# Do political events affect business cycles? The Maastricht treaty, the creation of the ECB and the euro economy*

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September 2006
Preliminary and incomplete, please do not quote

**Abstract**

This paper provides some evidence on the effect of the Maastricht treaty and the creation of the ECB on the dynamics of the European business cycles. With a panel VAR approach and quarterly data for seven European countries over the sample 1980:1-2004:4, the paper shows that: (i) an areawide cycle emerges in the 1990s; (ii) volatility, persistence and correlation of national cycles change over time, but not necessarily at the dates of the two political events; (iii) in an out-of-sample forecasting sense, neither event made a huge difference of European business cycles.

JEL classification: C15, C33, E32, E42

Key words: Business cycles, ECB, Panel VAR, structural changes

*We would like to thank D. Giannone, S. Eickmeier and the participants to the 8th Bundesbank Spring Conference for useful comments and suggestions. This paper should not be reported as representing the views of the Banco de España, the European Central Bank, or the Eurosystem.

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# 1 Introduction

There is abundant evidence that economic activity in developed countries displays a number of common characteristics. For the Euro area, Del Negro and Otrok (2003), Canova et al. (2005) and Giannone and Reichlin (2005) among others, have shown that the cyclical features of output and industrial production are similar within countries and between the area and the US.

There is now also mounting evidence that the cyclical characteristics of real fluctuations are changing over time. These variations involve the statistical features of the cycles, the nature of the phenomena and the causes of fluctuations. For example, Helbling and Bayoumi (2003) find a substantial increase in synchronization of cycles in advanced economies after 2000; Stock and Watson (2003) document changes in the volatilities of G-7 business cycles in the 1990s, while Canova et. all. (2005) show the presence of important and continuous time variations in G-7 countries since the beginning of the 1980s.

Why are the cyclical features of industrialized economies changing over time? At least two explanations come to mind. One possibility is that the transmission of shocks within and across countries has changed as a consequences of variations in the structural characteristics of the various economies. These could include variations in the preferences and objective functions of agents; shifts in the way expectations about future events are formed; as well as changes in the institutional and operational features of domestic and international goods, labor and financial markets. The first two of these options are typically invoked, for example, to explain the "Great inflation of the 1970s" and the subsequent period of more stable and predictable macroeconomic environment (see e.g. Lubik and Schorfheide (2004) or Cogley and Sargent (2005)), while the latter is used, for example, to explain the medium-long term dynamics of wage inequalities (see e.g. Greenwood and Yorokoglu (1997)). A second possibility is that the nature, the characteristics and the frequency of the shocks hitting developed economies has dramatically changed. For example, several authors have argued that changes in the volatility of structural macroeconomic shocks could be responsible for the changes in the persistence of output and inflation in the US (see e.g. Sims and Zha (2004) and Canova and Gambetti (2005)), and that they could also affect the magnitude and the direction of the correlation among international macroeconomic variables (see Stock and Watson (2003)). Others have claimed that common shocks are now more frequent than what used to be (see e.g. Helbling and Bayoumi (2003)).

Despite the relevance of the issue little has been done, to the best of our knowledge, to study how changes in the institutional and operational features of markets affect business cycle characteristics of industrialized economies. This seems to be an important shortcoming since OECD countries have witnessed major changes in the way labor markets function; financial developments have been substantial; and important features of goods markets, such as inventory management, have dramatically changed. We think there are good reasons for why the literature has shied away from addressing such questions and why standard statistical techniques may be non-informative about the issues of interest. First, institutions and the nature of markets typically change slowly making it difficult to pin down exactly a potential break point date and select subsamples over which to compare business cycle features. Second, institutional changes may affect cycles which have much longer periodicity than the ones typically associated with business fluctuations. However, business cycle analysis may be incapable of informing us on the nature and the direction of the relationship. Third, externalities and threshold effects are likely to be important, for example, when considering financial or labor market changes. Therefore, even if a specific break date could be found, nothing may occur to cyclical fluctuations until a critical mass of agents is aware of and/or takes advantage of the changes. Finally, changes of institutions never come alone and this makes it particularly difficult to attribute observed variations to one single factor.

This paper attempts to shed some light on the effect of changes in institutions on the dynamics of business cycles by focusing on the recent European experience and the two dramatic changes that have occurred in the 1990s: the Maastricht treaty and the creation of the ECB. Investigating the effects of these two events on the nature of business cycles in the Euro area is relevant from at least four different perspectives. First, both events were brought about by national politicians and were to a large extent exogenous with respect to the dynamics of the European economies and to their relationship with the rest of the world. Therefore, reverse causality effects can be excluded a-priori and this makes the experience unique from the common currency area point of view. Does the European experience disproof this literature, which often stresses that real convergence should naturally precede the establishment of common monetary institutions? Second, both events are monetary in nature. The ability of monetary factors to affect real variables at business cycle frequencies has been extensively studied over the last 50 years and although different ideas coexists, the common wisdom is that monetary changes have limited effects on real fluctuations. This common wisdom has been formed on the basis of studies which examine the effects of shocks to policy rules and, to a much lesser extent, the variations in the policy rules employed. However, the nature of events we are considering is substantially different and the effects they induce should a-priori be comparable to other major changes e.g. the establishment of the Fed, the breakdown of the gold standard, etc., which are known to have made quite a difference for cyclical fluctuations around the world (see e.g. Bergman, Bordo and Jonung (1998)). Third, for macroeconomic analysis it is common to separate cyclical movements from other types of fluctuations. To support such an approach researchers typically claim that the mechanism generating the two type of fluctuations is considerably different. If institutional changes, besides affecting the medium-long run tendencies of the economy, also exercise a significant impact on the business cycle features of the data, such a practice should be reconsidered. Furthermore, sources of business cycles fluctuations might have more long run causes. Finally, policymakers monitoring domestic cycles are typically concerned with the effects of national idiosyncrasies. When institutional changes cause similar variations in economic activity in countries with different economic structures, one should probably discount the role of national idiosyncrasies when trying to understand international comovements in economic activity and national policies designed to counteract these tendencies may be ineffective.

Since the subject is particularly vast and largely unexplored, we limit our attention to three particular questions. First, we would like to know whether there has been any tendency for areawide and national cycles to change after the Maastricht treaty and the ECB started operating. The focus in on the direction, the magnitude and the intensity of the variations which are observed after each episode on average. Rather than taking a single time series indicator, we will construct measures of national and areawide cycles which exploit both cross country and time series information and explicitly allow the indicators to drift over time. Second, we would like to know whether a clean structural break took place when these events occurred or whether changes occur more slowly and continuously over time. Third, we would like to know whether the two events had different relative impact on the cyclical characteristics of the data - our prior being that the Maastricht treaty, forcing a process of convergence in government spending and debt, should have had more sizable effects than the creation of the ECB.

To study these questions we use a panel VAR model of the same type employed by Canova and Ciccarelli (2004) and Canova, et al. (2005). Their setup is useful in our context for at least two reasons. First, the econometric methodology they develop is designed to handle large scale dynamic models displaying unit specific dynamics and cross country lagged interdependencies and flexibly allows for time variations in the correlation structure of cyclical fluctuations across variables and countries. Second, the parsimonious parametrization they use endogenously produces an index structure, where the posterior distribution of areawide and national specific cyclical indicators, which dynamically span the space of cross country interdependencies, can recursively be constructed and easily analyzed. Therefore, the specification is particularly suited to study the interrelationship between the structural changes and institutional events and to analyze the links through which the latter affect the nature of real cyclical fluctuations.

In our exercise we use data from seven European countries (Germany, France, Italy, Spain, Belgium, the Netherlands and Finland) for five different variables (output, employment, industrial production, consumption and investment) and construct distributional measures of area wide and national cycles using the sample 1980:1-2004:4. We perform unconditional and conditional forecasting exercises, with information available before and after the events and we exploit a Monte Carlo procedure to trace out the effect of shocks to interesting variables-country combination in the system.

Our results are as follows. First, we find that the features of the posterior distribution of our cyclical indicators have changed over time. In particular, we show that while an areawide cycle was largely absent in the 1980s, it became much more marked and significant over the last 15 years. At the same time, we show that national cycles have maintained most of their characteristics and that the scope for time variations is limited. While the timing of these changes do not necessarily line up with the two events of interest, we also find that the posterior median of our cyclical indicators shows decreasing volatility over the last 10 years, that areawide and national indicators become much more synchronic over the same period and, especially, over the last five years of the sample, and that persistence of the national indicators has somewhat increased after that date. Second, in an unconditional forecasting sense, the Maastricht treaty and the creation of the ECB did not make a huge difference. Using the information available before and after these two events, we are able to unconditionally predict both the direction and the magnitude of the changes in the five variables for the majority of the countries with unchanged precision. Third, in terms of conditional forecasting, we find that the transmission of shocks was somewhat altered. Once again conditioning on the information available before and after the two events, we demonstrate that the magnitude and the timing of the responses to a German shock is at times altered after both the Maastricht treaty and the creation of the ECB. Shocks originating outside the Euro area appear to produce a much stronger and quicker response of Euro area variables after the ECB was created.

All in all, while we find some evidence of structural changes in the features of business cycles and in the transmission of shocks over time, we do not find overwhelming support that these changes are related to the two political events we are interested in.

The rest of the paper is organized as follows: the next section presents the model specification, the technique used to construct various indicators, the procedure used to compute unconditional and conditional forecasts and the details of our empirical approach. Section 3 presents the results and section 4 concludes.

## 2 The empirical model

The model we employ has the form:

$$y_{it} = D_{it}(L)Y_{t-1} + F_{it}W_{it} + e_{it} \tag{1}$$

where $i = 1, ..., N$ refers to countries and $t = 1, ..., T$ to time. $y_{it}$ is a $G \times 1$ vector for each country $i$ and $Y_t = (y'_{1t}, y'_{2t}, \dots y'_{Nt})'$. $D_{it,j}$ are $G \times NG$ matrices for each lag $j = 1, \dots, p$, $W_{it}$ is a $Mq \times 1$ vector of exogenous variables and $F_{it,j}$ are $G \times M$ matrices each lag $j = 1, \dots, q$ and $e_{it}$ is a $G \times 1$ vector of random disturbances. Under the regularity conditions underlying the Wold theorem, a finite order VAR is a good characterization to any data one wants to analyze. Therefore, one can interchangeably think of the data or of its vector autoregressive representation without any loss of generality.

