Do political economy events affect business cycles? 
Evidence for the European Monetary Union*

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Abstract
This paper provides evidence on the effects of three political economy events (the Maastricht treaty, the creation of the ECB, and the Euro changeover) on the dynamics of European business cycles. We use a panel VAR approach and data from ten European countries - seven are part of the Euro area and three outside it. We find evidence of slow changes in the features of business cycles and in the transmission of shocks over time but fail to relate the timing of the changes with the three events. Changes appear to be linked to a general process of European convergence and synchronization.

JEL classification: C15, C33, E32, E42

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1 Introduction

There is abundant evidence supporting the view that real activity in developed countries displays common characteristics. For example, Del Negro and Otrok (2003), Giannone and Reichlin (2006) and Canova et al. (2007) have shown, with different techniques, that the cyclical features of output and industrial production are similar within Euro area countries and between the Euro 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 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 OECD cycles after 2000; Stock and Watson (2003) document changes in the volatilities of G-7 cycles in the 1990s, while Canova et al. (2007) document variations in the correlation structure of cyclical fluctuations of G-7 countries since the middle of the 1980s.

Why are the cyclical features changing? At least two explanations come to mind. First, variations in structural characteristics, such as preferences and objective functions of agents, the way expectations about future events are formed, or operational features of goods, labor and financial markets, have altered the transmission of shocks within and across countries. The first two options are typically invoked 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 possible explanation is that the nature, the characteristics and the frequency of the shocks hitting developed economies has dramatically changed. For example, Sims and Zha (2006) and Canova and Gambetti (2007) among others, 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. Stock and Watson (2003) have suggested that changes in the volatility of the shocks could also affect the magnitude and the direction of the correlation among international macroeconomic variables. Helbling and Bayoumi (2003) have claimed that
common shocks are now more frequent than used to be. Finally, political economy events may have altered the nature and the causes of fluctuations.

To the best of our knowledge, no one has tried to link changes in the nature of business cycles in developed countries with their political economy changes. This seems to be an important shortcoming since, at least in Europe, the political and institutional arena has witnessed important variations over the last 20 years. There are good reasons why the literature has largely shied away from the topic. Institutions and the nature of markets typically change slowly making it difficult to pin down a potential break point date and select subsamples over which to compare business cycle features. In addition, political economy or institutional changes may affect cycles which have much longer periodicity than the ones typically associated with business fluctuations and externalities and threshold effects are likely to be important when measuring these effects. 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 that political economy events have on the dynamics of business cycles by focusing on the three important ones occurred in the last 20 years in Europe: the Maastricht treaty, the creation of the ECB and the Euro changeover. Investigating the effects of these events is relevant from at least four different perspectives. First, since these events were brought about by national politicians and were, to a large extent, exogenous with respect to the dynamics of the European economies, the experience is unique to verify some well known implications of the common currency area literature. Does real convergence naturally precede the establishment of common monetary institutions or the reverse holds true? Second, two of the events are monetary in nature. The ability of monetary factors to affect real variables at business cycle frequencies has been extensively studied and limited effects have typically been found. However, the nature of the events substantially differs from those typically studied in the literature (where e.g. shocks to policy rules and, to a much lesser extent, the variations in the policy rules employed are analyzed) and their effects are a-priori comparable to the establishment of the Fed or the breakdown of the gold standard, which are known to have made quite a
difference for cyclical fluctuations around the world (see e.g. Bergman, Bordo and Jonung (1998)). Third, in macroeconomic analyses it is common to separate cyclical movements from other types of fluctuations claiming that the mechanism generating the two types of fluctuations is different. If political economy events, besides affecting the medium-long run tendencies of the economy, also exercise a significant impact on the business cycle, such a practice should be reconsidered. Finally, policymakers are typically concerned with the effects of national idiosyncrasies. If institutional changes cause similar variations in economic activity in countries with different structures, the role of national idiosyncrasies should be discounted and national policies put in proper perspective.

Since the subject is vast and largely unexplored, we limit our attention to three questions. First, has there been any tendency for European, Euro area, and national cycles to change after the Maastricht treaty, the ECB creation or the Euro changeover? Since we have data on both Euro area and non-euro countries, we have a natural controlled experiment which can be particularly informative about the cyclical consequences of these events. Second, do we observe a break in the cyclical relationships around the dates when these events took place or did the changes happen slowly and more continuously over time? Third, what is the relative impact of the three events on the cyclical features of the data? Our prior in this respect is that the Maastricht treaty, forcing a process of convergence in government spending and debt, should have had more sizeable effects than the other two events.

To study these questions we use a panel VAR model of the type employed by Canova and Ciccarelli (2004) and Canova et al. (2007). The setup is useful in our context for at least two reasons: it handles large scale models displaying unit specific dynamics and cross country lagged interdependencies, and flexibly allows for time variations in the correlation structure across variables and units; it automatically produces an index structure, where the posterior distribution of common, regional and national specific cyclical indicators can recursively be constructed. We use data from Germany, France, Italy, Spain, Belgium,

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1 The focus of the analysis is on the direction, the magnitude and the intensity of the variations which are observed after each episode on average across variables.
the Netherlands, Finland, the UK, Denmark and Sweden for output, employment, industrial production, consumption and investment, and construct distributional measures of common, regional and national cycles using the sample 1970Q1-2007Q3. We perform forecasting exercises with information available before each event and to trace out the effect of shocks to interesting variables/country combinations to the variables of the system before and after the events.

We find that the posterior distributions of our cyclical indicators have changed over time. Changes are slow and seem to take place largely in the 1980s and early 1990s. For example, a European-wide cycle, which was largely absent until mid-1980s, became much more significant over the last two decades. Furthermore, cyclical convergence occur earlier for the countries belonging to the EMU, but clearly takes places for all ten countries in the early 1990s. Moreover, estimates of our cyclical indicators show decreasing volatility since the early 1990s, and area wide and national indicators become highly synchronized around the same time.

