Long-Run Exchange Rates, Price Levels, and Purchasing Power Parity: Cointegration Tests of Five Korea Trading Partners' Currencies

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요 약

In this paper, we obtained some supportive evidence for the long-run PPP relationship concerning the Korean Won currency. Previous tests of PPP in the bilateral exchange rates of the Korean Won rate vis-a-vis the U.S. Dollar have been exposed to the lack of power problem. We argue that their failure to find PPP relation in Korean Won rates was due to the low power of Augmented Dickey-Fuller tests or the Engle-Granger two-step tests applied to the Korean exchange rate data with short sample period. In attempting to alleviate this low power problem, we used the error-correction model test and the Johansen test for bilateral long-run equilibrium relationships between exchange rates and price indices from Korea's major trading partners. It is surprising that our evidence supporting for long-run PPP in Korean Won rate contrasts sharply with Bahmani-Oskooee, Moshen and Rhee, Hyun-Jae(1992)'s.

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Ⅰ. Introduction

The purpose of this paper is to present empirical evidence of the purchasing power parity (PPP) hypothesis for five bilateral exchange rates involving the Korea trading partners' currencies. We examine the time series properties of the PPP relation for five countries, U.S., Japan, U.K., Germany, and Canada in relation to Korea during the period of 1980 to 1995. Employing the Johansen test as well as the error-correction model test, we do cointegration tests for bilateral long-run equilibrium relationships between exchange rates and price indices from Korea's major trading partners.

It is suggested that according to Cassel's idea, the nominal exchange rate should reflect the purchasing power of one currency against another. While this notion appears simple enough, many economists have introduced purchasing power parity as a long-run equilibrium condition in their open economy macroeconomic models. Therefore, testing the validity of PPP was a prerequisite to using the long-run determinant model of exchange rates in their models. A vast number of empirical tests has focused on PPP.

However, it has been widely reported that the empirical evidence of PPP did not hold in the short run. Generally, the exchange rate reacts quickly, while the price of goods adjusts slowly. And deviations from PPP can occur in the short run. The remaining question that must be solved concerns testing the validity of long-run PPP. Many empirical studies have investigated evidences for long-run PPP, presenting mixed results. For example, Frenkel(1978), Corbae and Ouliaris(1988), Enders(1988) and McNown and Wallace(1989, 1994) find supportive evidences for long-run PPP. Other studies by Frenkel(1981), Adler and Lehmann(1983) and Mark (1990) fail to find evidences in favor of the empirical validity of PPP.

Several explanations for the mixed evidences have been offered. One explanation is that in early studies such as Frenkel(1978), the nonstationarity of prices and exchange rates were not considered. A second explanation is that considerable amounts of observations are raquired to test such lon-run PPP (see Adler and Lehmann(1983) and Johnson(1990)). A third explanation is that the U.S. dollar rates as the base currency rate should be replaced with the counterpart rates (see Phylaktis(1992) and Pippenger(1993)).

The analysis conducted in this paper differs from the previous tests of PPP in three ways. First, we note that Dickey-Fuller test and Engle- Granger two-step method using unit root have the problem of low power unless they are applied to the large sample over a long horizon (see Dickey and Fuller(1979) and Hakkio and Rush(1991)). To overcome this low power problem, Johnson(1990) used longer span data and Oh(1996) used pooled panel data sets. Bahmani-Oskooee and Rhee(1992) applied the augmented Dicky-Fuller test to the Korean Won rate against the U.S. dollar rate. They found the results which are inconsistent with the hypothesis of long-run PPP. However, their failure to find favorable evidence of PPP may be thought to be due to lack of testing power which unit root tests have. In trying to alleviate these problems, we shall focus on the maximum likelihood estimation procedure offered by Johansen and Juselius (1990), as well as error-correction mechanism estimation. We shall follow the test procedures that Cheung and Lai(1993) and Kugler and Lenz(1993) used. Second, this paper conducts the PPP analysis on bilateral exchange rates and price indices from U.S., Japan, U.K., Germany and Canada, using the Korean Won rate as the base index. Third, this study will use monthly data covering the longer period from 1980-3 to 1995-12 in which the floating Korean Won rates are available. Price indices used in this analysis are both wholesale price index (WPI) and consumer price index (CPI). McNown and Wallace(1989) maintain that the choice of the price index matters in PPP analysis.

According to the absolute version of PPP, the equilibrium exchange rate between two countries is equal to the price ratio of those countries. The relative version of PPP states that the equilibrium exchange rate will be in proportion to the ratio of two countries' price levels. If two countries use the same weights and mixes of goods in constructing CPIs and WPIs, it follows that the law of one price will hold also for identical basket of goods and the absolute version of PPP will hold. In reality, since the two countries use differest consumption baskets, the relative version of PPP is widely used in the empirical analysis of PPP relationship. The absolute version of PPP denoting the law of one price can be written as

$$S_t = P(L)_t / P^*(L)_t$$

where S_t is the domestic-currency price of foreign currency, and $P(L)_t$ and $P^*(L)_t$ denote the domestic price index and the foreign price index, respectively. If a constant A, which is not equal to one, is multiplied to the right side of above equation, we have $S_t = A\,\frac{P(L)_t}{P^*(L)_t}$. We might expect the relative version of PPP to hold. The relative version of PPP states that changes in the exchange rate depend upon changes in relative price levels (That is, changes in domestic price levels minus changes in foreign price levels).

