# Real Exchange Rates and Productivity: Evidence From Asia

Yan, Isabel K. and Kakkar, Vikas

City University of Hong Kong

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Vikas Kakkar*and Isabel Yan †‡

Department of Economics and Finance

City University of Hong Kong

**Abstract** This paper examines a productivity-based explanation of the long run real exchange rate movements of six Asian economies. Using industry level data, we construct total factor productivities (TFPs) for the tradable and nontradable sectors. We find that (a) within each country the relative price of nontradable goods is cointegrated with the sectoral TFP differential, and (b) the real exchange rates are cointegrated with the home and foreign sectoral TFP differentials. Using the predicted real exchange rate as a measure of the "long-run equilibrium", we find that most Asian economies' real exchange rates are overvalued before the Asian Financial Crisis.

*Keywords*: Nontraded Goods, Balassa-Samuelson Model, Cointegration

*JEL Classification System*: F31, F41

*Corresponding author: Associate Professor, Department of Economics and Finance, City University of Hong Kong, Kowloon Tong, Hong Kong, E-Mail: efvikas@cityu.edu.hk; Tel: +852-34429707; Fax: +852-34420284

†Assistant Professor, Department of Economics and Finance, City University of Hong Kong, Kowloon Tong, Hong Kong, E-Mail: efyan@cityu.edu.hk; Tel: +852-34427315; Fax: +852-34420284

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# 1 Introduction

Of the several competing explanations for the persistent deviations of nominal exchange rates from their Purchasing Power Parities (PPPs), perhaps the earliest and most fundamental is the productivity differential hypothesis proposed by Balassa (1964) and Samuelson (1964). The Balassa-Samuelson hypothesis (henceforth B-S) asserts that different trends in tradable and nontradable sectors' productivity cause systematic departures of exchange rates from PPPs by changing the relative price of nontradable (to tradable) goods.[1] Since the B-S model relies on differential productivity growth rates, we would expect it to be especially relevant for determining the real exchange rates of the relatively fast growing Asian economies. However, the relatively sparse literature on Asian real exchange rates offers little support for the key predictions of the B-S model.

Ito, Isard and Symansky (1999) document a positive correlation between growth rates (relative to the U.S.) and real exchange rate appreciation for a group of East Asian economies. However, they find that the relationship between the real exchange rate and the relative price of nontradables seldom conforms to the B-S model. Chinn (1996) finds evidence of cointegration between relative prices of nontradables and real exchange rates for selected Asian economies with some exceptions. A later study by Chinn (2000) finds evidence of a cointegrating relationship between real exchange rates and labor productivity differentials for only three out of the nine Asian countries in his sample (Japan, Malaysia and the Philippines). Thomas and King (2008) extend Chinn's sample to include other Asian economies, but find similarly mixed evidence for cointegration between real exchange rates and labor productivity differentials despite including a host of other variables in their regressions.

To our understanding, studies in the extant literature have focused solely on labor

[1] Obstfeld and Rogoff (1996, Chapter 4) provides an excellent overview of the theory and evidence on the Balassa-Samuelson model.

productivity data since capital stock data for the Asian economies are generally not available. An important limitation of using labor productivity data is that one is unable to separate the impact of the supply-side effects from demand-side effects.[2] The B-S model is quintessentially about the impact of different trends in technological progress in the traded and nontraded goods sectors on the relative price of nontraded to traded goods and the real exchange rate.[3] Therefore, a priori, there is a greater likelihood of uncovering a link between real exchange rates and differential technological trends, if one exists, by using a theoretically more appropriate measure of technological progress.[4]

In this paper, we construct measures of sectoral total factor productivity (TFP) for six Asian economies (Hong Kong, Indonesia, Korea, Malaysia, Singapore, and Thailand) which are more consistent with the theory underlying the B-S model. We first construct estimates of the aggregate capital stock of each Asian economy using investment data. The gross capital stock is then allocated to the tradable and nontradable sectors in proportion to the share of capital income in that sector. The TFPs for the tradable and nontradable sectors of these economies are then computed as residuals from a Cobb-Douglas production function. The sectoral TFP data allows us to gauge the economic significance of the Balassa-Samuelson effect for the bilateral real exchange rates of these Asian countries against the U.S. dollar.

Given that most of the Asian countries in our sample had pegged their exchange

[2] For example, Drine and Rault (2002) do not find any evidence of cointegration between real exchange rates and labor productivity differentials for six Asian economies using Pedroni's (1999, 2004) panel cointegration tests. They attribute this failure to the fact that relative prices of nontradables within each country are not cointegrated with the domestic sectoral labor productivity differentials. Choudhri and Khan (2004), who use a larger panel of sixteen developing economies and similar panel cointegration methods as Drine and Rault (2002), uncover more favorable evidence for the B-S model.

[3] The real exchange rate is defined as the ratio of the domestic price level to the foreign price level multiplied by the nominal exchange rate. With this definition, deviations of nominal exchange rates from PPP are synonymous with changes in the real exchange rate.

[4] A similar point is made by De Gregorio, Giovannini and Krueger (1994) and Kakkar (2003) in the context of OECD countries.

rates to the U.S. dollar, it is also of interest to examine the implications of the productivity based model for real exchange rate misalignment prior to the Asian financial crisis. One noteworthy feature of this approach to measuring real exchange rate misalignment is that, since the real exchange rate is cointegrated with the productivity differentials, any deviation between the actual real exchange rate and its estimated equilibrium value is only temporary and will eventually vanish. This is a natural requirement for any measure of an "equilibrium" value but is not satisfied by the oft-used PPP-based measures of misalignment. Alba and Papell (2007) test for the stationarity of the U.S. dollar real exchange rates using panel unit root methods and find that they reject long-run PPP for groups of Asian and African countries. Cheung and Lai (2000) analyze 77 series of real exchange rates and they also uncover different persistence patterns between industrial countries and developing countries. Hence, it is important to allow for permanent changes in the real exchange rates of these countries when assessing real exchange rate misalignment.

We find that, with the exception of Indonesia, the real exchange rates of the other five Asian economies in our sample were overvalued in the three years prior to the financial crisis. These results are consistent with common economic intuition which suggests that overvalued currencies are likely to invite speculative attacks. They also conform to the literature on currency crises which indicates that a persistently overvalued real exchange rate is one of the key predictors of an impending currency crisis.

The rest of the paper is organized as follows. The next section describes the model and presents the two key predictions of the B-S model that are tested in this paper. Section 3 explains the data. Section 4 first presents the results of the HK-US case as a motivating example, followed by a discussion of the panel empirical results. Section 5 concludes.

# 2 Model

## 2.1 The Relative Price of Nontradables

Each country is divided into tradable and nontradable goods sectors: good $T$ is tradable and good $N$ is nontradable. The production side of the economy is summarized by the following Cobb-Douglas production functions:

$$Y_{T,it} = A_{T,it}(L_{T,it})^{\alpha_{T,i}}(K_{T,it})^{(1-\alpha_{T,i})}, \tag{1}$$

$$Y_{N,it} = A_{N,i}(L_{N,it})^{\alpha_{N,i}}(K_{N,it})^{(1-\alpha_{N,i})}. \tag{2}$$

Here $Y$ denotes output; $L$ and $K$ denote labor and capital, respectively; $A$ denotes TFP and $\alpha$ denotes the share of labor in production. Subscripts $i$ and $t$ refer to country $i$ and time $t$, respectively.

Under the standard assumptions of the B-S model[5], we have the following set of first-order conditions:

$$A_{T,it}(1-\alpha_{T,i})(k_{T,it})^{(-\alpha_{T,i})} = r_t = Q_{it}A_{N,it}(1-\alpha_{N,i})(k_{N,it})^{(-\alpha_{N,i})}, \tag{3}$$

$$A_{T,it}\alpha_{T,i}(k_{T,it})^{(1-\alpha_{T,i})} = w_{it} = Q_{it}A_{N,it}\alpha_{N,i}(k_{N,it})^{(1-\alpha_{N,i})}. \tag{4}$$

Here $r$ denotes the world real interest rate, which is determined in the world capital market; $w$ denotes the real wage rate; $k_T$ and $k_N$ denote the capital-labor ratios in the tradable and nontradable goods sectors, respectively; and $Q$ denotes the relative price of the nontradable good in terms of the tradable good. The tradable good is chosen to be the numeraire good, so that the real wage rate and the real interest rate are both measured in terms of tradables.

