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A RATIONAL EXPECTATIONS EQUILIBRIUM MODEL OF THE BUSINESS CYCLE: SOME FURTHER EVIDENCE FOR KOREA 1970. 1 - 1990. 4*

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This paper tests the rational expectations equilibrium model of the Korean business cycle in which unpredictable monetary growth affect the unemployment whereas both predictable and unpredictable monetary growth influence the price level. The model is estimated by using both OLS and efficient estimation procedures. The empirical results show that in the framework of rational expectations the unpredictable movements in the domestic money supply do cause fluctuations in the macroeconomic variables such as the unemployment and price level.

I. INTRODUCTION

Fluctuations in real economic activity are a feature of every economy and have been the focus of much theoretical work even before the Keynesian revolution.

In the 1960s though the emergence of high inflation and recession gave new impetus to the business cycle analysis and in particular suggested the need to devise models capable of providing a more statistically satisfactory and theoretically rigorous explanation of both price and output fluctuations than was provided in the early Keynesian models of, for example, Samuelson (1939), Kaldor (1940) and Hicks (1950).

An early and highly influential attempt to provide such a model was Friedman's (1968) statement of its so-called Natural Rate Hypothesis. This, particularly when later combined with the hypothesis of rational expectations by, for example, Lucas (1972), became a dominant model of the business cycle. The es-

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sential features of this model are that (i) individuals have limited information especially about the prices of goods and services; (ii) supplies respond to what they perceive as relative price movements by increasing or decreasing real output depending upon whether those relative price movements are positive or negative; (iii) agents use their limited information in a optimal, rational way when forming expectations; (iv) trade takes place at market clearing prices.

The essential, distinctive result of such models is that movements in aggregate demand variables such as the quantity of money or nominal spending itself cannot affect real variables, and therefore cannot be the source of the business cycle if they are predictable but they are unpredictable.

This result, which when applied by Barro (1976) became known as the policy ineffectiveness proposition was given empirical importance in a series of influential paper by Barro (1977, 1978) which will be discussed more fully in Section II. A number of studies have developed the Barro's equilibrium model and have estimated it with the data of many countries. These studies have been rare and less fruitful in explaining the business cycles of the developing economy. And comparatively little empirical works in the context of the Korean economy have been conducted up to date.

The purpose of this paper is to provide some further evidence on whether any success that this model have had in explaining business cycles in developed countries can be achieved when it is applied to a country at a somewhat different stage of economic development. In so doing I review the literature on the empirical evidence for the rational expectations equlibirum models which are developed up to date. Then I test the Barro-type model which is estimated for the unemployment and price level over the period 1970. 1 - 1990. 4.

This paper is organized as follows: In Section II I provide the empirical evidence review. In Section III I present the empirical result obtained by the two-step and efficient estimation procedures.

II. THE LITERATURE REVIEW ON EMPIRICAL EVIDENCE

2.1 The Basic Empirical Model

The central problem with testing the rational expectations equlibrium model is that of making empirically operational the theoretical distinction between unpredictable and predictable aggregate demand changes.

In a highly influential group of papers Barro (1977 & 1978) distinguished between predictable and unpredictable monetary growth - the variable he assumed influenced aggregate demand - in the following way: he postulates and estimates what can be seen as a policy reaction function in which U.S. monetary growth is taken as to be a function of a set of other macroeconomic variables. In general form this policy reaction function can be written as

$$DM_t = \underline{\alpha}\,\underline{X}_{t-1} + \mu_t \tag{2.1}$$

where DM_t is the rate of growth of a monetary aggregate, $\underline{X}_{t-1}$, a vector of lagged variables, $\underline{\alpha}$, a vector of coefficient, and μ_t, the error term. In Barro's first paper (1977) the monetary aggregate is $M1$; the X vector consists of monetary growth lagged once and twice, a lagged unemployment variable defined as $\log(U/1-U)$ (U is the annual average unemployment rate) and a measure of federal expenditure relative to its normal level (FEDV) which I shall examine more fully below. His data is annual and covers the period 1940 - 1973. If equation (2.1) is a satisfactory description of the process by which DM_t evolves then μ_t should be serially uncorrelated with any other variable dated prior to period t. Estimating this equation by OLS Barro obtains estimates of $\underline{\alpha}$ which we can denote by $\hat{\underline{\alpha}}$. Thus, using Barro's estimate of (2.1), the value of DM_t predicted by this regression, $DMH_t = \hat{\underline{\alpha}}\,\underline{X}_{t-1}$ and the component of monetary growth which the regression fails to predict, DMR_t is then simply,

$$DMR_t = DM_t - DMH_t \tag{2.2}$$

If equation (2.2) provides a satisfactory description of the process by which monetary growth evolved in the U.S. over the data period, then it follows from rational expectations that agents will form their expectation of monetary growth in accordance with it. Thus DMH_t can be taken as a measure of rational agents' predicted monetary growth over the data period, and similarly, DMR_t, can be taken as a measure of that component of actual monetary growth which rational agents could not predict. Thus Barro uses the OLS estimate of equation (2.1) to make empirically operational the distinction between predictable and unpredictable monetary growth: the predictable component of monetary growth in the period which we denote by DMH_t is DMH_t, the unpredictable component is DMR_t.

Given these series on DMH_t and DMR_t Barro proceeds to the second stage of his empirical test. Taking his unemployment variable, UN_t, as his measure of real economic activity, Barro estimates a relationship of the following general form:

$$UN_t = \underline{\beta}\,\underline{Z}_t + \sum_{i=0}^{n} \phi_i DMR_{t-i} + \varepsilon_t \tag{2.3}$$

where $\underline{Z}_t$ is a vector of variable assumed to influence the natural rate of unemployment, $\underline{\beta}$, a vector of coefficients, ϕ_i, a negative coefficient, and ε_t, random error term. The two variable which Barro allows to influence the natural rate of unemployment are a variable measuring the likelihood of being conscripted into the army if unemployed, and the minimum wage level. The first of these, if high, is thought to discourage people from registering as unemployed and hence to lower the natural rate of unemployment. The second captures the effects of minimum wage legistration on the natural rate of unemployment: the higher it is, the higher is likely to be the natural rate of unemployment.

A positive monetary shock can cause real economic activity to rise, i.e. can cause unemployment to fall below its natural rate and it is possible for such effects to be less than instantaneous - hence the presence of lagged DMR_s in equation (2.3). Barro allows the length of these lagged effects - i.e. the size of n - to be empirically estimated. That is, he estimates equation (2.3) with an initial, arbitrary value, for n and reduces n if $\hat{\phi}_n$, the estimate of ϕ_n, is insignificant.

