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Demand for International Reserves: A Case Study for Korea*

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This study investigates the Korean reserve demand and its structural change after the 1997 Asian financial crisis. The reserve dynamics for the pre- and post-crisis periods are reasonably specified by the error correction model. In general, our empirical results may provide the evidence that the Korean reserve demand became more sensitive to the adjustment cost and the openness, but less sensitive to the opportunity cost after the crisis, which is consistent with the rapid reserve accumulation of Korea after the crisis. However, our structural break tests provide mixed evidences for the change in the reserve demand after the crisis.

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Keywords: international reserve demand of Korea, buffer stock model, error correction model, dynamic ordinary least squares, structural break in the reserve demand

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1. INTRODUCTION

One of the most remarkable features of the recent trend in the international macro economy is a large international reserve (reserves hereafter) accumulation by East Asian countries.[1)] After the Asian crisis in 1997, East Asian countries built up huge amounts of reserves, so that by the end of 2005, China, Korea, Malaysia, Thailand, and Indonesia held roughly 28.1% of the world's total reserves.[2)] These countries have systematically increased their reserves over the last several years. Measured as a percentage of the total GDP of these five countries, their total reserves were about 10% by the end of the 1980s, but grew to 33% by 2005.

There is an ongoing debate about the need to hold such large reserves. On the one hand, holding large reserves can be viewed as a precautionary motive against a potential financial crisis resulting from a sudden stop of capital flow; on the other hand, it can be explained by a mercantilist view, in which reserves are accumulated to promote the exports by either depreciating or slowing the appreciation of the currency. Also, some critics point out that holding large reserves is costly. The yield on reserves invested in U.S. Treasury bonds is much lower than the opportunity cost of holding those reserves.[3)] On the contrary, proponents assert that the opportunity cost is smaller than the potential cost of another crisis.

For policymakers, holding large reserves seems to be a persuasive and safe choice. The more insecure the international monetary system is, the more reserves are needed to manage the currency and to defend against a financial crisis. A country facing a crisis might be shut out of the international capital markets because of sovereign risk concerns.

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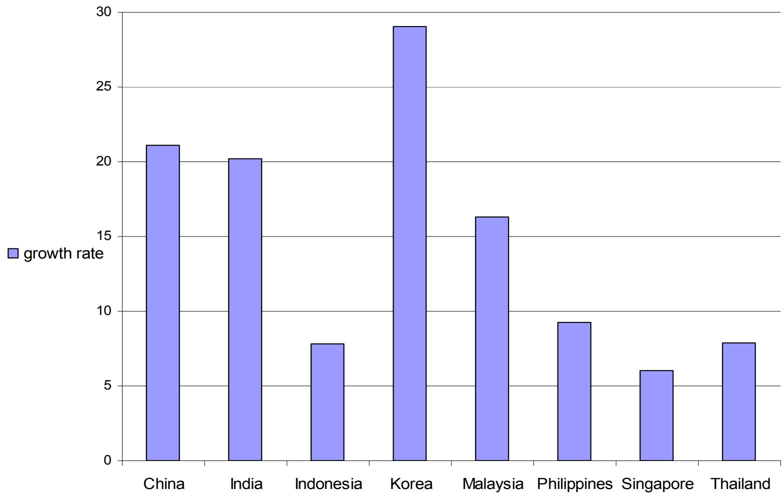
1) Reserves are usually measured as the sum of gold, convertible foreign exchange, the unconditional drawing right with the IMF, and special drawing rights.

2) The total reserves of the world are 4,041 billion U.S. dollars, and the total reserves of these five countries are 1,136 billion U.S. dollars (calculated from the International Financial Statistics (IFS) of IMF).

3) For central banks, the opportunity cost of holding current reserves is the best alternative that is given up; for example, reserves are usually invested in U.S. Treasury bonds, with a yield much lower than the expected return on local investments, giving a net opportunity cost of the difference between the two.

Figure 1 Annual Growth Rates of Reserves in Selected Asian Countries (1997-2005)



Holding large reserves has become typical in Korea since the Asian crisis. On February 15, 2005, the Bank of Korea announced that Korea's reserves exceeded U.S. $200 billion for the first time. Now, Korea ranks as the fourth largest reserve holder in the world, following China, Japan, and Taiwan. This is a remarkable turnaround from the U.S. $20 billion available at the end of 1997. Figure 1 shows the annual growth rate of the reserves of selected Asian countries. Korea marks the highest growth rate of reserves during the period 1997-2005 (29%). Especially when compared with the growth rates of three countries that were severely afflicted by the crisis (Indonesia, the Philippines, and Thailand), the growth rate of Korean reserves is remarkably high.

Also, Korean reserves have been continuously increasing since the 1960s, except for the period of the crisis in 1997, and there has been a more rapid increase in reserves since the crisis in 1997. In September 2001, the reserves surpassed U.S. $100 billion and climbed to U.S. $150 billion in November 2003. Finally they reached U.S. $200 billion in February 2005. It took less than 4 years to double the size. By the end of 2005, reserves accounted for about 28% of the GDP of Korea,[4] compared with about 7% at the start of the 1990s. Reserves covered 40.2 weeks of imports, up from 13 weeks at the start of the 1990s. Reserves, as a share of M2, almost doubled after 1990, to about 26%.

A debate on the optimal level of reserves for Korea is now under way also. Critics argue that Korea is paying unnecessarily high interest rates for reserves;[5] on the other hand, the BOK asserts that Korea, as a small open economy, needs to accumulate sufficient reserves to cope with an unexpected external shock like the Asian crisis (Korea Times, January 17, 2002). Also, Korea's dramatic change in regard to reserve holding and the rapid increase in reserves are introduced in several recent studies as a typical example of rapid reserve accumulation (Aizenman and Marion, 2003a, 2003b, and 2004; Aizenman and Lee, 2005). Aizenman, Lee, and Rhee (2004) examine the Korean case from the perspective of a precautionary motive. However, aside from these, there are few theoretical or empirical studies to explain this spectacular increase in Korean reserves.

