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Testing the joint hypothesis of rationality and neutrality under seasonal cointegration: The case of Korea

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TESTING THE JOINT HYPOTHESIS OF RATIONALITY AND NEUTRALITY UNDER SEASONAL COINTEGRATION: THE CASE OF KOREA

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To

My Beloved

Mother

(Lee, Myung Sook)

and

Father

(Chang, Doo Soon)

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ABSTRACT

The monetary policy of Korea of the last two decades was apparently conducted in a systematic and predictable manner, and believed to be effective. However, according to the monetary policy ineffectiveness proposition, predictable monetary policy has no real effect on output. The purpose of this research is to investigate whether the monetary policy of Korea has been effective or not.

The investigation is carried out by testing the so-called joint hypothesis of rationality and neutrality of monetary policy. The test is performed with both seasonally adjusted and seasonally unadjusted data, using a general error-correction multivariate system of money, interest rate, price and output.

Upon establishing that these four variables are integrated of order one, cointegrating relations were investigated with seasonally adjusted data, and one cointegrating relation was found. Similarly, with seasonally unadjusted data, one cointegrating relation at the zero frequency (long-run component) and one at the biannual frequency were found.

With seasonally adjusted data, the joint hypothesis cannot be rejected at the 5 percent confidence level, but the test result is ambiguous at the 10 percent level. With seasonally unadjusted data, the hypothesis cannot be rejected

at levels much higher than 10 percent level. The difference between the two test results, although somewhat marginal, is the consequence of exploiting the additional information contained in seasonally unadjusted data.

The result of this research is relevant for both macroeconomic policy and econometric practice. From the macroeconomic viewpoint, it provides evidence of money neutrality and rationality in Korean data. Its policy implication is that a fixed money supply rule such as Friedman's k percent rule is preferable to more complex countercyclical money supply rules aimed at fine-tuning the economy. From the econometric viewpoint, it offers evidence that seasonal adjustment procedures, such as the commonly used X-11 procedure, may eliminate useful information contained in the original data. It follows that using seasonally unadjusted data rather than adjusted data for empirical work may be more appropriate.

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INTRODUCTION

1.1 The Theoretical Debate

Since the publication of Keynes's (1936) "General Theory", it has been generally accepted that government policy is desirable to smooth fluctuations in the economy (fine-tuning), as it has significant effects on real output and unemployment. This acceptance has received further support from Phillips' (1958) empirical finding that there exists an inverse relationship between the growth rate of nominal wages and the rate of unemployment.

However, in the late 1960s many western industrial countries experienced high sustained levels of unemployment and inflation simultaneously, which the Phillips curve could not explain. This led to the theoretical attack on the Phillips curve by Phelps (1967) and Friedman (1968). Friedman argued that there is no true theoretical basis for the Phillips curve and that it is the real wage, and not the nominal wage, which matters in the labor market. He also argued that, as the existence of a natural rate of unemployment implies a vertical Phillips curve in the long-run, any inflation-unemployment trade-off must be temporary.

These arguments were further developed by the introduction of rational expectations into the debate.[1] The new-classical economists, for example, argue that a trade-off relationship between inflation and unemployment or output may not exist even in the short run, and that even an accelerating inflation cannot reduce unemployment below the natural rate as long as government policies are pre-announced or predictable.[2] Only unanticipated events may shock the economy away from the natural rate of unemployment. This leaves little scope for activist fine-tuning policies through monetary instruments, as long as these can be accurately anticipated. This is the essence of the so-called monetary policy ineffectiveness proposition (henceforth, the MPIP), proposed by Lucas (1972a, 1973), Sargent and Wallace (1975, 1976) and Barro (1976a) under the framework of market-clearing, rational expectations and imperfect information.

Regardless of whether the MPIP is derived from a general equilibrium model (Lucas 1972a, 1973; Barro 1976a) or from an

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[1] The idea of rational expectations was originally introduced by Muth (1961), but macroeconomists began to realize its implication only in the 1970s. The central idea of rational expectations is that agents are rational enough to avoid biased and predictable forecast errors. Its main implication is that one-step ahead forecast errors must have zero conditional mean, and must be uncorrelated with past errors.

[2] In particular, Lucas (1972b) argued that the hypothesis that agents form expectations adaptively does not lead to the hypothesis of a natural rate of output. On the contrary, the two hypotheses are mutually contradictory, whereas the hypothesis of rational expectations may lead to the natural rate theory.

ad hoc linear macroeconomic model (Sargent and Wallace 1975, 1976)[3], the arguments are similar and can be summarized as follows.

First, fully anticipated changes in the money supply have no effect on real output because if the public's predictions are accurate, these changes would already be incorporated into the agents' expectations. However, even if suppliers and demanders form their expectations rationally about future prices, imperfect information may prevent market participants from discriminating whether a price disturbance comes from real disturbance due to a shift in demand or from nominal disturbance due to a change in the money supply. Thus, in this framework, only changes in the money supply which are not fully anticipated as monetary disturbances, and therefore are partly interpreted as real disturbances, may affect real output.

Second, a higher variance of the money supply makes it more difficult for individuals to react appropriately to real changes. Suppliers are more likely to attribute a larger fraction of observed price movements to fluctuations in the general price level due to changes in the money supply (nominal disturbances) than to fluctuations in the relative price due to shifts in demand (real disturbances). Thus, the

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[3] Sargent and Wallace's model is ad hoc in that their model is not derived from a consistent set of assumptions about optimal behaviors of individuals and firms, as opposed to Lucas (1972a).

more unpredictable the money supply is, the smaller the effect on real output of any given unpredictable movement in the money supply (Lucas 1972a, 1973, 1975).

Third, a policy implication of the ineffectiveness proposition is that a fixed money supply rule such as Friedman's (1959) k percent rule is preferable to the complex discretionary money supply rule such as countercyclical money supply, because under rational expectations the former minimizes the variance of real output.[4]

Fourth, suppose the monetary authority has better information about the money supply rule than the public, which enables the authority to deceive the public systematically in the short-run.[5] Nonetheless, when the costs of providing that information to the public are negligible, the provision of the superior information to the public has identical implications for output as when the monetary authority lacks superior information.

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[4] Barro and Gordon (1983a, 1983b) argued that a fixed rule for money growth is better than a discretionary policy because the former can reduce the equilibrium rates of inflation and monetary growth. On the other hand, Buiter (1981) contended that contingent rules are superior to optimal fixed rules because the former permits new information to be considered when the policy instruments are selected, whereas fixed rules do not allow any response to new information.

[5] Taylor (1975) contended that a deceptive monetary policy can affect output during transition periods during which individuals gradually learn the policy rule. On the other hand, Sargent and Wallace (1976) maintained that there is no feedback rule for the government that can systematically deceive the public even in the short-run.

The MPIP has been criticized on several grounds. Among the many arguments and different contexts that have been proposed, two theories have caught most of the attention[6].

First, monetary policy can affect real output even under rational expectations when rigidities in the labor and/or goods markets prevent these markets from clearing (for a comprehensive survey, see Blanchard 1990). Rigidities can arise from staggered labor contracts (Fischer 1977a, 1977b), from a non-simultaneous determination of wages across firms (Taylor 1980), from wage indexation (Gray 1976, 1978), from price stickiness (Phelps and Taylor 1977; Parkin 1986), from slow responsiveness of inventory changes (Frydman 1981; Blinder 1982), or from menu costs (Blanchard and Kiyotaki 1987; Akerlof and Yellen 1985; Mankiw 1985).[7]

Second, the MPIP is derived from the basic assumption of the separated island economy paradigm, namely that current information set is not universally available to economic

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[6] Extensions and critiques include real money balance term (Jansen 1985), non-linear macroeconomic system (Snower 1984), different probability distributions (Otani 1985), agents' forward-looking behavior (Marini 1988), intertemporal substitution effects (Marini 1985; and Jansen 1990 for opposite arguments), temporary monetary disturbances (Waldo 1982; Goodfriend and King 1981).

[7] Against this critique, Barro (1977b) argued that the way by which employment is determined in Fischer's contracting procedure is not Pareto optimal, as there could exist room to improve the welfare of both households and firms. See also Mankiw (1988, 1990) for criticisms of long-term contracting, Barro (1976b) and Karni (1983) for arguments against wage indexation, and McCallum (1977, 1978, 1980, 1981) and Caplin and Spulber (1987) for arguments against price stickiness.

agents: local information is available, global information is not. Therefore, if a different information structure is built into the economic model, monetary policy can affect the behavior of real output. For instance, when individuals can access current information such as contemporaneous economy-wide interest rates (McCallum 1980), or when information of the investment profitability is distributed asymmetrically across private agents (Weiss 1980), or when expectations of demand and supply are based on different information sets (Scarth 1988, Ch. 6), countercyclical monetary policy can be effective. Against one of these critiques, Barro (1980) developed a model that includes an economy-wide interest rate and yet the MPIP still remains valid.

In parallel with theoretical developments, the MPIP has been tested empirically. The results are mixed, and differ for type of data and countries: some research corroborates the proposition; some reject it by finding that anticipated policy also has a significant effect on real variables; finally, some research finds that neither unpredicted nor predicted policy is significant (for summary, see Goodhart 1989, p310).

1.2 The Purpose of the Study

During the last three decades of export-oriented economic development of Korea, the government has depended on monetary

policy more often than on fiscal policy, partly because of the quicker response of the former and partly because of the great rigidity of the government budget, especially of defence expenditure. It has been generally believed, particularly by policy makers and business community, that monetary policy was effective during the course of economic development in Korea.

The Korean domestic financial market has been isolated from the international market under strict foreign exchange control: the foreign exchange system has operated under the fixed exchange rate regime until 1990. Thus, by and large, the Korean economy can be described as a small open economy under fixed exchange rate regime and with a capital market not integrated with the international capital market.

According to economic theory, under these circumstances, monetary policy does not affect output, as, for instance, an expansion of the money supply leads to a balance of payments deficit, which in turn, leads to a loss of international reserve and eventually to a reduction in the money supply. However, for the case of Korea, the chronic current account deficits until 1985 were partly offset by regulated capital inflows, and the sterilization of the money supply contractions through lower reserve requirement ratios and selective credit after arguments in support of a degree of effectiveness of monetary policy.

On the other hand, as discussed, for instance, in Cole and Park (1983), the monetary authorities appear to have

implemented a predictable policy: an expansion of the money supply when the economy was in a recession, and a contraction when the economy was in a boom. This implies, in our context, that the monetary policy of Korea should not have affected real output.

In summary, the monetary policy of Korea is believed to be effective, but at the same time it is believed to have been conducted in a systematic and predictable manner, thus implying ineffectiveness. The purpose of this research is to resolve this contradiction, by directly testing whether the MPIP holds or not for the Korean economy. As the anticipated component of the money supply dominates its unpredictable component, the test is expected to support the hypothesis of money neutrality or ineffectiveness. This study for Korea is new in the literature, although other researchers have tested related issues such as the Lucas hypothesis (Jung 1985; Kormendi and Meguire 1984; Alberro 1981) and money-income causality (Atesoglu and Tillman 1980).

A second purpose of this research relates to the type of data used. For consistency with econometric practice in Korea and elsewhere, the test will be carried out with seasonally adjusted data. However, as Korean data are published in a seasonally unadjusted form, the test will also be carried out with these data. Thus, a comparison of the two procedures will be available.

1.3 Organization of the Study

This dissertation is organized as follows: Chapter II critically reviews a number of empirical tests of the MPIP hypothesis for the U.S. economy. Chapter III provides a brief description of the monetary policy of Korea, a review of the literature dealing with Korean issues concerning the MPIP, and analyzes the data that will be used for the empirical tests. Chapter IV describes the model chosen to test the joint hypothesis of rationality and neutrality for Korea using seasonally adjusted data, and reports the empirical results. Chapter V analyzes the model by using seasonally unadjusted data and reports the empirical results. Chapter VI offers some concluding remarks and policy implications.

Chapter II

THE MPIP HYPOTHESIS FOR THE U.S. CASE

This chapter will review a number of empirical investigations on the monetary policy ineffectiveness proposition, or the MPIP, starting with the influential work of Barro (1977a, 1978, 1981) and Mishkin (1982a, 1982b, 1983). Barro's and Mishkin's models have essentially the same structure: the effects of monetary policy, if any, relate only to the transitory fluctuations of output around the historical trend; the monetary policy instrument is identified with the growth rate of the money supply; the transitory fluctuations of output are identified with the stationary process that results from detrending the historical series; agents make homogeneous forecasts of the growth rate of the money supply.

Barro and Mishkin tested the MPIP with the following model:

$$A(L)X_t = Z_{t-1}\gamma + u_t, \tag{2.1}$$

$$Y_t = Y_{nt} + B(L)(X_t - Z_{t-1}\gamma^*) + C(L)Z_{t-1}\gamma^* + \epsilon_t, \tag{2.2}$$

where X_t is the growth rate of the money supply, Z_{t-1} is a vector of macroeconomic variables used to forecast X_t; u_t is the forecast error (e.g., the unexpected component of X_t) and it is assumed to be uncorrelated with Z_{t-1}; Y_t is aggregate real output; Y_{nt} is the historical trend (or natural level) of

Y_t, and $Y_t - Y_{nt}$ is the transitory component of output. A(L), B(L) and C(L) are polynomials in the lag operator L (e.g., such that $X_{t-k} = L^k X_t$) and γ, γ^* are vectors of parameters.

The model (2.1) and (2.2) is interpreted as follows: rationality of expectations requires that $\gamma = \gamma^*$, i.e., the growth rate of money expected by the agents, $Z_{t-1}\gamma^*$, must be equal to the predictable component of the process that generates it, $Z_{t-1}\gamma$; neutrality requires that C(L) = 0, i.e., output fluctuations are affected only by the unanticipated component of the money supply.

In Barro's model, X_t is the log-difference of the annual average of the daily M_1 measure of money and A(L) is order of two. Z_{t-1} contains a measure of federal government expenditures to capture a fiscal link to money creation, and a lagged unemployment rate to capture the countercyclical response of monetary policy to the level of economic activity. Y_t is the log of real GNP to capture the apparently exponential growth of output and Y_{nt} is a linear time trend. The order of both B(L) and C(L) is three.

Barro estimated his model by using a two-step estimation procedure: equation (2.1) is first estimated by ordinary least squares; then the residuals u_t from this regression are used as the unanticipated component of money growth in equation (2.2). By estimating the latter with $\gamma = \gamma*$, Barro implicitly tested the joint hypothesis of rationality and neutrality

(Attfield et al. 1985, Ch. 6 and 7). His result was supportive of the policy ineffectiveness proposition.

In Mishkin's model, X_t is the log-difference of the quarterly average of daily M_1 and A(L) is order of four. Z_{t-1} contains four lagged values of 90-day treasury-bill rate and a measure of high-employment government budget balance.[1] Y_t is the log of real GNP and Y_{nt} is again a linear time trend. B(L) and C(L) have the same order, set at three different values: seven, seventeen and twenty. Criticizing Barro's two-step procedure[2], Mishkin estimated equations (2.1) and (2.2) simultaneously, using a joint non-linear estimation procedure based on Leiderman (1980). He rejected the joint hypothesis of rationality and neutrality, and contended that this is due primarily to a rejection of neutrality rather than rationality. This result supported the practice of assuming rationality of expectations in constructing macro models.

Barro's and Mishkin's work led to considerable research in this area during the past decade. Their models have been criticized on several grounds, especially in light of important developments in econometrics and time series analysis. This chapter will review a number of critiques.

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[1] The high-employment budget balance is an estimate of the budget that would prevail at 6 percent unemployment rate. Mishkin chose it as a proxy of discretionary fiscal policies.

[2] See also Pagan (1984), who argued that Barro's procedure produces consistent estimates, but t- and F-statistics are overstated.

First, the lag structures of Barro's and Mishkin's models were arbitrarily chosen. Second, their models assume that output is trend-stationary, - i.e., the series is made stationary by detrending, as opposed to more recent literature arguing that output may be difference-stationary, - i.e., the series is made stationary by differencing (for example, Nelson and Plosser 1982); if this is so, their results can be spurious (Nelson and Kang 1984). Third, recent literature has used a multivariate system approach - mostly in the form of vector autoregressive (VAR) models - that imposes fewer restrictions and allows for bidirectional or multidirectional causality among variables (Granger 1969; Sims 1972), whereas Barro and Mishkin only considered unidirectional causality from money (X_t) to the transitory component of output (Y_t-Y_{nt}). In general, if other variables are added to equation (2.2), the explanatory power of X_t on Y_t-Y_{nt} could be affected. Sims (1980b), for example, found that in a four-variables model with money, industrial production, wholesale prices and interest rate, unexpected money growth explained industrial production much less than in the three-variables model without interest rate. Fourth, some recent literature has indicated that VAR models in first-differences may be misspecified if the variables are cointegrated (Engle and Granger 1987; Johansen 1988). Fifth, both Barro and Mishkin, as well as many other researchers, used seasonally adjusted data in their empirical works. As pointed out in the literature, however,

seasonal adjustment might lead to spurious inference. Finally, Barro's and Mishkin's results can be criticized according to Lucas' (1976) critique, in the sense that they assume the polynomials B(L) and C(L) in (2.2) to be time-invariant, whereas in fact they may be affected by different policy rules.

These issues are now considered in detail.