If we compare the current-year exchange rate $S_t = A \cdot \frac{P(L)_t}{P^*(L)_t}$ with the base-year exchange rate $S_0 = A \cdot \frac{P(L)_0}{P^*(L)_0}$, we have

$$\frac{S_t}{S_0} = \frac{A \cdot \dfrac{P(L)_t}{P^*(L)_t}}{A \cdot \dfrac{P(L)_0}{P^*(L)_0}} = \frac{\dfrac{P(L)_t}{P(L)_0}}{\dfrac{P^*(L)_t}{P^*(L)_0}} = \frac{P(I)_t}{P^*(I)_t},$$

where $P(I)_t$ and $P^*(I)_t$ are the domestic price index and the foreign price index in relation to the base-year index, respectively.

Taking the natural logarithm of the above equation, we have the relative PPP relationship as follows.

$$s_t = \alpha + p_t - p^*_t$$

In order to study the above relative PPP relationship, Frenkel(1978) ran regressions of the following form.

$$s_t = \alpha + \beta(p_t - p^*_t) + \upsilon_t$$

Frenkel argued that estimates of β quite close to one indicate the empirical validity of PPP. Then, the PPP relationship might be used in any model of exchange rate determination.

However, a critical question was raised later in this test methodology. That is the possible nonstationarity of prices and exchange rates which can make the above test invalid.

This paper proceeds as follows. In Section Ⅱ, we explain the error- correction models including the PPP relationship and the Johansen maximum likelihood technique to search for multivariate cointegration. The empirical results are presented and discussed in section Ⅲ. Concluding remarks follow in the last section.

Ⅱ. Testing Models

1. Error-Correction Models

For purchasing power parity studies, we consider the matrix R_t=[S_t , P_t , P_t^*], each of which is nonstationary. The vector time series S_t , P_t , P_t^* denote a log exchange rate, a log domestic price index and a foreign price index, respectively. According to Engle and Granger(1987), the vector time series R_t is defined to have cointegration if some linear combination of the series produces a stationary series. The cointegrating regression can be written as follows.

$$S_t = \alpha + \beta_1 P_t + \beta_2 P_t^* + \nu_t$$

If ν_t follows a stationary process, then the vector time series R_t=[S_t , P_t , P_t^*] is cointegrated. In this regression, any restrictions are not imposed on the coefficient value vector β. McNown and Wallace(1989, 1991) impose a symmetry restriction $\beta_1 = -\beta_2$ on the coefficient values and use a bivariate cointegrating regression for the PPP test. If the proportionality restriction of α=0 and $\beta_1 = -\beta_2 = 1$ is imposed, the unit root test for the real exchange rate can be considerad as in Adler and Lehman(1983). However, Pippenger argues that the imposition of restrictions on the cointegrating vector are not appropriate, since price indices used in the test do not have the same weight across countries. In order to prevent the incorrect rejection of a PPP relationship, we shall not impose any restriction on the cointegrating vector.

The major flaw in the Engle-Granger two-step test for cointegrating regression is that it actually uses unit root test and lacks sufficient power to reject. Some argues that the issue of power can be improved by using a vast amount of data. But, we conclude that the Korean exchange rate sample starting in 1980 is far too short to reliably reject the null hypothesis of a PPP relationship.

In this paper, the statistical test for cointegration will be performed using an error-correction model. Engle and Granger(1987) derived the error correction representation theorem. It states that if the vector time series R_t is cointegrated, then there exists an error correction representation for the change in S_t. The error correction mechanism for the change in nominal exchange rate S_t can be expressed as follows.

$$\triangle S_t = \psi + \theta_1 \triangle P_t + \theta_2 \triangle P_t^* + \omega V_{t-1} + \varepsilon_t$$

The error correction term $V_{t-1} = (S_{t-1} - \alpha - \beta_1 P_{t-1} - \beta_2 P_{t-1}^*)$ denotes the estimated lagged deviation from PPP and forces the adjustment of the nominal exchange rate S_t towards its equilibrium value. The coefficient (ω) of V_{t-1}

indicates the speed at which the nominal exchange rate restores the long-run PPP relationship. The existence of cointegration in the vector time series R_t requires that the estimated coefficient (ω) of V_{t-1} should not be equal to zero at the statistically significant level. Therefore, if our test cannot reject $H_0 : \omega = 0$, it means that the nominal exchange rate and the domestic and foreign price indices are not cointegrated. We shall test the null hypothesis $H_0 : \omega = 0$ in the above equation of error correction mechanism, and we will investigate the existence of a cointegrating PPP relationship.