Equation (1) displays three important ingredient that Canova et. al. (2005) have found to be crucial in characterizing the similarities, the propagation and the structure of cyclical fluctuations in G-7 countries. First, the coefficients are allowed to vary over time. Second, the dynamic relationships are allowed to be unit specific. Third, whenever the $NG \times NG$ matrix $D_t(L) = [D_{1t}(L), \dots, D_{Nt}(L)]'$, is not block diagonal for some $L$, cross-unit lagged interdependencies are present. While these features add considerable realism to the specification, and a avoid the "incredible" short-cuts that the literature has often taken, they are not costless: the number of parameters is large (there are $k = NGp + Mq + 1$ parameters in each equation) and there is only one time period per unit to estimate them.

It is convenient to rewrite (1) in a simultaneous equations format:

$$Y_t = Z_t \delta_t + E_t \qquad E_t \sim N(0, \Omega) \tag{2}$$

where $Z_t = I_{NG} \otimes X'_t$; $X'_t = (Y'_{t-1}, Y'_{t-2}, \dots, Y'_{t-p}, W'_t, W'_{t-1}, \dots, W'_{t-q})$, $\delta_t = (\delta'_{1t}, \dots, \delta'_{Nt})'$and $\delta_{it}$ are $Gk \times 1$ vectors containing, stacked, the $G$ rows of the matrix $D_{it}$ and $F_{it}$, while $Y_t$ and $E_t$ are $NG \times 1$ vectors of endogenous variables and of random disturbances.

Since $\delta_t$ varies with cross–sectional units in different time periods, it is impossible to estimate it using unrestricted classical methods. However, even if $\delta_t$ were time invariant, its sheer dimensionality prevents unconstrained estimation. Our approach is to assume that $\delta_t$ has a flexible factor structure of the form:

$$\delta_t = \Xi_1\lambda_t + \Xi_2\alpha_t + \Xi_3\rho_t + \Xi_4\psi_t + u_t \tag{3}$$

where $\Xi_1,\ \Xi_2,\ \Xi_3, \Xi_4$ are matrices of dimensions $NGk \times 1$, $NGk \times N$, $NGk \times G$, $NGk \times 1$ respectively and $\lambda_t, \alpha_t, \rho_t, \psi_t$ are mutually orthogonal. Here $\lambda_t$ is a scalar process that captures movements in the coefficient vector which are common across units and variables; $\alpha_t$ captures movements in the coefficient vector which are common within countries and therefore its dimension equals to $N$, the number of countries in the panel; $\rho_t$ captures movements in the coefficient vector which are variable specific and its dimension is therefore equal to $G$, the number of variables in each country, while $\psi_t$ captures movements in the coefficients due to exogenous variables. Finally, $u_t$ captures all the unmodelled features of the coefficient vector, which may have to do with lag specific, time specific or other idiosyncratic effects.

Factoring $\delta_t$ as in (3) is advantageous in many respects. Computationally, it reduces the problem of estimating $NGk$ coefficients into the one of estimating $1 + N + G + 1$ factors characterizing their dynamic features. Therefore, even when the number of interdependent units is large estimation is feasible, noise is averaged out and reliable estimates of the features of interest can be obtained. Practically, the factorization (3) transforms an overparametrized panel VAR into a parsimonious SUR model where the regressors are averages of certain right-hand side variables of the VAR. In fact, substituting (3) into (2) we have

$$Y_t = \mathcal{Z}_{1t}\lambda_t + \mathcal{Z}_{2t}\alpha_t + \mathcal{Z}_{3t}\rho_t + \mathcal{Z}_{4t}\psi_t + v_t \tag{4}$$

where $\mathcal{Z}_{1t} = Z_t\Xi_1,\ \mathcal{Z}_{2t} = Z_t\Xi_2,\ \mathcal{Z}_{3t} = Z_t\Xi_3, \mathcal{Z}_{4t} = Z_t\Xi_4$ capture respectively, common, country specific, variable specific and exogenous specific information present in the data, and $v_t = E_t + Z_t u_t$. Economically, the decomposition in (4) is convenient since it allows us to measure the relative importance of areawide and country specific influences for fluctuations in $Y_t$ and therefore to examine whether the events we are interested in characterizing affect differently the two. In fact, $WLI_t = \mathcal{Z}_{1t}\lambda_t$ plays the role of an areawide indicator, while $CLI_t = \mathcal{Z}_{2t}\alpha_t$ plays the role of a vector of country specific indicators. Both coincident and leading versions of these indicators can be designed using either time $t$ or time $t-j$, $j = 1, 2, ...$ information, since $\mathcal{Z}_{it}$, $i = 1, \ldots, 4$ only contain information present in the predetermined and exogenous variables of the VAR (see Canova and Ciccarelli (2004) for this), and recursively constructed, given a law of motion of $\lambda_t$ and $\alpha_t$. Note that $WLI_t$ and $CLI_t$ are correlated by construction - the same variables enter in both $\mathcal{Z}_{1t}$ and $\mathcal{Z}_{2t}$ - but become uncorrelated as the number of countries becomes large.

To illustrate the structure of the matrices $\Xi$'s and the nature of $\mathcal{Z}_{1t}, \mathcal{Z}_{2t}, \mathcal{Z}_{3t}$, suppose there are $G = 2$ variables, $n = 2$ countries, $p = 1$ lags, no exogenous variables and an intercept of the form

$$\begin{bmatrix} y_t^1 \\ x_t^1 \\ y_t^2 \\ x_t^2 \end{bmatrix} = \begin{bmatrix} d_{1,1,t}^{1,y} & d_{2,1,t}^{1,y} & d_{1,2,t}^{1,y} & d_{2,2,t}^{1,y} \\ d_{1,1,t}^{1,x} & d_{2,1,t}^{1,x} & d_{1,2,t}^{1,x} & d_{2,2,t}^{1,x} \\ d_{1,1,t}^{2,y} & d_{2,1,t}^{2,y} & d_{1,2,t}^{2,y} & d_{2,2,t}^{2,y} \\ d_{1,1,t}^{2,x} & d_{2,1,t}^{2,x} & d_{1,2,t}^{2,x} & d_{2,2,t}^{2,x} \end{bmatrix} \begin{bmatrix} y_{t-1}^1 \\ x_{t-1}^1 \\ y_{t-1}^2 \\ x_{t-1}^2 \end{bmatrix} + \begin{bmatrix} c_1^y \\ c_1^x \\ c_2^y \\ c_2^x \end{bmatrix} + e_t \tag{5}$$

Here $\delta_t = [d_{1,1,t}^{1,y}, d_{2,1,t}^{1,y}, d_{1,2,t}^{1,y}, d_{2,2,t}^{1,y}, c_1^y, d_{1,1,t}^{1,x}, d_{2,1,t}^{1,x}, d_{1,2,t}^{1,x}, d_{2,2,t}^{1,x}, c_t^x, d_{1,1,t}^{2,y}, d_{2,1,t}^{2,y}, d_{1,2,t}^{2,y}, d_{2,2,t}^{2,y}, c_2^y, d_{1,1,t}^{2,x}, d_{2,1,t}^{2,x}, d_{1,2,t}^{2,x}, d_{2,2,t}^{2,x}, c_2^x]'$ is a $(20 \times 1)$ vector containing the time varying coefficients of the model and the typical element of $\delta_t$, $\delta_{l,s,t}^{i,j}$, is indexed by the country $i$, the variable $j$, the variable in an equation $l$ (independent of the country), and the country in an equation $s$ (independent of variable). If we are not interested in modelling all these aspects and call $u_t$ all unaccounted features, one possible factorization of $\delta_t$ is

$$\delta_t = \Xi_1\lambda_t + \Xi_2\alpha_t + \Xi_3\rho_t + u_t \tag{6}$$

where for each $t$, $\lambda_t$ is a scalar, $\alpha_t$ is a $2 \times 1$ vector, $\rho_t$ is a $2 \times 1$ vector, $\Xi_1$ is a $20 \times 1$ vector of ones ,and

$$\underset{(20\times2)}{\Xi_2} = \begin{bmatrix} \iota_1 & 0 \\ \iota_1 & 0 \\ 0 & \iota_2 \\ 0 & \iota_2 \end{bmatrix} \qquad \underset{(20\times2)}{\Xi_3} = \begin{bmatrix} \varkappa_1 & 0 \\ 0 & \varkappa_2 \\ \varkappa_1 & 0 \\ 0 & \varkappa_2 \end{bmatrix}$$

with $\iota_1 = \begin{pmatrix} 1 & 1 & 0 & 0 & 0 \end{pmatrix}'$, $\iota_2 = \begin{pmatrix} 0 & 0 & 1 & 1 & 0 \end{pmatrix}'$, $\varkappa_1 = \begin{pmatrix} 1 & 0 & 1 & 0 & 0 \end{pmatrix}'$ and

$\varkappa_2 = \begin{pmatrix} 0 & 1 & 0 & 1 & 0 \end{pmatrix}'$. Hence, the VAR (5) can be rewritten as

$$\begin{bmatrix} y_t^1 \\ x_t^1 \\ y_t^2 \\ x_t^2 \end{bmatrix} = \begin{bmatrix} \mathcal{Z}_{1t} \\ \mathcal{Z}_{1t} \\ \mathcal{Z}_{1t} \\ \mathcal{Z}_{1t} \end{bmatrix} \lambda_t + \begin{bmatrix} \mathcal{Z}_{2,1,t} & 0 \\ \mathcal{Z}_{2,1,t} & 0 \\ 0 & \mathcal{Z}_{2,2,t} \\ 0 & \mathcal{Z}_{2,2,t} \end{bmatrix} \alpha_t + \begin{bmatrix} \mathcal{Z}_{3,1,t} & 0 \\ 0 & \mathcal{Z}_{3,2,t} \\ \mathcal{Z}_{3,1,t} & 0 \\ 0 & \mathcal{Z}_{3,2,t} \end{bmatrix} \rho_t + v_t \tag{7}$$

where $\mathcal{Z}_{1t} = y_{t-1}^1 + x_{t-1}^1 + y_{t-1}^2 + x_{t-1}^2 + 1$, $\mathcal{Z}_{2,1,t} = y_{t-1}^1 + x_{t-1}^1$, $\mathcal{Z}_{2,2,t} = y_{t-1}^2 + x_{t-1}^2$, $\mathcal{Z}_{3,1,t} = y_{t-1}^1 + y_{t-1}^2$, $\mathcal{Z}_{3,2,t} = x_{t-1}^1 + x_{t-1}^2$ and $v_t = e_t + Z_t' u_t$. In the empirical application, all variables are measured in growth rates and therefore unweighted averaging will indeed be appropriate. When $\lambda_t$ is large relative to $\alpha_t$, $y_t^1$ and $x_t^1$ comove with $y_t^2$ and $x_t^2$. On the other hand, when $\lambda_t$ is zero, $y_t^1$ and $x_t^1$ may drift apart from $y_t^2$ and $x_t^2$. Note that, when $p > 1$, lags can be weighted using a decay factor as in Doan et al. (1984).