2.1 Lag Structure

Barro set B(L) at two/three lags and C(L) at three/four lags, arguing that these lag structures turned out to be empirically significant in explaining Y_t. He did not use more formal procedures to select lags, such as Schwarz's (1968) criterion (SC) or Akaike's (1969) final prediction error criterion (FPE).[3] Moreover, he considered a dynamic structure only for B and C in (2.2), and ignored the possibility of autocorrelations in the error term ϵ_t.

Mishkin arbitrarily set both B(L) and C(L) at seven, seventeen and twenty lags in his tests, arguing that a longer lag structure is a proper strategy for analyzing the

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[3] The criterion of SC and FPE is to minimize
SC = (SSR/T) + KlogT/T,
FPE = (SSR/T) x (T+k/T-k),
where SSR: sum of squared residuals,
T: sample size,
k: number of regressors (Judge et al. 1985).

robustness of the results, as a decrease in test power is compensated by the correctness of the test statistics.

Some authors have shown that the result of the MPIP test is vulnerable to lag structure. For example, Sephton (1990), using a four-variable VAR model, showed that the lag structure chosen through the SC supported the MPIP, but rejected it when using the FPE.[4] As no hard rule exists for the selection of lags in a multivariate setting, (see also Hafer and Sheehan 1991), careful attention will be given in this study to determine the lag structure of the model.

2.2 Stationary Representation of Variables

Both Barro and Mishkin have used a time trend to capture the historical mean of output (Y_{nt}). At the basis of this specification - a practice employed by most of the empirical research up to the early 1980s - is the assumption that money and output, as well as many other economic variables, are generated as trend-stationary processes (TSP), that is, as

[4] The FPE puts more importance on unbiasedness over efficiency while the SC places more weight on efficiency. In general, the FPE procedure suggested by Hsiao (1981) for use in Granger causality tests does not choose the true model as the sample size increases and the SC procedure tends to underfit the true model in small samples but perform well in case of more than two hundred observations (Geweke and Meese 1981).

zero-mean stationary process added to deterministic functions of time.

However, as discussed first in Nelson and Plosser (1982), most macro economic time series, including money and output, seem to be better characterized as difference-stationary processes (DSP), that is, as processes that need to be differenced once or more times to become stationary. As standard asymptotic distribution theory is based on stationarity, it is crucial to assess whether a non-stationary process can be transformed into a stationary one by time detrending or by differencing. Some literature (Nelson and Kang 1984; Phillips 1987; Durlauf and Phillips 1988) has shown that the use of one transformation, when in fact the process belongs to the other class, can seriously invalidate results.

To explain differences of behavior between the two types of process, consider the two following simple models:

$$y_t = bt + \omega_t, \tag{2.3}$$

$$\Delta y = b + \mu_t, \tag{2.4}$$

where Δ means first difference, i.e., $\Delta y_t = y_t - y_{t-1}$, and ω_t and μ_t are zero-mean uncorrelated processes (white noise). Equation (2.3) describes y_t as TSP, that is as a stationary process ω_t added to a linear time trend bt. Equation (2.4) describes y_t as DSP, that is as a process that becomes stationary (in this case, a white noise μ_t) after differencing. By integrating (2.4) from the starting time of

the process $t_0 = 0$ to the current time t, we get (assume the initial value to be zero):

$$y_t = bt + \sum_{j=0}^{t} \mu_{t-j}.$$

Thus, we see that even though the unconditional mean of both processes is the same, the unconditional variance is quite different: TSP has a constant and finite variance (the variance σ_ω^2 of the stationary process ω_t), while DSP exhibits a linearly increasing variance, e.g., $t\sigma_\mu^2$.

As the DSP (2.4) is transformed into a stationary process by differencing only once, it is called an integrated process of order one, or I(1). Processes that need to be differenced α times are said integrated of order α, or I(α). By extension, a stationary process is indicated as an integrated process of order zero, or I(0). I(α) processes are characterized by having α unit roots in their autoregressive structure. I(1) processes are very smooth, whereas I(0) processes look more irregular. The expected time for I(1) processes to cross their unconditional mean is infinite, whereas I(0) processes cross it frequently. I(1) processes have autocorrelations that never decay (long memory), while the autocorrelations of I(0) processes go to zero asymptotically (short memory). Thus, the regression of the current value on its distant past is significant for I(1) processes but not for I(0) processes. The spectral power at

the zero frequency is infinite for I(1) processes, and finite for I(0) processes. In practice, I(1) processes exhibit a tall and narrow peak at the low frequencies.

To distinguish between TSP or DSP, a unit root test was developed by Dickey-Fuller (1979, 1981) and extended by Perron (1989) for the case of structural breaks. Based on the evidence that output and money appear to be both I(1), most recent empirical work on the MPIP adopts first differencing instead of time-detrending (see, among others, Ermini and Chang 1993; Bohara 1991; Sephton 1990; Spencer 1989; Manchester 1989; Stock and Watson 1989; McGee-Stasiak 1985).

However, as recently noted by Ghysels and Perron (1993), when seasonally adjusted data is used, these tests may be biased toward rejection of the TSP hypothesis. On the other hand, when seasonally unadjusted data is used, unit roots can appear at frequencies different from zero, thus requiring an econometric treatment that has been developed only recently. This topic will be discussed in detail in section 2.5 and 3.3.3.

2.3 Multivariate Approach

Barro's and Mishkin's models implicitly assumed that money affects the transitory component of output, but not vice versa. Furthermore, other variables, such as interest rate

and unemployment rate, were introduced only as exogenous. In a more general formulation, these exogenous variables could in fact appear in the set of dependent variables as well, thus expanding Barro's and Mishkin's systems beyond the dimension of two. Both restrictions (unidirectional causality and two-dimensional framework) may limit the validity of Barro's and Mishkin's results.

One way to cope with these two problems is to use the vector-autoregressive (VAR) approach, whose application to macroeconomics has been initiated by Sims (1980a, 1980b). For a brief description of the VAR approach, consider a vector random process W_t and define $w_t = W_t - E(W_t | W_{t-i}, \text{ all } i > 0)$ as the innovation of W_t, where $E(W_t | W_{t-i})$ means expectation of W_t conditional on W_{t-i}. Under the conditions of the Wold decomposition theorem, W_t can be represented as an infinite moving average process of its own innovations (Granger and Newbold 1986):

$$W_t = D(L)\, w_t, \tag{2.5}$$

where D(L) is a matrix polynomial, in general of infinite order, and where $E(w_t) = 0$ and $E(w_t w_s') = 0$ for all $s < t$ by construction (e.g., w_t is a white-noise process). If the inverse of D(L) exists (call it E(L)), (2.5) can be reparametrized as a vector-autoregressive (VAR) process,

$$E(L)\, W_t = w_t, \tag{2.6}$$

where the order of the polynomial E(L) is not necessarily infinite. In what follows, (2.6) will be viewed as the generating mechanism of the random process W_t from its own innovations w_t.

Consider now the arbitrary decomposition of W_t into two subvectors W_{1t} and W_{2t}. Then (2.6) can be rewritten as

$$\begin{bmatrix} F(L) & -G(L) \\ -H(L) & I(L) \end{bmatrix} \begin{bmatrix} W_{1t} \\ W_{2t} \end{bmatrix} = \begin{bmatrix} w_{1t} \\ w_{2t} \end{bmatrix}, \tag{2.7}$$

where negative signs are introduced only to emphasize the distinction between dependent and independent variables in each equation.

Models usually estimated in the literature are often a special case of (2.7), obtained for H(L)=0:

$$F(L)\,W_{1t} = G(L)\,W_{2t} + w_{1t}, \tag{2.8}$$

$$I(L)\,W_{2t} = w_{2t}. \tag{2.9}$$

In this case, the vector W_{1t} is generated by its own innovations and by current and past values of W_{2t}, and it is called a VARX process (the additional X to indicate the presence of exogenous variables). The vector W_{2t} is generated only by its own innovations as a sub-VAR process. In the literature, usually only the VARX component (2.8) is estimated without simultaneously considering the presence of the sub-VAR structure (2.9).

Barro's and Mishkin's specifications (2.1) and (2.2) are special cases of (2.7), obtained by imposing a number of restrictions. It is readily seen that (2.1) and (2.2) can be written in the format of (2.7) as follows:

$$\begin{bmatrix} 1 & -B(L) & (B(L)-C(L))\gamma^* \\ 0 & A(L) & -\gamma \\ 0 & 0 & I(L) \end{bmatrix} \begin{bmatrix} Y_t \\ X_t \\ Z_{t-1} \end{bmatrix} = \begin{bmatrix} Y_{nt} \\ 0 \\ 0 \end{bmatrix} + \begin{bmatrix} \epsilon_t \\ u_t \\ \nu_t \end{bmatrix}, \tag{2.10}$$

where the block matrices and sub-vectors

$\begin{bmatrix} 1 & -B(L) \\ 0 & A(L) \end{bmatrix}$, $\begin{bmatrix} [B(L)-C(L)]\gamma^* \\ -\gamma \end{bmatrix}$, $[0\ 0]$, $[I(L)]$, correspond to F(L), -G(L), -H(L), I(L) in (2.7) respectively, and $\begin{bmatrix} Y_t \\ X_t \end{bmatrix}$ and Z_{t-1} relate to W_{1t} and W_{2t} respectively. In addition to imposing H(L) = 0, Barro and Mishkin ignored the sub-VAR structure $I(L)Z_{t-1} = \nu_t$ in estimating (2.10). In addition to these structural assumptions, Barro and Mishkin also imposed further restrictions such as the a priori setting of many coefficients of B(L) and C(L) to zero, and the orders of serial correlation and cross-serial correlation of the error terms. (For a detailed discussion, see Lucas and Sargent 1979).

Recent literature has taken up the two issues of unidirectional causality and two-dimensional system, by recasting the MPIP in a VAR framework. Frydman and Rappoport (1987) found in simultaneous regression of two equations of Mishkin type that money affected transitory output in the

short run, irrespective of whether money was rationally anticipated or not. McGee and Stasiak (1985) tested the MPIP in a log-difference three variable VAR model (money, output, price). Despite using a DSP representation of the three variables, McGee and Stasiak's results supported Mishkin's finding that anticipated policy affected transitory output. McGee and Stasiak's work was extended to four variable by Sephton (1990) by including interest rate. He questioned McGee and Stasiak's conclusion as being sensitive to the exclusion of the nominal interest rate, which McCallum (1983) argued should be included in a test of whether anticipated policy matters. Spencer (1989) tested with four variables, using various stationary specifications and lag structures. He found that money innovations affected output in a TSP framework, but not in DSP. Litterman and Weiss (1985) found similar results by including the real rather than the nominal interest rate in a four-variable VAR model in log levels.

More recently, Bohara (1991) tested the joint hypothesis of rationality and neutrality by extending Mishkin's model into a three-variable log-difference VARX model, where interest rate acted only as an exogenous variable. Even though Bohara's framework is different from Mishkin's in that he used a statistical procedure (FPE) to select the appropriate lag order, and considered output, money and interest rate as difference stationary, his results were similar to those of Mishkin: he rejected the neutrality while

he could not reject the rationality. Finally, Ermini and Chang (1993), criticizing Bohara's model for its lack of consideration of a possible long-run cointegrating relationship between money and output, tested the joint hypothesis with a trivariate error-correction model (see also next section), in which interest rate also acted as endogenous. Interestingly, the result is reversed, in that neutrality is no longer rejected. Ermini and Chang argued that the reversal of the test result is due to the inclusion of the error-correction term, which enhances the forecastability of money growth. This issue will be further discussed in the next section.

On a related topic, Christiano and Ljungqvist (1988) tested money-output causality in a bivariate VAR model to find out that the unanticipated component of money has statistical significance in forecasting real output. They compared results in levels and first differences and found that money-output causality was statistically significant when data was measured in log levels but not when they were measured in first differences of the logs.[5] Stock and Watson (1989) extended this VAR approach in testing money-output causality along various directions. They found that the deviation of money growth from a linear time trend Granger causes growth in

[5] Eichenbaum and Singleton (1986) also found remarkable reductions in the importance of money in three or four-variable VAR models when the tests were performed by using log differences rather than log levels with a linear time trend.

industrial production in a bivariate system. They also obtained similar results in a trivariate (price was added) and in a four-variable system (interest rate was added) as well.

The VAR approach has been criticized in the literature. Some authors have argued that VAR models are atheoretical (Cooley and LeRoy 1985; Leamer 1985) in that VAR approach does not use economic theory during the building process. Others argue that it yields inconsistent estimates when there is no restriction on contemporaneous correlation of residuals (Keating 1990). A crucial critique is that estimates of VAR models vary greatly according to alternative orderings of variables (Spencer 1989; Runkle 1987). Spencer observed that when money appeared before output in the ordering of variables, the proportion of the transitory component of output explained by unpredicted money growth would be greater than when output appeared before money. He also found that it was the relative ordering of money and output and not the relative ordering of money and other variables that was crucial to the explanation of unanticipated money on transitory output.

2.4 Cointegration and Error-Correction Models (ECM)

When a model is in levels but the variables are I(1), test results could be spurious due to the nonstationarity of the residuals (for example, Granger and Newbold 1986, p205). Furthermore, first differencing of I(1) variables to obtain stationarity can also be misleading if the variables are cointegrated. The error-correction model (ECM) has the merit of overcoming these two problems simultaneously.

To summarize briefly, in general, any linear combination of two integrated processes is also an integrated process; for example, if y_t and z_t are I(1), then in general $\alpha y_t + \beta z_t$ is also I(1), for arbitrary constants α and β. However, there may exist a particular linear combination (γ_1, γ_2) such that $\gamma_1 y_t + \gamma_2 z_t$ is stationary, i.e., it follows an integrated process of reduced order. In this case, the two processes are said to be cointegrated and (γ_1, γ_2) is called the cointegrating vector (for detail, see Granger and Newbold 1986; Engle and Granger 1987, 1991). Cointegration between I(1) processes can be interpreted as an equilibrium condition, as it implies that deviations between two processes are stationary even though the processes themselves are I(1), and therefore they move together in the long run without drifting too far apart.

The Granger representation theorem (Engle and Granger 1987) states that if two I(1) variables y_t and z_t are

cointegrated, then their VAR representation in levels can be equivalently rewritten as an error-correction representation in first differences, that is, a VARX model with an exogenous term (the error-correction term) which corresponds to their cointegrating relation. In symbols, if

$$[A^*(L)]\begin{bmatrix} y_t \\ z_t \end{bmatrix} = \begin{bmatrix} \varepsilon_t^y \\ \varepsilon_t^z \end{bmatrix} \tag{2.11}$$

is the generating mechanism of y_t, z_t, and if y_t and z_t are I(1) and cointegrated, with cointegrating relation $e_t = y_t - \alpha z_t$, then there exists the equivalent representation in first differences:

$$[A(L)]\begin{bmatrix} \Delta y_t \\ \Delta z_t \end{bmatrix} = \begin{bmatrix} -\rho_1 \\ -\rho_2 \end{bmatrix} e_{t-1} + \begin{bmatrix} \varepsilon_t^y \\ \varepsilon_t^z \end{bmatrix}, \tag{2.12}$$

where A(L) is in general different from $A^*(L)$. The negative sign for ρ_1 and ρ_2 indicates the error-correction mechanism: a positive deviation $e_{t-1} \neq 0$ from the equilibrium condition $y_{t-1} = \alpha z_{t-1}$ decreases y_t and/or z_t in the current period. Cointegration imposes that at least one of the two coefficients, ρ_1 or ρ_2, be different from zero. If y_t and z_t are not cointegrated, then both ρ_1 and ρ_2 must be zero, as now e_t would be I(1), and would make (2.12) unbalanced: I(0) in the left-hand side and I(1) in the right-hand side.

The Granger representation theorem establishes that the practice of modelling I(1) variables as a VAR in first differences is correct only if the variables are not cointegrated and if they are, then the VAR model in first differences must contain an additional term, the error-correction term, otherwise the model is misspecified. For the case of testing rationality and money neutrality, this point was made clear in Ermini and Chang (1993): by showing that money, interest rate and output are cointegrated and by correctly adding an error-correction term to the VAR model in first differences considered by Bohara (1991), neutrality is no longer rejected.

Cointegration analysis has been made "operational" for multivariate systems of any dimension by Johansen (1988, 1989, 1991) and Johansen and Juselius (1990), who developed a test procedure when cointegration occurs among more than two I(1) variables. To explain Johansen framework briefly, consider the following general VAR model

$$X_t = \Pi_1 X_{t-1} + \ldots + \Pi_k X_{t-k} + \mu + \upsilon_t, \quad t = 1,2,\ldots,T, \tag{2.13}$$

where X_t is a p-dimensional vector of non-stationary processes and υ_t is the vector of their innovations. This model can be rewritten in first differences as:

$$\Delta X_t = \Gamma_1 \Delta X_{t-1} + \ldots + \Gamma_{k-1} \Delta X_{t-k+1} + \Pi X_{t-k} + \mu + \nu_t, \quad t=1,\ldots,T, \tag{2.14}$$

where $\Gamma_i = -(I - \Pi_1 - \ldots - \Pi_i)$, $\quad i = 1,\ldots,k-1$,

and $\Pi = -(I - \Pi_1 - \ldots - \Pi_k)$; more compactly, this can also be rewritten as

$$\Gamma(L)\Delta X_t = \Pi X_{t-k} + \mu + \nu_t, \tag{2.15}$$

where $\Gamma(L)$ is a matrix polynomial of degree of k-1, i.e., $\Gamma(L) = I - \Sigma_{j=1}^{k-1} L_j \Gamma^j$. Johansen shows that the cointegration properties for the VAR model can be detected by focusing on the matrix Π. This matrix is decomposable into the product α and β of two pxr matrices ($r \leq p$), where r is the rank of Π. The r elements of $\beta' X_{t-k}$ are called the error-correction terms, and the columns of β are cointegrating vectors. These cointegrating vectors have the property that $\beta' X_t$ is stationary even though X_t itself is non-stationary. Therefore, (2.14) and (2.15) can be interpreted as an error-correction model; the error-correction terms identify the long-run cointegrating relations among the variables, and help forecast the vector X_t in addition to its autoregressive components. Note also that if the system is not cointegrated, then $\Pi = 0$ (i.e., $r = 0$), if Π has full rank, then X_t is I(0).

