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Monetary policy analysis in a small open economy using Bayesian cointegrated structural VARs

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**MONETARY POLICY ANALYSIS IN A
SMALL OPEN ECONOMY USING
BAYESIAN COINTEGRATED
STRUCTURAL VARS**

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DECEMBER 2003

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ABSTRACT. Structural VARs have been extensively used in empirical macroeconomics during the last two decades, particularly in analyses of monetary policy. Existing Bayesian procedures for structural VARs are at best confined to a severely limited handling of cointegration restrictions. This paper extends the Bayesian analysis of structural VARs to cover cointegrated processes with an arbitrary number of cointegrating relations and general linear restrictions on the cointegration space. A reference prior distribution with an optional small open economy effect is proposed and a Gibbs sampler is derived for a straightforward evaluation of the posterior distribution. The methods are used to analyze the effects of monetary policy in Sweden.

KEYWORDS: Structural, Vector autoregression, Monetary policy, Impulse responses, Counterfactual experiments.

JEL CLASSIFICATION: E44, C32, E52.

NON-TECHNICAL SUMMARY

Over the last 20 years structural vector autoregressions have been used extensively in empirical macroeconomics and have become one of the most often used tools for analysing the effects of monetary policy. Since the early studies, where structural relations were identified in recursive systems and the parameters were estimated with least squares, the methodology has developed in several directions. For example, techniques for estimating non-recursive systems have been proposed, Bayesian methods have emerged, which have also proven to be useful in forecasting exercises, and the link between VARs and cointegration has been established.

This paper integrates non-recursive structural systems and cointegration with the Bayesian approach to structural VARs. The framework we propose allows for more general forms of cointegration than has previously been suggested within the Bayesian approach to structural VARs; see, e.g., Sims and Zha (1998). In particular, we allow for an arbitrary number of cointegration relations and for general linear restrictions on the cointegration vectors. Our framework also extends some of Villani's (2000, 2003) work on reduced-form cointegrated VARs. Furthermore, we propose extensions to the Minnesota prior both in the direction of dealing with potential exogeneity of some variables and in the treatment of seasonality.

The methodology is then applied to quarterly data for the Swedish economy over the 1975-2001 period. The variables we consider are similar to those suggested by Kim and Roubini (2000) in their study of the G7 countries and our identifying restrictions are also nearly identical to those suggested by them. Within a 7 variable system of domestic output, prices, money (measured by M0), a short-term interest rate, the exchange rate and the foreign short-term interest rate and oil prices we find evidence of two cointegration relations. One stationary relation is the spread between the domestic and foreign short-term rates and the other involves the 5 domestic variables only. The identifying restrictions on the simultaneous relations imply, among other things, that domestic shocks have zero contemporaneous effect on the foreign variables and that domestic money, the short-term interest rate and the exchange rate are simultaneously determined, while output and prices can only react to these domestic variables with a lag.

For the over-identified system we find that monetary policy shocks generally have the expected effects on the domestic variables. Specifically, the price and exchange rate puzzles (rising prices and impact depreciation of the exchange rate following a monetary contraction) reported by Jacobson, Jansson, Vredin and Warne (2002) are no longer present. Furthermore, the monetary policy shock is mainly important for short run fluctuations in the domestic interest rate, but has only minor effects at business cycle frequencies.

We also examine a few counterfactual policy experiments using conditional forecasts, as suggested by Leeper and Zha (2003). Focusing on the conditioning period beginning in the last quarter of 1999, when the Riksbank first raised the repo rate by 35 basis points, and ending in the first quarter of 2000, when a further 50 basis point increase occurred, we consider 3 alternative scenarios. First, we study the case when the short-term rate follows the observed rate over these 2 quarters. Second, we assume that the short-term rate is constant and equal to its 1999 quarter 3 value. Finally, we examine the possibility that the short-term rate is lowered (relative to quarter 3 of 1999) by 1 percent over these two quarters. Based on these 3 scenarios we compute various statistics of interest (e.g., the probability of inflation becoming higher than 3 percent in the year 2000 and/or GDP growth being less than 1 percent in the same year) from the posterior distributions for conditional forecasts, beginning in the last quarter of 1999.

To evaluate the importance of the Lucas critique for counterfactual experiments of this type we apply the modesty metric suggested by Leeper and Zha (2003). For the first two scenarios we find that both are consistent with being viewed as modest policy interventions, while the third experiment is likely to be subject to the Lucas critique.

1. INTRODUCTION

Following the seminal paper by Sims (1980), structural vector autoregressions (SVAR) have become one of the most frequently used methods for empirical analysis of, e.g., the effects of monetary policy on output and prices. In these early days, SVARs were identified by assuming a recursive structure for the variables in the system and the parameters were estimated using least squares. The main tools for analysing the effects of structural shocks were impulse response functions and forecast error variance decompositions.

During the last two decades we have seen a number of important developments to the basic methodology. First, non-recursive identification procedures have been suggested (e.g., Bernanke (1986) and Sims (1986)), allowing variables to be simultaneously determined. When the vector of variables includes so called *fast* variables, such as interest rates and exchange rates, this is often a desirable feature. Second, Bayesian methods, often based on the so called Minnesota prior (e.g., Litterman, 1986), have emerged and proved useful mainly in forecasting exercises but also in policy analysis. Third, the link between VARs and cointegration in error correction models has been established (e.g., Engle and Granger, 1987) and is nowadays widely used to study long-run relations in data driven by unit-root (random walk) processes.

In contrast to the classical approach, with sampling distributions depending on whether the data generating process is stationary or not, the order of integration of the data has no bearing on Bayesian inference of time series data. Cointegration restrictions may nevertheless be useful for imposing long run structure on the otherwise loosely structured VARs, and they are especially important from a forecasting point of view as the imposed long run relations can be shown to be satisfied asymptotically at long forecast horizons (Engle and Yoo, 1987; Christoffersen and Diebold, 1998).

Sims and Zha (1998) propose a Bayesian approach to structural VAR modelling where the prior is approximated via the use of dummy observations similar to the mixed estimation approach of Theil and Goldberger (1961). Two of the dummy variables produce a shrinkage effect toward either none or a single cointegrating relation, without specifying the form of this relation.

In this paper we propose a more general framework for Bayesian analysis of cointegrated structural VARs (CSVAR) which allow for an arbitrary number of cointegrating relations and general linear over-identifying restrictions on the cointegration vectors. The approach extends some of Villani's (2000, 2003) work on reduced-form cointegrated VARs. In addition, we suggest extensions to the Minnesota prior both in the direction of dealing with potential exogeneity (as opposed to Cushman and Zha, 1997, who impose exact exogeneity) of certain variables and treatment of seasonality. Concerning the simultaneity of the model our approach is based on ideas presented in Sims and Zha (1999) and Waggoner and Zha (2003b).

As an illustration of our Bayesian CSVAR method we apply it to Swedish quarterly data covering 1975–2001. The identifying assumptions we consider are based on contemporaneous restrictions and, apart from imposing unit roots, are similar to those considered by Kim and Roubini (2000) for the non-US G7 countries. The vector of variables is also similar, but the foreign variables are TCW weighted rather than represented by the US.

We find that the identified monetary policy shock generally has the expected effects on the domestic variables. In particular, the price and exchange rate puzzles (rising prices and impact depreciation of the domestic currency following a monetary contraction) reported by Jacobson, Jansson, Vredin and Warne (2002) are no longer present. Moreover, the monetary policy shock is primarily important for short run fluctuations in the domestic interest rate (and to a lesser extent for the exchange rate), but has only negligible effects at business cycle frequencies. We also consider a few counterfactual policy experiments using conditional forecasts, as suggested

by Leeper and Zha (2003), and find that quite different interest rate paths are consistent with the idea of modest policy interventions.

The remainder of the paper is structured as follows. In the next section we briefly introduce the CSVAR model to establish notation. The prior distribution of its parameters is discussed in detail in Section 3, while the main results on conditional posterior distributions are presented in Section 4. In Section 5 we present the empirical application to the Swedish data, while Section 6 concludes. Finally, proofs of the theorems are given the Appendices.

2. COINTEGRATED STRUCTURAL VARs

The cointegrated structural VAR (CSVAR) is of the form

$$(2.1) \quad \Gamma'_0 \Delta x_t = \alpha \beta' x_{t-1} + \Gamma(L) \Delta x_t + \Phi' d_t + \varepsilon_t, \quad t = 1, \dots, T,$$

where x_t is p -dimensional vector containing the endogenous variables at time t and d_t is a q -dimensional vector of non-modelled exogenous variables (including lags), deterministic trends and seasonal components¹. Γ_0 is a $p \times p$ non-singular matrix of contemporaneous coefficients, $\Gamma(\lambda) = \sum_{i=1}^{k-1} \Gamma'_i \lambda^i$, is a $p \times p$ matrix polynomial with complex valued argument λ , and L is the usual lag operator such that $L^j x_t = x_{t-j}$. The columns of β ($p \times r$) are the cointegration vectors which defines r stationary long run relations between the endogenous variables and α ($p \times r$) is the matrix of adjustment coefficients. The structural shocks, ε_t , are assumed to be i.i.d. Gaussian with zero mean and identity covariance matrix, following the tradition in structural VAR modelling, see e.g. Leeper, Sims and Zha (1996) for a motivation.