[5] See, for instance, Obstfeld and Rogoff (1996, Chapter 4).

Equation (3) equates the marginal product of capital in each sector to the world real interest rate in terms of tradables, whereas Equation (4) equates the marginal product of labor in each sector to the real wage rate in terms of tradables. Since each competitive firm takes as given the world real interest rate $r$, the left-hand-side equation of (3) determines the capital-labor ratio in the tradable goods sector ($k_T$). Given $k_T$, the left-hand-side equation of (4) determines the real wage rate. Given the interest rate and the wage rate, the right-hand-side equations in (3) and (4) jointly determine the relative price of nontraded-goods ($Q$) and the capital-labor ratio in the nontradable goods sector ($k_N$).

Solving for the relative price of nontradables in terms of the sectoral TFPs and the world real interest rate and taking logs yields:

$$\ln(Q_{it}) = \lambda_i + \frac{\alpha_{N,i}}{\alpha_{T,i}} \ln(A_{T,it}) - \ln(A_{N,it}) + \frac{(\alpha_{T,i} - \alpha_{N,i})}{\alpha_{T,i}} \ln(r_t). \tag{5}$$

Here $\lambda_i \equiv \frac{\alpha_{N,i}(1-\alpha_{T,i})}{\alpha_{T,i}} \ln(1-\alpha_{T,i}) - (1-\alpha_{N,i}) \ln(1-\alpha_{N,i}) + \alpha_{N,i} \ln\left(\frac{\alpha_{T,i}}{\alpha_{N,i}}\right)$ is a constant that depends on the labor shares. Equation (5) yields the first key prediction of the B-S model by showing that the relative price of nontradables within each country depends on the labor-share adjusted sectoral TFP differential and the world real interest rate in terms of tradables.

It is important to emphasize here that although we have used this stylized model for exposition, the B-S effect is quite robust to the underlying assumptions used here. For instance, Obstfeld and Rogoff (1996) show that the assumptions of two factors and internationally mobile capital can both be relaxed without changing the basic relationship between the relative price of nontradables and sectoral TFP differentials. We are not concerned here with any specific version of the model but with its main predictions which are robust to the underlying assumptions.

As shown in Section 4.2, the relative price of nontradables and sectoral TFP differentials are both nonstationary variables. Since most economic models imply the world real interest rate to be stationary, we can interpret Equation (5) as implying that $\ln(Q_{it})$ should be cointegrated with the labor-share-adjusted sectoral TFP differential $d_{it} = (\alpha_{N,i}/\alpha_{T,i})\ln(A_{T,it})$-$\ln(A_{N,it})$ with the normalized cointegrating vector $(1, -1)'$. Various versions of the following cointegrating regression are estimated to test whether this implication of the model is supported empirically:

$$\ln\left(Q_{it}\right) = \lambda_i + \delta d_{it} + \varsigma_i \ln(r_t) + \varphi^*_{it}. \tag{6}$$

Here $\varphi^*_{it}$ is a zero-mean stationary random variable that captures any short run deviation of the relative price of nontradables from its long run equilibrium value. The predicted value of the coefficient of the sectoral TFP differential, $\delta$, is 1. Since $\ln(r_t)$ is not directly observable, we treat it as a common factor. Then eqt.(6) can be written as

$$\ln\left(Q_{it}\right) = \lambda_i + \delta d_{it} + \varphi_{it} \tag{7}$$

where $\varphi_{it} = \varsigma_i F_t + \varphi^*_{it}$ with $F_t$ denotes the common factor. The presence of this common factor invalidates the conventional panel cointegration tests, such as Kao (1999) and Pedroni (1999), by inducing cross-sectional dependence in the error term. Since the asymptotic critical values are no longer valid, we apply a bootstrap methodology to the conventional panel cointegration tests to obtain the appropriate critical values.[6] We turn next to the relationship between the relative price of nontradables, sectoral productivity differentials and the real exchange rate.

[6] We are very grateful to an anonymous referee for pointing this out and for suggesting the appropriate econometric framework for this case.

## 2.2 The Real Exchange Rate

Consider a world economy with two countries. We assume that the price level of each country, $P_{it}$, can be approximated by a geometric average of the prices of nontradable and tradable goods up to a stationary measurement error:

$$P_{it} = c_i(P_{N,it})^{\beta_i}(P_{T,it})^{1-\beta_i}. \tag{8}$$

Here $\beta_i$ is the share of nontradables in the overall price level of country $i$ and $c_i$ is a stationary measurement error that reflects factors which cause the general price level to deviate from the geometric average of the price of nontradable and tradable goods. Let $E_{it}$ denote the nominal exchange rate between country $i$ (the home country) and the U.S. (the foreign country) – $E_{it}$ units of the home country's currency buy one U.S. dollar at time $t$. The real exchange rate between country $i$ and the U.S., $E^r_{it}$, is the ratio of the home price level to the U.S. price level adjusted by the nominal exchange rate:

$$E^r_{it} = \frac{P_{it}}{E_{it} \cdot P_{USt}}. \tag{9}$$

The key to developing a link between the real exchange rate and the relative price of nontradables is the law of one price for tradable goods. In the presence of transportation costs and other frictions, goods market arbitrage is not likely to be instantaneous. We therefore assume that the law of one price holds for tradable goods in the long run, so that the real exchange rate for tradable goods, $(P_{T,it}/(E_{it} \cdot P_{T,USt}))$, is stationary.

Mathematically, we can write this assumption as

$$\ln(P_{T,it}) = \ln(E_{it}) + \ln(P_{T,USt}) + u_{it}, \tag{10}$$

where u is a stationary random variable. The stationarity of u ensures that deviations from PPP for tradable goods are transitory. Equations (8)-(10) imply that

$$\ln\left(E^r_{it}\right) = \theta_i + \beta_i \ln\left(Q_{it}\right) - \beta_{US} \ln\left(Q_{USt}\right) + \epsilon_{it}, \tag{11}$$

where $\theta_i = \{E(\ln(c_i)) - E(\ln(c_{US}))\}$ is a constant and $\epsilon_{it} = u_{it} + \{\ln(c_i) - E(\ln(c_i))\} - \{\ln(c_{US}) - E(\ln(c_{US}))\}$ is a zero-mean stationary random variable. Equation (11) shows that the real exchange rate depends on the relative price of nontradables in the home and foreign countries. To highlight the connection between real exchange rates and TFP differentials, we combine equations (7) and (11) to get

$$\ln\left(E^r_{it}\right) = \mu_{1,i} + \gamma d^C_{it} + \mu_{2,i} F_t + \epsilon_{1,it}, \tag{12}$$

where $\mu_{1,i} = (\theta_i + \beta_i \lambda_i - \beta_{US}\lambda_{US})$ is a constant, $\epsilon_{1,it} = (\epsilon_{it} + \beta_i \varphi^*_{it} - \beta_{US}\varphi^*_{USt})$ is a zero-mean stationary random variable, $d^C_{it} = (\beta_i d_{it} - \beta_{US} d_{USt})$ is the composite TFP differential between the home and foreign countries[7], and $\mu_{2,i} = (\beta_i \varsigma_i - \beta_{US}\varsigma_{US})$ represents the coefficient associated with the unobservable common factor. Equation (12) is the crux of the Balassa-Samuelson model as it implies that the real exchange rate is determined solely by the relative sectoral TFP differentials in the home and foreign countries in the long run. An increase in the home sectoral TFP differential, which means faster TFP growth in the tradable sector relative to the nontradable sector, is associated with a higher relative price of nontradables via equation (7) and an appreciating real exchange rate via equation (12). The predicted magnitude of the coefficient of the composite TFP differential $\gamma$ is 1.

Equations (7) and (12) are the key testable predictions of the B-S model and form the basis of the empirical work. Since the derivation of equation (12) from

[7] We construct the composite TFP differential by estimating the share of nontradables in the overall price index using data on the price of tradables, the price of nontradables and the overall price index in equation (8).

equation (7) requires the additional assumption of long run PPP for tradable goods, the evidence for this assumption is also tested.

