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Essays on Identification of Monetary Policy Shocks in Vector Autoregressive Models: Alternative Identification Schemes and Lag Structures.

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ESSAYS ON IDENTIFICATION OF MONETARY POLICY SHOCKS IN VECTOR AUTOREGRESSIVE MODELS: ALTERNATIVE IDENTIFICATION SCHEMES AND LAG STRUCTURES

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To my parents

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ABSTRACT

This dissertation is primary concerned with the sensitivity of the effects of monetary policy shocks across alternative identification schemes and lag structures. The four widely-cited identification schemes of Christiano, Eichenbaum, and Evans (1994; 1996), Strongin (1995), Bernanke and Mihov (1998), and the long-run restrictions approach pioneered by Blanchard and Quah (1989) are used. Also, three types of lag structures – symmetric, Keating-type, and Hsiao-type asymmetric lag structures – are employed.

The first essay focuses upon a closed economy framework. The results indicate that impulse response functions for macro variables are often sensitive to identification schemes and lag structures. For a given lag structure, the Strongin, Bernanke and Mihov, and long-run restrictions schemes generate similar results, while the Christiano, Eichenbaum, and Evans scheme often yields different responses from others. This essay also illustrates that the Christiano, Eichenbaum, and Evans and long-run restrictions schemes are relatively insensitive to the type of lag structures.

The second essay examines the effects of lag structure misspecification within a Monte Carlo framework. It is shown that the lag structure of a VAR model does matter when assessing the effects of monetary policy shocks. For most horizons, t-statistics support for the hypothesis that the responses from the misspecified VARs are significantly different from the assumed 'true' responses.

The dissertation is completed by the third essay in which the model is extended to an open economy framework. In general, the contemporaneous restriction schemes give reasonable results, but the magnitude and timing of the effects differ across

identification schemes. By contrast, the long-run restrictions approach is found to be not suitable for a relatively large system like our open economy framework. Also, in this essay, the responses for the open economy are contrasted with those for the closed economy. The results indicate that the quantitative effects are different, despite the similarity in the general patterns of the responses. In particular, all identification schemes considered in this essay showed either some degree of the 'price puzzle' or weaker price effects than in a closed economy framework, even in the presence of commodity prices and the exchange rate.

CHAPTER 1

INTRODUCTION

Vector autoregressive (VAR) models are widely used in the empirical analysis of the monetary policy transmission mechanism. A central feature of VAR analysis is the identification of monetary policy shocks or unanticipated shifts in monetary policy.[1] Certainly, to ensure the VAR analysis yields meaningful information on the effects of monetary policy, exogenous shocks to monetary policy must be separated from policy makers' systematic responses to nonmonetary developments in the economy; hence, fundamental identification problems must be solved. The huge literature on monetary VAR analysis explores three general strategies for identifying the monetary policy shocks in VAR models.[2]

The first strategy imposes a recursive causal structure (also called a Wold causal structure) on the contemporaneous relations among model variables to identify monetary policy shocks. In this approach, it is assumed that economic variables are determined in a block recursive way. Hence, one-way causation from variables higher in the ordering is assumed; all contemporaneous correlation between two variables is attribute to the variable higher in the order, while there is no contemporaneous feedback

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[1] Christiano, Eichenbaum, and Evans (1998) offer three interpretations of monetary policy shocks: (1) exogenous shocks to the preferences of the monetary authority, (2) shocks to private agents' expectations about the Federal Reserve policy, and (3) various technical factors like the measurement error in the preliminary data available to the Federal Open Market Committee (FOMC) at the time it makes decisions.

[2] We note that there is another strategy, the 'narrative approach', in which identifying monetary policy shocks does not involve explicitly modeling the monetary authority's feed back rule in a VAR model. For example, following Freidman and Schwartz (1963), Romer and Romer (1989) identify several episodes of big shifts in monetary policy based on their reading of the minutes of the FOMC. See also Boschen and Mills (1992). Refer to Leeper (1997), Hoover and Perez (1994), and Christiano, Eichenbaum, and Evans (1994) for discussion and critiques of the 'narrative approach'.

from variables lower in the ordering. Consequently, monetary policy shocks are estimated by decomposing variance-covariance matrices of the ordinary least squares residuals in VAR models in a triangular fashion (Choleski decomposition). The identification schemes of Sims (1980), Bernanke and Blinder (1992), Christiano, Eichenbaum, and Evans (1994;1996), and Strongin (1995), among others, are good examples of this approach.

The second strategy is to build structural VARs. Some authors like Sims (1986), Bernanke (1986), Gordon and Leeper (1994), and Bernanke and Mihov (1998) at least partially abandon the recursive assumptions. In this type of approach, an explicit structural model that relies on theoretical models is used to specify simultaneous interactions among variables in a system, although recursive structures are sometimes chosen for some variables in the system. For example, Bernanke and Mihov (1998a) develop a semi-structural VAR model which blends the Choleski decomposition with a structural model of the reserve market. This scheme imposes no restrictions on the relations among macro variables, but identifies monetary policy shocks by employing a simple structural model of the bank reserves market in which simultaneity among the structural shocks to the reserve market variables is allowed.

The last strategy identifies monetary policy shocks by assuming that they do not affect real variables in the long-run. This approach was pioneered by Blanchard and Quah (1989) and Shapiro and Watson (1988).[3] In this approach, no restrictions are placed on the contemporaneous relations among the variables, but identification is

[3] This approach has been used recently by Fackler and McMillin (1998) to identify monetary policy shocks and Lastrapes (1998a) to identify money supply shocks.

achieved by imposing long-run restrictions on the relations among the variables in the model.[4]

Although various economic and institutional arguments can be used to rationalize each identification scheme, there is little agreement on the preferred approach. In fact, the weakness of these approaches have been widely discussed in the literature. For example, Enders (1995) and Bernanke and Mihov (1998b) criticized the VARs with recursive assumptions in that the selection of ordering is generally ad hoc. By contrast, Christiano, Eichenbaum, and Evans (1998) claimed that, to identify monetary policy shocks in a structural VAR model, a broad set of economic relationships must be identified and the assumptions involved are also controversial. The limitations of the long-run restrictions approach are often discussed. Faust and Leeper (1997) argued that the estimates of the impulse response function might be distorted since this approach imposes infinite horizon restrictions in a VAR estimated with data from a finite sample.

Besides the identification scheme, another critical element in VAR analysis is determination of the lag structure of the VAR model. In fact, Braun and Mittnik (1993) showed that misspecification of lag length generates inconsistent coefficient estimates and hence results in distortions in impulse responses and variance decompositions. More recently, Lee (1996) also pointed out that underparamterization (lower order lag length than true lag length) results in estimation bias, while overparameterization (higher order lag length than true lag length) results in a loss of degree of freedom and

[4] Bernanke and Mihov (1998b) combined the semi-structural VAR model of Bernanke and Mihov (1998a) with the long-run restrictions approach. Also, Lastrapes and Selgin (1995) attempted to use combinations of short-run and long-run restrictions. See Lastrapes (1998b) in which Bayesian techniques are used to combine short-run and long-run restrictions.

estimation efficiency. Since the impulse response functions are functions of estimated reduced-form coefficients, both underparameterization and overparameterization may lead to less precise policy analysis. Thus, the determination of lag structure is a very important issue in assessing the effects of monetary policy shocks in VAR models.

In most VAR models, including the above-mentioned models, one maintained assumption is that the lag structure is symmetric in the sense that the same lag length is assumed for all variables in all equations of the model. Hsiao (1982), however, first examined the possibility of an asymmetric lag structure in a VAR model. He suggested a VAR model in which the lag length on each variable in each equation could differ. More recently, Keating (forthcoming) also suggested an asymmetric lag VAR model. In this Keating-type asymmetric lag VAR, the lag length potentially differs across the variables in the model, but is the same for a particular variable in each equation of the model. There is, however, no theoretical reason to believe that either a symmetric lag structure or an asymmetric lag structure is more appropriate in most VAR models. Indeed, Keating (forthcoming) showed that an asymmetric lag structure in a VAR is theoretically possible if a structural model is characterized by asymmetric lags. However, unfortunately, very seldom does theory provide any guidance as to the appropriate type of lag structure.

Given the uncertainty about the identification schemes and lag structures described above, the purpose of this dissertation is to investigate the sensitivity of impulse response functions of macroeconomic variables such as output, the price level, and the interest rate to monetary policy shocks associated with alternative identification schemes and lag structures in VAR models. More specifically, we try to answer the following three groups of questions that motivated this dissertation:

(1) How similar are estimates across symmetric and asymmetric lag structures for a given identification scheme? Is one identification scheme more sensitive to the type of lag structure than others? And, how similar are estimates for different identification schemes for a given lag structure?

(2) Are the impulse responses from a VAR model with a misspecified lag structure significantly different from those from the prespecified 'true' model in a Monte Carlo simulation framework?

(3) Do the identification schemes also generate reasonable impulse responses when the identification schemes are extended to an open economy framework? How do the results from an open economy framework compare to the results from a closed economy framework?

The study considers, in turn, each group of these questions in each of the subsequent chapters. Chapter 2 investigates, in a closed economy framework, the sensitivity of the effects of monetary policy shocks for alternative identification schemes on macroeconomic variables such as output, price, and interest rates across alternative lag structures. To answer the first group of questions presented above, we estimate and compare the impulse responses from the alternative identification schemes across alternative lag structures in a common VAR model over a particular sample period. Holding constant the variables in a VAR model and the sample period allows us to clearly observe the effects of identification schemes and lag structures. In this chapter, we employ the four widely-cited identification schemes of Christiano, Eichenbaum, and Evans (1994; 1996), Strongin (1995), Bernanke and Mihov (1998), and the long-run restrictions approach pioneered by Blanchard and Quah (1989). Also, three different lag structures–symmetric, Keating-type, and Hsiao-type asymmetric lag structures–are considered.

In Chapter 3, we investigate the distortions in the impulse responses due to lag structure misspecification in a VAR model. In a Monte Carlo experiment framework, we examine the results from two cases of misspecification. In the first case, the consequences of fitting Keating-type asymmetric lag VAR to the series generated by assuming a symmetric lag structure as the 'true' lag structure are examined. In the second case, the consequences of applying symmetric lag VARs to the series generated by using prespecified a Keating-type asymmetric lag structure are examined. The identification schemes and lag structures considered in this chapter are the same as the previous chapter except we do not considered the long-run restrictions approach and the Hsiao-type asymmetric lag structure. We note that, as will be explained later, the implementation of the long-run restrictions approach and the Hsiao-type lag structure are difficult in this Monte Carlo framework.

In Chapter 4, the model is extended to an open economy framework. The sensitivity of the effects of monetary policy shocks on the exchange rate and the trade balance as well as on output, the price level, and the interest rate across above alternative identification schemes is examined. The chapter also contrasts the effects of monetary policy shocks from the closed economy framework and an open economy framework. In addition, we investigate the effects of shocks to the exchange rate on macro variables including the trade balance. However, in this chapter, we do not consider asymmetric lag VARs since the Keating-type lag search process is almost impossible for the 11 variable monthly VAR model considered here. We also do not employ the Hsiao-type lag structure. Finally, Chapter 5 summarizes and concludes this dissertation.

CHAPTER 2

THE EFFECTS OF MONETARY POLICY SHOCKS AND LAG STRUCTURES: COMPARING SYMMETRIC AND ASYMMETRIC LAG STRUCTURES

2.1. Introduction

In the past two decades there has been substantial progress in assessing the effects of monetary policy shocks using statistical methods, especially vector autoregressive (VAR) models. The VAR approach certainly enables us to understand more about the effects of monetary policy shocks than we did twenty years ago. However, from a methodological point of view, we have not reached a consensus and still need to search for an appropriate way to identify monetary policy shocks.

A huge recent VAR literature has focused on identification assumptions, i.e. the determination of exogenous shocks to monetary policy following the tradition of Sims (1980). For example, Blanchard and Quah (1989), Bernanke and Blinder (1992), Gordon and Leeper (1994), Christiano, Eichenbaum, and Evans (1994; 1996), Strongin (1995), and Bernanke and Mihov (1998), among others, suggested their own identification schemes that can be rationalized by various economic and institutional arguments.

In most VAR models, including the above-mentioned models, one common assumption is that the lag structure is symmetric in the sense that the same lag length is assumed for all variables in all equations of the model. However, there is no theoretical reason for the lag length to be the same. This issue was first examined by Hsiao (1981). He suggested a VAR model in which the lag length on each variable in each equation could differ. In the VAR, Hsiao used a sequential procedure based on the concept of Granger-Causality and Akaike's final prediction error (FPE) criterion to choose appropriate lags for each variable in each equation. Recently, Keating (1995) re-examined the issue of an asymmetric lag VAR. He constructed a VAR model in which the lag length potentially differs across the variables in the model, but is the same for a particular variable in each equation of the model. Keating found that, using a small structural VAR model, an asymmetric lag VAR (AVAR) generates relatively fewer insignificant reduced-form parameters than traditional symmetric VAR models do. Based upon finding fewer insignificant parameters, Keating argued that an asymmetric VAR may more precisely estimate the effects of monetary policy shocks on macroeconomic variables since the impulse responses and variance decompositions are functions of estimated reduced-form coefficients.

Given uncertainty about the identification schemes and lag structures, the goal of this paper is to examine and compare the effects of monetary policy shocks for alternative identification schemes on macroeconomic variables such as output, price, and interest rates across alternative lag structures. The approach in this paper is similar in spirit to McMillin (1998) who compares the effects of shocks to monetary policy using contemporaneous and long-run restrictions approaches to identify policy shocks within a common model. It is, however, different from McMillin in that the current study extends the comparison of effects of monetary policy shocks across different lag structures. The identification schemes considered in this paper are the same as in McMillin (1998) who focused on the approaches suggested by Christiano, Eichenbaum and Evans (1994; 1996), Strongin (1995), Bernanke and Mihov (1998), and Blanchard and Quah (1989). Among these four identification schemes, the first three schemes impose restrictions on the contemporaneous relations among the variables, while the last scheme imposes long-run neutrality restrictions. Three different lag structures –symmetric, Keating-type, and Hsiao-type asymmetric lag structures–are considered. The effects of monetary shocks for the alternative identification schemes across lag structures are evaluated by estimating impulse responses for each scheme, using quarterly data.

The rest of this paper proceeds as follows. Section 2 briefly discusses the alternative identification schemes and lag structures and describes estimation methods and data. Section 3 provides the results for symmetric and asymmetric lag VARs that compare impulse response functions for the each identification scheme as in McMillin (1998). Section 4 gives a summary and conclusion.

2.2. Model Specification, Data, and Estimation

2.2.1. Identification Schemes

The first identification scheme considered in this paper is that of Christiano, Eichenbaum, and Evans (1994; 1996). For this scheme as well as the Strongin scheme, the importance of ordering is worth noting since these two schemes rely solely on the Choleski decomposition in which all contemporaneous correlation between two variables is attributed to the variable higher in the order. Consequently, it reflects basic assumptions about the contemporaneous causal relationships among a policy variable and other macroeconomic variables. The model employs six variables which are listed in the order used in the Choleski decomposition: output, the price level, commodity prices, nonborrowed reserves, total reserves, and the federal funds rate. Nonborrowed reserves, which are the variable most directly controlled by the Federal Reserve, are taken as the policy variable. This scheme, as the ordering implies, assumes that monetary policy affects output, the price level, and commodity prices only with a lag, while the Federal Reserve has full current information on the three variables. We note that above assumptions are more difficult to defend if one deals with high frequency data. The scheme also assumes that monetary policy has a contemporaneous effect on total reserves and the federal funds rate, although the Federal Reserve responds to movements in these variables only with a lag. The assumptions of the Christiano, Eichenbaum, and Evans scheme on the relationships among the policy variable and output, the price level, and commodity level can be also applied to the two schemes of Strongin (1995) and Bernanke-Mihov (1998). However, as we will see, the assumption about the relationship between the policy variable and total reserves is different from those schemes.

The second identification scheme considered in this paper is that of Strongin (1995) in which the policy variable is also nonborrowed reserves. Although Strongin constructed two sets of VARs with three variables and five variables, this paper employs the same six variables as in Christiano, Eichenbaum, and Evans (1994). However, the essential point of the Strongin scheme that shocks to total reserves reflect reserve demand shocks will be maintained. In this view, nonborrowed reserve shocks are viewed as a mixture of reserve demand shocks and policy shocks. When the Federal Reserve targets the federal funds rate, as it did over most of sample period used here, a reserve demand shock would tend to raise the federal funds rate unless the Federal Reserve expanded nonborrowed reserves. Thus, orthogonalized policy shocks can be extracted by placing total reserves prior to nonborrowed reserves in ordering. Consequently, the model has following the Wold causal ordering: output, the price level, commodity prices, total reserves, nonborrowed reserves, and the federal funds rate. Note that the causal link between nonborrowed reserves and total reserves is reversed compared to the Christiano, Eichenbaum, and Evans scheme.

The third identification scheme considered in this paper is Bernanke and Mihov's (1998) semi-structural VAR which comprises the same six variables as in Christiano, Eichenbaum, and Evans or Strongin. This scheme extracts monetary policy shocks from a model of the reserve market estimated from VAR residuals for nonborrowed reserves, total reserves, and the federal funds rate that are orthogonalized with respect to the other model variables. Bernanke and Mihov assumed the following structural model for bank reserves:

$$\mu_{tr} = -\alpha\mu_{ffr} + v^d \tag{2.1}$$

$$\mu_{br} = \beta(\mu_{ffr} - \mu_{disc}) + v^b \tag{2.2}$$

$$\mu_{nbr} = \phi^d v^d + \phi^b v^b + v^s \tag{2.3}$$

where the μ's represent the VAR residuals that are orthogonalized with respect to output, the price level, and commodity prices, and the v's are structural shocks. Subscripts tr, ffr, br, disc, and nbr represent total reserves, the federal funds rate, borrowed reserves, the discount rate, and nonborrowed reserves, respectively. Thus equation (2.1) describes the total reserve demand that depends negatively upon the federal funds rate, while equation (2.2) describes borrowed reserve demand that depends positively on the federal funds rate and negatively on the discount rate. Equation (2.3) represents the Federal Reserve's reaction function; hence v^s can be interpreted as the shock to monetary policy that we are interested in identifying. Equation (2.3) implies that the Federal Reserve has current information on the shocks to both total reserves and borrowed reserves. In this paper, we slightly modify above structural model, based upon Bernanke and Mihov's results and suggestions.

(2.1)′ $$\mu_{tr} = v^d$$

(2.2)′ $$\mu_{br} = \beta\mu_{ffr} + v^b$$

(2.3)′ $$\mu_{nbr} = \phi^d v^d + \phi^b v^b + v^s$$

Equation (2.1)′ imposes the restriction that α=0 on equation (2.1); the innovation in total reserves is assumed to reflect a demand shock, as in Strongin. This restriction is imposed because Bernanke and Mihov pointed out that a just-identified model with α=0 performs well. In equation (2.2)′, the discount rate shocks are set to zero in order to compare the Christiano, Eichenbaum, and Evans, and Strongin schemes that do not explicitly consider the discount rate.[1]

The long-run restrictions approach is the last identification scheme considered in this paper. This scheme, first introduced by Blanchard and Quah (1989) and Shapiro and Watson (1988), does not impose restrictions on contemporaneous relationship among the model variables as is done in the Christiano-Eichenbaum-Evans, Strongin, and Bernanke and Mihov schemes. In this paper, three assumptions are made to identify monetary policy shocks as in McMillin (1998).

(1) Shocks to monetary policy have no long-run effects on output.

[1] The structural model of reserve market variables is estimated by using a two-step efficient Generalized Methods of Moment (GMM) procedure. We used a RATS procedure, measure.src, provided by Bernanke and Mihov for estimation.

(2) Shocks to monetary policy have no long-run effects on the relative price of commodities.

(3) Shocks to monetary policy have no long-run effects on the interest rate.

The first and the third restrictions are familiar results of the IS-LM aggregate demand-aggregate supply model. A positive shock to monetary policy initially raises output above the natural level by raising real money balances and, in turn, shifting the LM curve and the aggregate demand curve. Consequently, as we move up the positively sloped short-run aggregate supply curve, output rises but the interest rate falls initially. However, in long-run equilibrium, as prices adjust and we return to the vertical long-run aggregate supply curve, real money balances return to their initial level as do output and the interest rate. The second restriction is another aspect of the assumption of neutrality. That is, monetary policy has no effect on long-run relative prices.

To implement these assumptions using a standard Choleski decomposition, we modified the model in following way. First, all the variables in the model are first differenced prior to estimation. In a VAR estimated in first difference form, the long-run effect of a shock to monetary policy on the level of model variables is the cumulative sum of the relevant part of the moving average representation. Note, in this case, that the moving average representation indicates the effect of the shock on the changes in the variables; hence to obtain the effect on the levels of the variables, the effects on changes must be cumulated. Consequently, in practice, one can easily impose neutrality restrictions by placing real variables prior to monetary variables in a Choleski decomposition.

Second, the model is specified as output, real commodity prices (=commodity prices deflated by the price level), commodity prices, and the three reserve market variables. With the above modification, we can identify shocks to monetary policy by a Choleski decomposition of the long-run relations with following ordering: output, real commodity prices, the federal funds rate, nonborrowed reserves, total reserves, and commodity prices. As noted earlier, the ordering implies that the shock to monetary policy has no long-run effect on output, real commodity prices or the interest rate, while the shock is allowed to affect total reserves and commodity prices in the long-run. Note that the impulse responses of the price level can be easily recovered from the difference in impulse responses between real commodity prices and commodity prices. An appealing feature of this approach is that it attempts to use less controversial long-run neutrality assumptions. It, however, is also not free from criticism. Faust and Leeper (1994) note that the estimates of the impulse response function might be distorted since this approach imposes infinite horizon restrictions in a VAR estimated with data from a finite sample.

2.2.2. Lag Structures

The first lag structure considered is symmetric in the sense that the same number of lags is assumed for each variable in each equation. For the symmetric lag structure, following Christiano, Eichenbaum, and Evans (1996), a lag of four quarters is used. The second lag structure is the asymmetric lag structure suggested by Keating (1995) in which the lag length potentially differs across the variables in the model but is the same for a particular variable in each equation of the model. Keating demonstrated that the asymmetric lag structure can be developed in the following way. Suppose a structural model has a form:

(2.4) $$\Phi_0 Y_t = C + \Phi_1 Y_{t-1} + \Phi_2 Y_{t-2} + \ldots + \Phi_p Y_p + v_t$$

where Φ_0 is the contemporaneous coefficient matrix, v_t is a vector of N white noise shocks, C is a $N \times N$ vector of constant terms, and Φ_t is an $N \times N$ coefficient matrix. Equation (2.4) can be rewritten as:

(2.5) $$\Phi(L) Y_t = C + v_t$$

where $\Phi(L)$ is $N \times N$ lag polynomial matrix in which its element at the i^{th} row and j^{th} column defined as

(2.6) $$\Phi_{ij}(L) = \phi_{0_{ij}} + \phi_{1_{ij}} L + \phi_{2_{ij}} L^2 + \ldots + \phi_{p_{ij}ij} L^{p_{ij}}$$

Premultiplying equation (2.5) by Φ_0^{-1} yields a reduced form.

(2.7) $$\Gamma(L) Y_t = D + e_t$$

where $\Gamma(L) = \Phi_0^{-1} \Phi(L)$ with the k^{th} and j^{th} element $\Gamma_{kj}(L) = \sum_{i=1}^{p} \Phi_0^{ki} \Phi_{ij}(L)$, $D = \Phi_0^{-1} C$, and $e_t = \Phi_0^{-1} v_t$. If each element in $\Gamma(L)$ has the same maximum number of lag, a symmetric lag structure is obtained. In this case, the lag length for the symmetric lag structure is the largest value of p_{1j}, p_{2j}, ..., p_{Pj}. However, if the structural model is characterized by asymmetric lag, i.e. if the p_{ij}'s in equation (2.6) differ for each element, the Keating-type asymmetric lag structure is theoretically possible.

We note that the Keating-type asymmetric lag VAR can be efficiently estimated by ordinary least squares because each equation has same set of explanatory variables. Given this type of asymmetric lag structures, Keating suggests a systematic search process in which statistical criteria are applied to every possible combination of lag length in order to determine the lag structure of the VAR. We note that the search process involves significant computational costs in terms of time; hence a maximum of eight lags was considered.[2] The lag selection criteria considered are Akaike's information criterion (AIC) and Schwarz's information criterion (SIC). As usual, the lag structure that generates the minimum AIC or SIC is selected as the optimal lag length. The lags selected for each variable in all identification schemes are reported in Table 2.1.

Table 2.1
Selected Lag Lengths for the Keating-type Asymmetric Lag VAR

(a) Christiano, Eichenbaum, and Evans, Strongin, Bernanke and Mihov Schemes

	LRGDP	LGDPD	LPCOM	NBREC1	FFR	TRL
AIC	7	2	6	5	2	3
SIC	1	2	1	2	1	2

(b) Long-Run Restrictions Approach

	DLRGDP	DLRPCOM	DFFR	DNBREC 1	DTRL	DLPCOM
AIC	1	3	5	1	1	6
SIC	1	1	1	1	1	1

Note: LRGDP: log of real gdp, LGDPD: log of gdp deflator
LPCOM: log of the commodity price index
NBREC1: nonborrowed reserves adjusted for reserve requirement change plus extended credit
FFR: the federal funds rate
TRL: total reserves adjusted for reserve requirement changes
DLRGDP: first difference in log of real gdp
DLRPCOM: first difference in (log of commodity price – log of gdp deflator)
DFFR: first difference in the federal funds rate
DNBREC1: first difference in nonborrowed reserves
DTRL: first difference in total reserves
DLPCOM: first difference in log of commodity indexes

[2] If the number of lags for the six variable model ranges from 1 to 8, there are 262,144 (=8^6) possible asymmetric lag VAR specifications. In this case, using a Pentium III processor, it took approximately one and half hours to complete the search.

Next, we ran Ljung-Box Q-tests for residuals from each equation for each selection criterion and found that the residuals based on the SIC suffer from severe serial correlation. We note that this is problematic since an assumption of the identification schemes used here is that VAR residuals are white noise; hence we report results only for lags determined by the AIC.

The last lag structure is an asymmetric lag structure in which the lags of a variable may differ in each equation of the system. This type of lag structure was first introduced by Hsiao (1981) and was employed by Caines, Keng, and Sethi (1982), and McMillin and Fackler (1984), among others. The procedure for lag selection in this type of lag structure is essentially equivalent to a stepwise procedure based on Granger-Causality and Akaike's final prediction error (FPE) criterion. In this paper, following McMillin and Fackler, we determine the appropriate lag length for each variable in each equation in the following way. First, construct an autoregression for each endogeneous variable, say y. Next calculate the FPE by varying the lag in the autoregression from zero to eight. Then find the lag length that minimizes the FPE.

(2.8) $$y_t = a_0 + a_{11}(L)y_t + e_t$$

(2.9) $$FPE_{(k)} = [(T + k + 1)(T - k - 1)][SSR_{(k)} / T]$$

where L = the lag operator, k = the lag length for k=1,...,8, T = number of observations in estimating the autoregression, and SSR = sum of squared residuals. Next, estimate all possible combinations of bivariate models by adding a variable denoted by a variable x in the following equation (2.10) to the autoregression with fixed $a_{11}(L)$. Find the lag length that minimizes the FPE for each bivariate model. The bivariate equation and $FPE_{(k,l)}$ can be described as following equations.

$$(2.10) \qquad y_t = a_0 + a_{11}(L)y_t + a_{12}(L)x_t + e_t$$

$$(2.11) \qquad FPE_{(k,l)} = [(T+k+l+1)(T-k-l-1)][SSR_{(k,l)}/T]$$

where $l =$ the lag length of variable x in bivariate equation. Then compare the minimum $FPE_{(k,l)}$ from each bivariate model with the minimum $FPE(k)$ from the autoregression in equation (2.8). If minimum $FPE_{(k,l)}$ < minimum $FPE(k)$, then the variable x is said to Granger-cause y and is included in the y equation. If not, the variable x is omitted from the y equation. Note that one should determine the order in which variables are added to the y equation if there is more than one variable that Granger causes y.[3] To deal with this problem, following Caines, Keng, and Sethi (1981), the specific gravity criterion is applied. That is, the variable with the lowest FPE from the bivariate equations is added first, holding constant its selected lag in the bivariate equation. A trivariate model is estimated holding constant these two variables, y and x with selected lags. This procedure is repeated until every variable is considered in each equation. The selected lag length for each variable in each equation across identification schemes is reported in Table 2.2.

2.2.3. Data and Estimation

As noted earlier, the model used in this study consists of output, the price level, a commodity price index, total reserves, nonborrowed reserves, and the federal funds rate. Nonborrowed reserves are specified as the policy instrument. All data are extracted from the DRI Basic Economics database. The variables, with their exact description and database name in parentheses, are as follows: output (real gdp: gdpq), the price level

[3] In Hsiao's procedure, the order that variables are considered is potentially important since the lag length for each variable in a equation is often sensitive to the other variables in the equation.

(the chain–weighted price index of gdp: gdpfc), commodity prices index (the Commodity Research Bureau's spot market price index for all commodities: psscom), total reserves (fmrra), nonborrowed reserves (fmrnbc), and the federal funds rate (fyff). The logs of output, the price level, and commodity prices are used, while the level of the federal funds rate is employed. These variables are referred to from now on as LRGDP, LGDPD, LPCOM, and FFR.

Table 2.2
Selected Lag Lengths for the Hsiao-type Asymmetric Lag VAR

(a) Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes

Equations	Variables					
	LRGDP	LGDPD	LPCOM	NBREC1	FFR	TRL
LRGDP	2	1	2	0	5	0
LGDPD	3	4	3	0	0	0
LPCOM	0	0	6	0	1	0
NBREC1	0	0	2	5	1	1
FFR	2	3	6	0	8	0
TRL	0	0	0	0	2	6

(b) Long-Run Restrictions Scheme

Equations	Variables					
	DLRGDP	DLRPCOM	DFFR	DNBREC1	DTRL	DLPCOM
DLRGDP	2	1	5	0	0	0
DLRPCOM	1	5	1	0	0	0
DFFR	1	4	7	1	0	5
DNBREC1	1	1	0	1	1	0
DTRL	0	1	3	0	1	1
DLPCOM	1	0	1	0	0	5

Note: see Table 2.1.

However, both total reserves and nonborrowed reserves are normalized by a 12-quarter moving average of total reserves. We use this type of normalization rather than taking logs since the Bernanke-Mihov model considered includes a linear model of the reserve market. Equilibrium in this model requires demand for total reserves equal to supply of total reserves. The structure of the model is based upon the fact that the supply of total reserves is the sum of nonborrowed reserves and borrowed reserves. Hence, using logarithms is not consistent with this type of linear model. Normalizing total reserves and nonborrowed reserves in this fashion is similar in spirit to both Strongin (1995) and Bernanke-Mihov (1998) who estimated models with monthly data. Strongin argued that, besides consideration of the linear reserve market structure, it would also be useful to have an explicit measure of the mix between nonborrowed reserves and total reserves; he normalized total reserves and nonborrowed reserves by the level of total reserves in the prior month. Bernanke and Mihov (1998) argued that Stongin's procedure is problematic in that it creates volatility in impulse response functions. They suggested a method that normalized total reserves and nonborrowed reserves by a 36-month moving average of total reserves. The normalized total reserves and nonborrowed reserves are referred to as NBREC1 and TRL from now on.

In terms of estimation technique, as we noted earlier, the Keating-type asymmetric lag VAR can be efficiently estimated by ordinary least squares as can the symmetric lag VAR. But, for the Hsiao-type asymmetric lag VAR, ordinary least squares is no longer efficient because the specification of each equation of the model is different. Consequently, we estimate the Hsiao-type asymmetric lag VAR using seemingly unrelated regression (SUR).

The models are estimated using quarterly data for the period 1962:1-1997:4. Data from 1962:1-1964:4 are used as pre-sample data since we construct the reserve measures using a 12-quarter moving average. The model is estimated over the period 1965:1-1997:4.

