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Dynamic Econometrics and Policy Analysis

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A thesis submitted for the Degree of Doctor of Philosophy

September 2001

UNIVERSITY OF SOUTHAMPTON

ABSTRACT

FACULTY OF SOCIAL SCIENCES
ECONOMICS

Doctor of Philosophy

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The objective of this thesis is to evaluate the effects of policy changes on economic activities in Korea. For this purpose, particular emphasis is placed on the aspects of econometric modeling. As the basic methodology, the research applies the dynamic modeling approach which combines two current extreme approaches - the theory-based structural approach and the data-based VAR approach - and focuses on formulating statistically valid empirical models through extensive specification tests with particular attention to the critics of Lucas (1976) and Sims (1980), both of which are the damaging criticisms of the current approaches. These two issues are formally tested in this study rather than assumed.

Econometrically and empirically, the thesis contributes to the existing literature. Econometrically, the thesis extends the current LSE methodology in a more practical way. First, the thesis empirically demonstrates that care is required to formulate marginal models for the test of super exogeneity in identifying observational equivalence in ECM type models, if the effects of regime changes on the constancy of marginal models do not exhibit substantial changes, because of the possibility of 'spurious' non-constancy in marginal models. Second, the thesis shows how the test of weak exogeneity can be used to identify policy actions within a VECM. Third, the thesis questions why currently available macroeconometric models separately focus on evaluating the effects of transitory and permanent policy changes in a divergent way and argues that both changes should be evaluated jointly within a model. Possible econometric advantages with this approach are additionally discussed in the context of the invariant property of the underlying model and the usefulness of obtaining data information through both parameter spaces and error terms. As the basic approach, the thesis extends Hendry and Mizon's (1998) work into a structural multivariate framework by formally incorporating the usefulness of weak exogeneity tests in identifying policy actions.

Empirically, to investigate the effects of policy changes, the study sequentially asks three questions. The questions are whether the demand for money is stable even under regime changes, whether policy variables measuring policy actions are exogenously determined, and how and whether transitory and permanent changes in policy affect the economy. The first two questions are necessary conditions for the last question to be effective. The main findings to these questions are that: the stability of the demand for money has not been broken down by financial deregulation; the stock of M2 money which has been used as the main policy variable is exogenously determined; and both transitory and permanent changes in policy substantially affect the economy. An unanticipated transitory shock on nominal money significantly affects output and prices, but in the case of an anticipated permanent change, 'real' money rather than 'nominal' money positively affects output. This implies the importance of price stability as a precondition for achieving the sustained economic growth of the economy.

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Acknowledgements

I wish to acknowledge grateful thanks to my supervisors Maozu Lu and Jan M Podivinsky for their careful guidance, encouragement, and generous advice during the writing of this thesis. I owed a tremendous debt to them. They have carefully read the whole draft and provided invaluable comments and suggestions. I also wish to thank John Aldrich and Grayham Mizon for their extremely helpful comments, suggestions, and criticisms on the various aspects of this work. The attitude of my econometric modeling in this work has been largely influenced by the four people mentioned above. My thanks also go to Stephen Hall who gave many valuable suggestions and comments.

Finally, I thank to John Driffill, Katarina Juselius, and Akos Valentinyi for their stimulating suggestions and discussions. Several parts of this thesis have benefited from their comments.

Chapter 1

Introduction

The main purpose of this thesis is to evaluate whether and how monetary policy affects the economy in Korea. To investigate these kind of issues, three models have been widely used in the literature: traditional structural models, vector autoregression (VAR) models, and non-quantitative narrative methods. The structural approach mostly developed by Cowles Commission econometricians specifies underlying models on the basis of economic theory, applies standard statistical procedures to estimate the parameters of models, and draws inferences about parametric and economic relationships between variables. In general, the studies following this line of research seek to corroborate the underlying theory rather than to exploit information from data. The main advantages of this approach are that the adopted models are very useful in measuring and analyzing structural interrelationships between economic variables and across different sectors of the economy. In contrast to the traditional structural approach, the VAR approach proposed by Sims (1980) eschews 'incredible' restrictions on parameters. Instead, the empirical studies in this line of research attempt to find some regularities from data by endogenizing all relevant variables, with 'minimal' identifying restrictions (see Canova (1995) for an excellent recent survey). Thus, the role of data rather than theory is largely emphasized in this atheoretic reduced-form approach. The 'non-quantitative narrative' method proposed by Romer and Romer (1989) identifies policy actions by examining the minutes and policy records published by monetary authorities and uses the identified information to investigate whether policy changes affect the economy. An advantage of this approach is that the direction of causation between monetary factors and real economic developments is uniquely identified.[1]

However, the empirical evidence derived from these three approaches on the effectiveness of policy is still inconclusive, even though a number of different sophisticated models have been used for different countries. An obvious reason of the difficulty to have convincing evidence lies in the intrinsic limitation of econometric methods to derive a simplified model which uniquely represents all the behavior of large numbers of individual agents all trying to achieve their own objectives. Another reason of the uncertainty of the empirical results arises from the fact that empirical models are constructed for different purposes, at different levels of aggregation, and with different relative weights on theory and data. Thus, it might be natural to have different empirical results. If so, an immediately following question is which evidence can be regarded as coherently characterizing the salient features of data. A plain truth is that for an empirical study to be reliable, at least the underlying model should not be mis-specified and all relevant hypotheses should be formally tested in all possible ways. Unfortunately, the approaches mentioned above suffer from this simple axiom, even though their contributions to the developments of current macroeconometric methods are enormous. First, the structural approach does not question the statistical adequacy of the underlying models. In this approach, econometric models are assumed to be correctly specified by economic theory, so the approach does not say much about the processes of specifying baseline models. Instead, the structural approach at-

[1] Recently, this approach has been criticized by many macroeconomists, because of a possible subjectivity of reviewer and the endogeneity of policy changes (see Hoover and Perez (1994), Bernanke *et al.*(1997), and Leeper (1997)).

tempts to verify economic theory in terms of signs, magnitudes, and significance of estimated parameters. However, it should be noted that no matter how elegant and sophisticated an economic theory is, there is no guarantee that the model implementing it is statistically well specified (see Hendry (1980), Leamer (1983), and Mizon (1995)). Equally, this critic on the lack of statistical tests in specifying econometric models is also applied to the data-based VAR approach. Most of the recent studies using VAR models assume that the baseline models are well specified, and seldom report the statistical tests of specifying models. Only lag selection criteria and identification issues are regularly discussed. However, it is well-known that VAR models are also very sensitive to the selection of lag lengths, overparameterization, and the treatments of data stationarity (see Spencer (1989), Todd (1990), Phillips (1998) and Abadir *et al.* (1999)).

Second, another damaging criticism to these approaches is the Lucas (1976) critique that entails the near-impossibility of constant conditional models when policy regimes change. Lucas (1976) argues that since economic agents rationally anticipate policy changes in the future, any econometric models which do not explicitly incorporate the rational behavior of the agents suffer from variant parameters and inaccurate predictions. Since the logic of this critic is very reasonable, the practical applicability should be formally tested. However, most of the empirical studies applied structural and VAR models rarely examine the invariant property of the underlying econometric models. The invariance is simply assumed with the justification either that policy regime changes are rare (Sims, 1982a) or that economic agents can't perceive policy changes. This *ad hoc* assumption is not desirable, since merely assuming that an econometric model is statistically valid without formal tests may lead to inaccurate policy analysis and forecasting.[2] Third, the empirical results obtained from the three approaches largely depend on *a priori* assumed identification of exogenous variables. Sims

[2]See Ericsson and Irons (1995) for an extensive survey on the empirical evidence of the critique. They argue that there is little evidence of the critique being empirically relavant.

(1980) strongly criticizes the traditional structural approach by arguing that the imposition of large numbers of restrictions is 'incredible', since they do not arise from economic theory but simply from the need of a modeler to have enough restrictions to secure identification. Ironically, the critique on this strong apriorism is also applicable to VAR models which explicitly place arbitrary or theory-based restrictions on either the covariance matrix of residuals or dynamic multipliers without statistical tests.[3] Indeed, an advantage of the VAR approach is minimal identifying assumption. Unfortunately, the empirical results are very sensitive to this weak-form apriorism (see Cooley and LeRoy (1985) and Cooley and Dwyer (1998)).

Considering the above deficiencies of the current popular methods, this thesis emphasizes valid model formulation in investigating the effects of policy changes on the economy and attempts to deliver statistically reliable analysis, by focusing on the issues of policy identification and parameter invariance under regime changes. To do this, the thesis applies the LSE methodology which combines structural and VAR approaches. The advantages of this approach over the conventional approaches are that: first, the methodology attempts to parsimoniously characterize data with the guidelines from economic theory and to provide a statistical model against which other models can be evaluated, so the drawbacks of data mining and theory calibration are not applied to this approach: second, the methodology is mainly concerned with *theory evaluation* rather than *theory confirmation*. In this context, the approach emphasizes rigorous model evaluation to have a reliable empirical result through extensive specification tests in all possible ways, before an economic interpretation is applied, and so at least seeks to avoid the critiques mentioned above by placing no *a priori* assumptions. Recently, a series of studies in this line have discussed the usefulness of the approach in analyzing policy-related issues. These include Hendry and Doornik (1994), Banerjee

[3]See, for examples, Bernanke (1986), Sims (1986), Blanchard and Quah (1989), and King *et al.* (1991) for the structural VAR approach.

et al. (1996), Ericsson *et al.* (1998), and Hendry and Mizon (1993, 1998, 1999a).

There is no doubt that the LSE methodology has already reached a considerable level in its theoretical developments for econometric modelling. Nonetheless, it is important to recall that the econometric concepts and tools in this approach are mainly designed for empirical purposes, so the usefulness of the approach depends on how the methodology can be practically used for such purposes. It has been recognized that there is still a gap between the theories and their empirical practice. This thesis attempts to fill the gap by extending the methodology in a more practical way and by applying the extension to real data, with a concomitant belief that extending along this line would lead to better performance. In this context, the thesis contributes to applying the current methodological tools of the LSE approach into more practical ways. More specifically, the thesis empirically demonstrates that if the effects of regime changes are not substantial, there is a possibility of 'spurious' structural breaks in formulating marginal models. In this special case, since the testing power of super exogeneity is not reliable, the thesis recommends to conduct both encompassing and super exogeneity tests in identifying observational equivalence in equilibrium correction (ECM) type models.[4] The thesis also shows how the concept of weak exogeneity can be usefully applied to identifying exogenous variables block-recursively within a vector equilibrium correction model (VECM) and rules out *a priori* assumed identification, being commonly used in the current VAR literature, by showing that the identified exogenous variables are ordered firstly for orthogonal impulse response functions. Finally, the thesis questions why currently available macroeconometric models separately focus on evaluating the effects of transitory and permanent policy changes in a divergent way and argues that both changes should be evaluated jointly within a model. Possible econometric advantages with this approach are

[4] In the literature, the term of 'equilibrium correction' is preferred to 'error correction', since the latter term only works within regimes (that is, unchanged equilibria) and does not correct towards a changed equilibrium (see, for a more detailed discussion, Hendry (1995)).

additionally discussed in the context of the invariant property of the underlying model and the usefulness of obtaining information on data generating processes (DGP) through both parameter spaces and error terms. For the practical application, the thesis extends the work of Hendry and Mizon (1998), which provides a reduced-form framework to assess the effects of policy changes and applies it to a bivariate model, into a structural multivariate model by formally incorporating the usefulness of the concept of weak exogeneity in identifying policy stance.

To examine the what-if type effects of changes in exogenous policy variables on the economy, the thesis sequentially asks the following three questions: (1) whether the demand function for money is stable even under regime changes, (2) whether policy variables measuring policy stance are exogenously determined, and (3) whether and how transitory and permanent changes in policy affect the economy. The reason for investigating these questions sequentially is that in order for policy changes to be effective, the demand function for money should be stable and policy variables such as money stocks or interest rates should reflect genuine policy changes, not endogenous responses to changes in the economy. When these necessary conditions are satisfied for the effectiveness of policy, the last issue can be legitimately analyzed. Thus the approach applied in this thesis is different from the previous studies which assume such necessary conditions *a priori*. The thesis formally tests those conditions with an expectation that the empirical results in this line be more robust than those in any other approaches.

The general structure of the thesis is organized as follows. Chapter 2 reviews the methodological framework of the LSE approach that underlies our applications in the subsequent chapters. The reviews include the representations of DGPs, the general form of empirical models, and the issues of model evaluation and design. The chapter further extends its discussion to the main advantages of the general-to-specific modelling strategy which is the major tool in the LSE approach. Following the general review of the methodology, Chapter 3 supplementarily explains several key concepts in the LSE methodology for the analysis of policy-related issues. The concepts considered are exogeneity, cointegration, and co-breaking. The discussions focus on how these concepts are important in formulating econometric models and are useful for policy analysis and forecasting. The relevant particular interests are the formulation of valid models and the critiques of Lucas (1976) and Sims (1980). The thesis then applies the ideas and techniques of the LSE methodology to analyzing the effects of policy changes in Korea. The discussion in Chapter 4 is related to the first question mentioned in the preceding paragraph. The chapter investigates whether the stability of the demand function for money has been changed or not under financial liberalization and identifies the observational equivalence of the underlying model. In doing so, the chapter first derives a simple ECM model using a general-to-specific approach and shows that the estimated parameters are constant even under regime changes. To examine the invariance and observational equivalence of the estimated model, the chapter performs the test of super exogeneity by formulating nonconstant marginal models for the DGPs of conditioning variables. Finally, to reinforce the reliability of the model, the chapter compares the model with a forward-looking model formulated on the basis of a multi-period quadratic loss function and rational expectations. In comparing these two models, the encompassing tests suggested by Mizon and Richard (1986) are applied. Methodologically, this chapter demonstrates that if the effects of regime changes are not substantial, the power of the tests of super exogeneity largely depends on the formulation of marginal models, and recommends, in this special case, to conduct both tests of encompassing and super exogeneity complimentarily in discriminating observational equivalence in ECM type models.

Chapter 5 shows how the concept of weak exogeneity can be usefully applied to identifying policy stance within a cointegrated VAR system. Particularly, the chapter focuses on the usefulness of Johansen's (1992) testing procedure for the identification of exogenous variables that should be ordered firstly in conducting orthogonal impulse response analysis. Then, the chapter examines the empirical feasibility of the approach using Korean macroeconomic data. In order to identify policy actions in a robust way, the chapter divides the stock of M2 money into outside and inside money and examines whether the possible endogeneity of inside money, as shown by King and Plosser (1984), is supported or not. Finally, the chapter discusses the main empirical findings by figuring out the differences from those in the current literature and emphasizes the importance of a statistical identification of policy actions. Chapter 6 questions why current macroeconometric models separately focus on evaluating the effects of transitory and permanent policy changes in a divergent way and argues that both changes should be evaluated jointly within a model. The necessity of the joint examination is discussed with econometric advantages that the 'invariant' property of the underlying model can be formally examined and information on data can also be obtained from both parameter spaces and error terms. For the application of the argument, the chapter extends Hendry and Mizon's (1998) work, which provides a reduced-form framework to assess the effects of policy changes and applies it to a bivariate model for the UK experience, into a structural multivariate framework by formally showing the usefulness of weak exogeneity in identifying policy shocks. Then, the empirical feasibility of the approach is examined using Korean data. In a structural dynamic model complementing the limitations of conventional structural and VAR models, impulse response techniques and the concept of co-breaking (Clements and Hendry, 1999) are applied to evaluate the effects of transitory and permanent changes, respectively. The importance of a statistical identification of policy stance and the invariance of estimated parameters is particularly emphasized in this chapter. Finally, Chapter 7 draws the conclusions and discusses some policy implications obtained through the research. Future directions of policy analysis in this line are also discussed with the methodological limitations in this work.

Chapter 2

Methodology of Dynamic Econometric Modeling

2.1 Introduction

Dynamic econometric modelling involves an attempt to match the lag reactions of a postulated theoretical model to the autocorrelation structure of the associated observed time-series data (see Hendry *et al.* (1984)). In this chapter we review the LSE methodology as a dynamic econometric modelling approach.[1] The initial developments of the LSE approach are primarily associated with Sargan (1964) and many of his students (see Mizon (1995) for an extensive survey on the historical development of this approach). In this approach, the economic structure characterizing the interrelationships between variables is neither assumed to be known *a priori* nor deemed to be unidentifiable. It is obtained through a series of simplications, statistically acceptable but guided by economic theory, from a general model formulated from numerous sources of information. Thus, in this approach, economic theory and data continuously interplay to find a good empir-

[1] Alternatively, the LSE approach is called as the general-to-specific modelling approach. Throughout this research we use both words interchangeably.

ical model. This distinctive feature of the 'middle' ground approach differentiates the methodology from the theory-driven structural approach and the data-driven VAR approach (Gilbert, 1986). As the main modelling strategy, the methodology prefers the general-to-specific approach to the specific-to-general approach. In this context, the approach recommends to start from a general model which contains all the variables the investigator regards as candidates for possible inclusion and passes the standard battery of diagnostic tests at reasonable significance levels. Sequential reductions from the general model are the main tool to derive a final model. The final model obtained in this way should not be worse than the general model in its fitting, but should encompass the general model and rival models. When these statistical properties are satisfied, the obtained model is then expected to be coherent with information available from economic theory and data.

This chapter is organized as follows. In Section 2, we discuss the general structure of the data generating processes (DGPs) which represent the joint probability distribution of all relevant variables. To represent the DGPs, a VAR model is provided as the underlying econometric model. Section 3 discusses several criteria for the evaluation and design of dynamic econometric models. In Section 4, the general-to-specific modelling approach advocated by the LSE methodology is discussed in comparison with the specific-to-general modelling approach. Finally, Section 5 concludes the chapter.

2.2 Statistical system

The statistical system is the Haavelmo distribution defined by specifying the variables of interest, their status, their degree of integration, data transformations, the history of the process, and the sample period (Hendry, 1995). In short, the system entails DGPs which incorporate all the necessary ingredients to characterize economic data. Consider the joint density function $D_X(\cdot)$ for T observations on a vector of n variables $x_t = (x_{1t}, \ldots, x_{nt})'$ for the complete sample X_T^1 :[2]

$$D_X(X_T^1|X_0,\theta), \quad \theta \in \Theta \subseteq \Re^s, \tag{2.1}$$

where X_0 denotes a set of initial conditions and θ denotes the paramter vector lying in the parameter space Θ which is a subset of s-dimensional real space $\Re^s$. Under the assumptions that the occurrence of one event is independent of the occurrence of the other event and that observations are conditionally generated by the past information, the data density function (2.1) can be sequentially factorized as follows:

$$D_X(X_T^1|X_0,\theta) = \Pi_{t=1}^{T} D_x(x_t|X_{t-1},\theta), \tag{2.2}$$

where $X_{t-1} = (X_0 : x_1, x_2, ..., x_{t-1}) = (X_0 : X_{t-1}^1)$. The second term in (2.2) shows that the interdependent joint distribution of $\{x_t\}$ is factorized into the product of T conditionally independent components $D_x(x_t|X_{t-1},\theta)$. The importance of this factorization is that; complicatedly intercorrelated DGPs are reduced to the one expressed in terms of mean-innovation errors; and that, because the innovation errors are constructed from data, it is generic and involves no loss of information in the sense that we can precisely recover the original data from the innovation errors and sequentially conditioned means (see Hendry, 1995). Since most econometric models involve describing conditional submodels suggested by economic theory, the data densities $D_x(x_t|X_{t-1},\theta)$ can be further factorized by partitioning $x_t' = (y_t' : z_t')$ and $X_T^{1\prime} = (Y_T^{1\prime} : Z_T^{1\prime})$ correspondingly as:

$$\Pi_{t=1}^{T} D_x(x_t|X_{t-1},\theta) = \Pi_{t=1}^{T} D_{y|z}(y_t|z_t, X_{t-1},\lambda_1) D_z(z_t|X_{t-1},\lambda_2), \tag{2.3}$$

[2] X_T^1 can be partitioned as $X_T^1 = (X_{t-1}^1 : x_t : X_T^{t+1})$ where X_{t-1}^1, x_t, and X_T^{t+1} represent the information sets of relative past, present, and future data, respectively.

where (λ_1, λ_2) is an appropriate reparameterization of θ with the parameters of interest ϕ. Note that $D_x(\cdot)$ is not variant with these parameter transformations. $D_{y|z}(y_t|z_t, X_{t-1,}\lambda_1)$ represents a conditional density describing the economy under consideration, whereas $D_z(z_t|X_{t-1,}\lambda_2)$ represents a marginal density describing the DGPs of exogenous variables. The exogeneity conditions for valid conditional transformations should be ascertained by examining the properties of weak, strong and super exogeneity of the underlying data (see Engle *et al.* (1983)).

2.3 The econometric model

The system (2.3) can be characterized by many types of linear or nonlinear models within which the associated hypotheses are tested. Although non-linear models are not ruled out from the system, the LSE approach prefers to use a VAR-type linear model as the 'baseline' model for formulating an econometric model (see Hendry and Richard, 1983). A simple VAR may be written as:

$$x_t = \sum_{i=1}^{p} \Pi_i x_{t-i} + \epsilon_t, \qquad \epsilon_t \sim IN(0, \textstyle\sum), \tag{2.4}$$

where $IN(0, \sum)$ denotes an independent normal density with mean zero and covariance matrix $\sum$. This model corresponds to the DGPs (2.2) and entails a time-dependent representation of endogenous variables in terms of available information. The model as a *closed*-form dynamic model could be formulated on the basis of simple lag selection criteria such as the Akaike Information Criterion or the Schwarz Bayesian Criterion. However, the initial specification is very important for the success of any empirical analysis since the dynamic system is the 'maintained' statistical model (Hendry, 1995). From (2.4), alternative *open* models can be derived by conditioning endogenous variables on exogenous variables which represent the marginal distributions of DGPs. The conditioning is conducted through a series of reductions on the basis of the tests of non-causality, exogeneity, and invariance. It is important that in order for the conditioning to be valid, at least conditioning variables should be weakly exogenous for the parameters of interest. An importance of (2.4) in econometric modelling is that the model is the baseline model within which empirical models are developed by imposing a structure, which intends to extract autonomous, parsimonious relationships between variables but is interpretable by economic theory. The identification of the structure normally takes the form of over-identifying restrictions that are guided by statistical tests and economic theory. A final model obtained in this way is expected to explain the underlying economic theory and to be congruent with data (see Spanos (1986) and Hendry and Richard (1983)). To satisfy the latter condition, the final model should have (1) a homoscedastic innovation error process, (2) weakly exogenous, conditioning variables, and (3) constant parameters.

Another alternative reformulation of (2.4) is the following VEqCM form:

$$\Delta x_t = \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \Pi x_{t-1} + \epsilon_t, \tag{2.5}$$

where $\Gamma_i = -\sum_{j=i+1}^{p} \Pi_j$ and $\Pi = \sum_{i=1}^{p} \Pi_i - I$ with $i = 1, ..., k-1$. This model is reparameterized from the unrestricted VAR (2.4), independently of whether the variables x_t are I(0) or I(1) (see Johansen (1988) and Hendry (1995)). If rank $(\Pi) = n$, the vector process $\{x_t\}$ is stationary; if rank $(\Pi) = 0$, the model is appropriate in first differences without the ECM term; if rank $(\Pi) = \gamma < n$, then $\Pi = \alpha\beta'$ where α and β are $n \times \gamma$ matrices of rank γ. In the last case, $\beta' x_t$ includes γ cointegrating I(0) relations, given the assumption of $I(1)$ data, so that (2.5) is rewritten as a restricted I(0) representation:

$$\Delta x_t = \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \alpha(\beta' x_{t-1}) + \epsilon_t. \tag{2.6}$$

Since all variables included are $I(0)$, the equation is stationary. Hence, this model can be used as an alternative statistical benchmark model for the derivation of a structural econometric model, if data are nonstationary. Deterministic components, such as intercepts or seasonal dummy variables, may be included in (2.6) and influence the resulting distributions. To formulate an econometric model, the restrictions on the short-run dynamics are mainly conducted on the basis of statistical tests, whereas the ECM term is formulated by taking account of economic theory which explains the long-run equilibrium relationship between variables. This is a unique feature of the ECM model which explicitly incorporates information about both data and economic theory (see Hylleberg and Mizon, 1989).

2.4 Model design and evaluation

In the previous section we have explained a VAR model as a particular form characterizing the underlying DGPs. However, we haven't discussed in detail how to evaluate and develop a reliable econometric model which is valid in terms of statistical tests and economic interpretations. The relevant issues are further discussed in this section. For this, our discussions are divided into two main categories: *congruence* and *encompassing*. Since the LSE methodology emphasizes that in order for an econometric model to be valid, the model should be congruent with the relevant available information and encompass alternative rival models, these two issues are central to this approach (see, for more details, Mizon (1995) and Bontemps and Mizon (2001)).

Congruence

Congruence is the property of a model that has fully exposited all the information implicitly available once an investigator has a set of variables to be used in modelling (Bontemps and Mizon, 2001). Thus the concept is related to model coherence with relevant available information. Following Hendry (1995), models are said to be congruent with their unknown DGPs, if several necessary conditions are satisfied; such as homoscedastic innovation errors (data coherence), weakly exogenous conditioning variables for the parameters of interest, constant and invariant parameters of interest, theory-consistent identifiable structures (theory consistency), and data admissible formulations on accurate observations (data admissibility). All these conditions are linked to particular types of available information: theory information, sample information, measurement information, and information from rival models. Detailed explanations about the relations are found in Hendry and Richard (1982).

Theory information entails that empirical models at least should be congruent with low-level theories and should also provide a framework within which the hypotheses of high-level theories can be tested.[3] It is not required that econometric models conform in all aspects to very detailed statements of theories, but required that models include a set of variables associated with the underlying theories. This requirement is arisen from the importance of the mutual interplay between economic theory and empirical evidence. The criterion relevant to this information is theory consistency. Sample information involves the requirement that econometric models should be congruent with observed random variables and their structure. Information in this type is closely associated with several important econometric concepts, such as exogeneity, Granger-causality and structural invariance. The sources of sample information can be categorized into three mutually exclusive information sets: relative past, present, and future sample information. *Relative past sample information* involves the lagged values of ob-

[3]Theories can be distinguished in terms of levels: low-level theories and high-level theories. In low-level theories, all econometric models result from the use of economic considerations, such as the choice of variables to include in the model, and of the functional forms to characterize the relationships between them. High-level theories include theories that imply very tight and specific forms for econometric models (such as models embodying rational expectations, intertemporal optimization, or Euler equation formations) and so embody testable hypotheses rather than specifying essential characteristics for any econometric models of the phenomena being studied (see Hendry (1995) and Mizon (1995)).

served data. For example, if an econometric model is congruent with past sample information, its residuals should not be significantly autocorrelated. This is the case when the residuals have a homoscedastic innovation process and so there is no further information that can be exploited from the ignored information. Hence, testing for the congruency of a model with information from this source is equivalent to examine whether the model has a homoscedastic innovation error. *Relative present sample information* involves current values of samples. The congruency of an econometric model with the information in this type is required to validly condition the parameters of interest. The relevant testable hypothesis is that conditioning variables are weakly exogenous. *Relative future sample information* involves checking the uncertainty of models in the future. An econometric model is congruent with future sample information, if the parameters are constant and so do not suffer from predictive failure. However, the more fundamental issue is invariance which requires that the parameters stay constant irrespective of how the underlying marginal processes change. In this context, the testable hypothesis is invariance which can be examined by applying the tests of super exogeneity in the literature.

Measurement information involves how variables in econometric models are measured and what their specific properties are. More specifically, these include the units of measurement, functional forms, variable transformations, and constructed identities, etc.,. These types of information are very useful for the specification of econometric models, which can be used to generate observed data (in-sample) and future data (out-of-sample). The required criteria related to this source of information are data measurement accuracy and data admissibility. Finally, information from rival models is related to the congruency with the information contained in competing models. Since each model has its own individual information, possibly with different functional forms, which is congruent to sample information, it is necessary for an econometric model to encompass rival models in order to exploit their information. This involves checking the validity of reduction from a general model that nests all competing models. Hence the testable null hypothesis relevant to this information is whether an econometric model parsimoniously encompasses the statistical model which embeds all relevant models, including rival models.

Encompassing

Encompassing is associated with the question as to whether a chosen model is robust to the extended information set. Good examples of extended information are newly developed theories which imply extensions of the underlying information set to incorporate the empirical models associated with the new theories. A related concern with this further information set is whether the chosen model can encompass the new rival models. In this context, this criterion is different from the preceding parsimonious encompassing which emphasizes the property that a model should be an acceptable reduction from the very general model.

Finally, it should be noted that even if a finally chosen model satisfies all the criteria suggested above, the model can't be regarded as a 'correct' model, since the model is a simplified representation of complex reality with the data which is at present available. However, reality changes over time. When new information (over time, variables, and interventions) arrives, we have to revise and improve the model currently preferred. In this sense the LSE approach prefers a *progressive* modelling strategy in which a current model is abandoned if a new better model is found. An important implication of this strategy is that (invariant) 'structure' characterizing the fundamental relationships between economic variables is determined in a progressive way. This is one of the main differences of the LSE approach from the theory-driven approach in which economic theories define structure in a permanent way (Hendry, 1995).

2.5 General-to-specific modelling

In the previous section we have discussed the importance of congruence and encompassing in designing and evaluating econometric models. While there is no unique way to find a data congruent model, this section discusses the usefulness of the general-to-specific modelling approach preferred by the LSE methodology (see, for more details, Mizon (1977) and Hendry and Mizon (1978)). The general-to-specific approach starts from the most general model which nests all models representing competing theories, and simplifies it to a parsimonious, but economically interpretable, model. The simplication involves 'reductions' through a series of marginalization and conditioning and is sequentially tested down by checking whether the corresponding reduction is statistically valid at every stage. Since a general model is overparameterized, there are many potential paths of the simplication. In this way, several competing models can be derived from the baseline model. To reduce the proliferation of rival models, the LSE methodology recommends to select a final model by applying the principle of encompassing (see Mizon and Richard (1986) and Hendry (1995)). The final model selected in this way is regarded as a congruent model approximating well the underlying true relationships between variables.[4] However, the main problems with this approach are that; (i) data may be limited because of the number of parameters growing with the square of the number of variables and so quickly exhausting degrees of freedom[5]; (ii) there is no unique sequence for simplifying the underlying general model and so, many competing models which are data congruent can be formulated (even though the problems can be eliminated by encompassing tests, in the case that competing models mutually encompass, there is still no unique

[4] In the context of this line of research, recently Hendry and Doornik (1994) further discuss ten interrelated reasons for commencing econometric analyses of economic time series from the joint data density.

[5] Note that the initial starting model in the LSE approach is a VAR.

solution.); and (iii) individual tests may have large type II error probabilities if the overall sequence is not to have a high type I error (Hendry, 1979).[6]

On the other hand, the specific-to-general modelling approach starts with a very specific, simple model based on economic theory and results in a model that is more general than the initial model. The main emphasis in this approach is the verification of the underlying economic theories, in terms of signs, magnitudes, and significance of estimated parameters, under the assumption that estimated models are correctly specified. The advantages of econometric modelling in this line are that selected models are not needlessly general and are relatively coherent with economic theories. However, the main problems with this approach are that models are excessively presimplified with inadequate diagnostic testing (Leamer, 1974); and that every test is conditional on arbitrary assumptions which are to be tested later, and if these are rejected, all earlier inferences are invalidated whether reject or not reject decisions (Hendry, 1979). For the purpose of a simple comparison between these two approaches, an extreme case may well illustrate an advantage of the general-to-specific methodology over the specific-to-general methodology. In the former case, specification errors are limited to inclusion of irrelevant variables, provided that the initial model covers all competing theories to be considered, while those in the latter case are omission of relevant variables because this approach assumes that the existing model is valid and so is almost certain to encounter 'omitted variable bias' early.[7] It is well known that in econometric models, the problems with omitted variables errors are more serious than those from irrelevant variable inclusions, since the former produces biased and inconsistent parameters but the latter merely produces inefficient estimates.

[6]The LSE approach often leads to an overfitted equation. Under a fixed type I error, the inclusion of irrelevant variables produces a large variance which does not reject a relevant null hypothesis, even though the null hypothesis is not true.

[7]See Hendry and Mizon (1978) for the problems of common factor.

2.6 Summary

In this chapter we have reviewed the basic methodology of the dynamic modelling approach associated with the LSE methods in econometrics. A distinctive feature of this approach is that the methodology provides a flexible tool which allows the complex interactions of economic theory and data. The proponents of the approach mainly focus on formulating statistically valid econometric models by examining whether the statistical models are congruent with available information and by evaluating whether the statistical models encompass rival models. In this context, the approach particularly emphasizes the importance of testing and evaluating econometric models. This theme of the LSE methodology is well represented by the words recommended by Hendry (1980): *'if something is testable, test it'* and '*the three golden rules of econometrics are test, test, and test'.* As the main modelling strategy, the LSE methodology advocates a general-to-specific approach rather than a specific-to-general approach. While several researchers criticize that the methodology is simply 'a combination of back-ward and forward step-wise regression' (Leamer, 1985); involves a certain amount of 'data (lag) mining' (Cuthbertson *et al.*, 1992); leads to an 'overfitted' model (Hess *et al.*, 1998); and is open to the charge of 'measurement without theory'(Koopmans, 1947; Darnell and Evans, 1990), the approach is supported by White (1990) in large samples and Abadir *et al.* (1999) and Hoover and Perez (1999) in small samples. The theoretical developments of the approach in econometric modelling have already reached to a certain high level, and are still evolving and being refined. However, compared to its long history, little research in this line has been applied to analyzing macroeconomic policy-related issues, except the recent work by Banerjee *et al.* (1996), Ericsson *et al.* (1998), and Hendry and Mizon (1998). In the next chapter we review several key econometric concepts in this approach for policy analysis.

