

Similarities and Convergence in G-7 Cycles*

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Abstract

This paper examines the properties of G-7 cycles using a multicountry Bayesian panel VAR model with time variations, unit specific dynamics and cross country interdependences. We demonstrate the presence of a significant world cycle and show that country specific indicators play a much smaller role. We detect differences across business cycle phases but, apart from an increase in synchronicity in the late 1990s, find little evidence of major structural changes. We also find no evidence of the existence of an Euro area specific cycle or of its emergence in the 1990s.

Key Words: Business cycle, G7, indicators, Panel Data, Bayesian methods
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1 Introduction

There is abundant evidence that economic activity in developed countries share a number of characteristics. For example, the real business cycle literature has demonstrated that macroeconomic fluctuations across industrialized countries are closely linked (see e.g. Backus, Kehoe and Kydland (1995), Baxter (1995), Canova and Marrinan (1998)) while more structured time series analyses have shown that a large portion of regional and country specific fluctuations are common (see e.g. Gregory et al.(1997), Del Negro (2000), Lumsdaine and Prasad (2003) and Kose et al. (2003)) and that a number of linear and non-linear business cycle features are similar (Harding and Pagan (2002)). Although all existing evidence suggests the presence of a common source of fluctuations in developed countries, results are typically derived using restrictive or conventional assumptions about the nature of the dynamic relationships. In fact, except for Del Negro (2000) or Kose et al. (2003), the issue of whether commonalities are present or not is examined within an empirical framework which does not explicitly allow for cross-country and cross-variables interdependencies.

Recently, the stability of the cross-country business cycle relationships has come under scrutiny. For example, Helbling and Bayoumi (2003) suggest that the increase in cyclical synchronization observed after 2000 in the advanced economies is the result of common shocks. This apparently represents a major shift relative to the 1980s where the increased similarities in macroeconomic fluctuations were the result of improved trade relationships (see e.g. Canova and Dellas (1992), Rose and Frankel (2000) for a quantification of the magnitude of trade links). It also represents a somewhat different propagation mechanism from the one used to explain the transmission of US shocks to Asia (see Mackowiak (2003)) or Latin America (Canova (2003)). On the other hand, Stock and Watson (2003) document changes in the volatilities of G-7 business cycles in the 1990s and indicate that such changes may have altered the correlation among international macroeconomic variables.

A third issue which has recently attracted attention is the behavior of European cycles in the period leading to the creation of the European Monetary Union, both within the area and in relationship with other developed countries. Casual observation indicates that European cycles have displayed more synchronized movements in the 1990s and, possibly, a larger transatlantic conformity. All in all, commentators suggest that national and/or European specific effects

may be slowly but surely vanishing.

Knowing whether fluctuations in the industrialized world are similar, understanding their sources and characterizing their time variations is important for both academics and policymakers. From an academic point of view one is interested in knowing whether business cycle links are the result of cross countries interdependencies or common shocks. Hence, one may welcome studies empirically documenting similarities in economic fluctuations since the presence of a common cycle facilitates the study of the relationship between national and international policy decisions and the state of the world economy. Policymakers monitoring domestic or regional cycles are typically concerned with the effect of national idiosyncrasies and with the consequences that their actions have on the working of international markets. However, if variations in economic activity in countries with different institutions, economic structures or economic policies are driven by a common cause, markets more than policies are the key to understanding the comovements in economic activity. Moreover, national or regional policies designed to counteract world tendencies may be ineffective. Finally, structural time variations may undermine the usefulness of policies which may have been effective in the past.

This paper breaks ground in the area by explicitly addressing two interrelated questions. First, we would like to know whether there has been any tendency for G-7 cycles to become more similar in the 1990s or if, on the contrary, they tend to be clustered along geographical, regional or other institutional characteristics. Second, we are curious as to whether there is any evidence that Euro cycles are different from those we observe in the rest of the G-7 or if they have become so in the recent past. In answering these questions we also provide some new evidence on the relative importance of world and country specific cycles, on their evolution over time and over business cycle phases.

To study these questions we employ a panel VAR model of the type developed in Canova and Ciccarelli (2002). Their approach is useful in our context for at least two reasons. First, the econometric methodology is designed for large scale dynamic models displaying unit specific dynamics and cross country lagged interdependencies and it is flexible enough to allow for time variations in the correlation structure of cyclical fluctuations across variables and countries. Second, the parsimonious parametrization they propose endogenously produces an index structure where indicators of world and national specific cycles are recursively constructed and dynamically span cross country interdependencies. Therefore the specification is particularly

suites to study the interrelationships and the structural changes present in G-7 cycles and to analyze what drives the common and the idiosyncratic components of G-7 fluctuations.

Our investigation confirms some of the existing evidence. For example, as in Kose, Otrok and Whiteman (2003) or Lumsdaine and Prasad (2003) we find evidence of a significant world business cycle, despite different empirical techniques and data sets. However, it also provides new and important insights in the phenomenon. For instance, our results indicate that the common (world) indicator accounts for about 30 percent of the fluctuations in sales, industrial production, output and employment of the seven most industrialized countries, that it captures the more persistent portions of G-7 fluctuations and that it has more information than simple average or principal component measures obtained using G-7 GDPs or IPs. On the other hand, country specific indicators are useful in explaining certain GDP and employment episodes across time, but fail to track cyclical movements in the four variables over the entire sample. Perhaps more interestingly, we find that both world and country specific fluctuations are much more synchronized in contractions than expansions. That is, the uncertainty surrounding estimates of both world and country specific indicators is an order of magnitude smaller in the former than in the latter. Expansions tend to have large idiosyncratic components, both across variables and countries, while declines in economic activity have common timing and similar dynamics, both within and across countries.

Regarding the questions of interest of this paper, we do not find evidence of structural breaks in country indicators in the 1990s. Hence, the often cited idea that national cycles are disappearing finds no support from our analysis. These indicators are as significant in explaining the differential growth rate of GDP across countries in the mid 1990s as they were in the mid 1980s and, if anything, slightly more important. We also find little support for the idea that Euro cycles are different from those of the rest of the world or that a Euro area cycle is emerging in the 1990s. This result should be contrasted with Lumsdaine and Prasad (2003) and Artis et al. (2003) who instead detected the presence of a EU cycle using IP data. Our analysis shows that the Euro signal is much weaker when one considers a broader set of variables and that regional causes have minor explanatory power for G-7 fluctuations throughout the sample.

