404**	–0.523***	–0.699***	–0.488*	–0.433***	–0.184†	–0.286
		(0.058)	(0.070)	(0.166)	(0.171)	(0.153)	(0.273)	(0.100)	(0.110)	(0.191)
$p^N - p^T$	(–)	–0.575	–0.464***	–0.761*	–0.280**	–0.703	–0.407	–0.010	2.203	–0.281
		(9.183)	(0.120)	(0.391)	(0.104)	(0.387)	(0.663)	(0.171)	(1.423)	(0.773)
Δq_{t-1}	(?)	0.467*		0.469*	0.556**	0.282	0.286	0.776	0.545**	0.452
		(0.236)		(0.237)	(0.226)	(0.149)	(0.284)	(0.144)	(0.206)	(0.299)
Δq_{t-2}	(?)				0.476*					
					(0.251)					
Δp_{t-1}	(?)		0.632**	–0.798	–0.207		–0.372	–0.405	0.292	–0.130
			(0.233)	(0.690)	(0.288)		(0.433)	(0.212)	(0.215)	(0.301)
Δp_{t-2}	(?)		0.486							
			(0.194)							
$\bar{R}^2$		0.97	0.94	0.70	0.85	0.55	0.51	0.80	0.88	0.70
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-value		[0.205]	[0.662]	[0.611]	[0.080]	[0.277]	[0.071]	[0.309]	[0.279]	[0.692]
Iterations		1	1	4	1	1	4	4	4	3
Sample		1973–92	1973–92	1972–92	1973–92	1973–84	1972–92	1972–92	1976–92	1972–90

Notes: †{*}(**)[***] indicates significance at the 20%{10%}(5%)[1%] MSL. (Sgn) is anticipated sign of coefficent. 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. "Iterations" is the number of iterations necessary for convergence. "Sample" is the sample period.

cases the estimated rate of reversion toward equilibrium is above 25%. For 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—free intersectoral factor mobility, and to a lesser extent, full capital mobility.[8] For Taiwan and Thailand, the failure to find reversion is more puzzling, although it is true that neither of these countries opened up their capital markets until the 1980s.

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]

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 an 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 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 (Chinn, 1996).

5. The Real Exchange Rate, Productivity, and Demand Shocks

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 et al. (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

$$p_t^N - p_t^T = a_t^T - a_t^N, \tag{10}$$

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 (1996) 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:

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

and the constrained specification is:

$$\Delta q_t = \mu + \sum_{i=1}^{k} \gamma_i \Delta q_{t-i} + \sum_{j=1}^{m} \delta_j \Delta \tilde{a}_{t-j} + \varphi[q_{t-1} + 0.5 \tilde{a}_{t-1}] + v_t. \tag{12}$$

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 (12) were estimated for all real exchange rates except those two currencies, in which case equation (11) was implemented. In general $k = j = 1$ (except where noted). The results in Table 2 indicate that the real exchange rate does appear to be cointegrated with relative traded/nontraded productivity levels for Japan, Malaysia, and the Philippines.[10] 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 productivity levels *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.

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 and public and private consumption are substitutes, then a change in government spending should induce a current appreciation in the real exchange rate. The results of this respecification are reported in Table 3. 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 either of two possibilities. First, most government spending falls on tradable, rather than nontradable, goods. Second, private and government consumption in these Asia–Pacific economies are complements, rather than substitutes. In contrast, US government spending, when it is significant, does appreciate the US$ (this is the Korean case), suggesting that the nature of government consumption might fulfill different roles in the two cases. These conclusions are not sensitive to the use of nominal ratios. Using the ratio of the deflated government consumption to real GDP does not alter these basic results.

