A Bayesian analysis of simultaneous equation models by combining recursive analytical and numerical approaches
Steel, M.F.J.

Publication date:
1991

Link to publication

Citation for published version (APA):
A Bayesian Analysis of Simultaneous Equation Models by Combining Recursive Analytical and Numerical Approaches

by Mark F.J. Steel


Reprint Series no. 48
CENTER FOR ECONOMIC RESEARCH

Research Staff

Helmut Bester
Eric van Damme
Frederick van der Ploeg

Board

Helmut Bester
Eric van Damme, director
Arie Kapteyn
Frederick van der Ploeg

Scientific Council

Eduard Bomhoff
Willem Buiter
Jacques Drèze
Theo van de Klundert
Simon Kuipers
Jean-Jacques Laffont
Merton Miller
Stephen Nickell
Pieter Ruys
Jacques Sijben
Erasmus University Rotterdam
Yale University
Université Catholique de Louvain
Tilburg University
Groningen University
Université des Sciences Sociales de Toulouse
University of Chicago
University of Oxford
Tilburg University
Tilburg University

Residential Fellows

Joseph Greenberg
Jan Magnus
Emmanuel Petrakis
Larry Samuelson
Jonathan Thomas
McGill University
Tilburg University
University of California at Los Angeles
University of Wisconsin
University of Warwick

Doctoral Students

Roel Beetsma
Hans Bloemen
Chuangyin Dang
Frank de Jong
Pieter Kop Jansen

Address: Hogeschoollaan 225, P.O. Box 90153, 5000 LE Tilburg, The Netherlands
Phone: +31 13 663050
Telex: 52426 kub nl
Telefax: +31 13 663066
E-mail: "center@htikub5.bitnet"

ISSN 0924-7874

1991
A Bayesian Analysis of Simultaneous Equation Models by Combining Recursive Analytical and Numerical Approaches

by

Mark F.J. Steel


Reprint Series no. 48
A Bayesian analysis of simultaneous equation models by combining recursive analytical and numerical approaches*

Mark F.J. Steel

Tilburg University, 5000 LE Tilburg, The Netherlands

Received February 1989, final version received January 1990

Inference on simultaneous equation models using a Bayesian approach is still far from a routine procedure: serious numerical problems may occur in most practical situations. In this paper a unifying framework is proposed using recent analytical [see Steel (1988)] and well-documented poly-t based methods in order to reduce the dimension of the final Monte Carlo integration. Different integration sequences are examined and the implications for various model and prior assumptions are tabulated. An application to Klein's model I clearly illustrates the computational feasibility and numerical stability of the hybrid method proposed.

1. Introduction

In spite of the numerous theoretical attractions of Bayesian methods, a stimulating discussion of which can be found in Poirier (1988), their implementation and wide-spread use in applied econometrics has developed at a much slower rate than could initially be expected from the pioneering work of Drèze (1962) and Zellner (1965).

One of the reasons for this lack of empirical applications undoubtedly lies in the rather heavy computational requirements of Bayesian methods in most 'serious' applied work. This could be interpreted as the price one has to pay

*This work has benefited from discussions with Luc Bauwens, who also kindly provided me with invaluable help in programming the application. Useful comments from Arnold Zellner, an associate editor, and a referee have improved the presentation of the paper. The usual disclaimer applies. The author gratefully acknowledges the support of the Royal Netherlands Academy of Arts and Sciences and a grant from the Netherlands Organization for Scientific Research (NWO/NFS-14).
for conducting exact marginal inferences, requiring integration rather than optimization as in the sampling-theoretical approach.

Straightforward analytical results have only been applicable to simple seemingly unrelated regression equation (SURE) models compatible with the extremely restrictive (matrix-form) structure induced by natural conjugate (NC) analyses. Recently, somewhat less straightforward, but, nevertheless, analytical results were obtained for SURE models under a more general type of restrictions, called the recursive limited information (RLI) structure, in Steel (1988). This RLI assumption essentially imposes a block-recursive structure on the model, in the sense that, broadly speaking, all variables appearing in previous equations should also enter in the present equation. Analytical results can then be obtained, under suitable stochastic prior assumptions, by using recursive techniques. Heuristically, this may be understood if we consider only the first $i$ equations and realize that equation $i$ will contain all regressors appearing in that block of equations, and can thus be treated as unrestricted with respect to this part of the model.

Although the RLI approach extends the realm of analytical Bayesian inference considerably, it is still not sufficient to cover the full range of so-called simultaneous equation models, which are used frequently in time-series econometrics and which are described in subsection 2.1.

In order to tackle such models two main lines of research have been developed in the literature:

(i) Numerical integration by Monte Carlo importance sampling methods was introduced in econometrics by Kloek and van Dijk (1978) and further developed in, e.g., van Dijk (1984), Bauwens (1984), Geweke (1989), and Zellner et al. (1988). The dimensionality of the integrals is not as crucial a problem as in the more common methods like Gaussian or Simpson's product rules, but it is still a factor to be reckoned with, certainly if we wish to use a personal computer for the calculations.

(ii) Using the specific properties of so-called poly-t densities as described in Drèze (1977), we can combine analytical and relatively simple numerical techniques in order to obtain posterior results for one equation, which are, however, conditional upon all remaining equations.

It seems only natural to combine different methods, as each method has its own specific merits (and difficulties). Van Dijk (1984) and Naylor and Smith (1988) explore the mixing of product rules and importance sampling, whereas Geweke (1988, 1989) focuses on antithetic acceleration of Monte Carlo procedures. Kloek and van Dijk (1978) mention the combination of analytical and Monte Carlo methods, following the precept found in Hammersley and Handscomb (1964, p. 74) that one should replace numerical techniques by analytical methods wherever possible.
Much in the same vein, it will be attempted here to find a hybrid strategy that combines:

(i) exact analytical methods based on the RLI approach and exploiting the recursive properties of a subset of the parameters,
(ii) existing techniques for analyzing [see Richard and Tompa (1980)] and drawing [see Bauwens and Richard (1985)] poly-t distributed random variables,
(iii) Monte Carlo integration by importance sampling of either regression coefficients or the covariance matrix, the relative merits of which have been discussed in Richard and Steel (1988) for models of the SURE type.

Moreover, these methods will be applied in the above order of preference.

Section 2 of the paper sets the stage by introducing the model, a recursive transformation of the latter as this is felt to clarify the analysis, and the prior structure, consisting of both exact prior restrictions and stochastic assumptions on the unrestricted parameters. The ensuing posterior analysis is considered in section 3, where the first three subsections assume a Monte Carlo analysis on a subset of the regression coefficients, and subsection 3.4 briefly sketches Monte Carlo integration of the covariance matrix. Finally, subsection 3.5 groups some results on computational complexity and existence of posterior moments. A fourth section applies the hybrid methodology to a small macroeconomic model (Klein’s model 1) with nine unknown coefficients. The pure Monte Carlo integration in (iii) is only required in four dimensions, however, which is shown to result in both impressive numerical stability and computational efficiency. The final section ventures some concluding remarks.

2. The statistical framework

2.1. The simultaneous equation model

In this paper we shall concentrate upon a widely used class of models in (time-series) econometrics, namely the simultaneous equation model (SEM), as introduced in Haavelmo (1943) and since then treated in a classical fashion in most textbooks, whereas a Bayesian analysis is provided in the seminal papers by Drèze (1962), Rothenberg (1963), and Zellner (1965), and later in, e.g., Drèze and Richard (1983).

A sequential model is assumed where the observations at time $t$ on $n$ endogenous variables in $y(t)$ will be described by a set of $n$ ‘structural’ equations. These equations serve to restrict the conditional expectation of
\( y_{(t)} \), given the information set at time \( t \), to some linear manifold:

\[
B' E(y_{(t)} | x_{(t)}) + \Gamma' x_{(t)} = 0,
\]

(2.1)

where it is assumed that the information set at time \( t \), containing weakly exogenous variables and lagged values of both endogenous and exogenous variables, counts a finite number \( m \) of variables that are relevant for the description of \( y_{(t)} \) and are grouped in \( x_{(t)} \).

Remark that the parameter matrices \( B \) and \( \Gamma \) in (2.1) are, respectively, of dimensions \( n \times n \) and \( m \times n \), and will typically be subject to exact prior restrictions. This issue will be dealt with later in this subsection. The fact that \( B \) is square means that (2.1) has a unique solution provided \( |B| \neq 0 \). We shall make this full rank assumption, implying that we are in the context of a 'complete' SEM, whereas Richard (1984) rightly argues that, in general, we shall have fewer structural equations than endogenous variables and shall, therefore, usually be in the realm of 'incomplete' SEM. We fully agree with this point of view and reinterpret the \( n \) equations in (2.1) as the 'completed' set of equations, containing both structural and 'reduced' (e.g., instrumental variables) types of equations. However, we shall still allow for general linear restrictions in these completing equations, and we shall, therefore, avoid distinguishing them from the truly structural equations in our notation.

Stochastics are introduced by putting a Gaussian structure on \( y_{(t)} | x_{(t)} \) with a constant covariance matrix \( V \), implying that

\[
YB + X\Gamma = U,
\]

(2.2)

with

\[
D(U) = f^{T \times n}_{MN}(U|0, \Sigma \otimes I_T),
\]

(2.3)

where

\[
\Sigma = B'VB,
\]

\[
Y = (y_{(t)} \cdots y_{(T)})',
\]

\[
X = (x_{(t)} \cdots x_{(T)})',
\]

and \( U \) is a \( T \times n \) matrix of innovations with respect to the information set at each time period \( t: 1 \rightarrow T \).

For the notation and definitions of density functions used here we refer to appendix A of Drèze and Richard (1983). A superscript plus sign will denote a Moore–Penrose inverse.

We wish to treat the exact prior restrictions that pertain to \( B \) and \( \Gamma \) in (2.2) explicitly. We assume, in line with most applications of SEM, that the
covariance matrix \( \Sigma \) is not subject to restrictions, other than being in the class of positive definite symmetric (PDS) matrices. Let us, for notational convenience, rewrite (2.2) as
\[
WA = U, \tag{2.4}
\]
where we have implicitly defined
\[
W = (Y: X), \tag{2.5}
\]
and where
\[
A = \begin{pmatrix} B \\ F \end{pmatrix} \tag{2.6}
\]
groups all coefficients in the model. Let us now restrict the \((m + n) \times n\) elements of \(A\) in an affine linear way by considering its column expansion
\[
vec A = \alpha = R\zeta + r, \tag{2.7}
\]
where \(R\) is a \((m + n)n \times l\) selection matrix of known constants and of full column rank, \(r\) is a vector of constants, and the \(l\)-dimensional vector \(\zeta\) groups all unrestricted parameters.

This characterization of the restrictions allows for linear restrictions both within and across equations and seems to imply enough flexibility to be empirically useful. Its general nature and, in particular, its advantages over restrictions expressed in matrix form (i.e., by directly restricting \(A\) instead of its column expansion) are discussed in more detail in Richard and Steel (1988) and Steel (1988). This matrix form can, of course, be written as a special case of (2.7), where \(R\) takes a Kronecker product structure, say,
\[
R = I_n \otimes F, \tag{2.8}
\]
which, indeed, leads to
\[
A = FZ + C, \tag{2.9}
\]
provided \(\zeta = \text{vec } Z\) and \(r = \text{vec } C\).

