

IMF Working Paper

The Transmission Mechanism of European Monetary Policy: Is There Heterogeneity? Is it Changing over Time?

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Abstract

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This paper investigates the transmission mechanism of monetary policy in the four largest euro area countries by means Bayesian estimation of dynamic econometric models. Based on pre-EMU evidence from Germany, France, Italy, and Spain, we show that: (i) there are differences in the timing of the effects of monetary policy on economic activity, but their cumulative impact after two years is rather homogeneous; (ii) the transmission mechanism seems to have changed over time in the run-up to EMU but its degree of heterogeneity has not decreased; (iii) the “European-wide” effects of monetary policy may have become faster in the second half of the 1990s. We interpret this evidence by conjecturing that the transmission mechanism of monetary policy had already become relatively homogenous in the second part of the 1990s.

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I. Introduction

The European Central Bank (ECB) has moved interest rates several times since it started to operate in January 1999, and yet we don't know what the magnitude and timing of the effects of its actions actually are. What are the effects on prices and output of a change in the common short-term interest rate? How long do these effects take to materialize? Are there differences in their impact across European countries and regions? Are these differences changing over time?

Most of these questions have already been asked in the literature. However, the answers provided so far are not entirely satisfactory. For instance, Monticelli and Tristani (1999) suggest consideration of the European Monetary Union (EMU) as a composite economic system rather than a collection of countries, and analyze the impact of monetary policy on what they call the 'EMU-wide economic system' by estimating a structural VAR with a GDP weighted average of individual time series of member countries.² If the transmission mechanism is similar across countries, this approach provides a measure of the European-wide effects of monetary policy that is as good as those obtained with alternative estimation methods. But if the transmission mechanism does differ across countries, because of different economic structures and/or institutions, the use of this approach cannot be justified. In this case, as shown by Pesaran and Smith (1995) for standard dynamic panel data models and discussed by Rebucci (2001) for general panel VAR specifications, aggregation of individual time series may bias the estimates obtained, and the European-wide impact of monetary policy must be measured either by aggregating individual time series estimates or by using other methods that allow for explicit variation of the parameters across countries. Before attempting to measure the system-wide effects of a 'synthetic' common European monetary policy by using aggregate time series models, one should therefore try to establish first whether there are cross-country differences in the transmission mechanism of monetary policy.

The existence of some degree of heterogeneity in the transmission mechanism of monetary policy in Europe is supported by a fairly large, albeit sometimes contradicting, body of empirical evidence.³ There are several methodological reasons why different studies might have come to very different conclusions. As noted by Guiso and others (2000), the specification of the econometric model sometimes differs across countries within the same study. It is difficult therefore to establish the extent to which different outcomes reflect true differences in the transmission mechanism or simply different econometric specifications. Second, most studies compare responses to monetary policy neglecting the contemporaneous and lagged interdependence between these economies. Obviously, this can provide only a

² A similar approach is followed by Peersman and Smets (2001) in studying whether monetary policy has asymmetric effects across business cycle states in European countries and by Ortega and Alberola (2000) in analysing the simulated impact of different kinds of shocks in the Euro-area.

³ See Guiso and others (2000) for a recent survey of this literature.

partial description of the transmission mechanism of monetary policy in open economies of a relatively well integrated area. In addition, this may distort the identification of country-specific monetary policy shocks. As noted by Dornbusch and others (1998), omitting the contemporaneous effect of German interest rates in the reaction function of other European central banks may erroneously lead to identifying as a country-specific monetary shock what in fact is an endogenous response to a German monetary shock. Third, as also noted by Guiso and others (2000), the kind of analysis typically conducted is not informative on what is likely to happen under EMU as most of these studies do not control for intra-Europe exchange rate movements and for heterogeneous preferences over inflation and output stabilization objectives in central banks' preferences—two potential sources of differences in the transmission mechanism of monetary policy that have disappeared under EMU.⁴

This literature is also potentially subject to the Lucas critique as it attempts to draw inference relevant for EMU based on econometric models estimated under a very different regime—the fixed but adjustable exchange rate regime of the European Monetary System (EMS). Indeed, the current 'consensus' view is that existing differences in the transmission mechanism are likely to decrease over time as real, and especially financial, convergence proceeds.⁵ However, there is no hard evidence that existing differences are decreasing over time. On the contrary, recent work by Cecchetti (1999) shows that they might persist for a long time because they are due to differences in the financial structure, which in turn are rooted in the legal framework of individual countries. If these differences were to persist for sometime, the ECB's life may become quite complicated, as pointed out by Dornbusch and others (1998), explicitly modelled by Giovannetti and Marimon (1998), and analyzed in simulations by Hughes-Hallet and Piscitelli (1999). It would be useful, therefore, to have some idea not only of the magnitude of these differences, but also of their degree of persistence over time.

This paper tries to overcome some of the methodological difficulties encountered in previous attempts to investigate the transmission mechanism of European monetary policy, and to analyse also its evolution over time, by rephrasing the key questions in the framework of a heterogeneous, time-varying panel *VAR* model recently proposed by Canova and Ciccarelli (2000). This is a flexible empirical framework that lets the parameters of the transmission mechanism differ both across countries and over time periods, and hence sets the stage for testing alternative homogeneity assumptions—including the extent to which parameter heterogeneity across countries has changed over time. By allowing for contemporaneous and lagged interdependence between open and integrated economies, it allows also for better

⁴ A notable early exception is represented by Dornbusch and others (1998), whose empirical evidence is based on a model allowing for limited interdependence between countries, controlling for intra-Europe exchange rate movements, and in which the impact of a 'coordinated' change in interest rates can be analysed. More recently, Clements, Kontolemis, and Levy (2001) and Sala (2001) have produced new evidence controlling more thoroughly for heterogeneous preferences, in addition to intra-Europe exchange rate movements and limited interdependence.

⁵ See Cecchetti (1999) for a summary of the arguments in favour of the 'convergence' hypothesis.

identification of monetary policy shocks and a more realistic description of their transmission mechanism—including their area-wide effects. The area-wide effects, in particular, can be recovered and measured regardless of the actual degree of heterogeneity present in the data.

Obviously, such a framework cannot be estimated without introducing restrictions on the model because of the very large number of parameters involved. Following Canova and Ciccarelli (2000), we specify the econometric model hierarchically (in a sense made clear in the next section) in terms of few hyper-parameters and take a Bayesian approach to estimation. In addition, we measure monetary policy by estimating a system of reaction functions *a-la* Clarida, Gali, and Gertler (1997), and then assess the impact of monetary policy on economic activity by estimating a system of output equations as done by Dornbusch and others (1998) and Peersman and Smets (1998). Thus, we do not model nominal exchange rates and inflation rates endogenously. The specification of the econometric model is the same for all countries considered, and is also similar to that used by Dornbusch and others (1998) even though it allows for a larger set of objectives to enter the reaction function of each central bank to take into account the varying degree of commitment to EMS of each country over time. The most novel feature of this paper, however, is the allowance of parameter variation over time. As far as we know, this is the first study of the transmission mechanism of European monetary policy which attempts to do so.

We consider a small group of European countries: Germany, France, Italy, and Spain.⁶ These are the four largest economies currently in the EMU, accounting for about 80 percent of the Euro-area GDP. France has been closely tied to Germany throughout the period considered and is widely regarded as a ‘core’ European economy (together with the Netherlands, Belgium, and Luxembourg), even though its legal structure is classified differently from that of Germany by Cecchetti (1999). Italy has often been singled out as a ‘divergent’ European partner mainly because of developments in its public finances until the mid-1990s and its dual economic structure. On the other hand, Spain has been catching up with the rest of Europe throughout the period considered. It is likely to have started from a very different position, but might have gone a long way toward closing the gap with the ‘core’ of Europe, including in terms of its sensitiveness to monetary policy. Hence, there is no strong reason to expect an *homogeneous* response to monetary policy across these economies from an ex-ante perspective.

Consistent in part with the rest of the literature, we show that there are some differences in the transmission mechanism of European monetary policy. But contrary to what previously thought, we show also that these differences seem a matter of *timing* rather than *magnitude* of the cumulative impact of monetary policy: the cumulative impact of country-specific monetary shocks after two years are rather homogenous in Germany, France, and Italy, and in

⁶ There is nothing in our empirical framework that would prevent us from including more than four countries except additional computing costs.

all countries considered if parameter variation over time is allowed for. We show also that the transmission mechanism of monetary policy was changing over time in these countries in the second part of the 1990s, but the existing degree of heterogeneity was not decreasing. This is mainly because the bulk of the change appears to have originated from changes in the common component of the transmission mechanism. Finally, we provide some evidence on the European-wide effects of monetary policy and show that, in the second part of the 1990s, the transmission mechanism might have become faster, peaking in between six and nine months. We interpret these results as evidence that the transmission mechanism of European monetary policy had already become relatively homogenous in the second part of the 1990s in these countries; thus, our results cast doubts on the presumption that the transmission mechanism of European monetary policy is heterogeneous and on the policy implications of the analysis of Cecchetti (1999), Giovannetti and Marimon (1998), and Hughes-Hallet and Piscitelli (1999).

The paper is organized as follows. The econometric framework is presented in section 2. Here we illustrate first the empirical model of the behaviour of these countries' central banks, and then the system of output equations. The rest of the paper reports and discusses the empirical results: the estimated monetary policy shocks in section 3; the empirical evidence on the effects of monetary policy shocks on economic activity and their degree of homogeneity across countries and stability over time in section 4; and the results on the European-wide impact of monetary policy in sections 5. Section 6 concludes. Details of the estimation techniques used are given in the appendix.

II. The Econometric Framework

Ideally, one would like to apply the empirical framework proposed by Canova and Ciccarelli (2000) to a small structural *VAR* for output, inflation, interest rates, and exchange rates, the set of variables usually considered in the literature. This approach is feasible in principle, but, while allowing for unconstrained interdependence between countries and parameter variation over time, it is, in practice, it is extremely demanding computationally.

Here, we follow the two-stage approach used by Dornbusch and others (1998) and Peersman and Smets (1998) and do not model inflation and the exchange rate explicitly. In the first stage, a measure of monetary policy is extracted from the data by estimating a system of reaction functions (one for each central bank) allowing for simultaneity and interdependence in short-term interest rates and parameter variation across countries and time periods. In the second stage, the impact of monetary policy on economic activity is analyzed by estimating a system of dynamic output equations allowing for parameter variation across countries and time periods, lagged interdependence and no contemporaneous simultaneity.

In the following two sub-sections, we present the econometric model of the reaction functions and output equations in turn.

A. Measuring Monetary Policy

1. Specification

The behaviour of the four European central banks considered is modelled as a *system* of reaction functions of the type discussed and estimated by Clarida, Gali, and Gertler (1997).