We fail to detect a clean and once-and-for-all structural break in the properties of cyclical fluctuations following the three events. In fact, we are able to unconditionally predict both the direction and the magnitude of the changes after the three events in the five variables for the euro area countries. This is consistent with the fact that, for the large euro area countries, the process of cyclical convergence had largely taken place prior to 1992. On the other hand, we do find a break in the properties of UK, Denmark and Sweden cycles after the Maastricht treaty, confirming that the business cycles of these countries converged toward the European average after the Maastricht Treaty.

The transmission of shocks was somewhat altered over time, but we fail to find a direct connection between these alterations and the political events of interest. The already mentioned convergence process taking place in all European countries, in general, makes them respond more to European shocks and less to non-European shocks since the mid 1990s. Some reversal of these tendencies took place in the 2000s but, once again, these changes do not appear to be related to the institutional changes we examine.

All in all, the changes we detect in the features of business cycles and in the transmis-
sion of shocks do not appear to be easily related to the political economy events we are interested in. Rather, they appear to be the result of a general process of slow convergence and increased synchronization which has taken place in European countries since the middle of 1980s.

The rest of the paper is organized as follows: the next section presents the model specification, the technique used to construct the indicators, and the procedure employed to compute unconditional and conditional forecasts. Section 3 presents the data and some specification checks. Section 4 has the results and section 5 concludes.

2 The empirical model

The model we employ has the form:

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

where \( i = 1, \ldots, N \) refers to countries and \( t = 1, \ldots, T \) to time. \( y_{it} \) is a \( G \times 1 \) vector for each country \( i \) and \( Y_t = (y_{1t}, y_{2t}, \ldots, y_{Nt})' \). \( D_{it,j} \) are \( G \times NG \) matrices for each lag \( j = 1, \ldots, p \), \( W_{it} \) is a \( Mq \times 1 \) vector of exogenous variables and \( F_{it,j} \) are \( G \times M \) matrices each lag \( j = 1, \ldots, q \) and \( e_{it} \) is a \( G \times 1 \) vector of random disturbances.

The model (1) displays three important ingredients for our analysis. First, coefficients are allowed to vary over time. We believe this is a crucial ingredient when examining the importance of institutional changes. If we do not allow for time variations in the coefficients, it is very much possible to attribute changes in business cycle features to the events of interest when in fact they occur slowly and smoothly over time. Second, the dynamic relationships are allowed to be unit specific. Without such a structure, heterogeneity biases may crucially distort our inference and convergence issues may not be examined. Third, whenever the \( NG \times NG \) matrix \( D_t(L) = [D_{1t}(L), \ldots, D_{Nt}(L)]' \), is not block diagonal for some \( L \), cross-unit lagged interdependencies are present. Such a feature allows dynamic feedback across units and greatly expands the type of interactions one can analyze in the model.

While these features add considerable realism to the specification, and avoid the
“incredible” short-cuts that the literature has often taken (see Canova and Ciccarelli 2004, for discussion), they are not costless: the number of parameters is large (there are \( k = NGp + Mq \) 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:

\[
Y_t = Z_t \delta_t + E_t \quad E_t \sim N (0, \Omega)
\]

where \( Z_t = I_{NG} \otimes X_t' \), \( X_t' = (Y_{t-1}', Y_{t-2}', \ldots, Y_{t-p}', W_{t}', W_{t-1}', \ldots, W_{t-q}') \), \( \delta_t = (\delta_{1t}', \ldots, \delta_{Nt}')' \) 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:

\[
\delta_t = \Xi_1 \lambda_t + \Xi_2 \alpha_t + \Xi_3 \rho_t + \Xi_4 \psi_t + u_t
\]

where \( \Xi_1, \Xi_2, \Xi_3, \Xi_4 \) are matrices of dimensions \( NGk \times s, NGk \times N, NGk \times G, NGk \times 1 \) respectively and \( \lambda_t, \alpha_t, \rho_t, \psi_t \) are mutually orthogonal. Here \( \lambda_t \) captures movements in the coefficient vector which are common across countries and variables (or groups of them) and is of dimension \( s \); \( \alpha_t \) captures movements in the coefficient vector which are common within countries and therefore its dimension equals to \( N \); \( \rho_t \) captures movements in the coefficient vector which are variable specific and its dimension is therefore equal to \( G \); while \( \psi_t \) is a scalar process which captures movements in the coefficients due to the \( M \) 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 \( s + N + G + 1 \) factors characterizing their dynamic features. Practically, the factorization (3) transforms an overparametrized panel VAR into a parsimonious SUR model where the regressors are

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averages of certain right-hand side variables of the VAR. In fact, substituting (3) into (2) we have

\[ Y_t = Z_{1t} \lambda_t + Z_{2t} \alpha_t + Z_{3t} \rho_t + Z_{4t} \psi_t + v_t \]

(4)

where \( Z_{1t} = Z_t \Xi_1 \), \( Z_{2t} = Z_t \Xi_2 \), \( Z_{3t} = Z_t \Xi_3 \), \( Z_{4t} = Z_t \Xi_4 \) 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 common 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, \( WLI_t = Z_{1t} \lambda_t \) plays the role of a common indicator, while \( CLI_t = Z_{2t} \alpha_t \) plays the role of a vector of country specific indicators. Both coincident and leading versions of these indicators can be designed using time \( t - j \), \( j = 0, 1, 2, \ldots \) information (see Canova and Ciccarelli (2004) for this), and are constructed recursively, 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 \( Z_{1t} \) and \( Z_{2t} \) - but become uncorrelated as the number of countries becomes large.

