# The impact of the Chinese renminbi on the exports of Korea and Japan to the U.S.

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This paper examines the impacts of the real exchange rates of the Chinese renminbi against the US dollar on the Japanese and the Korean exports to the US. Empirical test results, which have analyzed the quarterly data covering 1986Q1 to 2005Q2, show different long-run impacts of the renminbi in the export functions of the two countries. In particular, according to the estimation of cointegrating vectors, depreciation of the renminbi has a positive impact on the Japanese exports but a negative impact on the Korean exports. However, some stability tests indicate a structural break in the export functions. Different from the case of the estimation for the whole sample period, in empirical tests with recent sub-period data, depreciation of the Chinese renminbi turns out to have positive impact both on the Japanese exports and the Korean exports. In addition, the real GDP of the US turns out to have positive impacts on the exports of the two countries. The exchange rate volatility of the Korean won has negative impact on the Korean exports but a positive impact on the Japanese exports. The short-run dynamics examined by error correction models show similar impacts of the explanatory variables.

JEL Classification: C2, F1, F3
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## 1. Introduction

As is well known, China has maintained a de-facto fixed exchange rate of the Chinese renminbi against the U.S. dollar since 1994. Because the value of the renminbi against the US dollar has been fixed despite the remarkable economic growth and accumulating trade and current account surplus of China for the last decade, the US and other trading partners of China, firmly believing the renminbi was substantially undervalued, urged the Chinese government to revalue the renminbi or to shift to a more flexible exchange rate regime (Chang and Parker, 2004; Funke and Rahn, 2005). Entering 2005, the US government more strongly demanded a revaluation of the renminbi, and the Chinese government finally announced that it would appreciate the value of the renminbi by 2.1 percent on July 21st, 2005. In addition, the Chinese government announced that it would move to a managed float of the renminbi to a basket of currencies (*the New York Times*, 2005 July 22nd; *the Economist*, 2005 July 28th).

The appreciation of the renminbi is, in general, expected to have positive impacts on the exports of other East Asian countries because China is known as their major competitor in the world market. However, despite the hot debate surrounding the value of the renminbi and the importance of this issue, the effect of the value of the Chinese renminbi on the exports of other East Asian countries has rarely been explored.[1]

[1] Some papers investigate the impacts of the renminbi on the Chinese trade (Zhang , 2001; Chou 2000; Tang, 2003). Wang, Wang and Zhang (2003) examine the effects of Japanese yen's depreciation on Chinese exports. Bhattacharya, Ghosh, and Jansen (2001) investigate whether the emergence of China hurt Asian exports.

Against this background, this research aims to determine the effect of the value of the Chinese renminbi on the Japanese and the Korean export volume to the US. In fact, the US is one of the most important trading partners of Japan and Korea[2]. As of 2004, the share of the US market in the Japanese exports was 22.7 percent, exceeding that of the European Union market (15.78 %) or the Chinese market (13.1 %). In the meantime, the share of the US market in the Korean exports was 17.0 percent, slightly lower than only that of the Chinese market (19.6 %) and exceeding that of the European Union market (14.9 %) or the Japanese market (8.6).

To determine the impacts of the exchange rates of the renminbi on the Japanese and the Korean exports to the US, this paper analyzes the quarterly trade data for the period from 1986Q1 to 2005Q2. Specifically, following the work of Arize, Osang and Slottje (2000), Baak et al. (2006), Baum et al. (2002), Chou (2000), Chowdhury (1993) and Hassan and Tufte (1998) among others, this study examines the long-run relationship between the exports from one country to the other and other economic factors including the real exchange rate of the Chinese renminbi by performing cointegration tests. In addition, the short-run impacts of the real exchange rate are examined by estimating error-correction models, if the variables involved are cointegrated.

In particular, the volume of the real exports from Japan or Korea to the US is a function of the bilateral real exchange rates between the exporting county (Japan or Korea) and the importing country (the US) and other economic variables, such as a measure of economic activity of the US and exchange rate volatility. In addition, the

[2] The following numbers were calculated by the author from the data obtained from the *Direction of Trade Statistics (DOTS)* of the IMF.

exchange rate of the renminbi against the US dollar is also included as an explanatory variable in the two export functions.[3]

Empirical test results, which have analyzed the quarterly data covering 1986Q1 to 2005Q2, show different long-run impacts of the renminbi in the export functions of the two countries. In particular, according to the estimation of cointegrating vectors, depreciation of the renminbi has a positive impact on the Japanese exports but a negative impact on the Korean exports.