Chapter 3

Exogeneity, Cointegration and Co-breaking in Policy Analysis

3.1 Introduction

Following the brief review of the LSE methodology in Chapter 2, this chapter further discusses several important econometric concepts in this approach: exogeneity, cointegration, and co-breaking. In recent years, these concepts have brought a growing interest in the literature, particularly in the context of policy analysis and forecasting (see, Banerjee *et al.* (1996), Ericsson *et al.* (1998), and Hendry and Mizon (1998, 1999a)). Since the notions are central in the following chapters, we focus our discussions on characterizing the definitions of the concepts and on explaining how they can be used to analyse policy-related issues. *Exogeneity* in econometrics is defined as 'determined outside the system under analysis' (Koopmans, 1950). In econometric modelling, the issue arises when a subset of variables is analyzed by treating some variables as given. Following Engle *et al.*(1983), we consider three types of exogeneity - weak, strong and super exogeneity. Another important notion in the LSE methodology is *cointegration* that entails some stationary linear combinations between nonstationary time series. The concept defined by Granger (1981) and Engle and Granger (1987) is based on the work of Sargan (1964) and Davidson *et al.* (1978) and has been rapidly developed in the econometric literature over the last decade. The basic idea of cointegration is that even though some variables are nonstationary, linear combinations of those variables may eliminate the nonstationarity. Since cointegration is a statistical notion that entails the long-run behavior of economic time series, it plays an important role in linking economic theory to statistical dynamics in ECM models. This role makes cointegration be not separated from the LSE methodology which emphasizes the role of both data and theory in econometric modelling. Finally, *co-breaking* denotes some linear combinations of variables which remove deterministic shifts jointly (see, for more details, Clements and Hendry (1999)). The interdependence between economic variables with structural breaks - for example, policy and nonpolicy variables - provides a possibility of the existence of co-breaking. In the presence of regime changes, the concept is very useful in modelling statistically valid econometric models. Since parameters remain unchanged despite policy changes, co-breaking refutes the Lucas (1976) critique.

This chapter is organized as follows. Section 2 explains the definitions and properties of three types of exogeneity and discusses their role in analyzing economic policy and forecasting. Section 3 reviews the concept and testing procedure of cointegration, particularly by focusing on the Johansen (1988) procedure. Section 4 discusses the property of co-breaking and its usefulness in evaluating economic policy. A summary is provided in Section 5.

3.2 Exogeneity

Exogeneity has been one of the longest debated issues in the econometric literature. Since any econometric model inevitably involves conditioning some variables that are treated as 'given' for the explanation of the behavior of a subset of variables, valid conditioning is necessary for the formulation of an efficient econometric model without loss of relevant information. Mistreating a variable as exogenous when it is not leads to inefficient or inconsistent inferences and results in misleading forecasts and policy analysis. In this context, it is desirable to identify exogenous variables in a statistical way. Engle *et al.* (1983) dissatisfied with the conventional definitions of exogeneity, such as predeterminedness and strict exogeneity, propose three types of exogeneity: weak, strong, and super exogeneity. Briefly, weak exogeneity is related to the necessary conditions for formulating an efficient econometric model in which inferences about the parameters of interest can be conducted from a conditional density alone without loss of information; strong exogeneity is a combined concept of weak exogeneity and Granger non-causality, and is useful for conditional forecasts; and super exogeneity adds the requirement that parameters remain constant across policy changes to that weak exogeneity.[1]

Weak exogeneity The concept of weak exogeneity originally proposed by Richard (1980) and further analyzed by Engle *et al.* (1983) provides a sufficient condition for being able to conduct efficient conditional inferences, without loss of relevant sample information, by treating some variables as given. Thus, this notion is directly related to valid inferences and efficient estimations in formulating econometric models. Consider equation (2.2) $D_X(X_T^1|X_0,\theta) = \Pi_{t=1}^{T} D_x(x_t|X_{t-1},\theta)$ and partition $x_t' = (y_t' : z_t')$ and $X_T^1 = (Y_T^1 : Z_T^1)$, where y_t is an $n_1 \times 1$ vector of endogenous variables and z_t is an $n_2 \times 1$ vector of exogenous variables. If the conditional distribution of y_t given z_t is factorized without loss of information from the joint distribution of x_t :

$$\Pi_{t=1}^{T} D_x(x_t|X_{t-1},\theta) = \Pi_{t=1}^{T} D_{y|z}(y_t|z_t, X_{t-1},\lambda_1) D_z(z_t|X_{t-1},\lambda_2), \qquad (3.1)$$

[1]Ericsson (1994) provides an excellent exposition on exogeneity.

where $\lambda = f(\theta)$, *then* z_t *is weakly exogenous for* ϕ *(i) if the parameters of interest* ϕ *are a function of* λ_1 *alone and (ii)* λ_1 *and* λ_2 *are variation free (no cross-restrictions).* The conditional density in (3.1) characterizes that y_t is analyzed without information about the marginal process z_t. The former condition ensures that ϕ is uniquely learned from $\lambda_1$alone, whereas the latter condition guarantees that λ_1 and λ_2 are independently determined each other. No information about ϕ is available from the marginal distribution of $D_Z(z_t|X_{t-1}, \lambda_2)$. In formulating a conditional model, a failure of the condition (i) leads to inconsistency or (possibly) nonconstancy, whereas a failure of the condition (ii) results in inefficiency (Banerjee *et al.*, 1996). It should be noted that weak exogeneity depends on a set of the parameters of interest and so is not an intrinsic property of variables. Testing weak exogeneity is not easy since the power of the test largely depends on the formulation of marginal processes. Nevertheless, Engle (1982b,1984) provides various types of the conventional LM test. Recently, many studies discuss how the concept can be tested in the cointegrated system (see, for examples, Johansen (1992), Urbain (1992), Hendry and Mizon (1993), and Paruolo and Rahbek (1999)).

Strong exogeneity Strong exogeneity is related to the issue of whether or not the past information of the conditional variable Y_{t-1} affects the marginal variable z_t. *If* z_t *is weakly exogenous for the parameters of interest* ϕ *and* Y_{t-1} *does not Granger cause* z_t*,* z_t *is strongly exogenous for* ϕ. If this is the case, (3.1) is factorized as:

$$\Pi_{t=1}^{T} D_x(x_t|X_{t-1,}\theta) = \Pi_{t=1}^{T} D_{y|z}(y_t|z_t, X_{t-1,}\lambda_1) D_z(z_t|Z_{t-1,}\lambda_2). \tag{3.2}$$

Note that the last term $D_z(z_t|Z_{t-1,}\lambda_2)$ is a simplified form of $D_z(z_t|X_{t-1,}\lambda_2)$ in (3.1) without Y_{t-1}. The main difference between weak exogeneity and strong exogeneity is that the former primarily deals with the problems of static enquiry, related to the current information, for the efficient estimation of conditional models, whereas the latter extends the issue of exogeneity by adding dynamic aspects on the past information to weak exogeneity. This type of exogeneity is very useful for the valid analysis of conditional predictions by formally testing whether Y_{t-1} inflences z_t in the marginal density $D_z(z_t|Z_{t-1},\lambda_2)$. Thus, strong exogeneity avoids any feedback problems between conditional variables to be predicted and marginal variables.

Super exogeneity The issue of super exogeneity is related to the invariant property of conditional models to some classes of interventions affecting marginal processes.[2] Unlike the property of weak exogeneity entailing conditional inferences within a regime, super exogeneity extends the parameter constancy of conditional models to the case of regime shifts. *If z_t is weakly exogenous for the parameters of interest ϕ and the conditional model $D_{y|z}(y_t|z_t, X_{t-1},\lambda_1)$ in (3.1) is structurally invariant to a class of interventions which influence the DGP $D_x(x_t|X_{t-1},\theta)$, z_t is super exogenous for ϕ.* Given this definition of super exogeneity, λ_1 is invariant to the interventions affecting λ_2. Thus parameter constancy across regimes is central in testing the concept. Engle and Hendry (1993) provide several testing procedures for super exogeneity.[3] In the context of policy analysis, the concept is very useful for examining whether an econometric model is subject to the Lucas (1976) critique which concerns about the impossibility of invariant conditional models under regime changes.[4] In the presence of super exogeneity, an invariant conditional model refutes the critique. This distinctive feature of super exogeneity further implies that the invariant model can't be reversed because of the nonconstancy of the inverted equation. Engle and Hendry (1994) show

[2]See Aldrich (1989) for an elaborate account of the history of this concept and its relation with another similar concept, 'autonomy'.

[3]Psaradakis and Sola (1996) show that in finite samples, the power of Engle and Hendry's tests largely depends on the formulation of marginal processes. In a recent work, Hansen (2000) proposes a direct testing procedure without formulating marginal models.

[4]See, for more details related to this issue, Hendry (1988), Favero and Hendry (1992), and Ericsson and Irons (1995).

that if a given conditional model has both invariant parameters and invariant error variances across regimes, and if the conditioning process also varies across the regimes, then the reverse regression cannot have invariant parameters and should fail either or both constancy and invariance tests.

3.3 Cointegration

Cointegration is a statistical concept related to the fact that some economic time series are nonstationary but move together in a stationary manner in the long-run. Since the concept was defined by Granger (1981) and Engle and Granger (1987), it has pervasively affected the procedures of estimation and inference in the econometric literature. Given a huge amount of work in this area, we limit our discussions to a brief review of the literature and its role in policy analysis (see, for more details, Banerjee *et al.* (1993), Hamilton (1994), and Johansen (1995)). A time series is integrated, if differencing is required to make it stationary.[5] Nelson and Plosser (1982) show that most of macroeconomic data are nonstationary. The main problems with the presence of nonstationarity in data are that the use of standard normal tables for the tests of significance on conventional $t-$ and $F-$ratios in a regression leads to 'spurious' inferences (Granger and Newbold, 1974); and that the t-statistics calculated in this situation are diverged asymptotically and so the inferences on estimated parameters become worse as the sample size is increased (Phillips, 1986). However, these undesirable effects of nonstationary data on econometric models can be eliminated if there exist some linear combinations of integrated series. Engle and Granger (1987) show that if a set of integrated series are cointegrated, the non-stationarity of data is removed and so the long run static relation between such variables can be formalized. Given this property of cointegration, Stock (1987) further demonstrates that in the presence

[5] In the literature there are a number of the testing procedures for unit roots. See, for an excellent recent survey, Phillips and Xiao (1998).

of cointegration, OLS estimators are 'super-consistent' with a convergence rate at $O_p(T^{-1})$ rather than the usual $O_p(T^{-1/2})$.

To test the existence of cointegration, Engle and Granger (1987) suggest to check whether the cointegrated residuals based on a single equation follow a unit root process, using the Dickey-Fuller statistic (Dickey and Fuller, 1979) and the Durbin-Watson statistic (Sargan and Bhargava, 1983). While this approach is computationally simple, disadvantages of the single equation-based tests are that cointegrated parameters are biased because of serial correlation and nuisance parameters; only one cointegrating relationship between a set of variables is assumed; and a *priori* assumed exogenous independent variables may not be weakly exogenous for the parameters of interest. To overcome the deficiencies of this single equation-based test, numerous multivariate system-based approaches have been proposed.[6] Among them, the maximum likelihood procedure proposed by Johansen (1988) is conceptionally straightforward. The method is based on a maximum likelihood principle to determine the cointegrating rank of a linear dynamic system and generalizes the Dickey-Fuller statistic to a multivariate context for likelihood ratio statistics as a function of a reduced rank. One of the major advantages of this approach is that a whole battery of tests for the hypotheses related to economic theory can be performed in the cointegrating space, even though this advantage may be offset by several problems, such as the test results sensitive to the presence or absence of deterministic terms (a constant term, trend, and indicator variables), selected lag lengths, and sample sizes (see Banerjee *et al.* 1993).

In relation to policy analysis, cointegration plays several important roles. First, following the Granger representation theorem (Engle and Granger, 1987), the existence of cointegration among variables essentially implies an ECM repre-

[6]See Watson (1994) for an extensive survey. We limit our discussion to the Johansen (1988) procedure, since the approach is relatively straightforward and popularly used in applied econometrics. In the Appendix we briefly review the technical details of the procedure.

sentation between the relevant variables. The role of cointegration in this type of an ECM model is to link theoretical long-run relationships between economic variables to statistical short-run dynamics. Thus, the parameters (statistically identified) in cointegrating vectors measure the fundamental structure of how economic policy affects the economy in the long run. Second, if a pair of integrated variables are cointergrated, a causal relationship between them must exist at least in one direction. In an ECM model, there are two possible channels of causation of one variable to the other, either through differenced lag dynamics or through the ECM term. In the case of the former, the direction of causality is straightforward. However, the latter case is a little complicated. To see how causality between the variables of interest occurs through the ECM term, consider the case that both of y_t and z_t are $I(1)$ but cointegrated such that $x_t = y_t - Az_t \sim I(0)$. Then a consequence of the ECM model is that either Δy_t or Δz_t must be Granger-caused by x_{t-1} which is itself a function of y_{t-1} and z_{t-1}. This means that either $y_{t+1}$is Granger-caused in means by z_t or $z_{t+1}$by y_t. Note that cointegration is concerned with the long-run equilibrium between integrated variables, whereas causality in mean is concerned with short-run forecastability (Granger, 1988). This implies that in order for the pair of series to have an attainable equilibrium, there must be some causation between them to provide necessary dynamics. Third, the presence of reduced rank cointegration essentially involves common trends, just like cointegration involves an ECM term. While cointegration contains information about the adjustment processes of the economy when it deviates from its long-run steady states, common trends contain information about how the economy is driven out from its equilibrium. In the context of policy analysis, the former is useful for investigating the transitory effects of policy shocks and the latter is useful for investigating the permanent effects (see, for example, King *et al.* (1991)).

3.4 Co-breaking

Co-breaking is defined as some linear combinations of variables which remove deterministic shifts simultaneously. The basic idea of co-breaking is that some economic variables move together over time with structural breaks. A good example of co-breaking is the changes of nonpolicy variables following the changes in policy. Clements and Hendry (1999) show that in the presence of co-breaking, structural breaks between variables are eliminated, just as non-stationarity in an integrated system due to unit roots is removed by taking linear combinations of variables.[7] The usefulness of the concept is that if co-breaking holds between variables, the joint structural breaks between variables are not subject to the Lucas (1976) critique. Thus co-breaking is an important concept in analyzing policy related issues which involve comovements between policy and nonpolicy variables. Clements and Hendry (1999) propose two types of co-breaking: contemporaneous mean co-breaking (CMC) and intertemporal mean co-breaking (IMC). The former entails the case that structural breaks between variables occur at the same point in time but the breaks do not affect linear combinations of variables. The latter involves the case that the impacts of structural breaks are delayed. CMC is analogous to cointegration from I(1) to I(0), whereas IMC is more parallel to cointegration from I(2) to I(0) where timing matters.

Contemporaneous mean co-breaking CMC occurs in a solved form when breaks between variables are jointly removed. Consider an n-dimensional vector stochastic process $\{x_t\}$ over $\tau = \{1, ..., T\}$ with unconditional expectations around an initial parameter ψ at $t = 0$:

$$E[\mathbf{x}_t - \psi] = \boldsymbol{\mu}_t \in R^n, \tag{3.3}$$

[7] Structural breaks are 'permanent large shifts' that occur intermittently, as against 'permanent small shifts' occurring frequently and so generating I(1) effects.

where $|\boldsymbol{\mu}_t| < \infty$, but otherwise are unrestricted.

Definition 1. The $n \times s$ matrix $\mathbf{\Phi}$ of rank s $(n > s > 0)$ is said to be contemporaneous mean co-breaking of order s (denoted as CMC(s)) for $\{x_t\}$ in (3.3) if $\mathbf{\Phi}'\boldsymbol{\mu}_t = \mathbf{0}\ \forall t \in \tau$ (Clements and Hendry, 1999).

Under CMC, (3.3) can be rewritten as $E[\mathbf{\Phi}'\mathbf{x}_t - \mathbf{\Phi}'\psi] = \mathbf{\Phi}'\boldsymbol{\mu}_t = 0$, so that the parameterization of the reduced set of s linear transforms $\mathbf{\Phi}'\mathbf{x}_t$ is independent of deterministic shifts. Using the equivalence of reduced rank and linear dependence, it is easy to show the possible existence of the fixed matrix $\mathbf{\Phi}$ of which removes all changes in $\{\boldsymbol{\mu}_t\}$ which varies from period to period. To see this in more detail, let $\mathbf{M}_T^1 = (\mu_1, \mu_2, ..., \mu_T)$ be the $n \times T$ matrix where $T > n$, then $\mathbf{\Phi}'\boldsymbol{\mu}_t = 0$ entails $\mathbf{\Phi}'\mathbf{M}_T^1 = \mathbf{0}'$, and so a necessary and sufficient condition for $\mathbf{\Phi}'\boldsymbol{\mu}_t = \mathbf{0}$ is that rank $(\mathbf{M}_T^1) < n$. In a practical case where breaks are related across variables, CMC occurs for at least order $s = n - k$, if $\boldsymbol{\mu}_t = \boldsymbol{\alpha}\mathbf{l}_t\ \forall t$ where $\boldsymbol{\alpha}$ is $n \times k$ of rank $k < n$ and $\mathbf{l}_t$ is $k \times 1$. This is shown by rank $(\mathbf{M}_T^1) = \text{rank}(\boldsymbol{\alpha}\mathbf{l}_1, \boldsymbol{\alpha}\mathbf{l}_2, ..., \boldsymbol{\alpha}\mathbf{l}_T) = \text{rank}(\boldsymbol{\alpha}\mathbf{L}_T^1) \leq k$, where $\mathbf{L}_T^1 = (\mathbf{l}_1, \mathbf{l}_2, ..., \mathbf{l}_T)$.

Intertemporal mean co-breaking IMC occurs when breaks at different points of time are eliminated, such as $\sum_{i=1}^{n} \sum_{j=0}^{p} \phi_{i,j}\mu_{i,t-j} = 0$, where j denotes lagged periods. Under IMC, breaks are vanished between combinations of current and lagged values of variables. Let $\mathbf{\Phi}(L)$ denote an $n \times s$ polynomial matrix of degree $p > 0$, $\mathbf{\Phi}(L) = \sum_{i=0}^{p} \mathbf{\Phi}_i L^i$, and its associated $n(p+1) \times s$ matrix $\mathbf{\Phi}^* : \mathbf{\Phi}^{*\prime} = (\mathbf{\Phi}_0', \mathbf{\Phi}_1', ..., \mathbf{\Phi}_{p-1}', \mathbf{\Phi}_p')$. Then,

Definition 2. The $n \times s$ polynomial matrix $\mathbf{\Phi}(L)$ of degree $p > 0$, where the rank of $\mathbf{\Phi}^*$ is s $(n \geq s > 0)$, is said to be intertemporal mean co-breaking of order s (denoted as IMC(p, s)) for $\{x_t\}$ in (3.3) if $\mathbf{\Phi}'(L)\boldsymbol{\mu}_t = 0\ \ \forall t \in \tau$, and no $n \times s$ matrix polynomial of degree $p - 1$ and rank s annihilates $\boldsymbol{\mu}_t\ \forall t \in \tau$ (Clements and Hendry, 1999).

Given this definition, (3.3) can be rewritten as $E[\mathbf{\Phi}'(L)\mathbf{x}_t - \mathbf{\Phi}'(1)\psi] = \mathbf{\Phi}'(L)\boldsymbol{\mu}_t = 0$. The condition of $\mathbf{\Phi}'(L)\boldsymbol{\mu}_t = 0$ is much weaker than $\mathbf{\Phi}'\boldsymbol{\mu}_t = 0$, since the reduced

s-dimensional dynamic system of p lags in n variables $\mathbf{\Phi}'(L)\mathbf{x}_t$ does not depend on deterministic shifts immediately as in (3.3). Note that CMC(s) is parallel to IMC(1,s). Co-breaking discussed briefly in the above has an important implication for modelling invariant econometric models, especially in the context of policy analysis and forecasting. Economic policy commonly involves joint changes between policy (such as money stocks or interest rates) and nonpolicy variables (price levels, economic growth, and unemployment, etc,), whatever policy actions are exogenous or endogenous. This implies that policy and nonpolicy variables concomitantly move together with structural breaks. If this is the case, conventional conditional models are not invariant and so are subject to the Lucas (1976) critique. However, Hendry and Mizon (1998) show that conventional conditional models can be immunized from the critique, if there exist co-breaking relations between the parameters of conditional and marginal models, under the assumptions that agents are rational but act in a contingent manner and conditioning variables are weakly exogenous. Clements and Hendry (1999) further discuss the implications of co-breaking for predictive failure in forecasting.

3.5 Conclusions and summary

In this chapter, we have reviewed the notions of exogeneity, cointegration, and co-breaking and their roles in econometric modelling and policy analysis. Each of these concepts plays a central role in characterizing dynamic relations between variables and interactively provides a useful tool in analyzing the likely effects of economic policy. Weak exogeneity relates to the issue of how to formulate valid conditional models and plays a key role in examining all the other concepts. In the context of policy analysis, this concept is very useful to statistically identifying policy stance, particularly in VAR type models within which all variables are determined endogenously and so *ad hoc* arbitrary recursive orderings are used to identify exogenous policy variables. In VAR models, weak exogeneity can be used to provide a statistical legitimacy to block-recursively identify policy indicators, measuring policy actions, either as exogenous or as endogenous variables. Strong exogeneity which combines weak exogeneity and Granger causality can be used for the purpose of forecasts by avoiding any feed-back problems between endogenous variables to be forecast and exogenous variables. Super exogeneity and co-breaking are the major concepts for the evaluation of policy effectiveness. If the former holds, policy is ineffective, whereas if the latter holds, policy is effective. Thus, the two concepts provide different empirical conclusions for policy effectiveness. However, both rule out the applicability of the Lucas (1976) critique. Cointegration implies that there must be causal relationships between variables at least in one direction. The existence of cointegration which necessarily accompanies common trends provides a framework to be able to analyze the transitory and permanent effects of policy. In the next three chapters all these concepts are interactively used for policy analysis.

Appendix

In this appendix the Johansen (1988) maximum likelihood test is briefly reviewed. Consider an n-dimensional vector x_t with the p-th order VAR:

$$x_t = \sum_{i=1}^{p} \Pi_i x_{t-i} + \epsilon_t, \qquad \epsilon_t \sim IN(0, \sum). \tag{A3.1}$$

With $I(1)$ data, a useful reformulation of (A3.1) is the ECM form:

$$\Delta x_t = \Pi x_{t-1} + \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \epsilon_t, \tag{A3.2}$$

where $\Gamma_i = -\sum_{j=i+1}^{p} \Pi_j$, $\Pi = \sum_{i=1}^{p} \Pi_i - I$, and rank $(\Pi) = \gamma < n$, so $\Pi = \alpha\beta'$ where α and β are $n \times \gamma$ matrices. Let (A3.2) be rewritten as the following summarized form:

$$Z_{0t} = \Pi Z_{1t} + \Gamma Z_{2t} + \epsilon_t, \tag{A3.3}$$

where $Z_{0t} = \Delta x_t$, $Z_{1t} = x_{t-1}$, and $Z_{2t} = (\Delta x_{t-1}, ..., \Delta x_{t-k+1})$. Then, regress Z_{0t} and Z_{1t} on Z_{2t}, obtain the residuals $R_{0t} = Z_{0t} - Z_{2t}(Z_{2t}' Z_{2t})^{-1} Z_{2t}' Z_{0t}$ and $R_{1t} = Z_{1t} - Z_{2t}(Z_{2t}' Z_{2t})^{-1} Z_{2t}' Z_{1t}$ where Z_{0t}, Z_{1t}, R_{0t} and R_{1t} are $(T \times n)$ and Z_{2t} is $(T \times n(p-1))$,and compute their product moment matrices $S_{00} = T^{-1} R_{0t}' R_{0t}$, $S_{01} = T^{-1} R_{0t}' R_{1t}$, $S_{11} = T^{-1} R_{1t}' R_{1t}$, and $S_{10} = S_{01}'$. The system (A3.2) can be written as $R_{0t} = \Pi R_{1t} + \epsilon_t$ with its concentrated likelihood function (CLF):

$$\begin{aligned} \ell_c(\Pi) &= K_c - T/2 \log |T^{-1}(R_{0t} - \Pi R_{1t})(R_{0t} - \Pi R_{1t})'| \\ &= K_c - T/2 \log |S_{00} - \Pi S_{10} - S_{01}\Pi' + \Pi S_{11} \Pi'|. \end{aligned} \tag{A3.4}$$

Consider the case that Π is unrestricted. In this case the CLF is maximized by minimizing the sum of squares and then leads to

$$\ell_c(\Pi) = K_c - T/2 \log |S_{00} - S_{01} S_{11}{}^{-1} S_{10}| \tag{A3.5}$$

with $\widehat{\Pi} = S_{01}(S_{11})^{-1}$. If a reduced rank is imposed with $\Pi = \alpha\beta'$, where α is an $(n \times \gamma)$ and β' a $(\gamma \times n)$ matrix, the concentration of $\ell_c(\alpha, \beta)$ is minimized with respect to α, $\partial \ell_c(\alpha,\beta)/\partial\alpha = 0$, which implies that

$$\alpha_c(\beta) = S_{01}\beta(\beta' S_{11}\beta)^{-1}.$$

Substituting this into (A3.4) yields the following CLF $\ell_c^*(\beta)$:

$$\ell_c^*(\beta) = \; K_c - T/2 \log |S_{00} - S_{01}\beta(\beta' S_{11}\beta)^{-1}\beta' S_{10}|. \tag{A3.6}$$

Maximizing the likelihood function (A3.6) is equivalent to minimizing $|S_{00} - S_{01}\beta(\beta' S_{11}\beta)^{-1}\beta' S_{10}|$, which is the case when the generalized variance ratio is $|\beta'(S_{11} - S_{10}S_{00}^{-1}S_{01})\beta|/|\beta' S_{11}\beta|$.[8] With a normalization $\beta' S_{11}\beta = I$, this leads to the minimization of

$$|\beta'(S_{11} - S_{10}S_{00}^{-1}S_{01})\beta|. \tag{A3.7}$$

This factor is minimized by solving the characteristic equation $|\lambda S_{11} - S_{10}S_{00}^{-1}S_{01}| = 0$ with the γ largest eigenvalues $1 > \; \widehat{\lambda}_1 > \ldots > \widehat{\lambda}_n > 0$ and the corresponding eigenvectors:

$$(\widehat{\lambda}_i S_{11} - S_{10}S_{00}^{-1}S_{01})\widehat{\beta_i} = 0, \tag{A3.8}$$

where $\widehat{\lambda}_i S_{11}\widehat{\beta_i} = S_{10}S_{00}^{-1}S_{01}\widehat{\beta}_i$, $\widehat{\beta_i'} S_{11}\widehat{\beta}_j = 1$ if $i = j$, and $\widehat{\beta_i'} S_{11}\widehat{\beta}_j = 0$ if $i \neq j$. With the cointegrating rank γ and the corresponding eigenvector $\widehat{\beta}$, the restricted

[8] Just as the determinant of the following 2×2 matrix can be decomposed into two versions of submatrices $\begin{vmatrix} a & b \\ c & d \end{vmatrix} = ad - bc = d(a - bd^{-1}c)$ or $a(d - ba^{-1}c)$,the covariance matrix of the concentrated (A3.6) can be decomposed as $\begin{vmatrix} S_{00} & S_{01}\beta \\ \beta' S_{10} & \beta' S_{11}\beta \end{vmatrix} = |S_{00} - S_{01}\beta(\beta' S_{11}\beta)^{-1}\beta' S_{10}||\beta' S_{11}\beta|$ or $|S_{00}||\beta'(S_{11} - S_{10}S_{00}^{-1}S_{01})\beta|$. This implies that $|S_{00} - S_{01}\beta(\beta' S_{11}\beta)^{-1}\beta' S_{10}| = |S_{00}||\beta'(S_{11} - S_{10}S_{00}^{-1}S_{01})\beta|/|\beta' S_{11}\beta|$. Since $|S_{00}|$ is constant relative to β, the maximization of the likelihood function (A3.6) is equivalent to the minimization of $|\beta'(S_{11} - S_{10}S_{00}^{-1}S_{01})\beta|/|\beta' S_{11}\beta|$.

likelihood of (A3.7) and (A3.8) becomes $|\widehat{\beta}'(S_{11} - S_{10}S_{00}^{-1}S_{01})\widehat{\beta}| = |I_\gamma - \widehat{\Lambda_\gamma}| = \Pi_{i=1}^{\gamma}(1 - \widehat{\lambda}_i)$, where $\widehat{\beta}' S_{10}S_{00}^{-1}S_{01}\widehat{\beta} = \widehat{\Lambda\gamma} = diag(\widehat{\lambda}_1, ..., \widehat{\lambda}_\gamma)$. Then from (A3.6):

$$\ell_c^*(\widehat{\beta}) = K_c - T/2\log|S_{00}| - T/2\sum_{i=1}^{\gamma}\log(1 - \widehat{\lambda}_i) \qquad \text{(A3.9)}$$

corresponding to the γ largest eigenvalues. If Π is not restricted, the maximum of the likelihood is:

$$\ell_c^*(\widehat{\beta_0}) = K_c - T/2\log|S_{00}| - T/2\sum_{i=1}^{n}\log(1 - \widehat{\lambda}_i), \qquad \text{(A3.10)}$$

where $\widehat{\beta_0}$ is the $(n \times n)$ matrix of eigenvectors. By comparing (A3.9) and (A3.10), we can derive the trace statistic $\zeta_\gamma = -T\sum_{i=\gamma+1}^{n}\log(1 - \widehat{\lambda}_i)$. Alternatively, the maximum eigenvalue statistic for testing γ cointegrating vectors within $(\gamma + 1)$ can be obtained on the basis of the $(\gamma + 1)^{st}$ eigenvalue using $\eta_\gamma = -T\log(1 - \widehat{\lambda}_{\gamma+1})$. Critical values for the tests are tabulated in Johansen (1988), Johansen and Juselius (1990), and Osterwald-Lenum (1992). Reimers (1992) further discusses the adjustment of the above two statistics in finite samples.

Chapter 4

Regime Changes and Econometric Modeling of the Demand for Money[1]

4.1 Introduction

The stability of the demand for money is important for the effectiveness of monetary policy on real economic activities. Since Goldfeld's (1976) finding of 'missing money', the question of the instability of money demand has been extensively investigated in a number of empirical studies (see, for extensive surveys, Judd and Scadding (1982) and Laidler (1993)). One of the most common reasons for the instability includes regime changes. If this is the case, from the viewpoint of Lucas (1976) it is almost impossible to formulate constant conditional models for money demand, based on the backward looking behavior of economic agents. Lucas argues that since economic agents rationally integrate their knowledge about

[1] **An earlier version of this chapter was presented at the 1999 Far Eastern Meeting of the Econometric Society in Singapore and at Southampton University. A revised version will be forthcoming in Economic Modelling, 2002.**

policy changes, the coefficients of an econometric model which fails to account for the rational behavior of the agents are a mixture of 'deep' and expectational parameters and so are changed whenever policy changes, consequently providing misleading policy analysis and forecasting.

On the other hand, Hendry (1979) explains the instability of money demand function with mis-specifications in econometric models and establishes a constant error correction model (ECM) through a general-to-specific modelling approach. Extending this model with a slight modification, Hendry and Ericsson (1991b) also show that the model is invariant to regime changes and refutes the Lucas (1976) critique, by applying the test of super exogeneity defined by Engle *et al.* (1983). Following this approach, several empirical studies for developed countries have successfully formulated constant models, which are also invariant to regime changes, for the money demand behavior of economic agents.[2] Among others, these include Baba *et al.* (1992), Bårdsen (1992), Ericsson and Irons (1995), and Muscatelli and Spinelli (2000). In the case of developing countries, Gupta and Moazzami (1989), Tseng and Corker (1991), Arize (1994), and Qin (1998) also use ECM models to investigate money demand behavior. However, most of them, except Tseng and Corker (1991), do not examine parameter constancy across policy regimes, even though the countries they investigate have been under substantial structural changes due to a number of financial reform measures (namely financial liberalization) since the early 1980s.

The purpose of this paper is to investigate whether regime changes, particularly financial liberalization, in Korea have broken down the stability of the demand function for money. The Korean case may provide an interesting case study for developing countries. Like many other developing countries, during the past three decades the economy has experienced a number of instable economic events, such as oil price shocks and financial deregulation measures. Particularly,

[2]The main difference between constancy and invariance is that while the former is time-independence of parameters, the latter is constancy across interventions.

it has been widely recognized that financial liberalization undertaken from the early 1980s has caused the economy to be structurally changed (see Park (1994)). In this line, Tseng and Corker (1991) empirically show that the demand function for money in Korea has been instabilized by a series of financial deregulation. Unlike this finding, in this paper we present a contradictory evidence by developing a constant ECM model for money demand with non-constant marginal models for the data generating processes of conditioning variables and demonstrate that financial liberalization has caused the behavior of some exogenous variables to be changed, not the structural relationship between money and macroeconomic variables.

Since the robustness and usefulness of this finding depends on the statistical validity and appropriate interpretation of the underlying model, we develop the empirical model by applying extensive statistical tests, including misspecification, cointegration, and weak exogeneity, examine the properties of parameter constancy, invariance, and observational equivalence, and compare the model with a forward-looking model by applying encompassing tests.[3] The issues of invariance and observational equivalence are investigated by applying a testing strategy of super exogeneity tests suggested by Engle and Hendry (1993). Firstly, non-constant marginal models for conditioning variables are formulated; the determinants of the non-constancies are identified; and then the statistical significance of the identified determinants is tested in the estimated conditional model. The test results show that the obtained ECM model is invariant to regime changes, so refuting the Lucas (1976) critique, and identifies itself as a class of

[3]A well-known problem with an ECM type model is that estimated coefficients are observationally equivalent to the parameters of a feedforward model because of its data-based general specification (Nickell, 1985; Cuthbertson, 1988). In other words, an ECM model can be interpreted as a model representing the behavior of economic agents either who are adaptive to the past information or who rationally utilizes all relevant information. This implies that different policy prescriptions can be obtained from an estimated ECM model, depending on how to interpret it. This undesirable dual property necessitates identifying observational equivalence in ECM type models (see Ericsson and Hendry (1999) for a recent discussion on this issue).

backward looking models. To reinforce the credibility of the model, we compare the model, using encompassing tests, with a forward-looking model based on a multi-period quadratic loss function and rational expectations. The reasons for selecting this rival model are that the two models are the best rival models which provide invariant parameters under regime changes but are formulated in a completely different way, that is, theory-based and data-based, and that since the two models are observationally equivalent, the robustness of a direct discrimination by using the test of super exogeneity can be complementarily examined. The test results suggest that the estimated ECM model performs better than the theory-based expectations model and apparently discriminates itself from the interpretation as a class of feed-forward models. This evidence complementarily supports the results derived from the test of super exogeneity.