To illustrate the structure of the \( \Xi \)'s and the nature of \( Z_{jt} \)'s, suppose there are \( G = 2 \) variables, \( N = 2 \) countries, \( s = 1 \) common component, \( p = 1 \) lags, no exogenous variables and an intercept of the form

\[
\begin{bmatrix}
  y^1_t \\
  x^1_t \\
  y^2_t \\
  x^2_t
\end{bmatrix} =
\begin{bmatrix}
  d^1_{1,1} & d^1_{1,2} & d^1_{1,3} & d^1_{1,4} \\
  d^1_{2,1} & d^1_{2,2} & d^1_{2,3} & d^1_{2,4} \\
  d^2_{1,1} & d^2_{1,2} & d^2_{1,3} & d^2_{1,4} \\
  d^2_{2,1} & d^2_{2,2} & d^2_{2,3} & d^2_{2,4}
\end{bmatrix}
\begin{bmatrix}
  y^1_{t-1} \\
  x^1_{t-1} \\
  y^2_{t-1} \\
  x^2_{t-1}
\end{bmatrix} +
\begin{bmatrix}
  c^1_t \\
  c^2_t \\
  c^3_t \\
  c^4_t
\end{bmatrix} + e_t
\]

(5)

Here \( \delta_t = [d^1_{1,1,1}, d^1_{1,2,1}, d^1_{1,3,1}, d^1_{1,4,1}, d^1_{1,1,2}, d^1_{1,2,2}, c^1_t, d^1_{1,1,3}, d^1_{1,2,3}, d^1_{1,3,1}, d^1_{1,2,4}, c^1_t, d^1_{1,1,4}, d^1_{1,2,4}, c^1_t, d^2_{1,1,1}, d^2_{1,2,1}, d^2_{1,3,1}, d^2_{1,4,1}, d^2_{1,1,2}, d^2_{1,2,2}, d^2_{1,3,2}, d^2_{1,4,2}, c^2_t, d^2_{1,1,3}, d^2_{1,2,3}, d^2_{1,3,3}, d^2_{1,4,3}, c^2_t, d^2_{1,1,4}, d^2_{1,2,4}, d^2_{1,3,4}, d^2_{1,4,4}, c^2_t] \) 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

\[ \delta_t = \Xi_1 \lambda_t + \Xi_2 \alpha_t + \Xi_3 \rho_t + u_t \]

(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

\[
\Xi_2 = \begin{bmatrix}
\lambda_1 \\
\lambda_1 \\
0 \\
\nu_2 \\
0
\end{bmatrix}
\Xi_3 = \begin{bmatrix}
\nu_1 \\
0 \\
\nu_2 \\
0
\end{bmatrix}
\]

with \( \lambda_1 = (1 1 0 0 0)' \), \( \nu_2 = (0 0 1 1 0)' \), \( \nu_1 = (1 0 1 0 0)' \) and \( \nu_2 = (0 1 0 1 0)' \). Hence, the VAR (5) can be rewritten as

\[
\begin{bmatrix}
y_1^t \\
x_1^t \\
y_2^t \\
x_2^t
\end{bmatrix} = \begin{bmatrix}
Z_{1t} \\
Z_{1t} \\
Z_{2,t} \\
Z_{2,t}
\end{bmatrix} \lambda_t + \begin{bmatrix}
Z_{2,1,t} \\
0 \\
Z_{2,1,t} \\
0
\end{bmatrix} \alpha_t + \begin{bmatrix}
Z_{3,1,t} \\
0 \\
Z_{3,1,t} \\
0
\end{bmatrix} \rho_t + v_t
\]

where \( Z_{1t} = y_{t-1}^1 + x_{t-1}^1 + y_{t-1}^2 + x_{t-1}^2 + 1 \), \( Z_{2,1,t} = y_{t-1}^1 + x_{t-1}^1 \), \( Z_{2,2,t} = y_{t-1}^2 + x_{t-1}^2 \), \( Z_{3,1,t} = y_{t-1}^1 + y_{t-1}^2 \), \( Z_{3,2,t} = x_{t-1}^1 + x_{t-1}^2 \) and \( v_t = e_t + Z_t'u_t \). When \( \lambda_t \) is large relative to \( \alpha_t \), \( y_1^t \) and \( x_1^t \) comove with \( y_2^t \) and \( x_2^t \). On the other hand, when \( \lambda_t \) is zero, \( y_1^t \) and \( x_1^t \) may drift apart from \( y_2^t \) and \( x_2^t \). Note that, when \( p > 1 \), lags can be weighted using a decay factor as in Doan et al. (1984).

As the notation we used in the above example 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. The equal weighting scheme directly comes from (3) and the fact that all variables are measured in the same units, as all variables will be demeaned and standardized. Second, our indices dynamically span lagged interdependencies across countries and variables. Third, our indices are observable. Fourth, our loadings are allowed to be time varying. Finally, our averaging approach creates moving average terms of order \( p \) in the regressors of (4). Therefore, our indicators eliminate high frequency variability from the right hand side variables of the VAR.
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:

$$\delta_t = \Xi \theta_t + u_t \quad u_t \sim N(0, \Sigma \otimes V)$$

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

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

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 $\bar{B} = diag(B_1, B_2, B_3, B_4)$, and that $E_t$, $u_t$ and $\eta_t$ are mutually independent.

In (9) the factors evolve over time as random walks. 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. 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. In the empirical analysis we will assume an exact decomposition as a benchmark and set $\sigma^2 = 0$.