As the notation we have used makes it clear, the regressors in (4) are combinations of lags of the right hand side variables of the VAR, while $\lambda_t, \alpha_t, \rho_t, \psi_t$ play the role of time varying loadings. Using averages as regressors is common in the factor model literature (see e.g. Stock and Watson (1989) or Forni and Reichlin (1998)) and in the signal extraction literature (see e.g. Sargent (1989)). However, there are five important differences between (4) and standard factor models. First, the indices we construct equally weight the information in all variables while in factor models the weights generally depend on the variability of the components. The equal weighting scheme directly comes from (3) and the fact that all variables are measured in the same units. Since in the empirical applications we will be dealing with countries of different size, growth rates of the variables for each country are weighted by the relative size of the various economies (see also Pesaran et al., 2004). Second, our indices dynamically span lagged interdependencies across units and variables while in standard factor models they statically span the space of the variables of the system. Third, our indices are observable while in factor models they are estimated. Fourth, our loadings are allowed to be time varying while the loadings are typically treated as constant in factor models. Finally, our averaging approach creates moving average terms of order $p$ in the regressors of (4), even when $y_{it}$ are serially independent. Therefore, our indicators eliminate high frequency variability from the right hand side variables of the VAR. This feature makes them particularly useful in forecasting in the medium run and in detecting turning points of the actual data (see Canova et. al. (2005)). The emphasis that our transformed SUR model gives to the low frequencies movements present in the variables of the VAR will become evident in the next section.

To complete the specification we need to describe the evolution of $\lambda_t$, $\alpha_t$, $\rho_t, \psi_t$ over time and the features of their (prior) distribution. Write (3) compactly as:

$$\delta_t = \Xi\theta_t + u_t \qquad\qquad u_t \sim N(0, \Sigma \otimes V) \tag{8}$$

where $\Xi = [\Xi_1, \Xi_2, \Xi_3, \Xi_4]$, $\theta_t = [\lambda_t,\ \alpha_t', \rho_t', \psi_t]'$, and $V$ is a $k \times k$ matrix and let

$$\theta_t = \theta_{t-1} + \eta_t \qquad\qquad \eta_t \sim N(0, B_t) \tag{9}$$

Assume that $\Sigma = \Omega$ and $V = \sigma^2 I_k$, $\sigma^2$ unknown; that $B_t = \gamma_1 * B_{t-1} + \gamma_2 * \bar{B}$, $\gamma_1, \gamma_2$ known; that $\bar{B} = diag(\bar{B}_1, \bar{B}_2, \bar{B}_3)$, and that $E_t$, $u_t$ and $\eta_t$ are mutually independent.

In (9) the factors evolve over time as random walks. Alternative specifications which allow for more complex dynamics or exchangeability across units are possible (see e.g. Canova and Ciccarelli (2004)). We stick to this simple setup since experimentation with more complicated structures did not produce important improvements in our results. The spherical assumption on $V$ reflects the fact that the factors have similar units, while setting $\Sigma = \Omega$ is standard (see e.g. Kadiyala and Karlsson (1997)). The variance of the innovations in $\theta_t$ is allowed to be time varying to account for ARCH-M type effects and other generic volatility clustering that may appear in the coefficients of several, or all, series within and across units. Time invariant structures ($\gamma_1 = \gamma_2 = 0$), and homoskedastic variances ($\gamma_1 = 0$ and $\gamma_2 = 1$) are special cases of the assumed process. The block diagonality of $\bar{B}$ guarantees orthogonality of the factors, which is preserved a-posteriori, and hence their identifiability. Finally, independence among the errors is standard.

To summarize, our reparametrized empirical model has the state space structure:

$$\begin{aligned} Y_t &= (Z_t \Xi)\theta_t + v_t \\ \theta_t &= \theta_{t-1} + \eta_t \end{aligned} \tag{10}$$

where $v_t \sim (0, \sigma_t \Omega)$ and $\sigma_t = (1 + \sigma^2 X_t' X_t)$. To compute posterior distributions for the unknowns we need prior densities for $\phi_0 = (\Omega,\ \sigma^2,\ \bar{B}, \theta_0)$. Because we want to minimize the impact of our prior choices on the posterior distribution of the indicators, we specify loose but proper priors. Their exact form, the numerical approach used to compute posterior distributions and the details of the computations are in appendix A. While the model (10) can be estimated both with classical and Bayesian methods, we prefer the second approach since the exact small sample distribution of the objects of interest can be obtained even when $T$ and $N$ are relatively small - as it is the case here.

Besides characterizing the time profile of the posterior distribution of areawide and national indicators, we will be interested in computing predictive distributions for future $Y_{t+\tau}$, both unconditionally and conditionally. Both types of distributions can be obtained numerically using the structure of the model (10) and draws for the posterior of the parameters and/or the shocks. In particular, $f(Y_{t+\tau}) = \int f(Y_{t+\tau}|Y_t, \phi_{t+\tau})g(\phi_{t+\tau}|Y_t)d\phi_{t+\tau}$, is the unconditional predictive distribution, where $t$ takes different values and $\tau$ runs from 1 to 20 (quarters). To draw from the predictive density we don't condition on $\theta_{t+\tau} = \theta_t$, but rather use the law of motion of $\theta_t$ to get random draws for $\theta_{t+\tau}$ and average them out. In the second case we produce impulse responses obtained as the difference between two conditional forecasts: one where a particular variable (or set of variables) is shocked and one where the disturbance is set to zero. Formally, let $y^t$ be a history for $y_t$; $\theta^t$ be a trajectory for the coefficients up to $t$, $y_{t+1}^{t+\tau} = [y'_{t+1}, ...y'_{t+\tau}]'$ a collection of future observations and $\theta_{t+1}^{t+\tau} = [\theta'_{t+1}, ...\theta'_{t+\tau}]'$ a collection of future trajectories for $\theta_t$. Let $\mathcal{W}_t = (\Omega, \sigma^2, B_t)$; set $\xi'_t = [v'_{1t}, v'_{2t}, \eta'_t]$, where $v_{1t}$ are the shocks to the endogenous variables and $v_{2t}$ the shocks to exogenous variables. Let $\xi^{\delta}_{j,t+1}$ be a realization of $\xi_{j,t+1}$ of size $\delta$ and let $\mathcal{F}^1_t = \{y^t, \theta^t, \mathcal{W}_t, J_t, \xi^{\delta}_{j,t}, \xi_{-j,t}, \xi^{t+\tau}_{t+1}\}$ and $\mathcal{F}^2_t = \{y^t, \theta^t, \mathcal{W}_t, J_t, \xi_t, \xi^{t+\tau}_{t+1}\}$ be two conditioning sets, where $\xi_{-j,t}$ indicates all shocks, excluding the one in the j-th component and $J_t$ is an identification matrix satisfying $J_t J'_t = \sigma_t \Omega$. Then, responses at horizon $\tau$ to an impulse in $\xi^{\delta}_{j,t}$, $j = 1, \ldots$ are

$$IR^j_y(t, \tau) = E(y_{t+\tau}|\mathcal{F}^1_t) - E(y_{t+\tau}|\mathcal{F}^2_t) \qquad \tau = 1, 2, \ldots \tag{11}$$

In this paper, the shocks we consider domestic German disturbances (shocks which simultaneously affect all German variables) and US federal funds rate shocks. The German shock is identified with a block-Choleski decomposition of $\Omega$, placing German variables first with respect to the rest of the countries. In both cases we assume that the impulse lasts one period. For alternative responses to more persistent shocks, see Canova and Ciccarelli (2004). Note that when the coefficients are constant and shocks affect endogenous variables only, (11) collapses to the traditional impulse response function to unitary structural shocks.

# 3 The data

In the VAR we use quarterly growth rates of five variables (output, industrial production, employment, consumption and investment) for seven European countries (Germany, France, Italy, Spain, Belgium, Netherlands and Finland) for the sample 1980:1 to 2004:4. Among all possible variables and countries, we choose this combination since it maximizes the length of the data and since the same definitions can be maintained over the sample. Output is measured by Eurostat real GDP at 1995 prices, industrial production by its OECD index, employment by the OECD index of total employment, consumption by Eurostat total private consumption and investment by Eurostat gross fixed capital formation. All OECD series are in 1995 prices.

We use as exogenous variables the growth rates of the oil price, of a measure of world trade, of US GDP and of US stock prices and the US federal funds rate. Oil prices are obtained from the IMF Financial statistic series, world trade is measured by world trade in goods and services and it is from the OECD Main Economic Indicators. Finally the stock price series measures the NYSE index.