The chapter is organized as follows. Section 2 outlines the institutional characteristics of the Korean economy and their implications in modelling an econometric model for the demand for money. Section 3 discusses the basic function of money demand and examines the properties of the nonstationarity and weak exogeneity of data used. In Section 4 a data-based ECM model for M2 money is developed and its economical and statistical properties are discussed. Section 5 investigates the invariance and observational equivalence of the estimated model by applying the test of super exogeneity. Section 6 develops a theory-based feed-forward model and compares it with the data-based ECM model by applying encompassing tests. Section 7 concludes the chapter.

4.2 Institutional Characteristics and the Implications in Modelling the Demand for Money

In this section we review the institutional characteristics of the Korean economy over the sample period and discuss their implications in modelling an econometric model of the demand for money. The entire sample may be divided into two distinct sub-periods on the basis of degrees of government intervention: 1972-1980 and 1981-1997.[4] The former represents the period of financial repression and the latter represents the period of financial deregulation. In the early stage of its economic development in the 1970s, the economy can be characterized as the non-existence of well-developed money and capital markets, the predominance of government-owned banks in the financial system, a dualistic financial structure consisted of organized and unorganized curb markets, and high debt-equity ratios of business firms. Before 1980, the relative priority of government policy was given to economic growth rather than price stability. In order to maintain the rapid economic growth and industrial development of the economy, Korean monetary authorities had strictly controlled the whole economic system. All bank interest rates were fixed at lower levels than interest rates in markets, and most bank loans were directed to a few priority sectors designated by the government. This financially repressed development strategy had considerably contributed to the economic growth of the economy, but resulted in perpetuated high inflation and large inefficiency of the entire economic system.

In the early 1980s, the government which recognized the seriousness of the adverse effects of financial repression, changed its relative policy priority from rapid economic growth to price stability and began to undertake a series of deregulation measures to enhance the efficiency of the economic system and to improve the effectiveness of government policy. The measures include the liberalization of interest rates on interbank money transactions and prime commercial papers, the privatization of government-owned commercial banks, and the relaxation of direct controls on bank credits. Under this partially liberalized environment, the economy had achieved high economic growth and price stability for a while. However, after a decade of the favorable macroeconomic performance, the coun-

[4]See Park (1994) for more details.

try internally and externally came under pressure to internationalize its financial markets. From 1993 the government again began to undertake a wide range of substantial deregulation measures to open the economy to foreign investors. Regulated bank interest rates were fully liberalized through a step-by-step process during the period from 1993 to 1997, entry barriers to non-bank financial intermediaries were largely relaxed, and direct regulations on capital accounts transactions were considerably alleviated. Overall, this market-oriented financial liberalization since the early 1980s, even though there still exists a certain extent of regulation, has caused the economy to change structurally in various ways.

The institutional characteristics discussed above suggest several implications in formulating an econometric model of the demand for money in Korea. First, since the economy doesn't have well-developed money and capital markets which provide alternative financial assets to money and has been dominated by high inflationary expectations over the sample period, it is expected that like many other developing countries, inflation as a proxy of the opportunity cost to money is a dominant factor in holding real money balances. The long-run elasticity of the demand for money balances to the own rate of return on money may be small. Since the government has directly regulated deposit interest rates paid by banks at low levels, it is not expected that the money holdings of economic agents are sensitive to changes in the own rate of interest. Also, money is not expected to closely substitute a small range of alternative liquid financial assets because of its 'uniqueness' in the economy. The income elasticity of money demand, particularly defined in a broad sense, would be more than unity, because of a wealth effect resulted from high savings in the form of time deposits consisting of the majority of broad money balances and because of increased 'monetization' associated with the economy's rapid financial developments (Bordo and Jonung, 1981).

Second, a well-known identification problem in a typical money demand equation may not be serious in the case of Korea. Since interest rates have been continuously regulated over the sample period, they can be regarded as exogenous (Gordon, 1984).[5] This regulatory exogeneity of interest rates may rule out the possibility that an econometric model of the demand for money is interpreted as an equation for money supply. Third, since the country has experienced many regime changes over the sample period, careful examination of whether the structural relationship between money and economic variables has been changed or not is required. If the parameters of an econometric model for money demand are not constant, the model may be subject to the Lucas (1976) critique and then is not useful for policy analysis and forecasting. In this context, it is necessary to test the invariant property of the underlying model.

4.3 Statistical Properties of Data

Most theories of the demand for money explain the money holdings of economic agents by emphasizing the role of money as a medium of exchange and a store of value. In the former case that agents hold money to cover either current or unexpected future transactions, the aggregate balances of money are mainly determined by price levels and the volumes of current and future real transactions. On the other hand, the money holdings for speculative motives, as part of a store of value, primarily depend on returns on money itself, returns on alternative assets, and the level of total assets. To investigate these relationships between money and economic variables, a number of studies have used different empirical models for different definitions of money. However, the models used, whatever their theoretical backgrounds are, can be simplified as the following functional form under long-run price homogeneity:

$$(M_t/P_t) = f(Y_t, Ra_t, Rm_t), \tag{4.1}$$

[5]Following Cooley and LeRoy (1981), the statistical exogeneity of interest rates is further discussed in the next section.

where M_t is a measure of nominal money balances; P_t is the price level; Y_t is a scale variable, such as income or wealth; Ra_t is the rate of return on assets alternative to money; and Rm_t is the own rate of return on money.

To investigate equation (4.1), in this study we use seasonally unadjusted, quarterly data of $M2$ money, the consumer price index, real gross domestic product, a one-year time deposit rate, and a three-year corporate bond rate.[6] Except for interest rates, all variables were transformed into logarithms. Hereafter lower case letters denote the logs of the corresponding capitals. The data comprise the period from 1972(3) to 1997(4) and were obtained from the various issues of the Bank of Korea's monthly bulletin. As a preliminary inspection of the data, Figure 4.1(a) plots the time series of $(m-p)_t$, y_t, and p_t over the entire sample, with the latter two variables adjusted to match their means to the mean of real money. All the three series apparently exhibit a common upward trend, indicating that they may be nonstationary, while the movement of real income is largely dominated by seasonality. Figure 4.1(b) plots $\Delta_4 p_t$, Ra_t,and Rm_t, with the annual inflation adjusted to match the means and ranges of the latter two interest rates. They move together over the whole period with an evident break in the early 1980s, due to financial deregulation. Note that while $(m-p)_t$ steadily grows without any break, other macroeconomic variables which are potentially expected to be exogenous in a model for the demand for money function show apparent changes in their mean values in the early 1980s. With these asymmetric data series, our purpose is to derive a constant money demand model which is also invariant to the nonconstancy in sub-marginal models for economic variables.

Considering the nonstationarity of the data series, we first conducted augmented Dickey-Fuller (ADF) tests with up to five auxiliary lags. The test statistics reported in Table 4.1 suggest that $(m-p)_t$, y_t, Ra_t, Rm_t and $\Delta_4 p_t$ are $I(1)$,but m_t and p_t appear to be $I(2)$. Since all the variables we are concerned with

[6] The real GDP measure and the consumer price index are based at 1990 constant prices.

are integrated with the order of $I(1)$, we straightforwardly tested cointegration to examine the long-run equilibrium relationship between real money balances and other economic variables using Johansen's (1988) maximum likelihood procedure. A fifth-order vector autoregression (VAR) model was initially estimated with a constant term and seasonal dummies. The constant term was not restricted to lie in the cointegration space. The test results reported in Table 4.2(a) show that the null hypothesis of cointegrating rank $\gamma = 0$ in the trace test is rejected at the 5% level by the critical value taken from Osterwald-Lenum (1992). Therefore, we assume that there is only one cointegrating vector in the system. Table 4.2(b) reports the test statistics of long-run weak exogeneity proposed by Engle *et al.* (1983) and extended by Johansen (1992) into a cointegrated system. While $(m-p)_t$ can be treated as an endogenous variable in the system, y_t, Ra_t, Rm_t, and $\Delta_4 p_t$ are weakly exogenous for the parameters of the cointegrating vector. Hence, in the long run the weakly exogenous variables are not Granger-caused by real money.[7] As shown by Engle *et al.* (1983), an important methodological implication of this finding is that even though we conduct the analysis of the demand function for money through a single equation conditioned on weakly exogenous variables, we do not lose any information on the entire system.

Before we move on to single equation-based modelling, a long-run money demand equation is identified in the cointegration space for a later comparison. From the standardized eigenvectors of the cointegration test, an unrestricted long-run equilibrium relationship between variables was obtained by normalizing on real money balances: $\beta' x_t = (m-p)_t - 1.28 y_t + 0.01 Ra_t - 0.02 Rm_t + 0.78 \Delta_4 p_t$. The estimated parameters show expected signs, but the theoretical identification is still required for the equation to be interpreted from the economic point of view. A joint restriction of weak exogeneity and income homogeneity was initially tested, but the likelihood ratio (LR) test strongly rejects the null hypothesis at the 5%

[7] See Hunter (1992) who defines Granger non-causality in this type as 'co-integrating exogeneity'.

level with $\chi^2(5) = 23.12$. An alternative test for the joint restrictions of weak exogeneity on the loading matrix and of no opportunity cost from the inflation rate, under the assumption that the Fisher effect of the expected inflation rate on the level of the nominal interest rate exists in the long run, is accepted at the 1% level with $\chi^2(5) = 14.67$ and yields the following equation (the figures in parentheses are standard errors):

$$\begin{array}{lllll} \beta' x_t = & (m-p)_t & -1.31 y_t & +0.02 Ra_t & -0.01 Rm_t \\ & (-) & (0.03) & (0.01) & (0.008) \end{array} .$$

We may regard this equation as a money demand equation which represents the long-run static relationship between real money, real income and interest rates. The identified coefficients show expected signs consistent with economic theory and have similar magnitudes to those in the previous studies (see, for example, Arize (1994)).

4.4 An Econometric Model of the Demand for Money

Based on the results of cointegration and weak exogeneity tests in the previous section, we focus on modelling a single equation for money demand function. Initially, the following unrestricted ECM equivalent to a fifth-order autoregressive distributed lag (ADL) model was estimated:

$$\begin{aligned} \Delta(m-p)_t &= a_0 + \sum_{i=1}^{4} a_{1i} \Delta(m-p)_{t-i} + \sum_{i=0}^{4} a_{2i} \Delta Ra_{t-i} \\ &\quad + \sum_{i=0}^{4} a_{3i} \Delta Rm_{t-i} + \sum_{i=0}^{4} a_{4i} \Delta y_{t-i} \\ &\quad + \sum_{i=0}^{4} a_{5i} \Delta_4^2 p_{t-i} + a_6 (m-p)_{t-1} + a_7 y_{t-1} \end{aligned} \tag{4.2}$$

$$+a_8 Ra_{t-1} + a_9 Rm_{t-1} + a_{10}\Delta_4 p_{t-1}$$
$$+\sum_{i=1}^{3} a_{11i} S_{it} + u_t,$$

where S_t denotes seasonal dummies. For the level terms, the identified equation in the cointegrating vector can be directly used in a restricted way. But, here we estimate the terms unrestrictly and compare the obtained model with that derived from a multivariate VAR model in section 3. The initially estimated results show that most of the coefficients are not easily interpretable and statistically insignificant. So we sequentially simplified the model by eliminating insignificant coefficients and imposing other data-acceptable restrictions and finally obtained the following parsimonious dynamic ECM:[8]

$$\begin{aligned}
\Delta(m-p)_t = \; & \underset{(0.23)}{-0.82} \underset{(0.13)}{-0.57}\Delta_4^2 p_t \underset{(0.10)}{-0.36}\Delta_4^2 p_{t-4} \underset{(0.09)}{+0.16}\Delta(m-p)_{t-3} \\
& \underset{(0.001)}{-0.001}\Delta_2 Ra_t \underset{(0.03)}{-0.11}(m-p)_{t-1} \underset{(0.04)}{+0.15} y_{t-1} \underset{(0.001)}{-0.003} Ra_{t-1} \\
& \underset{(0.001)}{+0.002} Rm_{t-1} \underset{(0.01)}{-0.09} S_{1t} \underset{(0.01)}{-0.03} S_{2t} \underset{(0.01)}{-0.02} S_{3t}
\end{aligned} \tag{4.3}$$

$T = 74(4) - 97(4)$, $R^2 = 0.738$, $\hat{\sigma} = 0.0202$, $DW = 2.17$, $F_{AR}(5,76) = 2.44$, $F_{ARCH}(4,73) = 1.68$, $\chi^2_N(2) = 0.23$, $F_{RESET}(1,80) = 0.03$, $F_H(19,61) = 0.66$,

where F_{AR} is the Lagrange multiplier (LM) test for fifth-order autocorrelation; F_{ARCH} is the Engle (1982a) test for fourth-order autoregressive conditional heteroscedasticity; χ^2_N is the Doornik- Hansen (1994) test for nonnormal residual skewness and kurtosis; F_{RESET} is Ramsey's (1969) test for omitted variables and incorrect functional form; F_H is the White (1980) test for heteroscedasticity.

[8]See Hendry and Ericsson (1991a) for a similar empirical model.

The model satisfactorily performs in terms of mis-specification tests. All the test statistics are not significant at the 5% levels, except that the null hypothesis of serial correlation in residuals is marginally rejected. The fitted residuals have a standard deviation of 0.02. To evaluate the parameter constancy, we recursively reestimated the equation over the entire sample. Figure 4.2(a) graphs the recursive one-step residuals and the corresponding equation two standard errors (i.e., $\{y_t - \widehat{\beta}_t x_t\}$ and $\{0 \pm 2\widehat{\sigma}_t\}$ in a common notation). Only a residual value in 1993(4) is slightly outside the two-standard-error band. Figure 4.2(b) graphs the recursive break-point Chow (1960) statistics scaled by their 5% critical values for the sequences {1978(4)-1997(4), 1979(1)-1997(4), ..., 1997(3)-1997(4), 1997(4)}. None of the statistics rejects the parameter constancy.

Economically, the short-run changes in real money are largely affected by accelerated inflation and somewhat by changes in its own third lag and the opportunity cost variable Ra_t. The large negative influence from the changes in the annual inflation rate seems to reflect a higher degree of substitutability between money and real assets as inflation alters the relative returns on these assets in the absence of sophisticated financial assets alternative to money. The measured feedback coefficient $-\ 0.11$ is statistically significant, but indicates a slow adjustment to the past disequilibrium in money balances. By setting all changes in the short-run to zero, the long-run static equation of (4.3) is obtained: $(m-p)_t = -\ 7.45 + 1.36y_t - 0.03Ra_t + 0.02Rm_t$. However, in the case of a steady-state dynamic growth where $\Delta_4 m_t = \pi_1$ and $\Delta_4 p_t = \pi_2$ constantly grow but growth in interest rates is zero, the long-run equilibrium of (4.3) can be solved as $(m-p)_t = K + 1.36y_t - 0.03Ra_t + 0.02Rm_t$, where $K = -\ 7.45 - 1.90(\pi_1 - \pi_2)$. If $\pi_1 = \pi_2$, K is simply $-\ 7.45$, so that the equation is equivalent to the static case when the short-run dynamics are fixed over time. It is noticeable that the long-run parameter values are very close to those obtained from the VAR-based cointegration analysis in the previous section. All the elasticities show expected signs consistent with traditional priors and the magnitudes appropriately reflect the certain characteristics of the Korean economy, as explained in Section 2.

The long-run income elasticity exceeds unity, implying that money is a 'luxury' good in this economy and there are no economies of scales in holding money. A 1% increase in real income is associated with a 1.36% increase in real money balances and hence with a 0.36% decrease in income velocity. The semi-elasticities of interest rates indicate that an increase of 1% in Ra_t and Rm_t (in absolute value) is accompanied by a decrease of 3% and an increase of 2% in real money balances, respectively. The low elasticity of the demand for money with respect to the changes in the own rate of return on money seems to reflect the regulated nature of the interest rate paid by banks at low levels, while the low market interest elasticity implies that money doesn't closely substitute for alternative liquid financial assets because of its uniqueness. A well-known classical identification problem in a typical money demand equation does not seem to arise in this model, since the model has been formulated on the basis of the weak exogeneity of conditioning variables and the endogeneity of real money and since the obtained coefficients are interpreted as those of a typical equation for money demand. Furthermore, as briefly discussed in Section 2, interest rates have been administered by the government, so there would be no possibility of money supply shocks independent of shifts in money demand.

Overall, equation (4.3) is statistically acceptable and provides a sensible interpretation which is coincident with the conventional theory of the demand function for money. In particular, the estimated parameters are constant even under large regulatory and institutional changes, including two oil price shocks in the 1970's, a series of financial deregulation measures since 1981, and the introduction of a law for 'real name' use in financial transactions in 1993. This empirical evidence is contrasted with Tseng and Corker's (1991) finding that financial liberalization in Korea has broken down the stability of the demand function for money. Based on the constancy of our model, the observed non-constancy in the ECM used by Tseng and Corker seems to be resulted from mis-specifications of dynamics, possibly omitting relevant lags, rather than structural shifts in the relation between money and its determinants.

4.5 Invariance and Observational Equivalence

In the preceding section we presented a parsimoniously constant ECM model of the demand for M2 money under a number of regulatory and institutional changes. However, in order for the model to be useful for the purpose of policy analysis and forecasting, two econometric issues should be further clarified. One is the issue of invariance. If the model is variant to policy regime changes, it is subject to the Lucas (1976) critique and as a result, fails to represent the 'structure' of the economy. In this context, the invariant property of the model should be examined. The other is the issue of observational equivalence. Since the model is derived from a general specification summarizing the behavior of data and has constant parameters, it is difficult to discriminate the model either as a feed-forward model or as a feed-back model (Nickell, 1985; Cuthbertson, 1988). The main problem with this observational equivalence is that completely different policy prescriptions can be suggested, depending on the interpretation of the model, and as a result, mis-interpretation of the model might lead to incorrect conclusions. Fortunately, these two issues are simultaneously solved by using the concept of super exogeneity in the presence of regime changes.[9] Hendry (1988) and Favero and Hendry (1992) show that under regime changes, if conditioning variables in a conditional model are super exogenous - that is, conditioning variables are weakly exogenous and the parameters of interest are invariant to changes in the marginal processes for conditioning variables, then

[9] In the absence of structural changes in the marginal processes for conditioning variables, a direct discrimination, using the test of super exogeneity, of observational equivalence in ECM type models is very difficult. In this case, the encompassing tests suggested by Mizon and Richard (1986) can be used as an indirect alternative method to discriminate feed-back and feed-forward models.

the conditional model is certainly a feed-back model, encompasses a whole class of rational expectations models, and refutes the applicability of the Lucas critique. Thus, finding structural breaks and examining their statistical impacts on the conditional model provide a very powerful tool of discriminating between otherwise equivalent models and ascertaining structure.[10]

To examine whether the conditioning variables in (4.3) have such properties or not, we apply Engle and Hendry's (1993) testing strategies which emphasize the constancy of conditional models associated with the non-constancies in the marginal processes for conditioning variables. Initially, two marginal models for the contemporaneous variables $\Delta_4^2 p_t$ and $\Delta_2 Ra_t$ were estimated by using their own autoregressive (AR) terms as follows[11]:

$$\begin{aligned}\Delta_4^2 p_t = {} & \underset{(0.08)}{0.25\Delta_4^2 p_{t-2}} \; \underset{(0.08)}{+0.27\Delta_4^2 p_{t-3}} \; \underset{(0.08)}{-0.31\Delta_4^2 p_{t-4}} \; \underset{(0.06)}{-0.34\Delta_4^2 p_{t-8}} \\ & \underset{(0.01)}{+0.05 oil}\end{aligned} \tag{4.4}$$

$T = 75(4) - 97(4), R^2 = 0.49, \hat{\sigma} = 0.01, DW = 1.29, F_{AR}(5,79) = 2.75,$
$F_{ARCH}(4,76) = 3.57, \chi_N^2(2) = 7.70,\; F_{RESET}(1,83) = 6.82,$
$F_H(9,74) = 6.73,$

[10] See Hall, Mizon and Welfe (2000) for a recent exposition on this issue.

[11] *oil* and *d*974 are the dummy variables which capture the effects of the second oil shock on prices by taking unity for 1979(2) and 1980(1) and of the recent financial crisis on the interest rate by taking unity for 1997(4), respectively.

$$
\begin{aligned}
\Delta_2 Ra_t = \; & 1.12\Delta_2 Ra_{t-1} & -1.04\Delta_2 Ra_{t-2} & \; +0.71\Delta_2 Ra_{t-3} \\
& (0.10) & (0.14) & \; (0.14) \\
& -0.22\Delta_2 Ra_{t-4} & +4.99d974 & \\
& (0.11) & (1.51) &
\end{aligned} \tag{4.5}
$$

$T = 75(4) - 97(4), R^2 = 0.61, \widehat{\sigma} = 1.51, DW = 1.80, F_{AR}(5,79) = 2.73,$
$F_{ARCH}(4,76) = 8.50, \chi^2_N(2) = 13.45, \; F_{RESET}(1,83) = 4.67,$
$F_H(9,74) = 0.69.$

Figure 4.3(a) and (b) plot the recursive one-step and break-point Chow tests of equations (4.4) and (4.5), respectively. Although the break-point Chow tests of (4.4) show no changes in parameters, the one-step Chow tests of (4.4) and the one-step and break-point Chow tests of (4.5) reject parameter constancy, particularly in the early 1980s when financial liberalization began. These results indicate that the data generating processes of the two variables have been substantially influenced by financial liberalization over the sample period. To examine structural breaks in the marginal processes in detail, two dummy variables were developed to capture regime changes. The dummy DS, which is a zero/one shift dummy beginning at 1981(1), was created to capture the period of financial deregulation and added to equation (4.4). The dummy $d821$, which takes unity for 1982(1)-(2) and is zero otherwise, was created to capture the effect of a massive reduction of regulated bank interest rates on market interest rates and added to equation (4.5). The reestimated marginal models (4.6) and (4.7) indicate that the added dummies are significant and considerably improve the non-constancy of equations (4.4) and (4.5). They also fit the observed data well over the sample period and pass standard diagnostic tests for model specification, including parameter constancy. This evidence apparently suggests that the marginal processes has been influenced by regime shifts. This improvement of the parameter non-constancy of (4.4) and (4.5) complementarily supports the results of the preceding recursive Chow tests which have found substantial non-constancies in the marginal processes for the conditioning variables.

$$\begin{array}{lllll} \Delta_4^2 p_t = & 0.23\Delta_4^2 p_{t-2} & +0.26\Delta_4^2 p_{t-3} & -0.32\Delta_4^2 p_{t-4} & -0.34\Delta_4^2 p_{t-8} \\ & (0.08) & (0.08) & (0.08) & (0.06) \\ & +0.05oil & -0.003DS & & \\ & (0.01) & (0.001) & & \end{array} \quad (4.6)$$

$T = 75(4) - 97(4), R^2 = 0.51, \widehat{\sigma} = 0.01, DW = 1.31, F_{AR}(5, 78) = 2.17,$
$F_{ARCH}(4, 75) = 2.65, \chi_N^2(2) = 5.48,\ F_{RESET}(1, 82) = 5.51,$
$F_H(10, 72) = 4.96,$

$$\begin{array}{llll} \Delta_2 Ra_t = & 1.10\Delta_2 Ra_{t-1} & -0.95\Delta_2 Ra_{t-2} & +0.61\Delta_2 Ra_{t-3} \\ & (0.08) & (0.11) & (0.12) \\ & -0.29\Delta_2 Ra_{t-4} & -6.05d821 & +5.09d974 \\ & (0.09) & (0.94) & (1.24) \end{array} \quad (4.7)$$

$T = 75(4) - 97(4), R^2 = 0.74, \widehat{\sigma} = 1.24, DW = 1.77, F_{AR}(5, 78) = 1.36,$
$F_{ARCH}(4, 75) = 2.28, \chi_N^2(2) = 3.85,\ F_{RESET}(1, 82) = 0.19,$
$F_H(10, 72) = 0.96.$

With these effects of financial liberalization on prices and the interest rate, our major concerns are whether equation (4.3) is invariant to the determinants of in-sample non-constancies in equations (4.4) and (4.5) and whether the contemporaneous variables $\Delta_4^2 p_t$ and $\Delta_2 Ra_t$ are weakly exogenous. To examine the former condition, we added the created dummies to (4.3) and tested their statistical significance. The joint test $F(4, 71) = 1.70$ shows that they are insignificant at the 5% level and so don't affect the stability of the conditional model across regimes. To examine the weak exogeneity properties of $\Delta_4^2 p_t$ and $\Delta_2 Ra_t$, the residuals of the reestimated marginal models were calculated and their explanatory powers were tested in (4.3).[12] The test $F(2,71) = 0.86$ strongly rejects their significance at the 5% level, confirming that $\Delta_4^2 p_t$ and $\Delta_2 Ra_t$ are weakly exogenous. This combined evidence of invariance and weak exogeneity suggests that the conditioning variables $\Delta_4^2 p_t$ and $\Delta_2 Ra_t$ are super exogenous for the parameters of interest. Following Favero and Hendry (1992), the model (4.3) is interpreted as a feedback model which encompasses a whole class of feedforward models and refutes the Lucas (1976) critique.[13]

4.6 Comparison with a Feedforward Model

In the preceding section we investigated the invariance of the ECM model (4.3) and clarified it as a feedback model using the test of super exogeneity. However, the structural invariance doesn't mean that the model is the best one we can obtain. There are many alternative models congruent with data, particularly in the case of time series based modelling. One of the necessary conditions for equation (4.3) to be reliable is in its ability whether it can encompass the results obtained by rival models (Mizon and Richard, 1986). An appropriate rival model to (4.3) would be a type of rational expectations models, since these kinds of models are invariant to regime changes but are formulated in completely different ways - that is, data-based vs theory-based and backward-looking vs forward-looking.[14] In this section we formulate a feedforward model which has structural 'deep' parameters invariant to regime changes and compare it with (4.3) by conduct-

[12] Note that the long-run exogeneity properties of $\Delta_4 p_t$ and Ra_t in the cointegration analysis do not guarantee that the short-run parameters in (4.3) are also exogenous. Since there still exists the possibility of cross-equation restrictions on the short-run dynamics, it is necessary to examine the weak exogeneity properties of the short-run conditioning parameters.

[13] Recently, Psaradakis and Sola (1996) show that in finite samples, the power of tests in this type largely depends on the specification of marginal models.

[14] See Lucas (1976) for the invariance of rational expectations models.

ing encompassing tests. A model in this type can be formulated in a number of ways. But we follow two-step procedures applied by Cuthbertson and Taylor (1987), since our other aims are to look at whether AR forecasting equations for expected variables (which are expected to be equivalent to marginal models in the previous section) are modelled constantly and to show the importance of the formulation of marginal models for the test of super exogeneity when regime changes are not substantial.

Assume that the actual real money holdings of economic agents are divided into the sum of planned and unplanned balances $(m-p)_t = (m-p)_t^p + (m-p)_t^u + u_t$. The planned component $(m-p)_t^p$ would be chosen so as to minimize the expected discounted present value of a quadratic cost function conditioned on information available at time $t-1$:

$$\begin{aligned} Min\ C &= E_{t-1} \sum_{i=0}^{\infty} \delta^i \{ a_0 [(m-p)_{t+i} - (m-p)_{t+i}^{*}]^2 \\ &\quad + a_1 [(m-p)_{t+i} - (m-p)_{t+i-1}]^2 \}, \end{aligned} \tag{4.8}$$

where δ is the subjective discounting factor, $(m-p)_t^*$ is the desired long run money balance which could be written as $(m-p)_t^* = X_t \alpha'$ where $X_t = (y_t, Ra_t, Rm_t)$ and $\alpha = (\alpha_y, \alpha_{Ra}, \alpha_{Rm})$, and a_0 and a_1 are the relative cost weights of disequilibrium from the desired level and of adjusting money holdings, respectively. Minimizing (4.8) yields the following rational expectational forward-looking model:

$$\begin{aligned} (m-p)_t &= \lambda (m-p)_{t-1} + (1-\lambda)(1-\lambda\delta) \sum_{j=0}^{\infty} (\lambda\delta)^j \alpha' X_{t+j}^e \\ &\quad + \beta'(X_t - X_t^e) + u_t, \end{aligned} \tag{4.9}$$

where λ is the stable root of the system from Euler's equation and $X_{t+j}^e =$

$E(X_{t+j}|I_{t-1})$.[15] If data are $I(1)$ but are cointegrated, the equation is reparameterized into an $I(0)$ feedforward ECM model by ignoring the unanticipated surprise term $(X_t - X_t^e)$:

$$\begin{aligned} \Delta(m-p)_t &= (\lambda-1)[(m-p)_{t-1} - \alpha' X_{t-1}] \\ &\quad +(1-\lambda)\sum_{j=0}^{\infty}(\lambda\delta)^j \Delta\alpha' X_{t+j}^e + u_t. \end{aligned} \tag{4.10}$$

The main difference between (4.10) and (4.2) is that the latter augments the term $(m-p)_{t-1} - \alpha' X_{t-1}$ with first differenced lag terms ΔX_{t-j} which are modelled through data-based statistical tests, whereas the former augments the cointegrated error term with forward difference terms ΔX_{t+j}^e which are modelled through theory-based cross restrictions. If the expectational terms ΔX_{t+j}^e are generated by ΔX_{t-j}, the two models have identical residuals in infinite samples and so are observationally equivalent.

In order to estimate equation (4.10), two-step procedures being commonly used in the literature were applied.[16] Under the assumption that expectations are formed autoregressively, AR equations for $\Delta^2 p_t$, Δy_t,and ΔRa_t were first formulated using Wold's forecasting chain rule to predict expected values. Table 4.3 reports the results of the estimation.[17] All the equations are data acceptable, except for nonnormality in $\Delta^2 p_t$ and ΔRa_t. As noted by Hendry (1988), the parameter constancy of forecasting equations in this type is crucial to formulate a structurally constant feedforward model. The recursive break-point Chow tests reported in Figure 4.4 reveal that all the three equations are constant at the

[15] See Cuthbertson and Taylor (1987) for the derivation of this model.

[16] A main problem with this approach is that estimators may be serially correlated. To avoid this problem, alternatively the procedures of GMM or two-step, two-stages least squares can be applied. Since our estimated model doesn't show serial correlation and since our another purpose is to show similar dynamic structures in the marginal models for feedback and feedforward models, we use this estimation procedure.

[17] ΔRm_t is assumed to be perfectly anticipated since it has been regulated by the government.

5% levels over the sample period. Under the assumptions that economic agents forecast four-periods ahead for $\Delta^2 p_t$ and Δy_t and only the current period for ΔRa_t and that the discounting value of the future expectations terms is 0.90, corresponding to an annual rate of time preference of approximately 40 per cent, the following feedforward ECM was obtained:[18]

$$\begin{array}{lllll}
\Delta(m-p)_t = & -1.03 & -0.14(m-p)_{t-1} & +0.19y_{t-1} & -0.002Ra_{t-1} \\
 & (0.37) & (0.04) & (0.06) & (0.001) \\
 & +0.001Rm_{t-1} & -0.31se\Delta^2 p_t & +0.03se\Delta y_t & -0.01se\Delta Ra_t \\
 & (0.001) & (0.16) & (0.05) & (0.003) \\
 & -0.08S_{1t} & -0.02S_{2t} & -0.01S_{3t} & \\
 & (0.02) & (0.01) & (0.01) &
\end{array} \tag{4.11}$$

$T = 74(4) - 97(4), R^2 = 0.648, \widehat{\sigma} = 0.0232, DW = 1.88, F_{AR}(5,77) = 0.66,$
$F_{ARCH}(4,74) = 0.57, \chi^2_N(2) = 0.19,\ F_{RESET}(1,81) = 0.24,$
$F_H(17,64) = 1.13,$

where $se\Delta^2 p_t = \sum_{j=0}^{4}(\lambda\delta)^j \Delta^2 p^e_{t+j}, se\Delta y_t = \sum_{j=0}^{4}(\lambda\delta)^j \Delta y^e_{t+j},\ se\Delta Ra_t = \Delta Ra^e_t$, and S_{it} is seasonal dummies. The model accepts cross equation restrictions for the test of rationality with $F(8,74) = 1.34$ and performs well in terms of diagnostic tests for mis-specification. The estimated parameters are consistent with *a priori* expected signs and magnitudes. Furthermore, the recursive break-point Chow tests reported in Figure 4.5 clearly show parameter constancy. Overall, this theory-based feedforward model seems to coherently characterize the structural demand function for money. Note that the form of equation (4.11) is equivalent to that of equation (4.3). The only difference between them is in

[18] A low figure of the subjective discounting value δ reflects a large uncertainty concerning the expectations of $\Delta^2 p_t$, Δy_t,and ΔRa_t. Note that changes in δ in the range of 1 - 0.9 do not show any differences in the estimated results.

the forms of the dynamics included, which are due to the intrinsic difference between theory-driven models and data-driven models in finite samples but can be essentially reparameterized into the same form in infinite sample.