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

$$Y_t = (Z_t \Xi) \theta_t + v_t$$

$$\theta_t = \theta_{t-1} + \eta_t$$

where $v_t \sim (0, \Omega)$. To compute posterior distributions for the unknowns we need prior densities for $\varphi_0 = (\Omega, \bar{B}, \theta_0)$. 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 the appendix. 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.

Besides characterizing the time profile of the posterior distribution of interesting indicators, we will be interested in computing predictive distributions for future \( Y_{t+\tau} \), both unconditionally and conditionally. Both 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 above predictive density we condition on \( \theta_{t+\tau} = \theta_t \).

When we want to study conditional predictive distributions, 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 \( \theta_t \). Here too we condition on \( \theta_{t+\tau} = \theta_t \). Let \( \mathcal{W}_t = (\Omega, B_t) \); 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 \( \xi_{j,t+1}^\delta \) be a realization of \( \xi_{j,t+1} \) of size \( \delta \) and let \( \mathcal{F}^1_t = \{y^t, \theta^t, \mathcal{W}_t, J_t, \xi_{j,t}, \xi_{-j,t}, \xi_{t+1}^{t+\tau}\} \) and \( \mathcal{F}^2_t = \{y^t, \theta^t, \mathcal{W}_t, J_t, \xi_t, \xi_{t+1}^{t+\tau}\} \) 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 \( J_t J_t^\prime = I \). Then, responses at horizon \( \tau \) to an impulse in \( \xi_{j,t}^\delta, j = 1, \ldots \) are

\[
IR^\delta_y(t, \tau) = E(y_{t+\tau}|\mathcal{F}^1_t) - E(y_{t+\tau}|\mathcal{F}^2_t) \quad \tau = 1, 2, \ldots
\]

(11)

In this paper, we consider domestic German disturbances (shocks which simultaneously affect all German variables) and US short term interest 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. 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 and some specification checks

We use year-on-year demeaned and standardized quarterly growth rates of output, industrial production, employment, consumption and investment for Germany, France, Italy, Spain, Belgium, Netherlands, Finland, the UK, Denmark and Sweden for the sample 1970Q1 to 2007Q3. Since this set of countries covers most of the Euro Area and the three most relevant countries which declined joining the euro zone, we can properly test the relevance and the scope of the institutional changes for real fluctuations by measuring the difference across group of countries before and after the events.

The sample period is long enough to perform meaningful pre- and post-institutional changes exercises in all cases. The Maastricht Treaty was signed on February 7, 1992, but we take 1993Q4 as the cut-off point since it became effective only on November 1st, 1993; the ECB creation occurred on June 1st, 1998 (we take as cut-off point 1998Q3); and the Euro changeover occurred on January 1, 2002 (we take as cut-off point 2002Q1).

Industrial production is measured by its index and employment by the total employment index, both from OECD Main Economic Indicators. Output is measured by real GDP, consumption by total real private consumption expenditure and investment by real gross fixed capital formation, the three of them are measured in 2000 prices and taken from the OECD Economic Outlook database.

We use as exogenous variables the growth rates of non-energy commodity prices, oil prices, world trade, US GDP, US nominal interest rate and the NY stock market index. Non-energy commodity prices are world prices of primary commodities excluding energy from OECD Economic Outlook. Oil prices are quarterly average prices obtained from the IMF International Financial Statistic. World trade is measured by the total volume of world trade in goods and services in 2000 prices, and it is taken from the OECD Main Economic Indicators. US GDP data comes from the Bureau of Economic Analysis. US interest rate is measured by that on 3-month nationally traded certificates of deposit issued by commercial banks, obtained from the Federal Reserve Board. Since some variables
display seasonality despite being reported as seasonally adjusted at the source, we prefilter questionable series though TRAMO-SEATS. We use one lag of both endogenous and exogenous variables.\(^2\)

Before analyzing the questions of interest, it is useful to analyze the properties of the empirical model. Documenting the fit of the model is important because the outcomes of our exercises will be credible only if the model captures the data well and if our indicators reproduce important cyclical statistics of the data.

The model we use in the exercises was selected with a specification search where different nested and non-nested specifications were compared via marginal likelihood (ML). The marginal likelihood of model \(M_i\) is \(f(Y|M_i) = \int L(y|\phi_i, M_i)g(\phi|M_i)d\phi_i, \) where \(\phi_i = [\phi_{i1}, \ldots, \phi_{it}]\) is the vector of the parameters of \(M_i.\) \(M_i\) is preferred to \(M_{i'}\) if the Bayes factor \(BF(M_i, M_{i'}) = \frac{f(Y|M_i)}{f(Y|M_{i'})}\) substantially exceeds 1.

The alternative specifications we considered include a model with no country-specific dynamics (ML = -5723); a model where the variable factor is excluded from the specification (ML = -5486); a model where there is no factor for the exogenous variables and their coefficients are treated the same way as the coefficients on lagged endogenous variables (ML = -5343). The marginal likelihood of the model including unit-specific dynamics, one common factor and specific factors for the country, variable and exogenous components was found highest (ML = -5308).

We also experimented with two specifications for the common component: a single common cycle one \(s = 1;\) and an alternative one, \(s = 2,\) where we estimate two separate common cycles, one for the Euro countries and another one for the non-Euro countries. The latter model has ML = -5336, indicating that the evidence in favor of the single common cycle for both Euro and non-Euro area countries is overwhelming (log Bayes factor of 29). Figure 1, which plots the posterior median of the single common factor \(Z_{1t}\) (labelled ‘common all’), together with the posterior median of the two common factors \(Z_{11t}\) and \(Z_{12t}\) (labelled ‘common EMU’ and ‘common non EMU,’ respectively) shows

\(^2\)This implies that each of the 50 equations of the system has \(50 \times 1 + 6 \times 1 = 56\) coefficients and that the panel VAR has 2800 time varying coefficients.
why this is the case.

