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6 Private Consumption, Nontraded Goods, and Real Exchange Rate: Evidence from South Korea and Taiwan

Kenneth S. Lin

6.1 Introduction

There is little empirical evidence from the production perspective that any known fundamentals have reliable effects on real exchange rates.[1] According to Balassa (1964) and Samuelson (1964), real exchange rate movements reflect cross-country differences in the productivity differential between the traded and nontraded sectors. If higher productivity growth is expected to occur in the traded sector, there is a positive relation between real exchange rates and cross-country disparities in productivity growth.[2] Even though productivity differentials can account for long-run real exchange rate movements, a much higher productivity growth rate in the traded sector is required to justify the long-run movement.

This paper presents an empirical study of long-run real exchange rate movements from the consumption perspective. In most industrial countries, private consumption and the real exchange rate both have clear trends but exhibit different fluctuations. If the real exchange rate (or the relative price of nontraded goods) varies over time, aggregate consumption will respond to those price changes. Here I emphasize the role of risk aversion for nontraded goods consumption in accounting for long-run real exchange rate movements. When risk

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1. Examples include Adler and Lehmann (1983), Hsieh (1982), Huizinga (1987), Ito, Isard, and Symansky (chap. 4 in this volume), Kravis and Lipsey (1987), and Strauss (1996).

2. The real exchange rate has been a natural indicator of export competitiveness. Establishing the positive relation and underlying growth mechanism has become a central research topic in economic development (e.g., Ito et al., chap. 4 in this volume).

aversion is the inverse of intertemporal elasticity of substitution, the lower risk aversion is, the easier it is for private agents, in terms of utility, to forgo current consumption for future consumption, and thus the higher the consumption growth rate is. On the other hand, lower risk aversion decreases the value of diversification. Suppose that agents are more risk averse for nontraded goods than for traded goods. Even in a perfect international credit market, agents cannot fully diversify away preference and productivity shocks to nontraded goods through consumption smoothing. Those shocks could induce changes in real exchange rate movements. As a result, relatively higher risk aversion for nontraded goods implies a tighter relationship in trend properties between nontraded goods consumption and the real exchange rate.

Volatile and persistent movements of real exchange rates and small cross-country correlations of private consumption have been separate research topics in international macroeconomics.[3] However, few researchers have attempted to account for the comovement between private consumption and the real exchange rate. One exception is Backus and Smith (1993). They studied a dynamic exchange economy with one traded good, one nontraded good for each country, and an arbitrary number of countries. A main theoretical finding is that the private consumption ratio between the foreign country and the home country and the real exchange rate have similar fluctuations and are positively correlated over time. However, they found little evidence for the positive correlation in eight OECD countries. There are two possibilities for the discrepancy between theory and evidence. First, preference shocks are not admitted in their model. When endowment shock is the sole external shock, it can only generate positive correlation between changes in the consumption ratio and changes in the real exchange rate. Second, agents have identical preferences across countries.

In this paper, I adopt Ogaki and Park's (1989) cointegration-Euler equation approach. Given the assumption of stationary preference shocks, my model implies that the real exchange rate and private consumption in different countries have similar trend properties in the sense that they are cointegrated. Here preference shocks not only induce negative correlation between the real exchange rate and consumption in different countries but also provide an identifying assumption. Preference parameters and weights assigned to nontraded goods in the construction of a price index determine the similarity. Heterogeneous preferences across countries induce dissimilarity. For example, when agents' preferences and weights used in the construction of a price index are

3. Stulz (1987) analyzed the effect of nontraded goods on international portfolio allocation. Devereux, Gregory, and Smith (1992) used a different model assuming separable leisure that generates lower cross-country consumption correlations. Stockman and Tesar (1995) used the nonseparable utility function with respect to nontraded goods consumption to generate a low cross-country correlation of aggregate consumption growth rates. Lewis (1996) found that both nonseparabilities and certain capital market restrictions are necessary to explain international consumption comovements.

identical across two countries, the real exchange rate becomes positively related to the cross-country consumption disparity in traded goods, but negatively related to the cross-country consumption disparity in nontraded goods.[4] The cointegration-Euler equation approach has two advantages: (1) The regression relationship is not affected by the specification of an intertemporal budget constraint. (2) The consistency and asymptotic properties of coefficient estimates are unaffected by the presence of arbitrary stationary measurement error.

The remainder of the paper is organized as follows. In section 6.2, I derive the stationarity restriction on the trend properties of real exchange rates and private consumption from the Euler equation for the agent's optimization problem. These restrictions are the foundation for the cointegration-Euler equation approach. In section 6.3, I describe the econometric specifications concerning the trend property of individual series and their implications for the stationarity restriction. Section 6.4 explains the data and reports empirical results. The countries under consideration are Japan, South Korea, Taiwan, and the United States. Two sets of bilateral relations are examined, with South Korea and Taiwan each serving as the home country. The focus is on the role of private consumption of nontraded goods in accounting for long-run real exchange rate movements in South Korea and Taiwan. Recently, Froot and Rogoff (1991) found that the cross-country difference in government spending accounts for real exchange rate movements. When government consumption is concentrated in the purchase of nontraded goods, my model predicts real exchange rate appreciation in the country with a high growth rate of government consumption. I also investigate this alternative explanation. Section 6.5 contains concluding remarks.

6.2 A Cointegration-Euler Equation Approach

Consider two countries in a world economy. Imagine that each economy is populated with an infinitely lived representative household. The household in the home country in period t is endowed with X_t^* units of exportable goods, Y_t^* units of importable goods, and Z_t^* units of nontraded goods. Goods X_t and Y_t are costlessly traded in the world markets, while Z_t is only traded domestically.

The household ranks its consumption stream $\{(X_t Y_t Z_t)', t \geq 0\}$ according to its lifetime utility function

$$E_0\left[\sum_{t=0}^{\infty} \beta^t U(X_t,\ Y_t,\ Z_t)\right],$$

4. Lucas (1982) also studied a two-country model in which the representative agent ranks the exportable good and the importable good according to its preferences and must use currency to purchase the goods. The relative price between these two goods (terms of trade) is determined by the cross-country difference in the endowments of these two goods.

in which β is a constant discount factor with $0 < \beta < 1$ and E_t denotes the mathematical expectation conditioning on the information set available at the beginning of time t, Ω_t. The intraperiod utility is assumed to be the addilog utility function

$$U(X_t, Y_t, Z_t) \equiv \sigma_{xt} \frac{X_t^{1-\alpha_x} - 1}{1 - \alpha_x} + \sigma_{yt} \frac{Y_t^{1-\alpha_y} - 1}{1 - \alpha_y} + \sigma_{zt} \frac{Z_t^{1-\alpha_z} - 1}{1 - \alpha_z},$$

in which preferences take the constant relative risk aversion form for each good and $\alpha_i > 0$, for $i = x, y, z$. When $\alpha_x = \alpha_y = \alpha_z$, preferences are homothetic in the three consumption categories. Finally, preference shocks are allowed to influence the household utility via the stationary processes $\{\sigma_{xt}, \sigma_{yt}, \sigma_{zt}, t \geq 0\}$.

Let P_{xt}, P_{yt}, and P_{zt} be the prices of exportable goods, importable goods, and nontraded goods, respectively, in period t measured in units of domestic currency. Let b_t+_1 be the real value of international assets carried from period t to period $t + 1$ measured in units of exportables, and let r_t be the real interest rate measured in units of exportables. Without borrowing and lending restrictions in the international capital market, the household's budget constraint at time t is

$$b_{t+1} = \frac{P_{zt}}{P_{xt}}(Z_t^* - Z_t) + \frac{P_{yt}}{P_{xt}}(Y_t^* - Y_t) + (X_t^* - X_t) + (1 + r_{t-1})b_t.$$

The representative agent's intertemporal optimization problem is to maximize the lifetime utility function subject to the budget constraint, and the necessary first-order conditions for this problem are

$$\frac{\partial U}{\partial Y_t} = \frac{P_{yt}}{P_{xt}} \frac{\partial U}{\partial X_t},$$

$$\frac{\partial U}{\partial Z_t} = \frac{P_{zt}}{P_{xt}} \frac{\partial U}{\partial X_t},$$

$$E_t\left[\beta \frac{\partial U}{\partial X_{t+1}}(1 + r_t) - \frac{\partial U}{\partial X_t}\right] = 0,$$

and the budget constraint holds. Under my specification of the intraperiod utility function, Euler equations in the first-order conditions can be expressed as

$$\frac{\sigma_{xt} X_t^{-\alpha_x}}{\sigma_{yt} Y_t^{-\alpha_y}} = \frac{P_{xt}}{P_{yt}},$$

$$\frac{\sigma_{zt} Z_t^{-\alpha_z}}{\sigma_{xt} X_t^{-\alpha_x}} = \frac{P_{zt}}{P_{xt}}.$$

Taking the natural logarithm on both sides of the above equations yields

(1) $$p_{xt} - p_{yt} + \alpha_x x_t - \alpha_y y_t = u_{yt},$$

(2) $$p_{xt} - p_{zt} + \alpha_x x_t - \alpha_z z_t = u_{zt},$$

where $x_t \equiv \log X_t$, $y_t \equiv \log Y_t$, $z_t \equiv \log Z_t$, $p_{it} \equiv \log P_{it}$, for $i = x, y, z$, and $u_{it} = \log \sigma_{xt} - \log \sigma_{it}$, for $i = y, z$. In equilibrium, prices and consumption must satisfy equations (1) and (2).