Following Dornbusch and others (1998), we model this system of reaction functions empirically by specifying the following structural *VAR*:

$$A_t(L) R_t = B_t(L) W_t + D_t + U_t, \quad (1)$$

where $R_t = [r_t^1, \dots, r_t^4]'$ is a (4 x 1) vector of monetary policy instruments, $W_t = [w_t^1, \dots, w_t^4]'$ is a (4 x 1) vector of monetary policy final objectives, $A_t(L)$ and $B_t(L)$ are time-varying polynomial matrices in the lag operator L with lag length p , and D_t is a (4 x 1) vector of constants. Here, $U_t = [u_t^1, \dots, u_t^4]'$ is a (4 x 1) vector of monetary policy shocks assumed to be normally distributed and such that:

$$\begin{aligned} E[U_t | Z_{t-s}] &= 0, \text{ for all } t \text{ and } s > 0; \\ E[U_t U_t' | Z_{t-s}] &= I, \text{ for all } t \text{ and } s > 0; \\ E[U_t U_s'] &= 0, \text{ for all } t \neq s, \end{aligned} \quad (2)$$

where Z_t contains lagged R_t and contemporaneous and lagged W_t , and E denotes the expectation operator.

The monetary policy instrument is assumed to be a short-term interest rate. In each element of the (4 x 1) vector of final objectives, $w_t^i = [(\pi_{i,t} - \pi_i^*), (y_{i,t} - y_i^*), (e_{i,t} - e_i^*), \sigma_{i,t}]'$, we include contemporaneous and lagged inflation (π), output (y) and the nominal exchange rate (e), in percent deviation from the target (π^* , y^* , e^* , respectively), and a measure of the (unconditional) intra-month exchange rate volatility (σ) to control for shocks to exchange rate risk premia.⁷

⁷ While the inclusion of contemporaneous inflation and output gaps in the information set of policy makers is not controversial, because of the lags with which monetary policy affects activity and the presence of nominal rigidities, the inclusion of the nominal exchange rate—though not uncommon in the literature—might be questioned. Bagliano, Favero, and Franco (1998), however, show that this is not an empirically relevant problem (at least in the case of the U.S.) as they find that the contemporaneous correlation between an exogenous measure of the unexpected component of monetary policy and the DM/US dollar rate is statistically insignificant. Comforted by this evidence on the U.S. case, we include also contemporaneous exchange rate gaps in (1).

As a proxy for short-term interest rates we use 3-month Treasury bill rates. Output is measured by an industrial production index. Inflation is measured by the annual change in the consumer price index. We use the bilateral exchange rate vis-a-vis the deutsche mark (DM) for France, Italy, and Spain and the DM/US dollar rate for Germany. Bilateral rates vis-a-vis the DM are obtained as cross rates vis-a-vis the US dollar. The target variables (π^* , y^* , and e^*) are the fitted values of a linear regression of the actual variables ($\pi_{i,t}$, $y_{i,t}$, and $e_{i,t}$) on a constant and a linear trend, a constant and a quadratic trend, and a simple constant, respectively. σ is measured by the intra-month standard deviation of the nominal exchange rate in percent deviation from trend, where the trend is obtained by using an HP filter with smoothing parameter equal to 1600.⁸

The specification chosen imposes very few a-priori restrictions on the system of reaction functions: all parameters in $A_t(L^P)$ and $B_t(L^P)$, except those governing the contemporaneous causation among short-term interest rates, are unrestricted and can vary over time.⁹ This allows for the possibility of change in the behaviour of the central banks considered during the sample period analysed. In particular, given that the degree of each member's commitment to the EMS has varied over time, leaving $B_t(L^P)$ unrestricted lets the data reveal which objective was actually pursued in each period, while leaving $A_t(L^P)$ unrestricted for all $p \neq 0$ allows for lagged interdependence among short-term interest rates of different countries as well as varying degrees of interest rate smoothing over time.¹⁰ Nonetheless, we do impose an arguably strong lag length restriction, assuming $p = 1$ for all countries and variables considered, which we shall discuss in section 3 while presenting the estimation results.

Our assumptions on $A_t(L^P)$ and $B_t(L^P)$ are the key innovation in the specification of (1) compared to that used by Dornbusch and others (1998). Dornbusch and others assume that, throughout the period 1986-1996, Germany was targeting its own inflation, output, and exchange rate gaps, while France, Italy, and Spain, because of the EMS constraint, were targeting German variables; thus, they specify $A_t(L^P)$ and $B_t(L^P)$ as time invariant, and restrict them accordingly. While this is a good first approximation to a complex reality, one may want to take into account the fact that the Spanish peseta joined the EMS only in 1989, that the Italian lira has been floating more or less freely from September 1992 to November 1996, and that the fluctuation bands of all three currencies vis-a-vis the DM have changed several times during the period considered. Even the Bundesbank's focus might have shifted

⁸ All the data used are from the International Financial Statistics database of the IMF, except daily exchange rates which are courtesy of Marcello Pericoli of the Bank of Italy, whom we thank.

⁹ Assuming that the coefficient matrix of L^0 in $A_t(L)$ is constant over time renders the posterior distributions analytically tractable and is equivalent to assuming homoschedaticity of the structural residuals.

¹⁰ Note that the relative tightness of the prior distribution given on the elements of $A_t(L)$ and $B_t(L)$ distinguishes between own and other countries' monetary policy instruments (the endogenous variables) on the one hand, and between instruments and objectives on the other hand, but does not distinguish between own and other countries' objectives (the exogenous variables). See appendix for more details on this.

away from strictly domestic objectives, after the early years of the German unification, to pay closer attention to the level and/or the volatility of the DM vis-a-vis the currencies of its perspective EMU partners and the degree of synchronization between Germany's business cycle and the rest of Europe in the run up to EMU. It is evident that, if these policy changes are not accounted for by the estimated parameters of the system of reaction functions, they will end up in estimated residuals, thereby potentially undermining their interpretation of well-behaved (i.e., white noise) policy innovations as assumed in (2).

2. Identification

Identification of (1) may be achieved through exclusion restrictions on the coefficient matrix of L^0 in $A_t(L)$.

The identification scheme exploits the Bundesbank's presumed leading role under the ERM and the fact that other European countries considered are of comparable size.¹¹ More specifically, to identify the model, we place the German short-term interest rate first in the vector R_t , assuming that it affects other European interest rates contemporaneously without being affected by them, and then assume that the impact on country j of an increase in interest rates in country i is the same as the impact on country i of an increase in country j for $i, j = 2, 3, 4$.

Formally, the leader-follower behavior presumably characterizing the ERM is translated into the following block recursive structure for the coefficient matrix of L^0 in $A_t(L)$:

$$A(0) = \begin{bmatrix} A_{11}(0) & 0' \\ A_{21}(0) & A_{22}(0) \end{bmatrix} \quad (3)$$

where $A_{11}(0)$ is a scalar, $A_{21}(0)$ is 3×1 , and $A_{22}(0)$ is 3×3 . This gives us three restrictions. The remaining three restrictions needed to identify the model are obtained by imposing symmetry on $A_{22}(0)$. These six restrictions identify the model exactly irrespective of the order of the non-German interest rates in R_t .¹²

The structural VAR (1), therefore, can be rewritten as:

$$\begin{bmatrix} A_{11}(0) & 0' \\ A_{21}(0) & A_{22}(0) \end{bmatrix} \begin{pmatrix} R_t^1 \\ R_t^2 \end{pmatrix} = \begin{bmatrix} A_{11}(L) & A'_{12}(L) \\ A_{21}(L) & A_{22}(L) \end{bmatrix}_t \begin{pmatrix} R_{t-1}^1 \\ R_{t-1}^2 \end{pmatrix} + \begin{bmatrix} B_{11}(L) & B_{12}(L)' \\ B_{21}(L) & B_{22}(L) \end{bmatrix}_t \begin{pmatrix} W_t^1 \\ W_t^2 \end{pmatrix} + D_t + \begin{pmatrix} U_t^1 \\ U_t^2 \end{pmatrix} \quad (4)$$

¹¹ See Giavazzi and Giovannini (1989) and Kenen (1995) on this view of the functioning of the EMR from the mid-1980s onward.

¹² See Amisano and Giannini (1997, p. 166-67).

where R_t^1 , W_t^1 , and U_t^1 are the German monetary policy instrument, objectives, and shock, respectively, while R_t^2 , W_t^2 , and U_t^2 are the vectors containing the same variables for France, Italy, and Spain.

3. Estimation

Bayesian estimation of (4) exploits its block recursive structure.

Following Zha (1999), let k_j and G_j be the total number of right-hand-side variables per equation and the total number of equations in block j of (4), respectively, where the same set of variables enters the equations of each block j . If we pre-multiply (4) by the (4×4) matrix

$$A_d^{-1}(0) = \begin{bmatrix} A_{11}^{-1}(0) & 0' \\ 0 & A_{22}^{-1}(0) \end{bmatrix},$$

and rearrange terms, the model can be divided into two blocks:

$$R_t^j = Z_t^j \delta_t^j + v_t^j \quad j = 1, 2, \text{ for all } t. \quad (5)$$

Here, $Z_t^j = \text{diag} [Z_{1,t}^j, Z_{2,t}^j, \dots, Z_{G_j,t}^j]$ denotes a $(G_j \times k_j G_j)$ diagonal matrix whose elements are the $(1 \times k_j)$ vectors, $Z_{g,t}^j$, containing all contemporaneous (in our case only R_t^1 in block 2) and lagged endogenous variables, exogenous and deterministic variables, of equation g in block j for $g = 1, \dots, G_j$; $\delta_t^j = [\delta_{1,t}^j, \delta_{2,t}^j, \dots, \delta_{G_j,t}^j]$ denotes a $(k_j G_j \times 1)$ vector whose $(k_j \times 1)$ elements, $\delta_{g,t}^j$, contain the parameters of equation g in block j (for $g = 1, \dots, G_j$); and $v_t^j = A_{jj}^{-1}(0) U_t^j$ with

$$v_t^j \sim N(0, \Sigma_{jj}), \quad \Sigma_{jj} = A_{jj}^{-1}(0) A_{jj}^{-1}(0)', \text{ and } E[v_t^i v_t^{j'} | Z_{t-s}] = 0;$$

$$\text{for } i \neq j, \text{ all } t, \text{ and } s > 0.$$

As explained in more detail in the appendix, Bayesian estimation of the two blocks of (5) is obtained by using the Kalman filter and the Gibbs sampler to take into account the presence of time-varying parameters, as suggested by Chib and Greenberg (1995) and as done by Canova and Ciccarelli (2000). Intuitively, a joint prior on $(\delta_t^j, \Sigma_{jj})$ is combined with the likelihood of the data and suitable initial values of the model's hyperparameters to recover the joint posterior distribution of $(\delta_t^j, \Sigma_{jj})$. As the analytic integration of this distribution is difficult, this is implemented numerically by means of Monte Carlo simulation methods. Since the

matrices $A_{jj}(0)$ are exactly identified, and thus linked to Σ_{jj} with a one-to-one mapping, we can recover the posterior distribution of the structural parameters of the model, and hence the posterior distribution of the structural residuals (U_t), from the estimate of the model's reduced form for each iteration of the Gibbs sampler. The average of the empirical distribution of these residuals is then taken as our measure of the random or *unexpected* component of monetary policy.

