Unit Root and Cointegration Tests of Puraching Power Parity in the Pacific Rim

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The literature review about Purchasing Power Parity(PPP) suggests that there is controversy concerning the appropriateness of the PPP assumption. To date, the vast majority of the empirical tests have concerned the U.S., European, and Japanese economies. In a sense, such tests are most favorable to PPP; the implicit assumption is that PPP should perform best between economies with similar industrial structures. However, this assumption may not be valid. As shown by Mussa (1979), PPP works well for nations experiencing very different inflation rates. Enders' (1989) study of PPP during the greenback and gold-standard periods shows that PPP works well for nations experiencing very rapid growth rates.

In this light, it is interesting to consider the PPP relationship for the Korean economy. Korea represents a rapidly growing economy with strong trading ties to other Pacific Rim nations as well as to the U.S. and Europe. A comparison of the performance of PPP between Japan and Korea (both rapidly growing nations) might provide an interesting contrast to that of Korea versus other Pacific Rim nations and to Korea versus the U.S. The methodology that follows is that of enders (1988, 1989) and Corbae and Ouliaris (1988).

I. Unit Root Tests and the Real Exchange Rate

To test the PPP relationship, consider the following econometric model:

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$$(1)\quad e(t)p^*(t) - rp(t) = d(t)$$

where $e(t)$ = won price of the foreign currency in (t), $p^*(t)$ = foreign price level in (t), $p(t)$ = Korean price level in (t), $d(t)$ is a stochastic disturbance which represents a deviation from PPP, and r is a constant.

Long-run PPP implies that $r=1$ and $d(t)$ is stationary with mean zero. Note that e, p, and p^* are endogenous variables which are jointly determined; there is no obvious candidate for the left-hand side variable. To avoid the standard practice of estimating (1) by using instrumental variables, consider the reformulation of PPP in terms of the real exchange rate :

$$(2)\quad e(t)p^*(t)/p(t) = r + d_1(t)$$
$$r(t) = r + d_1(t)$$

where $d_1(t)$ is a stochastic disturbance and $r(t)$ is the real exchange rate $= e(t)p^*(t)/p(t)$.

In this formulation, long-run PPP holds if $d_1(t)$ is stationary; r is then the long-run value of the real exchange rate and $d_1(t)$ is the deviation of the real exchange rate from its long-run value.

The ARIMA model selection

If $d_1(t)$ is an indeterministic covariance stationary stochastic process, by the Wold decomposition theorem, $d_1(t)$ has an infinite order moving average representation which can be well approximated by a finite autoregressive representation under certain conditions.

If, for example, $d_1(t)$ is finite ARIMA (n, o, o), the underlying process for the real exchange rate movement is suggested by :

$$(3)\quad r(t) = a_o + a_1 r(t-1) + \cdots + a_n r(t-n) + e_1(t)$$

where $e_1(t)$ is a serially uncorrelated stochastic disturbance with mean equal to zero.

Given this specification, long-run PPP requires that all characteristic roots of (3) lie within the unit circle. Because we can test only the relative version of PPP, the data place no restrictions on the estimated value of r. Using monthly data from *International Fi-*

nancial Statistics, the real exchange rates for 6 of Korea's major trading partners--the U.S., Germany, Japan, India, the Philippines and Thailand--were constructed. The sample period is January 1973 to July 1987 (representing a period of flexible exchange rates).

The data series for the U.S. real exchange rate was constructed by multiplying the U. S. wholesale price index by the won price of the dollar and then dividing by the Korean wholesale price index. In the same way, we have obtained the real exchange rates for other countries.

Standard Box-Jenkins model selection procedures were used to characterize the nature of the $d_1(t)$ series. This Box-Jenkins modeling strategy consists of three stages (identification, estimation, and diagnostic checking). The maximum likelihood estimates of the "best" ARIMA models for each country are reported in Table 1.