As already explained in the introduction, the recursive analytical method of Steel (1988) will be used to deal with part of the parameters in the model. It was found that inference can be conducted analytically for the coefficients of a SURE model, provided it has a recursive limited information (RLI) structure and we possess diffuse prior information on the free parameters. Such an RLI assumption, in fact, says that the space spanned by the variables
corresponding to the free parameters in equation \( i \) should include all variables used in previous equations.

In practice, this will boil down to an assumption of block recursivity, where all variables in the previous \((i-1)\) equations also appear (with unrestricted coefficients) in the \( i \)th equation.

Heuristically, it can be understood that if this is the case, there are no extra restrictions implied by equation \( i \) when just focusing on the first block of \( i \) equations. Therefore, we can treat equation \( i \) as if it was unrestricted in this first block. Going through all the equations in a recursive manner, we can thus use analytical techniques for all coefficients in the SURE model that satisfy the RLI assumption.

In order to use these analytical techniques to reduce the dimension of the parameter space left for numerical treatment in a SEM, we single out those parameters for which these assumptions hold, i.e., they should not intervene in the Jacobian of the system (as in the SURE case) and they should, after possibly rearranging the equations, possess an RLI structure. In addition, their prior density has to belong to a special class as will be discussed later in subsection 2.3.

This means we shall rewrite the restrictions in (2.7) as

\[
\alpha = \alpha_1 + \alpha_2 = R_1^1 \xi^1 + R_2^2 \xi^2 + r^1 + r^2, \tag{2.10}
\]

where \( \xi^2 \) should be amenable to analytical treatment and both \( R_1^1 \) and \( R_2^2 \) are of full column rank, respectively \( l_1 \) and \( l_2 \), the sum of which will equal \( l \). The full column rank of the selection matrices implies, through basic theorems on triangularization [see, e.g., Strang (1980, p. 97)], that they can, without loss of generality, be rewritten in a block lower-triangular form, which induces the following recursive definition:

\[
R_{(i)}^j = \begin{pmatrix} R_{(i-1)}^j & 0 \\ R_{i(i-1)}^j & R_{ii}^j \end{pmatrix}, \tag{2.11}
\]

with \( R_{ii}^j \) of full column rank \( l_i^j \), and its analogue on the vector \( \xi \),

\[
\xi_{(i)}^j = \begin{pmatrix} \xi_{(i-1)}^j \\ \xi_i^j \end{pmatrix}, \tag{2.12}
\]

where \( j = 1, 2 \) and \( i: 1 \to n \) is the index\(^1\) referring to the equation in question; all this means that subscript \( i - 1 \) in parentheses refers to the

\(^1\) It proves notationally convenient to define these recursions for \( i: 1 \to n \) and thus allow \( R_{(0)}^j, R_{(0)}^{(0)}, \xi_{(0)}^j, \) and \( A_{(0)}^{(0)} \) in (2.24) \((j = 1, 2)\) to be of zero dimension.
previous \( i - 1 \) equations and subscript \( i \) pertains to equation \( i \) alone. Of course, we now have \( R^j = R^j_{(n)} \) and \( \zeta^j = \zeta^j_{(n)} \).

If we define
\[
W^j_i = WR^j_{ii},
\]
and denote by \( W^2_{(i)} \) the matrix of full column rank containing all the variables that have coefficients appearing in \( \zeta^2_{(i)} \), we can formally write down the requirements for \( \zeta^2 \) in order to ensure its analytical tractability; in particular, a sufficient condition is that
\[
P_{w_i}W^2_{(i)} = W^2_{(i)},
\]
i.e., the RLI assumption, where
\[
P_{w_i}^2 = W^2_i W^{2+} = W^2_i (W^2_i W^2_i)^{-1} W^2_i,
\]
provided \( W^2_i \) is of full column rank, which is assumed throughout the paper, and we require
\[
\|B\| = f(\zeta^1),
\]
i.e., that \( \|B\| \) is a function of \( \zeta^1 \) only and does not depend on \( \zeta^2 \). In subsection 2.3 a third defining characteristic of \( \zeta^2 \) is seen to be its prior dependence on \( \Sigma \) within a particular class of prior structures. For details and a proof of the sufficiency of (2.14), we refer to Steel (1988).

2.2. A recursive transformation

As argued in Steel (1988) a recursive transformation of our original model (2.2), (2.3), and (2.10) is not necessary for the argument, as all derivations can be done in both the original and the recursive parameterizations. It is, however, felt to somewhat clarify the analysis, and will, therefore, be introduced here.

Let \( \Lambda \) be the \( n \times n \) lower triangular matrix that diagonalizes \( \Sigma \), which was assumed to be PDS, and let us partition \( \Lambda \) as follows:
\[
\Lambda = \begin{pmatrix}
1 & & & 0 \\
-\lambda_{21} & 1 & & \\
-\lambda_3 & 1 & & \\
& & & \\
-\lambda_n & & 1
\end{pmatrix},
\]
where also \( \lambda_i = (\lambda_{i1}, \ldots, \lambda_{i,n}) \); now define

\[
\Lambda \Sigma \Lambda' = D_\omega = \begin{pmatrix}
\omega_1^2 & 0 \\
\vdots & \ddots \\
0 & \omega_n^2
\end{pmatrix}.
\]  

(2.18)

The transformation will, of course, also affect the coefficients \( A \) of the model, which will be mapped into, say, a matrix \( \Delta \),

\[
\Delta = A A',
\]  

(2.19)

again subject to restrictions of the form

\[
\delta \equiv \text{vec} \Delta = (\Lambda \otimes I_{m+n})(R\xi + r)
\]  

(2.20)

and, defining \( \delta^j \) consistently with \( \xi^j \), we write

\[
\delta^j = (\Lambda \otimes I_{m+n})(R^j\xi^j + r^j), \quad j = 1, 2.
\]  

(2.21)

The constant vectors \( r^j \) reflect, \textit{inter alia}, our choice of normalization, i.e., in practice the imposition of a unitary diagonal on \( B \). Since the elements of \( B \) are functions of \( \xi^1 \), it will prove convenient to use \( r^1 \) for this purpose. In order to prevent complicating an already heavy notation, let us, in the sequel, take \( r^2 = 0 \).

Provided that the restrictions in (2.10) then allow for a consistent solution in \( \xi^2 \) to exist, i.e., if

\[
R^2R^2+\alpha_2 = \alpha_2,
\]  

(2.22)

we can solve uniquely for \( \xi_i^2 \) in terms of \( \delta_i^2 \), \( \lambda_i \), and \( \xi_{(i-1)}^2 \) from (2.21)

\[
\xi_i^2 = R_i^2 + \left\{ \delta_i^2 + \left[ (\lambda_i' \otimes I) R_{(i-1)}^2 - R_{(i(i-1))}^2 \right] \xi_{(i-1)}^2 \right\},
\]  

(2.23)

where \( \delta^2 \) is partitioned as in (2.12).

Conformably with (2.10) we define, next,

\[
A = A^1 + A^2,
\]
or $\alpha_j = \text{vec } A^j = R^j \zeta^j + r^j$ (j = 1, 2), and we partition, as in (2.11) and (2.12),

$$A^j_{(i)} = \begin{pmatrix} A^j_{(i-1)} & a^j_i \end{pmatrix},$$

which allows us to write the likelihood of the model in the recursive parameterization as

$$\mathcal{L}(\xi, [\lambda, \omega^2]; Y, X) \propto \left[ f(\xi^1) \right]^T \prod_{i=1}^n (\omega_i^2)^{-T/2} \times \exp -\frac{1}{2} \omega_i^{-2} (W^1 - WA^1_{(i-1)}\lambda_i + W_0^2)' \times (W^1 - WA^1_{(i-1)}\lambda_i + W_0^2),$$

(2.25)

with $f(\xi^1) = \|B\|$ from (2.16), and the coefficients are still restricted by

$$\text{vec } A^j_{(i)} = R^j_{(i)} \xi^j_{(i)} + r^j_{(i)},$$

(2.26)

where $r^j_{(i)}$ just refers to the first $(m + n)i$ elements of $r^j$, and

$$\delta^2_i = R^2_{ii} \xi^2_i - \left[ (\lambda_i \otimes I) R^2_{(i-1)} - R^2_{ii} \right] \xi^2_{(i-1)},$$

(2.27)

Given the full column rank of $R^2_{ii}$, it is obvious from (2.27) that a vector defined as

$$\delta^2_i = R^2_{ii} \xi^2_i = R^2_{ii} \delta^2_i \quad (i: 1 \to n)$$

(2.28)

of dimension $l^2_i$, would not be subject to restrictions, given $\xi^2_{(i-1)}$. From (2.23) or (2.27) it is then clear that $\xi^2_i$ is fully determined by knowledge of $\delta^2_i$ and $\lambda_i$, given $\xi^2_{(i-1)}$, i.e., $\xi^2$ can be found recursively from all pairs $(\delta^2_i, \lambda_i)$. As argued in Steel (1988), it is precisely the RLI assumption (2.14) that ensures that the likelihood function can be written in terms of $\delta^2_i$, as it implies

$$W_0^2 \delta_i = W_0^2 \delta_i^2,$$

(2.29)

giving us the following representation of (2.25):

$$\mathcal{L}(\xi, [\lambda, \omega^2]; Y, X) \propto \left[ f(\xi^1) \right]^T \prod_{i=1}^n (\omega_i^2)^{-T/2} \times \exp -\frac{1}{2} \omega_i^{-2} (W^1 - Z_i \gamma_i)' \times (W^1 - Z_i \gamma_i),$$

(2.30)
where
\[ Z_i = \left( -W_i^2 : WA_{(i-1)}^1 \right), \]
and
\[ \gamma_i = \begin{pmatrix} \delta_i^2 \\ \lambda_i \end{pmatrix} \]
is an unrestricted vector, allowing us to conduct analytical inference on the elements of \( \zeta^2 \) in a recursive way, given, of course, the vector \( \zeta^1 \), which is still implicitly present in \( a_i^1 \) and \( Z_i \), both subject to the restrictions in (2.26).

2.3. The stochastic prior structure

As already mentioned in subsection 2.1 the recursive analytical approach in Steel (1988) is presented under diffuse prior information on the unrestricted coefficients of the model.

It is, however, not the only prior structure that allows for analytical inference, as we can use a specific type in the class of recursive extended natural-conjugate (RENC) prior densities introduced in Richard and Steel (1988), where we induce a linear dependence of one of the hyperparameters \( [F_{ii(i-1)}^0] \) in their notation] on \( \lambda_i \) in order to preserve prior independence across equations in terms of the recursive parameters \( \gamma_i \).

