DISCUSSION PAPER SERIES

No. 4535
CEPR/EABCN No. 6/2004

HAS THE TRANSMISSION MECHANISM OF EUROPEAN MONETARY POLICY CHANGED IN THE RUN-UP TO EMU?

Matteo Ciccarelli and Alessandro Rebucci

INTERNATIONAL MACROECONOMICS

€ABCN
Euro Area Business Cycle Network
www.eabcn.org

Centre for Economic Policy Research

www.cepr.org

Available online at: www.cepr.org/pubs/dps/DP4535.asp

ISSN 0265-8003

HAS THE TRANSMISSION MECHANISM OF EURO MONETARY POLICY CHANGED IN THE RUN-UP TO EMU?

Matteo Ciccarelli, European Central Bank (ECB)
Alessandro Rebucci, International Monetary Fund (IMF)

Discussion Paper No. 4535
September 2004

Centre for Economic Policy Research
90–98 Goswell Rd, London EC1V 7RR, UK
Tel: (44 20) 7878 2900, Fax: (44 20) 7878 2999
Email: cepr@cepr.org, Website: www.cepr.org

CEPR Discussion Paper No. 4535

September 2004

ABSTRACT

Has the Transmission Mechanism of Euro Monetary Policy Changed in the Run-Up to EMU?*

This Paper studies empirically the transmission mechanism of European monetary policy by means of time-varying, heterogenous coefficient models estimated in a numerical Bayesian fashion. Based on pre-EMU evidence from Germany, France, Italy, and Spain, we find that (i) the long-run cumulative impact on output of a common, homoskedastic monetary policy shock has decreased in all countries after 1991. These declines are statistically significant and accompanied by some changes in the conduct of monetary policy over the same period. At the same time, we also find that (ii) cross-country differences in the effects of the shock analysed have not decreased over time.

JEL Classification: C11, C33 and E52
Keywords: Bayesian estimation, European monetary policy, Gibbs sampling, time-varying coefficient model and transmission mechanism

Matteo Ciccarelli
European Central Bank
DG Research
Kaiserstrasse 29
D-60311 Frankfurt
GERMANY
Email: matteo.ciccarelli@ecb.int

For further Discussion Papers by this author see:
www.cepr.org/pubs/new-dps/dplist.asp?authorid=153879

Alessandro Rebucci
International Monetary Fund
700 19th Street NW
Washington DC 20431
USA
Tel: (1 202) 623 5873
Email: arebucci@imf.org

For further Discussion Papers by this author see:
www.cepr.org/pubs/new-dps/dplist.asp?authorid=135814

*This Paper is funded by the Euro Area Business Cycle Network (www.eabcn.org). This Network provides a forum for the better understanding of the euro area business cycle, linking academic researchers and researchers in central banks and other policy institutions involved in the empirical analysis of the euro area business cycle. We are grateful to Filippo Altissimo, Jordi Galí, Morten Ravn for useful comments, and to Fabio Canova and Chris Gilbert for discussions and comments on previous versions of this Paper (circulated under different titles and issued as Ente L. Einaudi Temi di Ricerca No. 23, Bank of Spain WP No. 0115, and IMF WP No. 02/54, respectively). We are also thankful to Luigi Guiso and Fernando Restoy for suggestions and encouragement, and to Ben Clements, Juan Dolado, Andy Hughes-Hallet, Zenon Kontolemis, Peter Ireland, Ron Smith, Javier Vallés, and seminar participants at the Bank of Spain, Boston College, CEMFI, ECB, 'Ente Einaudi', IMF, the Fifth International Conference on Macroeconomic Analysis and International Finance, the First Workshop of the Euro Area Business Cycle Network, and the 2003 Summer Meeting of the Econometric Society for comments and discussions. Ciccarelli's research was undertaken in part while he was at the University of Alicante (funding by DGIMCYT BEC2002-03097 and Generalitat Valenciana CTIDIB/2002/175 is gratefully acknowledged) and research fellow at the Ente Einaudi and the Bank of Spain. The views expressed in this Paper are exclusively those of the authors and not those of the International Monetary Fund or the European Central Bank. The usual disclaimer applies.

Submitted 16 September 2004

1 Introduction

Theory (e.g., Lucas, 1976), as well as a growing body of empirical evidence on the U.S. post-war monetary history—e.g., Boivin and Giannoni (2003, 2002), Cogley and Sargent (2002), and Primiceri (2002), among others—suggest that the transmission mechanism of monetary policy may change in response to expected or actual changes in the monetary policy regime. The European Monetary Union (EMU) was launched in January 1999, and the euro started to circulate in January 2001 (almost a decade after the adoption of the Maastricht treaty), replacing the exchange rate mechanism that had anchored the European Monetary System (EMS) for more than 20 years. It is thus natural to ask what has happened to the transmission mechanism of European monetary policy in the run-up to the EMU.

At the same time, a fairly large, albeit sometimes contradicting, body of empirical evidence based on pre-EMU data points to the presence of significant differences across countries (or heterogeneity) in the transmission mechanism of monetary policy in Europe—see Guiso et al. (2000) and Angeloni et al. (2002) for recent surveys of this literature. On the one hand, trade and financial integration may increase, and business cycles may become more synchronized, as a result of the currency union (Frankel and Rose, 1998). Hence, these differences might decrease over time if due to these factors. On the other hand, they could also persist for a long time if due to differences in the financial structures rooted in the legal frameworks of individual countries (Cecchetti, 1999). Thus, it would also be useful to have some idea on the time profile of these differences, but there is no hard evidence on whether they are decreasing or persisting over time.

This paper analyses the evolution of the transmission mechanism of European monetary policy by means of time-varying, heterogeneous coefficient models that allow us to test alternative stability and homogeneity assumptions, including particularly the extent to which the impact of monetary policy has changed and cross-country differences have decreased over time. As far as we know, this is the first empirical study of the transmission mechanism of European monetary policy that attempts to do so.

The "experiment" we design incorporates suggestions from the most recent contributions to the empirical literature on the transmission mechanism of European monetary policy and also innovates in one other respect by allowing for regional interdependence in the analysis. First, the specification of the econometric model is the same for all countries considered. Second, following Clements et al. (2001) and Sala (2001), we control for different central banks preferences and procedures as well as intra-Europe exchange rate movements. In addition, by allowing for contemporaneous and lagged interdependence among the open and integrated economies we consider, we avoid aggregation biases and provide for a more realistic description of the international transmission mechanism of monetary policy. Overall, these features bring our experiment closer to the "ideal" one described by Guiso et al. (2000).

Obviously, such a framework cannot be estimated without introducing restrictions on the model because of the very large number of parameters involved. We specify the econometric model in terms of a few hyper-parameters and take a Bayesian approach to estimation. In addition, we take a two-stage approach to the analysis. In the first stage, we measure monetary policy by estimating a system of reaction functions *a-la* Clarida et al. (1998, 2000). In the second stage, we assess the impact of monetary policy on economic activity by estimating a system of output equations as done by Dornbusch et al. (1998) and Peersman and Smets (1998). Thus, we do not model nominal exchange rates and inflation rates endogenously.

We also focus on a small group of European countries: Germany, France, Italy, and Spain. These are the four largest economies currently in the EMU, accounting for about 80 percent of the Euro-area GDP, and there are no strong reasons to expect an *homogeneous* or *stable* response to a common monetary policy shock across these economies. France has been closely tied to Germany throughout the period considered and is widely regarded as a "core" European economy, but its legal structure is different. 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. Spain, finally, joined the European Union and the EMS much later than the other three countries and has been catching up with the rest of Europe throughout the period considered.

Consistent with the predictions of the Lucas' critique and the evidence emerging from the empirical literature on the recent U.S. monetary history, we find that (i) the European transmission mechanism has changed after 1991, with the long-run cumulative impact of a common, homoskedastic monetary policy shock (identified with an innovation to the German reaction function) decreasing by about 10-20 percent in all countries. Consistent with Cecchetti's (1999), we also find that (ii) existing differences in the transmission mechanism of European monetary policy have not decreased over the same period.

We interpret these results as evidence that the transmission mechanism of European monetary policy is probably changing over time under the EMU, in all countries, but rather slowly and in a synchronized manner.

The paper is organized as follows. The econometric framework is presented in section 2. Here we illustrate first the empirical model for the system of reactions functions and then the model for the system of output equations. The rest of the paper reports and discusses the empirical results. In section 3, we present the estimated monetary policy shocks and reaction function parameters. In section 4, we present the estimated effects of monetary policy on economic activity as well as their degree of stability over times and heterogeneity across countries. Section 5 concludes. Details of the estimation techniques used are given in appendix.

2 The econometric framework

To address the questions we ask in the paper, ideally, one would like to estimate a structural, panel VAR for output, inflation, interest rates, and exchange rates—the set of variables usually considered in the literature—allowing for unconstrained interdependence among the countries considered and parameter time-variation. This is feasible, in principle, by taking a Bayesian approach to estimation as done, for instance, by Canova and Ciccarelli (2000) and Zha (1999). It is extremely demanding computationally in practice, though.[1]

Here, in addition to a Bayesian estimation approach (see below), we implement the analysis in two-stages, as done by Dornbusch et al. (1998) and Peersman and Smets (1998), and do not model inflation and the exchange rate endogenously. In the first stage, we extract a measure of the *unexpected* component of monetary policy from the data by estimating a system of reaction functions, allowing for both simultaneity and interdependence among short-term interest rates of different countries and parameter variation across countries and time periods. In the second stage, we analyze the impact of this measure of monetary policy on economic activity by estimating a system of reduced form dynamic output equations, also allowing for unrestricted, lagged interdependence and parameter variation across countries and over time periods.

Thus, the next two sub-sections present and discuss the econometric model of system of reaction functions and output equations, in turn.

[1]Coogley and Sargent (2002) and Primiceri (2002) do estimate a time-varying, structural VARs for output, inflation and the short-term interest rate, but focus only on the U.S. economy. Bovin and Giannoni (2003, 2002) not only focus on one country, the United States, but also take the break points they analyze as given.

2.1 Measuring monetary policy

This sub-section presents and discusses specification, identification, and estimation of the econometric model of the system of reaction functions used in the first stage of the empirical analysis. The definition of a common European monetary shock is also discussed, separately.

2.1.1 Specification

The behavior of the four European central banks considered is modelled empirically by means of the following time-varying structural VAR:

$$A_t(L)\,R_t = B_t(L)\,W_t + D_t + U_t, \tag{1}$$

where $R_t = [r_{1t}, \cdots, r_{4t}]'$ is a (4×1) vector of monetary policy instruments, $W_t = [w'_{1t}, \cdots, w'_{4t}]'$ is a (4×1) vector of vectors of monetary policy final objectives and exogenous variables, $A_t(L)$ and $B_t(L)$ are time-varying polynomial matrices in the lag operator L, with lag length p_1 and p_2 respectively, and D_t is a (4×1) vector of constants. Here, $U_t = [u_{1t}, \cdots, u_{4t}]'$ is a (4×1) vector of monetary policy shocks such that:

$$\begin{aligned} E\left[U_t \mid Z_{t-s}\right] &= 0, \text{ for all } t \text{ and } s > 0; \\ E\left[U_t U_s'\right] &= 0, \text{ for all } t \neq s; \\ E\left[U_t U_t' \mid Z_{t-s}\right] &= I, \text{ for all } t \text{ and } s > 0; \end{aligned} \tag{2}$$

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

More specifically, we assume that the monetary policy instrument is a short-term interest rate. In w_{it}, we include contemporaneous and lagged inflation (π), output (y), and lagged nominal exchange rate (e) in percent deviation from their targets (π^*, y^*, e^*, respectively). We compute π^* and y^* as the fitted values of a linear regression of $\pi_{i,t}$ and $\mathsf{y}_{i,t}$ on a constant and a linear trend.[2] For e^*, in the case of France, Italy and Spain, we take percent deviations from official, central parities vis-a-vis the DM, while we enter the DM/US dollar rate in monthly differences.[3] In addition, we

[2] We also compute the inflation and output targets by taking deviations from the German inflation rate and the HP filtered output trend, respectively, but find very similar results. These additional results, as well as all other results not reported, are available on request.

[3] The exchange rate gap for Italy and Spain is set to zero before the Spanish peseta joined the EMS in 1989 and during the period in which the Italian lira was floating, following the 1992 exchange rate crisis.

also include contemporaneous and lagged values of the U.S. Federal Fund Rate (r^{US}), an index of commodity prices (cp), and the lagged value of a broad monetary aggregate (m)—the latter two variables, in monthly differences, denoted with Δ. We include commodity prices and the Federal Fund rate to control for external shocks, while we include a domestic monetary aggregate because it is often argued that such a variable was monitored closely by some of the European central banks considered throughout the sample period, even after 1979. Thus,

$$w_{it} = [(\pi_{i,t} - \pi_i^*), (\mathsf{y}_{i,t} - \mathsf{y}_i^*), (e_{it-1} - e_i^*), \Delta m_{i,t-1}, \, \Delta cp_t, r_t^{US}]' \text{ for } i = 1, ...4.$$

All data used are from the *International Financial Statistics* database of the IMF. As a proxy for the short-term interest rate, following Bernanke and Mihov (1997), we use a money market rate. Output is measured by an industrial production index. Inflation is measured by the annual change in the consumer price index. The monetary aggregate chosen is a seasonally adjusted M3 series. As already noted, 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. All variables are in natural logarithm so that estimated coefficients can be interpreted as elasticities, taking the natural logarithm of the gross rate of interest for these series.

The specification of the structural VAR in equation (1) is consistent with the literature on the estimation of monetary rules (e.g. Clarida et al., 1998 and 2000), as well as the empirical analysis of monetary policy in open (European) economies—e.g., Bagliano and Favero (1998) and Favero and Marcellino (2001).[4] In addition, the specification presented imposes very few a-priori restrictions on the system of reaction functions, as all parameters in $A_t(L)$ and $B_t(L)$ are unrestricted and can vary over time, including particularly those governing the contemporaneous relations among short-term interest rates.

Leaving $B_t(L)$ unrestricted, while letting these parameters vary over time, allows for the behavior of the central banks considered to change during the sample period, letting the data reveal which objectives were actually being pursued in each period. For instance, the Spanish peseta joined the EMS only in May 1989, the Italian lira has been floating more or less freely from September 1992 to November 1996, and the official parities as well as the fluctuation bands of all three currencies

[4]As it is well known, equation (1) may be interpreted as the reduced form of a forward-looking model or as a system of backward-looking reaction functions. Estimating a system of forward looking reaction functions while also allowing for parameter variation over time is beyond the scope of this paper. We note, however, that Favero and Marcellino (2001) find very similar results when they compare a time-invariant, forward- and backward looking specification very similar to ours, for the same four countries considered here.

vis-a-vis the DM have changed several times during the sample period. Even the Bundesbank's conduct of monetary policy might have changed after the unification or in the run up to the EMU. It is evident that, if these actual or potential policy changes are not accounted for by allowing the system parameters to change over time, they would affect the estimated residuals, thereby undermining their interpretation of well behaved policy innovations (i.e., white noise residuals) as assumed in equation (2). (See Sims, 2001, on this point.).

Similarly, leaving $A_t(L)$ unrestricted and time-varying allows for lagged interdependence among short term interest rates of different countries as well as for varying degrees over time of interest rate smoothing. In addition, as we discuss below, our prior assumptions on the matrix of contemporaneous relations among the interest rates considered—i.e., the matrix $A_t(0)$—allow for the presence of outliers in U_t, possibly stemming from large exchange rate risk premium shocks and money demand shocks not fully accommodated by money supply, which have prevented other researchers from analyzing the unexpected component of monetary policy in Europe, as customarily done in the monetary VAR literature—See Dornbusch et al. (1998), and Favero and Marcellino (2001) for a discussion of this point. Again, it is evident that, if such a possibility is not allowed for in the model, it could result in fictitious parameter time variation, thereby erroneously inducing to conclude that the behavior of the monetary authorities has changed over time.[5]

Nonetheless, we do impose some lag-length restrictions by choosing p_1 and p_2 based on the Schwarz–Bayesian information criterion (BIC) and standard specification tests on the estimated posterior mean of the residuals. While this BIC criterion, which penalizes overparametrization, suggests to set $p_1 = 6$ and $p_2 = 1$, we find that our residuals pass almost all misspecification tests using $p_1 = 2$ and $p_2 = 1$ (see Section 3 below). Thus, to save computing time, the results are based on the shorter lag-length.

