An Empirical Investigation of Monetary Interaction in the Korean Economy

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May 8, 2009

Abstract

This paper conducts an econometric investigation of monetary interaction in the Korean economy over the past two decades. The study pays close attention to a critical role played by broad money and an interest rate term spread in the economy. A vector autoregression of Korea's time series data reveals two cointegrating relationships, both of which are consistent with macroeconomic theory: the first relationship corresponds to a broad money demand function, while the second represents a monetary policy rule function. All the variables in the cointegrated system, apart from the broad money and term spread, are judged to be weakly exogenous for parameters of interest, thereby allowing us to estimate a partial system with no loss of information. It is demonstrated that the preferred partial system characterizes interaction between money demand and monetary policy rule in the Korean economy.

Keywords: Monetary Interaction, Money Demand, Monetary Policy Rule, Cointegrated Vector Autoregression, Partial System.

JEL classification codes: C32, E41, E52.

1 Introduction

The objective of this paper is to pursue an econometric investigation into monetary aspects of the Korean economy over the past two decades. A cointegrated vector autoregressive (VAR) approach is adopted for the empirical exploration, and the study focuses on the role of broad money and an interest rate term spread in the VAR system. The introductory section presents a brief review of the empirical literature on money demand and monetary policy rules, and then describes the most significant aspect of this paper.

Numerous empirical studies in the macroeconomics literature have focused on determinants of the demand for money and its stability. While earlier studies relied on the

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traditional OLS method of estimation, many recent researches have used cointegration and equilibrium-correction models, among others. Johansen and Juselius (1990) is a seminal paper on money demand in Denmark and Finland, applying the cointegration method of Johansen (1988). Using a similar method, Hafer and Jansen (1991) and Hoffman and Rasche (1991), among others, find a stable relation for M1 and M2 of the US. Similar researches include Hendry and Mizon (1993), Hendry and Doornik (1994), and Ericsson (1998) on the UK M1; Ericsson and Sharma (1998), Peytrignet and Stahel (1998), Vega (1998), Juselius (1998), Brűggemann and Lűtkepohl (2006) on other European economies; Bahmani-Oskooee and Shabsigh (1996) and Miyao (1996a) on Japan; and Arize (1994) and Khalid (1999) on other Asian countries. Korean money demand has also been investigated in a number of empirical studies. Recently, some analyses have focused on the consequences of the financial crisis for money demand stability. The stability of Korean money demand for the broad aggregate is generally not rejected in the literature. Recent examples include Arize (1994), Bahmani-Oskooe and Rhee (1994), Lee and Chung (1995), Khalid (1999) and Bahmani-Oskooe and Shin (2002).

Taylor (1993) has triggered off an ongoing discussion about a monetary policy rule. See, for example, Ball (1997), Clarida *et al.* (1998, 2000), and Ball and Tchaidze (2002). Clarida *et al.* (1998) find that since 1979 each of the G3 (Germany, Japan, and the US) central banks has pursued an implicit form of inflation targeting. Clarida *et al.* (2000) show that in the US the interest rate policy after 1979 appears to have been more sensitive to expected inflation than before 1979. Hayo and Hofmann (2006), using the generalized method of moment (GMM), show that the ECB reacts similarly to expected inflation but significantly stronger to the output gap than the Bundesbank did. Gerdesmeier *et al.* (2007), using GMM and Kalman filter approach, find the Eurosystem, the US Federal Reserve, and the Bank of Japan do not show different policy reaction functions. Christensen and Nielsen (2009) find a stable long-term relationship between the Federal funds rate, unemployment and the bond rate, which is interpretable as a policy target. Cho (2007) studies the Korean monetary policy using GMM and so forth. Papers using the cointegrated VAR approach to the analysis of monetary policy rule are rare, while a few analyses (written in Korean) on the transmission mechanism of monetary policy using the structural VAR approach.

As reviewed above, the literature on the money demand and monetary policy rules seems to be separated in most cases, although both of these topics should represent important monetary aspects of an economy in question. An exceptional work is Brűggemann (2003), which is successful in simultaneously exploring money demand and monetary policy rule functions in Germany in a cointegrated VAR framework. It would also be very important, in the context of monetary policy in Korea, to investigate interaction between money demand and monetary policy functions in the Korean economy.

This paper, using time series data over the past two decades, pursues an econometric investigation into monetary aspects of the Korean economy. After the process of the model reduction, the paper is successful in achieving a data-congruent partial system for broad money and the interest rate term spread. Two long-run relationships, incorporated into the partial system, are consistent with what macroeconomic theory tells us: the first relationship corresponds to a broad money demand function, while the second is interpreted as a representation of the underlying monetary policy rule. The preferred system thus depicts interaction between money demand and monetary policy functions in Korea. To the best of our knowledge, the present paper is the first empirical study that

is successful in characterizing monetary interaction in the Korean economy over the past two decades. This paper is therefore regarded as a significant contribution to the research field of monetary economics and the Korean economy.

The organization of this paper is as follows. Section 2 provides a historical review of Korea's monetary policy, and Section 3 briefly explains a cointegrated VAR analysis and partial system. Section 4 then presents a canonical model focusing on money demand and monetary policy rule functions. Section 5 gives an overview of the data analyzed in this paper, and Section 6 performs a rigorous econometric analysis of the data in order to explore monetary interaction in the Korean economy. The overall summary and conclusion are given in Section 7. All the numerical analyses and graphics in this paper use *OxMetrics / PcGive* (Doornik and Hendry, 2006).

2 Historical Review of Korea's Monetary Policy

From 1957 to early 1998 Korea's monetary authorities adopted a form of money-growth targeting, and in the meantime the intermediate target varied — broadened in general — from M1 to domestic credit held by the central bank and to M2. Under the money-growth targeting the growth rate of the monetary aggregate was set as the intermediate target on the basis of the economic variables — both actual and projected — such as output growth, inflation rate, interest rate, and change in the velocity of money. Although the short-run interest rate was not explicitly managed, a smaller (larger) money supply would mean an increase (decrease) in the short-term interest rate.

Advance in financial liberalization and growth of non-bank financial intermediaries (NBFIs) in 1980s led to gradual portfolio shifts away from traditional monetary assets to new financial products, making M2 less useful as an intermediate target. The relationship between M2 and the economy appeared to have become unstable so that in late 1980s Korea's monetary authorities set another monetary aggregate target, which included short-term deposits held at NBFIs, in addition to M2.

In the process of financial liberalization in the 1980s, direct monetary policy instruments such as selective credit control and preferential interest rate were replaced by indirect instruments like government bonds repurchase facility (Repo) and Monetary Stabilization Bonds (MSBs). Markets for government bonds were not well developed, so that the Bank of Korea issued and used MSBs as an open-market operation means of controlling the liquidity of the economy. Meanwhile, legal reserve requirement ratio has always been used as another instrument, although it has not been altered frequently.

In 1998 inflation targeting replaced money-growth targeting, and the intermediate target shifted from M2 to M3. Note that a certain monetary aggregate continued to be used as an intermediate target even when money-growth targeting was officially no longer in place. The reason for the abandonment in Korea of money-growth targeting was not explicitly mentioned. One of the reasons for the adoption in Korea of inflation targeting could be the credibility benefits of a strict rule, which would match the newly gained independence of the central bank. Due to the reform in the financial sector in the wake of the financial crisis of 1997-98, demand for M2 was likely to become unstable, which led the Bank of Korea to focus on M3. And then in 2002 the composition of M3 was modified and it was replaced by the so-called liquidity aggregate of financial institutions, whose data series became available from 1986.

In general, inflation targeting does not preclude the use of monetary policy to stabilize output or other macroeconomic variables in the short run, although hitting the inflation target in the medium run is its first priority; and Korea has also been the case. Unlike Canada and New Zealand, which adopted some form of inflation targeting without explicitly setting any intermediate targets, Korea's monetary authorities have set the short-term interest rate as the operational target, the liquidity aggregate the intermediate target, and the inflation rate the ultimate goal. The Bank of Korea has pursued its policy for a while in the same way as the European Central Bank, which uses a modified inflation-targeting approach that retains some role for money-growth targets. As the Federal Reserve of the US, which has not officially adopted a strategy of inflation targeting but increased its emphasis on maintaining low and stable inflation, chooses a target for the Federal funds rate, so the Bank of Korea does the overnight call rate.