In particular, we choose for \( (\zeta^2, \Sigma) \) the product of the following densities over all equations \( i: 1 \to n \):

\[
D(\omega_i^2) = f_{i \gamma}(\omega_i^2 \mid \omega_{i0}^2, \nu_0 - n + i),
\]

\[
D(\lambda_i \mid \omega_i^2) = f_{N_i}^{-1}(\lambda_i \mid \lambda_{i0}^2, \omega_i^2 \Sigma_{(i-1)}^{0-1}),
\]

\[
D(\xi_i^2 \mid \lambda_i, \omega_i^2, \Sigma_{(i-1)}^{2}) = f_{\nu_i}^{2}\left(\xi_i^2 \mid \nu_{i0}, \omega_i^2 \Sigma_{(i-1)}^{0-1} + R_i^2 + ((\lambda_i \otimes I) R_{(i-1)}^2 - R_{ii(i-1)}^2)\right)
\times (\xi_{i-1}^2 - \xi_{i-10}^2),
\]

assuming \( N_i^0 \) to be PDS and partitioning \( \xi_0^2 \) conformably with \( \xi^2 \). To facilitate the analysis, we shall assume prior independence between \( (\zeta^2, \Sigma) \) and \( \zeta^1 \). Prior densities of \( \zeta^1 \) will be discussed at the end of this subsection. Remark that the prior densities on \( (\lambda_i, \omega_i^2) \) are compatible with an Inverted-Wishart prior structure on \( \Sigma \) of the form

\[
D(\Sigma) = f_{IW}^n(\Sigma \mid \Sigma_0, \nu_0),
\]

(2.34)
with $\nu_0 > n - 1$ and where both $\Sigma$ and $\Sigma_0$ are PDS matrices that decompose recursively into

$$
\Sigma(i) = \begin{pmatrix}
\Sigma(i-1) & \Sigma(i-1)\lambda_i \\
\lambda_i^t \Sigma(i-1) & \omega_i^2 + \lambda_i^t \Sigma(i-1)\lambda_i
\end{pmatrix},
$$

if we define $\Sigma = \Sigma(n)$. In view of (2.23) and (2.28), we can now write the prior assumptions in terms of the recursive parametrization $\gamma_i$ as (2.31) combined with

$$
D(\gamma_i|\omega_i^2, \xi_{i(i-1)}^2) = D(\gamma_i|\omega_i^2) = f_{N}^{2i-1}(\gamma_i|\gamma_i^0, \omega_i^2 H_i^{0-1}),
$$

where

$$
\gamma_i^0 = \begin{pmatrix}
\xi_{i0}^2 - R_{ii}^2 + [(\lambda_i^0 \otimes I) R_{i(i-1)}^2 - R_{i(i-1)}^2] \xi_{i(i-1)0} \\
\lambda_i^0
\end{pmatrix},
$$

and

$$
H_i^0 = \begin{pmatrix}
N_i^0 & N_i^0 R_{ii}^2 + A_{i(i-1)0}^2 \\
A_{i(i-1)0} R_{ii}^2 + N_i^0 & \Sigma(i-1) + A_{i(i-1)0} R_{ii}^2 + N_i^0 R_{i(i-1)}^2 + A_{i(i-1)0}^2
\end{pmatrix},
$$

defining $\text{vec} A_{i(i-1)0} = R_{i(i-1)}^2, \xi_{i(i-1)0}^2$. Note that prior independence across equations is imposed in terms of $\{\gamma_i, \omega_i^2\}$.

Prior elicitation can be performed using the same kind of recursive formulae as found in Steel (1988), in particular ($\forall i: 1 \rightarrow n$),

$$
E(\xi_i^2) = \xi_{i0}^2,
$$

$$
\text{cov}(\xi_i^2, \xi_j^2) = R_{ii}^2 + [(\lambda_i^0 \otimes I) R_{i(i-1)}^2 - R_{i(i-1)}^2] \text{cov}(\xi_{i(i-1)}, \xi_{j(i-1)}), \quad i > j,
$$

and

$$
\text{var}(\xi_i^2) = \frac{\omega_i^2}{\nu_0 - n + i - 2} \left\{ N_i^0 R_{ii}^2 + \sum_{k=1}^{i-1} \Sigma(i-1) R_{k-k}^2 \right\}
$$

$$
\times \text{cov}(\xi_{i(k)}^2, \xi_{i(q)}^2) R_{q(q)}^2 + R_{ii}^2 + [(\lambda_i^0 \otimes I) R_{i(i-1)}^2 - R_{i(i-1)}^2] \text{var}(\xi_{i(i-1)}) [(\lambda_i^0 \otimes I) R_{i(i-1)}^2 - R_{i(i-1)}^2],
$$

(2.38)
defining \( R_k^2 = (R_{k(k-1)}^2 : R_{kk}^2) \) and denoting by superscripts the relevant elements of the inverse of \( \Sigma_{(i-1)}^0 \). The restrictive features of natural-conjugate (NC) prior structures, i.e., proportionality of the prior covariance matrices across equations [see, e.g., Rothenberg (1963), Drèze and Richard (1983), and Steel (1988)], are avoided here by choosing a different precision matrix \( N_i^0 \) for every equation, breaking the proportionality between \( \text{var}(\zeta_i^2) \) and \( \text{var}(\zeta_j^2) \) \((i \neq j)\) from (2.38), and by reasoning in terms of \( \zeta^2 \) instead of \( A^2 \) while imposing linear restrictions in vec form, destroying the proportionality between \( \text{cov}(\zeta_i^2, \zeta_j^2) \) and \( \text{var}(\zeta_k^2) \) for \( k = \min(i, j) \). Indeed, under the matrix form restrictions in (2.9) for \( A^2 \) we notice that a structure as in (2.8) leads to \((i > j)\)

\[
\text{cov}(\zeta_i^2, \zeta_j^2) = (\lambda_i^0 \otimes I) \text{cov}(\zeta_{(i-1)}^2, \zeta_j^2),
\]

implying the latter type of proportionality, while (2.38) then simplifies to

\[
\text{var}(\zeta_i^2) = \frac{\omega_{i0}^2}{\nu_0 - n + i - 2} N_i^0 \quad \text{and}
\]

\[
+ \sum_{k, q = 1}^{i-1} \left( \frac{\omega_{i0}^2}{\nu_0 - n + i - 2} \Sigma_{(i-1)}^{0k} + \lambda_{ik}^0 \lambda_{iq}^0 \right) \text{cov}(\zeta_k^2, \zeta_q^2),
\]

still avoiding block-diagonal proportionality, which is imposed if we also give up the equation-specific precision matrices, i.e., we take \( N_i^0 = N_0 \), a PDS matrix, \( \forall i: 1 \to n \), leading to

\[
\text{var}(\zeta^2) = \frac{1}{\nu_0 - n - 1} \Sigma_0 \otimes N_0^{-1},
\]

the well-known NC result.

Apart from circumventing this NC problem, the specific class of prior densities in (2.31)-(2.33) has the advantage that its marginal prior moments are known analytically [which is generally not the case for other classes of extended natural-conjugate (ENC) prior densities, appearing, e.g., in Drèze and Richard (1983)], and thus its hyperparameters are easily elicited from assumptions on the prior moments. Remark, however, that there is a price to pay for analytical treatment here, which requires prior independence between \( \gamma_i \) and \( \gamma_j \) \((i \neq j)\), as this restricts the way in which \( \zeta_i^2 \) depends on \( \zeta_{(i-1)}^2 \) in prior mean [see (2.33)]. In fact, this means that the off-diagonal covariance structure in (2.37) is not as flexible as in the broader class of RENC densities in Richard and Steel (1988). Formally, we also need to check here whether
$N_i^0$, as elicited from our prior notions and (2.38) is PDS. On the other hand, the present structure can cope with general linear restrictions in vec form and might, in practice, prove sufficiently flexible.

Of course, if we wish to incorporate prior information on $\xi_i^2$, we should avoid being too diffuse on $\omega_i^2$, to prevent the well-known NC pitfall of inflating the relative prior precision [see Richard (1973, p. 181)]

Finally, we can always choose to remain vague about the unrestricted parameters of the model by letting $N_i^0$ and $\Sigma_0$ approach zero matrices, giving as a limit case

$$D(\xi^2, \Sigma) \propto |\Sigma|^{-\frac{1}{2}(\nu_0+n+1)} \prod_{i=1}^n (\omega_i^2)^{-\frac{1}{2}}$$

whereas, to simplify the comparison with other studies, the more usual class of diffuse prior densities

$$D(\xi^2, \Sigma) \propto |\Sigma|^{-\frac{1}{2}\mu_0} \tag{2.39}$$

is also considered, as in Steel (1988).

In fact, only in the special case where $l_i^2 = l/n, \forall i$, will the limit case exactly coincide with (2.39), implying $\mu_0 = \nu_0 + n + (l_i^2/n) + 1$. When performing the Monte Carlo analysis on the coefficients (subsections 3.1–3.3) the only consequence of choosing between the limiting density and (2.39) will be a change in the existence of posterior moments, discussed in subsection 3.5. Alternatively, a Monte Carlo analysis on the covariance matrix will be subject to slight changes in the formulae as indicated in subsection 3.4.

All that remains now is to specify a prior density for those coefficients that are not amenable to analytical treatment, grouped in $\xi^1$. Since we have to use numerical procedures for integrating $\xi^1$ anyway, as we shall discuss in the next section, we have a fully flexible choice in principle. It turns out, however, that matters are facilitated by choosing a prior structure such that $D(\xi_n^1 | \xi_{(n-1)}^1)$ takes on a Normal or a Student $t$ form (in the informative case). An easily elicited prior is, e.g., a product of independent Student densities:

$$D(\xi^1) = \prod_{i=1}^n f_{t_i}^{|\Sigma_1|(\xi_i^1 | \xi_{i0}^1, P_i^0, \nu_i^0)} \tag{2.40}$$

which has the advantage of being invariant with respect to normalization if $\nu_i^0 = 1$, i.e., the densities are of the Cauchy type; on the other hand, the latter choice for $\nu_i^0$ prevents the existence of prior moments, which might complicate the elicitation process.
For the sake of generality, we shall, in the sequel, allow for any prior density of $\zeta^1$ that is independent of $(\zeta^2, \Sigma)$ and that can be factorized as follows:

$$D(\zeta^1) = D(\zeta^1_{(n-1)} | \zeta^1_{(n-2)}) D(\zeta^1_{(n-2)}),$$

(2.41)

where both factors can, of course, be diffuse densities of the uniform type, and where the first factor is either a Student or a Normal density in informative cases.

In summary, our prior assumptions on the entire parameter space are expressed by

$$D(\zeta, \Sigma) = D(\zeta^2, \Sigma) D(\zeta^1),$$

(2.42)

where the first factor is the product of (2.33) and (2.34) [or, equivalently, of (2.31)-(2.33)] and the second factor is in the general class of densities (2.41).

