6

VAR APPROACH

6.1 Introduction: why VAR models?

The LSE methodology has interpreted the failure of the traditional Cowles Commission approach, heralded by the critiques of Lucas (1976) and Sims (1980*a*), as the result of using mis-specified and ill-identified models. The LSE methodology, however, does not question the potential of macroeconometric modelling for simulation and econometric policy evaluation. In fact, at the stage of simulation and policy evaluation, there is no difference between the traditional Cowles Commission approach and the LSE approach. The LSE solution to the problems of traditional macroeconometric modelling concentrates on the stages of identification and specification. The importance of estimation is de-emphasized, in that congruency of the specification is considered as a much higher priority than the choice of the most appropriate estimator. No innovation is proposed at the stage of simulation and policy evaluation: the traditional methods are applied, after having tested, tested, and tested.

The VAR approach shares with the LSE methodology the diagnosis of the problem of Cowles Commission models but also questions the potential of traditional macroeconometric modelling for policy simulation and econometric policy evaluation. VAR models of the monetary transmission mechanism differ from structural LSE models as to the purpose of their specification and estimation. In the traditional approach the typical question asked within a macroeconometric framework is 'What is the optimal response by the monetary authority to movement in macroeconomic variables to achieve given targets for the same variables?' The VAR approach fully recognizes the potential of the Lucas critique and acknowledges that questions like 'How should a central bank respond to shocks in macroeconomic variables?' are to be answered within the framework of quantitative monetary general equilibrium models of the business cycle. So the answer should rely on a theoretical model rather than on an empirical *ad hoc* macroeconometric model. Within this framework, there emerges a new role for empirical analysis, which is to provide evidence on the stylized facts to include in the theoretical model adopted for policy analysis and to decide between competing general equilibrium monetary models. The operationalization of this research program is very well described in a recent paper by Christiano, Eichenbaum and Evans (1998). There are three relevant steps:

1. monetary policy shocks are identified in actual economies;

2. the response of relevant economic variables to monetary shocks is then described;
3. finally, the same experiment is performed in the model economies to compare actual and model-based responses as an evaluation tool and a selection criterion for theoretical models.

LSE-type structural models and VAR models of the monetary transmission mechanism have a common structure which, using the notation of Chapter 3, is represented as follows:

$$\mathbf{A}\begin{pmatrix}\mathbf{Y}_t\\ \mathbf{M}_t\end{pmatrix} = \mathbf{C}(L)\begin{pmatrix}\mathbf{Y}_{t-1}\\ \mathbf{M}_{t-1}\end{pmatrix} + \mathbf{B}\begin{pmatrix}\boldsymbol{\nu}_t^Y\\ \boldsymbol{\nu}_t^M\end{pmatrix} \tag{6.1}$$

where $\mathbf{Y}$ and $\mathbf{M}$ are vectors of macroeconomic (non-policy) variables (e.g. output and prices) and variables controlled by the monetary policy-maker (e.g. interest rates and monetary aggregates containing information on monetary policy actions) respectively. Matrix $\mathbf{A}$ describes the contemporaneous relations among the variables and $\mathbf{C}(L)$ is a matrix finite-order lag polynomial. $\boldsymbol{\nu} \equiv \begin{pmatrix}\boldsymbol{\nu}^Y\\ \boldsymbol{\nu}^M\end{pmatrix}$ is a vector of structural disturbances to the non-policy and policy variables; non-zero off-diagonal elements of $\mathbf{B}$ allow some shocks to affect directly more than one endogenous variable in the system. The main difference between the two approaches lies in the aim for which models are estimated.

Traditional Cowles Commission structural models are designed to identify the impact of policy variables on macroeconomic quantities to determine the value to be assigned to the monetary instruments ($\mathbf{M}$) to achieve a given target for the macroeconomic variables ($\mathbf{Y}$). The policy variables in $\mathbf{M}$ are considered as exogenous on the grounds that these are the instruments controlled by the policy-maker. Identification in traditional structural models is obtained without assuming the orthogonality of structural disturbances. Dynamic multipliers are used to describe the impact of monetary policy variables on macroeconomic quantities. In the computation of dynamic multipliers the responses of macroeconomic variables to monetary policy can be, and usually is, obtained without decomposing monetary policy into its endogenous and exogenous components.

The assumed exogeneity of the monetary variables in the traditional approach makes the model invalid for policy analysis if monetary policy reacts endogenously to macroeconomic variables. The LSE methodology would recognize the problem of the invalid exogeneity assumption for $\mathbf{M}$, it would then proceed to the identification of an alternative enlarged model.[1] However, the new model would still be used for simulation and econometric policy evaluation, whenever the appropriate concept of exogeneity (respectively strong and super) where satisfied by the adopted specification.

[1] Presumably, such identification would be obtained through the imposition of a priori restrictions on the dynamics of the lagged variables.

VAR modelling would reject the Cowles Commission identifying restrictions as 'incredible' for reasons not very different from the ones pinned down by the LSE approach, however, VAR models of the transmission mechanism are not estimated to yield advice on the best monetary policy. They are rather estimated to provide empirical evidence on the response of macroeconomic variables to monetary policy impulses in order to discriminate between alternative theoretical models of the economy. It then becomes crucial to identify monetary policy actions using restrictions independent from the competing models of the transmission mechanism under empirical investigation, taking into account the potential endogeneity of policy instruments.

In a series of recent papers, Christiano, Eichenbaum and Evans (1996*a*)–(1996*b*) apply the VAR approach to derive 'stylized facts' on the effect of a contractionary policy shock, and conclude that plausible models of the monetary transmission mechanism should be consistent at least with the following evidence on price, output and interest rates: *(i)* the aggregate price level initially responds very little; *(ii)* interest rates initially rise, and *(iii)* aggregate output initially falls, with a j-shaped response, with a zero long-run effect of the monetary impulse. Such evidence leads to the dismissal of traditional real business cycle models, which are not compatible with the liquidity effect of monetary policy on interest rates, and of the Lucas (1972) model of money, in which the effect of monetary policy on output depends on price misperceptions. The evidence seems to be more in line with alternative interpretations of the monetary transmission mechanism based on sticky prices models (Goodfriend and King 1997), limited participation models (Christiano and Eichenbaum 1992) or models with indeterminacy–sunspot equilibria (Farmer 1997).

Having stated the objective of VAR models we are now in the position to assess how identification, estimation and simulation are implemented to analyse the monetary transmission mechanism.

VAR models concentrate on shocks. First the relevant shocks are identified, and the response of the system to shocks is described by analysing impulse responses (the propagation mechanism of the shocks), forecasting error variance decomposition, and historical decomposition.

6.2 Identification and estimation

We introduced the identification problem for VAR in Chapter 3.

The structural model (6.1) is not directly observable; however a VAR can be estimated as the reduced form of the underlying structural model:

$$\begin{pmatrix} \mathbf{Y}_t \\ \mathbf{M}_t \end{pmatrix} = \mathbf{A}^{-1}\mathbf{C}(L) \begin{pmatrix} \mathbf{Y}_{t-1} \\ \mathbf{M}_{t-1} \end{pmatrix} + \begin{pmatrix} \mathbf{u}_t^Y \\ \mathbf{u}_t^M \end{pmatrix}, \tag{6.2}$$

where $\mathbf{u}$ denotes the VAR residual vector, normally independently distributed with full variance-covariance matrix Σ. The relation between the residuals in $\mathbf{u}$ and the structural disturbances in $\boldsymbol{\nu}$ is therefore:

$$\mathbf{A}\begin{pmatrix}\mathbf{u}_t^Y\\ \mathbf{u}_t^M\end{pmatrix}=\mathbf{B}\begin{pmatrix}\boldsymbol{\nu}_t^Y\\ \boldsymbol{\nu}_t^M\end{pmatrix}. \tag{6.3}$$

Undoing the partitioning we have

$$\mathbf{u}_t=\mathbf{A}^{-1}\mathbf{B}\boldsymbol{\upsilon}_t,$$

from which we can derive the relation between the variance-covariance matrices of $\mathbf{u}_t$ (observed) and $\boldsymbol{\nu}_t$ (unobserved) as follows:

$$E\left(\mathbf{u}_t\mathbf{u}_t'\right)=\mathbf{A}^{-1}\mathbf{B}E\left(\boldsymbol{\upsilon}_t\boldsymbol{\upsilon}_t'\right)\mathbf{B}'\mathbf{A}^{-1}.$$

Substituting population moments with sample moments we have:

$$\widehat{\sum}=\widehat{\mathbf{A}}^{-1}\mathbf{BI}\widehat{\mathbf{B}}'\widehat{\mathbf{A}}^{-1}, \tag{6.4}$$

$\widehat{\sum}$ contains $n(n+1)/2$ different elements, which is the maximum number of identifiable parameters in matrices $\mathbf{A}$ and $\mathbf{B}$. Therefore, a necessary condition for identification is that the maximum number of parameters contained in the two matrices equals $n(n+1)/2$, such a condition makes the number of equations equal to the number of unknowns in system (6.4). As usual, for such a condition also to be sufficient for identification no equation in (6.4) should be a linear combination of the other equations in the system (see Amisano and Giannini 1996, Hamilton 1994). As for traditional models, we have the three possible cases of under-identification, just-identification and over-identification. The validity of over-identifying restrictions can be tested via a statistic distributed as a χ^2 with a number of degrees of freedom equal to the number of over-identifying restrictions. Once identification has been achieved, the estimation problem is solved by applying generalized method of moments estimation. We describe this class of estimators in the next chapter.

In practice, identification requires the imposition of some restrictions on the parameters of $\mathbf{A}$ and $\mathbf{B}$. This step has been historically implemented in a number of different ways. We concentrate on the most widely used strategies in the next subsections.

6.2.1 *Choleski decomposition*

In the famous article which introduced VAR methodology to the profession, Sims (1980a) proposed the following identification strategy, based on the Choleski decomposition of matrices:

$$\mathbf{A}=\begin{pmatrix}1 & 0\,0 & 0\\ a_{21} & 1\,0 & 0\\ . & .\,1 & .\\ a_{n1} & .\;a_{nn-1} & 1\end{pmatrix},\qquad \mathbf{B}\begin{pmatrix}b_{11} & 0 & 0 & 0\\ 0 & b_{22} & 0 & 0\\ . & . & b_{ii}. & .\\ 0 & 0 & 0 & b_{nn}\end{pmatrix}. \tag{6.5}$$

This is obviously a just-identification scheme, where the identification of structural shocks depends on the ordering of variables. It corresponds to a recursive economic structure, with the most endogenous variable ordered last.

6.2.2 *Structural models with contemporaneous restrictions*

In this identification scheme some a priori information is used to impose restrictions on the elements of matrices **A** and **B**, different from the Choleski ordering. If the objective of VAR is to provide evidence to choose between competing models, the identifying restrictions should be independent from the theoretical predictions of those models. The recent literature on the monetary transmission mechanism (see Strongin 1995, Bernanke and Mihov 1995, Christiano, Eichenbaum and Evans 1996, Leeper, Sims and Zha 1996), offers good examples on how these kind of restrictions can be derived. VARs of the monetary transmission mechanism are specified on six variables, with the vector of macroeconomic non-policy variables including gross domestic product (GDP), the consumer price index (P) and the commodity price level (Pcm), the vector of policy variables includes the federal funds rate (FF), the quantity of total bank reserves (TR) and the amount of non-borrowed reserves (NBR). Given the estimation of the reduced form VAR for the six macro and monetary variables, a structural model is identified by: (*i*) assuming orthogonality of the structural disturbances; (*ii*) imposing that macroeconomic variables do not simultaneously react to monetary variables, while the simultaneous feedback in the other direction is allowed, and (*iii*) imposing restrictions on the monetary block of the model reflecting the operational procedures implemented by the monetary policy-maker. All identifying restrictions satisfy the criterion of independence from specific theoretical models. In fact, within the class of models estimated on monthly data, restrictions (*ii*) are consistent with a wide spectrum of alternative theoretical structures and imply a minimal assumption on the lag of the impact of monetary policy actions on macroeconomic variables, whereas restrictions (*iii*) are based on institutional analysis. Restrictions (*ii*) are made operational by setting to zero an appropriate block of elements of the **A** matrix.

The contemporaneous relations among the Fed funds rate and the reserve aggregates are derived, as in Bernanke and Mihov (1995), from a specific model of the reserve market:

$$u^{TR} = -\alpha u^{FF} + \nu^{D} \tag{6.6}$$

$$u^{BR} = \beta u^{FF} + \nu^{B} \tag{6.7}$$

$$u^{NBR} = \phi^{D}\nu^{D} + \phi^{B}\nu^{B} + \nu^{S}. \tag{6.8}$$

Equations (6.6) and (6.7) describe banks' demand equations (expressed in innovation, i.e. VAR residual–form) for total, TR, and borrowed, BR, reserves (time subscripts are omitted): the federal funds rate affects negatively the demand for total reserves (6.6) and positively the demand for borrowed reserves.[2] ν^D and ν^B are disturbances to total and borrowed reserves respectively. The supply of non-borrowed reserves in (6.8) reflects the behaviour of the Federal Reserve. In

[2]We assume from the start that movements in the discount rate, which would enter (6.7) with a negative sign, are completely anticipated, so that the innovation in the Fed funds-discount rate differential is entirely attributable to the former rate.

particular, by means of open-market operations, the Fed can change the amount of NBR supplied to the banking system in response to (readily observed) disturbances to total and borrowed reserves demand. Moreover, variations in non-borrowed reserves may be due to monetary policy shocks unrelated to reserves demand behaviour. In (6.8) the coefficients ϕ^D and ϕ^B measure the reaction of the Fed to total and borrowed reserve demand movements respectively, and ν^S represents the monetary policy shock to be empirically identified. The market for reserves featuring the assumed simultaneous relations is described by Figure 6.1.

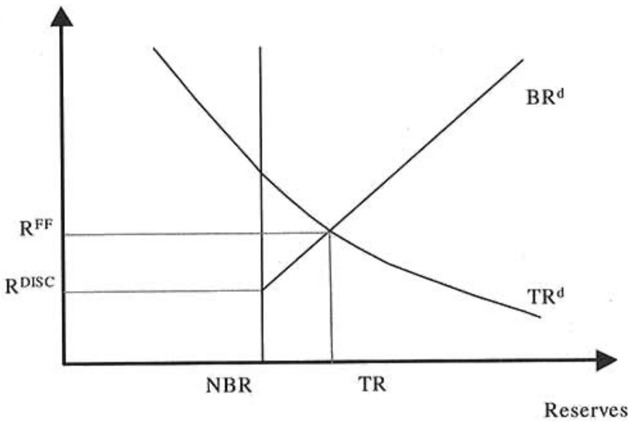

FIG. 6.1. The US market for bank reserves

Combining the market for reserves with the macroeconomic variables, we can explicitly rewrite (6.3) as follows:

$$\begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 & 0 \\ a_{31} & a_{32} & 1 & 0 & 0 & 0 \\ a_{41} & a_{42} & a_{43} & 1 & -\frac{1}{\beta} & \frac{1}{\beta} \\ a_{51} & a_{52} & a_{53} & \alpha & 1 & 0 \\ a_{61} & a_{62} & a_{63} & 0 & 0 & 1 \end{pmatrix} \begin{pmatrix} u_t^{GDP} \\ u_t^{P} \\ u_t^{Pcm} \\ u_t^{FF} \\ u_t^{TR} \\ u_t^{NBR} \end{pmatrix} = \begin{pmatrix} 1 & 0 & 0 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 & 0 & 0 \\ 0 & 0 & 1 & 0 & 0 & 0 \\ 0 & 0 & 0 & -\frac{1}{\beta} & 0 & 0 \\ 0 & 0 & 0 & 0 & 1 & 0 \\ 0 & 0 & 0 & \phi^B & \phi^D & 1 \end{pmatrix} \begin{pmatrix} \nu_{1t}^{NP} \\ \nu_{2t}^{NP} \\ \nu_{3t}^{NP} \\ \nu_t^{B} \\ \nu_t^{D} \\ \nu_t^{S} \end{pmatrix}. \tag{6.9}$$

Several features of (6.9) must be noted. First, VAR residuals from the first three equations, describing the non-policy part of the system, are orthogonalized simply by assuming a recursive (Choleski) structure for the corresponding block of the $\mathbf{A}$ matrix. This procedure yields orthogonal disturbances to which we do not attach a specific 'structural' interpretation, labelling them simply as ν_i^{NP} $(i = 1, 2, 3)$, where NP denotes a non-policy shock.

Second, as shown by Bernanke and Mihov (1995), the general formulation in (6.9) is still not identified, but identification can be completed by a careful analysis of the operational procedures followed by the central bank.