In principle the monetary authorities cannot simultaneously meet the targets of money supply and short-term interest rate, unless these targets are set to be consistent with each other. Thus in Korea the targets have not been — cannot be — strict with strong commitment at least with regard to the money supply. The liquidity aggregate became less important in its role, shifting from an intermediate target to a monitoring indicator in 2001 and then to just an information variable in 2003. We can assess that since 1998 Korea's monetary policy has been a flexible hybrid, which consists of inflation targeting with emphasis on the short-term interest rate, and phasing out money-growth targeting.

In summary, Korea's monetary authorities have been arguably concerned about price stability as well as output stability under both previous money-growth targeting and recent inflation targeting. Despite the regime change in Korea's monetary policy, some research papers (in Korean) find that interest rates might be useful as operational targets in the monetary policy even in 1980s and 1990s, when the money-growth targeting was adopted in Korea. See Ahn and Oh (1998), Oh (1998), and Song (1999), *inter alia*. Furthermore, Shin (1997) demonstrates that, although the monetary-growth targeting was officially adopted in Korea, the announced target of M2 growth rates was not accomplished in most cases. Thus, in spite of the regime change in Korea's monetary policy, there is still a possibility that we may succeed in estimating a stable econometric model over the past two decades, which characterizes interaction between money demand and monetary policy in Korea. This paper is an attempt to establish such a data-congruent econometric model.

3 Cointegrated VAR Analysis and Partial System

This section briefly explains the likelihood-based analysis of a cointegrated VAR model, together with a short review of a partial cointegrated VAR model and weak exogeneity. Macroeconomic time series data often exhibit non-stationary behavior and need to be treated as processes integrated of order 1 (denoted as $I(1)$ hereafter). Thus a cointegrated VAR model explored by Johansen (1988, 1996) has played an important role in applied macroeconomics. See Juselius (2006) and Kurita (2007), *inter alia*, for empirical research using cointegrated VAR models.

Consider an unrestricted VAR(k) model for a p-dimensional time series $X_{-k+1}, ..., X_T$:

$$\Delta X_t = (\Pi, \Pi_l) \begin{pmatrix} X_{t-1} \\ t \end{pmatrix} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \Phi D_t + \varepsilon_t, \quad \text{for } t = 1, ..., T, \tag{1}$$

where D_t is a s-dimensional vector of deterministic terms apart from linear trend and intercept, such as seasonal and impulse dummies, and innovations $\varepsilon_1, ..., \varepsilon_T$ have independent and identical normal $N(0, \Omega)$ distributions conditional on the starting values $X_{-k+1}, ..., X_0$. The parameters Π, Γ_i, $\Omega \in \mathbf{R}^{p\times p}$, Π_l, $\mu \in \mathbf{R}^p$ and $\Phi \in \mathbf{R}^{p\times s}$ vary freely and Ω is positive definite.

For the purpose of performing an $I(1)$ cointegration analysis, three regularity conditions need to be fulfilled. The first condition is that the characteristic roots obey the equation $|A(z)| = 0$, where

$$A(z) = (1-z)\, I_p - \Pi z - \sum_{i=1}^{k-1} \Gamma_j\, (1-z)\, z^i,$$

and the roots satisfy $|z| > 1$ or $z = 1$. This condition ensures that the process is neither explosive nor seasonally cointegrated. The second condition is given by

$$rank\,(\Pi, \Pi_l) \leq r \quad \text{or} \quad (\Pi, \Pi_l) = \alpha(\beta', \gamma'), \tag{2}$$

where $\alpha, \beta \in \mathbf{R}^{p\times r}$ and $\gamma' \in \mathbf{R}^r$ for $r < p$. Let $\beta^{*\prime} = (\beta', \gamma')$ and $X^*_{t-1} = \left(X'_{t-1}, t\right)'$ for future reference. A set of vectors α is referred to as adjustment vectors, while β^* is called cointegrating vectors. Condition (2) implies that there are at least $p - r$ common stochastic trends and cointegration arises when $r \geq 1$. The third condition is

$$rank\,(\alpha'_\perp \Gamma \beta_\perp) = p - r,$$

where $\Gamma = I_p - \sum_{i=1}^{k-1} \Gamma_i$, and $\alpha_\perp, \beta_\perp \in \mathbf{R}^{p\times p-r}$ are orthogonal complements such that $\alpha'\alpha_\perp = 0$ and $\beta'\beta_\perp = 0$ with $(\alpha,\, \alpha_\perp)$ and $(\beta,\, \beta_\perp)$ being of full rank. The final condition prevents the process from being $I(2)$ or of higher order. If these conditions are satisfied, an $I(1)$ cointegrated VAR model is defined as a sub-model of (1) as follows:

$$\Delta X_t = \alpha\beta^{*\prime} X^*_{t-1} + \sum_{i=1}^{k-1} \Gamma_i \Delta X_{t-i} + \mu + \Phi D_t + \varepsilon_t, \tag{3}$$

which is the basis for the subsequent cointegration analysis and model reduction.

Since the cointegrating rank r is usually unknown to investigators, it needs to be estimated from the data. A log-likelihood ratio ($\log LR$) test statistic is given by the null hypothesis of r cointegration rank $\mathsf{H}\,(r)$ against the alternative hypothesis $\mathsf{H}\,(p)$, and its asymptotic quantiles are provided by Johansen (1996, Ch.15). See also Nielsen (1997) and Doornik (1998) for the method of gamma approximations to calculate the quantiles. Determining the cointegrating rank in (3) allows us to test various restrictions on α and β^*. Cointegrating relationships, embodied by $\beta^{*\prime} X^*_{t-1}$, correspond to a set of stationary linear combinations, and they act as equilibrium correction mechanisms in (3). Thus, the relationships represent long-run economic linkages of variables in the system. It is therefore important to check if theory-consistent restrictions can be imposed on β^* estimated from the data.

Next, let the process be decomposed as $X_t = (Y'_t, Z'_t)'$ for $Y_t \in \mathbf{R}^m$, $Z_t \in \mathbf{R}^{p-m}$ and $m \geq r$. The set of parameters and the error terms are also given by

$$\alpha = \begin{pmatrix} \alpha_y \\ \alpha_z \end{pmatrix}, \Gamma_i = \begin{pmatrix} \Gamma_{y,i} \\ \Gamma_{z,i} \end{pmatrix}, \ \mu = \begin{pmatrix} \mu_y \\ \mu_z \end{pmatrix}, \ \Phi = \begin{pmatrix} \Phi_y \\ \Phi_z \end{pmatrix}, \ \varepsilon_t = \begin{pmatrix} \varepsilon_{y,t} \\ \varepsilon_{z,t} \end{pmatrix},$$

and the normal innovations have a variance-covariance matrix as follows:

$$\Omega = \begin{pmatrix} \Omega_{yy} & \Omega_{yz} \\ \Omega_{zy} & \Omega_{zz} \end{pmatrix}.$$

Model (1) is then decomposed into a conditional model for Y_t and a marginal model for Z_t, that is,

$$\Delta Y_t = \omega \Delta Z_t + (\alpha_y - \omega \alpha_z) \beta^{*\prime} X_{t-1}^* + \sum_{i=1}^{k-1} \widetilde{\Gamma}_{y,i} \Delta X_{t-i} + \widetilde{\mu}_y + \widetilde{\Phi}_y D_t + \widetilde{\varepsilon}_{y,t}, \tag{4}$$

$$\Delta Z_t = \alpha_z \beta^{*\prime} X_{t-1}^* + \sum_{i=1}^{k-1} \Gamma_{z,i} \Delta X_{t-i} + \mu_z + \Phi_z D_t + \varepsilon_{z,t}, \tag{5}$$

where

$$\begin{aligned} \omega &= \Omega_{yz} \Omega_{zz}^{-1}, \ \widetilde{\Gamma}_{y,i} = \Gamma_{y,i} - \omega \Gamma_{z,i}, \ \widetilde{\mu}_y = \mu_y - \omega \mu_z, \\ \widetilde{\Phi}_y &= \Phi_y - \omega \Phi_z, \ \widetilde{\varepsilon}_{y,t} = \varepsilon_{y,t} - \omega \varepsilon_{z,t}, \end{aligned}$$