However, stability tests such as the CUSUM test and the tests suggested by Hansen (1992a, 992b) indicate that the export functions should be very unstable, implying the presence of structural breaks. Accordingly, cointegration tests such as the test (S-L cointegration test, hereafter) suggested by Saikkonen and Lutkepohl (2000a, 2000b, 2000c) and the test (J cointegration test, hereafter) suggested by Johansen et al. (2000), which allow a structural break in the cointegrating vector, are performed. The results also confirm that the variables are cointegrated in each export function. Then, the export functions are re-estimated for the recent sub-period (1994Q1 to 2005Q2). The CUSUM tests show that the export functions are stable for this sub-period.[4]

Different from the case of the estimation for the whole sample period, in empirical tests with the sub-period data, depreciation of the Chinese renminbi turns out to have positive impact both on the Japanese exports and the Korean exports. In addition, the real GDP of the US turns out to have positive impacts on the exports of the two countries. The exchange rate volatility of the Korean won has negative impact on the Korean exports but a positive impact on the Japanese exports. The

[3]More detailed explanation will be provided in the following sections.

short-run dynamics examined by error correction models show similar impacts of the explanatory variables.

## 2. The models and the data

### 2.1. The export functions

This paper performs cointegration tests with the export functions (the export function of Japan to the US and the export function of Korea to the US) and estimates the coefficients of the functions to understand the long-run relationship between the export volumes and the explanatory variables. In addition, this paper examines the short-run dynamics of the export functions by estimating error-correction models, if the variables are cointegrated.[5]

Following the typical specification of other papers, an export function (or, a long-run equilibrium relation between exports and other economic variables) is assumed to have the following functional form:

$$Y_{ijt} = \xi_0 + \xi_1 . g_{jt} + \xi_2 p_{ijt} + \xi_3 \sigma_{ijt} + \xi_4 p_{cjt} + \varepsilon_{ijt} \qquad \text{------ (1)}$$

where $Y_{ijt}$ denotes real exports from country *i* to country *j*. Therefore, *i* denotes the exporting country and *j* the importing country. In this paper *i* is Japan or Korea, and

[4] The Hansen (1992a, 1992b) tests show mixed results.
[5] Similar methodological approaches regarding export or import functions of various countries can be found in the papers of Arize, Osang and Slottje (1999, 2000), Chowdhury (1993), Hassan and Tufte (1998), Chou (2000), Zhang (2001) and Tang (2003).

$j$ is the US. The variable $g_{jt}$ denotes the measure of economic activity of the importing country, $j$ (that is, the US).

The variables, $p_{ijt}$ and $p_{cjt}$, are real bilateral exchange rates. $p_{ijt}$ is the exchange rate of the exporting country $i$'s currency against the importing country $j$'s currency. Therefore, if $p_{ijt}$ rises, the products of exporting country $i$ becomes cheaper. $p_{cjt}$ is the exchange rate of a country $c$'s currency against the importing country $j$'s currency, and country $c$ is a competitor to country $i$ in the market of country $j$.[6] In this paper, country $c$ is China.

Finally, $\sigma_{ijt}$ denotes the volatility of the real bilateral exchange rates between country $i$ and country $j$, and $\varepsilon_{ijt}$ a disturbance term. All variables are in natural logarithm and the subscript $t$ symbolizes the time.

It is expected that the higher the economic activity in the importing country, the higher the demand for exports. Therefore, the value for $\xi_1$ is expected to be positive. Since a higher real exchange rate implies a lower relative price of the exporting products, the value for $\xi_2$ is also expected to be positive. In contrast, since low prices of the competitor's products (that is, higer $p_{cjt}$) will have negative impacts on the exports of country $i$, the value for $\xi_3$ is expected to be negative.

Extant theoretical and empirical papers have shown that exchange rate volatility may have either positive or negative influences on trade, depending on various

[6] The exchange rate of a competing country ( $p_{cjt}$ in this paper) is not included in the papers mentioned in footnote 5. However, various estimation experiments performed by the author showed its coefficients are significant and are not ignorable. The selection of country $c$ and more detailed reports are presented in sections 3 and 4.

economic and institutional environment.[7] However, if economic agents are moderately risk averse, it is generally expected that the impact of exchange rate volatility will be negative. In this case, the value for $\xi_4$ will be negative.

Subsection 2.3 shows more specifically how the data for the variables are computed.