This study investigates the Korean reserve demand. We set up the models for the Korean case based on the conventional buffer stock model and a model of dynamic adjustment. Also, we include a test for a structural change in the reserve demand patterns after the crisis. To our knowledge, ours is the first study to examine the Korean reserve demand for the past three decades, as well as its structural change after the crisis, from the perspectives of the buffer stock model and dynamic variants of this model.[6] Our empirical

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[4] This level presents a striking contrast to developed countries. For example, the U.S., the U.K., and Germany recorded 0.8%, 2.4%, and 1.3% respectively, in 2005.

[5] They point out that some reserves have been accumulated through government bond sales. The problem is that the reserves earn a current market rate of 2-3%, but the government bonds carry an interest rate of 9%.

[6] We note that there is another study of the Korean case: Aizenman, Lee, and Rhee (2004) account for the possibility that a sudden stop of short-term capital flow may trigger large output costs, due to the higher cost of credit or the banking crisis. According to their study, reserves may reduce the probability of a full-blown liquidity crisis, thereby increasing welfare. They suggest that equity inflows and short-term external debt have played a significant role in the rapid accumulation of reserves in the post-crisis period. However, the

results show that, after the crisis, the Korean reserve demand became more sensitive to the adjustment cost and the openness, but less sensitive to the opportunity cost of holding reserves, which is consistent with the rapid reserve accumulation of Korea after the crisis.

In the next section, we review relevant literature to explore theoretical and empirical studies of the reserve demand. In section 3, we set up the model for the Korean case and examine data for estimations. Section 4 presents the empirical results from the estimations. Finally, section 5 summarizes the main findings and concludes.

2. LITERATURE REVIEW

The empirical studies on reserve demand have been conventionally analyzed in terms of the cost-benefit approach. In this model, reserves as a buffer stock are accumulated in times of abundance and decumulated in times of scarcity, and the model has been quite successful in explaining reserve demand before the 1990s (Bahmani-Oskooee, 1985; Lehto, 1994; Bahmani-Oskooee and Brown, 2002). This approach is based on the inventory management principle, which optimizes the trade-off between flow holding costs and fixed restocking costs. It assumes that the central bank chooses an initial level of reserves that minimizes its total expected costs. Two costs are considered: the opportunity cost of holding reserves, and the adjustment cost that is incurred when reserves reach some lower limit.[7] The two costs are interrelated because a higher stock of reserves reduces the probability of having to adjust, which reduces the expected cost of

sample size is quite small (42 observations), and equity inflow and external debt data are not available until 1995.

[7] The adjustment cost is interpreted as the output or welfare forgone by having to take other policy measures to generate the external payments surplus necessary for reserve accumulation in times of actual reserves reaching some lower limit. An example of this kind of policy is the increase of the domestic interest rate that results in a decrease in investment and a reduction in the GDP.

adjustment, but at the cost of higher forgone earnings. Thus, the optimal level of reserves is determined when the expected cost is minimized.

Triffin (1947) argues that reserve demand would increase with the growth in international trade. Reserve adequacy would be determined by its external transactions. Machlup (1966) and Heller (1968) argue that the variability of trade is a much better measure of reserves demand than its level. In the following studies, both variance and levels of trade are used as explanatory variables. Based on Triffin's study, Heller (1966) formalizes the concept of reserve demand as an inventory control problem. In his study, reserves would be held to reduce the adjustment costs under the no-reserve scenario. However, the benefits from reserve holdings would need to be compared with the opportunity cost – for example, an alternative investment with a higher rate of return. Heller also develops the linkage between a country's propensity to import (*PI*) and its reserve demand. He argues that a higher *PI* would lower the marginal cost of adjustment. The external disequilibrium induced by a decline in exports earnings could be corrected by a decline in output. The smaller *PI*, the greater the output decline needed to bring about the correction. The cost of output adjustment could be reduced if the central bank finances the external deficit with its reserves. Thus, the cost of adjustment in the absence of reserves would be inversely related to *PI*. He predicts a negative relationship between *PI* and the reserve demand. However, Frenkel (1974) argues that *PI* reflects the openness and vulnerability of a country to external shocks. If reserves were held as a precautionary measure, this would imply a positive relationship between the import propensity and the reserve demand. Empirically, the average propensity to import (*API*) is used for *PI* because only *API* is available from the actual data. Based on Heller's work, Frenkel and Jovanovic (1981) model how optimal reserve holdings would increase as the reserve volatility increases. The study reveals that reserve volatility is a robust predictor of reserves demand. Flood and Marion (2002) extend Frenkel and Jovanovic's work by modifying the volatility measure. The demand for reserves (*R*) turns out to be a stable function of a few explanatory variables, which are the adjustment cost (C_1), the opportunity cost (C_2), and the size of international transactions (S).

PI is proposed as a variable for the adjustment costs, and the sign is expected to be negative. However, as mentioned above, *API* has been used as a substitute for *PI*, and its coefficient frequently turned out to be positive. Hence, *API* can be interpreted to measure the economy's openness and vulnerability to external shocks. The positive coefficient suggests that the reserve demand increases as the economy faces greater external vulnerability. Also, reserve volatility (σ_R) is proposed as the variable for the adjustment cost by Frenkel and Jovanovic (1981). The volatility is measured by the standard deviation of the trend adjusted changes in reserves over some previous period. Since the higher reserve volatility means that the reserves hit their lower bound more frequently, the central bank is willing to hold a larger stock of reserves to minimize the cost of restocking. The coefficient is expected to be positive. Regarding the opportunity cost, ideally, the cost should be measured as the difference between the highest possible marginal productivity forgone from an alternative investment and the yield on reserves. In general, it is computed as the difference between the country's own interest rate and the yield on U.S. Treasury bonds. If a proper measure for the opportunity cost were found, its coefficient in the model would be expected to be negative. Finally, the variable of the size of international transactions can be represented by imports, the real GDP, the real GDP per capita, or the population size. The coefficient is expected to be positive.