The Euro and non-Euro factors are very similar and they display fluctuations which are highly in phase with the single common factor. The only notable difference between Euro and non-Euro factors is that fluctuations in the former appear to be less volatile than those in the latter up to the early 1990s, but after that date no differences are noticeable. This result already provides negative evidence on the issues we care about. As the dynamics of business cycles in Euro area and Non-Euro area countries are roughly similar since the early 1990s, it is very unlikely that the creation of the ECB and the Euro changeover are crucial factors in understanding the variations of European business cycle characteristics.

4 Political economy events and real fluctuations

4.1 Preliminary evidence

To examine whether the Maastricht treaty, the creation of the ECB and the Euro changeover had anything to do with changes in the properties of European cycles, we first informally examine the timing of changes in dynamics of the common and of the country specific indicators we have estimated (see figures 2 and 3 and table 1). Lighter areas in figure 2 represent recessions according to the CEPR classification (www.cepr.org). Lighter areas in figure 3 represent official recessions dates as reported by the Economic Cycle Research Institute (ECRI)(www.businesscycle.com); for the Netherlands, Belgium, Denmark and Finland these are missing since no business cycle dating is available for these countries.

A number of interesting features are present in figures 2 and 3. First, all indicators are very precisely estimated and the common indicator has smaller variability than any of the national indicators. This is not entirely surprising, given that the common indicator averages information contained in the five variables for the ten countries.

Second, the time path of our common indicator shares important similarities with the synthetic Euro area GDP growth series (taken from the Area Wide Model dataset of the ECB). As shown in the last column in Table 1, the two indicators are highly correlated, show similar serial correlation but the synthetic Euro area GDP growth series displays higher volatility and seems to slightly lead the common indicator.
Third, the common indicator is characterized by different phases. Until mid-1980s, fluctuations were volatile and the indicator crossed the zero line often. Afterwards, the indicator stays away from the zero line for longer periods and fluctuates much less, particularly after the exchange rate crisis of early 1990s. Hence, while a “European cycle” can be identified throughout the period, it is only in the early 1990s that it acquires typical cyclical features.

The common indicator has four clear expansion phases (1985-90, 1995-96, 1998-2001, 2006-07), two strong recessions (1981-84, 1991-94) and a milder one (2001-05). The recession dates roughly correspond to those reported by the CEPR dating committee - no dating is available from that source after 2000. The business cycle phases of our national indicators are highly synchronized with those reported by ECRI. In fact, if we allow for one quarter (two quarters) of maximum discrepancy, the average coincidence of our growth cycle dating with that of the ECRI across countries is 58% (63%). Table 1 shows such coincidence rate for the individual countries. The length of expansions and recessions are similar, both for the common and the national indicators. Overall, our national indicators are quite heterogeneous in terms of timing, amplitude and duration of the fluctuations.

Fourth, the volatility of both the common and the national indicators falls in the early 1990s and falls even further after 2000 (the exception here is the German national indicator). This fall is in line with the reduction in the real business cycle volatility documented in the industrialized world (see e.g. Stock and Watson (2003)) and widely referred to as the Great Moderation. The analysis by subperiods in Table 1 reveals useful information in this respect. Interestingly, rather than happening in the early 1980s, the reduction in volatility appears to take place in early 1990s and therefore is distinct from the other well known phenomena. The reduction in volatility of the common indicator is accompanied by an increase in its persistence. On the contrary, national indicators display an initial reduction in persistence, which in some cases is reverted in the last five years, in particular for Germany and Denmark. Also, while business cycles phases appear to be roughly similar over time, the national indicators for Germany and Spain display longer expansions in the latter part of the sample. Finally, although it is hard to pin point an
exact date when this started, the national indicators of major countries tend to become more similar as time goes by.

These facts seem to square reasonably well with what is known in the literature. For example, in line with Canova et al. (2007), the strengthening of a common pattern in cyclical fluctuations does not imply that national cycles are disappearing. In fact, the stronger nature of the cyclical fluctuations showing up with time in the common indicator is not as one could think the result of an increase in the synchronization of business cycle phases, since the maximum correlation was the contemporaneous one in most cases for all subperiods (with some exception like the UK which moved from leading the EU-cycle to comoving with it in a synchronized fashion). Rather, it comes from a more intense comovement as reflected by the rising correlation coefficients. The contemporaneous correlation of all national indicators with the common indicator increases over time (Denmark indicator is the exception).

Artis and Zhang (1997) analyzed business cycles statistics computed using standard filtering methods before and after 1979 - the period of the first European Monetary System (EMS). They find an increase in the degree of conformity and the degree of synchronization in the fluctuations of the countries participating to the first monetary system, an increase which was not present in non-EMS countries. Table 1 shows that the post-Maastricht and post-ECB samples roughly display similar volatility and persistence of both common and national cycles, and very different to the pre-Maastricht sample. In line with our a-priori expectations, the magnitude of the changes is larger in the samples constructed after the Maastricht treaty. However, as table 1 shows, these changes in the characteristics of business cycles appear to be present even in the sample starting in 1985.

In sum, we find it hard to relate observed changes in the business cycle characteristics of our common and national indicators with the creation of the ECB as well as with the Euro changeover, both in absolute and in relative terms. In particular, we fail to find significant changes in the time series of our indicators around the time when these events occurred. Moreover, Euro area countries do not behave differently than non-Euro area countries in this respect. In theory, these two events should have affected differently the
business cycles of the two groups of countries. Our failure to find significant differences within countries across groups of countries casts some serious doubts about the real effects produced by these events.