2. Johansen Tests

In this paper, we also use the multivariate cointegration methodology proposed by Johansen(1988) and Johansen and Juselius(1990). This maximum likelihood approach can overcome some drawbacks in the previous studies using the unit root tests. It is designed to search for the multiple number of cointegrating relationships. We consider the vector time series R_t=[S_t, P_t, P_t^*] represented by a vector autoregression(VAR).

$$R_t = \mu + \sum_{i=1}^{i} \Pi_i R_{t-i} + \nu_t$$

where R_t is an $n \times 1$ time series vector, α is a constant vector, Π_i are coefficient matrices, and ν_t are i.i.d. $N(0, \Omega)$ distributed. By differencing the above equation, we can derive the following model of error correction representation.

$$\triangle R_t = \mu + \sum_{i=1}^{i-1} \Gamma_i \triangle R_{t-i} + \Gamma_j R_{t-j} + \upsilon_t$$

where $\Gamma_i = (-I + \Pi_1 + \cdots + \Pi_i)$ and Γ_j indicates the long-run effects of error term on R_t. The rank of Γ_j determines the number of cointegrating relationships.

The n by n matrix, Γ_j, of rank $r < n$ can be written as $\alpha \beta'$ (α and β' denote n by r matrix and r by n matrix, respectively). If r is greater than zero, $\beta' R_{t-1}$ becomes a stationary vector time series of dimension r. The, the matrix β' has the r cointegrating vectors.

The maximum likelihood estimator of β can be found by the following procedure. First, we solve the following equation for the eigenvalues $\lambda_1 > \lambda_2 > \cdots > \lambda_r$. Next, we find the r eigenvectors with respect to these eigenvalues. The r eigenvectors are the choice of β.

$$| \lambda S_{11} - S_{10} S_{00}^{-1} S_{01} | = 0$$

$$\text{where, } S_{pq} = T^{-1} \sum_{t=1}^{T} U_{pt} U'_{qt} \text{ , } 2p,q=0,1$$

$$\triangle R_t = a + \sum_{i=1}^{i-1} \Gamma_{0i} \triangle R_{t-i} + U_{0t}$$

$$R_{t-j} = a + \sum_{i=1}^{i-1} \Gamma_{1i} \triangle R_{t-i} + U_{1t}$$

The eigenvalues $\lambda_1 > \lambda_2 > \cdots > \lambda_n$ can be obtained by computing canonical correlations between residual vector R_{ot} and residual vector R_t.

Now, we want to test whether there are at most r cointegrating vectors, or whether a certain relationship can be found in the cointegrating space. For the hypothesis test of at most r cointegrating vectors, we use the following likelihood ratio test statistic which is called the trace statistics.

$$-2 \ell n \, Q_{r/0} = -T \sum_{i=r+1}^{n} \ell n \, (1-\lambda_i)$$

Similarly, we can test the null hypothesis of at most r cointegrating vectors against the alternative hypothesis of at most (r+1) cointegrating vectors. We compute the following likelihood ratio test statistic called the maximum eigenvalue test.

$$-2 \ell n \, Q_{r/r+1} = -T \ell n \, (1-\lambda_{r+1})$$

The asymptotic distributions and critical values for the above likelihood ratio test statistics are given in Johansen and Juselius(1990).

Ⅲ. Empirical Results

1. The Data

In this paper, we use monthly exchange rate and price data for the period March 1980 to December 1995. The exchange rates are monthly average nominal rates expressed as Korean Won per unit of foreign currency: the U.S. Dollar, Japanese Yen, British Pound, Deutsche Mark, and Canadian Dollar. These countries are Korea's major trading partners. The price series are consumer price indices and wholesale price indices taken form the IMF's International Financial Statistics. All the data series are seasonally unadjusted because seasonal adjustments may distort unit roots tests. We also work with the natural logarithms of all the data.

The sample in this study covers the period from March 1980 to December 1995, and the data with 190 monthly observations seems insufficient. However, Hakkio and Rush(1991) see that in attempting to gain more observations, some authors turn to more frequently sampled data such as daily observations. They argue that these efforts may distort the cointegration tests and do not improve the lack of power (see MacDonald (1995)). This forces us to stick to the sample data with monthly observations.

2. Test Results

In this paper, the series of exchange rates and price indices were each first checked for a unit root. Cointegration analysis assumes that individual variables are integrated of the same order and are nonstationary. In order to test for nonstationarity, we apply the Augmented Dickey-Fuller(ADF) test and the Phillips and Perron(1986) test procedures to both the level data and the first-differenced series of exchange rates and price indices.