2.3. Comparing Impulse Responses and Identified Shocks across Lag Structures

This section reports the empirical results. First, we compare the effects of monetary policy shocks in the aforementioned four identification schemes on output, price, and interest rates in the symmetric VAR framework. However, the results presented here are qualitatively similar to those of McMillin (1998), and hence are only briefly discussed. Second, we investigate the effects of monetary policy shocks in alternative identification schemes across three different lag structures. More specifically, we try to answer the following questions:

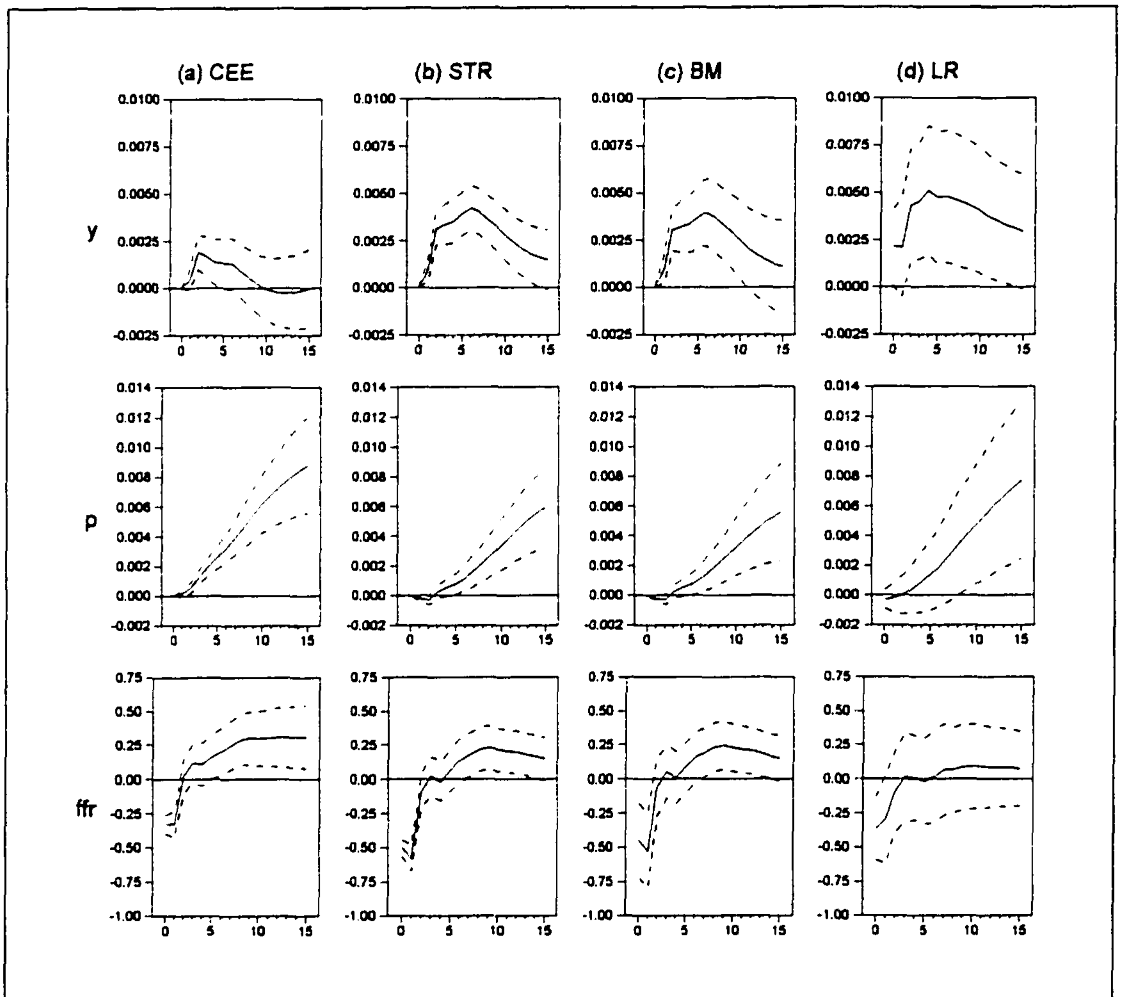
(i) How similar are estimates across lag structures for a given identification scheme?

(ii) Is one identification scheme more sensitive to lag length than others?

2.3.1. Comparing the Effects of Policy Shocks from the Alternative Identification Schemes in the Symmetric Lag VAR

Figure 2.1 graphs the impulse responses from the alternative identification schemes for output, the price level, and the federal funds rate. In each diagram, the solid lines represent the point estimates, while the dotted lines denote a plus and minus one standard deviation band that is constructed by Monte Carlo simulations with 1,000 replications. On the whole, the magnitudes and timing of the point estimates seem to be different across identification schemes although their basic patterns are consistent with our predictions based on economic theory. Several observations are worth noting. First, we observe a hump shaped response for output in all identification schemes. However, the point estimates for the Christiano-Eichenbaum-Evans scheme indicate relatively weaker and shorter lasting effects of monetary policy shocks compared to other schemes. Second, all identification schemes show a long-lasting effect on the price level. However, the magnitude of effects for the point estimates for the Christiano-

Eichenbaum-Evans and long-run restrictions schemes are stronger than for the Strongin and Bernanke-Mihov schemes. Third, all schemes show a strong liquidity effect, although the magnitude of the point estimates of the liquidity effect is somewhat stronger in the Strongin and Bernanke-Mihov schemes.

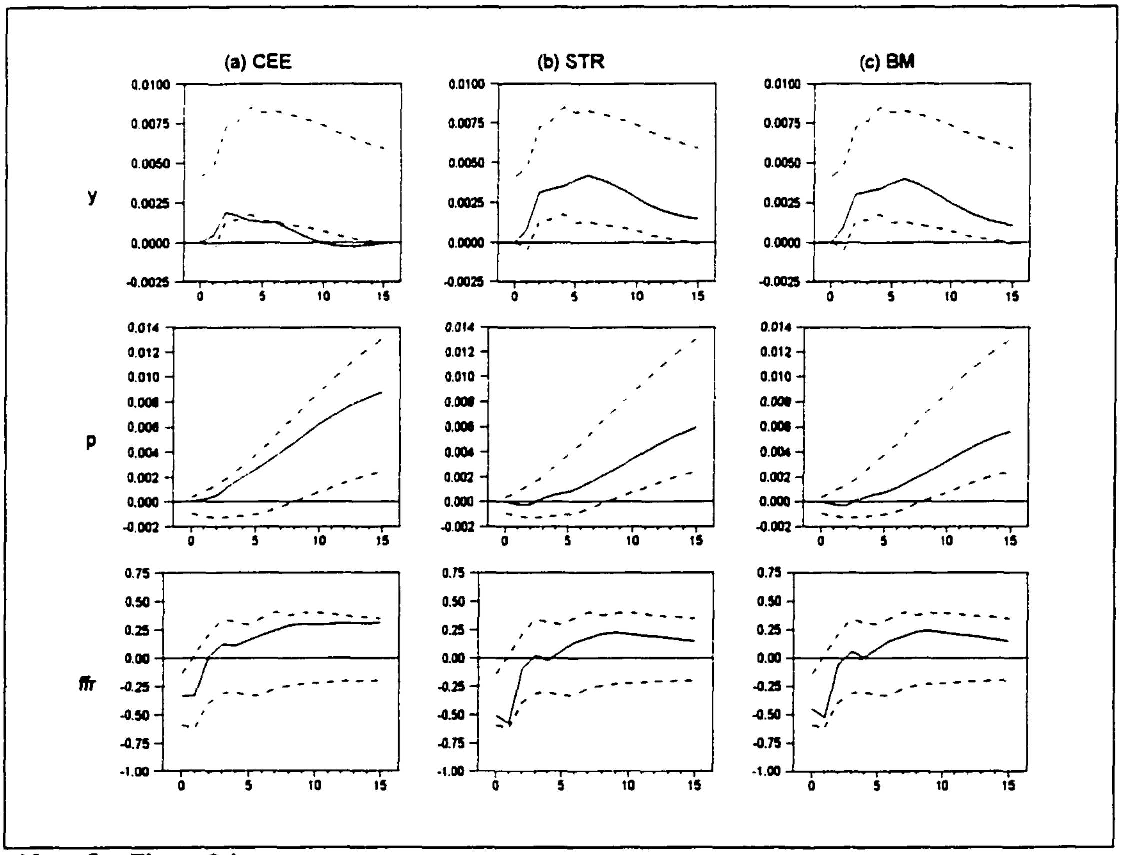


Note: CEE, STR, BM, and LR denote the Christiano, Eichenbaum, and Evans, Strongin, Bernanke and Mihov, and long-run restrictions schemes, respectively. Also, y, p, ffr denote output, the price level, and the federal funds rate.

Figure 2.1
Impulse Response Functions: Symmetric Lag VAR

To clearly see the different effects of monetary policy shocks across identification schemes, we plot the confidence bands of the long-run restriction approach along with the point estimates from other identification schemes in Figure 2.2.

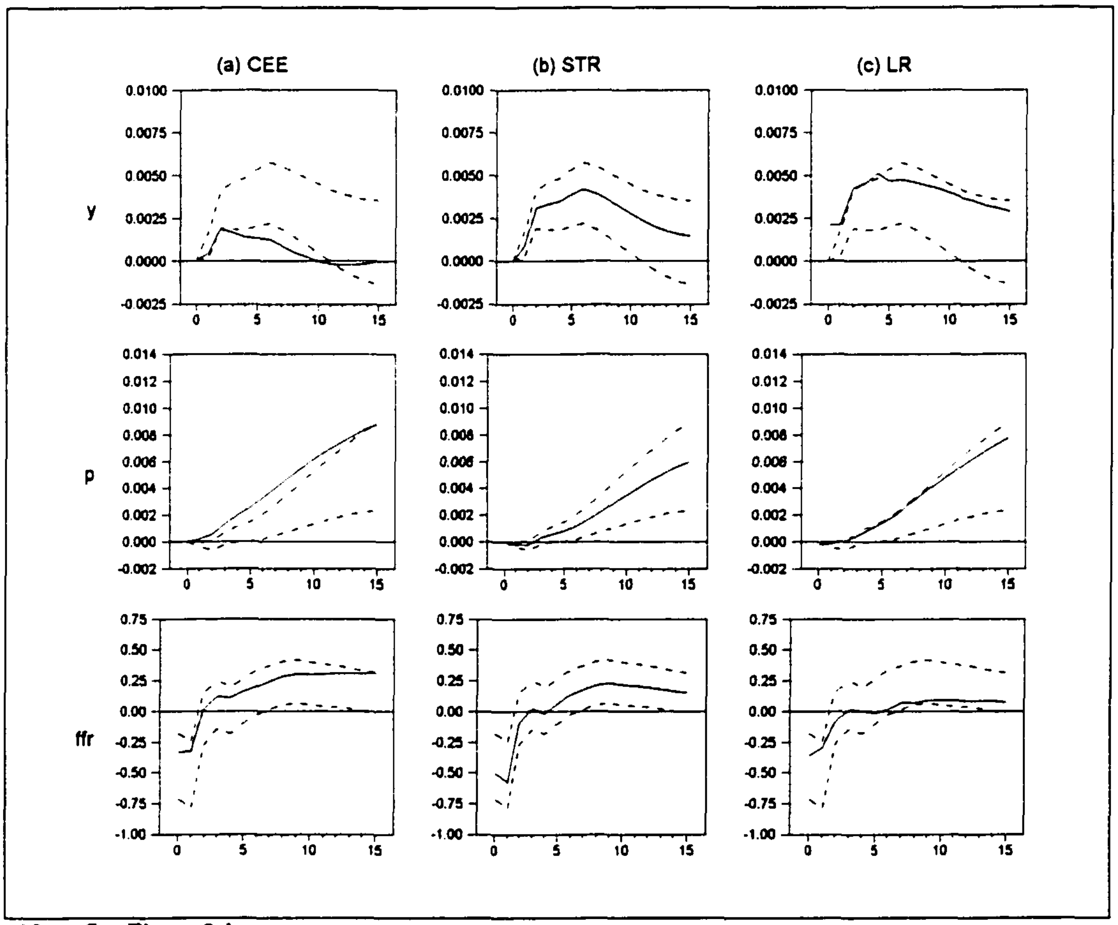
We observe that, for output, the point estimate for only the Christiano-Eichenbaum-Evans approach lies outside of the long-run restrictions confidence intervals, while the point estimates for the Strongin and Bernanke-Mihov schemes lie within the intervals. However, the point estimate for the Christiano-Eichenbaum-Evans approach is still very close to the lower bound of the long-run restrictions confidence intervals. For the price level, the point estimates for all identification schemes lie within the intervals. In the case of the federal funds rate, the point estimates for the Christiano-Eichenbaum-Evans scheme get close to the upper bound of the long-run restriction confidence intervals after approximately 14 quarters, although the point estimates for all identification schemes lie within the intervals.



Note: See Figure 2.1.

Figure 2.2
Long-Run Restrictions Confidence Intervals and Point Estimates from Other Identification Procedures: Symmetric Lag VAR

Figure 2.3 plots the confidence intervals for the Bernanke and Mihov scheme and the point estimates for the other identification schemes. For output, only the point estimates for the Strongin scheme lie entirely within the confidence intervals. The point estimates for the long-run restrictions scheme lie on or above the upper bound for the first 6 quarters, but are within the confidence bands thereafter. However, the point estimates for the Christiano-Eichenbaum-Evans scheme drop below the lower bound after approximately 2 quarters, but return and remain within the confidence bands. For



Note: See Figure 2.1.

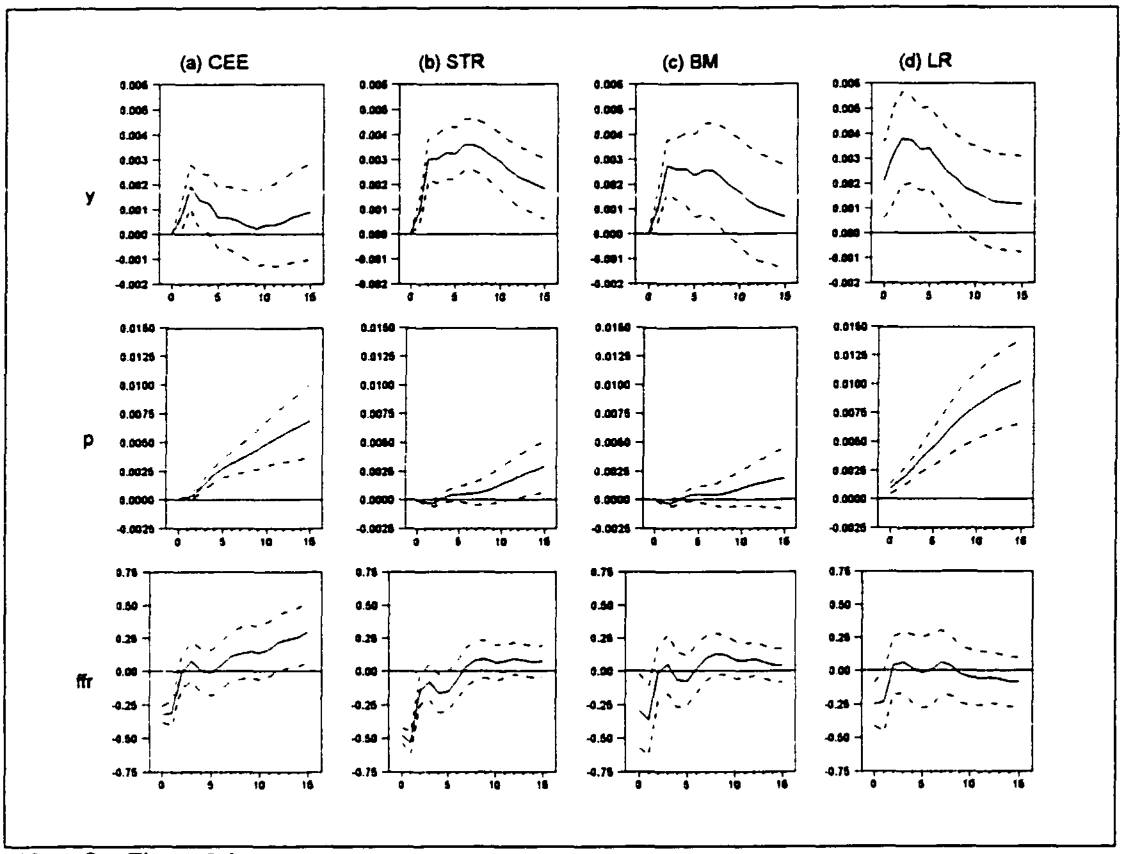
Figure 2.3
Bernanke-Mihov Restrictions Confidence Intervals and Point Estimates from Other Identification Procedures: Symmetric Lag VAR

the price level, the point estimates for the Strongin and long-run restrictions schemes lie within the confidence intervals, while the point estimates for the Christiano-

Eichenbaum-Evans scheme lie entirely above the upper bound of the intervals. In case of the federal funds rate, the point estimates for all identification schemes lie within the Bernanke-Mihov confidence intervals, although the point estimates for the long-run restrictions scheme are close to the lower bound.To summarize, in the symmetric lag structure, all identification schemes considered in this paper generally showed similar impulse responses for output, the price level, and the federal funds rate. We have observed, however, that the results for the Christiano, Eichenbaum, and Evans approach differ from the others in some degree. Note that the only difference between the Christiano, Eichenbaum, and Evans and Strongin schemes is the order of total reserves and nonborrowed reserves. As noted earlier, the causal relationships between these variables for the Strongin or Bernanke-Mihov approaches seem to be more consistent with the common belief that the Federal Reserve generally accommodated shocks to total reserves over most of sample period used here. Consequently, the Strongin and Bernanke-Mihov schemes are likely to be preferred to the Christiano, Eichenbaum, and Evans scheme.

2.3.2. Comparing the Effects of Monetary Policy Shocks for the Alternative Schemes across Lag Structures

Figure 2.4 displays the impulse responses for output, price, and the federal funds rate to monetary policy shocks for four alternative identification schemes in Keating-type asymmetric lag VARs. Overall, the magnitude of point estimates for the Strongin and Bernanke-Mihov schemes for price are clearly smaller than those of the other two schemes, while the magnitude of point estimates for the long-run restrictions approach for output and the price level are greater compared to other schemes. Several points are worth emphasizing. First, the point estimates for the Strongin and the Bernanke-Mihov schemes closely resemble those from the symmetric lag VAR even though the responses of price are weaker. Second, the impulse responses for the long-run restrictions approach are quite similar to those from the symmetric lag VAR. But, the

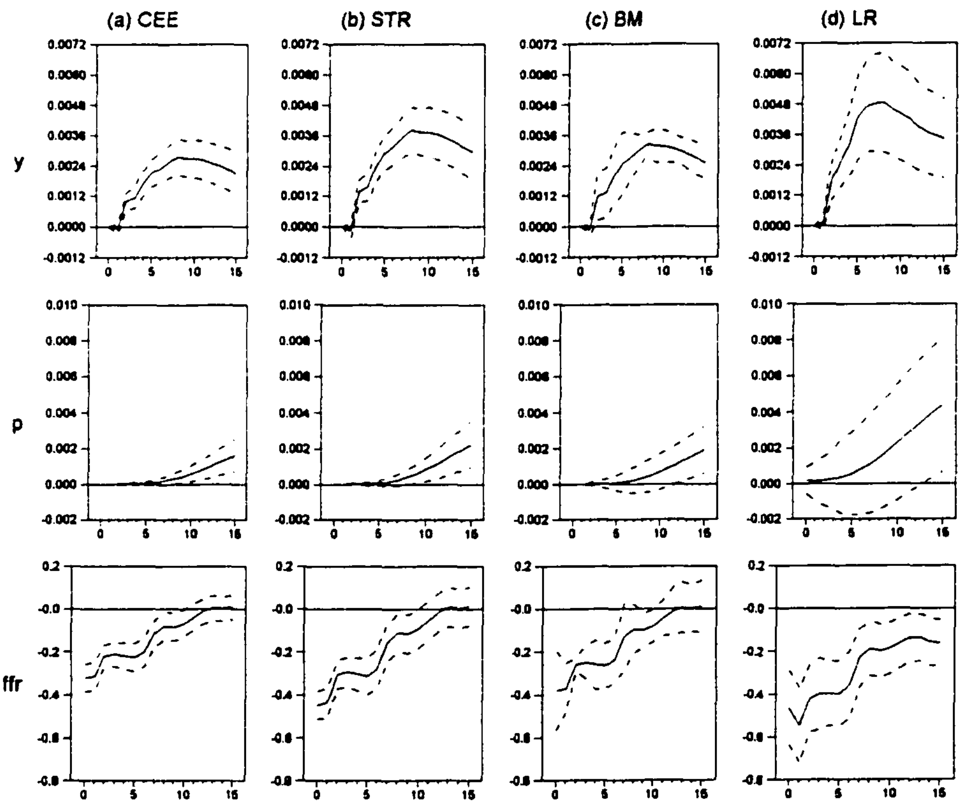


Note: See Figure 2.1.

Figure 2.4
Impulse Response Functions: Keating-type Asymmetric Lag VAR

point estimates and confidence intervals indicate shorter lasting effect of monetary policy shocks on output compared to the symmetric lag VAR. For example, the confidence bands for the asymmetric lag VAR span zero after approximately 9 quarters, while the bands for the symmetric lag VAR include zero after 14 quarters. Third, for the Christiano, Eichenbaum, and Evans scheme, a problematic feature of the impulse responses can be pointed out. The impulse responses of the federal funds rate rise only after an initial liquidity effect and the lower bound of confidence intervals rises somewhat above zero after approximately 12 quarters. Finally, we note that, for output the point estimates from the Christiano, Eichenbaum, and Evans scheme indicate shorter lasting effects compared to other schemes. For the liquidity effect, the Strongin scheme still shows a strong effect. But, for the other schemes, the magnitudes of the effects are relative weaker compared to the cases of symmetric lag VAR.

Figure 2.5 plots the impulse responses for the alternative identification schemes in a Hsiao-type asymmetric lag VAR. In general, the impulse responses seem to be quite different from those in the symmetric lag VAR or in a Keating-type asymmetric lag VAR. However, we note that the impulse responses of output, price, and the federal



Note: See Figure 2.1.

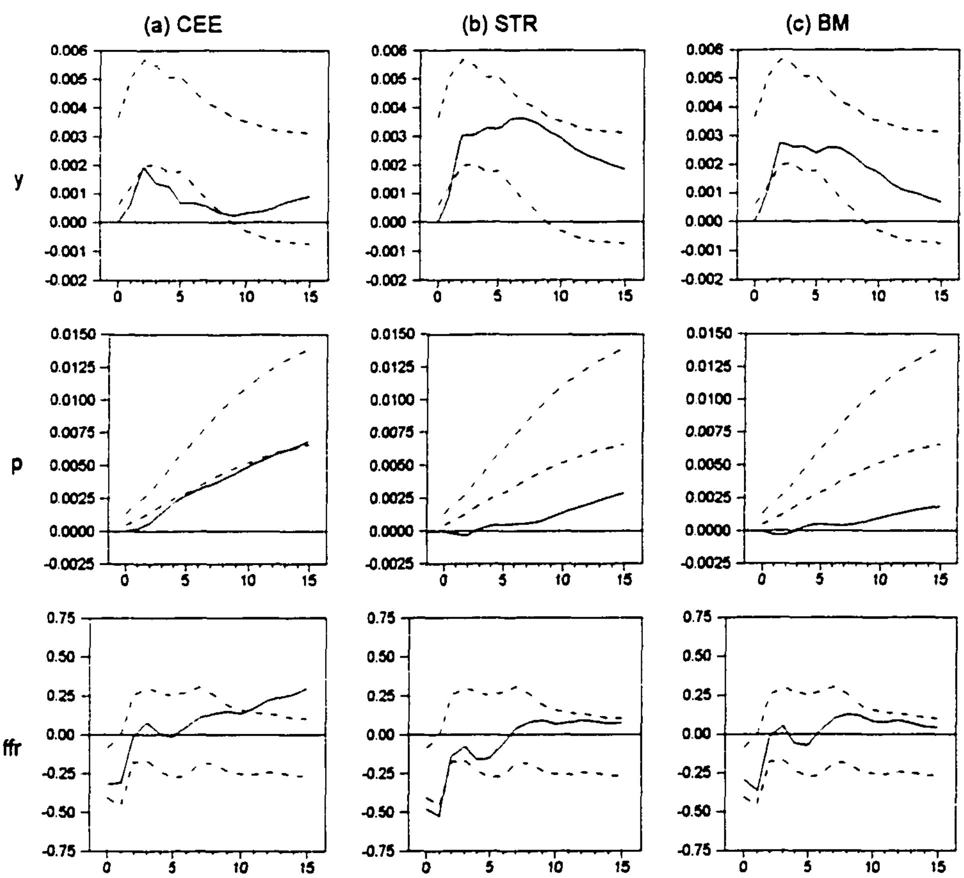
Figure 2.5
Impulse Response Functions: Hsiao-type Asymmetric Lag VAR

funds rate are similar across the alternative identification schemes within the Hsiao-type lag structure although the magnitudes of long-run restrictions seem to be greater.

In this lag structure, the problematic features of impulse responses of the federal funds rate for the Christiano, Eichenbaum, and Evans that appeared in the Keating-type VAR have disappeared. However, there is now another problematic feature for all schemes: for the federal funds rate, the point estimates for all schemes are below the initial value for a very extended time period (about 3 years for the Christiano-Eichenbaum-Evans, Strongin, and Bernanke-Mihov schemes). The point estimates for the long-run restrictions scheme always lies below zero.

Next, we conduct the same exercises as we did for the cases of the symmetric lag VAR in section 3.1. Figures 2.6 through 2.9 plot the confidence bands of a particular identification scheme along with the point estimates of the other identification schemes for the asymmetric lag VARs. Figures 2.6 and 2.7 are the results for the Keating-type asymmetric lag VAR, while the last two figures are for the Hsiao-type asymmetric lag VAR.

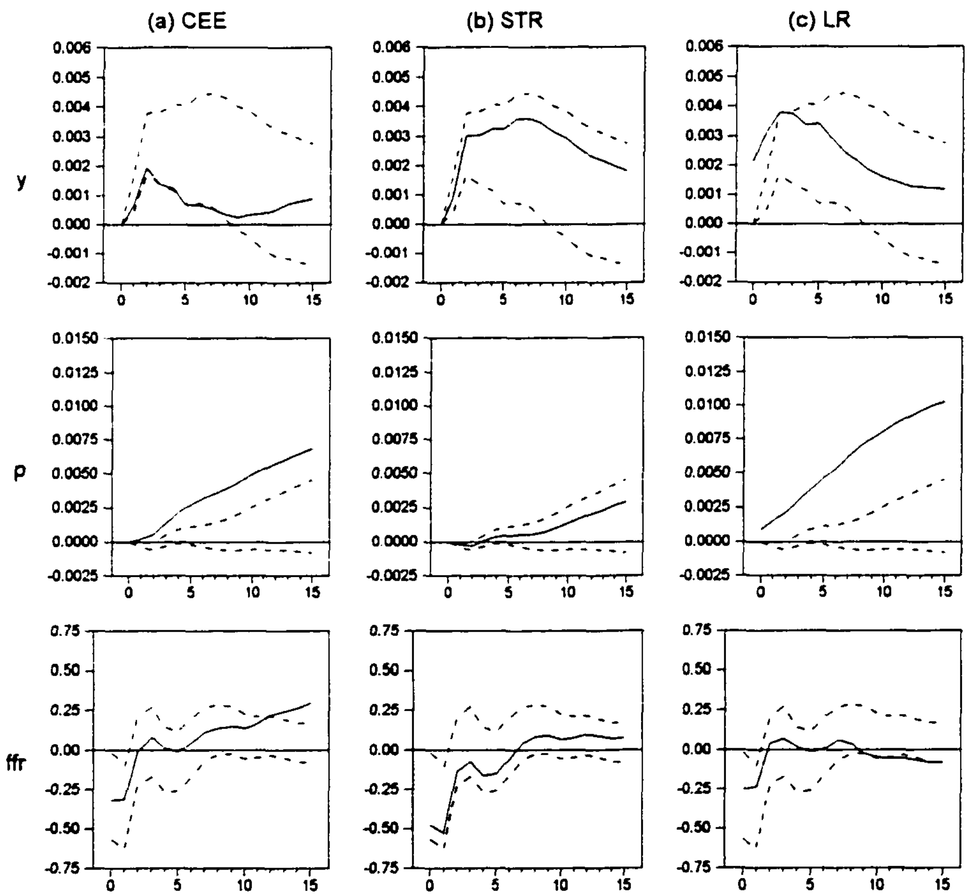
Figure 2.6 graphs the confidence bands of the long-run restrictions scheme and the point estimates from other identification schemes. We observe that, for output, the point estimates for the Christiano, Eichenbaum, and Evans scheme lie below the lower bound for the first 8 quarters, while the point estimates for the Strongin identification scheme lie outside of the intervals. The estimate for the Christiano, Eichenbaum, and Evans scheme are very close to the lower bound, but the estimates for the Strongin and Bernanke-Mihov schemes clearly lie below the lower bound. Certainly, we observe that, for the price level, the magnitudes of effects for the Strongin and Bernanke-Mihov schemes are weaker. For the federal funds rate, the point estimates for the Strongin and Bernanke-Mihov schemes virtually lie within the intervals. However, the estimates for the Christiano, Eichenbaum, and Evans scheme lie within the intervals for approximately 11 quarters and are above the upper bound of the confidence intervals after that.



Note: See Figure 2.1.

Figure 2.6
Long-Run Restrictions Confidence Intervals and Point Estimates from Other Identification Procedures: Keating-type Asymmetric Lag VAR

Figure 2.7 plots the confidence intervals of the Bernanke-Mihov scheme against the point estimates of the other schemes. It shows that, for output, the point estimates essentially lie within the intervals, although there are some deviations above the upper bound for the first two quarters for the long-run restrictions scheme. For the price level, only the point estimates from the Strongin scheme lie within the intervals. The estimates from the other two schemes are above the upper bound of the confidence intervals. We observe, however, the point estimates for all schemes lie within the confidence intervals for the federal funds rate initially, although the estimates for the Christiano, Eichenbaum, and Evans scheme lie above the upper bound after approximately 12 quarters.