In previous productivity-based studies of the real exchange rate, the terms of trade and real oil prices have entered significantly (Chinn and Johnston, 1996; DeGregorio

Table 2. Error Correction Regressions: Real Exchange Rates and Relative Productivity Levels

Coefficient	*(Sgn)*	*CH*	*IN*	*JP*	*KO*	*MA*	*PH*[a]	*SI*	*TA*	*TH*
Method		OLS	OLS	OLS	OLS	OLS	OLS	NLS	NLS	OLS
ECT	(–)	–0.059	–0.071	–0.260**	–0.050	–0.302*	–0.258*	0.000	–0.244†	–0.183
		(0.080)	(0.135)	(0.120)	(0.146)	(0.157)	(0.123)	(0.167)	(0.172)	(0.171)
$a^T - a^N$	(–)	–0.5	–0.5	–0.5	–0.5	–0.5	–0.5	–142.109	1.309†	–0.5
								(57,848.94)	(0.796)	
Δq_{t-1}	(?)	0.443†	0.150	0.310†	0.407†	0.412*	0.243	0.528***	0.643**	0.239
		(0.290)	(0.279)	(0.209)	(0.289)	(0.235)	(0.307)	(0.230)	(0.248)	(0.300)
Δq_{t-2}	(?)								–0.351	
									(0.295)	
Δa_{t-1}	(?)				0.348	0.605†		–0.261†	–0.218	
					(0.249)	(0.391)		(0.161)	(0.177)	
Δa_{t-2}	(?)							–0.289**		
								(0.124)		
$\bar{R}^2$		0.94	0.88	0.69	0.85	0.57	0.16	0.82	0.87	0.76
N		12	15	21	20	17	14	17	16	13
LM(1)		0.177	0.195	0.046	1.746	0.428	0.177	2.045	0.095	0.150
		[0.674]	[0.659]	[0.830]	[0.186]	[0.513]	[0.674]	[0.153]	[0.758]	[0.700]
Iterations		—	—	—	—	—	—	3	5	—
Sample		1979–90	1978–92	1972–92	1972–91	1973–89	1978–92	1976–92	1977–92	1978–90

Notes: See Table 1. LM(1) is the Breusch–Godfrey TR^2 test for serial correlation of order 1 [*p*-values in brackets].
[a] Philippines regression uses World Bank measure of manufacturing productivity. See text.

Table 3. Error Correction Regressions: Real Exchange Rates, Relative Productivity Levels, and Government Spending Changes

Coefficient	*(Sgn)*	*IN*	*JP*	*KO*	*MA*	*PH*[a]	*SI*	*TA*	*TH*
Method		OLS	OLS	OLS	OLS	OLS	NLS	OLS	OLS
ECT	(–)	–0.169†	–0.260*	–0.023	–0.077	–0.173	–0.128	0.032	0.087
		(0.103)	(0.126)	(0.137)	(0.122)	(0.133)	(0.179)	(0.071)	(0.372)
$a^T - a^N$	(–)	–0.5	–0.5	–0.5	–0.5	–0.5	0.310	–0.5	–0.5
							(1.004)		
Δq_{t-1}	(?)	0.193	0.247	0.329†	0.458***		0.535*	0.899***	0.095
		(0.229)	(0.232)	(0.282)	(0.157)		(0.257)	(0.227)	(0.457)
Δq_{t-2}	(?)								
Δa_{t-1}	(?)	0.190*		0.321	0.818**		–0.168†		
		(0.095)		(0.232)	(0.298)		(0.123)		
Δa_{t-2}	(?)						–0.289**		
							(0.124)		
Δg_t	(?)	–0.585	–1.089	–1.338*	0.409	–1.287†	0.168	0.084	–0.217
		(1.082)	(1.131)	(0.715)	(0.728)	(0.919)	(0.422)	(0.429)	(0.966)
Δg^*_t	(?)	0.741**	–0.307	0.127	–0.767***	0.116	–0.191†	–1.074***	–0.512
		(0.256)	(0.894)	(0.241)	(0.195)	(0.293)	(0.131)	(0.322)	(0.561)
$\bar{R}^2$		0.94	0.69	0.87	0.81	0.20	0.79	0.91	0.73
N		14	21	20	17	14	18	17	13
LM(1)		0.360	0.026	3.030	0.001	0.879	1.143	0.349	0.012
		[0.549]	[0.871]	[0.082]	[0.972]	[0.348]	[0.285]	[0.554]	[0.913]
Iterations		—	—	—	—	—	3	—	—
Sample		1979–92	1972–92	1972–91	1973–89	1978–92	1975–92	1976–92	1978–90