2.5 Seasonality

Stationary stochastic processes can be described both in time domain or frequency domain. These two are theoretically

equivalent through a Fourier transformation. The power spectrum of a stationary process is

$$s(\omega) = (2\pi)^{-1}\sum_{\tau=-\infty}^{\infty} \lambda_\tau e^{-i\tau\omega}, \tag{2.16}$$

where $\omega \epsilon [-\pi, \pi]$ is the frequency in radians, and λ_τ is the autocovariance generating function. The power spectrum contains the same information as the autocovariances but presented in a different form.

The previous discussions of I(1) processes and cointegration assume the existence of a unit root only at the zero frequency. The spectrum of such a process has infinite power at $\omega = 0$. As many economic time series, however, contain strong seasonalities, there is a definite possibility that there may be unit roots at the seasonal frequencies as well, that is, that the spectrum of the series has infinite power at $\omega = \pi$ (two cycles per year with quarterly data) and at $\omega = \pi/2$ (one cycle per year with quarterly data). With unit roots at the seasonal frequencies, the standard cointegration technique described in the previous section, and which focuses only on the zero frequency must be replaced by a more comprehensive technique that permits investigation of cointegration also at the seasonal frequencies. This technique, known as seasonal cointegration, was recently developed in Engle, Granger, Hylleberg and Lee (1993,

henceforth EGHL), Lee (1992) and Hylleberg, Engle, Granger and Yoo (1990, henceforth HEGY).

This possibility poses two orders of problems. First, it makes even more compelling the argument that in empirical work it is better to use seasonally unadjusted data, as the information contained at the seasonal frequencies - which would be lost through a seasonal adjustment procedure - may significantly improve both forecasts and inference. Secondly, as pointed out first in Engle, Granger and Hallman (1989), if seasonally unadjusted data are used, but the possible presence of unit roots at seasonal frequencies is not taken into account, cointegration tests and inference about unit roots at the zero frequency may be affected. Ghysels and Perron (1993) argued that unit root tests performed with seasonally adjusted data become less powerful against stationary alternatives as the seasonal adjustment filter induces a bias at the low frequency. It also causes a meaningful loss of valuable information on the seasonal behavior in data when seasonal fluctuations are a significant source of variation in the system.

Typically, empirical work on the Korean economy - as well as of other countries - is carried out with seasonally adjusted data, perhaps, as a consequence of the U.S. practice of using seasonally adjusted data. However, as Korean macroeconomic data is officially available in the unadjusted form, it is interesting to investigate the consequence of

losing potentially useful information contained at seasonal frequencies on testing the joint hypothesis of rationality and neutrality. For this reason, this dissertation will carry out the test using two data sets: (i) the official seasonally unadjusted data set of money, interest rate, price level and output; (ii) the seasonally adjusted data set obtained - in line with the common practice of research on Korea's economy - by applying to the unadjusted set the X-11 filter of the U.S. Bureau of Census. With the latter data set, the joint hypothesis will be tested within Johansen's framework and following the work of Ermini and Chang. With the former data set, the joint hypothesis will be tested within a seasonal cointegration framework, similar to Johansen's, recently developed by Lee (1992).

2.6 Time-Varying Parameters and Lucas Critique

As Lucas (1976) noted, the parameters of Z_{t-1} in equation (2.1) may be not time-invariant to changes in policy regime; correspondingly, the polynomials B(L) and C(L) in equation (2.2) may change according to different policy rules. However, only a few economists have taken up this problem in their empirical tests of the MPIP hypothesis. Cecchetti (1986) again rejected neutrality, even when the effect of unanticipated money on output varied over time. Kim and

Nelson (1989) estimated a time-varying-parameter model by using the Kalman-filter. Based on a two-step joint estimation procedure, they rejected the Lucas hypothesis: the conditional variance of monetary growth affects real output directly, not through the coefficients on the monetary shock. In addition to time-varying parameters, non-stationarities due to heteroscedasticity of residuals or structural breaks should be also taken into account, particularly in consideration of institutional changes.

Chapter III

INTRODUCTION TO THE KOREAN CASE

In this chapter, the monetary policy of Korea will be briefly discussed and some related literature will be reviewed. The data for the empirical test will also be described and analyzed.

3.1 Monetary Policy of Korea

During the last three decades of economic development of Korea, the government has utilized monetary policy more often than fiscal policy, partly because of the quicker response of monetary policy and partly because of the rigidity of the government budget, particularly of defense expenditure. As the priority of economic policies in the 1960s and the 1970s was placed on high growth rate to create new employment opportunities, the primary goal of monetary policy in those years was to finance the economic development. Given the low level of capital accumulation and the lack of well-organized money markets, funds were short and had to be supplied by expanding the money supply and by encouraging foreign capital. As a consequence, chronic inflation persisted alongside rapid economic growth.

As chronic inflation during the 1960s and the 1970s weakened the long-term growth potential of the economy, since 1981 the monetary authorities have adopted a relatively tight monetary policy: the annual growth rate of money (M_2), which had averaged over 30 percent in the 1970s, has decreased well below 20 percent since 1983 (Figure 3.1). Correspondingly, the rate of inflation, as high as 20 - 30% per annum in the 1970s, dropped to below 10% since 1982 (Figure 3.2).

In conducting monetary policy, the monetary aggregate M_2 has been generally chosen as an intermediate target, rather than the interest rate, as the latter was regulated by the monetary authorities and did not reflect market mechanisms.

The Korean domestic financial market has been isolated from the international market under strict foreign exchange control: the foreign exchange system has operated under the fixed exchange rate regime until 1990. Thus, by and large, the Korean economy can be described as a small open economy under fixed exchange rate regime and with a capital market not integrated with international capital market. In summary, the monetary policy of Korea seems not to be greatly disturbed by influences from abroad (Figure 3.3).

3.1.1 Characteristics of Money Supply (M_2)

For the purpose of this research, the M_2 measure of money is chosen, as it has been used as an intermediate target in conducting monetary policy since 1979. As depicted in Figure

Figure 3.1

Growth Rates of Money (M_2) and Real GNP

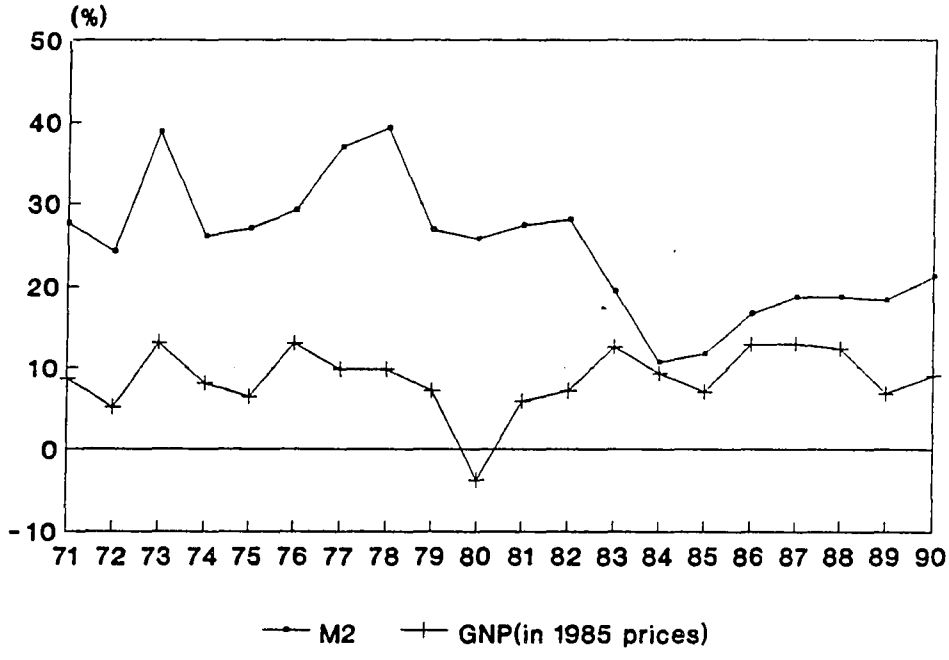


Figure 3.2

GNP Deflator

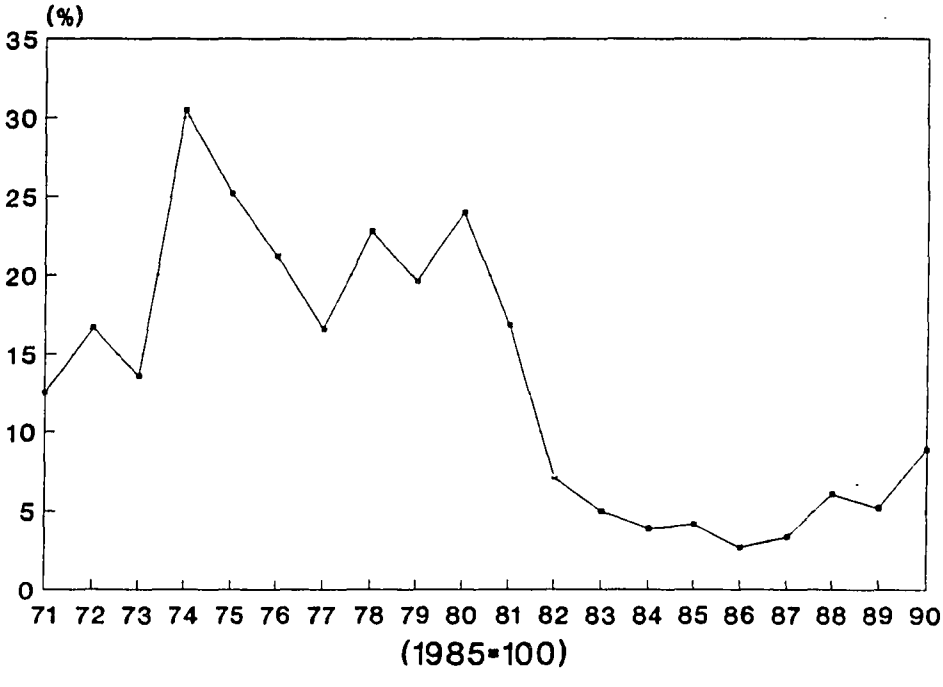


Figure 3.3

Foreign Exchange Rates

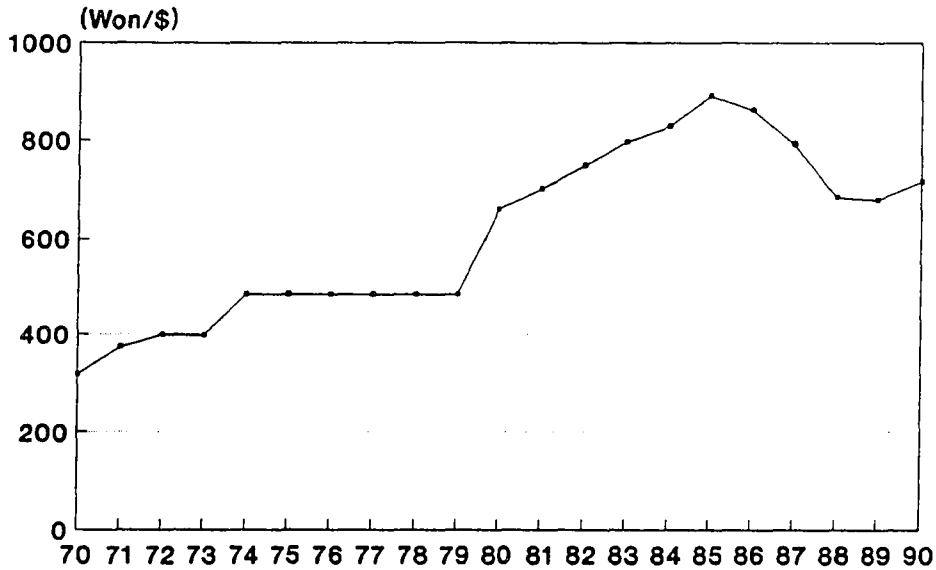
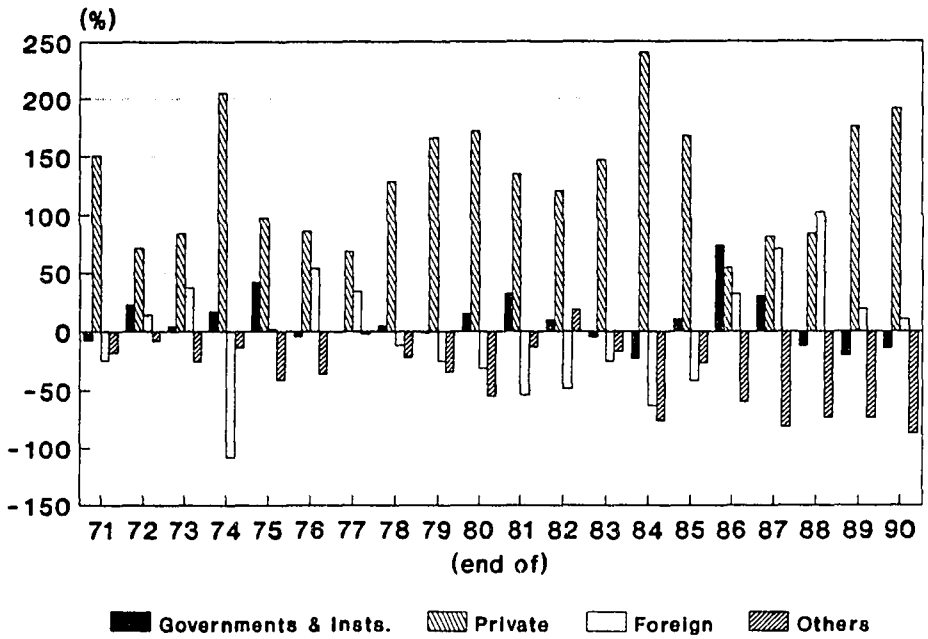


Figure 3.4

Contributions to Money Supply (M_2) by Sector



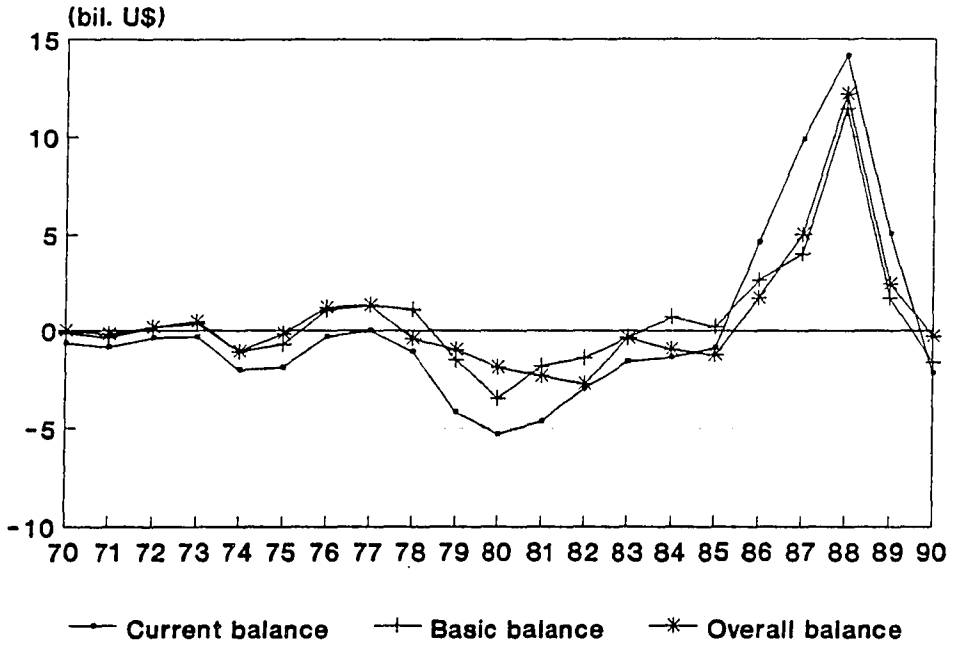
3.1, the growth rate of M_2 exhibits a very volatile pattern, similar to the one of real GNP growth rate. This can be seen as supporting the view that monetary authorities of Korea seemed to follow countercyclical monetary policies to offset business fluctuations rather than a fixed growth rule of money supply, as discussed in section 1.2.

As reported in Figure 3.4, most of M_2 has been supplied through the private sector and, in particular during the last two decades, the supplied amount exceeds the target level of M_2 set by monetary authorities for 12 years. This reflects the expansion of preferential policy loans such as loans for foreign trade, loans for capital equipment destined to the export industry and loans for small and medium size firms. Even though the Korean economy has experienced persistent current account deficit (Figure 3.5), the foreign sector has sometimes played a role of supplier of money rather than demander, although in a limited amount, as foreign borrowings often offset more than the current account deficit and added to the build-up of foreign reserves.