According to the test results of the ADF and Phillips-Perron statistics, for all the level data the hypothesis of a unit root cannot be rejected, but for the first-differenced series the hypothesis of a unit root is rejected. Thus, all the series used in this paper are said to be integrated of first order or to follow the process of I(1).

Specifically, the Augmented Dickey-Fuller test statistics could not reject the hypothsis of a unit root for the U.S. Dollar, Yen, Mark and Canadian Dollar, except for the British Pound. However, both the Z-statistic and the t-statistic by the Phillips-Perron procedure could not reject the hypothesis of a unit root for all the series. We applied the Phillips-Perron tests to the first-differenced series and found that the hypothesis of a unit root were rejected for all the series.

Unit root tests were applied also to the levels and the first-differenced series of the CPI and WPI for the countries concerned. The results show that the hypothesis of a unit root could not be rejected for all the level data but were rejected for all the first-differenced series. Thus, these results indicate that level variables of exchang rates and price indices show nonstationarity, but that they can achieve stationarity by using first- differences of each series.

The estimates of the error correction models (ECMs) are presented in Table 3. According to the Engle and Granger(1987)'s representation theorem, the adjustment mechanism to the long-run equilibrium can be derived from the cointegrating regression. Thus, the ECM estimates of the nominal exchange rates and the domestic and foreign prices can be used to test the hypothesis of error correction mechanism and cointegration for the PPP relationship.

More importantly, the statistical tests for the estimated coefficient of error correction term can decide the existence of cointegration. The insignificant coefficient cannot reject the null hypothesis of no cointegration. The slope coefficient on the error correction term measures the response of exchange rates and prices to restore the PPP equilibrium. Three error correction models are estimated for each cointegrating regression.

In this ECM estimation, we corrected standard errors using the covariance matrix as suggested by Newey and West(1987), and we need not to include the lagged variables of exchange rates and prices in order to adjust the possibility of serial correlation.

In Table 3, the null hypotheses, H_o: ω_1=0 or H_o: ω_2=0 or H_o: ω_3=0, mean that the exchange rates and the domestic and foreign prices are not cointegrated.

Panel A in Table 3 confirms the existence of cointegration only for the U.K. WPI series at the 5% significance level. Panel B shows estimation of the ECM for the cointegrating regression in which domestic prices are regressed as dependent variables. The coefficients on the error correction term are statistically significant for the U.S. WPI and the Japanese WPI. These results indicate cointegration of the exchange rates and prices for the U.S. and Japan, supporting the PPP ralationship. In Panel C, we found statistically significant estimates of error correction term coefficients when the Japanese WPI, the U.K. WPI and the U.K. CPI are used. These results can be interpreted as supportive evidences that the bivariate PPP relationship on the part of Korea can hold for Japan and U.K.

The Panel A of Table 4 shows that the Johansen test is performed in the context of vector autoregressions (VARs). In order to determine the appropriate lag length, we apply the Akaike's Information Criterion (AIC) procedure. We obtained the lag length of 4. We estimated VAR models including a constant term. Panel B presents the estimates of the Johansen trace statistic, $-2\ \ln Q_r$. r indicates the number of significant cointegrating vector.

The Johansen analysis in Panel A tests the null hypotheses that the number of cointegrating vector is at most 2 (H_0: $r \leq 2$), that the number of cointegrating vector is at most 1 (H_0: $r \leq 1$), and that there is no cointegrating vector (H_0:

r=0). Critical values for the Johansen trace statistic are given by 9.90 (H_0: $r \leqq 2$), 20.17 (H_0: $r \leq 1$), and 35.07 (H_0: r=0) at the 5% level.

In the Panel B of Table 4, the Johansen trace statistic could not reject the null hypotheses H_0:$r \leqq 2$ and H_0:$r \leqq 1$ for any currency or price index. However, the null hypothesis H_0:r=0, of no cointegrating vector could be rejected at the 2.5% significance level for the Japanese CPI and the German CPI. And the null hypothesis could be rejected at the 5% significance level for the U.S. CPI, the German WPI and the U.K. CPI.

To summarize the Johansen test results for three null hypotheses in Panel B, we found one cointegrating vector for U.S., Japan, Germany and U.K.. The evidences support the long-run PPP relationship with these four countries. For the Canadian Dollar, no evidence of cointegration is found when we use any price index. We could not find bilateral long-rup PPP relationship with Canada.