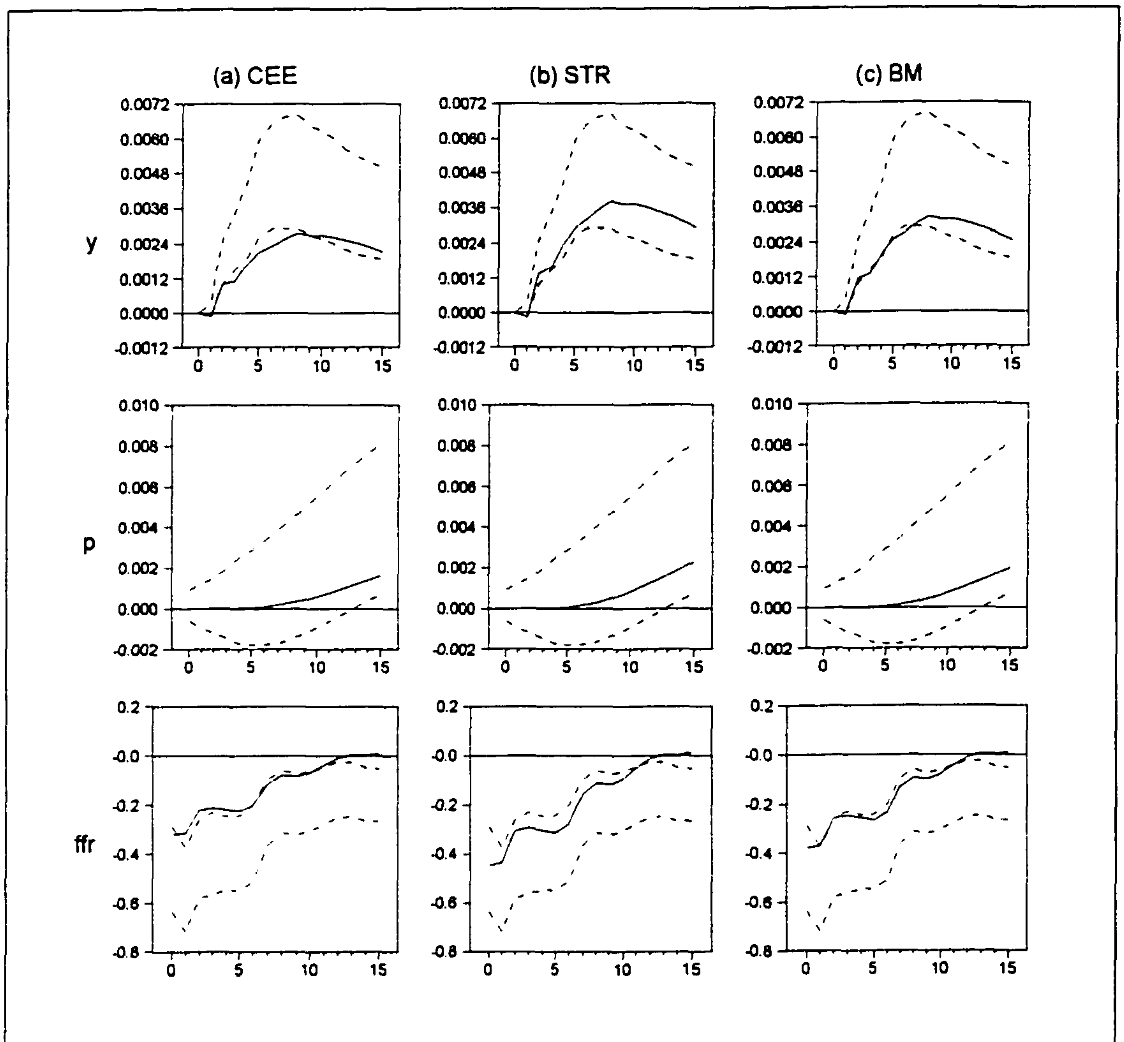


Note: See Figure 2.1.

Figure 2.7
Bernanke-Mihov Restrictions Confidence Intervals and Point Estimates from Other Identification Procedures: Keating-type Asymmetric Lag VAR

Before we move to the results for the Hsiao-type asymmetric lag VAR, one interesting point can be made: the difference between impulse responses for alternative identification schemes in the Keating-type asymmetric lag VAR are in general greater than in the symmetric lag VAR. Based on a comparison of Figure 2.2 and Figure 2.6, there is a bigger difference between impulse responses for the long-run restrictions scheme and for the other schemes in the Keating-type asymmetric lag VAR than in the symmetric lag VAR. Note that, in Figure 2.2, the point estimates for the Christiano-Eichenbaum-Evans, Strongin, and Bernanke-Mihov schemes lie within the confidence bounds of the long-run restrictions scheme; the only exception is the point estimates for the Christiano, Eichenbaum, and Evans scheme for output. But, in Figure 2.6, there are substantial deviations for the Christiano, Eichenbaum, and Evans scheme for output, the price level, and the federal funds rate, while the point estimates for the Strongin and Bernanke-Mihov schemes for the price level always lie below the lower bound. We can also see there is a bigger difference between impulse responses for the Bernanke-Mihov scheme and for the other schemes in the Keating-type asymmetric lag VAR than in the symmetric lag VAR.

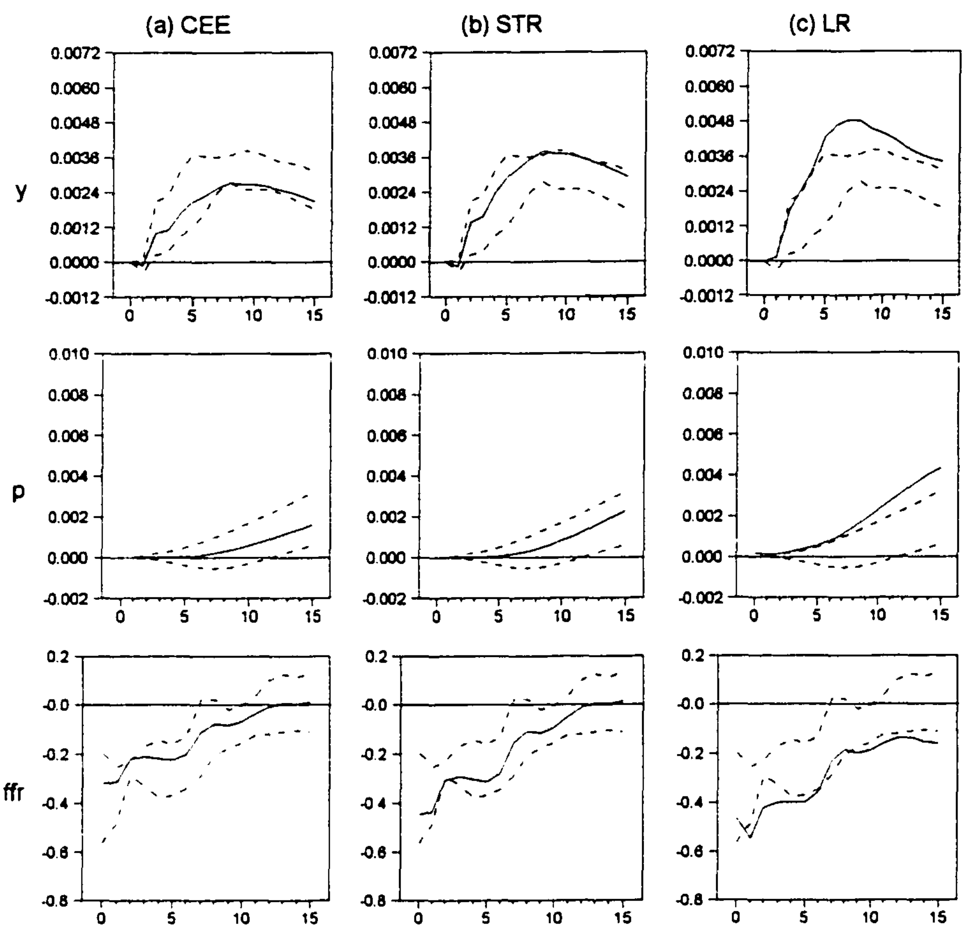
Figure 2.8 graphs the confidence intervals of the long-run restrictions scheme against the point estimates of the other schemes from the Hsiao-type asymmetric lag VAR. For output, the point estimates of the Strongin and Bernanke-Mihov schemes generally lie within the lower bound. The estimates from the Christiano, Eichenbaum, and Evans scheme are below the lower bound, but they are close to the bound. For the price level, the estimates of all schemes always lie within the interval. For the federal funds rate, the point estimates are on or above the upper bound.



Note: See Figure 2.1.

Figure 2.8
Long-Run Restrictions Confidence Intervals and Point Estimates from Other Identification Procedures: Hsiao-type Asymmetric Lag VAR

Finally, Figure 2.9 graphs the confidence intervals of the Bernanke-Mihov scheme and point estimates of the other schemes. For output, the price level, and the federal funds rate, the point estimates from the Christiano, Eichenbaum, and Evans, and Strongin schemes lie within the confidence bands. But the estimates from the long-run restrictions scheme for output, the federal funds rate and the price level often lie outside of the confidence regions.



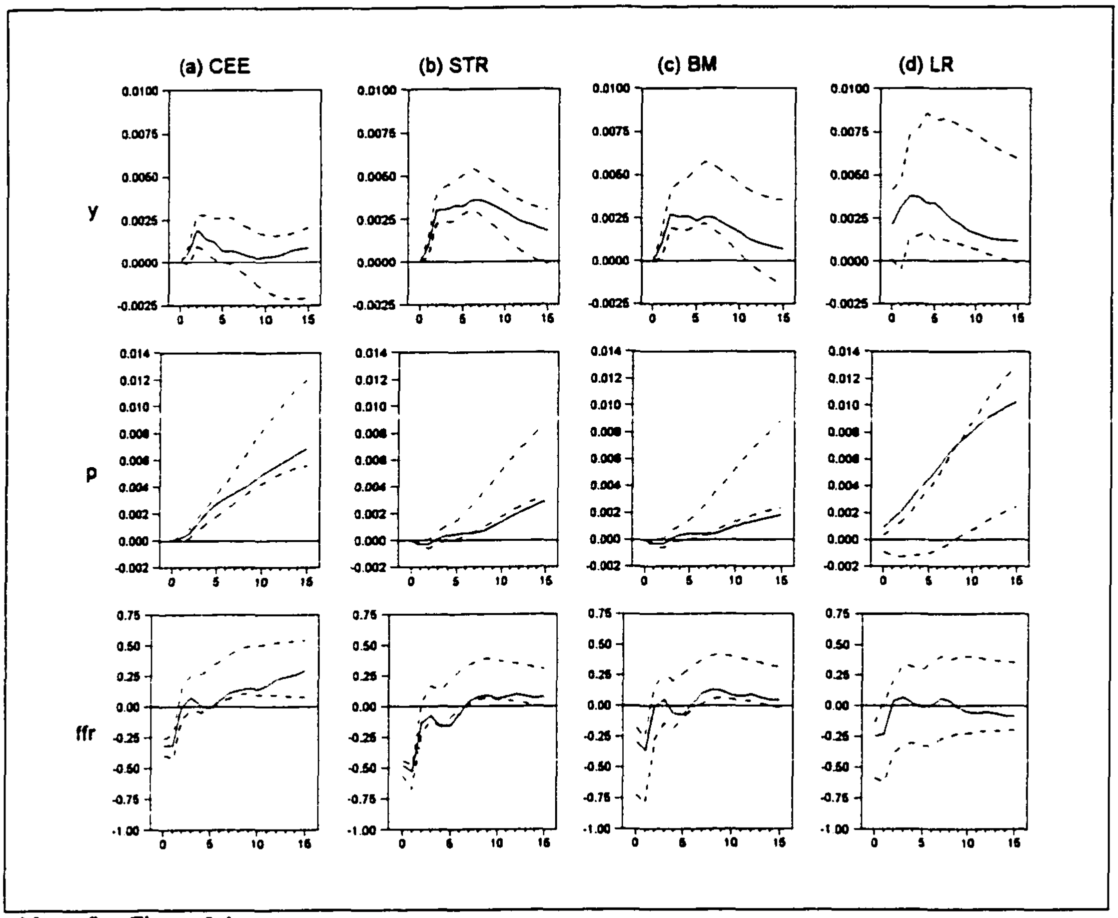
Note: See Figure 2.1.

Figure 2.9
Bernanke-Mihov Restrictions Confidence Intervals and Point Estimates from Other Identification Procedures: Hsiao-type Asymmetric Lag VAR

To summarize, from these figures, we have observed that the sensitivity of results from different identification schemes is more significant for the Keating-type asymmetric lag VARs than for the Hsiao-type asymmetric lag VAR. In fact, for the Hsiao-type lag VAR, the confidence intervals of the long-run restrictions scheme or Bernanke-Mihov scheme, in general, include the point estimates of the other identification schemes, although we observed some deviations from the confidence intervals of both schemes. There are substantial differences for the Keating-type asymmetric lag VAR, however. For example, for the price level, the point estimates of the Strongin and Bernanke-Mihov schemes deviate from the long-run confidence intervals, while the estimates of the Christiano, Eichenbaum, and Evans and long-run restrictions schemes lie outside the intervals of the Bernanke-Mihov scheme. In addition, it is interesting to note that the difference between impulse responses for alternative identification schemes in the Keating-type asymmetric lag VAR is in general greater than in the symmetric lag VAR.

Up to this point, we have seen the magnitudes and timing of effects are somewhat different for each scheme across lag structures. Consequently, it is useful to determine whether these differences are substantial, as we did before. We assume that the symmetric lag is the appropriate lag structure; hence we plot the confidence bands of the symmetric lag VAR along with the point estimates from the two types of asymmetric lag VARs.

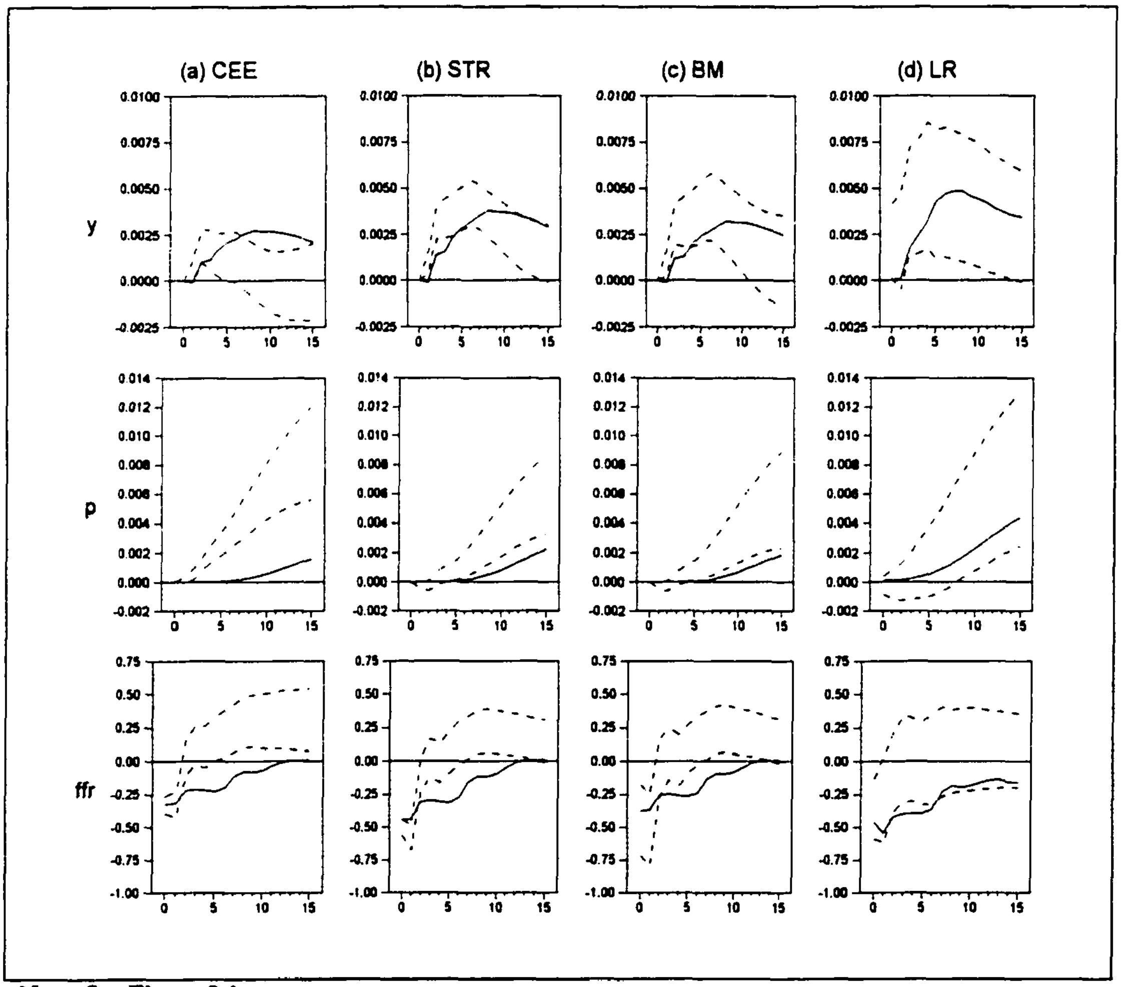
Figure 2.10 plots the confidence bands of the symmetric VAR along with the point estimates for the Keating-type asymmetric VAR. For output, the point estimates from all schemes in the asymmetric VAR lie within the confidence intervals. However, for the price level, the point estimates for only the Christiano, Eichenbaum, and Evans identification scheme lie entirely within the intervals over all horizons. We observe that the point estimates for the Strongin and Bernanke-Mihov drop and remain slightly below the lower bound after approximately seven quarters while the estimates from the long-run restrictions are slightly above the upper bound for the first eight quarters. In case of the federal funds rate, the point estimates from all identification schemes essentially lie within the confidence intervals.



Note: See Figure 2.1.

Figure 2.10
Symmetric Lag VAR Confidence Intervals with Keating-type
Asymmetric Lag VAR Point Estimates

Figure 2.11 plots the confidence bands from the symmetric lag VAR along with the point estimates for the Hsiao-type asymmetric lag VAR. In the case of output, the point estimates for the Christiano-Eichenbaum-Evans scheme are within the confidence intervals for first 6 quarters, but are above the upper bound after that. The point estimates for the Strongin and Bernanke-Mihov schemes are similar in pattern. They are initially outside the confidence bands, but are within the bands after approximately 4 quarters. The point estimates for the long-run restrictions approach are within the confidence bands at all horizons. For the price level, the point estimates for the long-run restrictions procedures lie within the confidence bands for the entire reported horizon, while the point estimates for the Christiano-Eichenbaum-Evans scheme lie outside the bands. We observe that the point estimates for the Strongin, and Bernanke-Mihov schemes lie below the lower bound of the symmetric lag VAR after approximately 5 quarters, although they are close to the bound. For the federal funds rates, the point estimates for the Christiano-Eichenbaum-Evans, Strongin, and Bernanke-Mihov schemes drop below the lower bounds after approximately 1 or 2 quarters, but approach the lower bound again after about 13 quarters. The point estimates for the long-run restrictions marginally lie within the confidence bands.



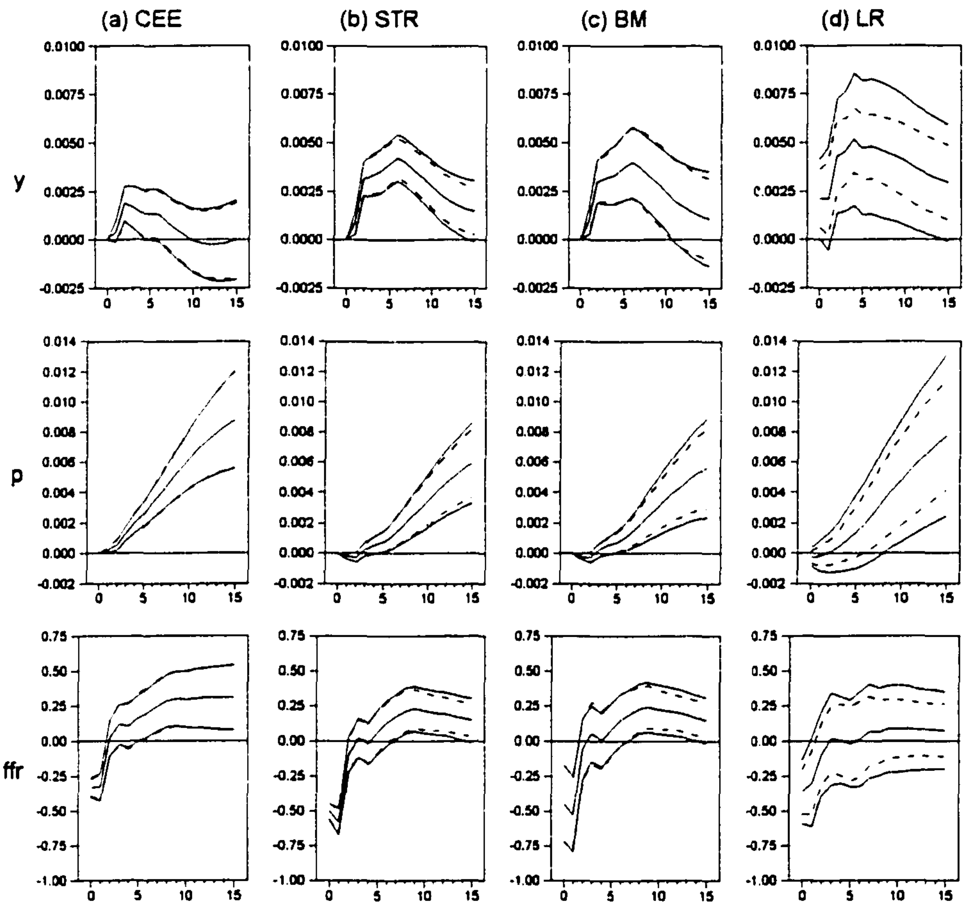
Note: See Figure 2.1.

Figure 2.11
Symmetric Lag VAR Confidence Intervals with Hsiao-type
Asymmetric Lag VAR Point estimates

To summarize, the Christiano-Eichenbaum-Evans scheme does seem to give similar results for both the symmetric and the Keating-type asymmetric lag VARs. In other words, the point estimates of the Keating-type asymmetric lag VAR are always within the confidence bands for the symmetric lag VAR. For the other schemes, the point estimates for output in the Keating-type asymmetric lag VAR are always within the confidence bands for the symmetric lag VAR. Similarly the point estimates for the federal funds rate virtually always lie within the bands. It is only with regard to the price level that the point estimates of the Keating-type asymmetric lag VAR are somewhat out of the confidence bands.

However, for the Hsiao-type asymmetric lag VAR, the results are somewhat different. For the Christiano-Eichenbaum-Evans, Strongin, and Bernanke-Mihov schemes, the point estimates of the Hsiao-type asymmetric lag VAR lie outside of confidence bands of the symmetric VAR for the price level and the federal funds rate. The point estimates for output in the Strongin and Bernanke-Mihov schemes are virtually within the bands. For the long-run restrictions scheme, the point estimates are within the confidence bands of symmetric lag VAR.

Before we conclude this section, we examine whether the confidence intervals for the asymmetric lag VAR are tighter than those for the symmetric lag VAR. As noted earlier, Keating (1995) found that an asymmetric lag VAR model tends to find smaller confidence intervals for impulse responses. In fact, he argued that the smaller confidence intervals along with fewer insignificant parameters suggest efficiency gains from asymmetric lag VAR to symmetric lag VAR. We investigate this point by plotting the confidence intervals from the symmetric lag VAR and the asymmetric lag VAR, simultaneously. In Figure 2.12, three solid lines are the upper bound, point estimates, and lower bound from the symmetric lag VARs. Two dashed lines represent the upper and lower bounds which are constructed by adding the standard deviations from the asymmetric lag VARs to the point estimates for the symmetric VARs.



Note: See Figure 2.1.

Figure 2.12
Confidence Intervals from Symmetric Lag VAR and Asymmetric Lag VAR

We observe that, for the Christiano-Eichenbaum-Evans, Strongin, and Bernanke-Mihov schemes, the confidence intervals from the symmetric lag VAR and asymmetric lag VAR are quite similar in magnitude, although the intervals from the asymmetric VAR are narrower than those from the symmetric lag VAR at longer horizons. There are substantial differences for the long-run restrictions scheme, however. In this case, confidence bounds for the asymmetric lag VARs are much narrower than the bounds for the symmetric lag VARs. We conclude that the confidence bands for the asymmetric lag VARs are at least not wider than those for the symmetric lag VARs. Consequently, we conclude that, for the long-run restrictions scheme, there are substantial efficiency gains from asymmetric lag VAR to symmetric lag VAR. For the other schemes, the gains are small or trivial, however.

2.3.3. Comparing the Alternative Policy Shocks across Lag Structures

We now turn our attention to identified shocks themselves from the alternative identification schemes across lag structures. We do this following two reasons. First, recently Rudebusch (1998) examined the correlation among monetary policy shocks measured by the orthogonalized federal funds rate equation innovations provided by Bernanke-Mihov (1998), Sims and Zha (1995), and Christiano, Eichenbaum, and Evans (1997). He found little or no correlation among those VAR shocks although these models generate quite similar impulse responses. Sims (1998) argued, using a simple supply-demand simultaneous equation model, that this phenomenon might result from different specifications in the VARs. For example, if each VAR model includes an exogenous shifter variable in the policy reaction function that is omitted in another VAR model and vice versa, its measured shocks include the other's shifter variables as well as the true monetary policy innovations. Consequently, a VAR could accurately estimate the impulse responses so long as the shifters are exogenous variables, although its measured shocks are quite different from those for other VARs. Sims, also, noted that for much of Rudebusch's sample period (1988-1995) the estimated shocks are small and that this may help explain the weak correlation among shocks. Therefore it is interesting to examine the correlations among the identified shocks in the four identification schemes, holding constant the model variables and extending the sample period to 1965:1-1997:4. Note that our policy variable is not the federal funds rate as in Rudebusch (1998) but is instead nonborrowed reserves. Second, we may clearly observe the sensitivity of each identification scheme across lag structures by examining the correlation between the identified shocks from a particular identification scheme across lag structures.

Table 2.3 reports the correlation among shocks from the alternative identification schemes in given lag structure. The first four columns of this table present the correlation of shocks for the symmetric lag VARs against shocks for the alternative lag structures. The remaining columns present analogous results for the Keating and Hsiao-type asymmetric VARs. For the correlation among shocks from the symmetric VAR, we observe that there is substantial correlation among shocks from the Christiano-Eichenbaum-Evans, Strongin and Bernanke-Mihov procedures. For example, the shocks for the Strongin and Bernanke-Mihov schemes are closely related as ρ=0.99. The correlation among shocks for the Christiano-Eichenbaum-Evans and for the other two schemes is around 0.76. However, correlation between shocks for the long-run restrictions scheme and shocks for the Strongin scheme or Bernanke-Mihov scheme is relative low (ρ is below 0.40), although their impulse responses are quite similar as we have seen section 2.3.1. For the Keating and Hsiao-type asymmetric lag VARs, shocks for the Strongin and Bernanke-Mihov procedures are still highly correlated. The correlation between shocks for the Christiano, Eichenbaum, and Evans

Table 2.3
Correlation among Shocks for the Alternative Identification Schemes

		Symmetric Lag VAR				Keating-type Asymmetric Lag VAR				Hsiao-type Asymmetric Lag VAR			
		CEE	STR	B-M	L-R	CEE	STR	BM	LR	CEE	STR	BM	LR
Symmetric VAR	CEE	1.000											
	STR	0.775	1.000										
	BM	0.772	0.996	1.000									
	LR	0.750	0.380	0.346	1.000								
Keating-type Asymmetric Lag VAR	CEE	0.959	0.706	0.702	0.758	1.000							
	STR	0.720	0.930	0.928	0.358	0.745	1.000						
	BM	0.686	0.887	0.909	0.242	0.707	0.951	1.000					
	LR	0.773	0.591	0.586	0.690	0.734	0.560	0.525	1.000				
Hsiao-type Asymmetric Lag VAR	CEE	0.880	0.672	0.666	0.684	0.875	0.656	0.616	0.742	1.000			
	STR	0.639	0.849	0.846	0.309	0.606	0.821	0.785	0.565	0.784	1.000		
	BM	0.628	0.834	0.839	0.276	0.599	0.809	0.799	0.548	0.779	0.994	1.000	
	LR	0.840	0.616	0.593	0.789	0.827	0.583	0.487	0.833	0.877	0.643	0.608	1.000

Note: CEE, STR, BM, and LR denote the Christiano, Eichenbaum, and Evans, Strongin, Bernanke and Mihov, and long-run restrictions schemes, respectively.

scheme and other schemes is similar to the cases of the symmetric lag VAR. However, the correlation between shocks for the long-run restriction scheme and shocks for the Strongin or Bernanke-Mihov procedure is somewhat higher compared to the symmetric lag VAR. The increases in the correlation between shocks are more substantial for the Hsiao-type asymmetric lag VAR than for the Keating-type asymmetric lag VAR.

To summarize, we have seen that the correlation between shocks for the Strongin and Bernanke-Mihov Schemes is quite high. We note that this is consistent with the results of the previous impulse response exercises. The shocks for the Christiano-Eichenbaum-Evans scheme also reveal relatively high correlation with those for the Strongin and Bernanke-Mihov scheme, although the shocks for the long-run restrictions scheme are not as highly correlated with shocks for those schemes. These results suggested a possibility that Rudebusch's claim that there is little or no correlation between identified shocks might not be a typical phenomenon in monetary VARs, although the results are difficult to generalize.

Finally, we, in general, observe that shocks from a particular identification scheme across alternative lag structures are highly correlated. For the Christiano, Eichenbaum, and Evans scheme, the correlation between shocks from symmetric lag VAR and Keating type asymmetric VAR is approximately 0.96. The correlation between shocks from the symmetric VAR and Hsiao-type asymmetric VAR drops to 0.88. The patterns are also similar for the Strongin and Bernanke-Mihov procedures. However, the correlation between shocks for the symmetric VAR and both types of asymmetric lag VARs is slightly lower for the long run restrictions scheme.

2.4. Summary and Conclusion

This paper has examined the sensitivity of impulse responses for four widely cited identification schemes across a symmetric and two asymmetric lag structures within the context of a six variable vector autoregressive model. The identification schemes considered are the Christiano-Eichenbaum-Evans, Strongin, Bernanke-Mihov, and long-run restrictions schemes, and the lag structures are the symmetric, Keating, and Hsiao-type asymmetric VARs. The sensitivity of identification schemes is examined by comparing impulse response functions and by computing the correlation among identified shocks.

For the symmetric lag structure, all identification schemes considered generally showed similar impulse responses although the results for the Christiano-Eichenbaum-Evans procedure differ from the others. Specifically, point estimates for the Christiano-Eichenbaum-Evans scheme indicate relative weaker output and liquidity effects. For the Keating-type asymmetric lag VAR, the impulse responses of the Strongin and Bernanke-Mihov schemes for the price level are clearly weaker than those of other schemes, while those of the Christiano-Eichenbaum-Evans scheme for the federal funds rate reveal somewhat problematic features. The impulse responses for the long-run restrictions scheme are quite similar to those from the symmetric lag VAR, although the point estimates show shorter lasting effect of monetary policy shocks on output compared to the symmetric lag VAR. For the Hsiao-type asymmetric lag structure, the impulse responses seem to be quite different from those in the symmetric or the Keating-type lag VAR.

As for the sensitivity of alternative identification schemes across the three types of lag structures, the point estimates for the long-run restrictions scheme in the asymmetric lag VARs generally are within the confidence intervals of the symmetric lag VAR. For the other schemes, the estimates from a Keating-type asymmetric lag VAR are within the intervals, but those from the Hsiao-type asymmetric lag VAR deviate from the intervals, especially for the price level and the federal funds rate.

In terms of correlation between identified shocks, there is substantial correlation among the shocks from the Christiano-Eichenbaum-Evans, Strongin, and Bernanke-Mihov schemes. The shocks for the long-run restrictions scheme are not highly correlated with shocks for the other schemes. When we consider shocks from a particular identification scheme across alternative lag structures, the correlations across identification schemes are high.

To conclude, the impulse responses of output, the price level, and the federal funds rate are often sensitive to identification schemes and lag structures. Therefore, this result suggests that one should pay more careful attention to the lag length selection procedure in order to ensure that identification schemes adequately account for the dynamic effects of shocks in monetary policy. Finally, we note that the long-run restrictions scheme showed relatively insensitive impulse responses across the alternative lag structures. However, the confidence bands for the long-run restrictions scheme are relatively wider than those for the other schemes, indicating less precise estimation. Consequently, it is useful to present the responses from the long-run restrictions scheme along with the responses from the Strongin or Bernanke and Mihov scheme.

CHAPTER 3

MONETARY POLICY SHOCKS AND LAG STRUCTURES IN VECTOR AUTOREGRESSIVE MODELS: A MONTE CARLO EXPERIMENT

3.1. Introduction

Traditionally, most vector autoregression (VAR) models have been estimated using symmetric lag structures; the same lag length is used for all variables in all equations of the model. An advantage of the symmetric lag structure is that ordinary least squares (OLS) yields consistent and efficient parameters. However, it is widely recognized that the VAR models estimated using a symmetric lag structure frequently generate a large number of statistically insignificant coefficients [Runkle (1987), Keating (1995), and Rudebusch (1998)].[1] This may be problematic in assessing the effects of monetary policy shocks within the context of the VAR models because the impulse responses and variance decompositions are functions of the estimated reduced-form coefficients.