Following Cushman and Zha (1997), the vector of observations x_t is partitioned as $x_t = (x'_{1,t}, x'_{2,t})'$, where x_2 is assumed to be exogenous with respect to x_1 . In the small open economy setting, x_2 represents the foreign variables and x_1 the domestic variables, whose effect on x_2 is assumed to be close to zero *a priori*, see Section 3.

A convenient form of the model in (2.1) for the whole sample of T time periods is obtained by transposing (2.1) for each time period and stacking the resulting row vectors in matrices, yielding

$$(2.2) \quad X \Gamma_0 = W \beta \alpha' + Q \Gamma + E,$$

where the t th row of X , W , Q and E is given by $\Delta x'_t$, x'_{t-1} , $(\Delta x'_{t-1}, \dots, \Delta x'_{t-k+1}, d'_t)$ and ε'_t , respectively, and $\Gamma = (\Gamma_1, \dots, \Gamma_{k-1}, \Phi)'$. $D = \{X, W, Q\}$ will be used as a short hand for the available data.

3. IDENTIFYING RESTRICTIONS AND PRIOR DISTRIBUTIONS

3.1. Prior distribution on $\Gamma_1, \dots, \Gamma_k$. Let $\gamma_{ij}^{(k)}$ be the (i, j) th element of Γ'_k , $i, j = 1, \dots, p$, i.e. $\gamma_{ij}^{(k)}$ is the coefficient on lag k of the j th endogenous variable in equation i . We will assume the $\gamma_{ij}^{(k)}$ to be independent (across i, j and k) with prior distribution

$$\gamma_{ij}^{(k)} \sim N[0, s^2(\gamma_{ij}^{(k)})], \quad i, j = 1, \dots, p; k = 1, 2, \dots,$$

where $s(\gamma_{ij}^{(k)})$ is the prior standard deviation of $\gamma_{ij}^{(k)}$.

The starting point for our prior is the so called Minnesota prior (Litterman, 1986; Robertson and Tallman, 1999; Sims and Zha, 1998) with

$$s(\gamma_{ij}^{(k)}) = \frac{\lambda_b}{k^{\lambda_l} \sigma_j},$$

¹Constant and trends may be restricted to the cointegration space simply by moving the relevant deterministic component from d_t to x_{t-1} as explained in Johansen (1995, Ch. 5).

where $\lambda_b > 0$ determines the overall tightness of the prior on $\Gamma_1, \dots, \Gamma_k$ around zero, $\lambda_l > 0$ controls the rate at which the coefficients are shrunk toward zero with increasing lag length and $\sigma_1, \dots, \sigma_p$ are scale factors which takes into account the possibly differing measurement units of the endogenous variables. Following Sims and Zha (1998), no scale factor is used in the numerator of $s(\gamma_{ij}^{(k)})$, as the structural shocks in ε_t , are normalized to have unit variance.

We extend the Minnesota prior in two directions. First, rather than imposing the restriction that the rest of the world (x_2) is exogenous with respect to the small economy (x_1) with probability one, as in Cushman and Zha (1997), we add an additional hyperparameter λ_e to control the tightness *around* the exogeneity restriction. Secondly, instead of using a uniform shrinkage with respect to the lag length, we use different shrinking pattern on the coefficients on seasonal and non-seasonal lags. Note that if the time series have a season of s time periods then the seasonal coefficients in the SCVAR model appear in $\Gamma_{s-1}, \Gamma_s, \Gamma_{2s-1}, \Gamma_{2s}, \dots$. The extended Minnesota prior sets

$$s(\gamma_{ij}^{(k)}) = \frac{\lambda_b \lambda_e^{I_i(j)} \lambda_s^{I_s(k)}}{f_s(k)^{\lambda_l} \sigma_j},$$

where $I_i(j)$ is an exogeneity indicator which equals 1 if variable j is assumed exogenous in equation i and zero otherwise, $0 < \lambda_e \leq 1$ determines the strength of the beliefs in the exogeneity assumption, $I_s(k)$ is a seasonality indicator which takes the value 1 for $k = s - 1, s, 2s - 1, 2s, \dots$ and zero otherwise, and

$$f_s(k) = \begin{cases} \frac{k}{s-1} & \text{if } (s-1) \mid k \\ \frac{k}{s} & \text{if } s \mid k \\ k & \text{otherwise,} \end{cases}$$

where $m \mid n$ denotes that m divides n , i.e. that $n = cm$ for some integer c . The hyperparameter λ_s ($0 < \lambda_s \leq 1$) determines the importance of coefficients on seasonal lags relative coefficients on non-seasonal lags. Note how the $f_s(k)$ function upgrades the importance of the seasonal lags compared to the non-seasonal ones.

3.2. Prior Distribution on Φ . The following prior will be used for Φ

$$\text{vec } \Phi \sim N_{pq}(0, \lambda_d^2 I_{pq}),$$

where λ_d is a shrinkage factor.

In summary, the prior on $\Gamma = (\Gamma'_1, \dots, \Gamma'_{k-1}, \Phi')'$ is

$$\text{vec } \Gamma \sim N_{p^2(k-1)+pq}(0, \Omega_\Gamma),$$

where

$$\Omega_\Gamma = P' \begin{pmatrix} \Psi & 0 \\ 0 & \lambda_d^2 I_{pq} \end{pmatrix} P$$

and Ψ is a diagonal matrix with typical element $s^2(\gamma_{ij}^{(k)})$ and P is the commutation matrix (Harville, 1997, Sec. 16.3) which maps $\text{vec}(\Gamma')$ into $\text{vec } \Gamma$.

3.3. Prior Distribution on Contemporaneous Coefficients in Γ_0 . Let $\gamma_i^{(0)}$ be the contemporaneous coefficients in the i th equation of the system, i.e. $\gamma_i^{(0)}$ is the i th column of Γ_0 . The assumed identity covariance matrix of the structural shocks fixes the scale of Γ_0 , but leaves the model invariant to rotations of Γ_0 . General linear restrictions on each column of Γ_0 may be used to settle this indeterminacy

$$(3.1) \quad \gamma_i^{(0)} = G_i \psi_i, \quad i = 1, \dots, p,$$

where G_i ($p \times s_i$) determines the restrictions and ψ_i is a s_i -vector of unrestricted coefficients.

Under the restrictions in (3.1), the model is still invariant to sign switches in the equations, as discussed in Sims and Zha (1999) and Waggoner and Zha (2003a). This may be settled by restricting one element in each ψ_i to be positive; see Section 4 for more details.

The fact that Γ_0 contains both location and scale components makes it difficult to specify a prior distribution on ψ_1, \dots, ψ_p . A way out of this problem would be to use a model where the location and scale components of Γ_0 are modelled as separate parameters. This leads to other difficulties, however, such as potentially improper posteriors (Sims and Zha, 1994). Here we settle with a uniform prior on ψ_1, \dots, ψ_p , leaving this aspect of the overall prior to future work. We note that sign restrictions on the contemporaneous coefficients advocated by Uhlig (1997) may be imposed simply by restricting the domain of Γ_0 . The posterior computations remain exactly the same apart from a rejection step in the algorithm following a Γ_0 -draw in conflict with the sign restrictions.

3.4. Prior Distribution on the Adjustment Coefficients in α . The prior distribution on α is modelled conditional on β as

$$\text{vec } \alpha | \beta \sim N_{pr}[0, (\beta' K_1 \beta)^{-1} \otimes K_2]$$

where $K_1 = \text{diag}(\sigma_1^2, \dots, \sigma_p^2)$, the σ 's are the scaling factors used in the Γ -prior, and

$$K_2 = \begin{pmatrix} \lambda_\alpha^2 I_{p_1} & 0 \\ 0 & \lambda_\alpha^2 \lambda_e^2 I_{p_2} \end{pmatrix}.$$

To understand the idea behind this prior, note that, conditional on β , α is the (partial) coefficient in the regression of $\Gamma_0' \Delta x_t$ on the predictor vector $\beta' x_{t-1}$. A shrinkage prior on a regression coefficient should take into account the scale of both the response variables and the predictors. x_{t-1} is not known when the prior is formulated, however, and the best one can do is to include the scale factors $\sigma_1, \dots, \sigma_p$. This explains the appearance of $(\beta' K_1 \beta)^{-1}$ in the prior conditional covariance matrix of α . Since the response variables are assumed to have unit variances and zero covariances, identity matrices are used in K_2 .

K_2 also contains the exogeneity parameter λ_e , which was also used in the Γ -prior, allowing the lower left block of $\alpha\beta'$ (which contains the effects of the domestic variables on the foreign variables) to be more heavily shrunk toward zero compared to the other elements of $\alpha\beta'$. The adjustment coefficients corresponding to an entirely foreign relation, i.e. a cointegration vector with zero restrictions on the coefficients on all domestic variables, should of course not be given the additional λ_e -shrinkage.

3.5. Prior Distribution on the Cointegration Vectors in β . It is well known that the cointegration vectors are only determined up to arbitrary linear combinations. Here we will use general linear identifying restrictions on each cointegration vector to identify β

$$(3.2) \quad \beta_i = h_i + H_i \phi_i, \quad i = 1, \dots, r,$$

where h_i is a vector containing the fixed part of β_i , including normalizations, H_i ($p \times d_i$) determines the restrictions and ϕ_i is a d_i -vector of free coefficients.