Since some variables display seasonality despite being reported as seasonally adjusted at the source, we have preliminary passed the questionable series though TRAMO-SEATS and used the output of the seasonal and working days adjustment procedure in our system. Data is demeaned and weighted by the share of GDP in the country. We use four lags of each of the endogenous variables and two lags of the exogenous ones. This implies that each of the 35 equations of the system has $35 \times 4 + 5 \times 2 = 150$ coefficients and that the panel VAR has 5250 coefficients.

## 4 Assessing the fit of the model

Before analyzing the relationship between cyclical fluctuations in real variables and the two events which are the focus of our investigation, we examine some of the properties of the empirical model we employ. Documenting the fit of the model is important because the outcomes of our exercises will acquire more credibility if the model captures the data well, if important features are not left unexplained and if our indicators can reproduce important cyclical statistics of the data.

To start with, it is worth mentioning that the exact model we use was selected with a specification search were different nested and non-nested models were compared via marginal likelihood. The marginal likelihood of model $i$ **is** $f(Y|M_i) = \int L(y|\phi_i, M_i)g(\phi_i|M_i)d\phi_i$, where $\phi_i = [\phi_{1i}, \dots \phi_{ti}]$is the vector of the parameters of the model $i$, and it is computed as in Chib (1995), letting both $\theta_t$ and $\sigma_t$ be vectors of latent variables. Model $i$ is preferred to model $i'$if the Bayes factor $BF(i, i') = \frac{f(Y|M_i)}{f(Y|M_{i'})}$substantially exceeds 1. We have considered the following specifications as alternative to the one we finally selected: a) a model with no cross country interdependencies; b) a model with no time variations, c) a model where the coefficient factorization is exact; d) a model with a more informative prior; e) a model with no unit-specific dynamics; f) a model where the variable factor is excluded from the specification; g) a model there is no factor for the exogenous variables (their coefficients are treated the same way as the coefficients on lagged endogenous variables); h) a model with no time variations but breaks at 1993:3 and 1998:4. In all cases the Bayes factor exceeded 100 and in some cases it is even larger, overwhelmingly favouring our selected specification for representing the data.

We graphically present the results of our estimation process in figure 1, where we plot the time path of the median of posterior and the 68 percent highest posterior band for the areawide indicator and the seven country specific indicators. Table 1 reports some basic statistics for the median of the various indicators and for synthetic Euro area GDP growth data. Shaded area in figure 1 represent recessions dates: the ones for the Euro area are from the CEPR; those for Germany, Italy, France, Spain and Finland are from ECRI (www.businesscyle.com) while no official dates exist for the Netherlands and Belgium and the dates are those reported by Harding and Pagan (2002). The figure and the table contain a number of interesting issues. First, the areawide indicator captures cycles which are smoother than national ones: the standard deviation of the median of the former is smaller and its autocorrelation is stronger than the standard deviations and the autocorrelations of the national indices. This is not entirely surprising, given that the Euro area indicator averages information contained in the variables of the seven countries. Second, the Euro area indicator is characterized by two important phases, roughly, pre 1990 and post 1990. In the first phase, fluctuations are of limited magnitude and often, the 68 percent posterior band includes the zero value. In the second phase, fluctuations are stronger, both in terms of magnitude and persistence, so that the whole 68 percent band is above or below zero at many dates. Hence, while in the first part of the sample areawide fluctuations are rather minor, an important areawide factor appears to emerge in the second part of the sample. Third, the areawide indicator has three clear expansion phases (1987-88, 1995, 1998-2000) and one strong recession (1992-93).

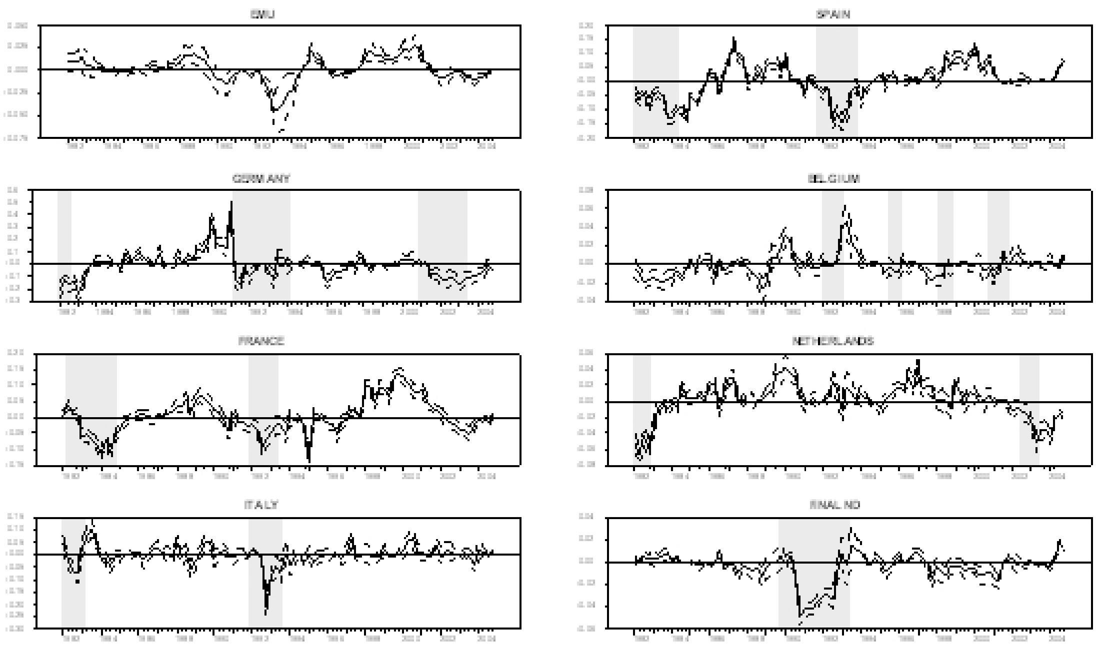

Figure 1: Cyclical Indicators, Posterior Medians and 68 percent bands

After 2000, the posterior band for the areawide indicator is entirely on the negative side of zero but, quantitatively speaking, the magnitude of the fall is small. Fourth, national indicators display dips in correspondence with the official recession dates and display cycles with well established features. For example, the German indicator shows a significant positive trend from the beginning of the 1980 to the end of the decade, with a marked expansion in the last two years of the decade; two recessions in the 1991-93 and 2001-2003 and a substantial stagnation in the 1990s. The French indicator displays three important recessions (1983-1984, 1992-93 and 2003) and two significant growth episodes, one in the late 1980s and one in the late 1990s. The Italian indicator is always around the zero line, except in two major recessions, while the Spanish one features important growth episodes at the time of EU membership, in the late 1990s and, again, since 2003. The national indicators for the three smaller countries also displays dips in correspondence with recessions dates. Interestingly, the Dutch indicator is highly correlated with the German indicator in the 1980s but not afterwards, and the Belgian one shows substantial stagnation in the 1990s. Finally, the emergence of an areawide cycle in the 1990s does not imply that national cycles are disappearing, in line with Canova et. al (2005). Instead, the presence of an areawide cycle is the result of a much higher synchronicity of timing and intensity in the national cycles of the major European countries.

Table 1: Statistics of the median of cyclical indicators

| | Euro | DE | FR | IT | ES | BE | NET | FIN | EuroGDP |
|---|---|---|---|---|---|---|---|---|---|
| | | | | | Full sample | | | | |
| Volatility | 1.42 | 9.31 | 5.32 | 3.83 | 6.06 | 1.24 | 2.06 | 1.14 | 1.30 |
| AR(1) | 0.91 | 0.57 | 0.81 | 0.39 | 0.82 | 0.70 | 0.77 | 0.68 | 0.85 |
| Correlation | | 0.25 | 0.44 | 0.33 | 0.54 | -0.61 | 0.09 | 0.31 | 0.77 |
| Max Correlation (lag) | | 0.25 (0) | 0.49 (-2) | 0.41 (-1) | 0.62 (-2) | -0.61 (0) | 0.24 (4) | 0.41 (-2) | 0.80 (-1) |
| | | | | | Pre Maastricht | | | | |
| Volatility | 1.67 | 11.55 | 4.36 | 4.81 | 6.80 | 1.61 | 2.26 | 1.44 | 1.47 |
| AR(1) | 0.94 | 0.51 | 0.82 | 0.42 | 0.80 | 0.73 | 0.76 | 0.67 | 0.88 |
| Correlation | | 0.19 | 0.46 | 0.32 | 0.52 | -0.66 | 0.07 | -0.28 | 0.82 |
| Max Correlation (lag) | | 0.25 (-1) | 0.49 (-2) | 0.41 (-1) | 0.62 (-2) | -0.61 (0) | 0.23 (4) | 0.41 (-2) | 0.83 (+1) |
| | | | | | Inter Maastricht-ECB | | | | |
| Volatility | 1.28 | 4.02 | 5.35 | 2.36 | 3.16 | 0.71 | 1.34 | 0.59 | 0.74 |
| AR(1) | 0.79 | 0.51 | 0.50 | 0.07 | 0.50 | 0.54 | 0.65 | 0.62 | 0.75 |
| Correlation | | 0.45 | 0.19 | 0.03 | 0.67 | -0.47 | -0.43 | -0.77 | 0.59 |
| Max Correlation (lag) | | 0.53 (+1) | 0.31 (-1) | 0.40 (+1) | 0.67 (0) | 0.10 (-4) | 0.31 (+4) | 0.37 (-4) | 0.90(+1) |
| | | | | | Post-ECB | | | | |
| Volatility | 0.88 | 6.3 | 5.90 | 2.07 | 4.09 | 0.47 | 1.73 | 0.65 | 1.34 |
| AR(1) | 0.93 | 0.88 | 0.93 | 0.26 | 0.81 | 0.79 | 0.76 | 0.66 | 0.84 |
| Correlation | | 0.82 | 0.89 | 0.48 | 0.77 | -0.33 | 0.72 | -0.51 | 0.79 |
| Max Correlation (lag) | | 0.82 (0) | 0.93 (-1) | 0.48 (0) | 0.85 (-1) | 0.04 (-4) | 0.72 (0) | -0.16 (-4) | 0.81 (-1) |

The median of our areawide indicator shares important similarities with synthetic Euro area GDP growth in terms of volatility and serial correlation. However, since our indicator contains information from series other than GDP, the correlation is far from perfect. In particular, the two series display similar cyclical fluctuations in the 1986-1998 period, but they are considerably different at the beginning and the end of the sample. Interestingly, while GDP growth is strongly negative since 2001, our areawide indicator presents a much less pessimistic view about the Euro area economy, at least until 2003 (see figure 2).