Instead, there are detectable changes around the date the Maastricht Treaty was implemented which could, at first sight, support the view that this event mattered for real fluctuations. For example, the common component of the fluctuations in European countries becomes more marked since the early 1990s and the volatility of all indicators declines since the early 1990s. Furthermore, both the decline in volatility and the increase in correlations show up in the cyclical indicator of all countries, and consistently with the expectations appear to be equally shared by both Euro area and non-Euro area countries (except for Denmark). However, a closer look at the data indicates that the beginning of most of these changes predate the event of interest. The common and the national indicators are, in fact, highly in phase since the mid-1980s; the general increase in the synchronicity of cyclical fluctuations in Europe begins roughly at the same time and appears to be the result of a larger commonality in the fluctuations rather than a change in the timing of the phases. Hence, the bulk of this preliminary evidence suggests that the institutional changes we considered followed rather than preceded the changes in the real fluctuations we have documented. While it seems that the creation of the ECB and the Euro changeover generated no major variations in business cycles, the evidence for the Maastricht treaty is more difficult to interpret.

Clearly, this informal analysis does not allow us to make strong causal statements one way or another. To acquire more evidence on the issue, we now turn to two forecasting exercises. In the first one, we try to see whether a structural change took place at the time when the political events occurred, by forecasting unconditionally the variables of interest using the information prior to these dates. If we can reasonably forecast the time path of these variables, then it is hard to say that the events of interest produced marked changes in the European business cycles. In the second one, we examine the transmission of two types of shocks over time. Again, if the events we are interested in matter for cyclical fluctuations, we should see significant changes in the shape, sign and magnitude of the
responses over time.

4.2 Forecasting unconditionally out of sample

We forecast the 5 variables of each country using the information available prior to the Maastricht treaty, the creation of the ECB and the Euro changeover. In particular, given the information available at 1993Q4, 1998Q3 and 2002Q1, we compute out-of-sample predictive distributions up to 5 years ahead. Notice that since future parameter uncertainty is averaged out when predictive distributions are constructed, the bands we present reflect data uncertainty, conditional on the exogenous variables in the model taking the values actually realized in the forecasting sample. We then check whether the actual path of the variable of interest falls within the tunnel constructed using the 90 percent predictive bands. If it does, at most of the horizons, no once-and-for-all change exist at these dates. If it does not, in addition to the slow moving time variations we have documented in the previous section, a break can be identified in correspondence with the three events of interest. Given that results do not depend on the variable we choose to forecast, we simply report results for GDP growth.

Figure 4 suggests that a clean and once-and-for-all structural break following the events is absent. In other words, our model with its slow moving time varying structure appears to adequately capture the dynamics of real variables in the ten countries around the political events we consider. Our predictive bands for GDP growth have the right direction and do not diverge from the actual values in most countries, at most horizons and for all three selected dates. In addition, the forecasting performance of the model is roughly unchanged over time. If anything, the performance seems to improve for Denmark, Sweden and Finland after 1998Q3, probably as a result of the convergence process that took place after the entry in the EU of the latter two in 1995, and for the UK - in this case the improvement is roughly uniform over the various subsamples. For Euro area countries, little measurable differences appear after Maastricht or the creation of the ECB, while after the Euro changeover the forecast accuracy of some country indicators worsens. This is particularly evident for Germany but this has more to do with the fact that the German
indicator increased its volatility in the last part of the sample rather than specific effects due to the Euro changeover.

These results are not sensitive to the dates we choose. For example, if we anticipate the dates at which we forecast by up to 4 quarters, no major changes are visible.

We would like to emphasize that the information contained in the figure nicely complements the one in table 1. Table 1 presents information constructed in-sample and on average over periods. Figure 4 instead presents case study exercises, where the out-of-sample predictive ability of the model is measured at particular dates. Hence, it is entirely possible that in-sample average variations coexist with unchanged unconditional forecasting ability of the model at particular dates, especially once it is taken into account that our model has time varying coefficients.

4.3 The Transmission of shocks

Unconditional forecasting exercises are a useful benchmark to detect breaks. However, by their nature they report “average” information and will not be particularly informative, for example, about the variations in the transmission mechanisms in response to particular types of shocks. In fact, an unchanged unconditional forecasting performance could be consistent with varying conditional forecasting performance, as long as the structural changes in the transmission approximately average out across shocks.

To gather more information about the issues of interest, we examine the transmission of a temporary German shock to the other countries, where by this we mean a shock which simultaneously increases all five German variables, as well as the transmission of an external shock - a temporary increase in the US nominal interest rate. These two shocks are chosen among the many potential options we have because, in addition to shedding light on the question of interest, they provide us with useful information about the nature of the intra-European and transatlantic transmission of disturbances, the magnitude of the synchronization and the qualitative nature of the heterogeneities present among European countries.

Once again, given the large number of variables in the system, we need to select which
responses to report and at which date. Figure 5 presents the responses of output growth to a German shock using the information available at 1993Q4, 1998Q3, 2002Q1 and at the end of sample date 2007Q3; Figure 6 the responses of output growth to a US interest rate shock using the same information.

Figure 5 shows that there are significant changes in the transmission of German shocks over time. With information up to 1993Q4, spillovers are small, quite idiosyncratic and for many countries statistically insignificant. This is true even for major players like the UK or Italy. We conjecture that the shock due to the German reunification may be responsible for this surprising result. Overall, cross country interdependencies within European countries appear to be reasonably small up to that date. With information up to 1998Q3, cross country responses appear to be much more synchronized, they are definitively larger, more persistent and statistically significant up to five quarters in Germany, Italy, France, Belgium. For Netherlands and Spain transmission takes one period but the effect is significant up to 5 quarters. The behavior of non-Euro countries relative to the earlier sample is mixed. Nevertheless, German shocks appear to still have little effect after the creation of the ECB. Domestic monetary and exchange rates policies designed to protect the economies against external shocks, could be the reason for this outcome. The responses after the creation of the ECB and the Euro changeover are very similar qualitatively. Quantitatively, one can notice that the spillover effects of German shocks is reduced in Euro countries. As we have noticed in previous subsections, national idiosyncrasies seem to matter more in the last 5 years of the sample. Nevertheless, responses computed with the information available in 2007Q3 reveal that these have little or no effect on the qualitative features transmission of German shocks.