If u_{it} is stationary for $i = y, z$, then equations (1) and (2) imply the stationarity of $p_{xt} - p_{yt} + \alpha_x x_t - \alpha_y y_t$ and $p_{xt} - p_{zt} + \alpha_x x_t - \alpha_z z_t$. This implication allows for different trend properties of consumption of various goods, depending on preference parameters. For example, when $\alpha_i > \alpha_x$, the restriction allows good i consumption to grow at a lower rate than good X consumption for any given path of relative price and preference shocks and $i = y, z$. This is because a given change in $p_{it} - p_{xt}$ induces a greater response of good i consumption.

Suppose that the general price index in the home country can be described by

$$p_t = \theta_x p_{xt} + \theta_y p_{yt} + \theta_z p_{zt} + u_{pt},$$

in which p_t is the logarithm of the domestic price index at time t and θ_i is the weight given to good i in the index with $\theta_i > 0$, for $i = x, y, z$, and $\theta_x + \theta_y + \theta_z = 1$. The error term, u_{pt}, captures the third-country effect and is assumed to be uncorrelated with p_{it}, for $i = x, y, z$. I use this definition to eliminate p_{zt} in equation (2):

(3) $$p_t = (\theta_x + \theta_z)p_{xt} + \theta_y p_{yt} + \alpha_x \theta_z x_t - \alpha_z \theta_z z_t - v_{zt},$$

in which $v_{zt} \equiv \theta_z u_{zt} - u_{pt}$. The foreign-country counterpart of equation (3) is

$$\hat{p}_t = (\hat{\theta}_x + \hat{\theta}_z)\hat{p}_{xt} + \hat{\theta}_y \hat{p}_{yt} + \hat{\alpha}_x \hat{\theta}_z \hat{x}_t - \hat{\alpha}_z \hat{\theta}_z \hat{z}_t - \hat{v}_{zt},$$

in which $\hat{v}_{zt} \equiv \hat{\theta}_z \hat{u}_{zt} - \hat{u}_{pt}$. Here and from now on, all variables and parameters pertaining to the foreign country are designated by a hat.

For my purpose, the real exchange rate at time t, denoted q_t, is defined as

(4) $$q_t = p_t - s_t - \hat{p}_t,$$

in which s_t is the logarithm of the nominal bilateral exchange rate. A decrease in s_t means an appreciation of the domestic currency. The purchasing power parity (PPP) doctrine states that the nominal exchange rate equals the ratio between domestic and foreign prices. Therefore, real exchange rate movements indicate deviations from PPP for p_t. To sharpen the focus on the role of nontraded goods, I assume that the law of one price holds for the goods that are traded between the two countries.[5] This is captured by the following relationship:

5. The law of one price obtains if (1) markets are competitive, (2) there are no transportation costs, and (3) there are no barriers to trade, such as tariffs or quotas. Hsieh (1982), Fisher and Park (1991), and Strauss (1996), among others, also adopted this assumption for traded goods.

$$p_{it} = s_t + \hat{p}_{it},$$

for $i = x, y$. The above assumption may not be as restrictive as it appears; we can easily abandon it by allowing movements in $p_{it} - s_t - \hat{p}_{it}$. If these deviations contain a trend component, that is, if PPP for either p_{xt} or p_{yt} does not hold in the long run, v_{zt} in equation (3) will contain a trend component. Hence, checking if the estimated residual in equation (3) is stationary provides a diagnostic analysis for possible misspecifications.

Substituting equation (3) and its foreign-country counterpart into equation (4) for p_t and $\hat{p}_t$, respectively, yields

$$(5)\quad q_t = (\hat{\theta}_y - \theta_y)(p_{xt} - p_{yt}) + \alpha_x\theta_z x_t - \hat{\alpha}_x\hat{\theta}_z\hat{x}_t - \alpha_z\theta_z z_t + \hat{\alpha}_z\hat{\theta}_z\hat{z}_t + v_t,$$

in which $v_t \equiv -v_{zt} + \hat{v}_{zt}$. It is clear from equation (5) that trade between two countries imposes an equilibrium relationship among the real exchange rate, the terms of trade, and private consumption in the two countries.[6] If v_t is stationary, equation (5) imposes the restriction regulating the comovement of q_t, $p_{xt} - p_{yt}$, x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$ that

$$q_t - (\theta_y - \hat{\theta}_y)(p_{xt} - p_{yt}) - \alpha_x\theta_z x_t + \hat{\alpha}_x\hat{\theta}_z\hat{x}_t + \alpha_z\theta_z z_t - \hat{\alpha}_z\hat{\theta}_z\hat{z}_t$$

be stationary. I call this restriction the stationarity restriction, which is the foundation of the cointegration-Euler equation approach. The derivation of this restriction does not require any use of budget constraint or first-order conditions relating to the intertemporal choice of consumption. Hence, the cointegration-Euler equation approach allows for the existence of liquidity constraints or other market imperfections.

The stationarity restriction has different long-run implications for the comovement of individual variables in equation (5), depending on the trend properties of those variables. For example, if PPP for p_t holds in the long run (i.e., q_t is stationary), then the stationarity restriction requires that $p_{xt} - p_{yt}$, x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$ be cointegrated with the cointegrating vector $(\hat{\theta}_y - \theta_y, \Pi')'$, in which $\Pi' = (\alpha_x\theta_z, -\hat{\alpha}_x\hat{\theta}_z, -\alpha_z\theta_z, \hat{\alpha}_z\hat{\theta}_z)$.[7] Suppose there is a change in the nominal exchange rate caused by nominal factors. Both traded and nontraded goods consumption in the two countries have significant influence on the general price index in each country. As a result, changes in consumption in both countries must manage to maintain the long-run relationship between price ratios in the two countries and the nominal exchange rate, and the nominal factors have effects only on the short-run movements of consumption. On the other hand, if q_t contains a trend component and $p_{xt} - p_{yt}$, x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$ are cointe-

6. Helpman and Razin (1982) also include export goods, import goods, and nontraded goods in their model, but they limit the discussion to a nonstochastic model.

7. Here I adopt the definition of cointegration given in Campbell and Perron (1991, 164). An $n \times 1$ vector of variables, S_t, is said to be cointegrated if there exists at least one nonzero n-element vector β such that $\beta' S_t$ is trend stationary. This definition does not require that each of the individual series in S_t contain a unit root; some or all series can be trend stationary.

grated with the cointegrating vector $(\hat{\theta}_y - \theta_y, \Pi')'$, then the stationarity restriction implies that private consumption in equation (5) cannot be a driving force for the trend component of q_t.

As argued in Hsieh (1982), different weights (θ_t) used in the construction of the price index can cause the movement of q_t. To see this, assume that the law of one price holds for both goods X and Y and that there are no nontraded goods in the world economy $(\theta_z = \hat{\theta}_z = 0)$. Then equation (5) becomes

$$q_t = (\hat{\theta}_y - \theta_y)(p_{xt} - p_{yt}) + v_t.$$

Clearly, it is private consumption of nontraded goods that creates a link between the real exchange rate and private consumption in the model. It is trade between the two countries that creates a link between the terms of trade and the real exchange rate. When $\hat{\theta}_y \neq \theta_y$ and $v_t = 0$, the terms of trade and the real exchange rate have similar dynamics. It is preference shocks that make the correlation between the real exchange rate and the terms of trade imperfect. If there is only one good, say good Y, in the world economy, then $\theta_y = \hat{\theta}_y = 1$ and equation (5) becomes $q_t = v_t$. That is, unlike the result obtained in Backus and Smith (1993), PPP for p_t does not necessarily hold exactly, due to the presence of preference shocks.

Even though the terms of trade can account for a significant fraction of real exchange rate movements, the real exchange rate (q_t) does not necessarily have positive correlation with the terms of trade. The sign of correlation is determined by that of $\hat{\theta}_y - \theta_y$. To see this, consider an increase in the terms of trade caused by a lower importables price. If consumption of importable goods is more important in the home country than in the foreign country in the sense that $\theta_y > \hat{\theta}_y$, then the value of a unit of domestic currency (in terms of a basket of goods) must rise relative to that of the equivalent units of foreign currency. When the real exchange rate appreciates, it is optimal for private agents to increase their consumption of importables. For this case, the terms of trade and the real exchange rate are negatively correlated.

To identify other sources for the movement of q_t, assume that households in the two countries have identical preferences ($\alpha_i = \hat{\alpha}_i$, for $i = x, y, z$) and that the weights used in the construction of the price index are the same for the two countries. Given those assumptions, equation (5) can be reduced to

$$q_t = \alpha_x \theta_z (x_t - \hat{x}_t) - \alpha_z \theta_z (z_t - \hat{z}_t) + v_t.$$

It is obvious that the cross-country consumption disparities for traded and nontraded goods account for the movement of q_t: q_t increases with the cross-country consumption disparity in traded goods but decreases with the cross-country consumption disparity in nontraded goods. A country that experiences real appreciation of its currency enjoys either more rapid growth in private consumption of traded goods or less rapid growth in private consumption of nontraded goods. Since nontraded goods will be relatively more expensive in a fast-growing economy, that country's currency will experience real appreciation.

Without preference shocks and the third-country effect on the demand side, I cannot derive the long-run restriction from equation (1). However, for productivity differential models, productivity shocks do not play such a role. For example, in Hsieh's (1982) model, the supply of labor is fixed but is mobile between the tradable goods, and labor is the only input factor in production. Then the real exchange rate is a deterministic function of the following variables: productivity differentials between the tradable and nontradable sectors in both countries and the cross-country disparity in unit labor costs of the traded goods.