## 2.2. The error-correction model

After observing the results of cointegration tests with equation (1), the following dynamic error correction (EC) model is constructed and estimated to see the short-run impacts of the explanatory variables on the exports:

$$\begin{aligned} \Delta Y_{ijt} = \alpha + \lambda EC_{ijt-1} + \sum_{h=0}^{nx} \beta_h \Delta Y_{ijt-h-1} + \sum_{h=o}^{np} \gamma_h \Delta p_{ijt-h} + \\ + \sum_{h=0}^{ng} \delta_h \Delta g_{jt-h} + \sum_{h=0}^{ns} \eta_h \Delta \sigma_{ijt-h} + \sum_{h=o}^{nc} \varphi_h \Delta p_{cjt-h} + u_{ijt} \end{aligned} \quad (2)$$

where *nx, np, ng, ns*, and *nc* are the lengths of included lags for each variable. If the variables in equation (1) are not cointegrated, the error correction term, $EC_{ijt-1}$, is eliminated from equation (2). In addition, lots of estimation experiments are performed to find a parsimonious structure of equation (2). In other words, variables which are insignificant and do not generate, even though omitted, any noticeable difference in the estimation results are eliminated from equation (2).

[7] See Secru and Uppal (2000) and their references.

## 2.3. The variables[8]

**Real exports ($Y_{ijt}$)**

The real export volume of country *i* to country *j* is defined as follows:

$$Y_{ijt} = \ln\left(\frac{EX_{ijt}}{EXUV_{it}} \times 100\right), \qquad (i = \text{Japan or Korea}; j = \text{the US})$$

where $Y_{ijt}$ denotes the log value of the real exports of country *i* to country *j*; $EX_{ijt}$ is the quarterly nominal exports of country *i* to country *j*; and $EXUV_{it}$ denotes the export unit value index of country *i*.

**Real GDP ($g_{jt}$)**

The real GDP of the importing country (country *j*) is commonly used as a proxy measure for economic activity of the importing country in much literature dealing with quarterly or annual data. Accordingly, the variable $g_{jt}$ in equation (1) is defined to be the real GDP of the US.

[8] In order to ensure consistency in data, variables, which were not seasonally pre-adjusted, were adjusted for seasonality prior to taking logarithm by applying the method Census X12 available in the software package E-views 4.

**Real bilateral exchange rates ( $p_{ijt}$ , $p_{cjt}$ )**

The real exchange rates are computed in the conventional way as follows:

$$p_{ijt} = \ln\left( E_{ijt} \times \frac{CPI_{jt}}{CPI_{it}} \right)$$

where $p_{ijt}$ symbolizes the real quarterly exchange rate in natural logarithm scale; $E_{ijt}$ is the nominal quarterly exchange rate of country *i*'s currency against country *j*'s currency; $CPI_{it}$ and $CPI_{jt}$ denote the quarterly consumer price index of an exporting country *i* and an importing country *j*, respectively.

The exchange rate of country *c*'s currency against the importing country *j*'s currency, $p_{cjt}$, is also computed in the same way, with the change that the subscript *i* is replaced by the subscript *c* in the formula above. Country c is a country which is competing with country *i* in the market of country *j*.

In the case of China, consumer price indices are not reported. Instead, the annual growth rates of monthly indices from 1986 are reported. The Chinese monthly consumer price indices are computed using these growth rates and the consumer price indices for the one year from December 2000 to November 2001.[9] Then, quarterly data are computed from these monthly data.

**Real exchange rate volatility ( $\sigma_{ijt}$ )**

[9] The Chinese consumer price indices from December 2000 to November 2001 were kindly provided by Yuqing Xing at the International University of Japan.

This present study applies the standard deviation of exchange rates as the measure of the exchange rate volatility.[10] Specifically, the real exchange rate volatility $\sigma_{ijt}$ is defined as the natural logarithm of the standard deviation of monthly real exchange rates for a certain time period:

$$\sigma_{ijt} = \ln\left(\sqrt{\frac{1}{n-1}\sum_{k=tm}^{tn}\left(RER_{ijk} - \overline{RER_{ij}}\right)^2}\right),$$

where *t* represents a quarter and *k* a month. $RER_{ijk}$ is a monthly real exchange rate, $\overline{RER}_{ij}$ is the mean of $RER_{ijk}$'s from *k*=*tm* to *k*=*tn*. *tm* and *tn* are the last and the first month included in the computation of $\sigma_{ijt}$, respectively. *k*=0 is defined to be the last month in quarter *t*, *k*=1 is one month earlier than that, and so on. If *t* is the first quarter of 2000, *tm* is 1, and *tn* is 4, for example, then *tm* represents February 2000 and *tn* November 1999. In empirical tests in section 4, *tm* and *tn* are set to be 0 and 5 respectively. Therefore, the exchange rate volatility of a quarter is computed by the standard deviation of monthly exchange rates of the current and the one lagged quarter.[11]

## 3. Empirical test results

[10] As Sercu and Uppal (2000) mention, this is one of the major ways to measure the exchange rate volatility. For example, see Akhtar and Hilton (1984), Côté (1994) and Baum et al. (2002).
[11] Lots of preliminary tests showed this setting generated the best results. For example, if we set *tm*=0 and *tn*=2, the volatility is computed using the monthly exchange rates of only the current quarter, but this change does not improve the test results at all.