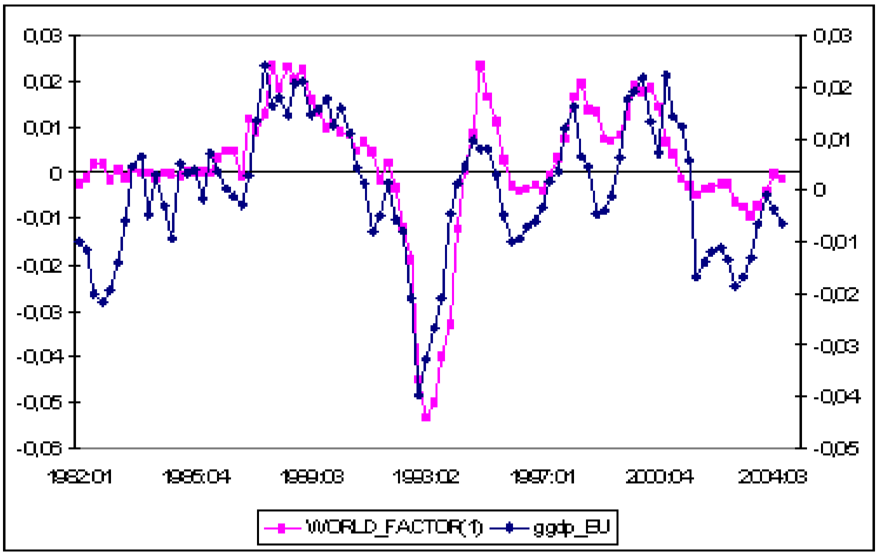


Figure 2

The reasonable picture that our indicators offer of the dynamics of national and of the area wide economies is also confirmed when looking at business cycle phases. Table 2 reports the average length in quarters of various business cycles phases for our indicators and for the synthetic GDP growth measure for the Euro area. Also in this case we use the posterior medians of our indicators in the comparison; employ the Bry-Boschen algorithm to date turning points and summarize the resulting information about business cycle phases reporting peak to peak (P-P) and trough to trough (T-T) average durations [1].

GDP growth tends to have longer peak to peak or trough to trough cycles than our EU indicator but this is expected since our indicator captures, among other things, fluctuations in the industrial production which are much shorter than those in GDP. The same pattern clearly emerges when we compare the our national indicators with national GDP growth. Despite these difference, the amplitude of the cycles is roughly the same and the concordance in the turning point dates is high. Interestingly, while trough to peak phases

[1] The rules used to date turning points are the following: the peaks and the throughs are the maximum and minimum value over 5 quarters, peaks and throughs must alternate, the minimum length of a P-P and a T-T phase is 5 quarters, the minumlength of a P-T and T-P pahse is 2 quarters.

are somewhat longer than peak to trough ones, the difference is considerably smaller than the difference typically found, for example, in the US.

In sum, the evidence we presented suggests that our procedure produces indicators which capture reasonably well the features of the data, track pretty much the ups and downs of synthetic Euro area and national GDP growth, and do not leave major recession or expansions episodes unexplained. This makes us confident that our setup can be used to answer the questions which are the focus of this investigation.

Table 2: Turning point statistics

| | Full sample | | Pre-Maastricht | | Inter Maastricht-ECB | | Post ECB | |
|---|---|---|---|---|---|---|---|---|
| | P-P | T-T | P-P | T-T | P-P | T-T | P-P | T-T |
| Euro | 10.5 | 10.8 | 10.0 | 10.4 | 12.0 | 12.0 | 5.0 | NA |
| Germany | 10.8 | 10.8 | 11.5 | 12.0 | 11.0 | 12.0 | 8.0 | 8.0 |
| France | 13.6 | 14.4 | 16.6 | 16.6 | 9.0 | 12.0 | 14.0 | NA |
| Italy | 11.7 | 12.3 | 12.5 | 12.6 | 18.0 | 15.0 | 7.0 | 6.0 |
| Spain | 10.5 | 10.5 | 11.5 | 13.3 | 11.0 | 9.0 | 6.0 | 8.0 |
| Belgium | 15.5 | 13.5 | 21.0 | 18.0 | 10.0 | 10.0 | NA | 16.0 |
| Netherlands | 20.5 | 19.0 | 20.6 | 18.0 | 20.0 | 28.0 | NA | 12.0 |
| Finland | 12.6 | 12.4 | 12.5 | 12.0 | 16.0 | 20.0 | 11.0 | 6.0 |
| EuroGDP | 20.0 | 17.0 | 24.0 | 26.0 | 24 | NA | 8.0 | 8.0 |

# 5 Political events and real fluctuations

## 5.1 Reduced form evidence

To examine whether the Maastricht treaty and the creation of the ECB had anything to do with the (reduced form) properties of Euro area real cycles, we first compare statistics of the posterior median estimates of the indicators in three subsamples (pre-Maastricht, inter Maastricht-ECB, post ECB), see table 1. While this evidence will not necessarily be conclusive, as the changes we are interested in characterizing do not necessarily show up in statistics computed over relatively short samples, it may help us to establish a number of stylized facts and compare them with those obtained in the full sample.

There are at least four interesting aspects which we would like to discuss. First, the volatility of the median of all indicators falls, when going from the early to the later subsamples and the decline in the post ECB sample is the largest of all. Hence, our indicators become progressively more stable, and, in a way, the creation of the ECB appear to have made a larger difference than the Maastricht treaty. Second, the contemporaneous correlation of national indicators and areawide indicator also dramatically increases in the latest subsample. The correlation increases also in the intermediate sample, but the magnitude of the change is smaller. This general increase in synchronicity, which we are already discussed in the previous section, is also evident from the smaller standard deviation bands around the median of the areawide indicator in the latter two subsamples. Since the posterior bands for the national indicators are roughly of the same magnitude throughout the sample, it must be the case that cyclical comovements across countries have increased, while those within countries have remained roughly unchanged. Third, the first order autocorrelation coefficient of the national indicators also generally increases after the creation of the ECB, except for the Italian and the Finnish indicators, while no change in the persistence of the national indices is recorded after the Maastricht treaty. Therefore, in the latter part of the sample, not only volatility declined but the persistence of national indices also increased. Finally, while national indicators of the four major countries leaded the areawide indicator in the first two subsamples, they all became roughly coincident with the Euro area indicator after that date. The opposite appears to be true for the indicators of the smaller countries, which go from lagging to leading after the creation of the ECB. One can also notice some changes in the lead-lag relationships after the Maastricht treaty, but no clear common pattern emerge. For example, the German indicator lagged the Euro area indicator in the pre 1993 sample and leaded afterwards, while the Italian and the French indicators went from leading one or two quarters in the pre-Maastricht sample to lagging one quarter after that date. In any case, one should be careful in evaluating changes of lead-lag relationships, as the last two subsamples are short and estimates of the coherence of various series are rather imprecise.

Artis and Zhang (1997) have analyzed business cycles statistics computed from standard filtering methods before and after 1979 - the period of the first European Monetary system (ERM). They find an increase in the degree of conformity and the degree of synchronicity in the fluctuations of the countries participating to the first monetary system, an increase which is not visible in non-ERM countries. Table 1 shows that the Maastricht treaty and the creation of the ECB had roughly similar cyclical consequences. Interesting, both political events appear to be associated with changes in the magnitude, the volatility and, to a much less extent, the persistence of both areawide and national cycles.

Table 2, which reports information for business cycle phases in the three subsamples of interest, also suggests the presence of some changes in the average length of each of the phases. In fact, one can notice a small decline in the length of both business cycle phases as we move from the first to the last subsample, but the decline is far from generalized. Such a pattern partly contrasts with the increased serial correlation of national cycles we observed in table 1. In any case, both changes should probably be taken with some care, once again because of the short samples we have available.

In conclusion, while it appears that important changes in the business cycle characteristics of our indicators have occurred in the sample, there is weak evidence that the Maastricht treaty and the creation of the ECB had anything to do with them. Our inability to make strong causal statements is due to the fact that only short samples are available, that the evidence reflects in-sample (ex-post) information and that many events took place since 1999, which make the post ECB period unique. To acquire more evidence on the issue, assess whether the changes we observe can be given a causal interpretation and, if this is the case, understand the reasons which may give rise to the existing structural changes, we now turn to two forecasting exercises which are relatively free from small sample size problems and are less prone to interpretation fallacies.

### 5.2 An unconditional forecasting exercise

In this subsection we report the results of a few unconditional forecasting exercises conducted using the information available prior to the Maastricht treaty and the creation of the ECB and afterwards. Given the information available at 1992:3, 1998:3, 2001:4 we compute the predictive distribution for forecasts up to 5 years ahead. Since future parameter uncertainty is averaged out the bands we present reflect the uncertainty present in the data, conditional on the exogenous variables in the model taking the values that actually realized in the forecasting sample. Since there are many variables in the system and three dates at which we forecast, the number of graphs one may present is very large. Given that results do not appear to depend on the variable we choose, we only report in figures 3 to 5 the 90 percent highest credible GDP growth predictive distributions an the actual GDP growth values at the three dates of interest and briefly discuss what happens when forecasting employment, consumption, investment and industrial production growth. The full set of forecasting results is presented in an appendix available on request.