These set of results taken together imply that movements in the world indicator have been the stable and consistent reason for the commonalities of the fluctuations in the G-7 economies over time and that structural breaks in both the pattern of transmission across

countries and in the sources of structural shocks are probably absent.

We attempt to characterize the informational content of the indicators we construct using simple correlation analysis. We document that our world indicator captures a variety of influences going from the magnitude of world trade, to the behavior of commodity prices, to the stance of monetary policy in the G-7 and to the spending power of consumers. However, oil shocks, US technology shocks, world financial or monetary factors and fiscal policy do not appear to be behind the movements in the world indicator. Because the indicator captures such a variety of sources of disturbances and endogenously allow their mix to change over time, it has been a very stable driver of international business cycles over the period. We also document that, country indicators capture primarily the differential stance of monetary policy but, in some cases, are also related to domestic money growth or US variables. Interestingly, their informational content is not related to the stance of local fiscal policy nor to the local spending capabilities, except for Canada.

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

2 The Panel VAR Model

The empirical model we consider has the form:

$$y_{it} = D_{it}(L)Y_{t-1} + c_{it} + e_{it} \quad (1)$$

where $i = 1, \dots, N$ refers to countries and $t = 1, \dots, T$ 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 , c_{it} is a $G \times 1$ vector of intercepts and e_{it} is a $G \times 1$ vector of random disturbances. We assume that there are p lags for the G endogenous variables.

Whenever $D_{it}(L)$ is not block diagonal, the model displays cross-unit lagged interdependencies. Lagged cross-country interdependencies add considerable realism to the specification but it is costly: the number of parameters is greatly increased (we have now $k = NGp + 1$ parameters in each equation). (1) displays two other important features. First, the coefficients are allowed to vary over time. Second, the dynamic relationships are allowed to be unit specific. All three ingredients are crucial when one wants to study similarities, time variations in the

transmission of business cycles across countries.

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

$$Y_t = X_t \delta_t + E_t \quad E_t \sim N(0, \Omega) \quad (2)$$

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

Since δ_t varies with cross-sectional units in different time periods, it is impossible to estimate it using classical methods. Two shortcuts are typically employed: it is assumed that the coefficient vector does not depend on the unit, apart from a time invariant fixed effect, or that there are no interdependencies across units (see e.g. Holtz-Eakin et al. (1988) or Binder et al. (2001)). Neither of these assumptions is appealing in our study. Instead, we assume that δ_t can be factored as:

$$\delta_t = \Xi_1 \lambda_t + \Xi_2 \alpha_t + \Xi_3 \rho_t + u_t \quad (3)$$

where Ξ_1, Ξ_2, Ξ_3 are matrices of ones and zeros of dimensions $NGk \times N_1 \ll N$, $NGk \times N$, $NGk \times G$, respectively and $\lambda_t, \alpha_t, \rho_t$ are assumed to be mutually orthogonal. λ_t captures movements in the coefficient vector which are common across units and variables - for most of this paper λ_t will be a scalar; in the later part it will be a 2×1 vector. α_t captures country specific movements in the coefficient vector (roughly speaking, the fixed effects). Its dimension is equal to the number of countries in the panel, N . ρ_t captures movements in the coefficient vector which are variable specific and its dimension is therefore equal to the number of variables in each country, G . Finally, u_t is an error term capturing all the unmodelled features of the coefficient vector (which may have to do with lag specific, time specific or other effects).

A simple example may clarify the structure of the Ξ matrices. Consider a model with two countries, two variables and one lag per variable. Let $y_t = [y_{11t}, y_{12t}, y_{21t}, y_{22t}]$. Then

$$\Xi = [1, 1, 1, 1]', \Xi_2 = \begin{bmatrix} 1 & 0 \\ 1 & 0 \\ 0 & 1 \\ 0 & 1 \end{bmatrix}, \Xi_3 = \begin{bmatrix} 1 & 0 \\ 0 & 1 \\ 1 & 0 \\ 0 & 1 \end{bmatrix}.$$

Factoring δ_t as in (3) is advantageous in many respects. Computationally, it reduces the problem of estimating NGk coefficients into the one of estimating $1+N+G$ or $2+N+G$ factors. Therefore, even when the number of interdependent cross sections of countries is large, noise

is averaged out and reliable estimates of the features of interest can be obtained. Practically, (3) transforms an overparametrized panel VAR model into a parsimonious SUR model where the regressors are appropriate “averages”. In fact, substituting (3) into (2) we have

$$Y_t = \mathcal{W}_t \lambda_t + \mathcal{A}_t \alpha_t + \mathcal{M}_t \rho_t + v_t \quad (4)$$

where $\mathcal{W}_t = X_t \Xi_1$, $\mathcal{A}_t = X_t \Xi_2$, $\mathcal{M}_t = X_t \Xi_3$ capture respectively, common, country specific and variable specific information present in the VAR and $v_t = E_t + X_t u_t$. As the notation makes it clear, the regressors in (4) are combinations of lags of the right hand side variables, while $\lambda_t, \alpha_t, \rho_t$ play the role of loadings. Using appropriate averages as regressors is common in factor model literature (see e.g. Stock and Watson (1989) or Forni and Reichlin (1998)) or the signal extraction literature (see e.g. Sargent (1989)). However, there are four main differences between (4) and these models: the indices here are linear combinations of right hand side variables of the panel VAR; they are therefore observable; they dynamically span lagged interdependencies across units and variables; and their loadings are time varying.

From an economic point of view, the decomposition in (4) is convenient since it naturally allows us to assess the relative importance of world and country specific influences for fluctuations in Y_t . In fact, $WLI_t = \mathcal{W}_t \lambda_t$ plays the role of world indicator, while $CLI_t = \mathcal{A}_t \alpha_t$ plays the role of a vector of country specific indicators. Note that both indicators are leading Y_t as they reflect information contained in the predetermined variables of the VAR and can be constructed recursively. Finally, note that while we treat (3) as a part of the prior, it can be also thought as a part of the model specification. If little prior information on $\lambda_t, \alpha_t, \rho_t$ is employed, then the small sample distributions of WLI and CLI we obtained are the same as those obtained with classical methods on a model composed of (2) and (3).

To complete the specification we need to describe how $\lambda_t, \alpha_t, \rho_t$ evolve over time and the features of their distribution at time zero.