Notes: See Tables 1 and 2.
[a] Philippines regression uses World Bank measure of manufacturing productivity. See text.

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 2 and 3 with the terms of trade or oil prices (either in levels or first differences). The terms of trade did not enter with statistical significance except in the case of the Philippines. In that instance, a 1% improvement in the Philippine terms of trade induces a long-run appreciation of the peso of 1.7%.[11] 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 re-estimating the error-correction models with oil prices included are reported in Table 4. Now all error-correction terms are statistically significant, although sometimes only at the 20% marginal significance level.

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.

The "Penn Effect" and Preferences

The finding that sectoral productivity differentials determine real exchange rates supports the HBS model; however, it is possible that sectoral productivity differentials just happen to be positively correlated with per capita income levels. If, as in the Linder (1961) hypothesis, rising incomes leads to a rising preference for nontradable goods, then the underlying variable of interest may have been omitted. While there *is* substantial correlation between the two variables, the correlation is *negative* in most cases, so that the cointegrating coefficient on relative income is incorrectly signed. The exceptions are Japan, Korea, and the Philippines (the latter is not significant). To further insure that the results are not driven by omitted variable bias, I augment the general error-correction specification including productivity, government spending, and oil prices with the relative income variable. In no case is the estimate on the preferences variable statistically significant, and in the correct direction. Hence, for most of these currencies, the role of relative productivity differentials in exchange rate determination is *not* as a proxy for preferences.

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, cross-section data on average real exchange rate appreciation and intercountry relative price differentials 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 these figures. In both cases, the negative relationship is apparent. For the exchange-rate/relative-price relationship, a simple OLS regression yields:

$$\begin{aligned}\Delta q &= \underset{(0.004)}{0.014} - \underset{(0.095)}{0.175} \times \Delta\tilde{p} + u_t \\ &\qquad \overline{R^2} = 0.26 \quad N = 8.\end{aligned}$$

Table 4. Error Correction Regressions: Augmented Specifications

Coefficient	*(Sgn)*	*IN*	*IN*	*JP*	*JP*	*KO*	*KO*	*PH*
Method		NLS	NLS	NLS	NLS	OLS	OLS	NLS
ECT	(–)	–0.601**	–0.361†	–0.216**	–0.209*	–0.150†	–0.130†	–0.275*
		(0.260)	(0.240)	(0.010)	(0.105)	(0.099)	(0.098)	(0.125)
$a^T - a^N$	(–)	–0.347**	–0.529*	–0.5	–0.5	–0.5	–0.5	–0.5
		(0.152)	(0.272)					
p^{oil}	(?)	0.386***	0.253	0.381†	–0.369	—	—	
		(0.096)	(0.191)	(0.275)	(0.298)			
*tot**	(+)							1.714**
								(0.732)
Δq_{t-1}	(?)	0.393†	0.263	0.401*	0.371†	0.792	0.665	
		(0.255)	(0.218)	(0.229)	(0.206)	(0.218)	(0.229)	
Δa_{t-1}	(?)	0.208*	0.222**					
		(0.105)	(0.087)					
Δp_t^{oil}	(?)			–0.210***	–0.205***	–0.106***	–0.092**[a]	
				(0.060)	(0.071)	(0.037)	(0.040)	
Δtot_{t-1}^*	(?)							0.066
								(0.215)
Δg_t			–0.657*		–0.461		–1.071†	
			(0.299)		(0.972)		(0.660)	
Δg_t^*			0.657*		0.148		0.216	
			(0.299)		(0.812)		(0.221)	
$\bar{R}^2$		0.91	0.94	0.79	0.81	0.89	0.89	0.53
N		14	14	21	21	20	20	14
LM(1)		0.242	0.002	1.248	1.302	0.716	2.576	0.064
		[0.623]	[0.967]	[0.264]	[0.254]	[0.398]	[0.108]	[0.806]
Iterations		6	6	6	6	—	—	1
Sample		1979–92	1979–92	1972–92	1972–92	1972–91	1972–91	1978–92