We suggest the following prior on ϕ_1, \dots, ϕ_r as a suitable reference prior

$$(3.3) \quad p(\phi_1, \dots, \phi_r) \propto |\beta' K_1 \beta|^{-p/2},$$

where $K_1 = \text{diag}(\sigma_1^2, \dots, \sigma_p^2)$ and the σ 's are the scaling factors used in the Γ -prior. The properties of this prior has been investigated in detail under just-identifying restrictions with $K_1 = I_p$ by Villani (2000, 2003), where it is shown that it assigns the same prior probability to each possible cointegration space of dimension r (more precisely, it implies a Haar invariant distribution over

the quotient space of all cointegration spaces). If $\sigma_i^2 \neq \sigma_j^2$ for any $i \neq j$, the prior on ϕ_1, \dots, ϕ_r no longer has this property, but the prior in (3.3) should nevertheless constitute a reasonable reference prior, especially if the time series have widely differing scales, and has the additional benefit of yielding straightforward posterior computations.

In summary, the overall prior distribution depends on the scaling factors $\sigma_1, \dots, \sigma_p$, and the following six hyperparameters

| | |
|------------------|---------------------------|
| λ_b | : Baseline shrinkage |
| λ_l | : Lag length shrinkage |
| λ_s | : Seasonality shrinkage |
| λ_α | : Adjustment shrinkage |
| λ_e | : Exogeneity shrinkage |
| λ_d | : Deterministic shrinkage |

The scaling factors $\sigma_1, \dots, \sigma_p$ are not hard to elicit given a proper understanding of the analyzed time series, but are for convenience usually estimated from data as the residual standard deviation in univariate autoregressive models fitted to the each of the series, see e.g. Litterman (1986) and Sims and Zha (1998).

4. POSTERIOR DISTRIBUTION

The joint posterior distribution of the SCVAR parameters is intractable. In this section we derive the necessary results for a numerical evaluation of the posterior distributions via the Gibbs sampler. The Gibbs sampler (e.g. Tierney, 1994) simulates from the joint posterior distribution of model parameters by iteratively generating draws from the full conditional posterior distributions, i.e. the posterior distribution of a group of model parameters conditional on all other parameters in the model. Denoting the full set of model parameters by θ , the end result of the Gibbs sampling is a sequence of draws $\theta^{(1)}, \theta^{(2)}, \dots, \theta^{(n)}$ from the posterior distribution. The draws are not independent but the following can be shown to hold under certain conditions (Tierney, 1994) which are satisfied here

$$g(\theta^{(i)}) \xrightarrow{d} p[g(\theta)|\mathcal{D}]$$

where \xrightarrow{d} denotes convergence in distribution and $g(\cdot)$ is any well-behaved real valued function. Thus, given the draws from the posterior of θ , the posterior distribution of any function of the model parameters, e.g. the impulse response functions or forecast error variance decompositions, are immediately available simply by applying the g -function to each posterior draw and subsequently using some density estimator.

The next theorem gives the full conditional posteriors of Γ , α and $\phi = (\phi'_1, \dots, \phi'_r)'$.

Theorem 1.

- Full conditional posterior of Γ

$$\text{vec } \Gamma | \Gamma_0, \alpha, \beta, \mathcal{D} \sim N_{p[p(k-1)+q]}(\bar{\mu}_\Gamma, \bar{\Omega}_\Gamma),$$

where $\bar{\Omega}_\Gamma^{-1} = (I_p \otimes Q'Q) + \Omega_\Gamma^{-1}$ and $\bar{\mu}_\Gamma = \bar{\Omega}_\Gamma \text{vec}[Q'(X\Gamma_0 - W\beta\alpha)']$.

- Full conditional posterior of α

$$\text{vec } \alpha | \Gamma_0, \Gamma, \beta, \mathcal{D} \sim N_{pr}(\bar{\mu}_\alpha, \bar{\Omega}_\alpha),$$

where $\bar{\Omega}_\alpha^{-1} = (\beta'W'W\beta \otimes I_p) + (\beta'K_1\beta \otimes K_2^{-1})$ and $\bar{\mu}_\alpha = \bar{\Omega}_\alpha \text{vec}[(X\Gamma_0 - Q\Gamma)'W\beta]$.

If $\lambda_e = 1$, then $\bar{\Omega}_\alpha^{-1}$ simplifies to $\lambda_\alpha^{-2}[\beta'(W'W + K_1)\beta] \otimes I_p$.

- Full conditional posterior of $\phi = (\phi'_1, \dots, \phi'_r)'$

$$\phi|\Gamma_0, \Gamma, \alpha, \mathcal{D} \sim N_d(\bar{\mu}_\phi, \bar{\Omega}_\phi),$$

where $d = \sum_{i=1}^r d_i$, $\bar{\Omega}_\phi^{-1} = H'AH$, $A = \alpha'\alpha \otimes W'W + \alpha'K_2^{-1}\alpha \otimes K_1$,

$\bar{\mu}_\phi = \bar{\Omega}_\phi H' \{\text{vec}[W'(X\Gamma_0 - Q\Gamma)\alpha] - Ah\}$, $H = \text{diag}(H_1, \dots, H_r)$ and $h = (h'_1, \dots, h'_r)'$. If $\lambda_e = 1$, then A simplifies to $\alpha'\alpha \otimes (W'W + \lambda_\alpha^{-2}K_1)$.

Proof. See the appendix. \square

The full conditional posterior distributions of the simultaneous effects, ψ_i , $i = 1, \dots, p$, do not belong to a known family of distributions and direct sampling is therefore not feasible. A solution to this problem for the structural VAR model without cointegration restrictions is given in Waggoner and Zha (2003b, Theorem 1). The Waggoner and Zha (2003b) solution cannot be applied directly when cointegration restrictions are allowed for, but a straightforward modification of their result is as follows. First we need the following definition.

Definition 1. A random variable X follows the *absolute normal distribution* $AN(\mu, \rho)$ if it has density function

$$f_{AN}(x; \mu, \rho) = c |x|^{\frac{1}{\rho}} \exp \left[-\frac{1}{2\rho}(x - \mu)^2 \right], \quad x \in \mathbb{R},$$

where c is a normalizing constant, $\rho \in \mathbb{R}^+$ and $\mu \in \mathbb{R}$.

In the next theorem, let B_{-i} equal the matrix B with the i th column deleted, B_\perp is the orthogonal complement of B , $\text{Chol}(B)$ is the Choleski root of B such that $B = \text{Chol}(B) \text{Chol}(B)'$, $\|\cdot\|$ is the usual Euclidean norm and $\stackrel{d}{=}$ denotes equality in distribution. With these preliminaries at hand we can prove the following extension of Theorem 1 in Waggoner and Zha (2003b).

Theorem 2. The full conditional posterior of the coefficients in the i th simultaneous relation is

$$\psi_i|\Gamma_{0,-i}, \alpha, \beta, \Gamma, \mathcal{D} \stackrel{d}{=} R_i \sum_{j=1}^{s_i} \xi_j v_j,$$

where $R_i = \text{Chol}[T(G'_i X' X G_i)^{-1}]$, $\xi_1 \sim AN(\hat{\xi}_1, T^{-1})$, $\xi_j \sim N(\hat{\xi}_j, T^{-1})$, for $j = 2, \dots, s_i$, $\hat{\xi}_j = \mu'_{\psi_i} R_i^{-1} v_j$, $v_1 = R'_i G'_i \Gamma_{0-i\perp} / \|R'_i G'_i \Gamma_{0-i\perp}\|$, $(v_2, \dots, v_{s_i}) = v_{1\perp}$ and $\mu_{\psi_i} = (G'_i X' X G_i)^{-1} G'_i X' z_i$, where z_i is the i th column of $W\beta\alpha' + Q\Gamma$.

Proof. See the appendix. \square

The sign restriction mentioned in Section 2 is easily imposed by replacing a generated ψ_i which does not satisfy the sign restriction with the ψ_i generated in the previous iteration of the Gibbs sampler. Thus, if the sign restriction is not satisfied in a given iteration, ψ_i is not updated at this iteration.² A good choice of normalizing variables (see Waggoner and Zha, 2003a, for an automatic method) usually gives a small number of the rejected draws. In general, an equation should be normalized on a variable whose coefficient is likely to have negligible posterior probability in a neighborhood of zero.

²This updating step of the algorithm then becomes a Metropolis-Hastings step, rather than a pure Gibbs step (Gilks, Richardson and Spiegelhalter, 1996).

In order to use Theorem 2 in a Gibbs sampler we need to be able to efficiently generate variates from the absolute normal distribution. A simple, very accurate and readily sampled approximation can be derived from the fact that the $AN(\mu, \rho)$ -distribution is bimodal with modes at $\frac{\mu}{2} \pm \frac{1}{2}\sqrt{(\mu^2 + 4)}$. Furthermore, the curvature at $x = x_0$ is

$$-\left[\frac{d^2}{dx^2} \ln f(x)\right]^{-1}\bigg|_{x=x_0} = \frac{x_0^2}{(1 + x_0^2)}\rho.$$

These two facts can be used to build the following mixture normal approximation to the $AN(\mu, \rho)$ distribution

$$f_{AN}(x; \mu, \rho) \approx vN(x; \mu_1, \sigma_1^2) + (1 - v)N(x; \mu_2, \sigma_2^2),$$

where $N(x; \cdot, \cdot)$ is used as a shorthand for the density of a normal distribution, $\mu_1 = \frac{\mu}{2} - \frac{1}{2}\sqrt{(\mu^2 + 4)}$, $\mu_2 = \frac{\mu}{2} + \frac{1}{2}\sqrt{(\mu^2 + 4)}$, $\sigma_i^2 = \frac{\mu_i^2}{(1 + \mu_i^2)}\rho$, $i = 1, 2$, and $v = [1 + \exp(2\mu\rho^{-1})]^{-1}$. It is easily seen that, for a given ρ , the approximation error is maximal when $\mu = 0$. The accuracy of this approximation increases inversely with ρ and even in the worst scenario when $\mu = 0$ it is already very accurate for $\rho = 0.1$. In our use of the absolute normal distribution, $\rho = T^{-1}$, where T is the length of the time series, so the approximation can, for all practical purposes, be taken as exact.