6.3 Econometric Specifications

The stationarity restriction summarizes the long-run equilibrium restrictions from the consumer's perspective. In a closed exchange economy, consumption equals production, and preference parameters can be identified from the stationarity restriction if the supply side exhibits much more volatility in the long run than the demand side. Ogaki (1992) and Ogaki and Park (1989) achieved the identification by assuming that productivity shocks have a stochastic trend.

Instead of modeling the production technology on the supply side, I consider an open exchange economy in the world markets. Trading opportunities imply that the consumption of goods X and Y in each country may not equal domestic production in equilibrium. For an open economy, the trend properties of private consumption of both exportable goods and importable goods are unlikely to be closely related to their domestic production. To achieve the identification of preference parameters, it is not sufficient to assume that the productivity shocks have a stochastic trend. However, preference parameters can be identified if the trend properties of export and import activities do not offset those of the corresponding production. Productivity shock is the dominant driving force in the long run.

In empirical investigation, it is difficult to obtain data on the consumption of exportables and importables. My focus will be on the two-good case: a traded good and a nontraded good. Let X_t denote the traded good. Since $p_{xt} = p_{yt}$ in the two-good case, equation (5) can be reduced to

$$q_t = \alpha_x \theta_z x_t - \hat{\alpha}_x \hat{\theta}_z \hat{x}_t - \alpha_z \theta_z z_t + \hat{\alpha}_z \hat{\theta}_z \hat{z}_t + v_t . \tag{6}$$

According to Campbell and Perron's (1991) definition of cointegration, the stationarity restriction does not require difference stationarity of all individual series. The stationarity restriction simply states that there exists at least a 5×1 vector $(1, -\Pi')$ for q_t, x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$ such that v_t is trend stationary. If all individual variables are trend stationary, then they are trivially cointegrated.

To assess the empirical significance of a heterogeneous utility function across countries, I follow the tradition in international trade and assume $\alpha_i = \hat{\alpha}_i$ for $i = x, z$ and $\theta_z = \hat{\theta}_z$. Then equation (6) can be further simplified to

$$q_t = \alpha_x \theta_z (x_t - \hat{x}_t) - \alpha_z \theta_z (z_t - \hat{z}_t) + v_t. \tag{7}$$

If both the real exchange rate and the cross-country consumption disparities, $x_t - \hat{x}_t$ and $z_t - \hat{z}_t$, contain different trend components, then the stationarity restriction implies that q_t, $x_t - \hat{x}_t$, and $z_t - \hat{z}_t$ are cointegrated. However, when q_t is stationary, the stationarity restriction does not necessarily imply the stationarity of $y_t - \hat{y}_t$ and $z_t - \hat{z}_t$. The disparities $y_t - \hat{y}_t$ and $z_t - \hat{z}_t$ can be cointegrated with the cointegrating vector $(\alpha_x \theta_z, -\alpha_z \theta_z)$ so that $\alpha_x \theta_z (x_t - \hat{x}_t) - \alpha_z \theta_z (z_t - \hat{z}_t)$ is stationary. For this case, if good X has a lower income elasticity than good Z ($\alpha_x > \alpha_z$), then the stationarity of q_t forces $z_t - \hat{z}_t$ to grow at a faster rate than $x_t - \hat{x}_t$, but private consumption does not have long-run effects on q_t. Finally, the estimates of α_x / α_z and $\hat{\alpha}_x / \hat{\alpha}_z$ can be identified here.

6.4 Data and Empirical Results

As displayed in figures 6.1 and 6.2, the consumption of both traded and nontraded goods and the bilateral real exchange rate all exhibit clear trends. I first present statistical tests for the trend properties of individual series and then estimate various cointegrating regressions under the two specifications of preference parameters and the weights given in the construction of p_t.

6.4.1 Data

The countries involved are Japan, South Korea, Taiwan, and the United States. Two sets of bilateral relations are examined, with South Korea and Taiwan each serving as the home country. Data on the exchange rates of the New Taiwan (NT) dollar against the U.S. dollar and the Japanese yen were taken from *Monthly Financial Statistics*, while the exchange rates of the Korean won against the two foreign currencies were taken from *International Financial Statistics* (*IFS*). To study the sensitivity of empirical results with respect to the use of the price index as a measure of general price level, the two selections of p_t are the consumer price index (CPI) and the wholesale price index (WPI; or producer price index, PPI). Japanese, South Korean, and U.S. price series were taken from *IFS*, and Taiwanese price series were taken from *National Income Accounts*. Let q_t^c denote the real exchange rate when the CPI is the measure of price index, and let q_t^w denote the real exchange rate when the WPI is the measure.

Following Kakkar and Ogaki (1993), real consumption expenditure on durables, semidurables, and nondurables is defined as the consumption of traded goods, while real consumption expenditure on services is defined as the consumption of nontraded goods. South Korean data on x_t and z_t were taken from *National Accounts*, published by the Bank of Korea, while Taiwanese series were taken from *National Income Accounts*. Japanese series for $\hat{x}_t$ and $\hat{z}_t$ were

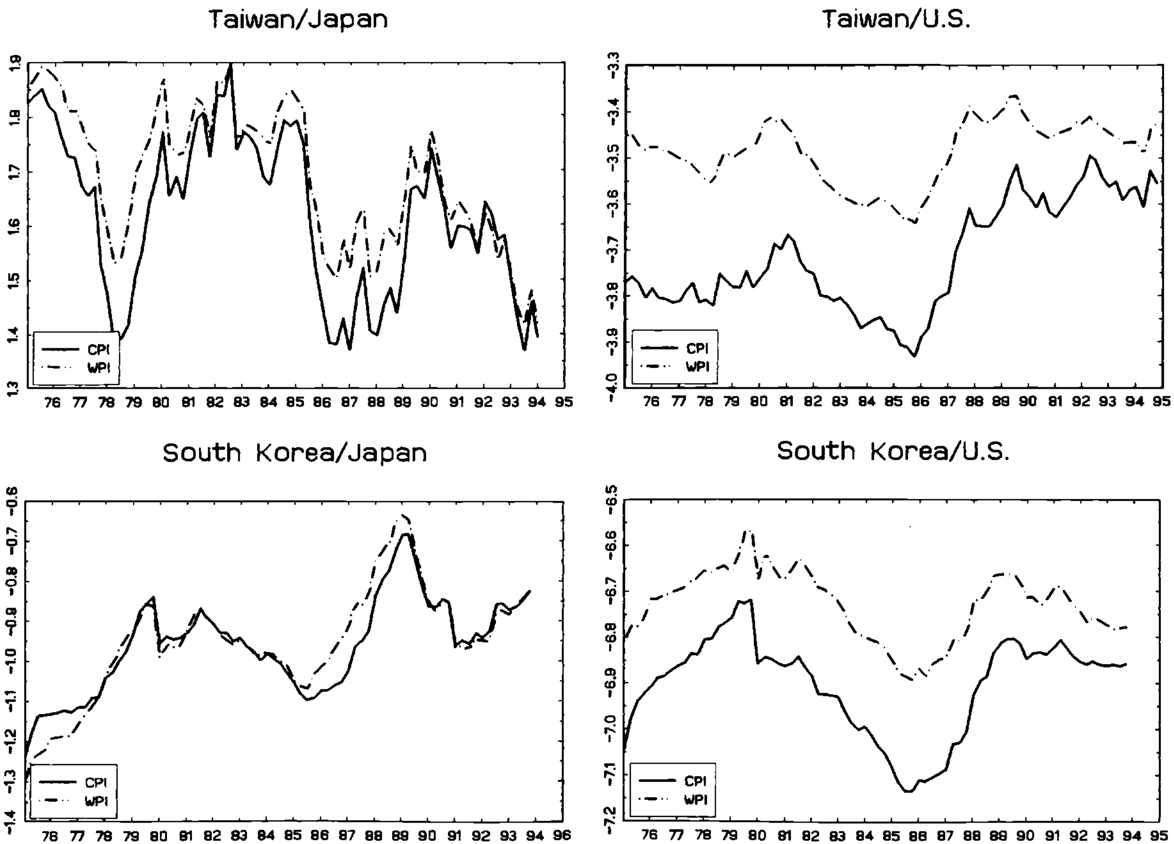

Fig. 6.1 Bilateral real exchange rate

Sources: Taiwan, *Financial Statistics Monthly* (Taipei: Central Bank of China, various issues); Japan, South Korea, and United States, *International Financial Statistics* (Washington, D.C.: International Monetary Fund, various issues).