2.1 The structure of time variations

Write (3) compactly as:

$$\delta_t = \Xi \theta_t + u_t \quad u_t \sim N(0, \Sigma \otimes V) \quad (5)$$

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

$$\theta_t = \theta_{t-1} + \eta_t \quad \eta_t \sim N(0, B_t) \quad (6)$$

We assume that $\Sigma = \Omega$ and $V = \sigma^2 I_k$, where σ^2 is known; that $B_t = \gamma_1 * B_{t-1} + \gamma_2 * \bar{B}$ where γ_1, γ_2 are known; that $\bar{B} = \text{diag}(\bar{B}_1, \bar{B}_2, \bar{B}_3)$, and that E_t, u_t and η_t are mutually independent. We describe our choices for $(\sigma^2, \gamma_1, \gamma_2)$ in appendix A.

In (6) the factors evolve over time as random walks. Alternative specifications allowing for more complex dynamics or exchangeability across units are possible (see e.g. Canova and Ciccarelli (2002)). We stick to this simple setup since experimentation with more complicated structures did not produce qualitatively important changes. The spherical assumption on V reflects the fact that the factors are measured in common units, while setting $\Omega = \Sigma$ is standard in the literature (see e.g. Kadiyala and Karlsson (1997)). The variance of the innovations in θ_t is allowed to be time varying to account for heteroskedasticity 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 variance ($\gamma_1 = 0$ and $\gamma_2 = 1$) are special cases of the assumed process. The block diagonality of \bar{B} is needed to guarantee the orthogonality across factors, which is preserved a-posteriori, and hence their identifiability. Finally, independence among the errors is standard.

To summarize, the model we will use in our exercises has the hierarchical structure:

$$\begin{aligned} Y_t &= X_t \Xi \theta_t + v_t \\ \theta_t &= \theta_{t-1} + \eta_t \end{aligned} \tag{7}$$

where $v_t \sim (0, \sigma_t^2 \Omega = (1 + \sigma^2 \mathbf{X}_t' \mathbf{X}_t) \Omega)$. To compute posterior distributions we need prior densities for $(\Omega, \sigma^2, \bar{B})$. Because we want to minimize the impact of our choices on the posterior distribution of the indicators, we specify rather loose priors. The exact form of these priors, the numerical approach used to compute posterior distributions and the details of the computations are provided in appendix A.

2.2 Verifying the hypotheses of interest

To evaluate the posterior support for the main issues of interest, i.e. whether there has been any tendency for G-7 cycles to become more similar in the 1990s and whether cycles in the Euro area are different from those in the other G-7 countries or have become so in the 1990s, we employ two types of evidence. First, we examine the behavior of the posterior distribution of WLI_t and of CLI_t over time. Second, we compare the out-of-sample forecasting performances of restricted and unrestricted specifications.

Since time series plots of the posterior distribution of the indicators are difficult to read we summarize their information by reporting the median of the distribution and a 68 percent posterior central band - the latter corresponding to one standard deviation around the mean in classical frameworks. Visual evidence in favor or against the first and the third hypotheses will be obtained by examining whether the 68 percent posterior band of the relevant indicator includes zero or not at most or all dates. Evidence in favor of time variations can be obtained when the 68 percent posterior bands includes zero in some time periods but not in others. We will complement this visual information by computing the percentage of variations in the endogenous variables due to the factors and by examining the importance of the two indicators in specific historical episodes.

Since the model has a natural leading indicator structure, we will also examine the posterior support for the three hypotheses by comparing the predictive ability of unrestricted and restricted specifications. Predictive distributions can be compared using Bayes factors. The predictive Bayes factor is $\frac{f(Y_{t+\tau}|y_t, M_i)}{f(Y_{t+\tau}|y_t, M_j)}$ where $f(Y_{t+\tau}|y_t, M_i) = \int L(y_{t+\tau}|\theta_t, y_t, M_i)p(\theta_t|M_i, y_t)d\theta_t$ is the predictive density for model M_i and $p(\theta_t|M_i, y_t)$ is the posterior on the parameters of model M_i . Model i is preferred if this statistics exceeds 1. Thus, for example, in examining the similarities of cycles across countries, M_j corresponds to a model with three observable indices and M_i to a model which leaves the country specific component CLI_t out; while in examining the differences of EU cycles from non-EU cycles, M_j is model where λ_t is a scalar so that WLI_t has one component and M_i is a model where λ_t is a 2×1 vector so that WLI_t has two distinct components, one for EU countries and one for non-EU ones.

3 The Results

3.1 The data and some forecasting features of the model

For each country we use quarterly growth rates of seasonally adjusted real GDP, employment, sales and industrial production as our basic variables. We choose this four series since they are among the variables used by the NBER when deciding the state of the business cycle in the US, by the CEPR when deciding the state of the business cycle in the Euro area and by the Economic Cycle Research Institute (ECRI) when measuring business cycles around the world¹.

¹The NBER uses real personal income instead of GDP, but such series are not available on a quarterly basis for all of the G-7 countries. The CEPR also looks at investment and at some other national variables like the unemployment rate.

The sample maximizes the amount of common data and covers the period 1979:1-2002:4.

Real GDP data is measured in constant 1995 prices, except for Canada (the base year is 1997). The Japanese series starts only in 1980 and has been extended backwards using the real GDP series measured in constant 1990 prices. Similarly, the Canadian real GDP was extended backwards from 1981 using real GDP data with base-year 1992. The source of all data is the Quarterly National Accounts of OECD, except for Germany whose real GDP comes from the Bundesbank database. We prefer this series since it explicitly takes into account the effects of German unification. Employment is measured by the civilian employment index, with base year 1995, and is from the OECD Main Economic Indicators. Sales are measured by the retail sales volume index, with base year 1995, and come from OECD Main Economic Indicators, except for the US, where the source is the Department of Commerce. The industrial production index also has 1995 as base year and comes from OECD Main Economic Indicators. Whenever these series are provided in non-seasonal adjusted form, we seasonally adjusted them using the TRAMO-SEATS program. In the panel VAR we use these four series in growth rates. While information about their long run properties may be lost, our choice avoids distortions in the construction of the indicators due to the different size of various aggregates across countries. This is important since the procedure cannot distinguish if a 2% growth is generated in the countries with a large level (say, the US) or with a smaller one (say, Canada).