B. The Transmission Mechanism of Monetary Policy

1. Specification

The impact of monetary policy on economic activity is modelled empirically through a system of output equations in which real output growth is regressed on our measure of the unexpected component of monetary policy and a set of control variables. For each country i , we specify the following output equation:

$$y_t^i = \sum_{j=1}^4 b_j^i y_{t-1}^j + b_\pi^i \pi_{t-1}^i + b_e^i e_{t-1}^i + (b_{u_1}^i \hat{u}_{t-1}^i + \dots + b_{u_{24}}^i \hat{u}_{t-24}^i) + \varepsilon_t^i,$$

that can be written in a more compact form as:

$$y_{it} = X'_{it} \beta_{it} + \varepsilon_{it}. \quad (6)$$

Here, y_{it} is the annual growth of industrial production, $X'_{it} = [\hat{u}_{it-l_i}, x_{it}]'$ is a $(1 \times k)$ vector of regressors with \hat{u}_{it-l_i} denoting lags of the series of estimated monetary policy shocks and x_{it} denoting the set of control variables, $\beta_{it} = [\beta_{it}^1, \beta_{it}^2]'$ is a $k \times 1$ vector of parameters with β_{it}^1 and β_{it}^2 denoting the coefficients of \hat{u}_{it-l_i} and x_{it} , respectively.

Following Dornbusch and others (1998), in x_{it} , we include lagged output growth of all countries considered to capture regional interdependencies, the first lag of the nominal exchange rate (vis-a-vis the DM for France, Italy, and Spain and vis-a-vis the US dollar for Germany) to hold constant the exchange rate channel of transmission of monetary policy, and the first lag of the annual inflation rate (of country i only) to control for domestic supply-side factors. The specification is the same for all countries considered and, in addition to the variables already mentioned, includes a constant and 24 lags of \hat{u}_{it-l_i} ($l_i = 1, 2, \dots, 24$), for a total of 30 regressors per equation.

The distinguishing feature of (6) is that it allows β_{it} to vary randomly both across countries and time periods, even though *a priori* only as different draws from the *same* distribution. This is achieved by assuming that β_{it} is a random variable drawn from a common

prior distribution, which changes also randomly over time, according to a given law of motion. Formally, for each country i and time t , we assume the following hierarchical structure of the prior distribution:

$$\beta_{it} = \theta_t + \zeta_{it} \quad \zeta_{it} \sim N(0, b_o), \quad (7)$$

$$\theta_t = \theta_{t-1} + \eta_t \quad \eta_t \sim N(0, B_1). \quad (8)$$

Here, b_o and B_1 denote the variance of the distribution of ζ_{it} and η_t , respectively. B_1 controls the time-variation of the prior mean of the parameters, whereas b_o controls their variation around the mean, both across countries and over time.¹³ If $B_1 = 0$, $\beta_{it} = \theta + \zeta_{it}$ for all t , and the parameters vary randomly across countries and over time around a constant mean. On the other hand, if $b_o = 0$, $\beta_{it} = \theta_{t-1} + \eta_t$ for all i . In this case, no cross-sectional heterogeneity is present, and β_{it} is shrunk towards a common time-varying mean. If both B_1 and b_o are zero, $\beta_{it} = \theta$ for all i and t and the prior distribution of the parameters collapse on a common constant. The prior variances of η_t and ζ_{it} , therefore, provide a means to control the degree of prior uncertainty introduced in the model with respect to how the parameters of interest may change over countries and time periods.

The assumptions (7-8) are only priors that must be combined with the data to generate posterior distributions of the parameters of interest. The moments of the posterior distribution of β_{it} do not need to be the same as those characterizing the prior, as the former are derived from a ‘mixture’ of the information contained in the data and that specified in the prior. Note in particular that, while the prior variance of β_{it} ($b_o + B_1$) is time-invariant, the posterior variance of β_{it} may change over time due to realizations of both η_t and ζ_{it} (see equation (21) in appendix on this). The assumptions in (7-8), therefore, clearly permit checking whether the degree of heterogeneity of the parameters of the transmission mechanism (i.e., the variance of the posterior distribution of these parameters) has changed over time.

2. Estimation

Stacking all equations by row and rewriting (6) as a standard system of seemingly unrelated regressions (SUR), we have:

$$y_t = X_t \beta_t + \varepsilon_t, \quad \varepsilon_t \sim N_g(0, \Omega). \quad (9)$$

¹³ The specification of the law of motion of θ_t in (8) implies that the parameters have an unconditional mean equal to zero. An alternative specification is:

$$\theta_t = \rho \theta_{t-1} + (1 - \rho) \bar{\theta} + \eta_t,$$

where $\bar{\theta}$ is the long run mean of θ_t . However, when we estimated the hyperparameter ρ by maximizing the sample likelihood in (9) below for each country i , we found values for ρ ranging from 0.9985 for Spain to 1 for France and Italy. Given this evidence, we decided to adhere to the computationally simpler specification in (7).

In this system $X_t = \text{diag} [X'_{1t}, \dots, X'_{Gt}]$ is of dimension $G \times h$, with $h = G * k$, $G = 4$ denoting the number of endogenous variables and $k = 31$ denoting the number of regressors in each equation, while $\beta_t = [\beta_{1t}, \dots, \beta_{gt}]'$ is of dimension $h \times 1$.

The assumptions on the prior distribution of the parameters' vector β_t can then be represented as:

$$\beta_t = M_o \theta_t + \zeta_t, \quad \zeta_t \sim N_h(0, B_o) \quad (10)$$

$$\theta_t = \theta_{t-1} + \eta_t, \quad \eta_t \sim N_m(0, B_1) \quad (11)$$

where the $(h \times k)$ matrix M_o is a column vector of G identity matrices of order k that relates β_t to the $(k \times 1)$ vector of common shift parameters θ_t , and Ω , B_o , and B_1 are unknown variance-covariance matrices of ε_t , ζ_t and η_t , respectively. The latter three random vectors are assumed mutually independent, implying that y_t is conditionally independent of θ_t , B_o , and B_1 .

As explained in the appendix, Bayesian estimation of the hierarchical model (9-11) is then performed in a manner similar to that described for the system of reaction functions. In addition to the assumptions stated in (9-11), prior assumptions are given on the hyperparameters of the model Ω , B_o , B_1 and combined with the information contained in the data (in the form of a likelihood function and initial conditions) to obtain posterior distributions.

There are two main differences between the system of reaction functions and that of output equations. First, rather than assuming a normal diffuse prior for the variance-covariance matrix of the residuals, Ω , here we assume a natural conjugate prior, a Wishart distribution, so that its posterior distribution is also Wishart. Second, we add one additional stage to the hierarchical structure of the model to allow for greater flexibility in studying the transmission mechanism of monetary policy and to obtain full Bayesian estimates of its parameters. As in the case of the estimation of the reaction functions, analytical integration is not feasible, and the Gibbs sampler is used to compute posterior distributions of the parameters of interest numerically.

3. Testing

Several hypotheses of parameter homogeneity can be performed on the posterior distribution of the parameters of interest. Of particular interest is the overall extent of time variation of the posterior distribution of the parameters of the transmission mechanism of monetary policy, their degree of heterogeneity across countries, and any tendency of this heterogeneity to change over time. More specifically, we want to test (i) the null (or the prior) assumption of parameter variation over time, and (ii) the null that the transmission

mechanism is homogeneous across countries, either over the entire sample period or in each yearly subperiod considered.

(i) Testing the Transmission Mechanism Instability Over Time. This hypothesis can be tested by letting B_1 depend upon two hyperparameters, ϕ_1 and ϕ_2 , the first controlling the time variation of the monetary policy parameters, and the second the time variation of other parameters. If the posterior distribution of ϕ_1 , which depends on both the prior assumption and the information content of the data, is concentrated around values closer to zero than its prior, then we can conclude that the evidence supporting a time-varying specification (at least for the parameters of the transmission mechanism of monetary policy) is weak (Chib and Greenberg, 1995, page 344).

Let us assume a proper prior distribution for ϕ_1 . If we were to find that the probability that ϕ_1 is arbitrarily small is larger under the *posterior* distribution than under the *prior* distribution, then we could conclude that the data have shifted the ‘odds’ in favour of a small ϕ_1 , and thus against a time-varying specification. Formally, following Chib and Greenberg (1995), the test is implemented by calculating the ratio

$$z = \left(\frac{\Pr(\phi_1 \leq \xi | y)}{\Pr(\phi_1 \leq \xi)} \right) \left(\frac{\Pr(\phi_1 > \xi | y)}{\Pr(\phi_1 > \xi)} \right), \quad (12)$$

for arbitrarily small values of ξ , where $\Pr(\phi \leq \xi | y)$ and $\Pr(\phi \leq \xi)$ denote the conditional *posterior* probability and unconditional *prior probability* that ϕ is less than ξ , respectively. The ratio in the first bracket compares the odds, under the posterior distribution, that $\phi_1 \leq \xi$, while the ratio of complementary probabilities in the second bracket is a weighting factor. The numerators of both ratios are obtained as relative frequencies from the Gibbs sampler, while the denominators are given by the prior assumption. Checking the value of this ratio for a number of arbitrarily small values of ξ , therefore, provides a way to implement a ‘specification test’ on the prior assumption that the variance of η_t is greater than zero, and hence provides a specification test for the null hypothesis that the parameters do change over time.

(ii) Testing the Homogeneity of the Transmission Mechanism Across Countries and Over Times. The presence of cross-country differences in the transmission mechanism of monetary policy is tested by using a procedure proposed by Ciccarelli (2001), which is an empirical-Bayesian analogous of the classical Wald-test, similar to that discussed by Hamilton (1994, pages 355-358). In the classical Wald test, one compares two quadratic forms: one *asymptotically* distributed as a $\chi^2_{(d)}$ with d degrees of freedom under the null hypothesis (which is assumed to hold exactly), and the other distributed as a *non-central* $\chi^2_{(d)}$ under the alternative. The greater the numerical value of the quadratic form in which the exact restrictions have been substituted, the more likely it is that the value drawn belongs to the distribution under the alternative hypothesis. The main difference with respect to the classical

Wald test is that, here, we know the *empirical* distribution of the quadratic form under the null assumption, while the null hypothesis is formulated as a *probabilistic* statement about the posterior distribution of a function of the parameters of interest. The distribution of the quadratic form under the null hypothesis becomes a ‘reference’ distribution, which can be compared to the distribution of the restricted quadratic form once both have been constructed empirically by means of the Gibbs sampler.

Let us write the null hypothesis of homogeneity of the parameters of interest as a general set of restrictions on the complete parameter vector β_t ¹⁴

$$R(\beta_t) = r, \quad \text{for each } t. \quad (13)$$

Conditional on other parameters of the model and given the specification (9)-(11), the posterior distribution of β_t is:

$$\beta_t \sim N(\hat{\beta}_t, \hat{\Omega}_t).$$

Therefore, the conditional posterior distribution of a linearized version of $R(\beta_t)$ is:

$$R(\beta_t) \sim N\left(R(\hat{\beta}_t), \nabla R(\hat{\beta}_t)' \hat{\Omega}_t \nabla R(\hat{\beta}_t)\right).$$

where $\nabla R(\hat{\beta}_t)$ denotes the gradient of the vector $R(\beta_t)$ computed at $\hat{\beta}_t$.