To compare this forward-looking model with the backward-looking model (4.3), non-nested encompassing tests were applied. If (4.3) encompasses the expectational model, the former has more information in explaining data than the latter and discriminates itself from the interpretation as a feedforward model, or alternatively, if the two models mutually encompass, then they are equivalent and share the same information on data (Mizon and Richard, 1986; Lu and Mizon, 1997). The Cox test extended by Pesaran (1974) into a linear model marginally accepts the null hypothesis of which the feedback model (H_0) encompasses the feedforward model (H_1) with an asymptotic normal distribution $N(0, 1) = -2.45$ at the 1% level, but strongly rejects the alternative hypothesis with $N(0, 1) = -11.10$ at the 5% level.[19] Similar results were obtained from the tests of parameter encompassing (Mizon and Richard, 1986). The null hypothesis of which H_0 encompasses H_1 is accepted with the test statistic $F(3, 78) = 1.19$ at the 5% level, but the reverse null is rejected with $F(4, 78) = 7.91$. Overall, the empirical results of encompassing tests suggest that equation (4.3) performs better than the theory-based expectations model and apparently discriminates itself from the interpretation as a feedforward model. This evidence complementarily supports the conclusion derived in the previous section.[20]

[19] Pesaran (1974) argues that the Cox type test often over-rejects both hypotheses in finite samples.

[20] Indeed, the conclusion derived from the previous section is based on the non-constancy of the underlying marginal models associated with the constant conditional model (4.3). However, while one-step Chow tests clearly show the non-constancy of the marginal models, break-point Chow tests accept the null hypothesis of parameter constancy in $\Delta_4^2 p_t$ and marginally reject it in the case of $\Delta_2 Ra_t$. This statistical weak evidence to find evident regime changes does not rule out that there may be many alternative ways by which the equations for marginal processes can be constructed constantly. This implies that the empirical results shown in the previous section is suggestive rather than conclusive. In the absence of a substantial exhibition of regime changes, encompassing tests between alternative models may provide a complementary evidence in discriminating observational equivalence.

4.7 Conclusions

This chapter has investigated the issue of whether financial liberalization in Korea has caused the demand function for money to be unstable, by applying a dynamic modelling approach. Since the valid inferences and reliable analysis depend on the statistical validity and appropriate interpretation of the underlying model, the paper extensively examined the statistical properties of the baseline ECM model. Particular attention was paid to issues of parameter constancy, invariance, observational equivalence and encompassing, under regime changes. The empirical results show that the estimated parameters are constant in spite of various financial reforms and institutional changes over the sample period. This evidence, which is contrasted with the finding of Tseng and Corker (1991), indicates that the stability of the demand for money has not been broken down by financial deregulation. The test of super exogeneity with non-constant marginal models shows that the estimated ECM model is invariant to policy changes and so is immune from the Lucas (1976) critique. As shown by Hendry (1988), this result implies that the model is a feedback model which encompasses a whole class of expectations models. Furthermore, encompassing tests show that our ECM model accounts for the results of an alternative forward-looking model which is formulated on the basis of a multi-period quadratic loss function. This evidence suggests that the ECM model delivers better information on data than the theory-based expectations model and complementarily confirms that it is a class of backward-looking models.

Overall, this chapter shows that financial liberalization has substantially changed the behavior of price levels, real income, and interest rates, but not the stability of money demand behavior, and highlights the importance of model specification in formulating econometric models. A methodological implication of super exogeneity and encompassing tests in this chapter is that a considerable care is required to formulate marginal models for the test of super exogeneity, if the effects of regime changes on the parameter non-constancy of marginal models are not substantial, because of a possibility of 'spurious' non-constancy in marginal models. In this special case, this study recommends to perform both encompassing and super exogenentiy tests to discriminate observational equivalence in ECM type models.

Table 4.1 Augmented Dickey-Fuller test statistics

Var	m_t	Δm_t	$\Delta^2 m_t$	p_t	Δp_t	$\Delta^2 p_t$	$\Delta_4 p_t$	$\Delta_4^2 p_t$
	c,t,s	c,s	s	c,t,s	c,s	s	c,s	s
lag	0	3	5	1	2	5	0	5
t value	-1.92	-2.41	-5.40	-3.06	-2.75	-5.08	-2.19	-2.89
c.v.5%	-3.46	-2.89	-2.89	-3.46	-2.89	-2.89	-2.89	-2.89
c.v.1%	-4.06	-3.50	-3.50	-4.06	-3.50	-3.50	-3.50	-3.50
Var	y_t	Δy_t	Ra_t	ΔRa_t	Rm_t	ΔRm_t	$(m-p)_t$	$\Delta(m-p)_t$
	c,t,s	c,s	c	–	c	-	c,t,s	c,s
lag	3	5	0	4	0	5	0	3
t value	-1.01	-3.48	-1.64	-3.58	-1.27	-2.67	-2.94	-3.54
c.v.5%	-3.46	-2.89	-2.89	-1.94	-2.89	-1.94	-3.46	-2.89
c.v.1%	-4.06	-3.50	-3.50	-2.59	-3.50	-2.59	-4.06	-3.50

Note: c,t, and s represent a constant term, a time trend, and seasonal dummies included in the testing equation.

Table 4.2 Cointegration analysis

(a) Cointegration test statistics

Eigenvalues	0.27	0.21	0.14	0.07	0.007
Hypotheses	$\gamma = 0$	$\gamma \leq 1$	$\gamma \leq 2$	$\gamma \leq 3$	$\gamma \leq 4$
Max statistic	29.73	22.97	15.11	7.44	0.68
95% c.v.	33.50	27.10	21.00	14.10	3.80
Trace statistic	75.94	46.20	23.23	8.20	0.68
95% c.v.	68.50	47.20	29.70	15.40	3.80

(b) Weak exogeneity test statistics

Variables	$(m-p)_t$	y_t	Ra_t	Rm_t	$\Delta_4 p_t$	$\{y_t, Ra_t, Rm_t, \Delta_4 p_t\}$
$\chi^2(\cdot)$	5.26	1.47	0.55	0.30	0.09	1.76
$p-value$	0.02	0.23	0.46	0.58	0.75	0.78

Note: The test statistics of weak exogeneity have a $\chi^2(1)$ distribution for $(m-p)_t$, y_t, Ra_t, Rm_t,and $\Delta_4 p_t$, respectively, and a $\chi^2(4)$ for the joint test $\{y_t, Ra_t, Rm_t, \Delta_4 p_t\}$.

Table 4.3 Autoregressive forecasting equations

$$\Delta^2 p_t = \underset{(0.10)}{-0.53\Delta^2 p_{t-1}} \; \underset{(0.10)}{-0.34\Delta^2 p_{t-2}} \; \underset{(0.10)}{-0.17\Delta^2 p_{t-3}} \; \underset{(0.08)}{-0.13\Delta^2 p_{t-4}} \; \underset{(0.01)}{+0.05 doil}$$

$T = 74(4) - 97(4), R^2 = 0.60, \widehat{\sigma} = 0.01, DW = 1.98, F_{AR}(5,79) = 2.87,$
$F_{ARCH}(4,76) = 1.97, \chi^2_N(2) = 10.57, \; F_{RESET}(1,83) = 4.59,$
$F_H(12,71) = 1.09.$

$$\Delta y_t = \underset{(0.08)}{-0.38\Delta y_{t-1}} \; \underset{(0.08)}{-0.36\Delta y_{t-2}} \; \underset{(0.08)}{-0.37\Delta y_{t-3}} \; \underset{(0.08)}{+0.58\Delta y_{t-4}} \; \underset{(0.03)}{-0.11 d804}$$

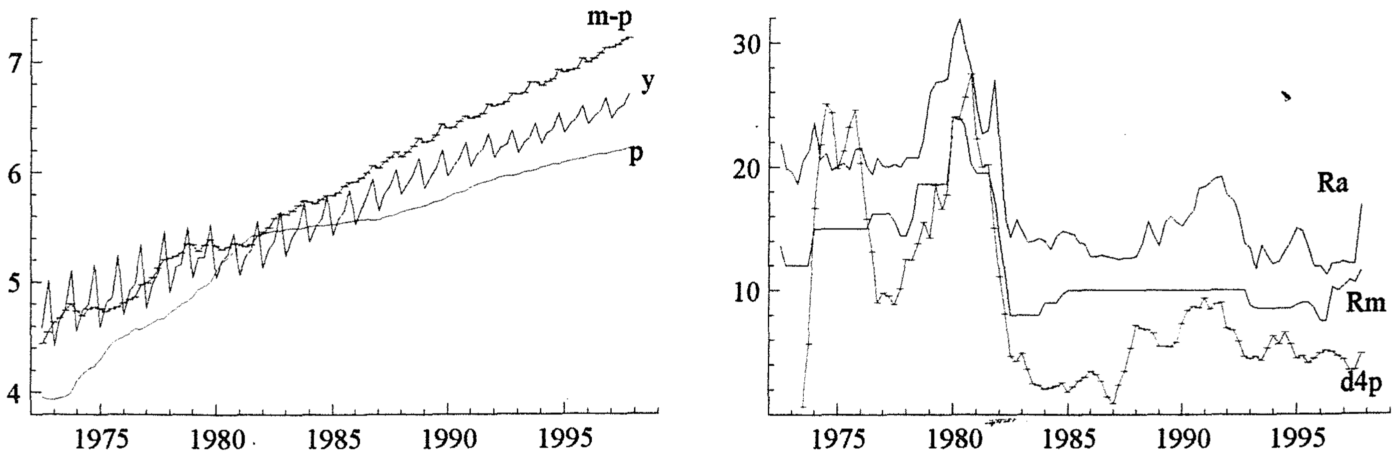
$T = 74(4) - 97(4), R^2 = 0.99, \widehat{\sigma} = 0.02, DW = 1.87, F_{AR}(5,79) = 0.87,$
$F_{ARCH}(4,76) = 1.82, \chi^2_N(2) = 2.28, \; F_{RESET}(1,83) = 0.01,$
$F_H(12,71) = 1.92.$

$$\Delta Ra_t = \underset{(0.08)}{0.21\Delta Ra_{t-1}} \; \underset{(0.82)}{-5.47 d821} \; \underset{(1.15)}{+4.80 d974}$$

$T = 74(4) - 97(4), R^2 = 0.44, \widehat{\sigma} = 1.15, DW = 1.85, F_{AR}(5,85) = 0.96,$
$F_{ARCH}(4,82) = 2.67, \chi^2_N(2) = 10.28, \; F_{RESET}(1,89) = 0.83,$
$F_H(4,85) = 0.12.$

Note: The equations of $\Delta^2 p_t$ and Δy_t include seasonal dummies, but not reported.[21]

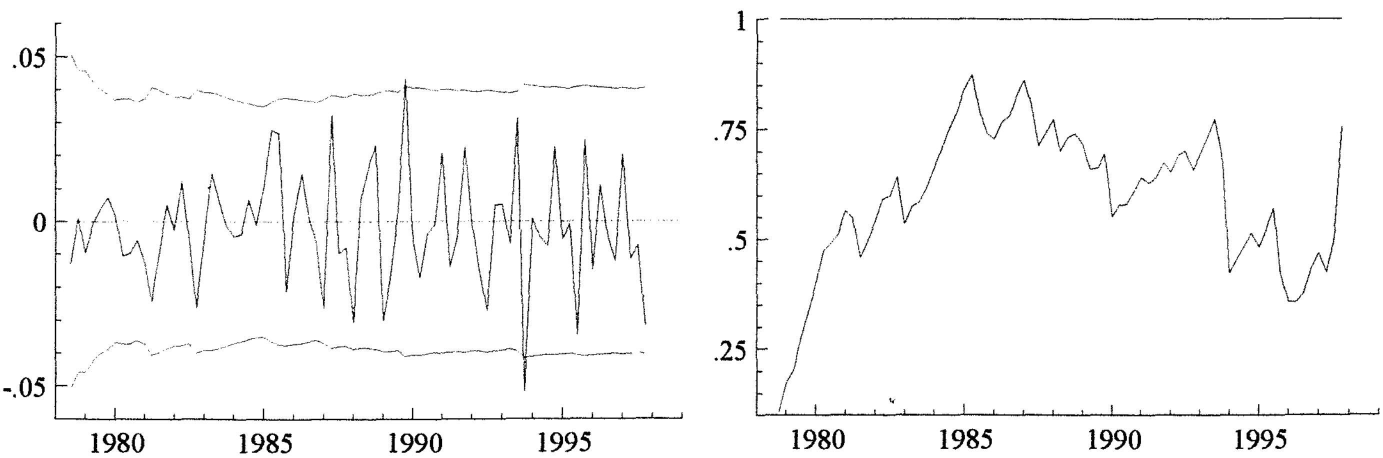
[21] Four dummies were added to consider outliers on the observed data. *doil, d804, d821*, and *d974* take value unity in 1980(1) to capture the second oil shock, in 1980(4) to capture a severe recession caused by political instability, in 1982(1)-(2) to capture the effect of a massive reduction of regulated bank interest rates on the market interest rate, and in 1997(4) to capture a financial crisis, respectively.



(a) real money (m-p), real income (y), and the price level (p).

(b) annual inflation (d4p), the bank deposit rate (Rm) and the three-year corporate bond rate (Ra).

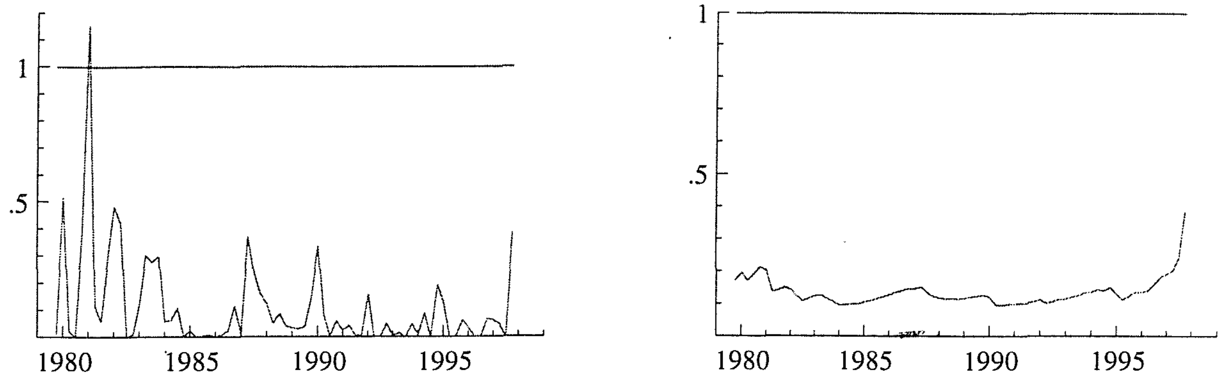
Figure 4.1 Graphs of data series.



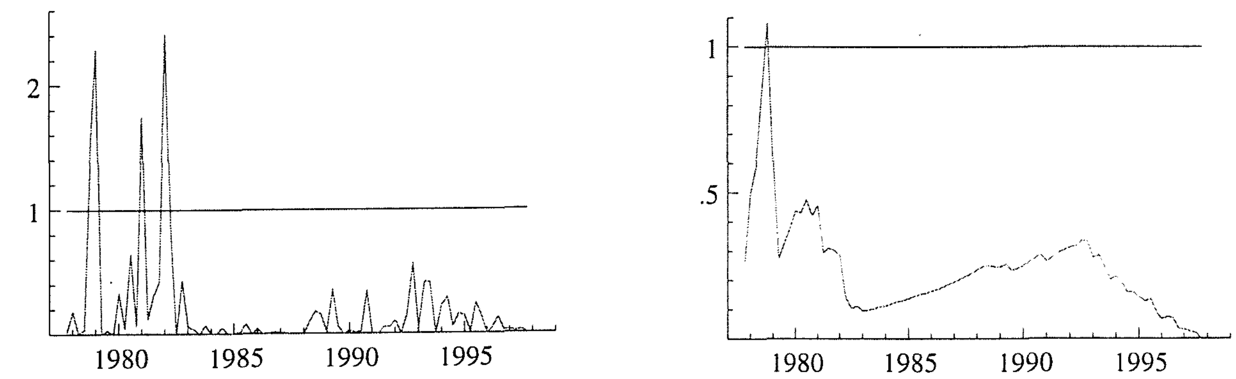
(a) one-step residuals with plus-or-minus two standard errors.

(b) break-point Chow statistics with one-off 5% critical values.

Figure 4.2 Tests of parameter constancy.

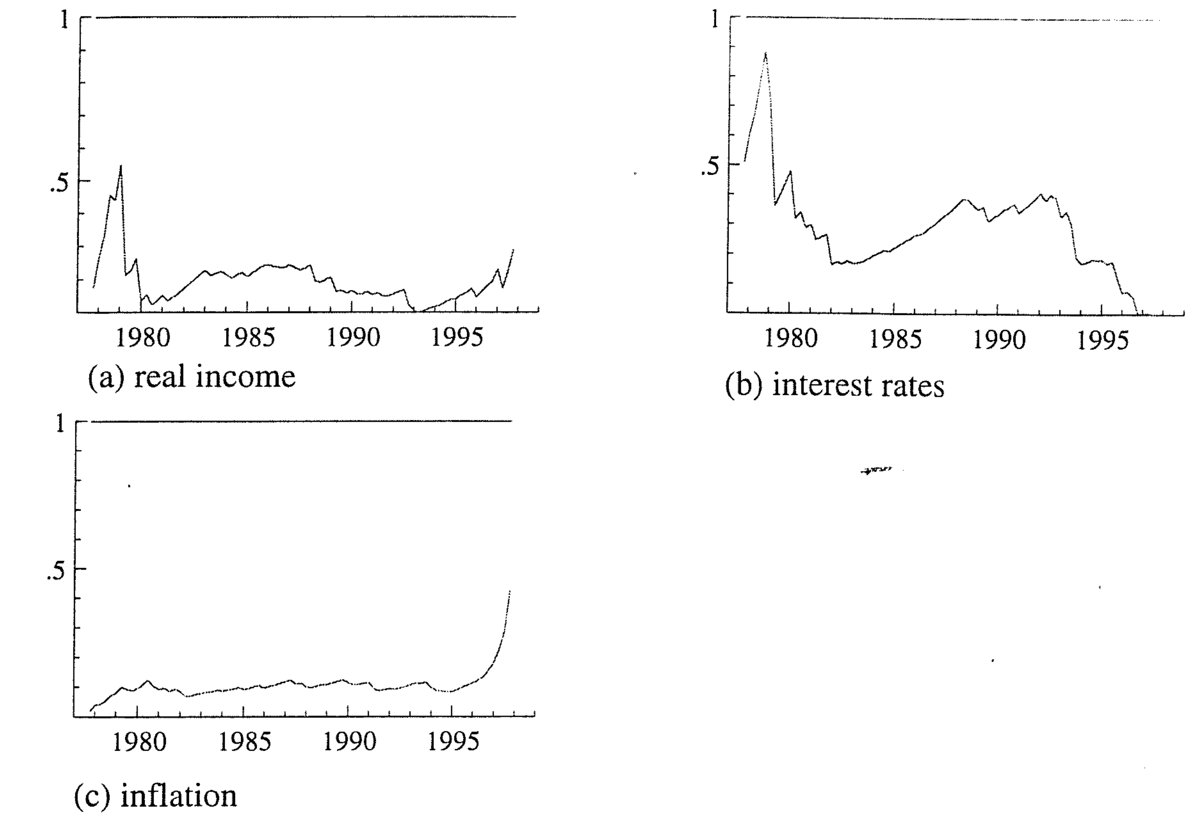


(a) one-period ahead and break-point Chow statistics of (4.4) with their 5% critical values.



(b) one-period ahead and break-point Chow statistics of (4.5) with their 5% critical values.

Figure 4.3 Recursive Chow statistics of marginal equations.

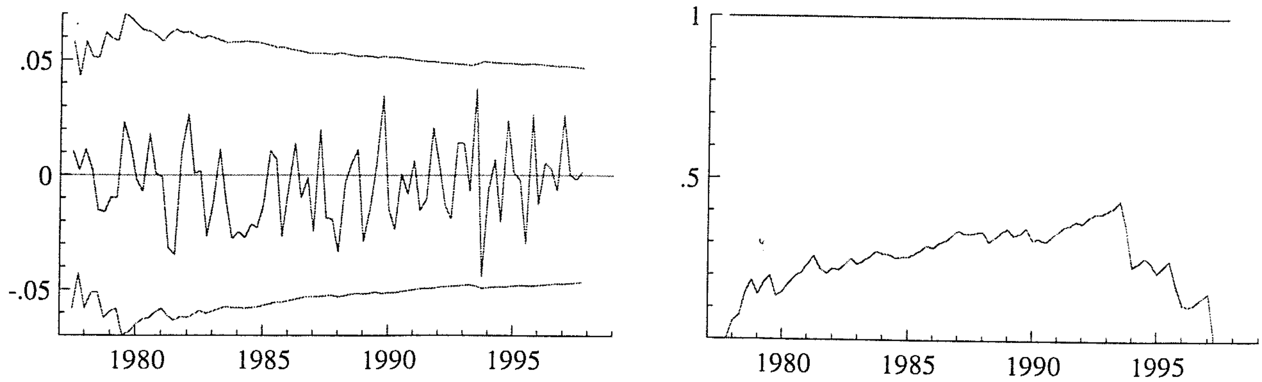


(a) real income

(b) interest rates

(c) inflation

Figure 4.4 Break-point Chow statistics of forecasting equtions.



(a) one-step residuals with plus-or-minus two standard errors

(b) break-point Chow statistics with one-off 5% critical values

Figure 4.5 Recursive statistics of the forward-looking ECM.

Chapter 5

Identification of Monetary Policy

5.1 Introduction

One of the longest debated issues in macroeconomics is how to treat policy variables. That is, is policy endogeous or exogenous? (see Goldfeld and Blinder (1972)). If policy contemporaneously *affects* the real sector but there is no feedback from the economy to policymakers' decisions within the current period, then policy is identified as exogenously determined. On the other hand, if policy contemporaneously *responds* to current developments in the economy but affects real variables such as prices and output only with lags, policy is identified as endogenously determined. A relevant importance of this issue in the case of orthogonalized impulse response analysis based on the Choleski factorization in vector autoregression (VAR) models is that if a policy variable measuring the stance of policy belongs to the former case, the exogeneity of the policy variable is identified by placing it first in the recursive ordering of variables. If the case is the latter, the reverse ordering is used in the literature.

However, a well-known problem with this approach is that unless the covariance matrix of the VAR residuals is diagonal, different orderings produce different Choleski factorizations of the matrix. As a consequence, the empirical results based on this approach are very sensitive to the ordering of the variables considered. Even under the uncorrelatedness of the VAR residuals (diagonal covariance matrix), the unreliability of the empirical results is not completely disappeared because of the lack of 'causal' interpretation between variables (see Cooley and LeRoy (1985)). What all of these arguments imply is the necessity of statistically identifying policy actions into either those that are endogenous or those that are exogenous in a VAR-based analysis for the effects of policy. Of course, the correct identification is impossible. However, at least the assumed property of policy stance should be formally tested, and then a reliable empirical result can be obtained. Unfortunately, in spite of the importance of this matter, most of the current studies in the literature do not attempt to identify policy variables by using formal statistical tests. Instead, *ad hoc* procedures, such as theory-based identification (structural VAR models), order changing, or Granger causality tests (even if the tests have nothing to do with causality), are commonly applied. Thus, the empirical results derived from these *ad hoc* procedures are very sensitive to the preliminary assumptions and, if the assumptions are not valid, lead to an invalid conclusion by treating policy variables as exogenous even if they are not.[1]

Considering the deficiencies of the current approach, in this chapter we discuss how monetary policy can be statistically identified in a VECM model by using *formal tests* and demonstrate that policy identification based on the Choleski factorization in the covariance matrix of the VECM residuals alone is not valid if proper corresponding restrictions are not imposed on the cointegration space. To do this, the concept of weak exogeneity proposed by Richard (1980) and Engle *et al.* (1983) is applied. Even though this concept has been popularly applied to formulate a single equation-based conditional model in the literature, little research has paid attention to the usefulness of the concept for the identification

[1] See Christiano *et al.* (1999) for a recent survey on the VAR literature.

of policy stance in VAR models. We focus our attention on the testing procedure suggested by Johansen (1992) within a VECM framework when data are nonstationary.[2] According to Engle *et al.* (1983), a set of variables is weakly exogeneous for the parameter of interest if it is possible to factorize the joint distribution such that the parameters of interest are determined in the distribution of the endogenous variables conditional on the weakly exogeneous variables and the parameters in the distribution for the weakly exogeneous variables are variation free in the sense that the parameters in conditional and marginal models are not subject to cross restrictions. A methodological usefulness of this concept in a VAR model is that the statisitcal properties of weak exogeneity provide a 'recursive structure' between the blocks of endogenous and exogenous variables. Thus, the exogeneity hypotheses of policy indicators is directly tested by applying weak exogeneity tests and the identified exogeneity is straightly used to interpret the 'causal' relation between policy and nonpolicy variables (see Richard, 1980). In this context, the empirical results based on this approach are expected to be more reliable than those obtained from the current approach which *a priori* assumes the *predeterminedness* of policy variables without formal statistical tests.

The practicability of the above discussion is experimented by applying the approach to the Korean experience. The economy is interested in applying the approach in several ways. Unlike many developed countries, the Korean monetary authority has officially used the stock of M2 money as the main policy variable over the past two decades, partly because of the non-existence of market interest rates for policy and partly because of a low interest rate policy. So, without any difficulties in finding alternative policy indicators, we can concentrate on the issue identifying the stance of policy actions.[3] Secondly, In the literature the broad M2

[2] See Ohanian (1988) and Phillips (1998) for the potential problems of using level variables in a VAR-based analysis when data are nonstationary.

[3] It is well-known that empirical results in VAR-based studies are also very sensitive to the choice of policy indicators. Those commonly used in the literature are the stock of M1 money (Sims, 1980), the Federal fund rate (Bernanke and Blinder, 1992), short-term interest rates

money is generally treated as endogeously determined. However, in the case of Korea, the money stock is a mixture driven by both economic agents' demand and the monetary authority's supply. Thus a reasonable method for the identification of the policy variable should be applied. Finally, a reason of using Korean data is related to the investigation of whether the data of developing countries support the view of real business cycle (RBC) theory that a broad money stock mostly consisted of inside money is endogenously determined and so the reverse causality from output to money holds. To the best of our understanding, there is no empirical study which investigates the suitability of the theory to developing countries. In this context, the empirical evidence of Korean data may provide a good benchmark example to other developing countries.

In conducting the research, we decompose the nominal M2 money stock into outside money which consists of monetary assets of the private sector in the form of claims on the public sector and inside money which consists of monetary assets held within the private sector which are also liabilities of members of that same sector. Then we examine how the two money stocks are determined and which form of money is more influential to the real economic activity. Our particular attention is in the statistical property of inside money, since the money has been determined by both economic agents' demand and the monetary authority's supply. For the M2 money to be a good policy variable, inside money consisting of a large portion of the broad money must be essentially unresponsive to changes in the real sector within the current period. The empirical results suggest that in general, the money stocks have been supplied independently of the demand for money; while exogenously identified inside and outside money significantly affect prices and output, the former has a bigger impact than the latter; and as a result, Korean data do not support the view of RBC theory. In addition to these main results, further interesting findings are that even if we use aggregate money

(Sims, 1992), and non-borrowed reserves (Strongin, 1995).

stocks, our study doesn't suffer from the common problems frequently appeared in earlier VAR-based studies, such as the liquidity puzzle associated with the positive initial responses of interest rates to a positive shock on money (Leeper and Gordon, 1992), the price puzzle related to the positive responses of the price level to a positive shock on interest rates (Sims, 1992), and the weak explanatory power of monetary aggregates on output when a VAR includes interest rates (Sims, 1982a; Litterman and Weiss, 1985).

This chapter is organized as follows. Section 2 analytically shows how the Johansen (1992) test for weak exogeneity can be applied to identifying policy in a VECM model. This approach is then applied using Korean data in Section 3. Section 4 concludes the paper.

5.2 Methodological Framework

In this section we demonstrate how the concept of weak exogeneity proposed by Engle *et al.* (1983) can be usefully applied to identifying policy. Particular attention is paid to the Johansen (1992) procedure in a VECM framework. To see this further, consider that x_t has an autoregressive process of the form

$$A(L)x_t = \epsilon_t, \quad \epsilon_t \sim IN(0, \sum), \tag{5.1}$$

where $A(L)$ is an $n \times n$ matrix polynomial in the lag operator L, $A(L) = \sum_{j=0}^{p} A_j L^j$, and ϵ_t is an $n \times 1$ vector of independently identically distributed zero mean errors and nonsingluar variance matrix $\sum$. If x_t is nonstationary with order $I(1)$ but cointegrated, (5.1) can be reformulated, using the well-known Beveridge-Nelson decomposition formula $A(L) = A^*(L)(1-L) + A(1)L$, into the following structural VECM:

$$A^*(L)\Delta x_t = -A(1)x_{t-1} + \epsilon_t, \tag{5.2}$$

where $A^*(L) = \sum_{i=0}^{p-1} A_i^* L^i$ with $A_i^* = \sum_{j=i+1}^{p} A_j$. Decompose $\Delta x_t = (\Delta y_t : \Delta z_t)'$, and assume that $A_0^* = \begin{pmatrix} A_{11.0}^* & A_{12.0}^* \\ A_{21.0}^* & A_{22.0}^* \end{pmatrix}$ has unit elements on the diagonal of $A_{11.0}^*$ and $A_{22.0}^*$ and that $A(1)$ has a reduced rank $(A(1)) = \gamma < n$ which is factorized as $A(1) = \alpha\beta'$ with $(n \times \gamma)$ matrices for both α and β. Then, with $p = 2$, (5.2) can be rewritten as the following system:

$$\begin{aligned} \begin{bmatrix} \Delta y_t \\ \Delta z_t \end{bmatrix} &= \begin{bmatrix} -\phi_1 \Delta z_t \\ -\phi_2 \Delta y_t \end{bmatrix} + \begin{bmatrix} \Gamma_{11} & \Gamma_{12} \\ \Gamma_{21} & \Gamma_{22} \end{bmatrix} \begin{bmatrix} \Delta y_{t-1} \\ \Delta z_{t-1} \end{bmatrix} \\ &\quad - \begin{bmatrix} \psi_{11} & \psi_{12} \\ \psi_{21} & \psi_{22} \end{bmatrix} \begin{bmatrix} y_{t-1} \\ z_{t-1} \end{bmatrix} + \begin{bmatrix} \epsilon_{1t} \\ \epsilon_{2t} \end{bmatrix}, \end{aligned} \tag{5.3}$$

where $\phi_1 = A_{12.0}^*$, $\phi_2 = A_{21.0}^*$, $\Gamma_{ij} = A_{ij.1}^* (i, j = 1, 2)$, $\psi_{11} = (\alpha_{11}\beta_{11}' + \alpha_{12}\beta_{21}')$, $\psi_{12} = (\alpha_{11}\beta_{12}' + \alpha_{12}\beta_{22}')$, $\psi_{21} = (\alpha_{21}\beta_{11}' + \alpha_{22}\beta_{21}')$, $\psi_{22} = (\alpha_{21}\beta_{12}' + \alpha_{22}\beta_{22}')$, and $\sum = \begin{pmatrix} \Sigma_{11} & \Sigma_{12} \\ \Sigma_{21} & \Sigma_{22} \end{pmatrix}$.

Let Δz_t and Δy_t denote a set of exogenous and endogenous variables, respectively, and consider the case that Δz_t is independent of the contemporaneous developments in Δy_t, possibly because of information lags.[4] If we assume that the parameters of interest are all the parameters of β, then a set of sufficient conditions for the efficient inference of the system (5.3) in which Δy_t is conditioned on Δz_t are $\phi_2 = 0$, $\Sigma_{12} \Sigma_{22}^{-1} - \phi_1 = 0$, and $\alpha_{2j} = 0$ $(j = 1, 2)$. The first two conditions imply that $\Sigma_{12} = \Sigma_{21}' = 0$, and the last condition ensures that the cointegrating vectors do not occur in the equations determining Δz_t, which is effectively an hypothesis of long-run causality from z_t to y_t (see Hendry and Mizon, 1999b). If these conditions hold, the equation for Δz_t is reduced to

[4]This type of 'inside lags' can arise because it takes time for policymakers both to recognize that a shock has occurred and to put appropriate policies into effect. Alternatively, 'outside lags' can be assumed if we believe that policies do not immediately affect slow-moving variables such as spending, income and prices within the current period (see Mankiw (1992)).

$$\Delta z_t = \Gamma_{21}\Delta y_{t-1} + \Gamma_{22}\Delta z_{t-1} + \epsilon_{2t}. \tag{5.4}$$

Substituting (5.4) into the equation Δy_t leads to the following reduced-form equation

$$\begin{aligned} \Delta y_t &= (\Gamma_{11} - \phi_1\Gamma_{21})\Delta y_{t-1} + (\Gamma_{12} - \phi_1\Gamma_{22})\Delta z_{t-1} \\ &\quad -\psi_{11}y_{t-1} - \psi_{12}z_{t-1} + v_{1t}, \end{aligned} \tag{5.5}$$

where $v_{1t} = \epsilon_{1t} - \phi_1\epsilon_{2t}$. Equations (5.4) and (5.5) consist of a standard reduced-form VECM system which identifies policy variables block-recursively, but the main differences from an unrestricted system is that the ECM term doesn't appear in (5.4) and the ECM term in (5.5) is restricted to isolate cross restrictions between variables. This structure shows that policy identification based on the Choleski factorization in the covariance matrix of the unrestricted VECM residuals alone is not valid if proper corresponding restrictions are not imposed on the ECM terms of the equations describing the generation of common trends. Note that while the error terms of the system are recursively correlated, the efficient inference is obtained since the condition of $\sum_{12}\sum_{22}^{-1} - \phi_1 = 0$ ensures zero covariances between the equations for Δz_t and Δy_t.[5]

An important implication of this block-recursive structure is that in an unrestricted reduced-form VECM, the effects of orthogonal shocks to exogenous innovations on endogenous variables can be measured with the impulse responses of variables in the block of Δy_t to changes in ϵ_{2t} by ordering exogenously identified variables firstly.[6] For example, if policy variables are exogenously identified, the variables are placed on the top in ordering, thus letting the variance in pol-

[5] This can be shown by $E[\epsilon_{2t}(\epsilon_{1t} - \phi_1\epsilon_{2t})] = E[(\epsilon_{1t}\epsilon_{2t}) - \phi_1 E(\epsilon_{2t}^2)]$
$= \sum_{12} - \sum_{12}/\sum_{22} \cdot \sum_{22} = 0$ since $\phi_1 = \sum_{12}\sum_{22}^{-1}$.