Estimating equation (2.3) by OLS Barro found a lag length for DMR_s of 2 and found that the unpredictable monetary shocks had significant negative effects on unemployment as predicted by the theoretical model. To test the second distinctive feature Barro added the predictable component of monetary growth - with the same lag length - equation (2.3) to obtain the estimating equation

$$UN_t = \underline{\beta}\,\underline{Z}_t + \sum_{i=1}^{n} \phi_i DMR_{t-i} + \sum_{i=1}^{n} \gamma_i DMH_{t-i} + \varepsilon_t \tag{2.4}$$

and tested the hypothesis that $\gamma_1 = \gamma_2 = \cdots = \gamma_n = 0$. Using standard 'F' tests Barro found that he could not reject this hypothesis: he could not reject the hypothesis that predictable monetary growth has no effect on real economic activity as measured by the level of unemployment. This result too, then appeared to confirm the empirical relevance of the rational expectation - imperfect information model of the business cycle.

In a subsequent empirical paper Barro (1978) adopted the same approach but used real output, Y and price level, P in log forms. He tested for the significance of unpredictable monetary growth in both the output and price equations and by adding current and lagged DMH_s to the estimating equations Barro tested for the influence of predictable monetary growth. His broad findings were that monetary shocks did affect output and price as his theoretical model suggested. They seemed to show that predictable money growth has no significant impact on real variables, only affecting prices. However, Barro himself expressed some dissatisfaction with his specification of the price equation and suggests that it should be modified to allow for some form of partial adjustment, perhaps by the inclusion of the lagged value of real money, or by allowing for a special response of money demand to temporary movements in real income.

2.2 Extensions to the Basic Empirical Model

Since Barro's original two articles a number of further studies for the U.S. have been made which make some minor amendments to the models. Barro and Rush (1980) attempt to estimate money growth, unemployment, output, and price level equations using quarterly data over the period 1941. I to 1978. I. Their results are little different from the Barro's original findings. Barro (1981) modifies his original unemployment equation by substituting government purchases per GNP variable, G_t / Y_t for a time trend and finds that the significance of unanticipated money growth recedes after one year; Sheehey (1984a) and Dukowsky and Atesoglow (1986) replace the military draft and minimum wage variables in Barro (1977) model with a time trend but find it makes little improvement to his original unemployment equation. Carns and Lombra (1983) reexamine Barro's real output and unemployment equations using different data sources covering the period 1970. 3 - 1979. 3. They finds that their DMR variables had more significant effects on real output and unemployment than did predictable monetary growth, but predictable monetary growth nevertheless had significant real effects. Sheehey (1984a) reappraises the Barro (1977) model using two different sample periods, 1946 - 1978 and 1948 - 1978. He finds that unpredictable money is of great significance and predicted money growth is of less when the years 1946 - 1947 are included.

A number of authors have criticised Barro's use of his federal expenditure variable (FEDV) in his DM equation. The central problem is that FEDV is defined as the time period t level of federal expenditure relative to its normal level which itself is estimated in a regression of actual expenditure on its own lagged values. As Mishkin (1982a) and Pesaran (1982) have both pointed out this variable cannot possibly be known to agents when making their one-period-ahead forecast of DM_t. Pesaran argues that a forecast of FEDV should replace actual FEDV in the money growth equation, and using this procedure Pesaran finds, using non-nested hypothesis testing procedure, that he can reject Barro's model in favour of a loosely specified Keynesian alternative.[1]

Barro's rationale for using actual FEDV rather than a forecast of it in his DM equation is that most of the movement in the series is due to high expenditure during period of war, and that these movements would be predictable once war had started, but they would not be predictable from the behaviour of the variable in years of peace. Rush and Waldo (1988) in reply to Pesaran's criticism

[1] Small (1979) criticises the inclusion of the FEDV variable on different grounds. He makes a distinction between temporary and permanent changes in federal expenditure and finds that the value of the coefficient on federal expenditure is larger in the case of temporary changes in federal expenditure. He suggests that the failure to allow for this may lead to overestimating money supply growth in those periods when changes in federal expenditure have been permanent.

have used a forecast of FEDV in the DM equation but one which, unlike Pesaran's forecast, allows for the effects of war by the use of dummy variables. In this case, Pesaran's Keynesian model can be rejected in favour of a Barro-type model. Perhaps the main lesson to be drawn from this is that if possible the DM equation should contain only variables which can in principle be known to agents at the time they are making their predictions of monetary growth.

Not surprisingly Barro's basic approach has been applied to many other countries. Attfield, Demery, and Duck (1981a & b) modify Barro's estimation procedure and apply the procedure to the U.K. using annual data for 1946 - 1977 and quarterly data for 1963 - 1978. Their broad findings are the same as those of Barro. Bellante, Morrel, and Zardkoodhi (1982) also apply a framework similar to Barro's, using the U.K. annual data over the period 1946 - 1977, and treating a union membership as an influence on the natural rate of unemployment. Their major findings are qualitively similar to Barro's. Furthermore a number of other studies obtain the same sort of results. For instance, Demery, Duck, and Musgrave (1984) for West Germany for the period 1964 - 1981; Wogin (1980) for Canadian data for the period 1927 - 1972; Alogoskoufis (1982) for Greek data for the period 1960 - 1977; Canarella and Pollard (1989) for annual data from 16 Latin American countries.

Since most countries are smaller and more open than the U.S., several paper have modified the basic model to reflect this openness. Sheehey (1984b) applies the Barro model to 16 Latin American countries, modifying the output equation to include terms of trade and relative price of foreign goods variables. Sheehey finds that contrary to the Barro model anticipated money growth does have real effects. Hanson (1980) in his study of 5 Latin American countries (Brazil, Chile, Colombia, Mexico and Peru) finds that if the output equation includes the influence of supply shocks such as crop failures, natural disaster and world price changes then unanticipated monetary growth has insignificant independent influence. The reverse of this result is found by Canarella and Pollard (1989) who employ a autoregressive integrated moving average (ARIMA) model of the money growth equations and estimate Barro's version of the output and price equation for 16 Latin American countries and report findings similar to Barro's. Leiderman (1979) using data for seven industrialized countries, Kimbrough and Koray (1984) for Canada, Chopra and Montiel (1986) for the Philippines, and Montiel (1987) for Mexico, all allow for the openness of the open economy by allowing for the influence on domestic ouput of both domestic and foreign money growth. They overall find that both domestic and foreign unpredictable money growth have positive effects on real output. Cho and Nakibullah (1994) finds these positive effects in the case of the output of non-traded goods for the quarterly data period 1970 to 1987. Dadkhah and Valbuena (1985) make a slight modification to the Barro model by including open economy variables such as the exchange rate and real autonomous exports in the real output equation. Using a non-nested hypothesis framework and using annual data from France, Germany, Italy and Spain they found that a Keynesian model is to be preferred in the case of Germany and Italy, but not in the case of France and Spain. Bryant (1991) reapplies the approach used by Dadkhah and Valbuena to the data from Australia, Canada, Japan, Sweden, and Switzerland. His broad results are identical to those of Valbuena.