# 3 Data

We collected industry level data on the output, the number of work hours, and labor income for six Asian economies – Hong Kong, Singapore, S. Korea, Thailand, Indonesia and Malaysia – from 1980 to 2001. The primary databases for the Asian countries were the CEIC database and the Statistical Yearbook published by UNESCO. These were supplemented by data published by various national statistical agencies. Since capital stock data were not available for the Asian economies, they were estimated from investment data using a perpetual inventory approach, similar to that used in Kim and Lau (1995), Chow (1993) and Feenstra and Kee (2004). The gross capital stock was then allocated to the tradable and nontradable sectors in proportion to the share of capital income in the sector. For the U.S., we utilized the STAN industrial database to construct the data on tradable and nontradable output, capital stock and labor hours. The following industries were classified as tradable: manufacturing; mining and quarrying; ocean and air transport; wholesale and retail trade; and financing, insurance and business services. The following industries were classified as nontradable: electricity, gas and water; construction; real estate; community, social and personal services; land transport and communication; and restaurants. [8] Sectoral TFPs were constructed as Solow residuals (Solow 1957) from constant-price domestic

[8] Our classification is very similar to that used for OECD countries by De Grogorio, Giovannini, and Wolf (1994) and Stockman and Tesar (1995). The only major difference is that we classify financial services as tradable, whereas they classify them as nontradable. Our choice was motivated by the observation that financial services are an important component of trade for Hong Kong and Singapore. We also conducted a sensitivity analysis in which financial services were allocated to the nontradable sector. This yielded qualitatively similar results which are available upon request from the authors.

currency series of output, capital, labor shares and hours worked.[9]

# 4 Empirical Results

## 4.1 A Univariate Example

Since we have a longer time dimension than cross-sectional dimension (N=6 and T=20), we mainly rely on time series asymptotics in our analysis. For this reason, it is instructive to build some insight by viewing the results for a single pair of countries (the HK-US case) using single-equation cointegrating regressions prior to delving into the panel empirical results. In particular, we use a sieve bootstrap (for both unit root and cointegration) to compare the asymptotic and bootstrap p-values for this single economy case. [10]

Table 1 reports the results of the unit root tests, including the average ADF test proposed by Im, Pesaran, and Shin (IPS) (1995) (denoted as $IPS_{95}$), the ADF-t and LM-bar tests suggested in Im, Pesaran, and Shin (1997) ($IPS_{97}$ and $IPS_{LM}$) as well as Breitung (2000)'s test. All tests allow for heterogeneous unit root coefficients and serial correlation in the error terms. IPS (2003) shows that the small sample performance of the IPS tests are generally better than that of the Levin and Lin (LL) (1993) test if a large enough lag order is selected for the underlying ADF regressions. Breitung (2000)'s test improves on the LL and IPS tests as the latter two test statistics contain bias correction terms which may result in losses of power. Overall, we cannot reject the null hypothesis of unit root for any of the series when the bootstrap p-values

[9] Gollin (2002) argues that officially reported "employee compensation" significantly understates total labor compensation, especially for developing countries, due to a significant proportion of workers who are self-employed or employed outside the corporate sector. We attempt to adjust for this missing component of labor income, which leads to an increase in the labor shares of Hong Kong, Thailand and Indonesia. Further details are provided in Appendix A of the working paper version of this paper, Kakkar and Yan (2011), which is available from the authors upon request.

[10] The details of the bootstrap methods are provided in Appendix B of the working paper version, which is available from the authors upon request.

are used. This contrasts to the asymptotic p-values, especially for the tradable price series $\ln\left(P_T^{US} \cdot E\right)$ and $\ln(P_T)$, which are biased towards the rejection of the null.

To test for cointegration in panel data with cross-sectional dependence, we bootstrap Kao (1999)'s ADF test statistic ($ADF$), the bias-corrected Dickey-Fuller rho and $t$ test statistics ($DF_\rho^*$ and $DF_t^*$), as well as Pedroni's (1999 and 2004)'s parametric Panel $t$-statistic and parametric Group $t$-statistic ($Panel_{t_p}$ and $Gr_{t_p}$). All test statistics are for testing the null hypothesis of no cointegration. Kao's tests are based on a model which assumes homogeneous autoregressive coefficients for the residuals. Kao's bias-corrected $DF_\rho^*$ and $DF_t^*$ tests have better size and power properties than the ADF test when the long run variance is small, but the ADF test dominates the others when the variance is large. Pedroni's tests allow for considerable heterogeneity among individual members of the panel, including heterogeneity in both the long-run cointegrating vectors as well as heterogeneity in the dynamics associated with short-run deviations from these cointegrating vectors. Pedroni's panel $t$-statistic is constructed by pooling the data along the within dimension of the panel, while the group $t$-statistic is by pooling along the between dimension[11]. The parametric version of the statistics are employed as they have better performance for small samples.

Table 2 presents the results of the cointegration estimation and tests. Table 2a contains the results for testing the predicted relationship between the relative price of nontradables and the (labor-share adjusted) sectoral TFP differential. We reject the null hypothesis of no stochastic cointegration at conventional significance levels based on both the bootstrap and asymptotic versions of the Kao and Pedroni tests. The estimated coefficient is 0.9016 which is strikingly close to the predicted the value of unity. This is evidence that the relationship between the relative price of nontradables and sectoral TFP differential for HK conforms to that implied by the B-S model.

[11] The within-dimension statistics are constructed by summing both the numerator and denominator terms over the N dimension separately, whereas the between-dimension statistics are constructed by first dividing the numerator by the denominator prior to summing over the N dimension.

Table 2b shows the results of testing the assumption of long run PPP for tradable goods between HK and the US. The estimated coefficient of U.S. tradables price is 1.2625, which has the correct sign and is reasonably close to unity. The bootstrap version of the Pedroni tests and Kao's ADF tests are all significant at the 1 percent significance level. This evidence provides support for the assumption of long run PPP for tradable goods.

Table 2c contains the results of the regression of the HK-U.S. bilateral real exchange rate on the composite TFP differential between HK and the U.S. The bootstrap version of Kao's and Pedroni's cointegration test statistics reject the null hypothesis of no stochastic cointegration at the 1 percent significance level. The coefficient of the composite TFP differential is 1.0978, which again is very close to the unity value implied by the B-S model.

Overall, the results for the HK-US case suggest that the key predictions of the B-S model are broadly supported empirically.

## 4.2 Trend Properties of Data

Table 3 reports the results of the bootstrap version of the panel unit root tests for all countries. None of the tests are significant for the relative price of nontradables within each country ($\ln Q$), the sectoral (labor-share adjusted) TFP differential ($d$), the domestic tradable price ($\ln(P_T)$), the tradable goods prices on U.S. tradable goods prices adjusted for the nominal exchange rate ($\ln\left(P_T^{US}\cdot E\right)$), the real exchange rates ($\ln(E^r)$) and the composite productivity differential ($d^C$). These results are consistent with much of the empirical literature in international finance which documents that relative prices of nontradables, real exchange rates and productivity differentials are well-approximated by processes that possess stochastic trends.

## 4.3 Relative Price of Nontradables

We turn next to the evidence for the first key prediction of the model, which relates to the relationship between the relative price of nontradables within each country and the sectoral (labor-share adjusted) TFP differential. Panel A of Table 4 reports the results of Kao and Pedroni's cointegration tests applied to the residuals from OLS (with homogeneous or heterogeneous cointegrating vectors) and Mark and Sul's (2003) PDOLS. All estimations allow for the presence of fixed effects. The homogeneous cointegration vector specification is of interest since the B-S theory suggests a homogeneous cointegrating vector of $(1, -1)'$. Under the homogeneity constraint, the cointegrating coefficient estimated by OLS is 0.6399, which is close to the PDOLS estimate of 0.688. The unit value of the coefficient is plausible based on the PDOLS standard errors. Moreover, five out of six cointegration tests based on the homogeneous OLS residuals reject the null hypothesis of no cointegration at the 1 percent significance level.