Recently, Keating (forthcoming) suggested an asymmetric lag VAR model, an alternative method of constraining the number of the insignificant reduced-form parameters.[2] In this Keating-type asymmetric lag VAR, the lag length potentially differs across the variables in the model but is the same for a particular variable in each equation of the model.[3] Keating argued that optimally selected asymmetric lag VARs

[1] Gordon and King (1982) also pointed out that VAR models usually contain only a limited number of variables since the symmetry in lags rapidly erodes the degree of freedom.

[2] Hsiao's (1981) autoregressive modeling and Litterman's (1986) Bayesian approach are two popular methods of constraining reduced form parameters.

[3] In fact, Hsiao (1981) first examined the possibility of asymmetric lag VAR models. Hsiao's asymmetric lag VAR models differ from the Keating's (forthcoming) in the sense that the lag length on each variable in each equation could differ. We do not consider this type of asymmetric lag VAR model

will probably have a smaller number of estimated parameters than symmetric lag VARs do. Keating (forthcoming) found that, using a small structural VAR model, an asymmetric lag VAR generates relatively fewer insignificant reduced-form parameters than symmetric lag VARs. In addition, he pointed out that the OLS estimates of this type of asymmetric lag VAR are also consistent and efficient as is true for the symmetric lag VARs.

There is, however, no theoretical reason to believe that either a symmetric lag structure or asymmetric lag structure is more appropriate in most VAR models. Indeed, Keating (forthcoming) showed that an asymmetric lag structure in a VAR is theoretically possible if a structural model is characterized by asymmetric lags. However, unfortunately, very seldom does theory provide any guidance as to the appropriate type of lag structure. Therefore, it would be interesting to investigate the distortions in the impulse responses associated with lag structure misspecification in a VAR model. This paper addresses this point using Monte Carlo simulations. This can be done by evaluating and comparing the impulse responses from both traditional symmetric and Keating-type asymmetric lag VARs in a common monetary VAR model.

To estimate the distortion in the impulse response functions, we first assume a particular lag structure – either asymmetric or symmetric – as the 'true' underlying lag structure. Next, we formulate a VAR model which follows the 'true' data generating process (DGP); actual economic data are used to obtain parameter settings and the

in our Monte Carlo study since an extensive iterative procedure is required to appropriately specify a lag structure and would take an acceptably long time to estimate. We also note that the lag structure of a Hsiao-type VAR model should be considered within a simultaneous equation framework. Even if the lag length is optimally selected using single equation methods for each equation of the system, it does not always guarantee the optimal lag structure of the VAR model.

variance-covariance matrix of errors for the Monte Carlo simulations. We, then, fit the VAR model with the alternative lag structure to the simulated series and estimate impulse responses for each replication. Finally, the possible inconsistencies in the impulse responses are investigated by comparing the impulse responses from the 'true' lag structure and from the other lag structure and calculating t-statistics under the null hypothesis that the difference between both impulse response functions is zero. In this paper, two types of lag structures are prespecified as the 'true' lag structure. A symmetric lag structure of autoregressive order 4 is employed when the symmetric lag structure is assumed to be the 'true' lag structure, while a Keating-type lag structure is employed when an asymmetric lag structure is assumed to be the 'true' lag structure.

In addition, although it is not the primary concern of this paper, another distortion in the impulse responses is possible if, given a particular lag structure type, the lag length of a VAR model is misspecified. In applied work, the lag lengths of most symmetric lag VARs are often assumed to be an arbitrary number (for example, 4 for quarterly data or 12 for monthly data), although they are sometimes selected by using explicit statistical criteria like the Akaike Information Criterion (AIC).[4] However, as is well-known, the determination of lag length is a critical element even when the lag structure of a VAR model is known to be symmetric. For example, within a theoretical

[4] We also note that there are several statistical criteria to determine the lag length in a VAR model. Schwarz's Information criteria (SIC) and Phillips' (1994) posterior information criterion (PIC), among others, are good examples. AIC, SIC, and PIC are defined as:

$$AIC = T \log|\Sigma| + 2N$$

$$SIC = T \log|\Sigma| + N\log(T)$$

$$PIC = \log|\Sigma| + (1/T)\log|\Sigma^{-1} \otimes X'X|$$

where $|\Sigma|$ is determinant of variance-covariance matrix of the residuals, N is total number of parameter estimates in all equations, T is number of usable observations, and $\otimes$ is the kronecker product operator. Alternatively, instead of employing a statistical criterion, Koray and McMillin (forthcoming) determined the lag length by examining the serial correlation properties for the VAR residuals for alternative lag length.

framework, Braun and Mittnik (1993) show that the estimators of a VAR whose lag length differs from the true lag length are inconsistent as are the impulse responses and variance decompositions. They also investigate the effects of lag length misspecification on the impulse responses and variance decompositions using Monte Carlo experiments and find that, indeed, the misspecification effects can be serious. In this paper, we also investigate the effects of this type of misspecification on the impulse response functions.

The rest of this paper is organized as follows: section 3.2 describes the empirical methodology, while section 3.3 reports the results. A brief summary and conclusion is presented in section 3.4.

3.2. Methodology

3.2.1. Design of Simulations

Monte Carlo simulations of 500 replications are used to evaluate the impulse responses for the alternative lag structures in a six variable VAR model.[5] The VAR is simulated using prespecified model parameters, prespecified lag structure, and a random number generator. To illustrate, consider a structural model with N variables which follows the true data generating process:

$$\Phi_0 y_t = C + \Phi_1 y_{t-1} + \ldots + \Phi_p y_{t-p} + v_t \tag{3.1}$$

where Φ_0 is the contemporaneous coefficient matrix, v_t is a N×1 vector of structural errors, which we want to identify, with covariance matrix $\sigma^2 I$, C is a N×1 vector of

[5] We choose a relatively small number of replications, 500, for the simulation because of computing time limitations. As illustrated in Essay 1, the Keating-type asymmetric lag search process for our six variable system with a maximum lag of 8 requires about one and half hours to finish an iteration using a PC with Pentium III processor.

constants, and Φ_i is an N×N coefficient matrix. By premultiplying both sides by Φ_0^{-1}, we obtain the VAR representation.

$$(3.2) \qquad y_t = \Phi_0^{-1}C + \Phi_0^{-1}\Phi_1 y_{t-1} + \ldots + \Phi_0^{-1}\Phi_p y_{t-p} + \Phi_0^{-1}v_t$$

For convenience, we can rewrite equation (3.2) as

$$(3.3) \qquad y_t = D + \beta_1 y_{t-1} + \ldots + \beta_{t-p} + e_t$$

where D is $\Phi_0^{-1}C$, β_i is a reduced-form coefficient matrix which equals $\Phi_0^{-1}\Phi_i$, and e_t is a vector of VAR residuals, i.e. $\Phi_0^{-1}v_t$, with variance-covariance matrix $\Sigma\,(=\sigma^2\Phi_0^{-1}\Phi_0^{'-1})$. Consequently, we can generate y_t using equation (3.3) by drawing e_t from N(0, $\sigma^2\Phi_0^{-1}\Phi_0^{'-1}$). Before we generate series for the simulations, we need to specify the matrix of β_i and the variance-covariance matrix of e_t, Σ.

In the spirit of Kennedy and Simons (1991), to obtain the parameter settings (namely the β_i matrix), the 'true' Keating type asymmetric lag structure, and the variance-covariance matrix Σ of the random errors for the simulation, we estimate a six variable quarterly VAR model using actual economic data from 1965:1 to 1997:4. The VAR model comprises output (y), the price level (p), commodity prices (cp), total reserves (tr), nonborrowed reserves (nbr), and the federal funds rate (ffr).[6] As in Essay 1, nonborrowed reserves are taken as the monetary policy variable.[7]

The series for the Monte Carlo simulations are constructed in the following way. First, we treat either the symmetric or the asymmetric lag structure as the 'true', underlying lag structure. When we treat the symmetric lag structure as the 'true'

[6] The exact descriptions of the data are presented in Essay 1.
[7] Although Bernanke and Blinder (1992) contend that the federal funds rate is a good measure of monetary policy, Eichenbaum (1992) argues that nonborrowed reserves are a preferred measure.

structure (Simulation I), it is simply assumed that the model has 4 lags on each variable in each equation. The impulse responses implied by this model are treated as the responses from the 'true' model for the symmetric lag structures. When we assume the asymmetric lag structure is the underlying lag structure (Simulation II), the 'true' lag structure is determined through a more complicated procedure. As suggested by Keating (forthcoming), we compute the AIC statistics using actual data for the possible asymmetric lag VAR specifications in which the lag length potentially differs across the variables in the model but is the same for a particular variable in each equation of the model.[8] In this paper, the maximum lag length, n, is set to 8. Consequently, to complete this search process, it requires 8^6 estimates of the VAR. As usual, the lag structure that generates the minimum AIC is selected as the optimal lag structure. The selected lag structure is 7 for output, 2 for the price level, 6 for commodity prices, 3 for total reserves, 5 for nonborrowed reserves, and 2 for the federal funds rate, respectively. We assume that the impulse responses from this lag structure represent the 'true' impulse responses for the asymmetric lag model. The prespecified lag structures and alternative lag structures for the simulations are summarized in Table 3.1.

Table 3.1
Prespecified and Alternative Lag Structures for Simulations (Summary)

	Lag Structures	
	Prespecified Lag Structure	Alternative Lag Structures
Simulation I	Symmetric Lag VAR(4)	Keating-type Asymmetric Lag VAR Symmetric Lag VAR(AIC)
Simulation II	Keating-Type Asymmetric Lag VAR	Symmetric Lag VAR(4) Symmetric Lag VAR(AIC)

Note: Symmetric Lag VAR(4) and Symmetric Lag VAR(AIC) refer to the symmetric lag VAR whose lag length is 4, and whose lag length is chosen by AIC, respectively.

[8] In fact, Keating (forthcoming) suggests the AIC and SIC as lag selection criteria in the asymmetric lag search process. However, as noted in Essay 1, the lag length selected using the SIC was found to frequently generate autocorrelations in VAR residuals. Hence, we only focus on the AIC in this paper.

Next, as noted earlier, the e_t were selected as random draws from N(0, Σ), and simulated series for y_t are constructed by using equation (3.3). For each draw of the simulation, 632 observations were generated in this fashion. However, to ensure the stationarity of the simulated y_t series, the first 500 observations were discarded; only last 132 observations (the length of the period 1965:1-1997:4) are used for the estimation of the impulse response functions.[9]

Finally, we estimate the VAR model with each alternative lag structure using the simulated series. For example, using the simulated series and assuming the symmetric lag structure with 4 lags as the 'true' lag structure, we first set the search process to determine the optimal lag structure for the Keating-type asymmetric lag VAR for each draw using the method described above. In addition, we also determine the optimal lag length of a symmetric lag VAR using the AIC for each draw. After that, for each draw of the Monte Carlo simulations, the impulse response functions of output, the price level, and the federal funds rate to nonborrowed reserves shocks for the alternative lag structures are computed using the optimal lag lengths selected in the previous step.

To identify the shocks to monetary policy, we consider the three widely-cited identification schemes of Christiano, Eichenbaum, and Evans (1994; 1996), Strongin (1995), and Bernanke-Mihov (1998).[10] For the Christiano, Eichenbaum, and Evans scheme, we consider following the Wold causal ordering of simulated series: y, p, cp, tr,

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[9] Ozciek and McMillin (forthcoming) employed a similar procedure.

[10] The exact specifications and rationales of these identification schemes are presented in Essay 1. We do not consider the long-run restrictions approach pioneered by Blanchard and Quah (1989) which is employed in Essay 1. Since the implementation of this scheme requires data in first-difference form as illustrated in Essay 1, the lag structure chosen in each draw of Monte Carlo simulation can not be directly comparable to the 'true' underlying lag structure described in equation (3.3). Consequently, we may not fully infer the effects of lag structure misspecification from the simulation.

nbr, and ffr. We also consider the ordering of y, p, cp, nbr, tr, and ffr for the Strongin scheme. For the Bernanke and Mihov scheme which blends the Choleski decomposition with a structural model of reserve market, we estimate a simple structural model of bank reserves as in Bernanke and Mihov (1998) using two-step efficient Generalized Method of Moment (GMM).[11] Finally, for the series generated by assuming the Keating-type asymmetric lag structure as the 'true' lag structure, analogous procedures are applied.

3.2.2. Evaluation of Impulse Response Functions

To evaluate the effects of the lag structure misspecification on the impulse responses, we employ two approaches. First, to provide convenient visual comparision, we plot the mean of point estimates from the misspecified models over 500 replications along with the point estimates from the 'true' model. Next, we use a formal approach to test the hypothesis that the differences between the 'true' point estimates and the point estimates from the alternative lag VAR are zero. That is, we calculate the mean-errors (me) for the difference between both impulse response functions and calculate t-statistics under the Ho: mean-error = 0. Specifically, the mean-error of impulse responses for horizon h is defined as:

$$me_h = \frac{1}{R}\sum_{i=1}^{R}(irf_h^i - trueirf_h) \qquad h = 0, 1, \ldots, 15$$

where R is the number of replications, i.e. 500. However, in order to conserve space, the results only for the horizons 1, 3, 5, 7, 9, 11, 13, and 15 are reported.

In addition, to examine the performance of alternative lag selection methods, we compute the mean-square-error (mse) where the error is the difference between the

[11] The detailed descriptions of the system for this scheme are presented in Essay 1.

impulse responses from the 'true' model and the impulse responses from the alternative lag models. Since the mse equals the square of the bias plus the variance of estimator, a lower mse indicates a lower bias or lower variance; hence, a smaller mse is desirable.

The mse of impulse responses for horizon h are defined as:

$$mse_h = \frac{1}{R}\sum_{i=1}^{R}(irf_h^i - trueirf_h)^2 \qquad h = 0, 1, \ldots, 15$$

where R is the number of replications, 500. In this case, we also report an overall mse measure that incorporates all 16 horizons:

$$mse = \frac{1}{R \times 16}\sum_{h=1}^{16}\sum_{i=1}^{R}(irf_h^i - trueirf_h)^2$$

3.3. Empirical Results

This section reports the empirical results. First, by treating the symmetric lag structure of order 4 as the 'true' underlying lag structure, we examine the inconsistencies in the impulse response functions associated with the lag structure misspecification. To obtain general information about the misspecification effects, we first graph the point estimates from the 'true' symmetric lag VARs along with the mean of the point estimates that are computed by fitting the Keating-type asymmetric lag structure to the series generated using the true lag structure. We also investigate the inconsistency associated with possible lag length misspecification when the lag length is not set to 4 but is determined by the AIC. Second, we also calculate the mean-errors where the error is the difference between the impulse responses from the 'true' model and the impulse responses from the Keating-type asymmetric lag VAR. The null hypothesis Ho: mean-error = 0 is tested against H_A: mean-error $\neq 0$ using a standard t-test. For simplicity, from now on, a symmetric lag VAR with autoregressive order 4 is referred as the symmetric lag VAR(4), while the symmetric lag VAR whose lag length is chosen by the AIC are referred as the symmetric lag VAR(AIC). Finally, we investigate the effects of misspecification on impulse responses when a Keating-type asymmetric lag is assumed to be the 'true' lag structure. We repeat analogous steps for this exercise.

3.3.1. Simulation I: Assuming the Symmetric Lag Structure of Autoregressive Order 4 as True

Before we investigate the effects of lag structure misspecification on impulse response functions, we briefly discuss the results of the lags selected in the Keating-type asymmetric lag search process and in the symmetric lag selection process using the AIC. Table 3.2 presents the percentage of lag lengths that each process has specified. The first column in this table is the lag length with the maximum lag 8. The next 6 columns present the results for the Keating-type asymmetric lag search process, while the last column reports the results for the AIC.

In the Keating-type lag search process, the lag lengths selected for each variable mostly fall in lags 3, 4, and 5. For example, for the first variable, the 'true' lag length, 4, is selected 33.2% of the time, while three lags are specified 34.6% of the time. For the second, third, fourth, and fifth variables, the 'true' lag length is selected 57.6%, 53.0%, 39.4%, and 49.4% of the times. However, for the sixth variable, four lags are selected only 22.0% of time. For this variable, three lags are specified 36.0% of the time, and two lags are selected 22.6% of the time. Finally, the mean of the specified lag length for each variable ranges from 3.5 to 4.3. The mean of the specified lag length across all variables is slightly less than 4; the mean is 3.8 (not reported in Table 3.2).[12]

[12] However, no case in the 500 replications correctly selected 4 lags for all six variables.

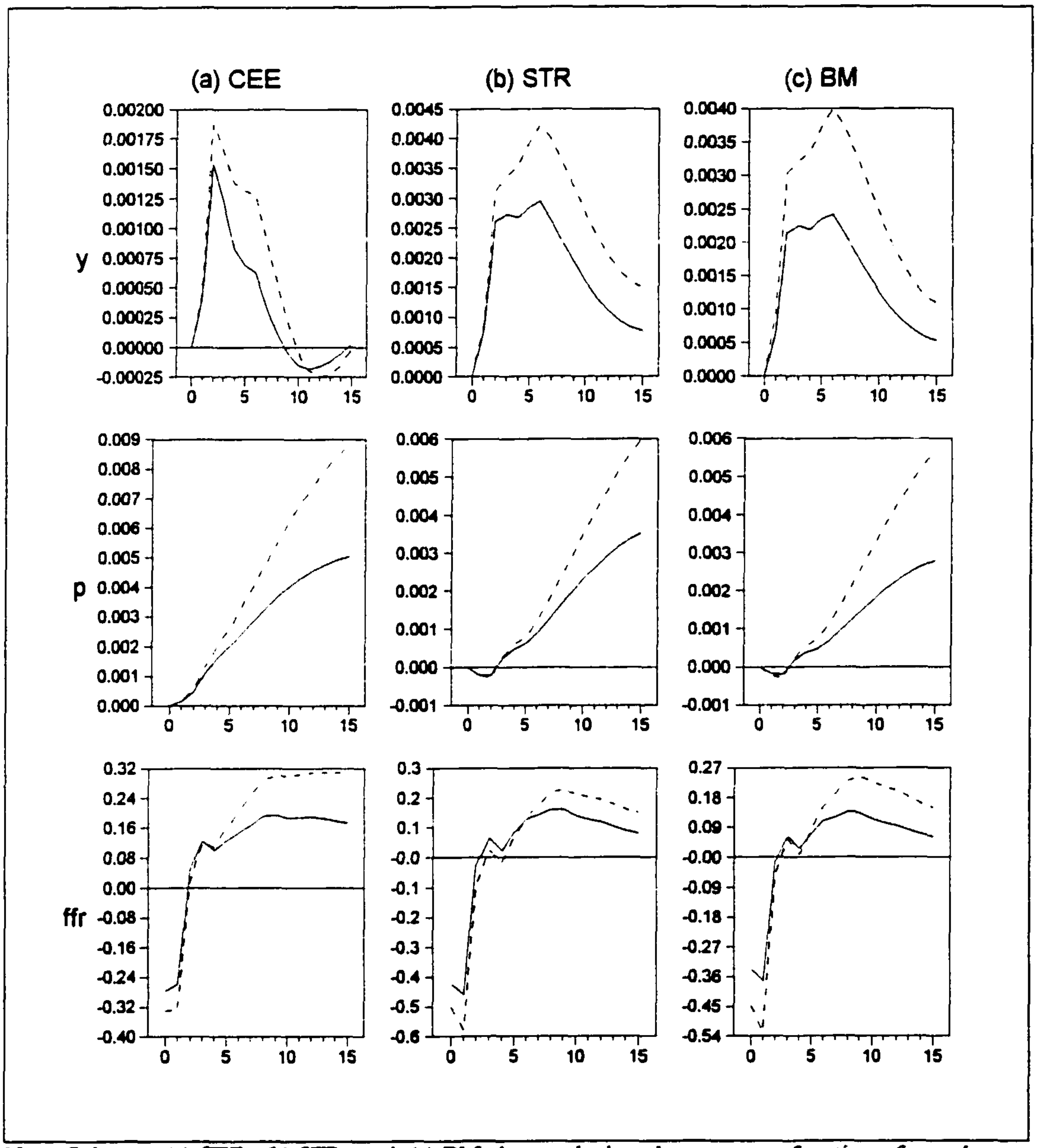
When the lag length is selected using a method whose criterion is to minimize the AIC, 90.4% of time the 'true' lag is specified; three and five lags are found approximately 4% of time. The mean of the specified lag length is about 4.0. Consequently, when the symmetric lag VAR(4) is the 'true' model, the AIC outperforms the Keating-type asymmetric lag search process.

Table 3.2
Percent of Time Lag Length Selected

	Keating-type Asymmetric Lag Search (AIC)						AIC
Lag	Variable 1	Variable 2	Variable 3	Variable 4	Variable 5	Variable 6	
1	2.0	0.6	0.8	0.4	0.6	0.8	0.0
2	9.2	7.8	6.2	15.4	9.8	22.6	0.2
3	34.6	11.2	11.4	24.0	15.6	36.0	4.2
4	33.2	57.6	53.0	39.4	49.4	22.0	90.4
5	9.6	11.4	13.8	9.4	12.2	8.4	4.0
6	5.2	4.0	6.8	4.2	6.4	4.6	0.6
7	3.4	4.8	4.4	4.0	2.4	3.2	0.2
8	2.8	2.6	3.6	3.2	3.6	2.4	0.4
Mean	3.8	4.2	4.3	3.9	4.1	3.5	4.0

Note: Variables 1, 2, 3, 4, 5, and 6 correspond to output, the price level, commodity prices, total reserves, nonborrowed reserves, and the federal funds rate in the actual model, respectively.

We now investigate the effects of the lag structure misspecification on impulse response functions. Consider first the case in which a Keating-type asymmetric lag VAR is fitted to the series simulated using the prespecified symmetric lag structure of autoregressive order 4. Figure 3.1 graphs the impulse responses from the 'true' model for output, the price level, and the federal funds rate as well as the responses from the Keating-type asymmetric lag structure models. The first column of this figure presents the results for shocks to nonborrowed reserves identified using the Christiano, Eichenbaum, and Evans scheme. The remaining columns present analogous results for the Strongin and Bernanke-Mihov schemes, respectively. In each diagram, the solid line is the mean of the point estimates for the Keating-type asymmetric lag VARs that are constructed from the Monte Carlo simulation with 500 replications. The dotted line represents the point estimates from the 'true' model.



Note: Columns (a) CEE, (b) STR, and (c) BM denote the impulse responses functions from the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke and Mihov schemes, respectively. Also, y, p, and ffr refer to output, the price level, and the federal funds rate.

Figure 3.1
Impulse Response Functions: The 'True' Symmetric Lag VAR(4) versus Keating-type Asymmetric Lag VAR

Overall, two points are worth nothing. First, we observe that fitting the Keating-type asymmetric lag structure to the series generated by using the symmetric lag structure of order 4 generally causes some changes to the responses for all identification schemes considered in this paper. As we will see, the responses from the Keating-type asymmetric lag VAR are somewhat weaker than from the 'true' model. Second, the general patterns in the responses from the Keating-type asymmetric lag VAR are similar to the 'true' model, although the magnitudes are different. For example, look at the results for the Christiano-Eichenbaum-Evans scheme. For output, the point estimates from the Keating-type asymmetric lag VAR lie below the 'true' estimates for the first 11 quarters, although they lie above the 'true' estimates after that. For the price level, the point estimates also deviate from the 'true' estimates and the deviation becomes larger as the horizon increases. In the case of the federal funds rate, the estimates are above the 'true' estimates for the first 5 quarters; the liquidity effect is slightly weaker than for the 'true' symmetric lag VAR. The results for the Keating-type asymmetric lag VARs in which monetary policy shocks are identified using the Strongin and Bernanke-Mihov schemes are qualitatively similar; they reveal somewhat weaker output, price, and liquidity effects than the 'true' symmetric lag VAR.

Although the general patterns of the impulse response functions from the true model and from the Keating-type asymmetric lag model are similar, the magnitude of the point estimates from both models are generally found to be different. Hence, as noted earlier, we examine whether these differences are significant using formal test statistics; we calculate mean-errors between the estimated impulse response functions and the 'true' impulse response functions across the 500 replications and t-statistics under the Ho: mean-error = 0 against H_A: mean-error $\neq$ 0 for each horizon.

The calculated mean-errors and their standard errors are presented in Table 3.3. In the table, Panels A, B, and C present the results for the Keating-type asymmetric lag VAR in which monetary policy shocks are identified using the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov Schemes, respectively. Also, the first column of the table denotes the horizons.

In general, the results indicate that, for most horizons, the differences between both impulse responses are significant. For the Christiano, Eichenbaum, and Evans scheme reported in Panel A, the t-ratio's indicate that, for the price level and the federal funds rate, the differences are significantly different from zero at the 1% significant level for all horizons reported; the only exception is horizon 4 for the federal funds rate. In the case of output, the differences are significant at shorter horizons but are typically not significant at longer horizons.

For the impulse responses for output, the price level, and the federal funds rate from the Strongin scheme in Panel B, the null hypothesis (Ho: mean-error=0) can be rejected for all horizons at conventional significant levels. Also, for the Bernanke and Mihov scheme, we strongly reject the null hypothesis for output and the price level, although for the federal funds rate, the hypothesis can not be rejected for horizon 3 even at 10% level.

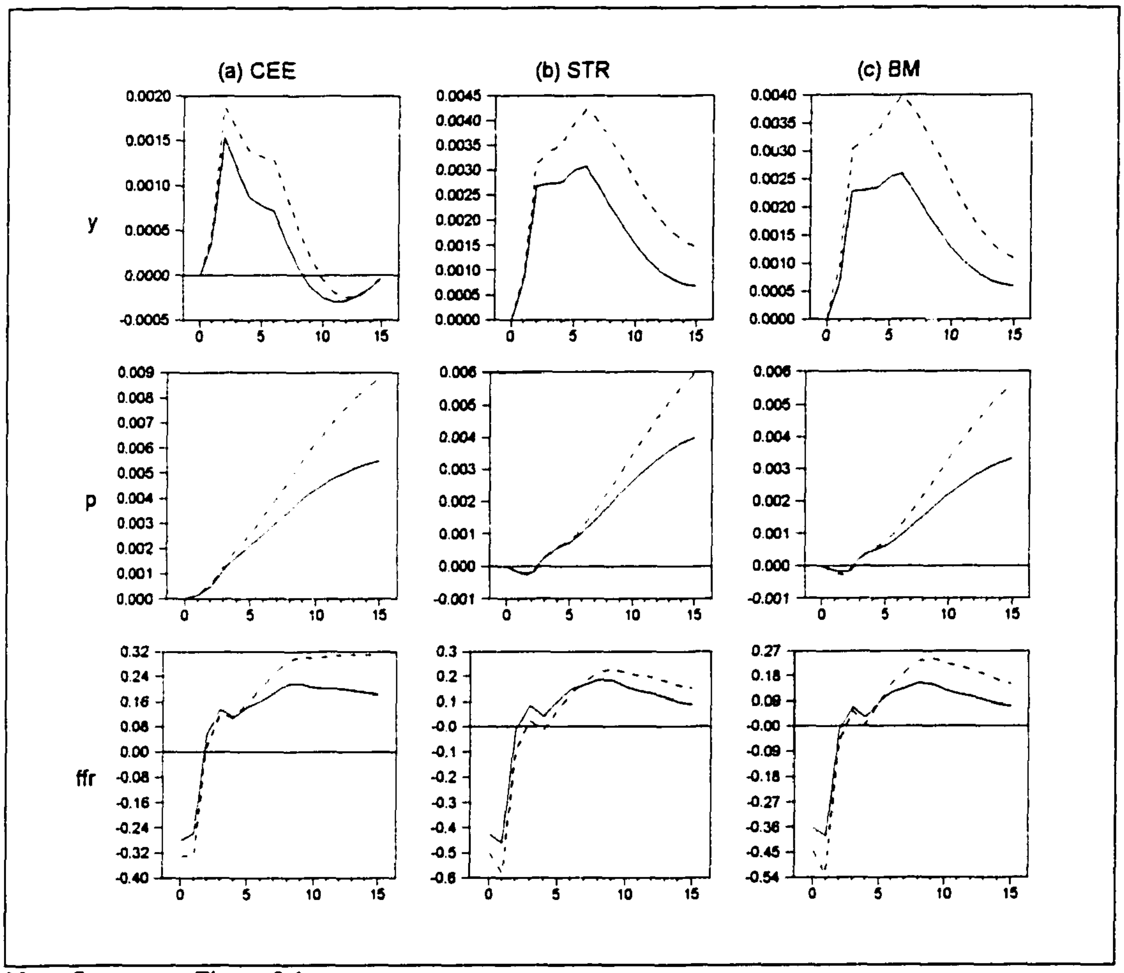
Next, we investigate the inconsistency in the impulse responses of output, the price level, and the federal funds rate when the possible misspecification results not from the lag structure, but from the lag length. As noted earlier, in this case, we assume

Table 3.3
Impulse Response Function mean-error (me):
Keating-type Asymmetric Lag VAR

Panel A: CEE	Output		Price Level		Federal Funds Rate	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-0.471	0.274[c]	-0.361	0.085[a]	0.684	0.044[a]
3	-4.213	0.478[a]	-2.382	0.204[a]	0.048	0.061
5	-6.288	0.544[a]	-6.116	0.341[a]	-0.344	0.070[a]
7	-5.637	0.568[a]	-11.742	0.514[a]	-0.804	0.069[a]
9	-2.595	0.600[a]	-18.481	0.707[a]	-1.070	0.067[a]
11	0.168	0.642	-25.497	0.886[a]	-1.151	0.066[a]
13	1 146	0.668[c]	-32.093	1.040[a]	-1.240	0.065[a]
15	0.361	0.677	-37.934	1.170[a]	-1.374	0.064[a]
Panel B: STR	**Output**		**Price Level**		**Federal Funds Rate**	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-1.139	0.271[a]	0.217	0.076[a]	1.176	0.041[a]
3	-6.302	0.443[a]	-0.403	0.175[b]	0.414	0.058[a]
5	-10.733	0.496[a]	-1.834	0.288[a]	0.185	0.063[a]
7	-13.509	0.519[a]	-4.990	0.431[a]	-0.334	0.059[a]
9	-12.689	0.536[a]	-9.350	0.589[a]	-0.687	0.052[a]
11	-10.416	0.549[a]	-14.513	0.666[a]	-0.765	0.048[a]
13	-8.336	0.545[a]	-19.851	0.802[a]	-0.749	0.045[a]
15	-7.128	0.527[a]	-24.773	0.916[a]	-0.704	0.045[a]
Panel C: BM	**Output**		**Price Level**		**Federal Funds Rate**	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-3.030	0.312[a]	0.630	0.085[a]	1.529	0.104[a]
3	-9.688	0.592[a]	-0.627	0.191[a]	0.094	0.074
5	-13.795	0.702[a]	-2.678	0.317[a]	-0.133	0.070[c]
7	-15.187	0.734[a]	-6.623	0.496[a]	-0.724	0.067[a]
9	-13.700	0.725[a]	-11.855	0.704[a]	-1.085	0.060[a]
11	-10.226	0.731[a]	-17.761	0.899[a]	-1.097	0.053[a]
13	-7.241	0.740[a]	-23.568	1.064[a]	-1.013	0.048[a]
15	-5.428	0.735[a]	-28.668	1.196[a]	-0.881	0.048[a]

Note: Panel A, B, C display the impulse response function mean-error (me) and its standard error (se) for the Christiano-Eichenbaum-Evans (CEE), Strongin (STR) and Bernanke-Mihov (BM) schemes.
[a] Significant at 1% level
[b] Significant at 5% level
[c] Significant at 10% level

that the 'true' process has a symmetric lag VAR(4) representation. But we fit a VAR in which the optimal lag length is chosen by the AIC [the symmetric lag VAR(AIC)]. Note that although the AIC is widely used in practice in order to determine the lag length in a VAR model, it does not always ensure the selection of the 'true' underlying lags.[13] The point estimates for the symmetric lag VAR(AIC) are presented in Figure 3.2. In each diagram, the dotted line represents the point estimates from the 'true' model.



Note: See note to Figure 3.1.