The behavior of money supply can be explained by looking at the determinants of money supply. Among the primary determinants of the monetary base, the actual reserve ratio and the currency ratio were stable during the 1970s but decreased significantly during the 1980s. Correspondingly, the money multiplier (M_2/monetary base) was stable during the 1970s, but increased markedly during the 1980s. This

Figure 3.5

Balance of Payments



significant rise in the money multiplier forced monetary authorities to reduce the growth rate of the monetary base to about 5 percent during the 1980s (except a few last years), from the 28% rate of increase per year on average during the 1970s.

One interesting aspect is that although the average growth rates of M_2 during the 1970s are twice as much higher than those during the 1980s, the average real GNP growth rate is almost the same in both periods (Figure 3.1). This difference is reflected in the price level (high inflation during the 1970s, low inflation during the 1980s), and seems to support the argument that the monetary policy of Korea did not affect real output.

3.1.2 Instruments of Monetary Policy

Under the circumstance of less-developed capital markets and persistent excessive demand for funds because of low loan rates compared with market rates, instruments of monetary policy were mostly confined to direct and selective credit controls: these include ceilings on the aggregate outstanding volume of loans of each commercial banks, guidelines on the allocation of banking funds, and preferential interest rates.

Among orthodox instruments, reserve requirements policy has been used more frequently than discount policy or open market operation. In particular, as opposed to developed countries, open market operations were utilized only

sporadically because of the shortage of marketable instruments. The government depended largely on borrowings from the central bank to finance fiscal deficit rather than on issuing of government securities.

Under this situation, the transmission mechanism of monetary policy in Korea seems to be better explained by the 'availability hypothesis': the expansion of money (M) increases the quantity of available loans (L) and this, in turn, induces the rise in investment spending (I) and thus of output(Y) (for a detailed discussion, see Mishkin 1989, ch. 23). This hypothesis contrasts the traditional Keynesian transmission mechanism based on the effect of money increase on the interest rate, and thus on investment.

3.2 Review of the Literature

No attempt has been made to test the MPIP hypothesis for Korea, although some researchers have tested the Lucas hypothesis (1972a, 1973, 1975) - effects of aggregate nominal shocks on the transitory component of real output are inversely related to the variability of such shocks - in the context of cross country comparisons.

Alberro (1981) extended Lucas (1973) by investigating whether there existed a positive correlation between the slope of the Phillips curve and the variance of nominal income in

Korea. He considered Korea as one of six countries with highly erratic aggregate demand and used annual data of nominal gross national products (GNP) spanning 1953-1969. He found that the Korean data supported the Lucas hypothesis, i.e., that the short-run Phillips curve was vertical. Jung (1985) used the same model of Alberro with an enlarged period of 1953-1980, but he found that there existed a positive Phillips curve for the case of Korea.

Kormendi and Meguire (1984) considered the Lucas hypothesis by using money supply as nominal shocks, like in Barro (1977a, 1978, 1981) and Mishkin (1982a, 1983). They estimated a money supply equation similar to (2.1), but with a time trend instead of policy variables for uniform implementation for each country, and estimated a real output equation similar to (2.2) with a time trend but without anticipated component of money. They found that the Korean data supported the hypothesis of long-run neutrality of money and confirmed the Lucas hypothesis.

Atesoglu and Tillman (1980) tested money-income causality in Granger sense with Korean quarterly data from 1960.I to 1974.IV, and found unidirectional causality from nominal income to broad money (M_2), and bidirectional causality between the narrow money (M_1) and nominal income. They argued that the money supply of Korea could not be treated as an exogenous variable with respect to changes in aggregate economic activity.

3.3 Data Analysis

3.3.1 Data Description

Money (M_t), interest rate (r_t), price (P_t) and output (Y_t) were chosen for the empirical work. The broadly defined money (M_2) is used for the money variable; composite average yield of government and public bonds whose maturity is not exceeding five years is chosen for the interest rate variable; GNP deflator and real GNP (1985=100) are employed for the variables of price level and output, respectively. All series are quarterly, from 1970.I to 1991.IV, seasonally unadjusted. For the test with seasonally adjusted data, the series were filtered by using the X-11 procedure of the SAS (1990) program. Both series are converted into logarithms. The data set was collected from various issues of the Monthly Statistical Bulletin published by the Bank of Korea.

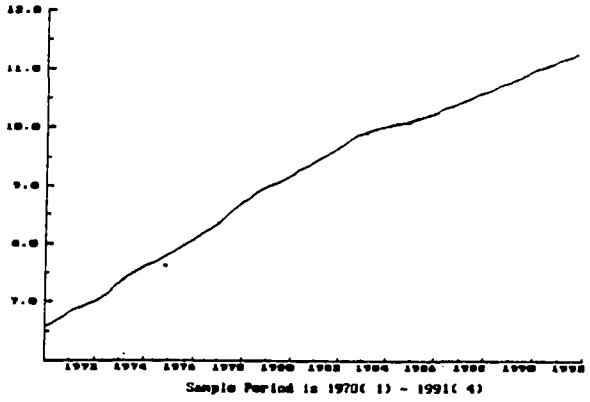
Figures 3.6 and 3.7 report the four seasonally adjusted and unadjusted series in log levels and in first differences, respectively.

3.3.2 Unit Root Tests

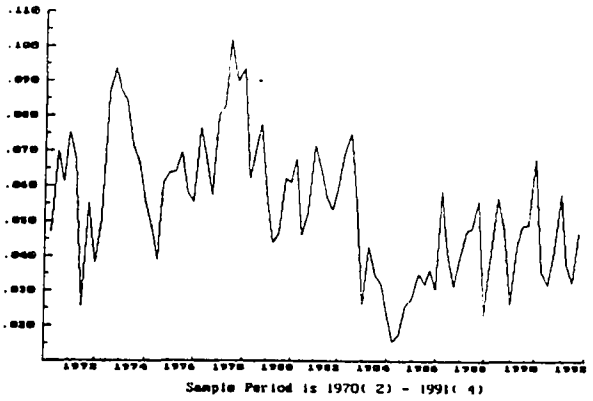
Unit root tests for these variables were performed in two ways following Dickey-Fuller (Dickey and Fuller 1979, 1981; Said and Dickey 1984) and Perron (1989).

Figure 3.6

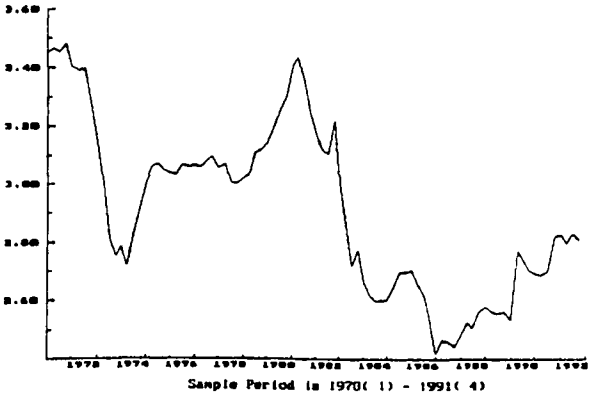
Plot of the Series: Seasonally Adjusted Data



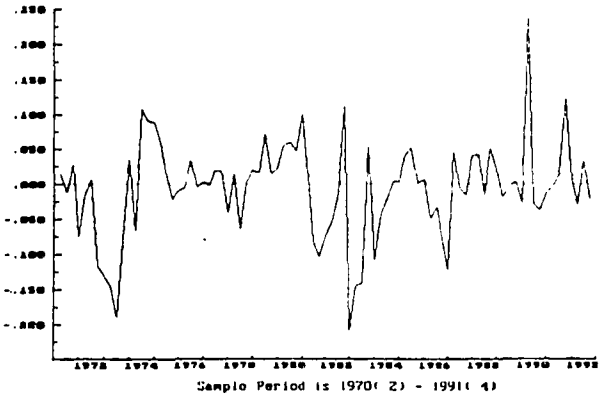
Log of Money



First Difference of Log Money

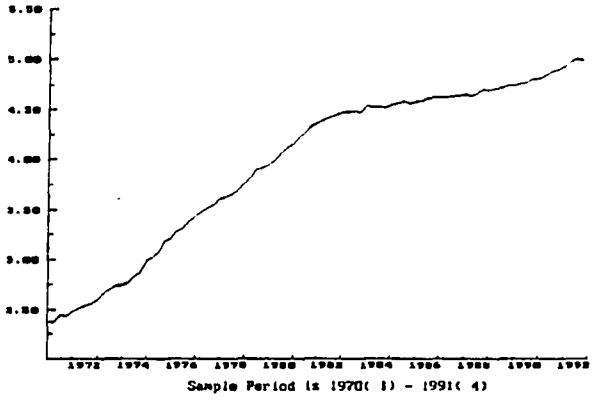


Log of Interest Rate

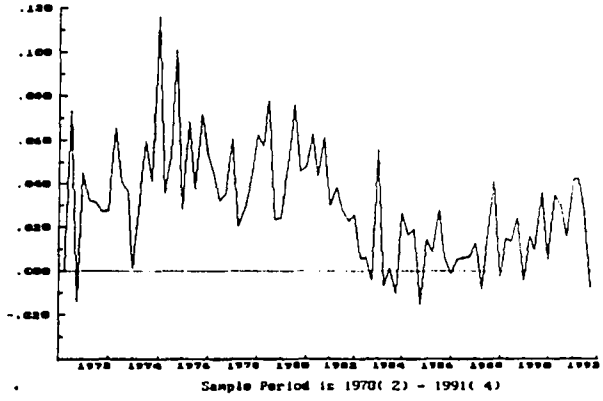


First Difference of Log Interest Rate

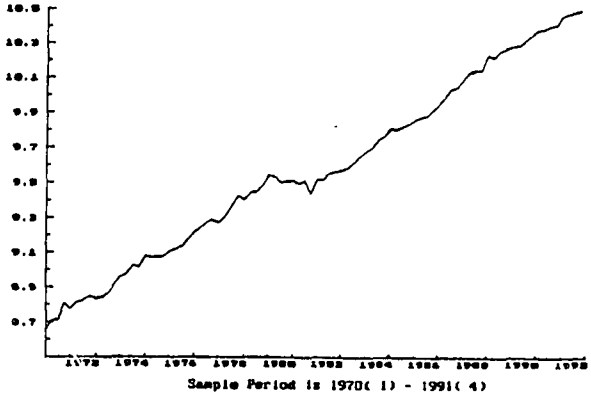
Figure 3.6 (continued)



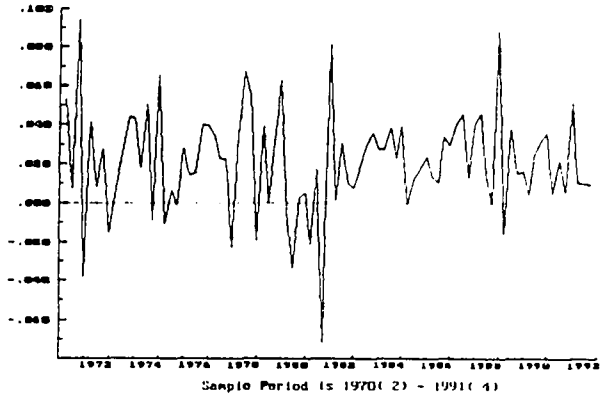
Log of GNP Deflator



First Difference of
Log GNP Deflator



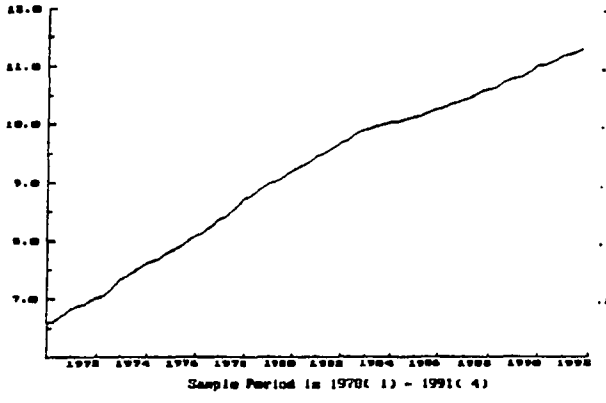
Log of Real GNP



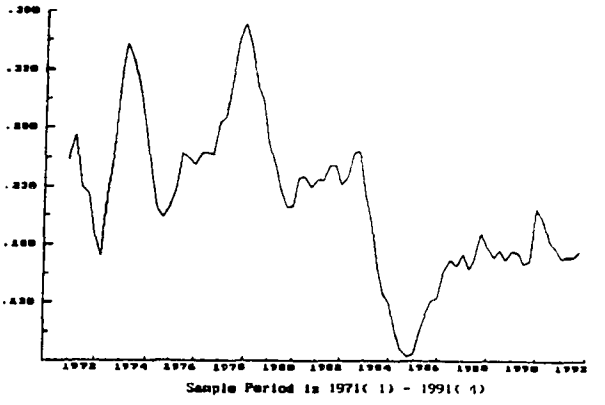
First Difference of
Log Real GNP

Figure 3.7

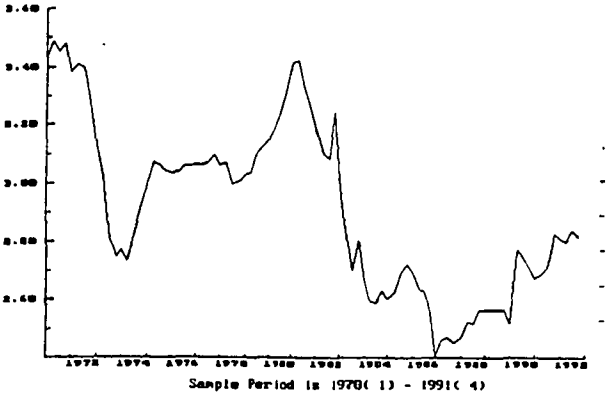
Plot of the Series: Seasonally Unadjusted Data



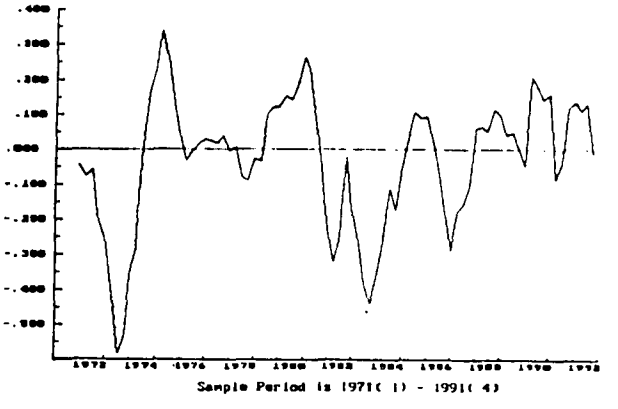
Log of Money



Fourth Difference of Log Money

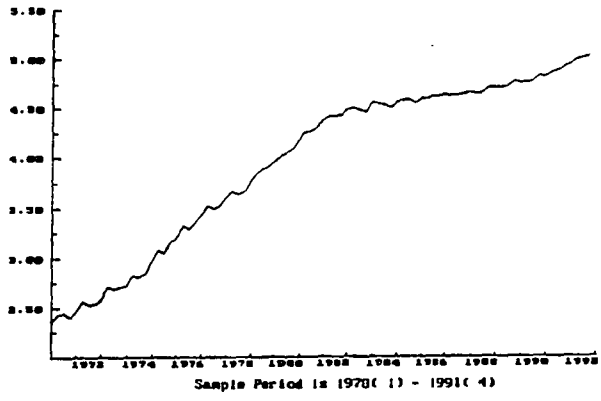


Log of Interest Rate

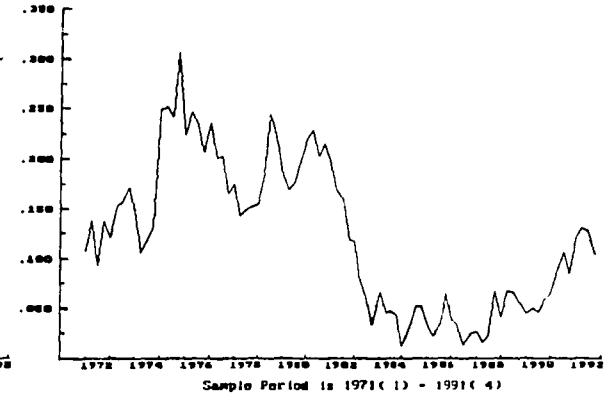


Fourth Difference of Log Interest Rate

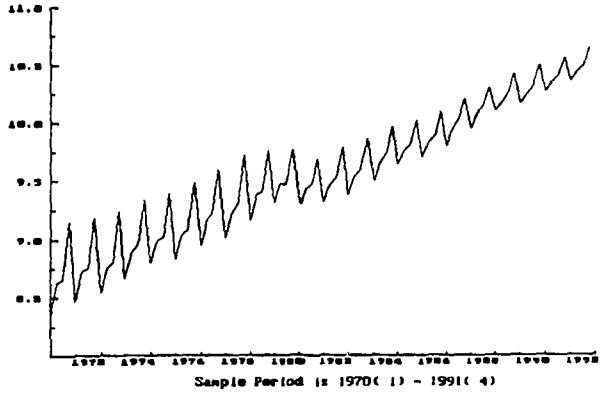
Figure 3.7 (continued)



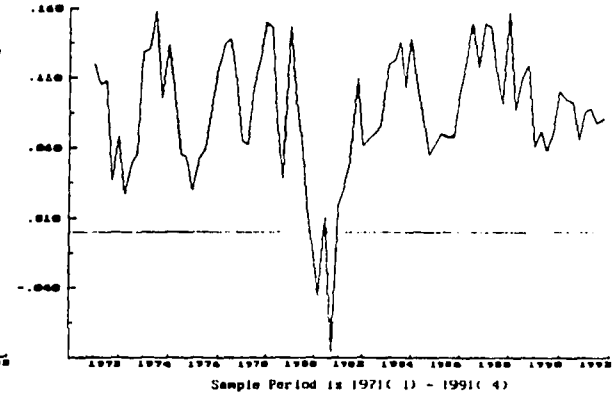
Log of GNP Deflator



Fourth Difference of
Log GNP Deflator



Log of Real GNP



Fourth Difference of
Log Real GNP

3.3.2.1 Dickey-Fuller Tests

The most common test for unit root is the Dickey-Fuller test based on the following augmented regression equation:

$$\Delta X_t = \alpha_0 + \alpha_1 t + \alpha_2 X_{t-1} + \sum_{i=1}^{k} \beta_i \Delta X_{t-i} + \epsilon_t, \qquad (3.1)$$

where ϵ_t is white noise; the autoregressive component is added to whiten the error term, if necessary. The null hypothesis of the test is that $\alpha_2 = 0$, so that X_t is I(1). The Dickey-Fuller test can be performed in two ways: investigate α_2 only (t-test) and α_0, α_1, α_2 simultaneously (F-test). AIC[1] and FPE criteria were used to select the appropriate order of lag (k). In checking the error term for uncorrelatedness, the Lagrange Multiplier test was employed rather than the frequently used Box-Pierce Q-statistic, as the latter is biased towards zero when lagged dependent variables are included in the regression equations (Doornik and Hendry 1992). This test is reported in Table 3.1.