When the homogeneity condition is not imposed, there is considerable variation in the individual estimates of coefficients of the sectoral TFP across countries. The coefficient of Hong Kong is 0.9, which is closest to the model's prediction, and the coefficients range from 0.13 for Korea to 0.76 for Indonesia among the other five countries. Kao's bias-corrected Dickey-Fuller rho-statistic and t-statistic as well as Pedroni's parametric panel and group $t$-statistics all reject the null of no cointegration in the relationship.[12]

Figure 1 plots the relative price of nontradables and the (labor-share-adjusted)

[12] We test the homogeneity restriction using the Wald-test proposed by Mark, Ogaki and Sul (2005). The homogeneity restriction is rejected. However, the Monte Carlo performance of these Wald tests documented by Mark, Ogaki and Sul (2005) indicates substantial size distortion in small samples. For example, with N = 5 and with T = 100, the effective (5%) size of the test is 0.23. Since T is much smaller than 100 for our dataset, the size distortion is likely to be even more severe and hence these results are not reported here.

sectoral TFP differential within each country. For HK and Indonesia, the two series move together very closely and virtually all of the medium to long-term changes in the relative price of nontradables are matched by similar changes in the sectoral TFP differentials. However, for Singapore and Malaysia comovements between relative prices and TFP differentials appear to be smaller. Overall, the visual evidence of Figure 1 appears to be consistent with the cointegration results documented above.

To summarize, the results of Table 4 provide reasonably strong confirmation of the first key prediction of the B-S model that the stochastic trend in sectoral TFP differentials can rationalize the stochastic trend in the relative price of nontradables. The null hypothesis of no cointegration between the relative price of nontradables and sectoral TFP differentials can be rejected based on most cointegration tests when the homogeneity assumption is maintained and by four out of six statistics when heterogeneity is allowed for. Moreover, the unit value of the coefficient of the sectoral TFP differential also appears to be plausible under the homogeneity restriction.

## 4.4 PPP for Tradables

Table 5 reports the results of the tests for the assumption of long run PPP for tradable goods. It is based on applying bootstrap cointegration tests to residuals obtained from various regressions of the Asian countries' tradable goods prices on U.S. tradable goods prices adjusted for the nominal exchange rate.

The point estimates of the homogeneous cointegration vector are 1.31 (homogeneous OLS) and 1.21 (PDOLS), and the unit value implied by the law of one price cannot be rejected based on PDOLS standard errors. The estimated coefficients based on heterogeneous OLS are positive for all countries except Singapore.

The upper section of Panel A reports the cointegration test results under the assumption of homogeneity implied by the law of one price. The null hypothesis of no cointegration is rejected by five out of six test statistics at conventional significance levels. The lower section of Panel A reports the results of cointegration tests when the homogeneous cointegrating vector assumption is relaxed. The null hypothesis of no cointegration is again rejected by most of the six test statistics except for Kao's Dickey-Fuller t statistic. The OLS estimates for HK, Thailand and Malaysia are 1.26, 0.95 and 0.91 respectively, which are relatively close to the predicted unit value. However, Singapore has a negative coefficient which contradicts the prediction of the PPP relationship.

Overall, the statistical evidence for PPP for tradable goods is quite supportive when the homogeneity restriction implied by the model is imposed but generally weaker under heterogeneity. However, it should be noted that aggregating micro data using CPI weights may increase the persistence of the median traded good. It is thus possible that using disaggregated data can provide more favorable evidence for PPP for tradable goods than is provided by our aggregated data.[13]

## 4.5 Real Exchange Rates

Table 6 reports the results for the second key prediction of the B-S model, which states that the bilateral real exchange rates should be cointegrated with the composite TFP differential between the home country and the U.S. Analogous to Tables 4 and 5, Panel A reports the cointegration test results while Panel B reports the estimates of the cointegrating vectors. Under the homogeneous cointegration vector assumption implied by the model, the estimated coefficients are 1.03 (homogeneous OLS) and 1.14 (PDOLS), which are remarkably close to the theoretically implied unit value.

[13] Crucini and Shintani (2008) document that the median traded good in the U.S. has a half-life of 17 months, which is significantly lower than the median nontraded good's half-life of 30 months. However, aggregating their micro data using CPI weights increases the persistence of the median traded good in the U.S. to 25 months and the median nontraded good to 50 months. This suggests that using disaggregated data may give more favorable evidence for PPP for tradable goods than using aggregated data.

Moreover, almost all cointegration tests reject the null hypothesis of no cointegration at the 1 percent significance level, except for Kao's $DF_\rho^*$.

For the heterogeneous cointegration vector case, there is considerable disparity across countries on the estimated coefficients of the composite productivity differential. The coefficient of HK (1.0978) is close to the predicted unit value but less so for other countries. Moreover, the results of the cointegration tests are rather mixed. While Pedroni's tests reject the null hypothesis of no cointegration, Kao's tests do not.[14]

To summarize the evidence for the second key prediction of the B-S model, there is strong evidence of cointegration between real exchange rates and the composite productivity differential when the assumption of homogeneous cointegrating vector is maintained. Moreover, the coefficient estimates are very close to the unit value implied by the model. However, there is less accord for the heterogeneous case.

### 4.6 Real Exchange Rate Misalignment

As mentioned in the introduction, a natural by-product of the productivity-based model is that it provides one with an estimate of the "long-run equilibrium real exchange rate" of the Asian real exchange rates against the U.S. dollar. Figure 2 plots the real exchange rates of the Asian countries against the U.S. dollar and the estimated long run equilibrium values based on the PDOLS cointegrating vector estimates reported in Panel B of Table 6. The first panel shows the results for Hong Kong. The actual real exchange rate moves quite closely together with the implied equilibrium value predicted by the model, although there is a modest undervaluation in the early 1990's and a modest overvaluation from 1993 onwards.

[14] The Wald test for the homogeneity restriction rejects the null hypothesis that the coefficients are identical across countries. However, as noted earlier, this test suffers from severe size distortion for small T.

The second panel of Figure 2 shows the results for Singapore. The B-S model predicts a sustained real depreciation of the Singapore dollar and it misses some of the big swings in the actual real exchange rate. These results are consistent with the earlier evidence suggesting that the basic ingredients of the B-S model – namely the PPP for tradables and the close relationship between the real exchange rate and the composite TFP differentials – appear not to hold for Singapore.

The third panel shows the real exchange rate and the fitted value for Korea. The model captures the major turning points of the actual real exchange rate, although it underestimates the volatility of the real exchange rate. The real exchange rate appears substantially overvalued in the years preceding the Asian financial crisis.

The fourth panel contains the results for Thailand. The model predicts a slight depreciation of the real exchange rate over the entire sample. However, the actual real exchange rate undergoes a continuous appreciation from the mid-1980's up to 1995, followed by a massive depreciation.

The fifth panel shows the actual and fitted real exchange rates for Indonesia. The model captures the secular depreciation of the real exchange rate over the entire sample quite well. In sharp contrast to the other countries, the real exchange rate appears to be undervalued in the years prior to the crisis.

The last panel shows the actual and fitted real exchange rates for Malaysia. The real exchange rate fluctuates around its long-run equilibrium value, exhibiting an undervaluation in the late 1980's and an overvaluation in the 1990's prior to the crisis.

Table 7 shows the estimated average overvaluation during the three year period prior to the crisis (1994 through 1996) and also at the end of 1996. At the eve of the crisis in 1996 all countries except Indonesia show overvalued real exchange rates, with Hong Kong being the least overvalued at 3.54% and Singapore the most overvalued at almost 26%. Korea and Malaysia also appear to be significantly overvalued, with the extent of overvaluation ranging between 14% to 16%. Figure 2 also shows that for all the countries except Indonesia, the real exchange rate overvaluation reached a peak near 1995 and then the downward adjustment towards equilibrium commenced. However, by 1996 panic had set in the region and the speculators were likely expecting large further declines. They therefore behaved in a way that resulted in the declines they were expecting. Hence the real and nominal exchange rates depreciated significantly more than the required adjustment indicated by the productivity based model. For instance, the real exchange rates of Korea, Thailand and Indonesia had depreciated below its implied equilibrium value by 1997.

Viewed through the lens of the B-S model, it therefore seems plausible that both fundamental factors and self-fulfilling expectations had a role to play in the Asian financial crisis. The productivity based fundamental factors indicate large and persistent overvaluations in the few years prior to the crisis.

# 5 Conclusions

This paper examined the evidence for a productivity-based explanation of the long run real exchange rate movements for six Asian economies in the context of the Balassa-Samuelson model. Relative to earlier studies, which are at best only weakly supportive of the Balassa-Samuelson effect, we find that sectoral TFP differentials play an important role in explaining the long term trends in both the relative price of nontradables and the real exchange rates of these Asian countries.