Table 1. Maximum likelihood ARIMA estimates

$$r(t) = a_0 + \alpha_1 r(t-1) + \alpha_2 r(t-2) + \alpha_3 r(t-3) + e1(t)$$

January 1973-July 1987	a_0	a_1	a_2	a_3	MU[a]
U.S.	0.0235	0.9867 (0.0738)[b]	0.1838 (0.1041)	-0.1933 (0.0739)	1.0325
Germany	0.0433	1.2499 (0.0722)	-0.2858 (0.0725)		1.2065
Japan	0.0288	1.2216 (0.0731)	-0.2427 (0.0740)		1.3705
Philippines	0.0464	0.9662 (0.0166)			1.3765
India	0.0233	1.2605 (0.0722)	-0.2858 (0.0722)		0.9232
Thailand	0.0406	0.9630 (0.0190)			1.1011

[a] MU indicates mean level of the real exchange rate.

[b] The standard errors are in parentheses.

We are now in a position to determine whether the real exchange rates are stationary. Mann and Wald (1943) proved that the vector of least squares estimators for the nth-order stationary time series converges in distribution to a vector normal random variable. For the nonstationary time series, the story is different. The special case of nonstationary time series with multiple unit roots has been discusded by Dickey and Fuller (1979), Hasza and Fuller (1982), and Dickey, Hasza, and Fuller (1984). In the presence of a single unit root, the standard Dickey-Fuller test is suggested by ARIMA representation.[1] The Dickey-Fuller test consists of rewriting equation (3) as :

$$(4)\quad r(t) = a_0 + b_1 r(t-1) + \sum_{i=2}^{n} b_i \mathrm{delr}(t+1-i) + e_1(t)$$

where : $b_1 = \sum_{i=1}^{n} a_i$; $b^i = \sum_{j=1}^{n} a_{j+1}$; $\mathrm{delr}(t+1-i) = r(t+1-i) - r(t-i)$

Dickey and Fuller show that the confidence intervals under the null hypothesis that $b_1 = 1$ are larger than the standard confidence intervals under the null of no unit root. To reject the null of no unit root, Dickey and Fuller calculate the $(\frac{b_1-1}{\mathrm{standarderror}})$ must be greater than :[2]

	Singificance Level		
Obs.	0.01	0.05	0.10
100	-3.51	-2.89	-2.58
250	-3.46	-2.88	-2.57

Unit root test results for the real exchange rate for each country are reported in Table 2.

Notice that these estimated parameters are in the stability set of the parameter space; the point estimate of the largest characteristic root always suggests convergence. Howerver, by using the Dickey-Fuller confidence intervals under the null of a single unit root, we canot reject the null at the 10% significance level for all countries but India. Even though we could not accept the random walk hypothesis for the real exchange rate for India, the point estimate for dominant root indicates that there is a great amount of persistence in any deviation from the PPP.

1) Given a stochastic difference equation, for example, $r(t) = a_0 + a_1 r(t-1) + a_2 r(t-2) + a_3 r(t-3) + e(t)$, if we suspect that there is a single unit root with the other roots less than 1, then the above stochastic difference equation can be rewritten in the following form : $r(t) = a_0 + b_1 r(t-1) + b_2 \mathrm{delr}(t-1) + b_3 \mathrm{delr}(t-2)$ where $b_1 = a_1 + a_2 + a_3$.

2) The stability condition is simply that the absolute value of b_1 be less than unity.

Table 2. Dickey-Fuller form of ARIMA estimates

$$r(t) = a_0 + b_1 r(t-1) + b_2 \mathrm{delr}(t-1) + b_3 \mathrm{delr}(t-2) + e_1(t)$$

January 1973-July 1987	a_0	b_1	b_2	b_3	t-stat[a]	R
U.S.	0.0284 (0.0147)[b]	0.9711 (0.0146)	0.0010 (0.0748)	0.1925 (0.0747)	-1.97	0.964
Germany	0.0408 (0.0204)	0.9652 (0.0171)	0.2649 (0.0734)		-2.03	0.951
Japan	0.0271 (0.0207)	0.9796 (0.0157)	0.2100 (0.0754)		-1.29	0.960
Philippines	0.0496 (0.0288)	0.9652 (0.0209)			-1.66	0.926
India	0.0327 (0.0132)	0.9619 (0.0145)	0.2620 (0.0725)		-2.61	0.963
Thailand	0.0463 (0.0218)	0.9562 (0.0198)			-2.19	0.932

[a] The t-statistic is for the hypothesis $b_1=1$.