### 3.1. Unit Root tests

As preparation for cointegration tests, the presence of unit roots in the variables included in equation (1) is examined using the augmented Dickey-Fuller (ADF) tests. Based on the visual examination of the time series, it is decided whether a trend in included in the test equation. The lengths of the lags included in the tests are determined by the Modified Akaike infomation criterion.

The ADF statistics for the levels of all the series are below the 5 percent critical values, implying the presence of unit roots. In contrast, the statistics obtained from the first differences of the variables reject the null hypothesis of a unit root at the 5 percent significance level with one exception. In the case of the first difference of the Korean export volume, the null is rejected at the 10 percent significance level. Tables 1-1 and 1-2 present the ADF test statistics for all the variables in equation (1) for both the Japanese and the Korean exports.

<Insert Table 1-1, Table 1-2 and Table 1-3>

However, because the dynamics of the exchange rate data and the volatility data, as is shown in Figure 1, illustrate drastic change, unit root tests with structural break (S-L unit root test, hereafter) suggested by Saikkonen and Lutkepohl (2002) are also performed.  The test statistics confirms the results of the ADF tests as is shown in Table 1-3

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### 3.2. Cointegration tests

Because all the variables involved have unit roots, cointegration tests are performed to examine whether the variables in each export function illustrated in equation (1) (the function for exports from Japan to the US and the function for exports from Korea to the US) have a long-run relationship.

Empirical test results[12], which have analyzed the quarterly data covering 1986Q1 to 2005Q2, detected a cointegrating relationship between the export volume and the explanatory variables, and the cointegrating vectors estimated by the OLS are presented in Table 3-1. The estimation results show different long-run impacts of the renminbi in the export functions of the two countries. In particular, depreciation of the renminbi has a positive impact on the Japanese exports but a negative impact on the Korean exports.

However, stability tests[13] such as the CUSUM test and the tests suggested by Hansen (1992a, 992b) indicated that the export functions should be very unstable, implying the presence of structural breaks. Accordingly, cointegration tests such as the test (S-L cointegration test, hereafter) suggested by Saikkonen and Lutkepohl (2000a, 2000b, 2000c) and the test (J cointegration test, hereafter) suggested by Johansen et al. (2000), which allow a structural break in the cointegrating vector, are performed. The results, reported in Table 2, also confirm that the variables are cointegrated in each export function. Then, the export functions are re-estimated for the recent sub-period (1994Q1 to 2005Q2). The estimation results are reported in Table 3-2 and Table 3-3. The OLS estimation and the fully modified OLS estimation

[12] The test results are not reported in this paper.
[13] The test results are not reported in this paper.

proposed by Phillips and Hansen (1990) show similar results. The CUSUM tests illustrated in Figure 2 show that the export functions are stable for this sub-period.[14]

<Insert Tables 3-1, 3-2, 3-3>

<Insert Figure 2>

Different from the case of the estimation for the whole sample period, in empirical tests with the sub-period data, depreciation of the Chinese renminbi turns out to have positive impact both on the Japanese exports and the Korean exports. In addition, the real GDP of the US turns out to have positive impacts on the exports of the two countries. The exchange rate volatility of the Korean won has negative impact on the Korean exports but a positive impact on the Japanese exports.

### 3.3. Error Correction models

Since the cointegration tests in the previous section detected one long-run equilibrium relationship for each of the export functions, error correction models illustrated in equation (2) are estimated to see the short-run dynamics of the export functions. The error correction terms are computed by the cointegration vectors reported in Tables 3-3.

Each error correction model is estimated in the first step with long lags of each explanatory variable, and the number of lagged variables is reduced in a way to

[14] The Hansen (1992a, 1992b) tests show mixed results.

increase the adjusted $R^2$'s. In other words, variables which are insignificant and do not generate, even though omitted, any noticeable difference in the estimation results are eliminated from equation (2) to find a parsimonious structure of the error correction models.