Slightly more formally the Barro model can be tested by computing the likelihood statistics from the restricted and unrestricted systems and comparing the resulting likelihoods ratio test statistics with the appropriate chi-square distribution. This method of testing has been used quite extensively, for example, Leiderman (1980), Barro and Rush (1980), Attfield, Demery, and Duck (1981a & b), Attfield and Duck (1983), and Rush (1986).

The Leiderman study is especially interesting in that it suggests a way of separately testing two distinct hypotheses incorporated in the Barro model: (i) the structural neutrality hypothesis that predicted money growth has no real effects; and (ii) the rational expectations hypothesis that predictable money growth equals predicted money growth. Applying this procedure to Barro's data, Leiderman found that the rational expectations hypothesis cannot be rejected by a likelihood ratio test and also that the structural neutrality hypothesis cannot be rejected by a likelihood ratio test of the joint model against the rational expectations model.[2] Barro and Rush (1980) and Rush (1986) attempt to estimate jointly the money growth, unemployment, output, and price level equations with U.S. data. In the case of the U.K. study Attfield, Demery and Duck (1981a, 1981b) employ simultaneous equation system and maximum likelihood estimation technique with full information and estimate the hypothesis of the joint rational expectations and structural neutrality with annual and quartely data. Their findings broadly confirm the Barro's results.

But, in fact, there is a growing empirical literature which uses the apporach explained above and yet finds little support for the model. An early series of papers by Mishkin (1982a & b) uses Barro's approach, but differs in two aspects. First, anticipated money terms are included in Barro's real output equation to help deal with the so-called observational equivalence problem.[3] Second, in his real output equation Mishkin allows lag lengths of twenty quarters on both the anticipated and unanticipated components of the aggregate demand variables.

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[2] Driscoll, Ford, Mullineux, and Sen (1983b) also use the testing procedure suggestd by Leiderman. They find that when applying it to UK data for the period 1946 - 1979 they can reject the joint restrictions implied by rational expectations and structural neutrality.

[3] Sargent (1976) and others (e.g. McCallum (1979), Pudney (1982) and Buiter (1983) have pointed out, the models so far described may have a different interpretation from the rational expectations competitive equilibrium interpretation from which they evolved. It is possible to devise classes of Keynesian models in which anticipated movements in money affect output and which are observationally equivalent to rational expectations models.

Then an attempt is made to estimate the aggregate demand and real output equations jointly imposing the appropriate cross equation restrictions, using quarterly U.S. data for the period 1954 to 1976. This allows the joint and individual tests of rational expectations and structural neutrality which are similar to those of Leiderman. Mishkin finds that lengthening the lag on the unanticipated money variables in output equation leads to rejection of the restriction of the rational expectations and structural neutrality model. In particular he finds that he can reject the joint and individual hypotheses of the rationality and structural neutrality when assuming 7 lags on the monetary shocks, but can reject them when assuming 20 lags.

Following Mishkin, Hoffman and Schlalenhauf (1982) pay particular attention to the lag length in their study of 6 industrial countries, Canada, Germany, Italy, Japan, U.K. & U.S., using quartely data for the period 1960 to 1980. They jointly estimate money growth and output equations imposing the relevant cross-equation restrictions, and consider two lag lengths, 7, 11 on the monetary shocks. Then they investigate whether the significance of unanticipated money in a real output equation is sensitive to the choice of lag length. To do that, the lag length, 7 is specified on the basis of excluding lagged unanticipated money growth which is statistically insignificant and adding the relevant variables in order to obtain more asymtotically efficient results. They test the hypotheses of rationality and neutrality jointly/individually, and find that the joint hypothesis of rationality and neutrality is rejected on a lag length of 11 for all countries except for Canada and Italy, and rejected on a lag length of 7 for all countries. The rational expectations hypothesis is rejected for only Germany with a lag length of 7, and for Germany, Japan, and U.S. with a lag length of 11. The neutrality hypothesis is not rejected for Canada and U.S. with a lag length of 7 and only for Canada with a lag length of 11. Gochoco (1986) applies Mishkin's model to Japanese data for the period 1973 - 1985 and finds that he can reject the joint hypothesis of rationality and neutrality with a choice of lag length of 11. Thoma (1989) tests for the non neutrality of anticipated future money for four industrial countries, Canada, Germany, U.K. and U.S. He explains one theoretical reason for this non-neutrality as follows: When the change in money growth is anticipated the public experts the change in inflation rate prior to the actual change in money growth rate which results in a change in real money balance. The change in real money balance results in nonneutral change in real output. The unanticipated change in money growth, however, is neutral, because it does not change the expected rate of inflation. Then he estimates the vector autoregressive (VAR) model of the money growth and output equations, and finds that for both lag length of 7 and 11 the individual and joint hypothesis of rationality and neutrality are rejected only for Canada.

Darat (1985) finds no real evidence of non neutrality of inflation for the Canadian economy 1960 - 1982, but Darat (1987) finds evidence of the non neutrality of money growth for the Danish economy over the annual data period 1953 - 1983, allowing a relatively short lag length of 2 years on monetary shocks. Choudhary and Parai (1991) also provide evidence for the non neutrality of money in the case of 13 Latin American countries. Beladi and Samanta (1988), using three different forms of the money growth equations - OLS regression, stepwise regression, and ARIMA models, find support for the rational expectations and structural neutrality model when the U.K. industrial production is the real variable but not the U.K. real output is used - their data period is 1952 - 1983 and their data is annual. The same testing procedures when applied by Beladi and Samanta (1988) to Indian data for the period 1955 - 1983, results in the rejection of the rational expectations and structural neutrality hypothesis for both real output and industrial production. The same rejection of the rational expectations and structural neutrality hypotheis is reported by Mohabbat and Ali-Saji (1991) using quarterly data of Iraq for 1961 - 1977. But Marashdel (1993), using autoregressive model - provides the mixed results of the structural neutrality of policy variables, for example money growth, inflation, government expenditure and balance of payment for real output in the case of the Malaysian economy over the quarterly data period 1970 to 1990. He presents that the neutrality of money growth and inflation is rejected but its rejections for govenment expenditure and balance of payment is not applied.

Ⅲ. EMPIRICAL RESULTS FOR KOREA

3.1 Specification for Monetary Growth

The starting point for testing the rational expectations equilibrium model of the business cycle is a satisfactory description of the process by which the money supply has grown over the data period. The objective is to use that description to make empirically operational the theoretically important distinction between predictable and unpredictable monetary growth.

To obtain a good description of the Korean money supply process I carried out a series of regression of monetary growth on a large number of variables.[4] These variables were: lags of monetary growth; government expenditure; the current account of the balance of payments; nominal income; the exchange rate; the inflation rate; the nominal interest rate; the unemployment rate; and a measure of the government deficit. I allowed each of the variables to enter the money growth equation with up to four lags and eliminated them if they appeared to be

[4] As in many industrialized countries, the monetary authorities in Korea used monetary policy over the period as one of many instruments to secure a broad set of objectives rather than linking it to the achievement of one objectives. So many macroeconomic variables have exerted a potential influence on monetary growth over the period.

insignificant unless droping them produced unsatisfactory characteristics for the resulting equation such as evidence of serial correlation.