The three figures suggest that a clean and once-and-for-all structural break following the Maastricht treaty and the creation of the ECB is absent. To put this conclusion in another way, the time varying structure of our model appears to adequately capture the dynamics of real variables in the seven countries around the two political events we consider without any need to single out these events in any way. Furthermore, comparing across figures, it is clear that the forecasting performance of the model is roughly unchanged, at least for the four major countries. That is to say, the variations in the in-samples statistics of the median indicators we reported in table 1, are not associated with changes in predictive ability of the model at the dates of interest.

Overall, our predictive distributions of GDP growth have the right direction and approximately the right magnitude in the four major countries at all three selected dates. Before the Maastricht treaty and after the creation of the ECB, they also replicate the persistence of the actual series. In Belgium, the Netherlands and Finland the sign of the one-step ahead prediction is right but the direction of the subsequent changes is off in the period before Maastricht. This pattern improves substantially up to the creation of the ECB and after the creation of the ECB the predictive distributions capture the persistent pattern in the growth rates of GDP of these countries.

The difference are even less noticeable and the pattern much more similar across countries when we look at employment, consumption or investment growth: the model is able to capture well both the sign and the magnitude of the initial impact in the seven countries in all three periods; tracks reasonably well the persistence of the actual data and it is able to account for the medium term trends in these variables at all three dates, except for Germany and Belgium employment growth before the Maastricht treaty. Relatively speaking, the performance is less appropriate for industrial production growth in the sense that the direction of the forecast is often wrong after a few periods, but once again this seems to be independent of the date we have chosen to compute predictive distributions.

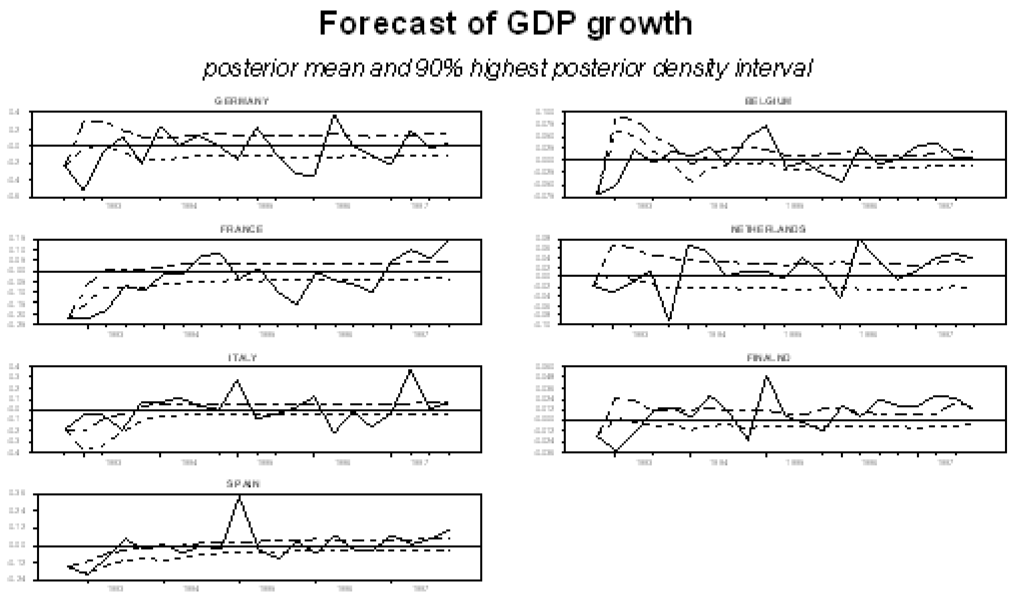


Figure 3: Pre Maastricht information

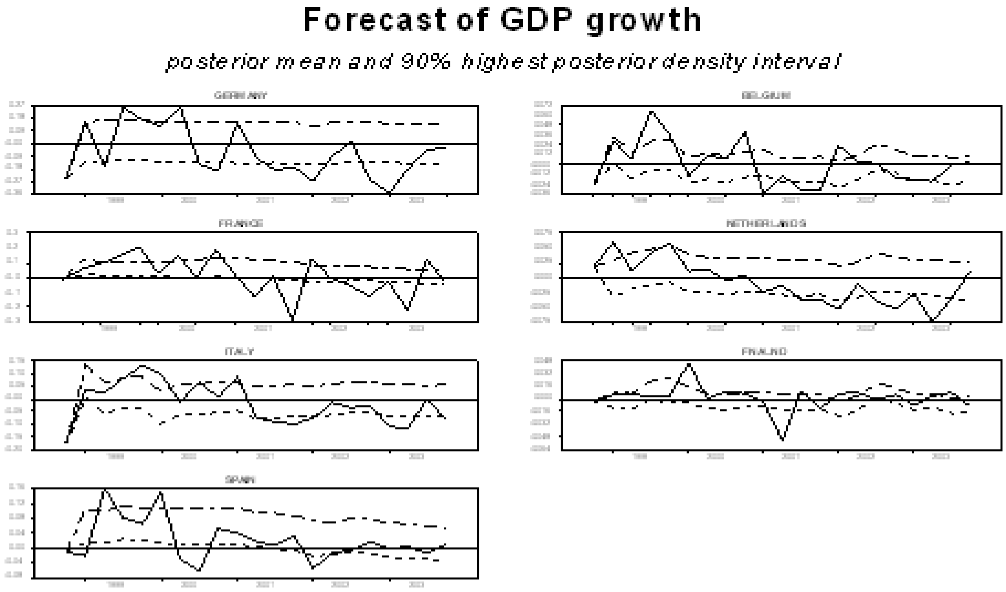


Figure 4: Pre ECB information

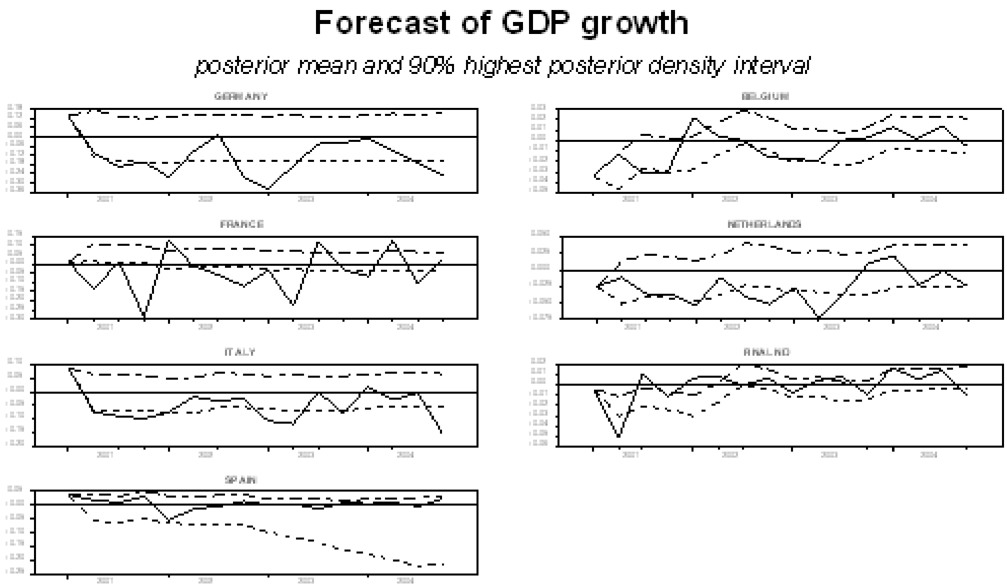


Figure 5: 2001 information

As an aside, we would like to stress that, while there is little serial correlation in output and industrial production growth and the model picks up that fact correctly, there is much stronger serial correlation in employment, consumption and investment growth data. This features allows our model, which emphasizes medium run trends, to correctly predict the dynamics of the growth rates of these three variables up to 3-4 years ahead.

To conclude, it is important to emphasize that the results of tables 1 and 2 and those of figures 3-5 are not necessarily in contrast with each other. Tables 1 and 2 present information constructed in-sample and on average over different time periods. What we present in figures 3 to 5 are instead case study exercises, where the out-of-sample predictive ability of the model is measured at particular dates. Hence, it is possible to observe in-sample average variations without affecting the unconditional forecasting ability of the model a particular dates. In fact, it is these changes in the features of available data that make the time varying structure of the model particularly appealing and allows to have an unchanged forecasting performance throughout the sample.

### 5.3 Two Conditional forecasting exercises

Unconditional forecasting exercises are a useful benchmark to understand the potential impact of the Maastricht treaty and the creation of the ECB on the dynamics of the cyclical fluctuations in the Euro area. However, such exercises will not be informative about the dynamics and the time variations obtained conditional on a particular type of shocks. In particular, an unchanged unconditional forecasting performance could be consistent with varying conditional forecasting performance, as long as the structural changes approximately average out across shocks.

To gather further information about the issues of interest, we have decided to undertake a couple of conditional forecasting exercises, one examining the transmission of a German shock on the variables of the other countries, and one examining the transmission of external shocks on the economies of the Euro area. Among all potential options, we choose these two since they may tell us something about the nature of the transatlantic transmission of disturbances, the magnitude of the synchronization and the qualitative nature of the heterogeneities which are present among the countries we examine. We first consider a shock which jointly and temporarily increase all German variables for one period; then we consider a temporary increase in the US Federal funds rate which lasts one period.

Once again, given the large number of variables in the system, we need to select which responses to report and at which date. Figures 6 to 8 present the responses of output growth in the seven countries to a German shock using the information available at 1992:3, 1998:4 and 2001:4. A brief discussion of what happens to the other four variables of system in response to these disturbances follows.