3. Posterior analysis

3.1. Recursive analytical results given $\zeta^1$

The defining characteristics of $\zeta^2$ were given in subsection 2.1 as the RLI structure (2.14), ensuring that the unrestricted $\gamma_i$ appears in the likelihood function (2.30) and the condition that $\zeta^2$ should not affect the Jacobian $\|B\|^T$ in (2.16). These requirements will lead to a separability of the likelihood function across equations in terms of the recursive parameters $\{\gamma_i\}$, from which $(\zeta^2, \lambda_i)$ can be found uniquely using (2.27) and (2.28), defining a one-to-one transformation.

A similar independence across equations was induced in the prior structure by the choices made in subsection 2.3. This implies that the analysis of the posterior densities of $(\zeta^2, \Sigma)$, conditional upon $\zeta^1$, will become very simple, as in the more specific case discussed in Steel (1988).

In particular, reasoning in terms of the recursive parameterization $\{\gamma_i, \omega^2_i\}$ we obtain, $\forall i: 1 \rightarrow n$,

$$D(\omega^2_i | \gamma_i, \zeta^1, Y, X) = f_{i\gamma}(\omega^2_i | \omega^2_i, \nu_0 + l_i^2 + 2i - n - 1),$$

(3.1)

with

$$\nu_0 = \nu + T,$$

$$\omega^2_i = \omega^2_{i0} + (\gamma_i - \gamma_i^0)H_i^0(\gamma_i - \gamma_i^0) + (Wa_i^1 - Z_i\gamma_i)(Wa_i^1 - Z_i\gamma_i),$$
and
\[ D(\gamma_i|\xi^1, Y, X) = f_i^{2+i-1}(\gamma_i|\gamma_i^*, q_i^{-1}H_i^*, \nu_* + i - n), \quad (3.2) \]
where the hyperparameters are given by well-known NC formulae as
\[
H_i^* = H_i^0 + Z_iZ_i, \quad \text{assumed PDS},
\]
\[
\gamma_i^* = H_i^{*-1}(H_i^0\gamma_i^0 + Z_iW_a^0),
\]
\[
q_i^* = \omega_i^2 + \gamma_i^0H_i^0\gamma_i^0 + a_iW'a_i^T W'W_a^0 - \gamma_i^*H_i^*\gamma_i^*, \quad \text{assumed positive},
\]
obviously depending on \( \xi^1 \), the posterior density of which then becomes
\[
D(\xi^1|Y, X) \propto \left[ f(\xi^1) \right]^T D(\xi^1) \prod_{i=1}^n q_i^{-\frac{1}{2}(\nu_* + i - n)|H_i^*|^{-\frac{1}{2}}}. \quad (3.3)
\]
The factors in the latter kernel originate from, respectively, the Jacobian of the likelihood function [see (2.16)], the prior density of \( \xi^1 \) in (2.41), and the integrating constants from the posterior densities in (3.2). Its numerical treatment will be the subject of the next subsection. If, instead of the RENC prior structure for \((\xi^2, \Sigma)\) (or its limit case), we use the popular class of diffuse prior densities in (2.39), the number of degrees of freedom in (3.2), and, consequently, the factor in brackets in the exponent of \( q_i^* \) in (3.3) are both changed to \( \nu_* + T - 2n - l_i^2 + i - 1 \).

It is now straightforward to use the densities in (3.2) in order to derive the moments of \( \xi^2 \) given \( \xi^1 \) by the same type of recursive formulae as found in Steel (1988). For example, the posterior mean can be obtained from
\[
E(\xi^2_i|\xi^1, Y, X) = \tilde{\delta}_i^2 + R_{ii}^{2+} \left[ (\lambda_i^{**} \otimes I) R_{(i-1)}^{2+} - R_{i(i-1)}^2 \right] \times E(\xi_{(i-1)}^2|\xi^1, Y, X), \quad (3.4)
\]
where \( \tilde{\delta}_i^2 \) and \( \lambda_i^{**} \) are the appropriate subvectors of \( \gamma_i^* \). The posterior covariance structure, given \( \xi^1 \), can be found from exactly the same formulae as its prior equivalent in (2.37) and (2.38), replacing \( \gamma_0^0, \omega_0^2, H_i^0, \) and \( \nu_0 \) by the posterior hyperparameters \( \gamma_i^*, q_i^*, H_i^*, \) and \( \nu_* \), respectively.

Moments of \( \lambda_i \) or certain elements of \( \Lambda^{-1} \), of interest for evaluating weak exogeneity, can be found through similar calculations as in Steel (1988), taking the appropriate hyperparameters, where it has to be stressed that all these posterior results for \( \xi^2 \) and \( \Lambda \) (or its transformations) are still conditional upon \( \xi^1 \). If our main interests do not involve \((\xi^2, \Lambda)\), this lack of marginal analytical moments is of minor concern; if \((\xi^2, \Lambda)\) do appear as
parameters of interest (possibly after some transformation), we can obtain the marginal moments as a 'by-product' of the Monte Carlo analysis on \( \zeta \), by evaluating the conditional moments at each drawing and computing the sample average with the usual weighting function [see, e.g., Kloek and van Dijk (1978) or Bauwens (1984, p. 17)]. Thanks to the known analytical properties of the conditional posterior densities (3.2), which can be interpreted as \( D(\zeta^2, \lambda_j | \zeta^2_{(i-1)}, \zeta^1, Y, X) \), this will not increase the dimension of the numerical Monte Carlo integration. Details of the actual strategy proposed here will be given in subsection 3.3.

### 3.2. Posterior inference on \( \zeta \)

The numerical part of the analysis will focus upon the \( J \) coefficients grouped in the vector \( \zeta \), the posterior density of which was described in (3.3).

In order to provide a link with results stated elsewhere in the literature, consider the case where \( A^2 \) is simply matrix restricted as in (2.9) and the corresponding NC prior structure is assumed for its unrestricted coefficients. This translates into

\[
R^2 = I \otimes F_2, \tag{3.5}
\]

and

\[
N_i^0 = N_0, \quad i: 1 \rightarrow n, \tag{3.6}
\]

which we assume to be PDS, just like its posterior counterpart

\[
N_* = N_0 + F_2 W' W F_2. \tag{3.7}
\]

Applying some algebra to (3.3) then gives

\[
D(\zeta^1 | Y, X) \propto [f(\zeta^1)]^T D(\zeta^1) \times |\Sigma_0 + Z_0^2 N_0 Z_0^2 + A^T W' W A - Z_*^2 N_* Z_*^2|^{-\nu_*/2}, \tag{3.8}
\]

One might think that evaluating the conditional moments for several values of \( \zeta \) (e.g., the posterior mean) could give a rough idea of marginal results. This can, however, be extremely deceptive, especially for moments of order greater than one. Only if the prior information on \( (\zeta^2, \Sigma) \) would strongly dominate the sample information, this might be a useful procedure, in view of the prior independence between \( (\zeta^2, \Sigma) \) and \( \zeta \) in (2.42). Of course, marginal variances are based upon the mean of the conditional variance plus the variance of the conditional mean.
with
\[
N \star Z_\star^2 = N_0 Z_0^2 - F_2^2 W' W A^1,
\]
and where we have defined \( A_0^2 = F_2 Z_0^2 \) and \( \text{vec } Z_0^2 = \zeta_0^2 \). Notice that this result is a direct generalization of formula (6.5) in Drèze and Richard (1983).\(^3\) In this specific case, the functional form of the posterior density on \( \zeta \) is not affected by integrating out the matrix-restricted coefficients \( \zeta^2 \). Whenever \( \zeta^2 \) has a more flexible RLI structure, the posterior density of \( \zeta^1 \) cannot be written as (3.8), but for the numerical analysis this will not make any difference, as the determinant in (3.8) is typically not the kernel of a multivariate Student t density of \( A^1 \), due to the general restrictions we can impose on these coefficients.

Let us, therefore, concentrate upon the posterior density as given in (3.3), which immediately allows for a factorization into \( \zeta_n^1 | \zeta_{(n-1)}^1 \) and \( \zeta_{(n-1)}^1 \). Indeed, the second factor, the prior density of \( \zeta^1 \), accommodates such a factorization by definition [see (2.41)], and the first factor, the Jacobian, can be decomposed as in appendix B.4 of Drèze and Richard (1983) into a factor that no longer contains \( \zeta_n^1 \) and a Student t kernel for \( \zeta_n^1 | \zeta_{(n-1)}^1 \), which is degenerate here if \( \zeta_n^1 \) also includes some coefficients from \( \Gamma \). As \( H_n^* \) does not depend on \( \zeta_n^1 \) from the definition of \( Z_i \) following (2.30), the only other factor in the conditional posterior density for \( \zeta_n^1 \) will be \((q_n^*)^{-\nu \cdot /2} \), again a Student kernel. We have, therefore, the following factorization:
\[
D(\zeta^1 | Y, X) = D(\zeta_n^1 | \zeta_{(n-1)}^1, Y, X) D(\zeta_{(n-1)}^1 | Y, X).
\]

If we take the Student option for \( D(\zeta_n^1 | \zeta_{(n-1)}^1) \) in (2.41), the first factor is a 2-1 poly-t density,\(^4\) exactly as in lemma 6.2 of Drèze and Richard (1983). The crucial difference with their analysis, however, is that we are now only reasoning in the space of \( \zeta^1 \), which is of dimension \( l_t \), instead of the full \( l \)-dimensional space of all unrestricted coefficients in \( \zeta \). As we shall see later, this can be a major advantage in performing our numerical integrations.

If our prior assumptions on \( \zeta^1 \) are changed to a diffuse (uniform) prior, the conditional posterior density for \( \zeta_n^1 \) simplifies to a 1-1 poly-t density, pseudo-random drawings from which are easily obtained following Bauwens

\(^3\)This becomes even more obvious for \( N_0 = 0 \), i.e., a diffuse prior structure on \( \zeta^2 \), when we obtain \( | \Sigma_0 + A^1 W' M_\star W A^1 | \) for the determinant in (3.8), with \( M_\star = I_f - W F_2 (F_2' W' W F_2)^{-1} F_2' W' \), where the inverse exists due to the full column rank of \( W_i^2 = W F_2, \forall i: 1 \rightarrow n \). Of course, (3.8) only becomes exactly equal to the result of Drèze and Richard (1983) if \( F_2 = 0 \), i.e., if we group all the free coefficients in \( \zeta^1 \), as they do.

\(^4\)For definitions of poly-t densities and an overview of their applications in econometrics, we refer the reader to the seminal paper of Drèze (1977). Suffice it to say here that the kernel of a general \( m-n \) poly-t density is defined as the ratio of the product of \( m \) Student t kernels over the product of \( n \) such kernels.
and Richard (1985). If, on the other hand, we wish to express our conditional prior opinions on $\xi_n^1$ by a Normal instead of a Student $t$ density, its posterior counterpart will be the product of a 1–1 poly-$t$ kernel and a Normal kernel, which can be analyzed with the same techniques as 2–1 poly-$t$ densities, using the results from section 3 in Richard and Tompa (1980). For the next subsection we shall, therefore, focus on the Student option, whereas the Normal prior density will be used in subsection 3.4.