The most parsimonious representation of the Korean money supply process I found was the following:

$$
\begin{array}{llllll}
DM_t = & -0.037 & +0.241\,DM_{t-1} & +0.187\,DM_{t-2} & +0.0002\,DDP_{t-1} & +0.022\,IR_{t-1} \\
& (0.020) & (0.117) & (0.101) & (0.0001) & (0.009) \\
& [1.816] & [2.006] & [1.851] & [1.557] & [2.565] \\
& +0.018\,UNR_{t-1} & +0.017\,GD_{t-1} & +0.066\,DEX_{t-1} & + \text{Seasonals} & \\
& (0.01) & (0.004) & (0.041) & & \\
& [1.693] & [4.010] & [1.608] & &
\end{array}
\tag{3.1}
$$

$$R^2 = 0.614;\ \text{D. W.} = 2.035;\ \text{Durbin } h = \text{not computable}$$

where: DDP_{t-1} is the change in the quarterly inflation rate between period $t-1$ and $t-2$; IR_{t-1} is the nominal interest in period $t-1$; GD_{t-1} is the real value of the government's budget deficit in period $t-1$; UNR_{t-1} is the log of the unemployment rate in period $t-1$; DEX_{t-1} is the change in the quarterly growth of the value of exports.

Although this equation explains a reasonably high percentage of quarterly monetary growth and has certain desirable statistical characteristics not all of the variables have the estimated coefficients I would initially expected them to have. For example, at least in the early part of the period the Korean authorities, in their attempt to maintain high growth rates, tended to finance government deficits through monetary expansion. Since the GD variable in the monetary growth equation is negative if there is deficit, one would have expected it to appear with a negative rather than a positive sign. One possible explanation for the negative sign is that the government deficit variable is entered with a lag of one quarter and the government deficit happens to be negatively related to the current government deficit but positively ralated to the lagged deficit. This conjecture is supported by a version of equation (3.1) which replaces the lagged with the current value of GD; the coefficient on the current value of GD is negative and significant. Besides, the point of equation (3.1) within this framework is to make operational the distinction between predictable and unpredictable monetary growth. So to some extent the interpretation of the coefficient is not crucial.

The DDP variable enters with a positive sign suggesting that a rise in inflation leads to an increase in monetary growth, though the coefficient is not particularly well determined. This may reflect the Korean authorities desire, at least in the early part of period, not to let inflation hinder their aim of high output growth.[5] The co-

[5] Inflation rates of 20% were quite typical of the 1970s in Korea, only in the 1980s was inflation markedly reduced.

efficients on the other variables in equation (3.1) are in line with what would have expected. High nominal interests and high levels of unemployment are likely to encourage a government whose aim is growth to relax monetary policy; similarly high export growth with its favourable implications for the current account of the balance of payments is likely to allow the authorities to follow a more relaxed monetary policy. So the signs on the IR, UNR, and DEX variables would be expected to be positive as they are.

For equation (3.1) to be an adequate model for distinguishing between rational agents' predictable and unpredictable monetary policy growth it requires at least three features: first the variables on the right hand side of the equation should be, in principle, known to agents in period $t-1$ when forecasting for period t; secondly, the error from the equation should be free of serial correlation so that no exploitable pattern exists in forecast errors; and thirdly it should be stable over the data period.

The first requirement is conventionally satisfied by equation (3.1) since all the right hand side variables are lagged rather than current varibles. In principle, then they are part of rational agents' information when, at the end of period $t-1$, agents are forecasting.

The Durbin-Watson statistic of 2.04 suggests that the null hypothesis of no first-order serial correlation cannot be rejected. However the Durbin-Watson statistic is known to be biased against rejection of the null by the presence of a lagged dependent variables and the alternative statistic - Durbin's 'h' statistic - cannot be computed in this case. In Table 1 I present the autocorrelations of the residuals from equation (3.1) for lags from 1 to 24.

Most of these autocorrelations are very low suggesting that the residuals are truely white noise. In addition, the Box-Ljung Pierce statistics to test the null hypothesis of zero serial correlation were 5.37 for up to 12th order and 26.43 from up to 24th order serial correlation. The critical chi-square variable at the 5% level is 21.03 with 12 degrees of freedom and 36.42 with 24 degrees of freedom. This indicates that the money growth equation is clearly free from serial correlation.

[Table 1] Autocorrelations of Residuals from Equation (3.1)

$\gamma1$	−0.02	$\gamma7$	−0.11	$\gamma13$	−0.19	$\gamma19$	0.08
$\gamma2$	−0.09	$\gamma8$	−0.10	$\gamma14$	−0.03	$\gamma20$	0.05
$\gamma3$	0.09	$\gamma9$	0.10	$\gamma15$	0.23	$\gamma21$	−0.05
$\gamma4$	−0.06	$\gamma10$	0.06	$\gamma16$	−0.21	$\gamma22$	0.01
$\gamma5$	0.02	$\gamma11$	−0.02	$\gamma17$	0.06	$\gamma23$	−0.12
$\gamma6$	0.00	$\gamma12$	0.06	$\gamma18$	0.14	$\gamma24$	−0.17

To test for the structural stability of equation (3.1) I applied the conventional Chow (1960) test to it splitting the period in two at (i) 1979. 4 and then at (ii) 1980. 4. The computed 'F' values were 0.954 and 0.990 respectively compared with a critical F value of 1.99 at the 5% level of significance with, the k. $n-2k$ degrees of freedom where k is the number of variables and n is the number of observations. The tests then imply that the hypothesis of stability in equation (3. 1) cannot be rejected.

All in all, equation (3.1) appears to provide a satisfactory way of distinguishing between predictable and unpredictable monetary growth over the period 1970 - 1990: it explains a reasonable proportion of the changes in monetary growth; it uses only lagged variables to do so; it appears not to suffer from first or high-order serial correlation; and it is stable. Accordingly, in what follows I use equation (3.1) as the basis for my tests of the rational expectations equilibrium approach to the business cycle.

3.2 Analysis of Unemployment and Price Level

In the first test of approach I use the two-step procedure of Barro (1977). I first define unpredictable monetary growth, DMR_t as the residual from equation (3.1), and predictable monetary growth, DMH_t as the difference between actual and unpredictable monetary growth, $DM_t - DMR_t$. Then I made the measure of real economic activity, unemployment which I shall now denote by UNR, (in logarithmic form) as trend statisfactory processes subject to single breaks in the trend function and which are thrown off their trend values by monetary shocks. The form of the OLS regressions for the UNR were:

$$UNR_t = \beta_0 + \sum_{i=0}^{8} \beta_{1+i}\, DMR_{t-i} + \beta_{10}\, T + \beta_{11}\, DU + \mu_{1t} \qquad (3.2)$$

where DMR_t is the error from the equation (3.1); T is a time trend; DU is a dummy allowing for a break in the level of the equation; the β_s are coefficient; β_0 includes any seasonal dummies; μ is mean zero, constant variance, serially uncorrelated error terms.