5. MONETARY POLICY IN A SMALL OPEN ECONOMY

5.1. Data and Model. Kim and Roubini (2000) studies the effects of monetary policy in the six non-US G7 countries using a structural VAR system with seven variables and non-recursive identifying restrictions on the contemporaneous coefficients. Their model avoids most of the puzzling results obtained in other studies and their conclusions seem to be rather robust across countries. Cushman and Zha (1997) use a similar scheme to identify Canadian monetary policy shocks in a larger model which also includes trade flows.

Following Kim and Roubini (2000), the data set analyzed here consists of quarterly observations from 1975:1 to 2001:4 on five Swedish variables: real GDP (y), CPI (p), M0 (m), the three month bills rate in annual terms (i), the nominal exchange rate (number of Swedish kronas needed to buy one unit of foreign currency), and two foreign variables: the oil price in US dollars (oil) and the TCW-weighted interest rate in annual terms (if). We shall also consider a model where CPI is replaced by a price index, where house mortgage interest expenditures, indirect taxes and subsidies have been excluded (UND1X). All variables except the two interest rates are in logs. The series are graphed in Figure 1.

Due to a break in the seasonal pattern in the late seventies, two different sets of seasonal dummies are used: one covering the period 1975:1 to 1979:4 and the other covering the period 1980:1 to 2001:4. To control for devaluations of the Swedish krona, we add five dummy variables at quarters 1976:4, 1977:2, 1977:3, 1981:3 and 1982:4. Two additional dummy variables are included in the analysis to account for the abandonment of the fixed exchange rate in 1992:4 followed by the introduction of an explicit inflation target by Sveriges Riksbank in 1993:1. The latter dummy takes the value zero during 1975:1-1992:4 and is equal to one from 1993:1 and onwards.

We use $k = 4$ lags in the SCVAR model. Standard information criteria suggested that this is probably excessive (SBC (Schwarz, 1978) and Hannan and Quinn's HQ criteria (Quinn, 1980) both suggested $k = 1$ whereas AIC (Akaike, 1974) preferred $k = 4$), but we found the seasonal component to be better modelled with four lags. Because of the prior on Γ , which shrinks longer lags more heavily toward zero, increasing the lag length is not that costly in terms of lost precision.

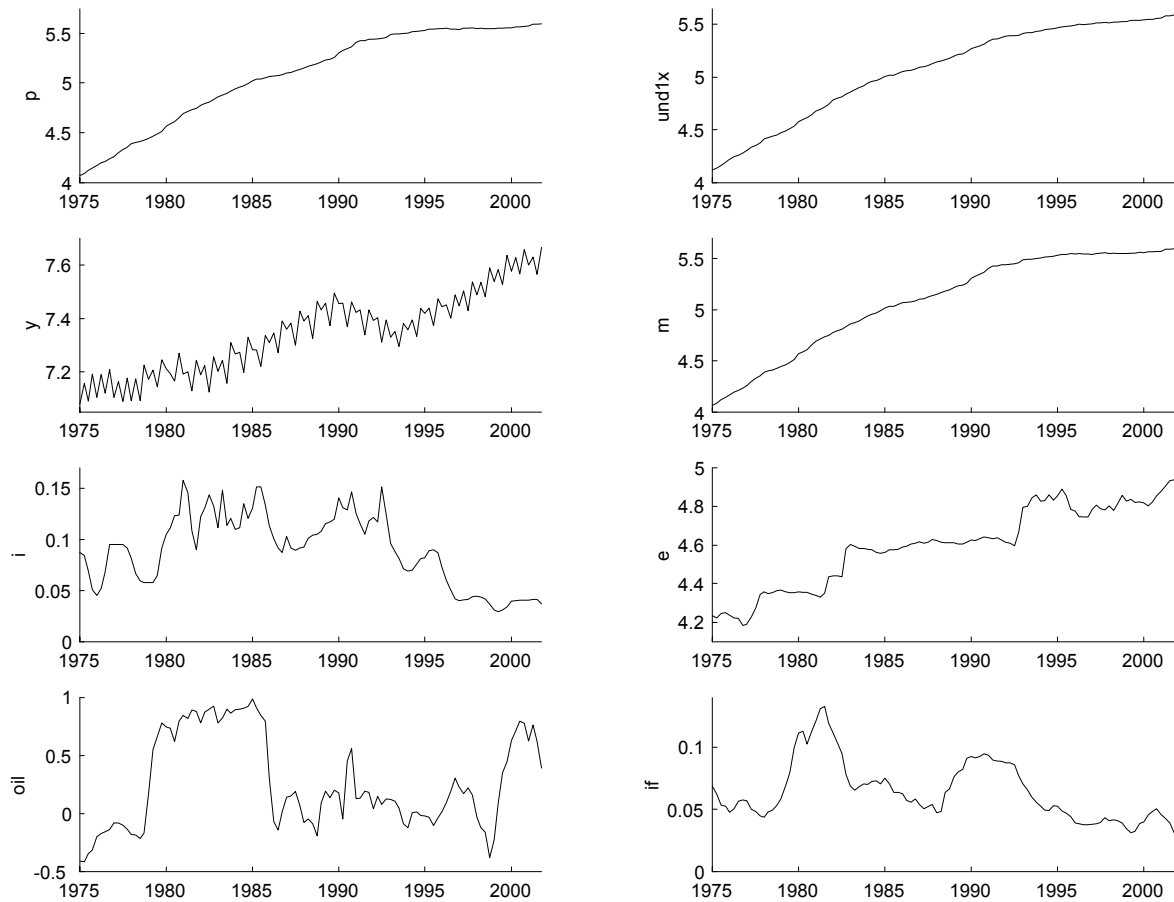


FIGURE 1. Swedish macro data 1975:1-2001:4.

To determine a value for the cointegration rank, the number of long run relations, we may use the Schwarz (1978) approximation of the posterior distribution of r : $\Pr(r = 0|\mathcal{D}) = 0.116$, $\Pr(r = 1|\mathcal{D}) = 0.863$, $\Pr(r = 2|\mathcal{D}) = 0.02$, and essentially zero for other r . The rank estimation procedure based on the trace test (Johansen, 1995) suggested that $r = 2$ on 1% significance level and $r = 3$ on the 5%-level using the asymptotic critical values. The corresponding test based on bootstrapped critical values suggested that $r = 1$ on 1% significance level and $r = 2$ on the 5%-level. The HQ criterion chose $r = 2$ whereas AIC gave $r = 5$ as optimal rank. In light of these results and the well known tendency of the Schwarz approximation to favor too small models, the ensuing analysis conditions on $r = 2$. Note, however, that all results conditional on a specific pair of lag length and cointegration rank (e.g. the impulse responses or forecasts) may be averaged over k and r with the posterior distribution as weight function (Draper, 1995).

One immediate candidate for a long run relation is the equality of the domestic and foreign interest rate, and we will take this as one of the cointegration vectors. The second cointegration vector is assumed to be a domestic relation normalized on m , with no other restrictions on the coefficients of the domestic variables. The specification of the two cointegration vectors is accomplished by six over-identifying restrictions. The Schwarz approximation of the posterior odds ratio of the restricted model to the unrestricted model yields the odds ratio 7.85, thus favoring the restricted cointegration space. The likelihood ratio test statistic is 19.27 which is

highly significant with respect to its asymptotic χ^2_6 -distribution, but the bootstrapped p -value of 0.3685 (Gredenhoff and Jacobson, 2001) confirms the finding of the Schwarz approximation.

5.2. Identification of the Structural Shocks. Successful application of structural VARs hinges on a proper identification of the structural shocks. Cushman and Zha (1997) and Kim and Roubini (2000) claim that non-recursive identifying restrictions are crucial for solving the price and exchange rate puzzles mentioned in the introduction. Their argument is that traditional recursive (Choleski) identification imposes restrictions which are incredible for a small open economy in that it requires either the exchange rate to be excluded from the monetary policy equation or the interest rate to be excluded from the exchange rate equation.

When applying the identifying scheme of Kim and Roubini (2000) to our data set, the interest and exchange rate equations turned out to be virtually inseparable, manifested in very uncertain impulse response functions. We therefore base our analysis on a different set of identifying restrictions, which still has many of the features of Kim and Roubini's specification, including the above mentioned simultaneity between the exchange and interest rate. The model used here is

$$(5.1) \quad \Gamma'_0 \Delta x_t = \begin{pmatrix} \gamma_{11} & 0 & 0 & 0 & 0 & \gamma_{61} & 0 \\ \gamma_{12} & \gamma_{22} & 0 & 0 & 0 & \gamma_{62} & 0 \\ \gamma_{13} & \gamma_{23} & \gamma_{33} & \gamma_{43} & 0 & 0 & \gamma_{73} \\ 0 & \gamma_{24} & \gamma_{34} & \gamma_{44} & \gamma_{54} & \gamma_{64} & \gamma_{74} \\ 0 & \gamma_{25} & 0 & \gamma_{45} & \gamma_{55} & \gamma_{65} & \gamma_{75} \\ 0 & 0 & 0 & 0 & 0 & \gamma_{66} & 0 \\ 0 & 0 & 0 & 0 & 0 & \gamma_{67} & \gamma_{77} \end{pmatrix} \begin{pmatrix} \Delta y_t \\ \Delta p_t \\ \Delta m_t \\ \Delta i_t \\ \Delta e_t \\ \Delta oil_t \\ \Delta if_t \end{pmatrix} = \Upsilon_t + \begin{pmatrix} \varepsilon_{y,t} \\ \varepsilon_{p,t} \\ \varepsilon_{md,t} \\ \varepsilon_{mp,t} \\ \varepsilon_{e,t} \\ \varepsilon_{oil,t} \\ \varepsilon_{if,t} \end{pmatrix},$$

where $\Upsilon_t = \alpha\beta'x_{t-1} + \Gamma(L)\Delta x_t + \Phi'd_t$.