Notes: See Tables 1 and 2.
[a] Oil price change lagged once.

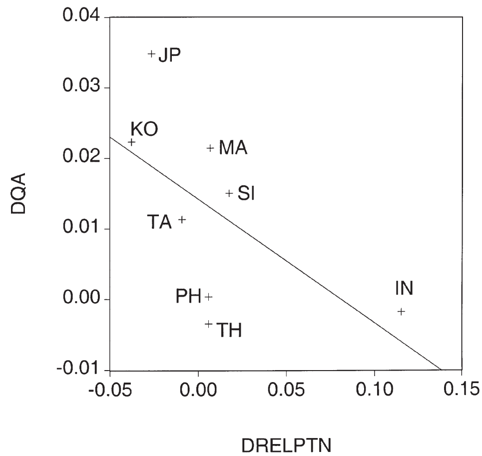

Figure 19. Cross-Section of Real Exchange Rates and Relative Prices

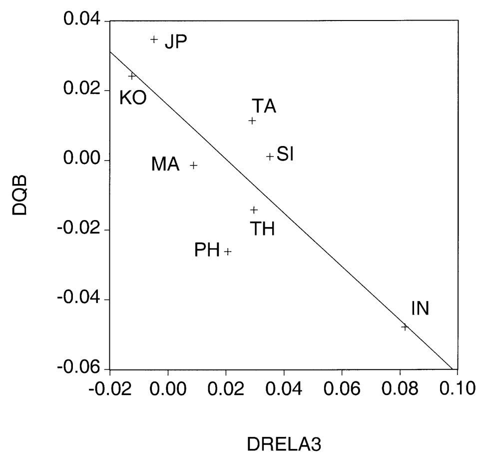

Figure 20. Cross-Section of Real Exchange Rates and Relative Productivity Differentials

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

$$\Delta q = 0.016 - 0.772 \times \Delta\tilde{a} + u_t$$
$$(0.007) \quad (0.209) \qquad \overline{R^2} = 0.64 \quad N = 8.$$

In contrast, the relative income *per capita* variable does not appear to be an important determinant of the real exchange rate. The scatterplot is shown in Figure 21. The regression estimates are:

$$\Delta q = 0.004 - 0.197 \times \Delta\tilde{y} + u_t$$
$$(0.010) \quad (0.292) \qquad \overline{R^2} = -0.08 \quad N = 8.$$

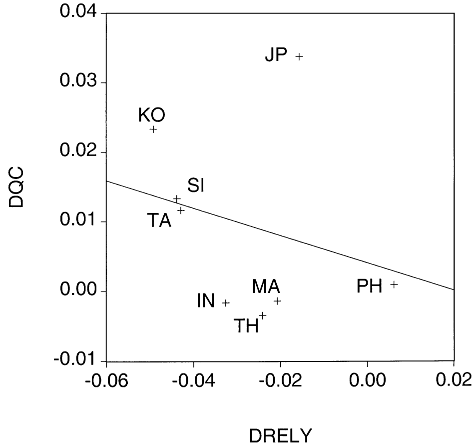

Figure 21. Cross-Section of Real Exchange Rates and Relative Incomes

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

The Relative Price of Tradables/Nontradables

In Table 5, 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 –0.091 (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 seven 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 $2\frac{1}{2}$ years. The coefficient on the relative price variable is again plausible in magnitude, and both key coefficients are highly significant.