We have conducted a number of specification searches to decide the structure of the model. A summary of these exercises appears in Appendix B. Overall, we found that a model which displays lagged cross country interdependencies and time variations in the factors, where the decomposition in (3) includes three elements and V is nondegenerate is preferable to specifications which exclude these features. We would like to emphasize two important aspects of our specification search. First, a model with lagged cross-country interdependencies has much higher predictive ability than a model without interdependencies (predictive Bayes factor is about 2.0). This implies that static factor models of the type employed e.g. by Stock and Watson (1989) may miss important sources of transmission of fluctuations across countries. Second, time variations in the factors play some role but not a very large one.

To demonstrate the ability of the model to approximate the data we present three types of statistics. In table 1 we report the Theil-U statistics (the ratio of the one step ahead MSE computed using the median of the posterior of our model to the one step ahead MSE of a naive

no-change univariate model) for IP and employment growth of the 7 countries. We choose to confront the forecast of our model to those of a no-change model because more complicated specifications, e.g. a VAR (with or without a prior) fail to improve over the no-change model for these two variables. Clearly, the smaller the reported numbers, the larger is the forecasting improvement of our model. The statistics are reported for two types of prior assumptions, a non informative and an informative one. Results for the latter are slightly better, but we prefer the non informative specification since it avoids interference with the data and does not alter much the results. Subsequent discussion and figures are therefore based on the non informative set-up. The values of the hyperparameters used in the two cases are in appendix A.

Table 1: 1 Step ahead Theil-U statistics, 1990:1-2002:2							
	US	Japan	Germany	UK	France	Italy	Canada
Non Informative priors							
IP growth	0.98	0.92	0.77	0.93	0.70	0.80	0.70
Employment growth	0.67	0.60	0.57	0.78	0.87	0.91	0.93
Informative priors							
IP growth	0.91	0.78	0.79	1.03	0.61	0.70	0.62
Employment growth	0.46	0.48	0.56	0.52	0.70	0.84	0.69

Figure 1 reports the predicted recursive one-step ahead 68 percent central posterior bands (dotted lines) together with the actual growth rates (solid line) for UK GDP growth, US employment growth and Japan IP growth. We choose these three series because the performance of the model for them is close to the median outcome in the sense that they were neither the best nor the worst tracked by our specification.

In figure 2, we report the record of turning point probabilities for US, UK, Canada and Germany GDP in levels. Probabilities of turning points are generated calculating the percentage of times our model generates the pattern predicted by the following simple rule: there is an upturn at t if $gdp_{t-2} < gdp_{t-1} < gdp_t > gdp_{t+1} > gdp_{t+2}$ and there is a downturn if $gdp_{t-2} > gdp_{t-1} > gdp_t < gdp_{t+1} < gdp_{t+2}$. While this rule is extremely simple and does not make the provision for minimum length of a cycle or for consecutive turning point signals as, for example, the Bry and Boschan (1971) procedure, it suffices to give an idea of the ability of the model to replicate important non-linear functions of the data. For readability, we present probabilities of downturns with positive sign and probabilities of upturns with a negative sign.

Superimposed on the graph are the level of the series and the downturn (peak to trough) phases as reported by ECRI.

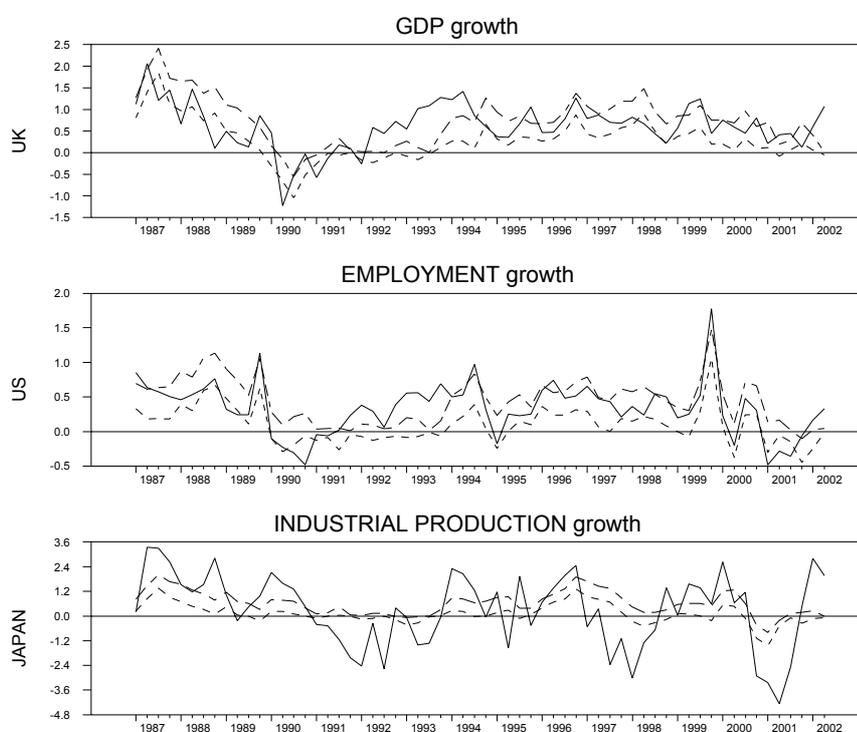


Figure 1. Predictive ability of the model

Table 1 shows that one-step ahead forecast for employment growth are slightly better than those for IP growth (which is considerably more volatile). However, for both series in all countries, our model is superior to the alternative. Gains are, on average, of the order of 20% for IP growth and exceed 25% for employment growth. We use a Diebold and Mariano (1995) test to check for the significance of the gain vis-a-vis a no-change univariate model: we confirm that the forecasts of the two models are different and that those of our model are better.

The forecasting ability of the model can also be graphically appreciated in figure 1. In general, forecasts are smooth and the model captures the direction of the movements in the three variables reasonably well. Moreover, there are episodes when it captures the timing and even the size of the changes actually observed.

While our model is designed to capture the features of cycles in the growth rates, it can also reproduce quite well turning points in the level of GDP of several of the G-7 countries.

Figure 2 shows that the model generates high probabilities of a turning point in correspondence of the actual ups and downs of the series. In fact, it does not miss any of the recession phases chosen by ECRI for the four countries. However, given the simplicity of the adopted rule, it may miss a particular turning point, like the most recent trough in the US GDP. Therefore, while improvements are possible it is comforting to see that even with this simple dating rule the model does not miss important recession signals present in the data.