The differences between our identifying assumptions and those used in Kim and Roubini (2000) are:

- The foreign interest rate appears in both the money demand and monetary policy equation (third and fourth equation, respectively).
- The CPI appears in the monetary policy equation. Since we use quarterly data and not monthly as in Kim and Roubini, the argument that the current periods price level is unobserved by the monetary authority at the time of the policy decision is highly questionable.
- Output is excluded from the exchange rate equation, motivated by the substantial time lags in the publication of swedish GDP followed by significant revisions.
- Money does not appear in exchange rate equation. This restriction is imposed to reduce the high degree of simultaneity between money, the interest and exchange rate.

Our identifying scheme imposes four over-identifying restrictions on Γ_0 . The posterior odds ratio comparing the model with these four over-identifying restrictions and the model with just-identifying restrictions is 0.359, indicating some preference for the just-identified model. Since all reasonable just-identifying schemes turns out to be unable to separate the money demand, monetary policy and exchange rate shocks, we nevertheless opt for the over-identified model.

Table 1 displays the presence or absence of contemporaneous effects of the structural shocks. Note that all shocks are allowed to influence money, the interest rate and the exchange rate contemporaneously.

5.3. Results. We use the following prior hyperparameters: $\lambda_b = 0.3$, $\lambda_l = 1$, $\lambda_s = 0.5$, $\lambda_\alpha = 1$, $\lambda_e = 0.1$, $\lambda_d = 10$. The scale factors $\sigma_1, \dots, \sigma_p$ are estimated from data as explained in Section 3. The results are insensitive to modest changes in the prior hyperparameters.

| | | Shock | | | | | | |
|----------|------------|----------|----------|-----------|-----------|----------|------------|-----------|
| | | <i>y</i> | <i>p</i> | <i>md</i> | <i>mp</i> | <i>e</i> | <i>oil</i> | <i>if</i> |
| Variable | <i>y</i> | * | 0 | 0 | 0 | 0 | * | 0 |
| | <i>p</i> | * | * | 0 | 0 | 0 | * | 0 |
| | <i>m</i> | * | * | * | * | * | * | * |
| | <i>i</i> | * | * | * | * | * | * | * |
| | <i>e</i> | * | * | * | * | * | * | * |
| | <i>oil</i> | 0 | 0 | 0 | 0 | 0 | * | 0 |
| | <i>if</i> | 0 | 0 | 0 | 0 | 0 | * | * |

TABLE 1. Contemporaneous effects of shocks. Stars indicate non-zero coefficients.

| Variable | | | | | | |
|-------------------------|--------------------------|-----------------------|-------------------------|-------------------------|-----------------------|---------------------------|
| <i>y</i> | <i>p</i> | <i>m</i> | <i>i</i> | <i>e</i> | <i>oil</i> | <i>if</i> |
| 75.74 (66.6, 84.6) | — | — | — | — | 0.11 (-1.3, 1.6) | — |
| -5.97 (-18.7, 6.2) | 149.18 (132.5, 166.7) | — | — | — | -0.81 (-2.3, 0.7) | — |
| -21.80 (-35.5, -8.2) | -18.81 (-45.1, 7.7) | 76.76 (42.5, 98.7) | -32.49 (-86.9, 17.9) | — | — | -46.15 (-98.8, 8.1) |
| — | -7.30 (-37.7, 23.1) | 51.65 (15.5, 86.3) | 90.03 (58.0, 114.7) | -21.42 (-41.5, -4.9) | -1.56 (-3.3, 0.2) | -67.26 (-116.8, -16.7) |
| — | -26.97 (-52.8, -2.3) | — | 52.28 (20.5, 90.4) | 51.02 (38.7, 59.8) | 0.70 (-0.7, 2.2) | 44.67 (-0.1, 90.9) |
| — | — | — | — | — | 7.62 (6.8, 8.5) | — |
| — | — | — | — | — | -2.25 (-3.5, -1.0) | 224.40 (199.1, 250.9) |

TABLE 2. Posterior means and 90% probability intervals (in parantheses) for the unrestricted elements in Γ_0 .

After 2500 burn-in draws, an additional 50.000 draws were simulated from the posterior distribution of the model parameters with the Gibbs sampler in Theorem 1 and 2 and every tenth draw was then saved for the inference. This subsampling reduces the autocorrelation of the draws and keeps storage demands manageable.

The first cointegration relation is, as mentioned above, the spread between the Swedish and the foreign interest rate. The second cointegrating relations is not as easily identified from data, although it can be concluded from tests of restrictions that it is probably a domestic relation. The posterior median estimate of the second cointegration relation, where the coefficients on the oil price and the foreign interest rate have been set to zero, is given by

$$m = \begin{pmatrix} 0.44 \\ (-0.08, 0.79) \end{pmatrix} y + \begin{pmatrix} +0.01 \\ (-0.55, 0.34) \end{pmatrix} p - \begin{pmatrix} 0.18 \\ (-2.13, 1.36) \end{pmatrix} i + \begin{pmatrix} +1.84 \\ (1.04, 3.40) \end{pmatrix} e,$$

with symmetric 90% probability intervals in parantheses. Additional restrictions, e.g. that m and p have the same coefficient but with opposite signs, were not supported by data and gave rise to unreasonable estimates of the remaining unrestricted parameters.

Next we turn to the estimated simultaneous relations between the variables in Table 2. For example, the fourth relation may be interpreted as the central bank's reaction function. According to the estimated coefficients, if domestic prices increase then the Riksbank raises the interest rate. Similarly, if the nominal exchange rate weakens or the foreign interest rate increases, then the

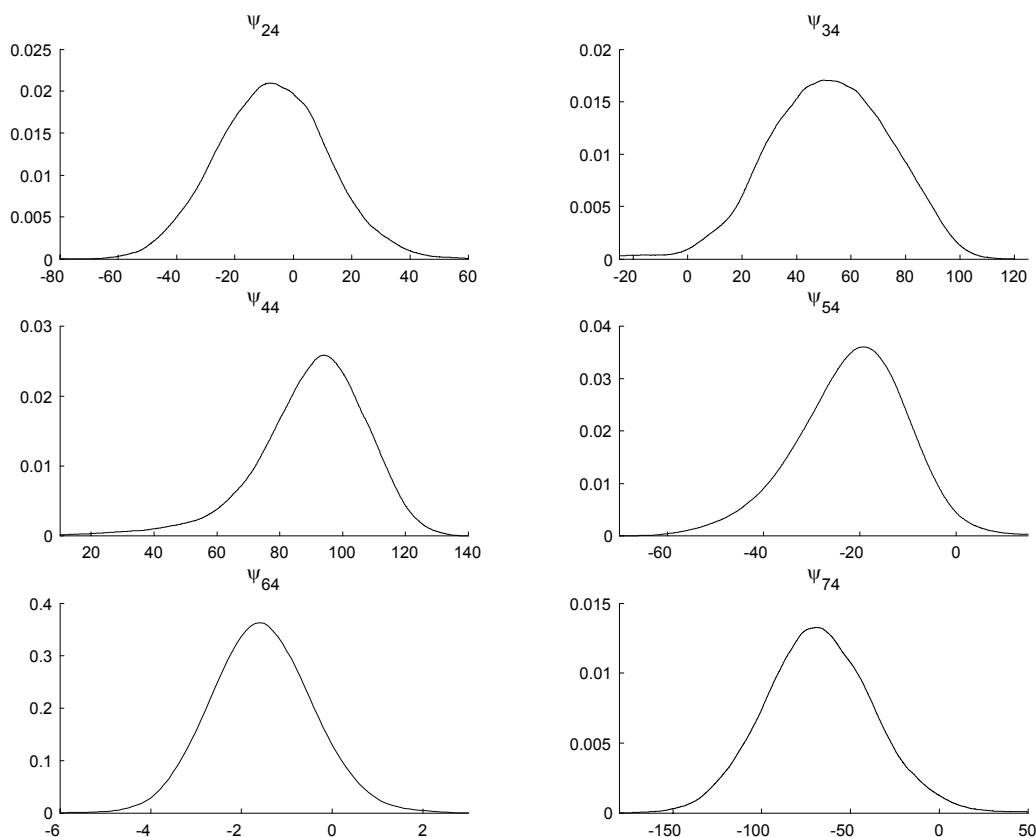


FIGURE 2. Posterior distributions of the simultaneous coefficients in the monetary policy equation: $\psi_{24}p_t + \psi_{34}m_t + \psi_{44}i_t + \psi_{54}e_t + \psi_{64}oil_t + \psi_{74}if_t$.