3.3. The Monte Carlo strategy

In order to perform the necessary numerical integrations, we shall use the method of Monte Carlo analysis with importance sampling, where we should like to perform our pseudo-random drawings according to a so-called importance function that closely approximates the density under scrutiny (i.e., the posterior density of $\xi^1$ in this case), yet allows for efficient algorithms to generate these drawings. As such techniques are treated in great detail in, e.g., Hammersley and Handscomb (1979), van Dijk (1984), and Bauwens (1984), we shall refrain from a more extensive account of the underlying theory here.

In an effort to minimize the dimension of the actual Monte Carlo problem, we shall make full use of the poly-$t$ structure of the conditional posterior for $\xi_n^1$, as in the PTST-$i$ (poly-$t$ Student, conditional posterior for equation $i$) method described in Bauwens (1984, p. 29). If we are interested in some integrable function $g(\Sigma, \xi)$ of the unrestricted parameters of our model, we first write its expectation as

$$E g(\Sigma, \xi) = E_{\xi_n^1} E_{\xi_n^1 | (\xi_{n-1}, \xi)} E_{\lambda, \xi_n^1} E_{\omega_i^2 | \lambda, \xi} \left[ g(\Sigma, \xi) \right] J ,$$

(3.11)

with the Jacobian $J = \prod_{i=1}^{n-1}(\omega_i^2)^{n-i}$ as derived in appendix C of Steel (1988), and where the subscripts indicate that we take the expected value with respect to the posterior density of those parameters. The last expectation is known analytically from (3.1), the third one can be evaluated analytically using recursive formulae and a one-to-one transformation on (3.2), and for the second one, corresponding to the first factor in (3.10), we can take advantage of the combined analytical and numerical techniques developed for poly-$t$ densities in, e.g., Richard and Tompa (1980) and Bauwens and Richard (1985). Finally, the only expectation that requires purely numerical Monte Carlo procedures for its evaluation is the first one, which is taken with respect to the second factor in (3.10), a density that defies analytical treatment altogether.

The evaluation of the second expectation in (3.11) can, in principle, be based on the algorithms for integrating poly-$t$ densities as found in the
computer software PTD, documented in Bauwens et al. (1981). The latter software can, however, not cope with moments of higher order than two or moments of nonlinear functions and requires that we write the result of the last two, analytical, expectations as an explicit function of $\xi_n^1$. Although the latter is not impossible, given the hyperparameters in (3.2) and the recursive formulae in (3.4) and as in (2.37) and (2.38), it might quickly lead to problems PTD cannot handle and might require considerable programming effort.

Therefore, we propose to draw values for $\xi_n^1$, given $\xi_{(n-1)}$, which can be done directly in the case of a 1--1 poly-t density and requires some importance sampling when we are faced with a 2--1 poly-t density, using a 1--1 poly-t density as importance function. The latter case would, of course, increase the overall dimension of the Monte Carlo integration using importance sampling, but, from the results in Bauwens and Richard (1985), it seems that this type of numerical integration can be performed rather efficiently.

In summary, we suggest the following strategy for conducting posterior inference on $g(\Sigma, \xi)$:

1. Draw a value for the $(l_1 - l_n^1)$-dimensional vector $\xi_{(n-1)}^1$, according to some importance function $D_{IF}(\xi_{(n-1)}^1 | Y, X)$. Denote the $k$th value drawn by $\xi_{(k)}^1$.  

2. Conditionally on $\xi_{(k)}^1$, draw a vector $\xi_n^1$ from a 1--1 poly-t density function in $l_n^1$ dimensions, say $D_{1-I}(\xi_n^1 | \xi_{(k)}^1, Y, X)$, which is either the actual conditional posterior density or an approximation to it constructed according to the principles in Bauwens and Richard (1985). We have then obtained a drawing for the entire vector $\xi^1$, which we call $\xi_{(k)}^1$.  

3. Given $\xi_{(k)}^1$, evaluate $g(\Sigma, \xi)$ analytically, using the conditional posterior densities in (3.1) and (3.2), and the recursive approach.

After $N$ replications of these three steps, we approximate the actual expected value of $g(\Sigma, \xi)$ by the weighted sample mean where now, for the sake of brevity, the analytical expectation is taken with respect to $\Sigma$,

$$
\hat{g} = \frac{\sum_{k=1}^{N} E_{\Sigma, \xi^2}(\xi_{(k)}^1)g(\Sigma, \xi_{(k)}^1, \xi_{(k)}^2)w(k)}{\sum_{k=1}^{N} w(k)},
$$

(3.12)

with the weight function

$$
w(k) = \frac{D(\xi_{(k)}^1 | \xi_{(n-1)}^1, Y, X)D(\xi_{(n-1)}^1 | Y, X)}{D_{1-I}(\xi_n^1 | \xi_{(k)}^1, Y, X)D_{IF}(\xi_{(n-1)}^1 | Y, X)},
$$

(3.13)
the ratio of the actual posterior densities in (3.10) and the importance functions drawn from, both evaluated at the \( k \)th value drawn. Note that, whenever the conditional posterior density of \( \xi_n^1 \) is of the 1–1 poly-t form, the weights will simplify to

\[
W^{(k)} = \frac{D(\xi_{n-1}^1, Y, X)}{D_{IF}(\xi_{n-1}^1, Y, X)},
\]

(3.14)

and only the numerical integration over \( \xi_{n-1}^1 \) will contribute to the estimated relative error bounds as defined in Kloek and van Dijk (1978) and Bauwens (1984, p. 18). Even in the 2–1 poly-t case for \( \xi_{n-1}^1 \), it is to be expected, given the encouraging results of Bauwens and Richard (1985), that the main difficulties in the Monte Carlo procedure will arise from integrating over \( \xi_{n-1}^1 \), so that the choice of its importance function seems critical.

Bauwens (1984, p. 25) suggests an overall Student t density

\[
D_{IF}(\xi_{n-1}^1, Y, X) = f_{t}^{(n-1)}(\xi_{n-1}^1, \xi_{n-1}^1, P_*, \mu_*),
\]

(3.15)

where the choice of hyperparameters can be based on, respectively, the posterior mode, the Hessian matrix, and the order of existence of moments. Provided our prior \( D(\xi_{n-1}^1) \) in (2.41) falls in the class of ENC prior densities, the mode and the Hessian can be evaluated by FIML techniques, as explained in Morales (1971), and the order up to which posterior moments are sure to exist is given in lemma 6.6 in Drèze and Richard (1983). A possible alternative method for choosing hyperparameters in (3.15) is to iteratively update them, based on the posterior moments resulting from small Monte Carlo runs, used only for ‘calibrating’ the importance function.

Of course, (3.15) is still very ‘well-behaved’, as it imposes both symmetry and unimodality. If such assumptions are strongly violated by the actual posterior density, we should consider using other classes of importance functions, possibly based on poly-t densities, along the lines of Bauwens’ (1984) further suggestions.

As the posterior analysis of the last equation’s coefficients in \( \xi^1 \) is based on the rather well-known (given some numerical effort) properties of its conditional posterior density, whereas this extra information is lacking in the case of \( \xi_{n-1}^1 \), one might expect that the results are not fully invariant with respect to the ordering of the equations. Remember, however, that in practice this ordering will, to a large extent, be directed by the effort to group as many coefficients as possible in \( \xi^2 \), which allows for (conditional) analytical treatment. Should this criterion still leave us with a choice, then one could think of putting an ‘equation of special interest’ last or one could be guided by
considerations of dimension or covariance structure as in Bauwens (1984, p. 30).

3.4. Monte Carlo analysis of the covariance matrix

In the previous subsections we have adopted the usual practice of first integrating out the covariance matrix, using a recursive decomposition as in the prior specification (2.31) and (2.32). Indeed, combining the likelihood function (2.25), in terms of \( \xi \), with the prior structure (2.31)–(2.33), we can deduce a similar factorization of the posterior density, given \( \xi \), as

\[
D(\omega^2|\xi, Y, X) = f_{i_0} \left( \omega^2|\omega^2_{i_0} + \lambda_i^0 \Sigma^0_{(i-1)} \lambda_i^0 \right)
+ \left( a_i^2 - a_{i_0}^2 \right)' R_{ii}^2 + N_i^0 R_{ii}^2 + (a_i^2 - a_{i_0}^2) + a_i^0 W' W a_i
- \lambda_i^* G_i^* \lambda_i^* \nu_\lambda - n + l_1^2 + i \right),
\]

(3.16)

\[
D(\lambda^2|\omega^2, \xi, Y, X) = f_{N}^{-1} \left( \lambda^2|\lambda^2_\lambda^* = \nu_\lambda^* - n + l_2^2 + i \right),
\]

(3.17)

where \( a_i \) is the \( i \)th column of \( A \), just like \( a_i^2 \) in (2.24), the prior equivalent of which also holds in prior mean with subscript 0, and we define

\[
G_i^* = \Sigma^0_{(i-1)} + \left( A_i^2_{(i-1)} - A_i^2_{(i-1)0} \right)' R_{ii}^2 + N_i^0 R_{ii}^2 + \left( A_i^2_{(i-1)} - A_i^2_{(i-1)0} \right)
+ A_i^0_{(i-1)} W' W A_i (i-1),
\]

(3.18)

a relative precision matrix which we assume to be nonsingular, and

\[
G_i^* \lambda_i^* = \Sigma^0_{(i-1)} \lambda_i^0 + \left( A_i^2_{(i-1)} - A_i^2_{(i-1)0} \right)' R_{ii}^2 + N_i^0 R_{ii}^2 + (a_i^2 - a_{i_0}^2)
+ A_i^0_{(i-1)} W' W a_i.
\]

(3.19)

If \( A^2 \) happens to be matrix-restricted as in (2.9), leading to (3.5) and (3.6), the densities (3.16) and (3.17) are compatible with an Inverted-Wishart density on the full covariance matrix, which then takes the form

\[
D(\Sigma|\xi, Y, X) = f_{\nu} \left( \Sigma|\Sigma_0 + (Z^2 - Z_0^2)' N_0 (Z^2 - Z_0^2) \right)
+ A' W' W A, \nu_\lambda^* + l_2/n \right),
\]

(3.20)
a straightforward generalization of formula (6.4) in Drèze and Richard (1983). It is, therefore, clear that only if the restrictions on \( A^2 \) are of a more general nature (i.e., vec form under the RLI assumption) do we gain by treating the model equation-by-equation in integrating out the covariance structure.

Note that the integrating constant from (3.20) will lead to the following posterior density for the matrix of unrestricted coefficients \( Z^2 \):

\[
D(Z^2|\xi^1, Y, X) = f_{M_{\delta}}^{1/\delta} n x n (Z^2|Z_{**}, \Sigma_0 + Z_0^2 N_0 Z_0^2 - Z_{**}^2 N_{**} Z_{**}) + A^1 W' W A^1, N_{**}, \nu_{**})
\]

(3.21)

using the definitions in (3.7) and (3.9), a simple NC outcome, as expected, under matrix restrictions on \( A^2 \). Finally, this gives us the same marginal posterior density as found in (3.8) for \( \xi^1 \).