The results of modelling the business cycle as due to monetary shocks were obtained when the break point was assumed to be 1973. 4 - the period of the first oil price shock.[6] The estimates presented in Table 2 provide some initial support for this approach to the modelling of the Korean business cycle. Several of the DMR variables are significant and they all have the expected sign - positive

[6] I allowed a single structural break in the trend function which is selected with a historical event in the data period. It is assumed that this break will help for the level of macroeconomic activity to have the trend stationary process. [See Perron (1989)]

[Table 2] OLS Estimates of the Unemployment and Price Level Equations

$$UNR_t = -2.65\,DMR_t - 2.58\,DMR_{t-1} - 2.56\,DMR_{t-2} - 1.71\,DMR_{t-3} - 3.00\,DMR_{t-4}$$
$$[2.19] \quad [2.16] \quad [2.11] \quad [1.43] \quad [2.60]$$

$$-2.39\,DMR_{t-5} - 2.11\,DMR_{t-6} - 1.54\,DMR_{t-7} - 2.29\,DMR_{t-8} - 0.009\,T$$
$$[1.94] \quad [1.66] \quad [1.16] \quad [1.79] \quad [7.10]$$

$$+0.19\,DU$$
$$[1.54]$$

$$R^2 = 0.69\ ;\quad \text{D. W.} = 1.32$$

$$\log(P_t) = -1.06\,DMR_t - 1.90\,DMR_{t-1} - 1.90\,DMR_{t-2} - 0.75\,DMR_{t-3} - 0.93\,DMR_{t-4}$$
$$[4.77] \quad [4.08] \quad [4.03] \quad [3.39] \quad [4.39]$$

$$-0.87\,DMR_{t-5} - 0.98\,DMR_{t-6} - 0.91\,DMR_{t-7} - 0.70\,DMR_{t-8} - 0.03\,T$$
$$[3.85] \quad [4.17] \quad [3.73] \quad [2.94] \quad [21.31]$$

$$+0.20\,DU + 1.11\log(M_t) - 0.05\,IR_t$$
$$[8.23] \quad [39.78] \quad [2.89]$$

$$R^2 = 0.998\ ;\ \text{D. W.} = 1.58$$

Notes : Each regression contains a constant and seasonal dummies; T is a time trend. DU is a break dummy allowing for a break in the level of the equation at 1973. 4; The figures in brackets are t statistic.

monetary shocks cause negative deviations in unemployment below its trend. The value of the F statistic for the test of null hypothesis that the coefficients on the current and lagged DMR_s are zero is 2.64 for the unemployment equation compared with a critical value at the 5% level for the appropriate degrees of freedom (9, 58 in this case) of 2.00, suggesting that the null hypothesis can be clearly rejected. The actual coefficients on the DMR_s suggest a much more jagged reaction of unemployment to monetary shocks than the triangular pattern found by Barro using annual data. They suggest that a positive 1% monetary shock will lower unemployment by between 1 and 2%.

I added additional lagged values of DMR to the unemployment equation and found that the null hypothesis that their coefficients were zero could not be rejected by a conventional 'F' test: the computed values was 0.56 compared with critical values of 2.57 at the 5% level with 4, 50 degrees of freedom.

The evidence from the Durbin-Watson statistics of the unemployment equation is indecisive: the statistics fall between the upper and lower critical values.

To explore the possibility of higher order serial correlation for the unemployment equation I computed the Box-Ljung Pierce statistic for 12th order autocorrelation. For the UNR equation it was 20.35. The critical chi-square variate at the 5% level with 12 degrees of freedom is 21.03. So it appears that the error is free from serial correlations.

I next added to the estimating equation current and up to 8 lagged values of anticipated monetary growth and computed a conventional 'F' test of their joint significance. The resulting 'F' statistic was 0.77 for the UNR equation. Since the critical 5% 'F' value with appropriate degrees of freedom (9,49) is 2.07 it follows that the null hypothesis that anticipated monetary growth has no effect cannot be rejected in the case of UNR.

I incorporated a price level equation into the model by inverting a log-linear demand for money equation of the form:

$$\log(M_t) - \log(P_t) = \theta_0 + \theta_2 \log(Y_t) - \theta_3 IR_t + \theta_4 T + \mu_{2t} \tag{3.3}$$

where M is the level of the nominal quantity of money; P is the implicit GNP deflation; Y_t is a measure of real GNP pertinent to money demand; IR is a measure of the nominal interest rate; T is a time trend. The θ_s are coefficients; θ_s includes a constant and seasonal dummies; μ is a mean zero constant variance and random error term. Rearranging equation (3.3) and substituting for $\log(Y_t)$ using a $\log(Y_t)$ equation where $\log(Y_t) = \gamma_0 + \sum_{i=0}^{8} \gamma_{1+i} DMR_{t-i} + \gamma_{10} T + \gamma_{11} DU + \mu_{3t}$ gives in general form the following OLS regression.

$$\begin{aligned}\log(P_t) = -\varphi_0 + \varphi_1 \log(M_t) - \sum_{i=0}^{8} \varphi_{2+i} DMR_{t-i} + \varphi_{11} IR_t - \varphi_{12} T \\ - \varphi_{13} DU - \mu_{4t}\end{aligned} \tag{3.4}$$

where the φs are coefficients; $\varphi_0 = \theta_0 + \theta_2\gamma_0$; $\varphi_1 = 1$; $\sum_{i=0}^{8} \varphi_{2+i} = \sum_{i=0}^{8} \theta_2\gamma_{1+i}$; $\varphi_{11} = \theta_3$; $\varphi_{12} = \theta_2\gamma_{10} + \theta_4$; $\varphi_{13} = \theta_2\gamma_{11}$; $\mu_{4t} = \theta_2\mu_{3t} + \mu_{2t}$

The results of estimating this regression are shown in Table 2. The estimated coefficients are all of the expected sign apart from the coefficient on the nominal interest rate. The DMR coefficients are all negative and apparently highly significant - an 'F' test of the null that all the coefficients are zero produced an 'F' value of 10.75 compared with a 5% critical value of approximately 2.0. An 'F' test on the null hypothesis that four additional lags of DMR were insignificant produced an 'F' value of 1.20 compared with a 5% critical value of 2.57 implying that the null hypothesis could not be rejected. The coefficient on the log of the quantity of money is close to one as the model predicts, though a strict test of its equality to one would reject the hypothesis that it is one.

The Durbin-Watson statistic is in the indeterminate region so tell us little about the presence of first-order serial correlation. To investigate the presence of higher order serial correlation I computed the Box-Ljung Pierce statistic for up to 12th-order autocorrelation. It was 6.56 compared with a 5% critical value of 21.03 so the null hypothesis of no serial correlation could not be rejected.