2.1.2 Identification

Identification of the structural VAR in equation (1) may be achieved through exclusion restrictions on the contemporaneous coefficient matrix $A_t(0)$. The specific scheme we use relies on the Bundesbank's presumed leading role in the EMS and the relative economic size of other countries.

[5] Again, on this point, see Sims (2001). Indeed, in a previous version of this paper, in which time-variation of the variance-covariance matrix of the residuals was not allowed for, we find much more marked changes in the conduct of monetary policy during the sample period than in this version that controls shocks heteroskedasticity (see *Temi di Ricerca n.* 23, Ente Einaudi, Roma).

More specifically, 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. We then assume that French and Italian interest rates affect contemporaneously the Spanish rate without being affected by it. This is plausible given that Spain's GDP was considerably smaller than that of France and Italy during much of the period considered and that Spain joined the EMS only in 1989. Finally, we assume that the impact of an increase in interest rates in Italy in France is the same as the impact in Italy of a rate increase in France.[6]

Formally, to identify the model, we need six restrictions on $A_t(0)$. The assumptions above provide the six restrictions needed to identify the model exactly and translate into the following block recursive structure for $A_t(0)$:

$$A_t(0) = \begin{bmatrix} A_{11}(0) & 0 & 0 \\ A_{21}(0) & A_{22}(0) & 0 \\ A_{31}(0) & A_{32}(0) & A_{33}(0) \end{bmatrix}_t \tag{3}$$

where $A_{11}(0)$, $A_{31}(0)$, and $A_{33}(0)$ are scalars, $A_{21}(0)$ and $A'_{32}(0)$ are 2×1, and $A_{22}(0)$ is 2×2 symmetric matrix.

Therefore, the structural VAR in equation (1) can be rewritten as:

$$\begin{aligned} \begin{bmatrix} A_{11}(0) & 0 & 0 \\ A_{21}(0) & A_{22}(0) & 0 \\ A_{31}(0) & A_{32}(0) & A_{33}(0) \end{bmatrix}_t \begin{pmatrix} R_{1t} \\ R_{2t} \\ R_{3t} \end{pmatrix} &= \begin{bmatrix} A_{11}(L) & A_{12}(L) & A_{13}(L) \\ A_{21}(L) & A_{22}(L) & A_{23}(L) \\ A_{31}(L) & A_{32}(L) & A_{33}(L) \end{bmatrix}_t \begin{pmatrix} R_{1t-1} \\ R_{2t-1} \\ R_{3t-1} \end{pmatrix} \\ &+ \begin{bmatrix} B_{11}(L) & B_{12}(L) & B_{13}(L) \\ B_{21}(L) & B_{22}(L) & B_{23}(L) \\ B_{31}(L) & B_{32}(L) & B_{33}(L) \end{bmatrix}_t \begin{pmatrix} W_{1t} \\ W_{2t} \\ W_{3t} \end{pmatrix} + D_t + \begin{pmatrix} U_{1t} \\ U_{2t} \\ U_{3t} \end{pmatrix} \end{aligned} \tag{4}$$

where R_{1t}, W_{1t} and U_{1t} are the German monetary policy instrument, objectives and shock, respectively, and R_{2t},W_{2t} and U_{2t} (R_{3t}, W_{3t}, and U_{3t}) are the vectors containing the same variables for France and Italy (and Spain).

2.1.3 Definition of a common European monetary policy shock

Before the adoption of the euro, as pointed out by Dornbusch et al. (1998) and Guiso et al. (2000), differences in the transmission mechanism across countries might have originated from differences in central banks behavior as well as differences in other transmission channels. The adoption of a

[6] See Amisano and Giannini (1997, page 166-67) for an example of identification by means of symmetry, as assumed here for France and Italy. Recall also that the U.S. Federal Fund rate enters contemporaneously all reaction functions, like an exogenous variable.

common currency, however, has removed the former potential source of heterogeneity. Thus, the relevant question from the present time perspective is whether the transmission mechanism was heterogeneous before the EMU launch once we account and hold constant potential heterogeneity in central banks behavior and other transmission channels that disappeared under the EMU.

A common European monetary policy shock may be defined as an innovation that affects all money markets considered simultaneously. By this definition, the only innovation with this property, in the model discussed above, is the structural residual of the Bundesbank reaction function. Indeed, the scheme discussed above identifies the German monetary shock by assuming that it is the only innovation that affects all short-term interest rates in the system. Thus, consistent with the presumed leading role of the Bundesbank under the EMS, as well as recent work by Sala (2001) and Clements et al. (2001), who focus on the identification of a common European monetary policy from per-EMU data, we identify and thus define a common European monetary policy shock with the innovation of the Bundesbank reaction function.

This is not the end of the story, though. Under the EMS, such a common innovation could have been offset or compounded by other central banks' reactions to this shock by regulating the supply of base money in the local money market, within the limits imposed by the exchange rate mechanism under the EMS. This degree of autonomy, though limited, was sufficient to induce short-term interest rate differentials with Germany, which were also part of the transmission mechanism of monetary policy under this monetary regime. Indeed, under the EMU, a surprise move by the European Central Bank affects all European money markets at the same time, but national central banks (including the Bundesbank) cannot react to these shocks anymore. Therefore, in order to approximate the conditions prevailing under the EMU as closely as possible, while using the EMS data to study the transmission mechanism, one would also like to hold constant these spreads, which have now disappeared.

To do so, following Clements et al. (2001), we shall include in the system of output equations discussed in the next section also the short-term differential with Germany. The rationale for the inclusion of these interest rate spreads may be seen by noting that, as pointed out by Clements et al. (2001), for any given reference country j (say Germany) and any other country in the system i, the following identity holds:

$$r_{it} \equiv r_{jt} + (r_{it} - r_{jt}).$$

Hence, all interest rates except the j^{th} can be split in two components: a common base interest rate (r_{jt}, the German interest rate) and a country-specific differential or spread ($r_{it} - r_{jt}$) over this reference rate. We can then further split the reference interest in expected and unexpected components ($\widehat{r}_{jt}$ and $\widehat{u}_{jt}$, respectively), and analyze the dynamic impact of $\widehat{u}_{jt}$ on the endogenous variables in the system. Obviously, there needs to be a good rationale for choosing the reference interest rate, but this follows naturally from the identification assumptions made in our case.

2.1.4 Estimation

Bayesian estimation of the structural VAR (4) exploits the block recursive structure of its reduced from. 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 enter the equations of each block j. If we pre-multiply (4) by the (4×4) matrix

$$A_{t,d}^{-1}(0) = \begin{bmatrix} A_{t,11}^{-1}(0) & \mathbf{0} & 0 \\ \mathbf{0} & A_{t,22}^{-1}(0) & \mathbf{0} \\ 0 & \mathbf{0} & A_{t,33}^{-1}(0) \end{bmatrix}, \tag{5}$$

and rearrange terms, the model can by divided in three blocks:

$$R_{jt} = Z_{jt}\delta_{jt} + v_{jt}; \qquad j = 1, 2, 3, \text{ for all } t. \tag{6}$$

Here, $Z_{jt} = diag\left[Z_{1,t}^j, Z_{2,t}^j, ..., Z_{G_j,t}^j\right]$ is a $(G_j \times k_j G_j)$ diagonal matrix whose elements are the $(1 \times k_j)$ vectors $(Z_{g,t}^j)$ that contains all contemporaneous (i.e., R_{1t} in blocks 2 and 3, and R_{1t} and R_{2t} in block 3) and lagged endogenous variables, exogenous and deterministic variables of equation g in block j, for $g = 1, ..., G_j$; $\delta_{jt} = \left[\delta_{1,t}^j, \delta_{2,t}^j, ..., \delta_{G_j,t}^j\right]$ is a $(k_j G_j \times 1)$ vector whose $(k_j \times 1)$ element $(\delta_{g,t}^j)$ contains the parameters of equation g in block j (for $g = 1, ..., G_j$); and $v_{jt} = A_{t,jj}^{-1}(0)\, U_{jt}$.

Parameter time-variation in (6) is random and assumes (*a priori*) that δ_{jt} follows a random walk without drift (i.e., $\delta_{jt} = \delta_{jt-1} + \eta_{jt}$). This assumption is a plausible way to describe permanent structural changes in behavior due to fundamental changes in policy regimes. It is standard in the literature on random time-varying coefficient models (e.g., Doan et al., 1984) and is also in the analysis of the recent U.S. monetary history (e.g., Cogley and Sargent, 2002, and Primiceri, 2002, among others). Moreover, the main alternatives available in the Bayesian literature, a stationary autoregressive process or a discrete regime-switching process with possibly absorbing states, seem less attractive, either because they imply that changes are reversible or because of the limited number of states the system can reach, on a priori basis.

Parameter time-variation in the contemporaneous coefficient matrix (5) requires modeling an heteroskedastic variance-covariance matrix of the residuals in (6). This is achieved, in a simple manner, by assuming (*a priori*) that (i) $U_{jt} \sim N(0, I)$ and (ii) $A_{t,jj}^{-1}(0) = \sqrt{\sigma_{jt}} \mathsf{A}_{jj}^{-1}(0)$, with (iii) σ_{jt} ~Inv-$\chi^2(\nu_j, 1)$, where $N(0, I)$ denotes a standard normal multivariate distribution and Inv-$\chi^2(\nu, 1)$ denotes a scaled, inverse-χ^2distribution with ν degrees of freedom. These assumptions allow for the presence of large realizations of v_{jt}, which may reflect exchange risk premium shocks, or large money demand shocks not fully accommodated by the monetary authorities, during periods of financial turbulence, and which might otherwise distort the posterior distributions of parameters of interest if not controlled for. These assumptions are also standard in the Bayesian literature (see, for instance, Gelman et al., 1995, p. 350), but differ from those made by Cogley and Sargent (2002) and Primiceri (2002), who assume an autoregressive structure for σ_{jt}, rather than a simpler, random coefficient specification as we do, to yield a stochastic volatility model similar to that of Uhlig (1992). But this is not necessary for the purpose of controlling for exchange premium and money demand shocks, which are our main concerns here.

To see how this works, note that assumptions (i)-(iii) are equivalent to the hypothesis that $v_{jt} \sim t_\nu(0, \Sigma_{jj})$, where $t_\nu(0, \Sigma_{jj})$ denotes a t-distribution with ν degrees of freedom and scale matrix Σ_{jj}. In fact, under (i) and (ii), $v_{jt} \sim N(0, \Upsilon_{t,jj})$, with $\Upsilon_{t,jj} = A_{t,jj}^{-1}(0) A_{t,jj}^{-1}(0)' = \sigma_{jt} \mathsf{A}_{jj}^{-1}(0) \mathsf{A}_{jj}^{-1}(0)' = \sigma_{jt}\Sigma_{jj}$. Thus, we also have that $v_{jt} \mid \sigma_{jt} \sim N(0, \sigma_{jt}\Sigma_{jj})$. From (iii), it then follows that $v_{jt} \sim t_\nu(0, \Sigma_{jj})$. And as it is well known, the t-student assigns higher probability than the normal to the tails of the distribution, and hence higher probability than the normal to extreme values or outliers.

As explained in more detail in the appendix, Bayesian estimation of the three blocks of (6) is obtained by using the Kalman filter and the Gibbs sampler as suggested by Chib and Greenberg (1995) for systems of seemingly unrelated regressions (SUR). Intuitively, a joint prior on the model parameters (i.e., a joint prior on the series of δ_t^j, their error variance, and the variance-covariance matrix of v_{jt}) is combined with the likelihood of the data, and suitable initial values for the model's hyper-parameters (i.e., the moments of the prior distributions), to recover the joint posterior distribution of the parameters of interest (i.e., the series of δ_t^j, their their error variance, the variance-covariance matrix v_{jt}, and any possibly non-linear function of them). As analytic integration of the joint posterior distribution is not feasible, this is implemented numerically by means of Gibbs Sampler, which is a widely used Monte Carlo Markov Chain simulation method (see, for instance, Gelfand et al. 1990).

The sample mean of the posterior distribution of the parameters and other quantities of interest (or their medians if the posteriors are symmetric) may then be interpreted as a classical point estimate, while the percentiles of the posterior distribution provide for the equivalent of classical confidence intervals. Note in particular that, since the matrix (5) is exactly identified, given the marginal posterior distribution of the variance-covariance matrix of v_{jt}, we can recover its posterior, and hence also the posterior distribution of the structural residuals and other model parameters. Specifically, as our measure of the *unexpected* common component of European monetary policy, in the second stage of the analysis, we use the the sample mean of the posterior distribution of the innovation of the German reaction function.

2.2 Modeling the transmission mechanism

This sub-section discusses specification, estimation and testing of the econometric model of the system of output equations that we use in the second stage of the analysis. This model shares many features with the model used in the first stage of the analysis. Hence, here, we focus primarily on those aspects of the model and the analysis that differ from the first stage.[7]

2.2.1 Specification

The transmission mechanism of a common European monetary policy shock to economic activity is modelled empirically by means of the following time-varying VAR system of output equations:

$$\mathsf{A}_t\,(L)\,\mathsf{Y}_t = \mathsf{B}_t\,(L)\,\mathsf{W}_t + \mathsf{C}(L)\widehat{u}_t + \mathsf{D}_t + \mathsf{U}_t, \tag{7}$$

where $\mathsf{Y}_t = [\mathsf{y}_{1t}, \cdots, \mathsf{y}_{4t}]'$ is the (4×1) vector of real output indices (in natural logarithms), $\mathsf{W}_t = [\mathsf{w}'_{1t}, \cdots, \mathsf{w}'_{4t}]'$ is a (4×1) vector of control variables, $\widehat{u}_t$ is the estimated residual of the German reaction function from the first stage of the analysis, $\mathsf{A}_t\,(L)$, $\mathsf{B}_t\,(L)$, and $\mathsf{C}_t(L)$ are time-varying polynomial matrices in the lag operator L of conforming dimension, with lag length p_1, p_2, and p_3, respectively. The (4×1) vector of residuals $\mathsf{U}_t = [\mathsf{u}_t^1, \cdots, \mathsf{u}_t^4]'$ has zero-mean and is serially uncorrelated, as in equation (2), but with an unrestricted and time-varying variance-covariance matrix specified more precisely below. Finally, the (4×2) matrix D_t includes country-specific

[7] Sans serif fonts distinguish the system of output equations from the system of reaction functions. All variables are defined as in the previous sub-section, unless explicitly noted.

constants and linear trends.[8]

In w_{it}, in all equations, we include the (lagged) nominal exchange rate (e^{ECU}) vis-a-vis the ECU, the fitted value of the German interest rate from the first stage of the analysis, and the spread over the German money market interest rate (the latter only in all equations except Germany). The first variable holds constant the intra-Europe exchange rate channel of transmission of monetary policy. The second holds constant the expected component of German monetary policy. The last holds constant both the expected and unexpected component of domestic monetary policy in the other three countries as we explained before. In addition, we include in all equations contemporaneous and lagged inflation (π), contemporaneous and lagged commodity prices (cp), and a U.S. output index (y^{US}). We include inflation to control for domestic supply side factors, and commodity prices and U.S. output to control for external shocks. Thus,

$$\begin{aligned}
\mathsf{w}_{it} &= \left[e^{ECU}_{1t-1}, \widehat{r}_t, \pi_{1,t}, cp_t, \mathsf{y}^{US}_t\right]' && \text{for } i = 1, \\
\mathsf{w}_{it} &= \left[e^{ECU}_{it-1}, \widehat{r}_t, sp_{it-1}, \pi_{i,t}, cp_t, \mathsf{y}^{US}_t\right]' && \text{for } i = 2, 3, 4.
\end{aligned}$$

As in the case of the system of reaction functions, we choose the lag-length based on the BIC criterion and a battery of misspecification tests on the estimated residuals, which lead us to select $\mathsf{p}_1 = 1$ and $\mathsf{p}_2 = \mathsf{p}_3 = 6$.