In Table 5, we derived cointegrating vectors with respect to r=1 on the basis of the Johansen trace statistic. The implication of r=1 is that there is one cointegrating vector linking the bilateral exchange rate and the domestic and foreign price indices. This cointegrating vector in turn represents a long-run linear equilibrium PPP relationship. he cointegrating vector is interpreted as an error-correction mechanism. For example, there is one cointegrating vector between Korea and U.S. on the basis of CPI. And this long-run equilibrium relationship between the exchange rate and the domestic and foreign prices can be expressed as β'=[1.000, 0.151, 0.085]. We let $R_t'=[S_t, P_t, P_t*]$, then the long-run PPP relationship between Korea and U.S. can be written as $\beta' R_t$= 0. Similarly, the long-run PPP relation- ship can be expressed as β'=[-0.190, -0.129, 1.000] for Japan, β'= [0.169, -0.944, 1.000] and [0.208, -0.626, 1.000] for Germany, and β'= [1.000, 0.213, -0.172] for U.K.. The long-run relationship in Table 5 shows evidences consistent with the PPP hypothesis, since domestic prices have the opposite signs as foreign prices except for U.S. Such a result for the U.S. seemed because a trend was not considered in a vector autoregression (VAR) model.

Table 5 presents also adjustment vector for $R_t'=[S_t, P_t, P_t*]$ with respect to cointegrating vector. Adjustment vector measures the response extent of exchange rate and prices. That is, the adjustment coefficients indicate the speed of adjustment with which each exchange rate and price variable restores the equilibrium relationship, when it deviates from the long-run PPP relationship by a certain shock. In Table 5, the estimated adjustment vector corresponding to each cointegrating vector shows that the estimated coefficients for nominal exchange rates are much higher than those for domestic and foreign prices. This result means that, once the PPP relationship deviates from the equilibrium, the role of exchange rate is much larger than that of prices in restoring equilibrium.

In other words, it means that exchange rate can adjust much faster than prices in order to achieve the PPP equilibrium relationship again.

Ⅳ. Concluding Remarks

In this paper, we obtained some supportive evidence for the long-run PPP relationship concerning the Korean Won currency. Previous tests of PPP in the bilateral exchange rates of the Korean Won rate vis-a-vis the U.S. Dollar have been exposed to the lack of power problem (see Bahmani-Oskooee, Moshen and Rhee, Hyun-Jae(1992)). This was because they applied unit root tests or the Engle-Granger two-step test to the Korean exchange rate data with short sample period.

For example, we argue that Bahmani-Oskooee, Moshen and Rhee, Hyun-Jae (1992)'s failure to find PPP relation in Korean Won rates was due to the low power of Augmented Dickey-Fuller technique applied to Korea's small amounts of data. They concluded that PPP might not hold in Korean Won rate relative to the U.S. Dollar. In attempting to alleviate this low power problem, we used more sophisticated techniques including the error- correction model test and the Johansen test. It is surprising that our evidence supporting for long-run PPP in Korean Won rate contrasts sharply with Bahmani-Oskooee, Moshen and Rhee, Hyun-Jae(1992)'s.

Our test results found in this paper are summarized as follows. First, it was shown that all the series of nominal exchange rates and price indices follow the process of I(1) integrated of first order. Second, the test result of error correction models showed that the PPP relationship might hold in the bilateral exchange rates of Korean Won rate relative to the U.S. Dollar, Japanese Yen and British Pound. Third, the Johansen test results supported that the Korean Won maintains the long-run PPP relationship with the U.S. Dollar, Japanese Yen, German Mark and British Pound, except for the Canadian Dollar. Finally, the estimated adjustment vector showed that exchange rate can adjust much faster than prices in order to restore the PPP equilibrium relationship.

APPENDIX

〈Table 1〉 Tests for Unit Roots in Nominal Exchange Rates

Panel A : Augmented Dickey-Fuller Tests

Model 1 : $\triangle X_t = \alpha + \delta^* X_{t-1} + \sum_{j=1}^{p} \beta_j \triangle X_{t-j} + \upsilon_t$

Model 2 : $\triangle X_t = \alpha + \gamma t + \overline{\delta} X_{t-1} + \sum_{j=1}^{p} \beta_j \triangle X_{t-j} + \upsilon_t$

S_t	$\gamma(\delta^*)$	$\Phi(\alpha\ \delta^*)$	$\gamma(\overline{\delta})$	$\Phi(\alpha\ \gamma\ \overline{\delta})$	$\Phi(\gamma\ \overline{\delta})$
	(-2.57)	(3.78)	(-3.13)	(4.03)	(5.34)
U.S. Dollar	-2.24	2.53	-2.24	1.68	2.51
Yen	-1.58	3.13	-2.54	3.62	3.52
Mark	-1.06	1.01	-2.58	2.56	3.38
Pound I	-3.19	5.13	-3.20	3.44	5.15
Can. Dollar	-2.31	2.69	-2.24	1.94	2.90

Panel B : Phillips-Perron Tests

Model 1 : $\triangle X_t = \alpha + \delta^* X_{t-1} + \upsilon_t$

Model 2 : $\triangle X_t = \alpha + \gamma(t - T/2) + \overline{\delta} X_{t-1} + \upsilon_t$