The broad observation of recent studies is that Frenkel and Jovanovic's model holds well for emerging countries, even in the era of high capital mobility (Flood and Marion, 2002; Aizenman and Marion, 2003). Frenkel and Jovanovic state that there is a dependence of the optimal reserve holdings on the variability of international transactions. Reserves are a buffer stock to accommodate fluctuations in external transactions, and it is expected that the optimal reserve stock depends positively on the fluctuations. They assume that the reserve demand is a stochastic process driven by payments and receipts, and define reserve movements in a continuous time period as an exogenous Wiener process. The model is based on the assumption that the stochastic process has no drifts, implying an unchanged reserve policy.

However, it is more common for a policy to change when the reserves reach either an upper or lower limit. Especially in the presence of speculative attacks and occasional restocking of reserves in the era of free flow capital, maintaining a static policy would not be a reasonable assumption. Indeed, Bar-Ilan *et al*. (2004) show that the change in policy reverses the direction of the drift in reserves. Also, if we do not consider the fact that the high frequency reserve data tend to have large and small errors in clusters, the estimated parameters of adjustment cost would be biased. [8)]

To overcome this bias problem, later studies (Ramachandran, 2004; Francisco and Domingos, 2004) devise alternative measures of volatility and derive the volatility from autoregressive conditional heteroskedasticity (ARCH) or generalized autoregressive conditional heteroskedasticity (GARCH) models proposed by Engle (1982) and Bollerslev (1986). The appropriate GARCH (p, q) model is selected by AIC (Akaike information criterion), SIC (Schwarz information criterion), and various diagnostic tests on error terms. The parameters are obtained by maximizing the likelihood function given in Engle (1982); then the conditional standard deviation (σ_{gt}) derived from the GARCH model can be used as volatility (adjustment cost) to estimate the reserve demand. Ramachandran (2004) works with ARCH and GARCH models for the study of India, and Francisco and Domingos (2004) discuss the optimal reserves holding for the Brazilian economy using the GARCH processed reserve volatility.

In addition, based on the recent advances in time series studies, i.e., the cointegration and error correction model, the emphasis is shifted to individual country studies using time series data. Elbadawi (1990) uses an error correction model and finds that the dynamic adjustment in response to a monetary disequilibrium is very slow in Sudan. For China, Huang (1995), also using the error correction model, shows that the speed of adjustment is slow in China's money market.

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8) See Appendix for more discussion.

3. MODEL AND DATA

In this section, starting with a simple buffer model using two variables (adjustment cost and opportunity cost) (following Frenkel and Jovanovic, 1981), we extend the model by adding openness and scale variables. If, in fact, some or all of the variables follow non-stationary processes, then estimation in levels may not be meaningful because of the spurious regression problem.

If the variables are cointegrated, the long-run relationship would be captured through cointegrating relationships in levels, and an appropriate dynamic model can be estimated in an error correction setting. We may consider cointegration estimation methods, such as a dynamic ordinary least squares method and Johansen's method – which are commonly used – to examine the cointegrating vectors. An error correction model is adopted to include the long-run equilibrium process in the short-run dynamics. We estimate three different time periods: the whole sample period (1973.5-2005.12), the pre-crisis period (1990.1-1997.11), and the post-crisis period (1998.3-2005.12). To deal with outliers during the crisis, we exclude the succeeding three months observations after the crisis occurred (1997.12, 1998.1, and 1998.2) from the estimations. As a robustness test, we exclude the succeeding four, five, and six month observations.[9)]

Finally, structural break tests are conducted using dynamic ordinary least square estimation. The specifications for each estimation follow.

3.1. Error Correction Model

We use dynamic models to examine the adjustment process of reserve changes. The partial adjustment model would be one specification to examine the adjustment process to the desired reserve level. Indeed, the partial adjustment model is a restricted form of the error correction model. If two variables, x and y, are cointegrated and the realized value y_t is linked to

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[9)] The results are not reported because the results are not significantly different.

its target value $y_t = \beta' x_t$, the simplest error correction form can be written as

$$\Delta y_t = \lambda_1 \Delta y_t^* + \lambda_2 (y_{t-1}^* - y_{t-1}), \tag{1}$$

where $\lambda_1 > 0$, $\lambda_2 > 0$. The last term represents the past equilibrium error. The partial adjustment model is given by

$$\Delta y_t = \lambda (y_t^* - y_{t-1}) = \lambda \Delta y_t^* + \lambda (y_{t-1}^* - y_{t-1}). \tag{2}$$

Thus, the partial adjustment model corresponds to the error correction model with $\lambda_1 = \lambda_2$. However, the null hypothesis of $\lambda_1 = \lambda_2$ is rejected in all cases at the 1% significance level, implying the more general error correction model is appropriate.

Also, an error correction model would be an appropriate specification if the variables are cointegrated. The error correction model has cointegration relations built into the specifications, so that it restricts the long-run behavior of the cointegrated variables to converge on their cointegrating relationships. The error correction term captures gradual adjustment of the model to the long-run equilibrium through a series of partial short-run adjustments. The coefficient of the error correction term measures the speed of adjustment of the variables towards equilibrium. Thus, we introduce the error correction model to capture both the short-term dynamics and the long-term relationship among the variables (Engel and Granger, 1987; Edwards, 1983 and 1984; Elbadawi, 1990; Ford and Huang, 1994).

For the model, we need to perform the unit root tests and determine whether the variables are integrated or not. If the variables involved are integrated, we perform cointegration tests to examine whether the variables in the models have a stable long-run relationship.

We estimate the cointegrating relationship using the dynamic ordinary least squares method developed by Saikkonen (1991), and Stock and Watson (1993), because we have relatively small sample size.[10] According to the

[10] In the experiments, the Johansen estimates have the smallest bias but the variance is much larger than the other efficient estimators. Since our results from DOLS would be sensitive

Monte Carlo experiments of Stock and Watson (1993) for evaluating the finite sample (100-300 observations) properties of six alternative estimators under cointegration, the dynamic ordinary least squares estimator had the smallest root mean squared error among the estimation methods. Thus, the dynamic ordinary least squares method would be appropriate for our study. Under cointegration, the estimating equation is given by

$$\log R_t = \alpha_k + \beta_k X_{k,t} + \sum_{i=0}^{p} \gamma_{k,-i} \Delta X_{k,t+i} + \sum_{i=1}^{p} \gamma_{k,i} \Delta X_{k,t-i} + v_{k,t}, \qquad (3)$$

where $X_{k,t}$ is a vector of the independent variables $\log \sigma_{gt}$, $\log r_t$, $\log Y_t$, and $\log API_t$.