[6] See Sims (1986) for an advantage of block recursive orderings.

icy variables be explained primarily by their own innovations. Conversely, if monetary authorities sensitively react to current economic developments and so policy variables are endogenously identified, then the reverse ordering is valid.[7] This block-recursive causal ordering should be done with the identification of the cointegrating vectors, which is normally achieved via a priori restrictions based on economic theory or previous empirical evidence. For example, the necessary order condition for β to be identified is the existence of γ^2 restrictions, including normalization, α and β. However, it is common for there to be more than γ^2 linearly independent restrictions and so the parameters are over-identified.

The main difference of this approach from the orthogonalized semi-structural VAR approach is that policy variables are formally identified through a statistical test. However, note that since the structure allows contemporaneous correlations between the error terms of equations within each block, the block-recursive identification doesn't provide complete ordering. The conventional problem of ordering can be occurred within each block, but is not the problem of the whole system. In this context, this approach is very useful when we consider that most of the current controversies in VAR-based policy analysis stem from a pre-assumed identification of unobservable policy stance. Instead of orthogonalized impulse responses based on the Choleski factorization of the covariance matrix, a unit or a standard-error-based impulse response, which is independent of ordering problems, can be alternatively applied. Even though these non-orthogonal impulse responses ignore potential correlatedness within variables, the block-recursively identified structure through statistical tests provides a legitimacy of the causal interpretation between variables. To identify policy variables in an unrestricted reduced-form VECM, the testing procedure suggested by Johansen (1992) is directly applicable. Johansen shows that given Δz_t, a necessary and sufficient

[7] Note that the problem of variable ordering is not immaterial even under a diagonal covariance matrix in residuals, because the contemporaneous uncorrelatedness of variables necessarily does not imply the 'causal' relations between variables (Cooley and LeRoy, 1985).

condition for weak exogeneity is simply $\alpha_{2j} = 0$ ($j = 1, 2$), if the parameters of interest are β. To test this type of restrictions, a conventional likelihood ratio (LR) test can be used by estimating the system with and without imposing $\alpha_{2j} = 0$. Alternatively, a Wald test can be used from the unrestricted VECM estimates of α_{2j} in the full system under the null hypothesis of $H_0 : \alpha_{2j} = 0$.[8]

5.3 Empirical Results

In this section we apply the approach discussed in the previous section using Korean data, with a particular attention on the relative importance of inside and outside money on economic activities.[9] Over the sample period, the Bank of Korea (BOK) has used the stock of M2 money as the main policy variable, which includes currency in circulation, demand deposits, and time and savings deposits. The alternatives such as interest rates and the monetary base have not been used as policy variables, partly because of the non-existence of market interest rates suitable for the implementation of policy and partly because of a low interest rate policy to promote a rapid economic growth of the economy, but used as supplementary indicators. To hit the predetermined M2 money targets, the bank has strictly controlled the broad money via direct credit controls, changes in reserve requirements, and banks' accessibility to rediscounting facilities (see Park (1994) for more details). Because of this strict control by the government, the policy variable has been often regarded as one of the typical exogenous variables in most previous empirical studies for policy effects. However, this may not be true in all cases. In the literature it is widely accepted that a significant portion of variance in a broad money stock is due to the policy actions accommodating demand-induced innovations rather than policy-induced supply innovations. Particularly,

[8] See Podivinsky (1992) who suggests an alternative approximate F-type test in small samples.

[9] The recent literature in this line can be found in Manchester (1989), Freeman and Huffman (1991), Chari *et al.* (1995), and Jefferson (1997).

this view is further pervasive in the case of the broad M2 money of which is mostly composed of inside money. King and Plosser (1984) argue that much of the contemporaneous correlation of economic activity and money is apparently with inside money, of which variations are resulted from future productivity disturbances. In this view, money is endogenously determined by demand forces. So the reverse causality from output to money holds. In this case, regarding the broad money stock as exogenously determined may lead to an invalid conclusion in policy analysis. An important implication of these two conflicting views on policy variables is that in examining policy effects, it is necessary to identify policy variables through a formal statistical test, either as the exogenous ones determined by discretionary policy actions or as the endogenous ones determined by accommodating contemporaneous macroeconomic developments.

To examine the observed positive correlation between money and output in Korea, the approach discussed in the previous section is applied. We use quarterly data spanning from 1972(3) to 1998(2). The data include real GDP (Y), the consumer price index (P), a three-year corporate bond rate (R), the stock of nominal monetary base (M_o) as outside money, and the stock of nominal M2 (M_i) excluding currency in circulation as inside money.[10] Except for R, all variables were seasonally adjusted by using the X-11 procedure and transformed into logarithms. Hereafter lower case letters denote the logs of the corresponding capitals. Data were obtained from the various issues of the BOK's monthly bulletin. Initially, to check the nonstationarity of the data series, augmented Dickey-Fuller (ADF) tests were conducted with up to five auxiliary lags. The test results suggest that all variables seem to be $I(1)$.[11] Based on these results, we tested

[10] In the literature, the use of short-term interest rates is preferred. However, in the case of Korea, the economy doesn't have any good market interest rates because of underdeveloped financial markets over the sample period. With this reason, we use the corporate bond rate which is considered as reflecting most adequately financial market conditions.

[11] The values of the ADF test statistics are -1.59(5), -3.07(5), -1.67(5), -1.81(5), and -2.02(5) for the levels of y_t, p_t, m_{ot}, m_{it}, and R_t, respectively, and -3.68**(3), -3.81**(1), -3.57**(4), -4.03**(1), and -3.88**(4) for the corresponding first differences. ** and $(\cdot)$ denote significance

cointegration using the Johansen (1988) procedure. A fourth-order VAR chosen from a series of system-based LR tests was initially estimated with a constant term and trend. Following Hendry (1995), the constant term was not restricted to lie in the cointegration space, but trend was forced to lie in the space in order to avoid of a quadratic deterministic trend in the levels of variables. The test results reported in Table 5.1(a) suggest that the null hypothesis of cointegrating rank $\gamma = 0$ is rejected at the 5% levels by both the max and trace tests with the critical values taken from Osterwald-Lenum (1992). Therefore, we assume that there is only one cointegrating vector.

Table 5.1(b) reports the test statistics of weak exogeneity (Johansen, 1992). The tests conducted by imposing zero restrictions on the loading matrices, corresponding to the cointegrating vector, show that while m_{ot}, m_{it}, and R_t can be treated as exogenous in the system, y_t and p_t are endogenously determined at the 10% and 5% levels, respectively.[12] This evidence shows that there are no contemporaneous effects from the non-policy variables y_t and p_t onto the exogenous variables m_{ot}, m_{it}, and R_t, and explains that the BOK has determined monetary aggregates at its discretion independently of the current developments of macroeconomic situations over the sample period. Interestingly, the exogeneity of inside money (m_{it}) is not consistent with its possible endogeneity in King and Plosser (1984). The statistical exogeneity of R_t, which is otherwise expected to be an endogenous intermediate variable under the official money target policy of the BOK,

at the 5% level and the longest significant lags in the augmentation of the test regression, respectively. In the cases of the former four variables, the tests for levels include a constant term and trend, but the tests for differenced series exclude trend. The tests for R_t and ΔR_t include only a constant term. The critical values were taken from MacKinnon (1991). In chapter 4, the seasonally *unadjusted* nominal money and prices were found to be I(2). However, when the variables are seasonally *adjusted,* they behave as I(1). This seems to reflect the weak testing power of conventional unit root tests.

[12]The finding that the nominal money stocks are exogenous is contrasted to the finding in Chapter 4 that the real money stock is endogenously determined. This evidence is coincide with Friedman's (1959) argument that 'it seems useful to regard the nominal quantity of money as determined primarily by conditions of supply, and the real quantity of money and the income velocity of money as determined primarily by conditions of demand.'

seems to reflect its regulated nature directly and indirectly. From the standardized eigenvectors of the cointegration test, a long run equilibrium equation between variables was derived: $\beta' x_t = y_t + 2.16p_t + 0.05m_{ot} - 1.79m_{it} - 0.03R_t + 0.02t$. However, it is difficult to have an economic meaning from this equation. To identify it economically, we imposed some hypothetical restrictions, following Johansen and Juselius (1992). This involves restrictions on the adjustment coefficients for the joint weak exogeneity of m_{ot}, m_{it},and R_t and on the cointegrating vector for a m_i velocity equation without trend and outside money. The conventional LR test strongly rejects the joint restrictions at the 5% level with the test statistic $\chi^2(7) = 35.13$. However, an alternative restriction without price homogeneity is accepted with $\chi^2(6) = 3.70$ at the 5% level and yields an economically interpretable equation $\beta' x_t = y_t - m_{it} + 1.39p_t - 0.03R_t$. This equation may be interpreted as an aggregate demand equation rather than a money demand equation, because of the exogeneity of m_{it}. Using this identified equation, we reparameterized the initial VAR into a VECM with three lags. As expected from the tests of weak exogeneity, the ECM terms appeared in the equations for Δm_{ot}, Δm_{it} and ΔR_t are not significant. A joint restriction on these variables is accepted with the LR statistic $\chi^2(3) = 1.96$ at the 5% level. Based on this result, we finally formulated a restricted VECM, as with equations (5.4) and (5.5), which implicitly represents a block-recursive conditional structure between economic variables ($\Delta\ y_t$ and Δp_t) and policy variables (Δm_{ot}, Δm_{it} and ΔR_t) and estimated it with the SURE method. Table 5.2 reports the estimated results.

To investigate the dynamic impulse responses of the system to initial shocks on each series, we use the one standard deviation of errors from the corresponding equations in the underlying model. Even though this non-orthogonal impulse response analysis doesn't consider the contemporaneous correlation between variables, An advantage of this approach is that the empirical results derived are invariant to the arbitrary ordering of variables. Furthermore, the casual relationship between variables can be explained from the empirical results of the weak exogeneity tests conducted in the previous section. Figure 5.1 displays the graphs of impulse responses that trace the impacts of specific shocks on levels of each series. The graphs show that shocks to individual variables do not die out, reflecting the non-stationarity of variables in the model, and so have permanent effects with non-zero convergence (see Lütkepohl (1991)). Shocks on outside and inside money stocks show similar effects on the economy, except price levels which positively react to the former but negatively to the latter. Both shocks substantially affect output, even if the empirical model includes an interest rate. This evidence is contrasted with the work of Sims (1982a) and Litterman and Weiss (1985) who find that if an interest rate is included in a VAR model, the relationship between money and output is weakened. The size of the long-run non-neutral effect of inside money on output is much bigger than that of outside money. The interest rate negatively reacts to shocks on both money stocks in the initial two and four quarters, thus showing the liquidity effects of increased money stocks, but it is eventually dominated by expected inflation and shows positive correlations with money shocks in the long run. This evidence is consistent with the recent findings of Strongin (1995) and Bernanke and Mihov (1998). However, it should be noted that our results were obtained from a statistically identified structure, while the previous two results were obtained from arbitrary identified structure.

The different impacts of inside and outside money stocks on prices seem to reflect the fact that the two money stocks have been governed by the BOK's different control methods. In this context, the opposite effects may be understood by carefully examining the institutional details of how monetary policy in Korea has been actually carried out. Over the sample period, the BOK has mainly supplied the monetary base through an automatic rediscounting facility on the commercial bills related to exports and heavy and chemical industries and through the provision of general loans to banks which are in the deficiency of required reserves (as a lender of last resort). Therefore, even though the money is under the control of the BOK, it has been procyclically supplied in accordance with economic situations. On the other hand, in order to maintain the broad M2 money stock close to its target values, the BOK has tightly controlled the supply of inside money, which is expected to be mainly determined by the behavior of economic agents, through direct control methods rather than indirect methods based on exogenous high-powered base-money multiplier. The main techniques include reserve requirements, direct credit ceilings, regulation on the payments of interest rates, and forced sale of monetary stabilization bonds. In this context, the supply of inside money is countercyclical to prices. That is, when the economy is under inflationary pressures, the central bank tightens the supply of bank credits. Then inside money (bank deposits) is contracted with rising price levels and falling output.

With these institutional characteristics of policy implementation, the strong negative response of prices to a shock to inside money may be interpretable from the view of a traditional monetarist framework, within which inside money innovations partially reflect 'expansionary' policy rather than 'demand-induced' policy reactions, when the economy is in recession. That is, policy authorities know that the economy is entering into a recession and expand to recover from the recession. Then prices would fall after the monetary expansion and output would rise because of the standard effects of nominal demand on real output. This explanation that deflationary pressures generate higher output mainly because of the reaction to them by the monetary authorities, is further evidenced by other supplementary findings, such as the positive responses of monetary base and output (Lacker, 1988), the initial negative response of the interest rate (Leeper and Gordon, 1992), and the initial sticky response of prices (Sims, 1992), in the wake of a shock of inside money. Note that the exogeneous property of policy variables in the cointegration analysis rules out an alternative possible interpretation, based on a RBC framework, that the observed close interrelationship between real and monetary variables is mainly resulted from the correlation of output with inside money; innovations in inside money reflect the arrival of information concerning future productivity disturbances; and as a result, the increases in the quantity of money associated with decreases in interest rates and with decreases in prices are due to beneficial productivity shocks, such as the development of new products or production methods and changes in government regulations affecting production.

An impulse on the interest rate leads outside and inside money stocks to initially decline for the first two and three quarters, respectively, but thereafter to rise sharply. On the other hand, output is negatively responded to a shock in this type. Such facts may be interpreted that since monetary policy disturbances are a leading factor in generating aggregate fluctuations, a money contraction to reduce rising aggregate demands causes the interest rate to rise sharply and output to decrease. The responses of prices to the interest rate shock are very negative. This pattern fits well with the conventional monetarist framework in which a monetary contraction reduces nominal aggregate demands and eventually causes output and interest rates to decline through the interactions of deflationary pressure.[13] An impulse on prices causes outside money and output to decline. But inside money reacts with initial negative responses for the first three quarters and then, rises sharply. This behavior of monetary aggregates may explain that the BOK has actively reacted to dampen unexpected price shocks by contracting the supply of money rather than by passively accommodating the increased demand for money. Finally, a shock on output negatively affects monetary aggregates and positively the interest rate and prices. The fact that outside and inside money stocks are negatively correlated with the innovation in output is suggestive of the hypothesis that the BOK has quickly moved to forestall inflationary pressures on the economy by contracting monetary aggregates.

To examine the empirical results robustly, we assume the identified equation

[13] For developed countries, Sims (1992) finds the positive comovements of prices with interest rates ('price puzzle') and explains the patterns with policy endogeneity. Recently, Christiano *et al.* (1996) and Kim (1999) solve the puzzle. However, their evidence is based on a pre-assumed identification. Thus, if the assumption is changed, the results will be changed.

$\beta' x_t = y_t - m_{it} + 1.39 p_t - 0.03 R_t$ in the cointegrating vector as a money demand equation by endogenously treating inside money. This assumption would be consistent with the view of RBC theory, but obviously violates the exogenous properties of the variables. In this case, the ECM terms are appeared in the equations for Δm_{ot}, Δm_{it} and ΔR_t, but not in the equations for $\Delta\ y_t$ and Δp_t. Based on the reverse conditional structure of the VECM, we reestimated the model and examined the dynamic impulse responses of each variable. Figure 5.2 reports the responses of each variable on a shock to output, which may be regarded as a real shock affecting the production function. The results do not support the standard RBC's scenario that a beneficial productivity shock increases price levels and money supply and decreases interest rates. This may illustrate the importance of valid identification of policy variables in a VAR-based empirical analysis. Overall, our empirical results show that exogenously identified monetary aggregates have significant effects on output with sticky prices, but the reverse does not hold. This evidence seems to be compatible with the view that money matters, even though the long-run neutrality of money is not found, but is inconsistent with the RBC view of the reverse causation that money is caused by output.[14]

5.4 Conclusions

In this chapter we have identified the policy stance in Korea and examined whether money matters. To do this, we emphasized a statistical identification of policy variables by applying the concept of weak exogeneity in a VECM model. A methodological usefulness of this technique is that the exogeneity of variables can be directly tested within the baseline model and used for the examination of the 'causal relation' between variables. Thus it is expected that the empirical

[14] Cagan (1969) argues that since in the real world expectations are imperfect and prices adjust slowly, changes in the money stock disturb the pattern of expenditures and output and so, neutral money is a figment of abstract theory in which changes in the money stock are fully anticipated and prices change proportionately.

results based on this approach are more reliable than those from the current approach which assumes *a priori* the predeterminedness of policy variables without a formal test. The cointegration analysis in this study shows that monetary aggregates which have been used by the BOK as the main policy instruments are exogenously identified, implying that the BOK has implemented its policy discretionarily rather than systematically responded to the movements of macroeconomic variables. The exogeneity of inside money is not consistent with King and Plosser's (1984) argument that money-output correlation reflects purely endogenous changes in money holdings formed in response to changes in current or expected future output. This rejection of the reverse causality is again confirmed by the dynamic interactions between variables which show considerable effects of monetary aggregates on output. The effects of inside and outside money on out are permanent, but the former has a bigger impact than the latter. The long-run non-neutrality of money might be due to the price stickiness stemmed from the strict regulation of the Korean government on prices over the sample period. This Korean evidence may shed some light on the incompatibility of the RBC hypothesis of endogenous money to developing countries. Finally, it is to be noted that our study does not suffer from search to explain the liquidity puzzle associated with the positive initial responses of interest rates to a positive shock on money, the price puzzle related to the positive responses of the price level to a positive shock on interest rates, and the weakened effect of money on output when a VAR model includes an interest rate, all of which have been commonly found in the previous VAR literature. This result seems to demonstrate the importance of a statistical identification of policy stance in a VAR-based approach.

TABLE 5.1 Cointegration analysis

(a) Cointegration tests					
Eigenvalues	0.37	0.23	0.15	0.12	0.06
Null hypothesis	$\gamma = 0$	$\gamma \leq 1$	$\gamma \leq 2$	$\gamma \leq 3$	$\gamma \leq 4$
Max	45.81**	25.72	16.58	12.59	6.55
Trace	107.2**	61.43	35.71	19.14	6.55

(b) Tests for weak exogeneity						
Variables	y_t	p_t	m_{0t}	m_{it}	R_t	$\{m_{0t}, m_{it}, R_t\}$
$\chi^2(\cdot)$	2.94*	19.72**	0.47	0.02	1.36	1.71

Notes: 1. ** and * denote significance at the 5% and 10% levels, respectively. 2. The test statistics of weak exogeneity are asymptotically distributed as $\chi^2(1)$ for y_t, p_t, m_{0t}, m_{it}, and R_t, respectively, and $\chi^2(3)$ for the joint test of $\{m_{0t}, m_{it}, R_t\}$.

TABLE 5.2 Model estimation

(a) Estimated results

Variables	Δm_{ot}	Δm_{it}	ΔR_t	Δp_t	Δy_t
Δm_{ot-1}	- 0.27	- 0.02	- 1.51	- 0.004	0.05
	(0.09)	(0.02)	(2.07)	(0.02)	(0.03)
Δm_{ot-2}	0.17	0.04	4.54	0.05	0.02
	(0.10)	(0.02)	(2.17)	(0.02)	(0.03)
Δm_{ot-3}	0.28	0.04	3.75	0.02	0.07
	(0.10)	(0.02)	(2.26)	(0.02)	(0.03)
Δm_{it-1}	0.89	0.16	- 4.04	- 0.06	0.43
	(0.46)	(0.10)	(9.65)	(0.08)	(0.13)
Δm_{it-2}	0.33	0.28	0.47	- 0.02	0.08
	(0.49)	(0.11)	(10.21)	(0.08)	(0.14)
Δm_{it-3}	- 0.12	0.13	- 4.68	- 0.22	0.04
	(0.49)	(0.11)	(10.34)	(0.08)	(0.14)
ΔR_{t-1}	- 0.004	- 0.001	0.16	- 0.002	- 0.001
	(0.005)	(0.001)	(0.11)	(0.001)	(0.002)
ΔR_{t-2}	0.001	0.003	- 0.22	- 0.001	- 0.0003
	(0.005)	(0.001)	(0.11)	(0.001)	(0.002)
ΔR_{t-3}	0.01	- 0.001	0.005	0.001	- 0.002
	(0.005)	(0.001)	(0.11)	(0.001)	(0.002)

Variables	Δm_{ot}	Δm_{it}	ΔR_t	Δp_t	Δy_t
Δp_{t-1}	- 0.69	- 0.02	8.64	0.20	-0.04
	(0.63)	(0.14)	(13.04)	(0.10)	(0.17)
Δp_{t-2}	0.73	- 0.06	0.75	0.06	-0.04
	(0.63)	(0.14)	(13.28)	(0.10)	(0.17)
Δp_{t-3}	- 0.63	0.19	- 10.89	0.09	-0.08
	(0.55)	(0.13)	(11.63)	(0.09)	(0.15)
Δy_{t-1}	- 0.07	- 0.11	3.29	0.13	-0.28
	(0.41)	(0.09)	(8.46)	(0.07)	(0.11)
Δy_{t-2}	- 0.12	- 0.20	5.64	0.08	-0.10
	(0.42)	(0.09)	(8.93)	(0.07)	(0.12)
Δy_{t-3}	0.64	- 0.06	8.24	- 0.09	-0.004
	(0.39)	(0.09)	(8.24)	(0.06)	(0.10)
$cons$	- 0.02	0.02	- 0.19	0.58	-0.31
	(0.03)	(0.01)	(0.58)	(0.09)	(0.16)
ecm_{t-1}	-	-	-	- 0.10	0.05
	-	-	-	(0.01)	(0.03)

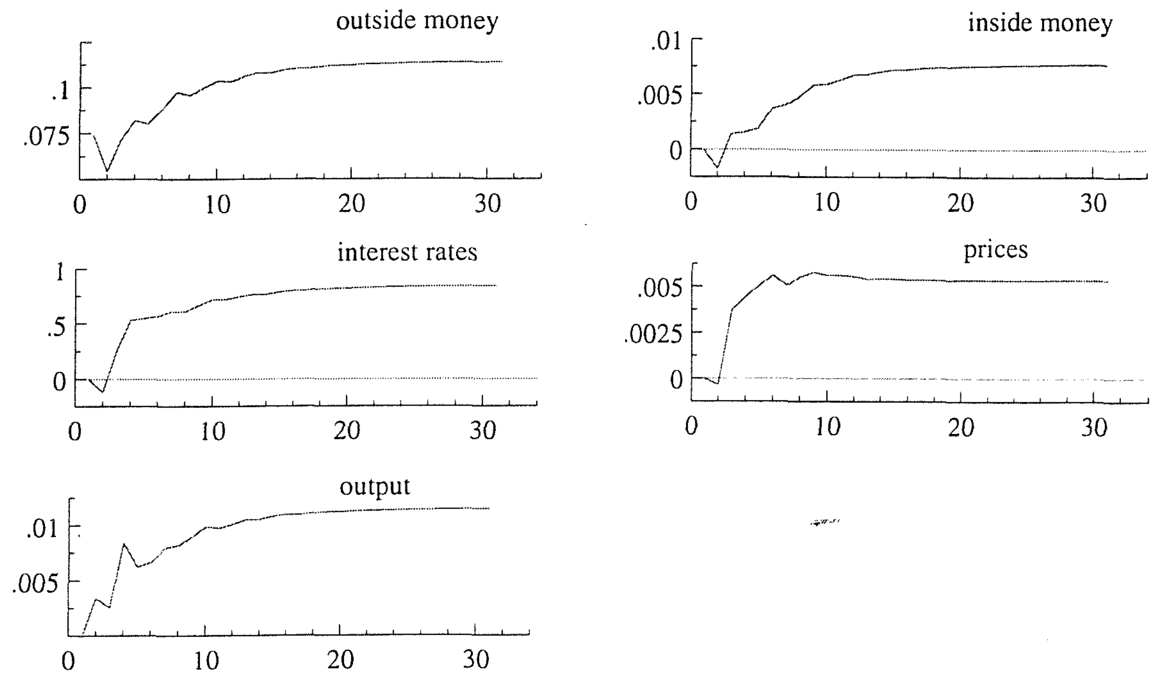
(b) Model evaluation

Statistics	Δm_{ot}	Δm_{it}	ΔR_t	Δp_t	Δy_t	model
$F_{AR}(5, 78)$	1.61	2.85*	1.10	1.52	0.58	-
χ^2_N (2)	2.57	1.35	20.51**	19.22**	8.05*	-
$F_{ARCH}(4, 75)$	0.48	0.43	1.42	2.27	1.14	-
$VEC_{AR}\, F(125, 275)$						1.27
$VEC_N\, \chi^2(10)$						49.24**
$VEC_{HET}\, F(480, 563)$						0.96

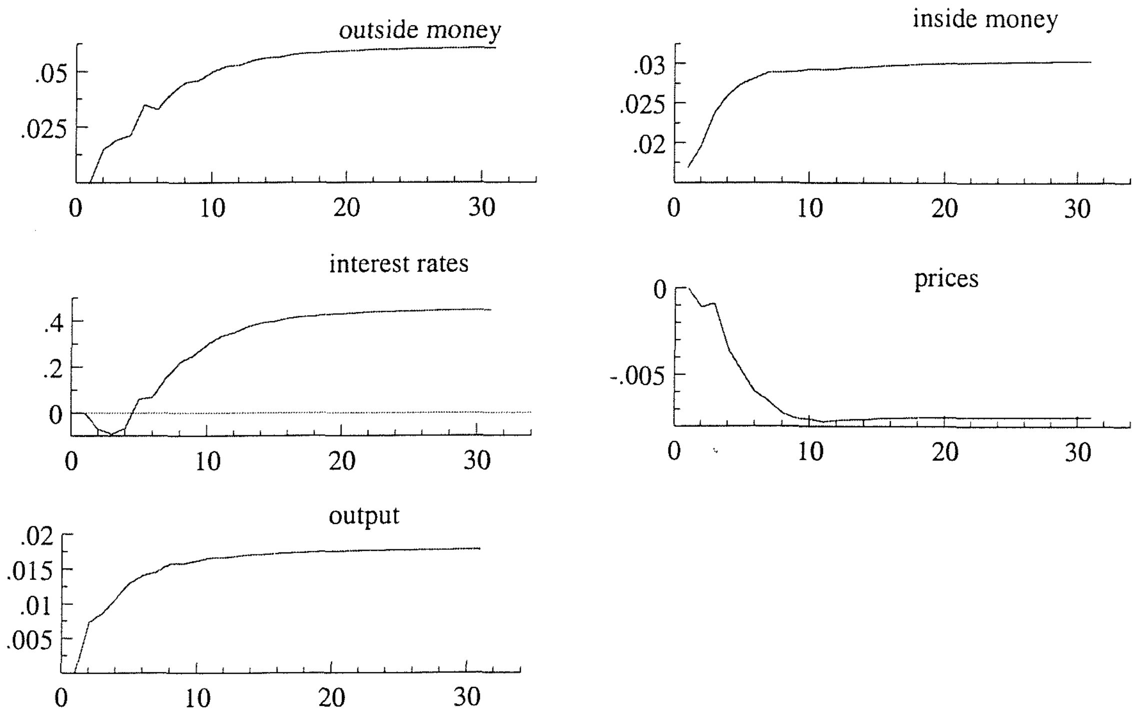
Notes: 1. ** and * denote significance at the 1% and 5% levels, respectively.

(c) Covariance matrix of the residuals

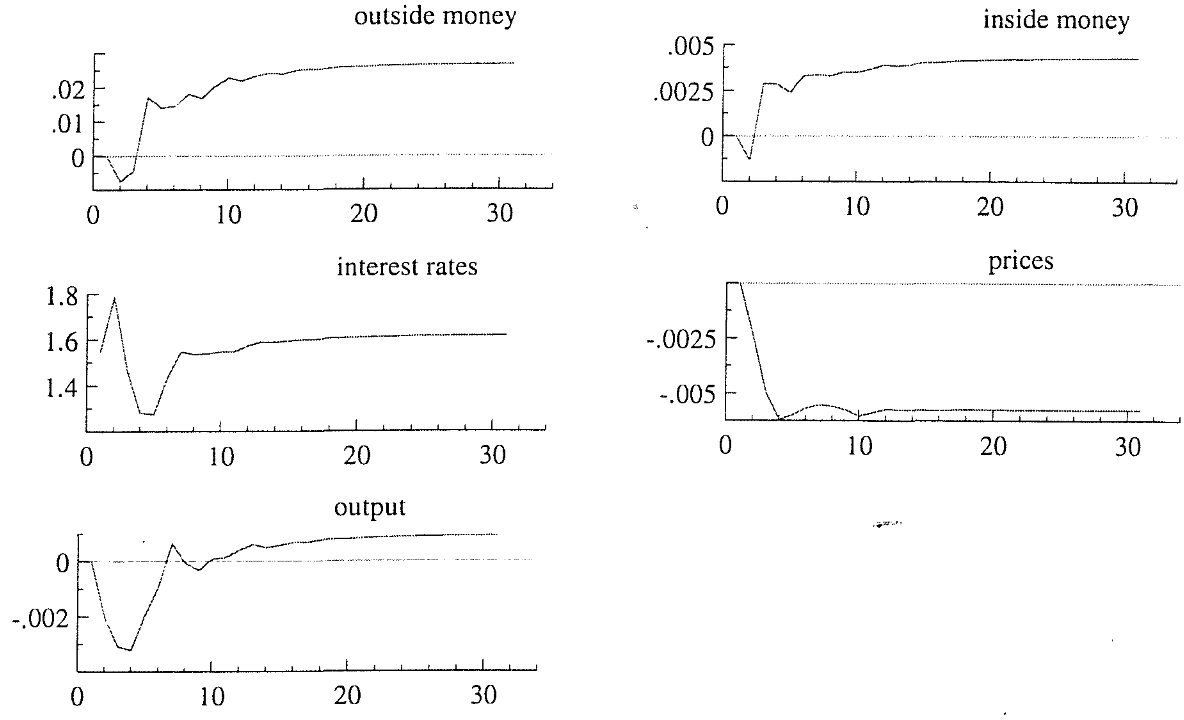
	Δm_{ot}	Δm_{it}	ΔR_t	Δp_t	Δy_t
Δm_{ot}	1.00				
Δm_{it}	0.05	1.00			
ΔR_t	- 0.19	- 0.03	1.00		
Δp_t	- 0.04	- 0.09	0.43	1.00	
Δy_t	0.01	0.06	- 0.06	- 0.12	1.00



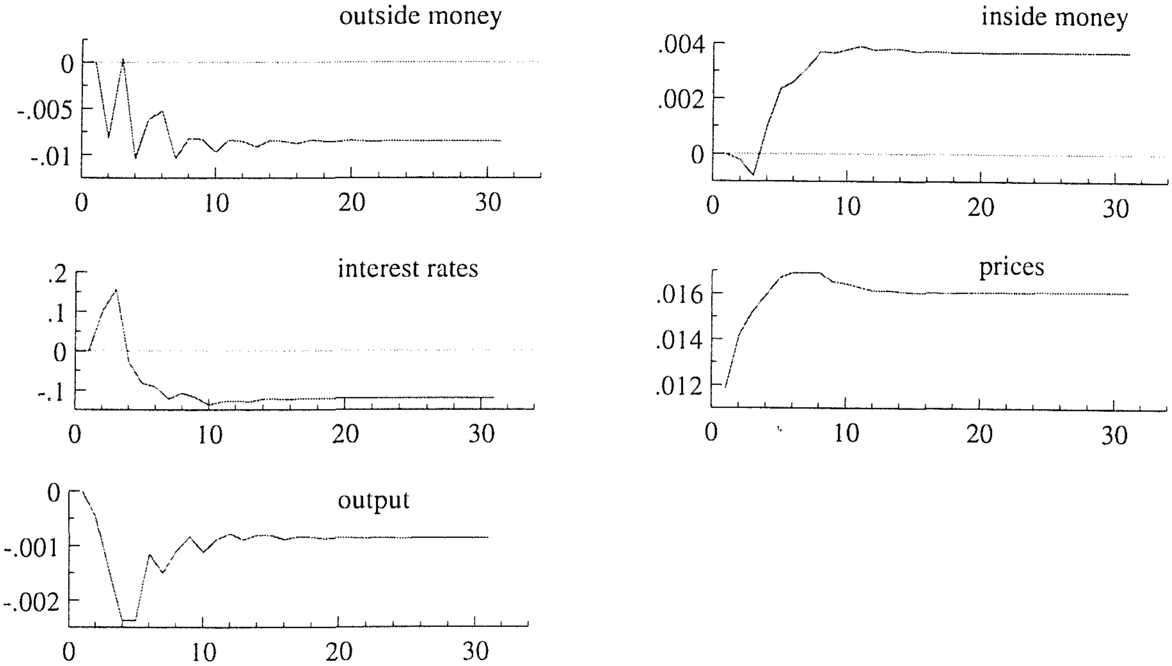
(a) Responses on a shock to outside money



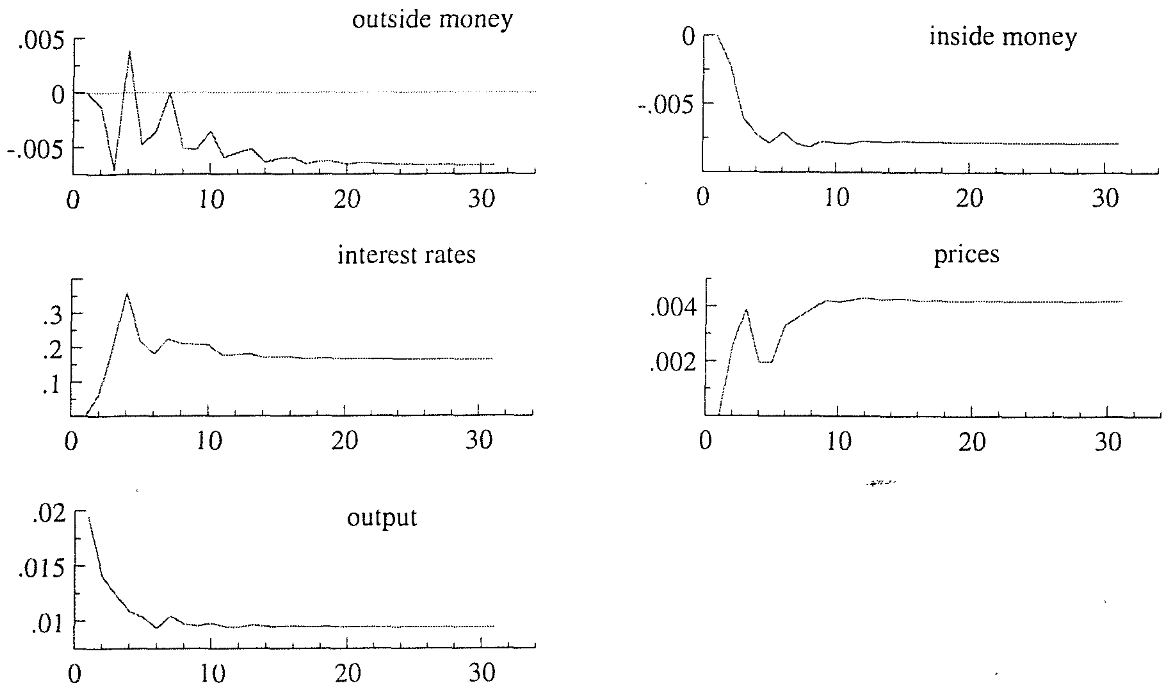
(b) Responses on a shock to inside money



(c) Responses on a shock to interest rates



(d) Responses on a shock to prices



(e) Responses on a shock to output

Figure 5.1 Impulse responses of the VECM model

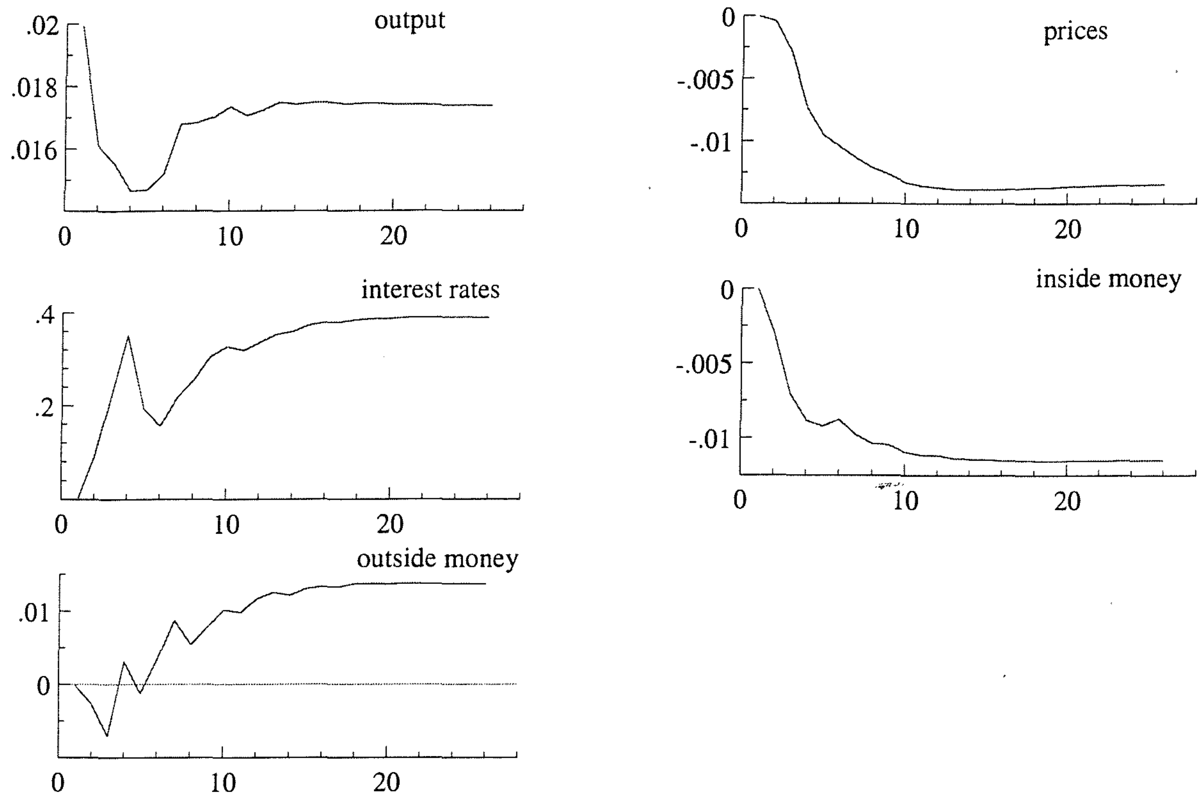


Figure 5.2 Impulse responses on an output shock.