In addition, to examine the stability of the estimates, the CUSUM statistics of the estimations of the error-correction models are computed and illustrated in Figure 3. As shown, the CUSUM statistics are within the 95 percent confidence bands, implying no structural break for the time period from 1994Q3 to 2005Q2.

<Insert Figure 3>

The estimated values of the error correction models are presented in Table 5.

<Insert Table 5>

As can be seen from the tables, the estimated coefficient values of the error-correction terms in all the models are negative and significant at the 5 percent significance level, confirming the presence of one long-run relationship among the variables involved. The short-run dynamics examined by error correction models show similar impacts of the explanatory variables.

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## Appendix

### Data Sources

Consumer Price Indices (CPI) of the US, annual growth rates of monthly CPI of China, the quarterly real GDP of the US, the annual nominal GDP of China, the Chinese annual GDP deflators, the US Export Unit Value Indices, and the US Import Unit Value Indices have been collected from the *International Financial Statistics (IFS)* of the International Monetary Fund (IMF).

The data for the US exports to China and the data for the US imports from China have been obtained from the *Direction of Trade Statistics (DOTS)* of the IMF.

The data starts from the first quarter of 1980 and ends at the fourth quarter of 2003, except for the annual growth rates of monthly CPI of China which are available only from 1986. Because of this restriction the empirical tests in this paper covers the period from the first quarter of 1986 to the fourth quarter of 2003.

**<Table 1-1> ADF Unit Root Test for the levels**

| Variable | Lags[1)] | Trend | ADF Statistic | P-Value |
|---|---|---|---|---|
| $Y_t^J$ | 1 | included | -2.384 | 0.386 |
| $Y_t^K$ | 0 | included | -1.802 | 0.697 |
| $g_t$ | 1 | included | -2.451 | 0.352 |
| $p_t^J$ | 1 | included | -1.617 | 0.779 |
| $p_t^K$ | 4 | included | -1.942 | 0.625 |
| $p_t^C$ | 0 | included | -2.452 | 0.351 |
| $\sigma_t^J$ | 5 | not included | -2.800 | 0.062[2)] |
| $\sigma_t^K$ | 4 | not included | -2.254 | 0.189 |

1) The lags were determined by the Modified Akaike Information Criterion.
2) AIC, lag length is 0. Then, no unit root.

**<Table 1-2> ADF Unit Root Test for the first differences**

| Variable | Lags | ADF Statistic | P-Value (5%) |
|---|---|---|---|
| $\Delta Y_t^J$ | 2 | -4.296 | 0.001 |
| $\Delta Y_t^K$ | 6 | -2.695 | 0.079[1)] |
| $\Delta g_t$ | 1 | -5.393 | 0.000 |
| $\Delta p_t^J$ | 2 | -4.177 | 0.001 |
| $\Delta p_t^K$ | 0 | -7.071 | 0.000 |
| $\Delta p_t^C$ | 1 | -5.503 | 0.000 |
| $\Delta \sigma_t^J$ | 0 | -14.471 | 0.000 |
| $\Delta \sigma_t^K$ | 0 | -12.601 | 0.000 |

1) sensitive to lag length. According to AIC, lag length is 0. Then, the p-value is 0.000.
2) only intercept is included in all tests.

**<Table 1-3> SL Unit Root Test with a structural break**

| Variable | Trend | Suggested break[1)] | SL Statistic | Critical values[2)] 1% | 5% | 10% |
|---|---|---|---|---|---|---|
| $p_t^J$ | included | 1995Q3 | -1.879 | -3.55 | -3.03 | -2.76 |
| $p_t^K$ | included | 1998Q1 | -1.480 | -3.55 | -3.03 | -2.76 |
| $p_t^C$ | included | 1994Q1 | -2.641 | -3.55 | -3.03 | -2.76 |
| $\sigma_t^J$ [3)] | not included | 1995Q1 | -1.657 | -3.48 | -2.88 | -2.58 |
| $\sigma_t^K$ | not included | 1997Q4 | -1.256 | -3.48 | -2.88 | -2.58 |

1) break suggested by JMulTi.
2) Critical values for the null hypothesis of unit root suggested by Lanne et al. (2002).
3) Depending the lag length, a different break is detected. But the result of unit root test is not affected.
4) Lag length is 4. Different standards suggest different lags. Changing the lag sometimes change the results, but evidence of unit root is stronger.