Figure 3.2
Impulse Response Functions: The 'True' Symmetric Lag VAR(4) versus Symmetric Lag VAR(AIC)

In general, the results for the symmetric lag VAR(AIC) are not very different from those for the Keating-type asymmetric lag VAR. The responses from the

[13] As we have seen at the beginning of this subsection, the AIC selected the true lag approximately 90% of the time. Also, using similar Monte Carlo experiments with a bivariate model, Ozciek and McMillin (forthcoming) found the AIC choose the true lag about 60% of the time.

symmetric lag VAR(AIC) reveal weaker effects than from the 'true' model.[14] In the case of the Christiano, Eichenbaum, and Evans scheme, the point estimates for output lie slightly blow the 'true' point estimates for the first 12 quarters, but the estimates are above the 'true' estimates after that. The estimates for the price level always lie below the 'true' point estimates, while the estimates for the federal funds rate lie above the 'true' estimates for shorter horizons. For the Strongin scheme, the responses for output, the price level, and the federal funds rate are similar to those for the Christiano, Eichenbaum, and Evans scheme. However, the responses for output and the price level are below the 'true' responses for most horizons. The responses for the federal funds rate lie above the 'true' estimates for the first 6 quarters. For the Bernanke and Mihov scheme, the patterns of the responses are quite similar to those for the Strongin scheme.

Again, we test whether the differences between the impulse responses from the 'true' model and from the symmetric lag VAR(AIC) model are significant using t-statistics. The results are presented in Table 3.4. Overall, like the results for the Keating-type asymmetric lag VARs, the responses from the symmetric lag VAR(AIC) are significantly different from those from the 'true' model. For the Christinao, Eichenbaum, and Evans scheme reported in Panel A, the responses for output are significantly different from the 'true' responses for shorter horizons up to 10 quarters. However, we cannot reject the null hypothesis for horizons after 12 quarters. For the price level and the federal funds rate, we can reject the null hypothesis for all horizons at conventional significant levels. In the case of the Strongin scheme (Panel B), for

[14] This is not a surprising result. Since we applied OLS to the series generated by assuming autoregressive order 4 to estimate VAR models in this simulation, there may be downward bias in estimated impulse response functions.

Table 3.4
Impulse Response Function mean-error (me):
Symmetric Lag VAR(AIC)

Panel A: CEE	Output		Price Level		Federal Funds Rate	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-1.026	0.257[a]	-0.229	0.080[a]	0.695	0.043[a]
3	-4.542	0.449[a]	-1.608	0.193[a]	0.155	0.058[a]
5	-5.443	0.527[a]	-4.649	0.326[a]	-0.173	0.068[b]
7	-5.220	0.562[a]	-9.561	0.192[a]	-0.595	0.067[a]
9	-3.314	0.606[a]	-15.387	0.682[a]	-0.869	0.067[a]
11	-1.032	0.646	-21.637	0.863[a]	-1.016	0.068[a]
13	0.046	0.667	-27.751	1.025[a]	-1.130	0.066[a]
15	-0.225	0.675	-33.327	1.165[a]	-1.267	0.064[a]
Panel B: STR	**Output**		**Price Level**		**Federal Funds Rate**	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-1.211	0.259[a]	0.286	0.072[a]	1.170	0.040[a]
3	-6.496	0.417[a]	0.200	0.167	0.601	0.055[a]
5	-9.440	0.469[a]	-0.566	0.270[b]	0.411	0.062[a]
7	-12.881	0.474[a]	-2.893	0.401[a]	-0.089	0.055
9	-13.226	0.502[a]	-6.254	0.551[a]	-0.460	0.050[a]
11	-11.538	0.520[a]	-10.574	0.693[a]	-0.609	0.047[a]
13	-9.486	0.525[a]	-15.353	0.817[a]	-0.641	0.044[a]
15	-7.889	0.508[a]	-19.983	0.922[a]	-0.618	0.045[a]
Panel C: BM	**Output**		**Price Level**		**Federal Funds Rate**	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-3.321	0.317[a]	0.730	0.083[a]	1.339	0.095[a]
3	-9.272	0.548[a]	0.031	0.183	0.158	0.068[b]
5	-11.994	0.652[a]	-1.311	0.298[a]	0.012	0.068
7	-14.625	0.655[a]	-4.370	0.461[a]	-0.576	0.066[a]
9	-13.380	0.658[a]	-8.512	0.653[a]	-0.926	0.060[a]
11	-10.199	0.689[a]	-13.476	0.832[a]	-0.971	0.053[a]
13	-7.075	0.724[a]	-18.623	0.985[a]	-0.906	0.047[a]
15	-4.870	0.728[a]	-23.304	1.109[a]	-0.777	0.046[a]

Note: See notes to Table 3.3.

output, we can reject the null hypothesis for all horizons reported. For the price level and the federal funds rate, t-ratios indicate that the differences between both impulse responses for most horizons are significantly different from zero at conventional significant level; exceptions are horizon 3 for the price level and horizon 7 for the federal funds rate. For the Bernanke and Mihov scheme, we strongly reject the null hypothesis that the responses of output from the 'true' model and from the symmetric lag model whose lag is chosen by the AIC are not different. In the cases of the price level and the federal funds rate, we also can reject the null hypothesis for most horizons. In sum, the point estimates from the two types of alternative lag VARs are significantly different from the assumed 'true' point estimates. The responses are weaker, and the differences are substantial for most horizons for output, the price level, and the federal funds rate.

So far, we have examined the effects of lag structure misspecification on impulse responses. A remaining question is whether impulse responses from the Keating-type asymmetric lag structure or from the symmetric lag structure whose lag length is chosen by AIC more closely resemble the 'true' impulse responses. We investigate this point by computing impulse response function mean-square-errors. Table 3.5 reports the results.

Regardless of the identification scheme or the response variable, the symmetric lag VAR(AIC) model produces impulse response functions that more closely resemble the 'true' responses than the Keating-type asymmetric lag VAR. This is not surprising results since, as we saw earlier, the symmetric lag search process using the AIC outperformed the Keating-type asymmetric lag search process.

For the Christiano, Eichenbaum, and Evans scheme, the overall impulse response function mean-square errors of output, the price level, and the federal funds rate from the Keating-type asymmetric lag VAR are 1.673($\times 10^{-6}$), 6.445($\times 10^{-6}$), and 2.812($\times 10^{-6}$), while the mean-square-errors for the symmetric lag VAR(AIC) are

Table 3.5
Impulse Response Function mean-square-errors (mse): Keating-type Asymmetric Lag VAR vs. Symmetric Lag VAR(AIC)

Panel A: CEE	Output mse($\times 10^{-6}$)		Price Level mse($\times 10^{-6}$)		Federal Funds Rate mse($\times 10^{-2}$)	
Horizons	AVAR	VAR(AIC)	AVAR	VAR(AIC)	AVAR	VAR(AIC)
1	0.378	0.341	0.037	0.032	1.463	1.407
3	1.321	1.215	0.265	0.212	1.865	1.742
5	1.873	1.685	0.957	0.747	2.572	2.357
7	1.932	1.853	2.701	2.124	3.083	2.661
9	1.868	1.945	5.912	4.688	3.450	3.033
11	2.059	2.096	10.419	8.402	3.527	3.348
13	2.244	2.224	15.700	12.953	3.652	3.472
15	2.290	2.280	21.224	17.883	3.946	3.705
Overall	1.673	1.629	6.445	5.282	2.812	2.593

Panel B: STR	Output mse($\times 10^{-6}$)		Price level mse($\times 10^{-6}$)		Federal Funds Rate mse($\times 10^{-2}$)	
Horizons	AVAR	VAR(AIC)	AVAR	VAR(AIC)	AVAR	VAR(AIC)
1	0.382	0.349	0.029	0.027	2.251	2.170
3	1.377	1.292	0.155	0.139	1.862	1.880
5	2.381	1.991	0.449	0.368	2.050	2.094
7	3.169	2.781	1.176	0.886	1.888	1.570
9	3.048	3.008	2.607	1.908	1.869	1.477
11	2.590	2.684	4.818	3.516	1.738	1.510
13	2.179	2.275	7.655	5.695	1.590	1.419
15	1.895	1.914	10.780	8.240	1.544	1.395
Overall	2.073	1.973	3.097	2.320	1.777	1.625

Panel C: CEE	Output mse($\times 10^{-6}$)		Price Level mse($\times 10^{-6}$)		Federal Funds Rate mse($\times 10^{-2}$)	
Horizons	AVAR	VAR(AIC)	AVAR	VAR(AIC)	AVAR	VAR(AIC)
1	0.578	0.613	0.040	0.039	7.779	6.317
3	2.691	2.363	0.187	0.167	2.774	2.389
5	4.364	3.561	0.575	0.462	2.471	2.343
7	5.194	4.281	1.667	1.252	2.792	2.535
9	4.500	3.954	3.880	2.853	2.995	2.698
11	3.714	3.410	7.188	5.271	2.655	2.382
13	3.261	3.117	11.206	8.317	2.222	1.951
15	2.992	2.888	15.368	11.568	1.932	1.699
Overall	3.347	2.938	4.502	3.355	3.253	2.846

Note : AVAR refer to the Keating-type asymmetric lag VAR.

$1.629(\times 10^{-6})$, $5.282(\times 10^{-6})$, and $2.593(\times 10^{-6})$, respectively. For the Strongin and Bernanke and Mihov schemes, the results also reveal similar patterns; the symmetric lag VAR(AIC) outperforms the Keating-type asymmetric lag VAR in the sense that the impulse response function mean-square-errors from the symmetric lag VAR(AIC) are smaller than those from the Keating-type asymmetric lag VAR.

3.3.2. Simulation II: Assuming the Keating-type Asymmetric Lag Structure as True

In this subsection, we investigate the effects of lag structure misspecification on the impulse response functions when the Keating-type asymmetric lag structure is assumed to be the 'true' lag structure. As noted earlier, the assumed 'true' Keating-type lag structure is 7 for output, 2 for the price level, 6 for commodity prices, 3 for total reserves, 5 for nonborrowed reserves, and 2 for the federal funds rate.

As in the previous section, we first discuss the percent of time the AIC selects a particular lag length.[15] As we can see from Table 3.6, the percent of time each lag is selected in the symmetric lag VAR(AIC) is as follows: 0.0% for lag 1, 0.8% for lag 2, 3.2% for lag 3, 4.2% for lag 4, 3.0% for lag 5, 45.8% for lag 6, 32.2% for lag 7, and 10.8% for lag 8. Consequently, the AIC selected lags longer than 6 approximately 89% of the time; hence, the loss in degrees of freedom is substantial for models with 6 or more lags compared to the 'true' Keating-type lag structure or to symmetric lag VAR(4).

Table 3.6
Percent of Time Lag Length selected: AIC model

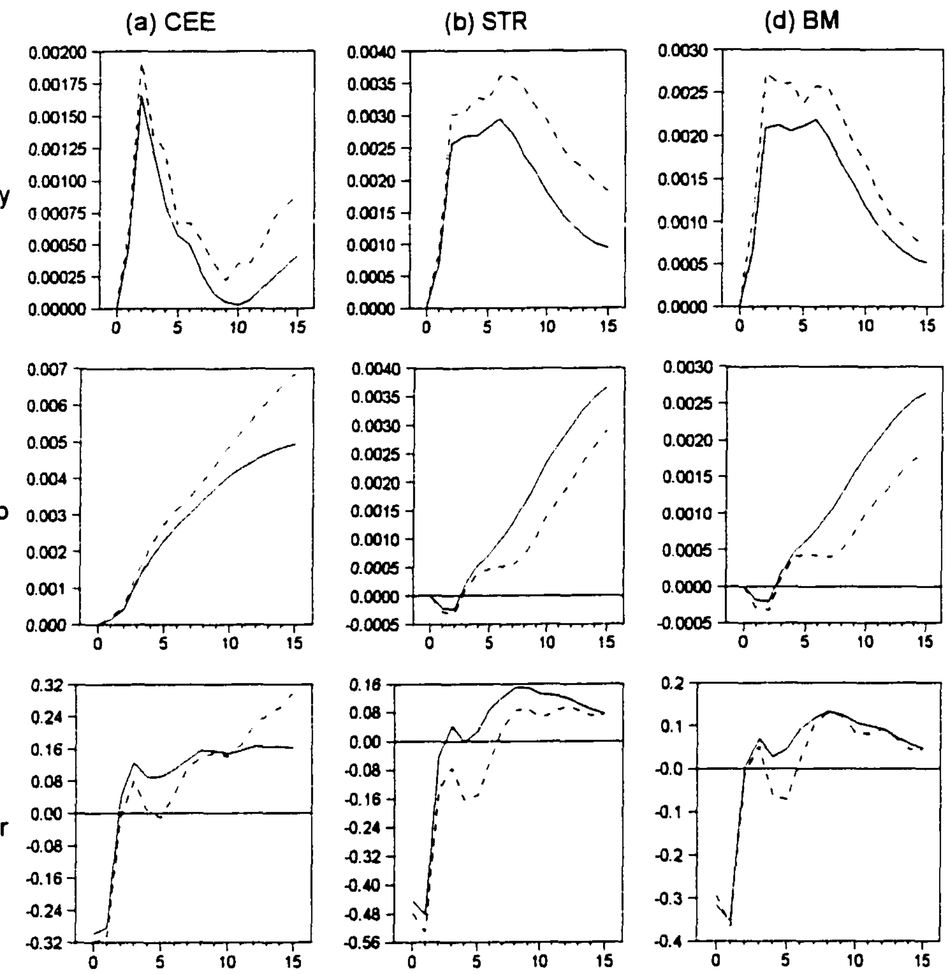
	Lag 1	Lag 2	Lag 3	Lag 4	Lag 5	Lag 6	Lag 7	Lag 8
AIC	0.0	0.8	3.2	4.2	3.0	45.8	32.2	10.8

[15] In this case, unlike the Simulation I reported in section 3.3.1, both the symmetric lag VAR(4) and the symmetric lag VAR(AIC) always lead to misspecification of the lag structure.

To investigate the effects of lag structure misspecification, first, we graph the mean of the point estimates from the symmetric lag VAR(4) along with the point estimates from the 'true' Keating–type asymmetric lag VAR. Second, we calculate the mean-errors between the impulse responses from the 'true' model and from the misspecified model over 500 replications for each horizon. Also, we test, using t-statistics, whether the mean-errors are significantly different from zero. We also repeat these steps for the symmetric lag VAR(AIC) model. Third, to compare the performance of the two alternative symmetric lag VARs, we estimate the mean-square-error (mse).

Now, we examine the effects of lag structure misspecification caused by fitting a symmetric lag VAR(4) to the series whose true data generating process (DGP) follows the Keating-type lag structure described above. Figure 3.3 plots the mean of the point estimates from the symmetric lag VARs in which the optimal lag length is set to 4 along with the point estimates of the 'true' Keating-type asymmetric lag VAR. In the diagrams, the solid lines are the means of the point estimates from the symmetric lag VAR(4)s, while the dotted lines denote the point estimates from the 'true' Keating-type asymmetric lag VARs. Overall, the point estimates from the symmetric lag VAR(4) are different from the 'true' model, although the differences are not large at shorter horizons.

For the Christiano, Eichenbaum, and Evans scheme, the point estimates of output and the price level from the symmetric lag VAR(4) always lie below the true point estimates, while the estimates of the federal funds rate lie above the true estimates for the first 10 quarters. In the case of the Strongin scheme, the point estimates for output from the symmetric lag VAR(4) lie below the true estimates, while the estimates for the price level and the federal funds rate are above the true estimates. However, for



Note: see note to Figure 3.1.

Figure 3.3
Impulse Response Functions: The 'True' Keating-type Asymmetric Lag VAR versus Symmetric Lag VAR(4)

the federal funds rate, the impulse responses from the symmetric lag VAR(4) recover their initial level only after 3 quarters, while the 'true' responses return to the initial level after 7 quarters. For the Bernanke-Mihov scheme, the point estimates of output for the symmetric VAR always lie below the true point estimates. The point estimates for the price level are similar to the case of the Strongin scheme; they lie above the true point estimates. The responses for the federal funds rate are more similar to the 'true' responses compared to other schemes at horizons of 3-8 quarters.

As in the previous section, in order to examine whether the differences between the impulse response functions are significant, the mean-errors and t-statistics (Ho: mean-error = 0) are presented in Table 3.7. In the table, Panels A, B, and C present the results for the Christiano-Eichenbaum-Evans, Strongin, and Bernanke-Mihov Schemes, respectively. Again, the first column of the table denotes the horizons. In general, we observe that the mean-errors are significantly different from zero regardless of the identification scheme in the sense that, in most horizons, we can reject the null hypothesis (Ho: mean error =0) at the 1% significant level. This implies that the distortions in the impulse responses are not trivial when a VAR model is fitted using a symmetric lag structure to the series whose true lag structures is asymmetric. As we will see momentarily, this result is also similar to those for the symmetric lag VAR(AIC) model.

First, look at the results for the Christiano, Eichenbaum, and Evans Scheme presented in Panel A. In the case of output, t-ratios indicate that the mean-errors for all horizons reported except horizon 5 are significantly different from zero at the 1% level. For the price level, we strongly reject the hypothesis that the differences between both impulse response functions are equal to zero for all horizons reported. In the case of the federal funds rate, we can reject the null hypothesis for horizons 1, 3, 5, 7, 13, and 15 at the 1% level; we also can reject the null hypothesis for horizon 11 at the 5% level.

However, for horizon 9, this hypothesis can not be rejected even at 10% level of significance.

Table 3.7
Impulse Response Function mean-error (me):
Symmetric Lag VAR(4)

Panel A: CEE	Output		Price Level		Federal Funds Rate	
	me($\times10^{-4}$)	se($\times10^{-4}$)	me($\times10^{-4}$)	se($\times10^{-4}$)	me($\times10^{-1}$)	se($\times10^{-1}$)
1	-1.424	0.251[a]	-0.415	0.091[a]	0.274	0.047[a]
3	-1.116	0.455[a]	-2.071	0.211[a]	0.438	0.060[a]
5	-0.849	0.529	-4.365	0.340[a]	1.024	0.066[a]
7	-2.550	0.562[a]	-4.679	0.490[a]	0.181	0.066[a]
9	-1.753	0.607[a]	-6.379	0.667[a]	0.062	0.063
11	-2.942	0.636[a]	-10.183	0.841[a]	-0.138	0.061[b]
13	-4.647	0.631[a]	-14.425	0.996[a]	-0.782	0.060[a]
15	-4.738	0.608[a]	-19.236	1.129[a]	-1.374	0.060[a]
Panel B: STR	**Output**		**Price Level**		**Federal Funds Rate**	
	me($\times10^{-4}$)	se($\times10^{-4}$)	me($\times10^{-4}$)	se($\times10^{-4}$)	me($\times10^{-1}$)	se($\times10^{-1}$)
1	-2.236	0.247[a]	0.582	0.080[a]	0.484	0.042[a]
3	-3.487	0.417[a]	1.100	0.185[a]	1.158	0.057[a]
5	-4.166	0.487[a]	2.811	0.292[a]	1.785	0.058[a]
7	-8.665	0.477[a]	7.220	0.425[a]	0.856	0.056[a]
9	-10.111	0.495[a]	9.683	0.580[a]	0.586	0.052[a]
11	-10.325	0.502[a]	9.707	0.732[a]	0.509	0.046[a]
13	-10.214	0.487[a]	9.087	0.866[a]	0.187	0.044[a]
15	-8.860	0.452[a]	7.550	0.978[a]	-0.005	0.043[a]
Panel C: BM	**Output**		**Price Level**		**Federal Funds Rate**	
	me($\times10^{-4}$)	se($\times10^{-4}$)	me($\times10^{-4}$)	se($\times10^{-4}$)	me($\times10^{-1}$)	se($\times10^{-1}$)
1	-3.791	0.311[a]	1.012	0.085[a]	0.136	0.118
3	-4.716	0.544[a]	0.818	0.190[a]	0.182	0.099[c]
5	-2.440	0.735[a]	1.728	0.306[a]	1.185	0.083[a]
7	-5.582	0.869[a]	6.139	0.458[a]	0.152	0.072[b]
9	-4.757	0.957[a]	8.163	0.644[a]	0.029	0.059
11	-3.685	0.960[a]	8.177	0.840[a]	0.203	0.050[a]
13	-3.128	0.908[a]	8.316	1.034[a]	0.033	0.048
15	-1.734	0.825[a]	8.058	1.210[a]	0.013	0.049

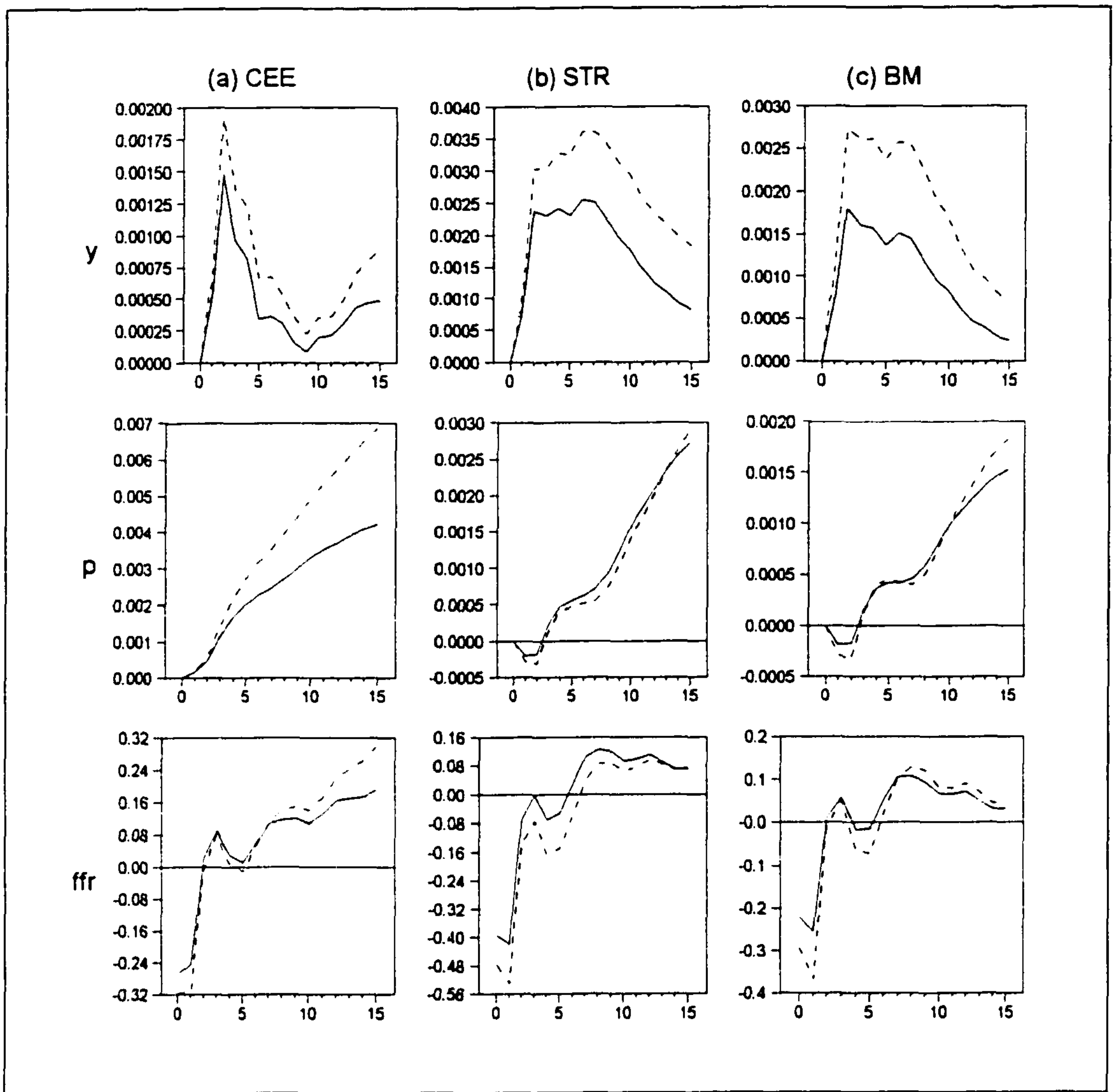
Note: see notes to Table 3.3.

The results for the Strongin Scheme reported in Panel B also show similar results. For output, the price level, and the federal funds rate, the t-statistics indicate that the difference between two impulse responses are significantly different from zero at the 1 % level for all reported horizons.

For the Bernanke-Mihov scheme in Panel C, we strongly reject the null hypothesis for the price level and the federal funds rate. However, in the case of the federal funds rate for horizons 1, 9, 13, and 15, we cannot reject the null hypothesis even at the 10% level of significance. For horizon 3, the difference is only marginally significant. Overall, these results indicate that the differences between the 'true' responses and the misspecified responses are generally significant when the Keating-type lag structure is assumed to be the 'true' lag structure.

We, next, investigate the effects of lag structure misspecification when a symmetric lag structure whose lag length is selected using the AIC is fitted. Figure 3.4 graphs the mean of the point estimates for impulse response functions estimated from the symmetric lag VAR specified using the AIC to determine the optimal lag length. The point estimates of the 'true' Keating-type asymmetric lag VAR are plotted with the dotted line.

In general, the results are similar to the case of the symmetric lag VAR(4), although, for the Strongin and Bernanke and Mihov schemes, the responses of the price level from the misspecified VAR are close to the 'true' responses. For the Christiano, Eichenbaum, and Evans scheme, the point estimates for output and the price level from the symmetric lag VAR(AIC) always lie above the 'true' point estimates. The point estimates for the federal funds rate are close to the 'true' point estimates for the first 6 quarters, although they deviate from the 'true' estimates for longer horizons.



Note: see notes to Figure 3.1.

Figure 3.4
Impulse Response Functions: The 'True' Keating-type Asymmetric Lag VAR versus Symmetric Lag VAR(AIC)

In the case of the Strongin scheme, the point estimates for output also reveal differences in the impulse responses; the estimates lie below the 'true' estimates for all reported horizons, indicating weaker effects on output. The point estimates for the price level are relatively close to the 'true' point estimates, however. The initial liquidity effects for the symmetric lag VAR(AIC) are weaker; the point estimates for the federal funds rate lie slightly above the 'true' estimates. For the Bernanke and Mihov scheme, the patterns are similar to the Strongin scheme. The point estimates for output indicate weaker effects, and the differences between the two impulse responses are large, while the point estimates for the price level and the federal funds rate are close to the 'true' impulse responses.

Next, we investigate the difference between the 'true' impulse responses and misspecified impulse responses by estimating the mean-errors and calculating t-statistics under the Ho: mean-error = 0. Table 3.8 presents the results. For the Christiano, Eichenbaum, and Evans scheme reported in Panel A, t-ratios indicate that, for most horizons, the mean-errors of output and the price level are significantly different from zero at the 1 % level, although, for output, the mean-errors for horizons 9 and 11 are significantly different from zero at the 5% level. For the federal funds rate, we reject the null hypothesis that the mean-error equals to zero for all horizons except horizon 7.

For the mean-errors from the Strongin scheme presented in Panel B, the mean-errors of output and the price level are significantly different from zero at the 1% level; the exceptions are the mean-errors of the price level for horizons 11 and 13. But, for the federal funds rate, the differences in the impulse responses between from the 'true' model and from the misspecified model for horizons 13 and 15 are not significant even at the 10% level.

In the case of the Bernanke-Mihov scheme reported in Panel C, for most horizons, we can strongly reject the hypothesis for output and the federal funds rate although the hypothesis for the federal funds rate cannot be rejected for horizons 3 and 7. For the price level, we cannot reject the hypothesis for horizons 5,7, 9, and 11.

Table 3.8
Impulse Response Function mean-error (me):
Symmetric Lag VAR(AIC)

Panel A: CEE	Output		Price Level		Federal Funds Rate	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-1.652	0.250[a]	-0.333	0.085[a]	0.670	0.044[a]
3	-3.946	0.460[a]	-2.784	0.205[a]	0.112	0.058[a]
5	-3.197	0.536[a]	-7.115	0.330[a]	0.227	0.066[a]
7	-2.263	0.587[a]	-10.699	0.474[a]	-0.068	0.067
9	-1.384	0.617[b]	-14.299	0.628[a]	-0.286	0.062[a]
11	-1.497	0.623[b]	-18.185	0.770[a]	-0.405	0.058[a]
13	-2.715	0.601[a]	-22197	0.895[a]	-0.737	0.057[a]
15	-4.089	0.567[a]	-26.398	1.003[a]	-1.096	0.058[a]
Panel B: STR	**Output**		**Price Level**		**Federal Funds Rate**	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-2.266	0.244[a]	0.777	0.074[a]	1.113	0.043[a]
3	-7.260	0.436[a]	1.041	0.172[a]	0.708	0.055[a]
5	-9.302	0.504[a]	0.825	0.281[a]	0.982	0.062[a]
7	-10.877	0.507[a]	1.603	0.423[a]	0.668	0.063[a]
9	-11.684	0.523[a]	1.945	0.562[a]	0.288	0.055[a]
11	-11.152	0.543[a]	1.266	0.687[c]	0.232	0.048[a]
13	-10.817	0.523[a]	-0.063	0.791	0.051	0.046
15	-10.090	0.479[a]	-1.774	0.881[b]	-0.043	0.045
Panel C: BM	**Output**		**Price Level**		**Federal Funds Rate**	
	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-4}$)	se($\times 10^{-4}$)	me($\times 10^{-1}$)	se($\times 10^{-1}$)
1	-3.621	0.268[a]	1.087	0.078[a]	1.125	0.105[a]
3	-9.931	0.556[a]	0.499	0.170[a]	0.075	0.078
5	-9.990	0.717[a]	-0.307	0.279	0.584	0.068[a]
7	-10.916	0.791[a]	0.592	0.432	0.038	0.064
9	-9.622	0.819[a]	0.484	0.600	-0.282	0.059[a]
11	-7.177	0.793[a]	-0.760	0.779	-0.134	0.052[a]
13	-5.808	0.749[a]	-2.036	0.955[b]	-0.201	0.051[a]
15	-4.497	0.689[a]	-3.113	1.115[a]	-0.124	0.053[b]

Note: See notes to Table 3.3.

In sum, the results indicate that the difference in the impulse responses between from the 'true' model and from the alternative lag structure models are significantly different from zero, although there are some exceptions, especially for the price level and the federal funds rate.

Up to now, we have examined the effects of lag structure misspecification on impulse responses when the Keating-type asymmetric lag structure is assumed to be the 'true' lag structure. As we did in the previous section, we investigate which impulse responses from the alternative symmetric lag structures, i.e. symmetric lag structure with 4 lags and symmetric lag structure whose lag length is chosen by AIC, more closely resemble the 'true' impulse responses. We investigate this point by computing impulse response function mean-square-errors (mse's) as in section 3.3.1.

The impulse response mse's are presented in Table 3.9. In the table, Panels A , B, and C present the mse's for the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke and Mihov schemes, respectively. In general, the impulse response mse's for both misspecified lag models tend to be smaller for the shorter horizons and larger for the longer horizons. This suggests that the lag length misspecification tend to be more serious problem in the long-run than in the short-run. More importantly, the symmetric lag VAR(AIC) generally outperforms the symmetric lag VAR(4) in the sense that the symmetric VAR(AIC) has smaller overall mses for 7 cases out of 9 responses. However, the differences between mses for both models are not very large.