[b] The standard errors are in parentheses.

SURE Estimates

If we could not rule out any possible correlation in the error terms across equations, it is natural for us to consider Zellner's Seemingly Unrelated Regressions (SURE) estimates of the real exchange rate. By doing this, we can improve the precision of our estimates.

In performing the SURE estimations, the ARIMA representation for the real exchange reate was used: and AR(1) for the Philippines and Thailand an AR(2) for Germany, Japan, and India, and an AR(3) fro the United States.The results are reported in Table 3. the column labeled 't-statistic' indicates the t-value for $b_1=1$.

From Table 3. we can now reject the null of a unit root for the U.S. as well as for India. It is surprising that, other than for Thailand, the Japanese-Korean real exchange rate is the most likely to be nonstationary.

Ⅱ. Cointegration and Error Correction Models

If two economic variables are nonstationary, it is still possible that a linear combination of the two is stationary. Following Granger and Engle (1984), we know that two time series—$e(t)p^*(t)$ and $p(t)$—are cointegrated of order (d, b) if :

Table 3. Unit root tests for SURE model

$$r(t) = a_0 + b_1 r(t-1) + b_2 delr(t-1) + b_3 delr(t-2) + e_1(t)$$

January 1973- July 1987	a_0	b_1	b_2	b_3	t-statistic[a]
U.S.	0.0364 (0.0107)[b]	0.9632 (0.0106)	0.1141 (0.0517)	0.0863 (0.0502)	-3.47
Germany	0.0340 (0.0151)	0.9709 (0.0125)	0.2795 (0.0526)		-2.32
Japan	0.0310 (0.0163)	0.9766 (0.0123)	0.2221 (0.0576)		-1.90
Philippines	0.0523 (0.0205)	0.9633 (0.0181)			-2.02
India	0.0308 (0.0092)	0.9641 (0.0200)	0.2636		-3.59
Thailand	0.0151 (0.0165)	0.9848 (0.0150)			-1.01

[a]The t-statistics are for the hypothesis $b_1=1$.

[b]The standard errors are in parentheses.

1) $e(t)p^*(t)$ and $p(t)$ are integrated of order d; thus, to have stationary stochastic processes, we have to difference both $p(t)$ and $e(t)^*(t)$ d times.

2) there exists a scalar $r(r \neq 0)$ so that the series $e(t)p^*(t) - r\ p(t)$ is integrated of order d-b.

Campbell and Shiller (1987) argued that vector autorgressive representation is not appropriate in the presence of a cointegrating vector; instead, an error-correction model is also recommended. As Granger and Engle (1984) show, if $p(t)$ and $e(t)p^*$ are integrated of order 1, then it is generally true that $z(t) = e(t)p^*(t) - s\ p(t)$ will also be I(1). However, it is possible that $d(t) = e(t)p^*(t) - r\ p(t)$ is intergrated of order zero.

By using this information, we can construct the error correcting model. A linear representation of the econometric model is as follows :

(1) $e(t)p^*(t) - r\ p(t) = d(t)$

(5) $e(t)p^*(t) - s\ p(t) = z(t)$

Note that $d(t)$ should be stationary because, by assumption, $p(t)$ and $e(t)p^*(t)$ are cointegrated of order (1, 0): the residual of (1) is stationary without differencing. Note that this is an assumption implied by PPP; unless d(t) is stationary, long-run PPP cannot hold. On the other hand, z(t) is assumed to follow a random walk; if z(t) is stationary, prices and the exchange rate must be stationary. The stationarity of z(t) is violated by the observed movements in prices and the exchange rate.

To formulate the error-correcting model, let the d(t) series exhibit first-order serial correlation. An AR representation for d(t) and a z(t) series could be written as :

$$(6)\quad d(t)-\rho d(t-1)=e_2(t)$$
$$(7)\quad z(t)-z(t-1)=e_3(t)$$

where $e_2(t)$ and $e_3(t)$ are uncorrelated white noise disturbances and $0<\rho<1$.