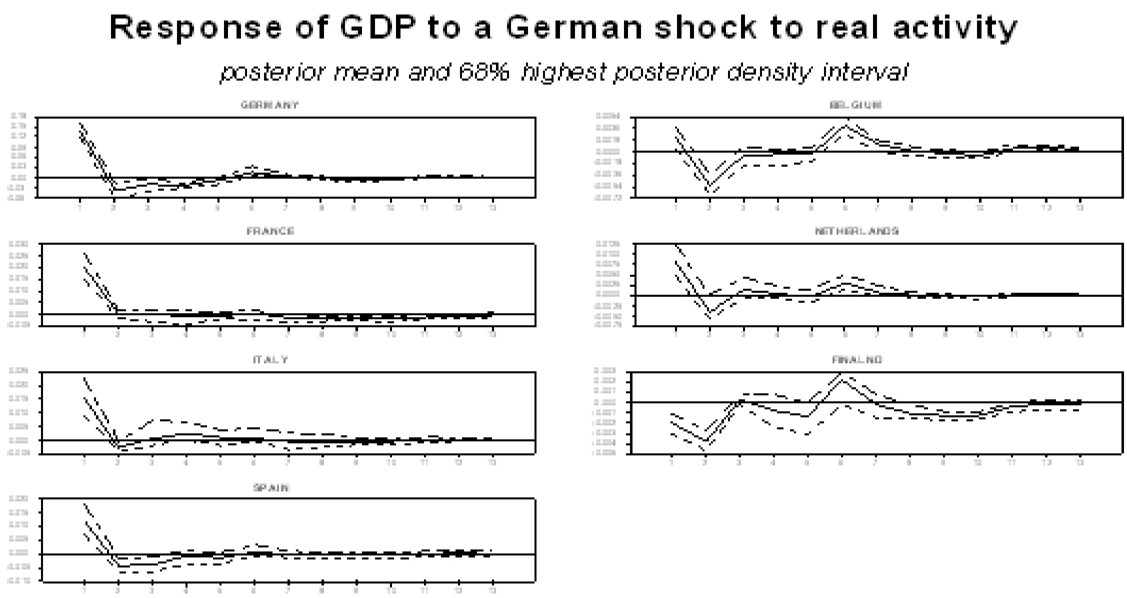


Figure 6: Pre Maastricht information

Figure 6, which reports the responses obtained at 1992:3 indicates that there is a lot of commonalities in the responses in the four large countries: a German impulse instantaneously displace GDP growth in Germany, Italy, France and Spain; the magnitude of the response is largest Germany and relatively similar in France, Italy and Spain; the effect lasts one period in France and Italy and about three periods in Spain and about 5 in Germany. The responses in Belgium and the Netherlands show a much more cyclical behavior but, the largest effects is also contemporaneous and little propagation takes place. Finally, a positive impulse in Germany, makes output growth fall in Finland and the effect appears to reverberate after about two years. Comparing figures 6 and 7 one can see that the shape of the responses are largely unchanged in 1998:4. Relatively speaking, the output growth effects in all countries is larger but the instantaneous international transmission weaker; effects are more persistent in France and Spain, where the positive effect lasts for about two years, and displays a much stronger commonalities in the timing of the responses - the exception still being Finland output growth who displays a positive response only after two quarters. Finally, comparing figure 6 and 7 with figure 8 one can notice both qualitative and quantitative changes in the responses after the creation of the ECB. The largest changes occurs in France, where GDP growth now peaks after 10 quarters, and in Belgium where GDP growth is persistently about the initial level for more than 4 years. However, also the responses of Spain and the Netherlands are altered - this time back to what they look like in 1992:3.

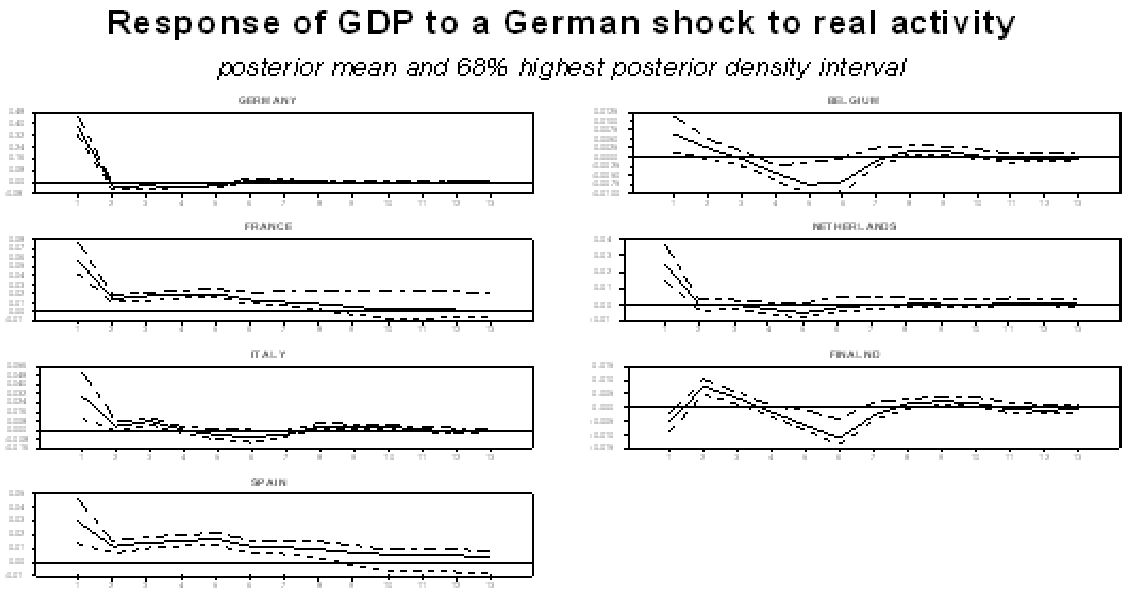


Figure 7: Pre ECB information

In general, changes up to 1998:4 appear to have limited and primarily concern the magnitude of the effects, even though a larger synchronicity in output responses is evident. After that date, both qualitative and quantitative changes are present in the responses of output growth across countries. Interestingly, these changes seem to go in the opposite direction of what one would a-priori expect and this make the dynamics in response to a German shock much more heterogeneous in 2001:4 than at any other date we consider.

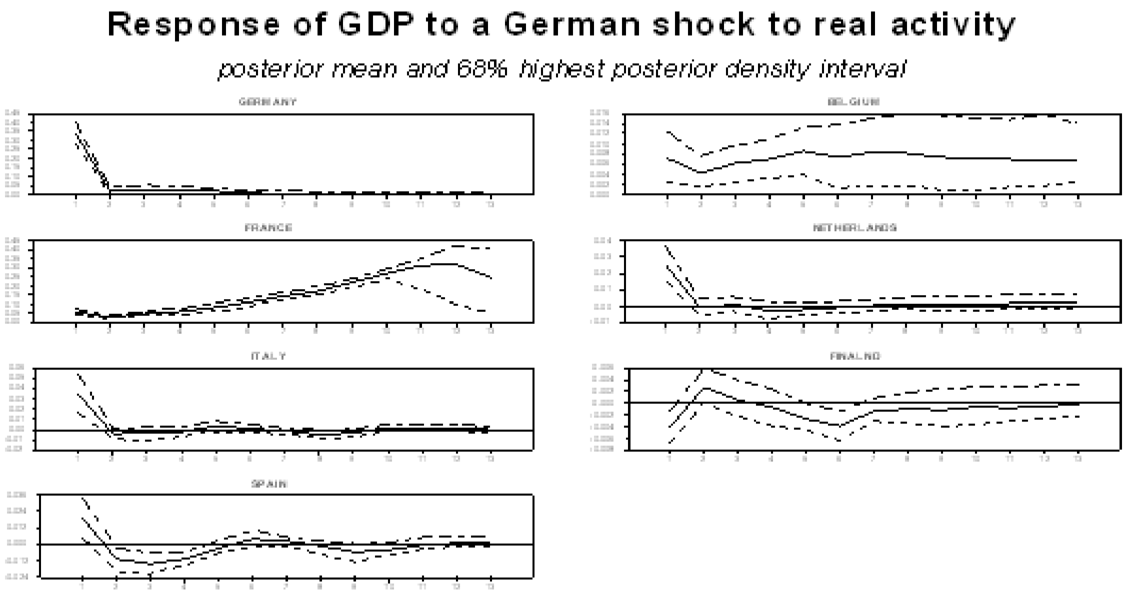


Figure 8: 2001 information

The responses of employment growth are somewhat different and their dynamic over time also appear to be different. In response to a German shock, employment growth in Germany, Italy and Spain all instantaneously increase at 1992:3 while the peak response in France is lagged by one period and the responses for the other three countries are insignificantly different from zero. In 1998:4 German and Italian employment growth instantaneously peak in response to the shock, but in all other five countries employment growth responses are delayed and typically peak after 2 to 5 periods even though the magnitude of the peak response is comparable across the two time periods. Finally, employment growth responses in 2001:4 track pretty closely output growth responses, with the exception of the Netherlands, where employment growth responses become significant only after 10 quarters. Relatively speaking at all three dates employment growth responses are always smaller than output growth responses except in Finland

The responses of the other three variables and their time profile are intermediate between those of output growth and employment growth. In particular, the pattern of responses of consumption growth is similar to those of employment growth while industrial production and investment growth closely track over time those of output growth.

While it is difficult to draw firm conclusions, it appears that the changes in the transmission of German shocks was smaller after the Maastricht treaty than after the ECB was created. However, different variables seem to have changed differently over time. In particular, while the transmission of disturbances to GDP growth has changed the least, the transmission of disturbances to employment growth display important differences with labor markets responding much strongly in the middle period and becoming much more synchronous with the good markets in the latter one.

(To Be Continued)

## 6 Conclusions

This paper attempts to shed some light on the effect of changes in institutions on the dynamics of business cycles by focusing on the recent European experience and the two dramatic changes occurred in the 1990s: the Maastricht treaty and the creation of the ECB. Given that this area of research is largely unexplored, we limit our attention in this paper to three particular questions. First, we have tried to provide evidence on whether there has been any tendency for areawide and national cycles to change after the Maastricht treaty and the creation of ECB. Second, we have attempted to assess whether a clean structural break took place in the European economy when these events occurred. Third, we tried to measure whether the two events had different relative impact on the cyclical characteristics of the data and which variables reacted most to the changes.