Look first at the case of the Christiano, Eichenbaum, and Evans scheme presented in Panel A. For output and the federal funds rate, the symmetric lag VAR(AIC) has smaller overall mse's than does the symmetric lag VAR(4), while for the price level the symmetric lag VAR(4) has a smaller overall mse. However, for individual horizons, the results are mixed. In Panel B, the results for the Strongin scheme are presented. The symmetric lag VAR(AIC) outperforms the symmetric lag VAR(4) for the price level and the federal funds rate in the overall mse sense. For

Table 3.9
Impulse Response Function mean-square-errors (mse): Symmetric Lag VAR(4) vs. Symmetric Lag VAR(AIC)

Panel A: CEE	Output mse($\times 10^{-6}$)		Price Level mse($\times 10^{-6}$)		Federal Funds Rate mse($\times 10^{-2}$)	
Horizons	VAR(4)	VAR(AIC)	VAR(4)	VAR(AIC)	VAR(4)	VAR(AIC)
1	0.336	0.340	0.043	0.037	1.179	1.449
3	1.047	1.212	0.265	0.288	1.992	1.711
5	1.407	1.538	0.769	1.050	3.281	2.233
7	1.643	1.775	1.418	2.268	2.241	2.307
9	1.872	1.922	2.630	4.013	2.040	2.053
11	2.109	1.962	4.567	6.267	1.890	1.874
13	2.206	1.877	7.031	8.925	2.458	2.222
15	2.070	1.772	10.061	11.997	3.741	2.885
Overall	1.534	1.498	3.010	3.956	2.223	1.988

Panel B: STR	Output mse($\times 10^{-6}$)		Price level mse($\times 10^{-6}$)		Federal Funds Rate mse($\times 10^{-2}$)	
Horizons	VAR(4)	VAR(AIC)	VAR(4)	VAR(AIC)	VAR(4)	VAR(AIC)
1	0.355	0.349	0.035	0.033	1.141	2.162
3	0.991	1.479	0.183	0.159	2.974	2.024
5	1.357	2.135	0.505	0.402	4.875	2.902
7	1.888	2.468	1.425	0.922	2.316	2.438
9	2.245	2.730	2.617	1.617	1.716	1.632
11	2.232	2.715	3.621	2.377	1.361	1.204
13	2.227	2.539	4.574	3.122	1.008	1.101
15	1.806	2.163	5.346	3.907	0.927	1.046
Overall	1.624	2.022	2.114	1.439	2.049	1.772

Panel C: BM	Output mse($\times 10^{-6}$)		Price Level mse($\times 10^{-6}$)		Federal Funds Rate mse($\times 10^{-2}$)	
Horizons	VAR(4)	VAR(AIC)	VAR(4)	VAR(AIC)	VAR(4)	VAR(AIC)
1	0.629	0.490	0.468	0.042	7.022	6.868
3	1.701	2.529	0.188	0.147	4.963	3.075
5	2.760	3.570	0.498	0.391	4.909	2.654
7	4.080	4.318	1.426	0.937	2.671	2.086
9	4.796	4.277	2.738	1.803	1.759	1.833
11	4.740	3.658	4.195	3.037	1.333	1.369
13	4.218	3.138	6.027	4.599	1.178	1.381
15	3.429	2.575	7.963	6.305	1.200	1.438
Overall	3.210	3.038	2.620	2.667	3.297	2.667

Example, the overall mse for the responses of the price level from the symmetric lag VAR(AIC) is $1.439(\times 10^{-6})$, while the overall mse from the symmetric lag VAR(4) is $2.114(\times 10^{-6})$. However, for output, the symmetric lag VAR(4) outperforms the symmetric lag VAR(AIC). Panel C gives the results when the Bernanke-Mihov scheme is employed to identify monetary policy shocks. For the overall mse, the symmetric lag VAR(AIC) outperforms the symmetric lag VAR(4) for output, the price level, and the federal funds rate.

We now make some summary remarks regarding the results of the simulations in which the Keating-type asymmetric lag structure is assumed as true. The results of t-tests indicate that, in general, the point estimates from the symmetric lag VAR(4) are significantly different from those of the 'true' Keating-type asymmetric lag VAR. Regardless of identification scheme, the responses for output from the misspecified models are significantly weaker in the sense that the calculated mean-errors are negative and significant. The weaker output effects of the misspecified VARs also can be seen in Figure 3.3. In the case of the price level, the responses from the misspecified VARs are weaker for the Christiano, Eichenbaum, and Evans scheme, while the responses are stronger for the Strongin and Bernanke-Mihov schemes. However, for all schemes, the differences between both impulse response functions are significant. For the federal funds rate, the responses from the misspecified VARs tend to be weaker and the differences between the two impulse responses are significantly different from zero for most horizons.

We also observe that point estimates from the symmetric lag VAR(AIC) significantly differ from the 'true' estimates across all identification schemes; for all identification schemes, the responses of output from the symmetric lag VAR(AIC) are weaker. For the price level, the responses are weaker for the Christiano-Eichenbaum-Evans scheme, while the responses are stronger for the Strongin scheme. In case of the federal funds rate, the responses from the symmetric lag VAR(AIC) are slightly weaker.

In addition, the results indicate that the impulse responses of the symmetric lag VAR(AIC) more closely resemble the 'true' impulse responses than the symmetric lag VAR(4); in the overall mse sense, the symmetric lag VAR(AIC) slightly outperforms the symmetric lag VAR(4) for 7 cases out of 9 responses. Thus, when the underlying lag structure is asymmetric, the determination of lag length using the AIC is weakly preferred.

3.4. Summary and Conclusion

In this paper, we have investigated the inconsistencies in the impulse response functions when misspecification of the lag structure is present. A symmetric lag structure of order 4, a symmetric lag VAR in which the optimal lag length is chosen by the AIC, and the Keating-type asymmetric lag structure are considered. To the identify the shocks to monetary policy, three widely-cited identification schemes, namely the Christiano-Eichenbaum-Evans (1994;1996), Strongin (1995), and Bernanke-Mihov (1998) schemes, are employed.

In general, we have observed that the responses from the misspecified VARs are different from the assumed 'true' responses. When the symmetric lag structure is assumed to be the 'true' lag structure, the responses from the Keating-type asymmetric lag VAR and from the symmetric lag VAR(AIC) seem to be weaker and are significantly different from the 'true' responses. In addition, the symmetric lag

VAR(AIC) outperforms the Keating-type asymmetric lag VAR. When the Keating-type asymmetric lag structure is assumed to be the 'true' lag structure, the point estimates from the symmetric lag VAR(4) and the symmetric lag VAR(AIC) also significantly deviate from the 'true' point estimates. Our empirical results suggest the following conclusions.

First, the lag structure of a VAR model does matter when assessing the effects of monetary policy shocks. For most horizons, the responses from the VARs with the misspecified lag structure are significantly different from the assumed 'true' responses, although the pattern of the effects is similar from the misspecified lag VARs to the pattern from the 'true' model. However, the quantitative effects are significantly different, and reliable estimates of the quantitative effects are important for policy evaluation. Thus, the determination of lag structure is essential for assessing the effects of monetary policy shocks.

Second, given inherent uncertainty about the lag structure in practice, it is important that one compare the impulse response functions from both symmetric lag and Keating-type asymmetric lag VARs in assessing the effects of monetary policy shocks. Since the differences between both responses are in general significant, employing a particular lag structure alone may result in misleading results. Consequently, this approach may lessen difficulties in specifying the appropriate lag structure in monetary VAR models.

Finally, the results suggest that a symmetric lag VAR whose lag length is chosen by the Akaike Information Criterion (AIC) is preferred to a symmetric lag VAR with an arbitrary autoregressive order, say 4. We note, however, that, to derive strong conclusions about which lag specification procedure or criterion is preferred, further experiments are required. For example, one can employ several different lag lengths in a given lag structure. However, this exercise is beyond the scope of this paper. Hence, we leave the exercise for future research.

CHAPTER 4

IDENTIFICATION OF MONETARY POLICY SHOCKS IN AN OPEN ECONOMY: COMPARING ALTERNATIVE IDENTIFICATION SCHEMES

4.1. Introduction

This paper reexamines the effects of U.S. monetary policy shocks on the exchange rate and the trade balance within a vector autoregression (VAR) model. To assess the effects of monetary policy shocks in an open economy framework, identifying monetary policy shocks is also a critical element as in a closed economy. However, unlike in a closed economy, monetary policy in an open economy may respond to the state of the foreign economy as well as the state of the domestic economy. Hence, identifying monetary policy shocks in an open economy leads to substantial complications relative to the closed economy. These complications may lead to different implications of the effects of shocks to monetary policy across various identification schemes such as the schemes suggested by Christiano, Eichenbaum, and Evans (1994; 1996), Strongin (1995), and Bernanke and Mihov (1998). We note that these identification schemes were originally proposed to identify monetary policy shocks in a closed economy, although Eichenbaum and Evans (1995) looked specifically at open economies.[1] We investigate, in this paper, the sensitivity of results across alternative identification schemes in an open economy framework.

Traditional open economy macroeconomic models including Mundell (1968) and Calvo and Rodriguez (1977) indicate that a positive monetary policy shock

[1] Cushman and Zha (1997) proposed a structural VAR model to identify the monetary policy shocks in an open economy framework. They used Canada as an example.

increases output and the price level, while it decreases the interest rate and depreciates the exchange rate. It also improves the trade balance in the short-run. In the long-run, however, output, the interest rate, and the trade balance are expected to return to their initial level. The price level is expected to be permanently higher.

In general, recent evidence supports this view. For example, Eichenbaum and Evans (1995) investigated the effects of monetary policy shocks on the U.S. bilateral exchange rates. They employed a seven-variable VAR model and relied solely on a Choleski decomposition of the variance-covariance matrix of residuals to identify monetary policy shocks. The main result of Eichenbaum and Evans' study is that contractionary shocks to U.S. monetary policy lead to persistent, significant appreciation of nominal and real exchange rates; the maximal impact of monetary policy shocks on nominal exchange rate takes 2 to 3 years to be felt. Eichenbaum and Evans argued that this finding is inconsistent with the exchange overshooting model [Dornbush (1976)] in which a contractionary monetary policy shock leads to a large initial appreciation followed by a depreciation in exchange rates. Koray and McMillin (forthcoming) extended the Eichenbaum and Evans' work to an 11 variable VAR model in which they adopt the Strongin scheme to identify monetary policy shocks, with a special focus on the trade balance. In contrast to Eichenbaum and Evans, they found the effects of contractionary monetary policy shocks on the exchange rate only last 7 months. Also, the maximal impact of the monetary policy shocks on the exchange rate occurs in 6 months. They argued that these results are consistent with the prediction of the asset market approach to exchange rate determination. They also inferred the typical J-curve effect in that, following a contractionary monetary policy shock, the trade balance improves initially and deteriorates at longer horizons. The J-curve effect refers to a phenomenon that a depreciation of the domestic currency against foreign currency initially worsens the trade balance, but it improves the trade balance over time.[2]

The main purpose of this paper is, as noted earlier, to examine the sensitivity of the effects of monetary policy shocks on the exchange rate and the trade balance across alternative identification schemes. The identification schemes considered in this paper are the approaches suggested by Christiano, Eichenbaum, and Evans (1994; 1996), Strongin (1995), Bernanke and Mihov (1998), and the long-run restrictions approach pioneered by Blanchard and Quah (1989). In addition, we investigate the effects of shocks to the exchange rate on macro variables including the trade balance.[3] This provides a more direct investigation of the J-curve effect than in Koray and McMillin (forthcoming), although the identification of shocks to the exchange rate is not easy.

The remainder of the paper is organized as follows. Section 2 describes the alternative identification schemes of this paper. Section 3 presents the empirical results and compares the results across alternative identification schemes. The results are summarized in the conclusion.

4.2. Methodology

4.2.1 Model Description and Data

We estimate an eleven-variable vector autoregression model using monthly data as in Koray and McMillin (forthcoming). The model comprises output (Y), the price

[2] For further discussion of this issue, see Koray and McMillin (forthcoming), Rose and Yellen (1989), Moffett (1989), and Krugman and Baldwin (1987).

[3] In general, an identified monetary policy shock represents an unanticipated action of the Federal Reserve given its information set. The exchange rate shocks might be interpreted as volatile movements in the exchange rate due to speculation in the currency market rather than factors like monetary policy.

level (P), commodity prices (CP), the federal funds rate (R), total reserves (TR), nonborrowed reserves (NBR), foreign output (Y*), the foreign price level (P*), a foreign short-term interest rate measure (R*), the nominal exchange rate (E), and a real trade balance measure (TB).[4] The index of commodity prices is included in order to capture additional information about future inflation. We expect the inclusion of the index may eliminate the well-known 'price puzzle'. The 'price puzzle' refers to the phenomenon that monetary tightening leads to a rising rather than falling price level in VAR models which do not include information variable about future inflation. Sims (1992) conjectured that the 'puzzle' appears since the information set of VAR models does not include a variable that proxies for the information of future inflation that is available to the Federal Reserve. Christiano, Eichenbaum and Evans (1994; 1996) reported that the inclusion of commodity prices has been found to eliminate the price puzzle.

Following Christiano, Eichenbaum, and Evans (1994), we consider nonborrowed reserves as the policy instrument. In fact, Eichenbaum and Evans (1995) and Koray and McMillin (forthcoming) considered two alternative measures of monetary policy variables, i.e. nonborrowed reserves and the federal funds rate, to identify monetary policy shocks in an open economy framework.[5] However, we do not consider the federal funds rate in this paper. As noted in Essay 1, the only difference between the Christiano, Eichenbaum, and Evans and Strongin schemes is the ordering of nonborrowed reserves and total reserves. Hence, with the federal funds rate as the

[4] We note that the impulse responses of the real exchange rate can be easily recovered, although the model does not explicitly include the real exchange rate.
[5] Bernanke and Blinder (1992) proposed that the federal funds rate is a good measure of monetary policy.

monetary policy variable, the difference between two schemes is not as clear as the case of nonborrowed reserves. In case of the Bernanke-Mihov scheme, it also seems to be more appropriate for nonborrowed reserves than for the federal funds rate as the policy variable. In addition, with the federal funds rate as the policy variable, applying the long-run restrictions approach implies that the Federal Reserve can set the level of the federal funds rate at any desired value in the long-run. We note that the assumption is more questionable than the case of nonborrowed reserves.

Trade-weighted measures of foreign output, the foreign interest rate, the foreign price level, and the exchange rate are constructed using data for the G-6 countries, i.e. the United Kingdom, Germany, France, Japan, Italy, and Canada. We focus on the G-6 countries because of following reasons. First, the G-6 countries include large industrial countries that are important trading partners of the United States. Consequently, U.S. monetary policy may respond to developments in these countries, and U.S. monetary policy may have important effects on the economies of these countries. Second, the quality of data for these countries is good, and consideration of these countries provides comparability to previous studies including Koray and McMillin (forthcoming).

For example, the trade-weighted exchange rate is calculated as follows:

$$E_w = S_m \sum_{i=1}^{6} (IM_i / \sum_{i=1}^{6} IM_i)\,(E_{it} / E_{i0}) + S_x \sum_{i=1}^{6} (EX_i / \sum_{i=1}^{6} EX_i)\,(E_{it} / E_{i0})$$

where E_w is the trade-weighted nominal exchange rate, S_m is the share of U.S. imports in total trade with the G-6, S_x is the share of U.S. exports in total trade with the G-6, IM_i is U.S. imports from country i, EX_i is U.S. exports to country i, E_{it} is the bilateral exchange rate which is expressed as foreign currency units per U.S. dollar at time t, and E_{i0} is the bilateral exchange rate at base period 0. The base period is set to 1974:2. The trade weighted measures of foreign output, the interest rate, and the price level are also calculated in a similar manner.

The model was estimated using log levels for all data except the interest rate variables, total reserves, and nonborrowed reserves. Given the linear structure of the reserve market in the Bernanke–Mihov scheme considered here, the log levels of total reserves and nonborrowed reserves are not appropriate and are not used. Consequently, as in Bernanke and Mihov (1998), both total reserves and nonborrowed reserves are normalized by a 36-month moving average of total reserves. The lag length for the VARs is set to 12.[6] We, however, do not consider the Keating-type asymmetric lag VARs in this paper as in Essay 1, since the Keating-type lag search process is almost impossible for an 11 variable monthly VAR model considered here. For example, we need to estimate 12^{11} VAR specifications to find an optimal Keating-type lag structure when the maximum lag length is set to 12. The model is estimated using monthly data from 1973:1 to 1997:12.[7] Further details on descriptions and sources of the data are in the Data Appendix.

4.2.2. Identification schemes

As noted earlier, we employ four widely-cited identification schemes to identify structural shocks to monetary policy in an open economy framework. In this sub-section, we briefly review the alternative identification schemes.

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[6] Following Koray and McMillin (forthcoming), the lag length for the VARs was determined by examining the serial correlation properties for VAR residuals for alternative lag length of 3, 6, 9, 12, and 13. To check serial correlation of the residuals, Ljung-Box Q-statistics were employed.
[7] To normalize nonborrowed reserves and total reserves with a 36-month moving average, we employed data for these variables starting at 1959:1.

The first two identification schemes considered in this paper are the Christiano, Eichenbaum, and Evans and Strongin schemes which rely solely on the Choleski decomposition of the variance-covariance matrix of residuals. The main and only difference between these two schemes is that the contemporaneous casual link between nonborrowed reserves and total reserves is reversed.

For the Christiano, Eichenbaum, and Evans scheme, we consider following the Wold causal ordering for decomposition: Y, P, CP, Y*, P*, NBR, R, TR, R*, TB, and E. This ordering implies that innovations to monetary policy affect output (Y and Y*) and prices (P and P*) only with a lag, while the Federal Reserve responds to current movements in these variables.[8] This scheme also assumes that monetary policy has a contemporaneous effect on total reserves, domestic and foreign interest rates, the trade balance, and the exchange rate.[9] In addition, to identify shocks to the exchange rate, it is assumed that innovations to the exchange rate have effects on the other variables including the trade balance only with a lag. That is, as we have seen above, we placed the exchange rate after all other variables in the ordering. This ordering reflects our assumptions: (1) the Federal Reserve responds only to sustained developments in foreign exchange markets, (2) current developments in financial markets alter the exchange rate, and (3) a shock to exports and imports has contemporaneous effects on the exchange rate.

[8] We considered an over-identified system for the contemporaneous restrictions schemes in which foreign output (y*) and the foreign price level (p*) are assumed not to have contemporaneous effects on monetary policy. However, the impulse responses from all contemporaneous identification schemes are not significantly different.

[9] We also considered an alternative ordering by placing TB just prior to NBR; this ordering implies monetary policy shocks affect TB only with a lag. However, the results are essentially unchanged from our primary ordering. Hence we report results only for the primary ordering.

For the Strongin Scheme, we consider following the Wold causal ordering: Y, P, CP, Y*, P*, TR, NBR, R, R*, TB, and E. Strongin (1995) viewed nonborrowed reserves shocks as a mixture of reserve demand shocks and policy shocks. He argued that under the policy procedure followed in our sample, the level of total reserves was primarily determined by Federal Reserve accommodation of the demand for reserves. Thus, an orthogonalized innovation to monetary policy that eliminates the contemporaneous effects of a total reserve demand shock can be extracted by placing total reserves just prior to nonborrowed reserves in a standard Choleski decomposition. The rationale for placing the exchange rate after other variables is the same as in the Christiano, Eichenbaum, and Evans scheme.

Next, we consider Bernanke-Mihov's semi-structural VAR which blends the Choleski decomposition with a structural model of the reserves market. This scheme extracts monetary policy shocks from a model of the reserves market estimated from VAR residuals for nonborrowed reserves, total reserves, and the federal funds rate that are orthogonalized with respect to non-policy variables such as output, the price level, and commodity prices.

Following Bernanke and Mihov (1998), we assume a specific model of the reserve market as follows:[10]

(4.1) $$\mu_{tr} = v^d$$

(4.2) $$\mu_{br} = \beta \mu_{ffr} + v^b$$

(4.3) $$\mu_{nbr} = \phi^d v^d + \phi^b v^b + v^s$$

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[10] **Bernanke-Mihov (1997) applied this model to identify monetary policy shocks in Germany. Bagliano and Favero (1998) also employed the model as a benchmark model.**

To identify the exchange rate shock, we further assume that the foreign interest rate, the trade balance, and the exchange rate also can be specified in innovation form as:

(4.4) $$\mu_{r*} = \delta^d v^d + \delta^b v^b + \delta^s v^s + v^{r*}$$

(4.5) $$\mu_{tb} = \theta^d v^d + \theta^b v^b + \theta^s v^s + \theta^{r*} v^{r*} + v^{tb}$$

(4.6) $$\mu_e = \eta^d v^d + \eta^b v^b + \eta^s v^s + \eta^{r*} v^{r*} + \eta^{tb} v^{tb} + v^e$$

where the μ's represent the observable VAR residuals that are orthogonalized with respect to domestic output (Y), the price level (P), commodity prices (CP), foreign output (Y*), and the foreign price level (P*), and the v's are unobservable structural shocks to be identified. Subscripts tr, ffr, br, nbr, r^*, tb, and e represent total reserves, the federal funds rate, borrowed reserves, nonborrowed reserves, the foreign interest rate, the trade balance, and the exchange rate, respectively.

As we noted in Essay 1, equation (4.1) describes banks' demand for total reserves which depends only upon a demand shock, while equation (4.2) denotes the demand for borrowed reserves that depends positively on the federal funds rate. Equation (4.3) reflects the Federal Reserve reaction function. The equation implies that the Federal Reserves responds to current shocks to total reserves and borrowed reserves. Equation (4.4) implies that U.S. monetary policy contemporaneously affects the foreign interest rate, but the Federal Reserve responds to shocks to the foreign interest rate only with a lag. In other words, we assume, in light of the highly integrated financial market for the G-7 countries, that the foreign interest rate is likely to respond shocks in the U.S. reserve market but the Federal Reserve will respond only to sustained developments in foreign financial markets. Equation (4.5) denotes that the trade balance depends upon the reserve market shock and the foreign interest rate shock, while equation (4.6) indicates the exchange rate depends upon shocks to the reserve market variables, the foreign interest rate, and the trade balance. Consequently, we assume that the trade balance responds only with a lag to a shock to the exchange rate.

Combining the market for reserves with equations (4.4) to (4.6), we can write the reduced form relationship between the VAR residuals μ and the structural shocks *ν* as:

(4.7) $\mu = G\mu + A\nu$

or

(4.8) $\mu = (I-G)^{-1}A\nu$

or in matrix form

$$\begin{pmatrix} \mu^{tr} \\ \mu^{nbr} \\ \mu^{ffr} \\ \mu^{r*} \\ \mu^{tb} \\ \mu^{e} \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ \phi^d & 1 & \phi^b & 0 & 0 & 0 \\ \frac{1}{\beta}(1-\phi^d) & -\frac{1}{\beta} & -\frac{1}{\beta}(\phi^b+1) & 0 & 0 & 0 \\ \delta^d & \delta^s & \delta^b & 1 & 0 & 0 \\ \theta^d & \theta^s & \theta^b & \theta^{r*} & 1 & 0 \\ \eta^d & \eta^s & \eta^b & \eta^{r*} & \eta^{tb} & 1 \end{pmatrix} \begin{pmatrix} \nu^d \\ \nu^s \\ \nu^b \\ \nu^{r*} \\ \nu^{tb} \\ \nu^{e} \end{pmatrix}$$

This model has twenty one unknown parameters to be estimated from the exact same number of residual variance and covariances; the model is just-identified. However, it might be argued that the foreign interest, the trade balance, and the exchange rate do not respond to contemporaneous shocks to total reserves and borrowed reserves. Thus, we also considered an over-identified system with $\delta^d = \theta^d = \eta^d = \delta^b = \theta^b = \eta^b = 0$. But the results were essentially unchanged from the just-identified system.[11]

[11] We also considered an over-identified system in which we assume shocks to reserve market and the foreign interest have no contemporaneous effects on the trade balance, i.e. $\theta^d = \theta^b = \theta^s = \theta^{r*} = 0$. But the results are essentially unchanged from the just-identified model.

To estimate this system, we employ a two-step efficient Generalized Method of Moments (GMM) procedure suggested by Bernanke and Mihov (1998). Specifically, we, first, estimate the VAR system by Ordinary Least Squares (OLS). Next, we match the second moments implied by the structural model (4.7) to the estimated covariance matrix of VAR residuals.

The last scheme considered in this paper is the long-run restrictions approach. Instead of imposing contemporaneous restrictions on model variables, this identification scheme employs less controversial long-run neutrality assumptions. In this paper, to identify shocks to monetary policy, we assume shocks to monetary policy have no effect on real variables such as domestic output, foreign output, the relative price of commodities (commodity prices deflated by the U.S. price level, RCP), the trade balance, and the real exchange rate (=E-(P-P*), RE) in the long-run. But, monetary policy shocks are allowed to affect the foreign price level, commodity prices, and total reserves in the long-run. As noted in introduction of this paper, these assumptions reflect familiar implications of open economy macroeconomic models. We further assume that monetary policy shocks have no effects on the federal funds rate and the foreign interest rate in the long-run. Following a positive shock to monetary policy, the interest rates initially fall because of liquidity effects. However, in the long-run equilibrium, as domestic and foreign prices rise and real money balances return to initial levels, the interest rates also rebound to their initial level.

We can implement these assumptions in a Choleski decomposition of long-run relations by specifying variables in a first difference form with following order: ΔY, ΔRCP, ΔY^*, ΔR, ΔR^*, ΔTB, ΔRE, ΔNBR, ΔTR, ΔP^*, and ΔCP. With the model estimated in first difference form, we can easily implement long-run restrictions in a VAR. As illustrated in Essay 1, the long-run effect of a shock to monetary policy on the level variables is the cumulative sum of the relevant part of the moving average representation. Consequently, we impose the restrictions by placing the variables which are not affected by shocks to monetary policy in the long-run just prior to a monetary policy variable. In addition, although the model does not explicitly include the U.S price level as a separate variable, the impulse responses of the price level can be recovered from the difference in impulse responses between real commodity prices and commodity prices. Similarly, the impulse responses of the nominal exchange rate also can be recovered by using the impulse responses of the real exchange rate, the price level, and the foreign price level.

To identify shocks to the exchange rate, we estimate a slightly different specification of VAR model, since the previous specification does not include the nominal exchange rate for which we want to identify shocks. We consider the following specification and order in a Choleski decomposition of long-run relations: ΔY, ΔRCP, ΔY^*, ΔR, ΔR^*, ΔTB, ΔE, ΔNBR, ΔTR, ΔP^*, and ΔCP. Notice that, in this specification, the real exchange rate (ΔRE) in the previous specification for identifying monetary policy shocks is replaced by the nominal exchange rate (ΔE). Therefore, it is assumed that a shock to the exchange rate has no effects on output, the relative price of commodities, interest rates, and the trade balance in the long-run, but it is allowed to affect nonborrowed reserves, total reserves, and the domestic and foreign price levels in the long-run. Typically, a shock to the exchange rate, which can be viewed as a negative shock to aggregate demand, affects the trade balance, output, the price level, and the interest rate in the short-run: these variables are expected to fall below their initial levels. However, in the long-run, as real money balances rise due to the fall in the price level, the trade balance, output, and the interest rate return to their initial levels. The price level is expected to be permanently higher.

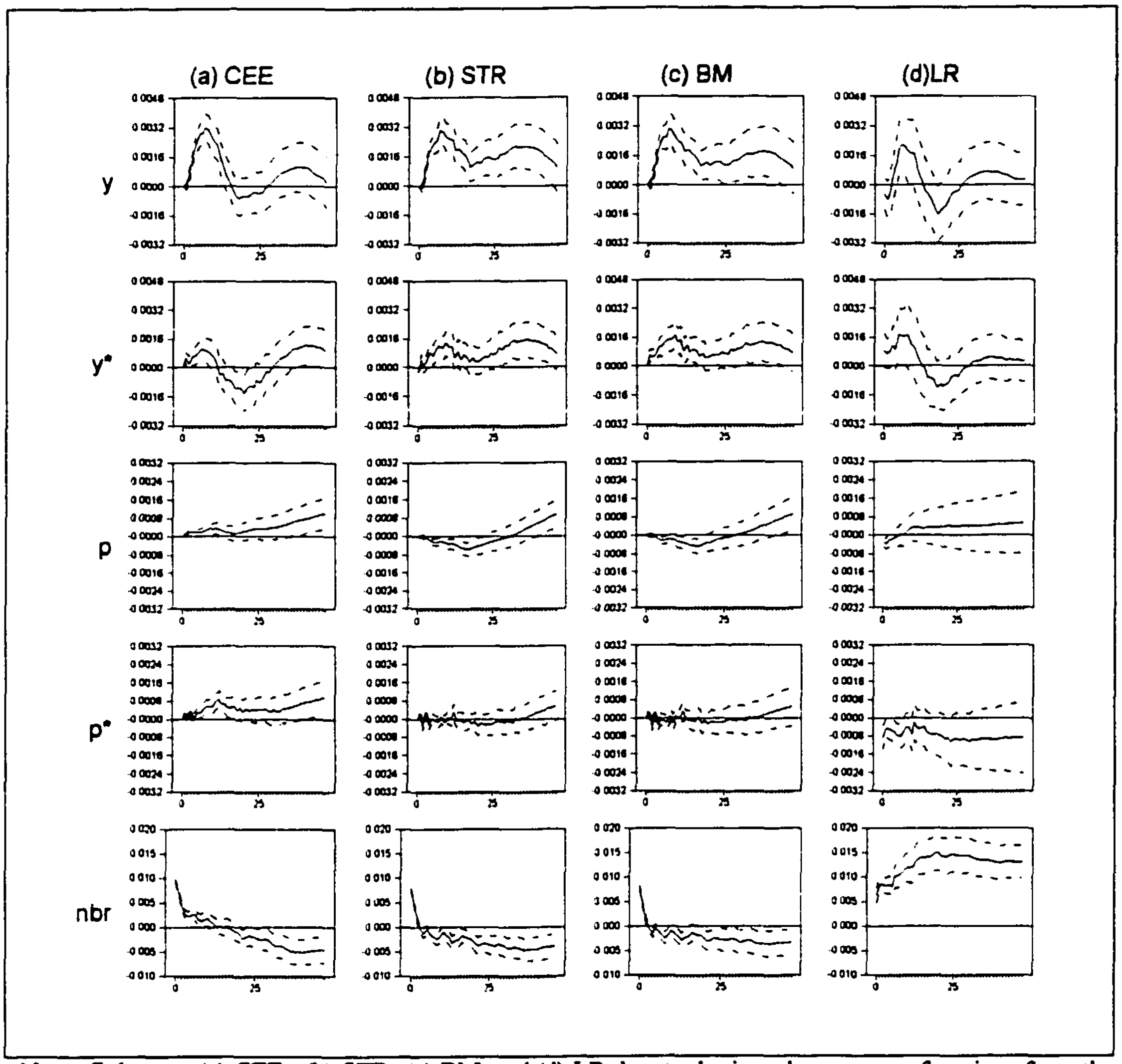
4.3. Empirical Results

In this section, we first investigate the effects of monetary policy shocks in the aforementioned four identification schemes. Then, we compare the effects across identification schemes by plotting confidence bands for a particular identification scheme with point estimates for another identification scheme. In addition, we briefly examine the effects of shocks to the exchange rate.

4.3.1. Comparing Impulse Responses across Identification Schemes: Shock to Monetary Policy

Figure 4.1 plots the impulse responses from the alternative identification schemes for domestic and foreign output, price levels, and nonborrowed reserves.[12] The first column of this figure presents the effects of monetary policy shocks identified using the Christiano, Eichenbaum, and Evans scheme. The remaining columns are results for the Strongin, Bernanke-Mihov, and long-run restrictions approach. In each diagram, the solid lines represent the point estimates, while the dashed lines denote one

[12] The impulse responses for the federal funds rate, the foreign interest rate, the nominal and real exchange rates, and the trade balance are presented in figure 4.2. However, in order to conserve space, we do not report the responses for total reserves and commodity prices. The responses of commodity prices from the contemporaneous restrictions schemes are significant and positive for longer horizons, although the responses are weaker for shorter horizons. The responses from the long-run restrictions scheme indicate no effects, however. For nonborrowed reserves, the responses from the Christiano-Eichenbaum-Evans and long-run restrictions schemes show significant and positive initial effects, while the responses from the Strongin and Bernanke-Mihov schemes are positive but insignificant. The responses from all schemes except the long-run restriction scheme are negative for longer horizons, although the upper bounds are close to zero after approximately 40-45 months.



Note: Columns (a) CEE, (b) STR, (c) BM, and (d) LR denote the impulse response functions from the Christiano, Eichenbaum, and Evans, Strongin, Bernanke and Mihov, and Long-run restrictions schemes, respectively. Also, y, y^*, p, p^*, and nbr refer to U.S. output, foreign output, the U.S. price level, the foreign price level, and nonborrowed reserves.