This specification provides only for a stylized description of the transmission mechanism of a common European monetary policy shock to real economic activity, nonetheless, it is consistent with standard VAR specifications for the empirical analysis of monetary policy in open economy and also with specifications previously used to analyze monetary policy in the countries we consider (e.g., Dornbusch et al., 1998, and Favero and Marcellino, 2001). In addition, this specification is the same for all countries considered, and also allows for unrestricted regional interdependence; an aspect of the transmission mechanism of European monetary policy generally overlooked, which might be important for the highly integrated economies we consider. Above all, as already discussed for the system of reaction functions, this specification allows the transmission mechanism to change over time, while controlling from shifts in the variance of the model residuals, and thus all those factors that are not explicitly accounted for in the model.

[8]We find similar results when we deternd the output series before estimation, or when we enter the output series in 12-month log-difference (i.e., in annual rate of growth). With annual growth, however, the estimated impact of monetary shocks is not neutral in the long run. For this reason, we prefer working with log-levels and a linear trend in the system; a specification that is also consistent with the first stage of the analysis. (See Leeper et al. (1996) for a similar treatment of trends in output series in a Bayesian VAR specification.)

2.2.2 Estimation

Stacking all equations by row and rewriting (7) as a SUR system we have:

$$\mathsf{Y}_t = \mathsf{X}_t \beta_t + \varepsilon_t. \tag{8}$$

In this system, $\mathsf{X}_t = diag\,[X'_{1t}, ..., X'_{4t}]$ is of dimension $4 \times h$, $\beta_t = [\beta'_{1t}, ..., \beta'_{4t}]'$ is of dimension $h \times 1$, where $h = k_{ger} + 3k$, with $k_{ger} = 33$ denoting the number of regressors in the German output equation and $k = 39$ denoting the number of regressors in the other equations.

As in the first stage of the analysis, parameters time-variation is random and follows:

$$\begin{aligned} \mathsf{Y}_t &= \mathsf{X}_t \beta_t + \varepsilon_t, & \varepsilon_t &\sim N_G\left(0, \varsigma_t \Omega\right), \\ \beta_t &= \beta_{t-1} + \zeta_t, & \zeta_t &\sim N_h\left(0, B_2\right). \end{aligned} \tag{9}$$

Here, $\varsigma_t \Omega$ and B_2 denote the variance-covariance matrix of ε_t and ζ_t, respectively. Thus, if $B_2 = 0$, then $\beta_t = \beta_{t-1} = \beta$, the model becomes time-invariant. Note however that parameters can still differ across countries.

To provide a benchmark for comparison, we also estimate the following time-invariant model:

$$\begin{aligned} \mathsf{Y}_t &= \mathsf{X}_t \beta + \varepsilon_t, & \varepsilon_t &\sim N_G\left(0, \Omega\right), \\ \beta &= M_o \overline{\beta} + \zeta, & \zeta &\sim N_s\left(0, B_o\right), \end{aligned} \tag{10}$$

where the $(h \times k)$ matrix M_o, which shrinks the $(h \times 1)$ vector of country specific parameters β to the $(k \times 1)$ vector of common parameters $\overline{\beta}$, is a column vector of four identity matrices of dimension k, with the first matrix taking into account of the different number of regressors entering the German equation. Thus:

$$M_o = \begin{bmatrix} \begin{pmatrix} I_{k_{ger}} & \mathbf{0}_{k_{ger} \times (k - k_{ger})} \\ \mathbf{0}'_{(k-k_{ger}) \times k_{ger}} & 0 \end{pmatrix} \\ I_k \\ \vdots \\ I_k \end{bmatrix}.$$

Here, Ω and B_o denote the variance-covariance matrices of ε_t and ζ, respectively. Thus, if $B_0 = 0$, then parameters are not only constant over time but also homogenous across countries.

As we describe in detail in appendix, Bayesian estimation of (10) and (9) is implemented by combining prior assumptions on the model parameters with the likelihood of the data, and suitable initial values for the model hyper-parameters, to obtain posterior distributions of the parameters and the objects of economic interest; in our case, the dynamic multipliers of the common monetary shock, $\widehat{u}_t$.

Dynamic multipliers with respect to an exogenous variable in the system, as impulse response functions to innovations in endogenous variables, are given by the coefficients of the moving-average representation of the model. Thus, they are non-linear functions of the elements of β_t. We compute them for our model following Lütkepohl (1991, pp. 338-39) and sample their posterior distributions within the Gibbs sampler because analytical integration is obviously not feasible.

2.2.3 Testing

Several hypotheses of stability and homogeneity can be tested on the posterior distributions of the parameters of interest. Testing for overall parameter stability over time, or homogeneity across countries, may be easily done by letting B_2 and B_0 depend upon only one hyper-parameter, say φ^2, which controls the random variation of *all* model coefficients, and then applying standard Bayesian model selection criteria.

One approach we take looks directly at the posterior distribution of φ^2, following Chib and Greenberg (1995). If this posterior is more skewed toward zero than its prior, then we can conclude that the data do not support a time-varying, or homogenous, specification. Specifically, if one finds that under the *posterior*, the probability that φ^2 is arbitrarily small is larger than under the *prior*, then we can conclude that the data shift the 'odds' in favour of a small φ^2, and thus against a random varying specification (Chib and Greenberg, 1995, p. 344). In practice, this test may be implemented by giving a proper prior distribution on φ^2 and calculating, for arbitrarily small values of α, the ratio

$$z = \left(\frac{\Pr\left(\varphi^2 \leq \alpha \mid y\right)}{\Pr\left(\varphi^2 \leq \alpha\right)} \right) \left(\frac{\Pr\left(\varphi^2 > \alpha \mid y\right)}{\Pr\left(\varphi^2 > \alpha\right)} \right), \tag{11}$$

where $\Pr\left(\phi^2 \leq \alpha \mid y\right)$ and $\Pr\left(\phi^2 \leq \alpha\right)$ denote the conditional *posterior* probability and unconditional *prior probability* that φ^2 is less than α, respectively.[9]

A complementary approach we take is the calculation of the Bayes factor (BF) of alternative models. This is computed as the ratio of the marginal data density under alternative assumptions.

[9] The ratio in the first bracket compares the odds that $\phi^2 \leq \alpha$ under the posterior distribution, while the ratio of complementary probabilities in the second bracket is a weighting factor.

Given two models, say M_1 (e.g., $\varphi^2 = 0$) and M_2 (e.g., $\varphi^2 \neq 0$),

$$BF = p\left(\mathsf{Y} \mid \mathsf{M}_1\right) / p\left(\mathsf{Y} \mid \mathsf{M}_2\right),$$

where

$$p\left(\mathsf{Y} \mid \mathsf{M}_j\right) = \int p\left(\mathsf{Y} \mid \psi_j, \mathsf{M}_j\right) p\left(\psi_j \mid \mathsf{Y}\right) d\psi_j, \qquad \text{for } j = 1, 2,$$

with $p\left(\mathsf{Y} \mid \psi_j, \mathsf{M}_j\right)$ denoting the likelihood under model M_j, given the specific parameter vector ψ_j, and $p\left(\psi_j \mid \mathsf{Y}\right)$ the posterior density of the latter. The larger is this ratio, the higher the posterior probability that the data comes from M_1. The marginal data densities needed for implementation of this test can then be computed by using the Gibbs sampler-based method developed by Chib (1995).

There is a problem with this testing strategy for the purpose of addressing the questions we ask in the paper, though. The objects of our economic interest are the dynamic multipliers of the common monetary shock, which are non-linear functions of all model parameters. Thus, small (large) changes over times, or differences across countries, in all parameters might be exacerbated (off-set) by their non-linear combination to yield large (small) changes in the dynamic multipliers of the transmission mechanism. In addition, the two testing procedures above compare the amount of random variation in the data relative to that assumed under the prior.

An alternative testing strategy we take is to look directly at the posterior distributions of the dynamic multipliers, at various time-horizons, computed within the Gibbs sampler. We are particularly interested in investigating (i) whether the transmission mechanism has changed over times in individual countries, (ii) whether it was homogeneous across countries, and (iii) whether any cross-country difference we might find has decreased or persisted over time. Therefore, in addition to the two tests of overall parameter variation described above, we also inspect the posterior distributions of the dynamic multipliers directly by means of standard statistical test. In particular, we shall focus on results based on t-tests for different sample means and the Kolmogorov-Smirnov (Mood et al., 1974, for instance) test for the overall distance between two sample distributions. We compare these test statistics across countries over the entire sample periods, over times for each country, and across countries and over time periods.

The discussion of alternative testing strategies for homogeneity and stability of the transmission mechanism of monetary policy concludes the presentation of the econometric framework we use. In the next two sections we report and discuss the empirical results.

3 Estimated monetary policy shocks and reaction functions

In this section we report the estimated reaction function residuals and parameters. While in the second stage of the analysis we look only at the transmission mechanism of a common (i.e., German) monetary policy shock, looking also at estimated residuals and parameters of other reaction functions helps assessing how widespread were changes over time in the conduct of monetary policy in the countries we consider. In turn, this helps placing the results we present in the next section on changes in the transmission mechanism in a broader context.

Figure 1 reports the posterior mean of the estimated residuals and the heteroskedasticity factor (σ_{jt}), while Figures 2-4 report the mean, the first and third quartile of the posterior distributions of the reaction function parameters. The data sample runs from January 1980 to December 1998, but the results starts only in February 1981 because we use 24 observations to initialize the estimation procedure.

Overall, these results show that the model fits the data well, both statistically and economically, and suggest that the behavior of all central banks changed somewhat in the run up to the EMU. In particular, the conduct of monetary policy in Germany displays a mild structural break in 1991, with weights attached to foreign variables and the domestic monetary aggregate becoming smaller after 1991. However, it is not clear whether these changes reflect a "unification" effect or a "Maastricht treaty" effect. The results also conform well to standard views on the functioning of the exchange rate mechanism under the EMS, portraying Bank of France as a close "follower" of the Bundesbank, and the Bank of Spain and the Bank of Italy as pursuing more independent, and more markedly changing over time, monetary policies.

3.1 Are estimated monetary policy shocks clean?

The estimated residuals of the system of reaction functions do not reveal any evidence of substantive model misspecification. As we can see from Figure 1, they appear remarkably well behaved for all countries considered: there are essentially no outliers and there is also little or no evidence of serial autocorrelation and/or heteroskedasticity, including particularly around the time of the 1992 exchange rate crisis.[10] As we can see from the posterior mean of the heteroskedasticity factor

[10]The 1992 crisis was has been a modeling obstacle to the analysis of the unexpected component of European monetary policy in many other studies. See in particular the discussion n Dornbusch et al. (1998) and Favero and Marcellino (2001).

in Figure 1, our prior assumptions on the variance-covariance matrix of the residuals appears to capture well that episode of instability for Italy and France as well as the high volatility of Spanish interest rates at the beginning of the sample period.

This assessment is confirmed by a battery of standard test statistics, reported in Table 1. The first two lines report the posterior mean of the residuals and the p-value on the null hypothesis that this is zero, respectively. The second two lines report a Kolmogorov-Smirnov statistics for the Durbin's (1969) cumulated periodogram test on the null hypothesis that the posterior means follow white-noise processes. The following six lines report the Ljung–Box's statistics for the null hypothesis of no serial correlation of order lower or equal than specified, with their respective p-values. Finally, the last two lines report Engle's test for the null of no autoregressive conditional heteroskedasticity of second order, also with their p-values in brackets.

As we can see from this table, the hypothesis of zero mean and white noise process cannot be rejected for all countries, and the *absence* of heteroskedasticity and autocorrelation cannot be rejected for Germany and France. The results for Spain, and to a much lesser extent also for Italy, point to the presence of some heteroskedasticity of order lower or equal than two and autocorrelation of order lower or equal than 12. This suggests that, in the case of Spain, our simple prior assumption (16) may be missing some higher-order conditional heteroskedasticity and a longer lag length might have done a better job. As in the second stage of the analysis we focus only on the German policy innovation, we do not purse these two issues further here. Thus, we can now focus on the estimated reaction function parameters.

3.2 Has the conduct of monetary policy changed in the run-up to the EMU?

The results show some evidence of a mild structural break in the Bundesbank reaction function after 1991, right after the German unification and the adoption of the Maastricht treaty, with a slight decrease in the weight attached to foreign variables and the domestic monetary aggregate considered. The other three countries' reaction functions show more pronounced changes, including in particular in the weight attached to the coefficient on the German exchange rate and the domestic monetary aggregate after 1992. The results also show substantial differences across countries in the conduct of monetary policy, despite of their EMS membership, thus confirming the wisdom of focusing on a common monetary shock in the second stage of the analysis.

The picture emerging from the estimated coefficients of the Bundesbank reaction function (Figure 2) fits well standard views on its behavior as the "leading central bank" in the EMS. The estimated coefficients show a high degree of interest rate smoothing, relatively large weight attached to domestic inflation and output targets, smaller weights attached on other European countries' targets (including particularly inflation and exchange targets), and considerable weight given to the U.S. Federal Funds rate and the DM/US dollar exchange rate.[11] In particular, the coefficient on the lagged interest rate is close to one throughout the sample period. Among domestic objectives, inflation appears the most important variable, although its effect is not estimated very precisely. The U.S. Federal Funds rate and the DM/U.S. dollar exchange rate also have notable impacts. The coefficient on the U.S. Federal Funds rate, in particualr, is comparable in size to that on the domestic inflation target. The coefficients of other European countries' variables, instead, are generally not significantly different from zero, except for the French and Spanish output gaps and the Italian exchange rate gap. But even these coefficients are smaller than those on domestic variables and declining over time. The coefficients of foreign inflations are clearly not significant statistically.

The parameters of the reaction function of the Bank of France also conform well to what one would expect from a "follower" central bank in the EMS (Figure 3): the German policy interest rate is by far the most important variable, and remains so throughout the sample period, while the Federal Fund Rate does not appear significant statistically; the parity versus the DM has a smaller importance, and declining markedly over time, but it is significant statistically throughout the sample period. The only domestic variable that has a comparable weight is the domestic monetary aggregate, but this declines considerably over time. Domestic and foreign inflation and output variables, instead, have weights which are either very small or not significant statistically, and changing to a varying degree over time.

The results for Spain, and to a lesser extent Italy, are consistent with the view that the Bank of Spain and the Bank of Italy were relatively less constrained by (or committed to) the exchange rate mechanism under the EMS than the Bank of France (Figures 4 and 5). Compared to France, both reaction functions show a much smaller, statistically insignificant, weight attached to the German and the U.S. interest rates, and notably larger weights attached to domestic output gaps. Interestingly, as in the case of France, the coefficient of the German exchange rate appears to have declined significantly after 1992. The weight attached to monetary growth in the reaction function of the Bank of Spain is larger than that appearing in the reaction function of the Bank of France,

[11] Note that an increase in the exchange rate denotes a depreciation.

while in the reaction function of the Bank of Italy is smaller. The weight attached to the domestic output target affects Spanish short-term interest rates to a much greater extent than in France, or even Italy, with an impact comparable to that of the German interest rate. The reaction function of the Bank of Spain also shows more instability, consistent with Spain's later entry in the EMS. Finally, interest rate persistence is the smallest in Spain.