To study these questions we employ a panel VAR model of the same type employed by Canova and Ciccarelli (2004) and Canova, et al. (2005), using data for five variables form seven different countries for the sample 1980:1-2004:4. We report how areawide and national indicators have evolved over time, some reduced form statistics over subsamples and the features of the cyclical phases they display over different subsamples. We also conducted two types of forecasting exercises, an unconditional one, using the information at 1992:3, 1998:4 and 2001:4, and a conditional one, tracing out the dynamics of the system in response to two different types of shocks at the same dates.

Three major conclusions come out of our work. First, we find some evidence of changes in the features of European business cycles. In particular, we find a stronger areawide cycle in the 1990s and a change in the volatility, the persistence and the correlation of national cycles over time. These changes however, do not seem to line up exactly with the two events we consider and,in fact, seem to predate them by a couple of years. Second, in an unconditional forecasting sense, the Maastricht treaty and the creation of the ECB did not make a huge difference. Using the information available before and after these two events we are able to unconditionally predict both the direction and the magnitude of the changes in the five variables of the system for the majority of the countries with similar precision. Hence, while the average in-sample evidence suggests some changes in the cyclical features of European data over the last 15 years, the out-of-sample evidence at selected dates indicates that business cycles have not displayed any structural breaks. Third, in a conditional forecasting sense, we find that the structure of the economy was somewhat altered after 1992 and after 1998. Once again conditioning on the information available before and after the two events we show that shocks originating in Germany induced first more synchronicity in cross country responses and then a larger heterogeneity. In general, while changes occur at all dates, the Maastricht treaty seems to have had smaller effects than the creation of the ECB.

Given the small sizes of the samples and the fact that the last subsample includes a number of events which can potentially account for the changes we observe, we tentatively conclude that our data do not overwhelmingly support the causality between the two political events we are interested in and the changes in business cycles we have detected.

**Appendices**

# A Estimation

## A.1 Prior information

We let $\bar{B}_i = b_i * I,\ i = 1, 2, 3,$ where $b_i$ is a parameter which controls the tightness of factor $i$ in the coefficients, and $p(\Omega^{-1}, \sigma^2, b_i, \theta_0) = p(\Omega^{-1})p(\sigma^2)p(\theta_0)\prod_i p(b_i)$ with

$$\begin{aligned} p(\Omega^{-1}) &= Wi(z_1, Q_1) \\ p(\sigma^2) &= IG\left(\frac{\zeta}{2}, \frac{\zeta s^2}{2}\right) \\ p(b_i) &= IG\left(\frac{\varpi_0}{2}, \frac{\delta_0}{2}\right) \\ p\left(\theta_0 \mid \mathcal{F}_{-1}\right) &= N\left(\bar{\theta}_0, \bar{\mathsf{R}}_0\right) \end{aligned} \tag{12}$$

where $N$ stands for Normal, $Wi$ for Wishart and $IG$ for Inverse Gamma distributions, and $\mathcal{F}_{-1}$ denotes the information available at time $-1$. The prior for $\theta_0$ and the law of motion for the coefficient factors imply the prior for $\theta_t$ is $p\left(\theta_t \mid \mathcal{F}_{t-1}\right) = N\left(\bar{\theta}_{t-1|t-1}, \bar{\mathsf{R}}_{t-1|t-1} + B_t\right)$.

We collect the hyperparameters of the prior in the vector

$$\mu = \left(z_1, \zeta, s^2, \varpi_0, \delta_0, \gamma_1, \gamma_2, vech(Q_1), \bar{\theta}_0, vech\left(\bar{\mathsf{R}}_0\right)\right)$$

where $vech\left(\cdot\right)$ denotes the column-wise vectorization of a symmetric matrix. We assume that the elements of $\mu$ are either known or can be estimated in the data, for example, splitting the sample in two pieces, using the first part ("training" sample) to estimate the $\mu$ and the second to estimate posterior distributions and to conduct inference. We have experimented with both informative and noninformative priors and report results obtained with the latter set of priors. Table A.1 present the hyperparameters values.

Table A.1: Prior hyperparameters

| $\zeta$ | $s^2$ | $z_1$ | $Q_1$ | $\varpi_0$ | $\delta_0$ | $\gamma_1$ | $\gamma_2$ | $\bar{\theta}_0$ | $\bar{\mathsf{R}}_0$ |
|---|---|---|---|---|---|---|---|---|---|
| 0.0 | $\hat{\sigma}^2$ | $N \cdot G + 1$ | $\hat{Q}_1$ | 0 | 0 | 1 | 0 | $\hat{\theta}_0$ | $\mathrm{I}_J$ |

Here $\hat{\sigma}^2$ is the average of the estimated variances of NG AR(p) models, $\hat{Q}_1$ is the estimated variance-covariance of the time invariant version of (1), $\hat{\theta}_0$ is obtained with a sequential OLS on (1), over the sample 1975-1980, and $J$ is the dimension of $\theta_t$. The values of the remaining hyperparameters have been chosen using previous experience.

### A.2 Posterior distributions

To calculate the posterior distribution of the unknowns $\phi = (\Omega^{-1}, b_i, \sigma^2, \{\theta_t\}_{t=1}^T)$, we combine the prior with the likelihood of the data, which is proportional to

$$L \propto \left(\prod_{t=1}^{T} \sigma_t^{-NG/2}\right) |\Omega|^{-T/2} \exp\left[-\frac{1}{2}\sum_t \left(Y_t - W_t \Xi \theta_t\right)' \left(\sigma_t \Omega\right)^{-1} \left(Y_t - W_t \Xi \theta_t\right)\right] \tag{13}$$

where $Y^T = (Y_1, ..., Y_T)$ denotes the data, and $\sigma_t = (1 + \sigma^2 X_t' X_t)$. Using Bayes rule, $p\left(\phi \mid Y^T\right) = \frac{p(\phi)L\left(Y^T|\phi\right)}{p(Y^T)} \propto p\left(\phi\right) L\left(Y^T \mid \phi\right)$. Given $p\left(\phi \mid Y^T\right)$, the posterior distribution for the components of $\phi$, $p\left(\Omega \mid Y^T\right)$, $p\left(b_i \mid Y^T\right)$, $p(\sigma^2 \mid Y^T)$ and $p\left(\{\theta_t\}_{t=1}^T \mid Y^T\right)$, can be obtained by integrating out nuisance parameters from $p\left(\phi \mid Y^T\right)$. Once these distributions are obtained, location and dispersion measures for $\phi$ and for any interesting continuous function of them can be obtained.

For the model we use, it is impossible to compute $p\left(\phi \mid Y^T\right)$ analytically. However, we can numerically simulate a sample from it using Monte Carlo techniques. A method which is particularly useful in our context is the Gibbs sampler since it only requires knowledge of the conditional posterior distribution of $\phi$. However, while the conditional posteriors of $\Omega^{-1}, b_i$ and $\{\theta_t\}_{t=1}^T$ are available in closed form, the conditional posterior of $\sigma^2$ is not and a Metropolis step within the Gibbs sampler is needed.

Denoting $\phi_{-\kappa}$ the vector $\phi$ excluding the parameter $\kappa$, the conditional distributions of interest are

$$\begin{aligned}
\theta_t \mid Y^T, \phi_{-\theta_t} &\sim N\left(\bar{\theta}_{t|T}, \bar{\mathsf{R}}_{t|T}\right) \quad t \leq T, \\
\Omega^{-1} \mid Y^T, \phi_{-\Omega} &\sim Wi\left(z_1 + T, \left[, \frac{\sum_t \left(Y_t - W_t \Xi \theta_t\right)\left(Y_t - W_t \Xi \theta_t\right)'}{\sigma_t} + Q_1^{-1}\right]^{-1}\right) \\
b_i \mid Y^T, \phi_{-b_i} &\sim IG\left(\frac{\varpi^i}{2}, \frac{\sum_t \left(\theta_t^i - \theta_{t-1}^i\right)'\left(\theta_t^i - \theta_{t-1}^i\right) + \delta_0}{2\xi_t}\right) \\
\sigma^2 \mid Y^T, \phi_{-\sigma^2} &\propto L\left(Y^T \mid \phi\right) \times p(\sigma^2)
\end{aligned} \tag{14}$$

where $\bar{\theta}_{t|T}$ and $\bar{\mathsf{R}}_{t|T}$ are the one-period-ahead forecasts of $\theta_t$ and the variance-covariance matrix of the forecast error, respectively, calculated with a simulation smoother, as described in Chib and Greenberg (1995), and $\varpi^1 = T+\varpi_0$, $\varpi^2 = Tg+\varpi_0$ and $\varpi^3 = TN+\varpi_0$.

The posterior for $\sigma^2$ is simulated using a Random Walk Metropolis algorithm, where, at each iteration $l$, we generate candidate draws according to $\left(\sigma^2\right)^* = \left(\sigma^2\right)^{(l-1)} + z$ where $z$ is a normal random variable with mean zero and variance chosen to ensure that the acceptance acceptance rate is about 0.3.

Under regularity conditions (see Geweke (2000)), cycling through the conditional distributions in (14) will produce in the limit draws from the joint posterior of interest. From these, marginal distributions can be computed averaging over draws nuisance dimensions. In particular, using the draws, the posterior distributions of $\lambda_t$ and $\alpha_t$ can be estimated using kernel methods and, in turns, the posterior distributions of $WLI_t$ and $CLI_t$ can be obtained. For example, a credible 68% interval is obtained ordering the draws of $WLI_t^h$ and $CLI_t^h$ for each $t$ and taking the 16th and the 84th percentile of the distribution.

Because we are not directly sampling from the posterior, it is important to monitor that the Markov chain induced by the sampler converges to the ergotic (posterior) distribution. We have check convergence in several ways: increasing the length of the chain, splitting the chain in two after a burn-in period and calculating whether the mean and the variances are similar; checking if cumulative means settle at some value. The result we present are based on chains with 33000 draws: 1100 blocks of 30 draws were made and the last draw for each block is retained after the discarding the first 3000. This means that a total of 1000 draws is used at each $t$ to conduct posterior inference.

Predictive distributions are obtained drawing $\phi_t$ from their posterior and the law of motion of the coefficients, averaging over $\theta_t$. Impulse responses are obtained as described in Canova and Ciccarelli (2004).

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