These results are consistent with the view espoused in recent research that real exchange rates possess both permanent and temporary components. For instance, Mark and Choi (1997) show that models in which the long-run real exchange rate is identified as the permanent component of the real exchange rate outperform models which assume long-run PPP holds in terms of out-of-sample forecasts. Engel (2000) finds that the real exchange rate contains an economically significant component associated with the relative price of nontraded goods. In conjunction with recent work that emphasizes the importance of nontradable goods in explaining long-run real exchange rate movements (e.g. Burstein, Eichenbaum and Rebelo 2005a, Burstein, Eichenbaum and Rebelo 2005b, Betts and Kehoe 2006, Crucini and Shintani 2008, Kakkar and Ogaki 1999, and Park and Ogaki 2007), these results suggest that productivity differentials may be an important factor in explaining the persistent departures of nominal exchange rates of these Asian countries from their purchasing power parities.

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Table 1: Hong Kong: Unit Root Tests of Im, Pesaran and Shin (1995, 1997) and Breitung (2000)

| | Im, Pesaran and Shin (1995, 1997)[d] | | | | | | Breitung[d] |
|---|---|---|---|---|---|---|---|
| | $IPS_{95}$ | $IPS_{95}^{trend}$ | $IPS_{97}$ | $IPS_{97}^{trend}$ | $IPS_{LM}$ | $IPS_{LM}^{trend}$ | (2000) |
| $\ln Q$ [b] | 0.8744 | 0.3551 | 0.8722 | 0.3688 | -1.0057 | -1.5523 | -0.4530 |
| bootstrap | (0.8300)[a] | (0.5950) | (0.8300) | (0.5950) | (0.1650) | (0.1650) | (0.5550) |
| asymptotic | (0.1909)[a] | (0.3612) | (0.1916) | (0.3561) | (0.1573) | (0.0603)* | (0.3253) |
| $d$ | 0.8626 | 0.4829 | 0.8602 | 0.4995 | -0.9979 | -1.5458 | -0.5264 |
| bootstrap | (0.7560) | (0.6690) | (0.7560) | (0.6690) | (0.2270) | (0.2270) | (0.5390) |
| asymptotic | (0.1942) | (0.3145) | (0.1948) | (0.3087) | (0.1592) | (0.0611)* | (0.2993) |
| $\ln\left(P_T^{US}\cdot E\right)^c$ | -1.5263 | -0.2133 | -1.5727 | -0.2124 | 1.7446 | 0.7661 | 1.2305 |
| bootstrap | (0.7340) | (0.8140) | (0.7340) | (0.8140) | (0.2660) | (0.2660) | (0.7790) |
| asymptotic | (0.0635)* | (0.4155) | (0.0578)* | (0.4158) | (0.0405)** | (0.2218) | (0.1092) |
| $\ln(P_T)$ | -1.5637 | 1.3803 | -1.6108 | 1.4171 | 1.7914 | 0.8056 | 2.3871 |
| bootstrap | (0.4050) | (0.7850) | (0.4050) | (0.7850) | (0.5950) | (0.5950) | (1.0000) |
| asymptotic | (0.0589)* | (0.0837)* | (0.0536)* | (0.0782)* | (0.0366)** | (0.2102) | (0.008)*** |
| $\ln(E^r)^c$ | 0.1353 | 0.9503 | 0.1194 | 0.9774 | -0.3445 | -0.9950 | 0.0866 |
| bootstrap | (0.6120) | (0.8500) | (0.6120) | (0.8500) | (0.3860) | (0.3860) | (1.0000) |
| asymptotic | (0.4462) | (0.1710) | (0.4526) | (0.1642) | (0.3652) | (0.1599) | (0.4655) |
| $d^C$ | 0.3757 | 0.8076 | 0.3643 | 0.8314 | -0.5949 | -1.2061 | -0.1620 |
| bootstrap | (0.6290) | (0.7830) | (0.6290) | (0.7830) | (0.3600) | (0.3600) | (0.5980) |
| asymptotic | (0.3536) | (0.2097) | (0.3578) | (0.2029) | (0.2759) | (0.1139) | (0.4356) |

Notes: [a] P-values are in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level respectively.

[b] $\ln Q$ stands for the log relative nontradable price. $d$ refers to the labor-share-adjusted sectoral TFP differential. $\ln\left(P_T^{US}\cdot E\right)$ refers to the log of the US tradable price times the nominal exchange rate. $\ln(P_T)$ refers to the home tradable price. $\ln(E^r)$ denotes the log real exchange rate, and $d^C$ denotes the composite TFP differential between the home and foreign countries.

[c] An Asian-crisis dummy is included to allow for a possible break in the nominal and real exchange Rate. The dummy equals 1 from 1997 onwards.

[d] **IPS$_{95}$** refers to the average ADF test proposed by Im, Pesaran, and Shin (1995). IPS considers the case that error terms are serially correlated.

**IPS$_{97}$ and IPS$_{LM}$** are the ADF t and LM-bar tests suggested in Im, Pesaran, and Shin (1997), respectively. The $IPS_{LM}$ statistics reported here are those that allow for serial correlation.

All IPS tests allow for heterogeneous unit root coefficients. The test statistics with superscript "trend" are performed on detrended data.

**Breitung (2000)** found the losses of power due to the bias correction terms in Levin and Lin (1993) and detrending bias in Im, Pesaran, and Shin (1997). Therefore, he suggested a new test without bias corrections. Breitung's test assumes homogeneous unit root coefficient.

Table 2a: Hong Kong — Kao's (1999) and Pedroni's (1999) Cointegration Tests on the Regression of the Relative Price of Nontradables on the Sectoral TFP Differentials

$$\ln Q_{HKt} = \alpha + \delta d_{HKt} + \epsilon_t$$

| Panel A: Cointegration Tests with OLS Estimation of the Cointegrating Vector | | | | | | |
|---|---|---|---|---|---|---|
| | Kao's Tests[a] | | | | Pedroni's Tests[b] | |
| | $DF_\rho^*$ | $DF_t^*$ | $ADF$ (1 lag) | $ADF$ (2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| | -3.0260 | -2.1379 | -3.3317 | -3.5643 | -2.6171 | -2.7374 |
| bootstrap | (0.008)[c]*** | (0.000)*** | (0.000)*** | (0.000)*** | (0.000)*** | (0.000)*** |
| asymptotic | (0.001)*** | (0.016)** | (0.000)*** | (0.000)*** | (0.004)*** | (0.003)*** |

| Panel B: OLS Estimation of the Cointegrating Vector | |
|---|---|
| | $\widehat{\delta}^{OLS}$ |
| Coefficient | 0.9016[d] |

Table 2b: Hong Kong — Kao's (1999) and Pedroni's (1999) Cointegration Tests on the Regression of the PPP for Tradable Goods

$$\ln P_{T,HKt} = \alpha' + \varphi D_{97,t} + \theta \ln \left(P_{T,USt} E_{HKt}\right) + \epsilon_t' \; ^{e}$$

| Panel A: Cointegration Tests with OLS Estimation of the Cointegrating Vector | | | | | | |
|---|---|---|---|---|---|---|
| | Kao's Tests[a] | | | | Pedroni's Tests[b] | |
| | $DF_\rho^*$ | $DF_t^*$ | $ADF$ (1 lag) | $ADF$ (2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| | -3.2129 | -1.9079 | -0.9676 | -0.9509 | -0.1938 | 0.1392 |
| bootstrap | (0.602)[c] | (0.159) | (0.000)*** | (0.000)*** | (0.000)*** | (0.000)*** |
| asymptotic | (0.001)*** | (0.028)** | (0.1666) | (0.1708) | (0.4232) | (0.5554) |

| Panel B: OLS Estimation of the Cointegrating Vector | |
|---|---|
| | $\widehat{\theta}^{OLS}$ |
| Coefficient | 1.2625[d] |