Of course, the order of the integrations could be reversed. In particular, Richard and Steel (1988) find that in SURE models it might be preferable to integrate out the coefficients analytically (under an RENC prior structure) given \( \Sigma \) and, subsequently, perform the Monte Carlo analysis on \( \Sigma \), using an Inverted-Wishart importance function. The latter allows for a simple iterative strategy to calibrate its hyperparameters, and, heuristically, it is felt that the vec form restrictions in (2.7) might induce a larger departure from the NC posterior results for the coefficients on which the restrictions are imposed than for the covariance matrix, which is not subject to restrictions in our analysis. Let us, now, briefly sketch the possibility of a similar treatment of the SEM discussed here.

The RENC prior density of \( \xi^2 \), given \((\lambda_i, \omega_i^2, \xi_i^{2|\xi_i^{1}})\) in (2.33), implies a Normal prior structure for the entire vector \( \xi^2 \), given \( \Sigma \). As our likelihood function in (2.25) can also be written in the Normal form for \( \xi^2 \), given \( \xi^1 \) and \( \Sigma \), this structure is preserved in the posterior density, from which we can simply integrate out \( \xi^2 \) analytically, conditional upon \( \xi^1 \) and \( \Sigma \).

The next step, however, is much less straightforward, as we then have to integrate out \( \xi^1 \), given \( \Sigma \). This is complicated by the presence of the Jacobian factor \( \|B\|^T \), which depends on \( \xi^1 \) [see (2.16)]. Only if we are in a triangular or a SURE system and provided the prior of \( \xi^1 \) is given a Uniform or Normal structure, \( \xi^1 \) can easily be handled (using the properties of a Normal density). We are, subsequently, left with an untractable posterior density for \( \Sigma \), which, following Richard and Steel (1988), can be well approximated by an importance function of the Inverted-Wishart type. The fact that, presently, this order of integration is limited to triangular or SURE models is a major

A triangular system is here taken to imply a triangular structure on the endogenous variables, i.e., \( B \) is upper- or lower-triangular, usually normalized to have ones on the diagonal, but the error covariance structure is still of a general form with \( \Sigma = B'VB \) as in (2.3), instead of the independence inherent in a recursive system as in subsection 2.2.
Table 1
Overview of posterior inference.

<table>
<thead>
<tr>
<th></th>
<th>Monte Carlo (MC) on the coefficients (subsections 3.1–3.3)</th>
<th></th>
<th>MC on $\Sigma$ (3.4)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td>$l_1 = 0$</td>
<td>$l_1 = l_n^1 &gt; 0$</td>
<td>$l_1 &gt; l_n^1 &gt; 0$</td>
</tr>
<tr>
<td>$D(\zeta_n^1</td>
<td>\zeta_{n-1}^1)$</td>
<td>—</td>
<td>NI</td>
</tr>
<tr>
<td>$D(\zeta_{n-1}^1)$</td>
<td>—</td>
<td>—</td>
<td>—</td>
</tr>
<tr>
<td>SEM</td>
<td>—</td>
<td>PTD or direct simulation</td>
<td>Direct simulation + MC [$l_1$]</td>
</tr>
<tr>
<td>$</td>
<td></td>
<td>B</td>
<td></td>
</tr>
<tr>
<td></td>
<td>St</td>
<td>2-0 poly-t</td>
<td>Direct simulation</td>
</tr>
<tr>
<td></td>
<td>Analytically</td>
<td>PTD or direct simulation</td>
<td>MC [$l_1 - l_n^1$] + MC [$l_1 - l_n^1$]</td>
</tr>
</tbody>
</table>

drawback of the approach in this subsection, which is, therefore, not described in detail.

Within the, rather narrow, bounds of its domain of application, however, the numerical performance of this approach may prove to be superior for the reasons outlined earlier. In addition, the dimension of the Monte Carlo integration may also be much smaller than $l$ when drawing on $\Sigma$, as discussed in Steel (1987). Some further research is required to investigate the possibility of using the efficient algorithms in Richard and Tompa (1980) for analyzing quadratic forms in Normal variables, in an effort to render this, numerically promising, approach feasible for SEM-type systems.

3.5. A survey of posterior results

Having presented a collection of posterior results under two distinct integration sequences, we shall now attempt to systematically indicate the implications of both strategies under different model and prior assumptions.

Table 1 groups these results, reporting the type of densities appearing as posterior densities for $\zeta_n^1$ given $\zeta_{n-1}^1$, whenever we draw on the coefficients, along with the method we would suggest for conducting inference, as well as the dimension of the integrations by Monte Carlo (in square brackets) if the latter proves necessary. The term ‘direct simulation’ is taken to imply that we can perform pseudo-random drawings from the actual posterior density, and ‘PTD’ indicates that we can use this computer software directly, provided our interest in $\zeta_n^1$ is limited to integrating constants, first- and second-order moments, fractiles, or complete univariate marginal density functions, and all
elements of $\zeta^1$ are grouped in the last equation, following the discussion in
subsection 3.3. As we know from the last subsection, the option of perform-
ing the Monte Carlo drawings on $\Sigma$ is not presently available for the general
SEM, indicated in the table by 'N.A.'. Other abbreviations used in the table
are 'NI' for noninformative, 'N' for a Normal, and 'St' for a Student $t$
density.

As we implicitly define SEM in the table as a system leading to a
nonconstant Jacobian $\|B\|^T$ and since $\zeta^1$ is defined as the vector of coeffi-
cients that $\|B\|$ can possibly depend on, we feel that SEM is incompatible
with the case where $l_1 = 0$, as this would essentially imply we are in the case
where $\|B\| = 1$ (or any other constant). The latter case is, computationally,
the most simple one and was analyzed in Steel (1988), where attention is
focused on the SURE model, i.e., $B$ is taken to be the identity matrix. From
our more general framework here it becomes obvious that these simple
analytical results carry over to the case where $B$ is triangular with constant
diagonal elements, provided, of course, the RLI assumption continues to
hold (i.e., $l_1 = 0$). The case where $l_n^1 > 0$ and we are noninformative on $\zeta^1_n$,
given $\xi_{(n-1)}$, is, in addition, equivalent to including $\zeta_n^1$ in $\zeta^2$ for triangular or
SURE systems, which means the two lower left entries in the table coincide.

Remark that the Monte Carlo strategy as given in subsection 3.3 strictly
applies only to SEM systems with $l_1 > l^1_n > 0$, the most complicated case, but
that it can easily be adapted to cover the other, less demanding, cases
appearing in table 1 (barring the last column, of course).

The different entries in the table will also entail different consequences for
the existence of posterior moments. The treatment of $\zeta^2$, however, is similar
for all cases where we draw on the coefficients, and from (3.2) we obtain that
the conditional posterior moments of $\zeta^2$, given $\zeta^1$, will exist at least up to
order

$$\rho < \nu_* - n + 1,$$

under the RENC prior specification or its limiting noninformative case, and
up to order

$$\rho < \mu_0 + T - 2n - \max_{l_i \in n} \{l_i^2 - i\} - 1,$$

if we choose the usual noninformative density in (2.39), which ensures the
existence of posterior moments of order

$$\rho < \mu_0 + T - 2n - l_n^2,$$

in view of the RLI assumption (i.e., $\max_i \{l_i^2\} = l_n^2$).
Table 2

Existence of posterior moments for $\xi_n^I \mid \xi_{(n-1)}^I$.

<table>
<thead>
<tr>
<th>$D(\xi^2, \Sigma)$</th>
<th>RENC (2.31)-(2.33)</th>
<th>NI (2.39)</th>
</tr>
</thead>
<tbody>
<tr>
<td>$D(\xi^I_n \mid \xi_{(n-1)}^I)$</td>
<td>NI</td>
<td>St</td>
</tr>
<tr>
<td>SEM</td>
<td>$\nu_0 - l_n^I$</td>
<td>$\tau^0_n + \nu_0$</td>
</tr>
<tr>
<td></td>
<td>$-l_n^I - 1$</td>
<td></td>
</tr>
<tr>
<td>$|B| = 1$</td>
<td>$\nu_0 - l_n^I + T$</td>
<td>$\tau^0_n + \nu_0 + T$</td>
</tr>
<tr>
<td></td>
<td>$-l_n^I + T - 1$</td>
<td></td>
</tr>
</tbody>
</table>

Of course, (3.2) is in terms of $\gamma_i$ instead of $\xi^2_i$, but it can easily be shown that moments for $\gamma_i$ and for $\xi^2_i$ should exist up to the same order. Following Drèze (1977), we can also give sufficient conditions for the existence of posterior moments for $\xi^I_n$, conditional on $\xi_{(n-1)}^I$, provided its prior density in (2.41) is of the Student $t$ form with $\tau^0_n$ degrees of freedom and a PDS precision matrix. Assuming that $q_n^*$ can be written as a nondegenerate kernel, we obtain the existence results grouped in table 2, where both the RENC prior structure as in (2.31)-(2.33) and the class of diffuse prior densities in (2.39) are considered, and we draw in the space of the coefficients.

Conditional posterior moments will exist up to any order smaller than the expressions recorded in the table, where we have used the same abbreviations as in table 1, and we have defined $l_n = l_n^I + l_n^2$.

Marginal moments for all coefficients will, under (2.39), at least exist for orders

$$\rho < \mu_0 - 2n - \max_{i: 1 \to n} \{l_i\},$$


4. An application to Klein's model I

This model has been the focus of several Bayesian analyses, in particular in van Dijk and Kloek (1980) and Bauwens (1984) under full information and in Zellner et al. (1988) in a limited-information framework. It is a small

---

6 This amounts to assuming that the relative precision matrix,

$$W_n^I \left[ I - Z_n H_n^{*-1} Z_n^* \right] W_n^I,$$

is nonsingular. From Drèze (1977) we know that only the nondegenerate kernels from the numerator will contribute to these sufficient conditions for the existence of moments.
macroeconomic model, introduced and estimated on annual U.S. data for 1921 through 1941 in Klein (1950), who also lists the series used.

In its original form, the model comprises seven equations, three of which are stochastic and contain twelve free coefficients. As in Bauwens (1984), we exclude the intercepts and measure the variables in deviations around their sample mean values. Including intercepts explicitly in the stochastic equations of the model does not really complicate the analysis, provided they are not subject to any general prior restrictions. Indeed, we then typically lack prior information about their values and since they do not affect the Jacobian $\|B\|^T$ either, they can be integrated out analytically, given the other coefficients of the model [see also van Dijk and Kloek (1980)]. Of course, this is not surprising as these intercepts can then clearly be classified as elements of $\xi^2$. Since they are generally not of interest to the model user, they can simply be ignored in the way described here. The choice between including or excluding the intercept terms can, however, influence the degrees of freedom. Taking a simple diffuse prior for $\Sigma$ as in (2.39), it can be shown that the prior degrees of freedom $\mu_0$ should take the form

$$\mu_0 = \mu'_0 + m,$$

(4.1)

where $\mu'_0$ does not vary with $m$, the number of variables in $X$, in order to obtain exactly the same posterior outcomes in both cases. Remark that Drèze's (1976) suggestion for $\mu_0$ satisfies (4.1), in contrast with the choices of Zellner (1971), Stewart (1987), and Malinvaud (1981).