- Case 1: Federal funds targeting.

In this case we have $\varphi^d = 1, \varphi^b = -1$. Central bank uses NBR to neutralize shocks coming from banks' and households' behaviour. We then have for the monetary block identification:

$$\begin{bmatrix} u_t^{TR} \\ u_t^{NBR} \\ u_t^{FF} \end{bmatrix} = \begin{bmatrix} 1 & \frac{\alpha}{\beta+\alpha} & 0 \\ 1 & 1 & -1 \\ 0 & -\frac{1}{\beta+\alpha} & 0 \end{bmatrix} \begin{bmatrix} \nu_t^D \\ \nu_t^S \\ \nu_t^B \end{bmatrix}.$$

The model is now over-identified. Note that this identification scheme is equivalent to a Choleski identification with the ordering Federal funds rate, total reserves and non-borrowed reserves Choleski plus some additional within equation and cross equations restrictions.

- Case II: targeting NBR.

$\varphi^d = 0, \varphi^b = 0.$

$$\begin{bmatrix} u_t^{TR} \\ u_t^{NBR} \\ u_t^{FF} \end{bmatrix} = \begin{bmatrix} \frac{\beta}{\beta+\alpha} & \frac{\alpha}{\beta+\alpha} & \frac{\alpha}{\beta+\alpha} \\ 0 & 1 & 0 \\ \frac{1}{\beta+\alpha} & -\frac{1}{\beta+\alpha} & -\frac{1}{\beta+\alpha} \end{bmatrix} \begin{bmatrix} \nu_t^D \\ \nu_t^S \\ \nu_t^B \end{bmatrix}$$

Monetary policy shocks now coincide with the VAR innovations for non-borrowed reserves, NBR, while the Federal funds rate reflects all type of shocks.

- Case III: Strongin identification (1994).

Shocks to reserves are demand shocks which the central bank has to accommodate. Therefore monetary policy shocks are the shocks to NBR orthogonal to shocks to TR. Moreover, the central bank does not react to borrowed reserves. $\alpha = 0, \varphi^b = 0,$

$$\begin{bmatrix} u_t^{TR} \\ u_t^{NBR} \\ u_t^{FF} \end{bmatrix} = \begin{bmatrix} 1 & 0 & 0 \\ \varphi^d & 1 & 0 \\ -\frac{\varphi^d - 1}{\beta} & -\frac{1}{\beta} & -\frac{1}{\beta} \end{bmatrix} \begin{bmatrix} \nu_t^D \\ \nu_t^S \\ \nu_t^B \end{bmatrix}$$

NBR is now the variable from which monetary policy shocks are extracted.

- Case IV: controlling borrowed reserves.

In this case TR – NBR is only a function of shocks ν^s. $\varphi^d = 1, \varphi^b = \frac{\alpha}{\beta}$. We have:

$$\begin{bmatrix} u_t^{TR} \\ u_t^{NBR} \\ u_t^{FF} \end{bmatrix} = \begin{bmatrix} 1 & \frac{\alpha}{\beta+\alpha} & \frac{\alpha}{\beta} \\ 1 & 1 & \frac{\alpha}{\beta} \\ 0 & -\frac{1}{\beta+\alpha} & -\frac{1}{\beta} \end{bmatrix} \begin{bmatrix} \nu_t^D \\ \nu_t^S \\ \nu_t^B \end{bmatrix}.$$

It is easily seen that an alternative regime would imply identification which is technically not far from Choleski's triangularization, with different ordering of the monetary variables. In a Fed funds targeting regime Federal funds rate does not react contemporaneously to the other monetary variables while in a non-borrowed reserves targeting regime it is NBR that does not react contemporaneously to the other two monetary shocks. Moreover, information on the operating procedure by the Fed are important in determining the appropriate identification scheme and, more importantly, VAR models of the monetary transmission mechanism should be estimated within a single policy regime. Bagliano and Favero (1998) provide evidence on the structural instability of VAR of the MTM estimated across different monetary policy regimes.

6.2.3 *Structural VARs with long-run restrictions*

Often long-run behaviour of shocks provide restrictions acceptable within a wide range of theoretical models. A typical restriction compatible with virtually all macroeconomic models is that in the long-run demand shocks have zero impact on output. Blanchard and Quah (1989) show how these restrictions can be used to identify VARs.

The structural model of interest is specified by posing $\mathbf{A}$ equal to the identity matrix and by imposing no restriction on the $\mathbf{B}$ matrix. We then have the following specification for a generic vector of variables $\mathbf{y}_t$:

$$\mathbf{y}_t = \sum_{i=1}^{p} \mathbf{A}_i \mathbf{y}_{t-i} + \mathbf{B}\mathbf{v}_t$$

from which one can derive the matrix which describes the long-run effect of the structural shocks on the variables of interest as follows:

$$\left(\mathbf{I} - \sum_{i=1}^{p} \mathbf{A}_i\right)^{-1} \mathbf{B}\mathbf{v}_t = -\Pi^{-1}\mathbf{B}\mathbf{v}_t.$$

Coefficients in Π are obtained from the reduced form, therefore, we are able to impose long-run restrictions given the estimation of the reduced form.

Two points are worth noting:

1. $(I - A_1)$ is -Π, for this matrix to be invertible the VAR must be specified on stationary variables;
2. the long-run restrictions are restrictions on the cumulative impulse response function.

Let us now consider the Blanchard and Quah (1989) dataset. The authors aim at separating demand shocks from supply shocks, they consider a VAR on two variables, the unemployment rate, *UN*, and the quarterly rate of growth of GDP, ΔLY. The original sample contains quarterly data from 1951:2 to 1987:4, we have retrieved the two series from Datastream and they are available only for the sample 1951:2–1987:4. The series are available in the file bq.wks. The VAR is specified with 8 lags, a constant, and a deterministic trend (in the original paper a break in the constant for ΔLY is also allowed but we do not allow it here) as follows:

$$\begin{pmatrix}\Delta LY_t\\ UN_t\end{pmatrix} = A_1\begin{pmatrix}\Delta LY_{t-1}\\ UN_{t-1}\end{pmatrix} + ... + A_8\begin{pmatrix}\Delta LY_{t-8}\\ UN_{t-8}\end{pmatrix} + A_9\begin{pmatrix}1\\ t\end{pmatrix} + \begin{pmatrix}u_{1t}\\ u_{2t}\end{pmatrix}.$$

The structure of interest is the following:

$$\begin{aligned}\begin{pmatrix}\Delta LY_t\\ UN_t\end{pmatrix} &= A_1\begin{pmatrix}\Delta LY_{t-1}\\ UN_{t-1}\end{pmatrix} + ...A_8\begin{pmatrix}\Delta LY_{t-8}\\ UN_{t-8}\end{pmatrix} + A_9\begin{pmatrix}1\\ t\end{pmatrix}\\ &\quad + \begin{pmatrix}b_{11}\ b_{12}\\ b_{21}\ b_{22}\end{pmatrix}\begin{pmatrix}v_{1t}\\ v_{2t}\end{pmatrix}.\end{aligned}$$

To obtain the identifying restrictions consider that

$$\begin{aligned}\begin{pmatrix}\Delta LY_t\\ UN_t\end{pmatrix} &= \left(\mathbf{I} - \sum_{i=1}^{p}\mathbf{A}_i\right)^{-1}\begin{pmatrix}b_{11}\ b_{12}\\ b_{21}\ b_{22}\end{pmatrix}\begin{pmatrix}v_{1t}\\ v_{2t}\end{pmatrix}\\ &= \begin{pmatrix}k_{11}\ k_{12}\\ k_{21}\ k_{22}\end{pmatrix}\begin{pmatrix}b_{11}\ b_{12}\\ b_{21}\ b_{22}\end{pmatrix}\begin{pmatrix}v_{1t}\\ v_{2t}\end{pmatrix}.\end{aligned}$$

Demand shocks are identified by imposing that their long-run impact on the level of output is zero:

$$k_{11}b_{11} + k_{12}b_{21} = 0.$$

Note that by imposing the restriction that the cumulative impulse response of the rate of output growth to a demand shock is zero we impose the restriction that the impulse response of the level of output to a demand shock is zero in the long run. As the variables are stationary the long-run response of ΔLY and UN to all shocks is zero by definition.

We implement the procedure on the data by using Malcolm (1998), a package written by Rocco Mosconi for RATS.

We start from the estimation of the VAR, we then implement the Johansen procedure on this VAR, we make sure that the null of stationarity is not rejected. We then retrieve the Π matrix. $\Pi = \begin{bmatrix}0.1451 & 0.2168\\ -0.5741 & -0.0693\end{bmatrix}$, then:

$$(-\Pi)^{-1} = \begin{bmatrix}-0.1451 & -0.2168\\ 0.5741 & 0.0693\end{bmatrix}^{-1} = \begin{bmatrix}0.60572 & 1.8949\\ -5.0179 & -1.2683\end{bmatrix},$$

and our long-run identifying restriction is

$$0.60572b_{11} + 1.8949b_{21} = 0.$$

Note the difference between this methodology and the Choleski decomposition, which would simply restrict b_{21} to zero.

6.2.4 *Identification in cointegrated VARs*

Let us consider now how the identification problem changes when we have a cointegrated VAR. Considering, for simplicity, only first-order dynamics, the cointegrated reduced form is:

$$\Delta \mathbf{y}_t = \Pi \mathbf{y}_{t-1} + \mathbf{v}_t,$$

where $\Pi = \boldsymbol{\alpha}\boldsymbol{\beta}'$. As we know, identification of the cointegrating vectors is a problem totally separated from identification of the structural shocks of interest. Therefore, having solved the identification of the cointegrating relationships, we still have to deal with the problem of posing appropriate restrictions on the parameters of the $\mathbf{B}$ matrix in order to pin down the shocks $\mathbf{u}_t$

$$\Delta \mathbf{y}_t = \Pi \mathbf{y}_{t-1} + \mathbf{B}\mathbf{u}_t.$$

In the context of cointegration, the identification problem can be solved in a very natural way. Consider, for simplicity, the case of a bivariate model $\mathbf{y}_t = (y_t, x_t)$, in which variables are non-stationary $I(1)$ but cointegrated with a cointegrating vector $(1, -1)$, so the rank of the Π matrix is 1 and we use the following representation of the stationary reduced form:

$$\begin{pmatrix} \Delta y_t \\ \Delta x_t \end{pmatrix} = \begin{pmatrix} \alpha_{11} \\ \alpha_{21} \end{pmatrix} (1 \; -1) \begin{pmatrix} y_{t-1} \\ x_{t-1} \end{pmatrix} + \begin{pmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} v_{1t} \\ v_{2t} \end{pmatrix}. \tag{6.10}$$

Model (6.10) can be re-written as follows Mellander, Vredin and Warne (1993):

$$\begin{aligned} \begin{pmatrix} -1 & 1 \\ 0 & 1 \end{pmatrix} \begin{pmatrix} (1-L) & 0 \\ 0 & 1 \end{pmatrix} \begin{pmatrix} (y_t - x_t) \\ \Delta x_t \end{pmatrix} &= \begin{pmatrix} \alpha_{11} & 0 \\ \alpha_{21} & 0 \end{pmatrix} \begin{pmatrix} (y_{t-1} - x_{t-1}) \\ \Delta x_{t-1} \end{pmatrix} \\ &\quad + \begin{pmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} v_{1t} \\ v_{2t} \end{pmatrix}. \end{aligned} \tag{6.11}$$

The two representations are completely identical (they feature the same residuals). The second representation has been widely used in research based on present value models. The cointegrating properties of the system suggest the presence of two types of shocks: a permanent one (related to the single common trend shared by the two variables) and a transitory one (related to the cointegrating relation). It seems therefore natural to identify one shock as permanent and the other as transitory. Given that we have a stationary system, the identification of shocks is obtained by deriving long-run responses of the variables of interest to relevant shocks. From (6.11) we have:

$$\left(\begin{pmatrix} -1 & 1 \\ 0 & 1 \end{pmatrix} \begin{pmatrix} (1-L) & 0 \\ 0 & 1 \end{pmatrix} - \begin{pmatrix} \alpha_{11}L & 0 \\ \alpha_{21}L & 0 \end{pmatrix} \right) \begin{pmatrix} (y_t - x_t) \\ \Delta x_t \end{pmatrix} = \begin{pmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} v_{1t} \\ v_{2t} \end{pmatrix},$$

from which long-run responses are obtained by setting $L = 1$ and by inverting the matrix pre-multiplying variables in the stationary representation of VAR

$$\begin{pmatrix} (y_t - x_t) \\ \Delta x_t \end{pmatrix} = \begin{pmatrix} -\alpha_{11} & 1 \\ -\alpha_{21} & 1 \end{pmatrix}^{-1} \begin{pmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} v_{1t} \\ v_{2t} \end{pmatrix} \tag{6.12}$$

$$\begin{pmatrix} (y_t - x_t) \\ \Delta x_t \end{pmatrix} = \begin{pmatrix} \frac{-b_{11}+b_{21}}{\alpha_{11}-\alpha_{21}} & -\frac{b_{12}-b_{22}}{\alpha_{11}-\alpha_{21}} \\ \frac{-\alpha_{21}b_{11}+\alpha_{11}b_{21}}{\alpha_{11}-\alpha_{21}} & \frac{-\alpha_{21}b_{12}+\alpha_{11}b_{22}}{\alpha_{11}-\alpha_{21}} \end{pmatrix} \begin{pmatrix} v_{1t} \\ v_{2t} \end{pmatrix}. \tag{6.13}$$

Thus v_{2t} can be identified as the transitory shock by imposing the following restriction:

$$-\alpha_{21}b_{12} + \alpha_{11}b_{22} = 0$$

which, given knowledge of the $\boldsymbol{\alpha}$ parameters from the cointegration analysis, provides the just-identifying restriction for the parameters in $\mathbf{B}$. Note that, there is one case in which this identification is equivalent to the Choleski ordering, the case in which $\alpha_{11} = 0$. Note that this is the case in which Δy_t is weakly exogenous for the estimation of b_{21}. An application of this identifying scheme is provided in Favero, Giavazzi and Spaventa (1997) where the procedure is implemented to separate international from local factors in the determination of interest rates fluctuations.

6.3 Why shocks?

Having identified the 'monetary rule' by proposing an explicit solution to the problem of the endogeneity of money, the VAR approach concentrates on deviations from the rule. Deviations from the rule are obtained either by changing the systematic component of monetary policy or by considering exogenous shocks, which leave systematic monetary policy unaltered. In the former case the deviation from the rule is obtained by changing some parameters in the $\mathbf{A}$ matrix describing the simultaneous relations among variables, while in the latter case the parameters in the matrices $\mathbf{A}$ and $\mathbf{B}$ are left unaltered. Consider for example the case of Federal funds targeting. The first type of deviations is obtained by modifying the response of the Federal funds rates to macroeconomic conditions, i.e. fluctuations in output, commodity prices and the consumer price index, while the second type of deviations is obtained by considering an exogenous shock which does not alter the response of the monetary policy-maker to macroeconomic conditions. VAR modelers have exclusively concentrated on simulating shocks, leaving the systematic component of monetary policy unaltered.

The VAR approach to the monetary transmission mechanism has been criticized on the basis that it views central banks as 'random number generators'. This is incorrect, since monetary policy rules are explicitly estimated in structural VAR models. However, the focus is not on rules but on deviations from rules, since only when central banks deviate from their rules it becomes possible to collect interesting information on the response of macroeconomic variables to monetary policy impulses, to compare with the predictions of the alternative theoretical models. In fact, deviations from monetary policy rules provide researchers with the best opportunity to detect the response of macroeconomic variables to monetary impulses unexpected by the market. The first chain of most models of the monetary transmission mechanism links the policy rates to the term structure of the interest rates and the most popular model of the term structure, the expectational model, predicts that the term structure does not generally react to expected monetary impulses. The monetary impulses relevant to the transmission analysis are therefore structural shocks in (6.1).

Recently, McCallum (1999) has criticized the choice of VAR modelers of concentrating on shocks which leave the systematic component of monetary policy unaltered. It is argued that the emphasis on the shock component is misplaced because the unsystematic portion of policy-instrument variability is small in relation to the variability of the systematic component.

> ...Indeed, it is conceivable that the policy behaviour could be virtually devoid of any unsystematic component. In the limit, that is, the variance of the shock component could approach zero. But this would not imply that monetary policy is unimportant for price level behaviour, central bank's main responsibility... (McCallum 1999:5).

The simulation of systematic monetary policy requires, for robustness to the Lucas critique, the specification of a forward-looking model in which 'deep parameters' are identified independently from nuisance parameters describing expectations formation and dependent on the policy regime. This is what McCallum (1999) effectively does in a series of papers where the impact of modifications in the monetary policy-maker reaction function is dynamically simulated.