The test then is based on the comparison of these two quadratic forms:

$$q_t = \left(R(\beta_t) - R(\hat{\beta}_t)\right)' \left(\nabla R(\hat{\beta}_t)' \hat{\Omega}_t \nabla R(\hat{\beta}_t)\right)^{-1} \left(R(\beta_t) - R(\hat{\beta}_t)\right) \quad (14)$$

and

$$q_{1t} = (R(\beta_t) - r)' \left(\nabla R(\hat{\beta}_t)' \hat{\Omega}_t \nabla R(\hat{\beta}_t)\right)^{-1} (R(\beta_t) - r). \quad (15)$$

If the posterior distribution of $R(\hat{\beta}_t)$ is centered on r —that is, in the limit, the restrictions (13) are true with probability 1 and $R(\hat{\beta}_t) \equiv r$ —then q_{1t} must have the same distribution as q_t ; otherwise, it is conditionally distributed as a non-central distribution with respect to the distribution of q_t . In order to construct a rejection region for the null hypothesis, therefore, it

¹⁴ In the specific case of linear restrictions, the restriction matrix $R=[R_{i,j}]$ has dimension $d \times Gk$, where G and k are defined as before, $d = (G - 1) p_m$, and p_m is the number of monetary policy coefficients restricted to be the same across countries. In particular, the null hypothesis that all parameters of the transmission mechanism are equal implies $p_m = 24$. In this case, R has 72 rows, whose values are 1 when $i = j$, -1 when $j = i + k$, and 0 otherwise. The hypothesis that the impact of monetary policy at specific lags, or that its cumulative effect after one or two years, are equal across countries can also be easily accommodated designing R accordingly.

is enough to compare these two distributions. Specifically, the larger the distance between q and q_1 , the more likely is that the restrictions imposed are converting the reference distribution in a *non-central* distribution, and thus the greater is the probability, *a posteriori*, that the null hypothesis is false. The empirical posterior distributions of q and q_1 are easily obtained from the Gibbs sampler. The distance between these two distributions can then be quantified using a standard Kolmogorov-Smirnov statistics.

Now, if the model is specified with time-varying parameters, we can easily compute the empirical distributions of q and q_1 and quantify their distance for each subperiod considered. Thus, we can test the null hypothesis of parameter homogeneity across countries for each subperiod considered. The time profile of the Kolmogorov-Smirnov statistic measuring the distance between the two distributions, therefore, can provide a clear indication of the direction of change of the differences across countries in the parameters of the transmission mechanism, if any is found.

The illustration of the procedures used to test for homogeneity and stability of the transmission mechanism of monetary policy concludes the presentation of the econometric framework. The next three sections discuss the empirical results.

III. Estimated Monetary Policy Shocks

In this section we report and discuss the estimated posterior mean of the structural residuals derived from estimation of (4) that will be used in the rest of the paper as our measure of the unexpected component of monetary policy.¹⁵

The estimated structural residuals of equation (4), are plotted in Figure 1, from after the German unification in January 1991 to the eve of the EMU launch in December 1998.¹⁶ Interestingly, these residuals look remarkably well behaved: there is only one large outlier (for

¹⁵ The posterior distributions of the parameters of the reaction function of the four central banks considered are not reported here because of space constraints, but are available on request and are discussed in Ciccarelli and Rebucci (2001). These distributions are symmetric and generally their means have the expected signs. They display also significant parameter time variation, especially until 1992-93 for Germany and 1994-1995 for other countries, and relatively high persistence with an autoregressive coefficient ranging between 0.7 and 0.9 in all countries considered in the second part of the 1990s. Exchange rate volatility appears to matter for all countries considered. Germany's seems to have reacted mainly to domestic objectives, even though the Bundesbank's attention appears to have shifted in the run up to EMU from the dollar value of the DM to the external value of the DM vis-a-vis other European currencies. France, Italy, and Spain seem to have had different reaction functions. All three central banks, however, reacted strongly to contemporaneous movements in German interest rates. The behavior of the central bank of Spain is the most peculiar, appearing to be the least constrained by EMS, with its own output gap affecting short term interest rates throughout the period considered.

¹⁶ Even though we use data from January 1985 to December 1998 in the estimation, the first five years of monthly observations are used to initialise the estimation procedure (see appendix). Reported estimates, therefore, run only from January 1991 to December 1998.

France in April 1993), and there is also little or no evidence of serial autocorrelation and/or heteroscedasticity.¹⁷

In light of the arguably strong lag length assumption made in the specification of the econometric model, this result may be surprising and warrants some explanation. Note first that, when we reestimate (4) without exchange rate volatility and restricting $B(L)$ in (1) as done by Dornbusch and others (1998), we find residuals very much like theirs with several large outliers at about the same dates (result not reported).¹⁸ This confirms that adding exchange rate volatility and allowing for $B(L)$ to be unrestricted has helped to obtain better residuals, and thus ‘cleaner’ monetary policy shocks. Secondly, as noted in footnote 13, the inspection of lagged interest rate coefficients reveals the presence of relatively high persistence (i.e., a relatively strong preference for and interest rate smoothing) in all reaction functions. In the presence of unrestricted lagged interdependence among highly persistent variables it is not that surprising to find that one lag is sufficient to deliver well-behaved residuals. This in turn explains why we find very similar residuals when we experiment with a higher number of lags (up to six) only for Germany (results not reported).

The estimated residuals of (4) can be used to compare across countries the transmission mechanism of *country-specific* monetary shocks. These shocks reflect each central bank’s preference over the set of possible monetary policy objectives, as represented by the reaction functions specified and estimated. However, a key feature of EMU is that individual members’ preferences and reaction functions have been substituted by, or aggregated into, those of the ECB and its policymaking bodies. Further, it is possible to show that heterogeneous preferences over the objectives of monetary policy, as measured by different reaction functions, are sufficient to induce a heterogeneous response to monetary policy disturbances, as formally shown by Dornbusch and others (1998) and Clements, Kontolemis, and Levy (2001). As noted by Guiso and others (2000), in order to approximate the conditions prevailing under the EMU as closely as possible, one would also like to investigate the response of these economies to *common* shocks—that is, shocks that reflect the aggregation of countries’ preferences over the possible objectives of monetary policy. As a matter of fact, however, and as we shall see in the next section, we do find evidence of *homogeneity* even before controlling for this potential source of heterogeneity. For this reason, in this paper, we do not analyse the effects of common monetary policy shock, even though such a shock could be easily identified with the residual of the German reaction function in our econometric model along the lines pursued by Sala (2001) and Clements, Kontolemis, and Levy (2001).

¹⁷ This outlier coincides with the beginning of an aggressive, but short-lived, reduction of French official interest rates in the midst of the financial turbulence following the 1992 ERM crisis, not captured by the volatility variable. See Kenen (1995, page 154) on this episode.

¹⁸ All results not reported in this and the next two sections of the paper are available on request.

IV. The Impact of Monetary Policy on Economic Activity

In this section we report parameter estimates of the impact of monetary policy on economic activity in *individual* countries, and test statistics on the heterogeneity across countries and stability over time of these effects, derived from the estimation of the system of output equations (9). The next section will discuss our estimates of the *European-wide* effects of monetary policy. The results cover the period 1994-1998, from after the EMR crisis and the announcement of the Maastricht treaty to the EMU launch.¹⁹

We present two sets of estimation and testing results in this section: one based on the estimation of (9) specified without parameter time variation to compare our results with those previously found in the literature, and one based on (9) estimated with time-varying parameters. To facilitate the interpretation of the results, the time-varying specification actually estimated allows the parameter vector β_t to change only yearly, while we use monthly data (see appendix). In fact, we do not want to isolate changes at monthly frequency as the type of behavioural change we are interested in, presumably induced by anticipation of and preparation to the EMU, is likely to have taken place over time rather slowly. Hence, some time aggregation in estimating the parameters of the transmission mechanism of monetary policy may be desirable per se.

A. Are There Cross-Country Differences?

Table 1 and 2 provide an answer to this question which is comparable to those provided in other studies of the transmission mechanism of European monetary policy, while Table 3 reports such a comparison. The robustness of the results presented to a time varying specification of the econometric model is then analysed in the next subsection.

For all countries considered, Table 1 reports the mean and the median (which may be interpreted as classical point estimates), and the inter quartile range (which may be interpreted as a classical 50 percent confidence interval) of the posterior distribution of the coefficient of selected lags of \hat{u}_{it} , and the cumulative impact of this variable on output growth after 12 and 24 months (which we denote ‘cumul 12’ and ‘cumul 24’, respectively). As we can see, there are some cross-country differences in the impact of monetary policy at particular lags, but apparently not much quantitative difference with respect to the cumulative impact after 24 months as far as Germany, France, and Italy are concerned. In Spain, however, the effects of monetary shocks on output growth seem different in terms not only of their timing but also of their cumulative impact after 24 months, which is lower than in other countries. The

¹⁹ The series of estimated monetary policy shocks reported in Figure 1 run from January 1991 to December 1998, but we include 24 lags of this variable in the system and we need an additional year of monthly observations to initialize estimation (see appendix).

difference in the transmission mechanism of monetary policy between Spain and the other countries analysed is also statistically significant, as indicated by formal homogeneity tests.

Table 2 reports a set of Kolmogorov-Smirnov statistics (henceforth, KS) on the distance between the posterior distribution of q and q_1 under the corresponding null hypothesis.²⁰ As we can see, testing the null hypothesis of equality of all parameters of the transmission mechanism, either among all countries considered or through pair-wise comparisons (see the column of p-values under ‘all lags’ in Table 2), we reject the null decisively. This points to the existence of statistically significant differences in the *timing* of the effects of monetary policy among all countries considered. Running the same test on the cumulative impact of monetary policy after 12 and 24 months (see the corresponding columns of p-values in Table 2) between each pair of countries considered, however, we find that the difference among the four countries analysed is mainly due to Spain. Thus, in the case of Spain, when the model is estimated without time-variation, there appears to be a difference in terms of both *timing* and *magnitude* of the effects of monetary policy.

A direct comparison of our results with those obtained in other studies is difficult because of different econometric specifications, estimation methods, sample periods, type of shock, etc. Nonetheless, Table 3 attempts to do this, to the extent possible, by contrasting the ranking implied by some of the studies surveyed by Guiso and others (2000) and some more recent studies with that implied by our estimates.²¹ As is evident from the first part of Table 3, our estimates of the impact of monetary policy after 12 months produce a ranking that is consistent with that implied by most studies surveyed by Guiso and others—namely, Barron, Coudert, and Mojon (1996), Gerlach and Smets (1995, variant 1), Ehrmann (1998), and Dedola and Lippi (1999). Interestingly, comparing our results with those obtained in more recent studies, which attempt to control for also heterogeneity in the reaction function of different central banks, we continue to find broad consistency. As we can see from the second part of Table 3, the ranking implied by our estimate of the cumulative impact of a monetary policy shock after 24 months (the most comparable measure with these studies) is essentially the same as that implied by Sala (2001) and Ortega and Alberola (2000). This suggests that the difference between Spain and the other countries analysed is unlikely to be due to different preferences of the Bank of Spain over the objectives of monetary policy.²²

²⁰ Recall from section 2.2.3 that a posterior distribution of q_1 far apart from that of q can be interpreted as evidence against the null of equality of the relevant parameters of interest.

²¹ A comparison with the point estimates of Dornbusch and others (1998) based on a comparable specification is reported by Ciccarelli and Rebucci (2001), showing that none of the estimates is far away from those reported by Dornbusch and others (1998). This gives us confidence that the results reported here are not systematically distorted by any feature of the empirical framework used.

²² Ortega and Alberola explain the different response of Spain to a *common* monetary shock with a different sensitiveness to the wealth effect of interest rate changes compared to Germany, France, and Italy.