Chapter 6

Dynamic Econometrics and Policy Analysis[1]

6.1 Introduction

Traditionally, monetary policy has been used one of the conventional macro instruments of overall policies. In particular, the policy is popular to policymakers because it can be changed on short notice, with little or no legislative delay. Thus, one of the major issues in macroeconomics is whether and how changes in monetary policy affect output, prices, and other economic variables. Policy changes are generally divided into two types of changes: transitory and permanent changes (see Lucas (1976) and Taylor (1988)).[2] *Transitory* changes are defined as the changes associated with unanticipated 'shocks' on the variations around permanent components. These types of policy changes occur when policymakers decide what kind of appropriate policy actions should be undertaken in current economic situations, without a necessary connection between the choices of dif-

[1] **An early version of this chapter was presented at the 6th Spring Meeting of Young Economists 2001 in Copenhagen.**

[2] New classical economists prefer to divide policy changes into anticipated and unanticipated changes by giving more attention to the role of economic agents' expectations.

ferent periods. So their influence on the economy is unsystematic. On the other hand, *permanent* changes in policy, including institutional or regulative changes in a broad sense (for examples, the switches from fixed to flexible exchange rates and from interest rate to money targets) or rule changes in a narrow sense, take place when policymakers decide how much the average values of policy variables should take and how long the values should remain unchanged.[3] In this context, permanent changes in policy are defined as the changes associated with 'maintained' changes in the mean, variance or growth rates of policy variables. Since these types of 'one-time' changes are generally announced or enacted prior to their implementation, permanent policy changes can be anticipated in advance, but may cause the economy to change structurally in the long run. Thus their effects on the economy may be permanent and systematic.

To investigate the effects of these two types of change in policy on the economy, a large number of empirical studies have applied different types of models for different countries. Most of the studies can be categorized into two major quantitative approaches: the structural approach and the vector autoregression (VAR) approach.[4] The former approach addresses the effects of policy changes by measuring the structural impacts of exogenous policy variables on target variables. However, the analysis in this approach is mainly conducted through parameter spaces on which identifying restrictions are directly imposed on the basis of economic theory. Little attention is paid to the interrelations of error terms. Thus, empirical studies in this line of research largely focus on evaluating the effects of

[3]Policy changes in this type are analogous with the changes of the time consistent 'operating regime' of Bryant *et al.* (1993). Operating regime is defined as a prescribed general guide for conduct, which need not be simple and rigid and which might or might not allow policymakers to have a substantial scope for discretion in the future. This concept is different from 'rule' which is a determined, often simple and rigid, procedure indicating how policy should be implemented and so gives little or no scope to policymakers for activist discretion today or in the future.

[4]In addition to these two approaches, the 'narrative approach' proposed by Romer and Romer (1989) can be regarded as a third line of research. In this paper, however, we don't discuss the approach because policy analysis in this non-quantitative method is mainly conducted within the framework of either a structural model or a VAR model with identified changes in policy.

systematic permanent changes in policy rather than the unsystematic changes. On the other hand, the latter approach analyzes the effects of policy changes on the economy by capturing the impulse responses of nonpolicy variables to transitory shocks on the innovation errors of policy variables. In this approach all variables are treated as endogenous, and *a priori* 'minimal' restrictions for identification are imposed in the covariance matrix of error terms rather than parameter spaces. Thus, the VAR-based studies mainly focus on analyzing the effects of unsystematic transitory changes (see Christiano *et al.* (1999) for an extensive recent survey).

Note that none of the approaches look at the effects of transitory and permanent changes jointly. One reason for failing to take account of both effects is in the intrinsic limitation of the econometric techniques which emphasize their own procedures in a divergent way. However, a further fundamental reason seems to lie in theoretical differences on characterizing the nature of policy making. Early new classical macroeconomists shift policy analysis away from single policy actions to policy rules and divide policy into systematic and unsystematic parts, whether or not it is intentionally based on rules, because of the rational expectations of economic agents (see, for examples, Lucas (1972, 1976) and Sargent and Wallace (1976)).[5] They argue that the systematic parts representing permanent changes, such as rule changes, are anticipated by rational economic agents and so is not effective, whereas the unsystematic parts representing transitory changes, such as surprise shocks, are unanticipated but effective. In analyzing the effects of these changes in policy, however, they regard the latter as simply

[5] In traditional Keysian policy evaluation, policy changes are regarded as interventions in the present and future values of exogenous forcing variables to maximize a social objective function which is subject to the constraints imposed by the economic system. However, Lucas (1976) criticizes this approach as selecting the sequences of forcing variables 'arbitrary', in the sense that they are not characterized stochastically, and so failing to account for economic agents who maximize utilities subject to well-defined constraints. Instead, he views policy changes as parameter shifts rather than variable shifts and argues that policy should be governed by simple rules.

noise and so not useful for policy analysis. Instead, they emphasize the importance of the systematic part in *a priori* identified behavioral equations, which can be regarded as structural, and concentrate on analyzing the effects of permanent changes in policy through parametric characterization with a particular emphasis of rational expectations. On the contrary, Sims (1982b), who is even a new classical economist, is reluctant to accept the possible effects of permanent regime changes on the economy because they are rare events and argues that 'normal policymaking' merely selects errors in a stable policy rule. With this view, he prefers to analyze the effects of policy changes through the unsystematic transitory part, instead of the systematic part, and suggests a VAR approach as an alternative way for policy analysis.

Recently, several macroeconomists emphasize the importance of both policy changes to the economy. Cooley and LeRoy (1985) and LeRoy (1995) argue that in analyzing policy, the effects of an intervention in parameters, which is associated with a question 'what should we do now, given that we did the same in the past?' and an intervention in variables, which is associated with a question 'what should we do here and now, taking the past as given?' must be investigated separately. Bernanke *et al.* (1997), who apply a VAR-oriented approach to investigate the effects of systematic policy changes, also argue that looking only at unanticipated policy changes in VAR models begs the question of how systematic monetary policy changes affect the economy and empirically find that a large fraction of the U.S economy's real effects from oil price shocks since 1970s has resulted from the systematic changes in policy to these shocks rather than from the shocks themselves. McCallum (1999), who applies a structural approach to investigate the effects of policy changes, further argues that more emphasis should be given to the systematic portion of policy behavior and correspondingly less to random shocks, because shocks account for a very small fraction of policy-instrument variability, and stresses the importance of structural analysis for the effects of policy changes.[6] The main theme of these arguments is that the effects of the permanent systematic component of monetary policy on the economy is at least as important as those of the transitory unsystematic component.

In addition to the above empirical importance of both policy changes, there are several advantages in the context of econometric modelling if we account for the changes jointly within a simultaneous model. First, permanent changes in policy which are associated with regime changes imply variant parameters in econometric models. Unfortunately, most of macroeconomic data involve a number of regime changes (see, for an example, Judd and Rudebusch (1998)). If this is the case, following Lucas (1976) it is almost impossible to formulate invariant econometric models without incorporating the rational behavior of economic agents. Of course, as investigated by Ericsson and Iron (1995), the force of the critique may be less powerful in practice than theory. However, the logic of the critique makes sense. Thus, we should empirically examine whether or not an econometric model considered is invariant to different policy regimes, if the model doesn't formally incorporate rational expectations. In this context, useful information about the stability property of a model can be obtained by choosing events that represent permanent regime changes and by evaluating the invariance of the model under such interventions - that is, whether the underlying model is subject to the Lucas critique or not. Second, the likely effects of permanent changes are normally analyzed through identified behavioral parameters which are regarded as structural, while the effects of transitory changes are analyzed through unanticipated shocks on error terms. Hence, analyzing both changes within a model implies that we can obtain information on data from both parameter spaces and error terms. This point is well explained by Pagan (1994) who argues that 'it seems ridiculous to assume we know something about con-

[6]Note that the studies of Bernanke *et al.* (1997) and McCallum (1999), who use a VAR and a structural model respectively to investigate the effects of systematic policy changes, do not formally examine the parameter constancy of their baseline models and so are open to the Lucas (1976) critique.

temporaneous relations and nothing about dynamics as in the VAR; equally silly is the idea that we would wish to leave covariance unrestricted at the expense of saying something about dynamics and contemporaneous relations'.

In this chapter we examine the empirical feasibility of the above arguments using Korean data. To do this, we extend Hendry and Mizon's (1998) work, which provides a reduced-form framework on how to assess the effects of policy changes and applies it to a bivariate model, into a structural framework by formally incorporating the usefulness of weak exogeneity in identifying policy actions and by applying the approach to a multivariate model. Following Hendry and Mizon (1998), the effects of transitory changes are analyzed by using impulse response functions. Unlike the conventional approach which assumes policy indicators a priori either as endogenous or as exogenous, we identify them block-recursively by using the test of weak exogeneity suggested by Engle *et al.* (1983) and extended by Johansen (1992) into a cointegrated model. This identification of policy variables provides a straightforward 'causal' interpretation between variables, so avoiding Cooley and LeRoy's (1985) critic on the conventional VAR-based technique. On the other hand, the effects of permanent changes are evaluated by examining whether economic variables move together after deterministic changes in policy variables. We do this by using the concept of co-breaking suggested by Hendry and Mizon (1998) and Clements and Hendry (1999). The usefulness of this approach is that even policy and nonpolicy variables systematically move together, the underlying model maintains its invariant property and so is not subject to the Lucas (1976) critique, since co-breaking cancels out structural breaks. To apply this approach empirically, we first specify a structural dynamic model using extensive specification tests which include tests of encompassing, exogeneity, and parameter constancy. Distinctive features of the model are that no single restriction is placed *a priori* without formal tests and that the effects of transitory and permanent changes in policy are evaluated jointly within the model. The comparative advantages of the model to structural and VAR models are that it doesn't suffer from the Lucas (1976) critique which is concerned with the near-impossibility of estimating structural equations if data involve regime changes, from the Sims (1980) critique which is concerned with *a priori* 'incredible' restrictions for identification, and from mis-specification which invalidates the model itself.[7]

The chapter is organized as follows. Section 2 discusses the methodological aspects employed in this study. Section 3 briefly overviews the historical backgrounds of monetary policy in Korea and provides a small dynamic macroeconometric model. Based on the model, Section 4 evaluates the responses of the economy to transitory policy changes using impulse response functions and to permanent changes in policy using the concept of co-breaking. Section 5 finally concludes the chapter.

6.2 Econometric Methodology

This section explains the econometric techniques employed in this chapter. For the expositional clarity, we discuss our arguments in an *I(0)* VAR framework. The similar analysis in the *I(1)* system is provided in the Appendix. Consider an $n-$dimensional VAR model as follows:

$$A(L)x_t = \epsilon_t, \qquad \epsilon_t \sim IN(0, \textstyle\sum), \tag{6.1}$$

where $A(L)$ is an $n \times n$ matrix polynomial in the lag operator L, $A(L) = \sum_{j=0}^{p} A_j L^j$, and ϵ_t is an $n \times 1$ vector of independently identically distributed

[7]The traditional structural approach does not question the adequacy of the models considered, rather taking it for granted that economic theory do that. Thus the approach assumes that the estimated models are correctly specified and seeks to verify the underlying economic theory in terms of signs, magnitudes, and significance of estimated parameters. The lack of statistical tests in the structural approach can also be applicable to the VAR approach. Most of the research papers based on VAR models seldom report specification tests. However, it is well-known that VAR models are very sensitive to the selection of lag lengths and overparameterization (see Spencer, 1989; Phillips, 1998; Abadir *et al.* 1999).

zero mean errors and nonsinglar covariance matrix $\sum$. The main characteristic of this model is that it coherently represents complex data generating processes through dynamic interactions by treating all variables as endogenous. In the approach of dynamic econometric modelling, the model is used as a benchmark model to derive a structural model by applying sequential reductions and transformations based on extensive specification tests and economic theory (Hendry, 1995). The test of over-identifying restrictions against the unrestricted VAR provides a legitimacy of valid reduction for a final model which can be regarded as adequately characterizing data information and accounting for the results obtained by rival models. A distinctive feature of the model obtained in this way is that it is an intermediate structural model between traditional structural and VAR models, within which policy analysis can be conducted through parameter spaces and error terms jointly. The former is useful for analyzing the effects of permanent changes in policy and the latter is useful for the effects of transitory changes. By partitioning x_t as $x_t = (y_t' : z_t')'$, where y_t represents a subset of nonpolicy variables and z_t represents a subset of policy variables, such effects may be represented as the partial responses of nonpolicy variables y_{t+h} to changes in the means of policy variables μ_{zt} $(\partial y_{t+h} / \partial \mu_{zt})$ and to shocks in the innovations of policy variables ϵ_{zt} $(\partial y_{t+h} / \partial \epsilon_{zt})$, respectively.

6.2.1 Effects of transitory changes

Transitory changes in policy take the form of 'surprise' shocks on innovation terms. In the VAR literature the effects are normally evaluated by tracing out the dynamic effects of random shocks to policy variables on economic variables, using orthogonalized impulse response techniques. However, it is well-known that one of the major problems in this approach is the ordering of variables which in turn depends on identifying exogenous variables. In this subsection we show how this problem can be eliminated 'block-recursively' by using the concept of weak exogeneity (Engle *et al.* 1983) and how the effects of transitory changes work in this block-recursively identified system. To simplify the analysis, consider a first-order structural VAR model from (6.1) with diagonal unit elements on the contemporaneous correlation matrix:

$$\begin{bmatrix} y_t \\ z_t \end{bmatrix} = \begin{bmatrix} -A_{12.0}z_t \\ -A_{21.0}y_t \end{bmatrix} + \begin{bmatrix} A_{11.1} & A_{12.1} \\ A_{21.1} & A_{22.1} \end{bmatrix} \begin{bmatrix} y_{t-1} \\ z_{t-1} \end{bmatrix} + \begin{bmatrix} \epsilon_{yt} \\ \epsilon_{zt} \end{bmatrix}, \quad \sum = \begin{bmatrix} \sum_{yy} & \sum_{yz} \\ \sum_{zy} & \sum_{zz} \end{bmatrix}. \tag{6.2}$$

Assume that policy variables z_t are independent of the contemporaneous developments of nonpolicy variables y_t and so $A_{21.0} = 0$, possibly because of information lags.[8] Then, equation z_t is reduced to

$$z_t = A_{21.1}y_{t-1} + A_{22.1}z_{t-1} + \epsilon_{zt}, \tag{6.3}$$

and substituting this into equation y_t leads to a reduced-form equation

$$y_t = (A_{11.1} - A_{12.0}A_{21.1})y_{t-1} + (A_{12.1} - A_{12.0}A_{22.1})z_{t-1} + v_t, \tag{6.4}$$

where $v_t = \epsilon_{yt} - A_{12.0}\epsilon_{zt}$.

Equations (6.3) and (6.4) represent a standard reduced-form VAR system with recursively correlated error terms and the following properties:

$$\begin{bmatrix} v_t = \epsilon_{yt} - A_{12.0}\epsilon_{zt} \\ \epsilon_{zt} \end{bmatrix} \sim IN(0, \Omega),$$

$$\Omega = \begin{bmatrix} \sum_{yy} + 2(-A_{12.0}) \sum_{yz} + (-A_{12.0})^2 \sum_{zz}, & \sum_{yz} + (-A_{12.0}) \sum_{zz} \\ \sum_{zy} + (-A_{12.0}) \sum_{zz}, & \sum_{zz} \end{bmatrix},$$

and the conditional equation y_t given z_t is $y_t = \alpha_1 z_t + \alpha_2 y_{t-1} + \alpha_3 z_{t-1} + u_t$, $u_t \sim IN(0, \sum^2)$, where $\alpha_1 = -A_{12.0} + \sum_{yz} \sum_{zz}^{-1}$, $\alpha_2 = A_{11.1} - A_{21.1} \sum_{yz} \sum_{zz}^{-1}$, $\alpha_3 =$

[8] This type of 'inside' lags can arise because it takes time for policymakers to recognize that a shock has occurred and to put appropriate policies into effect.

$A_{12.1} - A_{22.1} \sum_{yz} \sum_{zz}^{-1}$, and $\sum^2 = \sum_{yy} - \sum_{yz}^{2} \sum_{zz}^{-1}$.

If $\sum_{yz} = 0$, then z_t is weakly exogenous for the parameters of y_t in (6.2). In this case, the latter is recovered from only the parameters of α_1, α_2, α_3, and $\sum^2$ in the conditional equation y_t (see Engle *et al.* 1983). An important implication of the block-recursive structure of the error terms in the equations of (6.3) and (6.4) is that in an unrestricted reduced-form VAR, exogenously identified z_t can be placed firstly in the ordering of variables for orthogonalized impulse response analysis.[9] This shows the usefulness of weak exogeneity in identifying exogenous variables, which provides a straightforward 'causal' interpretation. However, note that this type of exogeneity depends on the choice of the parameters of interest and is not an intrinsic property of z_t. Thus, the usefulness of this approach is limited to the case that orthogonal transformations based on the Choleski decomposition are legitimately tested. Even though the block-recursive identification in this approach doesn't provide complete ordering, it is very useful when we consider that most of the current controversies in VAR-based analysis stem from a pre-assumed identification of unobservable policy stance, by transforming the VAR into orthogonal errors. Alternatively, instead of the orthogonalized impulse responses based on the Choleski factorization of the covariance matrix, a unit or a standard-error-based impulse response, which is independent of variable ordering, can be used. Even though these non-orthogonal impulse responses ignore the potential correlatedness within variables, identified exogeneity provides a legitimacy of the causal interpretation between variables. The importance of this analysis is that exogeneity in this framework is formally tested rather than merely assumed *a priori* as in the current VAR-based approach. To test weak exogeneity, Engle's (1982b, 1984) LM test procedures can be used. In the case of a cointegrated VAR, Johansen's (1992) procedure is directly applicable.

[9] Note that the problem of ordering in a VAR approach is not immaterial for policy analysis even under the contemporaneous uncorrelatedness of reduced-form errors, since the diagonal covariance matrix necessarily does not imply 'causal' relations between variables (see Cooley and LeRoy (1985)).

To see how the effects of transitory changes in policy work in this system, let a transitory change in ϵ_{zt} be $d\epsilon_{zt}$. In this case, the response of ϵ_{yt} is $d\epsilon_{yt} = \delta d\epsilon_{zt}$ from the identity $\epsilon_{yt} = \delta\epsilon_{zt} + v_t$, where $\delta = A_{12.0} = \sum_{yz}\sum_{zz}^{-1}$, leaving $dv_t = 0$. Then the partial response of y_{t+h} to ϵ_{zt} is obtained from equation (6.4) as follows:

$$
\begin{aligned}
\left(\frac{\partial y_{t+h}}{\partial \epsilon'_{zt}}\right)_{d\epsilon_{yt}=\delta d\epsilon_{zt}} &= (I_{n1}:0)\begin{pmatrix} A_{11.1}-A_{12.0}A_{21.1} & A_{12.1}-A_{12.0}A_{22.1} \\ A_{21.1} & A_{22.1}\end{pmatrix}^h \quad (6.5)\\
\begin{pmatrix}\delta \\ I_{n2}\end{pmatrix} &= \left((I_{n1}:0)\begin{pmatrix} D_h & E_h \\ F_h & G_h\end{pmatrix}\begin{pmatrix}\delta \\ I_{n2}\end{pmatrix}\right)\\
&= D_h\delta + E_h.
\end{aligned}
$$

The first term represents the dynamic response of the model to the contemporaneous correlation δ between ϵ_{yt} and ϵ_{zt}, while the second term represents the Granger causality of z_t to y_t and so, if z_t does not cause y_t, then $E_h = 0$. Hence, (6.5) shows that the major determinants of impulse responses to transitory changes in policy are contemporaneous correlations and Granger causality between policy and nonpolicy variables. Note that the impulse responses of (6.4) to a unit shock on z_t are equivalent to orthogonalized impulse responses in a conventional VAR. If there is no contemporaneous correlation between two blocks with $\delta = 0$ and so $d\epsilon_{yt} = 0$, then the partial response of y_{t+h} to ϵ_{zt}, which is equivalent to unit impulse responses in a conventional VAR, is $\partial y_{t+h}/\partial\epsilon'_{zt} = E_h$.

6.2.2 Effects of permanent changes

In this subsection, a static system is considered by assuming $A_{ij.k} = 0$ $(i, j, k = 0, 1, 2)$ and by introducing mean terms $E(y_t, z_t) = (\mu_{yt}, \mu_{zt})$ in the reduced-form system (6.3) and (6.4). Within this framework, the effects of permanent changes in policy, which are generally associated with 'maintained one-time' changes such as regime changes, are analyzed through the immediate interdependence of the means of policy and nonpolicy variables. The symbolic representation of the interdependence, which is the partial responses of y_{t+h} to changes in μ_{zt}, is:

$$\left(\frac{\partial y_{t+h}}{\partial \mu_{zt}'}\right) = \left(\frac{\partial y_{t+h}}{\partial \mu_{yt}'}\right)\left(\frac{\partial \mu_{yt}}{\partial \mu_{zt}'}\right) = I_{n1}\Theta_t = \Theta_t, \qquad \forall h \geq 0, \tag{6.6}$$

under a schematic behavioral relationship between the means of y_t and z_t:[10]

$$E[y_t|I_{t-1}] = \mu_y^* + \Theta_t E[z_t|I_{t-1}], \tag{6.7}$$

which can be rewritten as $\mu_{y,t} = \mu_y^* + \Theta_t \mu_{z,t}$, where μ_y^* denotes an initial value. If the response Θ_t is not zero, this indicates that permanent changes in z_t are effective because the mean of y_t has been affected. However, the main problem with this systematic change is that whenever economic agents rationally change expectations by integrating their knowledge about policy changes, the parameters of the behavioral model may be changed and so are subject to the Lucas (1976) critique.

Recently, Hendry and Mizon (1998) provide necessary conditions for a conditional model associated with changes in marginal processes to be immunized from the classical problem using the concepts of weak exogeneity and co-breaking (see, for more details, Engle *et al.* (1983) for weak exogeneity and Clements and Hendry (1999) for co-breaking). Co-breaking is defined as the cancellation of deterministic breaks across linear combinations of variables. Consider that $\{x_t\}$ has unconditional expectation around an initial parameter ψ at time $t = 0$ such

[10] The schematic behavioral relation characterizing the plans on which agents base their actions is derived from an assumption of agents' weak rationality. Let an observed outcome y_t deviate from the planned value by an innovation ϵ_t and so $y_t = y_t^p + \epsilon_t$. Taking conditional expectations to this equation yields $E(y_t|I_{t-1}) = E(y_t^p|I_{t-1}) + E(\epsilon_t|I_{t-1}) = y_t^p + E(\epsilon_t|I_{t-1})$. If economic agents are rational and so there is no systematic error $E(\epsilon_t|I_{t-1}) = 0$, then the equation becomes $E(y_t|I_{t-1}) = y_t^p = \mu_y^* + \Theta_t z_t^e$, which can be rewritten as $E(y_t|I_{t-1}) = \mu_y^* + \Theta_t E(z_t|z_{t-1})$ or $\mu_{y,t} = \mu_y^* + \Theta_t \mu_{z,t}$, under the assumption that the plan y_t^p formed for time t depends on the value z_t^e which z_t is expected to take at the time the plan is implemented and on μ_y^*. Here, $\mu_{y,t}$ denotes agents' plans about the variables they control and $\mu_{z,t}$ policy makers' plans about which agents hold expectations (see Hendry and Mizon, 1998).

that $E[x_t - \psi] = \tau_t \in R^n$ where $|\tau_t| < \infty$. If there is any unique vector Φ which makes $\Phi'\tau_t = 0$, then the $n \times s$ matrix Φ of rank s $(n > s > 0)$ is said to be contemporaneous mean co-breaking of order s for $\{x_t\}$. Given this definition of co-breaking, the fixed matrix Φ removes all changes in $\{\tau_t\}$ which varies from period to period, such that $E[\Phi' x_t - \Phi'\psi] = \Phi'\tau_t = 0$. This indicates that the parameterization of the reduced set of s linear transforms $\Phi' x_t$ is independent of deterministic shifts. Note that co-breaking occurs in a solved form when breaks between parameters are removed at the same point in time. This is shown that if $\Theta_t = \Theta$ is constant in equation (6.7), z_t and y_t co-break since μ_y^* is constant by construction:

$$\Phi'\mu_{xt} = (I_{n1} : -\Theta)(\mu_{yt},\ \mu_{zt})' = \mu_y^*, \tag{6.8}$$

where $\Phi' = (I_{n1} : -\Theta)$ represents a constant vector independent of breaks. To see how this co-breaking works in an econometric model, consider a simple conditional model

$$E(y_t|z_t, I_{t-1}) = \mu_{yt} + \delta(z_t - \mu_{zt}) = (\mu_{yt} - \delta\mu_{zt}) + \delta z_t, \tag{6.9}$$

where $\delta = \sum_{yz} \sum_{zz}^{-1}$. If $\Theta_t = \Theta$, combining (6.7) and (6.9) yields:

$$E[y_t|z_t, I_{t-1}] = \mu_{yt} + \delta\epsilon_{z,t} = \mu_y^* + (\Theta - \delta)\mu_{zt} + \delta z_t. \tag{6.10}$$

This equation is estimated efficiently, when $\Theta = \delta$ for all t and so z_t is weakly exogenous for Θ. If this holds, (6.10) becomes:

$$E[y_t|z_t, I_{t-1}] = \mu_y^* + \Theta z_t. \tag{6.11}$$

Using $v_t = y_t - E[y_t|I_{t-1}]$ in (6.4), (6.11) can be rewritten as $y_t = \mu_y^* + \Theta z_t + v_t$ with $E[z_t v_t] = 0$, which in turn, implies $E[y_t|I_{t-1}] = \mu_y^* + \Theta E\ [z_t|I_{t-1}]$ in (6.7) with $\Theta_t = \Theta$, where Θ is co-breaking for the expectations form. This final model has an expectational interpretation, but entails that economic agents act in a contingent manner. An important feature of the model is that it still maintains the constancy of its parameters even under regime changes in marginal processes and so refutes the Lucas (1976) critique. Note that this has been occurred under the assumptions that Θ_t is constant, z_t is weakly exogenous for Θ, and $(I_{n1} : -\Theta)$ is a co-breaking matrix. Hence, it is concluded that if there exist co-breaking relations between the parameters of conditional and marginal models and if conditioning variables are weakly exogenous for the parameters of interest in the former, then conditional models can be efficiently estimated even under regime changes and are immune to the Lucas critique.

6.3 Formulation of a Dynamic Model

6.3.1 Institutional backgrounds

During the sample period, the conduct of monetary policy in Korea may be divided into two distinctive sub-periods on the basis of degrees of government intervention: 1972-1980 and 1981-1997.[11] The former represents the period of the financial repression and the latter represents the period of the financial deregulation. In the early stage of its economic development before 1980, the peculiar characteristics of the economy can be described as the existence of underdeveloped money and capital markets, the predominance of government-owned banks in the financial system, a dualistic financial structure of organized and unorganized curb markets, and high debt-equity ratios of business firms. In this period, the relative priority of government policy was economic growth rather than price stability. In order to maintain the rapid economic growth and industrial development of the economy, Korean monetary authorities strictly controlled the whole economic system. Interest rates were regulated at low levels and most

[11] See Park (1994) for more details.

bank loans were directed to a few priority sectors designated by the government. This financially repressed development strategy had considerably contributed to the rapid economic growth, but resulted in perpetuated high inflation and large inefficiency of the whole economy. Meanwhile, the Bank of Korea (BOK), as the central bank, initially used the stock of M1 money as the main policy variable, but in 1979 changed its official policy variable from M1 to M2. In controlling the stock of money, the BOK largely relied on direct quantitative control methods, such as reserve requirements, credit ceilings, and formal or informal directives. The indirect methods, like open market operations, were not regarded as an important policy tool, partly because of the absence of well-developed money markets and partly because of a low interest rate policy.

In the early 1980s, when the inefficiency of government intervention apparently led the economy to serious macroeconomic imbalances and retard economic growth, the government changed its relative policy priority from the rapid economic growth to price stability and began, in a limited way, to undertake a series of deregulation measures, in order to enhance the efficiency of the financial system and to improve the effectiveness of government policy. These include the liberalization of interest rates on interbank money transactions and prime commercial papers, the privatization of government-owned commercial banks, and the reduction of direct controls on bank credits. With this partially deregulated environment, the economy had enjoyed high economic growth with stable price levels for a while. Remarkably reduced inflation kept fixed bank interest rates positive in real terms. However, after a decade of the favorable macroeconomic performance, the country internally and externally came under pressure to internationalize its financial markets. From 1993 the government again began to undertake a wide range of substantial deregulation measures to open the economy to foreign investors. Regulated interest rates were fully liberalized through a step-by-step process during the period from 1993 to 1997, entry barriers to nonbank financial intermediaries were substantially relaxed, and controls on capital accounts transactions were much reduced. As a consequence, non-bank financial intermediaries which were less regulated than banks considerably expanded and a large amount of foreign money flew into the domestic capital markets. To comply with this rapidly changing financial structure, the BOK changed its technique for policy implementation from a direct to indirect control in the late 1980s and its monetary target from the stock of M2 money to MCT in 1997.[12] Overall, it has been generally accepted that regulatory changes in the early 1980s have caused the economy to be more competitive and market oriented than before and altered the nature and extent of the monetary policy's influence on the real sector of the economy, even though there still exists a certain extent of direct regulation (see Park (1994) and Sikorski (1996) for more details).