and

$$\begin{pmatrix} \widetilde{\varepsilon}_{y,t} \\ \varepsilon_{z,t} \end{pmatrix} = N \left[\begin{pmatrix} 0 \\ 0 \end{pmatrix}, \begin{pmatrix} \Omega_{yy.z} & 0 \\ 0 & \Omega_{zz} \end{pmatrix} \right],$$

for

$$\Omega_{yy.z} = \Omega_{yy} - \Omega_{yz} \Omega_{zz}^{-1} \Omega_{zy}.$$

If $\alpha_z = 0$, the conditional and marginal models are simplified to

$$\Delta Y_t = \omega \Delta Z_t + \alpha_y \beta^{*\prime} X_{t-1}^* + \sum_{i=1}^{k-1} \widetilde{\Gamma}_{y,i} \Delta X_{t-i} + \widetilde{\mu}_y + \widetilde{\Phi}_y D_t + \widetilde{\varepsilon}_{y,t}, \tag{6}$$

$$\Delta Z_t = \sum_{i=1}^{k-1} \Gamma_{z,i} \Delta X_{t-i} + \mu_z + \Phi_z D_t + \varepsilon_{z,t}, \tag{7}$$

and Z_t is then said to be weakly exogenous for the estimation of the following parameters:

$$\alpha_y, \ \beta^*, \ \omega, \ \widetilde{\Gamma}_{y,i}, \ \widetilde{\mu}_y, \ \widetilde{\Phi}_y \text{ and } \Omega_{yy.z}. \tag{8}$$

Note that $\beta^{*\prime} X_{t-1}^*$ is not embedded in the marginal model (7). If the condition for weak exogeneity, $\alpha_z = 0$, is satisfied, the parameters (8) can then be estimated from the conditional model (6) alone without loss of information, with no need for the estimation of the marginal model (7). The conditional model (6) is often referred to as a partial system. See Johansen (1996, Ch.8) for further details of a partial system and weak exogeneity.

4 Canonical Model for Long-Run Economic Linkages

This section, based on recent developments in macroeconomics, presents a canonical model for money demand and monetary policy rule functions. These functions may be subject to a long-run econometric analysis in the cointegrated VAR framework. We consider the two functions in turn.

A standard postulate for the money demand is

$$M_t^d = f\left(P_t, Y_t, R_t\right),$$

where M_t^d denotes the nominal broad money demand, P_t is the price level, Y_t is a measure of scale variable (usually, output), R_t is a vector of rates of return within and outside money, and $f(\cdot)$ denotes a function of a set of variables embedded in the bracket. If the money market clears and the observed broad money stock M_t corresponds to equilibrium, it is then possible to replace M_t^d with M_t. The variable R_t may include several kinds of opportunity costs of holding money as against alternative interest-bearing assets and physical assets, and thus we have

$$M_t = f\left(P_t, Y_t, i_t^s, i_t^l, \pi_t\right),$$

where i_t^s is a short-term interest rate and measures the own rate of money, i_t^l is a long-term interest rate represented by yields on public bonds, and π_t is an annual inflation rate, which proxies for the change in the nominal value of physical assets held. We impose price homogeneity, so that the relationship becomes

$$\frac{M_t}{P_t} = f\left(Y_t, i_t^s, i_t^l, \pi_t\right). \tag{9}$$

A log-linear specification of (9) is often adopted to find

$$m_t - p_t = \gamma_0 + \gamma_1 y_t + \gamma_2 i_t^s - \gamma_3 i_t^l - \gamma_4 \pi_t,$$

where m_t, p_t and y_t denote the logarithms of M_t, P_t, and Y_t, respectively, and $\gamma_i > 0$ for $i = 1, \ldots, 4$.

With regard to the financial market opportunity costs of holding money, we may test the two interest rates, i_t^s and i_t^l, for equal coefficients with opposite sign. The interest rate spread has been considered in Hetzel and Mehra (1989), Johansen and Juselius (1990), Hendry and Doornik (1994), Doornik *et al.* (1998), Ericsson (1998), Juselius (1998), Lűtkepohl and Wolters (1998), and Vega (1998), among others. We may further take unity for the income elasticity, that is, $\gamma_1 = 1$. Thus we have

$$m_t - p_t = \gamma_0 + y_t + \gamma_2(i_t^s - i_t^l) - \gamma_4 \pi_t.$$

Furthermore, it is of interest to test $\gamma_2 = \gamma_4$ as in the following equation:

$$m_t - p_t = \gamma_0 + y_t - \gamma_2(i_t^l - i_t^s + \pi_t). \tag{10}$$

In the empirical analysis conducted in this paper, we consider the money demand function (10) as a candidate for one of the long-run economic relationships embedded in the data.

Let us turn to a monetary policy rule function. Taylor (1993) suggests that the monetary policy process can be summarized by a simple policy rule, in which the short-term policy rate responds to deviations of unemployment and inflation from their policy targets:

$$i_t^s = \pi_t + \delta_0 - \delta_1\left(u_t - u_t^*\right) + \delta_2\left(\pi_t - \pi_t^*\right),$$

where u_t and u_t^* denote the unemployment rate and the natural rate of unemployment respectively, π_t and π_t^* denote inflation and its monetary policy target respectively, δ_0 is interpretable as the target real short-term interest rate, and $\delta_i > 0$ for $i = 0, 1, 2$.

Replacing in the above equation the unemployment with an output on the grounds that the two variables can be closely related, we have

$$i_t^s = \pi_t + \eta_0 + \eta_1 (y_t - y_t^*) + \eta_2 (\pi_t - \pi_t^*), \tag{11}$$

where y_t^* is the natural logarithm of a potential output, and again η_0 is interpretable as the target real short-term interest rate, and $\eta_i > 0$ for $i = 0, 1, 2$. For instance, Clarida *et al.* (1998) use an output gap, instead of unemployment, for the analysis of a monetary policy rule.

We can view the above Taylor rule as a complement to the inflation targeting. While inflation targeting offers a medium-run plan, the Taylor rule may be a short-run operating procedure for a medium-run inflation target. The rule responds also to the output gap, which can be viewed as a measure of inflationary pressure. If the output target is assumed to show a (log) linear deterministic trend *i.e.* $y_t^* = \theta_1 + \theta_2 t$ with $\theta_i > 0$ for $i = 1, 2$ and if the inflation target is further assumed constant *i.e.* $\pi_t^* = \theta_3 > 0$, equation (11) then collapses to

$$i_t^s = \mu_0 + \mu_1 y_t + (1 + \mu_2) \pi_t + \mu_3 t, \tag{12}$$

where $\mu_i = \eta_i$ for $i = 1, 2$, $\mu_0 = \eta_0 - \mu_1\theta_1 - \mu_2\theta_3$, and $\mu_3 = -\mu_1\theta_2$.

Following Clarida *et al.* (1998, 2000) and Christensen and Nielsen (2009), we may stress the forward-looking nature of monetary policy and emphasize the role of expectations. Thus, based on Christensen and Nielsen (2009), we have the following alternative form for (12):

$$i_t^s = \mu_0^* + \mu_1^* \mathsf{E}(y_{t+h} | I_t) + (1 + \mu_2^*) \mathsf{E}(\pi_{t+h} | I_t) + \mu_3^* t, \tag{13}$$

where $\mathsf{E}(\cdot | I_t)$ denotes rational expectation conditional on the information set I_t at time t, and h denotes a forecast horizon. In a linearized model, rational expectations can be linear functions of realized values of the variables concerned:

$$\mathsf{E}(y_{t+h} | I_t) = \rho_{yy} y_t + \rho_{y\pi} \pi_t \quad \text{and} \quad \mathsf{E}(y_{t+h} | I_t) = \rho_{\pi y} y_t + \rho_{\pi\pi} \pi_t.$$

Substituting these expressions into equation (13) yields

$$\begin{aligned} i_t^s &= \mu_0^* + \mu_1^* \left(\rho_{yy} y_t + \rho_{y\pi} \pi_t\right) + (1 + \mu_2^*) \left(\rho_{\pi y} y_t + \rho_{\pi\pi} \pi_t\right) + \mu_3^* t \\ &= \mu_0^* + \left[\mu_1^* \rho_{yy} + (1 + \mu_2^*) \rho_{\pi y}\right] y_t + \left[\mu_1^* \rho_{y\pi} + (1 + \mu_2^*) \rho_{\pi\pi}\right] \pi_t + \mu_3^* t \\ &= \mu_0^* + \mu_1^{**} y_t + \mu_2^{**} \pi_t + \mu_3^* t, \end{aligned} \tag{14}$$

for

$$\mu_1^{**} = \mu_1^* \rho_{yy} + (1 + \mu_2^*) \rho_{\pi y} \quad \text{and} \quad \mu_2^{**} = \mu_1^* \rho_{y\pi} + (1 + \mu_2^*) \rho_{\pi\pi}.$$