The dependent variable at time t is regressed upon the independent variables at time t plus an appropriate number of lead and lag differences of the independent variables (including the contemporaneous difference). The estimate for the parameter vector, β_k, typically the main parameters of interest, is super-consistent if the system of $I(1)$ variables are cointegrated. However, the lead and lag differenced terms in equation (3), while eliminating the correlation between the lead and lag differenced terms and the error terms (v_t), do not remove the serial correlation in v_t. To accommodate the serial correlation in v_t, the standard errors need to be properly scaled upward.[11)]

Once the cointegrating relationship has been estimated, the following error correction model is constructed and estimated to determine the short-run impact of the explanatory variables on the reserve demand

$$\begin{aligned} \Delta \log R_t = \alpha + \lambda ECT_{t-1} + \sum_{i=1}^{l} \beta_i \Delta \log R_{t-i} + \sum_{i=0}^{m} \gamma_i \Delta \log \sigma_{gt-i} \\ + \sum_{i=0}^{n} \delta_i \Delta \log r_{t-i} + \sum_{i=0}^{p} \eta_i \Delta \log Y_{t-i} + \sum_{i=0}^{q} \varphi_i \Delta \log API_{t-i} + u_t, \end{aligned} \qquad (4)$$

due to the small sample size, we also estimate Johansen's method, but the results are not significantly different from the previous results. The results are not reported.

11) See Hayashi (2003, pp. 653-657) for more discussion.

where *ECT* is the error correction term, and *l, m, n, p,* and *q* are the lengths of included lags for each variable.

3.2. Dynamic Ordinary Least Squares Structural Break Test

Normally the estimated cointegrating vectors are taken to represent stable long-run relationships among the variable, and the estimated parameters are taken as constant over time. In this study, we would like to test whether the crisis in 1997 caused detectable changes in the reserve demand relationship.

There are a number of alternative tests for structural change under cointegration. For instance, Quintos and Phillips (1993) develop tests for parameter constancy in cointegrating relations in a single-equation setting. Gregory and Hansen (1996) develop the residual-based, single-equation methods; however, they have low power like residual-based tests, because they tend to ignore equation dynamics (Maddala and Kim, 1998). Full information maximum likelihood methods based on the multivariate Johansen (1995) procedure, such as Hansen (2003), may be superior to single equation methods for addressing problems of simultaneity. However, the performance is typically poor in small samples (Gangnes and Parson, 2004). Thus, the dynamic ordinary least squares would be more appropriate for our study based on the same reasoning as the previous section. Hayashi (2003) shows that the dynamic ordinary least squares system can be augmented to allow for structural breaks by including dummy variables. Then, the equation will be

$$\begin{aligned}\log R_t = \alpha_k + \alpha_{k,T+} Dum_{k,T+} + \beta_k X_{k,t} + \beta_{k,T+} XDum_{k,T+} \\ + \sum_{i=0}^{p} \gamma_{k,-i} \Delta X_{k,t+i} + \sum_{i=1}^{p} \gamma_{k,i} \Delta X_{k,t-i} + v_{k,t}{}',\end{aligned} \tag{5}$$

where $X_{k,t}$ is a vector of the independent variables $\log \sigma_{gt}$, $\log r_t$, $\log Y_t$, and $\log API_t$, $\alpha_{k,T+}$ is the coefficient of a dummy variable having unit values beginning in period *T*, and $\beta_{k,T+}$ is a vector of slope coefficients on dummy variables $XDum_{k,T+}$ with non-zero values from period T onward. We use the data 1990.1-1997.11 for the pre-crisis and 1998.3-2005.12 for the post-crisis periods. Thus, T will be 1998.3.

3.3. Data

For the estimation all of the data for the variables are collected or derived from the International Financial Statistics of the IMF and the Bank of Korea. GDP data start from 1960. The data of yields on government bonds (i_G) and U.S. bonds (i_G^*) for the spread (the opportunity cost) starts from 1973 and 1963 respectively. Imports data starts from 1957. Reserves (R), GDP (Y), and imports (IM) are in real terms, calculated as the nominal terms deflated by the CPI of Korea. All absolute values are in U.S. million dollars, calculated using the average market exchange rate if necessary. Also, obvious seasonality is found in the GDP (Y) and in the imports (IM). Thus, we adjust the data to eliminate the seasonality.[12] We include gold because the portion of gold is relatively large in the 1970s. A scale variable is usually chosen to be imports (IM) or GDP (Y). We use the GDP as a scale variable following the conventional method.[13] The average propensity to import (API), the degree of openness, is calculated by dividing the imports (IM) by GDP (Y). The reserve volatility is obtained from the GARCH specification of the change in reserves. We denote the volatility as σ_g. For the variable of the opportunity cost, we use the spread (r), which is the difference between the yield on a one year domestic government bond and the yield on a one year U.S. Treasury bond, $(i_G - i_G^*)$.[14]

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[12] We use the X-11 procedure developed by the U.S. Census Bureau.

[13] We transform the quarterly GDP data to monthly data by dividing the quarterly GDP by 3.

[14] In the case of Korea, it seems to be difficult to find the appropriate opportunity cost because the long-term capital market was not developed until recently and interest rates have been controlled by the government. The liberalization of the capital market and interest rates starts after the crisis in 1997. There are other interest rates that can be considered, such as the money market rate, the lending rate, the deposit rate, and the corporate bond rate, but the government bond rate may be more suitable for the counterpart of the yield on the U.S. Treasury bond.

4. EMPIRICAL RESULTS

4.1. Stationarity and Cointegration Tests

The Augmented Dickey-Fuller and Phillip-Perron unit root tests are performed to test on the stationarity of the data.[15] Neither unit root test rejects the unit root in level, but both reject the unit root in the differenced data. Thus, the variables have an $I(1)$ process, which means the data are non-stationary in levels.