Figure 4.1
Impulse Response Functions: U.S. Output, Foreign Output, the U.S. Price Level, the Foreign Price Level, and Nonborrowed Reserves

standard error confidence bands around the point estimates. The standard errors are generated from Monte Carlo simulations with 1,000 replications. In general, the point estimates from the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes are similar in pattern, while the impulse responses from the long-run restrictions approach are quite different from others. We note that the impulse response functions of the long-run restrictions approach may be estimated less precisely than those of others. This can be seen informally since the confidence bands for the long-run restrictions approach are much wider than those of others.

For output, the impulse responses from the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes are similar. However, the impulse responses from the Christiano, Eichenbaum, and Evans scheme indicate relatively shorter lasting effects of monetary policy shocks compared to the other two schemes. The confidence bands for the scheme span zero after approximately 13-14 months. The responses from the Strongin and Bernanke-Mihov scheme reveal persistent monetary policy effects on output in that the lower bounds for these schemes include zero after about 43-44 months. However, the confidence bands for the long-run restrictions approach span zero for most horizons except 4-9 months, indicating that monetary policy shocks have little effect on output. Moreover, although the point estimates for first two periods are negative (which is contradictory to our prediction based on open economy macroeconomic models), the confidence bands for these periods span zero indicating these effects are not significant.

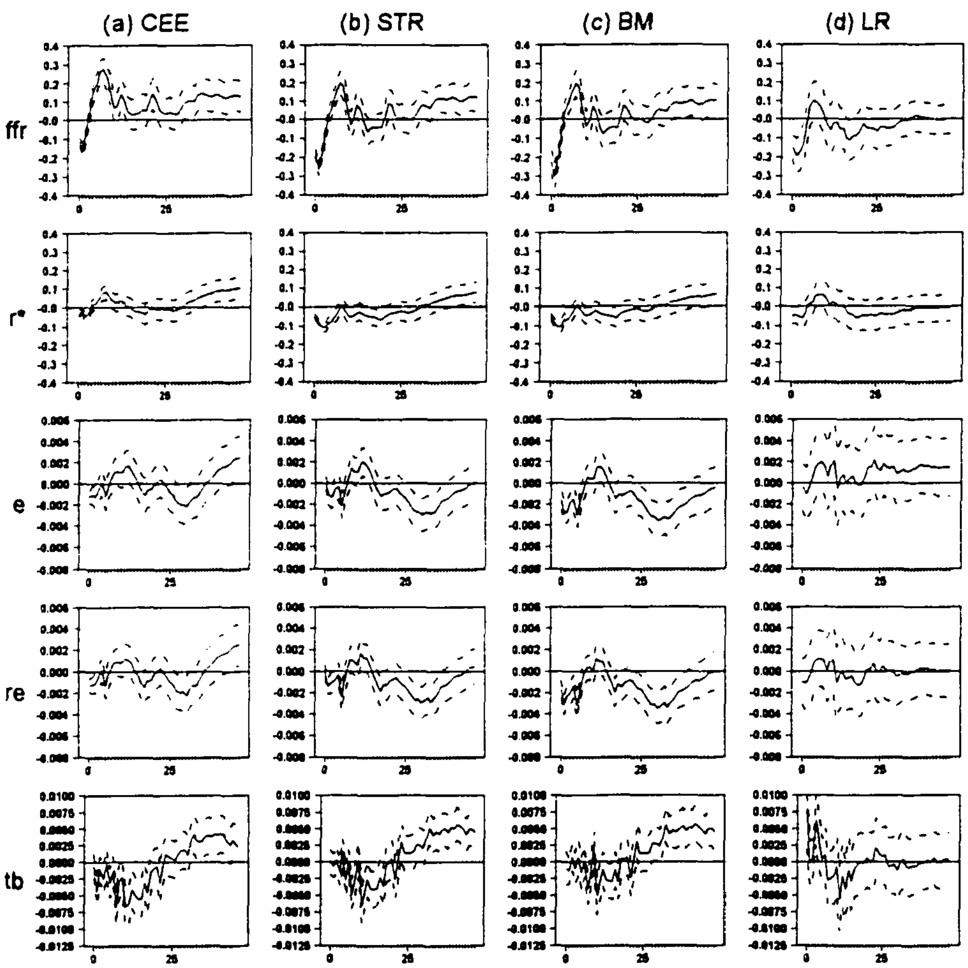
The responses of foreign output for the identification schemes using contemporaneous restrictions are similar to the results for domestic output, although the magnitude is smaller and the effects are shorter lasting. The confidence bands for the Christiano, Eichenbaum, and Evans scheme span zero after approximately 8 months, while the bands for the Strongin and Bernanke-Mihov schemes include zero after about 13-15 months. In contrast to domestic output, the initial responses for the long-run restrictions approach are positive, although the confidence bands span zero.

In the case of the domestic and foreign price levels, the point estimates for the Christiano, Eichenbaum, and Evans scheme are always positive and persistently rise, although the confidence bands include zero for considerable periods of time. However, the point estimates for the Strongin and Bernanke-Mihov schemes reveal some degree of a 'price puzzle' in that the price level declines following a positive shock to nonborrowed reserves, despite the presumption that an increase in nonborrowed reserves represents an expansionary monetary policy. The point estimates for the long-run restrictions scheme also show a similar 'price puzzle', but the confidence bands include zero after 1 month. For the foreign price level, almost the same results are emerged. The only big difference is the point estimates for the long-run restrictions. It seems to be problematic in that the point estimates for the long-run restrictions lie below zero over the first two years. The confidence bands for the Strongin, Bernanke-Mihov, and long-run restrictions scheme include zero for almost all horizons.

For nonborrowed reserves, the point estimates from all schemes but the long-run restrictions approach reveal immediate, sharp and significant rises in nonborrowed reserves. Within several months, the estimates drop actually below the initial level and remain there for the entire reported horizons.[13] However, the impulse responses for the long-run restrictions approach reveal persistent effects on nonborrowed reserves. This is not surprising in that the long-run neutrality assumptions are made only for the domestic and foreign output, relative price of commodities, the trade balance, and the real exchange rate.

[13] If the horizons for the impulse responses are extended, the confidence bands for these identification schemes include zero after approximately 50 months.

In figure 4.2 the impulse responses from the alternative identification schemes for the federal funds rate, foreign interest rate, nominal and real exchange rates, and the trade balance are presented. For all schemes, the initial responses of the federal funds rate to a positive monetary policy shock are strongly negative, indicating a strong liquidity effect. However, after approximately 3 months, the point estimates for the



Note: Columns (a) CEE, (b) STR, (c) BM, and (d) LR denote the impulse response functions from the Christiano, Eichenbaum, and Evans, Strongin, Bernanke and Mihov, and Long-run restrictions schemes, respectively. Also, ffr, r*, e, re, and tb refer to the federal funds rate, foreign interest rate, nominal exchange rate, real exchange rate, and trade balance.

Figure 4.2
Impulse Response Functions: the Federal Funds Rate, Foreign Interest Rate, Nominal and Real Exchange Rates, and Trade Balance

Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes rise sharply and are above zero for a while, possibly due to expected inflation, output, and price level effects, and return to the initial level after approximately 8-9 months. However, the lower bounds of these identification schemes, especially for the Christiano, Eichenbaum, Evans and Strongin schemes, lie above zero for the longer horizons. The impulse responses for the long-run restrictions approach are different; the responses return to the initial level after 3 months without rising above zero. Within a year, confidence bands for all schemes include zero. As for the magnitude of the point estimates of liquidity effect, the Bernanke-Mihov approach indicates stronger effects. For the foreign interest rate, the patterns of response are very similar to those for the federal funds rate, although the magnitudes are much smaller. However, unlike the federal funds rate, the responses of the foreign interest rate for the Strongin and Bernanke-Mihov schemes no longer rise above zero after initial drops.

For the nominal exchange rate, the point estimates from all schemes report initial depreciation following an expansionary monetary policy shock. This result is consistent with other previous research including Eichenbaum and Evans (1995) and Koray and McMillin (forthcoming). The confidence bands for the Christiano, Eichenbaum, and Evans scheme spans zero after approximately 3 months, while the confidence bands for the Bernanke and Mihov schemes include zero after about six months.[14] However, the confidence bands for the Strongin scheme include zero even for

[14] When we employed a lag length of 6, we found the impulse responses of the exchange rate to monetary policy shocks are similar to those of Eichenbaum and Evans (1995). However, the confidence bands for the VAR are much wider. Also, the results of Ljung-Box Q-tests show that the 6 month lag length yields serial correlation in some of equations in the model. We note that this is problematic in that an assumption of the identification scheme used here is that VAR residuals are white noise.

the first several months. This indicates that the scheme gives less evidence of initial depreciation than do the other contemporaneous restrictions schemes. The point estimates for the long-run restrictions approach also indicate the initial depreciation, although the confidence bands for the approach always include zero. The responses of the real exchange rate to an expansionary monetary policy shock are very similar to those of the nominal exchange rate regardless of identification scheme. This is not surprising in that the effects on domestic and foreign prices are small as shown in Figure 4.1.

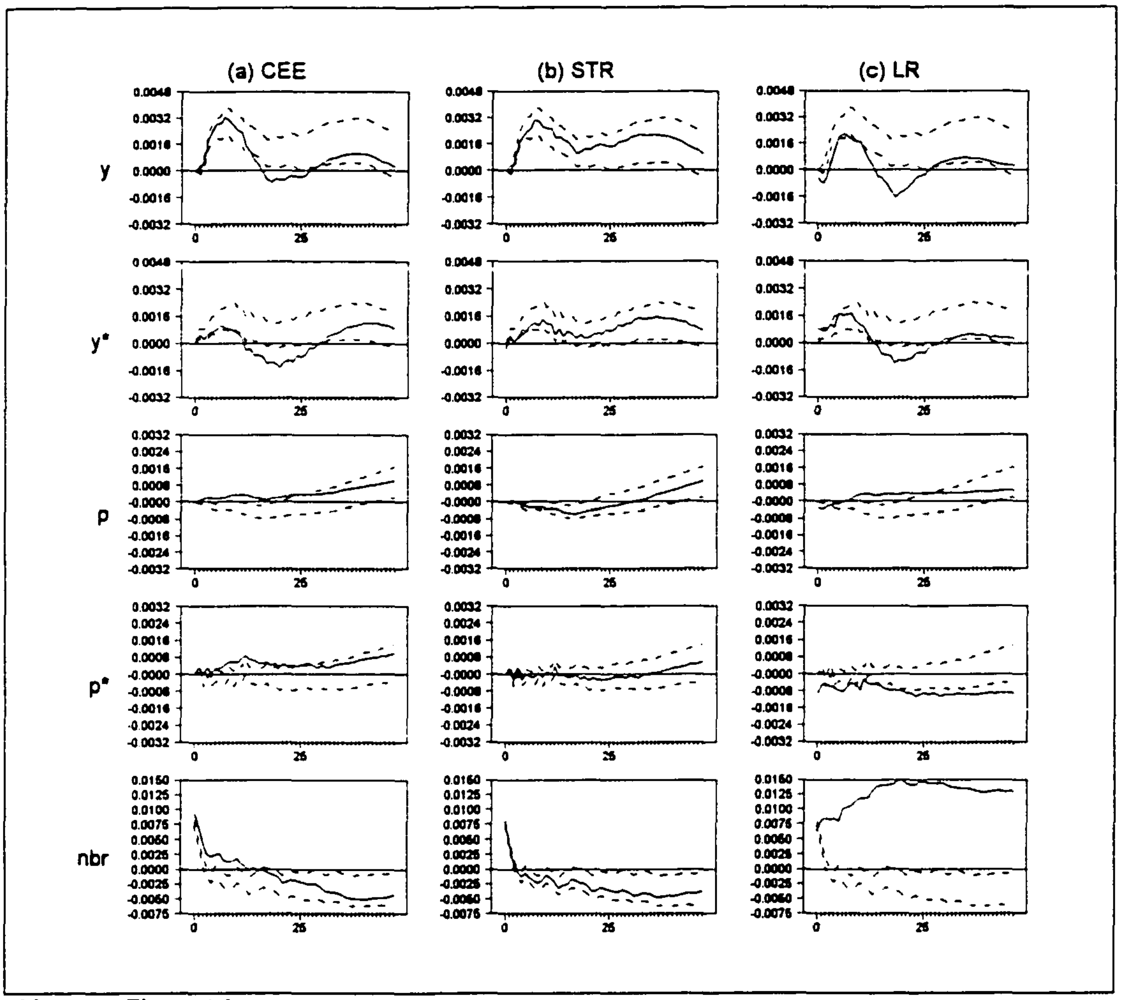
The responses of the trade balance to monetary policy shocks for the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes are generally negative for about 20 months and rebound above zero. For example, for the Christiano, Eichenbaum, and Evans scheme, the point estimates generally remain below zero for the first 22 months, but they rebound above zero after that. Although the point estimates for the Strongin and Bernanke-Mihov schemes show similar response patterns, their effects are slightly weaker than those of the Christiano, Eichenbaum, and Evans scheme. The confidence bands of the Strongin and Bernanke-Mihov approaches include zero for the first half of horizons. These results from the contemporaneous restrictions approaches are inconsistent with the prediction of traditional open economy models in which the trade balance improves following expansionary monetary policy shocks. One explanation of this phenomenon is that it results primary from the asymmetry in the effects of monetary policy shocks on domestic and foreign output and partially from the J-curve effect. Since the responses of domestic output to monetary policy shocks are much greater than those of foreign output for about first 20 months, the increase in exports is less than the increase in imports, indicating temporal deterioration in the trade balance. Moreover, the J-curve effects also may lead to deterioration in the trade balance. As the asymmetry in output effect is eliminated, the trade balance starts to improve. In contrast, the point estimates for the long-run restrictions scheme indicate sharp, strong, and positive initial effects, although the confidence bands span zero after 5 months. The explanation for this difference in responses between for the contemporaneous restrictions schemes and for the long-run restrictions scheme is straightforward. As we have seen previously, monetary policy shocks have little effect on domestic and foreign output in the long-run restrictions approach. Hence, the type of asymmetric output effects in the contemporaneous restrictions schemes does not appear in the long-run restriction scheme.

Overall, the empirical results indicate that the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes generate similar impulse responses. However, it is worth noting that the magnitude and timing of the point estimates differ across these schemes. The responses from the long-run restrictions approach are sometimes quite different from others. We investigate this point by plotting the point estimates for other identification schemes with the confidence bands for the Bernanke-Mihov scheme. This provides additional information on whether the differences in magnitude and timing of responses across alternative identification schemes are substantial.

Figure 4.3 plots the confidence bounds from the Bernanke-Mihov scheme and point estimates from the other schemes for output, foreign output, the domestic and foreign price levels, and nonborrowed reserves. For U.S. output, the point estimates from the Strongin scheme essentially lie within the confidence intervals. The point estimates of the Christiano, Eichenbaum, and Evans scheme drop below the lower bound after 12 months and remain there for a year, indicating significant shorter lasting effects compared to the Bernanke-Mihov scheme. The point estimates for the long-run restriction approach lie on or slightly below the low bound for the first 10 months, but the estimates lie below the lower bound for the periods of 11-26 months. Over time, the estimates lie within the intervals. In the case of foreign output, the point estimates for the Christiano, Eichenbaum, and Evans and Strongin schemes reveal similar patterns to domestic output. The estimates for the long-run restrictions scheme initially lie within the intervals and drop below the lower bound for a while.

For the price level, the point estimates for the Christiano, Eichenbaum, and Evans scheme lie above the upper bound for the first 25 months and within the bounds thereafter, indicating somewhat stronger effects on the price level. The point estimates for the Strongin scheme lie within the bands. The point estimates for the long-run restrictions scheme initially lie below the lower bound, but lie above the upper bound for the period from 7 to 25 months. Over time, the estimates lie within the confidence bands. For the foreign price level, the point estimates for the Strongin scheme lie within the intervals, while the point estimates for the long-run restrictions lie on or below the intervals. The point estimates for the Christiano, Eichenbaum, and Evans scheme lie above the upper bound for first 26 months and within the intervals thereafter.

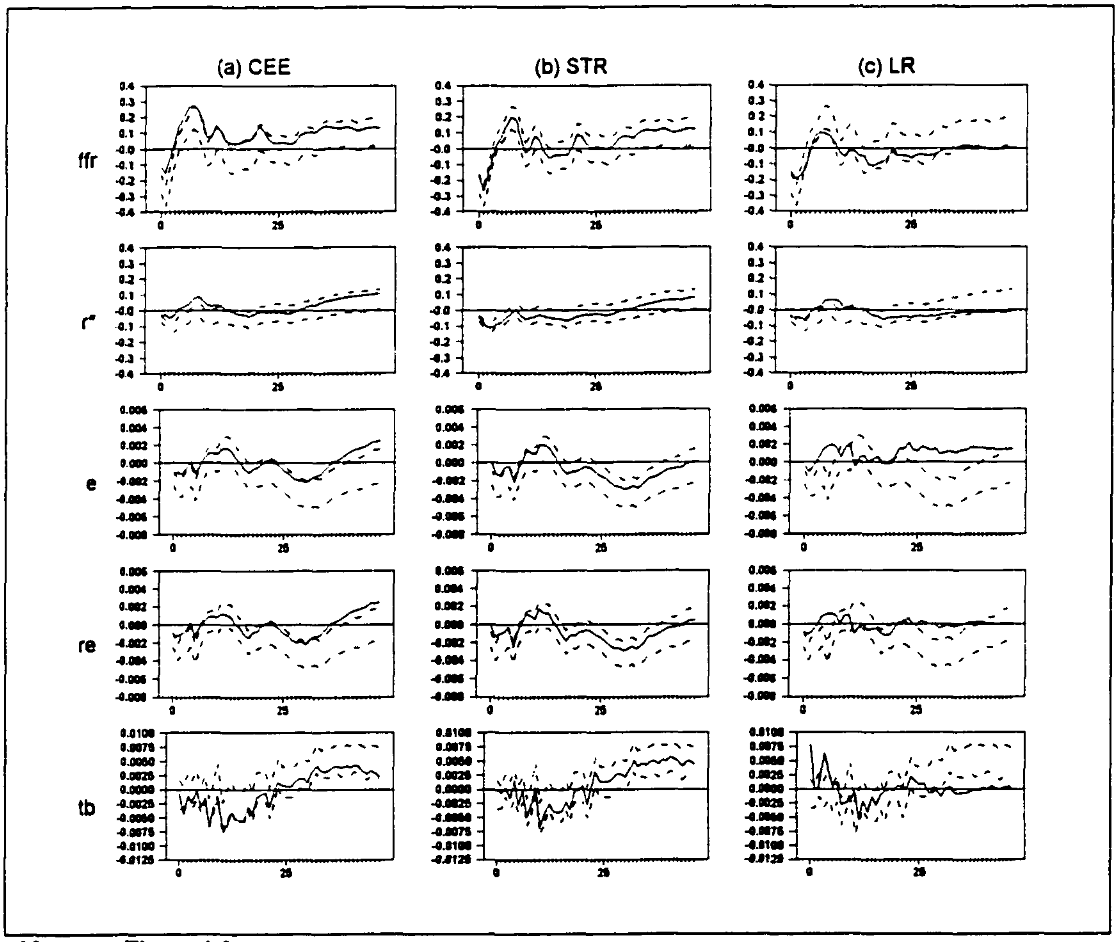
For nonborrowed reserves, we observe that the point estimates for the Christiano, Eichenbaum, and Evans scheme lie above the confidence bands for the first 16 months, while the estimates for the Strongin scheme lie within the intervals for the entire horizons. The point estimates for the long-run restrictions approach indicate big differences; the estimates lie above the upper bounds for the entire horizon.



Note: see Figure 4.1.

Figure 4.3
Bernanke-Mihov Restrictions Confidence Intervals and Point Estimates from Other Identification Procedures: U.S. Output, Foreign Output, U.S. Price Level, the Foreign Price Level, and Nonborrowed Reserves

Figure 4.4 reports the confidence bounds for the Bernanke-Mihov procedure and the point estimates from the other approaches for the federal funds rate, the foreign interest rate, the nominal and real exchange rate, and the trade balance. In case of the federal funds rate, we observe that the point estimates from the Christiano, Eichenbaum, and Evans scheme are slightly above the upper bound during first 6 months, but are within the confidence bands thereafter. The point estimates from the long-run restrictions approach lie slightly above or on the lower bound, while the point estimates from the Strongin scheme always lie within the intervals. For the foreign interest rate, similar response patterns are found. Only exception is that the point estimates for the long-run restriction scheme lie slightly above or on the upper bound for the first several months.



Note: see Figure 4.2.

Figure 4.4
Bernanke-Mihov Restrictions Confidence Intervals and Point estimates from other Identification Procedures: the Federal Funds Rate, Foreign Interest Rate, Nominal and Real Exchange Rates, and Trade Balance

In the case of the nominal and real exchange rate, the point estimates for the Christiano, Eichenbaum, and Evans and Strongin schemes lie within the confidence bands, although the estimates for the Christiano, Eichenbaum, and Evans scheme lie slightly above the upper bounds for the last 10 months. The point estimates for the nominal exchange rate from the long-run restriction approach lie above the upper bound, except the periods of 7-20 months after shocks. Also, the estimates for the real exchange rate from the long-run restrictions approach lie above the upper bound for two periods, 1 to 7 months and 22 to 40 months.

For the trade balance, the point estimates for the Christiano, Eichenbaum, and Evans scheme marginally lie above the lower bound for the first 18 months and lie within the confidence bands thereafter. The point estimates for the Strongin scheme lie within the intervals for the entire horizon. The estimates for the long-run restrictions approach lie above the upper bounds for the first month, but they are within the intervals for the periods, 5-26 months. Over time, the estimates lie below the lower bound.

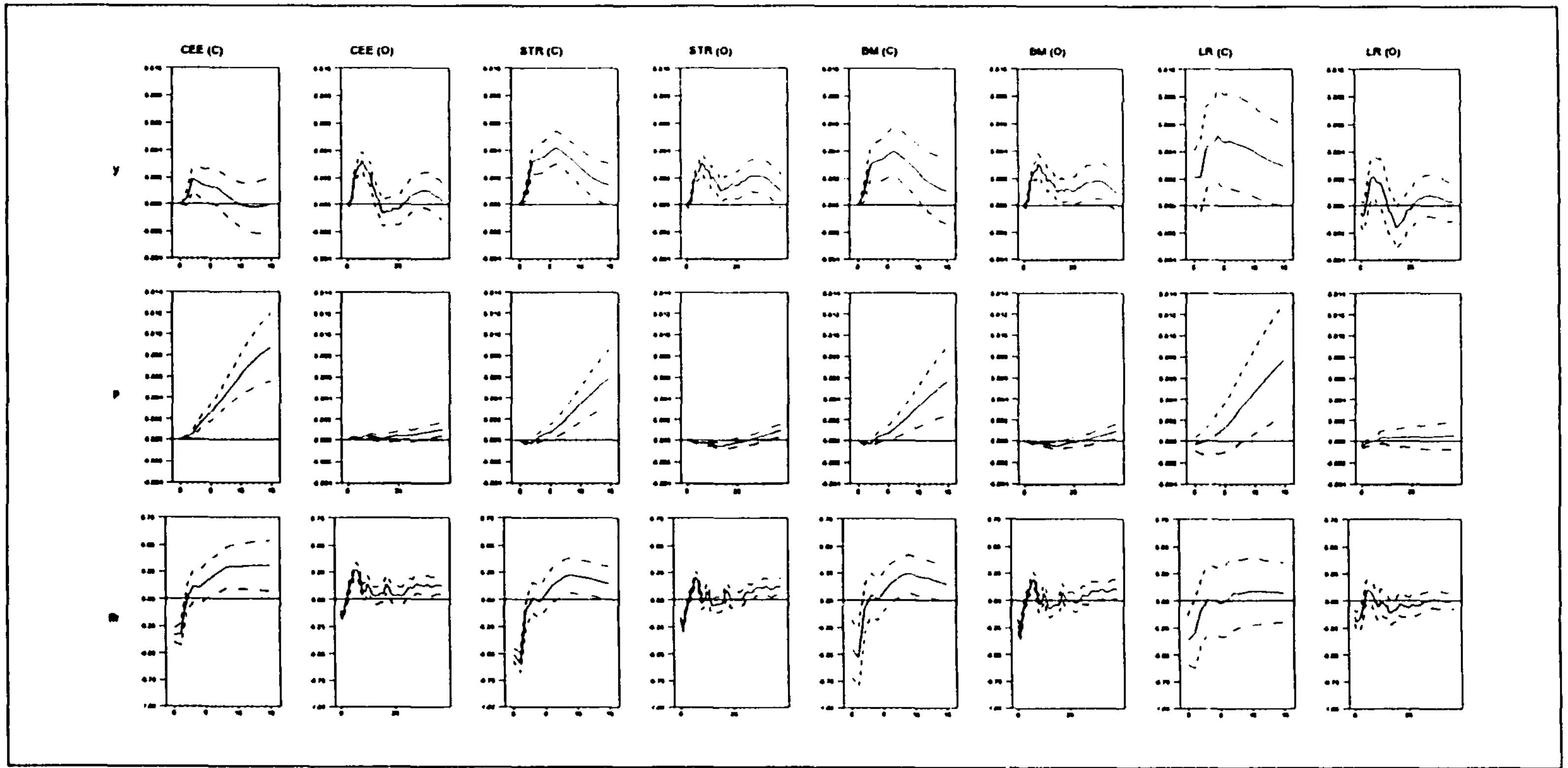
In sum, we observe that the impulse responses for the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes give generally reasonable results for output, interest rate, the trade balance and the exchange rate. The shocks to monetary policy lead to positive but transitory rises in output, sharp initial falls in the interest rate, depreciation in the exchange rate, and initial deterioration and subsequent improvement in the trade balance. However, the responses of the price level for the Strongin and Bernanke-Mihov schemes seem to be problematic in that these schemes generate the well-known 'price puzzle'. By contrast, the responses of the price level for the Christiano, Eichenbaum, and Evans scheme do not generate the 'puzzle', although the effects are much weaker than the closed economy model illustrated in Essay 1. The major difference in results for the long-run restrictions approach compared to other schemes is that the nonborrowed reserved shock can be interpreted as a permanent shock to the level of nonborrowed reserves. The responses of nonborrowed reserves continuously rise after shock. Also, the statistical uncertainty about responses is quite large in that the confidence bands for the long-run approach are much wider than those for other schemes.

As for the sensitivity of effects of monetary policy shocks across alternative identification schemes, the impulse responses for the Christiano, Eichenbaum, and Evans scheme reveal a relatively shorter lasting effect for output than for the Strongin and Bernanke-Mihov schemes. Also, the impulse responses for the Christiano, Eichenbaum, and Evans scheme reveal weaker initial effects on the exchange rate and stronger effects on the price level, compared to the Strongin and Bernanke-Mihov schemes. Finally, the impulse responses, especially for the trade balance and nonborrowed reserves, from the long-run restrictions approach are different from other schemes.

Up to now, we have discussed the effects of monetary policy shocks for the open economy framework. A natural question is how do the impulse response results from the open economy models compare to those from the closed economy models which are described in Essay 1? To answer the question, we now compare the impulse responses of output, the price level, and the federal funds rate from the closed economy model and from the open economy model. Although differences in data frequencies and sample periods constrain direct comparisons of impulse responses, this exercise will provide a big sketch of differences and similarities in patterns of impulse response functions from both frameworks. Recall that the closed economy models in Essay 1 are estimated using quarterly data for the period 1965:1-1997:4, while the open economy models in this paper are fitted using monthly data for the period 1973:1-1997:12.

Figure 4.5 plots the impulse responses of output, the price level, and the federal funds rate across the alternative identification schemes from both frameworks. The first column of this figure presents the effects of monetary policy shocks identified using the Christiano, Eichenbaum, and Evans scheme in a closed economy framework. The second column shows the effects of monetary policy shocks identified using the same identification scheme in an open economy framework. The remaining columns are analogous results for the Strongin, Bernanke-Mihov, and long-run restrictions approaches.

Overall, the hump-shaped patterns of responses for output in the open economy models are similar to those in the closed economy models. However, for the Strongin, Bernanke-Mihov, and long-run restrictions approaches, the effects of monetary policy shocks on output are somewhat weaker than those in the closed economy models. In contrast, for the Christiano, Eichenbaum, and Evans scheme, the effects are slightly greater. As for the timing in restoring the initial level after shocks, the responses for the contemporaneous restrictions schemes from the open economy framework are roughly similar to those from the closed economy framework. However, the timing for the long-run restrictions approach is quite different. For the closed economy framework, it takes



Note: See Figure 4.1. Also, letters C and O in brackets [ ] represent the results from a closed economy model and open economy model. Horizontal scale for the closed economy model [C] is in quarters, while the horizontal scale for the open economy model [O] is in months.

Figure 4.5
Impulse Response Functions: Closed Economy Model [C] versus Open Economy Model [O]

approximately 3.5 years for output to return to the initial level, but, for the open economy framework, there are no significant effects on output except 4-9 months.

In the case of the price level, the difference between impulse responses from the two frameworks seems to be clear. The responses from the closed economy models are positive and significant for most horizons, indicating that there is no significant 'price puzzle'. However, the responses from the open economy models are problematic. The responses are clearly weaker than the closed economy counterparts and show some degree of the 'price puzzle' for the Strongin and Bernanke-Mihov schemes. Although the responses of the price level for the Christiano, Eichenbaum, and Evans scheme do not generate the 'puzzle', the effects are much weaker than the closed economy model.

We conclude this subsection by comparing the liquidity effects from the open economy framework and from the closed economy framework. Regardless of identification schemes considered in this paper, the liquidity effects for the open economy model are clearly weaker than for the closed economy model; the effects are about one half of the effects from the closed economy model. However, in spite of the differences in magnitude, both frameworks generate significant liquidity effects for all identification schemes.

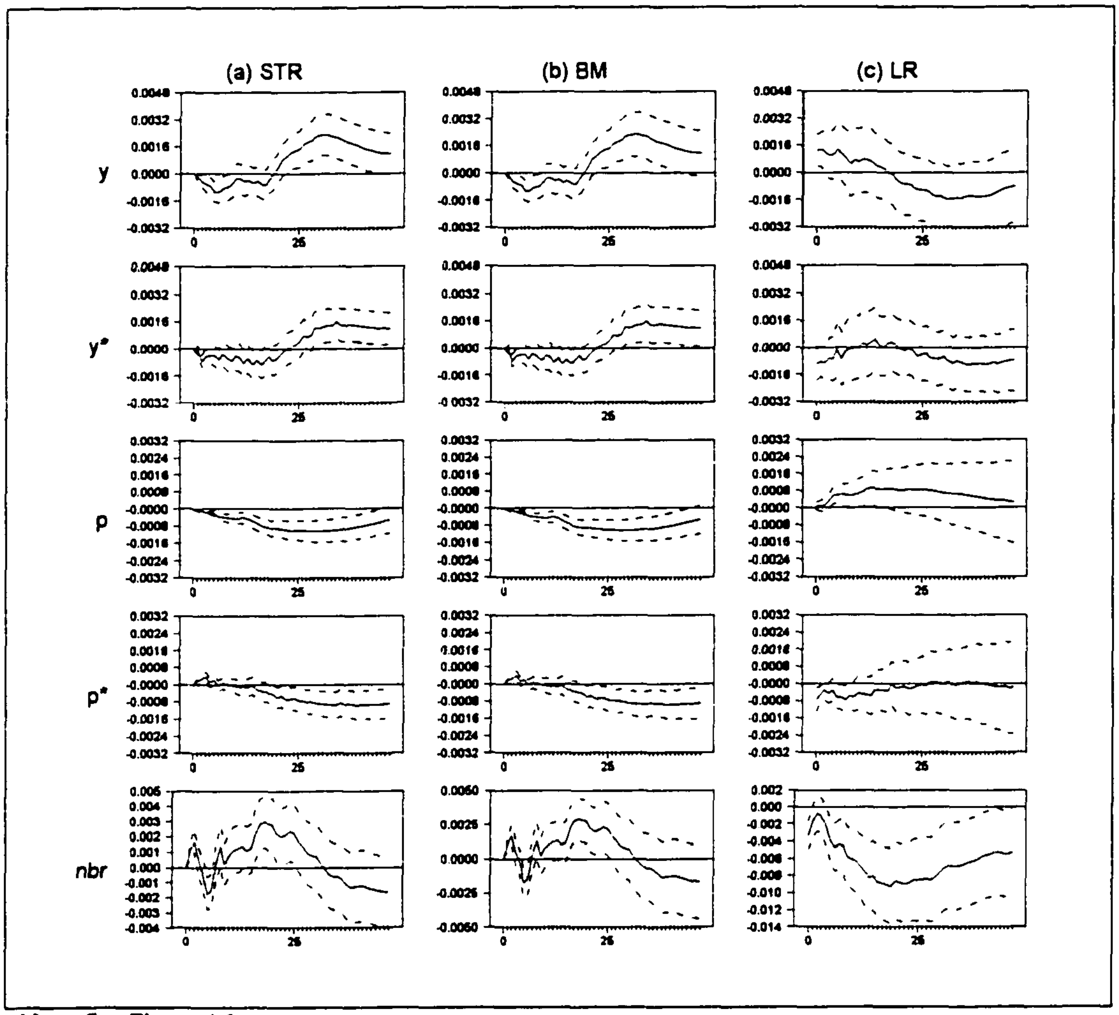
4.3.2. Comparing Impulse Responses across Identification Schemes: Shock to the Exchange Rate

In this subsection, we investigate the effects of shocks to the exchange rate on output, prices, interest rates, the exchange rate, and the trade balance. However, before we report our empirical results, one point is worth noting: the identified shocks to the exchange rate for the Christiano, Eichenbaum, and Evans and Strongin schemes are exactly the same. This is because, in our Choleski ordering, the only difference between these two schemes is the causal relationship between total reserves and nonborrowed reserves. Consequently, it does not affect the identification of shocks to the exchange rate since the exchange rate is placed after the reserve market variables in the ordering.[15] Hence, we only report the impulse responses for the Strongin, Bernanke-Mihov, and long-run restrictions schemes.