S_t	$Z(\delta^*)$	$\gamma(\delta^*)$	$\Phi(\alpha\ \delta^*)$	$Z(\overline{\delta})$	$\gamma(\overline{\delta})$	$\Phi(\alpha\ \gamma\ \overline{\delta})$	$\Phi(\gamma\ \overline{\delta})$
	(-11.2)	(-2.57)	(3.78)	(-18.2)	(-3.13)	(4.03)	(5.34)
U.S. Dollar	- 5.02	-3.58	8.70	- 4.06	-2.89	8.34	10.09
Yen	- 2.08	-1.55	4.36	- 7.76	-2.01	3.73	2.46
Mark	- 1.49	-0.72	0.94	- 8.54	-2.14	2.04	2.38
Pound	-11.20	-2.36	2.82	-11.35	-2.38	1.90	2.84
Can. Dollar	- 8.95	-3.16	5.29	- 7.63	-2.72	5.37	7.75

Note : (a) The test statistics $Z(\delta^*)$ and $\gamma(\delta^*)$ are the Z-statistic and t-statistic to test the null hypothesis H_0: $\delta=0$. The test statistic $\Phi(\alpha\ \gamma\ \overline{\delta})$ is the F-statistic to test the null hypothesis H_0: $\alpha=\gamma=\overline{\delta}=0$.

(b) Critical values at the 10% significance level are in parentheses.

(c) In the Augmented Dickey-Fuller Tests, lag length is given by p=q according to the Schwartz criterion. The Z-statistic is not considered since it partly depends upon lag length p.

〈Table 2〉 Tests for Unit Roots in the WPIs and the CPIs

Panel A : Augmented Dickey-Fuller Tests

Model 1 : $\triangle X_t = \alpha + \delta^* X_{t-1} + \sum_{j=1}^{p} \beta_j \triangle X_{t-j} + \upsilon_t$

Model 2 : $\triangle X_t = \alpha + \gamma t + \overline{\delta} X_{t-1} + \sum_{j=1}^{p} \beta_j \triangle X_{t-j} + \upsilon_t$

P_t / P_t^*		$\gamma(\delta^*)$	$\Phi(\alpha\delta^*)$	$\gamma(\overline{\delta})$	$\Phi(\alpha\gamma\overline{\delta})$	$\Phi(\gamma\overline{\delta})$
		(-2.57)	(3.78)	(-3.13)	(4.03)	(5.34)
Korea	WPI	-0.79	1.09	-3.36	4.35	5.70
	CPI	0.88	3.67	-3.09	6.21	5.83
U.S	WPI	0.44	2.46	-2.12	3.48	2.79
	CPI	-0.90	3.04	-2.61	4.24	3.65
Japan	WPI	-1.30	1.47	-1.83	1.60	1.79
	CPI	-1.02	1.63	-1.98	2.23	2.22
Germany	WPI	-2.61	7.32	-2.88	6.20	5.33
	CPI	-0.35	0.78	-3.03	3.58	4.62
U.K.	WPI	-1.96	5.44	-2.88	6.17	5.61
	CPI	-0.84	2.09	-2.07	2.73	2.33
Canada	WPI	-0.66	2.82	-2.32	3.57	2.70
	CPI	-3.09	14.92	-2.65	11.50	6.94

Panel B : Phillips-Perron Tests

Model 1 : $\triangle X_t = \alpha + \delta^* X_{t-1} + \upsilon_t$

Model 2 : $\triangle X_t = \alpha + \gamma(t - T/2) + \overline{\delta} X_{t-1} + \upsilon_t$

P_t / P_t^*		$Z(\delta^*)$	$\tau(\delta^*)$	$\Phi(\alpha\delta^*)$	$Z(\overline{\delta})$	$\tau(\overline{\delta})$	$\Phi(\alpha\gamma\overline{\delta})$	$\Phi(\gamma\overline{\delta})$
		(-11.2)	(-2.57)	(3.78)	(-18.2)	(-3.13)	(4.03)	(5.34)
Korea	WPI	-3.80	-3.89	20.67	-11.12	-4.59	18.35	13.48
	CPI	-1.11	-2.55	44.75	-7.59	-3.15	33.61	7.14
U.S	WPI	-1.43	-1.39	9.87	-8.60	-2.47	8.21	3.32
	CPI	-1.21	-4.85	120.78	-9.91	-4.46	94.77	20.90
Japan	WPI	-0.80	-0.62	1.26	-3.22	-1.26	1.24	0.80
	CPI	-2.34	-2.57	11.65	-17.72	-3.60	11.29	8.46
Germany	WPI	-6.46	-2.45	4.27	-24.76	-4.09	6.57	8.66
	CPI	-0.62	-1.65	51.32	-3.09	-1.67	34.96	2.29
U.K.	WPI	-1.15	-3.34	62.08	-10.002	-3.36	45.48	10.24
	CPI	-1.30	-3.08	40.15	-8.75	-2.61	28.13	7.21
Canada	WPI	-1.70	-2.89	34.81	-7.02	-3.59	27.15	8.46
	CPI	-3.25	-2.10	4.95	-32.10	-4.56	8.93	11.02

Note : (a) The test statistics $Z(\delta^*)$ and $\gamma(\delta^*)$ are the Z-statistic and t-statistic to test the null hypothesis H_o: δ =0. The test statistic $\Phi(\alpha\gamma\overline{\delta})$ is the F-statistic to test the null hypothesis H_o: $\alpha = \gamma = \overline{\delta} = 0$.