However, it is important to note that McCallum's work aims not at model selection but rather at model simulation. The question of using the empirical evidence to judge between different theoretical models is not addressed in McCallum's work, based on a specific model.

If VAR models are, instead, used to describe the empirical evidence relevant to the choice between alternative theoretical models, then there is a possible defence of the choice of concentrating on shocks rather than on the systematic components of monetary policy. Such a defence is related to the Lucas critique.

Consider the following data generating process:

$$\begin{aligned} y_t &= a_1 m_{t+1}^e + a_2 y_{t-1} + u_{1t} \\ m_t &= b_0 + b_1 y_{t-1} + b_2 m_{t-1} + u_{2t}, \end{aligned}$$

where y is the macroeconomic variable and m is the monetary policy variable.

The DGP is the relevant theoretical model, which is unknown to the empirical researcher. The empirical researcher tries instead to describe the empirical relation between the monetary instruments and the macroeconomic variables by specifying the following structural VAR:

$$
\begin{aligned}
y_t &= c_o + c_1 m_t + c_2 y_{t-1} + v_{1t} \\
m_t &= b_0 + b_1 y_{t-1} + b_2 m_{t-1} + v_{2t},
\end{aligned} \tag{6.14}
$$

in which the restrictions $c_o = a_1 b_0$, $c_1 = a_1 b_2, c_2 = a_2 + a_1 b_1$ hold.

Equation (6.14) is not viable for econometric evaluation of systematic monetary policy, in that the parameters in the equation for y cannot be kept constant when the systematic component of the monetary policy rule is altered. However, the simulation of the dynamic impact of a monetary policy shock identified *à la* Choleski ordering m first is still viable in that it is performed while keeping all parameters constant.

Note that this small example reiterates the importance of estimating parameters in structural VAR models by concentrating on a single policy regime, in fact regime shifts require different parameterization. This reflects the fact, noted by Keating(1990), that a forward-looking DGP imposes cross-equation restrictions on parameters in a VAR.

Last, but certainly not least, a crucial assumption in structural VAR modelling is that structural shocks are linear combinations of the residuals in reduced form VAR models. Lippi and Reichlin (1993) argue that modern macroeconomic models which are linearized into dynamic systems tend to include non-invertible moving average components and structural shocks are therefore not identifiable. In fact, the linearized modern macroeconomic models of the monetary transmission mechanism, which we shall describe in the next chapters, deliver short VARs. In such models structural shocks are combinations of the residuals in the reduced form VARs (the Wold innovations) and the Lippi–Reichlin critique does not seem to be applicable (for a further discussion of this point see Amisano and Giannini 1996).

6.4 Description of VAR models

After the identification of structural shocks of interest, the properties of VAR models are described using impulse response analysis, variance decomposition and historical decomposition. Consider a structural VAR model for a generic vector $\mathbf{y}_t$, containing m variables:

$$\mathbf{A}_0 \mathbf{y}_t = \sum_{i=1}^{p} \mathbf{A}_i \mathbf{y}_{t-i} + \mathbf{B}\mathbf{v}_t,$$

which we can rewrite as:

$$
\begin{aligned}
\left[\mathbf{A}_0 - \mathbf{A}(L)\right] \mathbf{y}_t &= \mathbf{B}\mathbf{v}_t \\
\mathbf{A}(L) &= \sum_{i=1}^{p} \mathbf{A}_i L^i.
\end{aligned}
$$

By inverting $[\mathbf{A}_0 - \mathbf{A}(L)]$ (under the assumption of invertibility of this polynomial) we obtain the moving average representation for our VAR process:

$$\begin{aligned} \mathbf{y}_t &= \mathbf{C}(L)\,\mathbf{v}_t, \\ \mathbf{y}_t &= \mathbf{C}_0\mathbf{v}_t + \mathbf{C}_1\mathbf{v}_{t-1} + \ldots + \mathbf{C}_s\mathbf{v}_{t-s}, \\ \mathbf{C}(L) &= [\mathbf{A}_0 - \mathbf{A}(L)]^{-1}, \\ \mathbf{C}_0 &= \mathbf{A}_0^{-1}\mathbf{B}. \end{aligned} \tag{6.15}$$

To illustrate the concept of an *impulse response function,* we interpret the generic matrix $\mathbf{C}_s$ within the moving average representation as follows:

$$\mathbf{C}_s = \frac{\partial \mathbf{y}_{t+s}}{\partial \mathbf{v}_t}.$$

The generic element $\{i, j\}$ of matrix $\mathbf{C}_s$ represents the impact of a shock hitting the j-th variable of the system at time t on the i-th variable of the system at time $t+s$. As s varies we have a function describing the response of variable i to an impulse in variable j. For this function of partial derivative to be meaningful we must allow that a shock to variable j occurs while all other shocks are kept to zero. Of course this is allowed for structural shocks, as they are identified by imposing they are orthogonal to each other. Note, however that the concept of an impulse response function is not applicable to reduced form VAR innovations, which, in general, are correlated to each other.

Historical decomposition is obtained by using the structural MA representation to separate series in the components (orthogonal to each other) attributable to the different structural shocks.

Finally *forecasting error variance decomposition* (FEVD) is obtained from (6.15) by deriving the error in forecasting $\mathbf{y}_s$ period in the future as:

$$(\mathbf{y}_{t+s} - E_t\mathbf{y}_{t+s}) = \mathbf{C}_0\mathbf{v}_t + \mathbf{C}_1\mathbf{v}_{t-1} + \ldots + \mathbf{C}_s\mathbf{v}_{t-s}$$

from which we can construct the variance of such forecasting error as:

$$Var\,(\mathbf{y}_{t+s} - E_t\mathbf{y}_{t+s}) = \mathbf{C}_0 I\mathbf{C}_0' + \mathbf{C}_1 I\mathbf{C}_1' + \ldots + \mathbf{C}_s I\mathbf{C}_s'$$

from which we can compute the share of the total variance attributable to the variance of each structural shock. Note again that such composition makes sense only if shocks are orthogonal to each other. Only in this case we can write the variance of the total forecasting error as a sum of variances of the single shocks (as the covariance terms are zero following the orthogonality property of structural shocks).

To illustrate the three concepts consider the following bivariate VAR, in which structural parameters have been identified and estimated via a Choleski decomposition:

$$\begin{pmatrix} y_{1t} \\ y_{2t} \end{pmatrix} = \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix} \begin{pmatrix} y_{1t-1} \\ y_{2t-1} \end{pmatrix} + \begin{pmatrix} b_{11} & 0 \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} v_{1t} \\ v_{2t} \end{pmatrix}.$$

The MA representation is

$$\begin{pmatrix} y_{1t} \\ y_{2t} \end{pmatrix} = \begin{pmatrix} b_{11} & 0 \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} v_{1t} \\ v_{2t} \end{pmatrix} + \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix} \begin{pmatrix} b_{11} & 0 \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} v_{1t-1} \\ v_{2t-1} \end{pmatrix}$$
$$+ ... + \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix}^s \begin{pmatrix} b_{11} & 0 \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} v_{1t-s} \\ v_{2t-s} \end{pmatrix},$$

from which impulse response functions, historical decomposition and forecasting error variance decomposition are immediately obtained.

6.5 Monetary policy in closed economies

Cumulative work on the analysis of the monetary transmission mechanism in the US led to the specification of a VAR system which has by now become the standard reference model. We have already seen and discussed this benchmark specification which contains six variables: gross domestic product (GDP), the consumer price index (P) and the commodity price level (Pcm) together with the federal funds rate (FF), the quantity of total bank reserves (TR) and the amount of non-borrowed reserves (NBR).

It is interesting to see how the specification of the benchmark model has developed over time. Initially models were estimated on a rather limited set of variables, i.e. prices, money and output, and identified imposing a diagonal form on the matrix **B** and a lower triangular form on the matrix **A** with money coming last in the ordering of the variables included in the VAR (Choleski identification). This first type of models is discussed in Leeper, Sims and Zha (1996), we replicate their results on the dataset lszusa.wf1. The underlying structural model is specified as follows:

$$\mathbf{A}_0 \mathbf{y}_t = \sum_{i=1}^{k} \mathbf{A}_i \mathbf{y}_{t-i} + \mathbf{B} \boldsymbol{v}_t \tag{6.16}$$

$$\mathbf{y}_t = \begin{bmatrix} p_t \\ y_t \\ m_t \end{bmatrix}, \qquad \mathbf{A}_0 = \begin{bmatrix} 1 & 0 & 0 \\ a_{21} & 1 & 0 \\ a_{31} & a_{32} & 1 \end{bmatrix}$$

$$\mathbf{B} = \begin{bmatrix} b_{11} & 0 & 0 \\ 0 & b_{22} & 0 \\ 0 & 0 & b_{33} \end{bmatrix}, \qquad \boldsymbol{\nu}_t \sim n.i.d.\,(0, I).$$

All variables are expressed in logarithms. Identification is Choleski-type with money ordered last. This is a model geared to deliver monetary policy shocks, so the identification of shocks to LP and LY does not matter. As Leeper, Sims and Zha did, we have estimated the model on the sample 1960:1–1996:3, including six lags of each variable. The impulse responses obtained are reported in Figure 6.2.

Prices slowly react to monetary policy, output responds in the short run, in the long run (from two years after the shock onwards) prices start adjusting

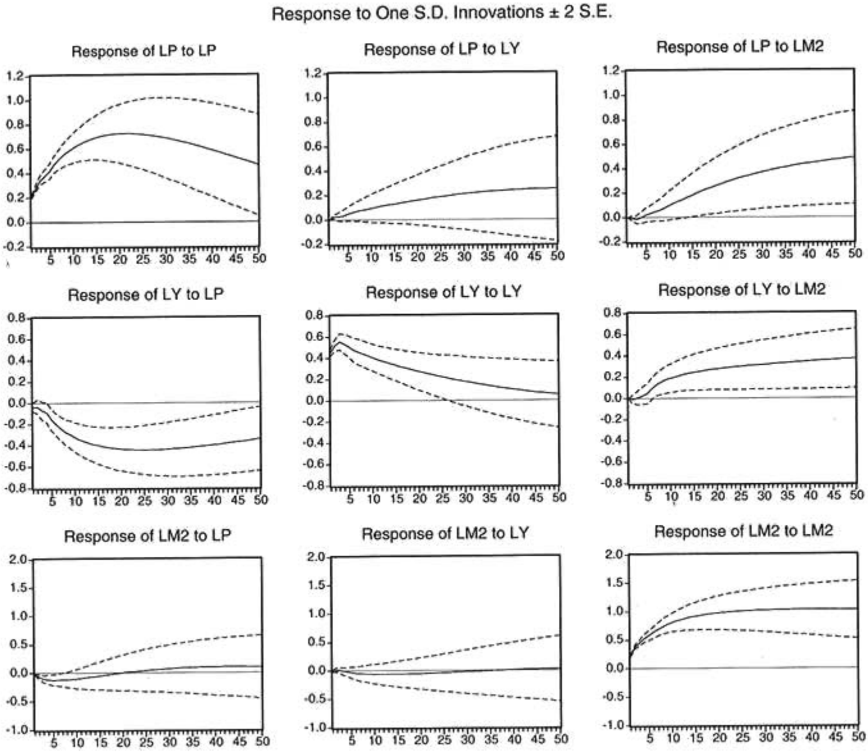

FIG. 6.2. Impulse response functions in a three-variables VAR of the MTM

and the significant effect on output vanishes. There is no strong evidence for the endogeneity of money. This is easily checked by looking at the estimated parameters in $\mathbf{A}_0$ and by analysing FEVD in Figure 6.3.

Macroeconomic variables play a very limited role in explaining the variance of the forecasting error of money, while money instead plays an important role in explaining fluctuations of both the macroeconomic variables.

Sims (1980) extended the VAR to include the interest rate on Federal funds ordered just before money as a penultimate variable in the Choleski identification. The idea is to see the robustness of the above results after identifying the part of money which is endogenous to the interest rate. Impulse response functions are modified as in Figure 6.4, while FEVD is modified as in Figure 6.5.

Impulse response functions and FEVD raise a number of issues.

1. Though little of the variation in money is predictable from past output and prices, a considerable amount becomes predictable when past short-term interest rates are included in the information set.

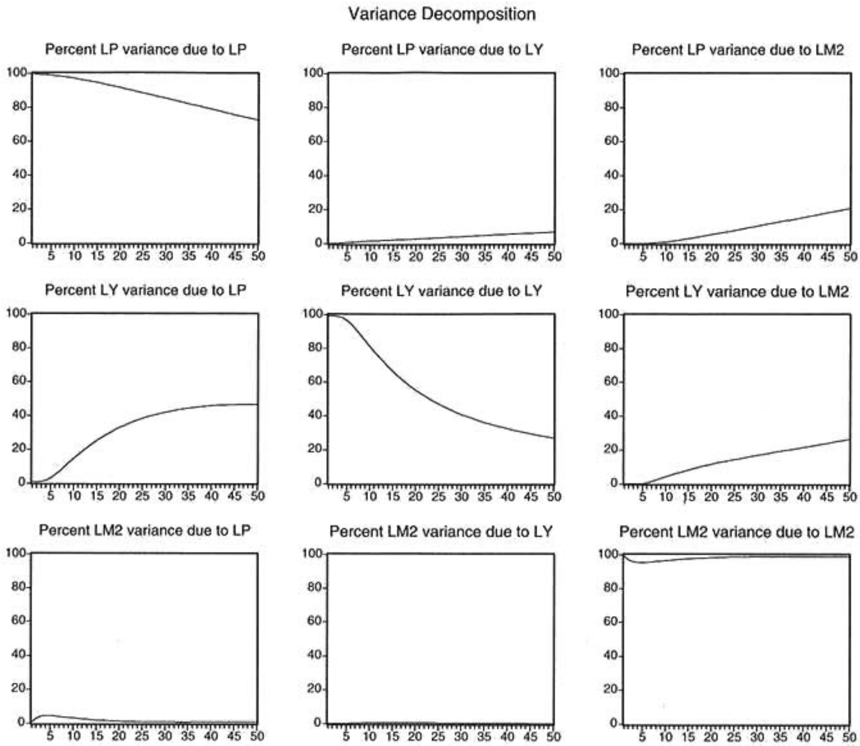

FIG. 6.3. FEVD in a three-variables VAR of the MTM

2. It is difficult to interpret the behaviour of money as driven by money supply shocks. The response to money innovations gives rise to the 'liquidity puzzle': the interest rate declines very slightly contemporaneously in response to a money shock to start increasing afterwards.

3. There are also difficulties with interpreting shocks to interest rates as monetary policy shocks. The response of prices to an innovation in interest rates gives rise to the 'price puzzle': prices increase significantly after an interest rate hike. An accepted interpretation of the liquidity puzzle relies on the argument that the money stock is dominated by demand rather than supply shocks. Moreover, the interpretation of money as demand shocks driven is consistent with the impulse response of money to interest rates. Note also that, even if the money stock were to be dominated by supply shocks, it would reflect both the behaviour of central banks and the banking system. For both these reasons the broad monetary aggregate has been substituted by narrower aggregates, bank reserves, on which it is easier to identify shocks mainly driven by the behaviour of the monetary policy

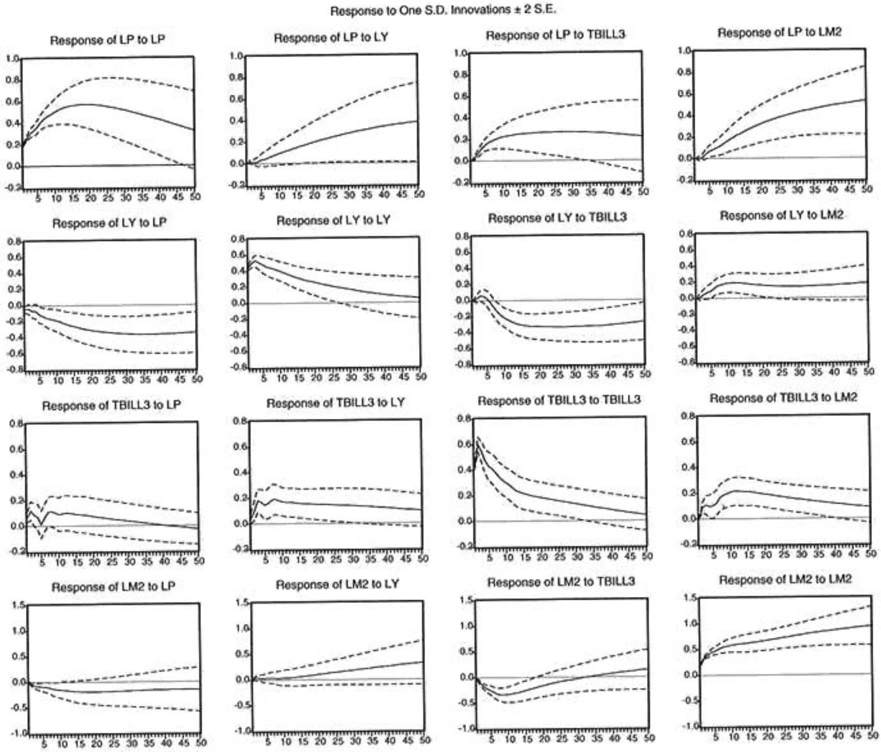

FIG. 6.4. Impulse responses in a four-variables VAR of the MTM

maker. The 'price puzzle' has been attributed to mis-specification of the four-variables VAR used by Sims. Suppose that there exists a leading indicator for inflation to which the Fed reacts. If such a leading indicator is omitted from the VAR, then we have an omitted variable positively correlated with inflation and interest rates. Such omission makes the VAR mis-specified and explains the positive relation between prices and interest rates observed in the impulse response functions. It has been observed (see Christiano, Eichenbaum and Evans 1996, Sims 1996) that the inclusion of a Commodity Price Index in the VAR solves the 'price puzzle'.