**<Table 2> Cointegration tests with a structural break**

| Statistic | $H_0$: / $H_A$: | $r = 0$ / $r \geq 1$ | $r \leq 1$ / $r \geq 2$ | $r \leq 2$ / $r \geq 3$ | $r \leq 3$ / $r = 4$ | $r \leq 4$ / $r = 5$ |
|---|---|---|---|---|---|---|
| Japanese Exports to US | | | | | | |
| Johansen Statistic[(3)] | | 125.35* | 71.64 | 38.63 | 20.72 | 5.00 |
| (p-value) | | 0.006 | 0.294 | 0.733 | 0.765 | 0.947 |
| S-L Statistic[(4)] | | 67.97* | 27.50 | 12.22 | 3.39 | 0.02 |
| (p-value) | | 0.034 | 0.779 | 0.917 | 0.969 | 0.999 |
| Korean Exports to US | | | | | | |
| Johansen Statistic[(3)] | | 109.01* | 56.66 | 30.41 | 14.15 | 5.98 |
| (p-value) | | 0.093 | 0.833 | 0.962 | 0.979 | 0.895 |
| S-L Statistic[(4)] | | 63.79* | 33.46 | 20.14 | 4.20 | 1.15 |
| (p-value) | | 0.079 | 0.448 | 0.383 | 0.929 | 0.753 |

Notes: (1) *r* denotes the number of co-integrating vectors. (2) The asterisk (*) indicates the rejection of the null hypothesis of no cointegration at the 10% significance level. (3) Johansen et al. (2002). (4) Saikkonen and Lutkepohl (2000a,b,c)

**<Table 3-1> Estimates of the contegrating vectors by OLS for the whole period**

| | c | $g_t$ | $p_t$ | $p_t^c$ | $\sigma_t$ | trend | $R^2$ |
|---|---|---|---|---|---|---|---|
| Japanese Exports to US | | | | | | | |
| Coeff. | -19.44*** | 3.295*** | 0.120 | 0.253*** | 0.007 | -0.019*** | 0.94 |
| Std. error | 5.819 | 0.668 | 0.090 | 0.076 | 0.008 | 0.005 | 0.93[2)] |
| Korean Exports to US | | | | | | | |
| Coeff. | -49.07*** | 5.984*** | 0.791*** | -0.109 | -0.034 | -0.025* | 0.94 |
| Std. error | 14.12 | 1.699 | 0.187 | 0.226 | 0.039 | 0.013 | 0.94[2)] |

Notes: (1) The whole period is from 1986Q2 to 2005Q2. (2) Adjusted R-square. (3)Standard errors were computed by the method of Newey and West (1987). (4) The asterisks (*), (**) and (***) indicate the rejection of the null hypothesis of zero coefficient at the 10%, 5% and 1% significance level, respectively.

**<Table 3-2> Estimates of the contegrating vectors by OLS for the second time period**

| | c | $g_t$ | $p_t$ | $p_t^c$ | $\sigma_t$ | trend | $R^2$ |
|---|---|---|---|---|---|---|---|
| Japanese Exports to US | | | | | | | |
| Coeff. | -28.97*** | 4.306*** | 0.225** | 0.544*** | 0.010 | -0.029*** | 0.95 |
| Std. error | 4.320 | 0.485 | 0.102 | 0.168 | 0.006 | 0.004 | 0.94[2)] |
| Korean Exports to US | | | | | | | |
| Coeff. | -40.35*** | 5.013*** | 0.398*** | 0.700* | -0.037*** | -0.010 | 0.98 |
| Std. error | 10.77 | 1.267 | 0.112 | 0.359 | 0.013 | 0.010 | 0.98[2)] |

Notes: (1) The second time period is from 1994Q3 to 2005Q2. See also the notes of table 3-1.

**<Table 3-3> Estimates of the contegrating vectors by fully modified OLS for the second time period**

| | c | $g_t$ | $p_t$ | $p_t^c$ | $\sigma_t$ | trend | $L^c$ |
|---|---|---|---|---|---|---|---|
| Japanese Exports to US | | | | | | | |
| Coeff. | -34.39*** | 4.615*** | 0.267*** | 0.935*** | 0.009 | -0.034*** | 0.655 |
| Std. error | 3.316 | 0.350 | 0.067 | 0.171 | 0.006 | 0.003 | 0.2[1)] |
| Korean Exports to US | | | | | | | |
| Coeff. | -49.11*** | 5.858*** | 0.397*** | 1.009*** | -0.044*** | -0.018*** | 1.275 |
| Std. error | 7.592 | 0.854 | 0.112 | 0.361 | 0.016 | 0.007 | 0.019 |