6.3.2 A small dynamic macroeconometric model

To formulate a dynamic macroeconometric model for Korea, we use quarterly data spanning from 1972(3) to 1997(4). The data obtained from the various issues of the BOK's monthly bulletin include real GDP (Y), the consumer price index (P), a three-year corporate bond rate (R), and the stock of M2 money (M). Except for R, all variables were seasonally adjusted by using the X-11 procedure and transformed into logarithms. Hereafter lower case letters denote logs of the corresponding capitals. As a preliminary inspection, Figure 6.1 plots the basic data series in levels. m_t, y_t, and p_t are strongly trended, possibly suggesting that they may be nonstationary. However, it is noticeable that each of the variables moves together over the whole sample period with an evident break in the early 1980s. R_t also shows a similar pattern. This graphical examination of the data indicates that as discussed in the preceding sub-section, financial liberalization started from the early 1980s has severely influenced the macroeconomic developments of the economy. To check the nonstationarity of the data series, augmented

[12] MCT is defined as the stock of M2 money plus CDs and money in trust institutions.

Dickey-Fuller (ADF) tests were conducted with up to five auxiliary lags. The test results suggest that all variables appear to be $I(1)$.[13]

Based on the results of the ADF tests, we further tested for cointegration to examine the long-run relationship between these integrated series using Johansen's (1988) maximum likelihood procedure. A fifth-order VAR was initially estimated with a constant term and four impulse dummies.[14] Following Doornik *et al.* (1998), we didn't restrict the intercept term and dummies to lie in the cointegration space. The test results reported in Table 6.1(a) suggest that there is only one cointegrating vector in the system. Table 6.1(b) reporting the results of the test statistics for weak exogeneity shows that m_t and R_t are weakly exogenous with $\chi^2(1) = 0.99$ and $\chi^2(1) = 2.35$, respectively, and $\chi^2(2) = 4.96$, jointly.[15] This evidence of no contemporaneous feedbacks from the nonpolicy variables y_t and p_t to the policy variables m_t and R_t implies that over the sample period, the BOK has exogenously determined the latter control variables independently from the current developments of macroeconomic situations and so the long-run Granger causality between the variables runs from the monetary sector to the real sector.[16] The statistical exogeneity of the long-term interest rate R_t, which is otherwise expected to be an endogenous intermediate variable under the official operating target of money stocks by the BOK, seems to reflect its regulated nature directly and indirectly. From the standardized eigenvectors of the

[13]The values of the ADF test statistics are -1.62(5), -3.04(5), -2.37(5), and -2.09(5) for the levels of m_t, p_t, y_t, and R_t, respectively, and -4.14**(1), -3.78**(1), -3.91**(4), and -3.58**(4) for the corresponding differences. ** denotes significance at the 5% level and $(\cdot)$ represents the longest significant lag in the augmentation of the testing equation. In the case of m_t, y_t, and p_t, the tests for levels include a constant term and trend, but the tests for differenced series exclude trend. The tests for R_t and ΔR_t include only a constant term. The critical values were taken from MacKinnon (1991).

[14]The dummies included are D1, D2, D3, and D4 to capture the historical economic events of the economy, such as two oil price shocks in the 1970s, a massive reduction of regulated bank interest rates in 1982, and the financial crisis in 1997, respectively. All take zero except for unity in 1974(1) and 1975(1)(2) for D1, 1980(1) for D2, 1982(1) for D3, and 1997(4) for D4.

[15]For a robust result, we performed weak exogeneity tests without the dummies which unrestrictedly enter the initial VAR system. The same conclusions were obtained from the tests.

[16]See Toda and Phillips (1993) for the test of long-run Granger causality in the I(1) system.

cointegration test, a long-run equilibrium relation between variables is derived as $\beta' x_t = p_t + 1.54 y_t - 1.12 m_t - 0.006 R_t$. However, this equation does not give any economic meaning and the coefficient of y_t is not significant as well. To identify it as an interpretable equation, some hypothetical restrictions were imposed (Johansen and Juselius,1992). This involves the joint weak exogeneity of m_t and R_t on the loading matrix and a price equation in the cointegration space with no effect from output. The conventional likelihood ratio (LR) test accepts the restrictions at the 5% level with the test statistic $\chi^2(3) = 5.47$ and yields an aggregate supply equation $\beta' x_t = p_t - 0.47 m_t - 0.07 R_t$ which is interpretable that in the long run the price level is mainly determined by the stock of money and interest rates.

With this identified cointegrating vector, the initial VAR was reparameterized into a three-lag *I(0)* vector error correction model (VECM). Since the unrestricted model is so highly overparameterized that the estimated model may capture accidental features of the sample, we sequentially eliminated insignificant parameters to reduce the sample dependence of the model and finally derived a dynamic structural model estimated by full information maximum likelihood (FIML).[17] The obtained models are:

$$\begin{array}{rcccc}
\Delta y_t = & 0.36\Delta m_{t-1} & +0.18\Delta m_{t-2} & +0.11\Delta m_{t-3} & -0.12\Delta p_{t-3} \\
& (0.09) & (0.10) & (0.10) & (0.11) \\
& -0.41\Delta y_{t-1} & -0.14\Delta y_{t-2} & -0.002\Delta R_{t-3} & +0.03 ecm_{t-1} \\
& (0.10) & (0.09) & (0.001) & (0.01) \\
& -0.03D4 & +0.06 & & \\
& (0.02) & (0.01) & &
\end{array}$$

$$\hat{\sigma} = 0.018,\ F_{AR}(5,75) = 2.32,\ \chi^2_N(2) = 15.85,\ F_{ARCH}(4,72) = 0.74,$$

[17]Structure has many meanings in econometrics (see Hendry (1995)). In dynamic econometrics, if a model is 'invariant' and directly characterizes the relations of the economy under analysis, the model is regarded as containing some structure.

$F_H(30,49) = 1.10,$

$$\begin{array}{lllll}
\Delta p_t = & 0.001\Delta R_{t-3} & +0.11\Delta m_{t-2} & +0.12\Delta p_{t-1} & +0.13\Delta p_{t-2} \\
& (0.001) & (0.05) & (0.07) & (0.07) \\
& +0.05\Delta y_{t-1} & -0.14\Delta y_{t-3} & +0.002\Delta R_t & -0.03ecm_{t-1} \\
& (0.05) & (0.04) & (0.001) & (0.01) \\
& +0.04D1 & +0.04D2 & -0.02D3 & -0.03 \\
& (0.01) & (0.01) & (0.01) & (0.01)
\end{array}$$

$\hat{\sigma} = 0.009,\ F_{AR}(5,75) = 2.61,\ \chi^2_N(2) = 14.46,\ F_{ARCH}(4,72) = 4.01,$
$F_H(30,49) = 1.17,$

$$\begin{array}{lllll}
\Delta m_t = & 0.23\Delta m_{t-2} & +0.27\Delta m_{t-3} & -0.004\Delta R_{t-1} & +0.003\Delta R_{t-2} \\
& (0.09) & (0.09) & (0.001) & (0.001) \\
& +0.17\Delta p_{t-1} & -0.08\Delta y_{t-2} & +0.03D3 & -0.03D4 \\
& (0.10) & (0.09) & (0.02) & (0.02) \\
& +0.02 & & & \\
& (0.01) & & &
\end{array}$$

$\hat{\sigma} = 0.018,\ F_{AR}(5,75) = 1.67,\ \chi^2_N(2) = 3.17,\ F_{ARCH}(4,72) = 1.49,$
$F_H(30,49) = 0.39,$

$$\begin{array}{lllll}
\Delta R_t = & -1.63\Delta m_{t-1} & +17.41\Delta p_{t-2} & -18.40\Delta p_{t-3} & +0.27\Delta R_{t-1} \\
& (5.85) & (8.97) & (8.86) & (0.09) \\
& -0.26\Delta R_{t-2} & +17.01\Delta y_{t-2} & +11.54\Delta y_{t-3} & +5.29D2 \\
& (0.08) & (6.58) & (6.01) & (1.29) \\
& -6.20D3 & +4.85D4 & -0.51 & \\
& (1.27) & (1.20) & (0.37) &
\end{array}$$

$\hat{\sigma} = 1.195,\ F_{AR}(5,75) = 1.99,\ \chi^2_N(2) = 11.87,\ F_{ARCH}(4,72) = 0.96,$

$F_H(30, 49) = 1.92$,

where $ecm_t = ecm_{t-1} + \Delta p_t - 0.47\Delta m_t - 0.07\Delta R_t$. $\hat{\sigma}$ denotes the standard deviation of residuals; F_{AR} denotes a LM test for fifth-order autocorrelation; χ^2_N denotes a test statistic for normality (Doornik and Hansen, 1994); F_{ARCH} denotes a LM test for fourth-order ARCH (Engle, 1982a); F_H denotes the White (1980) test for heteroscedasticity.

The diagnostic tests of mis-specification indicate that the empirical models perform well except for non-normality in the equations for ΔR_t, Δp_t, and Δy_t, mainly caused by excess kurtosis.[18] The test of overidentification, following Hendry and Mizon (1993), accepts parameter restrictions at the 5% level with the LR statistic $\chi^2(30) = 13.07$. This result of the system-based parsimonious encompassing test ensures valid reduction from the benchmark VECM without any loss of information. To check parameter constancy, the system was reestimated by recursive FIML. Figure 6.2(a) plots the recursive one-step residuals of each equation and the corresponding two standard errors (i.e., $\{Y_t - \widehat{\beta}'_t X_t\}$ and $\{0 \pm 2\hat{\sigma}_t\}$ in a standard system notation). All the residuals are within the two-standard-error-bands, showing no evidence of large parameter changes. This is complimentarily confirmed by Figure 6.2(b) which plots the system-based breakpoint Chow (1960) test statistics scaled by their 5% critical values for the sequences $\{79(4) - 97(4), 80(1) - 97(4), ..., 97(3) - 97(4), 97(4)\}$.[19] None of the tests reject parameter constancy. As expected, the error correction terms appear only in the equations for Δy_t and Δp_t, reflecting the endogenous property of these variables in the system. All variables interact in a dynamic way through lags. Only an exception is ΔR_t which contemporaneously affects prices. The

[18]Error normality is the issue related to asymtotic theory. Thus, we do not regard this problem seriously, because of our finite samples and the non-existence of serial correlation and heteroscedasticity in the models (see Hansen (1999)).

[19]The system-based breakpoint Chow tests are the $F-$statistics approximated by the R^2 computed from $1 - \exp(-2\widehat{l}_{t-1} + 2\widehat{l}_T)$, $t = M, ..., T$, where M is the first observed sample and $\widehat{l}_t = -1/2 \log |T^{-1}\widehat{V}_t'\widehat{V}_t|$ from a standard system equation $Y_t = \beta X_t + V_t$.

coefficient of Δm_{t-1} in the equation ΔR_t is insignificant, but it was included to capture dynamic interactions between money and interest rates in impulse response analysis.

Changes in output are substantially influenced by money. In the short-run, the combined effect from money is 0.55, but the effects from prices and interest rates are negative with small magnitudes. The antilog coefficient of about - 3% in $D4$ shows a strong negative impact of the financial crisis in 1997 on the economy. Changes in prices are mostly driven by money and its own lags, and somewhat by output and interest rates. Two oil price shocks in 1970s caused quarterly inflation to rise by about 8%. About 2% decrease in Δp_t from $D3$, which is an intervention dummy to capture a massive reduction of regulated bank interest rates in 1982, seems to reflect a mark-up pricing behavior of the Korean business firms by passing reduced interest rate costs to commodity prices. The speed of adjustment to equilibrium is very slow, 3% of any disequilibrium being removed each quarter. Changes in money which is the main policy variable are largely influenced by its own past history. Nevertheless, there is, to some extent, a positive effect from prices. The long-run elasticity of Δp_t on the changes in money is 0.34. The combined short-run effect of changes in interest rates is statistically significant, but negligible in terms of magnitude. Short-run changes in the interest rate are mainly determined by both prices and output, whereas the effect from money is statistically insignificant with a small coefficient. The long-run elasticity of output in this semilog form model is about 0.29. The coefficients of intervention dummies indicate that the economic events of $D3$, $D4$, and $D2$ caused ΔR_t to fall by 6.20% and to rise by 4.85% and 5.29%, respectively, in absolute value. Overall, the estimated dynamic system seems to well characterize the basic structure and historical events of the economy. This model is now regarded as a baseline model to examine the effects of transitory and permanent changes in policy on economic activities.

6.4 Evaluation of Policy Changes

6.4.1 Transitory changes

This section considers the effects of unanticipated transitory changes in policy on the economy. To do this, we examine the dynamic impulse responses of the system to initial shocks on each series by applying the one standard deviation of errors from the corresponding equations in the VECM model. Even though this non-orthogonal impulse response analysis doesn't consider the correlation between equation residuals, the empirical results derived from this approach are invariant to the arbitrary ordering of variables. Furthermore, the casual relationship between variables is explained from the empirical results of the weak exogeneity tests conducted in the previous section. All the dummies in the system were eliminated, so that they were treated as fixed in the following impulse response analysis. Figure 6.3 displays the graphs of impulse responses that trace the impacts of specific shocks on levels of each series. The graphs show that shocks to individual variables do not die out, reflecting the non-stationarity of variables in the model, and so have permanent effects with non-zero convergence (see Lütkepohl (1991)). The dynamic interactions are largely in line with *a priori* expectations. A shock to the stock of money, which has been used as the main policy variable by the BOK, positively affects prices and output. It is noteworthy that even if the empirical model includes an interest rate, the effect of money on output is large and permanent, thus showing the long-run monetary non-neutrality. This evidence is contrasted with a common finding in the literature that if an interest rate is included in a VAR model, the relationship between money and output is weakened since the interest rate predicts a considerable fraction of movement in the stock of money (see Sims (1982a) and Litterman and Weiss (1985)). In the initial four quarters, the interest rate negatively reacts to the shock on the money stock because of a liquidity effect, but it is eventually dominated by expected inflation and shows positive correlations with the money shock in the long run. These short-run negative responses of the interest rate are consistent with the recent findings of Strongin (1995) and Bernanke and Mihov (1998). However, it should be noted that our results were obtained from a statistically identified structure, while the previous two results were obtained from an arbitrary identified structure.

An impulse on the interest rate leads money to decline quicker than output which responds with a slightly slow decline. This pattern seems to be consistent with Sims' (1992) view that monetary policy disturbances are important in generating aggregate fluctuations; interest rates surprises represent monetary policy shocks and monetary contraction generates declining money and output. Prices to the interest rate shock negatively respond. This may be because monetary contraction reduces nominal aggregate demand and eventually causes output and interest rates to decline through the interactions of deflationary pressure with price stickiness.[20] An impulse on prices causes money to rise and output to initially fluctuate at its equilibrium level but after seven quarters to be risen sharply. The interest rate positively reacts to the shock over the first three quarters because of the Fisher effect of prices, but after that, rapidly declines. A shock on output negatively affects money. The negative responses of money seem to reflect a tightened policy to decrease the increased aggregate demand and to show that money is not determined endogenously by following either technology or real stochastic shocks from the real sector, as argued by King and Plosser (1984). This argument may be further evidenced by the responses of the interest rate and prices to the output shock. The interest rate is positively responded to the output shock at first because of the increased aggregate demand. Then a tightened policy causes

[20]For developed countries, Sims (1992) finds the positive comovements of prices with interest rates ('price puzzle') and explains the patterns with policy endogeneity. Recently, Christiano *et al.* (1996) and Kim (1999) solve the puzzle. However, their evidence is based on a pre-assumed identification. Thus, if the assumption is changed, the results will be changed.

the interest rate to be further accelerated. Also, prices are positively responded to an output shock at first, but after three quarters rapidly declined, possibly because of the monetary contraction to cool down the heated economy. Overall, the results of the impulse responses discussed so far suggest that an unexpected transitory shock on money has a substantial effect on the economy. In spite of the long-run non-neutrality of money, this empirical result seems to be largely consistent with the conventional economic theory which emphasizes the role of money and also complies with the policy scenarios assumed by Korean monetary authorities.[21]

6.4.2 Permanent changes

In this subsection we investigate whether permanent changes in policy have their expected results as planned. To do this, we primarily focus on financial liberalization with two reasons. First, it is widely accepted that financial liberalization has caused the economy to be changed structurally. Second, it is still controversial whether financial liberalization has contributed to economic growth. According to the literature of economic development, financial liberalization is advocated since deregulated high real interest rates and financial deepening increase overall savings and investable funds, finally leading to economic growth (McKinnon, 1973; Kapur, 1976). Thus, a policy prescription for a financially repressed economy is either to raise regulated low bank interest rates or to reduce the rate of inflation. In the case of Korea, as shown in Figure 1, all data series are structurally broken in the early 1980s. The breakpoint is coincided with the inauguration of the fifth regime in 1981, which changed the relative priority of monetary policy from the rapid economic growth to the stability of price levels and began to lib-

[21] Cagan (1969) argues that since in the real world expectations are imperfect and prices adjust slowly, changes in the money stock disturb the pattern of expenditures and output and so, neutral money is a figment of abstract theory in which changes in the money stock are fully anticipated and prices change proportionately.

eralize the economy. This clearly indicates that the regulatory change in policy has significantly affected the real sector. In this context, our major concern is whether financial liberalization in Korea has contributed to stabilizing price levels and maintaining economic growth, as stated. This issue is evaluated in this subsection by using the concept of co-breaking explained in Section 2.

There are many possible ways to demonstrate co-breaking between variables. A simple way to do this is to extend the underlying dynamic system by adding a step dummy which represents financial liberalization as a permanent change in policy and to check the statistical significance and parameter constancy (see Hendry and Mizon, 1998). Following this strategy, the dummy DS, which is 0 before 1981(1) and 1 thereafter, was introduced to the initial unrestricted VAR and its statistical significance in the system was checked. Under the constancy of the conditional models in the dynamic system in Section 6.3, the step dummy DS is regarded as a kind of omitted variables. The conventional F-test shows that the variable is significant at the 5% level with the test statistic $F(4, 68) = 4.58$. Furthermore, even with this variable the Johansen cointegration test still suggests one cointegrating vector, and R_t and m_t also maintain their initial exogeneity properties with $\chi^2(1) = 2.63$ and $\chi^2(1) = 0.001$, respectively, and with $\chi^2(2) = 2.64$, jointly. Identifying restrictions on the cointegrating vector, which are similar to those in Section 3, are accepted with $\chi^2(3) = 2.76$ at the 5% level. Reformulating the $I(1)$ system into a stationary system with an identified long-run equilibrium equation $\beta' x_t = p_t - 0.73m_t - 0.17R_t$ and conducting sequential reductions, we obtained the following conditional models for Δy_t and $\Delta p_t$with an overidentified test statistic $\chi^2(31) = 22.76$ which is insignificant at the 5% level:

$$
\begin{array}{rllll}
\Delta y_t = & 0.38\Delta m_{t-1} & +0.20\Delta m_{t-2} & +0.12\Delta m_{t-3} & -0.06\Delta p_{t-3} \\
& (0.09) & (0.10) & (0.10) & (0.11) \\
& -0.43\Delta y_{t-1} & -0.16\Delta y_{t-2} & -0.001\Delta R_{t-3} & +0.02 ecm_{t-1} \\
& (0.10) & (0.10) & (0.001) & (0.004) \\
& -0.03 D4 & +0.08 & +0.01 DS & \\
& (0.02) & (0.02) & (0.006) &
\end{array}
$$

$\hat{\sigma} = 0.018,\ F_{AR}(5, 74) = 3.55,\ \chi^2_N(2) = 12.76,\ F_{ARCH}(4, 71) = 0.64,$
$F_H(31, 47) = 1.19,$

$$
\begin{array}{rlllll}
\Delta p_t = & 0.001\Delta R_{t-3} & +0.11\Delta m_{t-2} & +0.15\Delta p_{t-1} & +0.13\Delta p_{t-2} & \\
& (0.001) & (0.05) & (0.08) & (0.07) & \\
& +0.05\Delta y_{t-1} & -0.13\Delta y_{t-3} & +0.002\Delta R_t & -0.007 ecm_{t-1} & \\
& (0.05) & (0.05) & (0.001) & (0.002) & \\
& +0.04 D1 & +0.04 D2 & -0.02 D3 & -0.03 & -0.009 DS \\
& (0.01) & (0.01) & (0.01) & (0.01) & (0.003)
\end{array}
$$

$\hat{\sigma} = 0.009,\ F_{AR}(5, 74) = 2.82,\ \chi^2_N(2) = 19.19,\ F_{ARCH}(4, 71) = 3.32,$
$F_H(31, 47) = 1.41.$

All the reestimated coefficients and diagnostic tests are closely matched to those of the initial model in Section 3, and none of the system-based breakpoint Chow test statistics reported in Figure 6.4 reject parameter constancy.[22] The significance of the regime shift dummy DS in each equation is summarized as follows

[22]Note that the parameter constancy has been achieved with the presence of several dummies representing structural breaks. This indicates that in the absence of the dummies parameter non-constancy is apparent.

	Δm_t	ΔR_t	Δp_t	Δy_t
coefficient	-0.02	-0.70	-0.009	0.01
t-value	3.33	1.80	3.00	1.67

The coefficients are significant at the 5% levels in the equations for Δm_t and Δp_t and at the 10% levels in the equations for ΔR_t and Δy_t. This evidence of co-breaking shows that financial liberalization has clearly caused the economy to change structurally. Park (1994) argues the possible endogeneity of financial liberalization from the real sector. However, the evidence of weak exogeneity tests shows that the causality runs from policy variables Δm_t and ΔR_t to economic variables Δy_t and Δp_t. This rules out the endogenous view of financial liberalization. The step-shift dummy negatively affects price levels, interest rates, and money growth, and positively output. This may be interpreted in a way that a tight money supply for the stabilization of the economy has lowered the (expected) inflation rate; the reduced price levels have in turn raised the real rates of fixed bank interest, the demand for real money, and the supply of real money (bank credits); and eventually, the increased availability of the real bank credits has led to economic growth.[23] Note that even though the growth rate of nominal money supply has been lowered, the average growth rate of real money supply in the period of 1981-1997 has been increased by 3% from that of the period 1972-1980. Furthermore, the real rates of regulated bank interests were negative until 1980 because of perpetuated high inflation, but since then, the rates became positive even though the nominal rates have been scaled down by the government in the course of a major disinflation. This has also forced market interest rates, which are closely tied to bank interest rates, to fall (see McKinnon (1989)).

[23]Kapur (1976) shows that the direct effect of a reduced money supply is negative on the rate of economic growth, but the indirect effect through the increase in real money demand caused by lowered inflation is positive.

From this analysis, it is concluded that financial liberalization in Korea, even if it has been gradually undertaken in a limited way, has contributed to maintaining price stability and sustaining economic growth. This finding is largely consistent with the view of financial liberalization proponents, but is contrasted with a recent work of Demetriades and Hussein (1996) who find a bi-directional causality between financial liberalization and economic growth. Econometrically, the comovements between policy and nonpolicy variables may induce the Lucas critique which concerns the impossibility of constant conditional models associated with non-constant marginal processes under the rational behavior of agents. However, our empirical findings on the weak exogeneity of policy variables and co-breaking between policy and nonpolicy variables with no parameter changes in behavioral equations ensure that the initial conditional models still maintain their invariant property even under regime changes and so are not subject to the Lucas (1976) critique.

6.5 Conclusions

This chapter has questioned why currently available macroeconometric models separately focus on evaluating the effects of transitory and permanent policy changes in a divergent way and shown that both changes can be evaluated jointly within a simultaneous dynamic model. The econometric advantages of the joint examination are that the 'invariant' property of the underlying model is examined and that information on data is obtained from both parameter spaces and error terms. Methodologically, the study extended Hendry and Mizon's (1998) work into a structurally multivariate framework by formally incorporating the usefulness of weak exogeneity to identifying policy actions. Then, the empirical feasibility of the approach was examined using Korean data. In a structural dynamic model complementing the limitations of traditional structural and VAR models, the effects of transitory and permanent changes in policy were evaluated by using impulse response functions and the concept of co-breaking, respectively. Distinctive features of this approach are that even under regime shifts the baseline model does not suffer from the classical Lucas (1976) critique since co-breaking cancels out structural breaks between policy and nonpolicy variables and from the Sims (1980) critique since identifying restrictions are formally tested.

The empirical results show that both changes substantially affect the economy. An unanticipated transitory shock on 'nominal' money positively affects output with the long-run non-neutrality, but in the case of the permanent change in which prices are fully adjusted, 'real' money rather than nominal money positively affects output. Financial liberalization as the proxy for a permanent regime change in policy clearly contributed to stabilizing price levels and to sustaining economic growth. This evidence of the Korean case supports the view of financial liberalization proponents in the literature of economic development and may shed some light to the development strategy of other developing countries. An important policy implication of the results in this study is that as long as price stability maintains, an expansion of money, whether it is anticipated or not, may contribute to enhancing economic growth. However, in the case that money is nonneutral, there may exist an inverse relationship between inflation and economic growth as captured in the Phillips curve. That is, reducing inflation may require the loss of a certain portion in economic growth (see Okun (1978) and Gordon and King (1982)). In this context, changes in policy should be conducted in such a way that the optimal trade-off is achieved.

Future research in this direction may widen this approach in a more practical way. For example, this study assumes that regime changes are known and immediately affect the economy. However, this might not be a realistic approach. Alternatively, the suggestions of Andrews (1993) and Hansen (2000) to find unknown regime changes can be used in a way that the regimes detected may intertemporarily affect the economy. In this context, intertemporal mean co-breaking between policy and nonpolicy variables may be more appropriate to assess the effects of permanent changes than the contemporaneous mean co-breaking used in this chapter.

Appendix A. Effects of Transitory and Permanent Changes in Policy in the Cointergrated Structural VECM

In this appendix we discuss how the effects of transitory and permanent changes in policy on the economy can be analyzed in a structural vector equilibrium (VECM) model. The analysis is similar to that employed in Section 2.

A1. Transitory changes

Consider the following unrestricted vector autoregression (VAR) as in (6.1):

$$\mathbf{x}_t = \sum_{j=1}^{p} \mathbf{A}_j \mathbf{x}_{t-j} + \boldsymbol{\epsilon}_t, \quad \text{where} \quad \boldsymbol{\epsilon}_t \sim \mathsf{IN}_n\left[\mathbf{0}, \boldsymbol{\Sigma}\right], \tag{A1-1}$$

which can be re-parameterized as a vector equilibrium correction model (VECM):

$$\Delta \mathbf{x}_t = \sum_{j=1}^{p-1} \boldsymbol{\Gamma}_j \mathbf{x}_{t-j} + \mathbf{A}(1)\mathbf{x}_{t-1} + \boldsymbol{\epsilon}_t, \quad \text{where} \quad \boldsymbol{\epsilon}_t \sim \mathsf{IN}_n\left[\mathbf{0}, \boldsymbol{\Sigma}\right], \tag{A1-2}$$

when

$$\mathbf{A}(1) = -\left(\mathbf{I}_n - \sum_{j=1}^{p} \mathbf{A}_j\right), \quad \boldsymbol{\Gamma}_j = -\sum_{i=j+1}^{p} \mathbf{A}_i.$$

In the case that the data are integrated of order $\mathsf{I}(1)$ with γ cointegrating vectors $\mathbf{A}(1)$ has rank γ and takes the form $\mathbf{A}(1) = \boldsymbol{\alpha}\boldsymbol{\beta}'$ with $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$ being $(n \times \gamma)$ matrices of rank γ. Let $\mathbf{x}_t' = (\mathbf{y}_t' : \mathbf{z}_t')$ where $\mathbf{y}_t$ and $\mathbf{z}_t$ are $n_1 \times 1$ and $n_2 \times 1$ vectors with $n_1 + n_2 = n$. Then for the case that $p = 2$ (A1-2) can be written as:

$$\begin{bmatrix} \Delta \mathbf{y}_t \\ \Delta \mathbf{z}_t \end{bmatrix} = \begin{bmatrix} \boldsymbol{\Gamma}_{11} & \boldsymbol{\Gamma}_{12} \\ \boldsymbol{\Gamma}_{21} & \boldsymbol{\Gamma}_{22} \end{bmatrix} \begin{bmatrix} \Delta \mathbf{y}_{t-1} \\ \Delta \mathbf{z}_{t-1} \end{bmatrix} \tag{A1-3}$$

$$- \begin{bmatrix} \mathbf{A}_{11}(1) & \mathbf{A}_{12}(1) \\ \mathbf{A}_{21}(1) & \mathbf{A}_{22}(1) \end{bmatrix} \begin{bmatrix} \mathbf{y}_{t-1} \\ \mathbf{z}_{t-1} \end{bmatrix} + \begin{bmatrix} \epsilon_{y,t} \\ \epsilon_{z,t} \end{bmatrix},$$

where $\mathbf{A}_{11}(1) = (\boldsymbol{\alpha}_{11}\boldsymbol{\beta}'_{11} + \boldsymbol{\alpha}_{12}\boldsymbol{\beta}'_{21})$, $\mathbf{A}_{12}(1) = (\boldsymbol{\alpha}_{11}\boldsymbol{\beta}'_{12} + \boldsymbol{\alpha}_{12}\boldsymbol{\beta}'_{22})$, $\mathbf{A}_{21}(1) = (\boldsymbol{\alpha}_{21}\boldsymbol{\beta}'_{11} + \boldsymbol{\alpha}_{22}\boldsymbol{\beta}'_{21}), \mathbf{A}_{22}(1) = (\boldsymbol{\alpha}_{21}\boldsymbol{\beta}'_{12} + \boldsymbol{\alpha}_{22}\boldsymbol{\beta}'_{22})$ and the error covariance matrix is $\boldsymbol{\Sigma} = \begin{bmatrix} \boldsymbol{\Sigma}_{yy} & \boldsymbol{\Sigma}_{yz} \\ \boldsymbol{\Sigma}_{zy} & \boldsymbol{\Sigma}_{zz} \end{bmatrix}$. Johansen (1992) shows that a necessary and sufficient condition for $\mathbf{z}_t$ to be weakly exogenous for $\boldsymbol{\beta}$ in (A1-3) is $\boldsymbol{\alpha}_{2j} = 0$ for $j = 1, 2$. Hendry and Mizon (1993) provide an alternative set of sufficient conditions for weak exogeneity when the parameters of interest are a subset of $\boldsymbol{\beta}$.

A simultaneous equations model of the distribution of $\Delta\mathbf{y}_t$ and $\Delta\mathbf{z}_t$ can be written in VECM format (often called a structural VECM) as:

$$\begin{bmatrix} \mathbf{B}_{11} & \mathbf{B}_{12} \\ \mathbf{B}_{21} & \mathbf{B}_{22} \end{bmatrix} \begin{bmatrix} \Delta\mathbf{y}_t \\ \Delta\mathbf{z}_t \end{bmatrix} = \begin{bmatrix} \mathbf{C}_{11} & \mathbf{C}_{12} \\ \mathbf{C}_{21} & \mathbf{C}_{22} \end{bmatrix} \begin{bmatrix} \Delta\mathbf{y}_{t-1} \\ \Delta\mathbf{z}_{t-1} \end{bmatrix} - \begin{bmatrix} \mathbf{A}^*_{11}(1) & \mathbf{A}^*_{12}(1) \\ \mathbf{A}^*_{21}(1) & \mathbf{A}^*_{22}(1) \end{bmatrix} \begin{bmatrix} \mathbf{y}_{t-1} \\ \mathbf{z}_{t-1} \end{bmatrix} + \begin{bmatrix} \nu_{y,t} \\ \nu_{z,t} \end{bmatrix} \tag{A1-4}$$

with $\mathbf{A}^*_{11}(1) = (\boldsymbol{\alpha}^*_{11}\boldsymbol{\beta}'_{11} + \boldsymbol{\alpha}^*_{12}\boldsymbol{\beta}'_{21})$, $\mathbf{A}^*_{12}(1) = (\boldsymbol{\alpha}^*_{11}\boldsymbol{\beta}'_{12} + \boldsymbol{\alpha}^*_{12}\boldsymbol{\beta}'_{22})$, $\mathbf{A}^*_{21}(1) = (\boldsymbol{\alpha}^*_{21}\boldsymbol{\beta}'_{11} + \boldsymbol{\alpha}^*_{22}\boldsymbol{\beta}'_{21}), \mathbf{A}^*_{22}(1) = (\boldsymbol{\alpha}^*_{21}\boldsymbol{\beta}'_{12} + \boldsymbol{\alpha}^*_{22}\boldsymbol{\beta}'_{22})$, when the error covariance matrix $\boldsymbol{\Omega} = \begin{bmatrix} \boldsymbol{\Omega}_{yy} & \boldsymbol{\Omega}_{yz} \\ \boldsymbol{\Omega}_{zy} & \boldsymbol{\Omega}_{zz} \end{bmatrix}$ and $\begin{bmatrix} \mathbf{B}_{11} & \mathbf{B}_{12} \\ \mathbf{B}_{21} & \mathbf{B}_{22} \end{bmatrix}$ are nonsingular matrices. The parameters of (A1-4) are unidentified, with identification usually achieved via a priori restrictions based on economic theory or previous empirical evidence. For example, the necessary order condition for $\boldsymbol{\beta}$ to be identified is the existence of γ^2 restrictions, including normalizationon, $\boldsymbol{\alpha}$ and $\boldsymbol{\beta}$. However, it is common for there to be more than γ^2 linearly independent restrictions and so the parameters are over-identified. Sufficient conditions for $\Delta\mathbf{z}_t$ to be independent of the contemporaneous developments in $\Delta\mathbf{y}_t$ are $\mathbf{B}_{21} = 0, \mathbf{B}_{12} = \boldsymbol{\Sigma}_{yz}\boldsymbol{\Sigma}_{zz}^{-1}$, and $\boldsymbol{\alpha}^*_{2j} = 0$ for $j = 1, 2$ (see Johansen (1992) and Urbain (1992)). The first two conditions imply that $\boldsymbol{\Omega}_{yz} = \boldsymbol{\Omega}'_{zy} = 0$, and the final conditions ensure that the cointegrating vectors do not occur in the n_2 equations determining $\Delta \mathbf{z}_t$ which is effectively an hypothesis of long run causality from $\mathbf{z}_t$ to $\mathbf{y}_t$ (see Hendry and Mizon, 1999b). If these restrictions are satisfied then $\mathbf{z}_t$ is weakly exogenous for the parameters $\boldsymbol{\beta}$ in (A1-4).