4 The impact of a common monetary policy shock on economic activity

In this section, we report the estimated impact of a common and homoskedastic monetary policy shock on real economic activity in individual countries, together with test statistics on its stability over time, heterogeneity across countries, and on how heterogeneity might (or might not) have changed over time. We first provide a benchmark for comparison by looking at the estimation results based on the time-invariant model (10). We then report the time-varying results based on (9).

All results refer to the sample period 1984-1998, since the series of monetary shocks reported in Figure 1 runs from 1981 to 1998, but we use 36 observations to initialize the estimation procedure. We report results for end-1985 (before widespread exchange rate controls were finally lifted), end-1991 (before the German unification and the adoption of the Maastricht treaty), end-1995 (after the entry into force of the Maastricht treaty, and end-1998 (right before the EMU launch).

Overall, the results show that the long-run cumulative effect of a monetary policy shock on economic activity has decreased after 1991, in all countries considered, and that cross-country differences have persisted during this period. These results are broadly consistent with the main thrust of the existing literature on the transmission mechanism of European monetary policy and are in the ballpark of those emerging from the literature on the recent U.S. monetary policy history.

4.1 Were there cross-country differences?

There appears to be some cross-country differences in the transmission mechanism of European monetary policy, even after controlling for all those channels that have now disappeared under the EMU, consistent with most of the existing literature on this issue.[12] Yet these differences seem

[12]See see Clements (2001), Guiso et al.(2000), and Ciccarelli and Rebucci (2001) for attempts at comparing quantitative results among different contributions in the literature.

more a matter of *timing* than *magnitude* of the cumulative effects of monetary policy on output, as far as Germany, Italy, and Spain are concerned, consistent with the results reported by Favero and Marcellino (2001), who find little difference among these countries. In the case of France, instead, there also seems to be significant quantitative differences.

Figures 6 reports the posterior mean, together with the posterior first and the third quartile, of the dynamic multipliers of the series of common monetary policy shocks at different time horizons. In all countries, including France, the dynamic effect of the common monetary policy shock is hump-shaped, becoming negative in a statistically significant manner within six-months, peaking within a year, and vanishing within approximately two years. Interestingly, these multipliers are estimated quite precisely for all four countries, even holding constant country-specific channels of transmission disappeared under EMU.

The different impact of monetary policy on output can be seen clearly from Figure 6 in the case of France, while the timing differences among the other three countries are less evident. Cumulating the dynamic multipliers plotted in Figure 6 brings them out more clearly. We present these results in Table 2, which reports the mean, the median, the first and the third quartile of the posterior distribution of the accumulated dynamic multiplier after six, 12, 24 months, and a theoretically infinite number of months (henceforth, called the "long-run impact"). These results confirm that there are cross-country differences in the impact of monetary policy on the level of output in the short run (i.e., say within six to 12 months), but apparently much smaller differences in the longer run (i.e., say after 24 months), as far as Germany, Italy, and Spain are concerned. Only in the case of France the cumulative impact of monetary policy after 24 months is about 25 percent lower than in the other three countries.

Despite the similarities in the cumulative impact of monetary policy among Germany, Spain and Italy, Chib's (1995) specification test and the Bayes factor criterion discussed before reject a pooled model, i.e., a model with identical regression coefficients across countries as in (10) (see Figure 7a). But this is not surprising because, as we noted previously, dynamic multipliers, which are the parameters of economic interest, are highly non-linear functions of the VAR parameters, and small (large) differences in the latter may result in large (small) difference in the former. For the same reason, we cannot draw firm conclusions from the evidence reported in Figure 7b, showing increased heterogeneity at the level of VAR parameters when we omit those control variables that hold constant channels of transmission disappeared under EMU.[13]

Similarities among Germany, Italy and Spain on the one hand, and differences between France and the other three countries on the other hand, however, appear statistically significant based on standard statistical tests applied directly to the posterior distributions of the cumulative dynamic multipliers. This is brought out clearly by Table 3, which reports a set of bilateral comparisons based on a t-test for equal sample means and a Kolmogorov-Smirnov test for equal distributions (henceforth, denoted KS). These results show that the only detectable difference in the cumulative impact of monetary policy after two years is that between France and the other three countries, while the equality of the posterior distribution of these quantities among other three countries cannot be reject by the data.

4.2 Has the transmission mechanism changed over time?

The transmission mechanism of European monetary policy appears to have changed in the run-up to EMU, but not dramatically, and differences across countries do not seem to have decreased. While the hypothesis that all VAR coefficients are stable over time cannot be rejected by the data, we find that the long-run cumulative impact of monetary policy on output has declined in all four countries in the 1990s, by about 10-20 percent. By the same token, differences across countries remained broadly stable; thus, suggesting that these changes might have been driven by common factors rather than a convergence process.

These conclusions may be drawn in part from Figure 8, which plots the posterior mean of the dynamic multipliers at different time horizons, at the end of selected years in the sample period. As we can see, there are some changes over time, especially between 1995 and 1998, with the peak of the negative impact of monetary policy on output decreasing slightly in all countries. This change is more notable in the case of Germany and less marked in the case of Italy. Even in the case of Italy, however, the peak effect in 1995 is smaller than in 1985 and 1991. From this figure, we can also see that the impact of monetary policy continues to differ markedly from that in other countries in the case of France throughout the period considered.[14]

[13]Comparing dynamic multipliers estimated by including progressively smaller sets of control variables, we find that the intra-Europe exchange rate channel appears the most important source of heterogeneity. This is consistent with the results of Sala (2001) and Clements (2001), who continue to find considerable heterogeneity even after controlling for different central bank reaction functions, but either do not control for the exchange rate channel (Sala, 2001), or use different model specifications for different countries (Clements et al, 2001), and don't allow for regional interdependence.

[14]While a quantitative comparison with other studies would not be meaningful, the percent declines in the peak

Our main conclusions are easier to see by inspecting Table 4, which reports the cumulative impact after 6, 12, 24 months, and in the long-run, for the same selected end-years. This confirms that the long-run cumulative impact of monetary policy has decreased in all countries by a similar percentage compared to 1985. As a result, cross-country differences do not appear to have decreased in the run up to EMU.

Most of these changes in the transmission mechanism of monetary policy also appear statistically significant, despite the fact that the hypothesis that all VAR coefficients are stable over time cannot be rejected by the data (see Figure 9). This is supported by a direct inspection of the posterior distributions of the dynamic multipliers and their cumulative sums: a battery of t-tests for equal means and KS statistics for equal distributions on the posterior distributions of the cumulative multipliers, for each country and pair of end-periods considered, reported in Table 5, does show that the changes over time in the long-run cumulative impact of monetary policy are significant statistically in all countries, except Italy between 1995 and 1998.

A similar battery of tests (reported in Table 6), for the same bilateral comparisons analyzed in Table 3 at the same selected end-years analyzed in Table 5, show that we continue to reject the null hypothesis of equality of the short-run impact of a common monetary shock among all countries (see p-values on cumulative impacts within six and 12 months in the upper part of Table 6), as well as the null of equality of the long-run impact between France and the other three countries (see p-values on cumulative impacts within six and 12 months in the lower part of Table 6), right up to end-1998 before the EMU launch.

5 Conclusions

In this paper we have studied empirically the transmission mechanism of a common, homoskedastic monetary policy shock, identified as an innovation to the reaction function of the Bundesbank, in the four largest European countries currently in the EMU by using time-varying, heterogenous models, estimated in a Bayesian fashion based on per-EMU data.

The analysis documented in the paper shares several features of the 'ideal experiment' described by Guiso et al. (2000): the model specification is the same for all countries considered; intra-Europe

of the output response to the shock are in the ballpark of those reported for the United States before and after the Volcker experiment: they are smaller than those reported by Bovin and Giannoni (2002 and 2003) but larger than those reported by Primiceri (2002).

exchange rate movements as well as differences in central banks objectives and procedures, which have disappeared under EMU, are controlled for; regional interdependence, through which monetary policy in part operates, is also allowed for; and, most importantly, both model parameters and shock variances are allowed to change over time.

We have found that the transmission mechanism of monetary policy has changed in the run-up to EMU, in all countries considered, albeit not dramatically, with the long-run cumulative impact of a common monetary shock on output decreasing by about 10-20 percent in all countries after 1991. These changes are statistically significant and associated with some shifts in the behavior of the monetary authorities in all countries considered. They are particularly evident in the case of Germany and less clear cut in the case of Italy. Interestingly, we have also found that cross-country differences in the effects of this shock have not decreased over the same period.

These results are not only consistent with the predictions of the Lucas' critique, but also with two separate bodies of existing empirical evidence. First, they are consistent with the now large, albeit sometime contradicting, literature on cross-country differences in the transmission mechanism of European monetary policy. Second, our results are also in line with the most recent literature on the evolution of U.S. monetary policy and its effects on the economy.

We conclude from this evidence that the transmission mechanism of European monetary policy is probably changing over time, albeit slowly and in all countries at the same time. Thus, the remaining differences might be here to stay, as suggested by Cecchetti (1999).

A Appendix: Estimation procedures

In this appendix we present in more details the Bayesian procedures used for the estimation of the systems of reaction functions (6) and output equations (8).

A.1 Reaction functions

Denoting R_j^T the data sample for each block j of equation (6), the data density conditional on the lagged endogenous and exogenous variables (Z_{jt}), the initial observation (R_{j0}) and the parameters of the model (δ_{jt}, Σ_{jj}, and σ_{jt}), i.e., the likelihood function $L\left(R_j^T \mid Z_{jt}, R_{j0}, \delta_{jt}, \Sigma_{jj}, \sigma_{jt}\right)$, is proportional ($\propto$) to:

$$|\sigma_{jt}\Sigma_{jj}|^{-T/2} \exp\left[-\frac{1}{2}\sum_t \left(R_{jt} - Z_{jt}\delta_{jt}\right)' \left(\sigma_{jt}\Sigma_{jj}\right)^{-1} \left(R_{jt} - Z_{jt}\delta_{jt}\right)\right]. \tag{12}$$

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 prior (e.g. Doan et al. 1984) for the slope coefficients (δ_{jt}) is combined with a diffuse prior on the variance-covariance matrix of the residuals (Σ_{jj}), and a proper prior for the heteroskedasticity factor (σ_{jt}), with prior independence. Thus:

$$p\left(\delta_{jt}, \Sigma_{jj}, \sigma_{jt}\right) = p\left(\delta_{jt}\right) p\left(\Sigma_{jj}\right) p\left(\sigma_{jt}\right), \tag{13}$$

where

$$\begin{aligned} p\left(\Sigma_{jj}\right) &\propto |\Sigma_{jj}|^{-(G_j+1)/2}, & (14)\\ \delta_{jt} &= \delta_{jt-1} + \eta_{jt} \qquad \eta_{jt} \sim N\left(0, \Phi_j\right), & (15)\\ \sigma_{jt} &\sim \text{Inv-}\chi^2\left(\nu_j, 1\right). & (16) \end{aligned}$$

Here Φ_j governs the variance of η_{jt}, which in turn controls the time variation of δ_{jt}.

In order to use the Gibbs sampler, we need to derive the *conditional* posterior distributions of all parameters from the product of the likelihood and the priors. The posterior distribution of Σ_{jj}^{-1} conditional on the entire history of δ_{jt} for $t = 0, ...T$ (which we denote $\{\delta_{jt}\}$), σ_{jt}, and the data, is obtained from the combination of the likelihood (12) with the prior $p\left(\Sigma_{jj}\right)$ as the following Wishart distribution with $T - G_j - 1$ degrees of freedom and scale matrix S (Zha, 1999, p. 299):

$$\Sigma_{jj}^{-1} \mid \{\delta_{jt}\}, \sigma_{jt}, R_j^T, R_{j0} \sim W\left(T - G_j - 1, S\right), \tag{17}$$

where

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

The conditional posterior distribution of σ_{jt} is obtained by combining the likelihood (12) with the prior $p(\sigma_{jt})$ as the following Inverse-χ^2 distribution with $\nu_j + T$ degrees of freedom and scale matrix s_j^2 (Gelman et al., 1995, p. 350):

$$\sigma_{jt} \mid \{\delta_{jt}\}, \Sigma_{jj}, R_j^T, R_{j0} \sim \text{Inv-}\chi^2\left(\nu_j + T, s_j^2\right) \tag{18}$$

where

$$s_j^2 = \frac{\left[\nu_j + (R_{jt} - Z_{jt}\delta_{jt})' \Sigma_{jj}^{-1} (R_{jt} - Z_{jt}\delta_{jt})\right]}{\nu_j}.$$

Finally, the joint conditional posterior distribution of $\{\delta_{jt}\}_t$ is obtained in two steps as shown by Chib and Greenberg (1995, pp. 349-350).

First, we initialize $\{\delta_{jt}\} \mid \Sigma_{jj}, \sigma_{jt}$ by Kalman Filter and save its output $\left\{\widehat{\delta}_{jt}, \hat{\Omega}_{jt|t}, F_t\right\}$ for each t:

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

where

$$\begin{aligned}
\hat{\delta}_{jt|t-1} &= \hat{\delta}_{jt-1|t-1} \\
\hat{\Omega}_{jt|t-1} &= \hat{\Omega}_{jt-1|t-1} + \Phi_j.
\end{aligned}$$

Second, we sample $p(\{\delta_{jt}\} \mid \Sigma_{jj}, \sigma_{jt})$ *in reverse time order* from

$$\begin{aligned}
\delta_{jT} &\sim N\left(\hat{\delta}_{jT|T}, \hat{\Omega}_{jT|T}\right) \\
\delta_{jT-1} &\sim N\left(\hat{\delta}_{jT-1}, \hat{\Omega}_{jT-1}\right) \\
&\vdots \\
\delta_{j0} &\sim N\left(\hat{\delta}_{j0}, \hat{\Omega}_{j0}\right)
\end{aligned} \tag{19}$$

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

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

Given (17), (18) and (19), and initial values for Φ_j, $\hat{\Omega}_{j0}$, and $\hat{\delta}_{j0}$, the *marginal* posterior distributions of Σ^{-1}_{jj}, σ_{jt} and $\{\delta_{jt}\}$ can then be obtained from the Gibbs sampler by drawing alternately from these conditional distributions. We obtain initial values of Φ_j, $\hat{\Omega}_{j0}$, and $\hat{\delta}_{j0}$ from a classical SUR estimation of (6). The degrees of freedom ν_j is set equal to 5, which allow for the maximum degree of departure from normality.[15]

The Gibbs sampler cycles through (17)-(19) 10,000 times, thus generating 10,000 draws from the marginal distributions. We select 500 draws (i.e., one every 20 cycles) and discard the first 100. The remaining 400 draws are then used for inference. We check convergence to the ergodic distribution with standard criteria.

A.2 Output equations

A.2.1 Time-invariant model

The likelihood of model (10) is proportional to:

$$|\Omega|^{-T/2} \exp\left\{-\frac{1}{2}\sum_{t=1}^{T}(\mathsf{Y}_t - \mathsf{X}_t\beta)'\,\Omega^{-1}\,(\mathsf{Y}_t - \mathsf{X}_t\beta)\right\}.$$

Assuming prior independence, the prior assumptions we make on this model's hyperparameters are

$$p\left(\Omega^{-1}, \bar{\beta}, B_o\right) = p\left(\Omega^{-1}\right) p\left(\bar{\beta}\right) p\left(B_o\right), \tag{20}$$

where

$$\Omega^{-1} \sim W(\omega_o, \Theta), \tag{21}$$

$$\bar{\beta} \sim N(\mu, B_1) \text{ with } B_1^{-1} = 0 \tag{22}$$

$$B_o = \tau^2 I_h, \qquad \tau^2 \sim \text{Inv-}G(\kappa_o/2, \xi_o/2), \tag{23}$$

Here, $W(\omega_o, \Theta)$ denotes a Wishart distribution with ω_o degrees of freedom and scale matrix Θ, Inv-$G(\kappa_o/2, \xi_o/2)$ is an Inverse-gamma distribution with shape $\kappa_o/2$ and scale $\xi_o/2$, with I_h denoting

[15] The extent to which v_{jt} in (6) departs from normality depends on the number of degrees of freedom, ν, the more so the larger ν. In fact, $v_{jt} \mid \sigma_{jt}$ converges in distribution to $N(0, \Sigma)$ as ν approaches infinity, because the mean of σ_{jt} tends to one and its variance tends to zero in the limit if $\sigma_{jt} \sim$Inv-$\chi^2(\nu_j, 1)$.

an identity matrix of dimension h. Note that by assuming $B_1^{-1} = 0$, we give a diffuse prior on the variance of $\bar{\beta}$ at the last stage of the hierarchy (22), without having to make any assumption on μ.