Equation (14) is observationally equivalent to (12). The Fisher hypothesis allows us to assume that the long-term interest rate moves along with expectations of inflation. Inflation is already present in (12), thus we add to (12) the information as measured by the real long-term interest rate, $i_t^l - \pi_t$, to find

$$\begin{aligned} i_t^s &= \mu_0 + \mu_1 y_t + (1 + \mu_2) \pi_t + \mu_3 t + \mu_4 \left[\left(i_t^l - \pi_t\right) + \overline{r}^l\right] \\ &= \mu_0^* + \mu_4 i_t^l + \mu_1 y_t + (1 + \mu_2 - \mu_4) \pi_t + \mu_3 t, \end{aligned} \tag{15}$$

where $\overline{r}^l$ is the mean of the real long-term interest rate, assumed to be time-invariant, and $\mu_0^* = \mu_0 + \mu_4 \overline{r}^l$. If there is a one-to-one effect from the long-term interest rate to the

short-term interest rate, then $\mu_4 = 1$, and we obtain from (15) a monetary policy target for the interest rate spread as follows:

$$i_t^s - i_t^l = \mu_0^* + \mu_1 y_t + \mu_2 \pi_t + \mu_3 t,$$

It follows from $\mu_3 = -\mu_1 \theta_2$ that

$$i_t^s - i_t^l = \mu_0^* + \mu_1 (y_t - \theta_2 t) + \mu_2 \pi_t. \tag{16}$$

Both Laurent (1988) and Bernanke and Blinder (1992) suggest the spread between the funds rate and a long-term bond rate as a useful monetary indicator, on the grounds that the long-term rate incorporates the inflationary expectations component of all interest rates but is relatively insensitive to short-run variations in monetary tightness or ease. Mehra (2001) also includes the bond rate as an additional variable in the Taylor rule function. In the empirical analysis performed in this paper, we also regard the monetary policy rule function (16) as a candidate for one of the underlying long-run economic linkages in the data, in addition to the money demand function given by (10).

5 Data Overview

This section presents an overview of quarterly time series data for Korea's macroeconomic series, and discusses possible long-run economic linkages in line with the canonical model described in the previous section. The sample period runs from the first quarter in 1986 to the first quarter in 2008 (denoted 1986.1-2008.1 hereafter). See the Appendix for details of the data.

Figure 1(a) displays time series data of $m_t - p_t$ and y_t. The scale of the figure given along the vertical axis is normalized for $m_t - p_t$. Both of the series exhibit smooth trending features in a similar fashion, suggesting the presence of a close and stable linkage between the two variables. Such a close linkage indicates that the velocity of money, or $m_t - p_t - y_t$, may play an important role in the underlying cointegrating relationships. A large decrease is observed in y_t around the early 1998, corresponding to the aftermath of the Asian currency crisis in 1997. Data plots of $i_t^s - i_t^l$ and π_t are presented in Figure 1(b). Both $i_t^s - i_t^l$ and π_t tend to fluctuate in a synchronized manner, although the level of π_t varies according to the pre and post currency crisis periods.

Figure 1(c) presents data of two combined series, $m_t - p_t - y_t$ and $\left(i_t^l - i_t^s\right) + \pi_t$. The former corresponds to the velocity of money, while the latter is interpreted as the overall opportunity cost in holding money. See the canonical model in the last section for details of the opportunity cost. The scale of the figure given along the vertical axis is normalized for $i_t^s - i_t^l$. Both of the series exhibit trending features in the opposite direction, suggesting that their linear combination leads to a cointegrating linkage interpretable as a money demand function, which is given by (10) in the canonical model.

Lastly, Figure 1(d) displays time series data of $i_t^s - i_t^l$, in addition to those of $y_t - y_t^*$, which denotes a real income or output adjusted for linear trend and constant. The scale of the figure given along the vertical axis is normalized for $i_t^s - i_t^l$. The linear trend may schematically represent potential output, thus $y_t - y_t^*$ could be interpreted as so-called the output gap. It appears that $i_t^s - i_t^l$ and $y_t - y_t^*$ move closely together over time, indicating that they may share common stochastic trends. Figures 1 (b) and (d) suggest that the

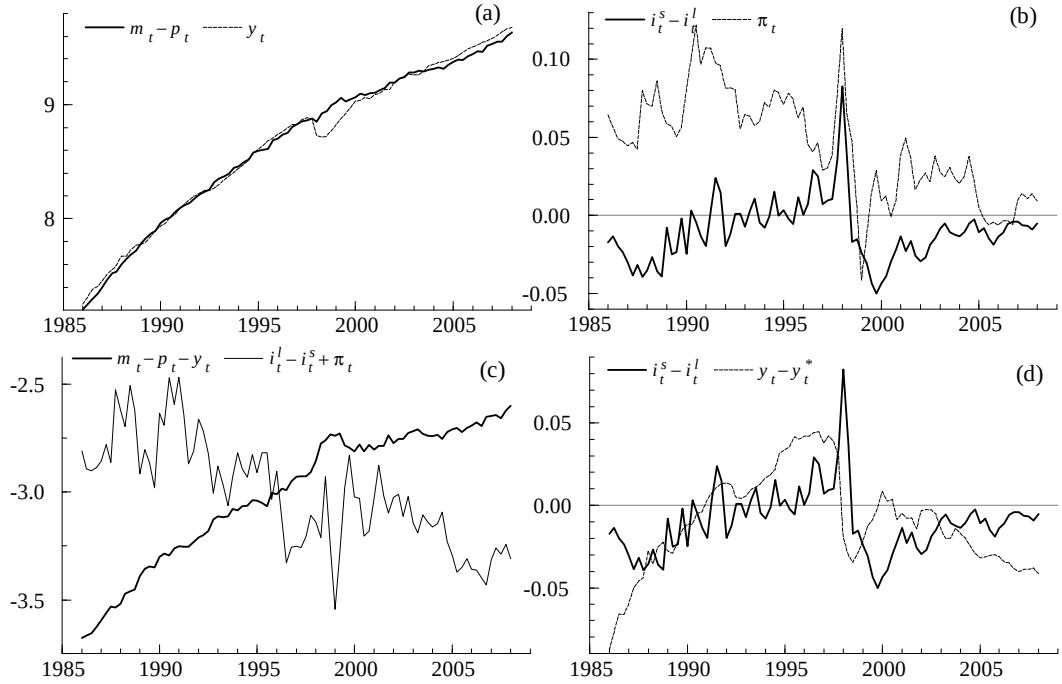

Figure 1: An Overview of the Data

monetary policy rule function given by (16) could correspond to one of the underlying cointegrating combinations.

The overview of the data indicates the possibility of theory-consistent long-run linkages between Korea's macroeconomic series over the past two decades. The next section, using the cointegrated VAR analysis, aims to reveal interpretable long-run economic relationships embedded in the Korean data.

6 Empirical Analysis of Monetary Interaction

We are now in a position to conduct a comprehensive cointegration analysis of Korea's time series data. A set of variables to be analyzed is given by

$$X_t = \left(m_t - p_t, y_t, i_t^s - i_t^l, \pi_t\right)',$$

which yields a four-dimensional VAR system formulated as (1).

This section is composed of four sub-sections. Section 6.1 estimates an unrestricted VAR model paying attention to various structural breaks observed in the data. Section 6.2 determines the number of the cointegrating vectors using Johansen's procedure. Section 6.3 then tests a set of hypotheses on the adjustment and cointegrating vectors so as to investigate weak exogeneity and theory-consistent long-run relationships. Finally, in Section 6.4, a partial equilibrium correction system is estimated conditional on a set of weakly exogenous variables.

6.1 Unrestricted VAR Model

This sub-section estimates an unrestricted VAR model and examines its diagnostic test statistics. The sample period for estimation is 1986.1-2008.1, in line with Figure 1. The lag length of the VAR model is first set to be 4, and an overall F test statistic then rejects the model reduction from length 4 to length 3. Thus the lag length 4 is chosen, and the effective number of observations is therefore 106. The VAR model also includes centered seasonal dummy variables.