Riksbank also reacts by raising the domestic interest rate. Output changes, on the other hand, can only influence the domestic interest rate with a lag. The posterior distributions of the simultaneous coefficients, graphed in Figure 2, show that the money stock, domestic and foreign interest rate, the exchange rate and the oil price probably enter the reaction function whereas the coefficients on CPI may well be zero; a possible explanation for why the Riksbank does not react contemporaneously to CPI may be the inability of the Riksbank to see current periods CPI due to one month time lag in the publication in combination with the revisions following the preliminary figure.

To learn more about what this model suggest about the effects of monetary policy shocks we turn to Figure 3. The unconditional responses in the domestic variables along with 68 and 90 percent confidence intervals suggests that a positive monetary policy shock which raises the domestic interest rate during the first year affects the CPI after roughly 9 months when it begins to fall. Domestic output and the nominal exchange rate are affected more quickly, with a strengthening of the exchange rate and a weakening of output. There is a rather large posterior probability of a liquidity puzzle with an instantaneous increase of M0 from a contractionary monetary policy shock. This picture is reversed after a few months and after a approximately two years the money stock has fallen by roughly the same percentage as the CPI.

In the same figure we have also plotted the responses in the domestic variables from a monetary policy shock when the CPI measure has been replaced with the price variable used by the Riksbank for the inflation forecasts (UND1X). The main difference between this model and the former is

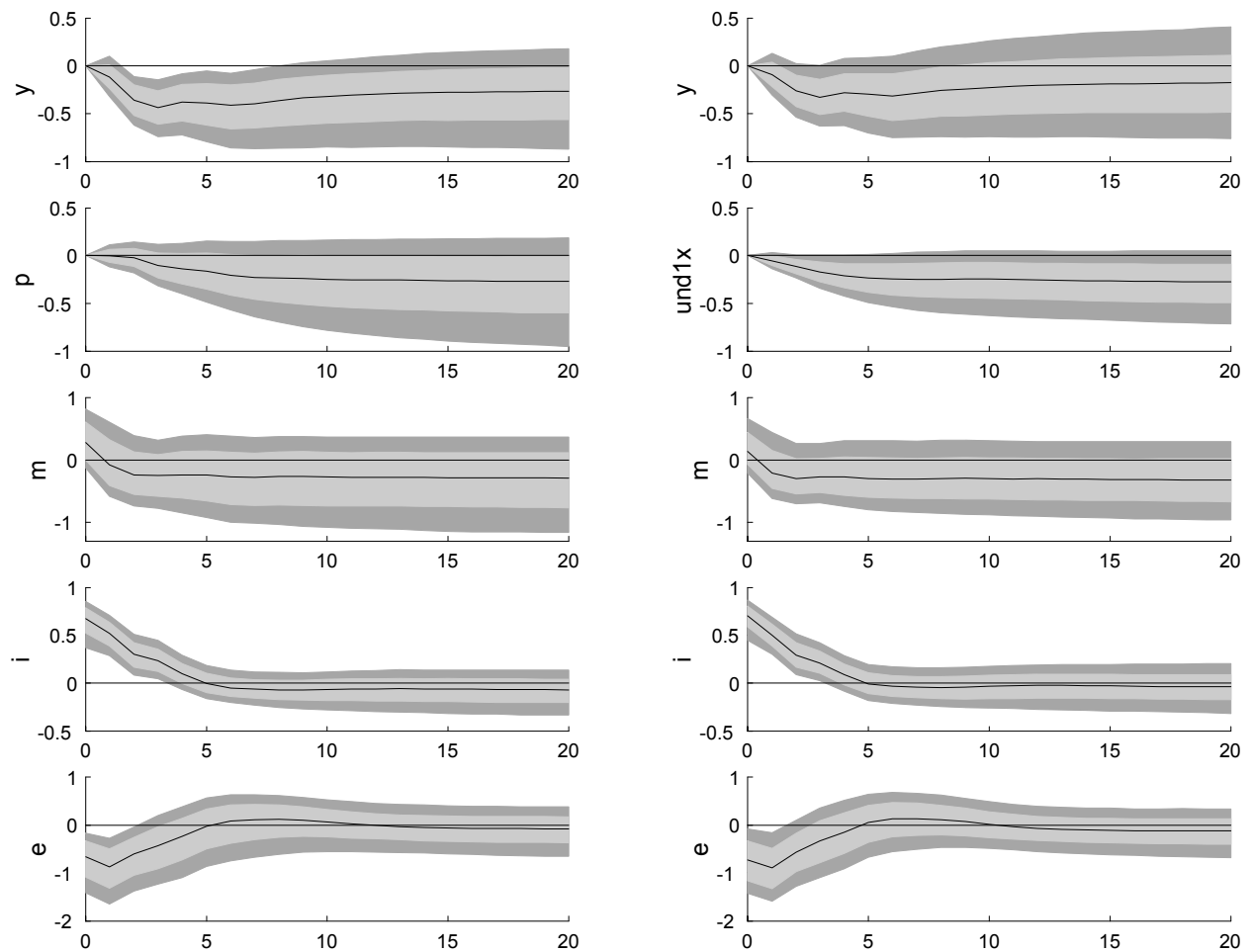


FIGURE 3. Impulse response functions over 20 quarters to a unit monetary policy shock.

that prices start to fall already after 3 months and that the probability of a liquidity puzzle is slightly smaller.

The forecast error variance decompositions for the monetary policy shock are given in Table 3. As expected, such shocks accounted for very little of the forecast uncertainty in the foreign interest rate and the oil price. Moreover, while they account for nearly half of the uncertainty in the domestic interest rate 1 quarter ahead, they seem to be important only for the nominal exchange rate. The uncertainty in the price level and output are only marginally affected by such shocks, suggesting that monetary policy surprises or mistakes are unimportant in the sense of being a sufficiently small part of the random behavior in these variables.

To further evaluate the empirical implications about monetary policy within our model, we turn to conditional forecasts. In particular, we focus on a period beginning in 1999:4 when the Riksbank first raised the repo rate from 2.90 to 3.25 on November 17, with a further increase by 50 basis points to 3.75 on February 9, 2000. Relative to the actual path of the repo rate we will consider three alternative interest rate paths generated exclusively by monetary policy shocks. That is, the interest rate path during 1999:4-2000:1 is fixed at the experiment's target by feeding a sequence of monetary policy shocks into the system. All other shocks are set to zero during these two quarters. In the subsequent quarters, the monetary policy shocks are set to zero (being

| | | <i>y</i> | <i>p</i> | <i>m</i> | <i>i</i> | <i>e</i> | <i>oil</i> | <i>i_f</i> |
|---------|----|--------------------|--------------------|--------------------|----------------------|---------------------|-------------------|----------------------|
| Horizon | 1 | 1.0 (0.0, 3.5) | 0.4 (0.0, 1.6) | 8.8 (1.4, 29.3) | 45.2 (15.1, 71.6) | 18.6 (1.6, 52.4) | 0.1 (0.0, 0.4) | 0.0 (0.0, 0.2) |
| | 4 | 8.0 (1.2, 18.3) | 1.6 (0.1, 5.6) | 9.6 (1.6, 26.5) | 30.4 (9.6, 53.5) | 15.6 (1.6, 45.8) | 0.8 (0.0, 2.8) | 0.4 (0.0, 1.5) |
| | 8 | 9.3 (1.1, 23.0) | 3.0 (0.1, 10.3) | 9.9 (1.1, 30.9) | 16.2 (4.9, 31.2) | 14.2 (2.2, 39.5) | 1.5 (0.0, 5.5) | 0.9 (0.0, 3.5) |
| | 12 | 8.4 (0.8, 22.1) | 3.5 (0.1, 12.4) | 9.8 (0.8, 32.5) | 10.8 (3.1, 20.6) | 13.0 (2.3, 35.6) | 1.8 (0.0, 6.7) | 1.2 (0.0, 4.7) |
| | 16 | 7.5 (0.7, 20.8) | 3.8 (0.1, 13.2) | 9.6 (0.7, 33.2) | 7.7 (2.2, 15.9) | 11.8 (2.1, 32.1) | 2.0 (0.0, 7.3) | 1.4 (0.0, 5.5) |

TABLE 3. Forecast error variance decomposition for the monetary policy shocks. Posterior means and 90% probability intervals (in parentheses)

controlled by the central bank) and all others shocks are drawn from their distributions. The first forecast path for the interest rate replicates its actual path in 1999:4-2000:1. The second forecast path assumes a constant interest rate, i.e., 2.90 for the two quarters. Thirdly, we study a scenario where the repo rate is lowered by 1 percent for 2 quarters. Since we do not use the repo rate in our model we impose the conditions on the closely related 3-month rate. Data up to 1999:3 are used in the estimation.

In Figure 4 we present the unconditional and the three conditional forecast paths for GDP and CPI inflation in annual growth terms, the domestic interest rate and the exchange rate. We find that the interest rate forecasts for the three conditional paths converge rather quickly and are roughly equal within a year and a half. All three conditional forecasts overshoot the realized interest rate after the first year and onwards, suggesting that the relatively low interest rate in 1999:4 (compared to the unconditional forecasts) will have to be compensated later on by an aggressive tightening of policy. The unconditional interest rate forecasts are fairly accurate over the whole forecasting horizon.