As in the previous case, we obtain the conditional posterior distributions of the parameters of interest by combining the likelihood of the data with the prior distributions above. Denote $\mathbf{Y} = (\mathsf{Y}_1, ..., \mathsf{Y}_T)$ the data sample, $\psi = (\beta, \Omega, \bar{\beta}, \tau^2)$ the model parameters, and with $\psi_{-\gamma}$ the model parameters excluding the generic parameter γ. It can be shown (Chib and Greenberg, 1995, pp. 348-349) that, under the assumptions made, the conditional posterior distributions of the model parameters are given by:

$$\beta \mid \psi_{-\beta}, \mathbf{Y} \sim N\left(\hat{\beta}, V_T\right); \tag{24}$$

$$\Omega^{-1} \mid \psi_{-\Omega}, \mathbf{Y} \sim W\left(\omega_o + T, \Theta_T\right); \tag{25}$$

$$\bar{\beta} \mid \psi_{-\bar{\beta}}, \mathbf{Y} \sim N\left(\left(M_o' B_o^{-1} M_o\right)^{-1}\left(M_o' B_o^{-1} \beta\right), \left(M_o' B_o^{-1} M_o\right)^{-1}\right); \tag{26}$$

$$\tau^2 \mid \psi_{-\tau^2}, \mathbf{Y} \sim \text{Inv-}G\left(\frac{(\kappa_o + h)}{2}, \frac{\xi_o + (\beta - M_o\bar{\beta})'(\beta - M_o\bar{\beta})}{2}\right); \tag{27}$$

where

$$\hat{\beta} = V_T\left(B_o^{-1} M_o \bar{\beta} + \sum_t \mathsf{X}_t' \Omega^{-1} \mathsf{Y}_t\right), \qquad V_T = \left(B_o^{-1} + \sum_t \mathsf{X}_t' \Omega_t^{-1} \mathsf{X}_t\right)^{-1}, \tag{28}$$

$$\Theta_T = \left[\Theta^{-1} + \sum_{t=1}^{T} (\mathsf{Y}_t - \mathsf{X}_t\beta)(\mathsf{Y}_t - \mathsf{X}_t\beta)'\right]^{-1}. \tag{29}$$

To initialize the Gibbs sampler, all hyperparameters of the model (i.e., $\omega_o, \Theta, \kappa_o$, and ξ_o) must be known. We set $\omega_o = G + 1$, and initialize Θ with the variance-covariance matrix of a classical SUR estimate of (8). The parameters of the gamma distributions of τ^2 are $\kappa_o = 6$ and $\xi_o = 1$, implying that the prior mean and the standard deviation of τ^2 are 0.25 and 0.25, respectively. With this assumption the amount of time variation assumed a prior is remarkable. Finally, we also set $\tau^2 = 0.5$ and $\Omega = I_g$, initially.

At each iteration of the Gibbs sampler, dynamic multipliers of $\widehat{u}_t$ are computed as non-linear functions of the elements of β, and used to compute the cumulative effects of this shock after 6, 12 and 24 months, and a theoretically infinite number of months.

The Gibbs sampler runs through (24)-(29) 16,000 times, thus generating 16,000 draws from the marginal distributions. We select 500 draws (one every 20 cycles), and discards the first 300. The remaining 400 draws are then used for inference, after checking convergence.

A.2.2 Time-varying model

The likelihood of the time-varying model (9) is:

$$|\varsigma_t \Omega|^{-T/2} \exp\left\{-\frac{1}{2}\sum_{t=1}^{T}(\mathsf{Y}_t - \mathsf{X}_t\beta_t)'(\varsigma_t\Omega)^{-1}(\mathsf{Y}_t - \mathsf{X}_t\beta_t)\right\}.$$

The prior assumptions for the hyperparameters of this model are:

$$p\left(\Omega^{-1}, \varsigma_t, B_2\right) = p\left(\Omega^{-1}\right) p(\varsigma_t)\, p(B_2),$$

where

$$\begin{aligned}
\Omega^{-1} &\sim W(\omega_o, \Theta), \\
\varsigma_t &\sim \text{Inv-}\chi^2(\varrho_o, 1), \\
B_2 &= \phi^2 I_h \qquad \phi^2 \sim \text{Inv-}G(\kappa_o/2, \xi_o/2).
\end{aligned}$$

There are two differences between the procedure for the estimation of the system of reaction functions and the system of output equations. First, here we assume a natural-conjugate prior for Ω (i.e., a Wishart distribution), rather than a normal-diffuse prior, so that its posterior distribution is also Wishart. Notice, however that a diffuse prior could be set by taking $\omega_o \to 0$, which is the case adopted in the reaction functions. Here anyway we set ω_o to the smallest number which ensure a prior which is proper, though providing vague initial information relative to that to be provided by the data. Second, here we assume a random variance-covariance matrix ($B_2 = \phi^2 I_h$) for ζ_t, because we want a proper prior to test for overall stability of β_t, as explained in the text.

The *conditional* posterior densities of the parameters of interest are obtained by combining the likelihood of the data with the prior distributions above. It can be shown that the posterior distributions of Ω^{-1}, ς_t, and ϕ^2 are, respectively, given by:

$$\Omega^{-1} \mid \psi_{-\Omega}, \mathbf{Y} \sim W(T + \omega_o, \Theta_T); \tag{30}$$

$$\varsigma_t \mid \psi_{-\varsigma}, \mathbf{Y} \sim \text{Inv-}\chi^2\left(\varrho_o + T, \bar{\varsigma}^2\right) \tag{31}$$

$$\phi^2 \mid \psi_{-\phi^2}, \mathbf{Y} \sim \text{Inv-}G\left(\frac{(T\kappa_o + h)}{2}, \frac{\xi_o + \sum_t(\beta_t - \beta_{t-1})'(\beta_t - \beta_{t-1})}{2}\right) \tag{32}$$

where

$$\Theta_T = \left[\Theta^{-1} + \frac{\sum_t(\mathsf{Y}_t - \mathsf{X}_t\beta_t)(\mathsf{Y}_t - \mathsf{X}_t\beta_t)'}{\varsigma_t}\right]^{-1} \tag{33}$$

and

$$\bar{\varsigma}^2 = \frac{\left[\varrho_o + (\mathsf{Y}_t - \mathsf{X}_t\beta_t)' \, \Omega^{-1} \, (\mathsf{Y}_t - \mathsf{X}_t\beta_t)\right]}{\varrho_o}. \tag{34}$$

The joint posterior distribution of $\{\beta_t\}_t$ conditional on all other parameters and the data is then obtained in two steps, as discussed for the system of reaction functions.

To initialize the hyperparameters of the model $(\omega_o, \Theta, \varrho_o, , \kappa_o, \xi_o)$, we set $\omega_o = G+1$ and assign to Θ the variance-covariance matrix of a classical SUR estimate of (8). The parameters of the gamma distribution of ϕ^2 are $\kappa_o = 6$ and $\xi_o = 1$, implying that the prior mean and standard deviation of ϕ^2 are 0.25 and 0.25, respectively. The degrees of freedom ϱ_o is set equal to 5. To initialize the Gibbs sampler, we also set $\phi^2 = \varsigma_t = 0.5$ for each t and $\Omega = I_g$, while the initial conditions β_0 are drawn from a Normal distribution, centered on the estimates of the time invariant version of the model.

In this case, the Gibbs sampler runs through (30)-(34) and the recursive scheme for β_t 15,000 times, and we select 500 (independent) draws, one every 30 cycles, retaining only the last 300 iterations, after having checked convergence with standard criteria.

References

[1] Amisano G. and C. Giannini (1997), *Topics in Structural VAR Econometrics* (Second ed.), Springer Verlag, Berlin.

[2] Angeloni, I., A. Kashyap, B. Mojon, and D. Terlizzese (2002), "Monetary Transmission in the Euro Area: Where Do We Stand?," European Central Bank, Working Paper No. 114, (Frankfurt, Germany).

[3] Bernanke, B. and I. Mihov (1997), "What does the Bundesbank target?," *European Economic Review*, Vol. 41, pp. 1025–53.

[4] Boivin, J. (2001), "The Fed's Conduct of Monetary Policy: has it Changed and Does it Matter?", Unpublished manuscript (Columbia University: New York).

[5] Boivin, J. and M. Giannoni (2003), "Has Monetary Policy Become More Effective?," NBER Working Paper No. 9459, (Cambridge, Massachusetts: National Bureau of Economic Research).

[6] Boivin, J. and M. Giannoni , (2002), "Assessing Changes in the Monetary Transmission Mechanism: A VAR Approach," Economic Policy Review, Federal Reserve Bank of New York, (May).

[7] Canova F. and M. Ciccarelli (2003), 'Forecasting and Turning Point Prediction in a Bayesian Panel VAR Model', *Journal of Econometrics*, forthcoming.

[8] Cecchetti S.G. (1999), 'Legal Structure, Financial Structure and Monetary Policy', NBER WP no. 7151.

[9] Chib S. (1995), Marginal Likelihood from the Gibbs Output, *Journal of the American Statistical Association*, Vol. 90, pp. 1313-1321.

[10] Chib S. and E. Greenberg (1995), 'Hierarchical Analysis of SUR Models with Extensions to Correlated Serial Errors and Time-Varying Parameter Models', *Journal of Econometrics*, Vol. 68, pp. 409–431.

[11] Ciccarelli M. and A. Rebucci (2001), 'Asymmetries in the Transmission Mechanism of European Monetary Policy', *Temi di Ricerca n.* 23, Ente Einaudi (Roma: Italy).

[12] Ciccarelli M. and A. Rebucci (2001), "*The Transmission Mechanism of European Monetary Policy: Is There Heterogeneity? Is It Changing over Time*?" Bank of Spain WP No. 0115 and IMF WP No. 02/54.

[13] Clarida R., J. Gali, and M. Gertler (1998), 'Monetary Policy Rules in Practice. Some International Evidence', *European Economic Review,* Vol. 42, pp. 1033-1067.

[14] Clarida R., J. Gali, and M. Gertler (2000), Monetary Policy Rules and Macroeconomic Stability: Evidence and Some Theory, *Quarterly Journal of Economics,* Vol. 115, No. 1, pp. 147-80.

[15] Clements B., Z. Kontolemis and J. Levy (2001), 'Monetary Policy under EMU: Differences in the Transmission Mechanism?', IMF WP/01/102, Washington DC.

[16] Cogley T. and T.J. Sargent (2002), "Drifts and Volatilities: Monetary Policies and Outcomes in the Post WWII U.S.," unpublished manuscript*,* Arizona State University.

[17] Doan T., Litterman R., and C. Sims (1984), 'Forecasting and Conditional Projections Using Realist Prior Distributions', *Econometric Review*, Vol. 3, No. 1, pp. 1-100.

[18] Dornbusch R., C. Favero, and F. Giavazzi (1998), 'A Red Letter Day', CEPR Discussion Paper No. 1804. Published also as 'The Immediate Challenge for the European Central Bank', *Economic Policy*, 26, 17-64.

[19] Favero C. A. and M. Marcellino (2001), Large Datasets, Small Models and Monetary Policy in Europe,IGIER Working Paper No. 208 (Bocconi University: Milan, Italy).

[20] Frankel J. and A. Rose (1998) "The Endogeneity of the Optimum Currency Area Criteria", *Economic Journal.*

[21] Gelfand A.E., S. E. Hills, A. Racine-Poon, and A.F.M. Smith (1990), 'Illustration of Bayesian Inference in Normal Data Models Using Gibbs Sampling', *Journal of the American Statistical Association*, Vol. 85, pp. 972-985.

[22] Gelman A., J.B. Carlin, H.S. Stern and D.B. Rubin (1995), *Bayesian Data Analysis*, Chapman & Hall, London.

[23] Guiso L., A.K. Kashyap, F. Panetta, and D. Terlizzese (2000), 'Will a common European Monetary Policy Have Asymmetric Effects?', *Economic Perspectives,* Federal Reserve Bank of Chicago.

[24] Leeper, E. M., C. A. Sims, and T. Zha (1996), "What Does Monetary Policy Do?," *Brookings Papers on Economic Activity*, Vol. 2, pp. 1–78.

[25] Lucas, R. E., Jr. (1976), "Econometric Policy Evaluation: A Critique," in "The Phillips Curve and Labor Markets," edited by Karl Brunner and Alan Meltzer, *Carnegie-Rochester Series on Public Policy*, Vol. 1.

[26] Lütkepohl, H., *Introduction to Multiple Time Series Analysis*, Berlin, Springer–Verlag, 1990.

[27] Peersman G. and F. Smets (1998), 'The Taylor Rule: A Useful Monetary Policy Guide for the ECB?', paper presented at the conference *'Monetary Policy of the ESCB: Strategic and Implementation Issues*', Bocconi University, Milan.

[28] Primiceri, G. E. (2002), "Time Varying Structural Vector Autoregressions and Monetary Policy," (unpublished; New Jersey, N.J.: Princeton University).

[29] Sala L. (2001), 'Monetary Transmission in the Euro Area: A Factor Model Approach', unpublished manuscript (Bocconi University: Milan).

[30] Sims, C. A. (2001), "Comment on Sargent and Cogley's 'Evolving Post World War II U.S. Inflation Dynamics,' " *NBER Macroeconomics Annual* No. 16, pp. 373-379.

[31] Uhlig H., (1992), "Bayesian Vector Autoregressions with Stochastic Volatilities", *Econometrica*, Vol. 65, No. 1, pp. 59-73.

[32] Zha T. (1999), 'Block Recursion and Structural Vector Autoregressions', *Journal of Econometrics,* Vol. 90, 291-316.

[33] Zellner A. (1971), *An Introduction to Bayesian Inference in Econometrics*, Wiley, New York.