The unrestricted VAR model is a purely statistical representation, so that the estimated coefficients are not necessarily subject to an economic interpretation. The VAR model should provide a basis for the subsequent cointegration analysis and thus needs to pass a battery of diagnostic tests. It turns out, however, the estimated model shows strong evidence for non-normal residuals. These are due to outliers caused by policy and regime changes occurring in the sample period. The following dummy variables corresponding to several historical events therefore need to be introduced in the model:

$$D_{1,t} = 1 \text{ (1987.4)}, \; D_{2,t} = 1 \text{ (1997.4)}, \; D_{3,t} = 1 \text{ (1998.1)},$$

and 0 otherwise. The first dummy variable, $D_{1,t}$, captures an outlier found in the residuals of the equation for $m_t - p_t$. A favorable international environment around 1987, which was ushered in by low world interest rate, low oil price, and depreciation of the US dollar against the Japanese yen, led to an improvement in the current and capital account in Korea. The consequent inflow of foreign money into Korea inflated the domestic money stock, resulting in a jump in the inflation rate, particularly late in 1987. The second and third dummy variables, $D_{2,t}$ and $D_{3,t}$, pick out outliers in 1997.4 and 1998.1, corresponding to impacts of the Asian currency crisis in 1997. The currency crisis had an enormous negative effect on the real Korean economy, as shown in the behavior of y_t in Figure 1 (a). The Korean government was forced to request a financial support from the IMF during the crisis, and then started to pursue such policies as abiding by the IMF's recommendations.

	$m_t - p_t$	y_t	$i_t^s - i_t^l$	π_t
Autocorr.[F_{ar}(4,57)]	0.71 [0.59]	2.43 [0.06]	0.18 [0.95]	2.21 [0.08]
ARCH [F_{arch}(4,53)]	1.34 [0.26]	0.85 [0.50]	1.46 [0.23]	1.93 [0.12]
Hetero.[F_{het}(34,26)]	0.49 [0.98]	0.33 [0.99]	1.27 [0.27]	0.87 [0.66]
Normality [$\chi^2_{nd}(2)$]	3.22 [0.20]	2.79 [0.25]	1.42 [0.49]	1.43 [0.49]

Note. Figures in the square brackets are p-values.

Table 1: Diagnostic Tests for the Unrestricted VAR Model

Table 1 presents a set of diagnostic tests for the unrestricted VAR model including the above dummy variables. Most of the test results are given in the form $F_j(k, T-l)$, which denotes an approximate F-test against the alternative hypothesis j: kth-order serial correlation (F_{ar}: see Godfrey, 1978, Nielsen, 2007), kth-order autoregressive conditional heteroskedasticity or ARCH (F_{arch}: see Engle, 1982), heteroscedasticity (F_{het}: see White, 1980). A chi-square test for normality (χ^2_{nd}: see Doornik and Hansen, 1994) is also

provided. All the mis-specification test statistics in the table are insignificant at the 5% level, leading to the judgement that the unrestricted VAR model provides a satisfactory representation of the data. The unrestricted VAR model can therefore be subject to the subsequent cointegration analysis and model reduction.

6.2 Determination of the Cointegrating Rank

This sub-section discusses the choice of cointegrating rank r in the VAR model. Table 2 presents the $\log LR$ test statistics for cointegrating rank, in addition to the modulus of the six largest roots of a companion matrix for the model.

	$r=0$	$r \leq 1$	$r \leq 2$	$r \leq 3$
$-2\log Q\left(\mathsf{H}\left(r\right)\left\|\mathsf{H}\left(p\right)\right.\right)$	145.27[0.00]**	68.55[0.00]**	17.07[0.42]	2.86[0.88]

mod (unrestricted)	0.98	0.91	0.82	0.82	0.80	0.80
mod $(r=2)$	1.00	1.00	0.95	0.83	0.83	0.81

Note. ** denotes significance at the 1% level.

Table 2: Determination of the Cointegration Rank

The $\log LR$ test statistics for the cointegrating rank, denoted by $-2\log Q\left(\mathsf{H}\left(r\right)\left|\mathsf{H}\left(p\right)\right.\right)$ in the first panel of Table 2, sequentially reject the null hypotheses of $r=0$ and $r \leq 1$, thus supporting $r=2$. The second panel of Table 2 provides modulus (denoted *mod*) of the six largest eigenvalues of the companion matrix for two cases, unrestricted and restricted with $r=2$. These are the reciprocal values of the roots of $A(z)$ discussed in the last section. Neither of the two cases contains any eigenvalues greater than 1.0, suggesting that no explosive root is incorporated in the model. All the eigenvalues apart from the first and second ones are distinct from a unit root in the restricted case, although the third eigenvalue is close to unity, which may indicate the presence of some $I(2)$ properties in the data, possibly in $m_t - p_t$ or y_t, according to Figure 1(a). Both $m_t - p_t$ and y_t are not nominal but real variables (real money and income), and it is justifiable to treat them as $I(1)$ variables rather than $I(2)$ on the grounds that $I(2)$ trends are often synonymous with nominal trends (see Kongsted, 2005). In addition, using a testing procedure suggested by Johansen (1996, p.74), we have also checked if each variable is judged to be individually stationary. This test corresponds to a unit root test in the VAR framework. The hypothesis of stationarity is rejected with respect to each variable in the system at the 5% significance level, suggesting that all the variables should be treated as non-stationary series.

We can therefore arrive at the overall conclusion that the model is expressed as an $I(1)$ cointegrated VAR model with the number of the cointegrating relations given by $r=2$. An $I(1)$ cointegrated VAR analysis is conducted with the restriction of $r=2$ in the following sub-sections.

6.3 Fully Specified Adjustment and Cointegrating Vectors

The determination of the cointegrating rank enables us to estimate (3), in which we need to examine joint restrictions on the adjustment and cointegrating vectors. As demonstrated

by Johansen (1996, Ch.3), cointegrating relations represent attractor sets embedded in the VAR system. It is therefore important to identify interpretable cointegrating relations and how adjustment mechanisms work in the system. A series of trials of hypothetical restrictions, motivated by the canonical model in Section 4 and the data overview in Section 5, have led to the following acceptable restrictions:

$$X_t^* = \begin{pmatrix} m_t - p_t \\ y_t \\ i_t^s - i_t^l \\ \pi_t \\ t \end{pmatrix} : \widehat{\alpha} = \begin{pmatrix} * & 0 \\ 0 & 0 \\ * & * \\ 0 & 0 \end{pmatrix}, \ \widehat{\beta}^* = \begin{pmatrix} 1 & 0 \\ -1 & * \\ -b & 1 \\ b & * \\ 0 & * \end{pmatrix}.$$

Asterisks in the α and β^* matrices denote coefficients free from any restrictions, while b indicates a coefficient on which a homogeneity restriction is imposed. Note that the two cointegration relations are normalized with respect to $m_t - p_t$ and $i_t^s - i_t^l$, respectively.

$$\begin{bmatrix} \widehat{\alpha} & [1] & [2] \\ m_t - p_t & \underset{(0.011)}{\mathbf{-0.093}} & 0 \\ y_t & 0 & 0 \\ i_t^s - i_t^l & \underset{(0.016)}{\mathbf{-0.087}} & \underset{(0.125)}{\mathbf{-0.996}} \\ \pi_t & 0 & 0 \end{bmatrix}, \begin{bmatrix} \widehat{\beta}^* & [1] & [2] \\ m_t - p_t & \mathbf{1} & 0 \\ y_t & \mathbf{-1} & \underset{(0.024)}{\mathbf{-0.195}} \\ i_t^s - i_t^l & \underset{(-)}{\mathbf{-3.967}} & \mathbf{1} \\ \pi_t & \underset{(0.536)}{\mathbf{3.967}} & \underset{(0.070)}{\mathbf{-0.528}} \\ t & 0 & \underset{(0.0004)}{\mathbf{0.00153}} \end{bmatrix}$$

Log LR Test Statistic for the Set of Restrictions: $\chi^2(11) = 10.275$ [0.174]

Table 3: Fully Specified Adjustment and Cointegrating Vectors

Table 3 presents a set of restricted estimates as well as the corresponding log LR test statistic. The p-value of the test statistic is 0.17, thus allowing us to conclude that the null hypothesis for the joint restrictions is not rejected at the 5% level. The condition for identifying restrictions, given in Johansen (1996, Ch.7), is satisfied, therefore it is possible to interpret that the restricted cointegration relations represent interpretable long-run economic linkages. Significant coefficients are expressed in bold in Table 3, together with the corresponding standard errors in parentheses.