Our brief historical record of the empirical analysis of closed economy VAR of the MTM has brought us to the justification of the six variables included in what is by now known as the benchmark VAR model. We have already discussed its identification, let us now examine impulse response functions derived by using the Federal fund targeting identifying restrictions and reported in Figure 6.6.

The evidence reported in the impulse response functions represents the relevant stylized facts to include in theoretical models of the MTM. This kind

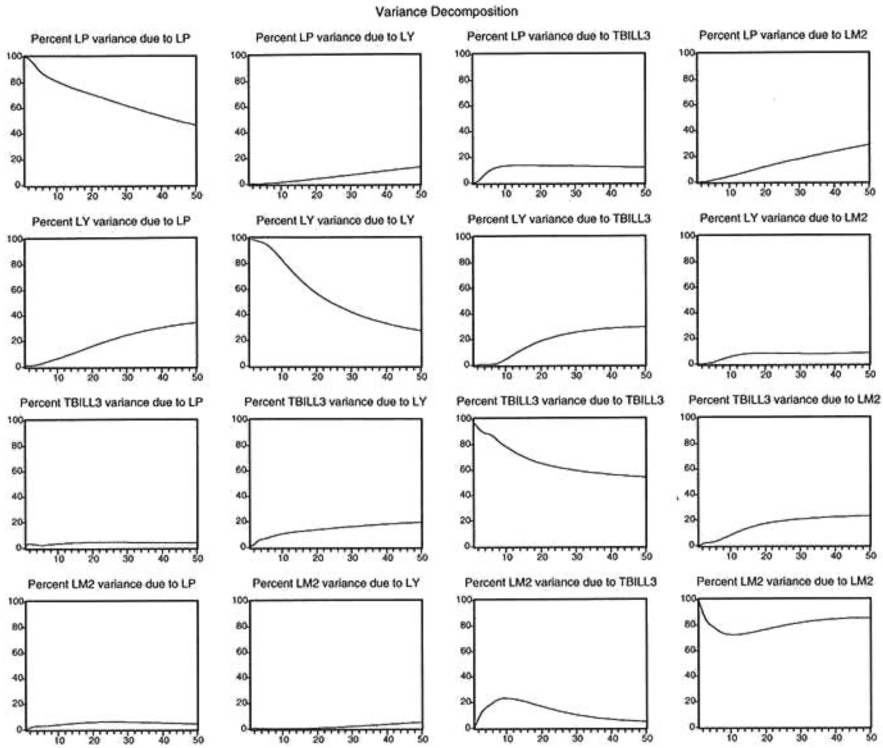

FIG. 6.5. FEVD in a four-variable VAR of the MTM

of evidence has established that plausible models of the monetary transmission mechanism should be consistent at least with the following evidence on price, output and interest rates: *(i)* the aggregate price level initially responds little; *(ii)* interest rates initially rise, and *(iii)* aggregate output initially falls, with a j-shaped response, and a zero long-run effect of the monetary impulse.

6.6 Monetary policy in open economies

Various papers have examined the effects of monetary shocks in open economies, but this strand of literature has been distinctly less successful in providing accepted empirical evidence than the VAR approach in closed economies.

The first results have been provided by Eichenbaum and Evans (1995), using an open-economy VAR with the following structure:

$$\mathbf{A}_0\mathbf{y}_t = \sum_{i=1}^{k}\mathbf{A}_i\mathbf{y}_{t-i} + \mathbf{B}\boldsymbol{v}_t, \tag{6.17}$$

where $\mathbf{y}_t = \left\{\mathrm{Y}_t^{US}, \mathrm{P}_t^{US}, \mathrm{NBRX}_t^{US} \text{ (or } \mathrm{FF}_t), \mathrm{Y}_t^{FOR}, \mathrm{P}_t^{FOR}, \mathrm{R}_t^{FOR}, \mathrm{e}_t \text{ (or } \mathrm{q}_t)\right\}$.

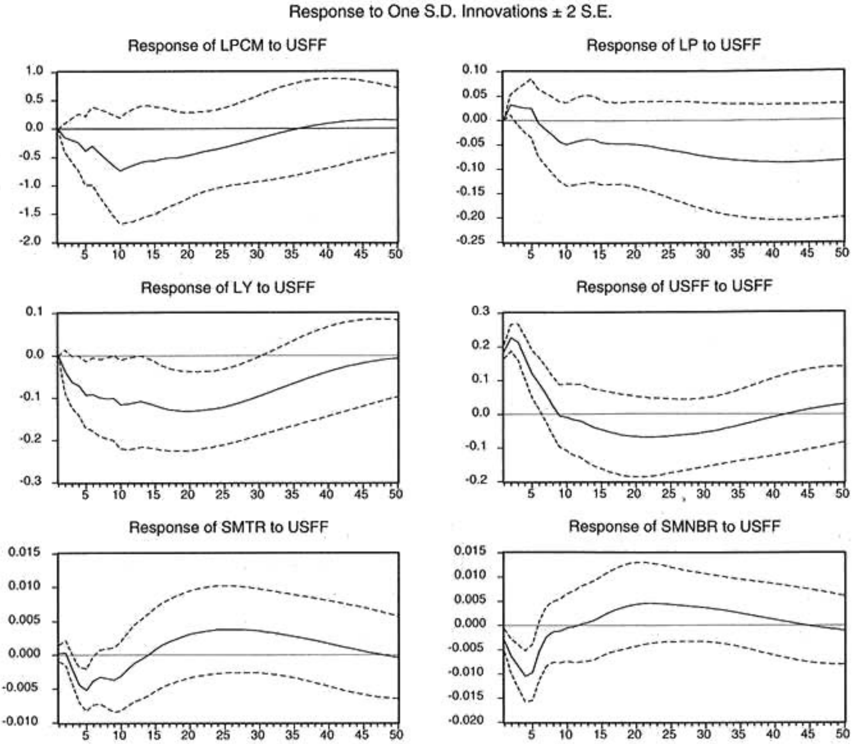

FIG. 6.6. Impulse response functions in a six-variables VAR of the MTM

$\mathrm{Y}^{US}, \mathrm{P}^{US}$ are logarithms of US output and price, NBRX^{US} is the ratio of non-borrowed to total reserves (the appropriate variable from which one can extract monetary policy shocks under a regime of non-borrowed reserves targeting). FF is the Federal Funds rate, which is considered as an alternative to NBRX^{US}, and it is an informative variable for the extraction of monetary policy shocks under the regime of interest rate targeting; $\mathrm{Y}^{FOR}, \mathrm{P}^{FOR}$, and R^{FOR} are respectively the logarithms of output, prices, and the level of short-term interest rate in the foreign country; e is the nominal bilateral exchange rate, while q is the real bilateral exchange rate. The matrix $\mathbf{B}$ is diagonal and $\mathbf{A}$ is lower-triangular. The empirical analysis is implemented by considering in turn, as a foreign country, each of the G7 countries on a sample of monthly data from 1974:1 to 1990:5. The following evidence emerges: (i) a restrictive US monetary policy shock generates a significant and persistent appreciation of the US dollar; (ii) a restrictive US monetary policy shock generates a significant and persistently larger effect on the domestic interest rate with respect to the foreign rate; (i) and (ii) imply a sharp deviation from the uncovered interest parity condition in favour of

US dollar-denominated investments (the 'forward-discount puzzle'); (*iii*) identified US monetary policy shocks are not different from the shocks derived within closed-economy VARs (*iv*) the closed-economy response of US prices and output to monetary policy shocks is robust to the extension of the VAR to the open economy; (*v*) a restrictive foreign monetary policy shock generates an appreciation of the US dollar (the 'exchange-rate puzzle'); and (*vi*) the response of the real exchange rate to the US and foreign monetary policy shocks does not differ significantly from the response of the nominal exchange rate. Such evidence is substantially confirmed by the work of Schlagenhauf and Wrase (1995), who consider a very similar specification for the G5 countries over the sample 1972:2–1990:2, using quarterly data. Some considerations should help to interpret the above results.

First, the empirical models are estimated over samples including shifts in the US and foreign monetary policy regimes, therefore, parameter instability is a potential problem.

Second, the extension to the open economy features the omission from the VAR of the commodity price index and of the monetary variables irrelevant to the extraction of the policy shocks. While the simplification of the monetary block is sustainable in the light of the absence of contemporaneous feedback between the informative variables and the other monetary variables under the chosen identification schemes, the omission of the commodity price index is unjustifiable as it leads to the same mis-specification as in the closed economy model for US monetary policy shocks. Moreover, such an omission might well also bias the identification of the foreign monetary policy shocks if the commodity price index is regarded as a leading indicator of inflation by the foreign policy-maker.

In conclusion, it is passible to argue that the observed puzzles might depend on the incorrect specification of the VAR.

Third, on the identification scheme, while some rationale can be provided for a quasi-recursive scheme in closed economies, a similar justification is much harder to provide in open economies. In fact, the recursive identification scheme with the exchange rate ordered last implies that neither the US nor the foreign monetary authority react contemporaneously to exchange rate fluctuations. This assumption seems sustainable for the US (the Fed benign neglect for the dollar) but it is certainly heavily questionable when the foreign countries are considered, as they are much more open economies than the US.

The failure of the recursive identification scheme could also be responsible for the observation of the puzzles.

In fact, most of the recent empirical work aims at breaking such a recursive structure.

Kim and Roubini (1997) have attempted this by introducing a structural identification by the explicit consideration of a money demand and supply functions. Their specification for the generic non-US country is:

$$\mathbf{A}_0\mathbf{y}_t = \sum_{i=1}^{k}\mathbf{A}_i\mathbf{y}_{t-i} + \mathbf{B}\boldsymbol{v}_t \tag{6.18}$$

$$\mathbf{y}_t = \begin{bmatrix} \mathrm{OPW}_t \\ \mathrm{FF}_t \\ \mathrm{Y}_t^{FOR} \\ \mathrm{P}_t^{FOR} \\ \mathrm{M}_t^{FOR} \\ \mathrm{R}_t^{FOR} \\ \mathrm{e}_t \end{bmatrix}, \quad \mathbf{A}_0 = \begin{bmatrix} 1 & 0 & 0 & 0 & 0 & 0 & 0 \\ a_{21} & 1 & 0 & 0 & 0 & 0 & 0 \\ a_{31} & 0 & 1 & 0 & 0 & 0 & 0 \\ a_{41} & 0 & a_{43} & 1 & 0 & 0 & 0 \\ 0 & 0 & a_{53} & a_{54} & 1 & a_{56} & 0 \\ a_{61} & 0 & 0 & 0 & a_{65} & 1 & a_{67} \\ a_{71} & a_{72} & a_{73} & a_{74} & a_{75} & a_{76} & 1 \end{bmatrix} \tag{6.19}$$

$$\boldsymbol{\nu}_t \sim n.i.d.\,(0, I) \tag{6.20}$$

with R^{FOR} denoting the short-term non-US policy rate, M^{FOR} a monetary aggregate (M0 or M1), P^{FOR} the log of consumer price index, Y^{FOR} the log of industrial production, OPW the world index of oil price in dollars, FF the Federal funds rate, and e the nominal exchange rate against the dollar. $\mathbf{B}$ is a diagonal matrix. The model is estimated over the sample 1974:7–1992:5, on monthly data. The main differences between the proposed structural identification and the recursive identification schemes can be understood by analysing the $\mathbf{A}_0$ matrix. Some elements under the principal diagonal are set to zero to allow the introduction of simultaneous feedbacks between demand and supply for money and central bank behaviour and exchange rates. The estimated model is over-identified with 23 parameters estimated in the $\mathbf{A}_0$ and $\mathbf{B}$ matrix, out of a possible maximum of 28. The over-identifying restrictions are tested and not rejected. The identifying restrictions are rather standard. The US economy is taken as exogenous and the exchange rate does not enter in the Fed reaction function, US output and prices are not included in the VAR, while a simultaneous feedback is allowed between money demand and supply (the central bank rule). According to this rule, contemporaneous US interest rate movements are relevant to the foreign central bank only if they affect the exchange rate. Only the exchange rate is allowed to contemporaneously react to news in all the other variables.

Unfortunately the coefficients in the $\mathbf{A}_0$ matrix are estimated rather imprecisely. If we consider the case of the US versus Germany, the only significant parameters in the matrix are a_{53} and a_{72}. The first parameter is difficult to interpret, given that the identification scheme does not explicitly address aggregate demand and supply shocks, while the point estimate of the second parameter implies an appreciation of the dollar against the Deutschmark in response to a US restrictive monetary policy. The potential simultaneous feedback between foreign monetary policy and the exchange rate does not seem to be empirically relevant. However, all the puzzles disappear and the empirical results for the impulse response functions are broadly in line with results from the US closed economy model. Given that this VAR included some proxy for the commodity price index the evidence cannot be decisive on the source of the 'puzzles', although the fact that the simultaneous feedback between foreign interest rates and the exchange rate is not significant is consistent with attributing a substantial role to commodity prices.

Also in this case the sample considered spans different regimes, moreover, this methodology brings broad monetary aggregates back into the specification. Interestingly, money is now used to extract demand rather than supply shocks, however, the specification of money demand implicit in the VAR might not be rich enough to capture the dynamic in the data. As pointed out by Faust and Whiteman (1997), single equation work by Hendry and colleagues on money demand has clearly shown the importance of including in the model the opportunity cost of holding money, which is often a spread between the interest rates. Interest spreads capturing the opportunity cost of holding money are never included in VAR models of the MTM. An identification similar to the one adopted by Kim and Roubini (1997) is the one proposed for the Canadian case by Cushman and Zha (1997), who aid the structural identification by explicitly introducing the trade sector into the model.

An interesting alternative approach to the identification of the simultaneous feedback between non-US interest rates and exchange rates is proposed by Smets (1996)–(1997). Smets considers the following structural model for non-US countries:

$$\mathbf{A}_0\mathbf{y}_t = \sum_{i=1}^{k}\mathbf{A}_i\mathbf{y}_{t-i} + \mathbf{B}\boldsymbol{v}_t \tag{6.21}$$

$$\mathbf{y}_t = \begin{bmatrix} \Delta y_t \\ \Delta p_t \\ R_t \\ \Delta e_t \end{bmatrix}, \quad \mathbf{B} = \begin{bmatrix} b_{11} & b_{12} & 0 & 0 \\ \frac{-(k_{11}b_{11}+k_{13}b_{31}+k_{14}b_{41})}{k_{12}} & b_{22} & 0 & 0 \\ b_{31} & b_{32} & 1 & 0 \\ b_{41} & b_{42} & b_{43} & b_{44} \end{bmatrix}$$

$$\mathbf{A}_0 = \begin{bmatrix} 1 & 0 & 0 & 0 \\ 0 & 1 & 0 & 0 \\ 0 & 0 & 1-\omega & \omega \\ 0 & 0 & 0 & 1 \end{bmatrix}$$

where Δy_t is output growth, Δp_t is inflation, R is a short-term interest rate and Δe_t is the exchange rate appreciation. No US variables are introduced, and the commodity price index is also excluded. However, Smets is more ambitious and aims at identifying both macroeconomics and monetary shocks. Three type of restrictions are imposed. First the semi-structural restrictions, macro variables do not react contemporaneously to monetary variables. Second, macroeconomic supply shocks are identified for macroeconomic demand shocks by assuming that the long-run effect of demand shocks on output is zero. Third, monetary policy shocks are identified from exchange rate shocks by assuming that the central bank reacts proportionally to interest rate and exchange rate developments (short-term Monetary Condition Indexes). Macroeconomic shocks are separated into demand and supply shocks by noting that the long-run response of output to a demand shock is given by element $(1,1)$ of the matrix $\left(A_0 - \sum_{i=1}^{k} A_i\right)^{-1} B$. Given the block-recursiveness assumption on A_0, the elements of $\left(A_0 - \sum_{i=1}^{k} A_i\right)^{-1}$ relevant to determine the element (1,1) of $\left(A_0 - \sum_{i=1}^{k} A_i\right)^{-1} B$ are not functions of the A_0 matrix and therefore from the estimation of the reduced-form VAR one can retrieve all the elements in the A_i matrix and generate an identifying restriction for the structural parameters in the B matrix by setting the element (1,1) $\left(A_0 - \sum_{i=1}^{k} A_i\right)^{-1} B$ to zero. In practice, given that

$$\left(A_0 - \sum_{i=1}^{k} A_i\right)^{-1} = \begin{bmatrix} k_{11} & k_{12} & k_{13} & k_{14} \\ k_{21} & k_{22} & k_{23} & k_{24} \\ k_{31} & k_{32} & k_{33} & k_{34} \\ k_{41} & k_{42} & k_{43} & k_{44} \end{bmatrix},$$

one can show that $k_{11}, k_{12}, k_{13}, k_{14}$ are determined independently from the parameters in A_0, therefore restricting to zero the long-run effect of demand shock on output, we have $b_{21} = -\left(k_{11}b_{11} + k_{13}b_{31} + k_{14}b_{41}\right)/k_{12}$.