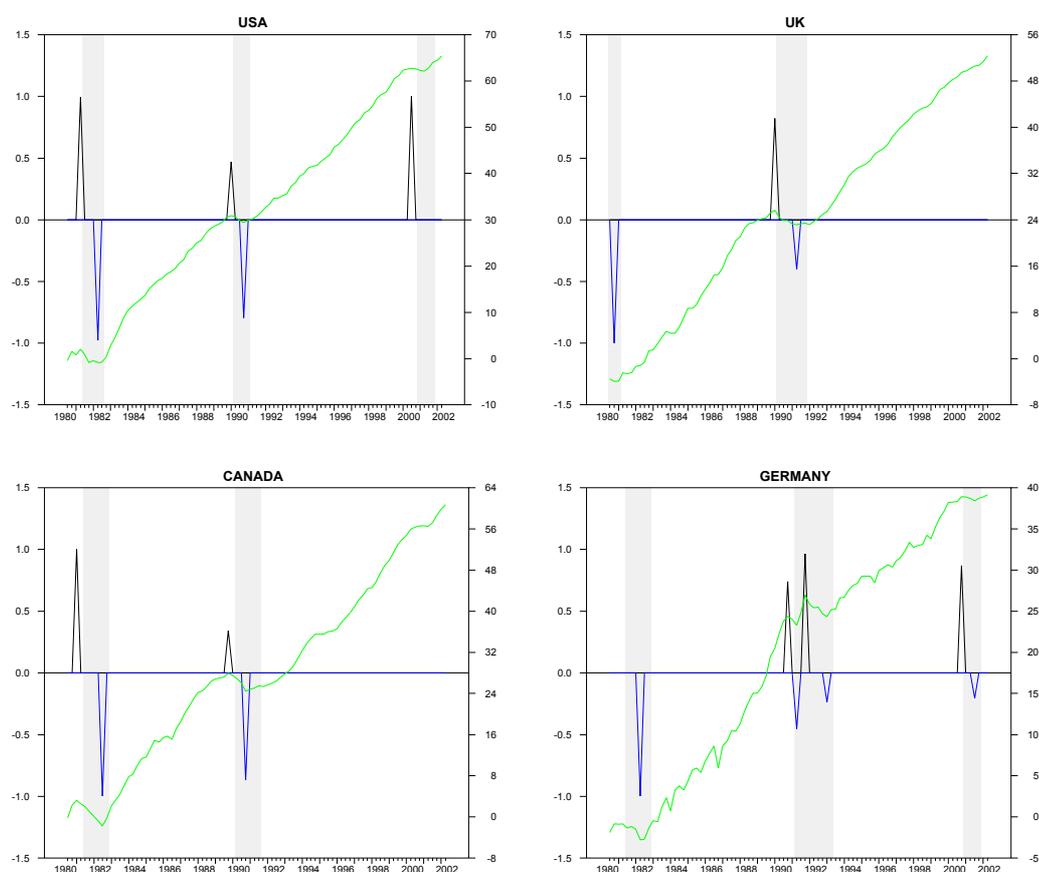


Figure 2. GDP levels, recession phases and Turning Point Probabilities

3.2 Are G-7 cycles (more) similar?

To answer this question we plot the median and the 68 percent posterior central band for the world indicator WLI_t and for the seven country specific indicators in figure 3, all in growth rates. The country indicators displayed include the information common to all four variables

in each country, i.e., the one shared with the rest of the variables in the model, WLI_t , plus the information specific to that particular country, CLI_{it} .

There are several features of figure 3 that are worth commenting upon. First, the world indicator captures major events occurred in the last 20 years: it displays the double dip experienced by growth rate of output of several countries at the beginning of the 1980s and a sustained growth thereafter; it captures both the US recession of the beginning of the 1990s and the European one a couple of years later; it shows a positive trend growth in the 1990s and drops below zero in the first quarter of 2001. According to this indicator, the 1982 recession is the deepest and, apparently, the longest of all while the (common) growth in the 1980s was slightly larger, on average, than the one experienced in 1990s (Japan's poor performance in the 1990s being partially responsible for this outcome). Second, posterior uncertainty (as measured by the size of the 68 percent band) depends on the state of the world economy. Bands are tighter at the beginning, in the middle and at the end of the sample and these happened to be periods when the world indicator was either negative or very close to zero. Hence, the timing and the size of comovements across variables and countries appears to be more similar in contraction than in expansions since posterior uncertainty in λ_t appears to be independent of the state of the economy.