First, let us consider implications of the two restricted cointegrating vectors. The restricted cointegrating relations are denoted by $c_{1,t}$ and $c_{2,t}$, respectively. According to the right panel of Table 3, the first cointegrating relation is expressed as

$$c_{1,t} = m_t - p_t - y_t + 3.967 \left(i_t^l - i_t^s + \pi_t \right),$$

which exactly matches the money demand function given by (10) in the model. The first relationship, $c_{1,t}$, is therefore interpreted as an empirical representation of money demand in Korea. The second cointegrating relation is, according to the right panel of Table 3, expressed as

$$c_{2,t} = i_t^s - i_t^l - 0.195 \left(y_t - 0.008t \right) - 0.528\pi_t,$$

which coincides with the monetary policy rule function given by (16) in the canonical model. Hence, the second cointegrating relation, $c_{2,t}$, is treated as an empirical representation of the monetary policy rule in Korea. It is noteworthy that these restricted cointegrating relations, estimated from the real-life time series data in Korea, exactly match the theoretical relationships derived in the canonical model. These cointegrating linkages appear to give an overview of the money market in Korea: $c_{1,t}$ represents the demand side while $c_{2,t}$ manifests a policy rule relating to the supply side.

Next, we inspect the structure of restricted adjustment vectors corresponding to $c_{1,t-1}$ and $c_{2,t-1}$, reported in the left panel of Table 3. The first column displays that the adjustment coefficients for $m_t - p_t$ and $i_t^s - i_t^l$ are highly significant, while the remaining coefficients are all insignificant and therefore set to be zero. These results indicate that only $m_t - p_t$ and $i_t^s - i_t^l$ react to disequilibrium errors represented by $c_{1,t-1}$. According to the second column in the panel, the adjustment coefficient of $i_t^s - i_t^l$ is significant with all the other coefficients being zero *i.e.* the variable $i_t^s - i_t^l$ exclusively adjusts to $c_{2,t-1}$.

Let us re-examine the adjustment structure in terms of the equations for $\Delta(m_t - p_t)$ and $\Delta\left(i_t^s - i_t^l\right)$ in the cointegrated VAR system (3). First, we pay attention to $m_t - p_t$, which reacts only to $c_{1,t-1}$, so that the equation for $\Delta(m_t - p_t)$ in the system (3) is expressed as

$$\Delta(m_t - p_t) = -0.093c_{1,t-1} + \cdots,$$

where $\cdots$ represents a set of omitted short-run dynamics. This adjustment structure is in line with the normalization of $c_{1,t-1}$, lending weight to the validity of the interpretation that $c_{1,t-1}$ represents Korea's money demand function. Turning to $i_t^s - i_t^l$, we find that $i_t^s - i_t^l$ adjusts to both $c_{1,t-1}$ and $c_{2,t-1}$, although the coefficient for $c_{1,t}$ is much smaller than that for $c_{2,t-1}$. That is, the equation for $\Delta\left(i_t^s - i_t^l\right)$ in the system (3) is given by

$$\Delta\left(i_t^s - i_t^l\right) = -0.087c_{1,t-1} - 0.996c_{2,t-1} + \cdots,$$

The cointegrating combination $c_{2,t-1}$ is normalized for $i_t^s - i_t^l$, thus the adjustment of $i_t^s - i_t^l$ towards $c_{2,t-1}$ is in accord with the interpretation that $c_{2,t-1}$ corresponds to Korea's monetary policy rule function. The adjustment towards $c_{1,t-1}$ in the equation for $\Delta\left(i_t^s - i_t^l\right)$ seems to indicate the presence of influences of the money demand function on monetary policy, that is, the possibility that Korea's monetary authority allows for the behavior of the overall money demand in the conduct of monetary policy. Hence, the existence of feedback towards $c_{1,t-1}$ could imply interaction between money demand and monetary policy rule.

Furthermore, all the adjustment coefficients for y_t and π_t are set to be zero in Table 3, suggesting that these two variables are judged to be weakly exogenous for parameters of interest given by (8). It is therefore feasible, as discussed in Section 3, to estimate a partial system for $m_t - p_t$ and $i_t^s - i_t^l$, given y_t and π_t, with no loss of information for inference purposes. Using the notational conventions in Section 3, $m_t - p_t$ and $i_t^s - i_t^l$ belong to Y_t while y_t and π_t to Z_t. The two restricted cointegrating relations, $c_{1,t-1}$ and $c_{2,t-1}$, act as equilibrium correction mechanisms in the partial system, which is pursued in the next sub-section.

6.4 A Partial System for $\Delta(m_t - p_t)$ and $\Delta\left(i_t^s - i_t^l\right)$

In this sub-section we report a partial equilibrium correction system for $\Delta(m_t - p_t)$ and $\Delta\left(i_t^s - i_t^l\right)$ and also presents various figures for the purpose of checking the validity of the

model specification.

First, the data are mapped to the $I(0)$ space by differencing and using the restricted cointegrating combinations. A general unrestricted equilibrium correction system for $\Delta(m_t - p_t)$ and $\Delta\left(i_t^s - i_t^l\right)$ is then estimated conditional on Δy_t and $\Delta\pi_t$. Insignificant regressors are eliminated from the system step by step, and it turns out that Δy_t is insignificant in all the equations in the system and therefore removed. As a result of the model reduction, a parsimonious equilibrium correction system is attained as follows:

$$
\begin{aligned}
\Delta(m_t - p_t) \;=\; & -\underset{(0.01)}{0.09}\, c_{1,t-1} - \underset{(0.07)}{0.20}\, \Delta(m_{t-1} - p_{t-1}) - \underset{(0.08)}{0.22}\, \Delta(m_{t-3} - p_{t-3}) \\
& - \underset{(0.10)}{0.34}\, \Delta y_{t-3} - \underset{(0.09)}{0.39}\, \Delta\left(i_{t-3}^s - i_{t-3}^l\right) - \underset{(0.09)}{0.31}\, \Delta\pi_t + \underset{(0.08)}{0.30}\, \Delta\pi_{t-2} \\
& + \underset{(0.08)}{0.17}\, \Delta\pi_{t-3} - \underset{(0.01)}{0.04}\, D_{1,t} - \underset{(0.01)}{0.03}\, D_{2,t} - \underset{(0.01)}{0.04}\, D_{3,t} - \underset{(0.02)}{0.21}\,, \\
& \widehat{\sigma} = 0.01,\ F_{ar}(4{,}58) = 1.65[0.17],\ \chi^2_{nd}(2) = 5.40[0.07], \\
& F_{arch}(4{,}65) = 1.23[0.31],\ F_{het}(38{,}34) = 0.77[0.78], \\
\Delta\left(i_t^s - i_t^l\right) \;=\; & -\underset{(0.01)}{0.09}\, c_{1,t-1} - \underset{(0.10)}{0.93}\, c_{2,t-1} + \underset{(0.05)}{0.10}\, \Delta(m_{t-3} - p_{t-3}) - \underset{(0.09)}{0.21}\, \Delta y_{t-1} \\
& - \underset{(0.08)}{0.34}\, \Delta y_{t-3} - \underset{(0.08)}{0.12}\, \Delta\left(i_{t-2}^s - i_{t-2}^l\right) + \underset{(0.07)}{0.13}\, \Delta\pi_t \\
& - \underset{(0.06)}{0.13}\, \Delta\pi_{t-3} - \underset{(0.01)}{0.05}\, D_{1,t} - \underset{(0.26)}{2.33}\,, \\
& \widehat{\sigma} = 0.009,\ F_{ar}(4{,}58) = 4.01[0.01],\ \chi^2_{nd}(2) = 3.59[0.17], \\
& F_{arch}(4{,}65) = 1.41[0.24],\ F_{het}(38{,}34) = 0.79[0.77],
\end{aligned}
$$

where $\widehat{\sigma}$ denotes the standard error of the regression. A set of seasonal dummy variables significant in the equation for $\Delta(m_t - p_t)$ is omitted for the sake of simplicity. Most of the diagnostic tests are insignificant at the 5% level, suggesting that the partial system appears to be a satisfactory representation of the data.