Lastly in the monetary block, monetary policy shocks are identified from exchange rate shocks by assuming that the appropriate indicator of exogenous monetary stance is a short-term MCI where the exchange rate and interest rate are appropriately weighted. The weights can be estimated or imposed given the knowledge of the relative weights in several central banks MCIs. This approach encompasses as a special case, the pure interest-rate targeting and the pure exchange-rate targeting. The main empirical problems with this procedure are the instability of the estimated ω and the potentially disruptive implications of mis-specification for the identification of aggregate demand and supply shocks (see Faust and Leeper 1997).

6.6.1 *Replicating the empirical evidence*

The dataset berlin.wf1 contains the relevant variables to replicate the open-economy VAR models discussed so far.

We first estimate a benchmark open-economy model for the US and the German economy without including the Commodity Price Index. The model is estimated on monthly data over the sample 1983:1–1997:12. The VAR is specified by including six lags of US industrial production, the US consumer price index, Federal funds rate, German industrial production, German consumer price index, German call money rate, and the US-dollar/Deutschmark nominal exchange rate (unit of DM for one US dollar). The choice of the sample is motivated by the need of having a single monetary policy regime for the US, featuring Fed funds targeting; see Bagliano and Favero (1998). Impulse responses for all variables to US and German monetary policy shocks are reported in Figures 6.7 and 6.8.

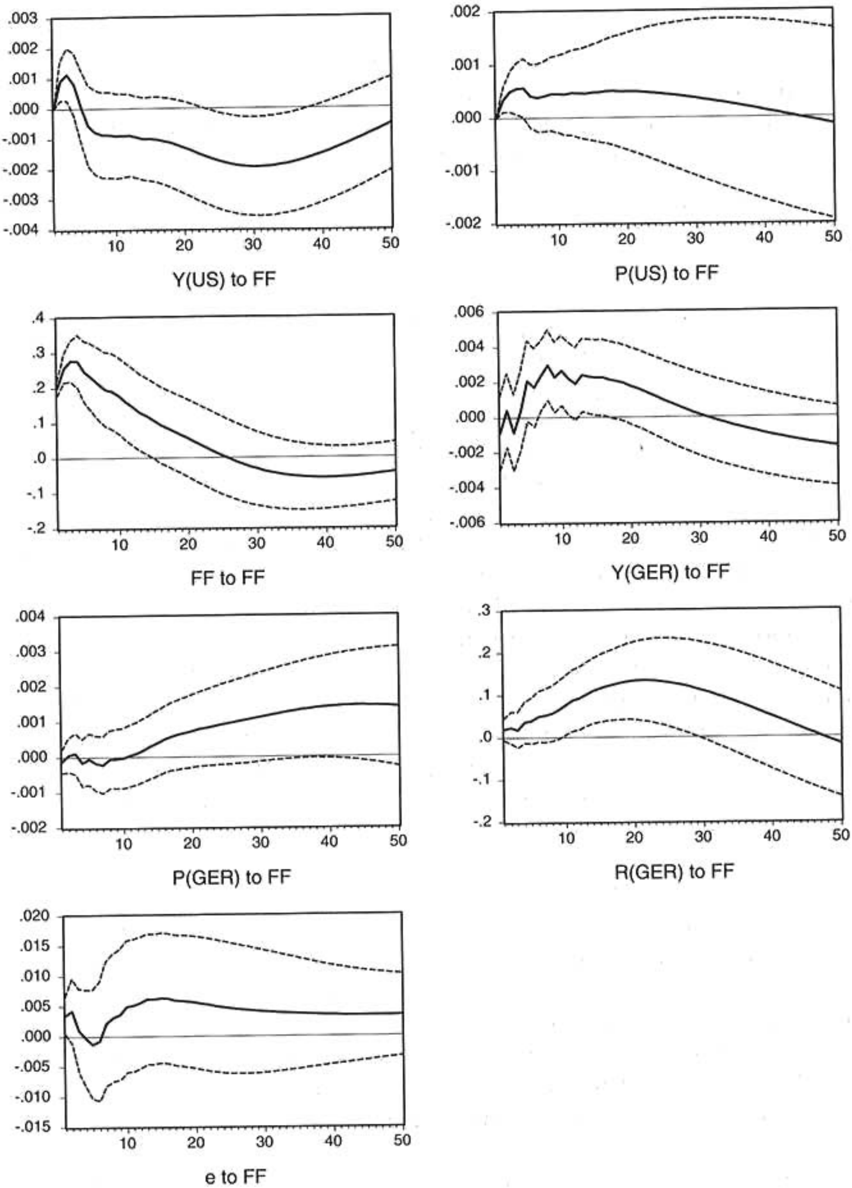

FIG. 6.7. Impulse responses to US monetary policy shock in the benchmark VAR open-economy model (dashed lines: 2 standard error bands)

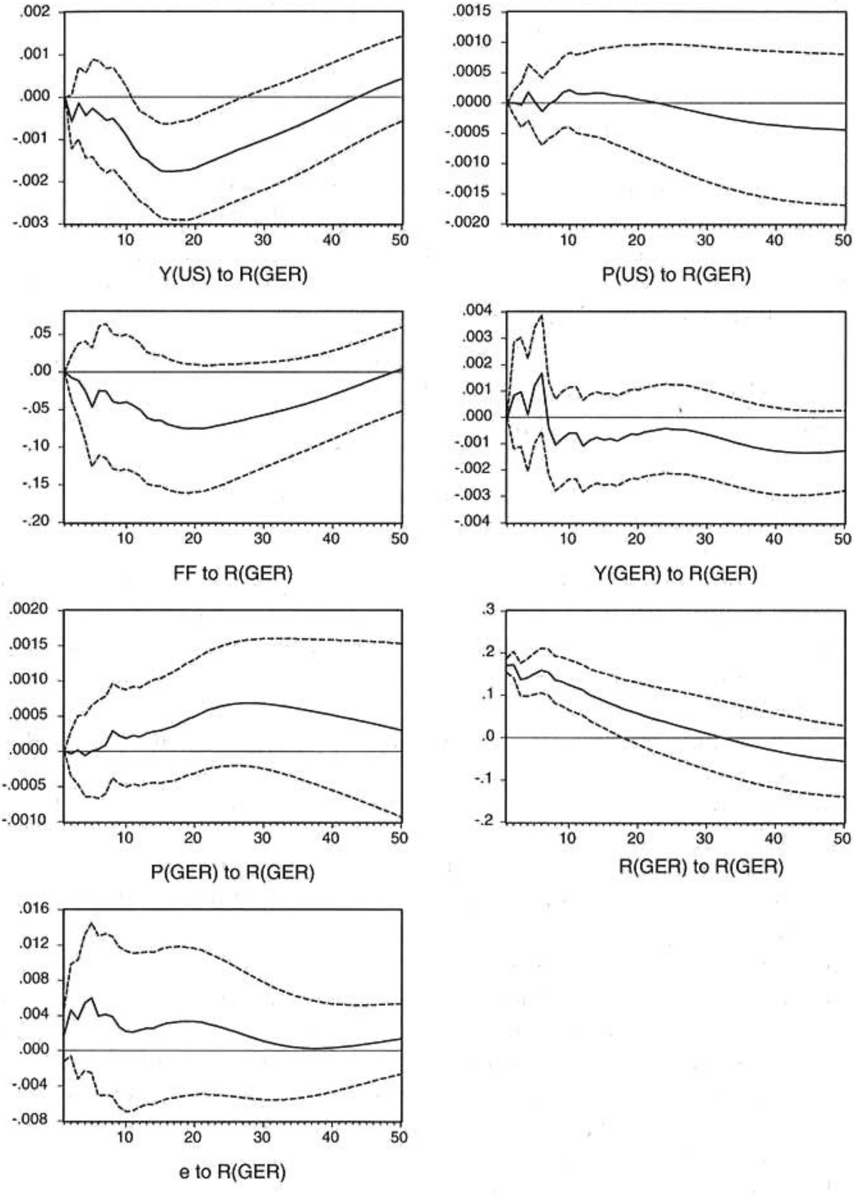

FIG. 6.8. Impulse responses to German monetary policy shock in the benchmark VAR open-economy model (dashed lines: 2 standard error bands)

We have confirmation of all the facts and puzzles observed in the literature. The analysis of the contemporaneous feedback between variables within the recursive specification provides evidence on the endogeneity of US monetary policy, which reacts significantly to internal conditions, and of the German monetary policy which reacts to both internal conditions and, less significantly to US monetary policy. The exchange rate reacts contemporaneously significantly to US monetary policy (positive interest rate shock in the US induces appreciation of the US dollar *vis-à-vis* the DM) and to macroeconomic conditions in the US and Germany (a positive shock in US industrial production and in German price lead contemporaneously to an appreciation of the US dollar) both are not contemporaneously significantly affected by German monetary policy.

The analysis of the responses to monetary impulses in the US and Germany confirms all the main findings of the literature namely:

- a significant U-shaped response of US output to US monetary policy;
- the existence of a price puzzle both for the US and Germany;
- the existence of a forward-discount puzzle generated by US monetary policy restriction;
- exchange-rate puzzle for German monetary policy shock.

6.6.2 *Omitted variables*

Our analysis of VAR models of the monetary transmission mechanism has shown the crucial importance of the Commodity Price Index in the derivation of monetary policy shocks. The arguments made for the inclusion of this variable in closed-economy VAR of the monetary transmission mechanism are also compelling for open-economy VAR. Possibly puzzles observed in open economies are related to mis-specification, via the omission of a commodity price index in the benchmark open-economy VAR. We consider this potential explanation, by concentrating on an open-economy VAR model linking the US and the German economy.

We include a commodity price index by keeping the Choleski identification and considering Pcm as a macroeconomic variables influencing both US and German monetary policy. The new impulse responses are reported in Figures 6.9 and 6.10. The results in the two figures show that the inclusion of commodity prices solves the price puzzle for both countries, moreover also the forward-discount bias puzzle and the exchange-rate puzzle tend to disappear. Finally, although we do not observe a symmetric contemporaneous effect of the US and German monetary policy on the exchange rate, the impulse response functions of the exchange rate to the two monetary shocks over a horizon of four year show a remarkable degree of symmetry.

Although the inclusion of commodity prices seems clearly relevant, we have opened the issue of potential simultaneity between the exchange rate and the policy rate in a small open economy and not yet addressed it. We shall consider this issue by looking at the more general problem of assessing the reliability of the measurement of monetary policy with VAR models.

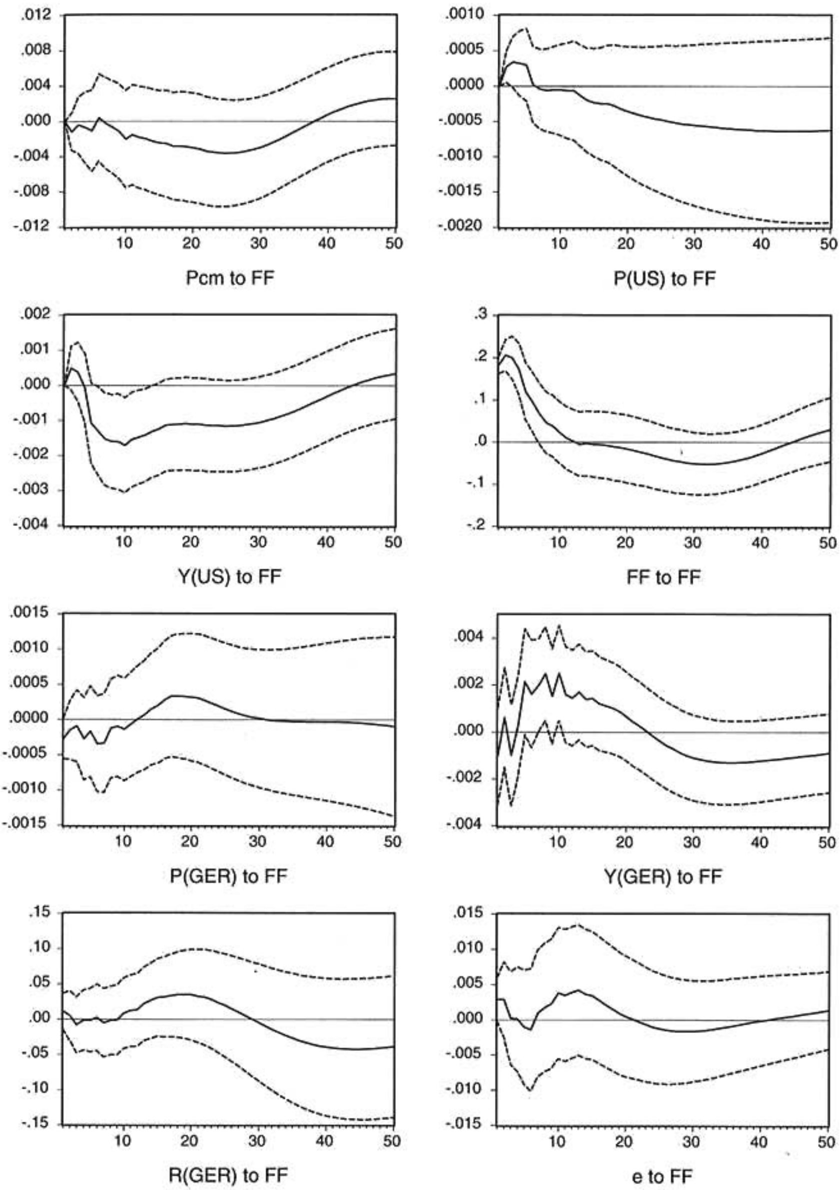

FIG. 6.9. Impulse responses to the US monetary policy shock in the benchmark VAR with commodity price index (dashed lines: 2 standard error bands)