In Figure 4.6, we plot the impulse responses for U.S. and foreign output and price, and for nonborrowed reserves across three alternative identification schemes. Overall, two points are worth noting. First, the impulse responses from the Strongin and Bernanke-Mihov approaches are quite similar. However, the responses from the long-run restrictions scheme are different from others. Second, the responses are, in general, reversed in pattern compared to the responses to monetary policy shocks, although the magnitude and timing are different.

A positive shock to the exchange rate, which is identified by using either the Strongin scheme or the Bernanke-Mihov Scheme, has initial significant but transitory negative effects on U.S. output. The point estimates for these schemes are negative for the first 19 months and rebound above zero thereafter. Finally, the confidence bands include zero about 42 months after shock, indicating no long-run effects. In the case of foreign output, the responses for the Strongin and Bernanke-Mihov schemes reveal similar patterns, although the effects are weaker. However, the responses of U.S. output from the long-run restriction scheme are quite different. The point estimates are initially positive and significant, but the point estimates drop below zero after approximately 17 months. The confidence bands span zero for almost all horizons. The responses of

[15] For further discussion of this issue, see Keating (1994) and Christiano, Eichenbaum, Evans (1998).



Note: See Figure 4.1.

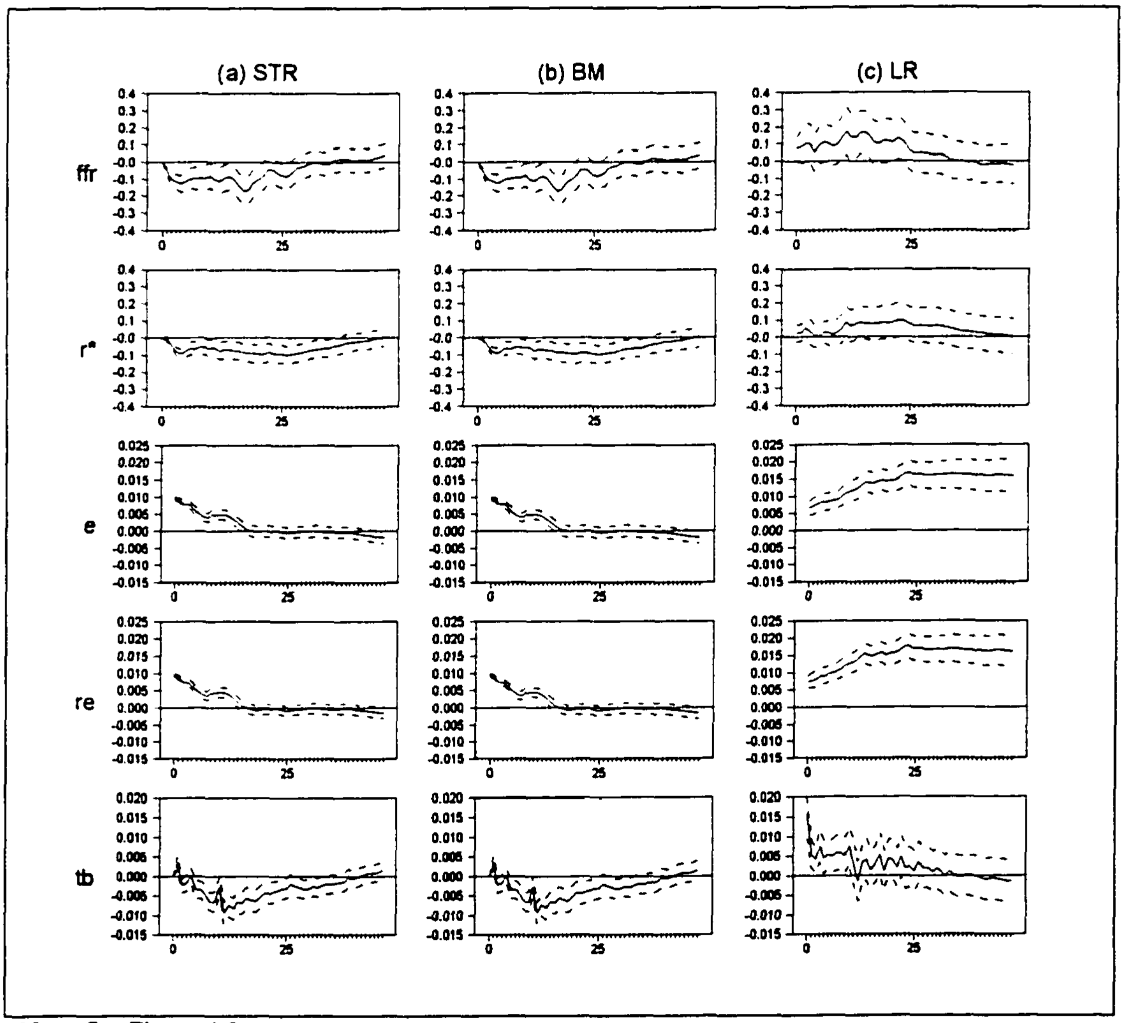
Figure 4.6
Impulse Response Functions: Shocks to Exchange Rate
U.S. output, Foreign Output, the U.S. Price Level, the Foreign Price Level, and Nonborrowed reserves

foreign output for the long-run restrictions approach are different from the case of U.S. output. The point estimates are negative and return to the initial level after 9 months, although the confidence bands always include zero.

For the price level, the responses for the U.S. price level from the Strongin and Bernanke-Mihov schemes are always negative. However, a problematic feature of the responses from the long-run restrictions approach can be pointed out: following a positive shock to the exchange rate, the responses from the long-run restrictions approach are always positive. But, it is inconsistent with the prediction of open economy macroeconomics in which, following a positive shock to the exchange rate (hence a negative aggregate demand shock), the price level eventually falls rather than rises. The responses of the foreign price level for the Strongin and Bernanke-Mihov scheme are initially positive, although they drop below zero after approximately 6 months. The responses for the long-run restrictions approach are initially negative and over time return the initial level.

In the case of nonborrowed reserves, the results for the contemporaneous restrictions approaches and for the long-run restrictions approach are contradictory. The confidence bands of the Strongin and Bernanke-Mihov procedure span zero for almost all horizons except the first two months, indicating that there is no substantial effect on nonborrowed reserves. This implies that the Federal Reserve does not respond strongly to the exchange rate shocks. However, the long-run restrictions approach generates a very different result. The impulse responses for the approach are negative and significant for almost all horizons. It suggests that the Federal Reserve responds to a positive exchange rate shock by decreasing nonborrowed reserves for a substantial period of time. This is problematic in two points. First, if the Federal Reserve is interested in offsetting a positive shock to exchange rate, it would increase rather decrease nonborrowed reserves. Decreasing nonborrowed reserves in this fashion might worsen the situation. Second, since the Federal Reserve typically views the aggregate demand shocks as transitory shocks, its prolonged response to the exchange rate shocks is unrealistic.

In Figure 4.7, the impulse responses for the federal funds rate, foreign interest rate, nominal and real exchange rates, and the trade balance are presented. For the interest rate, the response of U.S. and foreign interest rates are negative for the Strongin and Bernanke-Mihov schemes, although they eventually return to the initial level. In sharp contrast to the Strongin and Bernanke-Mihov schemes, the responses for the long-run restrictions scheme are positive and eventually return to initial level. For the exchange rate, the responses for the Strongin and Bernanke and Mihov schemes reveal



Note: See Figure 4.2.

Figure 4.7
Impulse Response Functions: Shocks to Exchange Rate
The Federal Funds Rate, Foreign Interest Rate, Nominal and Real Exchange Rate, and Trade Balance

sharp rises (appreciation) in both nominal and real exchange rates, but the responses eventually return to the initial level after approximately 16 months. However, the responses for the long-run restriction scheme indicate persistent effects on both nominal and real exchange rates.

For the Strongin and Bernanke-Mihov schemes, the immediate responses of the trade balance to a positive exchange rate shock are positive, but they drop below zero after 2 months, indicating the J-curve effects last only for a month. The maximal effect of shocks to the exchange rate occurs 11 months after the shock, although the effects of the shock are prolonged for 33-34 months after the shock. However, for the long-run restrictions approach, the responses of the trade balance are positive and eventually return to the initial level after about 10 months.

Before concluding this sub-section, we notice that we do not repeat the same exercise as in the preceding sub-section in which we draw the confidence bands for the Bernanke-Mihov schemes with the point estimates for the other schemes. Since the similarities and differences among the effects of the exchange rate shocks across alternative identification schemes are clearly seen in Figures 4.6 and 4.7, doing an exercise as in the previous sub-section provides no additional information.

To summarize, like the effects of monetary policy shocks, the responses to an exchange rate shock for the Strongin (and hence Christiano-Eichenbaum-Evans) and Bernanke-Mihov schemes are similar, while the responses for the long-run restrictions scheme are quite different. For the Strongin and Bernanke-Mihov schemes, deterioration in the trade balance following positive shocks to the exchange rate is persistent. For the long-run restrictions approach, the trade balance is improved for about 10 months following positive shocks to the exchange rate. We note this is inconsistent with the implication of traditional open economy macroeconomic models.

4.3. Summary and Conclusion

This paper investigated the sensitivity of the effects of monetary policy and exchange rate shocks across alternative identification schemes in an open economy framework. For the monetary policy shocks, we have observed that the impulse responses for the contemporaneous restriction schemes, i.e. the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes, give, in general, reasonable results for output, interest variables, and the trade balance.

However, the magnitude and timing of the effects differ to some degree among these three schemes. The impulse responses for the Christiano, Eichenbaum, and Evans scheme reveal a relatively shorter lasting effect for output, a weaker initial effect for the exchange rate, and a larger initial negative effect for the trade balance compared to the Strongin and Bernanke-Mihov schemes. The Strongin and Bernanke-Mihov schemes give quite similar results. One problematic feature of these schemes can be seen in the responses of domestic and foreign price levels. In particular, the responses for the Strongin and Bernanke-Mihov schemes indicate some degree of 'price puzzle', which was not appeared in the closed economy models in Essay 1, for the domestic price level. The responses for the long-run restriction scheme are a good bit different from the contemporaneous restrictions schemes, especially for nonborrowed reserves and exchange rate. However, the point estimates for the long-run restrictions approach seem to be less precisely estimated.

The effects of exchange rate shocks, like the effects of monetary policy shocks, are similar for the Strongin and Bernanke-Mihov schemes, while the responses for the long-run restrictions scheme are quite different. The deterioration in the trade balance following positive shocks to the exchange rate is persistent in the Strongin and Bernanke-Mihov schemes. For the long-run restrictions approach, the trade balance is improved for first 10 months following positive shocks.

We note that, on the basis of the impulse response functions presented above, there is little basis to choose among the Christiano, Eichenbaum, and Evans, Strongin, and Bernanke-Mihov schemes. However, the long-run restrictions approach might not be suitable for a relatively large system like our 11-variable open economy framework. In particular, all identification schemes considered here showed either some degree of the 'price puzzle' or weaker price effects than in a closed economy framework (at least for the U.S. economy), even in the presence of commodity prices and the exchange rate.[16] This result suggests that we need more careful attention to the identification of monetary policy shocks in an open economy framework.

[16] Sims (1992) reported positive innovations in the foreign interest rate in Japan, France, and Germany, which indicate contractionary monetary policy shocks, are associated with persistent increases in price for the countries.

CHAPTER 5

CONCLUSIONS

This dissertation investigates the sensitivity of the effects of monetary policy shocks across alternative identification schemes and lag structures within vector autoregressive models. The four widely-cited identification schemes of Christiano, Eichenbaum, and Evans (1994; 1996), Strongin (1995), Bernanke and Mihov (1998), and the long-run restrictions approach pioneered by Blanchard and Quah (1989) are used.[1] Also, three different lag structures, namely symmetric, Keating-type asymmetric, and Hsiao-type asymmetric lag structures are employed.[2] The first essay focuses upon the sensitivity of the effects of monetary policy shocks within a closed economy framework, while the second essay is an attempt to clarify the effects of lag structure misspecification in assessing the effects of monetary policy shocks within a Monte Carlo experiment framework. In the third essay, the model is extended to an open economy framework.

In the first essay, using the above mentioned four identification schemes and three lag structures, the study found that the impulse response functions for output, the price level, and the federal funds rate are often sensitive to identification schemes and lag structures. For a given lag structure, the Strongin and Bernanke-Mihov Schemes generate quite similar results, while the Christiano, Eichenbaum, and Evans scheme often yields different responses from others. The responses from the long-run

[1] As explained earlier, the long-run restrictions approach is omitted in Chapter 3.
[2] Symmetric and Keating-type asymmetric lag structures are considered in Chapter 3, while only the symmetric lag structure is considered in Chapter 4.

restrictions approach are in general not substantially different from those for other schemes. When a symmetric lag structure is employed, all identification schemes considered generally showed similar impulse responses, although the results for the Christiano-Eichenbaum-Evans procedure indicate weaker output and liquidity effects. When a Keating-type asymmetric lag VAR is used, the Strongin and Bernanke-Mihov schemes reveal clearly weaker price effects than those of other schemes, while the Christiano-Eichenbaum-Evans scheme indicates somewhat problematic features for the federal funds rate. Finally, when a Hsiao-type asymmetric lag structure is used, the impulse responses from all identification schemes seem to be quite different from those in the symmetric or the Keating-type lag VAR.

As for the sensitivity of alternative identification schemes across the lag structures, the Christiano, Eichenbaum, and Evans and long-run restrictions schemes are relatively insensitive to the type of lag structures compared to the Strongin and Bernanke and Mihov schemes. For example, the Christiano, Eichenbaum, and Evans scheme seems to be insensitive to changes in lag structures between the symmetric lag structure and the Keating-type asymmetric lag structure, while the long-run restrictions approach is relatively insensitive between the symmetric and Hsiao-type lag structures. Finally, the Strongin and Bernanke-Mihov schemes are found to be somewhat sensitive to the type of lag structure.

In the second essay, it is shown that the lag structure of a VAR model does matter when assessing the effects of monetary policy shocks. For most horizons, t-statistics indicate that the responses from the VARs with the misspecified lag structure are significantly different from the assumed 'true' responses, although the pattern of the effects from the misspecified lag VARs is similar to the pattern from the 'true' model. When a symmetric lag structure is assumed to be the 'true' lag structure, the responses from a Keating-type asymmetric lag VAR and from a symmetric lag VAR(AIC) seem to be significantly weaker than the 'true' responses. This is also true for the case when a Keating-type asymmetric lag structure is assumed to be the 'true' lag structure; in most horizons, the mean of the point estimates from the symmetric lag VAR(4) and the symmetric lag VAR(AIC) also deviate significantly from the 'true' point estimates.

In the last essay, the sensitivity of the effects of monetary policy shocks across alternative identification schemes is investigated in an open economy framework. We found that the contemporaneous restriction schemes give, in general, reasonable results for output, interest variables, and the trade balance, although the long-run restriction scheme gives results that are a good bit different from those for the contemporaneous restrictions schemes. However, even for the contemporaneous restrictions schemes, the magnitude and timing of the effects differ to some degree across identification schemes. For example, the impulse responses for the Christiano, Eichenbaum, and Evans scheme reveal a relatively shorter lasting effect for output, a weaker initial effect for the exchange rate, and a larger initial negative effect for the trade balance compared to the Strongin and Bernanke-Mihov schemes. The Strongin and Bernanke-Mihov schemes reveal very similar results. These schemes generate some degree of the 'price puzzle'. The long-run restrictions approach might not be suitable for a relatively large system like our 11-variable open economy framework. The results from this approach indicate that monetary policy shocks have little effect on output. Moreover, the estimated confidence intervals are relatively large, indicating less precise estimation.

The results for the open economy framework are clearly contrasted with the results for the closed economy. Although the general patterns of the impulse responses are similar to those for the open economy model, the magnitude of responses are different. In the case of output, the hump-shaped patterns in the open economy models are similar to those in the closed economy models. However, for the Strongin, Bernanke-Mihov, and long-run restrictions approaches, the effects of monetary policy shocks on output are somewhat weaker than those in the closed economy models. In contrast, for the Christiano, Eichenbaum, and Evans scheme, the effects are slightly greater. For the price level, all identification schemes considered here showed either some degree of the 'price puzzle' or weaker price effects than in a closed economy framework (at least for the U.S. economy), even in the presence of commodity prices and the exchange rate. Regardless of identification scheme considered in this paper, the liquidity effects for the open economy model are clearly weaker than for the closed economy model; the effects are about one half of the effects from the closed economy model. However, in spite of the differences in magnitude, both frameworks generate significant liquidity effects for all identification schemes.

Several further remarks are in order. First, for a closed economy framework, although the responses from the long-run restrictions scheme are relatively insensitive to the type of lag structures, the responses seem to be less precisely estimated compared to other contemporaneous restrictions schemes. Consequently, it is useful to present the response from the long-run restrictions scheme along with the response from a contemporaneous restrictions scheme, especially either the Strongin or Bernanke and Mihov schemes.

Second, given inherent uncertainty about the lag structure in practice, it is important that one compare the impulse response functions from both symmetric lag and Keating-type asymmetric lag VARs in assessing the effects of monetary policy shocks. Since the differences between both responses are in general significant, employing a particular lag structure alone may result in misleading results. We note that even though the qualitative effects of monetary policy shocks are similar, reliable estimates of the quantitative effects are important for policy evaluation. Consequently, this approach may lessen difficulties in specifying the appropriate lag structure in monetary VAR models.

Finally, we should pay more careful attention to the identification of monetary policy shocks in an open economy framework. Although a closed economy framework generally gives reasonable responses, an identification scheme which incorporates international linkages between the U.S. and other industrial countries may be required for accurate estimates of the effects of monetary policy.

REFERENCES

Bagliano, Fabio C. and Carlo A. Favero. "Measuring Monetary Policy with VAR models: An Evaluation." *European Economic Review* 42 (June 1998), 1069-1112.

Bernanke, Ben S. "Alternative Explanation of Money-Income Correlation." *Carnegie Rochester Conference Series on Public Policy* 25 (Autumn 1986), 49-100.

Bernanke, Ben S. and Alan S. Blinder. "The Federal Funds Rate and the Channels of Monetary Transmission." *American Economic Review* 82 (September 1992), 901-921.

Bernanke, Ben S. and Ilian Mihov. "Measuring Monetary Policy." *Quarterly Journal of Economics* 113 (August 1998a), 869-902.

Bernanke, Ben S. and Ilian Mihov. "The Liquidity Effect and Long-run Neutrality." *Carnegie-Rochester Conference Series on Public Policy* 49 (1998b), 149-94.

Bernanke, Ben S. and Ilian Mihov. "What does the Bundesbank Target?" *European Economic Review* 41 (June 1997), 1025-1053.

Blanchard, Oliver Jean and Danny Quah. "The Dynamic Effects of Aggregate Demand and Supply Disturbances." *American Economic Review* 79 (September 1989), 655-73.

Boshen, John and Leonard Mills. "The Relation Between Narrative and Money Market Indicators of Monetary Policy." *Economic Inquiry* 33 (January 1995), 24-44.

Braun, P. A. and S. Mittnik. "Misspecification in Vector Autoregressions and Their Effects on Impulse Responses and Variance Decompositions." *Journal of Econometrics* 59 (October 1993), 319-41.

Caines, P.E., C.W. Keng, and S.P. Sethi. "Causality Analysis and Multivariate Autoregression Modelling with an Application to Supermarket Sales Analysis." *Journal of Economic Dynamics and Control* 3 (August 1981), 267-98.

Calvo, Guillermo A. and Carlos A. Rodriguez. "A Model of Exchange Rate Determination under Currency Substitution and Rational Expectations." *Journal of Political Economy* 85 (June 1977), 617-625.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans. "Monetary Policy Shocks: What Have We Learned and to What End?" NBER working paper 6400 (February 1998), 16-34.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans. "The Effects of Policy Shocks: Evidence from the Flow of Funds." *Review of Economics and Statistics* *78* (February 1996), 16-34.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans. "Identification and the Effects of Monetary Policy Shocks." Federal Reserve Bank of Chicago Working Paper 94-7, May 1994.

Christiano, Lawrence J., Martin Eichenbaum, and Charles L. Evans. "Sticky Price and Limited Participation Models: A Comparsion." *European Economic Review* 41 (June 1997), 1201-1249.

Cushman David O. and Tao Zha. "Identifying Monetary Policy in a Small Open Economy under Flexible Exchange Rates." *Journal of Monetary Economics* 39 (February 1996), 16-34.

Dornbush, Rudiger. "Expectations and Exchange Rate Dynamics." *Journal of Political Economy* 134 (1976), 1161-1176.

Eichenbaum, Martin. "Comments on Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy." *European Economic Review* 36 (June 1992), 1001-11.

Eichenbaum, Martin and Charles L. Evans. "Some Empirical Evidence on the Effects of Shocks to Monetary Policy on Exchange Rate." *Quarterly Journal of Economics* 110 (November 1995), 975-1009.

Enders, Walter. *Applied Econometric Time Series.* First Edition, New York: Wiley, 1995.

Fackler, James S. and W. Douglas McMillin. "Historical Decomposition of Aggregate Demand and Supply Shocks in a Small Macro Model" *Southern Economic Journal* 64 (January 1998), 648-64.

Faust, Jon and Eric M. Leeper. "When Do Long-Run Identifying Restrictions Give Reliable Results?" *Journal of Business and Economic Statistics* 15 (July 1997), 345-53.

Friedman, Milton and Anna J. Schwartz (1963), *A Monetary History of the United States: 1867-1960.* Princeton, N. J.: Princeton University Press.

Gordon, D.B. and E.M Leeper. "The Dynamic Impacts of Monetary Policy: An Exercise in Tentative Identification." *Journal of Political Economy* 102 (December 1994), 1228-47.

Gordon, Robert J. and Stephen King. "The Output Cost of Inflation in Traditional and Vector Autoregressive Models." *Brooking Papers on Economic Activity 1* (1982), 205-44.

Hoover, K.D. and S.J. Perez. "Post Hoc Ergo Propter Hoc Once More: An Evaluation of "Does Monetary Policy Matter?" in the Spirit of James Tobin." *Journal of Monetary Economics* 34 (August 1994), 89-99.

Hsiao, Cheng. "Autoregressive Modelling and Money-Income Causality Detection." *Journal of Monetary Economics* 7 (January 1981), 85-106.

Keating, John W. "Vector Autoregressive Models with Asymmetric Lag structures." *Journal of Macroeconomics* (forthcoming).

Keating, J. "Structural Information in Recursive VAR Orderings." *Journal of Economics and Control* 20 (September-October 1996), 1557-1580.

Kennedy, P. and D. Simons. "Fighting the Teflon Factor: Comparing Classical and Bayesian Estimators for Autocorrelated Errors." *Journal of Econometrics,* 48 (April/May 1991), 15-27.

Koray, Faik and W. Douglas McMillin. "Monetary Policy Shocks, the Exchange Rate, and the Trade Balance." *Journal of International Money and Finance* (forthcoming).

Krugman, Paul R. and Robert E. Baldwin. "The Persistence of the U.S. Trade Deficit." *Brooking Papers on Economic Activity* 1 (1987), 1-43.

Lastrapes, William D. "International Evidence on Equity Prices, Interest Rates, and Money." *Journal of International Money and Finance*, 17 (June 1998a), 377-406.

Lastrapes, William D. "The Dynamic Effects of Money: Combining Short-Run and Long-Run Identifying Restrictions Using Bayesian Techniques." *Review of Economics and Statistics*, 80 (November 1998b), 588-99

Lastrapes, William D. and Selgin, Gerorge. "Testing Long-Run Neutrality: Identifying Short-Run Interest Rate Dynamics Using Long-Run Restrictions." *Journal of Macroeconomics*, 17 (Summer 1995), 387-404.

Lee, Jim. "Money, Income and Dynamic Lag Patterns." *Southern Economic Journal*, 64 (July 1997), 97-103.

Leeper, Eric. "Narrative and VAR Approaches to Monetary Policy: Common Identification Problems?" Indiana University Working Paper, December 1996.

Litterman, Robert B. "Forecasting with Bayesian Vector Autoregressions-Five Years of Experience." *Journal of Business and Economic Statistics* 4 (January 1986), 25-38.

Lucas, Robert E. Jr. "Econometric Policy Evaluation: A Critique." *Journal of Monetary Economics* 2 (Supplement 1976), 19-46.

McMillin, W. Douglas and James S. Fackler. "Monetary vs. Credit Aggregates: An Evaluation of Monetary Policy Targets." *Southern Economic Journal* 50 (January 1984), 711-723.

McMillin, W. Douglas."The Effects of Monetary Shocks: Comparing Contemporaneous Versus Long-Run Identifying Restrictions." Working Paper (May 1998), Louisiana State University.

Moffett, Michael H. "The J-Curve Revisited: An Empirical Examination for the United States ." *Journal of International Money and Finance* 8 (September 1989), 425-44.

Mundell, Robert A. "Capital Mobility and Stabilization Policy under Fixed and Flexible Exchange Rates." Chapter 18 in his *International Economics*, New York: McMillan, 1968.

Ozcicek, Omer and W. Douglas McMillin. "Lag Length Selection in Vector Autoregressive Models: Symmetric and Asymmetric Lags." *Applied Economics* (forthcoming).

Phillips, P.C.B. "Model Determination and Macrroeconomic Activity." Cowles Foundation Working Paper No. 1083 (1994).

Romer, C.R. and D.H. Romer. "Does Monetary Policy Matter? A New Test in the Spirit of Friedman and Schwartz." in O.J. Blanchard and S. Fisher, eds. *NBER Macroeconomics Annual 1989*, Cambridge: MIT Press, 1989, 121-69.

Rose, Andrew K. and Janet L. Yellen. "Is There a J-curve?" *Journal of Monetary Economics* 24 (July 1989), 53-68.

Rudebusch, Glenn D. "Do Measures of Monetary Policy in a VAR Make Sense?" *International Economic Review* 39 (November 1998), 907-32.

Runkle, David E. "Vector Autoregressions and Reality." *Journal of Business and Economic Statistics* (October 1987), 437-54.

Shapiro, Matthew and Mark Watson. "Sources of Business Cycle Fluctuations." *NBER Macroeconomics Annual* 3 (1989), 111-48.

Sims, Christopher A. "Comment on Glenn Rudebusch's 'Do Measure of Monetary Policy in a VAR Make Sense?'" *International Economic Review* 39 (November 1998), 933-41.

Sims, Christopher A. "Interpreting the Macroeconomic Time Series Facts: The Effects of Monetary Policy" *European Economic Review* 36 (June 1992), 975-1011.

Sims, Christopher A. "Are Forecasting Models Usable for Policy Analysis?" *Federal Reserve Bank of Minneapolis Quarterly Review* (Winter 1986), 3-16.

Sims, Christopher A. "Macroeconomics and Reality." *Econometrica* 48 (January 1980), 1-48.

Strongin, Steven. "The Identification of Monetary Policy Disturbances: Explaining the Liquidity Puzzle." *Journal of Monetary Economics* 35 (August 1995), 463-97.

APPENDIX: DATA DESCRIPTIONS AND SOURCES

This appendix provides a description and sources of the data used in Chapter 4 in detail. All data are extracted from the DRI database: especially, DRI Basic Economics and International Monetary Fund (IMF) databases. The consumer price index, exports and imports, and commodity price were seasonally adjusted using the X-11 procedure.

Table A.1
Data Descriptions and Sources

Variables	Code	Unit	SA/NSA	Source
US Variables				
Industrial Production	IP	1992=100	SA	DRI Basic
Personal Consumption Deflator	GMDC	1987=100	SA	
Commodity Prices	PSCCOM	1987=100	NSA	
Nonborrowed Reserves	FMRNBA	Mil.$	SA	
Total Reserves	FMRRA	Mil.$	SA	
the Federal Funds Rate	FYFF	Mil.$	NSA	
Bilateral Exchange Rates				
France	EXRFR	franc/$	NSA	DRI Basic
Germany	EXRGER	DM/$	NSA	
Japan	EXRJAN	yen/$	NSA	
Italy	EXRITL	lira/$	NSA	
U.K.	EXRUK	c/pound	NSA	
Canada	EXRCAN	C. $/$	NSA	
US exports to G6				
France	FZEXFR	Mil.$	NSA	DRI Basic
Germany	FZEXG	Mil.$	NSA	
Japan	FZEXJP	Mil.$	NSA	
Italy	FZEXIT	Mil.$	NSA	
U.K.	FZEXUK	Mil.$	NSA	
Canada	FZEXCA	Mil.$	NSA	
U.S imports to G6				
France	FZIMFR	Mil.$	NSA	
Germany	FZIMG	Mil.$	NSA	
Japan	FZIMJP	Mil.$	NSA	
Italy	FZIMIT	Mil.$	NSA	
U.K.	FZIMUK	Mil.$	NSA	
Canada	FZIMCA	Mil.$	NSA	

(Table continued)

Variables	Code	Unit	SA/NSA	Source
Foreign Industrial Production				
France	IPFR	1987=100	SA	DRI Basic
Germany	IPWG	1990=100	SA	
Japan	IPJP	1990=100	SA	
Italy	IPIT	1987=100	SA	
U.K.	IPUK	1987=100	SA	
Canada	IPCA	1992=100	SA	
Foreign CPI				
France	PC6FR		NSA	DRI Basic
Germany	PC6WG		NSA	
Japan	PC6JP		NSA	
Italy	PC6IT		NSA	
U.K.	PC6UK		NSA	
Canada	PC6CA		NSA	
Foreign Interest rates				
France	L60B@132	Percent per annum		IMF
Germany	L60B@134			
Japan	L60B@158			
Italy	L60B@136			
U.K.	L60B@112			
Canada	L60C@156			

Note: SA denotes seasonally adjusted series, while NSA represents not seasonally adjusted series at sources.

VITA

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DOCTORAL EXAMINATION AND DISSERTATION REPORT

Candidate: Keuk-Soo Kim

Major Field: Economics

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Approved:

Major Professor and Chairman

Dean of the Graduate School

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