Figure 1. Monetary Shocks and Their Volatility
(January 1981-December 1998)

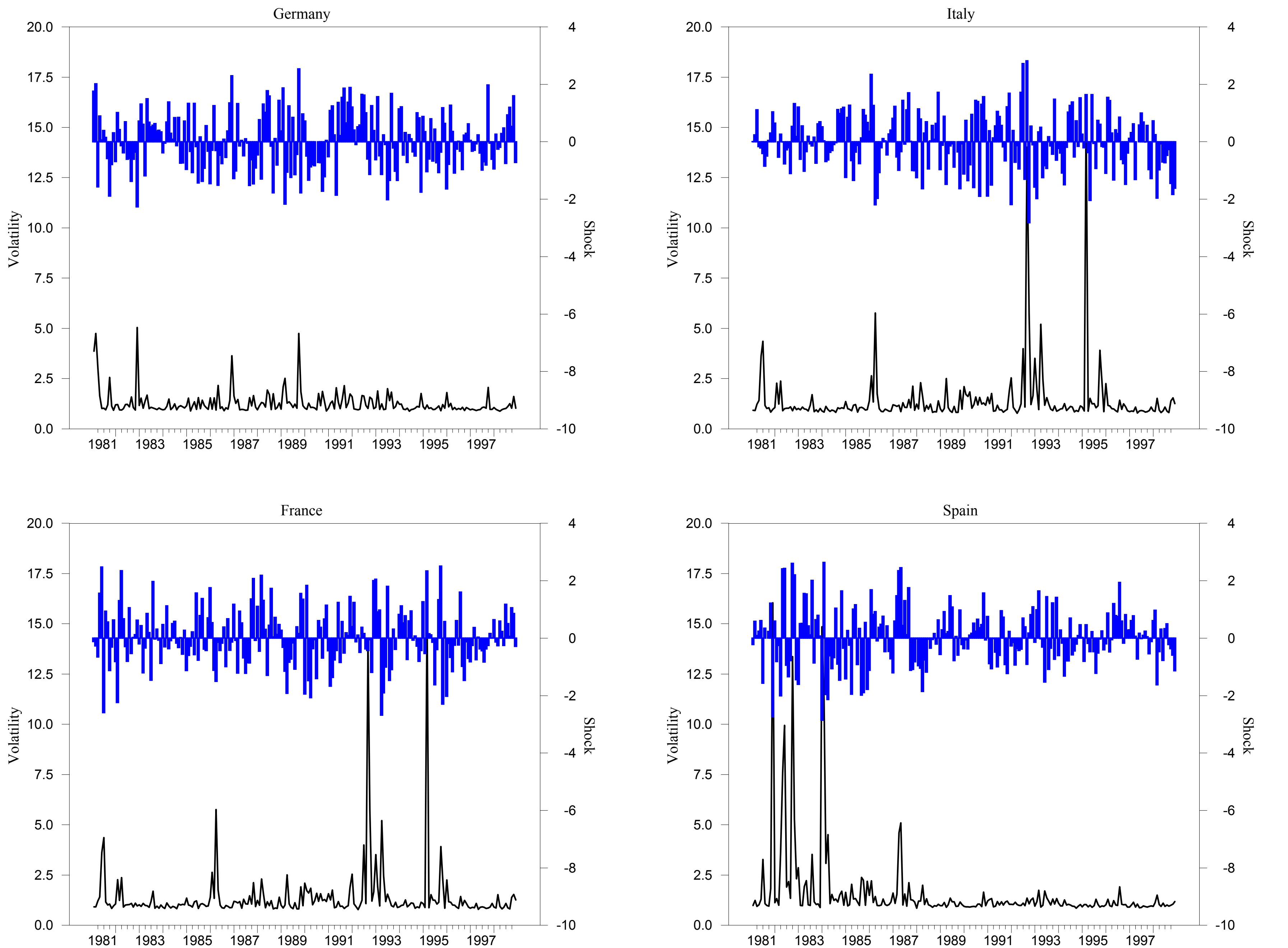

Figure 2. Germany: Reaction Function Parameters
(Elasticities, January 1981-December 1998)

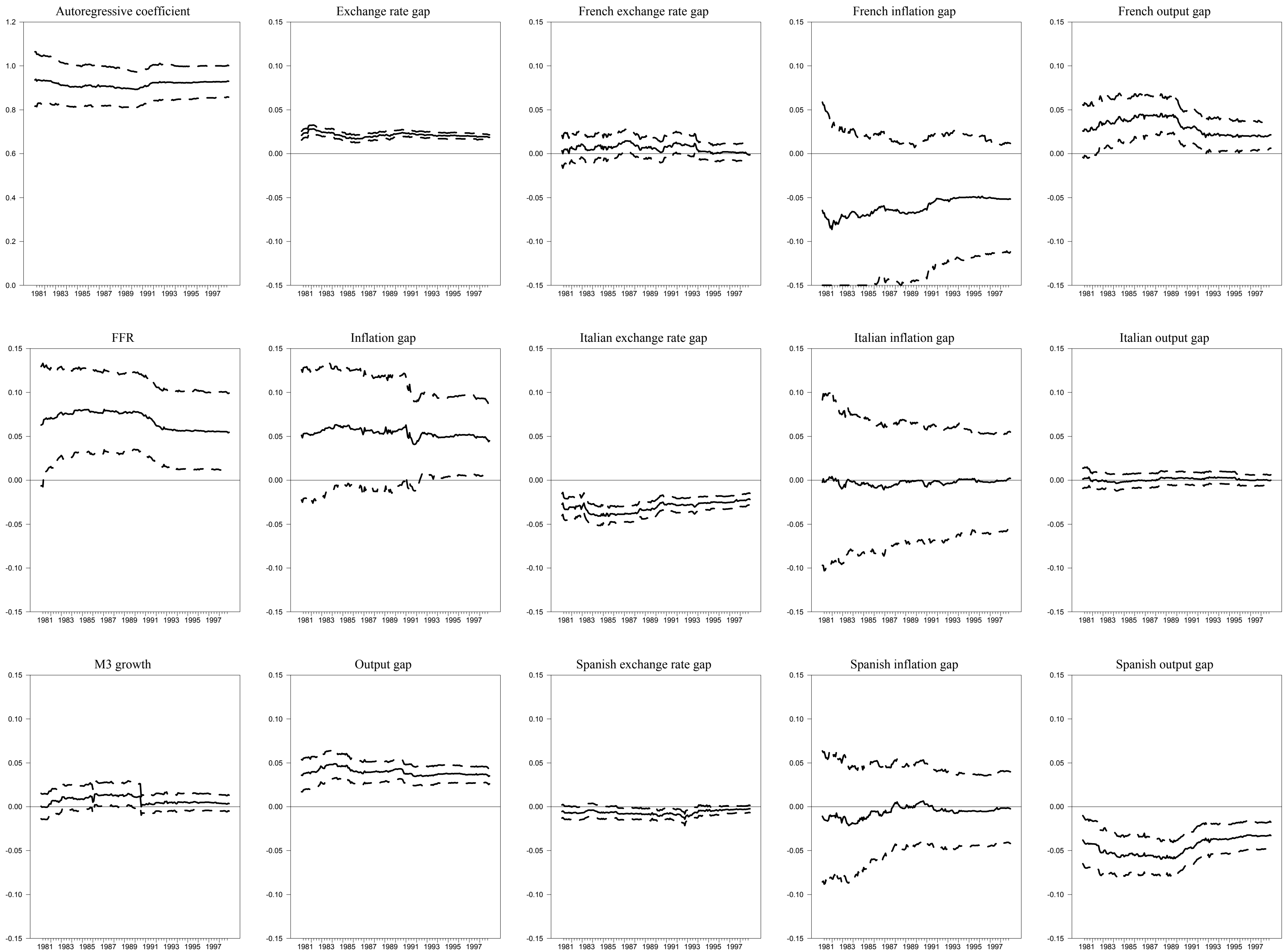

Figure 3. France: Reaction Function Parameters
(Elasticities, January 1981-December 1998)

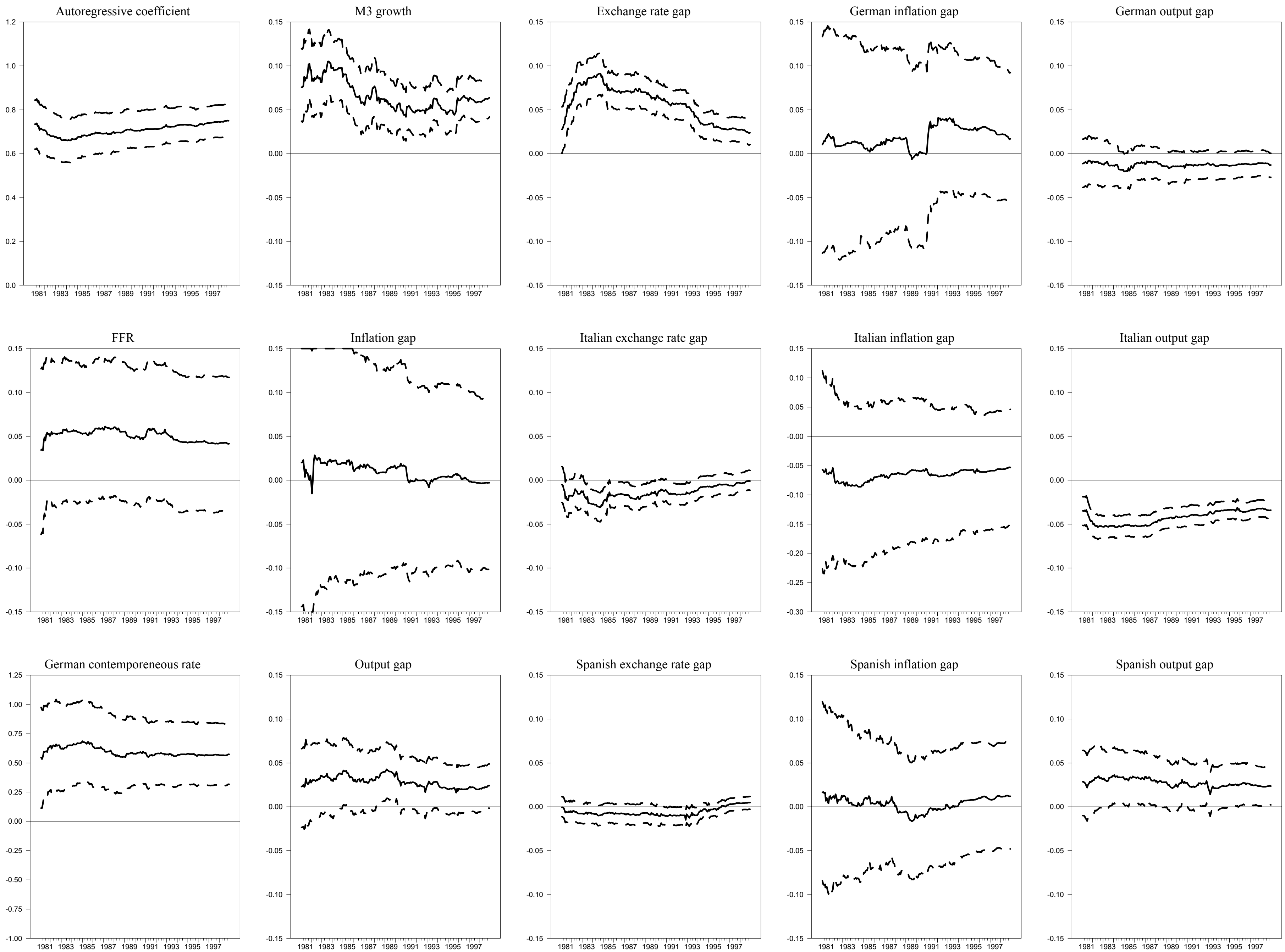

Figure 4. Italy: Reaction Function Parameters
(Elasticities, January 1981-December 1998)

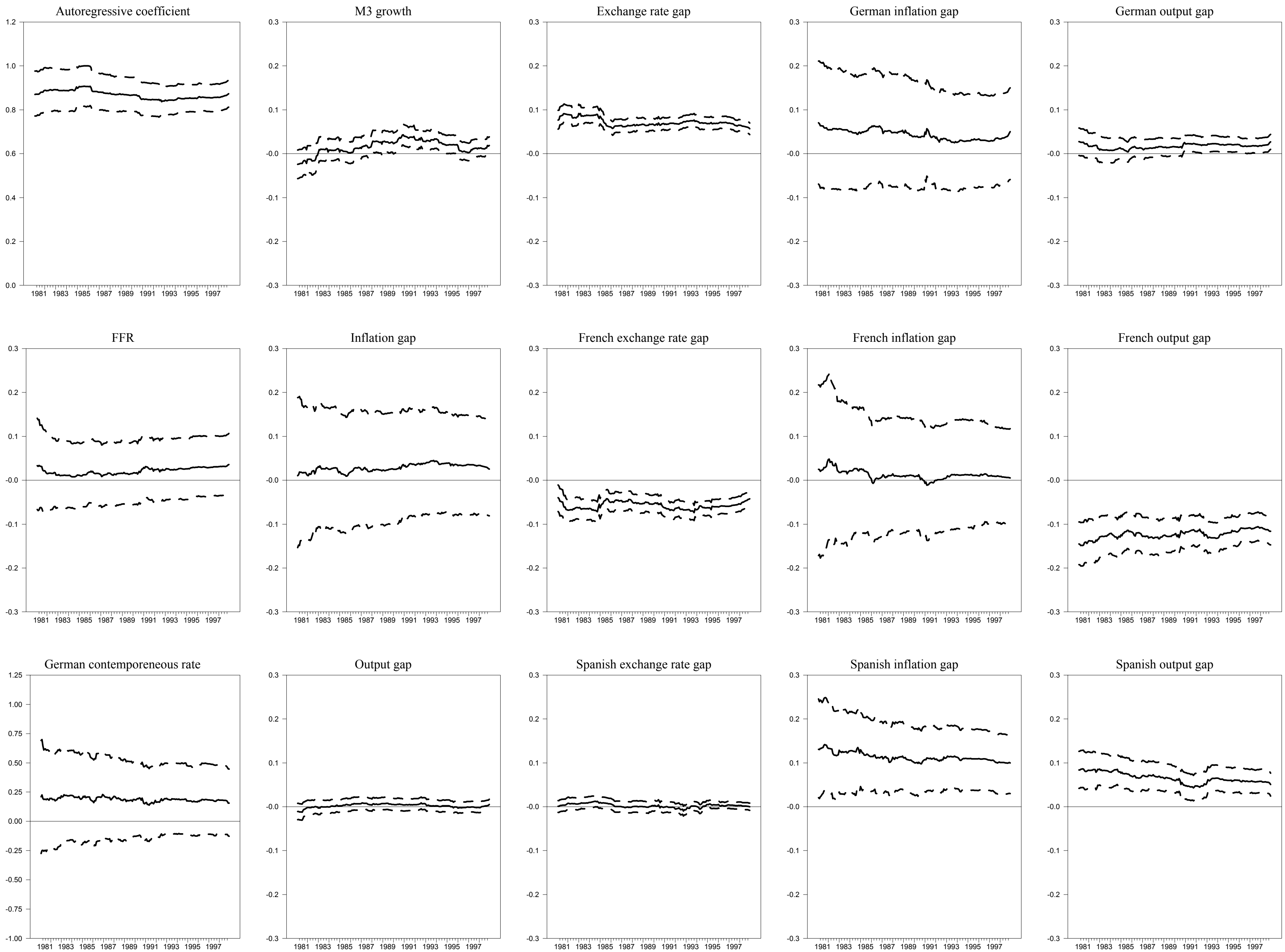

Figure 5. Spain: Reaction Function Parameters
(Elasticities, January 1981-December 1998)