The $t(\alpha_2)$ statistics for all the variables in (3.1) are all less in absolute term than the 5 percent critical value (-3.45) for the sample size of 100, as reported in Table 8.5.2 of Fuller (1976, p373). F-ratios for the null hypotheses H_0[1]

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[1] The criterion of AIC is to minimize
AIC = ln(SSR/T) + 2k/T,
where SSR: sum of squared residuals,
T: sample size,
k: number of regressors (Judge et al. 1985).

Table 3.1
Results of Dickey-Fuller Unit Root Tests

	k	$\chi^2(5)$	F-Form	$t(\alpha_2)$	$F(\phi_3)$	$F(\phi_2)$
Money	6	10.48	1.99	-1.42	4.28	4.21
Interest Rate	1	4.16	.78	-2.15	2.75	2.02
Price	5	2.24	.39	-1.91	3.36	2.69
Output	4	1.81	.32	-1.99	2.04	6.20*

1. * indicates significant at the 5 percent confidence level.

2. The error term is examined by using up to five lags.

3. The critical value at the 5 percent significance level of $F(\phi_3) = 6.49$, $F(\phi_2) = 4.88$, $\chi^2(5) = 11.07$; also $F(5, 67) = 2.35$ for money, $F(5, 77) = 2.33$ for interest rate, $F(5, 69) = 2.35$ for price, $F(5, 71) = 2.35$ for output.

Table 3.2
Results of Perron Unit Root Tests

	k	λ	$\chi^2(5)$	F-Form	$t(\alpha)$
Money	6	.5	7.12	1.23	-3.10
Interest Rate	1	.5	6.04	1.12	-2.45
Price	5	.5	5.44	.94	-3.24
Output	4	.4	9.02	1.66	-3.97

1. The error term is examined by using up to five lags.

2. The critical value at the 5 percent significance level of $\chi^2(5) = 11.07$; also $F(5, 64) = 2.37$ for money, $F(5, 73) = 2.35$ for interest rate, $F(5, 66) = 2.35$ for price, $F(5, 68) = 2.35$ for output.

: $(\alpha_0, \alpha_1, \alpha_2) = (\alpha_0, 0, 0)$ and H_0^2 : $(\alpha_0, \alpha_1, \alpha_2) = (0, 0, 0)$ are all below the critical value, 6.49 for H_0^1 and 4.88 for H_0^2 at the 5 percent significance level but output for H_0^2. The critical value for F-test statistic corresponds to Φ_3 and Φ_2 in Fuller, respectively. These results indicate that all the four variables are I(1).

3.3.2.2 Perron Tests

Perron (1989) has argued that in case of structural breaks in the time series, the Dickey-Fuller test is biased toward the hypothesis of unit root. As each series seems to exhibit a structural break in the early 1980s (see Figure 3.6), Perron's test was also used here. Perron's test is available for several types of structural breaks. In our case, Perron's model C (see Perron 1989, p1380) was used, which allows a one-time change in time series at a time T_B ($1 < T_B < T$). This test is carried out by using the following regression equation:

$$x_t = \mu + \theta DU_t + \beta t + \gamma DT_t + dDTB_t + \alpha X_{t-1} + \sum_{i=1}^{k} \Delta X_{t-i} + e_t, \tag{3.2}$$

where e_t is the error, and

$DU_t = 1$	if $t > T_B$,	0	otherwise;
$DT_t = t$	if $t > T_B$,	0	otherwise;
$DTB_t = 1$	if $t = T_B + 1$,	0	otherwise.

Therefore, DU_t is a dummy variable for a change in intercept; DT_t for a change in growth; DTB_t for a crash.

The null hypothesis of the test is that $\alpha = 1$, so that X_t has a unit root. Thus, when α is significantly different from 1, the null hypothesis of unit root is rejected. The test results are reported in Table 3.2. As every t-statistic of α is less than the critical value of -4.24 in absolute value at the 5 percent of confidence level with a ratio of pre-break sample size to total sample size, $\lambda = .5$ (-4.22 with $\lambda = .4$), the null hypothesis cannot be rejected. The result of Perron's unit root tests confirms that of Dickey-Fuller. Both results, therefore, suggest that each time series of M_t, r_t, P_t and Y_t is suitably described as a difference-stationary process.

3.3.3 Seasonal Unit Root Tests

Tests of unit roots at seasonal frequencies were performed based on HEGY (1990), which adopts the seasonal differencing operator $1-L^4$ suggested by Box and Jenkins (1976) for quarterly data. The test is based on the following regression:

$$D(L)\,y_{4,t} = \Pi_1 y_{1,t-1} + \Pi_2 y_{2,t-1} + \Pi_3 y_{3,t-1} + \Pi_4 y_{3,t} + \varepsilon_t, \qquad (3.3)$$

where

$$y_{1,t} = S_1(L)\,x_t = (1+L+L^2+L^3)\,x_t,$$

$$y_{2,t} = S_2(L)\,x_t = (1-L+L^2-L^3)\,x_t,$$

$$y_{3,t} = S_3(L)x_t = (L-L^3)x_t = L(1-L^2)x_t,$$

$$y_{4,t} = (1-L^4)x_t = \Delta_4 x_t.$$

$S_1(L)$ is a seasonal filter that eliminates unit roots at all seasonal frequencies (i.e., $\omega=\pi$ and $\omega=\pi/2$), $S_2(L)$ removes unit roots at the zero frequency ($\omega=0$) and seasonal frequency $\omega=\pi/2$ and $S_3(L)$ eliminates unit roots at the zero frequency and seasonal frequency $\omega=\pi$. Thus, the filtered series $y_{1,t}$, $y_{2,t}$ and $y_{3,t}$ have only one unit root at frequency 0, π and $\pi/2$, respectively. The null hypothesis of the test of unit root at the zero frequency is $\Pi_1 = 0$, at the seasonal frequency $\omega = \pi$ is $\Pi_2 = 0$, and at the seasonal frequency $\omega=\pi/2$ is the joint test, $\Pi_3 \cap \Pi_4 = 0$. Therefore, if Π_2 and $\Pi_3 \cap \Pi_4$ are different from zero, there will be no seasonal unit roots. The test statistic is the t-value for Π_1 and Π_2, and the F-value for $\Pi_3 \cap \Pi_4$.

The result in Table 3.3 indicates that log money has a unit root at all seasonal frequencies as well as at the zero frequency, while log interest rate has a unit root only at the zero frequency. The log of price and output also have a unit root at the zero frequency. However, the presence of a unit root at seasonal frequencies for both series depends on the inclusion of a drift, and/or a trend and/or seasonal dummies in the regression equation (3.3): the series of price has a seasonal unit root at both $\omega=\pi$ and $\omega=\pi/2$ when seasonal dummies are included and no seasonal unit root at either $\omega=\pi$ or $\omega=\pi/2$,

Table 3.3
Results of Seasonal Unit Root Tests

1. Money

< >	k	$\chi^2(5)$	F-form	$t(\Pi_1)$	$t(\Pi_2)$	$t(\Pi_3)$	$t(\Pi_4)$	$F(\Pi_3 \cap \Pi_4)$
<1>	7	5.58	1.02(5,69)	1.31	-1.49	- .78	-2.20*	2.74
<2>	7	5.94	1.08(5,68)	-2.77	-1.56	- .92	-1.98*	2.39
<3>	15	7.64	1.14(5,49)	-1.85	-1.52	-1.03	-2.84*	4.57
<4>	7	6.39	1.15(5,67)	-1.25	-1.51	- .93	-2.01*	2.47
<5>	15	8.10	1.20(5,48)	-1.22	-1.47	-1.06	-2.87*	4.70

2. Interest Rate

< >	k	$\chi^2(5)$	F-form	$t(\Pi_1)$	$t(\Pi_2)$	$t(\Pi_3)$	$t(\Pi_4)$	$F(\Pi_3 \cap \Pi_4)$
<1>	4	6.68	1.30(5,75)	- .91	-5.51*	-4.46*	-6.14*	43.26*
<2>	4	6.50	1.24(5,74)	-2.44	-5.61*	-4.81*	-5.82*	43.93*
<3>	4	2.88	.50(5,71)	-2.41	-6.26*	-4.32*	-5.36*	33.83*
<4>	4	7.65	1.46(5,73)	-2.51	-5.67*	-4.94*	-5.71*	44.50*
<5>	0	3.42	.59(5,70)	-2.64	-6.36*	-4.49*	-5.23*	34.51*

1. Critical values for the tests are taken from HEGY (1990).

2. * indicates significant at the 5 percent confidence level.

3. The critical value of $\chi^2(5)$ = 11.07 at the 5 percent confidence.

4. Number in < > indicates whether intercept(I), seasonal dummies(SD), and/or trend(T) are included or not in the regression equations. <1> is for no I, SD, T; <2> for I, no SD, T; <3> for I, SD, no T; <4> for I, T, no SD; <5> for I, SD, T.

Table 3.3 (continued)

3. Price

< >	k	$\chi^2(5)$	F-form	$t(\Pi_1)$	$t(\Pi_2)$	$t(\Pi_3)$	$t(\Pi_4)$	$F(\Pi_3 \cap \Pi_4)$
<1>	7	5.30	.97(5,69)	.79	-2.01*	-2.31*	-1.21	3.48*
<2>	7	5.27	.95(5,68)	-2.53	-2.10*	-2.24*	-1.17	3.24*
<3>	7	4.61	.78(5,65)	-2.39	-2.05	-2.21	-1.41	3.50
<4>	7	5.94	1.06(5,67)	-1.89	-2.04*	-2.20*	-1.16	3.15*
<5>	7	5.24	.89(5,64)	-1.89	-1.99	-2.21	-1.45	3.54

4. Output

< >	k	$\chi^2(5)$	F-form	$t(\Pi_1)$	$t(\Pi_2)$	$t(\Pi_3)$	$t(\Pi_4)$	$F(\Pi_3 \cap \Pi_4)$
<1>	5	8.84	1.74(5,73)	3.30	-3.32*	-2.20*	-1.19	3.11
<2>	5	8.88	1.72(5,72)	.22	-3.30*	-2.18*	-1.18	3.07
<3>	5	7.49	1.37(5,69)	.16	-1.52	-1.76	- .68	1.75
<4>	5	8.41	1.60(5,71)	-2.26	-3.37*	-2.19*	-1.17	3.09*
<5>	5	6.89	1.23(5,68)	-2.25	-1.53	-1.83	- .56	1.82

1. Critical values for the tests are taken from HEGY (1990).

2. * indicates significant at the 5 percent confidence level.

3. The critical value of $\chi^2(5)$ = 11.07 at the 5 percent confidence.

4. Number in < > indicates whether intercept(I), seasonal dummies(SD), and/or trend(T) are included or not in the regression equations. <1> is for no I, SD, T; <2> for I, no SD, T; <3> for I, SD, no T; <4> for I, T, no SD; <5> for I, SD, T.

otherwise; the series of output has a seasonal unit root at $\omega=\pi$ only when seasonal dummies are included and has a seasonal unit root at $\omega=\pi/2$ in all cases except that both an intercept and a trend were included in the regression equation.

Chapter IV

THE KOREAN CASE I: SEASONALLY ADJUSTED DATA

The test of the joint hypothesis of rationality and neutrality will be discussed in this chapter with seasonally adjusted data. For consistency with common practice among Korean researchers, the data is adjusted with the X-11 procedure of the U.S. Bureau of Census. As anticipated, the test procedure will follow Ermini and Chang (1993), which allows for the explicit recognition of error-correction terms.

4.1 The Model

As in Barro's and Mishkin's original work, the test procedure is based on a forecasting equation for money and on a generating mechanism for output. For the former, instead of the restricted form (2.1), the following more general parametrization is used:

$$A(L)M_t = B(L)Z_{t-1} + u_t. \tag{4.1}$$

Here M_t is the log of the monetary aggregate, Z_t is a vector of the log of macro-variables that have predictive power for M_t, and u_t is the one-step ahead forecasting error. A(L) is a polynomial in the lag operator L, and B(L) is a matrix of

polynomials in L. Both A(L) and B(L) may have unit roots. In this analysis, Z_t contains interest rate (r_t), price (P_t) and output (Y_t).

As we cannot exclude a priori multidirectional causality among these variables, the equation for money is considered to be part of a larger VAR system:

$$C(L)X_t = \mu + \upsilon_t, \tag{4.2}$$

where X_t is the vector [M_t, r_t, P_t, Y_t] and C(L) is a matrix of polynomials. As these time series are I(1) (see section 3.3.2), to allow for possible long-run cointegrating relation, the VAR model (4.2) will be estimated following Johansen's procedure described in section 2.4. That is, (4.2) is rewritten as

$$D(L)\Delta X_t = \Pi X_{t-k} + \mu + \nu_t, \tag{4.3}$$

where D(L) is a matrix polynomial of degree of k-1, with D_0 equal to the identity matrix. Note that the first row of (4.3) is the forecasting equation for money. When the 4x4 matrix Π has full rank, the vector process X_t is stationary; when r = 0, Π is the null matrix and (4.3) becomes a differenced vector time series model without the term ΠX_{t-k}, implying that the I(1) vector X_t is not cointegrated; when $0 < \text{rank}(\Pi) = r < 4$, there are pxr matrices α and β such that $\Pi = \alpha\beta'$, and thus there exist r cointegrating relations.

The anticipated component of money is then used to build the generating mechanism of output. Generalizing equation (2.2) of Barro's and Mishkin's, consider the following output equation:

$$E(L)\Delta Y_t = F(L)(\Delta M_t - P_{t-1}[\Delta M_t]) + G(L)P_{t-1}[\Delta M_t] + \nu_{4,t}, \quad (4.4)$$

where $P_{t-1}[.]$ represents the one-step ahead prediction and $\nu_{4,t}$ is the output innovation. The first term on the right-hand side identifies the effect on output of unanticipated money, and the second term the effect of anticipated money. Neutrality imposes the restriction $G(L) \equiv 0$, and rationality implies that the forecast of money is identical to its mathematical expectation, that is $\Delta M_t - P_{t-1}[\Delta M_t] = \nu_{1,t}$, where $\nu_{1,t}$ is the innovation term of the first equation of (4.3) (under the hypothesis that (4.3) is correctly specified). The degree of the polynomials E(L), F(L) and G(L) will be determined in the empirical analysis.

Following previous work in this area, the prediction of money, $P_{t-1}[\Delta M_t]$, is obtained from the first equation of (4.3), that is

$$P_{t-1}[\Delta M_t] = (1-D_{11}^*(L))\Delta M_t - D_{12}^*(L)\Delta r_t - D_{13}^*(L)\Delta P_t \quad (4.5)$$
$$-D_{14}^*(L)\Delta Y_t + \alpha_1^* W_{t-k},$$

where $D_{11}^*(L)$, $D_{12}^*(L)$, D_{13}^* and $D_{14}^*(L)$ are components of a matrix $D^*(L)$ in general different from D(L) of (4.3), and W_{t-k} is the

error-correction term, that is $\beta' X_{t-k}$. Combining (4.5) with (4.4) and replacing in (4.3), the following four-variable "unrestricted"[1] VAR model is obtained:

$$\begin{bmatrix} D_M(L) & D_r(L) & D_P(L) & D_Y(L) \\ N_1^*(L) & N_2^*(L) & N_3^*(L) & N_4^*(L) \end{bmatrix} \begin{bmatrix} \Delta M_t \\ \Delta r_t \\ \Delta P_t \\ \Delta Y_t \end{bmatrix} = [\mu] + \begin{bmatrix} \alpha \\ \alpha^*(L) \end{bmatrix} W_{t-k} + [\nu_t], \tag{4.6}$$

where

$$N_1^*(L) = (1 - D_{11}^*(L))\ [F(L) - G(L)] - F(L), \tag{4.7}$$

$$N_2^*(L) = -\ D_{12}^*(L)\ [F(L) - G(L)],$$

$$N_3^*(L) = -\ D_{13}^*(L)\ [F(L) - G(L)],$$

$$N_4^*(L) = E(L) - D_{14}^*(L)\ [F(L) - G(L)],$$

$$\alpha^*(L) = \alpha_1^*\ [G(L) - F(L)].$$

Note that D_{ij} (i=1,2,3; j=1,2,3,4) are polynomials of degree k-1; the order of the polynomials $N_1^*(L)$, $N_2^*(L)$, $N_3^*(L)$ and $N_4^*(L)$ and the value of k will be determined in the empirical work.