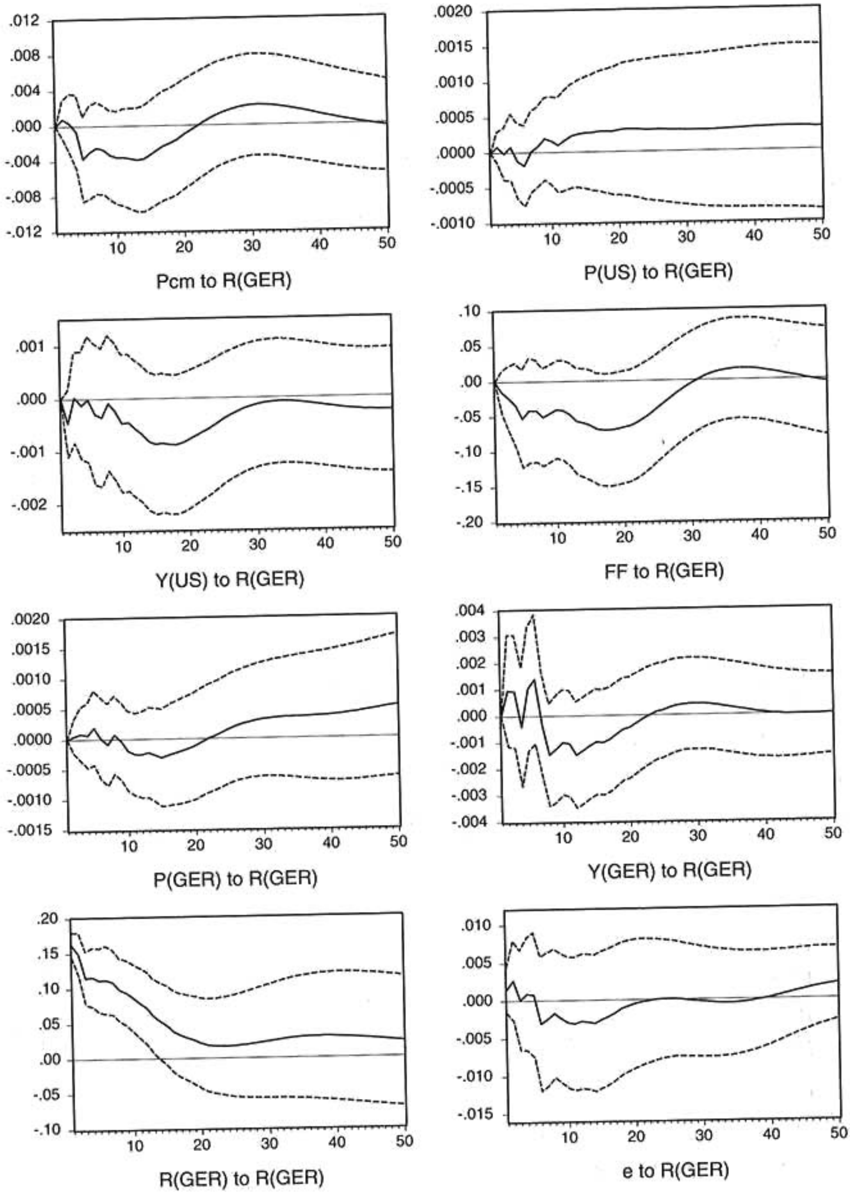

FIG. 6.10. Impulse responses to German monetary policy shock in the benchmark VAR with commodity price index (dashed lines: 2 standard error bands)

6.7 VAR and non-VAR measures of monetary policy

Econometric measurement of monetary policy has always been a debated issue. VAR models are linear, constant-parameter autoregressive distributed lag models, bound to include a very limited number of variables with a parsimonious lag parameterization. The crucial step to deriving evidence from the data using VARs, is the possibility of posing identifying restrictions independently from theoretical models. We have illustrated how a consensus has been reached in the case of closed economies and how the same result has not yet been achieved for open economies. We have provided an interpretation of this difference in the light of the difficulties in identifying monetary policy shocks in open economies. Recently, VAR based monetary policy shocks have been compared with monetary policy shocks measured by alternative approaches. We think that these developments can be useful not only to evaluate VAR methodologies, but also to help identification when, as in the case of the open economy models, the traditional VAR methods have difficulties in delivering the necessary number of identifying restrictions.

6.7.1 *Non-VAR measures of monetary policy shocks*

Historically, alternatives to econometric measurement of monetary policy have always been considered, think for example of qualitative indicators of monetary policy derived adopting the 'narrative approach' of Romer and Romer (1989; 1994). In a recent paper, Leeper (1996) shows that even the dummy variable generated by the 'narrative approach' (identifying episodes of deliberate monetary contractions) is predictable from past macroeconomic variables, thus reflecting the endogenous response of policy to the economy, and the estimated coefficients cannot provide an unbiased estimate for the response of the macroeconomic variables to a monetary impulse.

Recently the attention of monetary economists has turned to financial markets, which are a potential source of powerful information and measurement of monetary policy. We consider a variety of measures of monetary policy derived from financial markets and assess their role in the evaluation of VAR-based monetary policy shocks in open and closed economies.

A first alternative was proposed by Rudebusch (1998) and further analysed by Brunner (1996). Monetary policy shocks are derived from the thirty-day Federal funds future contracts, which have been quoted on the Chicago Board of Trade since October 1988, and are bets on the average overnight Federal funds rate for the delivery month, the variable included in benchmark VARs. Shocks are constructed as the difference between the Federal funds rate at month t and the thirty-day federal funds future at month $t-1$. Such a choice is based on the evidence that the regression of the Federal funds rate (FF) at t on the thirty-day Federal funds future (FFF) at $t-1$ produces an intercept not significantly different from zero, a slope coefficient not significantly different from one, and serially uncorrelated residuals:

$$\mathrm{FF}_t = -\underset{(0.0436)}{0.037} + \underset{(0.007)}{0.999}\mathrm{FFF}_{t-1} + \widehat{u}_t$$
$$R^2 = 0.99, \qquad \sigma = 0.145, \qquad \mathrm{DW} = 1.86.$$

Note that this procedure produces shocks, labelled FFF, comparable to the reduced form innovations from the VAR and not to the structural monetary policy shocks, because surprises relative to the information available at the end of month $t-1$ may reflect endogenous policy responses to news about the economy that become available in the course of month t. However, if an identification scheme is available, then innovations derived from the future contracts can be transformed in the relevant shocks by applying to them the standard VAR identification procedure. A non-trivial problem with this procedure comes from the fact that Federal funds future are available from 1988 onwards. Future contracts on the one-month Eurodollar are available on a more extended sample. Given that the properties of the series generated by the one-month Eurodollar are very close to the properties of Federal funds future, the direct measurement based on one-month Eurodollar could be used on an extended sample.

A second non-VAR measure of policy shocks is based on the work of Skinner and Zettelmeyer (1996). They derive a measure of unanticipated monetary policy shocks by following a two-step methodology: first, using information from central bank reports and newspapers they make a list of days on which monetary policy announcements occurred; second, monetary policy shocks are identified with the changes in the three-month interest rate on the days of policy announcements. The validity of such a procedure requires that:

1. short rates (e.g. the overnight rate) are affected by policy;
2. arbitrage is effective between the overnight and the three-month interest rates;
3. the impact of other news affecting the three-month rate on the day of the policy decision is negligible;
4. policy actions are not endogenous responses to information that becomes available on the day when the decision is made.

To ensure that conditions 3 and 4 are applicable, Skinner and Zettelmeyer go through reports of the policy actions and exclude from their sample those which do not conform to requirements 3 and 4. The main problem with the index obtained this way is that it can only pin down shocks associated to monetary policy decisions reflected in some action on controlled variables, whereas shocks associated with *no* action (while some action was expected by the markets) are neglected.

An alternative approach which might overcome this problem has been proposed by Bagliano and Favero (1998), applying the methodology set out in Svensson (1994) and Soderlind and Svensson (1997). The methodology is based on the use of instantaneous forward rates as monetary policy indicators. Forward rates are interest rates on investments made at a future date, the settlement date, and expiring at a date further into the future, the maturity date. Instantaneous forward interest rates are the limit as the maturity date and the settlement date approach one another.

To illustrate our derivation of the spot rate let us start by the consideration of a *zero-coupon* bond issued at time t with a face value of 1, maturity of m years and price P_{mt}^{ZC}. The simple yield Y_{mt} is related to the price as follows:

$$P_{mt}^{ZC} = \frac{1}{(1+Y_{mt})^{m}}. \tag{6.22}$$

Defining the spot rate r_{mt} as $\log(1+Y_{mt})$, which is the continuously compounded yield, and the discount function D_{mt} as the price at time t of a zero coupon that pays one unit at time $t+m$, we then have:

$$P_{mt}^{ZC} = \exp(-mr_{mt}) = D_{mt}. \tag{6.23}$$

Consider now a *coupon* bond that pays a coupon rate of c per cent annually and pays a face value of 1 at maturity. The price of the bond at trade date is given by the following formula:

$$P_{mt} = \sum_{k=1}^{m} cD_{kt} + D_{mt}. \tag{6.24}$$

Given the observation of prices of coupon bonds, spot rates on the zero coupon equivalent are derived by fitting a discount function based on the following specification for the spot rates:

$$\begin{aligned} r_{kt} = \beta_0 + \beta_1 \frac{1-\exp\left(-\frac{k}{\tau_1}\right)}{\frac{k}{\tau_1}} + \beta_2 \left(\frac{1-\exp\left(-\frac{k}{\tau_1}\right)}{\frac{k}{\tau_1}} - \exp\left(-\frac{k}{\tau_1}\right)\right) \\ +\beta_3 \left(\frac{1-\exp\left(-\frac{k}{\tau_2}\right)}{\frac{k}{\tau_2}} - \exp\left(-\frac{k}{\tau_2}\right)\right). \end{aligned} \tag{6.25}$$

Such a specification was originally introduced by Svensson (1994) and it is an extension of the parameterization proposed by Nelson and Siegel (1987). Implied forward rates can be calculated from spot rates. A forward rate at time t with trade date $t+t'$ and settlement date $t+T$ is calculated as the return on an investment strategy based on buying zero-coupon bonds at time t maturing at time $t+T$ and selling at time t, zero-coupon bonds maturing at time $t+t'$. The forward rate is related to the spot rate by the following formula:

$$f_{t+T,t+t',t} = \frac{Tr_{T,t} - t'r_{t',t}}{T-t'} \tag{6.26}$$

so the forward rate for a one-year investment with settlement in two years and maturity in three years is equal to three times the three-year spot rate minus twice the two-year spot rate. The instantaneous forward rate is the rate on a forward contract with an infinitesimal investment after the settlement date:

$$f_{mt} = \lim_{T \to m} f_{t+T,t+m,t}. \tag{6.27}$$

In practice, we identify the instantaneous forward rate with an overnight forward rate, a forward rate with maturity one day after the settlement. The relation between instantaneous forward rate and spot rate is then:

$$r_{mt} = \frac{\int_{\tau=t}^{t+m} f_{\tau t} d\tau}{m}$$

or, equivalently

$$f_{mt} = r_{mt} + m \frac{\partial r_{m,t}}{\partial m}. \tag{6.28}$$

Given specification (6.25) for the spot rate, the resulting forward function is:

$$f_{kt} = \beta_0 + \beta_1 \exp\left(-\frac{k}{\tau_1}\right) + \beta_2 \frac{k}{\tau_1} \exp\left(-\frac{k}{\tau_1}\right) + \beta_3 \frac{k}{\tau_2} \exp\left(-\frac{k}{\tau_2}\right). \tag{6.29}$$

Therefore, as k goes to zero, the spot and the forward rates coincide at $\beta_0 + \beta_1$ and as k goes to infinity, the spot and the forward rates coincide at β_0. The forward rate function features a constant, an exponential term decreasing when β_1 is positive, and two 'hump-shaped' terms. In principle $\beta_0+\beta_1$ can be restricted to match the observed overnight rate, but we do not follow this strategy. In fact, by definition a monetary policy shock implies a jump in, at least, the short end of the term structure. Forcing the smooth instantaneous forward rate curve to fit exactly the observed overnight rate would not allow us to seize an eventual expected monetary policy action. For this reason, we exclude the overnight rate from the information used for estimation. Then, exploiting the continuity of the functional form, we reconstruct the very short end of the term structure allowing for a gap between the estimated overnight and the observed overnight rates. Such a gap represents the jump in the very short end of the term structure associated with expectations of intervention by the central bank. An example can clarify matters. On the occasion of the meeting held on the 2nd of December 1993 the Bundesbank reduced the repurchase rate by 25 basis points. On the close of the markets before the meeting we observed the structure of spot rates relevant to the estimation of our yield curves reported in Table 6.1.

TABLE 6.1. German interest rates

Date	30/11/93	2/12/93
o/n	6.70	6.35
seven-days	6.44	6.31
one-month	6.44	6.31
three-month	6.19	6.06
six-month	5.81	5.75
one-year	5.37	5.25
two-year	5.08	5.03
three-year	5.05	5.02
four-year	5.16	5.15
five-year	5.3	5.29
seven-year	5.69	5.68
ten-year	6.16	6.17

In Figure 6.11 we report Nelson–Siegel interpolants. More precisely we report the two instantaneous forward curves associated, respectively, to the spot curve estimated excluding the overnight rate (IFW) and to the spot curve estimated including the overnight rate (IFOY).

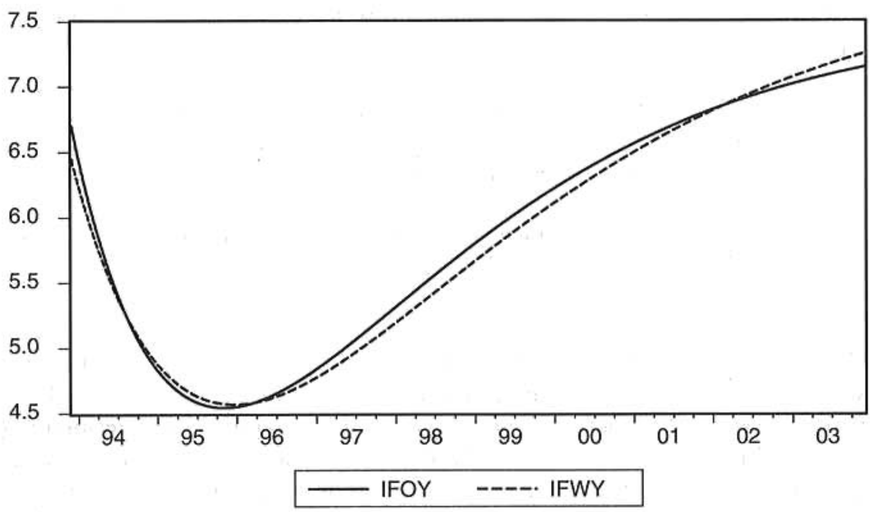

FIG. 6.11. Estimated forward-rate curves on the 30 Novermber 1993 with and without the overnight rate

Fitting the curve on the data including the overnight without allowing for a jump in the term structure from the date of the Bundesbank Council meeting afterwards would have spuriously generated an interest rate shock.

If the pure expectational model is valid and there is no term premium, then instantaneous forward rates at future dates can be interpreted as the expected spot interest rates for those future rates. The observable equivalent of the instantaneous forward rate is the overnight rate.

The following strategy for identification of monetary policy shocks directly exploits the relation between spot rates and instantaneous forward rates:

1. exploiting the fact that intervention on policy rates for Germany and the US takes place on the occasions of regular meetings of the Bundesbank Council and of the Federal Open Market Committee (FOMC) (since 1994), collect data on the term structure of interest rates the day before the monetary policy meetings. Observations on one-day, seven-day rate, one-month euro, three-month euro, six-month euro, twelve-month euro, three-, five-, seven-, and ten-year fixed interest rate swap are available on Datastream and other databases;

2. estimate a term structure for spot rates and the associated curve of instantaneous rates;

3. interpret the instantaneous rate as the overnight rate, and from the curve derive the expected implicit overnight rate for the day after the monetary policy meeting;

4. derive monetary policy shocks, subtracting from the observed overnight rate the day after the policy meeting the overnight rate implicit in the curve the day before the policy meeting;

5. aggregate the above daily measures (concentrated in a few special days) to construct monthly measures of shocks.

There are several difficulties that one should overcome in constructing this measure of monetary policy shocks. Following Bagliano and Favero (1999), we illustrate examples of monetary shocks generated by unanticipated action or by unanticipated inaction by the Bundesbank, and examples of markets' anticipation of Bundesbank behaviour when expectations on monetary policy turned out to be correct and no shocks were observed. We consider the sample 1984–1997.

Consider first July 1988. In this month the Bundesbank Council met twice, on the 14th and on the 28th. During the first Council the Bundesbank didn't take any action, during the second Council it was decided to raise the Lombard Rate by 50bp. In Figure 6.12 we report the weekly and the overnight rate, along with the monetary policy action (PMA).