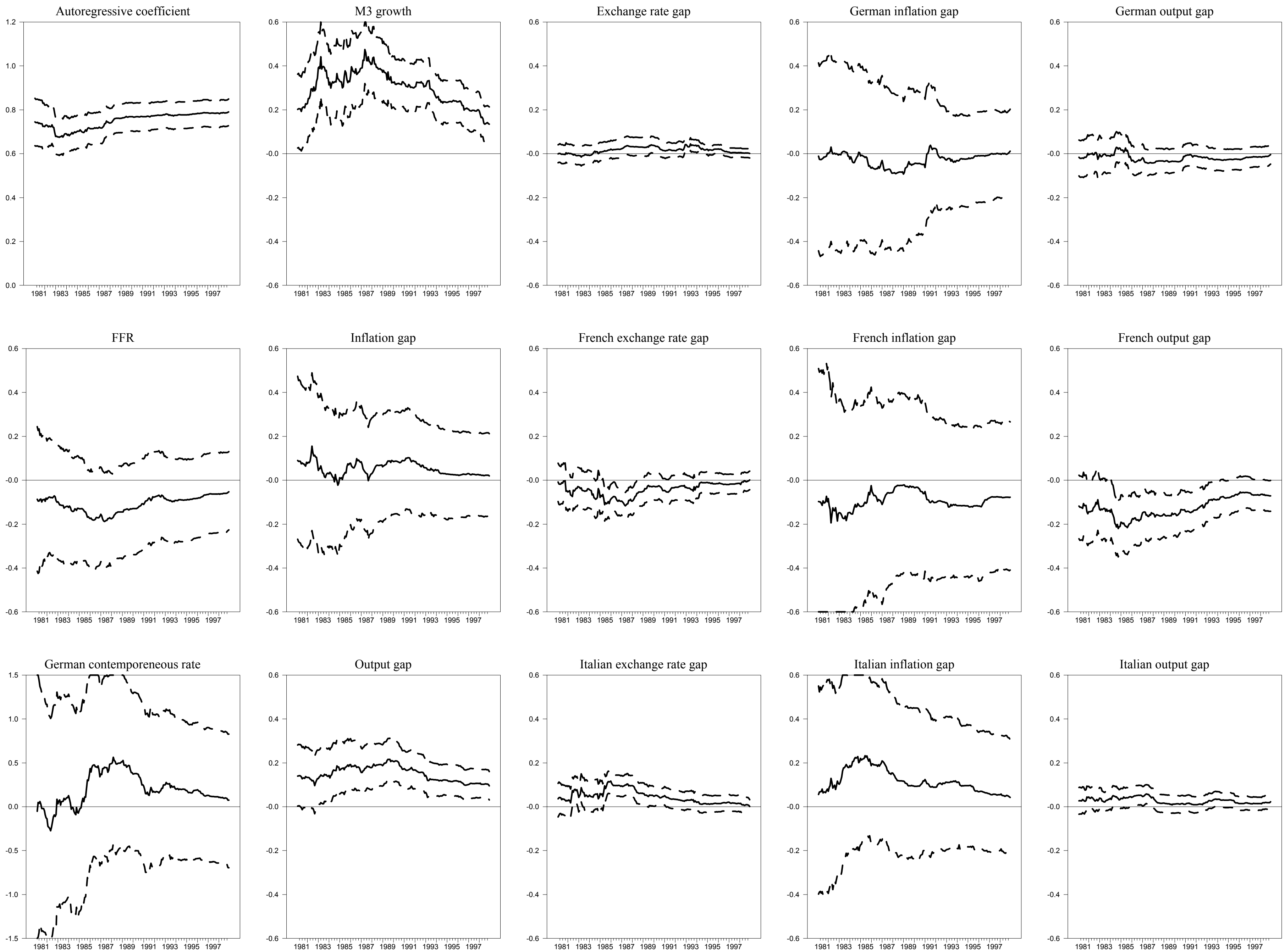

Figure 6. Responses to a Common Monetary Policy Shock
(Central 50 percent of the posterior distribution)

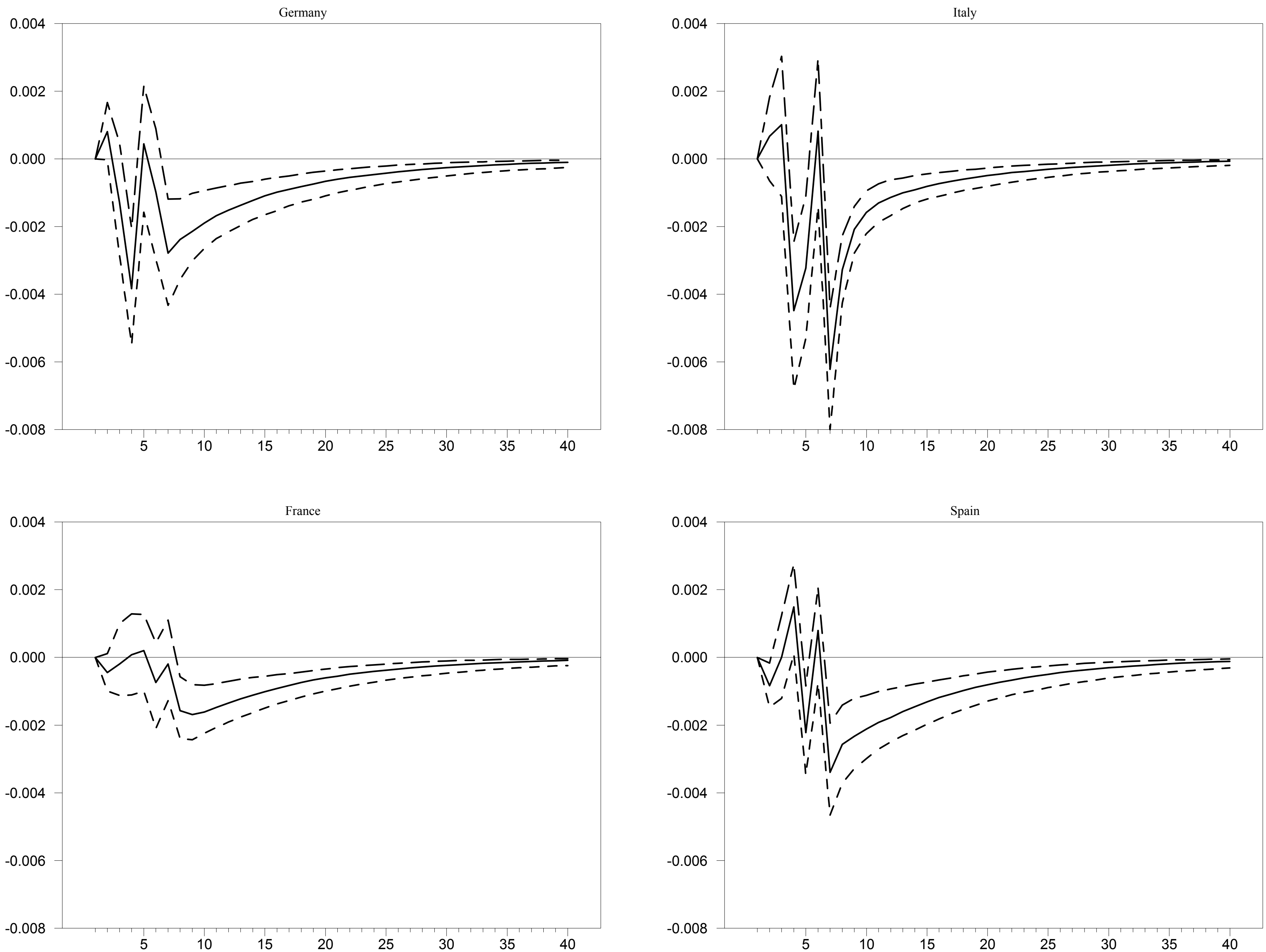

Figure 7a. Testing Homogeneity Across Countries

(Prior and Posterior Distribution of τ^2)

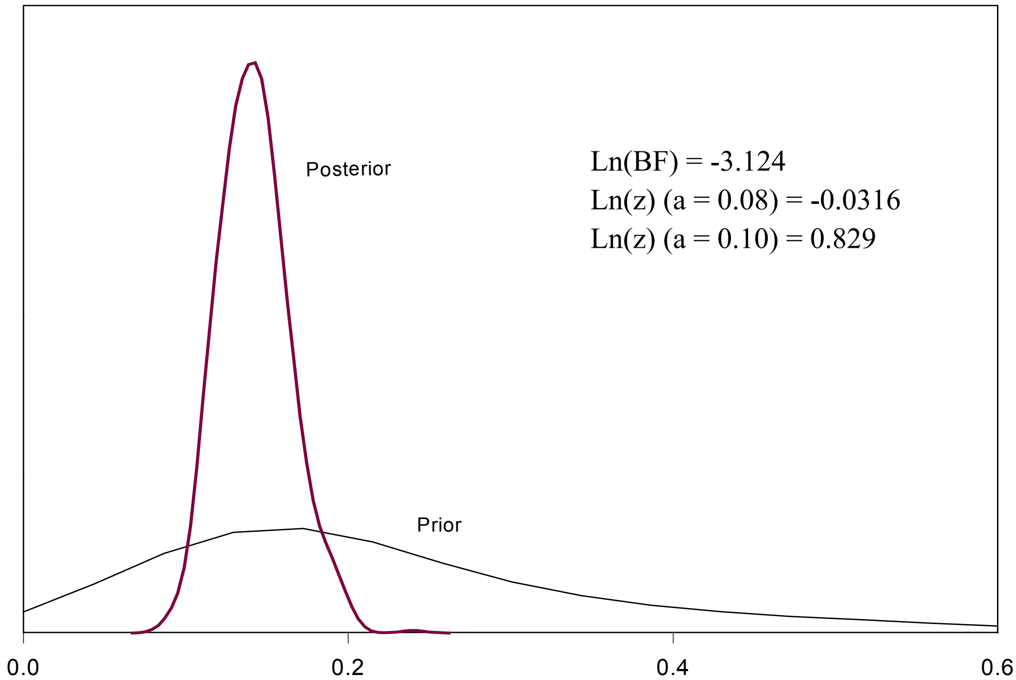

Figure 7b. Testing Homogeneity Across Countries Without Control Variables

(Prior and Posterior Distribution of τ^2)

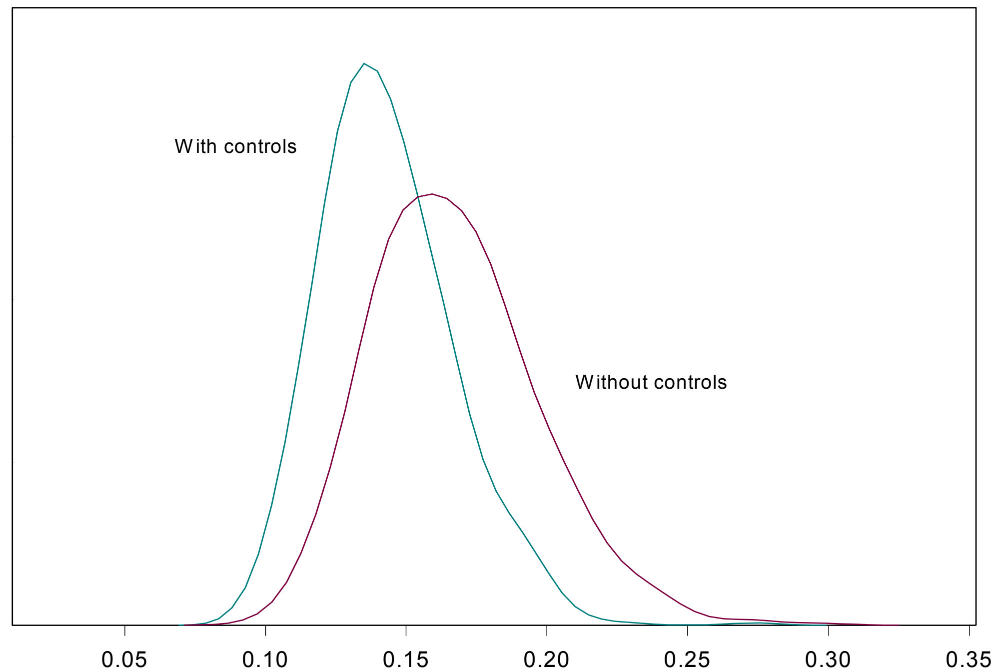

Figure 8. Time-Varying Responses to a Common Monetary Policy Shock
(Posterior medians. End of selected years: 1985, 1991, 1995, 1998)

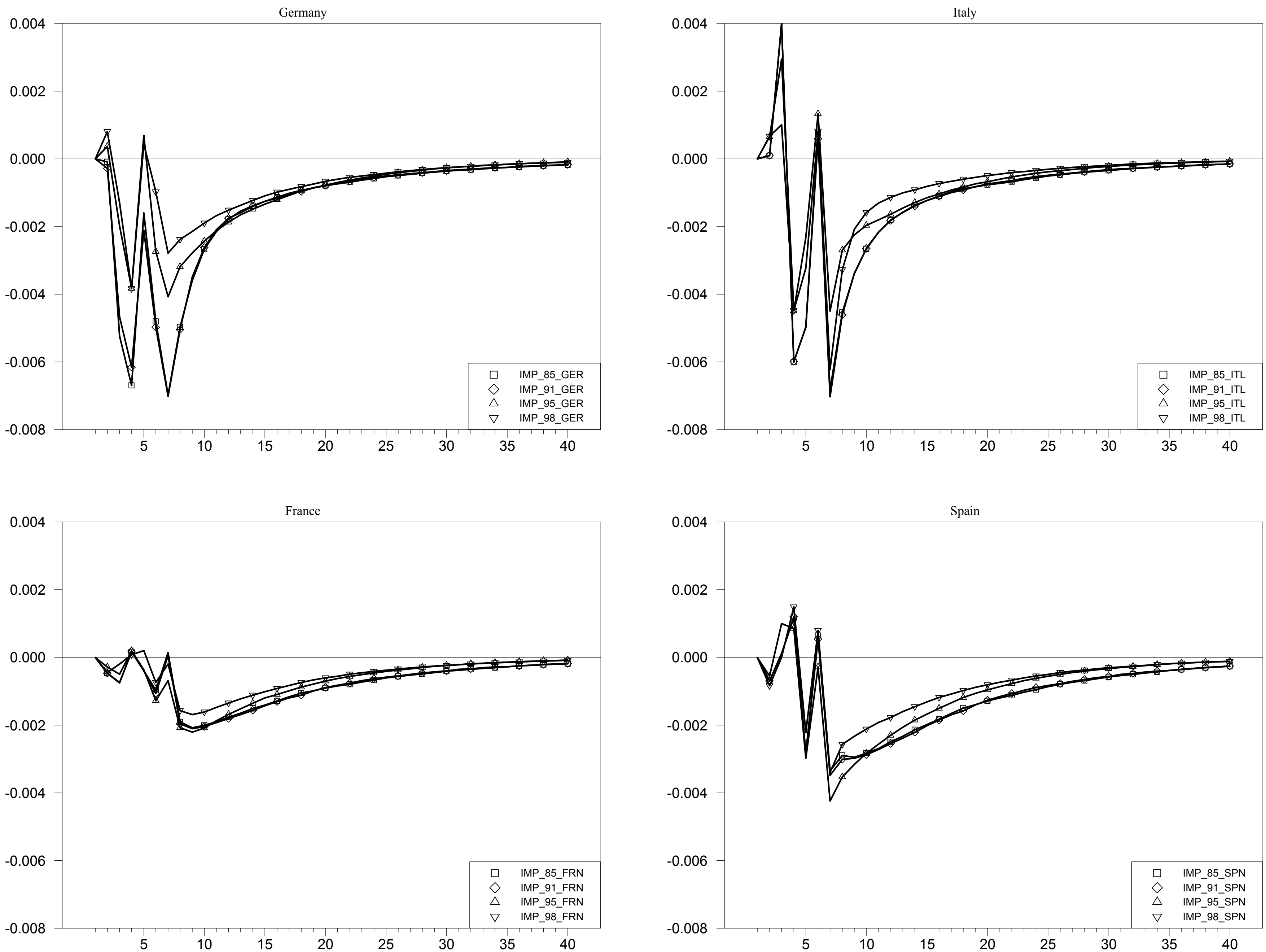

Figure 9. Testing Stability Over Time
(Prior and Posterior Distribution of ϕ^2)

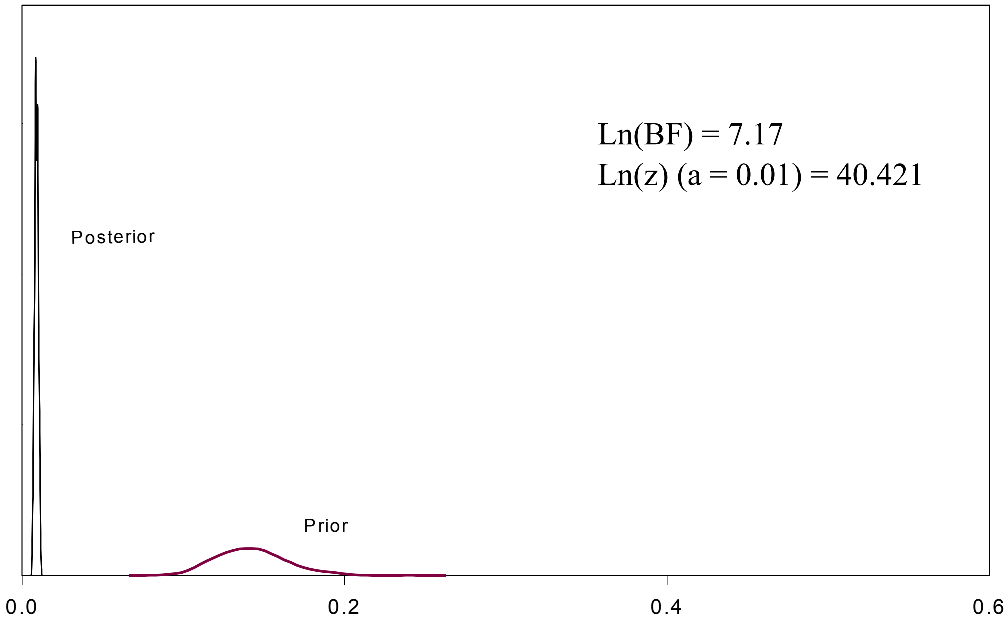

Table 1. Monetary Policy Shocks: Test Statistics

	Germany	France	Italy	Spain
Sample Mean	0.0053	0.0062	-0.0301	-0.0044
	(0.937)	(0.927)	(0.659)	(0.948)
Cum. Period. 1/	0.0816	0.0665	0.0617	0.0997
	(0.0981)	(0.0981)	(0.0981)	(0.0981)
Q(4)	6.2035	1.4762	2.3753	5.3435
	(0.184)	(0.831)	(0.667)	(0.254)
Q(8)	7.1921	8.3869	6.5074	12.1860
	(0.516)	(0.397)	(0.591)	(0.143)
Q(12)	7.8025	10.4300	16.4072	23.8513
	(0.800)	(0.578)	(0.173)	(0.021)
Arch(2)	2.7708	3.5593	7.1603	33.4121
	(0.250)	(0.169)	(0.029)	(0.000)

1/ Maximum gap between the actual and theoretical cumulated periodogram and approximate rejection limit at the 5 percent significance level (p-values in brackets).