To test for the irrelevance of anticipated money in this regression (other than in the form in which it enters equation (3.4) I added current and up to 8 lagged values of anticipated monetary growth. The result was an 'F' value of 2.68 which is between the critical 5% and 1% F value with the appropriate degrees of freedom, (9, 49). So the hypothesis that predictable monetary growth affects the price level in ways not captured by equation (3.4) cannot be decisively rejected.

Because the two-step procedure I have used up to now is consistent but not asymtotically efficient, I present the results of estimating the same basic model efficiently, a procedure which allows me to test some of the restrictions which the model implies.

To explain the estimating procedure I shall first write the monetary growth equation - equation (3.1) in compact form as,

$$DM_t = \alpha_0 + \sum_{i=1}^{10} \alpha_i X_{it-1} + \varepsilon_{1t} \tag{3.5}$$

where X_{1it-1} is one of the variables assumed to influence DM_t (including seasonals), and α_i is its coefficient; α_0 is a constant; μ_{1t} is the unpredictable error term associated with monetary growth.

Defining unpredictable monetary growth in period $t-j$ from equation (3.5) above as $\varepsilon_{1t-j} = DM_{t-j} - \alpha_0 - \sum_{i=1}^{16} \alpha_i X_{it-1-j}$, the equations for unemployment and the price level can be rewritten as

$$UNR_t = \beta_0 + \sum_{j=1}^{8} \beta_{j+1} [DM_{t-j} - \alpha_0 - \sum_{i=1}^{10} \alpha_i X_{1it-j}] + \sum_{j=10}^{14} \beta_j X_{2jt} + \beta_1 \varepsilon_{1t} + \varepsilon_{2t} \tag{3.6}$$

$$\begin{aligned} \log(P_t) = \theta_1 \log(M_t) - \theta_0 - \theta_2 [\sum_{j=1}^{8} \beta_{j+1} [DM_{t-j} - \alpha_0 - \sum_{i=1}^{10} \alpha_i X_{1it-j}] \\ + \sum_{j=10}^{14} \beta_j X_{2jt} + \beta_{1t} + \varepsilon_{2t}] - \theta_3 IR_t - \theta_4 T - \varepsilon_{3t} \end{aligned} \tag{3.7}$$

where X_{2jt} variables are the time trend, the break dummy and the three seasonals; ε_{1t}, ε_{2t} and ε_{3t} are error terms.

Efficient estimation involves estimating these equations, or some of them, jointly imposing all the cross equation restrictions implied by the model. Notice that in the above equation I treat the effect on unemployment and the price level of the current monetary shock, ε_{1t}, as part of the error term of the equation. Had I not done so the β_1 coefficient could not have been indentified since it would be attached to an endogenous variable, DM. Estimating the above equations, or a subset of them, through non-linear full information maximum likelihood allows me to compute an estimate of β_1 from the variance-covariance matrix of residuals.[7]

In Tables 3-4 I present the results of estimating equation (3.5) jointly with equation (3.6), and with equation (3.6) and (3.7).

[Table 3] Joint Estimation of Equations (3.5) and (3.6)

$$\underset{}{DM_t} = \underset{(0.11)}{0.15\,DM_{t-1}} + \underset{(0.10)}{0.30\,DM_{t-2}} - \underset{(0.0001)}{0.0001\,DDP_{t-1}} + \underset{(0.006)}{0.025\,IR_{t-1}} + \underset{(0.008)}{0.031\,UNR_{t-1}}$$

$$+ \underset{(0.004)}{0.015\,GD_{t-1}} + \underset{(0.035)}{0.009\,DEX_{t-1}}$$

$R^2 = 0.61$; D. W. = 2.05 ; BL(12) = 10.8

$$UNR_t = \underset{(-)}{-0.34\,DMR_t} - \underset{(1.00)}{2.23\,DMR_{t-1}} - \underset{(1.01)}{2.42\,DMR_{t-2}} - \underset{(1.04)}{1.26\,DMR_{t-3}} - \underset{(0.98)}{3.37\,DMR_{t-4}}$$

$$\underset{(1.08)}{-1.56\,DMR_{t-5}} - \underset{(1.12)}{1.63\,DMR_{t-6}} - \underset{(1.10)}{1.21\,DMR_{t-7}} - \underset{(1.05)}{2.64\,DMR_{t-8}} - \underset{(0.002)}{0.007\,T}$$

$$+ \underset{(0.106)}{0.124\,DU}$$

$R^2 = 0.73$; D. W. = 1.49 ; BL(12) = 11.6

Notes: All the regressions included a constant and seasonal dummies; T is a time trend. DU is a break dummy allowing for a break in the level of the equation at 1973. 4; The figures in brackets are asymtotic standard errors; BL(12) is the Box-Ljung Pierce statistic for 12th order correlation.

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[7] For exmaple, the variance-covariance matrix of residuals from equations (5) and (6) is

$$\begin{bmatrix} \sigma_{\varepsilon_1\varepsilon_1} & \beta_1\sigma_{\varepsilon_1\varepsilon_1} \\ \beta_1\sigma_{\varepsilon_1\varepsilon_1} & \beta_1\sigma_{\varepsilon_1\varepsilon_1} + \sigma_{\varepsilon_1\varepsilon_1} \end{bmatrix}$$

. An unique estimate of β_1 can be obtained from the ratio of $\beta_1\sigma_{\varepsilon_1\varepsilon_1} + \sigma_{\varepsilon_1\varepsilon_1}$. See Attfield, Demery and Duck (1981a)

[Table 4] Joint Estimation of Equations (3.5), (3.6) and (3.7)

$$\begin{aligned} DM_t = & -0.073\,DM_{t-1} + 0.398\,DM_{t-2} - 0.0002\,DDP_{t-1} + 0.022\,IR_{t-1} + 0.021\,UNR_{t-1} \\ & \quad (0.096) \qquad (0.080) \qquad (0.00007) \qquad (0.004) \qquad (0.006) \\ & +0.005\,GD_{t-1} - 0.049\,DEX_{t-1} \\ & \quad (0.003) \qquad (0.022) \end{aligned}$$

$R^2 = 0.515$; D. W.= 1.907 ; BL(12) = 15.34

$$\begin{aligned} UNR_t = & -2.789\,DMR_t - 1.008\,DMR_{t-1} - 1.780\,DMR_{t-2} - 0.971\,DMR_{t-3} - 2.975\,DMR_{t-4} \\ & \quad (-) \qquad (0.390) \qquad (0.451) \qquad (0.378) \qquad (0.451) \\ & -1.331\,DMR_{t-5} - 2.397\,DMR_{t-6} - 2.168\,DMR_{t-7} - 2.086\,DMR_{t-8} - 0.009\,T \\ & \quad (0.357) \qquad (0.463) \qquad (0.411) \qquad (0.431) \qquad (0.001) \\ & +0.173\,DU \\ & \quad (0.071) \end{aligned}$$

$R^2 = 0.717$; D. W. = 1.418 ; BL(12) = 16.67

$$\begin{aligned} \log(P_t) = & \ 1.052\log(M_t) - 0.371\log(Y_t) - 0.082\,IR_t + 0.024\,T - 0.126\,DU \\ & \quad (0.031) \qquad (0.063) \qquad (0.017) \qquad (0.002) \quad (0.034) \end{aligned}$$

$R^2 = 0.998$; D. W.= 1.989 ; BL(12) = 13.59

Notes : All the regressions included a constant and seasonal dummies; T is a time trend. DU is a break dummy allowing for a break level of the equation at 1973. 4; The figure in brackets are asymtotic standard errors; BL(12) is the Box-Ljung Pierce statistic for 12th order serial correlation.