Figure 2(a) presents the actual and fitted values of $\Delta(m_t - p_t)$ and Figure 2(b) displays the scaled residuals of the $\Delta(m_t - p_t)$ equation. Similarly, the actual and fitted values of $\Delta\left(i_t^s - i_t^l\right)$ are displayed in Figure 2(c) and the scaled residuals of the $\Delta\left(i_t^s - i_t^l\right)$ equation are plotted in Figure 2(d). The overall tracking seems to be satisfactory. The residual density functions are displayed in Figures 3 (a) and (b), and the residual correlograms are also presented in Figures 3 (c) and (d). Consistent with the above test results, no graph indicates any clear evidence for model mis-specification. Parameter constancy is also required in a congruent representation. Figures 3 (e) and (f) show recursive break-point Chow tests (see Chow, 1960), supporting the parameter stability of the preferred partial system.

The overall evidence allows us conclude that the parsimonious model is a data-congruent representation. It should be noted that the data-congruent system has been achieved using the data over the past two decades, without splitting the sample period for estimation. In the parsimonious equilibrium correction system, the first-order differences of all the variables in X_t are significant in various manners, indicating the presence of monetary interaction in the short-run dynamics of the model, in line with the long-run dynamics embodied by $c_{1,t-1}$ and $c_{2,t-1}$.

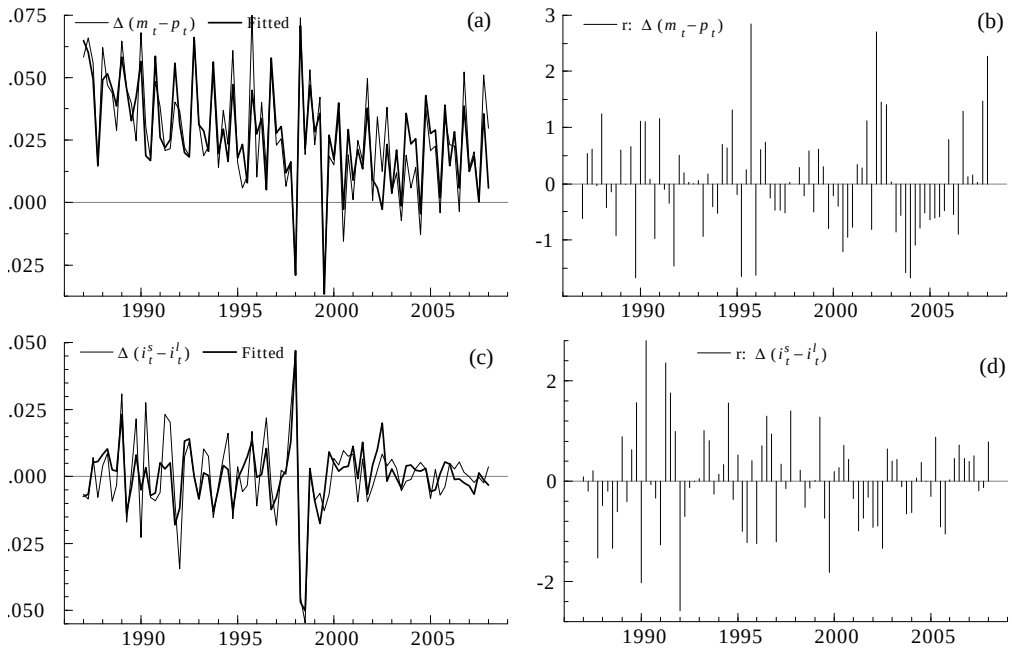

Figure 2: Fitted Values and Residuals Plots

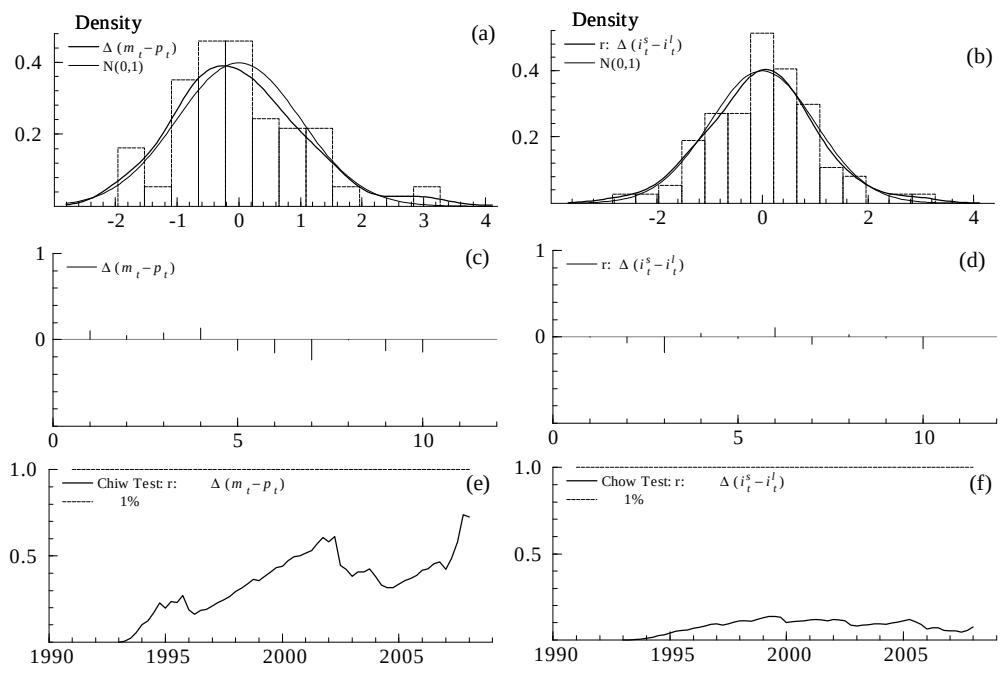

Figure 3: Residual Density, Correlogram and Parameter Constancy

7 Summary and Conclusion

This paper pursues an econometric investigation of monetary interaction in the Korean economy over the past two decades. The study pays close attention to an important role played by broad money and an interest rate term spread in the economy. A vector autoregression of Korea's time series data reveals two cointegrating relationships, both of which are in line with macroeconomic theory: the first relationship corresponds to a broad money demand function, while the second represents a monetary policy rule function. All the variables in the cointegrated system, apart from the broad money and term spread, are judged to be weakly exogenous for parameters of interest, thereby allowing us to estimate a partial system with no loss of information. It is demonstrated that the preferred partial system characterizes interaction between money demand and monetary policy rule in the Korean economy.

Acknowledgements:

The first author started this paper while he was staying as a visiting professor at the Center for Advanced Economic Study, Fukuoka University, during his sabbatical leave from Sejong University. This work was supported by the Faculty Research Fund of Sejong University in 2008. The second author gratefully acknowledges a financial support from JSPS KAKENHI (19830111).

Appendix: Data Definitions and Sources

(Data Definitions)

$m_t - p_t$ = the log of the average of monthly liquidity aggregate of financial institutions (a modification of the previous M3) − the log of the GDP deflator,
y_t = the log of the real GDP,
$i_t^s - i_t^l$ = the overnight call rate − the yield on the national housing bond,
π_t = the percentage change in the GDP deflator over the previous four quarters *i.e.* $\Delta^4 p_t$.

(Sources)

International Financial Statistics, International Monetary Fund.
Economic Statistics System, Bank of Korea.

References

[1] Ahn, S. Oh, S., 1998. The usefulness of interest rates as a monetary policy measures (In Korean), Economic Analysis 4(1), Bank of Korea, 1-30.

[2] Arize, A.C., 1994. A re-examination of the demand for money in small developing economies. Applied Economics 26, 217-228.

[3] Artis, M., Beyer, A., 2004. Issues in money demand: The case of Europe. Journal of Common Market Studies 42, 717-736.

[4] Bahmani-Oskooee, M., Lee, H.J., 1994. Long-run elasticity of the demand for money in Korea: Evidence from cointegration analysis. International Economic Journal 8, 83-93.

[5] Bahmani-Oskooee, M., Shabsigh, G., 1996. The demand for money in Japan: evidence from cointegration analysis. Japan and the World Economy 8, 1-10.