Because all of the variables have unit roots, cointegration tests are performed to examine whether the variables have a stable long-run relationship. The presence of a long-run relationship between the variables (the cointegrating vector) can be detected by performing unit root tests with the residuals of the OLS estimation of reserve demand. If the $I(1)$ variables are cointegrated, it is known that the OLS estimates are super-consistent (Davidson and MacKinnon, 1993; Hamilton, 1994). However, the standard t-statistics or F-statistics are not valid. Different critical values should be used to test for the significance of the estimates. Significance levels are based on the critical tau values, as computed by Engel and Granger (1987).[16] To confirm the results obtained from the single-equation OLS estimations, Johansen (1988 and 1991) cointegration tests are also applied.[17]

The results show that the null hypothesis of the unit root is rejected at the 1 or 5% significance level.[18] It implies that the variables in reserve demand are cointegrated. The Johansen cointegration tests also confirm the presence of cointegrating vectors.[19] Both trace statistics and maximum eigenvalue statistics indicate the presence of one cointegrating vector.

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15) The results of are summarized in Appendix, tables A2 and A3.

16) They are well summarized in Davidson and MacKinnon (1993). Further, we acknowledge that the residual-based test has low power because it ignores equation dynamics and concentrates on error dynamics (Maddala and Kim, 1998, pp. 203-205).

17) An intercept is included, but a trend is not included in the cointegration equations.

18) See Appendix, table A3 for the results of the unit root tests using the residuals of the OLS estimations.

19) See Appendix, tables A4 (the simple model) and A5 (the extended model) for the results of the Johansen cointegration tests.

4.2. Estimation of the Buffer Stock Model

Based on the cointegrating test, the estimates of the coefficients of the reserve demand (the cointegrating vector) by the dynamic ordinary least squares method are reported in table 1.[20] Since the estimation equation is a log-linear form (equation (3)), each coefficient of the variable represents the elasticity. For instance, in table 2, the coefficient of the opportunity cost (spread), –1.381 from the pre-crisis period (extended model), implies that a 1% increase in the spread brings a 1.381% (U.S. $13,810) decrease in the reserve demand.

Table 1 Estimates of the Cointegrating Vectors by the DOLS

Variables	Simple			Extended		
	Whole Period	Pre-crisis	Post-crisis	Whole Period	Pre-crisis	Post-crisis
	log R_t	log R_t	log R_t	log R_t	log R_t	log R_t
log $\sigma_{g\,t}$	0.958^{***} (0.047)	0.033 (0.069)	0.711^{**} (0.309)	0.219^{***} (0.057)	-0.866^{***} (0.151)	0.046 (0.042)
log r_t	-0.267^{***} (0.053)	-0.500^{***} (0.150)	0.672^{***} (0.075)	-0.126^{***} (0.042)	-1.381^{***} (0.176)	0.082^{**} (0.039)
log Y_t	–	–	–	1.326^{***} (0.089)	0.120 (0.120)	2.414^{***} (0.080)
log API_t	–	–	–	2.512^{***} (0.341)	0.893^{***} (0.152)	2.178^{***} (0.392)
Adjusted R^2	0725	0.890	0.716	0.851	0.974	0.994
F-statistics	31.528	24.647	7.300	45.039	45.037	28.675
Time Period	1973.5-2005.12	1990.1-1997.11	1998.3-2005.12	1973.5-2005.12	1990.1-1997.11	1998.3-2005.12

Notes: Significance levels are 10% *, 5% **, and 1% ***. Coefficients on constant terms, and lead and lag differenced terms are not reported. The number in parenthesis is the scaled standard error.

[20] The lag lengths are determined following information based rules. We select lag 2 and lag 3 for the simple and extended models. We also follow Hayashi's method to measure the adjusted standard errors.

For the simple model, both of the coefficients of the adjustment cost and the opportunity cost are consistent with the theoretical predictions for the pre-crisis period, but the coefficient of the adjustment cost is not statistically significant. For the post-crisis period, the coefficient of the adjustment cost is positive (the expected sign) and significant at the 5% significance level. However, the coefficient of the opportunity cost turns to positive, which is different from the theoretical prediction, and significant at the 1% significance level.

For the extended model, the coefficients of all four variables are consistent with the theoretical predictions for the pre-crisis period. Only the coefficient of the adjustment cost presents an unexpected sign (negative) and is significant at the 1% significance level. However, after the crisis, the coefficient of the adjustment cost turns to positive (but is not statistically significant), and the coefficient of the opportunity cost presents a different sign (positive) from the theoretical prediction and is significant at the 1% significance level.

Overall, if we compare the pre- and post-crisis periods, the sign of the adjustment cost turns to positive (but is not statistically significant for the extended model) from insignificant (the simple model) or negative (the extended model). The signs of the opportunity cost change from negative to positive and are statistically significant. The Korean reserve demand became more sensitive to volatility, but less sensitive to the opportunity cost.

Since the previous cointegration tests detect one long-run equilibrium relationship for each model, the error correction models illustrated in equation (4) are estimated to determine the short-run dynamics of the reserve demand.[21] The error correction terms are computed by the cointegration vectors. Following Hendry's general to specific strategy, each error correction model is estimated with long lags (here, twelve lags for a one year interval) of each explanatory variable. Variables that are insignificant are excluded from the equation to find a parsimonious structure of the model and

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[21] The statistics for the whole sample period are not reported because the adjusted R^2s are extremely small (around 0.03 to 0.04).