4.2. Empirical Results

The first step is to estimate (4.3) and to perform a cointegration analysis of the system is based on the rank of the matrix Π. As reported in Table 4.1, the hypothesis that

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[1] The model is "unrestricted" with respect to the hypothesis of rationality and neutrality.

Table 4.1
Results of Cointegration Test

	Eigenvalues(λi)			
	(.002693	.109562	.183597	.345656)
	Standardized ß' Eigenvectors			
Money	1.00000	1.49766	-.88245	-1.20229
Interest Rate	10.46769	1.00000	-8.53099	-13.48115
Price	-1.01437	.12418	1.00000	1.33847
Output	-6.09493	-2.32016	8.85826	1.00000
	Standardized α Coefficients			
Money	-.00528	-.00825	.01109	.00016
Interest Rate	-.10734	.00590	.06284	-.00236
Price	.01436	-.00573	-.05066	-.00059
Output	-.03890	.00257	-.03987	.00056

Test Statistics

H_1	λ_{max}	$\lambda_{max}(.05)$	Trace	Trace(.05)
$r \leq 3$	.226530	8.083	.226530	8.083
$r \leq 2$	9.747510	14.595	9.974040	17.844
$r \leq 1$	17.039209	21.279	27.013249	31.256
$r = 0$	35.626286	27.341	62.639535	48.419

rank of Π is $r \leq 3$, or $r \leq 2$, or $r \leq 1$ could not be rejected, while $r = 0$ was rejected at the 5 percent significance level. This implies that there exists only one significant cointegrating relation. The estimated error-correction term is

$$W_{t-4} = \beta' X_{t-4} = M_{t-4} + 1.50 r_{t-4} - .88 P_{t-4} - 1.20 Y_{t-4}. \qquad (4.8)$$

Note that error correction term has strong predictive power for each variable, except for money as reported in Table 4.2.

On the basis of t- and F-statistics, output was found to have no predictive power for money, except for its presence in the error correction term. To parsimonialize the four-variable system, output was eliminated from the money equation, i.e., the restriction $D_{14}(L) = 0$ in (4.6) and the corresponding term in the output equation, $D_{14}^{*}(L) = 0$, was imposed. Regarding the interest rate, the first lag is significant, and thus it was left in the money equation.

Following Ermini and Chang (1993) and Bohara (1991), the order of dynamic component of output, E(L), was set at two, and the order of F(L) and G(L) was set at one. The system was then reestimated in the more parsimonious form:

$$\begin{bmatrix} D_{11}(L) & D_{12}(L) & D_{13}(L) & 0 \\ D_{21}(L) & D_{22}(L) & D_{23}(L) & D_{24}(L) \\ D_{31}(L) & D_{32}(L) & D_{33}(L) & D_{34}(L) \\ N_1^*(L) & N_2^*(L) & N_3^*(L) & E(L) \end{bmatrix} \begin{bmatrix} \Delta M_t \\ \Delta r_t \\ \Delta P_t \\ \Delta Y_t \end{bmatrix} = \qquad (4.9)$$

$$\begin{bmatrix}\mu_1\\ \mu_2\\ \mu_3\\ \mu_4\end{bmatrix} + \begin{bmatrix}\alpha_1\\ \alpha_2\\ \alpha_3\\ \alpha^*(L)\end{bmatrix} W_{t-4} + \begin{bmatrix}\nu_{1,t}\\ \nu_{2,t}\\ \nu_{3,t}\\ \nu_{4,t}\end{bmatrix},$$

where

$$D_{11}(L) = \Sigma_{j=0}^{3} d_{11}^{j} L^j, \quad d_{11}^{0} = 1, \tag{4.10}$$

$$D_{12}(L) = d_{12}^{1} L, \quad d_{12}^{0} = 0,$$

$$D_{13}(L) = \Sigma_{j=0}^{3} d_{13}^{j} L^j, \quad d_{13}^{0} = 0,$$

$$D_{21}(L) = \Sigma_{j=0}^{3} d_{21}^{j} L^j, \quad d_{21}^{0} = 0,$$

$$D_{22}(L) = \Sigma_{j=0}^{3} d_{22}^{j} L^j, \quad d_{22}^{0} = 1,$$

$$D_{23}(L) = \Sigma_{j=0}^{3} d_{23}^{j} L^j, \quad d_{23}^{0} = 0,$$

$$D_{24}(L) = \Sigma_{j=0}^{3} d_{24}^{j} L^j, \quad d_{24}^{0} = 0,$$

$$D_{31}(L) = \Sigma_{j=0}^{3} d_{31}^{j} L^j, \quad d_{31}^{0} = 0,$$

$$D_{32}(L) = \Sigma_{j=0}^{3} d_{32}^{j} L^j, \quad d_{32}^{0} = 0,$$

$$D_{33}(L) = \Sigma_{j=0}^{3} d_{33}^{j} L^j, \quad d_{33}^{0} = 1,$$

$$D_{34}(L) = \Sigma_{j=0}^{3} d_{34}^{j} L^j, \quad d_{34}^{0} = 0,$$

$$\begin{aligned}
N_1^*(L) &= [1 - D_{11}^*(L)][F(L) - G(L)] - F(L) \\
&= - \Sigma_{j=1}^{3} d_{11}^{j*} L^j [\Sigma_{j=0}^{1} (f_j - g_j)] L^j - \Sigma_{j=0}^{1} f_j L^j, \\
&= - f_0 - [d_{11}^{1*}(f_0 - g_0) + f_1] L - [d_{11}^{2*}(f_0 - g_0) + d_{11}^{1*}(f_1 - g_1)] L^2 \\
&\quad - [d_{11}^{3*}(f_0 - g_0) + d_{11}^{2*}(f_1 - g_1)] L^3 - [d_{11}^{3*}(f_1 - g_1)] L^4,
\end{aligned}$$

$$\begin{aligned}
N_2^*(L) &= - D_{12}^*(L)[F(L) - G(L)] \\
&= - d_{12}^{1*} L [\Sigma_{j=0}^{1} (f_j - g_j)] L^j \\
&= - [d_{12}^{1*}(f_0 - g_0)] L - [d_{12}^{1*}(f_1 - g_1)] L^2,
\end{aligned}$$

$$\begin{aligned}
N_3^*(L) &= - D_{13}^*(L)[F(L) - G(L)] \\
&= - \Sigma_{j=1}^{3} d_{13}^{j*} L^j [\Sigma_{j=0}^{1} (f_j - g_j)] L^j
\end{aligned}$$

$$= - [d_{13}^{1*}(f_0-g_0)]L - [d_{13}^{2*}(f_0-g_0)+d_{13}^{3*}(f_1-g_1)]L^2$$

$$- [d_{13}^{3*}(f_0-g_0)+d_{13}^{2*}(f_1-g_1)]L^3 - [d_{13}^{3*}(f_1-g_1)]L^4,$$

$$E(L) = \Sigma_{j=0}^{2} e_j L^j = e_1 L + e_2 L^2, \quad e_0 = 1,$$

$$\alpha^*(L) = \alpha_1^*[G(L) - F(L)] = \alpha_1^* \Sigma_{j=0}^{1} (g_j - f_j) L^j$$

$$= \alpha_1^*(g_0 - f_0) + \alpha_1^*(g_1 - f_1)L.$$

Note that D_{ij}^k (i = 1, 2, 3; j = 1, 2, 3, 4; k = 1, 2, 3) is the (i,j)-th element of the 4x4 matrix of coefficients at lag k.

The chosen parsimonialization is conservative in that it imposes zero restrictions only on the money equation, (and indirectly on the output equation) but not on the equations of interest rate and price. The rationale for this procedure is that the purpose of this study is not to develop an efficient model to represent the system of four variables, but rather to test the non-linear restrictions between money and output imposed by the joint hypothesis of rationality and neutrality. In this sense, the focus here is more on the overall Wald statistic for the restrictions imposed by rationality and neutrality, than on the significance of each individual coefficient of the model.

The "unrestricted" reduced form corresponding to (4.9) is:

$$\begin{bmatrix} \Delta M_t \\ \Delta r_t \\ \Delta P_t \\ \Delta Y_t \end{bmatrix} = \begin{bmatrix} -d_{11}^1 & -d_{12}^1 & -d_{13}^1 & 0 \\ -d_{21}^1 & -d_{22}^1 & -d_{23}^1 & -d_{24}^1 \\ -d_{31}^1 & -d_{32}^1 & -d_{33}^1 & -d_{34}^1 \\ n_1^{0*} d_{11}^1 + n_1^{1*} & n_1^{0*} d_{12}^1 + n_2^{1*} & n_1^{0*} d_{13}^1 + n_3^{1*} & e_1 \end{bmatrix} \begin{bmatrix} \Delta M_{t-1} \\ \Delta r_{t-1} \\ \Delta P_{t-1} \\ \Delta Y_{t-1} \end{bmatrix} \qquad (4.11)$$

$$+\begin{bmatrix} -d_{11}^{2} & 0 & -d_{13}^{2} & 0 \\ -d_{21}^{2} & -d_{22}^{2} & -d_{23}^{2} & -d_{24}^{2} \\ -d_{31}^{2} & -d_{32}^{2} & -d_{33}^{2} & -d_{34}^{2} \\ n_1^{0*}d_{11}^{2}+n_1^{2*} & n_2^{2*} & n_1^{0*}d_{13}^{2}+n_3^{2*} & e_2 \end{bmatrix}\begin{bmatrix} \Delta M_{t-2} \\ \Delta r_{t-2} \\ \Delta P_{t-2} \\ \Delta Y_{t-2} \end{bmatrix}$$

$$+\begin{bmatrix} -d_{11}^{3} & 0 & -d_{13}^{3} & 0 \\ -d_{21}^{3} & -d_{22}^{3} & -d_{23}^{3} & -d_{24}^{3} \\ -d_{31}^{3} & -d_{32}^{3} & -d_{33}^{3} & -d_{34}^{3} \\ n_1^{0*}d_{11}^{3}+n_1^{3*} & 0 & n_1^{0*}d_{13}^{3}+n_3^{3*} & 0 \end{bmatrix}\begin{bmatrix} \Delta M_{t-3} \\ \Delta r_{t-3} \\ \Delta P_{t-3} \\ \Delta Y_{t-3} \end{bmatrix}$$

$$+\begin{bmatrix} 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ n_1^{4*} & 0 & n_3^{4*} & 0 \end{bmatrix}\begin{bmatrix} \Delta M_{t-4} \\ \Delta r_{t-4} \\ \Delta P_{t-4} \\ \Delta Y_{t-4} \end{bmatrix}+\begin{bmatrix} \alpha_1 \\ \alpha_2 \\ \alpha_3 \\ -n_1^{0*}\alpha_1+\alpha_1^{0*} \end{bmatrix}W_{t-4}+\begin{bmatrix} 0 \\ 0 \\ 0 \\ \alpha_1^{1*} \end{bmatrix}W_{t-5}.$$

To detect parameter constancy, the Chow (1960) tests over the forecast period of 16 quarters were performed for each equation. This test statistic was an F distribution with (n, T-k) degree of freedom under the null hypothesis of no structural change, where n is the length of out-of-sample period, T sample size and k the number of explanatory variables. As reported in Table 4.2, the null hypothesis for each variable was not rejected. F-tests for residual heteroscedasticity were also conducted. This test statistic follows an F distribution with (2k-2, T-3k+1) degree of freedom under the null hypothesis of homoscedasticity. The null hypothesis for each variable was not rejected either.

Table 4.2

Results of Estimation of Unrestricted Model

	ΔM_{t-1}	Δr_{t-1}	ΔP_{t-1}	ΔY_{t-1}	ΔM_{t-2}	Δr_{t-2}
ΔM_t	.3899 (3.50)	-.0367 (-1.47)	.0422 (.54)	0	.0750 (.63)	0
Δr_t	.0189 (.03)	.1400 (1.19)	.2449 (.60)	.1536 (.42)	-.0436 (-.07)	.0257 (.21)
ΔP_t	.0635 (.38)	.0656 (1.82)	-.0442 (-.36)	-.0687 (-.62)	.3505 (1.91)	.1001 (2.69)
ΔY_t	.6884 (3.70)	-.0326 (-.81)	-.1056 (-.79)	-.4162 (-3.34)	-.2101 (-1.01)	-.0586 (-1.41)

	ΔP_{t-2}	ΔY_{t-2}	ΔM_{t-3}	Δr_{t-3}	ΔP_{t-3}	ΔY_{t-3}
ΔM_t	.0455 (.61)	0	.0419 (.39)	0	.1276 (1.66)	0
Δr_t	.2573 (.67)	.2229 (.57)	.2453 (.44)	-.0548 (-.43)	.2690 (.72)	.4272 (1.20)
ΔP_t	.1458 (1.25)	.0657 (.55)	.0140 (.08)	.0471 (1.21)	.0455 (.40)	.1195 (1.10)
ΔY_t	.0225 (.17)	.0098 (.08)	-.2779 (-1.37)	0	.0501 (.40)	0

Numbers in parentheses are t-statistics.

Table 4.2 (continued)

	ΔM_{t-4}	Δr_{t-4}	ΔP_{t-4}	ΔY_{t-4}	W_{t-4}	W_{t-5}
ΔM_t	0	0	0	0	.0053 (1.03)	0
Δr_t	0	0	0	0	-.0688 (-2.38)	0
ΔP_t	0	0	0	0	.0262 (2.96)	0
ΔY_t	.1569 (.87)	0	-.1586 (-1.27)	0	-.0614 (-2.15)	.0411 (1.56)

	Adj-R^2	Chow	F-test
ΔM_t	.4606	1.24(16,59)[1]	1.31(16,58)
Δr_t	.1244	1.24(16,54)	.86(26,43)[2]
ΔP_t	.3543	.62(16,54)	.84(26,43)
Y_t	.1943	.54(16,51)	.61(28,39)

Numbers in parentheses are t-statistics.

[1] The critical value of F(15,60) = 1.84 and F(15,50) = 1.87 at the 5 percent confidence level.

[2] The critical value of F(30,40) = 1.75 at the 5 percent confidence level.

Rationality implies $d_{1j}^{k*} = d_{1j}^{k}$ (j,k = 1,2,3) and neutrality implies G(L) ≡ 0. By imposing this joint restriction to (4.11), one obtains the reduced "restricted" VAR:

$$\begin{bmatrix}\Delta M_t\\ \Delta r_t\\ \Delta P_t\\ \Delta Y_t\end{bmatrix} = \begin{bmatrix}-d_{11}^1 & -d_{12}^1 & -d_{13}^1 & 0\\ -d_{21}^1 & -d_{22}^1 & -d_{23}^1 & -d_{24}^1\\ -d_{31}^1 & -d_{32}^1 & -d_{33}^1 & -d_{34}^1\\ f_1 & 0 & 0 & e_1\end{bmatrix}\begin{bmatrix}\Delta M_{t-1}\\ \Delta r_{t-1}\\ \Delta P_{t-1}\\ \Delta Y_{t-1}\end{bmatrix} \tag{4.12}$$

$$+\begin{bmatrix}-d_{11}^2 & 0 & -d_{13}^2 & 0\\ -d_{21}^2 & -d_{22}^2 & -d_{23}^2 & -d_{24}^2\\ -d_{31}^2 & -d_{32}^2 & -d_{33}^2 & -d_{34}^2\\ f_1 d_{11}^1 & f_1 d_{12}^1 & f_1 d_{13}^1 & e_2\end{bmatrix}\begin{bmatrix}\Delta M_{t-2}\\ \Delta r_{t-2}\\ \Delta P_{t-2}\\ \Delta Y_{t-2}\end{bmatrix}$$

$$+\begin{bmatrix}-d_{11}^3 & 0 & -d_{13}^3 & 0\\ -d_{21}^3 & -d_{22}^3 & -d_{23}^3 & -d_{24}^3\\ -d_{31}^3 & -d_{32}^3 & -d_{33}^3 & -d_{34}^3\\ f_1 d_{11}^2 & 0 & f_1 d_{13}^2 & 0\end{bmatrix}\begin{bmatrix}\Delta M_{t-3}\\ \Delta r_{t-3}\\ \Delta P_{t-3}\\ \Delta Y_{t-3}\end{bmatrix}$$

$$+\begin{bmatrix}0 & 0 & 0 & 0\\ 0 & 0 & 0 & 0\\ 0 & 0 & 0 & 0\\ f_1 d_{11}^3 & 0 & f_1 d_{13}^3 & 0\end{bmatrix}\begin{bmatrix}\Delta M_{t-4}\\ \Delta r_{t-4}\\ \Delta P_{t-4}\\ \Delta Y_{t-4}\end{bmatrix} + \begin{bmatrix}\alpha_1\\ \alpha_2\\ \alpha_3\\ 0\end{bmatrix} W_{t-4} + \begin{bmatrix}0\\ 0\\ 0\\ -f_1\alpha_1\end{bmatrix} W_{t-5}.$$

This restricted model imposes the following set of eight non-linear cross-equation restrictions:

$$\phi_{11}{}^1\phi_{41}{}^1 + \phi_{41}{}^2 = 0, \tag{4.13}$$

$$\phi_{12}{}^1\phi_{41}{}^1 + \phi_{42}{}^2 = 0,$$

$$\phi_{13}{}^1\phi_{41}{}^1 + \phi_{43}{}^2 = 0,$$

$$\phi_{11}{}^2\phi_{41}{}^1 + \phi_{41}{}^3 = 0,$$

$$\phi_{13}^{2}\phi_{41}^{1} + \phi_{43}^{3} = 0,$$

$$\phi_{11}^{3}\phi_{41}^{1} + \phi_{41}^{4} = 0,$$

$$\phi_{13}^{3}\phi_{41}^{1} + \phi_{43}^{4} = 0,$$

$$\psi_{1}^{0}\phi_{41}^{1} + \psi_{4}^{1} = 0,$$

where ϕ_{ij}^{k} is the (i,j)-th element of the 4x4 matrix of coefficients at lag k of the "unrestricted" model of (4.11), and ψ_{j}^{s} is the j-th element of the vector of coefficients associated with the error-correction terms W_{t-4} (s = 0) and W_{t-5} (s=1).