A common approach to policy analysis is to estimate the response of $\Delta \mathbf{y}_t$ to impulse or transitory changes in policy variables. When such impulse response analysis is done by estimating the response of $\Delta \mathbf{y}_t$ to an impulse in $\epsilon_{z,t}$ in (A1-3) this ignores the correlation between $\epsilon_{y,t}$ and $\epsilon_{z,t}$. Transforming the VECM to achieve orthogonal errors (as in the Choleski transformation) avoids this problem but introduces others: it violates weak exogeneity for most orderings of the variables in $\mathbf{y}_t$ and $\mathbf{z}_t$; and the measured responses are no longer those of $\Delta \mathbf{y}_t$ to the policy impulse, but those of a linear combination of variables. Indeed, the Choleski decomposition corresponds to a transformation that imposes, without testing, a complete causal sequencing of variables that depends on the ordering of the variables, which is often arbitrary. Further, Hendry and Mizon (1998) showed that the effects of a change in an error and one in an intercept are indistinguishable in the model, and yet this will only be so in reality if the policy variables are weakly exogenous. Hence it is important to test the validity of over-identifying restrictions and establish that the policy variables are weakly exogenous for at least the equilibrium parameters $\boldsymbol{\beta}$, as is done in this chapter.

A2. Permanent changes

Given the conditions of weak exogeneity in Appendix A1, the system (A1-3) can be rewritten as the following conditional and marginal models by introducing intercept terms;

$$\begin{bmatrix} \Delta y_t \\ \Delta z_t \end{bmatrix} = \begin{bmatrix} (\omega_1 - \delta_1 \omega_2) \\ \omega_2 \end{bmatrix} + \begin{bmatrix} \delta_1 \Delta z_t \\ 0 \end{bmatrix} \tag{A2-1}$$

$$
-\begin{bmatrix} \pi_{11} & \pi_{12} \\ 0 & 0 \end{bmatrix} \begin{bmatrix} y_{t-1} \\ z_{t-1} \end{bmatrix}
+\begin{bmatrix} (\Gamma_{11}-\delta_1\Gamma_{21}) & (\Gamma_{12}-\delta_1\Gamma_{22}) \\ \Gamma_{21} & \Gamma_{22} \end{bmatrix} \begin{bmatrix} \Delta y_{t-1} \\ \Delta z_{t-1} \end{bmatrix}
+\begin{bmatrix} v_{yt} \\ \epsilon_{zt} \end{bmatrix},
$$

where

$$
\begin{aligned}
\delta_1 &= \mathbf{\Sigma}_{yz}\mathbf{\Sigma}_{zz}^{-1}, \ \pi_{11} = (\alpha_{11}\beta'_{11} + \alpha_{12}\beta'_{21}) - \delta_1(\alpha_{21}\beta'_{11} + \alpha_{22}\beta'_{21}), \\
\pi_{12} &= (\alpha_{11}\beta'_{12} + \alpha_{12}\beta'_{22}) - \delta_1(\alpha_{21}\beta'_{12} + \alpha_{22}\beta'_{22}), \\
v_{yt} &= \varepsilon_{yt} - \delta_1\varepsilon_{zt},
\end{aligned}
$$

such that

$$
\begin{aligned}
E[\Delta y_t|\Delta z_t, I_{t-1}] &= (\omega_1 - \delta_1\omega_2) + \delta_1\Delta z_t - \pi_{11}y_{t-1} \\
&\quad -\pi_{12}z_{t-1} + (\Gamma_{11} - \delta_1\Gamma_{21})\Delta y_{t-1} \\
&\quad +(\Gamma_{12} - \delta_1\Gamma_{22})\Delta z_{t-1}.
\end{aligned} \tag{A2-2}
$$

Fulfilling the condition of equation (6.8) in Section 2 implies :

$$
\Gamma_{11} = \Theta\Gamma_{21}, \ \Gamma_{12} = \Theta\Gamma_{22},
$$

so that (A2-2) leads to:

$$
\begin{aligned}
E[\Delta y_t|\Delta z_t, I_{t-1}] &= (\omega_1 - \delta_1\omega_2) + \delta_1\Delta z_t - \pi_{11}y_{t-1} - \pi_{12}z_{t-1} \\
&\quad +(\Theta - \delta_1)(\Gamma_{21}\Delta y_{t-1} + \Gamma_{22}\Delta z_{t-1}),
\end{aligned} \tag{A2-3}
$$

which implies that z_t is weakly exogenous for Θ if and only if $\Theta = \delta_1$, in which case y_t and z_t co-break, and then becomes;

$$E[\Delta y_t|\Delta z_t, I_{t-1}] = (\omega_1 - \Theta\omega_2) + \Theta\Delta z_t - \pi_{11}y_{t-1} - \pi_{12}z_{t-1}.$$

This equation shows that the effects of permanent changes in Δz_t on Δy_t can be analyzed by measuring the immediate responses of Δy_t in the intercept term without being subject to the Lucas critique.

Appendix B. Infinite Effects of Transitory and Permanent Changes in Policy

This appendix discusses the infinite effects of transitory and permanent changes in policy on the economy in a reduced-form VECM framework. With the assumption of $A_0 = I_n$, equation (A1-1) can be rewritten as follows:

$$\Delta x_t = \sum_{i=1}^{p-1} \Gamma_i \Delta x_{t-i} + \alpha\beta' x_{t-1} + \epsilon_t, \qquad \text{(B1-1)}$$

where $\Gamma_i = -\sum_{j=i+1}^{p} A_j$. Assume that Δx_t has a Wold representation:

$$\begin{aligned} \Delta x_t &= C(L)(\epsilon_t - \Pi\gamma - \alpha\eta) \qquad \text{(B1-2)} \\ &= [C(1) + \Delta C^*(L)]\,(\epsilon_t - \Pi\gamma - \alpha\eta) \\ &= C(1)(\epsilon_t - \Pi\gamma) + C^*(L)\Delta(\epsilon_t - \Pi\gamma - \alpha\eta), \end{aligned}$$

where $\Pi = \sum_{i=1}^{p-1}\Gamma_i - I_n$, $\gamma = E[\Delta x_t]$, $\eta = E[\beta' x_t]$, $C(1) = \beta_\perp(\alpha'_\perp \Pi \beta_\perp)^{-1}\alpha'_\perp$ ($\alpha_\perp$ and $\beta_\perp$ are the orthogonal matrices of α and β, respectively), $\beta' C(1) = 0$, and $C(1)\alpha = 0$, and which is partitioned as:

$$\begin{pmatrix} \Delta y_t \\ \Delta z_t \end{pmatrix} = \begin{pmatrix} C_{yy}(1) \; C_{yz}(1) \\ C_{zy}(1) \; C_{zz}(1) \end{pmatrix} \left[\begin{pmatrix} \epsilon_{yt} \\ \epsilon_{zt} \end{pmatrix} - \begin{pmatrix} \Pi_{yy} \; \Pi_{yz} \\ \Pi_{zy} \; \Pi_{zz} \end{pmatrix} \begin{pmatrix} \gamma_y \\ \gamma_z \end{pmatrix} \right] \quad \text{(B1-3)}$$
$$+ \begin{pmatrix} C^*_{yy}(L) \; C^*_{yz}(L) \\ C^*_{zy}(L) \; C^*_{zz}(L) \end{pmatrix}$$
$$\Delta \left(\begin{pmatrix} \epsilon_{yt} - \alpha_y \eta \\ \epsilon_{zt} - \alpha_z \eta \end{pmatrix} - \begin{pmatrix} \Pi_{yy} \; \Pi_{yz} \\ \Pi_{zy} \; \Pi_{zz} \end{pmatrix} \begin{pmatrix} \gamma_y \\ \gamma_z \end{pmatrix} \right).$$

The solution of (B1-2) has a representation $x_t = K_{\beta\perp} x_0 + C(1) \sum_{i=1}^{t} (\epsilon_t - \Pi\gamma) + C^*(L)\epsilon_t - C^*(1)(\Pi\gamma + \alpha\eta)$, where $K_{\beta_\perp} = \beta_\perp(\beta'_\perp\beta_\perp)^{-1}\beta'_\perp$. Multiplying this equation by β' yields the corresponding moving average representation of the cointegrating vectors under the stationarity condition $\beta' C(1) = 0$:

$$\beta' x_t = \beta' C^*(L)\epsilon_t - \beta' C^*(1)(\Pi\gamma + \alpha\eta). \quad \text{(B1-4)}$$

From (B1-3) and (B1-4), the h-period impulse responses of Δy_t to transitory changes in ϵ_{zt} can be derived as follows, as $h \to \infty$:

$$\left(\frac{\partial \Delta y_{t+h}}{\partial \epsilon'_{zt}} \right)_{d\epsilon_{yt} = \delta_1 d\epsilon_{zt}} = \left(C^*_{yy,h} - C^*_{yy,h-1} \right) \delta_1 + \left(C^*_{yz,h} - C^*_{yz,h-1} \right) \to 0, \quad \text{(B1-5)}$$

$$\left(\frac{\partial \beta' x_{t+h}}{\partial \epsilon'_{zt}} \right)_{d\epsilon_{yt} = \delta_1 d\epsilon_{zt}} = \beta'_y \left(C^*_{yy,h} + C^*_{yz,h} \right) \delta_1 + \beta'_z \left(C^*_{zy,h} + C^*_{zz,h} \right) \to 0. \quad \text{(B1-6)}$$

and

$$\left(\frac{\partial \Delta y_t}{\partial \epsilon_{zt}'}\right)_{d\epsilon_{yt}=\delta_1 d\epsilon_{zt}} = \left(C_{yy}(1) \ + C^*_{yy,0}\right)\delta_1 + \left(C_{yz}(1) \ + C^*_{yz,0}\right) = \delta, \tag{B1-7}$$

where det $C^*(L) = 0$ has all its roots outside the complex unit circle. Equations (B1-5) and (B1-6) show that the responses of Δy_{t+h} and $\beta' X_{t+h}$ to changes in ϵ_{zt} die out rather than persist indefinitely in the long run. On the other hand, (B1-7) shows a convergent sum of the responses of Δy_t to the changes of ϵ_{zt}. Imposing $\delta_1 = 0$ leads to conventional impulse response functions.

In the case of permanent changes, there are two possible channels within which the changes can be evaluated; one is the changes in the equilibrium means η resulting from shifts in the attitude of policy makers to acceptable levels of disequilibria $\beta' x_t$ and the other is the shifts in policy variables γ_z. The partial responses of Δy_{t+h} and $\beta' x_{t+h}$ to changes in η and γ_z can be derived from (B1-3) and (B1-4), as $h \to \infty$:

$$\begin{aligned}\left(\frac{\partial \Delta y_{t+h}}{\partial \eta'}\right) &= -\left(C^*_{yy,h}\ \alpha_y + C^*_{yz,h}\alpha_z\right) \\ &= -C^*_{y,h}\alpha \ \to 0,\end{aligned} \tag{B1-8}$$

$$\left(\frac{\partial \beta' x_{t+h}}{\partial \eta'}\right) = -\beta' C^*(1)\alpha \ \to 0, \tag{B1-9}$$

$$\begin{aligned}\left(\frac{\partial \Delta y_{t+h}}{\partial \gamma_z'}\right) &= -\left((C_{yy}(1) \ + C^*_{yy,h})\Pi_{yz} + (C_{yz}(1) \ + C^*_{yz,h})\Pi_{zz}\right) \\ &= -(C_{yy}(1)\Pi_{yz} + C_{yz}(1)\Pi_{zz}).\end{aligned} \tag{B1-10}$$

and

$$\left(\frac{\partial \beta' x_{t+h}}{\partial \gamma'_z}\right) = -\left((\beta'_y C^*_{yy}(1) \ + \beta'_z C^*_{zy}(1))\Pi_{yz} + (\beta'_y C^*_{yz}(1) + \beta'_z C^*_{zz}(1))\Pi_{zz}\right) \to 0. \tag{B1-11}$$

where det $C^*(L) = 0$ has all its roots outside the complex unit circle. The responses show that while changes in the equilibrium mean η do not affect Δy_t in the long-run, changes in the mean growth rate of z_t have a long run effect on Δy_t.

Table 6.1 Cointegration analysis

(a) Cointegration tests					
Eigenvalues	0.37	0.18	0.14	0.12	
Null hypothesis	$\gamma = 0$	$\gamma \leq 1$	$\gamma \leq 2$	$\gamma \leq 3$	
Max	38.12**	16.86	12.65	10.18	
Trace	77.82**	39.69	22.84	10.17	
(b) Tests for weak exogeneity					
Variables	y_t	m_t	p_t	R_t	$\{m_t, R_t\}$
$\chi^2(\cdot)$	24.87**	0.89	11.57**	0.01	0.96

Notes: 1. The Max and Trace statistics were adjusted for degrees of freedom, following Reimers (1992), and the critical values were taken from Osterwald-Lenum (1992). 2. ** denotes significance at the 5% level. 3. The test statistics for weak exogeneity are asymptotically distributed as $\chi^2(1)$ for y_t, m_t, p_t, and R_t, respectively, and $\chi^2(2)$ for the joint test of $\{m_t, R_t\}$.

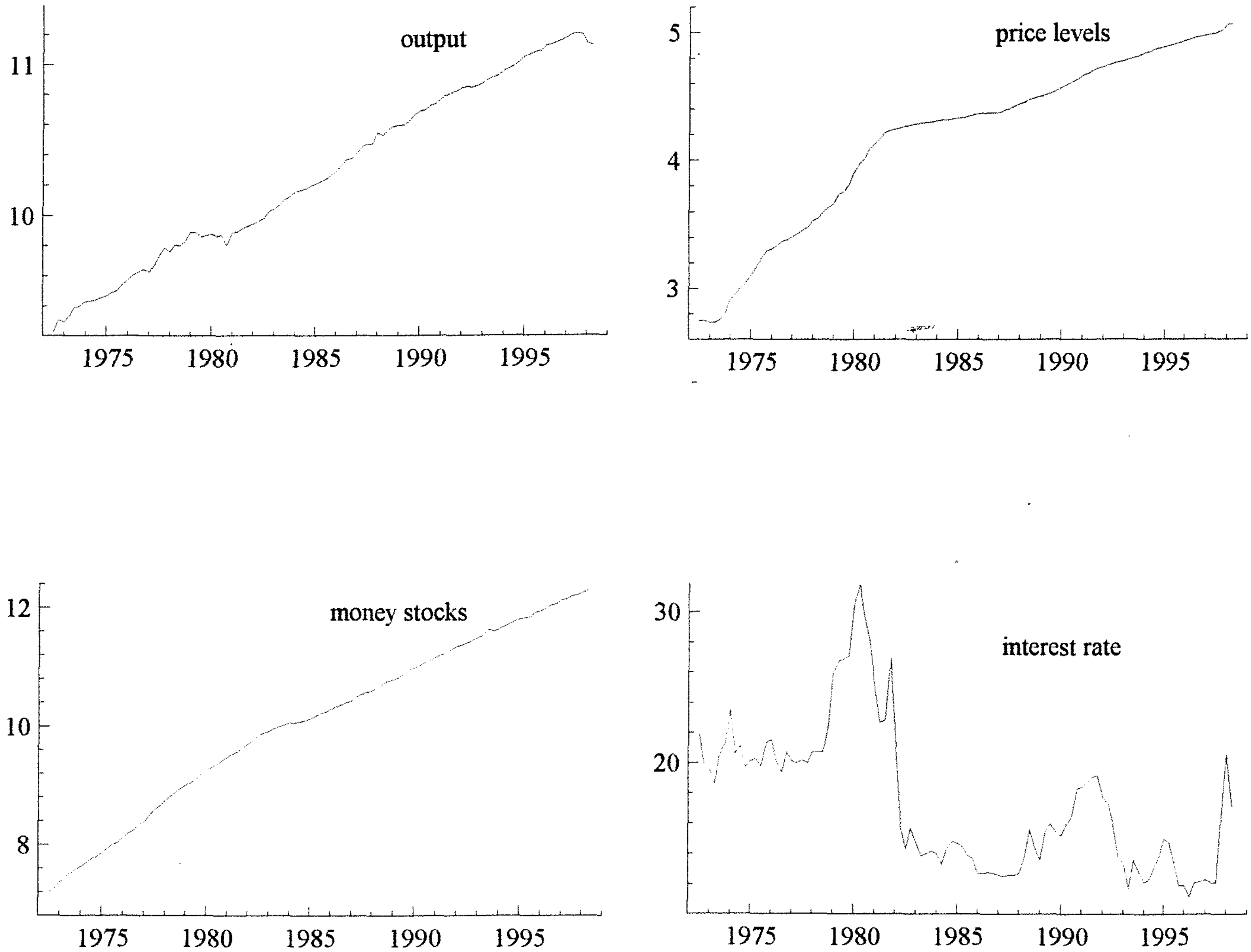
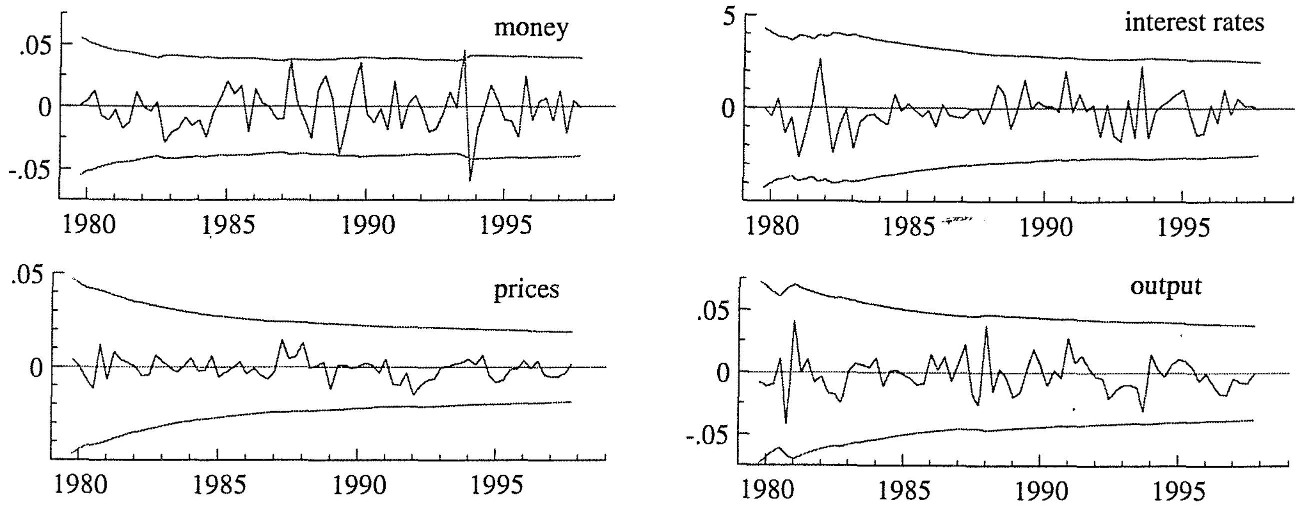
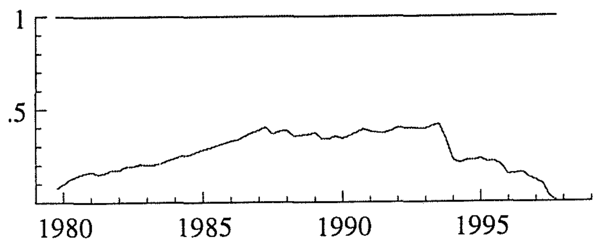


Figure 6.1 Data in levels.



(a) one-step residuals with +/-2*standard errors.



(b) break-point Chow statistics with their 5% critical values.

Figure 6-2. Statistics for parameter constancy.

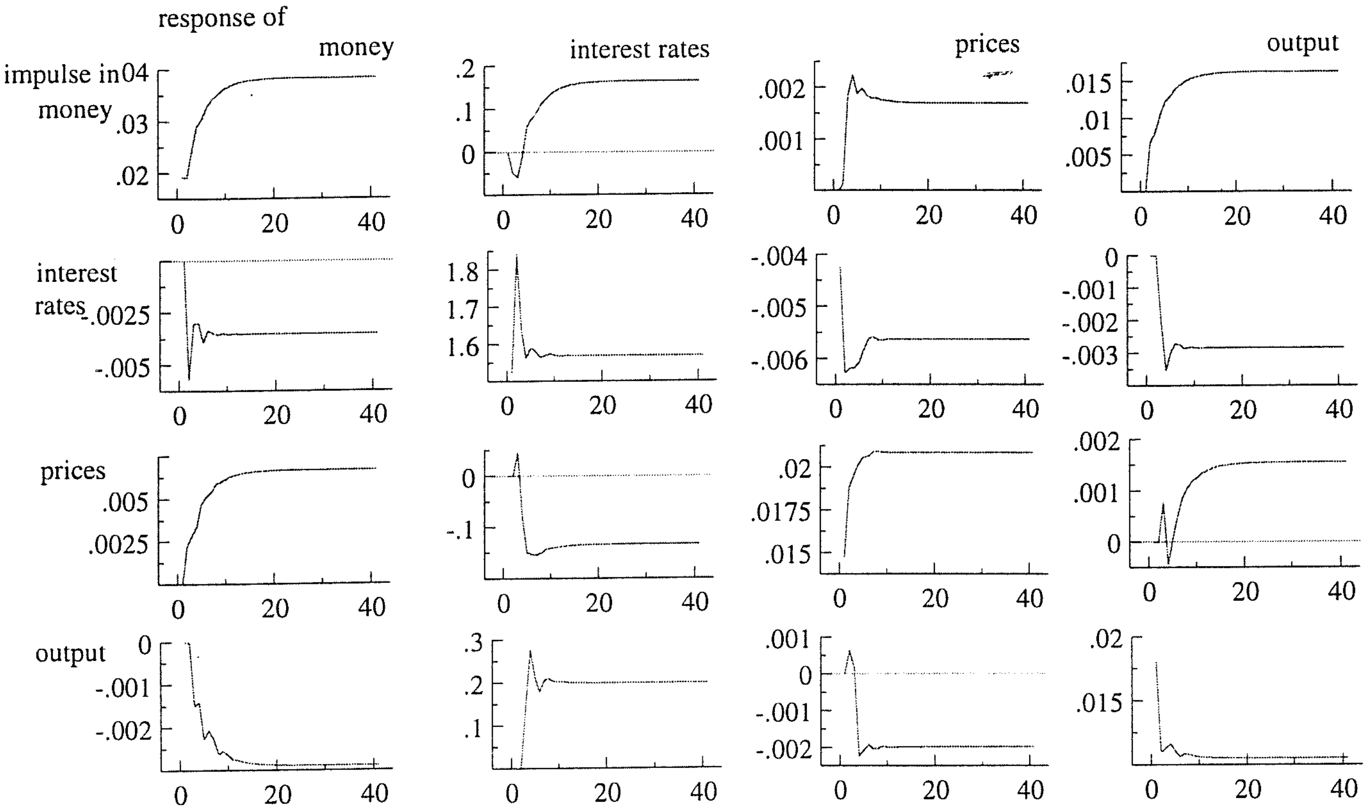
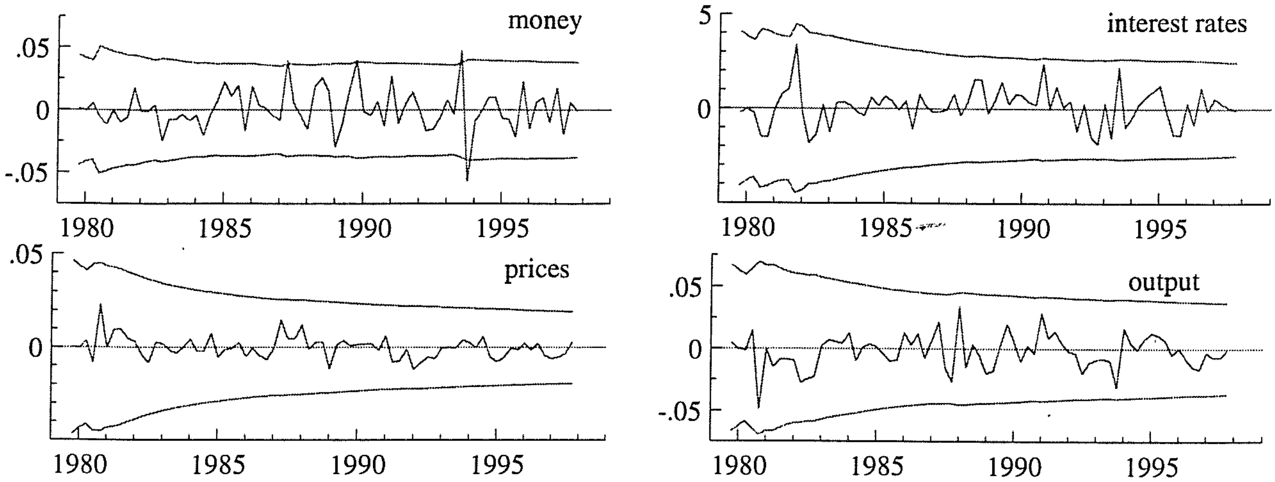
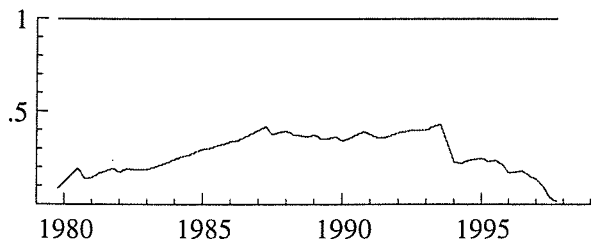


Figure 6.3 Impulse responses



(a) one-step residuals with +/-2*standard errors.



(b) break-point Chow statistics with their 5% critical values.

Figure 6.4 Statistics for parameter constancy under co-breaking.

Chapter 7

Conclusions

This thesis has evaluated the effects of policy changes on economic activities in Korea with emphasis on the aspects of econometric modelling. As the basic methodology, the research has applied the dynamic modelling approach which combines current two extreme approaches - the theory-based structural approach and the data-based VAR approach - and focused on formulating the statistically well-defined, underlying models through extensive specification tests. Particular attention was paid to how to avoid the critiques of Lucas (1976) and Sims (1980), both of which are damaging criticisms to the current macroeconometric approaches. Unlike the previous approaches, these issues were formally tested in this research rather than assumed. Thus a distinguished feature of the research is that no assumption has been made *a priori* and so the empirical results are more legitimate and reliable than those from structural and VAR approaches. Econometrically and empirically, the thesis contributes to the existing literature. Econometrically, the thesis has attempted to discuss and extend the current LSE methodology in a more practical way. First, the thesis has demonstrated that care is required to formulate marginal models for the test of super exogeneity to identifying observational equivalence in ECM type models, if the effects of regime changes on the constancy of marginal models do not exhibit substantial changes, because of the possibility of 'spurious' nonconstancy in marginal models. In this special case, this thesis recommends to complementarily perform both encompassing and super exogeneity tests to supplement the weak power of the latter test. Second, the thesis has shown how the concept of weak exogeneity can be used to identify policy variables within a VECM. This concept is very useful to find exogenous variables which is ordered firstly in orthogonalized impulse response analysis. Our approach rules out *a priori* assumption being commonly used in the current VAR literature to identify exogenous policy variables. Third, in the last chapter the thesis has questioned why currently available macroeconometric models separately focus on evaluating the effects of transitory and permanent policy changes in a divergent way and argued that both changes should be jointly evaluated within a model. Possible econometric advantages with this approach have been additionally discussed in the context of the invariant property of the underlying model and the usefulness of gaining data information through both parameter spaces and error terms. Methodologically, the thesis extends the work of Hendry and Mizon (1998), who provides a reduced-form framework to assess the effects of policy changes and applies it to a bivariate model, into a structural multivariate framework by formally incorporating the usefulness of weak exogeneity to identifying policy actions.

Empirically, the thesis has assessed the effects of policy changes. To investigate this issue, the study has sequentially asked three questions. The questions were whether the demand for money is stable under regime changes, whether policy variables measuring policy stance are exogenously determined, and how and whether transitory and permanent changes in policy affect the economy. Our study differs from the earlier studies which *a priori* assume the first two necessary conditions for the last question. The main findings to these questions are that: (1) the stability of the demand function for money has not been broken down by financial deregulation. This finding is contrasted with a previous work which finds a considerable degree of parameter instability during the period of financial liberalization; (2) the stock of M2 money which has been used as the main policy variable is exogenously determined, implying that the BOK has implemented its policy discretionarily rather than systematically responded to the movements of current macroeconomic developments. Particularly, the exogeneity of disaggregated inside money is not consistent with the assumption of endogenous money of King and Plosser (1984); (3) both transitory and permanent changes in policy substantially affect the economy. An unanticipated transitory shock on nominal money significantly affects output and prices, but in the case of an anticipated permanent change, 'real' money rather than 'nominal' money positively affects output. This implies the importance of price stability as a precondition for the sustained economic growth of the economy. Another important finding is that financial liberalization as a proxy for permanent changes in policy has clearly contributed to stabilizing price levels and to sustaining economic growth. These empirical results support the view of the proponents of financial liberalization in the development literature and sheds some light to the development strategy of developing countries. The overall results are contrasted with the view of the new classicals that only unanticipated policy shocks matter, and with the theory of real business cycles that policy changes, whether short-or long-run and anticipated or unanticipated, do not affect the real economy but the reverse causation from output to money holds.

The empirical findings in the above suggest two major implications for the conduct of monetary policy in Korea. First, the empirical results show that both transitory and permanent changes in policy significantly affect the economy, but price stability is a required precondition to continue the sustained economic growth of the economy. The importance of this finding is that as long as price stability is maintained, an expansion of money, whether anticipated or not, may contribute to enhancing economic growth. However, in the case that money is nonneutral and so a money increase is dissipated partly in inflation and partly in output, there may exist a Phillips curve type inverse relationship between inflation and economic growth. That is, reducing inflation may require sacrificing a certain portion of economic growth (see Okun (1978) and Gordon and King (1982)). In this context, policy should be conducted in such a way that the optimal trade-off between inflation and economic growth is achieved. One way to reduce 'the sacrifice ratio', which measures the output loss required to eliminate permanently one point of inflation, would be that policymakers credibly commit their policy to reduce inflation and so let people believe their commitment. Under this credible policy, the costs of reducing inflation may be much lower than those under an alternative policy because of agents' lowered expectations on inflation. Second, even though financial liberalization in Korea has been much limited in its scope and degree, it has largely contributed to enhancing economic growth over the sample period. While this thesis is being written, lots of deregulation measures have been undertaken, including the deregulation of interest rates and the adoption of a floating exchange rate system. However, large parts in all sectors are still regulated. This necessitates undertaking further deregulation measures. The recent financial crisis in 1997 has blamed financial liberalization as one of the major causes of the crisis. However, as shown in this research, financial liberalization has achieved its goals, such as price stability and economic growth, in terms of macroeconomic policy. Thus the crisis has nothing to do with mis-implementation of monetary policy. Instead, the causes of the crisis could be found from the failure of structural policy, such as the failure of the supervision of bank intermediaries, ill-handling the bankruptcy of large domestic companies, and the uncertainty created by the government's indecisive policy attitude. The main suggestion of this study is not a 'complete' deregulation all at once. The experiences with financial reforms and liberalization in the Latin American in 1980s and the Asian countries in 1990s suggest that capital account liberalization may bring a disastrous result to developing countries, mainly because of market failures in the banking sector (Fry, 1995). However, the main theme of our argument is that government regulation should be reduced to its 'minimum' levels which are required for the optimal mix of markets and government intervention.

Finally, the research reported in this thesis has many limitations in applying the econometric techniques we used in this study and the limitations can be widened in a more realistic way. Future research in the following directions may refine the methodology and enhance the credibility and pervasiveness of the conclusions obtained in this chapter. First, in this thesis we have applied traditional Chow (1960) tests to detect regime changes. However, a well-known deficiency of this type of tests is the assumption that the dates of regime changes are known. In an empirical practice, however, it is difficult to find the precise dates of structural breaks, so the inferences based on the Chow tests might lead to invalid conclusions. Recently, the literature provides several data-based testing procedures for the null hypothesis of constant coefficients when the break points of regime changes are not known (see, for examples, Andrews (1993) and Hansen (2000)). Particularly, Hansen (2000) provides a realistic test procedure which accommodates structural changes in conditioning variables. This is a modified version of Andrews' (1993) test, which assumes the stationarity of conditioning variables. By applying the test to identify structural breaks, we may expect more reliable empirical results. Second, concerning with the basic form of underlying models, this thesis has assumed linear relationships between variables. This would be a narrow approach, since most of major macroeconomic variables have non-linear relationships. If this is the case, non-linear models are appropriate to characterize actual data generating processes, even though there are some costs to pay for the complexity of finding and estimating the underlying non-linear functions. Recently, a variety of non-linear time-series models have been proposed (see, for extensive surveys, Granger and Teräsvirta (1993) and Tong (1990)). In relation to policy analysis, Weise (1999) and Choi (1999) apply non-linear models, such as smooth transition and threshold VARs. However, none of these studies have paid to the evaluation of the models considered. In this context, the theme of the LSE methodology, which emphasizes statistically valid model formulation, may be directly applicable to this line of research. In a recent paper, Jansen and Teräsvirta (1996) show how to test super exogeneity in a smooth transition autoregressive model and Teräsvirta (1998) applies it to UK data for house prices. Finally, the thesis has identified all the data considered as I(1). However, the seasonally unadjusted data for nominal money and price levels in Chapter 4 appear to be I(2), as in the findings of Johansen (1992) for the UK data. It is well-known that failing to account for the proper order of non-stationarity in data causes serious econometric problems, since statistical inferences for I(2) variables and the corresponding economic interpretation are completely different from those in the case of I(1) variables. Recently it has become increasingly important in macroeconomics to deal with I(2) non-stationary data (see Johansen (1995) and Haldrup (1998)). A good example can be found in Juselius (1998) who stresses the aggregate demand for money and the role of I(2) components and structural changes. This line of research may adequately represent the fundamental relationships of macroeconomic variables.

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