The three behavioural equations are then formulated as

$$c_t = \tilde{\alpha}_1 \tilde{p}_t + \tilde{\alpha}_2 \tilde{p}_{t-1} + \tilde{\alpha}_3 \tilde{w}_t + u_{1t},$$

(4.2)

$$i_t = \tilde{\beta}_1 \tilde{p}_t + \tilde{\beta}_2 \tilde{p}_{t-1} + \tilde{\beta}_3 \tilde{k}_{t-1} + u_{2t},$$

(4.3)

$$\tilde{w}_{1t} = \tilde{\gamma}_1 \tilde{x}_t + \tilde{\gamma}_2 \tilde{x}_{t-1} + \tilde{\gamma}_3 \text{time} + u_{3t},$$

(4.4)

where Normal errors are assumed as in (2.3). In (4.2) consumption expenditure at time $t$ ($c_t$) is explained by profits ($\tilde{p}_t$), lagged profits, and total wages ($\tilde{w}_t$). The investment ($\tilde{i}_t$) equation also contains the capital stock at the beginning of the year ($\tilde{k}_{t-1}$), and, finally, the wage bill of the private sector ($\tilde{w}_{1t}$) depends on net private production ($\tilde{x}_t$), the latter one period lagged, and a trend term. Four identities introduce three more exogenous variables, namely the government wage bill ($\tilde{w}_{2t}$), government nonwage expenditure (including the net foreign balance) $\tilde{g}_t$, and business taxes $\tilde{t}_t$. Tildes were added in order to avoid confusion with the notation used in previous subsections.
These identities, which can be found in any of the references cited above for Klein’s model I, do not contain any unknown parameters and we can, therefore, simply restrict our attention to a ‘solved out’ system of three equations, corresponding directly to (2.2)–(2.3).

As there are different possible choices in substituting the identities, we focus on a choice that implies a maximum value for \( l_2 \), i.e., the purely analytical dimension. Such a specification is

\[
\begin{pmatrix}
\tilde{c} & \tilde{e} & \tilde{w}_1
\end{pmatrix}
\begin{pmatrix}
1 - \tilde{\alpha}_1 & -\tilde{\beta}_1 & -\tilde{\gamma}_1 \\
-\tilde{\alpha}_1 & 1 - \tilde{\beta}_1 & -\tilde{\gamma}_1 \\
\tilde{\alpha}_1 - \tilde{\alpha}_3 & \tilde{\beta}_1 & 1
\end{pmatrix}
\begin{pmatrix}
\tilde{y}_1 \\
\tilde{y}_2 \\
\tilde{y}_3
\end{pmatrix}
\]

\[
+ \begin{pmatrix}
\tilde{g} & \tilde{e} & \tilde{p}_{-1} & \tilde{w}_2 & \tilde{k}_{-1} & \tilde{w}_{-1} & \tilde{t}_{-1} & \text{time}
\end{pmatrix}
\begin{pmatrix}
-\tilde{\alpha}_1 & -\tilde{\beta}_1 & -\tilde{\gamma}_1 \\
\tilde{\alpha}_1 & \tilde{\beta}_1 & 0 \\
-\tilde{\alpha}_2 & -\tilde{\beta}_2 & -\tilde{\gamma}_2 \\
-\tilde{\alpha}_3 & 0 & 0 \\
0 & -\tilde{\beta}_3 & 0 \\
0 & 0 & -\tilde{\gamma}_2 \\
0 & 0 & -\tilde{\gamma}_3
\end{pmatrix}
\begin{pmatrix}
u_1 \\
u_2 \\
u_3
\end{pmatrix},
\] (4.5)

where the absence of time indices indicates we are treating \( T \)-dimensional vectors of observations and error terms, and lagged values are denoted by the subscript \(-1\).

From (4.5) we can single out four parameters that do not appear in \( B \) and enter \( \Gamma \) in such a way that subsequent equations do not become more restricted (the RLI assumption), associated with the variables \( \tilde{p}_{-1} \) and time. The specification thus suggests the following classification of the nine free coefficients:

\[
\zeta^1 = \begin{pmatrix}
\tilde{\alpha}_1 & \tilde{\alpha}_3 & \tilde{\beta}_1 & \tilde{\beta}_3 & \tilde{\gamma}_1
\end{pmatrix}',
\] (4.6)

\[
\zeta^2 = \begin{pmatrix}
\tilde{\alpha}_2 & \tilde{\beta}_2 & \tilde{\gamma}_2 & \tilde{\gamma}_3
\end{pmatrix}',
\] (4.7)
where $\xi_n^1$ comprises $\tilde{y}_1$, implying that the remaining 'pure' Monte Carlo integration will be in four dimensions instead of nine as in the previous studies. We shall analyze the model under two distinct prior densities. The first choice is a diffuse prior structure as in (2.39) (denoted by 'NI') with $\mu_0 = m + n + 1 = 12$, following Drèze’s (1976) invariance argument and satisfying (4.1). This is sufficient to ensure the existence of second-order marginal posterior moments for $\xi$, fourth-order conditional moments for $\xi^1_n$ given $\xi_{(n-1)}$, and allows posterior moments for $\xi^2$ conditional on $\xi^1$ to exist up to the order $T + 3$ (remember that conditionally on $\xi^1$ we are really in a SURE system, which explains the extra $T$ moments). We refer to subsection 3.5 for details concerning the existence of moments.

A second prior structure is a combination of (2.39) with truncation of the parameter space for the drawn coefficients ($\xi^1$) to the interval $(0, 1)$, except for $\beta_3$ which is restricted to lie in $(-1, 0)$. Let us call this prior 'TR' and, again, we take $\mu_0 = 12$.

Table 3 summarizes our results. The skewness coefficients calculated in the table are defined as the ratio of the third central moment to the cubed standard deviation. Estimated relative error bounds are calculated at a 95% confidence level for the posterior means in table 3 ($\epsilon_\xi$) and for the integrating constant in table 4 ($\epsilon_{\text{int}}$). These error bounds are expressed in percentage points, and we refer to Kloek and van Dijk (1978) or Bauwens (1984) for their discussion.

For the Monte Carlo part of our analysis, we used a simple multivariate Student $t$ importance function with three degrees of freedom under NI, so that the same order of moments exist as we know to exist for the posterior (under TR we take five degrees of freedom; the far tails do not matter here anyway). We then used a very simple iterative calibration by performing some

---

Table 3

<table>
<thead>
<tr>
<th>Prior</th>
<th>$\tilde{\alpha}_1$</th>
<th>$\tilde{\alpha}_2$</th>
<th>$\tilde{\alpha}_3$</th>
<th>$\tilde{\beta}_1$</th>
<th>$\tilde{\beta}_2$</th>
<th>$\tilde{\beta}_3$</th>
<th>$\tilde{\gamma}_1$</th>
<th>$\tilde{\gamma}_2$</th>
<th>$\tilde{\gamma}_3$</th>
</tr>
</thead>
<tbody>
<tr>
<td>Mean</td>
<td>-0.18</td>
<td>0.33</td>
<td>0.86</td>
<td>0.14</td>
<td>0.62</td>
<td>-0.14</td>
<td>0.39</td>
<td>0.18</td>
<td>0.19</td>
</tr>
<tr>
<td>Std. dev.</td>
<td>0.18</td>
<td>0.14</td>
<td>0.05</td>
<td>0.16</td>
<td>0.15</td>
<td>0.03</td>
<td>0.05</td>
<td>0.04</td>
<td>0.04</td>
</tr>
<tr>
<td>NI</td>
<td>Skewness</td>
<td>-1.4</td>
<td>-0.76</td>
<td>-1.5</td>
<td>-0.55</td>
<td>-0.67</td>
<td>-0.22</td>
<td>-0.37</td>
<td>-0.37</td>
</tr>
<tr>
<td>$\epsilon_\xi$ (in %)</td>
<td>2.3</td>
<td>0.86</td>
<td>0.15</td>
<td>2.5</td>
<td>0.47</td>
<td>0.46</td>
<td>0.22</td>
<td>0.37</td>
<td>0.37</td>
</tr>
<tr>
<td>Mean</td>
<td>0.05</td>
<td>0.18</td>
<td>0.83</td>
<td>0.19</td>
<td>0.58</td>
<td>-0.15</td>
<td>0.38</td>
<td>0.19</td>
<td>0.20</td>
</tr>
<tr>
<td>Std. dev.</td>
<td>0.04</td>
<td>0.07</td>
<td>0.03</td>
<td>0.10</td>
<td>0.10</td>
<td>0.03</td>
<td>0.04</td>
<td>0.03</td>
<td>0.03</td>
</tr>
<tr>
<td>TR</td>
<td>Skewness</td>
<td>1.2</td>
<td>-0.09</td>
<td>0.10</td>
<td>-0.22</td>
<td>-0.44</td>
<td>-0.17</td>
<td>-0.26</td>
<td>-0.24</td>
</tr>
<tr>
<td>$\epsilon_\xi$ (in %)</td>
<td>1.4</td>
<td>0.44</td>
<td>0.07</td>
<td>0.92</td>
<td>0.26</td>
<td>0.29</td>
<td>0.17</td>
<td>0.26</td>
<td>0.24</td>
</tr>
</tbody>
</table>

These are not given for $\xi^2$, since only the conditional skewness (i.e., given $\xi^1$) was calculated for these coefficients.
Table 4

Estimated variation coefficients and relative error bounds of reciprocal of integrating constant.

<table>
<thead>
<tr>
<th>Prior</th>
<th>Number of drawings</th>
<th>Percentage of rejections</th>
<th>Variation coefficient</th>
<th>$\varepsilon_{\text{int}}$ (in %)</th>
</tr>
</thead>
<tbody>
<tr>
<td></td>
<td></td>
<td></td>
<td></td>
<td></td>
</tr>
<tr>
<td>NI</td>
<td>5,000</td>
<td>-</td>
<td>1.036</td>
<td>5.7</td>
</tr>
<tr>
<td></td>
<td>10,000</td>
<td>-</td>
<td>1.052</td>
<td>4.1</td>
</tr>
<tr>
<td></td>
<td>20,000</td>
<td>-</td>
<td>1.030</td>
<td>2.9</td>
</tr>
<tr>
<td></td>
<td>50,000</td>
<td>-</td>
<td>1.041</td>
<td>1.8</td>
</tr>
<tr>
<td></td>
<td>100,000</td>
<td>-</td>
<td>1.046</td>
<td>1.3</td>
</tr>
<tr>
<td>TR</td>
<td>5,000</td>
<td>15.5</td>
<td>0.8310</td>
<td>4.6</td>
</tr>
<tr>
<td></td>
<td>10,000</td>
<td>16.1</td>
<td>0.8441</td>
<td>3.3</td>
</tr>
<tr>
<td></td>
<td>20,000</td>
<td>15.6</td>
<td>0.8502</td>
<td>2.4</td>
</tr>
<tr>
<td></td>
<td>50,000</td>
<td>16.0</td>
<td>0.8518</td>
<td>1.5</td>
</tr>
<tr>
<td></td>
<td>100,000</td>
<td>15.8</td>
<td>0.8572</td>
<td>1.1</td>
</tr>
</tbody>
</table>

small Monte Carlo runs. Since the weights only depend on $\zeta_{(n-1)}^1$ as in (3.14), we can limit ourselves to just drawing from a four-dimensional Student $t$ density in this calibration, which makes it very fast indeed (an 8 MHz 80287 processor took about 0.035 sec. per drawing with our MS-Fortran programmes, whereas a 25 MHz 80387 reduced that to 0.005 sec. per drawing). Due to the very low dimension of the actual Monte Carlo integration, the numerical stability of the results is quite exemplary. Where Bauwens (1984, p. 49) commented that a Student $t$ importance function in nine dimensions without truncation led to exploding variation coefficients of the reciprocal of the posterior integrating constant (6.622 for 10,000 drawings and 16.50 for 50,000 drawings), we find a very constant value of around one. See table 4 for details.