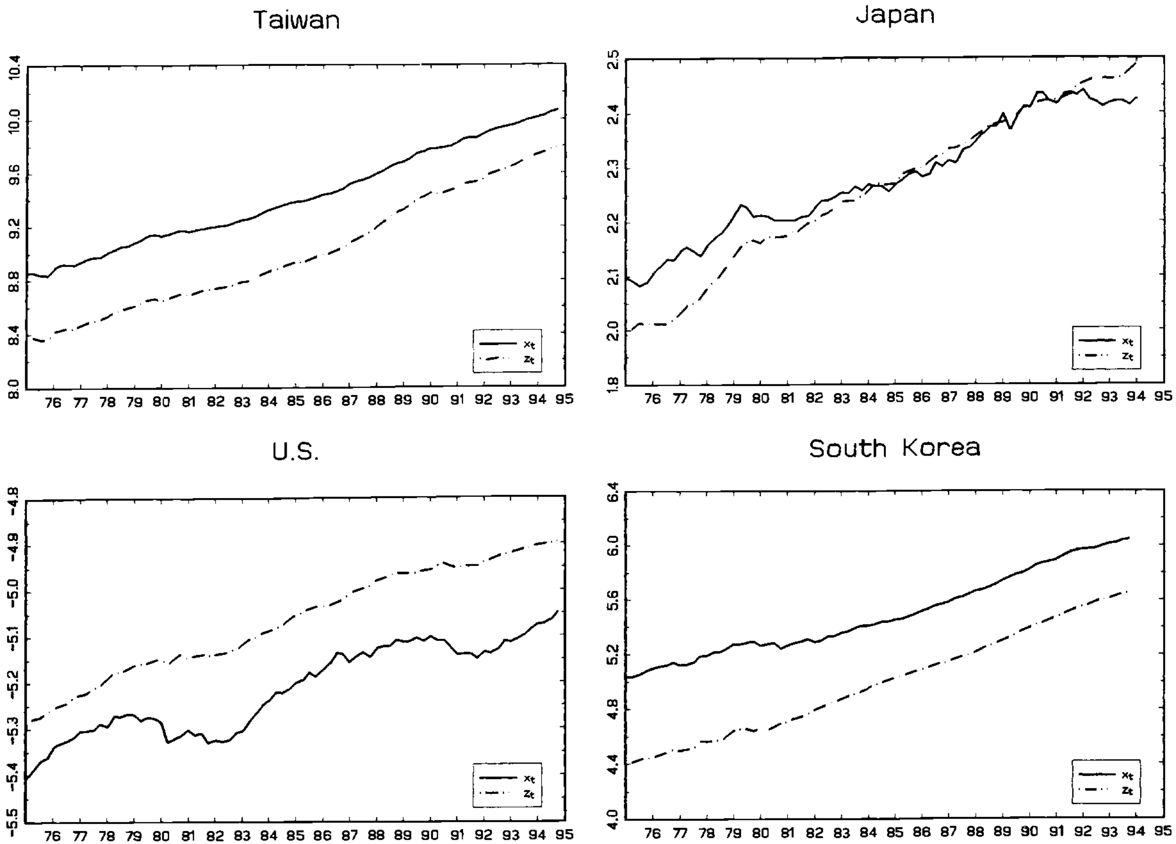

Fig. 6.2 Private consumption of traded and nontraded goods

Sources: Taiwan, *National Income Accounts* (Taipei: Directorale-General of Budget, Accounting and Statistics, various issues); South Korea, *National Accounts* (Seoul: Bank of Korea, various issues); Japan, *OECD Quarterly National Accounts* (Paris: Organization for Economic Cooperation and Development, various issues); United States, *Survey of Current Business* (Washington, D.C.: U.S. Departtaken from *OECD Quarterly National Accounts,* and U.S. series were taken from the *Survey of Current Business,* published by the Department of Commerce. Per capita real consumption of goods and services is constructed as follows. I deflate nominal consumption expenditure by the appropriate price index and then divide the resulting number by total population. All data are quarterly series. The sample period is 1975:1–1994:4 for Taiwan and the United States, and 1975:1–1993:4 for Japan and South Korea.

6.4.2 Evidence from Time-Series Data

Real bilateral exchange rates are displayed in figure 6.1, plots of x_t, z_t, $\hat{x}_t$, and $\hat{z}_t$ in figure 6.2, and those of $x_t - \hat{x}_t$ and $z_t - \hat{z}_t$ in figure 6.3. Two points are worth mentioning. First, Taiwan generally experienced a real appreciation of its currency against U.S. dollars during the sample period. The nominal depreciation of NT dollars against U.S. dollars caused a real depreciation of Taiwan's currency from 1981 to 1986, and then the real value of NT dollars was pushed up under pressure from the United States when Taiwan had a sizable current account surplus in the 1986–89 period. The bilateral exchange rate of Korean won against U.S. dollars exhibits a less clear upward trend. The real depreciation of the Korean won in 1980 and in 1982–86 was caused by the continuing nominal depreciation of Korean won against U.S. dollars. When South Korea began to enjoy a sizable current account surplus in 1986, the Korean won also came under pressure from the United States to have an unprecedented appreciation against the U.S. dollar through 1989. After 1989, a mild real depreciation of the won against the U.S. dollar was mainly due to two factors: the deterioration of South Korea's international payment position and the appreciation of the Japanese yen against the U.S. dollar since 1991. As a result, the real value of Korean won against U.S. dollars fell to the level of the late 1970s in 1993–94. The bilateral real exchange rate between South Korea and Japan exhibits a similar clear upward trend in the 1975–94 period. However, the real exchange rate between Taiwan and Japan exhibits a downward trend with volatile fluctuations.

Second, real per capita private consumption expenditures on traded goods and nontraded goods contain different trend components in the four countries. As a result, the cross-country disparities in private consumption of traded goods and nontraded goods exhibit upward trends for the four pairs of countries. The cross-country evidence in figure 6.2 indicates higher growth in the per capita real consumption of services in the course of economic development.

6.4.3 Testing for the Purchasing Power Parity Doctrine

I first test the trend property of the bilateral real exchange rate between the home country and the foreign country. If the real exchange rate does not contain a trend component, then the PPP doctrine for p_t holds in the long run. Otherwise, it does not hold in the long run. Therefore, testing the trend

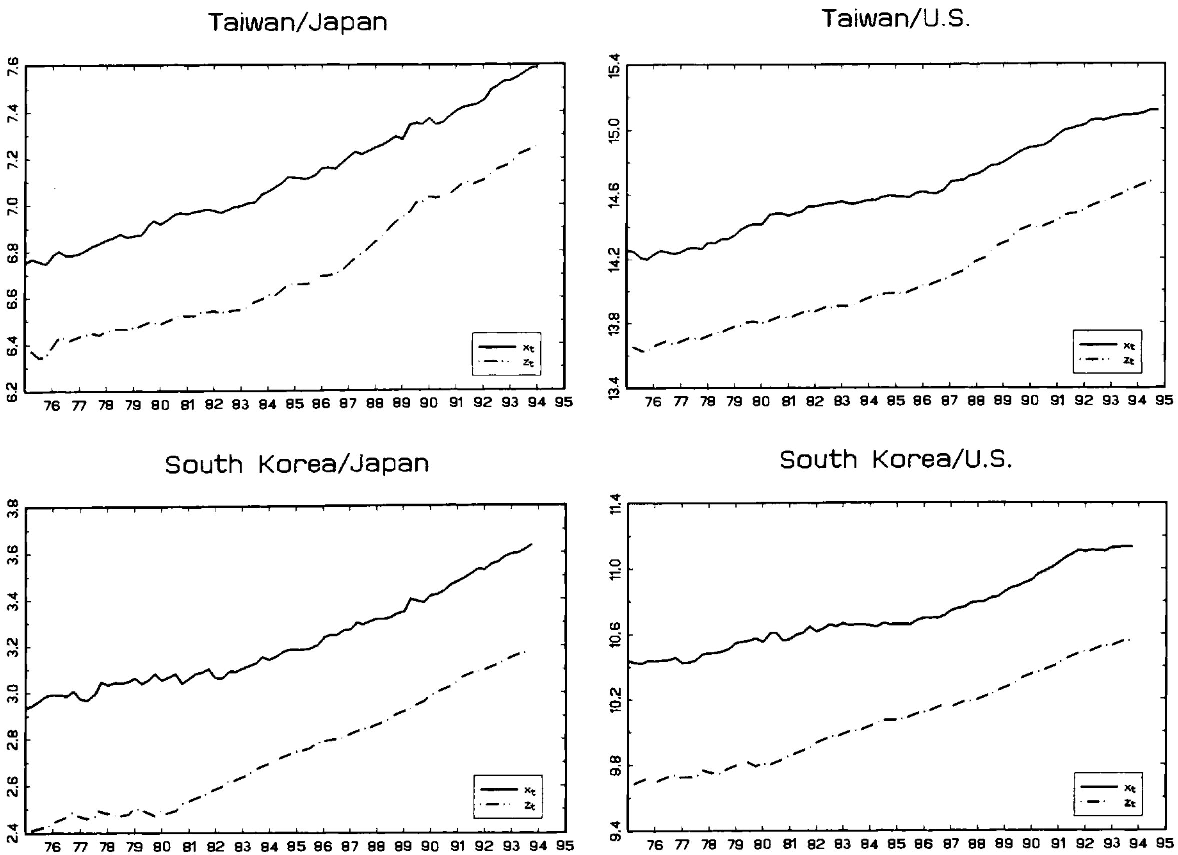

Fig. 6.3 Cross-country consumption disparity
Sources: See fig. 6.2.

property of the real exchange rate is equivalent to testing the PPP doctrine. Here, I use Park and Choi's (1988) $J(p, q)$ and $G(p,q)$ tests. I reject the null of difference stationary around a linear time trend when the $J(1,q)$ statistic is smaller than the critical values tabulated in Park and Choi (1988)[8] and reject the null of trend stationarity when the $G(1,q)$ statistic is larger than the critical values.[9]

Table 6.1 displays test results for the trend property of bilateral real exchange rates. For q_t^c and q_t^w between Taiwan and the United States, the $J(1,q)$ tests with $q = 2,3,4$ cannot reject the null of difference stationarity around a linear time trend at the 10 percent significance level. There is evidence against the trend stationarity of q_t^c at the 5 percent significance level in terms of the $G(1,2)$ and $G(1,4)$ tests. On the other hand, the $G(1,q)$ tests with $q = 2,3,4$ yield weaker evidence against the trend stationarity of q_t^w.