Table 2c: Hong Kong — Kao's (1999) and Pedroni's (1999) Cointegration Tests on the Regressions of the Real Exchange Rate on the Composite TFP Differentials

$$\ln E_{HKt}^r = \alpha'' + \varphi'' D_{97,t} + \gamma d_{HKt}^C + \epsilon_t'' \; ^{e}$$

| Panel A: Cointegration Tests with OLS Estimation of the Cointegrating Vector | | | | | | |
|---|---|---|---|---|---|---|
| | Kao's Tests[a] | | | | Pedroni's Tests[b] | |
| | $DF_\rho^*$ | $DF_t^*$ | $ADF$ (1 lag) | $ADF$ (2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| | -3.2435 | -2.4852 | -2.5899 | -3.6291 | -1.8893 | -1.8735 |
| bootstrap | (0.7880)[c] | (0.3480) | (0.000)*** | (0.000)*** | (0.000)*** | (0.000)*** |
| asymptotic | (0.000)*** | (0.006)*** | (0.005)*** | (0.000)*** | (0.0294)** | (0.031)** |

| Panel B: OLS Estimation of the Cointegrating Vector | |
|---|---|
| | $\widehat{\gamma}^{OLS}$ |
| Coefficient | 1.0978[d] |

[a] $DF_\rho^*$ and $DF_t^*$ denote the bias-corrected Dickey-Fuller rho and t statistics of Kao (1999) respectively.
[b] $Panel_{t_p}$ and $Gr_{t_p}$ denote Pedroni's (1999 and 2004) parametric panel $t$-statistic and parametric group $t$-statistic, respectively. The number of lags for each cross section is calculated according to the Akaike Information Criterion or Bayesian Information Criterion (AIC/BIC). The length of kernel window is calculated a la Andrews or Newey-West. For the $Panel_{t_p}$ test, we use the estimate of the long-run variance.
[c] P-values are in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level respectively.
[d] Since the OLS standard errors are not valid for conducting inference, we do not report them here.
[e] An Asian-crisis dummy $D_{97}$ is included to allow for a possible break in the nominal exchange rate. The dummy equals 1 from 1997 onwards.

Table 3: Panel Unit Root Tests of Im, Pesaran and Shin (1995, 1997) and Breitung (2000)

| | Im, Pesaran and Shin (1995, 1997)[d] | | | | | | Breitung[d] |
|---|---|---|---|---|---|---|---|
| | $IPS_{95}$ | $IPS_{95}^{trend}$ | $IPS_{97}$ | $IPS_{97}^{trend}$ | $IPS_{LM}$ | $IPS_{LM}^{trend}$ | (2000) |
| $\ln Q$ [b] | -0.6795 | 1.1762 | -0.7369 | 1.2166 | -2.8261 | -4.1084 | -0.4039 |
| | (0.9600)[a] | (0.9620) | (0.9600) | (0.9620) | (0.4390) | (0.4390) | (0.9940) |
| $d$ | -0.1953 | -0.8073 | -0.2438 | -0.8115 | -2.4916 | -3.8264 | -0.8943 |
| | (0.5400) | (0.1570) | (0.5400) | (0.1570) | (0.5940) | (0.5940) | (1.0000) |
| $\ln\left(P_{T,US} \cdot E\right)^c$ | -2.4585 | 0.4158 | -2.5486 | 0.4392 | -1.6133 | -3.0856 | -2.7027 |
| | (0.7120) | (0.8050) | (0.7120) | (0.8050) | (0.4560) | (0.456) | (0.9410) |
| $\ln(P_T)$ | 0.6236 | 2.8670 | 0.5902 | 2.9455 | -2.6584 | -3.9671 | 3.7599 |
| | (0.6600) | (0.9200) | (0.6600) | (0.9200) | (0.2590) | (0.2590) | (1.0000) |
| $\ln(E^r)^c$ | -0.8566 | 1.5906 | -0.9173 | 1.6404 | -2.9624 | -4.2234 | -0.5328 |
| | (0.6870) | (0.9520) | (0.6870) | (0.9520) | (0.2410) | (0.2410) | (0.030)** |
| $d^C$ | 2.8903 | 1.2485 | 2.8987 | 1.2905 | -2.1066 | -3.5019 | -0.1939 |
| | (0.8110) | (0.7990) | (0.8110) | (0.7990) | (0.3430) | (0.3430) | (1.0000) |

Notes: [a] Bootstrap p-values are in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level respectively.

[b] $\ln Q$ stands for the log relative nontradable price. $d$ refers to the labor-share-adjusted sectoral TFP differential. $\ln\left(P_{T,US} \cdot E\right)$ refers to the log of the US tradable price times the nominal exchange rate. $\ln(P_T)$ refers to the home tradable price. $\ln(E^r)$ denotes the log real exchange rate, and $d^C$ denotes the composite TFP differential between the home and foreign countries.

[c] An Asian-crisis dummy is included to allow for a possible break in the nominal and real exchange rate. The dummy equals 1 from 1997 onwards.

[d] **$IPS_{95}$** refers to the average ADF test proposed by Im, Pesaran, and Shin (1995). IPS allows for a heterogeneous coefficient of $y_{i,t-1}$ and considers the case that error terms are serially correlated with different serial correlation coefficients across cross-sectional units.

**$IPS_{97}$ and $IPS_{LM}$** are the ADF t and LM-bar tests suggested in Im, Pesaran, and Shin (1997), respectively. The $IPS_{LM}$ statistics reported here are those that allow for serial correlation.

All IPS tests allow for heterogeneous unit root coefficients. The test statistics with superscript "trend" are performed on detrended data.

**Breitung (2000)** found the losses of power due to the bias correction terms in Levin and Lin (1993) and detrending bias in Im, Pesaran, and Shin (1997). Therefore, he suggested a new test without bias corrections. Breitung's test assumes homogeneous unit root coefficient.

Table 4: Kao's (1999) and Pedroni's (1999) Cointegration Tests on the Regression of the Relative Price of Nontradables on the Sectoral TFP Differentials

Panel A: Cointegration Tests

| | Based on OLS Estimation with Homogeneous Cointegrating Vector[e] $\ln Q_{it} = \alpha_i + \delta d_{it} + \epsilon_{it}$ | | | | | |
|---|---|---|---|---|---|---|
| | Kao's Tests[a] | | | | Pedroni's Tests[b] | |
| | $DF^*_\rho$ | $DF^*_t$ | $ADF$(1 lag) | $ADF$(2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| statistic | -4.3765 | -6.2549 | -6.0048 | -3.2566 | 0.3553 | -1.3632 |
| p-value | (0.000)***[c] | (0.106) | (0.000)*** | (0.000)*** | (0.003)*** | (0.000)*** |
| | Based on OLS Estimation with Heterogeneous Cointegrating Vector[e] $\ln Q_{it} = \alpha_i + \delta_i d_{it} + \epsilon_{it}$ | | | | | |
| | Kao's Tests | | | | Pedroni's Tests | |
| | $DF^*_\rho$ | $DF^*_t$ | $ADF$ (1 lag) | $ADF$ (2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| statistic | -18.4535 | -9.2445 | -5.9251 | -1.9814 | -2.3103 | -2.9531 |
| p-value | (0.006)*** | (0.000)*** | (0.451) | (0.999) | (0.000)*** | (0.000)*** |

Panel B: Estimation of the Cointegrating Vector

| | OLS with Homogeneous Cointegrating Vector[e] $\ln Q_{it} = \alpha_i + \delta d_{it} + \epsilon_{it}$ | | | | | |
|---|---|---|---|---|---|---|
| $\widehat{\delta}^{OLS}$ | 0.6399[d] | | | | | |
| | OLS with Heterogeneous Cointegrating Vector[e] $\ln Q_{it} = \alpha_i + \delta_i d_{it} + \epsilon_{it}$ | | | | | |
| | $HK^f$ | $SIN$ | $KOR$ | $THA$ | $IND$ | $MAL$ |
| $\widehat{\delta}_i^{OLS}$ | 0.9016[d] | 0.4368 | 0.1303 | 0.3159 | 0.7569 | 0.5457 |
| | Mark and Sul (2003)'s PDOLS $\ln Q_{it} = \alpha_i + \delta d_{it} + \epsilon_{it}$ | | | | | |
| $\widehat{\delta}^{PDOLS}$ | 0.688 | | | | | |
| S.E. | (parametric s.e.: 0.235, Andrews s.e.: 0.196) | | | | | |

[a] $DF^*_\rho$ and $DF^*_t$ denote the bias-corrected Dickey-Fuller rho and t statistics of Kao (1999) respectively.
[b] $Panel_{t_p}$ and $Gr_{t_p}$ denote Pedroni's (1999 and 2004) parametric panel $t$-statistic and parametric group $t$-statistic, respectively.
[c] Bootstrap p-values are in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level respectively.
[d] Since the OLS standard errors are not valid for conducting inference, we do not report them here.
[e] The cointegrating vectors are estimated using OLS with country-specific fixed effects.
[f] "HK", "SIN", "KOR", "THA", "IND" and "MAL" refer to Hong Kong, Singapore, S. Korea, Thailand, Indonesia and Malaysia respectively.