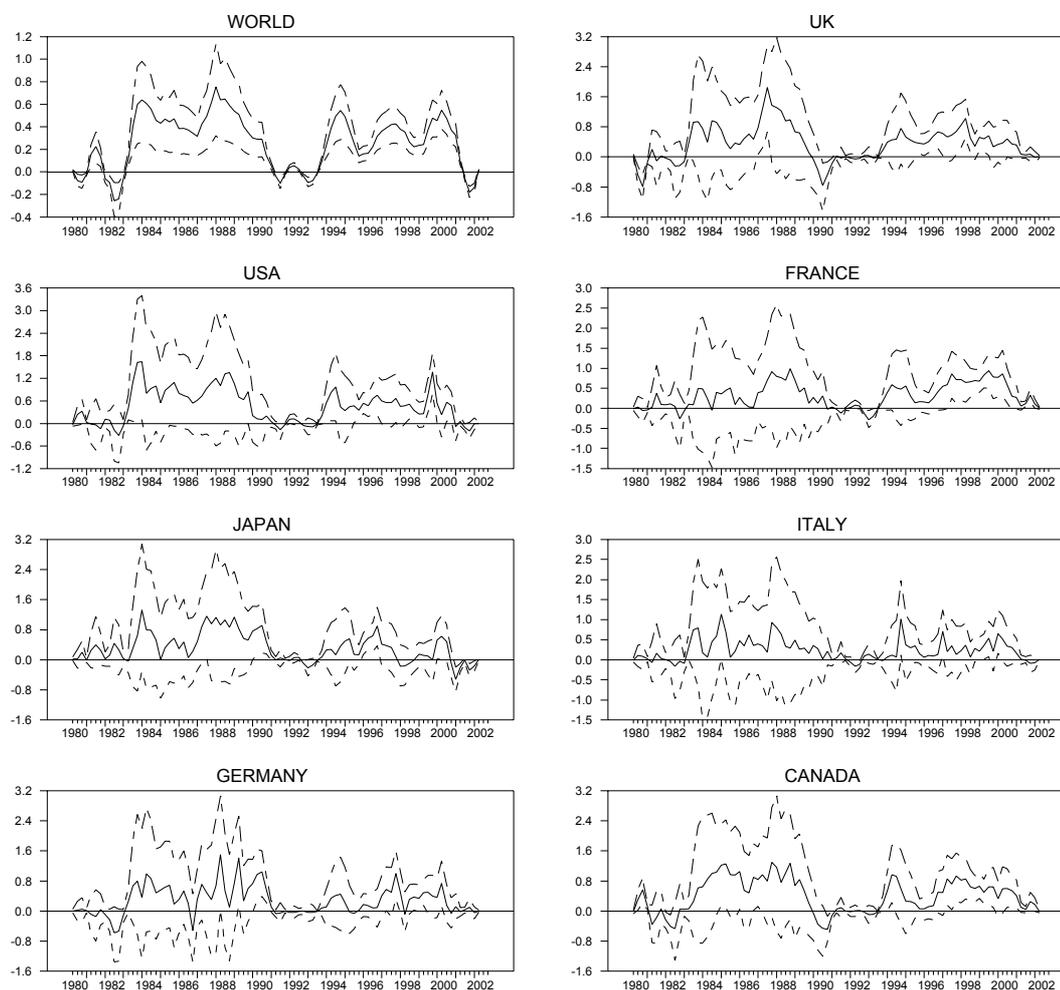


Figure 3. World and Country Indicators

The pattern of comovements observed after 2001 has drawn the attention of several researchers (see e.g. Doyle and Faust (2002), Perlsman (2003), Helbling and Bayoumi (2003)). In particular, the high synchronicity of the downturn in the industrialized world has been contrasted with the much slower transmission processes experienced early on and commentators have suggested that the sources of international business cycle must have changed. While our analysis confirms the importance of world influences in determining the extent of such a slowdown (the world indicator is entirely responsible for the similarities in the comovements of G-7 GDPs after that date), it also suggests that the contribution of world disturbances is not

unusual when compared with other recessionary episodes of the last 20 years. What appears to have drawn the attention of researchers is not so much a (permanent) structural change in the relationship but the switch in the uncertainty present in recession and expansion phases.

Third, at each t , the world indicator has little posterior probability mass in the symmetric region centered around zero. Therefore, there is posterior evidence that it significantly contributes to the comovements of the 28 variables we consider. A simple way to numerically measure this contribution is to check how much of the variance of each of the four series is explained by the posterior median of the world indicator. We find that, on average across countries, the median value of the world indicator explains about 33% of the fluctuations in the growth rate of GDP, about 56% of the fluctuations in employment growth, 17% of the fluctuations of IP growth and 15% of the fluctuations of sales growth. Similarly, we find that 30% of the fluctuations of the four variables within countries are explained by the indicator, with France (46%) and Japan (23%) being the two opposite extremes. The numbers for GDP are surprisingly similar to those reported by Kose et al (2003) for G-7 countries (36%), despite the fact that the sample and the frequency of the data they use is different. They are also remarkably stable. In fact, considering the 1980's and the 1990's separately we find that the world indicator explains 32 percent of GDP fluctuations in the first and 36 percent in the second sample. Hence, the changes experienced by the world economy over the last 20 years have not altered the importance of common influences in shaping international business cycle dynamics.

Fourth, the 68 percent posterior band of country indicators includes zero at most dates. This means that their explanatory power for G-7 fluctuations in the last 20 years is probably minor. To put this observation in another way, national variables display important fluctuations but the source does not appear to be distinctively national; rather, world wide influences are the reason behind this commonality. Also for country indicators, posterior uncertainty is significantly smaller in downturns than in upturns. Hence, asymmetries in business cycle phases are also present at national level, with synchronicity being stronger in downturns than upturns.

To explicitly evaluate the contribution of country indicators to cyclical fluctuations in the G-7 economies, we plot in figure 4 the growth rate of GDP (GDPG), the median value of WLI_t (WLI) and the median value of $WLI_t + CLI_t$ for each country (WCLI). Sizeable differences

between the two indicators will emerge when country specificities are of relevance. One can observe that in the US, Japan and Canada, country specific idiosyncrasies are important in explaining movements in the growth rate of domestic GDP for the period 1983-1990 and, to a smaller extent, in the late 1990s. Confirming conventional wisdom, national specific reasons appear to be behind the poor GDP growth performance of Japan in the late 1990s.



Figure 4. Actual GDP growth and posterior medians of the indicators

The country specific indicator adds little explanatory power to GDP growth movements in Germany and Italy: roughly speaking, it captures high frequency movements in GDP growth but little else. In France, it significantly deviates from the world indicator in the middle of the 1980s and especially in the late 1990s - in the earlier episode on the negative side, in the more recent one on the positive side. In the UK, country specific disturbances appear to be crucial in capturing the downward trend in GDP growth experienced over the 1986-1992 period and

have some power in explaining the above average growth rate experienced in the mid-1990s. Finally, the picture shows that the world indicator captures low frequency comovements in G-7 countries, while country specific indicators replicate more high frequency type fluctuations. In fact, the AR(1) coefficient of the posterior median of the world indicator is 0.91, while the AR(1) coefficient of the posterior median of the country indicators varies from 0.60 for Germany to 0.83 for Canada, with Italy and France close to the lower end. Very similar conclusions are obtained examining employment growth fluctuations across countries.

To formally examine whether G-7 cycles are similar, we compare the predictive ability of two models: one without country specific components (M_i) and one which includes them (M_j). The predictive Bayes factor is 0.90 suggesting that country indicators have some role in explaining the dynamics of the data. To understand the meaning of this number, note that one should have a prior odds in excess of 1.11 (attributing 0.53 or more a-priori probability on the model without country specific indicators) to have mild posterior evidence against the existence of country specific cycles. The predictive power of country specific indicators comes almost entirely from the mentioned above episodes. In fact, if we use only data for the 1990's, Bayes factor drops to 0.86, indicating a higher role for the country component in this sample. This evidence is corroborated when we compute the ratio of the share of GDP variability explained by country factors and the world factor. For the whole sample, the posterior 68 percent range for this ratio is [0.31, 0.65] while for the 1990's the range is moved to [0.45, 0.79].