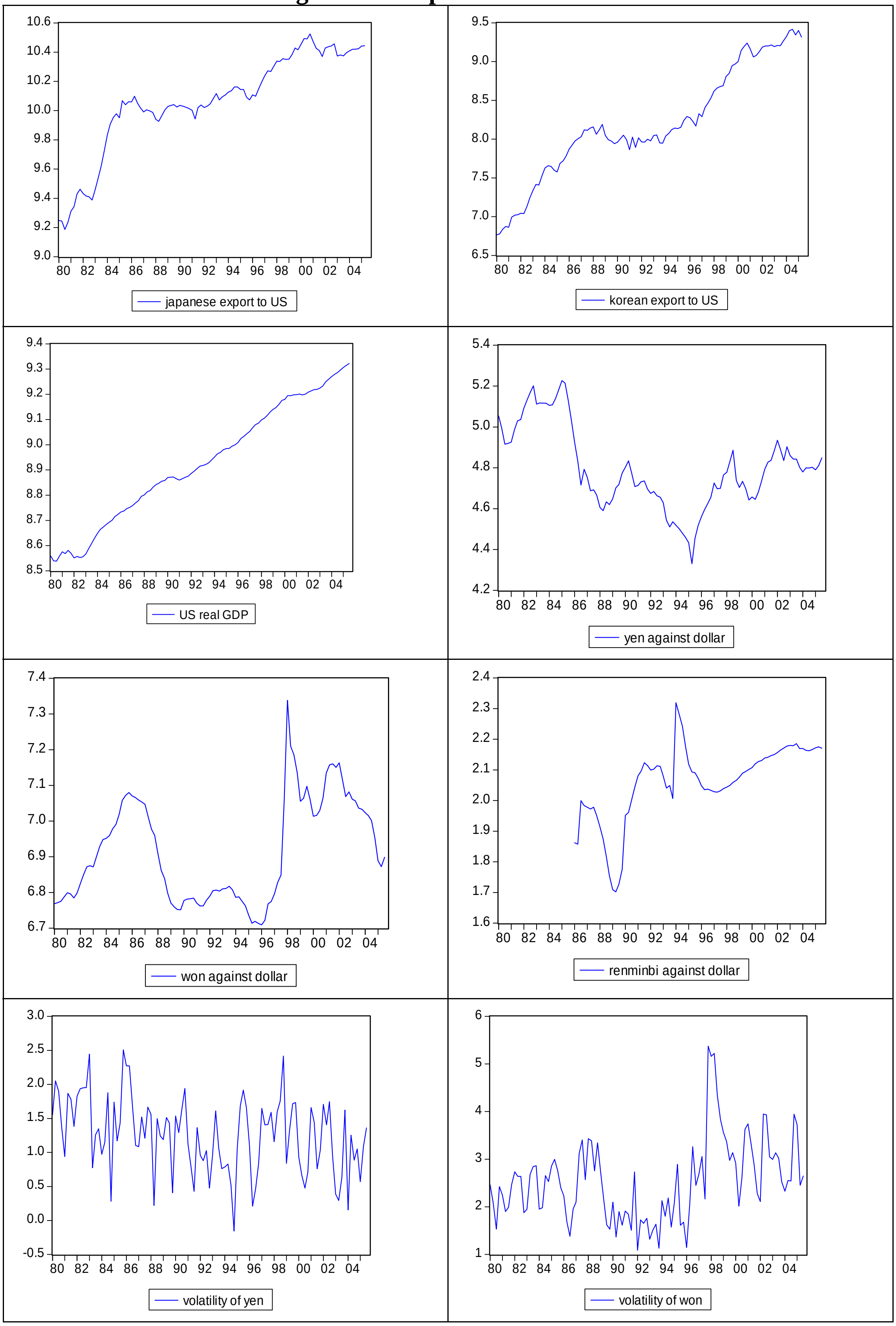
Notes: See the notes of table 3-1.
The second time period is from 1994Q3 to 2005Q2.
1) p-value is higher than 0.2. therefore, the null of stability is accepted.