In summary, this first set of results appears broadly consistent with those found in previous studies. The results reported point to the existence of some degree of heterogeneity in the transmission mechanism of monetary policy across the countries considered, even though in the case of Germany, France, and Italy, according to our results, these differences seem more a matter of *timing* rather than *magnitude* of the cumulative impact of monetary policy. The differences in the cumulative impact of monetary policy on economic activity after 24 months in these three countries, in fact, are quantitatively small and statistically insignificant. Instead, in the case of Spain, the difference appears quantitatively sizable and statistically significant. It is also unrelated to differences in central banks' preferences over the objectives of monetary policy, judging by a comparison with the studies that have controlled explicitly for this potential source of heterogeneity.

B. Are These Differences Changing Over Time?

To address this question, first we reestimate the system of output equations (9), allowing for parameter variation over time, and test the null hypothesis that the posterior variance of the third stage of the hierarchy (9-11) is zero, i.e., the hypothesis that $\phi_1 = 0$, and thus that the hyperparameter tightening the time variation of the coefficients describing the transmission mechanism of monetary policy (β_{it}^1) equals zero.

This is done by using the test statistic (12) explained in section 2.2.3. As already noted, if the posterior distribution of ϕ_1 is less concentrated on values close to zero than the prior distribution, then we can reject the null of overall parameter stability, and thus reject a time-invariant specification of (9). The actual value of z in (12) is 0.47 for $\xi = 0.03$, and 1.84 for $\xi = 0.05$.²³ Small values of z for arbitrarily small values of ξ imply that the posterior distribution of ϕ_1 is located farther away from zero than the prior distribution, and thus provide evidence in favour of a time-varying specification. This, in turn, suggests that the transmission mechanism of monetary policy has changed over time during this period. This result may be appreciated also by direct inspection of Table 4 in which the mean and the inter-quartile range of the posterior distribution of the coefficient of selected lags of \hat{u}_{it} , as well as the cumulative impact of this variable on output growth after 12 and 24 months, is reported for all countries considered over the period 1994-1998.

Once it is established that the transmission mechanism of monetary policy has changed over time during this period, we check whether its overall degree of heterogeneity across countries has also changed in the run up to EMU, either decreasing or increasing. This is done by running a battery of KS statistics, one for each yearly subperiod from 1994 to 1998, on the posterior distributions of q and q_1 under the null hypothesis that the posterior mean of all coefficients of the transmission mechanism of monetary policy is the same across

²³ Values for ξ have been chosen to be arbitrarily small, as in Chib and Greenberg (1995).

all four countries considered. Table 5 reports the results (see the statistic ‘all lags’ which tests the same hypothesis as ‘joint’ ‘all lags’ in Table 2), showing that there is some weak evidence of a decreasing distance between the benchmark distribution and the posterior one, as measured by a decreasing value of the KS statistic. However, the overall picture is one of neither decreasing nor increasing heterogeneity but rather simply persistence. In fact, the p-values of these statistics are zero throughout the period considered. This suggests that the bulk of the change in the transmission mechanism of monetary policy during the period considered is due to changes in the common component (θ_t), while our prior assumptions allow also for time changes due to country specific factors (see section 2.2.1). This latter result is supported by a direct inspection of the posterior distribution of the parameters of the transmission mechanism at selected lags reported in Table 4: for instance, the location of the posterior distribution of the 18th lag has shifted from a clearly negative value in 1994-1995 to a value not significantly different from zero in 1997-1998 in *all* countries considered. As a consequence, the cross-country difference in the location of the posterior distribution of the parameters remains broadly constant over time. But this is not the end of the story.

As we can see from Table 5, in fact, if the model is estimated with time-varying parameters, we do accept the null hypothesis of equality of the cumulative impact of monetary policy after 12 and 24 months for all countries considered, including Spain, whereas these two hypothesis were decisively rejected by the data if tested on the time-invariant model (compare p-values of the statistics ‘joint’ ‘cumul 12’ and ‘cumul 24’ in Table 2 and 5). Further, the p-values of the statistic for the null hypothesis of equality of the cumulative impact after 24 months fluctuates slightly over time, but remains always greater than 10 percent (see statistic ‘cumul 24’ in Table 5). This evidence confirms that the heterogeneity in the transmission mechanism of monetary policy across all these countries, Spain included, was mainly a matter of timing of the effects of monetary policy rather than magnitude of their cumulative impact.

Note finally from Table 4 that the cumulative impact of monetary policy after 12 months has increased slightly over time while the impact after two years has decreased slightly in all countries considered. This suggests that the length of the transmission mechanisms might have become shorter in the second half of the 1990s in all countries considered, arguably as a result of capital markets developments and gradually increasing labor market flexibility at the regional level.

In summary, this second set of estimation results show that the hypothesis of parameter stability over time is rejected by the data. The transmission mechanism of monetary policy seems to have changed in these countries in the second half of the 1990s, possibly becoming shorter mainly because of a gradual change in its common component. Its degree of heterogeneity, however, has neither increased nor decreased during this period. On the other hand, the null hypothesis of equality of the cumulative impact of monetary policy after one and two years among all countries considered cannot be rejected by the data if the econometric

model is estimated allowing for parameter variation over time, even though it was clearly rejected (mainly because of Spain) when tested over the same sample period but without allowing for parameter time variation. This evidence confirms that the existing differences in the transmission mechanism among all countries considered were more a matter of *timing* than *magnitude* of the cumulative impact of monetary policy.

C. Interpreting the Estimation Results

In this subsection we elaborate on the interpretation of the results presented so far before moving on to discuss the area-wide effect of monetary policy briefly and then conclude.

The estimation results based on the model specified with time-varying parameters are *qualitatively* different both from those found by other studies in the literature and from what we have seen earlier in this paper based on the model estimated without time-varying parameters, as they suggest that the transmission mechanism of monetary policy in the four European countries considered might be more homogeneous than previously thought.

We interpret these results as an indication that the transmission mechanism of monetary policy in these countries might have started to change much before the EMU launch and that, by the mid-1990s, had already become relatively homogeneous. This interpretation would require a strong acceleration in the rate of change of the transmission mechanism of monetary policy at the beginning of the 1990s, and thus a fast private sector behavioural response to the announcement of the new monetary regime (on the occasion of the signature of the Maastricht treaty on February 1992) and the observation of relatively marked changes in the monetary authorities' own behaviour thereafter. Such a response would be consistent with a standard rational expectations hypothesis of private sector behaviour and hence is plausible, at least in principle.

According to this interpretation, other studies might have found a greater degree of heterogeneity because they are based on the estimation of time-invariant models on data-generating processes that were actually changing over time because of progress with European economic, monetary, and financial integration—a process that started back in 1958 with the Treaty of Rome, joined later by Spain in 1986. For instance, Dornbusch and others (1998), whose study is the most comparable with ours, estimate the system of output equations over the period July 1987/July 1996. It is evident that, if the heterogeneity of the transmission mechanism of monetary policy was relatively high at the beginning and relatively low at the end of this period, on average, one could find that heterogeneity is present but is not quantitatively sizable and/or statistically significant, as essentially both our and their studies find in a model estimated without time-varying parameters.

Indeed, it is possible that our results are qualitatively different from those reported in other studies also because of other features of the ‘experiment’ conducted in this paper. Among possible explanations (different specification, estimation method, sample period, type of shock, etc.), however, the specification of the econometric model used is the most plausible because none of the studies whose results are summarized in Table 3 allows for parameter variation over time. In any case, model specification is the only explanation for the different results found in this paper by estimating the model *with* and *without* time-varying parameters.

V. The Euro Area Impact of Monetary Policy

The evidence presented so far suggests that the effects of monetary policy on economic activity in these European countries might differ in terms of their timing, though not much in terms of their cumulative effects. Nonetheless, because of these timing differences, a study of the European-wide effects of monetary policy based on averages of country-specific time series along the lines of Tristani and Monticelli (1999) may be potentially biased. Moreover, we have seen that, in the specific case of Spain, a time-invariant specification of the econometric model gives rise to more heterogeneous estimates of the parameters of interest than what was found by allowing the parameters to change over time. Thus, parameter instability may further complicate the dynamic analysis of models based on area-wide averages of individual countries’ time series.

Within the empirical framework used in this study, the European-wide effects of monetary policy are measured by the posterior distribution of θ_t , the cross-sectional mean or common component of β_{it} . The mean, the median, and the inter-quartile range of the posterior distribution of the common component of the coefficient of selected lags of \hat{u}_{it} , and their evolution over time for each yearly subperiods from 1994 to 1998, can be appreciated from Table 6: monetary policy shocks appear to have had a system-wide effect peaking between 12 and 18 months in the mid-1990s. Toward the end of the 1990s, instead, they seem to have peaked earlier, between six and nine months. This evidence is consistent with what we saw in the previous subsection and confirms that the European-wide transmission mechanism of monetary policy may have become shorter in the second part of the 1990s.

VI. Conclusions

In this paper we have studied empirically the transmission mechanism of monetary policy in the four largest European countries that are currently members of the EMU by using dynamic heterogeneous models estimated in a Bayesian fashion with EMS data. The ‘experiment’ documented in this paper shares several features of an ‘ideal’ one: the model specification is the same for all countries considered; no strong a priori restrictions are imposed on the behaviour of the central banks studied; intra-Europe exchange rate movements

that have disappeared under EMU are controlled for; regional interdependence, through which monetary policy in part operates, is also allowed for; and, most importantly, the parameters of the transmission mechanism of monetary policy are allowed to change over time. Introducing parameter variation over time has proven to be crucial for our results. By allowing parameters to change over time, everything else being equal, we have found relatively more homogeneity in the transmission mechanism of European monetary policy than what was previously found in the literature and what we have found in this paper by restricting parameters to be constant over time.

More specifically, the empirical results based on the time-varying model show that (i) there were differences in the timing of the effects of monetary policy across these European countries in the second part of the 1990s, and (ii) these differences had not disappeared by the time EMU was launched, even though (iii) the parameters of the transmission mechanism were changing over time during this period. This evidence suggests that these changes were mainly due to shifts in the common component of the transmission mechanism, which can be interpreted as the area-wide (or regional) impact of monetary policy in our framework. According to our results, (iv) the area-wide impact of monetary policy became faster in the second part of the 1990s and, by the time EMU was launched, its effects were peaking between six and nine months. At the same time, we have shown also that (v) the cumulative impact of monetary policy after two years was relatively homogenous across these countries in the second half of the 1990s.

These results are partially consistent with those previously found in the literature to the extent that they point to some degree of heterogeneity in the transmission mechanism of European monetary policy. Unlike the results found in previous studies, however, they suggest that these differences are more a matter of timing rather than magnitude of the cumulative impact monetary policy. To interpret these results, we conjectured that the transmission mechanism of European monetary policy had already become relatively homogeneous by the mid-1990s.