Once the calibration completed (in fact, two or three runs of 1000 drawings proved sufficient), we included drawing $\zeta_n^1$, given $\zeta_{(n-1)}^1$, from a 1-1 poly-$t$ density (its actual posterior density under our prior assumptions) and calculated first- and second-order moments for $\zeta^2$, given $\zeta^1$. Computation time$^8$ per drawing was then around 0.7 sec. for an 8 MHz 80287 machine and 0.11 sec. for a 25 MHz 80387 equipped PC.

The results in tables 3 and 4 reflect very satisfactory numerical accuracy; $\varepsilon_{\zeta}$ is always under 1%, except for $\tilde{\alpha}_1$ and $\tilde{\beta}_1$ (with NI), coefficients that are notoriously difficult to determine [see van Dijk and Kloek (1980)]. In addi-

$^8$The bulk of this increase in computation time comes from obtaining a 1-1 poly-$t$ distribution with different hyperparameters (depending on $\zeta_{(n-1)}^1$) for each drawing, which requires tabulating the distribution function anew at each drawing (a direct method is used as $I_n^1 = 1$). Execution time may decrease substantially by doing the Monte Carlo over all five elements in $\zeta^1$, but at the expense of some numerical accuracy. Nevertheless, remark that Bauwens (1984) suggests the use of a poly-$t$ based importance function (PTST $- 2$) with NI and reports CPU-times per drawing on a DG-MV 8000 mini-computer that are roughly twice as high as what we require on a PC 386!
Fig. 1. Marginal posterior densities for Klein’s model 1; diffuse prior: ---, truncated prior: ·······.
tion, the integrating constants are very precisely calculated (\(\epsilon_{\text{int}}\) is quite small) and the procedure is not at all upset by adverse tail behaviour,\(^9\) as is evidenced by table 4.

\(^9\)This refers to directions in which the far tail of the posterior density is located above the importance function. This situation may lead to tail ratios of posterior density over importance function values [i.e., the weights as defined in (3.13) or (3.14) here] that are many orders of magnitude larger than the ratios in regions of higher posterior density, attributing an overwhelming weight to a small number of drawings (sometimes only one) in the far tails. Of course, if the number of drawings goes to infinity, even such low probability drawings will cancel out, but in practice the size of the Monte Carlo runs is finite and serious biasing of the results may occur.
A final indication of the accuracy of the results is the fact that we have found very little variation across replications of similar runs and that with NI posterior moments hardly change from 1000 drawings on for means (defined here as having the same first two decimal places) and from 10,000 drawings on for standard deviations, whereas skewness coefficients require about 50,000 drawings to stabilize to two decimal places. If we truncate, such stability is already achieved at, respectively, 500, 1000, and 20,000 drawings.

Fig. 1 gives the marginal posterior densities for $\zeta^1$ under both prior specifications. Performing 100,000 drawings was largely sufficient to obtain very smooth plots.

Overall, results on posterior moments are comparable to those found in Bauwens (1984). However, even when using a complicated poly-$t$ based importance function $PTST - 2$, he finds a variation coefficient of 2.91 for NI (compare our table 4).

Van Dijk and Kloek (1980) do not report results without truncation, but their prior $2$ is close to our TR, with the exception that $\mu_0 = 4$ in their case [Zellner's (1971) suggestion] and that they truncate all the coefficients in $\zeta$ to the unit interval (they reverse the sign of $\beta_3$). Accordingly, their results for prior $2$ roughly correspond to what we find with TR.

5. Concluding remarks

The aim of the present paper was to increase the reliability and, at the same time, reduce the numerical complexity of the multivariate integrations required for Bayesian inference outside the restrictive natural conjugate framework. This seems an essential requirement if we wish to extend the domain of application of Bayesian methods to models of some empirical relevance, like the popular simultaneous equation model discussed here.

Provided our prior assumptions are formulated within a Normal and Inverted-Wishart framework, we can treat some of the parameters analytically and we typically have a choice as to which parameters we integrate out first in this analytical fashion.

(i) Given prior independence between $\zeta$ and $\Sigma$ and an Inverted-Wishart prior structure on $\Sigma$, we know its posterior density will be of the same functional form and we can treat it analytically. In the RENC prior density assumed here, however, $\zeta^2$ does depend on $\Sigma$, leading to an Inverted-Wishart result for $\Sigma$ only under matrix-form restrictions on $A^2$ (i.e., the natural conjugate case). Nevertheless, an RLI structure on $A^2$ proves to be sufficient for analytical integration of $\Sigma$, exploiting its recursive decomposition and treating the system equation by equation.
The specific RENC prior density for $\zeta^2$ in (2.33) then allows for its analytical treatment, using the recursive techniques explained in Steel (1988), given that $\zeta^2$ does not intervene in the Jacobian of the model. Of course, we shall try to handle as many regression coefficients as possible by this analytical method.

The next step exploits the poly-$t$ structure of the remaining parameters in one equation, given those in the rest of the model.

Finally, the latter are treated by Monte Carlo procedures using importance sampling.

(ii) An alternative integration sequence proceeds from first using the Normal prior structure on $\zeta^2$ to integrate it out analytically, given all other parameters, and then attempt to handle $\zeta^1$ in an analytical way, given $\Sigma$. The snag of this strategy is that the latter integration can, at least presently, only be performed for triangular or SURE systems, seriously restraining its applicability.

An advantage, on the other hand, might be that the resulting Monte Carlo analysis of $\Sigma$ could prove rather reliable and efficient, as suggested by Richard and Steel (1988).

Note that, in this approach, the RLI assumption becomes irrelevant as recursivity is not exploited here. Only in the limit case of matrix-form restrictions (natural conjugate\(^\text{10}\)) will simplifications occur (to the well-known analytically tractable natural conjugate densities).

Naturally, the proof of the pudding is in the eating, so we have applied the techniques in (i) to Klein’s (1950) model I. Of its nine free coefficients, four can be integrated out analytically and one can be treated with the help of poly-$t$ algorithms, which leaves only four dimensions for the Monte Carlo integration. This considerable dimensional gain renders our hybrid method much more efficient than performing Monte Carlo integration on the space of all nine structural coefficients. The dimension argument will, typically, also hold for the method in (ii) [see Steel (1987)], but additional work is required to apply this approach to SEM systems.

An additional reason for preferring a partly analytical approach to a fully numerical one is, of course, the reliability of the inference, which, in a numerical Monte Carlo approach can always be subject to unexpected changes due to undominated tails of the posterior density or an insufficient number of drawings. The latter argument is valid irrespective of computational considerations and may, in smaller models, even be the main attraction of the hybrid approach proposed here.

\(^{10}\)In order to find the exact natural conjugate analytical results we also need to impose a prior dependence of $\zeta^1$ on $\Sigma$, if we wish to be informative on $\zeta^1$. 


Appendix

Table 5

<table>
<thead>
<tr>
<th>Model</th>
<th>Prior</th>
<th>Posterior</th>
<th>Assumption</th>
</tr>
</thead>
<tbody>
<tr>
<td>$B$ [2.1]</td>
<td></td>
<td></td>
<td>Nonsingular</td>
</tr>
<tr>
<td>$W_i^2$ [2.1]</td>
<td></td>
<td></td>
<td>Full column rank</td>
</tr>
<tr>
<td>$W_i^2$ [2.1; 3.2]</td>
<td></td>
<td></td>
<td>Full column rank</td>
</tr>
<tr>
<td>$R_i, R_i^2$ [2.1]</td>
<td></td>
<td></td>
<td>Full column rank</td>
</tr>
<tr>
<td>($\Rightarrow R_i^2$ [2.1])</td>
<td></td>
<td></td>
<td>Full column rank</td>
</tr>
<tr>
<td>($\Rightarrow F_i, F_i^2$ [2.1; 3.2])</td>
<td></td>
<td></td>
<td>Full column rank</td>
</tr>
<tr>
<td>$\Sigma_j$ [2.1; 2.2]</td>
<td>$\Sigma_j^b$ [2.3]</td>
<td>$G_i^* [3.4]$</td>
<td>PDS</td>
</tr>
<tr>
<td>$N_{1i}^b$ [2.3]</td>
<td></td>
<td></td>
<td>PDS</td>
</tr>
<tr>
<td>($\Rightarrow F_i^2$ [2.3])</td>
<td>$H_i^* [3.1]$</td>
<td></td>
<td>PDS</td>
</tr>
<tr>
<td>$N_0$ [2.3; 3.2]</td>
<td>$N_{1i}^*$ [3.2]</td>
<td></td>
<td>PDS</td>
</tr>
</tbody>
</table>

*Numbers in square brackets denote the relevant subsection(s). Implications of assumptions are given in parentheses. Indices $j$ can be either 1 or 2 and $i: 1 \rightarrow n$.

References

Bauwens, L., 1984, Bayesian full information analysis of simultaneous equation models using integration by Monte Carlo (Springer Verlag, Berlin).


Drèze, J.H., 1962, The Bayesian approach to simultaneous equations estimation, ONR research memorandum no. 67 (Technological Institute, Northwestern University, Evanston, IL).


Rothenberg, T., 1963, A Bayesian analysis of simultaneous equation systems, Econometric Institute report no. 6315 (Erasmus Universiteit, Rotterdam).


Steel, M.F.J., 1988, Seemingly unrelated regression equation systems under diffuse stochastic prior information: A recursive analytical approach, CentER discussion paper 8805 (Katholieke Universiteit Brabant, Tilburg).


van Dijk, H.K., 1984, Posterior analysis of econometric models using Monte Carlo integration (Erasmus Universiteit, Rotterdam).


Reprint Series, CentER, Tilburg University, The Netherlands:


No. 5 Th. ten Raa and F. van der Ploeg, A statistical approach to the problem of negatives in input-output analysis, Economic Modelling, Vol. 6, No. 1, 1989, pp. 2 - 19.


No. 8 Th. van de Klundert and F. van der Ploeg, Wage rigidity and capital mobility in an optimizing model of a small open economy, De Economist 137, nr. 1, 1989, pp. 47 - 75.