Table 2 Results of the Error Correction Model for the Simple Model

DOLS Cointegration Pre-crisis		DOLS Cointegration Post-crisis	
Variables	$\Delta \log R_t$	Variables	$\Delta \log R_t$
C	0.002 (0.004)	C	0.007^{***} (0.003)
Error correction term	–0.024 (0.028)	*Error correction term*	-0.017^{*} (0.009)
$\Delta \log \sigma_{g\,t}$	-0.057^{**} (0.027)	$\Delta \log \sigma_{g\,t}$	0.028^{**} (0.011)
$\Delta \log \sigma_{g\,t\text{-}2}$	-0.097^{**} (0.050)	$\Delta \log \sigma_{g\,t\text{-}2}$	0.026^{*} (0.015)
$\Delta \log \sigma_{g\,t\text{-}5}$	-0.061^{**} (0.030)	$\Delta \log \sigma_{g\,t\text{-}10}$	0.039^{**} (0.019)
$\Delta \log \sigma_{g\,t\text{-}7}$	-0.067^{**} (0.030)	$\Delta \log \sigma_{g\,t\text{-}11}$	0.029^{*} (0.015)
$\Delta \log r_t$	-0.136^{*} (0.079)	$\Delta \log r_t$	0.015^{*} (0.011)
$\Delta \log r_{t\text{-}1}$	-0.077^{*} (0.040)	$\Delta \log R_{t\text{-}1}$	0.601^{**} (0.088)
$\Delta \log r_{t\text{-}6}$	-0.109^{**} (0.048)		
$\Delta \log R_{t\text{-}1}$	0.222^{**} (0.092)		
Adjusted R^2	0.257	Adjusted R^2	0.527
$Q(15)$	16.765	$Q(15)$	8.784
$Q^2(15)$	4.163	$Q^2(15)$	4.918
Jarque-Bera	73.184^{***}	Jarque-Bera	72.256^{***}
AR*F*	1.342	AR*F*	0.659
ARCH-*F*	1.310	ARCH-*F*	0.076
Heteroskedasticity *F*	1.666^{*}	Heteroskedasticity *F*	2.369^{***}
Observations	95	Observations	94

Note: Significance levels are 10% *, 5% **, and 1% ***. Δ denotes the first difference.

to avoid depriving us of too many degrees of freedom.

Table 2 shows the results from the reduced model. Overall, for the pre-crisis period, the changes in both the adjustment cost and the opportunity cost have negative effects on the change in reserves at the current and lag terms. The changes in the opportunity cost also have negative effects at the current terms. On the contrary, for the post-crisis period, the effects of changes in the adjustment cost and the opportunity cost turn to positive. The changes in the adjustment cost have positive effects at the current and lag terms. The changes in the opportunity cost also have positive effects at the current term. For the simple model, the chane in reserves adjusts to the past disequilibrium by 1.7% in a month (20.4% in a year), meaning that the adjustment speed is substantially slow. The coefficients of the error correction terms are negative in both periods and statistically significant in the post-crisis period.

Further, the table presents the diagnostic tests on the residuals. $Q(15)$ and AR*F* statistics present no autocorrelation, and $Q^2(15)$ and ARCH-*F* present no ARCH effect, implying that the model is correctly specified. We note that the Jarque-Bera test rejects the normality, and White's heteroskedasticity test presents the existence of heteroskedasticity in the residuals, which may be due to the small sample size. Overall, our diagnostic tests indicate that there are no significant concerns about the specification.

Next, table 3 shows the results for the extended model. The empirical results are similar to those of the simple model. For the pre-crisis period, the changes in both the adjustment cost and the opportunity cost have a negative effect on the change in reserves. The effects of the scale variable (*Y*) are positive but the effects of openness (*API*) are insignificant for all lags. On the other hand, for the post-crisis period, the effects of the changes in the adjustment cost and the opportunity cost turn to positive. However, the effects of the scale variable (*Y*) show an unexpected result (negative). The effects of openness (*API*) are positive as expected. The coefficients of the error correction terms are negative and statistically significant in both cases. Further, our diagnostic tests indicate that there are no significant concerns about the specification.

Our results from the error correction models may imply a change in the Korean reserve demand. The results from both the long-run and the short-run relationships among the variables provide the evidence that, after the crisis, the Korean reserve demand became more sensitive to the adjustment cost and the openness, but less sensitive to the opportunity cost.

Table 3 Results of the Error Correction Model for the Extended Model

DOLS Cointegration Pre-crisis		DOLS Cointegration Post-crisis	
Variables	Δ log R_t	Variables	Δ log R_t
C	–0.003 (0.004)	C	0.018*** (0.003)
Error correction term	–0.192*** (0.060)	*Error correction term*	–0.047*** (0.009)
Δ log $\sigma_{g\ t\text{-}7}$	–0.078*** (0.038)	Δ log $\sigma_{g\ t\text{-}4}$	0.043** (0.018)
Δ log r_t	–0.137*** (0.049)	Δ log $\sigma_{g\ t\text{-}5}$	0.034** (0.013)
Δ log $r_{t\text{-}8}$	–0.109*** (0.060)	Δ log $r_{t\text{-}2}$	0.028*** (0.008)
Δ log $Y_{t\text{-}1}$	0.884*** (0.146)	Δ log $r_{t\text{-}7}$	0.024** (0.009)
Δ log $Y_{t\text{-}2}$	0.396* (0.205)	Δ log $r_{t\text{-}9}$	0.024* (0.013)
		Δ log $Y_{t\text{-}3}$	–0.212*** (0.048)
		Δ log $API_{t\text{-}1}$	0.114*** (0.027)
		Δ log $API_{t\text{-}2}$	0.145*** (0.037)
		Δ log $R_{t\text{-}1}$	0.479** (0.059)
Adjusted R^2	0.483	Adjusted R^2	0.668
Q(15)	15.203	Q(15)	15.351
Q^2(15)	22.203	Q^2(15)	3.880
Jarque-Bera	3.992	Jarque-Bera	4.051
ARF	2.325*	ARF	1.020
ARCH-F	2.295*	ARCH-F	0.101
Heteroskedasticity F	1.020	Heteroskedasticity F	0.348
Observations	95	Observations	94

Note: Significance levels are 10% *, 5% **, and 1% ***. Δ denotes the first difference.

Briefly, the experience and lessons from the financial crisis may have provoked the Korean monetary authority to change its reserve demand decision-making guidelines to favor the adjustment cost rather than the opportunity cost. In other words, our results may reflect the change in the decision rule of the Bank of Korea regarding the reserve policy. This is consistent with the reasoning for the rapid accumulation of reserves after the crisis; that is, Korea needs to be more cautious about the external shock, and the benefit of reserve holdings is large enough to offset the opportunity cost.