Figure 1. Monetary Policy Shocks (1991-1998)

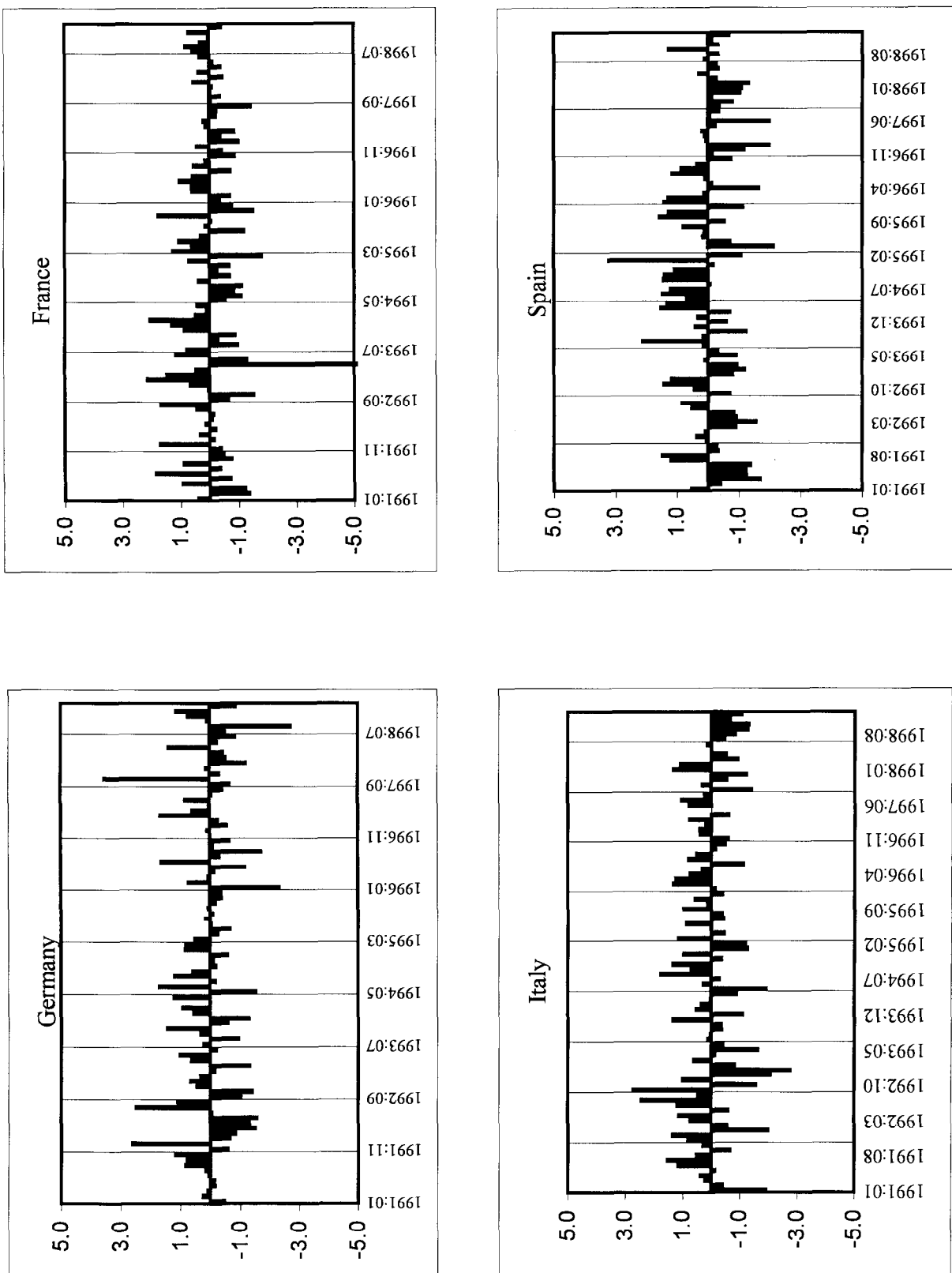


Table 1. Comparing the Impact of Monetary Policy Shocks Across Countries (1994-1998)
(Selected lags and cumulative impact)

		Germany	France	Italy	Spain
Lag (6th month)	1st Quartile	-0.17	-0.04	-0.07	-0.07
	Mean	-0.06	0.06	0.05	0.04
	Median	-0.06	0.06	0.05	0.03
	3rd Quartile	0.06	0.16	0.17	0.14
Lag (12th month)	1st Quartile	0.06	-0.04	0.10	0.09
	Mean	0.17	0.06	0.22	0.18
	Median	0.17	0.06	0.22	0.19
	3rd Quartile	0.27	0.16	0.33	0.28
Lag (18th month)	1st Quartile	-0.33	-0.21	-0.38	-0.30
	Mean	-0.22	-0.12	-0.26	-0.20
	Median	-0.22	-0.12	-0.26	-0.20
	3rd Quartile	-0.11	-0.02	-0.14	-0.10
Lag (24th month)	1st Quartile	-0.25	-0.14	-0.29	-0.22
	Mean	-0.14	-0.05	-0.17	-0.13
	Median	-0.14	-0.05	-0.17	-0.13
	3rd Quartile	-0.03	0.05	-0.06	-0.03
Cumulative Impact (12-month)	1st Quartile	-0.88	-0.84	-0.72	-0.54
	Mean	-0.41	-0.37	-0.26	-0.15
	Median	-0.42	-0.37	-0.27	-0.15
	3rd Quartile	0.08	0.11	0.20	0.25
Cumulative Impact (24-month)	1st Quartile	-2.11	-2.03	-2.18	-1.38
	Mean	-1.41	-1.35	-1.51	-0.90
	Median	-1.40	-1.35	-1.50	-0.89
	3rd Quartile	-0.70	-0.64	-0.84	-0.39

Table 2. Testing Homogeneity Across Countries (1994-1998) 1/

	All lags	Cumulative Impact (12 months)	Cumulative Impact (24 months)
All countries (Joint)	0.5020 (0.0000)	0.0788 (0.0001)	0.1530 (0.0000)
Germany versus France	0.3370 (0.0000)	0.0152 (0.7280)	0.0138 (0.8327)
Germany versus Italy	0.2528 (0.0000)	0.0370 (0.0079)	0.0172 (0.5781)
Germany versus Spain	0.3058 (0.0000)	0.0600 (0.0000)	0.1045 (0.0000)
France versus Italy	0.3223 (0.0000)	0.0192 (0.4372)	0.0198 (0.4051)
France versus Spain	0.3162 (0.0000)	0.0767 (0.0000)	0.1135 (0.0000)
Italy versus Spain	0.2388 (0.0000)	0.0342 (0.0175)	0.1700 (0.0000)

1/ The null hypothesis is $F(q) = F(q1)$; the numbers reported are Kolmogorov-Smirnov statistics;
P-values in brackets.

Table 3 - Comparing the Estimated Impact of Monetary Policy Shocks with Other Studies

	Strength of response	Germany	France	Italy	Spain
This study 1/ Ramaswamy and Sloek (1998) Barron, Coudert, and Mojon (1996) Gerlach and Smets (1995), variant 1 Ehrmann (1998) Dedola and Lippi (1999)		Effects after one year			
	S<I<F<G	-0.41	-0.37	-0.26	-0.15
	F<I<G	-0.60	-0.40	-0.50	na
	I<F<G	-0.60	-0.40	-0.30	na
	F=I<G	-0.30	-0.20	-0.20	na
	I<F<G	-0.90	-0.50	-0.10	na
	I<F<G	-2.20	-1.40	-1.10	na
This study 2/ Sala (2001) Ortega and Alberola (2000), variant 1 Clements, Kontolemis, and Levy (2001)		Effects after two years			
	I<F<G<S	-0.66	-0.66	-0.48	-1.24
	I<F<S<G	-0.60	-0.30	-0.16	-0.60
	F<I<G<S	-0.40	-0.26	-0.32	-0.48
	S<G<I<F	na	na	na	na

1/ Posterior mean of "Cumulative impact, 12-month" in Table 1.

2/ Posterior mean of "Cumulative impact, 24-month" in Table 1.

Table 4. Comparing the Impact of Monetary Policy Shocks Across Countries and Over Time
(Selected lags and cumulative impact)

		Lag 6th month					Lag 12th month				
		1994	1995	1996	1997	1998	1994	1995	1996	1997	1998
Germany	1st Quartile	-0.16	-0.17	-0.17	-0.39	-0.22	-0.39	-0.32	-0.35	-0.19	-0.26
	Mean	0.00	-0.02	-0.02	-0.22	-0.03	-0.21	-0.14	-0.18	-0.01	-0.07
	3rd Quartile	0.16	0.14	0.13	-0.05	0.16	-0.04	0.02	-0.03	0.16	0.12
France	1st Quartile	-0.12	-0.22	-0.12	-0.40	-0.20	-0.45	-0.31	-0.36	-0.26	-0.24
	Mean	0.04	-0.07	0.03	-0.23	0.00	-0.29	-0.13	-0.21	-0.09	-0.05
	3rd Quartile	0.21	0.07	0.17	-0.06	0.20	-0.13	0.02	-0.07	0.08	0.14
Italy	1st Quartile	-0.12	-0.20	-0.18	-0.36	-0.20	-0.41	-0.28	-0.30	-0.23	-0.20
	Mean	0.04	-0.05	-0.01	-0.18	-0.01	-0.23	-0.10	-0.13	-0.05	0.00
	3rd Quartile	0.20	0.11	0.15	0.00	0.20	-0.06	0.06	0.02	0.12	0.19
Spain	1st Quartile	-0.09	-0.18	-0.21	-0.37	-0.20	-0.43	-0.29	-0.31	-0.24	-0.22
	Mean	0.07	-0.04	-0.05	-0.20	0.01	-0.25	-0.11	-0.16	-0.07	-0.03
	3rd Quartile	0.22	0.11	0.09	-0.03	0.20	-0.07	0.05	-0.02	0.09	0.16
		Lag 18th month					Lag 24th month				
Germany	1st Quartile	-0.44	-0.38	-0.32	-0.23	-0.27	-0.33	-0.10	-0.09	-0.14	-0.26
	Mean	-0.28	-0.23	-0.15	-0.05	-0.09	-0.16	0.05	0.06	0.01	-0.09
	3rd Quartile	-0.12	-0.09	0.01	0.12	0.10	0.01	0.19	0.22	0.16	0.08
France	1st Quartile	-0.40	-0.37	-0.28	-0.12	-0.25	-0.32	-0.04	-0.08	-0.17	-0.14
	Mean	-0.25	-0.22	-0.13	0.04	-0.07	-0.15	0.09	0.08	-0.03	0.01
	3rd Quartile	-0.11	-0.08	0.03	0.20	0.12	0.02	0.22	0.24	0.10	0.18
Italy	1st Quartile	-0.47	-0.36	-0.28	-0.23	-0.29	-0.30	-0.12	-0.15	-0.18	-0.21
	Mean	-0.32	-0.20	-0.12	-0.06	-0.09	-0.13	0.03	0.02	-0.03	-0.04
	3rd Quartile	-0.17	-0.05	0.04	0.12	0.10	0.03	0.18	0.19	0.12	0.13
Spain	1st Quartile	-0.46	-0.37	-0.31	-0.18	-0.29	-0.31	-0.16	-0.08	-0.14	-0.19
	Mean	-0.30	-0.21	-0.15	-0.01	-0.11	-0.14	-0.01	0.08	0.00	-0.03
	3rd Quartile	-0.15	-0.06	0.00	0.16	0.08	0.02	0.13	0.25	0.13	0.14
		Cumulative impact 12-month					Cumulative impact 24-month				
Germany	1st Quartile	-2.14	-2.36	-2.30	-2.46	-2.00	-4.68	-4.74	-4.30	-3.88	-2.91
	Mean	-1.14	-1.46	-1.32	-1.52	-0.96	-2.93	-2.96	-2.59	-1.98	-0.97
	3rd Quartile	-0.28	-0.61	-0.54	-0.71	-0.02	-1.49	-1.57	-1.33	-0.64	0.56
France	1st Quartile	-2.15	-2.45	-2.36	-2.65	-1.96	-4.83	-4.81	-4.35	-4.02	-2.85
	Mean	-1.16	-1.56	-1.37	-1.72	-0.98	-3.05	-3.02	-2.61	-2.12	-0.98
	3rd Quartile	-0.28	-0.73	-0.58	-0.91	-0.07	-1.62	-1.72	-1.38	-0.74	0.53
Italy	1st Quartile	-2.16	-2.42	-2.31	-2.41	-1.94	-4.80	-4.75	-4.32	-3.83	-2.78
	Mean	-1.17	-1.49	-1.32	-1.49	-0.89	-3.04	-2.95	-2.56	-1.95	-0.87
	3rd Quartile	-0.30	-0.65	-0.56	-0.71	0.03	-1.64	-1.61	-1.31	-0.63	0.66
Spain	1st Quartile	-2.12	-2.47	-2.31	-2.36	-1.94	-4.73	-4.83	-4.30	-3.72	-2.76
	Mean	-1.12	-1.53	-1.32	-1.43	-0.90	-2.98	-3.06	-2.58	-1.79	-0.83
	3rd Quartile	-0.29	-0.70	-0.54	-0.66	0.01	-1.63	-1.69	-1.35	-0.44	0.69