Regarding the transmission of external shocks to Europe, Figure 6 shows that an increase in US interest rates has roughly the same qualitative effects at all dates we examine. After an increase in US interest rates the dollar appreciates. This increases the price-competitiveness of European countries and hence rises their GDP growth after a few quarters. The responses are typically hump-shaped and long lasting. Quantitatively, the magnitude of the responses changes over time. With information up to 1993Q4, responses
are large in Italy, Belgium and Spain, but also quite large in the non-Euro countries. With
information up to 1998Q3, all countries responses are significantly reduced. After the Euro
changeover, the magnitude of the responses is further reduced, while no more changes are
visible comparing responses by the Euro changeover and at the end of the sample. Notice
that the reduction in the magnitude of the responses are quite visible in inflation targeting
non-Euro countries; for Euro area countries they are much more contained.

Overall, the evidence of this subsection seems to indicate that an important conver-
gence process has taken place among all European countries. Furthermore, it appears that
the transmission of shocks became more similar for countries that now belong to the Euro
area. Once again we find it hard to relate the beginning of this convergence process with
the three events of interest. Similarly, its speed was not significantly changed by the three
events we consider. The analysis of this subsection has brought to light two interesting
facts which may have important policy implications. First, the elimination of the idio-
syncrasies in response to a German shock is not a monotonic process. Second, external
shocks have smaller effects on European economies in the latter part of the sample but it
is non-Euro countries that appear to be more insulated with respect to these shocks.

5 Conclusions

This paper attempts to shed some light on the effect of political economy events on the
dynamics of business cycles by focusing on the recent European experience and the three
important institutional changes occurred related to the European Monetary Union: the
implementation of the Maastricht treaty in 1993, the creation of the ECB in 1998 and the
Euro changeover in 2002. 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 these events. Second, we have attempted to assess whether a clean structural
break took place in the European economy when they occurred. Third, we tried to measure
whether the three 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 estimate a panel VAR model of the same type employed by Canova and Ciccarelli (2004) and Canova et al. (2007), using data for five variables from ten European countries, including the main EMU and non-EMU ones, for the sample 1970Q1-2007Q3. We report how areawide and national indicators have evolved over time and some reduced form statistics over different subsamples. We also conduct two types of forecasting exercises, an unconditional one, using the information at 1993Q4, 1998Q3, 2002Q1 and at the end of the sample, 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 European-wide cycle since early 1990s, common to both EMU and non-EMU countries (although it seems to emerge somewhat earlier among EMU countries), and a decreasing volatility, changing persistence and increasing correlation between country cycles and the common one. These changes, however, do not seem to line up exactly with the political events we consider.

Second, in an unconditional forecasting sense, the Maastricht treaty, the creation of the ECB and the Euro changeover did not make a huge difference. Using the information available before and after these 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 EMU countries with similar precision, while the non-EMU country cycles are better predicted the more updated the information. Hence, while the average in-sample evidence suggests some changes in the cyclical features of all European data over the last 15 years, the out-of-sample evidence at selected dates does not indicate any clearly identified structural break in European business cycles.

Third, in a conditional forecasting sense, we find that an important convergence process has taken place among all European countries: it appears that the transmission of shocks became more similar for countries that now belong to the Euro area, although we find it hard to relate the beginning of this convergence process or a change in its speed with the three events of interest. Once again conditioning on the information available before
and after the events we show that the transmission of German shocks was substantially smaller before the Maastricht treaty than before and after the ECB was created or the Euro changeover occurred. The larger relevance of the German reunification episode in the data up to 1993Q4 than in more updated information sets may be behind the pre-Maastricht results of this conditional forecast exercise. The responses to an extra-European shock also changed with time, with significant reductions observed in EMU countries around mid1990s while non-EMU ones are further reduced when using more updated information.

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. A convergence process has unmistakably taken place among european countries, which makes them respond in the last decade more to european shocks and less to out-of-Europe shocks, and this convergence process seems to have occurred somewhat faster between EMU countries. However, it seems hard to identify this convergence process with the particular moment of the implementation of the Maastricht treaty or with the ECB creation.
Appendices

A Estimation

A.1 Prior information

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

$$
p(\Omega^{-1}) = Wi(z_1, Q_1)
$$

$$
p(b_i) = IG\left(\frac{\omega_0}{2}, \frac{\delta_0}{2}\right)
$$

$$
p(\theta_0 | \mathcal{F}_{-1}) = N(\bar{\theta}_0, \bar{R}_0)
$$

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

$$
p(\theta_t | \mathcal{F}_{t-1}) = N(\bar{\theta}_{t-1}|t-1, \bar{R}_{t-1}|t-1 + B_t).
$$

We collect the hyperparameters of the prior in the vector

$$\mu = (z_1, \omega_0, \delta_0, \gamma_1, \gamma_2, vech(Q_1), \bar{\theta}_0, vech(\bar{R}_0))$$

where $vech(\cdot)$ 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:

| $z_1$ | $Q_1$ | $\omega_0$ | $\delta_0$ | $\gamma_1$ | $\gamma_2$ | $\bar{\theta}_0$ | $\bar{R}_0$ | $N \cdot G + 50$ | $\hat{Q}_1$ | $10^6$ | $1.0$ | $1.0$ | $0.0$ | $\bar{\theta}_0$ | $I_J$ |
|-------|-------|------------|------------|------------|------------|----------------|----------|----------------|--------|--------|--------|--------|--------|--------|--------|--------|
| 0.0   | 0.0   | 1.0        | 0.0        | 0.0        | 0.0        | 0.0           | 0.0      | 0.0            | 0.0    | 0.0    | 0.0    | 0.0    | 0.0    | 0.0    | 0.0    |