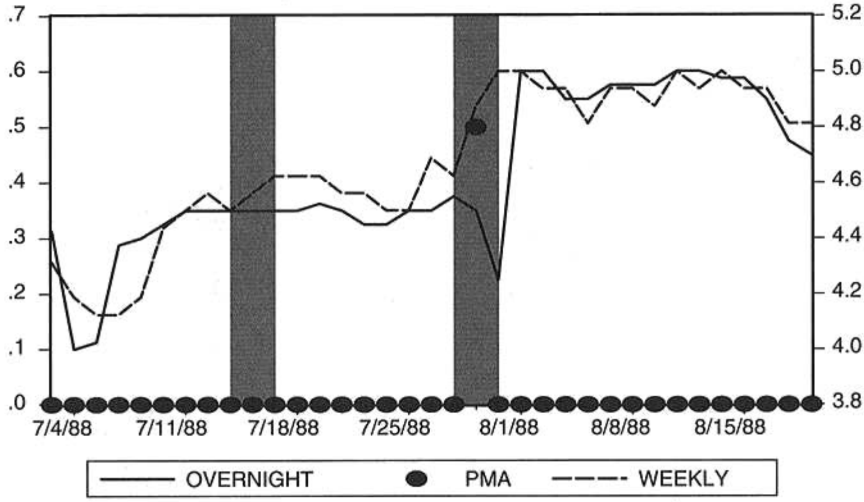

FIG. 6.12. Monetary policy interventions and short-term interest rates in Germany, July 1988

We shade areas of the three days centred around the meetings. Note that no monetary policy action was expected during the first meeting, while some action was expected before the second one. Six days before the meeting the weekly rate contains the first six days of maturity which doesn't include the action and the seventh one which, instead, does include it, so the weekly rate should start to 'reflect' the monetary policy action six days before the meeting. The weight of the seventh day is one-seventh, so the information doesn't appear clearly six days before, but as we approach the date of the council the weight of the action becomes greater and the expectation discloses itself. In fact, the weekly rate starts reacting three days before the meeting. It is also possible that the market realizes that the Bundesbank will act only a few days before the Council (say less than six days before), in this case the weekly rate starts reacting later than six days before the Council. The weekly rate should be the best observed interest rate to identify expectations on monetary policy actions. In fact, Council meetings take place fortnightly and the one-month rate immediately before any meeting reflects expectations on the outcome of the following two meetings.

The second episode we consider is the tightening of monetary policy that occurred after German reunification in January–February 1991. Two meetings were held in this period, on the 17th of January and the 2nd of February. As Figure 6.13 clearly shows, the weekly rate increased sharply just before the first Council revealing an expected increase in interest rates.

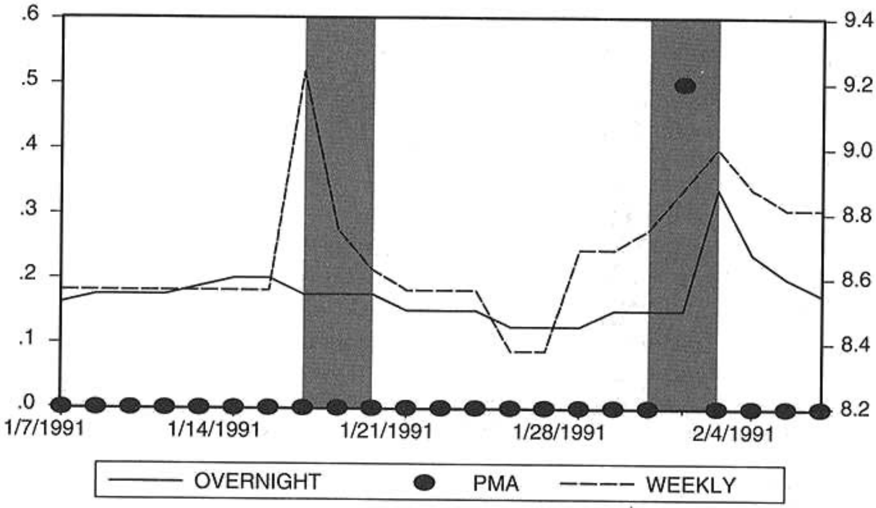

FIG. 6.13. Monetary policy interventions and short-term interest rates in Germany, January 1991

The Bundesbank didn't act on that meeting. We immediately observe that the expected tightening happened during the following Council meeting, when the Bundesbank raised the discount rate and the Lombard rate by 50 bp. To summarize, on the 14th of January we observed a monetary policy shock arising from an anticipated action that didn't occur, meanwhile on the 2nd of February there is no shock as the policy move has been correctly anticipated.

The third episode we single out occurred in December 1991 (see Figure 6.14) when the Bundesbank tightened the monetary policy, once again raising the discount rate and the Lombard rate by 50 bp. The dates of the Bundesbank Councils are the 5th and the 19th of December. During the latter meeting the Bundesbank surprised the market, and we observe a shock arising from an unexpected policy action.

The main strength of the methodology based on forward interest rate curves is its flexibility and capability to capture shocks independently from the specification and parameterization of a linear autoregressive model. The main limitation of this approach is caused by the volatility of very short-term rates not related to expectations on monetary policy. Figures 6.15 and 6.16 report daily observations on the overnight and the weekly rates for the estimation sample period used in the VAR.

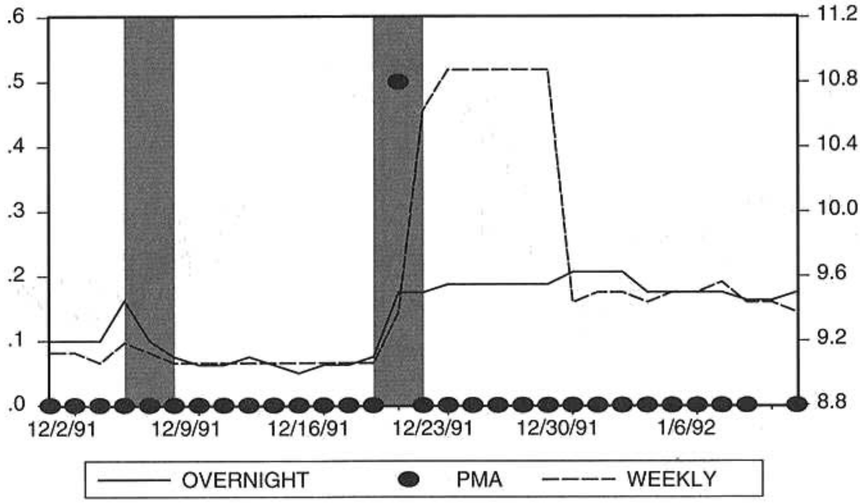

FIG. 6.14. Monetary policy interventions and short-term interest rates in Germany, December 1991

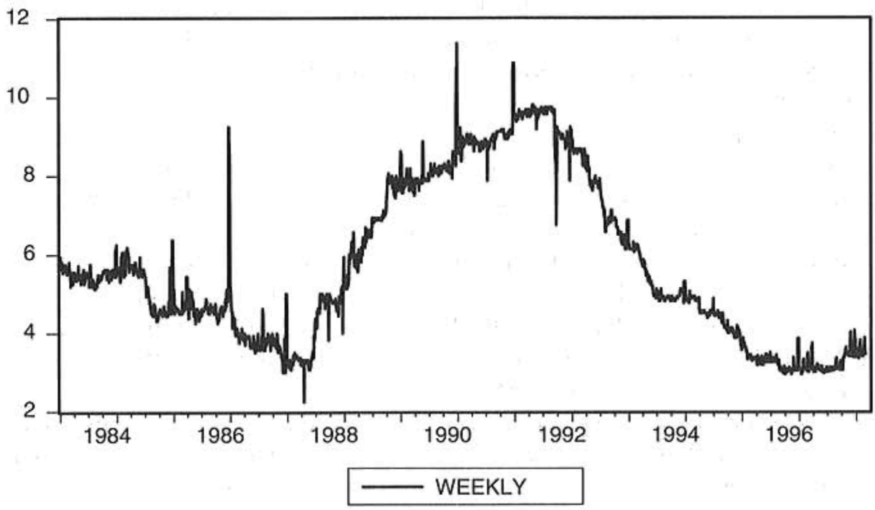

FIG. 6.15. The German seven-day rate

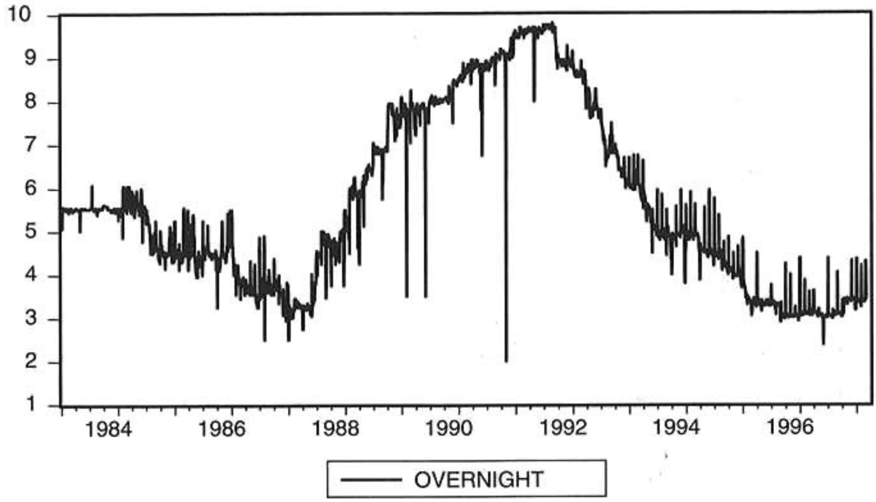

FIG. 6.16. The German overnight rate

We immediately note a number of blips in the series. They can be damaging to our methodology whenever they happen on occasion of a Bundesbank Council meeting.

Most of those blips are generated by banks' reserves' management which run into a non-perfectly liquid market, such as on the occasion of the last day of the average reserves maintenance period. We make an effort to render our inference robust to blips.

In fact, we have estimated our curves starting from the seven-day rather than the overnight rate, and our methodology of estimation considers the information contained in the whole term structure. However, we have run a further check and avoid labelling as policy shocks all unexpected movements in policy rates which have disappeared within a week after the Council meeting. Such correction led us to single out two outliers in 1988:9 and 1991:12. The 1988:9 outlier, whose determination is described in Figure 6.13, is the only one of a relevant magnitude.

In Figure 6.17 we report the behaviour of the seven-day and the one-month rate in the course of September 1988.

No policy intervention was decided in September 1988, however, just before the meeting of mid-September we observe a hike in the seven-day rate. Such a hike is not reflected in the term structure for longer maturities (we report one-month for reference). This hike would have been labelled as a shock by the methodology; however, as it is reversed within the week after the meeting this episode should be considered as a monetary policy shock.

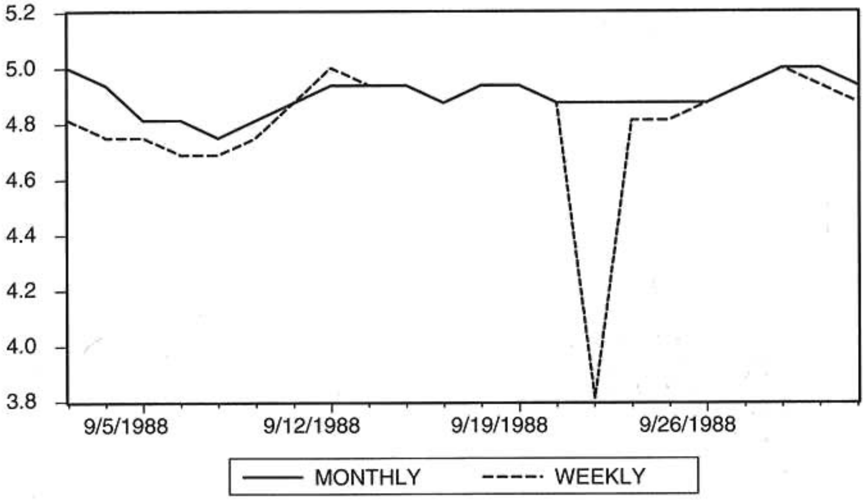

FIG. 6.17. The German seven-day and one-month rate in September 1988

6.8 Empirical results

Non-VAR measures of monetary policy can be directly compared with VAR measures; they can also be used to assess the robustness of the VAR-based descriptions of the monetary transmission mechanism, finally, they can be exploited, within a VAR, to aid identification of other structural shocks. To illustrate these possibilities we consider in turn the closed economy (US) case and the open economy (US–Germany) case.

6.8.1 *Closed economy (US)*

To evaluate the role of non-VAR based measures of monetary policy shock, we first estimate the closed-economy four-variable version of the VAR model for the US and compute impulse response functions of all variables to a shock in the Federal funds rate. Our model is specified as follows:

$$\mathbf{A}\begin{pmatrix} Y_t^{US} \\ Pcm_t \\ P_t^{US} \\ FF_t \end{pmatrix} = \mathbf{C}(L)\begin{pmatrix} Y_{t-1}^{US} \\ Pcm_{t-1} \\ P_{t-1}^{US} \\ FF_{t-1} \end{pmatrix} + \mathbf{B}\begin{pmatrix} \nu_t^1 \\ \nu_t^2 \\ \nu_t^3 \\ \nu_t^{FF} \end{pmatrix} \tag{6.30}$$

where $\mathbf{A}$ is lower-triangular and $\mathbf{B}$ is diagonal.

The ordering chosen allows for a contemporaneous response of the policy rate to innovations in output, consumer prices and the commodity price level. The orthogonalized residual of the Federal funds rate equation, ν^{FF}, is identified as a monetary policy shock. No structural interpretation is given to the (orthogonalized) residuals from the other equations in the system. We consider two non-VAR measures of monetary policy, derived respectively from one-month

Eurodollar forward rate (EUR\$) and from the estimation of the instantaneous forward rate curve on the occasions of FOMC meetings (IFS^{US}). These alternative shocks are plotted with the VAR-based shocks in Figures 6.18 and 6.19.

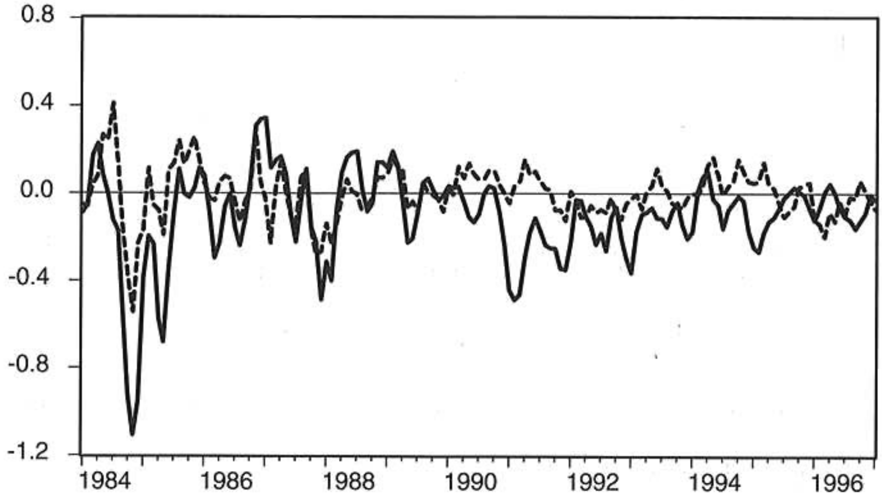

FIG. 6.18. Three-month centred moving averages of EUR\$ shocks (solid line) and closed-economy VAR monetary policy shocks (dotted line)

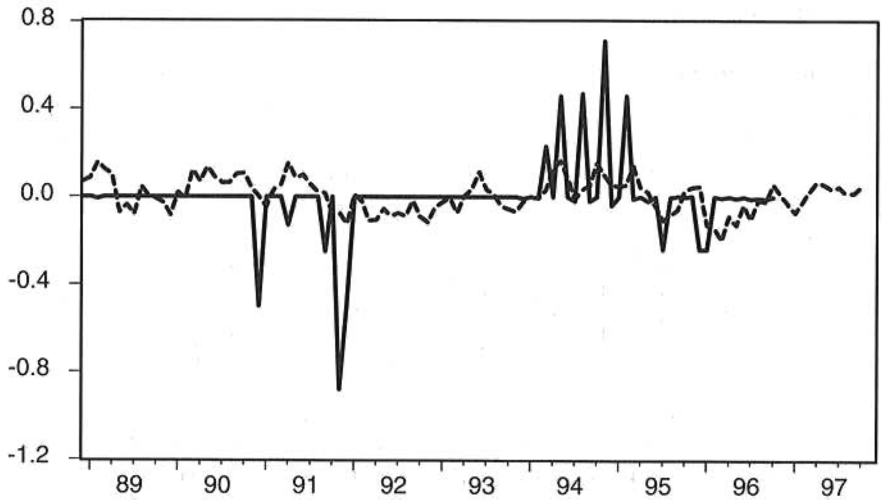

FIG. 6.19. IFS^{US} shocks (solid line) and closed-economy VAR monetary policy shocks (dotted line)

Note that the EUR\$ measure is available on a larger sample than the IFS^{US} measure as the practice of modifying monetary policy rates on the occasions of given and announced dates started only in the 1990s. We report in Table 6.2 the correlations of VAR and non-VAR measures of monetary policy.