Are national cycles disappearing? Figure 3 indicates that there are significant time variations in the posterior distributions of both the world and the country specific indicators. The most evident one concerns the varying level of uncertainty present in recessions and expansions. Another appears to be the decline in the uncertainty surrounding both types of indicators over time². Finally, country specific indicators have become slightly more synchronized and, marginally, more significant in the last decade - probably as a result of greater synchronization in the ups and downs of domestic variables.

While some of these observations have important policy implications - the decline in national uncertainty suggests that the information contained in GDP may suffice to characterize the state of the local economy - and others require further study to sort out alternative

²It is unclear if this drop is due to increase synchronization or simply better recursive estimates. Our conjecture is that a bit of both contribute to the outcome.

explanations - is the slight increase in domestic synchronization the result of policy actions, of domestic specialization, of increased similarities in the sectorial shocks hitting the economies or simply of reduced uncertainty connected with a specific growth episode? - the evidence denies the presence of significant breaks in the structure of comovements across the two parts of the sample. The 1980s are not different from the 1990s as far as sources of business cycles are concerned: national cycles played some role in both decades; their importance does not seem to be vanishing over time.

This result is not necessarily in contrast with any idea concerning the globalization of markets (see also Perri and Heathcote (2001)). Conventional wisdom suggests that more globalization should bring a stronger synchronicity in the business cycles of the industrialized world following an increase in trade. However, the possibility of easy trade may also make local production more specialized (along the lines of standard static or dynamic comparative advantage). Specialization may therefore bring less synchronicity in production. There is no reason to expect one of the two effects to dominate the other, and one can envision scenarios where increased globalization leaves the synchronicity of G-7 cycles unaffected.

3.3 Are European cycles different?

One reason for the increased significance of country indicators in the latter part of the sample could be the emergence of a regional element in the fluctuations. For example, if European cycles tend to deviate from say, Rest-of-the-world (ROW) cycles in the late 1990s, the presence of one world indicator may result in significant country specific indicator, because of omitted variable biases. This suspect is partially confirmed by the fact that France and the UK are the countries which display the most significant country specific indicators in the late 1990s and by the analyses of Lumsdaine and Prasad (2003), who find high correlations of EU industrial production indices with the EU factor they construct.

We therefore repeated our exercise allowing for 2×1 vector of world indices: one for Euro area countries (France, Italy and Germany) and one for the Anglo-Saxon countries. The rest of the model is unchanged. We decided to leave Japan out of the two groups since the synchronicity of Japan and the other G-6 cycles dramatically declines from the mid-1980s when other East Asian countries (Korea, Taiwan, China) became the major trading partners of Japan. We also experimented including Japan in a rest of the world (ROW) cycle and

switching the UK into the European group, despite evidence of some diversity of Euro area and UK cycles (see e.g. Harding and Pagan (2002) or Artis et al. (2003)). We settled on the above specification since Bayes factors always preferred it and, in both cases, by a large margin.

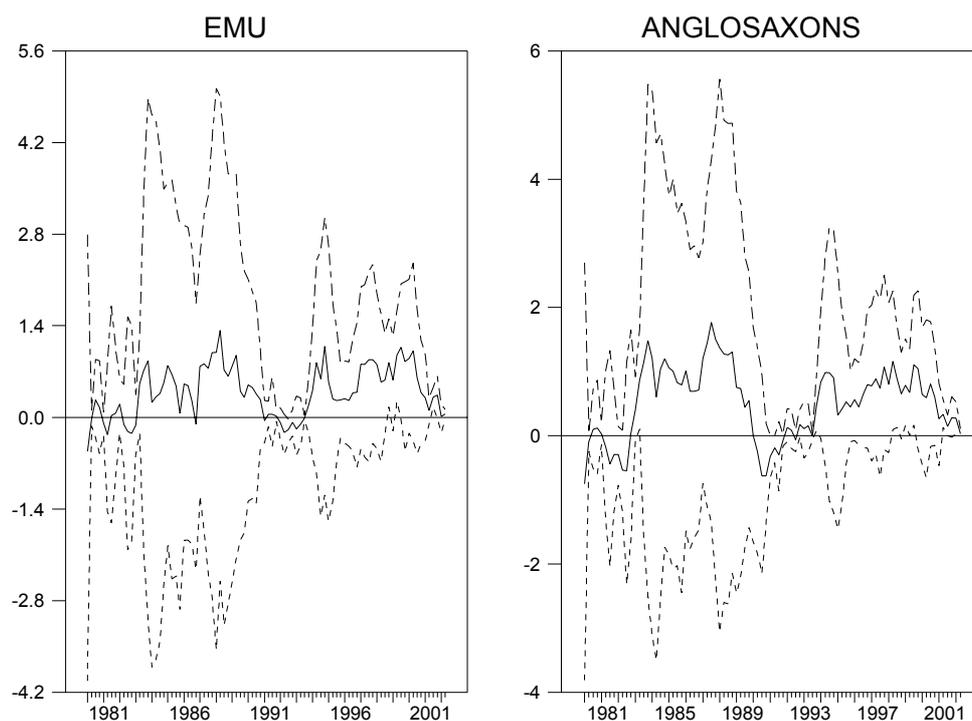


Figure 5. EMU and Anglo-Saxons Indicators

Figure 5 presents the time path of the median and 68 percent posterior central bands for the two regional indicators. Four features of the figure stand out. First, the posterior band of the Euro area indicator almost always includes zero. In fact, out of the 92 data points in the sample, the posterior 68% band excludes zero only once (at the beginning of the 1990), roughly one percent of the times. Hence, there is little posterior evidence of a significant Euro area indicator. Second, there also little posterior evidence that Euro area cycles differ from the Anglo-Saxon ones: the overlap of the two posterior distributions at each t is substantial. Third, the evolution of the median of the posterior of the two indicators is similar in both the 1980s and the 1990s. Therefore, there is no posterior support for the idea that Euro area cycles

became more important in the later part of the sample. Finally, also with this specification, posterior uncertainty is very much reduced during contraction phases. This result is important: the large posterior uncertainty present in the two indicators for most of the sample cannot be due to the lack of information in the two aggregates. However, it could be that enlarging the number of countries in the Euro area group, for example including Portugal or Ireland, will make the EMU indicators "more" different from the Anglo-Saxons ones.