Relative Productivity Levels

In the remaining columns of Table 5, the role of productivity and other factors in a sample *excluding* China is examined. 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.

Table 5. Panel NLS Error-Correction Model Regressions

	Full sample (1)	*No China* (2)	*No China* (3)	*No China* (4)	*No China* (5)
ECT	–0.091***	–0.243***	–0.153**	–0.146**	–0.149
	(0.030)	(0.047)	(0.057)	(0.057)	(0.057)
$p^N - p^T$	–0.679**	–0.411***			
	(0.300)	(0.093)			
$a^T - a^N$			–0.238	–0.231	–0.219
			(0.309)	(0.329)	(0.325)
p^{oil}					0.111
					(0.712)
Δq_{t-1}	0.181**	0.241***	0.257***	0.275***	0.258**
	(0.084)	(0.084)	(0.100)	(0.101)	(0.104)
Δp_{t-1}	0.169†	0.197†			
	(0.127)	(0.131)			
Δa_{t-1}			0.192***	0.190***	0.188***
			(0.067)	(0.068)	(0.068)
Δp^{oil}_t					0.095**
					(0.043)
Δg_t				–1.010	–2.924
				(0.880)	(2.206)
Δg^*_t				–0.014	–0.000
				(0.115)	(0.117)
Wald(*ECT*)	0.206	4.583			
	[0.650]	[0.032]			
Wald($p^N - p^T$)	0.632	3.588			
	[0.427]	[0.058]			
Wald($a^T - a^N$)			0.138	0.011	
			[0.710]	[0.918]	
Wald(Δg^*)				0.073	
				[0.786]	
Wald(p^{oil})					2.389
					[0.122]
Wald(Δp^{oil})					102.04
					[0.000]
$\bar{R}^2$	0.99	0.99	0.99	0.99	0.99
N	166	146	131	131	131

Notes: 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. Oil price change lagged once.

The influence of productivity may be obscured by the omission of relevant variables. In columns (4) and (5) 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 6, 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 ensure 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.

Table 6. Panel NLS Error-Correction Model Regressions on Restricted Samples

No China, Singapore, Taiwan and Thailand			
ECT	–0.166**	–0.164**	–0.167**
	(0.081)	(0.081)	(0.081)
$a^T - a^N$	–0.620†	–0.631†	–0.614†
	(0.394)	(0.407)	(0.400)
p^{oil}			0.129
			(0.769)
Δq_{t-1}	0.266**	0.260*	0.237*
	(0.133)	(0.135)	(0.139)
Δa_{t-1}	0.280***	0.279***	0.275***
	(0.095)	(0.095)	(0.096)
Δp^{oil}_t			0.100*
			(0.056)
Δg_t		–1.077	–3.198
		(1.119)	(2.756)
Δg^*_t		0.040	0.072
		(0.195)	(0.200)
Wald($a^T - a^N$)	0.001		
	[0.977]		
Wald(Δg^*)		0.004	
		[0.952]	
Wald(p^{oil})			2.204
			[0.138]
Wald(Δp^{oil})			76.460
			[0.000]
Wald(Δq)	3.136	2.732	0.056
	[0.077]	[0.098]	[0.098]
$\bar{R}^2$	0.99	0.99	0.99
N	84	84	84

Notes: 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.

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. Furthermore, I am able to go beyond previous studies' findings to conclude that a rising preference toward services is not at the heart of the secular appreciation in most East Asian real exchange rates. However, the evidence on the productivity-based model is by no means conclusive, especially as it relates to some of the countries now growing the most rapidly—China and Thailand.

Finally, 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 non-tradables, or relative productivity levels. 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. Carolan et al. (1998) have confirmed at the four-digit SITC level that the East Asian evolution of trade composition has been remarkably rapid. Hence, explanation for the long swings in this variable may require a more disaggregated approach.