Table 5: Kao's (1999) and Pedroni's (1999) Cointegration Tests on the Regression of the PPP for Tradable Goods

Panel A: Cointegration Tests

| | Based on OLS Estimation with Homogeneous Cointegrating Vector[e] $\ln P_{T,it} = \alpha_i' + \varphi D_{97,it} + \theta \ln \left(P_{T,USt} E_{it}\right) + \epsilon_{it}'$ [g] | | | | | |
|---|---|---|---|---|---|---|
| | Kao's Tests[a] | | | | Pedroni's Tests[b] | |
| | $DF_\rho^*$ | $DF_t^*$ | $ADF$ (1 lag) | $ADF$ (2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| statistic | -11.556 | -8.9464 | -5.3346 | -5.0509 | 0.1302 | 0.7766 |
| p-value | (1.000)[c] | (0.017)** | (0.000)*** | (0.000)*** | (0.009)*** | (0.000)*** |
| | Based on OLS Estimation with Heterogeneous Cointegrating Vector[e] $\ln P_{T,it} = \alpha_i' + \varphi_i D_{97,it} + \theta_i \ln \left(P_{T,USt} E_{it}\right) + \epsilon_{it}'$ | | | | | |
| | Kao's Tests | | | | Pedroni's Tests | |
| | $DF_\rho^*$ | $DF_t^*$ | $ADF$ (1 lag) | $ADF$ (2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| statistic | -15.424 | -9.1361 | -5.1827 | -3.6028 | -1.7039 | -0.4768 |
| p-value | (0.050)** | (0.978) | (0.050)** | (0.021)** | (0.062)* | (0.000)*** |

Panel B: Estimation of the Cointegrating Vector

| | OLS with Homogeneous Cointegrating Vector[e] $\ln P_{T,it} = \alpha_i' + \varphi D_{97,it} + \theta \ln \left(P_{T,USt} E_{it}\right) + \epsilon_{it}'$ | | | | | |
|---|---|---|---|---|---|---|
| $\widehat{\theta}^{OLS}$ | 1.3121[d] | | | | | |
| | OLS with Heterogeneous Cointegrating Vector[e] $\ln P_{T,it} = \alpha_i' + \varphi_i D_{97,it} + \theta_i \ln \left(P_{T,USt} E_{it}\right) + \epsilon_{it}'$ | | | | | |
| | $HK$[f] | $SIN$ | $KOR$ | $THA$ | $IND$ | $MAL$ |
| $\widehat{\theta}_i^{OLS}$ | 1.2625[d] | -0.1427 | 0.6589 | 0.9530 | 1.9459 | 0.9102 |
| | Mark and Sul (2003)'s PDOLS $\ln P_{T,it} = \alpha_i' + \varphi D_{97,it} + \theta \ln \left(P_{T,USt} E_{it}\right) + \epsilon_{it}'$ | | | | | |
| $\widehat{\theta}^{PDOLS}$ | 1.213 | | | | | |
| S.E. | (parametric s.e.: 0.364, Andrews s.e.: 0.207) | | | | | |

[a] $DF_\rho^*$ and $DF_t^*$ denote the bias-corrected Dickey-Fuller rho and t statistics of Kao (1999) respectively.
[b] $Panel_{t_p}$ and $Gr_{t_p}$ denote Pedroni's (1999 and 2004) parametric panel $t$-statistic and parametric group $t$-statistic, respectively.
[c] Bootstrap p-values are in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level respectively.
[d] Since the OLS standard errors are not valid for conducting inference, we do not report them here.
[e] The cointegrating vectors are estimated using OLS with country-specific fixed effects.
[f] "HK", "SIN", "KOR", "THA", "IND" and "MAL" refer to Hong Kong, Singapore, S. Korea, Thailand, Indonesia and Malaysia respectively.
[g] An Asian-crisis dummy $D_{97}$ is included to allow for a possible break in the nominal exchange rate. The dummy equals 1 from 1997 onwards.

Table 6: Kao's (1999) and Pedroni's (1999) Cointegration Tests on the Regression of the Real Exchange Rate on the Composite TFP Differentials

Panel A: Cointegration Tests

| | Based on OLS Estimation with Homogeneous Cointegrating Vector[e] $\ln E^r_{it} = \alpha''_i + \varphi'' D_{97,it} + \gamma d^C_{it} + \epsilon''_{it}$ [g] | | | | | |
|---|---|---|---|---|---|---|
| | Kao's Tests[a] | | | | Pedroni's Tests[b] | |
| | $DF^*_\rho$ | $DF^*_t$ | $ADF$ (1 lag) | $ADF$ (2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| statistic | -9.8236 | -8.6873 | -5.0046 | -6.9259 | -0.1493 | -0.0276 |
| p-value | (1.000) | (0.001)*** | (0.000)*** | (0.000)*** | (0.000)*** | (0.000)*** |

| | Based on OLS Estimation with Heterogeneous Cointegrating Vector[e] $\ln E^r_{it} = \alpha''_i + \varphi''_i D_{97,it} + \gamma_i d^C_{it} + \epsilon''_{it}$ | | | | | |
|---|---|---|---|---|---|---|
| | Kao's Tests | | | | Pedroni's Tests | |
| | $DF^*_\rho$ | $DF^*_t$ | $ADF$ (1 lag) | $ADF$ (2 lags) | $Panel_{t_p}$ | $Gr_{t_p}$ |
| statistic | -13.6931 | -8.9268 | -4.8115 | -4.7265 | -1.5939 | -1.5522 |
| p-value | (0.466) | (0.486) | (0.827) | (0.178) | (0.020)*** | (0.000)*** |

Panel B: Estimation of the Cointegrating Vector

| | OLS with Homogeneous Cointegrating Vector[e] $\ln E^r_{it} = \alpha''_i + \varphi'' D_{97,it} + \gamma d^C_{it} + \epsilon''_{it}$ |
|---|---|
| $\widehat{\gamma}^{OLS}$ | 1.0296[d] |

| | OLS with Heterogeneous Cointegrating Vector[e] $\ln E^r_{it} = \alpha''_i + \varphi''_i D_{97,it} + \gamma_i d^C_{it} + \epsilon''_{it}$ | | | | | |
|---|---|---|---|---|---|---|
| | $HK^f$ | $SIN$ | $KOR$ | $THA$ | $IND$ | $MAL$ |
| $\widehat{\gamma}_i^{OLS}$ | 1.0978[d] | -0.6896 | -0.9880 | 2.0016 | 4.7100 | 0.2750 |

| | Mark and Sul (2003)'s PDOLS $\ln E^r_{it} = \alpha''_i + \varphi'' D_{97,it} + \gamma d^C_{it} + \epsilon''_{it}$ |
|---|---|
| $\widehat{\gamma}^{PDOLS}$ | 1.144 |
| S.E. | (parametric s.e.: 0.390, Andrews s.e.: 0.236) |

[a] $DF^*_\rho$ and $DF^*_t$ denote the bias-corrected Dickey-Fuller rho and t statistics of Kao (1999) respectively.
[b] $Panel_{t_p}$ and $Gr_{t_p}$ denote Pedroni's (1999 and 2004) parametric panel $t$-statistic and parametric group $t$-statistic, respectively.
[c] Bootstrap p-values are in parentheses. *, ** and *** denote significance at the 10%, 5% and 1% level respectively.
[d] Since the OLS standard errors are not valid for conducting inference, we do not report them here.
[e] The cointegrating vectors are estimated using OLS with country-specific fixed effects.
[f] "HK", "SIN", "KOR", "THA", "IND" and "MAL" refer to Hong Kong, Singapore, S. Korea, Thailand, Indonesia and Malaysia respectively.
[g] An Asian-crisis dummy $D_{97}$ is included to allow for a possible break in the nominal exchange rate. The dummy equals 1 from 1997 onwards.