For q_t^c between Taiwan and Japan, the $J(1,q)$ tests all reject the null of difference stationary. The $J(1,3)$ and $J(1,4)$ tests even reject it at the 1 percent significance level. When the WPI is the measure of P_t, there is slightly improved evidence for the difference stationarity of q_t^w. The $J(1,3)$ and $J(1,4)$ tests still reject the null, and only $J(1,2)$ fails to reject it at the 10 percent significance level. On the other hand, I did not find significant evidence against the null of trend stationarity for both q_t^c and q_t^w in terms of the $G(1,q)$ tests.[10]

There is conflicting evidence for the trend property of q_t^c between South Korea and Japan. I found that the $J(1,2)$ and $J(1,4)$ tests cannot reject the null of difference stationarity. But results of the $G(1,q)$ tests with $q = 2,3,4$ support the null of trend stationarity. On the other hand, there is more consistent evidence for the difference stationarity of q_t^w. The $J(1,q)$ tests with $q = 2,3,4$ all fail to reject the null of difference stationarity for q_t^w at the 10 percent significance level. Only the $G(1,3)$ test fails to reject the null of trend stationarity for q_t^w at the 10 percent significance level. Finally, for q_t^c between South Korea and the United States, both $J(1,q)$ and $G(1,q)$ tests with $q = 2,3,4$ provided significant evidence for the null of difference stationarity. However, there is slightly weaker evidence for the difference stationarity of q_t^w in terms of the $J(1,q)$ tests. Only the $G(1,2)$ and $G(1,3)$ tests fail to reject the null of trend stationarity.

The above findings can be summarized as follows. First, the bilateral ex-

8. The $J(p, q)$ test does not require the estimation of the long-run variance and has an advantage over Phillips and Perron's Z_α (Z_t) test and the augmented Dickey-Fuller (ADF) test in that neither the bandwidth parameter nor the order of autoregression needs to be chosen. Monte Carlo experiments also show that the $J(p, q)$ test has a stable size and is not dominated by the ADF test in small samples in terms of powers.

9. Kahn and Ogaki (1992) recommend small q when the sample size is small, according to their Monte Carlo simulations. Here I chose $q = 2$, 3, and 4. For estimation of the long-run variance, I use Andrews's (1991) quadratic spectral kernel with the automatic bandwidth parameter estimator based on AR(1).

10. I report the ADF test in table 6.3 because it was widely used in the literature. None of the ADF tests reject the null of difference stationary for both q_t^c and q_t^w in the Taiwan/Japan and Taiwan/U.S. cases. In the following discussion, I only present the $J(p, q)$ and $G(p, q)$ test results when there is no conflicting evidence between these tests and the ADF test.

Table 6.1 **Tests for Trend Property of Real Exchange Rates**

Null and Statistic[a]	Taiwan/ Japan	Taiwan/ U.S.	South Korea/ Japan	South Korea/ U.S.
		Price Index: CPI		
Difference stationarity				
$J(1, 2)$	0.007*	0.397	0.026	0.257
$J(1, 3)$	0.008***	0.414	0.098*	0.484
$J(1, 4)$	0.048***	1.144	0.341	3.194
Difference stationarity				
ADF(1)	−2.352	−1.798	−2.204	−1.295
ADF(2)	−2.628	−1.936	−2.571	−1.583
ADF(3)	−2.683	−1.895	−2.554	−1.770
Trend stationarity				
$G(1, 2)$	0.135	5.586**	0.487	3.628*
$G(1, 3)$	0.171	5.750*	1.710	5.794*
$G(1, 4)$	0.951	10.483**	4.869	13.520***
		Price Index: WPI		
Difference stationarity				
$J(1, 2)$	0.053	0.168	0.193	0.028
$J(1, 3)$	0.073*	0.248	0.226	0.164
$J(1, 4)$	0.096**	0.622	0.500	1.648
Difference stationarity				
ADF(1)	−2.612	−1.894	−2.331	−1.822
ADF(2)	−2.429	−2.274	−2.537	−1.775
ADF(3)	−2.673	−2.469	−2.694	−1.885
Trend stationarity				
$G(1, 2)$	1.128	2.806*	3.098*	0.511
$G(1, 3)$	1.509	3.871	3.538	2.598
$G(1, 4)$	1.947	7.463*	6.395*	11.487***

[a] $J(p, q)$ and $G(p, q)$ denote Park and Choi's (1988) tests with a time polynomial of order p in the null hypothesis and a time polynomial of order q in the fitted regression. ADF(p) denotes Dickey and Fuller's (1984) test with a time polynomial of order 1 in the null hypothesis and p lagged first-difference terms in the fitted regression.

*Significant at the 10 percent level.

**Significant at the 5 percent level.

***Significant at the 1 percent level.

change rates contain a unit root and linear time trend in the South Korea/Japan, South Korea/U.S., and Taiwan/U.S. cases. And q_t^c and q_t^w between Taiwan and Japan are stationary around a linear time trend. Second, the measure of p_t chosen in testing the trend property of the real exchange rate does not matter for the long-run deviation of PPP for q_t. Recently, based on data in other countries, Kim (1990) and Kakkar and Ogaki (1993) found more favorable evidence for long-run PPP when the WPI is used as the measure of p_t than when the CPI is used. They argued that the large weight given to nontraded goods in the CPI could be the reason the long-run PPP doctrine based on the CPI did not receive much empirical support.

6.4.4 Testing for the Trend Property of Private Consumption

Given the trend property of the real exchange rate, private consumption in different countries must exhibit trends in order to account for long-run real exchange rate movements under the stationarity restriction. Table 6.2 presents test results for the trend property of x_t, z_t, $\hat{x}_t$, and $\hat{z}_t$.

First, for both x_t and z_t in Taiwan, the null of difference stationarity around a linear time trend cannot be rejected at the 10 percent significance level in terms of the $J(1,q)$ tests with $q = 2,3,4$. The $G(1,q)$ tests with $q = 2,3,4$ significantly reject the null of trend stationarity in favor of difference stationarity at the 1 percent significance level. Second, the null of difference stationarity for both x_t and z_t in South Korea received strong support from the $J(1,q)$ tests. But the $G(1,q)$ tests yield significant evidence against trend stationarity for these two series. In light of the above results, I assume that both x_t and z_t in South Korea and Taiwan contain a unit root around a linear time trend.

For the U.S. series of $\hat{x}_t$, I found weaker evidence for the null of difference stationarity around a linear trend. Even though the $G(1,q)$ tests with $q = 2,3,4$ reject the null of trend stationarity at the 10 percent significance level, both the $J(1,2)$ and $J(1,4)$ tests reject the null of difference stationarity at the 10 percent significance level. On the other hand, there is significant evidence for the null of difference stationarity for the U.S. series of $\hat{z}_t$. These results are confirmed by results of the $G(1,q)$ tests. For $\hat{x}_t$ and $\hat{z}_t$ in Japan, there is mixed evidence for difference stationarity. First, both the $J(1,2)$ and $J(1,3)$ tests reject the null of difference stationarity for $\hat{x}_t$ at the 10 percent significance level. Second, the $J(1,q)$ tests with $q = 2,3,4$ cannot reject the null of difference stationarity for $\hat{z}_t$ at the 10 percent significance level. They are consistent with the results of the $G(1,q)$ tests in table 6.2. Since the $G(p,q)$ test tends to overreject the null when the autoregressive root is close to one, the above findings can be viewed as conclusive evidence for the trend stationarity of $\hat{x}_t$ in Japan and the United States. And I assume that $\hat{z}_t$ in Japan and the United States contains a unit root and linear time trend.

Recently, Ogaki and Park (1989) have found significant evidence for the null of difference stationarity for the U.S. data on $\hat{z}_t$ and evidence against the trend stationarity of $\hat{z}_t$. They use seasonally adjusted monthly data on durables, nondurables, and services in the National Income and Product Accounts. The sample period is from January 1959 to December 1986. When a shorter sample period is used (February 1968 to December 1986), the null of trend stationarity for $\hat{z}_t$ cannot be rejected. Given the mixed evidence on the null of difference stationarity for the consumption of durables and nondurables, their findings are generally consistent with my results.

6.4.5 Testing for Cross-Country Consumption Disparity

If preference parameters and weights used in the construction of p_t are identical across the home country and the foreign country, the cross-country con-

Table 6.2 **Tests for Trend Property of Private Consumption**

Consumption[a]	Null: Difference Stationarity						Null: Trend Stationarity		
	$J(1, 2)$	$J(1, 3)$	$J(1, 4)$	ADF(1)	ADF(2)	ADF(3)	$G(1, 2)$	$G(1, 3)$	$G(1, 4)$
				Japan					
$\hat{x}_t$	0.008*	0.025**	1.573	−1.993	−2.156	−2.206	0.175	0.530	13.184***
$\hat{z}_t$	1.063	1.153	1.323	−1.760	−2.296	−2.066	10.105***	10.502***	11.165**
				South Korea					
x_t	2.331	2.346	9.972	−1.041	−1.126	−1.438	12.407***	12.432***	16.115***
z_t	1.986	3.135	3.175	−2.487	−2.357	−2.241	12.829***	14.622***	14.668***
				Taiwan					
x_t	1.065	1.072	3.601	−1.590	−0.925	−1.276	9.980***	10.013***	15.144***
z_t	2.131	2.179	6.300	−1.106	−0.769	−1.446	12.723***	12.814***	16.132***
				United States					
$\hat{x}_t$	0.000***	0.164	0.172*	−1.563	−1.922	−2.106	0.000	2.730	2.843
$\hat{z}_t$	0.427	0.550	0.831	−1.744	−1.639	−1.957	5.951**	7.048**	9.021**

[a]x_t and z_t denote per capita real consumption on traded and nontraded goods, respectively.