Manipulating (1), (5), (6), and (7), we can derive the error correction representation.[3]

$$(8)\quad (1-L)e(t)p^*(t)=-s(1-\rho)/(s-r)d(t-1)+s/(s-r)e_2(t)-r/(s-r)e_3(t)$$
$$(9)\quad (1-L)p(t)=-(1-\rho)/(s-r)d(t-1)+1/(s-r)e_2(t)-1/(s-r)e_3(t)$$

Equations (8) and (9) show how exchange rates and/or prices [e(t)p*(t), p(t)] can be explained by the previous deviation--d(t-1)--from equilibrium. Notice that the error-correction model would be appropriate if there is a cointegrating vector which makes the linear combination of economic variables stable. But we can not exclude the possibility of no cointegrating vector.

Cointegration Tests

Engle and granger (1987) argued that the estimated cointegrating vector is a consistent estimator in a large sample. Regressing e(t)p*(t) on p(t), we obtain an estimate of r which is a consistent estimator proviede that d(t) is stationary. In the same way, the re-

3) Equations (1) and (6) can be combined to obtain : $(1-\rho L)e(t)p^*=r(1-\rho L)p(t)+e_2(t)$, where L deontes the lag operator. We can obtain the following equation, (A), by adding and subtracting Le(t)p*(t) and r p(t-1).

$$(A)\quad (1-L)e(t)p^*(t)=r(1-L)p(t)+r(1-\rho)p(t-1)-(1-\rho)e(t-1)p^*(t-1)+e_2(t)$$

Equation (7) yields (B).

$$(B)\quad (1-L)e(t)p^*(t)=s(1-L)p(t)+e_3(t)$$

Solving (A) and (B) simultaneously yields (8) and (9).

gression of p(t) on e(t)p*(t) yields a consistent estimator of 1/r.

To perform the cointegration tests, the residuals of this equilibrium regression should be checked for stationarity by uisng a Dickey-Fuller test; if p(t) and e(t)p*(t) are cointegrated, the residuals must be stationary. The 'equilibrium' relationship was estimated as follows :

$$e(t)p^*(t) = r\ p(t) + d(t).$$

If d(t) is stationary, a finite AR representation for d(t) is possible. This AR representation for d(t) could well be rewritten for a Dickey-Fuller test.

Table 4. Cointegration test

January 1973- July 1987		t-statistic[a]
No lagged changes; $(1-L)d(T) = a_0 + d(t-1)$		
United States	-0.0105 (0.0110)[b]	0.9545
Germany	0.0 (0.0148)	0.0
Japan	-0.005 (0.0134)	-0.073
Philippines	-0.0363 (0.0216)	-1.6805
India	-0.0425 (0.0221)	-1.92
Thailand	-0.0935 (0.0325)	-2.87
Four lags; $(1-L)d(t) = a_0 + \theta d(t-1) + \sum_{i=1}^{4} (1-L)d(t-i)$		
United States	-0.0135 (0.0112)	-1.2053
Germany	0.0031 (0.0140)	0.2214
Japan	-0.0174 (0.0143)	-1.2167
Philippines	-0.0344 (0.0225)	-1.5288
India	-0.0578 (0.0226)	-2.5575
Thailand	-0.0147 (0.0364)	-4.05

a) The t-statistic is for the hypothesis $\theta = 0$. To reject the null of no unit root, Dickey and Fuller show that the t-statistic should be greater than -2.58 (with 100 observation) at 10% significance level.

b) The standard errors are in parentheses.

Specifically : $(1-L)\hat{d}(t) = \theta \hat{d}(t-1) + \sum_{i=1}^{n} (1-L)\hat{d}(t-i)$

where the d(t) series is the estimated residual of (1). If the estimated residuals are stationary, the estimated value of θ will be significantly different from zero. Cointegration test results are reported in Table 4.