The inflation forecasts overshoot the realized path throughout the forecasting period, but are rather accurate in 2001. The tightening of policy does not exert a large downward pressure on prices. Regarding output growth, all forecasts indicates a general decline in growth, but the timing could have been better. The exchange rate forecasts suggest a weakening of the Krona, but not nearly by as much as was actually experienced in late 2000 and early 2001.

As a complement to Figure 4, we present probabilities of certain outcomes on GDP growth and CPI inflation for the three interventions and the unconditional forecast. Due to the relatively modest effects of the monetary policy shocks, there are no drastic differences between the scenarios. The probability of low GDP growth and the joint probability of *both* low growth and high inflation (3% is the upper bound of the Riksbanks inflation targeting interval) in 2000 is much larger for *Actual* scenario compared to the *Easing* scenario, however.

Counterfactual experiments of this type may, of course, be subject to the well known Lucas critique. Leeper and Zha (2002), who first introduced the concept of *modest* policy interventions, have suggested a simple metric for evaluating how unusual a conditional projection is relative to the typical size of the direct effects of policy. Their metric attempts to reflect how large the direct effects of a hypothetical intervention are and how likely the intervention is to trigger changes in agents' behavior. In Table 5 we report the Leeper-Zha modesty metric for the three policy scenarios and each domestic variable over the forecast period 1999:4–2001:4. When the metric, in absolute terms, is not greater than two, then the intervention is deemed modest and thus there is no reason for agents to change their beliefs about the policy regime and expectation formation effects may be assumed to be insignificant.

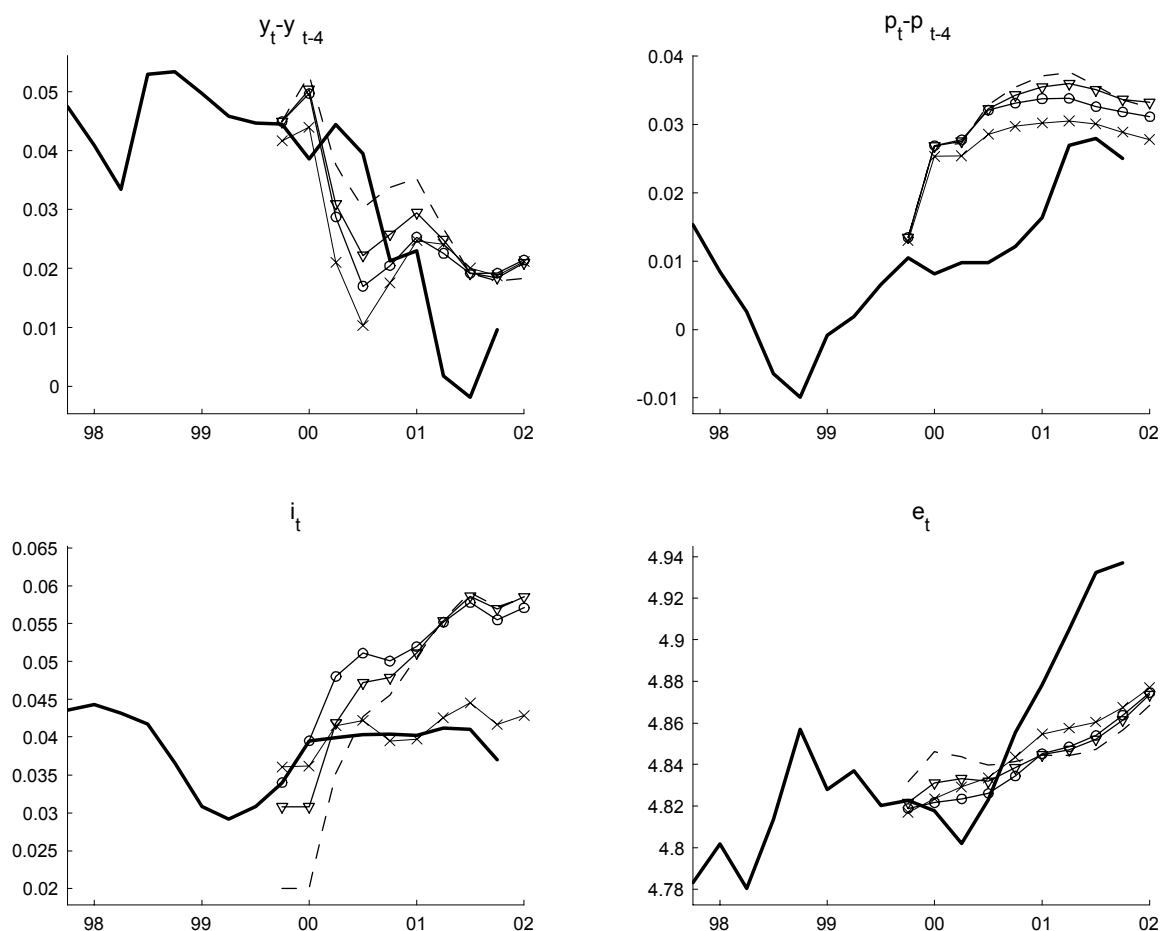


FIGURE 4. Forecasts conditional on different paths for the Swedish interest rate in the two quarters 1999:4 and 2000:1 generated exclusively by monetary policy shocks. Solid lines are the realized values.

Unconditional forecast ($- \times -$), Actual path: $i_{99:4} = 0.0340$, $i_{00:1} = 0.0395$ ($-o-$).
 No change: $i_{99:4} = i_{00:1} = 0.0308$, $i_{00:1} = 0.0308$ ($- \nabla -$). Easing: $i_{99:4} = i_{00:1} = 0.02$ ($- - -$).

| Event | Scenario | | | |
|---------------------------------------|----------|--------|----------|--------|
| | Uncond. | Actual | Constant | Easing |
| <1% GDP growth 2000 | 0.391 | 0.292 | 0.196 | 0.098 |
| <1% GDP growth 2001 | 0.388 | 0.347 | 0.352 | 0.366 |
| <1% inflation 2000 | 0.204 | 0.051 | 0.044 | 0.039 |
| <1% inflation 2001 | 0.266 | 0.164 | 0.137 | 0.141 |
| >3% inflation 2000 | 0.496 | 0.584 | 0.623 | 0.654 |
| >3% inflation 2001 | 0.483 | 0.531 | 0.567 | 0.565 |
| <1% GDP growth and >3% inflation 2000 | 0.219 | 0.190 | 0.136 | 0.074 |
| <1% GDP growth and >3% inflation 2001 | 0.211 | 0.209 | 0.223 | 0.236 |

TABLE 4. Posterior probability of certain events conditional on different monetary policy scenarios.

| | y | p | m | i | e |
|----------|--------|---------|--------|---------|--------|
| Actual | 0.6528 | -0.6528 | 0.4956 | -0.2478 | 0.4108 |
| Constant | 1.0393 | -1.0393 | 0.0467 | -1.0822 | 1.1724 |
| Easing | 2.3436 | -2.3436 | 0.4548 | -2.1165 | 2.3997 |

TABLE 5. Leeper-Zha modesty metric for the three monetary policy scenarios.

From Table 5 we find that the *Actual* and *Constant* scenarios result in modesty metrics below two for the five domestic variables. In the case of the one percent decrease, however, the metrics are above two for all variables but money. Hence, the actual path and the constant interest rate scenario may both be viewed as modest policy interventions and thus not likely to be subject to the Lucas critique. The third case, however, does not seem to be a modest intervention. Given that it implies an interest rate at 1.4 and 1.95 percent lower than the actual path over the two conditioning quarters, this does not seem unreasonable. However, the fact that the constant interest rate scenario means that the interest rate is 0.32 and 0.95 percent lower than the actual paths scenario suggests that a fairly broad range of hypothetical interest rate paths may be deemed as being consistent with modest policy interventions.

6. CONCLUDING REMARKS

This paper introduces a practicable Bayesian analysis of structural VARs with an arbitrary number of long run relations structured by general linear restrictions on the cointegration space. The structural part of the model is identified by restricting the contemporaneous relations among variables, following the tradition in the Bayesian structural VAR literature. There may be situations where long run identifying restrictions as in Blanchard and Quah (1989) are preferred, e.g. the restriction that monetary policy shocks have no long run impact on output. The related common trends approach introduced by King, Plosser, Stock and Watson (1991) and further developed by Warne (1993) connects the structural part of the model to its long run (cointegration) properties and provides a natural framework for the class of cointegrated processes studied here. A Bayesian analysis of models with long run restrictions is by no means straightforward and poses a real challenge for future methodological work.

APPENDIX A. PROOF OF THEOREM 1

A.1. The joint posterior. The joint posterior distribution of Γ_0, Γ, α and β conditional on the data D is

$$p(\Gamma_0, \alpha, \beta, \Gamma | D) \propto |\Gamma_0|^T \text{etr} [(X\Gamma_0 - W\beta\alpha' - Q\Gamma)'(X\Gamma_0 - W\beta\alpha' - Q\Gamma)] p(\Gamma)p(\alpha, \beta)$$

where $\text{etr}(B) = \exp(-\frac{1}{2} \text{tr } B)$, for any quadratic matrix B ,

$$p(\Gamma) \propto \exp[-\frac{1}{2}(\text{vec } \Gamma)' \Omega_\Gamma^{-1}(\text{vec } \Gamma)]$$

and

$$\begin{aligned} p(\alpha, \beta) &= p(\alpha|\beta)p(\beta) \\ &\propto |(\beta' K_1 \beta)^{-1} \otimes K_2|^{-1/2} |\beta' K_1 \beta|^{-p/2} \exp\left(-\frac{1}{2}(\text{vec } \alpha)'(\beta' K_1 \beta \otimes K_2^{-1})(\text{vec } \alpha)\right) \\ &\propto \text{etr}(\alpha' K_2^{-1} \alpha \beta' K_1 \beta). \end{aligned}$$

A.2. Full conditional of Γ . Follows easily from the treatment of the multivariate regression in Zellner (1971).