The coefficients estimated by this procedure in each case are generally quite similar to the OLS estimates. In each estimates of the money growth equation the interest and unemployment rate variables exert a positive and significant influences as does the twice lagged money growth term: the once money growth term is less significant and its sign switches; the same is true of the inflation and export variables; the government deficit variable appears to exert a generally significant and positive influence on money growth. The explanatory power of the equation is much the same as before. The D. W. statistics shown in the Tables (which, as I have said before, are biased in the presence of a lagged dependent variable, so not weight can be given to it) and the Box-Ljung Pierce statistics computed for up to 12th order autocorrelation were all consistent with zero serial correlation.

The results in Tables 3 and 4 all continue to support the prediction that unpredictable monetary growth exerts a significant negative effect on the unemployment rate under this efficient estimation procedure, and once again the coefficients are generally of the same size. The estimates of the price equation also seem to support the findings of the OLS procedure. The log of the quantity of money generally has a coefficient of unity in the price equation but current and monetary shocks exert a negative influence on the price level. For all the equations I computed the Box-Ljung Pierce statistic for up to 12th order autocorrelation and for no case could the null hypothesis of zero serial correlation be rejected.

I turn now to a set of tests on the restrictions implied by the different version of the model estimated above. The type of model I have been estimating incorporates two sets of restrictions which can be separately tested.[8] These restrictions are (i) those implied by rationality of expectations; and (ii) those implied by what is known as structural neutrality - in this context the hypothesis that anticipated monetary growth has no real effect.

Because the number of varibles and lags in my model is quite large, I shall explain the tests of these two hypothesis which I carried out using just the version of the model which consists of equation (3.4) and (3.6). The extension to the other models is straightforward though algebraically cumbersome.

Consider a model which consists of the money growth equation (3.5) and a modified version of the unemployment equation (3.6) in which anticipations of money growth are found rationally so that predicted money growth equals $\alpha_0 + \sum_{i=1}^{10} \alpha_i X_{it-1}$ and that anticipated money is allowed a significant role in influencing unemployment with the same number of lags as I have allowed for unpredictable money growth. This model can be written:

$$DM_t = \alpha_0 + \sum_{i=1}^{10} \alpha_i X_{it-1} + \varepsilon_{1t} \tag{3.5a}$$

$$\begin{aligned} UNR_t = \beta_0 + \sum_{j=1}^{8} \beta_{j+1} DM_{t-j} + \sum_{j=1}^{8} [[\beta^h_{j+1} - \beta_{j+1}][\alpha_0 + \sum_{i=1}^{10} \alpha_i X_{1it-j-1}]] \\ + \sum_{j=10}^{14} \beta_j X_{2jt} + \beta_j \varepsilon_{1t} + \xi_{1t} \end{aligned} \tag{3.6a}$$

where the coefficient β^h_{j+1} shows the influence of predictable monetary growth on the unemployment rate: if it is zero then jth lagged predictable monetary growth has no effect on the unemployment. The two equations (3.5a) and (3.6a) form a two-equation model which does not assume structural neutrality - since the β^h_{j+1} are not forced to zero, but does assume rational expectations since anticipated

[8] See Leiderman (1980) and Driscoll, Ford, Mullineux and Sen (1983a & b)

monetary growth in period t is assumed to equal $\alpha_0 + \sum_{i=1}^{10} \alpha_i X_{it-1}$.

Despite the relaxation of the restrictions imposed by the structural neutrality hypothesis equations (3.5a) and (3.6a) still constitute a model with cross equation restrictions. To see this, first notice that in the system consisting of equations (3.5a) and (3.6a) the number of parameters is 33: 11 αs, and 14 βs(β_0, and $\beta_2 - \beta_{14}$) and 8 β^hs.

Now write a version of the two equations in which the DM equation is written as an unrestricted function of the X_{it-1} variables and UNR is written as an unrestricted function of all the variables on the left hand side of equation (3.6a). This procedure to make the unrestricted function is slightly function complicated by the fact that some lags of DM appear twice in equation (3.6a): explicitly and as element of X_{2it-1}. Eliminating this double presence gives a two equation system.

$$DM_t = \pi_0 + \sum_{i=1}^{10} \pi_{1i} X_{it-1} + \eta_{1t} \tag{3.5b}$$

$$UNR_t = \pi_{20} + \sum_{j=1}^{8} \pi_{2j+1}\, DM_{t-j} + \sum_{j=9}^{13} \pi_{2j} X_{2jt-1} + \sum_{j=14}^{55} \pi_{2j} X_{3jt-1} + \eta_{2t} \tag{3.6b}$$

where the πs are freely estimated coefficients and the η_t terms are error terms. The X_{3j} terms contain all the lagged terms in equation (3.6b) not already included in (3.6a). There will be 42 such terms: the 5 variables other than lags of DM which influence monetary growth, each lagged 1 to 8, plus the two extra lags of monetary growth, DM_{t-9} and DM_{t-10}, not already included in the first summation term.

So, in this unrestricted version of the model, there are 67 coefficients to be estimated including the two constants. Clearly then the hypothesis of rational expectations - the only difference between the model represented by equations (3.5a) and (3.6a) and the model represented by equations (3.5b) and (3.6b) - variables in the two models is the same but the number of parameters is less when rational expectations is imposed. The restrictions can be formally tested by comparing the likelihoods obtained from the two models. In large samples twice the difference between the log of likelihoods of the two models is distributed as a chi-square with the degrees of freedom given by the number of restrictions.

If the test fails to reject the restrictions implied by rational expectations then a further test can be carried out for structural neutrality by rewriting equations (3.5a) and (3.6a) after setting all the β^h terms to zero. The result is a restricted version of the model formed from equations (3.5) and (3.6). This can be seen as a restricted version of the model formed from equations (3.5a) and (3.6a) with the new restrictions solely the result of the assumption of structural neutrality. They can be tested by comparing the likelihoods of the two equations (3.5a) and (3.6a). However, if the first of restrictions is rejected then strictly the second set cannot be tested since there is no valid model against which to test them.