Table 2. Comparing the Impact of a Common Monetary Policy Shock Across Countries
(1982-1998)

1/	Posterior Moments	Germany	France	Italy	Spain
6-Month	1st Quartile	-0.013	-0.005	-0.016	-0.009
	Mean	-0.008	-0.001	-0.012	-0.004
	Median	-0.008	-0.001	-0.012	-0.004
	3rd Quartile	-0.002	0.002	-0.006	0.001
12-Month	1st Quartile	-0.028	-0.017	-0.029	-0.025
	Mean	-0.018	-0.010	-0.022	-0.016
	Median	-0.019	-0.011	-0.022	-0.017
	3rd Quartile	-0.010	-0.003	-0.015	-0.006
24-Month	1st Quartile	-0.042	-0.030	-0.039	-0.042
	Mean	-0.028	-0.019	-0.029	-0.028
	Median	-0.029	-0.019	-0.029	-0.028
	3rd Quartile	-0.015	-0.007	-0.020	-0.013
Long-Run	1st Quartile	-0.048	-0.036	-0.044	-0.051
	Mean	-0.033	-0.024	-0.033	-0.034
	Median	-0.032	-0.022	-0.032	-0.032
	3rd Quartile	-0.017	-0.010	-0.021	-0.016

1/ Posterior Moments of the Cumulative Impact After a Selected Number of Months.

Table 3. Testing Homogeneity Across Countries
(1982-1998)

Test Statistics	6-Month		12-Month		24-Month		Long-Run	
	1 /	2 /	1 /	2 /	1 /	2 /	1 /	2 /
Germany versus France	-16.892 (0.0000)	0.396 (0.0000)	-12.377 (0.0000)	0.301 (0.0000)	-8.818 (0.0000)	0.236 (0.0000)	-7.151 (0.0000)	0.187 (0.0000)
Germany versus Italy	9.799 (0.0000)	0.223 (0.0000)	5.909 (0.0000)	0.164 (0.0000)	1.442 (0.1497)	0.106 (0.0007)	0.046 (0.9635)	0.091 (0.0051)
Germany versus Spain	-8.644 (0.0000)	0.197 (0.0000)	-2.828 (0.0048)	0.087 (0.0087)	-0.145 (0.8851)	0.046 (0.4294)	0.576 (0.5647)	0.040 (0.5983)
France versus Italy	29.459 (0.0000)	0.577 (0.0000)	21.034 (0.0000)	0.429 (0.0000)	11.952 (0.0000)	0.281 (0.0000)	8.447 (0.0000)	0.250 (0.0000)
France versus Spain	8.162 (0.0000)	0.220 (0.0000)	9.091 (0.0000)	0.257 (0.0000)	8.158 (0.0000)	0.237 (0.0000)	7.231 (0.0000)	0.209 (0.0000)
Italy versus Spain	-19.574 (0.0000)	0.400 (0.0000)	-8.990 (0.0000)	0.223 (0.0000)	-1.513 (0.1306)	0.141 (0.0000)	0.606 (0.5444)	0.116 (0.0001)

1 / Modified t-test for equal mean of two distributions (P-values in brackets).

2 / Kolmogorov-Smirnov test for equality of two distributions (P-values in brackets).

Table 4. Comparing the Impact of a Common Monetary Policy Shock Across Countries and Over Times (Selected Years) 1/

Countries	Posterior Moments	6-Month				12-Month			
		1985	1991	1995	1998	1985	1991	1995	1998
Germany	1st Quartile	-0.018	-0.019	-0.017	-0.013	-0.032	-0.034	-0.035	-0.028
	Mean	-0.013	-0.014	-0.012	-0.008	-0.026	-0.027	-0.025	-0.019
	Median	-0.014	-0.015	-0.012	-0.008	-0.026	-0.027	-0.026	-0.019
	3rd Quartile	-0.008	-0.009	-0.006	-0.003	-0.019	-0.021	-0.017	-0.011
France	1st Quartile	-0.002	-0.002	-0.007	-0.005	-0.012	-0.013	-0.021	-0.017
	Mean	-0.001	-0.001	-0.003	-0.001	-0.011	-0.011	-0.014	-0.010
	Median	-0.001	-0.001	-0.003	-0.001	-0.011	-0.011	-0.015	-0.011
	3rd Quartile	-0.001	-0.001	0.001	0.002	-0.009	-0.009	-0.007	-0.003
Italy	1st Quartile	-0.012	-0.012	-0.012	-0.016	-0.027	-0.027	-0.027	-0.029
	Mean	-0.011	-0.011	-0.006	-0.012	-0.025	-0.025	-0.018	-0.022
	Median	-0.011	-0.011	-0.007	-0.012	-0.025	-0.025	-0.018	-0.022
	3rd Quartile	-0.010	-0.010	-0.001	-0.006	-0.022	-0.023	-0.009	-0.015
Spain	1st Quartile	-0.005	-0.005	-0.011	-0.009	-0.020	-0.020	-0.033	-0.025
	Mean	-0.004	-0.004	-0.006	-0.004	-0.017	-0.018	-0.022	-0.016
	Median	-0.004	-0.004	-0.006	-0.004	-0.018	-0.018	-0.023	-0.017
	3rd Quartile	-0.003	-0.003	0.000	0.001	-0.015	-0.015	-0.012	-0.006

Countries	Posterior Moments	24-Month				Long-Run			
		1985	1991	1995	1998	1985	1991	1995	1998
Germany	1st Quartile	-0.042	-0.043	-0.049	-0.042	-0.048	-0.048	-0.054	-0.049
	Mean	-0.035	-0.036	-0.037	-0.029	-0.039	-0.041	-0.042	-0.033
	Median	-0.035	-0.036	-0.037	-0.029	-0.039	-0.040	-0.041	-0.033
	3rd Quartile	-0.027	-0.029	-0.024	-0.017	-0.032	-0.033	-0.027	-0.019
France	1st Quartile	-0.024	-0.024	-0.035	-0.030	-0.030	-0.030	-0.040	-0.037
	Mean	-0.021	-0.021	-0.024	-0.019	-0.026	-0.027	-0.029	-0.023
	Median	-0.021	-0.021	-0.025	-0.020	-0.026	-0.026	-0.028	-0.023
	3rd Quartile	-0.018	-0.018	-0.014	-0.007	-0.023	-0.023	-0.016	-0.010
Italy	1st Quartile	-0.037	-0.037	-0.041	-0.039	-0.042	-0.042	-0.046	-0.044
	Mean	-0.033	-0.034	-0.028	-0.030	-0.038	-0.038	-0.033	-0.033
	Median	-0.033	-0.034	-0.028	-0.029	-0.038	-0.038	-0.031	-0.032
	3rd Quartile	-0.030	-0.031	-0.016	-0.019	-0.034	-0.035	-0.018	-0.021
Spain	1st Quartile	-0.036	-0.036	-0.052	-0.042	-0.044	-0.044	-0.060	-0.051
	Mean	-0.032	-0.032	-0.037	-0.028	-0.040	-0.040	-0.043	-0.034
	Median	-0.032	-0.032	-0.037	-0.029	-0.039	-0.040	-0.042	-0.034
	3rd Quartile	-0.028	-0.028	-0.022	-0.012	-0.035	-0.035	-0.025	-0.016

1/ Posterior Moments of the Cumulative Impact After a Selected Number of Months.

Table 5. Testing Stability Over Time
(Selected Years)

Countries	Test Statistics	6-Month		12-Month		24-Month		Long-Run	
		1 /	2 /	1 /	2 /	1 /	2 /	1 /	2 /
Germany	1998 versus 1985	8.883	0.305	7.814	0.253	5.028	0.238	4.152	0.243
		(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
France		-0.5548	0.345	1.069	0.298	2.089	0.303	2.487	0.340
		(0.5793)	(0.0000)	(0.2855)	(0.0000)	(0.0372)	(0.0000)	(0.0132)	(0.0000)
Italy		-1.886	0.373	4.036	0.323	4.523	0.333	4.659	0.350
		(0.0600)	(0.0000)	(0.0001)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Spain		-1.398	0.395	1.563	0.290	3.318	0.310	3.847	0.360
		(0.1629)	(0.0000)	(0.1188)	(0.0000)	(0.001)	(0.0000)	(0.0001)	(0.0000)
		6-Month		12-Month		24-Month		Long-Run	
Countries	Test Statistics	1 /	2 /	1 /	2 /	1 /	2 /	1 /	2 /
Germany	1998 versus 1991	11.183	0.348	9.612	0.300	6.251	0.278	5.034	0.288
		(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
France		-0.183	0.3475	1.630	0.310	2.475	0.333	2.667	0.358
		(0.8549)	(0.0000)	(0.1039)	(0.0000)	(0.0137)	(0.0000)	(0.0079)	(0.0000)
Italy		-1.395	0.365	4.792	0.350	5.085	0.353	4.993	0.373
		(0.1638)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Spain		-1.061	0.393	2.135	0.318	3.722	0.335	4.034	0.378
		(0.2891)	(0.000)	(0.0333)	(0.000)	(0.0002)	(0.000)	(0.0001)	(0.0000)
		6-Month		12-Month		24-Month		Long-Run	
	Test Statistics	1 /	2 /	1 /	2 /	1 /	2 /	1 /	2 /
Germany	1998 versus 1995	6.008	0.218	6.462	0.220	5.605	0.195	4.584	0.183
		(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
France		3.3112	0.15	4.913	0.170	4.377	0.168	3.488	0.155
		(0.001)	(0.0002)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0005)	(0.0001)
Italy		-9.487	0.260	-4.915	0.175	-1.035	0.098	-0.223	0.080
		(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.3008)	(0.0388)	(0.8233)	(0.1379)
Spain		2.888	0.105	5.596	0.200	5.222	0.165	4.310	0.138
		(0.004)	(0.0209)	(0.000)	(0.0000)	(0.000)	(0.0000)	(0.000)	(0.0009)

1 / Modified t-test for equal mean of two distributions (P-values in brackets).

2 / Kolmogorov-Smirnov test for equality of two distributions (P-values in brackets).

Table 6. Testing Homogeneity Across Countries and Over Times
(Selected years)

	6-month								12-month							
	1985		1991		1995		1998		1985		1991		1995		1998	
Test Statistics	1/	2/	1/	2/	1/	2/	1/	2/	1/	2/	1/	2/	1/	2/	1/	2/
Germany versus France	-28.474	0.880	-33.891	0.920	-16.879	0.493	-13.369	0.425	-25.181	0.800	-29.841	0.845	-12.360	0.410	-9.785	0.323
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Germany versus Italy	-5.045	0.470	-7.744	0.515	-8.679	0.278	6.726	0.218	-1.660	0.245	-3.388	0.285	-7.238	0.238	4.109	0.153
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0977)	(0.0000)	0.0008	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0001)
Germany versus Spain	-22.462	0.780	-27.193	0.833	-9.830	0.305	-7.022	0.223	-13.671	0.545	-16.779	0.595	-2.757	0.118	-2.484	0.095
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0060)	(0.0068)	(0.0132)	(0.0472)
France versus Italy	96.260	0.998	105.167	1.000	6.777	0.283	21.842	0.583	57.205	0.958	63.380	0.983	4.883	0.193	15.753	0.435
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
France versus Spain	29.737	0.730	32.338	0.770	5.942	0.223	6.319	0.228	27.308	0.698	30.347	0.738	8.491	0.328	6.929	0.278
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Italy versus Spain	-67.062	0.985	-73.276	0.993	-0.919	0.078	-14.502	0.393	-26.484	0.683	-29.119	0.725	3.937	0.155	-6.800	0.213
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.3583)	(0.1619)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0001)	(0.0001)	(0.0000)	(0.0000)

	24-month								Long-run							
	1985		1991		1995		1998		1985		1991		1995		1998	
Test Statistics	1/	2/	1/	2/	1/	2/	1/	2/	1/	2/	1/	2/	1/	2/	1/	2/
Germany versus France	-19.482	0.648	-23.253	0.705	-8.895	0.330	-7.030	0.245	-17.032	0.575	-20.396	0.633	-7.164	0.315	-5.834	0.198
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Germany versus Italy	-1.798	0.198	-3.342	0.220	-5.937	0.248	0.872	0.098	-2.036	0.188	-3.476	0.210	-5.001	0.235	-0.108	0.090
	(0.0727)	(0.0000)	(0.0009)	(0.0000)	(0.0000)	(0.0000)	(0.3838)	(0.0388)	(0.0422)	(0.0000)	(0.0005)	(0.0000)	(0.0000)	(0.0000)	(0.9137)	0.0689
Germany versus Spain	-3.813	0.203	-5.504	0.245	-0.066	0.070	-0.389	0.060	0.025	0.120	-1.170	0.125	0.521	0.060	0.178	0.045
	(0.0002)	(0.0000)	(0.0000)	(0.0000)	(0.9471)	(0.2543)	(0.6972)	(0.4303)	(0.9803)	(0.0053)	(0.2423)	(0.0032)	(0.6023)	(0.4303)	(0.8589)	(0.7742)
France versus Italy	32.680	0.785	36.532	0.823	2.986	0.115	9.109	0.268	25.565	0.688	28.663	0.718	2.233	0.098	6.672	0.240
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0029)	(0.0086)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0259)	(0.0388)	(0.0000)	(0.0000)
France versus Spain	25.372	0.680	28.393	0.713	7.965	0.295	6.246	0.253	25.060	0.678	27.897	0.713	6.985	0.288	5.643	0.215
	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)	(0.0000)
Italy versus Spain	-3.451	0.155	-3.772	0.150	5.319	0.198	-1.257	0.150	2.992	0.123	3.223	0.138	5.048	0.205	0.300	0.123
	(0.0006)	(0.0001)	(0.0002)	(0.0002)	(0.0000)	(0.0000)	(0.2091)	(0.0002)	(0.0029)	(0.0042)	(0.0013)	(0.0009)	(0.0000)	(0.0000)	(0.7646)	(0.0042)

1 / Modified t-test for equal mean of two distributions (P-values in brackets).
2 / Kolmogorov-Smirnov test for equality of two distributions (P-values in brackets).