(b) Critical values at the 10% significance level are in parentheses.

(c) In the Augmented Dickey-Fuller Tests, lag length is given by p=q according to the Schwartz criterion. The Z-statistic is not considered since it partly depends upon lag length p.

〈Table 3〉 Results of Error-Correction Model Estimation

Panel A : ECM Model $\triangle S_t = \psi_1 + \theta_{11} \triangle P_t + \theta_{21} \triangle P_t^* + \omega_1 V_{t-1} + \varepsilon_t$

S_t/P_{t*}		$\hat{\psi}_1$	$\hat{\theta}_{11}$	$\hat{\theta}_{21}$	$\hat{\omega}_1$	R^2	DW
U.S.	WPI	-0.002 (-2.34)	0.84 (4.69)	0.64 (3.63)	-0.0006 (-0.05)	0.99	0.86
	CPI	-0.004 (-5.82)	0.13 (1.39)	1.42 (15.94)	0.003 (0.45)	0.99	0.74
Japan	WPI	-0.001 (-0.65)	2.53 (5.00)	-1.12 (-2.46)	-0.01 (-0.26)	0.99	1.30
	CPI	0.003 (1.22)	0.25 (0.67)	1.01 (2.96)	-0.02 (-1.43)	0.99	1.26
Germany	WPI	-0.0008 (-0.34)	1.28 (10.60)	0.06 (0.58)	-0.01 (-0.94)	0.99	1.25
	CPI	-0.00006 (-0.02)	-0.04 (-0.11)	1.37 (3.4)	-0.01 (-0.59)	0.99	1.29
U.K.	WPI	-0.006 (-2.47)	1.06 (3.67)	0.66 (2.16)	-0.07 (-2.17)	0.99	1.16
	CPI	-0.009 (-3.63)	0.72 (2.26)	1.09 (3.4)	-0.05 (-1.59)	0.99	1.23
Canada	WPI	-0.003 (-3.35)	0.83 (4.33)	0.63 (3.27)	0.01 (0.72)	0.99	1.48
	CPI	-0.007 (-6.73)	1.49 (31.88)	0.08 (1.80)	0.01 (0.89)	0.99	1.89

Panel B : ECM Model $\triangle P_t = \psi_2 + \theta_{12}\triangle S_t + \theta_{22}\triangle P_t^{*}* + \omega_2 V_{t-1} + \varepsilon_t$

S_t/P_{t*}		$\hat{\psi}_2$	$\hat{\theta}_{12}$	$\hat{\theta}_{22}$	$\hat{\omega}_2$	R^2	DW
U.S.	WPI	0.001 (2.59)	0.28 (7.51)	0.55 (9.73)	-0.06 (-3.10)	0.99	1.48
	CPI	0.002 (3.93)	0.10 (1.43)	0.79 (7.19)	-0.01 (-1.30)	0.99	1.85
Japan	WPI	0.002 (5.50)	0.10 (6.95)	0.77 (43.8)	-0.04 (-3.33)	0.99	1.41
	CPI	0.003 (6.39)	0.01 (0.47)	0.88 (26.17)	-0.004 (-0.31)	0.99	1.81
Germany	WPI	0.001 (1.08)	0.24 (1.95)	0.64 (3.85)	-0.09 (-1.54)	0.99	2.05
	CPI	0.003 (5.51)	-0.009 (-0.37)	0.91 (26.20)	-0.01 (-1.01)	0.99	1.69
U.K.	WPI	-0.001 (-1.61)	0.06 (1.73)	0.94 (14.69	-0.02 (-1.53)	0.99	1.86
	CPI	0.001 (1.49)	0.07 (1.36)	0.87 (9.08)	-0.01 (-0.91)	0.99	2.29
Canada	WPI	0.0004 (0.93)	0.18 (6.05)	0.73 (16.39)	-0.01 (-1.25)	0.99	1.63
	CPI	0.004 (6.30)	0.60 (22.48)	0.03 (0.91)	-0.008 (-0.68)	0.99	1.80

Panel C : ECM Model $\triangle P_t* = \psi_3 + \theta_{13}\triangle P_t + \theta_{23}\triangle S_t* + \omega_3 V_{t-1} + \varepsilon_t$