Table 4 Results of the Dynamic Ordinary Least Squares Structural Break Test

Variables	Simple	Extended
	$\log R_t$	$\log R_t$
C	13.063*** (1.265)	0.195 (1.905)
$\log \sigma_{gt}$	–0.140 (0.156)	0.102 (0.089)
$\log r_t$	–1.019*** (0.144)	–0.466*** (0.093)
$\log Y_t$	–	0.920*** (0.101)
$\log API_t$	–	–0.371 (0.296)
Dum	1.748 (2.260)	–7.951*** (1.477)
Dum $\log \sigma_{gt}$	0.216 (0.290)	0.046 (0.078)
Dum $\log r_t$	0.353** (0.153)	0.169** (0.085)
Dum $\log Y_t$	–	0.837*** (0.100)
Dum $\log API_t$	–	0.007 (0.390)

Note: Significance levels are 10% *, 5% **, and 1% ***. The number in parenthesis is the scaled standard error.

4.3. Dynamic Ordinary Least Squares Structural Break Test

Based on the previous cointegration tests, the results of the dynamic ordinary least squares structural break tests with dummy variables are reported in table 4.[22] We note that under this specification (equation (5)), the coefficients on the variables in levels represent the long-run elasticity of reserves with respect to the independent variables, and the dummy variables test for the structural changes in the reserve demand.

The results from the simple model demonstrate that the coefficients of the opportunity cost are consistent with the theoretical predictions (a negative sign), and the significance levels are high (significant at the 1% significance

[22] The lag lengths are determined using information based rules. We select lag 3 and lag 6 for the simple and the extended models. Also, we follow the Hayashi's method to rescale the standard errors. The results of selected various lag lengths are not significantly different from this.

level). However, the coefficients of the adjustment cost are negative, which is different from the theoretical prediction, and they are not different from zero statistically. Our structural break tests provide mixed evidences for the change in reserve demand. The coefficients of the $Dum \log r_t$ are positive and significant at the 5% significance level suggesting that the Korean reserve demand became less sensitive to the opportunity cost after the crisis. The coefficients of the *Dum* and the $Dum \log \sigma_t$ are also positive but not statistically significant.

For the extended model, the coefficients of the variables are consistent with the theoretical predictions except for the $\log API_t$. However, the coefficients of the $\log \sigma_t$ and the $\log API_t$ are insignificant. Further, the structural break in the reserve demand after the crisis presents mixed results. The results demonstrate an increase in the effects of the adjustment cost, the scale, and the openness, and a decrease in the effects of the opportunity cost after the crisis. However, only three dummy variables out of five (*Dum*, $Dum \log r_t$, and $Dum \log Y_t$) are statistically significant.

Briefly, our structural break test results provide us with mixed evidences for the change in the Korean reserve demand after the crisis for both the simple and extended models.

5. CONCLUSION

This study investigates the Korean reserve demand and examines its structural change after the financial crisis in 1997. For the past 15 years (1990.1-2005.12) the dynamics of the reserve demand during the pre- and post-crisis periods are reasonably specified by the error correction model. The cointegrating vectors, which show the long-run relationship between the reserve demand and the variables, are estimated by the dynamic ordinary least squares method. The effects of the adjustment cost and the openness on the reserve demand get larger. (For example, the parameters of the adjustment cost turn positive after the crisis.) On the contrary, the effects of the opportunity cost get smaller. The short-run relationship is also consistent with the long-run relationship.

Furthermore, the dynamic ordinary least squares structural break tests present the increase in the effects of the adjustment cost and the scale variable, and the decrease in the effects of the opportunity cost after the crisis. However, only the cases of the opportunity cost and the scale variable are statistically significant. Thus, our structural break tests provide mixed evidences for the change in the Korean reserve demand.

Our empirical results may provide the evidence that the Korean reserve demand became more sensitive to the adjustment cost and the openness, but less sensitive to the opportunity cost after the crisis. This may imply the change in the decision-making rules of the reserve policy. The results are consistent with the reasoning of the rapid accumulation of reserves after the crisis; that is, that Korea needs to be more cautious about the external shock, and that the benefit of reserve holdings is large enough to offset the opportunity cost.

APPENDIX

The Bias Problem of the Adjustment Cost and the Correction for the Bias

According to Frenkel and Jovanovic, the reserve follows a random walk process with a drift. However, as mentioned above, this assumption is not appropriate in the presence of speculative attacks and the occasional restocking of reserves. For example, let the reserve volatility be one period rolling variance: $\sigma_t^2 = \varepsilon_{t-1}^2$. If Frenkel and Jovanovic are correct in assuming that the observed reserves, R_t, are around their optimal level, R_0, the least square coefficient of reserve volatility is defined as

$$\beta_1 = \mathrm{cov}(R_0, \sigma_t^2) / \mathrm{var}(\sigma_t^2). \tag{6}$$

Table A1 The Augmented Dickey-Fuller Unit Root Tests on the Variables

Variables	Whole Period		Pre-crisis		Post-crisis	
	Level	First difference	Level	First difference	Level	First difference
$\log R$	–0.053	-18.669^{***}	–0.872	-7.279^{***}	–2.554	-5.837^{***}
$\log \sigma_g$	-2.632^{*}	-19.676^{***}	–1.301	-9.668^{***}	–1.843	-9.503^{***}
$\log r$	–1.413	-14.941^{***}	–1.775	-7.535^{***}	–0.419	-8.991^{***}
$\log Y$	–1.677	-8.129^{***}	–1.226	-9.779^{***}	–1.482	-10.624^{***}
$\log API$	–2.550	-22.808^{***}	–0.777	-15.085^{***}	0.115	-10.078^{***}
Time Period	1973.5-2005.12		1990.1-1997.11		1998.3-2005.12	

Note: Significance levels are 10% *, 5% **, and 1% ***. We selected the augmentation lags for each Dickey-Fuller regression in order to minimize the Schwarz Information Criterion (SIC). Each regression contains an intercept but no time trend.