Appendix: Data and Sources

Real Exchange Rates

Calculated as in equation (2) using nominal exchange rates (period average) line *rf* from IMF*, International Financial Statistics*, November 1996 CD-ROM, and for Taiwan, Bank of China, *Financial Statistics*, various issues, and aggregate GDP deflators from World Bank, *World Tables, 1995* CD-ROM.

Sectoral Price Deflators

Calculated by dividing nominal GDP by real GDP. Aggregate and sectoral nominal and real output: 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. "Relative prices" are calculated as "other" sector deflator/manufacturing sector deflator (in log terms).

Sectoral Labor Productivity

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. 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. Alternate manufacturing labor productivity measure: World Bank, *World Tables, 1995* CD-ROM. "Relative productivity" is calculated as manufacturing sector productivity/other sector productivity (in log terms).

Fiscal Variable

Log government spending to GDP ratio. Calculated using government consumption, from World Bank, *World Tables 1995* CD-ROM. Alternate log government spending variable: government consumption deflated by the service plus construction sector deflator, divided by real GDP.

Terms of Trade

Price of exports divided by price of imports, both in US$ terms. IMF, *International Financial Statistics*, CD-ROM. (*IFS* lines 74..d and 75..d, respectively).

Oil Prices

Log oil price deflated by US CPI. Oil price, *IFS* line 76aad, IMF, *International Financial Statistics*, CD-ROM.

Preferences Variable

Income per capita: GDP per capita in Summers and Heston "1985 International" dollars (the chain-weighted variable, RGDPCH). The data were drawn from the Penn World Tables 5.5, on diskette. "Relative Preferences" calculated as log(RGDPCHUS) – log(RGDPCH##).

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Notes

1. 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 (DeGregorio and Wolf, 1994; Chinn and Johnston, 1996).
2. Bergstrand (1991) tests a model that nests both the HBS and Linder (1961) hypotheses, for a set of developed and less developed countries. Accomplishing this task requires more specific data than is available on a time-series basis for these particular countries.
3. Note that this ratio omits the agricultural sector. Hence, unlike the approach adopted in Engel (1999) 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 these statistical

results hinges upon this alternative definition, although the time-series pattern for Thailand, and the trends for China, Indonesia and Korea, do change.
4. In fact, if the aggregate price index is a Cobb–Douglas function of tradable and nontradable prices, then the equation using aggregate productivity can represent only a cointegrating relation if the cointegrating coefficients are allowed to time-vary nonlinearly with a^T/a^N.
5. These categorizations are the same as those used in DeGregorio and Wolf (1994). The former did not find sensitivity of their results to the exact criterion for defining tradables versus nontradables. While it would be useful to test the sensitivity of the results to alternative categorization, the data are not reported in a form that easily allows for this.
6. The sensitivity of the empirical results to the choice of the productivity proxy can be examined in the case of Japan, where both series are available from the *ISDB*. In fact, the regressions coefficients are *more* in line with the Harrod–Balassa–Samuelson model when using the total factor productivity data. This suggests that use of the labor productivity numbers biases the results against the maintained hypothesis. See the Appendix to Chinn (1997) for detailed sensitivity analyses.
7. There are at most 22 observations. If one adjusts the critical values for the loss of degrees of freedom owing to estimating the parameters, then one cannot in general reject the null hypothesis of no cointegrating vectors. See Cheung and Lai (1993).
8. For instance, if the currency is not freely convertible, then there is no particular reason for PPP to hold even for traded goods prices. This would imply no finding of cointegration.
9. An alternative interpretation is that the Taiwanese deflators are very mismeasured. Young (1995, pp. 662–3) 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. 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 to Chinn (1997)). 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.
11. I could not include this variable for Indonesia and Taiwan, since the former does not report import prices, and the latter does not report the terms of trade in dollar terms.