These restrictions were tested through the Wald test (Granger and Newbold 1986, pp261-62), obtaining a test statistic W = 13.20, with a p-value of approximately .10. Thus, the joint restriction of rationality and neutrality cannot be rejected at confidence levels less than 10%. This result seems to corroborate the conjecture that anticipated monetary policy of Korea has no effect on the real economic activity, as changes in monetary policy are incorporated into the economic agents' expectations. The policy implication of this ineffectiveness result is that a fixed money supply rule such as Friedman's k percent rule may be preferable to the countercyclical money supply rule for fine-tuning the economy, as the former minimizes the variance of real output and reduces the equilibrium rates of inflation. As the joint hypothesis cannot be rejected, other possibly relevant macro

variables, such as exports, were not included into the system, as their presence can only strengthen the result.

Note, however, that, as the Wald test statistic has a p-value of about 10%, this ineffectiveness result would be reversed for confidence levels higher than the usual 5-10%. In this case, the Korean data would support the new-Keynesian view, rather than the new-classical view.

Chapter V

THE KOREAN CASE II: SEASONALLY UNADJUSTED DATA

This chapter will discuss the test of the joint hypothesis of rationality and neutrality using seasonally unadjusted data. The analysis will be carried out within the general framework of seasonal cointegration as anticipated in section 2.5.

5.1 The Model

A seasonal time series can be described as one with a spectrum having a tall and narrow peaks at the seasonal frequencies $\omega_s \equiv 2\pi j/s$, $j = 1,\ldots, s/2$, where s is the number of observations in a year. If the power at the seasonal frequencies is infinite, then the series exhibits unit roots at these frequencies, and it is called an integrated seasonal process. For example, the seasonal differencing operator suggested by Box and Jenkins (1976) for quarterly series, $(1 - L^4)$, has four unit roots: ± 1 and $\pm i1$. This corresponds to a unit root at the zero frequency, one at frequency π (i.e., $2\pi/4 \times 2$: two cycles per year), one at frequency $\pi/2$ (i.e., one cycle per year), and one at frequency $3\pi/2$ (again one cycle per year, as a result of aliasing three cycles per year).

Given the evidence (section 3.3.3) that the time series used in this study have seasonal unit roots at π and $\pi/2$, except the interest rate that appears to be stationary at the seasonal frequencies, their stochastic properties can be analyzed under the conjecture that these series are generated by a seasonal process of the form:

$$A(L)(1-L^4)x_t = \varepsilon_t. \tag{5.1}$$

To allow for cointegration analysis at the seasonal frequencies following Johansen framework (see also Lee 1992), this model can be rewritten as

$$D(L)\Delta_4 x_t = \Pi_1 y_{1,t-1} + \Pi_2 y_{2,t-1} + \Pi_3 y_{3,t-1} + \Pi_4 y_{3,t} + \mu + \varepsilon_t, \tag{5.2}$$

where

$$y_{1,t} = S_1(L)x_t = (1+L+L^2+L^3)x_t,$$

$$y_{2,t} = S_2(L)x_t = (1-L+L^2-L^3)x_t,$$

$$y_{3,t} = S_3(L)x_t = (L-L^3)x_t = L(1-L^2)x_t.$$

Note that the filtered series $y_{1,t}$, $y_{2,t}$ and $y_{3,t}$ have only one unit root at frequency 0, π and $\pi/2$, respectively.

When the matrix Π_i (i=1,2,3,4) has full rank, the series does not have unit root at that frequency and when the rank of Π_i is zero, the series are not seasonally cointegrated at that frequency. In the intermediate case when $0 < \text{rank}(\Pi_i) = r < p$, it can be shown that Π_i can be decomposed into the product of two suitable pxr matrices α_i and β_i, $\Pi_i = \alpha\beta_i'$, such that $\beta_i' y_{i,t-1}$

(i=1,2,3,4) is stationary even though $y_{i,t-1}$ itself is non-stationary.[1] The series $\beta_i' y_{i,t-1}$ are called the error-correction terms, and the columns of the matrices β_i identify cointegrating vectors. The hypothesis of seasonal cointegration can be formulated as the condition that rank(Π_i) $\leq$ r for the corresponding frequency against the alternative that rank(Π_i) > r.

Model (5.2) dealing with seasonally unadjusted data is different from the model (4.3) dealing with seasonally adjusted data: while the former allows for four cointegrating relations (at the zero frequency and at the seasonal frequencies), the latter allows for one cointegrating relation only at the zero frequency; while the former consists of fourth differenced variables, the latter consists of first differenced variables.

As in the previous chapter, the first equation of the system (5.2) identifies the forecasting equation for money, from which the anticipated component of money, $P_{t-1}[M_t]$, is taken to be replaced in the output equation. The latter, identical to equation (4.4) of the previous chapter, can be equivalently rewritten in fourth-differences, for consistency with (5.2). It is easily shown that from (4.4), one gets:

$$A(L)\Delta_4 Y_t = B(L)(\Delta_4 M_t - P_{t-1}[\Delta_4 M_t]) + C(L)P_{t-1}[\Delta_4 M_t] + \varepsilon_{4,t}. \tag{5.3}$$

---

[1] In this notation, for convenience, $y_{3,t}$ is rewritten as $y_{4,t-1}$.

Now the prediction of money, $P_{t-1}[\Delta_4 M_t]$ is replaced with the linear forecast:

$$P_{t-1}[\Delta_4 M_t] = (1-D_{11}^*(L))\Delta_4 M_t - D_{12}^*(L)\Delta_4 r_t - D_{13}^*(L)\Delta_4 P_t \quad (5.4)$$
$$- D_{14}^*(L)\Delta_4 Y_t + \delta_{11}^* Z_{1,t} + \delta_{21}^* Z_{2,t} + \delta_{31}^* Z_{3,t} + \delta_{41}^* Z_{4,t}$$

where $D_{11}^*(L)$, $D_{12}^*(L)$, $D_{13}^*(L)$ and $D_{14}^*(L)$ are components of a matrix $D^*(L)$ in general different from $D(L)$ of (5.2) and $\delta_{i1}^* Z_{i,t}$ (i=1,2,3,4) are the error-correction terms. Combining (5.4) with (5.3) and substituting in (5.2), one obtains a four-variable "unrestricted" VAR model of the form

$$\begin{bmatrix} D_M(L) & D_r(L) & D_P(L) & D_Y(L) \\ N_1^*(L) & N_2^*(L) & N_3^*(L) & N_4^*(L) \end{bmatrix} \cdot \begin{bmatrix} \Delta_4 M_t \\ \Delta_4 r_t \\ \Delta_4 P_t \\ \Delta_4 Y_t \end{bmatrix} = [\mu] + [\Psi]\, D_t \quad (5.5)$$

$$+ \begin{bmatrix} \delta_1 \\ \delta_1^*(L) \end{bmatrix} Z_{1,t} + \begin{bmatrix} \delta_2 \\ \delta_2^*(L) \end{bmatrix} Z_{2,t} + \begin{bmatrix} \delta_3 \\ \delta_3^*(L) \end{bmatrix} Z_{3,t} + \begin{bmatrix} \delta_4 \\ \delta_4^*(L) \end{bmatrix} Z_{4,t} + [\varepsilon_t],$$

where D_t are seasonal dummies and $D_M(L)$, $D_r(L)$, $D_P(L)$, $D_Y(L)$, δ_1, δ_2, δ_3, δ_4 are all 3x1 vectors, μ and Ψ are 4x1 vectors, and

$$N_1^*(L) = (1 - D_{11}^*(L))[B(L) - C(L)] - B(L), \quad (5.6)$$
$$N_2^*(L) = - D_{12*}(L)[B(L) - C(L)],$$
$$N_3^*(L) = - D_{13}^*(L)[B(L) - C(L)],$$
$$N_4^*(L) = A(L) - D_{14}^*(L)[B(L) - C(L)],$$
$$\delta_1^*(L) = \delta_{11}^*[C(L) - B(L)],$$
$$\delta_2^*(L) = \delta_{21}^*[C(L) - B(L)],$$

$$\delta_3^*(L) = \delta_{31}^*[C(L) - B(L)],$$

$$\delta_4^*(L) = \delta_{41}^*[C(L) - B(L)].$$

The order of polynomials of D_{ij} (i=1,2,3; j=1,2,3,4) and $N_1^*(L)$, $N_2^*(L)$, $N_3^*(L)$ and $N_4^*(L)$ will be determined in the empirical work.

5.2 Empirical Results

As it is desirable to include seasonal dummies in the regression equation when seasonally unadjusted data is used, the empirical test is carried out based on the following model:

$$D(L)\Delta_4 x_t = \Pi_1 y_{1,t-1} + \Pi_2 y_{2,t-1} + \Pi_3 y_{3,t-1} + \Pi_4 y_{4,t-1} + \Psi D_t + \mu + \varepsilon_t. \quad (5.7)$$

The testable hypothesis of cointegration is formulated as

$$H_1 : \Pi_1 = \alpha_1\beta_1', \quad (5.8)$$

$$H_2 : \Pi_2 = \alpha_2\beta_2',$$

$$H_3 : \Pi_3 = \alpha_3\beta_3',$$

where H_1 corresponds to cointegration at the zero frequency (ω=0), H_2 at the biannual frequency ($\omega=\pi$) and H_3 at the annual frequency ($\omega=\pi/2$).[2]

---

[2] As pointed out by Lee (1992), Π_4 can be ignored with little effect on the test of seasonal cointegration at frequency $\omega = \pi/2$.

With two autoregressive lags for $\Delta_4 X_t$, two cointegrating relations were found at the 5 percent confidence level: one at the long-run or zero frequency (ω=0) and the other at the biannual frequency ($\omega=\pi$) (see Table 5.1); no cointegrating relation was found at the annual frequency ($\omega=\pi/2$). The corresponding error-correction terms are

$$Z_{1,t} = \Delta_a M_{t-1} + .38\Delta_a r_{t-1} - .88\Delta_a P_{t-1} - 1.27\Delta_a Y_{t-1}, \qquad (5.9)$$

$$Z_{2,t} = \Delta_b M_{t-1} - .40\Delta_b r_{t-1} + .01\Delta_b P_{t-1} - .05\Delta_b Y_{t-1}, \qquad (5.10)$$

where $\Delta_a = (1+L+L^2+L^3)$ and $\Delta_b = (1-L+L^2-L^3)$.

Next step is to rewrite the money equation in the fourth-difference with the appropriate error-correction terms. After eliminating some insignificant variables (the second lag of interest rate, price and output), one gets:

$$\Delta_4 M_t = -\sum_{j=1}^{2} d_{11}\Delta_4 M_{t-j} - d_{12}^1\Delta_4 r_t - d_{13}^1\Delta_4 P_t \qquad (5.11)$$

$$- d_{14}^1\Delta_4 Y_t + \delta_{11}Z_{1,t} + \delta_{21}Z_{2,t} + \varepsilon_{1,t},$$

from which the anticipated component of money, $P_{t-1}[\Delta_4 M_{t4}]$, is derived and replaced into the output equation. By setting the order of A(L) at two, and of B(L) and C(L) at one, the final system to be estimated becomes:

$$\begin{bmatrix} D_{11}(L) & D_{12}(L) & D_{13}(L) & D_{14}(L) \\ D_{21}(L) & D_{22}(L) & D_{23}(L) & D_{24}(L) \\ D_{31}(L) & D_{32}(L) & D_{33}(L) & D_{34}(L) \\ N_1^*(L) & N_2^*(L) & N_3^*(L) & N_4^*(L) \end{bmatrix} \begin{bmatrix} \Delta_4 M_t \\ \Delta_4 r_t \\ \Delta_4 P_t \\ \Delta_4 Y_t \end{bmatrix} = \begin{bmatrix} \mu_1 \\ \mu_2 \\ \mu_3 \\ \mu_4 \end{bmatrix} + \begin{bmatrix} \Psi_1 \\ \Psi_2 \\ \Psi_3 \\ \Psi_4 \end{bmatrix} D_t \qquad (5.12)$$

Table 5.1

Results of Seasonal Cointegration Tests

1. Cointegration at frequency $\omega=0$

Eigenvalues($\lambda_{1,i}$)

(.0000726 .176365 .2025911 .4124265)

Standardized β_1' Eigenvectors

Money	1.000000	.382488	-.877462	-1.272482
Interest Rate	-34.052260	1.000000	24.569955	46.079802
Price	-.985450	.405314	1.000000	1.234288
Output	-2.733788	-1.212528	3.497089	1.000000

Standardized α_1 Coefficients

Money	-.032049	.000369	.003094	.000019
Interest Rate	-.103535	-.001481	-.018194	-.000278
Price	.007436	.000607	-.016694	-.000044
Output	-.023910	-.000588	-.017326	.000109

Test Statistics

H_1	λ_{1max}	Trace	Trace(.05)[1]
$r \leq 3$	.005953	.005953	8.6
$r \leq 2$	15.910280	15.916233	19.3
$r \leq 1$	18.563790	34.480023	34.5
$r = 0$	43.603823	78.083846	48.42[2]

[1] Critical values cited from Lee and Siklos (1992).

[2] Critical value cited from Johanssen and Juselius (1990).

Table 5.1 (continued)

2. Seasonal cointegration at frequency $\omega=\pi$

Eigenvalues($\lambda_{2,i}$)

(.0201999 .085669 .1681971 .2832053)

Standardized β_2' Eigenvectors

Money	1.000000	-.399556	.012118	-.054734
Interest Rate	7.907871	1.000000	1.541847	-.050374
Price	-1.636604	-.106067	1.000000	-.195242
Output	-2.247642	-.714841	.894008	1.000000

Standardized α_2 Coefficients

Money	.149656	.044915	.020324	-.000775
Interest Rate	-1.580823	.069518	-.095472	-.004498
Price	-.136873	.006856	.127366	.004642
Output	.043275	.011093	-.099508	.015498

Test Statistics

H_1	λ_{2max}	Trace	Trace(.05)[1]
$r \leq 3$	1.673350	1.673350	8.6
$r \leq 2$	7.344136	9.017486	19.3
$r \leq 1$	15.101101	24.118587	34.4
$r = 0$	27.303197	51.421784	48.42[2]

[1] Critical values cited from Lee and Siklos (1992).

[2] Critical value cited from Johanssen and Juselius (1990).

Table 5.1 (continued)

3. Seasonal cointegration at frequency $\omega=\pi/2$

Eigenvalues($\lambda_{3,i}$)

(.0019733 .0492798 .1389254 .2723608)

Standardized β_3' Eigenvectors

Money	1.000000	.359141	-.193428	-.047499
Interest Rate	3.136121	1.000000	7.572862	2.550857
Price	48.571826	-8.663445	1.000000	-11.806800
Output	5.405124	-.524842	.112133	1.000000

Standardized α_3 Coefficients

Money	.084346	.000236	.001010	.004893
Interest Rate	.999384	.048844	-.002340	-.014290
Price	-.087439	.040870	.001372	-.000215
Output	-.025848	.007150	-.004051	.004208

Test Statistics

H_1	λ_{3max}	Trace	Trace(.05)[1]
$r \leq 3$	.161971	.161971	11.9
$r \leq 2$	4.143909	4.305880	24.3
$r \leq 1$	12.265079	16.570959	40.6
$r = 0$	26.071897	42.642856	..[2]

[1] Critical values cited from Lee and Siklos (1992).