In reality, however, the observed reserves remain far from the optimal level. For instance, $R_{t-2} = R_0$. Due to the random walk process, $R_{t-1} = R_0 + \mu + \varepsilon_{t-1}$ and $R_t = R_0 + 2\mu + \varepsilon_{t-1} + \varepsilon_t$. If so, then

$$\operatorname{cov}(R_t, \sigma_t^2) = \operatorname{cov}(R_0, \sigma_t^2) + E(\varepsilon^3). \tag{7}$$

If ε is positively skewed due to reserve restocking, then β_1 would be upwardly biased and vice versa.

To eliminate the bias discussed in Appendix, we construct the volatility measure by modeling the variance of reserve changes through the GARCH specifications. If we define ε having a conditional variance of the GARCH(1, 1) process following Engle (1982), then

$$\Delta R_t = \omega_t + \sqrt{\sigma_t^2}\,\upsilon_t, \tag{8}$$

$$\sigma_t^2 = \delta + a\varepsilon_{t-1}^2 + b\sigma_{t-1}^2, \tag{9}$$

Table A2 The Phillips-Perron Unit Root Tests on the Variables

Variables	Whole Period		Pre-crisis		Post-crisis	
	Level	First difference	Level	First difference	Level	First difference
$\log R$	–0.079	–18.669***	–0.964	–7.354***	–4.309***	–4.795***
$\log \sigma_g$	–2.647*	–19.686***	–1.301	–9.668***	–1.856	–9.532***
$\log r$	–0.877	–18.453***	–1.389	–7.396***	–0.972	–8.667***
$\log Y$	–1.896	–20.753***	–1.681	–9.994***	–1.555	–10.636***
$\log API$	–4.221	–69.357***	–0.518	–23.392***	0.110	–12.224***
Time Period	1973.5-2005.12		1990.1-1997.11		1998.3-2005.12	

Note: Significance levels are 10% *, 5% **, and 1% ***. Each regression contains an intercept but no time trend.

Table A3 The Cointegration unit Root Tests on the OLS Residuals

Unit Root	Simple			Extended		
	Whole Period	Pre-crisis	Post-crisis	Whole Period	Pre-crisis	Post-crisis
ADF	–2.858***	–2.730***	–2.261**	–2.372**	–4.019***	–2.259**
Phillips-Perron	–2.717***	–2.818***	–2.279**	–2.455**	–4.136***	–2.038**
Time Period	1973.5-2005.12	1990.1-1997.11	1998.3-2005.12	1973.5-2005.12	1990.1-1997.11	1998.3-2005.12

Note: Significance levels are 10% *, 5% **, and 1% ***. Significance levels are based on the critical tau values, as computed by Engel and Granger (1987). Each test contains an intercept but no time trend.

where υ_t follows normal distribution. If we define the conditional standard deviations from the above equation as σ_{gt}, then

$$\mathrm{cov}(R_t, \sigma_{gt}) = \mathrm{cov}(R_0, \sigma_{gt}). \tag{10}$$

Hence, β_1 is not biased.

Table A4 The Johansen Tests for Cointegration (The Simple Model)

Period	Hypothesized Number of Cointegrating Vectors	Eigenvalues	Trace Statistics	Maximum Eigenvalue Statistic
Whole Sample	0	0.061	42.522** (35.193)	24.146** (22.300)
	≤ 1	0.026	18.376 (20.262)	10.311 (15.892)
	≤ 2	0.021	8.065 (9.165)	8.065 (9.165)
Pre-crisis	0	0.180	35.871** (35.193)	18.815 (22.300)
	≤ 1	0.130	17.056 (20.262)	13.222 (15.892)
	≤ 2	0.040	3.834 (9.165)	3.834 (9.165)
Post-crisis	0	0.261	41.810** (35.193)	26.567** (22.300)
	≤ 1	0.131	15.244 (20.262)	12.341 (15.892)
	≤ 2	0.032	2.902 (9.165)	2.902 (9.165)

Note: Significance levels are 5% * and 1% **. The variables of cointegration test are $\log R$, $\log \sigma_g$, and $\log r$. The number in parenthesis is a critical value at the 0.05 level. Each specification for the whole, pre-, and post-crisis period includes no lag, one lag, and three lags, respectively, assuming a trend in the series but not in the cointegrating relationships.

Table A5 The Johansen Tests for Cointegration (The Extended Model)

Period	Hypothesized Number of Cointegrating Vectors	Eigenvalues	Trace Statistics	Maximum Eigenvalue Statistic
Whole Sample	0	0.111	93.417** (76.973)	44.587** (34.806)
	≤ 1	0.064	48.830 (54.079)	24.916 (28.588)
	≤ 2	0.036	23.914 (35.193)	13.744 (22.300)
	≤ 3	0.021	10.170 (20.262)	7.860 (15.892)
	≤ 4	0.006	2.310 (9.165)	2.310 (9.165)
Pre-crisis	0	0.290	86.417** (76.973)	32.513* (34.806)
	≤ 1	0.196	53.905 (54.079)	20.689 (28.588)
	≤ 2	0.154	33.216 (35.193)	15.928 (22.300)
	≤ 3	0.115	17.287 (20.262)	11.652 (15.892)
	≤ 4	0.058	5.636 (9.165)	5.636 (9.165)
Post-crisis	0	0.330	77.049** (69.819)	34.090** (33.877)
	≤ 1	0.224	42.959 (47.856)	21.589 (27.584)
	≤ 2	0.152	21.370 (29.797)	14.053 (21.132)
	≤ 3	0.080	7.316 (15.495)	7.123 (14.264)
	≤ 4	0.002	0.193 (3.841)	0.193 (3.841)

Note: Significance levels are 5% * and 1% **. The variables of cointegration test are $\log R$, $\log \sigma_g$, $\log r$, $\log Y$, and $\log API$. The number in parenthesis is a critical value at the 0.05 level. Each specification for the whole, pre-, and post-crisis period includes three lags, four lags, and three lags, respectively, assuming a trend in the series but not in the cointegrating relationships.

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