Table 5. Testing Homogeneity Across Countries and Over Time 1/
(All countries jointly)

	1994	1995	1996	1997	1998
All lags	0.3615 (0.0000)	0.3195 (0.0000)	0.3117 (0.0000)	0.3027 (0.0000)	0.2533 (0.0000)
Cumulative impact (12-month)	0.0153 (0.7280)	0.019 (0.4538)	0.0158 (0.6907)	0.0543 (0.0000)	0.0215 (0.3042)
Cumulative impact (24-month)	0.0158 (0.6907)	0.0155 (0.7094)	0.0183 (0.5055)	0.0268 (0.1100)	0.0112 (0.9565)

1/ The null hypothesis is $F(q) = F(q1)$; the numbers reported are Kolmogorov-Smirnov statistics;
P-values in brackets.

Table 6. The Euro Area Impact of Monetary Policy Shocks
(Selected lags)

		1994	1995	1996	1997	1998
Lag (6th month)	1st Quartile	-0.11	-0.17	-0.16	-0.35	-0.20
	Mean	0.03	-0.04	-0.02	-0.19	-0.02
	Median	0.03	-0.04	-0.04	-0.20	-0.02
	3rd Quartile	0.17	0.09	0.10	-0.04	0.17
Lag (9th month)	1st Quartile	-0.46	-0.55	-0.41	-0.62	-0.42
	Mean	-0.30	-0.41	-0.25	-0.46	-0.24
	Median	-0.31	-0.41	-0.25	-0.45	-0.23
	3rd Quartile	-0.14	-0.26	-0.10	-0.30	-0.05
Lag (12th month)	1st Quartile	-0.40	-0.29	-0.30	-0.22	-0.21
	Mean	-0.24	-0.13	-0.16	-0.06	-0.04
	Median	-0.25	-0.14	-0.17	-0.07	-0.04
	3rd Quartile	-0.09	0.01	-0.03	0.09	0.13
Lag (18th month)	1st Quartile	-0.42	-0.35	-0.28	-0.18	-0.25
	Mean	-0.28	-0.22	-0.14	-0.03	-0.08
	Median	-0.29	-0.22	-0.14	-0.03	-0.08
	3rd Quartile	-0.15	-0.09	0.00	0.13	0.09
Lag (24th month)	1st Quartile	-0.28	-0.10	-0.08	-0.14	-0.18
	Mean	-0.14	0.03	0.06	-0.01	-0.03
	Median	-0.14	0.02	0.06	-0.02	-0.03
	3rd Quartile	0.01	0.15	0.20	0.11	0.12

Appendix I. Estimation Procedures

In this appendix we present in more details the Bayesian procedures used for the estimation of the systems of reaction functions (4) and output equations (9).

A. Reaction Functions

The pdf of the data for each block j of (5), $L(R_{j1}, \dots, R_{jT} \mid Z_{jt}, \delta_{jt}, \Sigma_{jj})$, conditional on the exogenous variables, the initial observations of R_{jt} , and the parameters of the model (δ_{jt} and Σ_{jj}) is proportional (\propto) to:

$$|\Sigma_{jj}|^{-T/2} \exp \left[-\frac{1}{2} \sum_t (R_{jt} - Z_{jt}\delta_{jt})' \Sigma_{jj}^{-1} (R_{jt} - Z_{jt}\delta_{jt}) \right]. \quad (16)$$

The prior assumptions on the model's parameters generalize those introduced by Zellner (1971, Chapter 8) to take into account the presence of time-varying coefficients: a time-varying Minnesota-type of prior (Doan, Litterman, and Sims, 1984) for the slope coefficients (δ_{jt}) is combined with a diffuse (i.e., non-informative) prior on the variance-covariance matrix of the residuals, Σ_{jj} , assuming prior independence. Thus:

$$p(\delta_{jt}, \Sigma_{jj}) = p(\delta_{jt}) p(\Sigma_{jj}),$$

where

$$\begin{aligned} p(\Sigma_{jj}) &\propto |\Sigma_{jj}|^{-(G_j+1)/2}, \\ \delta_{jt} &= P_j \delta_{jt-1} + (I - P_j) \bar{\delta}_j + \eta_{jt} \quad \eta_{jt} \sim N(0, \Phi_j). \end{aligned} \quad (17)$$

Here, P_j is a $(G_j k_j \times G_j k_j)$ matrix governing the law of motion of δ_{jt} , $\bar{\delta}_j$ is the unconditional mean of δ_{jt} , Φ_j governs the time variation of δ_{jt} , and η_{jt} is assumed to be independent from v_{jt} .

The posterior distribution of Σ_{jj}^{-1} conditional on the entire history of δ_{jt} for $t = 0, \dots, T$ (denoted $\{\delta_{jt}\}$) and R_{j0} is easily obtained combining (16) with $p(\Sigma_{jj})$ as the following Wishart distribution with T degrees of freedom and scale matrix S :

$$\Sigma_{jj}^{-1} \mid \{\delta_{jt}\}, R_{j0} \sim W(T, S), \quad (18)$$

where

$$S = \sum_t [(R_{jt} - Z_{jt}\delta_{jt})(R_{jt} - Z_{jt}\delta_{jt})']^{-1}.$$

The joint posterior distribution of $\{\delta_{jt}\}$ *conditional* on Σ_{jj} is obtained in two steps as shown by Chib and Greenberg (1995, pp. 349-350). First, we initialize $\{\delta_{jt}\} \mid \Sigma_{jj}$ by Kalman Filter and its output $\{\hat{\delta}_{jt}, \hat{\Omega}_{jt|t}, F_t\}$ is saved for each t :

$$\begin{aligned}\hat{\delta}_{jt|t} &= \hat{\delta}_{jt|t-1} + \hat{\Omega}_{jt|t-1} Z'_{jt} F_t \left(R_{jt} - Z_{jt} \hat{\delta}_{jt|t-1} \right), \\ \hat{\Omega}_{jt|t} &= \hat{\Omega}_{jt|t-1} - \hat{\Omega}_{jt|t-1} Z'_{jt} F_t Z_{jt} \hat{\Omega}_{jt|t-1}, \\ F_t &= \left(Z_{jt} \hat{\Omega}_{jt|t-1} Z'_{jt} + \Sigma_{jj} \right)^{-1},\end{aligned}\tag{19}$$

where

$$\begin{aligned}\hat{\delta}_{jt|t-1} &= P_j \hat{\delta}_{j,t-1|t-1} + (I - P_j) \bar{\delta}_j, \\ \hat{\Omega}_{jt|t-1} &= P_j \hat{\Omega}_{j,t-1|t-1} P'_j + \Phi_j.\end{aligned}$$

Second, $p(\{\delta_{jt}\} \mid \Sigma_{jj})$ is sampled in reverse time order from

$$\begin{aligned}\delta_{jT} &\sim N\left(\hat{\delta}_{jT|T}, \hat{\Omega}_{jT|T}\right) \\ \delta_{jT-1} &\sim N\left(\hat{\delta}_{jT-1}, \hat{\Omega}_{jT-1}\right) \\ &\vdots \\ \delta_{j0} &\sim N\left(\hat{\delta}_{j0}, \hat{\Omega}_{j0}\right)\end{aligned}\tag{20}$$

where

$$\begin{aligned}\hat{\delta}_{jt} &= \hat{\delta}_{jt|t} + M_t \left(\delta_{jt+1} - \hat{\delta}_{jt|t} \right) \\ \hat{\Omega}_{jt} &= \hat{\Omega}_{jt|t} - M_t \hat{\Omega}_{jt+1|t} M'_t,\end{aligned}$$

with $M_t = \hat{\Omega}_{jt|t} \hat{\Omega}_{jt+1|t}^{-1}$.

Given (18) and (20), the marginal posterior distributions of Σ_{jj}^{-1} and $\{\delta_{jt}\}$ can be obtained by means of Gibbs sampling (i.e., numerical integration) drawing alternately from (18) and (20) given initial values for P_j , Φ_j , $\hat{\Omega}_{j0}$, and $\hat{\delta}_{j0}$.²⁴ To make operational the updating scheme in (18)-(20), therefore, we need initial values for P_j , Φ_j , $\hat{\Omega}_{j0}$, and $\hat{\delta}_{j0}$.

Following Litterman (1986), we define the matrices P_j , Φ_j , $\hat{\Omega}_{j0}$, $\hat{\delta}_{j0}$ in terms of six hyperparameters and then maximise the likelihood of the data as a function of this smaller set of parameters to obtain numerical values that are fed into the Gibbs sampler. More specifically, for each equation of block j , we assume that each $(k_j \times 1)$ vector δ_{j0}^g (the g^{th} element of $\hat{\delta}_{j0}$) depends only on one hyperparameter ($\pi_{1,g}^j$) so that

²⁴ On the Gibbs sampler see Gelfand et al. (1990) among others.

$$\delta_{j0}^g = (0, \dots, 0, \pi_j^{1g}, 0, \dots, 0),$$

where π_j^{1g} is the prior mean of the coefficient of the lagged dependent variable in equation g of block j . The elements of $\hat{\delta}_{j0}$ are assumed to be mutually independent and independent from analogous components in other equations of the block j , so that $\hat{\Omega}_{j0}$ is diagonal. The diagonal elements of $\hat{\Omega}_{j0}$ are then defined so that, for each block j , the relative tightness of the prior of the coefficient of the lagged dependent variable, of other lagged endogenous variables and of deterministic and exogenous variables are respectively controlled by $\pi_j^{2g}, \pi_j^{3g}, \pi_j^{4g}$ (all scalar values). In practice, the prior variances of the parameters of equation g in block j are specified as follows:

$$Var(\delta_{jg}^0) = \begin{cases} (\pi_j^{2g}/l) & \text{for lagged dependent variables} \\ (\pi_j^{2g}\pi_j^{3g}/l)(\sigma_g/\sigma_i) & \text{for other lagged endogenous variables} \\ (\pi_j^{2g}\pi_j^{4g}\sigma_g) & \text{for exogenous and deterministic variables} \end{cases}$$

where l denotes the lag length, and σ_g/σ_i is a scaling factor which takes into account the range of variation of different variables. Hence, the overall tightness of the model's parameters is controlled by π^2 . If π^2 goes to infinity, the prior becomes diffuse. The tightness of the coefficients of the lagged dependent variable relative to that of other lagged endogenous variables in the equation is controlled by π_3 , and if $\pi_3 = 0$, the prior defines a set of univariate autoregressive processes of order p . Finally, π_4 controls the degree of uncertainty with respect to the coefficients of exogenous and deterministic variables.