*Significant at the 10 percent level.

**Significant at the 5 percent level.

***Significant at the 1 percent level.

sumption disparity must be nonstationary in order to account for long-run real exchange rate movements. For this purpose, I test for cointegration between private consumption in different countries. If domestic consumption and foreign consumption of traded goods (nontraded goods) are not cointegrated with the normalized cointegrating vector $(1, -1)$, then the cross-country consumption disparity for traded goods (nontraded goods) contains a trend component.

Here I use Park's (1992) $H(p,q)$ statistics in testing the cointegrating relationship.[11] In particular, the $H(0,1)$ statistic can be used to test the deterministic cointegrating restriction. According to the $H(p,q)$ statistics in table 6.3, I found much evidence against cointegration between x_t and $\hat{x}_t$ (and between z_t and $\hat{z}_t$) for all possible pairs of home and foreign countries: the deterministic cointegration restriction was rejected by the $H(0,1)$ test, while the stochastic cointegration restriction was rejected by the $H(1,q)$ tests with $q = 2,3,4$ at the 1 percent significance level. These results are consistent with visual impressions obtained from figure 6.3. Both test results and visual impressions clearly indicate that the cross-country consumption disparities for traded and nontraded goods contain a trend component.

6.4.6 Testing for the Stationarity Restriction

Given the difference stationarity of q_t, the stationarity restriction simply implies that private consumption series in different countries are not cointegrated with the cointegrating vector Π'. It is still possible that private consumption series in different countries are cointegrated with other cointegrating vectors. The hypothesis-testing strategy is to conduct cointegration tests for private consumption in different countries with and without the real exchange rate included. Suppose that the test results reject the null of cointegration for the set of variables excluding q_t but fail to reject the null for the set of variables including q_t. Then the long-run movements of q_t are driven by private consumption in different countries.

Table 6.3 reports the $H(0,q)$ and $H(1,q)$ tests for the null of cointegration for the four private consumption series. When Taiwan (the United States) is designated as the home (foreign) country, the $H(0,1)$ test fails to reject the null of deterministic cointegration for x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$, and the $H(1,q)$ tests with $q = 2,3,4$ also provide strong evidence for the null of stochastic cointegration restriction. When Japan is the foreign country, the $H(0,1)$ test rejects the null of deterministic cointegration for x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$ at the 10 percent significance level. However, the $H(1,q)$ tests with $q = 2,3,4$, strongly favor the stochastic cointegration restriction.

I found much evidence against the null of cointegration for x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$ in the South Korea/Japan and South Korea/U.S. cases. Only the deterministic cointegration restriction in the South Korea/Japan case cannot be rejected by

11. Park (1992) showed that the $H(p, q)$ statistic converges in distribution to a $\chi^2(p-q)$ random variable under the null of cointegration.

Table 6.3 **Tests for Cointegration**

Variable[a]	Null: Cointegration				
	$H(0, 1)$	$H(0, 2)$	$H(1, 2)$	$H(1, 3)$	$H(1, 4)$
			South Korea/Japan		
$(x_t, \hat{x}_t)$	3.154*	19.104***	15.529***	15.538***	16.676***
$(z_t, \hat{z}_t)$	14.361***	16.584***	6.900***	10.029***	10.356**
$(x_t, \hat{x}_t, z_t, \hat{z}_t)$	1.231	10.621***	11.737***	11.761***	17.770***
$(x_t - \hat{x}_t, z_t - \hat{z}_t)$	0.862	12.060***	14.586***	14.690***	14.004***
$(q^c_t, x_t, \hat{x}_t, z_t, \hat{z}_t)$	1.136	5.078*	3.325*	3.339	3.491
$(q^w_t, x_t, \hat{x}_t, z_t, \hat{z}_t)$	2.872*	5.232*	1.834	2.085	2.367
$(q^c_t, x_t - \hat{x}_t, z_t - \hat{z}_t)$	4.671**	4.833*	0.552	0.973	4.234
$(q^w_t, x_t - \hat{x}_t, z_t - \hat{z}_t)$	7.275***	7.326**	0.019	1.695	4.309
			South Korea/United States		
$(x_t, \hat{x}_t)$	14.152***	17.612***	14.819***	14.932***	16.226***
$(z_t, \hat{z}_t)$	14.189***	16.994***	12.207***	14.243***	14.632***
$(x_t, \hat{x}_t, z_t, \hat{z}_t)$	4.117**	9.964***	7.764***	9.424***	12.677***
$(x_t - \hat{x}_t, z_t - \hat{z}_t)$	1.136	5.629*	4.003**	8.967**	11.601***
$(q^c_t, x_t, \hat{x}_t, z_t, \hat{z}_t)$	0.637	0.688	0.009	0.225	2.287
$(q^w_t, x_t, \hat{x}_t, z_t, \hat{z}_t)$	0.639	0.642	0.000	0.279	2.396
$(q^c_t, x_t - \hat{x}_t, z_t - \hat{z}_t)$	0.030	1.936	1.281	1.291	4.383
$(q^w_t, x_t - \hat{x}_t, z_t - \hat{z}_t)$	3.747*	4.253	0.597	0.961	3.491
			Taiwan/Japan		
$(x_t, \hat{x}_t)$	12.466***	20.772***	14.266***	14.357***	15.436***
$(z_t, \hat{z}_t)$	9.408***	16.151***	11.388***	11.396***	15.564***
$(x_t, \hat{x}_t, z_t, \hat{z}_t)$	2.842*	3.004	0.108	0.130	5.350
$(x_t - \hat{x}_t, z_t - \hat{z}_t)$	13.770***	14.074***	4.987**	7.893**	8.720**
$(q^c_t, x_t, \hat{x}_t, z_t, \hat{z}_t)$	0.397	1.320	0.493	0.651	1.438
$(q^w_t, x_t, \hat{x}_t, z_t, \hat{z}_t)$	0.818	1.811	0.527	1.071	2.671
$(q^c_t, x_t - \hat{x}_t, z_t - \hat{z}_t)$	0.069	0.133	0.019	0.083	2.839
$(q^w_t, x_t - \hat{x}_t, z_t - \hat{z}_t)$	0.320	0.752	0.068	0.371	2.978
			Taiwan/United States		
$(x_t, \hat{x}_t)$	16.968***	17.812***	13.980***	13.980***	15.998***
$(z_t, \hat{z}_t)$	9.136***	17.143***	13.803***	13.843***	15.588***
$(x_t, \hat{x}_t, z_t, \hat{z}_t)$	1.608	2.056	0.744	1.494	1.764
$(x_t - \hat{x}_t, z_t - \hat{z}_t)$	6.534**	9.456***	3.393*	5.130*	5.132
$(q^c_t, x_t, \hat{x}_t, z_t, \hat{z}_t)$	1.316	2.159	0.360	0.720	0.723
$(q^w_t, x_t, \hat{x}_t, z_t, \hat{z}_t)$	1.306	2.808	0.798	1.848	2.594
$(q^c_t, x_t - \hat{x}_t, z_t - \hat{z}_t)$	6.756***	6.914**	0.488	1.666	1.828
$(q^w_t, x_t - \hat{x}_t, z_t - \hat{z}_t)$	2.953*	2.966	0.054	2.153	2.170

[a] q_t denotes real exchange rate; x_t and z_t denote per capita real consumption on traded and nontraded goods, respectively.

*Significant at the 10 percent level.

**Significant at the 5 percent level.

***Significant at the 1 percent level.

the $H(0,1)$ test. There is more than a single source of nonstationarity in generating the long-run movements of x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$ here.

Next, I apply the $H(p,q)$ tests to q_t, x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$; the results are also given in table 6.3. Using both measures of p_t, I found little evidence against the stationarity restriction in the Taiwan/Japan and Taiwan/U.S. cases: the deterministic cointegration restriction was not rejected by the $H(0,1)$ test, nor was the stochastic cointegration restriction rejected by the $H(1,q)$ tests with $q = 2,3,4$. Even if the four individual private consumption series are cointegrated, the above finding clearly suggests that private consumption in different countries can account for long-run movements of the real exchange rate. And the private consumption series are cointegrated with a cointegrating vector other than Π'. In previous subsections, I found evidence for the trend stationarity of q_t between Taiwan and Japan and of $\hat{x}_t$ in Japan and the United States. These results apparently did not affect the test results for the stationarity restriction.

There is mixed evidence for the stationarity restriction in the South Korea/Japan case. When the CPI is the measure of p_t, the $H(0,1)$ test fails to reject deterministic cointegration for q_t, x_t, $\hat{x}_t$, z_t, and $\hat{z}_t$. On the other hand, the stochastic cointegration restriction was rejected by the $H(1,2)$ test at the 10 percent significance level. When the WPI is the measure of p_t, the stationarity restriction was rejected by the $H(0,1)$ test but cannot be rejected by the $H(1,q)$ tests with $q = 2,3,4$. For the South Korea/U.S. case, I found little evidence against the stationarity restriction. The difference stationarity of q_t and the stationarity restriction together imply that private consumption accounts for the long-run movement of the real exchange rate.

When I assumed that preference parameters and weights used in the construction of p_t are identical across the home country and the foreign country, I found significant evidence for the stationarity restriction in the Taiwan/Japan case, and weaker evidence for the stationarity restriction in the Taiwan/U.S. case. The stochastic cointegration restriction cannot be rejected by the $H(1,q)$ tests with $q = 2,3,4$ in the South Korea/Japan and South Korea/U.S. cases.