Here $\hat{Q}_1$ 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 1975-1990, and $J$ is the dimension of $\theta_t$. The values of the remaining hyperparameters have been chosen using previous experience. Priors are all informative and quite loose. The prior time variation is tightly small, and comparable to the usual Litterman assumption.
A.2 Posterior distributions

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

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

where $Y^T = (Y_1, ..., Y_T)$ denotes the data.

Using Bayes rule, $p(\phi | Y^T) = \frac{p(\phi) L(Y^T | \phi)}{p(Y^T)} \propto p(\phi) L(Y^T | \phi)$. Given $p(\phi | Y^T)$, the posterior distribution for the components of $\phi$, $p(\Omega | Y^T)$, $p(b_i | Y^T)$, and $p(\{\theta_t\}_T | Y^T)$, can be obtained by integrating out nuisance parameters from $p(\phi | Y^T)$. 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 $p(\phi | Y^T)$ 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 $\phi$.

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

$$\theta_t \mid Y^T, \phi_{-\theta_t} \sim N(\tilde{\theta}_t^T, \tilde{R}_t^T) \quad t \leq T,$$

$$\Omega^{-1} \mid Y^T, \phi_{-\Omega} \sim Wi \left( z_1 + T, \sum_t (Y_t - W_t \Xi \theta_t) (Y_t - W_t \Xi \theta_t)' + Q^{-1}_1 \right)^{-1}$$

$$b_i \mid Y^T, \phi_{-b_i} \sim IG \left( \frac{\bar{\omega}^i}{2}, \sum_t (\theta_t^i - \bar{\theta}^i_{t-1})' (\theta_t^i - \bar{\theta}^i_{t-1}) + \delta_0 \right)$$

where $\tilde{\theta}_t^T$ and $\tilde{R}_t^T$ 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 $\bar{\omega}^1 = T + \bar{\omega}_0$, $\bar{\omega}^2 = T g + \bar{\omega}_0$ and $\bar{\omega}^3 = T N + \bar{\omega}_0$.

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 150000 draws: 3000 blocks of 50 draws were made and the last draw for each block is retained after the discarding the first 1000. This means that a total of 2000 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).
References


Table 1. Basic statistics of the cycle indicators.

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S.D. is the unconditional standard deviation, AR(1) the first autoregressive coefficient, Corr(lag) the largest correlation with the common cyclical indicator and the lag at which it occurs, Coin is how coincident are the turning points of our indicators with the turning points reported by ECRI, P-T and T-P are the average length in quarters of recessions and expansions, respectively.
| PERIOD | Peak or Trough | COMMON | DE | DE | FR | FR | IT | IT | ES | ES | BE | NL | FI | UK | UK | DK | SE | SE |
|--------|----------------|--------|----|----|----|----|----|----|----|----|----|----|----|----|----|----|----|----|----|
| 71-73  | P              | T      | Q4 | Q1 | Q1 | Q3 | Q3 | Q3 | Q2 | Q2 | Q2 | Q1 | Q4 | Q3 | Q1 | Q4 | Q1 | Q4 | Q3 |
| 73-75  | P              | T      | Q3 | Q4 | Q4 | Q4 | Q5 | Q3 | Q4 | Q4 | Q5 | Q3 | Q4 | Q5 | Q4 | Q4 | Q5 | Q4 | Q5 |
| 75-77  | P              | T      | Q4 | Q4 | Q3 | Q5 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 77-79  | P              | T      | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 79-81  | P              | T      | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 81-83  | P              | T      | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 83-85  | P              | T      | Q4 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 85-87  | P              | T      | Q4 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 87-89  | P              | T      | Q4 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 89-91  | P              | T      | Q4 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 91-93  | P              | T      | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 93-96  | P              | T      | Q4 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 96-98  | P              | T      | Q4 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 98-00  | P              | T      | Q4 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 00-02  | P              | T      | Q4 | Q4 | Q3 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 02-04  | P              | T      | Q4 | Q3 | Q3 | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 | Q3 | Q4 |
| 04-06  | P              | T      | Q4 | Q3 | Q3 | Q4 | Q4 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
| 06-07  | P              | T      | Q4 | Q3 | Q3 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 | Q4 |
|        | Avg. duration, Q | P      | 6.3 | 7.5 | 6.1 | 5.6 | 7.6 | 5.6 | 5.4 | 6.2 | 5.9 | 4.7 | 5 | 8.5 |
|        |                 | T      | 5.7 | 9.4 | 4.7 | 5.4 | 9 | 5.8 | 5.7 | 7.4 | 6.4 | 6 | 6.6 |
|        | TP as in ECRI (Coin ±1Q) |       | 15/22=59.1% | 8/19=42.1% | 12/50=40% | 14/21=66.7% | 8/22=36.4% | 9/23=39.1% |
|        |                 |       | 14/22=60.8% | 12/19=63.1% | 15/50=50% | 15/21=71.4% | 15/22=68.2% | 13/23=56.5% |
Figure 1. Common indicators and GDP growth
Figure 2. Common cyclical indicator
Figure 3. Country cyclical indicators
Figure 4. Forecast of GDP growth
Figure 5. GDP growth responses to a German real shock
Figure 6. GDP growth responses to a US interest rate shock