**<Table 5> Estimates of the Error Correction Models**

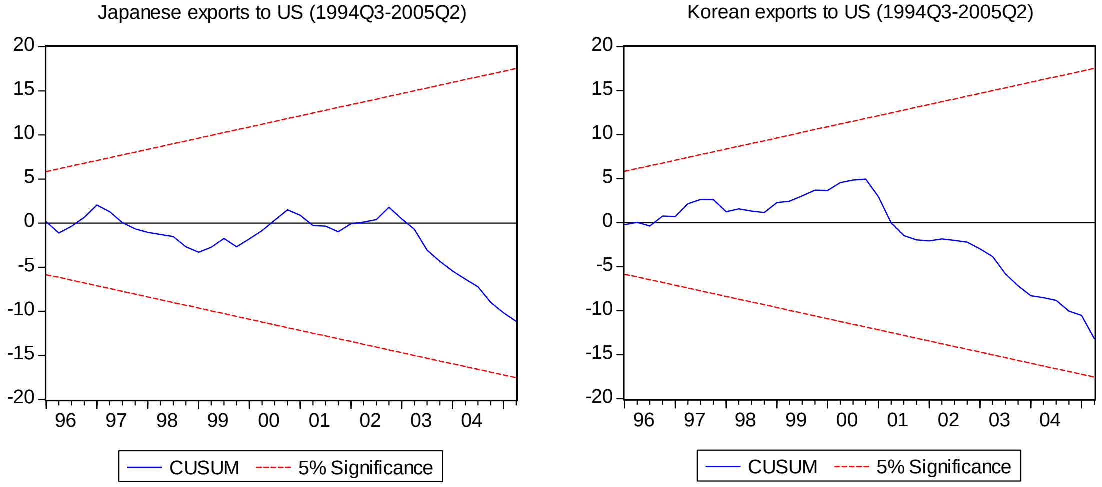
| Variables | [1] Japanese exports to US | | [2] Korean exports to US | |
|---|---|---|---|---|
| | coefficient | Std. error | coefficient | Std. error |
| C | 1.649*** | 0.314 | 0.413** | 0.182 |
| $EC_{t-1}$ | -0.862*** | 0.162 | -0.386** | 0.164 |
| $\Delta Y_{t-1}$ | 0.233* | 0.121 | -0.358* | 0.175 |
| $\Delta Y_{t-2}$ | 0.300*** | 0.104 | 0.413*** | 0.142 |
| $\Delta Y_{t-3}$ | 0.269*** | 0.091 | | |
| $\Delta Y_{t-4}$ | -0.308*** | 0.094 | -0.337** | 0.135 |
| $\Delta g_t$ | 2.672*** | 0.684 | | |
| $\Delta g_{t-1}$ | -1.443 | 0.919 | | |
| $\Delta g_{t-2}$ | | | 4.522** | 1.634 |
| $\Delta g_{t-3}$ | | | 4.900** | 1.826 |
| $\Delta g_{t-4}$ | | | -2.508 | 1.588 |
| $\Delta g_{t-5}$ | | | -4.753*** | 1.468 |
| $\Delta p_t$ | 0.227*** | 0.062 | 0.193* | 0.101 |
| $\Delta p_{t-1}$ | -0.117* | 0.062 | | |
| $\Delta p_{t-2}$ | | | 0.485*** | 0.149 |
| $\Delta p_{t-3}$ | -0.120** | 0.058 | -0.300** | 0.113 |
| $\Delta p_{t-4}$ | 0.180*** | 0.057 | | |
| $\Delta p_t^C$ | | | 2.459** | 0.975 |
| $\Delta p_{t-1}^C$ | | | 3.913*** | 1.078 |
| $\Delta p_{t-2}^C$ | | | | |
| $\Delta p_{t-3}^C$ | 0.360** | 0.159 | -3.197*** | 0.568 |
| $\Delta p_{t-4}^C$ | -0.195*** | 0.062 | | |
| $\Delta \sigma_t^K$ | 0.012*** | 0.004 | -0.028*** | 0.008 |
| $\Delta \sigma_{t-1}^K$ | | | -0.028** | 0.013 |
| $\Delta \sigma_{t-2}^K$ | | | -0.023** | 0.011 |
| $\Delta \sigma_{t-3}^K$ | | | -0.016 | 0.011 |
| $\Delta \sigma_{t-4}^K$ | | | 0.023** | 0.009 |
| $R^2$ | 0.801 | | 0.862 | |
| Adjusted $R^2$ | 0.698 | | 0.731 | |
| Breusch-Godfrey (p-value) | 4.795 (0.309) | | 6.065 (0.194) | |

(1) The asterisks (*), (**) and (***) indicate the rejection of the null hypothesis of a zero coefficient at the 10%, 5% and 1% significance level, respectively.

# <Figure 1> Graphs of the variables



## <Figure 2> CUSUM Tests for the Cointegrating vectors estimated by OLS (1994Q3-2005Q2)




## <Figure 3> CUSUM Tests for the Error Correction Models (1994Q3-2005Q2)