6.4.7 Cointegrating Regression Results

In addition to the stationarity restriction, our model imposes restrictions on the signs of coefficients in the cointegrating regressions. In this subsection, I investigate the signs of coefficient estimates as a way to evaluate the economic significance of the model. Table 6.4 reports the cointegrating regression results using Park's (1992) canonical cointegrating regression (CCR) procedure and Phillips and Hansen's (1990) fully modified (FM) estimation procedure. When heterogeneous utility functions are assumed in estimation, coefficient estimates are generally inconsistent with the predictions of the model. These results make at least two points clear. First, private consumption can account for the long-run movement of the real exchange rate. Second, if we take the restrictions on the signs of coefficients imposed by the model seriously, it is necessary to refine the specifications of preferences so that private consumption will have consistent effects on the real exchange rate.

Table 6.4 **Cointegrating Regressions of Real Exchange Rates on Private Consumption**

Price Index and Equation	$\alpha_x\theta_z$	$\hat{\alpha}_x\hat{\theta}_z$	$\alpha_z\theta_z$	$\hat{\alpha}_z\hat{\theta}_z$
		South Korea/Japan		
CPI				
Eq. (6)	1.502*/1.462*	−0.686/−0.804	2.778*/2.775*	1.159/1.057
Eq. (7)	1.669*/1.628*		2.488*/2.452*	
WPI				
Eq. (6)	1.341*/1.267*	−0.591/−0.786	3.546*/3.543*	1.298/1.113
Eq. (7)	1.241*/1.209*		2.756*/2.728*	
		South Korea/United States		
CPI				
Eq. (6)	2.657*/2.653*	1.477*/1.503*	1.864*/1.826*	2.662*/2.767*
Eq. (7)	2.061*/2.057*		1.625*/1.601*	
WPI				
Eq. (6)	1.686*/1.682*	1.233*/1.236*	1.728*/1.707*	2.425*/2.439*
Eq. (7)	1.427*/1.424*		1.721*/1.700*	
		Taiwan/Japan		
CPI				
Eq. (6)	−8.731*/−8.476*	0.032/−0.073	−5.225*/−5.087*	3.770*/3.749*
Eq. (7)	−1.222/−1.232		−0.342/−0.331	
WPI				
Eq. (6)	−5.414*/−5.331*	−0.476/−0.498	−3.114*/−3.068*	2.923*/2.964*
Eq. (7)	−1.364/−1.343		−0.270/−0.255	
		Taiwan/United States		
CPI				
Eq. (6)	−0.791/−0.943	1.462*/1.475*	−1.841*/−1.945*	2.784*/2.824*
Eq. (7)	0.657*/0.646*		−0.673*/−0.700*	
WPI				
Eq. (6)	0.149/−0.064	0.698*/0.741*	−0.815/−0.961	2.078*/2.213*
Eq. (7)	0.226/0.215		−0.487*/−0.512*	

Note: In each entry A/B, A denotes Park's (1992) CCR estimate, and B denotes Phillips and Hansen's (1990) FM estimate.

*Significant at the 5 percent level.

When utility functions are identical across two countries, I found that estimates of $\alpha_x\theta_z$ and $\alpha_z\theta_z$ have theoretically correct signs in the South Korea/Japan and South Korea/U.S. cases. Note that α_x/α_z measures the ratio of income elasticities of z_t and x_t in South Korea. The implied value of α_x/α_z is less (greater) than one in the South Korea/Japan (South Korea/U.S.) case. The unstable ratio across the two cases indicates that the model does not perform well in this respect. As revealed in figures 6.1 and 6.3, South Korea experienced mild real appreciation against both the U.S. dollar and the Japanese yen, and $x_t - \hat{x}_t$ and $z_t - \hat{z}_t$ both exhibit clear upward trends in these cases. To account for the more significant upward trend in real exchange rate, the risk aversion for nontraded good consumption must be higher in the South Korea/Japan case.

For the Taiwan/Japan case, I had theoretically wrong signs for the estimates of $\alpha_x\theta_z$ and $\alpha_z\theta_z$. The bilateral real exchange rate between Taiwan and Japan exhibits a downward trend, which reflects the depreciation of NT dollars against Japanese yen in the sample period. Since Taiwan experienced relatively more rapid growth in x_t and z_t, as displayed in figure 6.3, coefficient estimates for $x_t - \hat{x}_t$ and $x_t - \hat{z}_t$ must switch sign to account for the declining pattern of the real exchange rate. Facing continuing real appreciation of the Japanese yen in the two-country world economy, private agents in Japan are expected to increase their consumption of traded goods by increasing imports from Taiwan, while those in Taiwan are expected to substitute relatively cheaper nontraded goods for more expensive traded goods. Since Taiwan had increasing trade deficits with Japan in the sample period, the substitution effects in the two-country world economy cannot be a crucial element in the determination of real exchange rate movements. Finally, I found that the coefficient estimates of $\alpha_z\theta_z$ have wrong signs in the Taiwan/U.S. case. When Taiwan experienced a significant real appreciation against the U.S. dollar, my model predicts that private agents in Taiwan enjoyed less rapid growth in the consumption of nontraded goods. When the upward trend in cross-country disparity in traded good consumption is not significant enough in accounting for the real appreciation, it forces the sign of the $\alpha_z\theta_z$ estimate to change.

6.4.8 Private Consumption versus Government Consumption

An alternative explanation of the long-run movement of the real exchange rate was recently proposed by Froot and Rogoff (1991). The channel linking government consumption expenditure and the real exchange rate can be described as follows. When a larger fraction of government consumption falls on nontraded goods than does private consumption, an increase in government consumption increases the real appreciation of domestic currency against foreign currency. Those countries that experienced real appreciation against foreign currency enjoyed relatively more rapid growth in government consumption expenditure.

Table 6.5 shows the results of cointegrating regressions of the real exchange rate on private consumption and government consumption:

$$(8) \qquad q_t = \alpha_x\theta_z x_t - \hat{\alpha}_x\hat{\theta}_z\hat{x}_t - \alpha_z\theta_z z_t + \hat{\alpha}_z\hat{\theta}_z\hat{z}_t + \gamma g_t - \hat{\gamma}\hat{g}_t + v_t',$$

in which g_t and $\hat{g}_t$ are per capita real government consumption expenditure in the home country and foreign country, respectively. If government consumption expenditure is assumed to fall totally on nontraded goods, then the movement of private consumption of nontraded goods completely reflects that of government consumption spending. Hence, we expect that the coefficient estimates of γ and $\hat{\gamma}$ are insignificantly different from zero once private consumption of traded and nontraded goods is a regressor in the cointegrating regressions. In general, we expect that $\gamma > 0$ and $\hat{\gamma} > 0$. The evidence in table 6.5 indicates that the empirical relationships between the real exchange rate and private consumption are not significantly affected by the presence of government consumption expenditure in the cointegrating regressions. The data show

Table 6.5 **Cointegrating Regressions of Real Exchange Rates on Private Consumption and Government Consumption**

Price Index and Equation	$\alpha_x\theta_z$	$\hat{\alpha}_x\hat{\theta}_z$	$\alpha_z\theta_z$	$\hat{\alpha}_z\hat{\theta}_z$	γ	$\hat{\gamma}$
			South Korea/Japan			
CPI						
Eq. (8)	1.671*/1.655*	−0.774/−0.832	2.886*/2.743*	1.011/0.834	−0.076/−0.067	−0.235/−0.468
Eq. (9)	1.214*/1.272*		2.335*/2.348*		0.227/0.182	
WPI						
Eq. (8)	1.564*/1.513*	−0.629/−0.746	3.513*/3.285*	0.948/0.665	−0.067/−0.051	−0.586/−0.934
Eq. (9)	0.785*/0.839*		2.644*/2.658*		0.233/0.192	
			South Korea/United States			
CPI						
Eq. (8)	2.304*/2.327*	1.367*/1.306*	1.110*/1.171*	3.749*/3.585*	0.005/0.003	0.912*/0.910*
Eq. (9)	1.611*/1.666*		1.292*/1.343*		0.299*/0.260*	
WPI						
Eq. (8)	1.536*/1.562*	1.311*/1.241*	1.437*/1.513*	3.091*/2.894*	−0.001/−0.002	0.239/0.240
Eq. (9)	1.279*/1.295*		1.677*/1.685*		0.104*/0.091*	
			Taiwan/Japan			
CPI						
Eq. (8)	−9.457*/−9.333*	3.070*/2.803*	−5.555*/−5.495*	4.775*/4.690*	1.694*/1.588*	0.542/0.515
Eq. (9)	0.440/0.286		0.838/0.785		1.066/1.029	
WPI						
Eq. (8)	−5.715*/−5.635*	1.534*/1.365*	−3.316*/−3.255*	3.591*/3.471*	0.818*/0.794*	0.362/0.347
Eq. (9)	0.148/0.051		0.619/0.584		0.633/0.621	
			Taiwan/United States			
CPI						
Eq. (8)	−0.451/−0.569	2.111*/2.114*	−1.645*/−1.755*	3.653*/3.716*	−0.163/−0.231	−0.672*/−0.611*
Eq. (9)	0.912*/0.908*		−0.620*/−0.635*		−0.260/−0.257	
WPI						
Eq. (8)	0.693/0.604	1.847*/1.830*	−0.644/−0.708	3.561*/3.565*	−0.406*/−0.426*	−1.029*/−0.992*
Eq. (9)	0.783*/0.775*		−0.352/−0.370		−0.557*/−0.529*	

Note: In each entry A/B, A denotes Park's (1992) CCR estimate, and B denotes Phillips and Hansen's (1990) FM estimate. Sample period is 1975:1–1993:4.
*Significant at the 5 percent level.

no evidence of government consumption effects on real exchange rates. Some of the coefficients on government consumption in the home country and foreign country are not statistically different from zero and are even of the wrong sign. The inclusion of government consumption regressors in equation (8) has little effect on the estimates of $\alpha_x\theta_x$, $\hat{\alpha}_x\hat{\theta}_z$, $\alpha_z\theta_z$, and $\hat{\alpha}_z\hat{\alpha}_z$. These remain as statistically significant as before, with the signs for coefficient estimates unchanged.