TABLE 6.2. VAR and non-VAR monetary policy shocks

Sample: 1988:11–1996:10			
Correlation coefficients (standard errors on the diagonal)			
	$EUR\$$	IFS^{US}	ν^{FF}
$EUR\$$	0.185	-	-
IFS^{US}	0.203	0.169	-
ν^{FF}	0.352	0.319	0.123
Sample: 1988:11–1996:10			
	$EUR\$$	ν^{FF}	
$EUR\$$	0.277	-	
ν^{FF}	0.500	0.211	

Rudebusch (1996) , using the Federal funds future contracts, obtains similar results to those obtained in our shorter sample and concludes that VAR-based measures of monetary policy do not make sense. We note that much better results are obtained in the enlarged sample. To provide further evidence we specify a VAR augmented by the non-VAR measures of monetary policy shocks, considered as an exogenous variable. Following Amisano and Giannini (1996), we represent the estimated system as follows:

$$\mathbf{A}\begin{pmatrix} \mathrm{Y}_t^{US} \\ \mathrm{Pcm}_t \\ \mathrm{P}_t^{US} \\ \mathrm{FF}_t \end{pmatrix} = \mathbf{C}(L)\begin{pmatrix} \mathrm{Y}_{t-1}^{US} \\ \mathrm{Pcm}_{t-1} \\ \mathrm{P}_{t-1}^{US} \\ \mathrm{FF}_{t-1} \end{pmatrix} + \begin{pmatrix} g_1 \\ g_2 \\ g_3 \\ g_4 \end{pmatrix} x_t + \mathbf{B}\begin{pmatrix} \nu_t^1 \\ \nu_t^2 \\ \nu_t^3 \\ \nu_t^{FF} \end{pmatrix} \tag{6.31}$$
$$x = \mathrm{EUR\$}, \mathrm{IFS}^{US}$$

where $\mathbf{A}$ is lower-triangular and $\mathbf{B}$ is diagonal. The estimated values of the coefficients g_i are reported in Table 6.3.

TABLE 6.3. Coefficients on $EUR\$$ and IFS^{US} in the benchmark VAR

Sample 1988:11–1997:11				
	Y^{US}	Pcm	P^{US}	FF
$EUR\$$	0.0061 (0.0032)	0.0055 (0.0121)	0.0013 (1.0633)	0.468 (0.097)
IFS^{US}	0.0025 (0.0031)	0.0082 (0.0116)	0.0009 (0.0013)	0.356 (0.099)
Sample: 1984:1–1997:11				
	Y^{US}	Pcm	P^{US}	FF
$EUR\$$	0.0026 (0.0016)	0.0007 (0.0006)	0.0058 (0.0063)	0.552 (0.062)

We note that none of the macroeconomic variables responds to the non-VAR monetary policy shocks, while the Federal fund rate does. As suggested by the correlations between shocks, results are stronger on the larger sample. We then concentrate on this sample and compare impulse responses to monetary policy shocks in the traditional benchmark VAR specification with impulse responses to non-VAR monetary policy shocks in our augmented specification.

The results, shown in Figure 6.20, illustrate that a contractionary monetary policy shock produces the expected negative effect on output and a persistent effect on the Federal funds rate.

The inclusion of the commodity price index is successful in solving the price puzzle: the consumer price level does not show a 'perverse' response to restrictive policy.

The pairs of impulse response functions, based on the VAR and the non-VAR shocks, describe a very similar transmission mechanism, supporting the evidence already provided by Brunner (1996) and Bagliano and Favero (1998) with different exogenous measures.

Despite a correlation of 0.5 between EUR\$ and the measure of policy shock obtained from the benchmark VAR, the dynamic effects of monetary policy show very close features.

The different measures capture unexpected variations in the policy rate related to monetary policy and the existence of other non-policy disturbances does not change the basic features of the response to a policy shock.

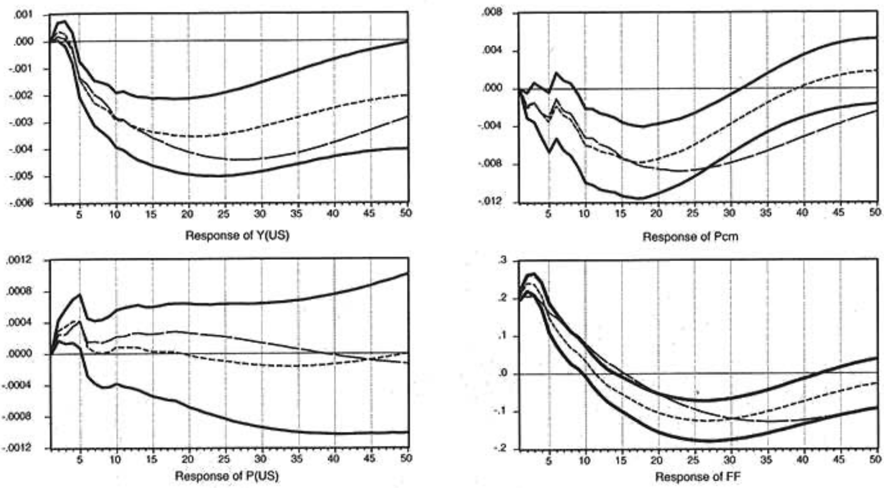

FIG. 6.20. Responses to EUR\$ shocks (solid line) and to VAR-based structural shocks ν^{FF} (dotted line) with one standard deviation confidence intervals from the benchmark VAR

6.8.2 *Open economy (US–Germany)*

Let us now consider the open-economy version of the VAR system. We begin by a baseline specification which includes the commodity price index:

$$\mathbf{A}\begin{pmatrix} \mathrm{Y}_t^{US} \\ \mathrm{Pcm}_t \\ \mathrm{P}_t^{US} \\ \mathrm{FF}_t \\ \mathrm{Y}_t^{GER} \\ \mathrm{P}_t^{GER} \\ e_t \\ \mathrm{R}_t^{GER} \end{pmatrix} = \mathbf{C}(L)\begin{pmatrix} \mathrm{Y}_{t-1}^{US} \\ \mathrm{Pcm}_{t-1} \\ \mathrm{P}_{t-1}^{US} \\ \mathrm{FF}_{t-1} \\ \mathrm{Y}_{t-1}^{GER} \\ \mathrm{P}_{t-1}^{GER} \\ e_{t-1} \\ \mathrm{R}_{t-1}^{GER} \end{pmatrix} + \mathbf{B}\begin{pmatrix} \nu_t^1 \\ \nu_t^2 \\ \nu_t^3 \\ \nu_t^{FF} \\ \nu_t^5 \\ \nu_t^6 \\ \nu_t^e \\ \nu_t^{RGER} \end{pmatrix} \tag{6.32}$$

where **A** is lower-triangular and **B** is diagonal. As we have done for the closed economy case, we compare the orthogonalized residual of the German call money rate equation (ν^{RGER}) with the non-VAR measure of German monetary policy shocks IFS^{GER}, derived from instantaneous forward rates. Figure 6.21 and Table 6.4 confirm the results for correlations obtained in the closed-economy case.

TABLE 6.4. VAR and non-VAR monetary policy shocks

Sample: 1984:1–1997:11

	IFS^{GER}	ν^{RGER}
IFS^{GER}	0.194	-
ν^{RGER}	0.163	0.169

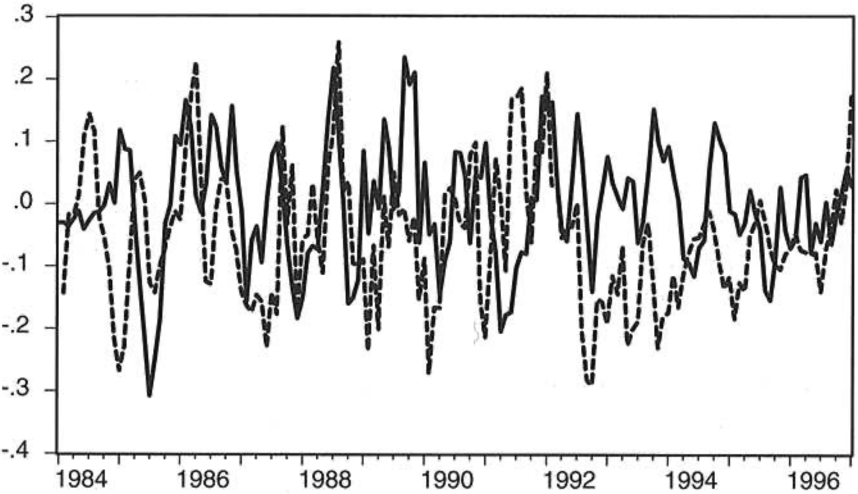

FIG. 6.21. Three-month centred moving averages of IFS^{GER} shocks (solid line) and open-economy VAR German monetary policy shocks (dotted line)

As in the closed-economy case, we augment the previously estimated system by including the exogenous measure of German monetary policy shocks IFS^{GER} described in the previous section.

The open-economy VAR is now the following:

$$\mathbf{A}\begin{pmatrix} \mathrm{Y}_t^{US} \\ \mathrm{Pcm}_t \\ \mathrm{P}_t^{US} \\ \mathrm{FF}_t \\ \mathrm{Y}_t^{GER} \\ \mathrm{P}_t^{GER} \\ e_t \\ \mathrm{R}_t^{GER} \end{pmatrix} = \mathbf{C}(L)\begin{pmatrix} \mathrm{Y}_{t-1}^{US} \\ \mathrm{Pcm}_{t-1} \\ \mathrm{P}_{t-1}^{US} \\ \mathrm{FF}_{t-1} \\ \mathrm{Y}_{t-1}^{GER} \\ \mathrm{P}_{t-1}^{GER} \\ e_{t-1} \\ \mathrm{R}_{t-1}^{GER} \end{pmatrix} + \begin{pmatrix} g_1 \\ g_2 \\ g_3 \\ g_4 \\ g_5 \\ g_6 \\ g_7 \\ g_8 \end{pmatrix} IFS_t^{GER} + \mathbf{B}\begin{pmatrix} \nu_t^1 \\ \nu_t^2 \\ \nu_t^3 \\ \nu_t^{FF} \\ \nu_t^5 \\ \nu_t^6 \\ \nu_t^e \\ \nu_t^{RGER} \end{pmatrix}. \quad (6.33)$$

Using our exogenous measure of monetary policy shocks in combination with a Choleski ordering with the German policy rate coming last, we are able to directly address the issue of simultaneity between German monetary policy and the exchange rate. The contemporaneous effect of a monetary policy shock on the exchange rate is given by the coefficient on IFS^{GER} in the exchange rate equation (g_7), while the response of the German interest rate to innovations in the exchange rate is endogenized by the ordering chosen. As shown in Table 6.5, we do not observe a significant contemporaneous feedback between the German interest rate and the exchange rate in any direction. In our framework, this is a testable proposition rather than an assumed identified restriction. We note that our measure of monetary policy shocks significantly enters in the German policy rate equation and that the contemporaneous response of the US output to German monetary policy shocks is small but marginally significant.[3]

TABLE 6.5. Coefficients on IFS^{GER} and simultaneous responses of e

	Y^{US}	Pcm	P^{US}	FF	Y^{GER}	P^{GER}	e	R^{GER}
IFS^{GER}	−0.007 (0.002)	−0.01 (0.008)	−0.0013 (0.0008)	−0.0892 (0.1146)	0.00002 (0.0011)	0.0029 (0.007)	0.0084 (0.0127)	0.230 (0.097)
e	1.36 (0.037)	−0.15 (0.11)	−0.15 (1.09)	0.022 (0.0083)	0.045 (0.126)	2.44 (0.79)	−0.007 (0.002)	−0.008 (0.01)

The pair of impulse response functions shown in Figure 6.22, along with one standard deviation bands, confirm qualitatively the results obtained for the closed US economy: measuring monetary policy shocks using financial market data does not alter the main features of the monetary transmission mechanism for Germany.

6.9 Conclusions

The VAR approach to the monetary transmission mechanism aims at the derivation of stylized facts to help the selection of the theoretical model to use for simulating the effects of monetary policy. The identification of parameters in this type of model does not allow us to separate deep parameters describing taste and technology from expectational parameters dependent on policy regimes. However, by estimating this model on a single policy regime and by describing the responses of variables to structural shocks of interest, it is hoped to derive some stylized facts on the monetary transmission mechanism.

Unfortunately, structural shocks are not directly observable and the imposition of a set of identifying restrictions is a necessary prerequisite for the analysis. Given the aim of the analysis, it is essential that identifying restrictions are posed

[3] We report impulse responses based on restricting such a coefficient to zero; relaxing this restriction does not affect the shape and magnitude of the impulse responses.

independently from specific theories. All the developments of the Choleski ordering that we have discussed in the chapter provide the researcher with tools for achieving this aim. In particular, we have shown how information from financial markets can be used both to assess robustness of results derived within traditional VAR models of the monetary transmission mechanism and to aid identification in cases when traditional analysis does not deliver a sufficient number of restrictions. Within this framework it is also natural that the number of identifying restrictions is kept at a minimum.

VAR models of the monetary transmission mechanism are very rarely cointegrated VARs. We have seen that multivariate cointegration analysis requires the solution of a long-run identification problem, and that imposing cointegrating restrictions on a VAR in levels increases efficiency in the estimator at the cost of potential inconsistency, when the incorrect identifying restrictions are imposed. The monetary transmission mechanism is a short-run phenomenon and this explains why researchers prefer to employ unrestricted VARs to evaluate impulse response analysis over a short to medium horizon. Cointegrated VARs are however an almost inevitable choice when the relevant, theory-neutral restrictions, are long-run ones.

As VAR models are the natural empirical counterparts of dynamic general equilibrium models, their statistical adequacy is not as closely scrutinized as the adequacy of reduced-form specifications within the LSE approach. In particular, in numerous applications outliers are not removed and non-normality of residuals is a common feature. Parameter stability is also an issue in the debate.

Our analysis of VAR-based empirical evidence of the monetary transmission mechanism has shown that the main problem with the empirical evidence provided by these models is uncertainty. Large standard errors are associated with the point estimates of impulse response functions. The more so in the case of VAR in open economies, where practitioners have developed the habit of reporting one standard deviation bands rather than two standard deviation bands. The main consequence of such uncertainty is that the aim of the exercise, once again model selection, is difficult to achieve in practice.

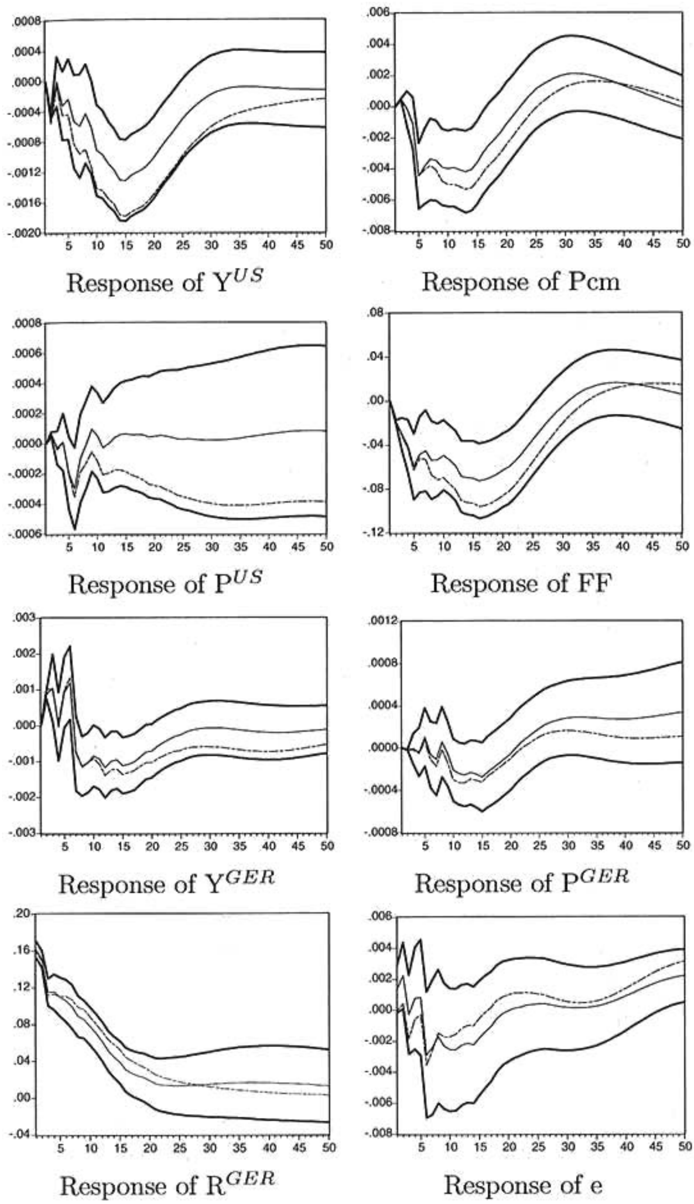

FIG. 6.22. Responses to IFS^{GER} shocks (solid line) and to VAR-based structural shocks ν^{RGER} (dotted line) with one standard deviation confidence intervals from the benchmark VAR

6.10 References

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