Table 4. indicates that the cointegration tests fail for all nations except for the case of Thailand. All the t-statistics except for Thailand's are sufficiently small that we can not reject the null hypothesis of no cointegrating vector at a 10% significance level.
Notice, however, that cointegration tests for PPP between Korea and India are borderline insignificant at the 10% level. Given the point estimates, there is some evidence supportive of PPP; in general, howeve, it is hard to argue that long-run PPP holds for Korea.

Error correcting model

Given that the Thai and Korean price levels are cointegrated, it is possible to estimate the error-correcting model. Consider (10) and (11) :

$$(10)\quad (1-L)e(t)p^{*}(t) = \underset{(0.0093)}{0.0298} - \underset{(0.0364)}{0.0712}[e(t-1)p^{*}(t-1) - rp(t-1)]$$

$$(11)\quad (1-L)p(t) = \underset{(0.0050)}{0.0307} + \underset{(0.0194)}{0.0227}\{e(t-1)p^{*}(t-1) - rp(t-1)]$$

where r is the estimate of the long-run real exchange rate obtained from the equilibrium regression.

The Thailand price level multiplied by the won price of the baht declined in response to a positive deviation from PPP. The point estimate of the slope coefficient in (10) says that approximately 7% of the previous deviation from the equilibrium relationship was adjusted within one month.

Note that the Korea price level does not seem to be responsive to deviations from previous equilibrium relationship with Thailand; the point estimate of the adjustment coefficient for the Korean price level is well within a standard deviation from zero. This result would be expected if we consider that Thailand is the minor trade partner for

Korea.

Consider the error-correcting models for the Korea major trade partners, the U.S. and Japan. Even though they failed the formal test for cointegration, the error-correcting representations are instructive.

Consider the error-correcting model for the U.S. :

$$(12)\quad (1-L)e(t)p^*(t)=\underset{(0.0047)}{0.0321}-\underset{(0.0098)}{0.0498}[e(t-1)p^*(t-1)-rp(t-1)]$$

$$(13)\quad (1-L)p(t)=\underset{(0.0048)}{0.0303}-\underset{(0.0100)}{0.0368}[e(t-1)p^*(t-1)-rp(t-1)]$$

In response to a positive deviation from PPP, the U.S. price level multiplied by the won price of the dollar decreased. the point estimate of the slope coefficient in (12) implies that about 5% of the previous deviation from the PPP was corrected within one month.

The Korean price level actually declined in response to a positive deviation from the PPP; the point estimate of the adjustment coefficient for the Korean price level is significantly different from zero.

Consider also the error-correcting model for Japan :

$$(14)\quad (1-L)e(t)p^*(t)=\underset{(0.0143)}{0.0533}-\underset{(0.0130)}{0.0092}[e(t-1)p^*(t-1)-rp(t-1)]$$

$$(15)\quad (1-L)p(t)=\underset{(0.0050)}{0.0306}-\underset{(0.0045)}{0.0065}[e(t-1)p^*(t-1)-rp(t-1)]$$

Contrary to the U.S., Japan's price level multiplied by the won price of the yen was not responsive to a positive deviation from the PPP. The point estimate of the slope coefficient in (14) is well within a standard deviation from zero.

Moreover, the Korean price level did not appear to be responsive to deviations from the real exchange rate movement, the point estimate of the adjustment coefficient for the Korean price level is well within a standard deviation from zero.

To see the adjustment between the exchange rate and the PPP in a different way, we have repeated the Engle and Granger (1987) procedure by using the foreign price level and the Korean price level divided by the won price of the foreign exchange.

The estimated error-correcting models for Thailand are :

$$(16)\quad (1-L)p^*(t) = \underset{(0.0020)}{0.0120} + \underset{(0.0157)}{0.0220}[p(t-1)/e(t-1) - (1/r)p^*(t-1)]$$

$$(17)\quad (1-L)p(t)/e(t) = \underset{(0.0041)}{0.0126} - \underset{(0.0323)}{0.0728}[p(t-1)/e(t-1) - (1/r)p^*(t-1)]$$

The point estimate of the slope coefficient for (16) shows that Thailand's price level did not seem to be responsive to a positive deviation from the previous equilibrium regression; the point estimate of the adjustment coefficient for the Thailand price level is well within a standard deviation from zero.