To access the empirical significance of the cross-country disparity in real government consumption, $g_t - \hat{g}_t$, in the cointegrating regression of equation (7), table 6.5 also presents the results of the following cointegrating regression:

$$(9) \qquad q_t = \alpha_x\theta_z(x_t - \hat{x}_t) - \alpha_z\theta_z(z_t - \hat{z}_t) + \gamma(g_t - \hat{g}_t) + v_t''.$$

I obtain results for the effect of the cross-country disparity in government consumption on the real exchange rate similar to those above. The coefficients on domestic and foreign private consumption become larger and even more statistically significant when $g_t - \hat{g}_t$ is included. But the wrong signs for the estimates of $x_t - \hat{x}_t$ and $z_t - \hat{z}_t$ remain quite severe. Thus accounting for government consumption does not seem to overturn the result that private consumption affects the long-run movement of the real exchange rate.

6.5 Concluding Remarks

The empirical evidence suggests that private consumption in home and foreign countries provides a significant component of the explanation of long-run movements in the real exchange rate in South Korea and Taiwan. Based on the signs of coefficient estimates in the cointegrating regressions, it seems that private consumption may not be a reliable fundamental that has reliable effects on the real exchange rate.

It is useful to incorporate supply-side elements such as productivity differentials in a general equilibrium model of real exchange rate determination and explore the trend and cyclical implications from equilibrium relationships obtained in the model. Since fluctuations in the relative price of traded goods account for a significant fraction of real exchange rate movements, another interesting topic for future research is to estimate equation (5).

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Comment *Yun-Wing Sung*

This paper is interesting because it approaches the long-run behavior of real exchange rates from the demand side instead of the more usual supply side. The paper also contains a vast amount of empirical tests and statistical results.

In the empirical tests, four different sets of stationarity assumptions were used for identification. Each assumption was tested for four different cases of real exchange rate movements between (i) South Korea and Japan, (ii) South Korea and the United States, (iii) Taiwan and Japan, and (iv) Taiwan and the United States. For each of these cases, two measures of the real exchange rate were used: one by the CPI and the other by the WPI. There were thus a total of 32 cases ($4 \times 4 \times 2$).

The empirical results were disappointing. In quite a lot of cases, the stationarity assumptions required for identification were not satisfied. In the cases where the stationarity assumptions were satisfied, the regression coefficients were often the wrong signs or were insignificant. Among the 32 cases, only two cases (the real exchange rate between South Korea and the United States measured by the CPI and by the WPI under one set of assumptions) gave good results, that is, significant regression coefficients with the right signs.

While a researcher always hopes for good empirical results, he or she may not find them, and the fault may not be with the researcher. Maybe the real world is too complicated for even the best methodology, or maybe the data are deficient. A paper should not be judged merely by its empirical results.

In terms of exposition, the paper could certainly be improved. The paper is rich in technical details and statistical tests but short on economics and interpretation of results. For instance, the author did not give the motivation for the demand-side approach. Though I am not familiar with this approach, I think it may be superior to the more traditional supply-side approach of Balassa and Samuelson in several ways. In the supply-side approach, the behavior of the real exchange rate hinges on differential changes in productivity between tradables and nontradables, and productivity change in nontradables (services) is notoriously difficult to measure. In the demand-side approach, the real exchange rate is positively (negatively) related to the ratio of consumption of traded (nontraded) goods in the two countries, and the consumption of traded and nontraded goods is much easier to measure than productivity change in these goods, especially in nontraded goods.

Some assumptions of the demand-side approach are less stringent than those of the supply-side approach, and this can be an important advantage. The demand-side approach requires the equalization of the marginal rates of substitution in consumption across countries. Under free trade, this is generally true except when there are quantitative restrictions on consumption, which is rare.

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In general, it can be claimed that market imperfections on the demand side are much less likely than those on the supply side. However, the demand-side approach requires an additive utility function, and the author did not make clear what limitations this would imply.

The author summarized the results of his cointegrating regressions in about one page, and very little economic interpretation was given for any of the results. The paper would benefit from more discussion of the pros and cons of the demand-side approach, the economics behind the assumptions used in identification, and careful interpretations of the economics of the results.

Comment Gian Maria Milesi-Ferretti

In recent years a number of theoretical and empirical studies have examined the determinants of real exchange rates (see the excellent survey by Froot and Rogoff 1995). Most of these studies have focused on supply-side factors, such as differences in productivity growth rates within a country (between the traded and the nontraded goods sectors) and across countries. This paper focuses instead on demand-side determinants of the real exchange rate and tests the implications of a simple theoretical model on bilateral real exchange rate data between Taiwan and Korea, on the one side, and Japan and the United States, on the other side.

The theoretical analysis links real exchange rate changes to the dynamics of the terms of trade and of relative consumption growth and shows in particular that the real exchange rate should appreciate (depreciate) when traded (nontraded) goods consumption growth in the home country increases relative to consumption growth in the foreign country.

The link between the real exchange rate, the terms of trade, and cross-country consumption ratios of traded and nontraded goods can be easily understood in the following simplified setting. Assume that preferences and the weights of traded and nontraded goods in the price index are the same across countries and that the utility function is separable in the two goods and exhibits constant intertemporal elasticity of substitution. In this case the relative price of nontraded goods in terms of traded goods is equal to the ratio of marginal utilities, which in turn is inversely proportional to the consumption ratio of nontraded and traded goods:

$$\frac{P_N}{P_T} = \frac{C_N^{-\alpha_N}}{C_T^{-\alpha_T}}, \tag{1}$$

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where the α terms are the inverse of the intertemporal elasticities of substitution. An analogous expression obtains for the foreign country. The (logarithm of the) real exchange rate q is defined as

$$(2) \qquad q = (p_T - s - p_T^*) + \theta_N[(p_N - p_T) - (p_N^* - p_T^*)],$$

where lowercase variables indicate logs, an asterisk indicates "foreign" variables, s is the nominal exchange rate between the domestic and the foreign currency, and θ_N is the weight of nontraded goods in the price index of both countries. Taking logs of equation (1) and inserting it into equation (2) we obtain the result that the real exchange rate is a function of the relative price of traded goods across countries (itself a function of the terms of trade) and of the relative consumption of nontraded and traded goods across countries:

$$(3) \qquad q = (p_T - s - p_T^*) + \theta_N[\alpha_T(c_T - c_T^*) - \alpha_N(c_N - c_N^*)].$$

The paper derives more general forms of equation (3), considering the case in which there are two traded goods, and tests the theoretically implied restrictions on the time-series properties of the real exchange rate and consumption series. The author uses an impressive array of state-of-the-art tests to characterize these time-series properties. I have nothing useful to say about the tests, other than ritually recalling their limited power when the time series is short. I feel, however, that the paper could be improved by (i) providing better links between its different parts and (ii) integrating this demand-side approach with supply-side considerations.

With regard to the first point, the econometric specification section should rely more clearly on the results of the univariate time-series analysis of the real exchange rate and consumption series. For example, if unit root tests provide evidence of level stationarity for a series (e.g., consumption of tradables in Japan or the United States), then it seems inappropriate to run cointegrating regressions that treat that same variable as a difference-stationary one. Also, some economic interpretation of the empirical results would help to improve the link between theoretical and quantitative analysis and would provide the reader with a feel for the performance of the model. Finally, since we know the composition of the CPI and WPI, it seems logical and straightforward to use the results of the empirical analysis to draw inferences about the underlying preference parameters.

With regard to the second point, it should not be difficult in future work to incorporate supply-side considerations into this paper's basic theoretical structure and empirical analysis. For example, theoretically it is sufficient to extend the endowment model to allow for production and productivity growth. Indeed, a number of authors have studied theoretically and empirically both demand- and supply-side determinants of real exchange rates (see, e.g., Froot and Rogoff 1991; De Gregorio, Giovannini, and Krueger 1994), although they did not rely on the time-series properties of consumption series.

On a more general note, a nice aspect of this paper's approach is that it establishes links between the real exchange rate and consumption variables, which are in general more readily available than productivity data. The issue, of course, is how far one can get in explaining real exchange rate behavior relying solely on these variables. In the empirical analysis carried out in the paper, terms-of-trade fluctuations are not explicitly considered, and the focus is on variables explaining the relative price of nontraded goods in terms of traded goods within countries, under the assumption that the law of one price holds for traded goods. A general problem plaguing real exchange rate analysis based on the supply-side, Balassa-Samuelson approach is that intersectoral productivity differentials are good predictors of the relative price of nontraded goods in terms of traded goods but poor predictors of real exchange rate behavior (see, e.g., Asea and Mendoza 1994). This empirical result reflects the widely documented fact that fluctuations in the relative price of traded goods account for a significant fraction of real exchange rate changes. It seems therefore that future research in this area cannot abstract from the examination of determinants of the terms of trade. This paper's theoretical analysis is a step in this direction; the task is now to make it empirically implementable.

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