A.3. Full Conditional of α . Using the same manipulations as for the multivariate regression model (Zellner, 1971) we obtain

$$p(\mathcal{D}|\Gamma_0, \Gamma, \alpha, \beta) \propto \exp\left[-\frac{1}{2}(\text{vec } \alpha - \text{vec } \hat{\alpha})'(\beta' W' W \beta \otimes I_p)(\text{vec } \alpha - \text{vec } \hat{\alpha})\right],$$

where $\hat{\alpha} = (X\Gamma_0 - Q\Gamma)'W\beta(\beta'W'W\beta)^{-1}$. The conditional prior of α is

$$p(\alpha|\beta) \propto \exp\left[-\frac{1}{2}(\text{vec } \alpha)'(\beta' K_1 \beta \otimes K_2^{-1})(\text{vec } \alpha)\right],$$

Thus,

$$\text{vec } \alpha|\Gamma_0, \beta, \Gamma, \mathcal{D} \sim N_{pr}(\bar{\mu}_\alpha, \bar{\Omega}_\alpha),$$

where $\bar{\Omega}_\alpha^{-1} = (\beta' W' W \beta \otimes I_p) + (\beta' K_1 \beta \otimes K_2^{-1})$ and

$$\bar{\mu}_\alpha = \bar{\Omega}_\alpha[(\beta' W' W \beta \otimes I_p) \text{vec } \hat{\alpha}] = \bar{\Omega}_\alpha \text{vec}(\hat{\alpha} \beta' W' W \beta) = \bar{\Omega}_\alpha \text{vec}[(X\Gamma_0 - Q\Gamma)'W\beta].$$

A.4. Full conditional of β . Let $U = X\Gamma_0 - Q\Gamma$. Straight-forward calculations (Zellner, 1971) give

$$\begin{aligned} p(\beta|\Gamma_0, \Gamma, \alpha, \mathcal{D}) &\propto \text{etr}[(U - W\beta\alpha')'(U - W\beta\alpha')] \text{etr}(\alpha' K_2^{-1} \alpha \beta' K_1 \beta) \\ &= \exp\left\{-\frac{1}{2}[\text{vec } U - (\alpha \otimes W) \text{vec } \beta]'[\text{vec } U - (\alpha \otimes W) \text{vec } \beta]\right\} \\ &\quad \times \exp\left(-\frac{1}{2}(\text{vec } \beta)'(\alpha' K_2^{-1} \alpha \otimes K_1)(\text{vec } \beta)\right) \\ &\propto \exp\left\{-\frac{1}{2}(\text{vec } \beta - \text{vec } \tilde{\beta})' A (\text{vec } \beta - \text{vec } \tilde{\beta})\right\}, \end{aligned}$$

where $A = \alpha' \alpha \otimes W' W + \alpha' K_2^{-1} \alpha \otimes K_1$, $\text{vec } \tilde{\beta} = A^{-1}(\alpha' \alpha \otimes W' W) \text{vec } \hat{\beta}$ and

$$\text{vec } \hat{\beta} = [(\alpha' \alpha)^{-1} \otimes (W' W)^{-1}](\alpha' \otimes W') \text{vec } U = \text{vec}[(W' W)^{-1} W' (X\Gamma_0 - Q\Gamma) \alpha (\alpha' \alpha)^{-1}].$$

Inserting $\text{vec } \hat{\beta}$ into the formula for $\text{vec } \tilde{\beta}$ gives

$$\begin{aligned} \text{vec } \tilde{\beta} &= A^{-1}(\alpha' \alpha \otimes W' W) \text{vec}[(W' W)^{-1} W' (X\Gamma_0 - Q\Gamma) \alpha (\alpha' \alpha)^{-1}] \\ &= A^{-1} \text{vec}[W' (X\Gamma_0 - Q\Gamma) \alpha]. \end{aligned}$$

Using that $\text{vec } \beta = h + H\phi$, yields

$$\begin{aligned} p(\beta|\Gamma_0, \Gamma, \alpha, \mathcal{D}) &\propto \exp\left\{-\frac{1}{2}(h + H\phi - \text{vec } \tilde{\beta})' A (h + H\phi - \text{vec } \tilde{\beta})\right\} \\ &\propto \exp\left\{-\frac{1}{2}[A^{1/2}(\text{vec } \tilde{\beta} - h) - A^{1/2} H\phi]'[A^{1/2}(\text{vec } \tilde{\beta} - h) - A^{1/2} H\phi]\right\} \\ &= \exp\left\{-\frac{1}{2}(\phi - \bar{\mu}_\phi)' H' A H (\phi - \bar{\mu}_\phi)\right\}, \end{aligned}$$

where $\bar{\Omega}_\phi^{-1} = H' A H$, $\bar{\mu}_\phi = \bar{\Omega}_\phi H' A (\text{vec } \tilde{\beta} - h) = \bar{\Omega}_\phi H' \{\text{vec}[W' (X\Gamma_0 - Q\Gamma) \alpha] - Ah\}$.

APPENDIX B. PROOF OF THEOREM 2

B.1. Full Conditional of ψ_i . Let $Z = W\beta\alpha' + Q\Gamma$.

$$\begin{aligned}
 p(\psi_i | \Gamma_{0,-i}, \Gamma, \alpha, \beta, \mathcal{D}) &\propto |\Gamma_0|^T \text{etr}[(X\Gamma_0 - Z)'(X\Gamma_0 - Z)] \\
 &\propto |\Gamma_0|^T \text{etr}[(\Gamma_0 - \hat{\Gamma}_0)'X'X(\Gamma_0 - \hat{\Gamma}_0)] \\
 &= |\Gamma_0|^T \exp\left(-\frac{1}{2} \sum_{k=1}^p (G_k\psi_k - \hat{\gamma}_k)'X'X(G_k\psi_k - \hat{\gamma}_k)\right) \\
 &\propto |\Gamma_0|^T \exp\left(-\frac{T}{2}(\psi_i - \mu_{\psi_i})'\Omega_{\psi_i}^{-1}(\psi_i - \mu_{\psi_i})\right)
 \end{aligned}$$

where $\Omega_{\psi_i}^{-1} = T^{-1}G_i'Y'YG_i$, $\hat{\Gamma}_0 = (\hat{\gamma}_1, \dots, \hat{\gamma}_p) = (X'X)^{-1}X'Z$ and $\mu_{\psi_i} = (G_i'X'XG_i)^{-1}G_i'X'X\hat{\gamma}_i$. Note that $\hat{\gamma}_i = (X'X)^{-1}X'z_i$, where z_i is the i th column of $W\beta\alpha' + Q\Gamma$, so that $\mu_{\psi_i} = (G_i'X'XG_i)^{-1}G_i'X'z_i$.

Let $R_i = \text{Chol}(\Omega_{\psi_i})$ such that $R_iR_i' = \Omega_{\psi_i}$ and, following Waggoner and Zha (2003b), decompose ψ_i as $\psi_i = R_i \sum_{j=1}^{s_i} \beta_j w_j$. Then

$$\begin{aligned}
 p(\psi_i | \Gamma_{0,-i}, \Gamma, \alpha, \beta, \mathcal{D}) &\propto |\Gamma_0|^T \exp\left(-\frac{T}{2}(\psi_i'\Omega_{\psi_i}^{-1}\psi_i - 2\mu_{\psi_i}'\Omega_{\psi_i}^{-1}\psi_i + \mu_{\psi_i}'\Omega_{\psi_i}^{-1}\mu_{\psi_i})\right) \\
 &\propto |\Gamma_0|^T \exp\left(-\frac{T}{2}\left[\left(\sum_{j=1}^{s_i} \beta_j w_j\right)' \left(\sum_{j=1}^{s_i} \beta_j w_j\right) - 2\mu_{\psi_i}'R_i'^{-1}\left(\sum_{j=1}^{s_i} \beta_j w_j\right)\right]\right) \\
 &= |\Gamma_0|^T \exp\left(-\frac{T}{2}\left[\sum_{j=1}^{s_i} (\beta_j^2 - 2\beta_j \mu_{\psi_i}'R_i'^{-1}w_j)\right]\right) \\
 &\propto |\Gamma_0|^T \exp\left(-\frac{T}{2}\left[\sum_{j=1}^{s_i} [(\beta_j - \hat{\beta}_j)^2]\right]\right),
 \end{aligned}$$

where $\hat{\beta}_j = \mu_{\psi_i}'R_i'^{-1}w_j$. As $|\Gamma_0| = |(\gamma_1, \dots, G_iR_i \sum_{j=1}^{s_i} \beta_j w_j, \dots, \gamma_p)| \propto |\beta_1|$ (Waggoner and Zha, 2003b), we have

$$p(\psi_i | \Gamma_{0,-i}, \Gamma, \alpha, \beta, \mathcal{D}) \propto |\beta_1|^T \exp\left(-\frac{T}{2}(\beta_1 - \hat{\beta}_1)^2\right) \prod_{j=2}^{s_i} \exp\left(-\frac{T}{2}(\beta_j - \hat{\beta}_j)^2\right).$$

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