The matrices P_j and Φ_j are defined as:

$$P_j = diag(P_{j1}, \dots, P_{jG_j})$$

$$\Phi_j = diag(\Phi_{j1}, \dots, \Phi_{jG_j}) \hat{\Omega}_{j0}$$

where $P_{jg} = diag(\pi_{5,g})$ are $(k_j \times k_j)$ matrices with $\pi_{5,g}$ controlling the coefficients of the law of motion of each element of δ_{jg} , and $\Phi_{jg} = diag(\pi_{6,g})$ are $(k_j \times k_j)$ matrices with $\pi_{6,g}$ controlling the amount of variance around these values actually introduced in the model, for $g = 1, \dots, G_j$. Finally, the model's parameters $(\delta_{jt}, \Sigma_{jj})$ are initialized with a classical SUR estimate of the entire model.²⁵

Given the values of the model's hyperparameters, (19) is run. Then the Gibbs sampler starts iterating, switching between (18) and (20), taking the estimated values of π_1, \dots, π_6 as given. The Gibbs sampler runs 5,000 times yielding 4,000 draws from the joint and marginal

²⁵ Estimated hyperparameters not reported but available on request. Note that as the first block of the model contains only one equation, (18) becomes an inverted gamma for $j = 1$ and the equation's parameters can be initialized by OLS.

posterior distributions of the parameters of interest after discarding the first 1,000 draws. All the numerical integrations and the statistics presented are based on these last 4,000 draws.

A. Output Equations

1. Time-varying Model

Denote annual output growth in month s of year τ with $y_{i,\tau}^s$. For each country i , $y_{i,\tau}^s$ is modelled as:

$$y_{i\tau}^s = X_{i\tau}^{s'} \beta_{i\tau} + \varepsilon_{i\tau}^s$$

$$i = 1, \dots, G; \quad \tau = 1, \dots, T_1; \quad s = 1, \dots, S.$$

In our sample, the number of years (T_1) is 6, the number of countries or endogenous variables (G) is 4, and the number of subperiods for each year (S) is 12. Hence, the total number of observations for each variable is $T = T_1 * S = 72$, while we have 30 regressors for each equation. As noted in the text, stacking countries by row, this system can be written as:

$$y_\tau^s = X_\tau^s \beta_\tau + \varepsilon_\tau^s$$

$$\tau = 1, \dots, T_1; \quad s = 1, \dots, S.$$

The likelihood of the data is proportional to:

$$|\Omega|^{-T/2} \exp \left\{ -\frac{1}{2} \sum_{\tau} \sum_s (y_\tau^s - X_\tau^s \beta_\tau)' \Omega^{-1} (y_\tau^s - X_\tau^s \beta_\tau) \right\}.$$

The prior assumptions are:

$$\begin{aligned} y_\tau^s &= X_\tau^s \beta_\tau + \varepsilon_\tau^s, & \varepsilon_\tau^s &\sim N_G(0, \Omega), \\ \beta_\tau &= M_o \theta_\tau + \zeta_\tau, & \zeta_\tau &\sim N_s(0, B_o), \\ \theta_\tau &= \theta_{\tau-1} + \eta_\tau, & \eta_\tau &\sim N_m(0, B_1). \end{aligned}$$

with

$$\begin{aligned} \Omega^{-1} &\sim W(\omega_o, \Theta), \\ M_o &= e_G \otimes I_k, \\ B_o &= I_G \otimes \Upsilon, & \Upsilon^{-1} &\sim W(\sigma_o, \Psi_o), \\ B_1 &= \text{diag}(\phi_1 I_{k_1}, \phi_2 I_{k-k_1}), \\ \phi_1 &\sim \text{diffuse}, \\ \phi_1 &\sim IG(\kappa_o/2, \xi_o/2). \end{aligned}$$

Here, $W(\omega_o, \Theta)$ denotes a Wishart distribution with ω_o degrees of freedom and scale matrix Θ , $IG(\kappa_o/2, \xi_o/2)$ is an inverted gamma distribution, e_G is a $(G \times 1)$ vector of ones, I_k denotes an identity matrix of dimension k , and k_1 is the number of monetary policy parameters. Note that ϕ_1 tightens the time variation of monetary policy parameters, while ϕ_2 tightens the time variation of other parameters.

The posterior densities of the parameters of interest are obtained by combining the likelihood of the data with the prior distributions above in the form of *conditional* posterior distributions, as before. Denote $y = (y_1, \dots, y_T)$ the data sample and $\psi = (\{\beta_\tau\}_\tau, \Omega, \{\theta_\tau\}_\tau, \Upsilon, \phi_1, \phi_2)$ the set of parameters of interest whose joint posterior distribution needs to be determined, and ψ without the parameter γ , with $\psi_{-\gamma}$. It can be shown (Chib and Greenberg, 1995, page 348-349) that the conditional posterior distributions of the parameters of interest are given by:

$$\begin{aligned}\beta_\tau &| \psi_{-\beta_\tau}, y \sim N(\hat{\beta}_\tau, V_\tau), \text{ for all } \tau \leq T_1; \\ \Omega^{-1} &| \psi_{-\Omega}, y \sim W(T + \omega_o, \Theta_T); \\ \Upsilon^{-1} &| \psi_{-\Upsilon}, y \sim W(GT_1 + \sigma_o, \Psi_{T_1}); \\ \phi_1 &| \psi_{-\phi_1}, y \sim IG\left(\frac{(k_1 T_1 + v_o)}{2}, \frac{\zeta_o + \Upsilon_\tau (\theta_\tau^1 - \theta_{\tau-1}^1)' (\theta_\tau^1 - \theta_{\tau-1}^1)}{2}\right); \\ \phi_2 &| \psi_{-\phi_2}, y \sim IG\left(\frac{T_1(k - k_1)}{2}, \frac{\Upsilon_\tau (\theta_\tau^2 - \theta_{t-1}^2)' (\theta_\tau^2 - \theta_{t-1}^2)}{2}\right);\end{aligned}$$

where

$$\begin{aligned}\hat{\beta}_\tau &= V_\tau \left(B_o^{-1} M_o \theta_\tau + \sum_s X_\tau^{s'} \Omega^{-1} y_\tau^s \right), \\ V_\tau &= \left(B_o^{-1} + \sum_s X_\tau^{s'} \Omega^{-1} X_\tau^s \right)^{-1}, \\ \Theta_T &= \left[\Theta^{-1} + \sum_\tau \sum_s (y_\tau^s - X_\tau^s \beta_\tau) (y_\tau^s - X_\tau^s \beta_\tau)' \right]^{-1}, \\ \Psi_{T_1} &= \left[\Psi_o^{-1} + \sum_\tau \sum_i (\beta_{i\tau} - \theta_\tau) (\beta_{i\tau} - \theta_\tau)' \right]^{-1},\end{aligned}\tag{21}$$

with θ_τ^1 and θ_τ^2 denoting the time-varying common component of the parameters of the transmission mechanism of monetary policy and other parameters, respectively. Instead, the

posterior distribution of $\{\theta_\tau\}_{\tau=0}^{T_1}$, conditional on the other parameters is obtained using an updating scheme as in (20) above.

All hyperparameters of the model $(\omega_o, \Theta, \sigma_o, \Psi_o, \kappa_o, \xi_o)$ are assumed to be known. We set $\omega_o = g + 1$, $\sigma_o = k + 1$, and $\Psi_o = \text{diag}(1.0)$, and initialize Θ with the variance-covariance matrix of a classical SUR estimate of (9). The parameters of the gamma distribution of ϕ_1 are $\kappa_o = 6$ and $\xi_o = 1$, implying that the prior mean and the standard deviation of ϕ_1 are 0.25 and 0.25, respectively. To initialize the Gibbs sampler we also set $\phi_1 = \phi_2 = 0.5$, $\Omega = I_g$, and $\Upsilon = I_k$, while β_τ (for all $\tau = 1, \dots, T_1$) is initialized with the posterior mean of the parameters of the model estimated without time-variation (see below).

Given these values, the Gibbs sampler starts by generating $\{\theta_\tau\}_{\tau=0}^{T_1}$, and then continues with all the other parameters. The Gibbs sampler is run 5,000 times and yields 4,000 draws from the joint limiting posterior distribution, after discarding the first 1,000 draws as before.

2. Time-invariant Model

The model is also estimated by restricting the coefficients to be constant over time. In this case, we used the following hierarchy:

$$\begin{aligned} y_t &= X_t \beta + \varepsilon_t, & \varepsilon_t &\sim N_G(0, \Omega) \\ \beta &= M_o \theta + \zeta, & \zeta &\sim N_s(0, B_o) \\ \theta &= M_1 \mu + \eta, & \eta &\sim N_m(0, B_1) \end{aligned}$$

where now $t = 1, \dots, T$.

The likelihood now becomes proportional to:

$$|\Omega|^{-T/2} \exp \left\{ -\frac{1}{2} \sum_{t=1}^T (y_t - X_t \beta)' \Omega^{-1} (y_t - X_t \beta) \right\}.$$

All hyperparameters, including μ and B_1 , are assumed to be known as before. In particular, we set $B_1^{-1} = 0$, i.e., the last stage of the hierarchy is degenerate.

Using the same priors and notation as before, the conditional posterior distributions now become:

$$\beta \mid \psi_{-\beta}, y \sim N(\hat{\beta}, V_T);$$

$$\Omega^{-1} \mid \psi_{-\Omega}, y \sim W(\omega_o + T, R_T);$$

$$\theta \mid \psi_{-\theta}, y \sim N(\Delta_1 (B_1^{-1} M_1 \mu + M_o' B_o^{-1} \beta), \Delta_1);$$

$$\Upsilon^{-1} \mid \psi_{-\Upsilon}, y \sim W(\sigma_o + G, \Psi_G);$$

where

$$\hat{\beta} = V_T \left(B_o^{-1} M_o \theta + \sum_t X_t' \Omega^{-1} y_t \right), \quad V_T = \left(B_o^{-1} + \sum_t X_t' \Omega^{-1} X_t \right)^{-1},$$

$$R_T = \left[R_o^{-1} + \sum_{t=1}^T (y_t - X_t \beta) (y_t - X_t \beta)' \right]^{-1},$$

$$\Delta_1 = (B_1^{-1} + M_o' B_o^{-1} M_o)^{-1},$$

$$\Psi_G = \left[\Psi_o^{-1} + \sum_{i=1}^G (\beta_i - \theta) (\beta_i - \theta)' \right]^{-1}.$$

The Gibbs sampler is then initialized and run as was done in the case of the time-varying model.

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