Table 7
Estimated Real Exchange Rate Misalignment

| Country | Average Overvaluation during 1994-96 | Overvaluation in 1996 |
|---|---|---|
| Hong Kong | 3.79% | 3.54% |
| Singapore | 24.92% | 25.92% |
| S. Korea | 19.16% | 14.10% |
| Thailand | 11.70% | 5.26% |
| Indonesia | -6.00% | -6.50% |
| Malaysia | 17.12% | 16.22% |

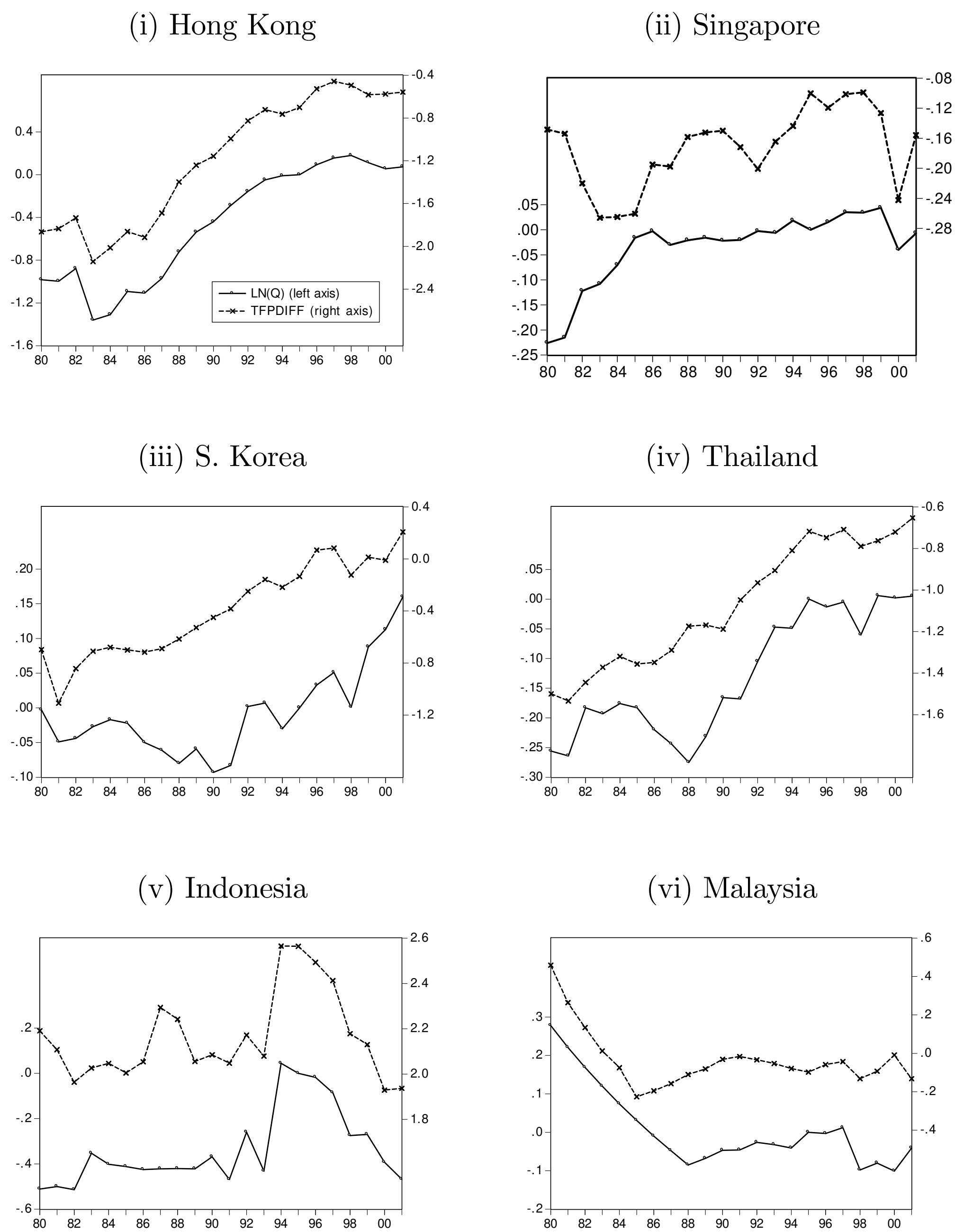


Figure 1: Plot of the (log) relative nontradable price of nontradables and the (log) labor share adjusted TFP-differentials.

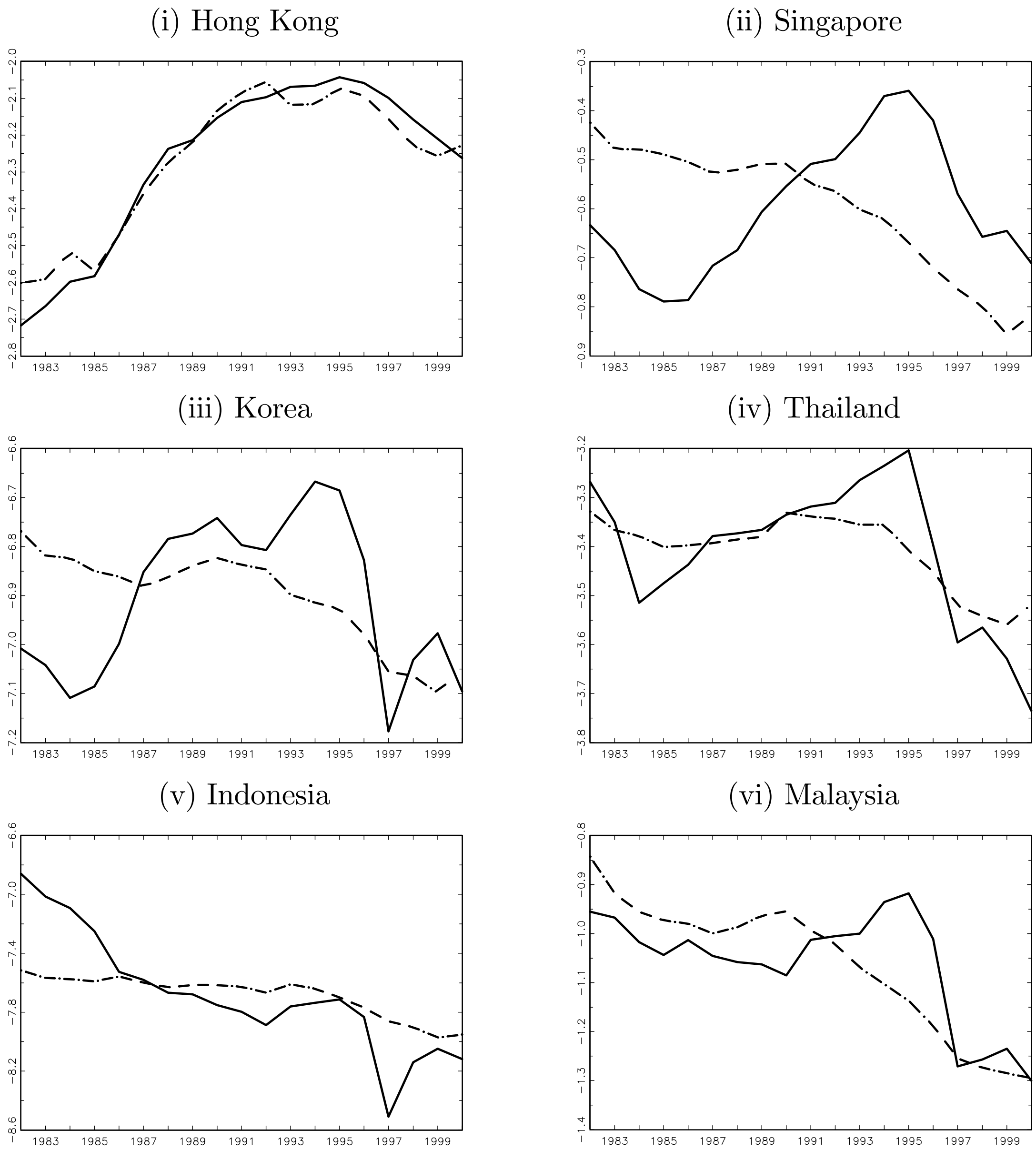


Figure 2: Plot of the actual and predicted Asian real exchange rates against the U.S. dollar based on the panel dynamic OLS estimates with one lag. The solid line is the observed real exchange rate and the dashed line is the predicted value based on the B-S model.