On the other hand the Korean price level divided by the won price of the baht eliminated almost 7% of the deviation within one month.

The estimated error-correcting models for the U.S. are :

$$(18)\quad (1-L)p^*(t) = \underset{(0.0010)}{0.0095} + \underset{(0.0048)}{0.0226}[p(t-1)/e(t-1) - (1/r)p^*(t-1)]$$

$$(19)\quad (1-L)p(t)/e(t) = \underset{(0.0030)}{0.0090} + \underset{(0.0137)}{0.0057}[p(t-1)/e(t-1) - (1/r)p^*(t-1)]$$

The U.S. price level was adjusted to eliminate almost 2% of the deviation from PPP. On the contrary, the slope coefficient of (19) implies that the Korean price level divided by the won price of the dollar was not corrected to eliminate the previous deviation from the PPP.

The estimated error-correcting models for Japan are :

$$(20)\quad (1-L)p^*(t) = \underset{(0.0012)}{0.0042} + \underset{(0.0075)}{0.0325}[p(t-1)/e(t-1) - (1/r)p^*(t-1)]$$

$$(21)\quad (1-L)p(t)/e(t) = \underset{(0.0038)}{0.0013} + \underset{(0.9234)}{0.0095}[p(t-1)/e(t-1) - (1/r)p^*(t-1)]$$

Surprisingly, we have gotten the same result with the U.S. The Japanese price level moved in the correct direction in response to the previous deviation from PPP. The price level divided by the won price of the yen did not appear to adjust to a positive deviation from the equilibrium relationship.

Ⅲ. Conclusions

Point estimates of ARIMA models of the real exchange rate for Korea and her major trading partners indicate convergence; this result is in accord with long-run PPP. However, by using Dickey-Fuller tests, we could not reject the null hypothesis of a single unit root for any nation except India. SURE estimates indicated that both the U.S. and Indian real exchange rates were convergent.

Engle and Granger (1987) argued that if there is an equilibrium relationship between economic variables, these time series might be cointegrated with each other. Cointegration tests for all nations but Thailand failed to indicate PPP. The overall impression is that PPP cannot be said to hold for the Koean economy.

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〈국문초록〉

환태평양 지역에서의 구매력평가설에 관한 단위근 및 공적분 검정

김 진 옥

구매력평가설(Thory of Purchasing Power Parity)에 관한 문헌을 개관하여 보면 구매력평가설이 성립하기 위한 가정의 적합성에 관하여 상당한 논란의 여지가 있음을 알 수 있다. 지금까지 행해진 경험적 분석들은 미국 · 유럽 및 일본경제에 있어서 구매력 평가설의 성립여부를 검정한 것 들이었다. 이러한 검정결과들은 구매력평가설이 성립하는 것을 옹호하는 편이지만, 그 암묵적 가정은 구매력평가설이 유사한 산업구조를 갖고 있는 국가들 사이에 성립한다는 것이다. 그러나 이러한 가정은 적절하지 못하다.

이러한 관점에서, 한국경제를 주축으로 한 환태평양지역에서 구매력평가설의 성과를 규명코자 시계열분석의 일환인 공적분 검정(Cointegration Test)과 단위근 검정(Unit Root Test)을 하였다. 공적분 검정결과에 의하면 태국을 제외한 국가들(미국, 독일, 일본, 필리핀, 인도)에 있어서 구매력평가설은 성립하지 않았다. 즉 한국경제에 있어서 구매력평가설은 성립하지 않음을 알 수 있다.

전형적인 이론에 의하면, 단기에 있어서 통화공급충격은 구매력평가의 편차를 야기하지만, 장기에 있어서 이러한 편차는 소멸된다.

반면에 실질충격(Oil Shock, 정부지출)은 단기 또는 장기에서 구매력평가의 편차를 야기한다. 따라서 본 논문의 향후 연구과제는 환태평양지역에서 구매력평가설이 성립하지 않음을 이론적으로 재규명하고 동시에 경험적 분석을 하는 데 있다.