[2] Critical value is not available.

$$+ \begin{bmatrix} \delta_1 \\ \delta_1^*(L) \end{bmatrix} Z_{1,t} + \begin{bmatrix} \delta_2 \\ \delta_2^*(L) \end{bmatrix} Z_{2,t} + [\varepsilon_t],$$

where

$$D_{11}(L) = \Sigma_{j=0}^{2} d_{11}^{j} L^j, \quad d_{11}^{0} = 1, \tag{5.13}$$

$$D_{12}(L) = d_{12}^{1} L, \quad d_{12}^{0} = 0,$$

$$D_{13}(L) = d_{13}^{1} L, \quad d_{13}^{0} = 0,$$

$$D_{14}(L) = d_{14}^{1} L, \quad d_{14}^{0} = 0,$$

$$D_{21}(L) = \Sigma_{j=0}^{2} d_{21}^{j} L^j, \quad d_{21}^{0} = 0,$$

$$D_{22}(L) = \Sigma_{j=0}^{2} d_{22}^{j} L^j, \quad d_{22}^{0} = 1,$$

$$D_{23}(L) = \Sigma_{j=0}^{2} d_{23}^{j} L^j, \quad d_{23}^{0} = 0,$$

$$D_{24}(L) = \Sigma_{j=0}^{2} d_{24}^{j} L^j, \quad d_{24}^{0} = 0,$$

$$D_{31}(L) = \Sigma_{j=0}^{2} d_{31}^{j} L^j, \quad d_{31}^{0} = 0,$$

$$D_{32}(L) = \Sigma_{j=0}^{2} d_{32}^{j} L^j, \quad d_{32}^{0} = 0,$$

$$D_{33}(L) = \Sigma_{j=0}^{2} d_{33}^{j} L^j, \quad d_{33}^{0} = 1,$$

$$D_{34}(L) = \Sigma_{j=0}^{2} d_{34}^{j} L^j, \quad d_{34}^{0} = 0,$$

$$\begin{aligned} N_1^*(L) &= [1 - D_{11}^*(L)][B(L) - C(L)] - B(L) \\ &= - \Sigma_{j=1}^{2} d_{11}^{j*} L^j [\Sigma_{j=0}^{1} (b_j - c_j)] L^j - \Sigma_{j=0}^{1} b_j L^j \\ &= - b_0 - [d_{11}^{1*}(b_0 - c_0) + b_1] L - [d_{11}^{2*}(b_0 - c_0) + d_{11}^{1*}(b_1 - c_1)] L^2 \\ &\quad - [d_{11}^{2*}(b_1 - c_1)] L^3, \end{aligned}$$

$$\begin{aligned} N_2^*(L) &= - D_{12}^*(L)[B(L) - C(L)] \\ &= - d_{12}^{1*} L [\Sigma_{j=0}^{1} (b_j - c_j)] L^j \\ &= - [d_{12}^{1*}(b_0 - c_0)] L - [d_{12}^{1*}(b_1 - c_1)] L^2, \end{aligned}$$

$$\begin{aligned} N_3^*(L) &= - D_{13}^*(L)[B(L) - C(L)] \\ &= - d_{13}^{1*} L [\Sigma_{j=0}^{1} (b_j - c_j)] L^j \\ &= - [d_{13}^{1*}(b_0 - c_0)] L - [d_{13}^{1*}(b_1 - c_1)] L^2, \end{aligned}$$

$$
\begin{aligned}
N_4^*(L) &= A(L) - D_{14}^*(L)[B(L) - C(L)] \\
&= a_0 + a_1L + a_2L^2 - d_{14}^{1*}L^1[\Sigma_{j=0}^{1}(b_j - c_j)]L^j \\
&= 1 + [a_1-d_{14}^{1*}(b_0-c_0)]L + [a_2-d_{14}^{1*}(b_1-c_1)]L^2, \\
\delta_1^*(L) &= \delta_{11}^*[C(L) - B(L)] = \delta_{11}^*\Sigma_{j=0}^{1}(c_j-b_j)L^j \\
&= \delta_{11}^*(c_0-b_0) + \delta_{11}^*(c_1-b_1)L, \\
\delta_2^*(L) &= \delta_{21}^*[C(L) - B(L)] = \delta_{21}^*\Sigma_{j=0}^{1}(c_j-b_j)L^j \\
&= \delta_{21}^*(c_0-b_0) + \delta_{21}^*(c_1-b_1)L.
\end{aligned}
$$

D_{ij}^k (i=1,2,3; j=1,2,3,4; k=1,2) is the (i,j)-th element of the 4x4 matrix of coefficients at lag k. Note that for generality, no zero restrictions were imposed on the equations of the interest rate and price.

The corresponding "unrestricted" reduced form is:

$$
\begin{bmatrix} \Delta_4 M_t \\ \Delta_4 r_t \\ \Delta_4 P_t \\ \Delta_4 Y_t \end{bmatrix} = \begin{bmatrix} -d_{11}^1 & -d_{12}^1 & -d_{13}^1 & -d_{14}^1 \\ -d_{21}^1 & -d_{22}^1 & -d_{23}^1 & -d_{24}^1 \\ -d_{31}^1 & -d_{32}^1 & -d_{33}^1 & -d_{34}^1 \\ n_1^{0*}d_{11}^1+n_1^{1*} & n_1^{0*}d_{12}^1+n_2^{1*} & n_1^{0*}d_{13}^1+n_3^{1*} & n_1^{0*}d_{14}^1+n_4^{1*} \end{bmatrix} \begin{bmatrix} \Delta_4 M_{t-1} \\ \Delta_4 r_{t-1} \\ \Delta_4 P_{t-1} \\ \Delta_4 Y_{t-1} \end{bmatrix} \tag{5.14}
$$

$$
+ \begin{bmatrix} -d_{11}^2 & 0 & 0 & 0 \\ -d_{21}^2 & -d_{22}^2 & -d_{23}^2 & -d_{24}^2 \\ -d_{31}^2 & -d_{32}^2 & -d_{33}^2 & -d_{34}^2 \\ n_1^{0*}d_{11}^2+n_1^{2*} & n_2^{2*} & n_3^{2*} & n_4^{2*} \end{bmatrix} \begin{bmatrix} \Delta_4 M_{t-2} \\ \Delta_4 r_{t-2} \\ \Delta_4 P_{t-2} \\ \Delta_4 Y_{t-2} \end{bmatrix}
$$

$$
+ \begin{bmatrix} 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ n_1^{3*} & 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} \Delta_4 M_{t-3} \\ \Delta_4 r_{t-3} \\ \Delta_4 P_{t-3} \\ \Delta_4 Y_{t-3} \end{bmatrix}
$$

$$+\begin{bmatrix}\delta_{11}\\ \delta_{12}\\ \delta_{13}\\ -n_1^{0*}\delta_{11}+\delta_1^{0*}\end{bmatrix}Z_{1,t}+\begin{bmatrix}0\\0\\0\\ \delta_1^{1*}\end{bmatrix}Z_{1,t-1}$$

$$+\begin{bmatrix}\delta_{21}\\ \delta_{22}\\ \delta_{23}\\ -n_1^{0*}\delta_{21}+\delta_2^{0*}\end{bmatrix}Z_{2,t}+\begin{bmatrix}0\\0\\0\\ \delta_2^{1*}\end{bmatrix}Z_{2,t-1}.$$

The results of estimating (5.14) are reported in Table 5.2. The null hypothesis of no structural change was not rejected in all equations and F-tests for heteroscedasticity of residuals show that the null hypothesis of homoscedasticity was not rejected in every equation. While the error correction terms at the zero frequency, $Z_{1,t}$ and $Z_{1,t-1}$, are significant except in the price equation, the error-correction terms at the biannual frequency, $Z_{2,t}$, are insignificant except in the interest rate equation.

By imposing the joint restriction of rationality, $d_{1j}^{k*} = d_{1j}^{k}$ (j,k = 1,2,3), and of neutrality, C(L) ≡ 0, one obtains the "restricted" reduced model:

$$\begin{bmatrix}\Delta_4 M_t\\ \Delta_4 r_t\\ \Delta_4 P_t\\ \Delta_4 Y_t\end{bmatrix}=\begin{bmatrix}-d_{11}^1 & -d_{12}^1 & -d_{13}^1 & -d_{14}^1\\ -d_{21}^1 & -d_{22}^1 & -d_{23}^1 & -d_{24}^1\\ -d_{31}^1 & -d_{32}^1 & -d_{33}^1 & -d_{34}^1\\ b_1 & 0 & 0 & a_1\end{bmatrix}\begin{bmatrix}\Delta_4 M_{t-1}\\ \Delta_4 r_{t-1}\\ \Delta_4 P_{t-1}\\ \Delta_4 Y_{t-1}\end{bmatrix} \qquad (5.15)$$

Table 5.2

Results of Estimation of Unrestricted Model

	$\Delta_4 M_{t-1}$	$\Delta_4 r_{t-1}$	$\Delta_4 P_{t-1}$	$\Delta_4 Y_{t-1}$	$\Delta_4 M_{t-2}$
$\Delta_4 M_t$	1.119 (10.27)	-.026 (-1.37)	.093 (2.50)	-.153 (-2.39)	-.263 (-2.50)
$\Delta_4 r_t$	.710 (1.16)	.553 (3.87)	.426 (1.35)	-.618 (-1.88)	.415 (-.71)
$\Delta_4 P_t$	.111 (.54)	-.034 (-.71)	.637 (6.06)	-.061 (-.55)	.155 (.80)
$\Delta_4 Y_t$	.804 (3.15)	.100 (1.57)	-.340 (-2.62)	0 (0)	-.645 (-2.22)

	$\Delta_4 r_{t-2}$	$\Delta_4 P_{t-2}$	$\Delta_4 Y_{t-2}$	$\Delta_4 M_{t-3}$	$Z_{1,t}$
$\Delta_4 M_t$	0	0	0	0	-.027 (-3.30)
$\Delta_4 r_t$	-.163 (-1.55)	-.380 (-1.25)	.477 (1.51)	0	-.137 (-3.10)
$\Delta_4 P_t$	.061 (1.74)	.159 (1.56)	.134 (1.27)	0	.016 (1.05)
$\Delta_4 Y_t$	-.042 (-1.09)	-.022 (-.20)	.086 (.69)	.286 (1.54)	-.359 (-4.03)

Numbers in parentheses are t-statistics.

Table 5.2 (continued)

	$Z_{1,t-1}$	$Z_{2,t}$	$Z_{2,t-1}$	Chow	F-test
$\Delta_4 M_t$	0	.100 (1.37)	0	1.11(16,55)[1]	.66(17,53)[2]
$\Delta_4 r_t$	0	-1.557 (-4.06)	0	1.03(16,52)	1.03(23,44)
$\Delta_4 P_t$	0	-.170 (-1.33)	0	.41(16,52)	1.20(23,44)
$\Delta_4 Y_t$	.322 (3.42)	.046 (.33)	0 (0)	.83(16,50)	.87(25,40)

[1] The critical value of F(15,60) = 1.84 and F(15,50) = 1.87 at the 5 percent confidence level.

[2] The critical value of F(20,40) = 1.84 at the 5 percent confidence level.

Numbers in parentheses are t-statistics.

$$+\begin{bmatrix} -d_{11}^2 & 0 & 0 & 0 \\ -d_{21}^2 & -d_{22}^2 & -d_{23}^2 & -d_{24}^2 \\ -d_{31}^2 & -d_{32}^2 & -d_{33}^2 & -d_{34}^2 \\ b_1 d_{11}^1 & b_1 d_{12}^1 & b_1 d_{13}^1 & a_2 + b_1 d_{14}^1 \end{bmatrix} \begin{bmatrix} \Delta_4 M_{t-2} \\ \Delta_4 r_{t-2} \\ \Delta_4 P_{t-2} \\ \Delta_4 Y_{t-2} \end{bmatrix}$$

$$+\begin{bmatrix} 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ 0 & 0 & 0 & 0 \\ b_1 d_{11}^2 & 0 & 0 & 0 \end{bmatrix} \begin{bmatrix} \Delta_4 M_{t-3} \\ \Delta_4 r_{t-3} \\ \Delta_4 P_{t-3} \\ \Delta_4 Y_{t-3} \end{bmatrix}$$

$$+\begin{bmatrix} \delta_{11} \\ \delta_{12} \\ \delta_{13} \\ 0 \end{bmatrix} Z_{1,t} + \begin{bmatrix} 0 \\ 0 \\ 0 \\ -b_1 \delta_{11} \end{bmatrix} Z_{1,t-1} + \begin{bmatrix} \delta_{21} \\ \delta_{22} \\ \delta_{23} \\ 0 \end{bmatrix} Z_{2,t} + \begin{bmatrix} 0 \\ 0 \\ 0 \\ -b_1 \delta_{21} \end{bmatrix} Z_{2,t-1}.$$

This model imposes the following set of six non-linear cross-equation restrictions:

$$\begin{aligned} \phi_{11}{}^1 \phi_{41}{}^1 + \phi_{41}{}^2 &= 0, \\ \phi_{12}{}^1 \phi_{41}{}^1 + \phi_{42}{}^2 &= 0, \\ \phi_{13}{}^1 \phi_{41}{}^1 + \phi_{43}{}^2 &= 0, \\ \phi_{11}{}^2 \phi_{41}{}^1 + \phi_{41}{}^3 &= 0, \\ \psi_{11}{}^0 \phi_{41}{}^1 + \psi_{14}{}^1 &= 0, \\ \psi_{21}{}^0 \phi_{41}{}^1 + \psi_{24}{}^1 &= 0, \end{aligned} \tag{5.16}$$

where ϕ_{ij}^k is the (i,j)-th element of the 4x4 matrix of coefficients at lag k of the "unrestricted" model of (5.14), and ψ_{ij}^s is the j-th element of the column vector (i=1 for $Z_{1,t}$ and i=2 for $Z_{2,t}$) of coefficients associated with the error-correction terms $Z_{1,t}$, $Z_{2,t}$ (s=0) and $Z_{1,t-1}$, $Z_{2,t-1}$ (s=1) of (5.14).

By testing these restrictions through the Wald test, a test statistic W = .21 was obtained, compared with a χ^2 critical value for five degrees of freedom of 11.07 at the 5% level and 9.24 at the 10% level. Thus, the Wald statistic under seasonal cointegration turned out to be quite smaller than the similar statistic under standard cointegration. The ambiguity of interpreting the previous test at the 10% confidence level disappears with seasonally unadjusted data. Now one cannot reject the joint hypothesis of rationality and neutrality even at levels of much higher than 10%.

Although the difference in test results between the two cases is somewhat marginal, it is nonetheless sufficient to confirm that seasonally unadjusted data may contain useful information. This information, if properly used, for example, through the procedure of seasonal cointegration, may be enough to reverse test result or to improve forecast, compared to the use of seasonally adjusted data. In essence, this study shows that seasonal cointegration yields a better representation of the statistical properties of seasonal time series than the commonly used X-11 filter.

Chapter VI

CONCLUSIONS

The joint hypothesis of rationality and money neutrality was tested with Korean time series data in two ways: (i) by using data seasonally adjusted through the X-11 filter and (ii) by using seasonally unadjusted data. The data set consists of money, interest rate, price and output and spans the period from 1970.I to 1991.IV. For the test, a general error-correction multivariate framework was adopted.

As a preliminary step for the development of the model, unit root tests on each variable were conducted based on Dickey-Fuller (1979, 1981) and Perron (1989). Upon establishing that these four variables are integrated of order one, cointegrating relationships among them were investigated following Johansen (1988, 1989, 1991; Johansen and Juselius 1990). With seasonally adjusted data, one cointegrating relation was found. With seasonally unadjusted data, one cointegrating relation at the zero frequency ($\omega=0$) and one at the biannual frequency ($\omega=\pi$) were found under the seasonal cointegration framework developed by Lee (1992). In each case, the error-correction terms corresponding to each of these cointegrating relations were incorporated into the multivariate model.

The money equation estimated with this four variable model was then substituted into an output equation, where output is affected by both unanticipated and anticipated money. This procedure yields a 4x4 "unrestricted" model, in which no restriction is imposed by the joint hypothesis of rationality and neutrality. In selecting the dynamic structure of the model, various criteria such as SC, AIC, HQC and FPE were employed. Finally, by imposing the joint hypothesis yielded a number of non-linear cross-equation restrictions which were tested with the Wald test procedure.

The p-value of the test statistic is about .10 with seasonally adjusted data, so that the joint hypothesis cannot be rejected at the 5 percent confidence level, but the test result is ambiguous at the 10 percent level. With seasonally unadjusted data under seasonal cointegration, the joint hypothesis cannot be rejected even at levels much higher than 10 percent. The difference between the two cases, although somewhat marginal, lies in the different test framework used with different data set. With seasonally unadjusted data under seasonal cointegration, additional information is available compared to the case with seasonally adjusted data: the error-correction term at the seasonal frequency $\omega=\pi$. This information improves the forecast of money, and thus makes the effect of the anticipated component of money stronger than in the previous case.

The result of this study is relevant in two ways. First, it provides evidence of money neutrality and rationality in Korean data. This evidence implies that only when the monetary disturbances are not fully anticipated, they affect real output. The policy implication of this ineffectiveness result is that a fixed money supply rule such as Friedman's k percent rule may be preferable to the countercyclical money supply rule for fine-tuning the economy. In fact, under rational expectations the former policy minimizes the variance of real output, and reduces the equilibrium rates of inflation (Barro and Gordon 1983a, 1983b).

Second, it offers evidence that valuable information contained in the data set may be lost when seasonally adjusted data are used rather than seasonally unadjusted data. Given the current practice of most researchers in Korea to seasonally adjust economic data with the X-11 filter for their empirical work, this evidence is quite compelling: the recent development of econometric procedures, such as seasonal cointegration, suitable to analyze seasonally unadjusted data with unit roots at the seasonal frequencies makes the filtering with X-11 unnecessary and inefficient.

However, this study has a limitation: the number of regressors to be estimated is too large compared with the number of observations. This may cause multicollinearity among variables, which in turn, makes the t-value of regressors quite low; it also decreases the test power. This

problem, however, is not serious, as estimates still hold the properties of unbiasedness and consistency, and the decrease in test power is compensated by the correctness of the test statistics. The problem of the relatively insufficient observations is not easily solved, since some data set such as the broad money M_2 is only available from 1970 in Korea. It will be interesting to reinvestigate the robustness of the joint hypothesis test with a larger number of observations several years from now.

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