In the light of this discussion it is possible to think of two versions of the combination of equation (3.5). The first is the fully restricted version where rational expectations and structural neutrality are imposed. The second is where the rational expectations restrictions are imposed but the structural neutrality restrictions are not. And the third is where neither the rational expectations nor the structural neutrality restricions are imposed. For any combination of the equation I denote the likelihood obtained under the first set of restrictions by the index '*c*', the likelihood obtained under the second set of restrictions by the index '*b*' and the likelihood obtained under the neither set of restrictions by the index '*a*'. For each model it is possible to test each set of restrictions separately and to test them jointly as explained in the previous two paragraphs.

In Table 5 I present in the column (i) the log likelihoods, L of various models; in column (ii) the number of parameters in that model; in column (iii) under the heading L. r. the results of various tests of one model against one another: the test statistic shown in each case is minus twice the difference in the log likelihoods of the two models. This, as I have said, is distributed as a chi-square variate with degrees of freedom given by the difference between the number of parameters in the two models: the degrees of freedom are given in column (iv) under the heading d. f.; in column (v) an adjusted likelihoods ratio test which equals the unadjusted statistic shown in column (iii) times an adjustment factor to take account of the size of the sample.[9]

Models 1a-c are based on joint estimation of the DM and UNR equation, equation (3.5) and (3.6). The likelihood ratio test statistics for these models are supportive of the model. The rational expectations restrictions cannot be rejected at the 5% level; nor does the (valid) test of the structural neutrality restrictions lead to a rejection of these restrictions; and the test of the joint set of restrictions produces a likelihood ratio test statistic well below its critical 5% level suggesting that the restrictions cannot be rejected.

Models 2a-c are based on joint estimation of the DM, UNR and P equations. In this model the case for rejecting the rational expectations restrictions is again marginal; assuming the validity of the test the structural neutrallity restrictions are marginally rejected; nor can the joint set of restrictions be rejected.

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[9] Attfield, Demery and Duck (1981b) and Demery, Duck and Musgrave (1984) apply this adjustment factor to a similar problem citing Sims (1980). The adjustment factor k is defined as $(T-n/m)/T$ where n is the number of parameters, T the number of observations where T is 73 for all the joint estimation with certain restrictions.

[Table 5] Likelihoods and Likelihood Ratio Tests of Various Models

Equations	Restrictions (i) None	(ii) RE	(iii) RE & SN
(3.5) & (3.6)	Model 1a	Model 1b	Model 1c
(3.5) & (3.6) & (3.7)	Model 2a	Model 2b	Model 2c

Likelihood and Likelihood Ratio Tests

Model	Likeli-hood L (i)	No.of Coeffs (ii)		Ratios L.r. (iii)	d.f (iv)	Adj.L.r. (v)	Tests Critical X^2 5% (vi)	1% (vii)
1a	251.50	67	1a v 1b	40.16	32	21.74	46.17	53.45
1b	231.2	35	1b v 1c	8.98	10	6.82	18.31	23.31
1c	226.93	25	1a v 1c	49.13	42	26.58	58.11	68.74
2a	493.17	125	2a v 2b	164.79	81	70.73	103.0	113.5
2b	410.77	44	2b v 2c	23.62	10	18.87	18.31	23.21
2c	398.93	34	2a v 2c	188.47	91	80.89	114.2	125.3

Notes: RE = rational expectations; RE & SN = rational expectations & structural neutrality; L.r = likelihood ratios; d.f = a degree of freedom; Adj. L.r. = adjusted likelihood ratios.

Ⅳ. CONCLUDING REMARKS

In this paper I have applied the rational expectations equilibrium model for the Korean economy which sees the source of the business cycle as monetary shocks and which assumes price flexibility and rational expectations. The major departure of this model is marked from the idea that is involved in Lucas's (1972) model. He was concerned about the prices of goods which individuals can observe. They confuse movements in the overall price level with movements in relative prices. An unpredictable inflation leads individuals to infer that the relative price of the goods they produce is temporarily high, which induces them to increase the quantity supplied.

The Lucas's model is developed by Barro (1977, 1978) whose model emphasises the role of imperfect information regarding monetary surprises and has primarily been applied to industrialised countries. I have tested the Barro-type model using quarterly data for Korea which is estimated by the two-step OLS and non-linear joint estimation procedures. The OLS estimates appear to confirm a number of its predictions - notably of the role of monetary shocks in generating business cycles. Further testing of this model - notably of its rational expectations and structural neutrality restrictions - suggests that its model is also satisfactory in explaining the fluctuations of unemployment and price level. And these fluctuations are independent of the response to the systematic or predictable monetary policy.

In this respect it is suggested that my overall findings for Korea are favourable to the equilibrium model in which monetary shocks generate the business cycle but not real shocks cause it. Although the rational expectations equilibrium model has some validity for Korean economy it is noted that there are other possible routes of the propagation mechanism such as the channel of the inventory shocks which may contaminate dynamics with the rational expectations and market clearing.

DATA APPENDIX

Definition and Sources

1. The Nominal Quantity of Money, M

This is defined as M2, that is M1 plus quasi-money according to the conventional definitions of the IMF. The data are quarterly, seasonally unadjusted and run from 1970(1) to 1990(4). They are taken from the January Issue of the IMF publication International Financial Statistics, lines 34 and 35.

2. The Real Government Deficit, GD

This is defined as the nominal value of the government budget deficit deflated by the price level and multiplied by 1000. The nominal value of the deficit was taken from the IMF publication International Financial Statistics, (various issues), line 80 entitled government budget deficit. A deficit implies a negative value, a surplus a positive value. The data are quarterly, seasonally unadjusted and run from 1970(1) to 1990(4)

3. The Nominal Interest rate, IR

This is defined as the interest rate on time deposits taken from various issues of the Bank of Korea's Monthly Bulletin. The data are quarterly, seasonally unadjusted and run from 1970(1) to 1990(4).

4. The Real Value of Exports of Goods and Service, EX

This is the value of the exports of goods and services taken from the Bank of Korea's natioanl accounts 1990 issue, Table 2-1. This source provides data only from 1970(1) to 1989(4). The data from 1990(1) to 1990(4) were taken from the IMF publication, International Financial Statistics, (September 1991 issue), line 9c entitled Exports of Goods and Services. The data are quarterly, seasonally unadjusted and run from 1970(1) to 1990(4).

5. The Unemployment Rate, UNR

This is defined as the percentage of the workforce unemployed and is taken from the United Nations publication, The Monthly Bulletin of Statistics, various issues Table 8. This source provided only monthly data from 1970 to 1990, which were seasoanlly unadjusted. The data used are the average of 3 months data.

6. The Price Level, P

This is defined as the GNP deflator multiplied by 100. The GNP deflator is taken from the 1990/91 issue of the Bank of Korea publication National Accounts. The data are seasonally unadjusted and run from 1970(1) to 1990(4).

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