S_t/P_{t*}		$\hat{\psi}_3$	$\hat{\theta}_{13}$	$\hat{\theta}_{23}$	$\hat{\omega}_3$	R^2	DW
U.S.	WPI	-0.0001 (-0.27)	0.65 (6.51)	0.24 (3.57)	-0.04 (-1.5)	0.99	1.34
	CPI	0.001 (2.81)	0.33 (4.80)	0.43 (10.22)	0.02 (1.36)	0.99	1.11
Japan	WPI	-0.003 (-5.67)	1.20 (36.34)	-0.07 (-2.85)	-0.09 (-2.62)	0.99	1.31
	CPI	-0.004 (-6.49)	0.98 (22.82)	0.08 (2.82)	-0.03 (-0.78)	0.99	1.70
Germany	WPI	0.001 (-1.06)	0.88 (12.28)	0.11 (2.08)	-0.31 (-1.62)	0.99	2.41
	CPI	-0.002 (-5.41)	0.96 (21.39)	0.09 (3.14)	-0.02 (-0.79)	0.99	1.67
U.K.	WPI	0.001 (3.61)	0.83 (21.26)	0.06 (2.90)	-0.01 (-3.24)	0.99	1.99
	CPI	0.0006 (1.11)	0.75 (16.25)	0.13 (5.24)	-0.03 (-2.28)	0.99	2.31
Canada	WPI	0.0007 (1.57)	0.76 (9.55)	0.15 (2.80)	-0.009 (-0.88)	0.99	1.51
	CPI	0.001 (0.75)	0.40 (2.30)	0.38 (3.54)	-0.20 (-1.39)	0.99	2.70

Notes : (a) The WPI (or CPI) for domestic price index P_t is used with respect to foreign price index.

(b) In the tests of error-correction models, standard errors are adjusted using the covariance matrix as suggested by Newey and West(1987). The covariance matrix helps to correct the possible problems of heteroscedasticity and serial correlation.

(c) Corrected standard errors are in parentheses.

<Table 4> Results of the Johansen Trace Test

Panel A :

Model 1 : $\triangle R_t = \mu + \sum_{i=1}^{i-1} \Gamma_i \triangle R_{t-i} + \Gamma_j R_{t-j} + \upsilon_t$

where, $R_t' = [S_t, P_t, P^*_t]$

$\Gamma_i = (-I + \Pi_1 + \cdots + \Pi_i)$

μ is a constant vector, $\Gamma_j = \alpha \beta'$,

and β has the r cointegrating vectors.

Panel B : Johansen Trace Statistic

S_t/P_t^*		H_0 : $r \leq 2$	H_0 : $r \leq 1$	H_0 : $r=0$
U.S.	WPI	0.057	8.725	26.382
	CPI	2.710	9.717	38.902
Japan	WPI	2.069	8.419	26.030
	CPI	1.938	11.114	43.675
Germany	WPI	1.175	14.330	36.907
	CPI	0.253	14.950	44.091
U.K.	WPI	6.680	15.966	30.044
	CPI	3.170	11.333	36.038
Canada	WPI	0.000	7.103	34.925
	CPI	0.226	7.259	28.874

Notes : (a) The WPI (or CPI) for domestic price index P_t is used with respect to foreign price index.

(b) The Johansen trace statistic for the null hypothesis that there are at most r cointegrating vectors, can be written as follows.

$$-2 \, ln \, Q_r = -T \sum_{i=r+1}^{n} \ln(1-\hat{\lambda}_i),$$

where $\hat{\lambda}_i$ are eigenvalue estimates.

(c) Critical values for the trace statistic are given by Johansen and Juselius(1990). They tabulated the distributions of the trace by simulation. VAR models consider a constant, but a drift term is not included.

(d) At the 5% significance level, critical values for the trace statistic are given by 9.90 (H_0: $r \leq 2$), 20.17 (H_0: $r \leq 1$), and 35.07 (H_0: $r=0$). At the 2.5% significance level, critical values are given by 10.71 (H_0: $r \leq 2$), 22.20 (H_0: $r \leq 1$) and 37.60 (H_0: $r=0$).

〈Table 5〉 Estimates for the Cointegrating Vectors and the Adjustment Vectors in the Johansen Test

S_t/P_{t*}		r	Eigen-value	Cointegrating Vector β			Adjustment Vector α		
				S_t	P_t	P_t^*	S_t	P_t	P_t^*
U.S.	CPI	1	0.145	1.000	0.151	0.085	-0.001	-0.0009	-0.0006
Japan	CPI	1	0.161	-0.190	-0.129	1.000	0.006	-0.00008	-0.001
Germany	WPI	1	0.114	0.169	-0.944	1.000	0.0001	0.0008	-0.003
	CPI	1	0.145	0.208	-0.626	1.000	-0.003	0.001	-0.0006I
U.K.	CPI	1	0.125	1.000	0.213	-0.172	0.007	-0.001	0.001

Notes : (a) The WPI (or CPI) for domestic price index P_t is used with respect to foreign price index P_t^*.

(b) In the vector autoregressions (VARs), $\Gamma_j = \alpha \beta'$ and eigenvector β has the r cointegrating vectors.

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