[6] Bahmani-Oskooee, M., Shin, S., 2002. Stability of the demand for money in Korea. International Economic Journal 16, 85-95.

[7] Ball, L., 1997. Efficient rules for monetary policy. NBER Working Paper Series 5952.

[8] Ball, L., Tchaidze, R.R., 2002. The FED and the new economy. American Economic Review 92, 108-114.

[9] Bernanke, B.S., Blinder, A.S., 1992. The Federal funds fate and the channels of monetary transmission. American Economic Review 82, 901-921.

[10] Brűgemann, I., 2003. Measuring monetary policy in Germany: A structural vector error correction approach. German Economic Review 4, 307-339.

[11] Brűgemann, R., Lűkepohl, H., 2006. A small monetary system for the euro area based on German data. Journal of Applied Econometrics 21, 683-702.

[12] Cho, S., 2007. An empirical assessment of the Korean monetary policy since the foreign exchange crisis. The Korean Economic Review 23, 329-352.

[13] Chow, G.C., 1960. Tests of equality between sets of coefficients in two linear regressions. Econometrica 28, 591-605.

[14] Christensen, A.M., Nielsen, H.B., 2009. Monetary policy in the Greenspan era: A time series analysis of rules vs. discretion. Oxford Bulletin of Economics and Statistics 71, 69-89.

[15] Clarida, R., Galí, J., Gertler, M., 1998. Monetary policy rules in practice: some international evidence. European Economic Review 42, 1033-1067.

[16] Clarida, R., Galí, J., Gertler, M., 2000. Monetary policy rules and macroeconomic stability: Evidence and some theory. Quarterly Journal of Economics 115, 147-180.

[17] Doornik, J.A., 1998. Approximations to the asymptotic distribution of cointegration tests. Journal of Economic Surveys 12: 573-593.

[18] Doornik, J.A., Hansen H. 1994. A practical test for univariate and multivariate normality. Discussion paper, Nuffield College, University of Oxford.

[19] Doornik, J.A., Hendry, D.F., 2006. Modelling Dynamic Systems Using PcGive: Volume 2. Timberlake Consultants LTD.

[20] Doornik, J.A., Hendry, D.E., Nielsen, B., 1998. Inference in cointegrating models: UK Ml revisited. Journal of Economic Surveys 12, 533-572.

[21] Eichengreen, B. 2004. Monetary and exchange rate policy in Korea: Assessments and policy issues, Research on Financial Economy 194, Bank of Korea.

[22] Engle, R.F., 1982. Autoregressive conditional heteroscedasticity, with estimates of the variance of United Kingdom inflation. Econometrica 50: 987-1007.

[23] Engle, R.F., Hendry ,D.F., Richard, J.F., 1983. Exogeneity. Econometrica 51: 277-304.

[24] Ericsson, N.R., 1998. Empirical modeling of money demand. Empirical Economics 23, 295-315.

[25] Ericsson, N.R., Sharma, S., 1998. Broad money demand and financial liberalization in Greece. Empirical Economics 23, 417-436.

[26] Gerdesmeier, D., Mongelli, F.P., Roffia, B., 2007. The Eurosystem, the U.S. Federal Reserve, and the Bank of Japan: Similarities and differences. Journal of Money, Credit and Banking 39, 1785-1819.

[27] Godfrey, L.G., 1978. Testing for higher order serial correlation in regression equations when the regressors include lagged dependent variables. Econometrica 46: 1303-1313.

[28] Hafer, R.W., Jansen, D.W., 1991. The demand for money in the United States: Evidence from cointegration tests. Journal of Money, Credit and Banking 23, 155-168.

[29] Hansen, G., Kim, J-R., 1996. Money and inflation in Germany: A cointegration analysis. Empirical Economics 21, 601-616.

[30] Hayo, B., Hofmann, B., 2006. Comparing monetary policy reaction functions: ECB versus Bundesbank. Empirical Economics 31, 645-662.

[31] Hendry, D.F., Mizon, G., 1993. Evaluating dynamic econometric models by encompassing the VAR. In Phillips, P.C.B. ed., Models, Methods, and Applications of Econometrics: Essays in Honor of A.R. Bergstrom. 272-300, Basil Blackwell: Oxford, UK and Cambridge, Massachusetts.

[32] Hendry, D.F., Doornik, J.A., 1994. Modelling linear dynamic econometric systems. Scottish Journal of Political Economy 41, 1-33.

[33] Hetzel, R., Mehra, Y., 1989. The behavior of money demand in the 1980s. Journal of Money, Credit and Banking 21, 455-463.

[34] Hoffman, D.L., Rasche, R.H., 1991. Long-run income and interest elasticities of money demand in the United States. The Review of Economics and Statistics 73, 665-674.

[35] Johansen, S., 1988. Statistical analysis of cointegration vectors. Journal of Economic Dynamics and Control 12, 231-254.

[36] Johansen, S., 1992. Cointegration in partial systems and the efficiency of single-equation analysis. Journal of Econometrics 52, 389-402.

[37] Johansen, S., 1996. Likelihood-Based Inference in Cointegrated Vector Auto-Regressive Models, 2nd printing, Oxford University Press.

[38] Johansen, S., Juselius, K., 1990. Maximum likelihood estimation and inference on cointegration–with applications to the demand for money. Oxford Bulletin of Economics and Statistics 52, 169-210.

[39] Juselius, K., 1998. Changing monetary transmission mechanisms within the EU. Empirical Economics 23, 455-481.

[40] Juselius, K., 2006. The Cointegrated VAR Model. Oxford University Press.).

[41] Khalid, A.M., 1999. Modeling money demand in open economies: The case of selected Asian countries. Applied Economics 31, 1129-1135.

[42] Kongsted, H.C., 2005. Testing the nominal-to-real transformation. Journal of Econometrics, 124, 205-225.

[43] Kurita, T., 2007. A dynamic econometric system for the real yen-dollar rate. Empirical Economics 33, 115-149.ds.

[44] Laurent, R.D., 1988. An interest rate-based indicator of monetary policy. Economic Perspectives (Federal Reserve Bank of Chicago) 12, 3-14.

[45] Lee, T.H., Chung, K.J., 1995. Further results on the long-run demand for money in Korea: A cointegration analysis. International Economic Journal 9, 103-113.

[46] Lűkepohl, H., Wolters, J., 1998. A money demand system for German M3. Empirical Economics 23, 371-386.

[47] Mehra, Y.P., 2001. The bond rate and estimated monetary policy rules. Journal of Economics and Statistics 53, 345-358.

[48] Miyao, R., 1996a. Does a cointegrating M2 demand relation really exist in Japan? Journal of the Japanese and International Economies 10, 169-180.

[49] Miyao, R., 1996b. Does a cointegrating M2 demand relation really exist in the United States? Journal of Money, Credit and Banking 28, 365-380.

[50] Nielsen, B., 1997. Bartlett correction of the unit root test in autoregressive models. Biometrika 84, 500-504.

[51] Nielsen, B., 2007. Order determination in general vector autoregressions. In: H-C. Ho, C-K. Ing and T.L. Lai, Time Series and Related Topics: In Memory of Ching-Zong Wei. IMS Lecture Notes and Monograph 52.

[52] Oh, J. 1998. Inflation targeting and transmission mechanism of interest rate (In Korean), Economic Analysis 4(4), Bank of Korea, 52-91.

[53] Peytrignet, M., Stahel, C., 1998. Stability of money demand in Switzerland: A comparison of the M2 and M3 cases. Empirical Economics 23, 437-454.

[54] Shin, I. 1997. Declining forecasting power of M2 and unstable policy target: a study on monetary and credit policy in Korea since 1980s (In Korean), Journal of Economic Policy 19(3), Korea Development Institute.

[55] Song, W., 1999. How to pursue a monetary policy under inflation targeting, (In Korean), Economic Analysis 5(1), Bank of Korea, 1-30.

[56] Taylor, J.B., 1993. Discretion versus policy rules in practice. Carnegie-Rochester Conference Series on Public Policy 39, 195-214.

[57] Vega, J.L., 1998. Money demand stability: Evidence from Spain. Empirical Economics 23, 387-400.

[58] White, H., 1980. A heteroskedastic-consistent covariance matrix estimator and a direct test for heteroskedasticity. Econometrica 48, 817-838.