We formally examine whether the inclusion of a Euro area indicator improves the performance of the model by constructing predictive Bayes factors. If European cycles are substantially different from the Anglo-Saxon ones, the predictive power of a model with two indices should be substantially superior to our original specification. It turns out that it is not: Bayes factor is 0.989 suggesting that there is only a very marginal improvement in the predictive power of the model with separate European and Anglo-Saxons indicators. Given the evidence in figure 5, this is not surprising: even though one extra degree of freedom is gained, the uncertainty in the estimates is so large that the predictive ability does not necessarily improve.

While the lack of a Euro area cycle may appear surprising, it is worth stressing that our result is not unique. For example, Kose et al.(2003) find no evidence of a second world factor, in general, and of a EU factor, in particular. Interestingly, the absence of this second factor is demonstrated using non-Bayesian techniques. How should one then interpret the evidence? Our favorite interpretation is the following. While Euro area aggregates display common fluctuations across variables and economies, their source is not distinctively European. Euro area and Anglo-Saxon fluctuations are similar in timing, size and amplitude because they are driven by the same source of disturbances. Hence, since regional causes play a minor role in explaining comovements in cyclical fluctuations in a large cross section of industrialized and less developed countries, Euro area policymakers should closely monitor the state of the international business cycle and de-emphasize Euro area (and national) cycles.

4 What drives our indicators?

So far the analysis has been primarily statistical. However, to go beyond the simple documentation of the time series properties of the world and country indicators and study how they relate to the more structural evidence presented e.g. in Perri and Heathcothe (2001), it is necessary to study the informational content of the indicators we constructed.

To start with, we examine how the world indicator relates to simple and easily computable measures of common statistical fluctuations such as arithmetic averages of GDP, IP, employment or sales growth in the G-7 or their principal components. If the behavior of the world indicator we construct can be reproduced with such simple measures, our more complicated setup can be clearly dismissed. We find that arithmetic averages share some informational content with our indicator but the overlap is far from perfect. In fact, the point estimate of the bivariate correlation between the posterior median of the world indicator and the simple average at each point in time of the GDPs, IPs, employments and sales growth rates are 0.58, 0.64, 0.67, 0.37 respectively.

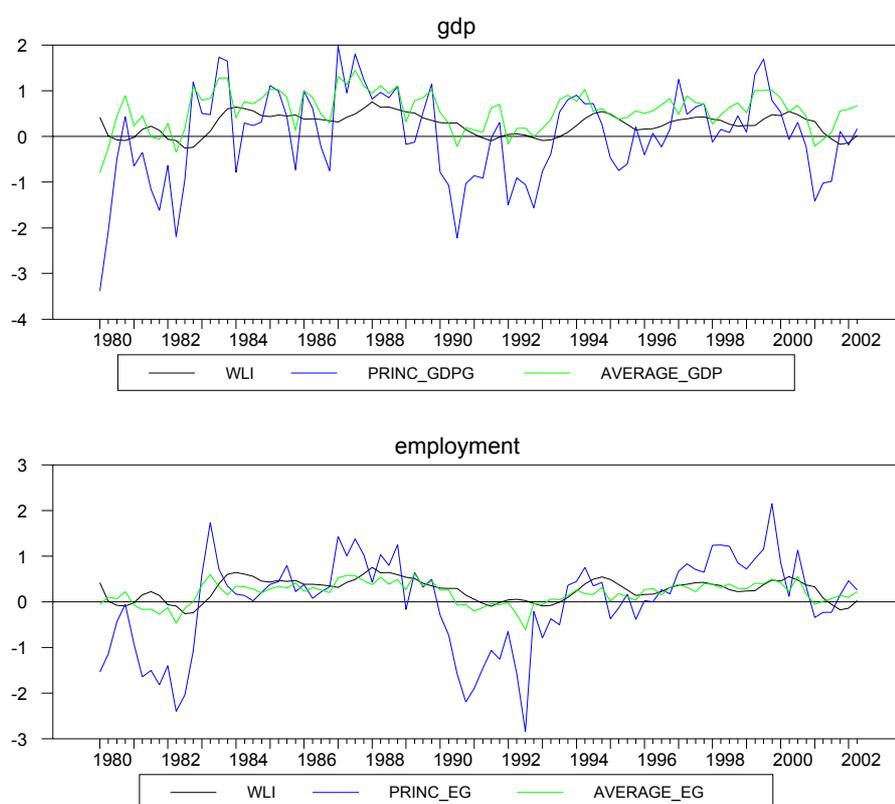


Figure 6. World Indicator, principal components and averages

Similarly, we find that our indicator is correlated with the principal component of GDPs, IPs, employment and sales growth although the association is less strong (point estimates 0.37, 0.35, 0.54, 0.01). Figure 6 graphically provides evidence of this association for GDP and

employment growth: the principal component is typically much more volatile than our indicator while simple arithmetic average measures occasionally miss important cyclical movements in the data. Our world indicator is the smoothest and mimics the local trend present in the data better than these standard measures.

To study the informational content of the indicators one could proceed as in the recent factor-VAR literature (see e.g. Bernanke et al. (2003)) and measure the contribution of structural shocks to their fluctuations. Since structural shocks are difficult to identify in our international context, we prefer to compute a number of simple correlations. Since these correlations are non-structural they are only suggestive of the possible sources of cyclical movements captured by our indicators. We have computed correlations of the world indicator with the growth rate of the US real personal non-agricultural income, the growth rate of world commodity price index, the NYSE stock return index, the growth rate of the world market crude petroleum price, the growth rate of the world goods trade, the quarterly US technology shock extracted by Gali, López-Salido and Vallés (2003), the average spread between US and other G-7 short term interest rates, both real and nominal, the average growth rates of real private consumption to GDP ratio, real effective exchange rates and M3, and average G-7 government deficit.

Is the world indicator a stand-in for oil shocks, technology shocks or other types of supply side disturbances? The answer is mixed. While we find that the correlation between the world indicator and commodity prices is significant³ although not very large (point estimate 0.27), the one of oil prices and US technology shocks is small and insignificant (point estimates 0.05 and -0.05). Similarly, the correlation with real exchange rates is low and insignificant. Therefore, while oil shocks may be an important source of national disturbances, and technology shocks in the US may explain an important portion of US fluctuations, it is necessary to go beyond these disturbances to explain and interpret existing world business cycles.

Does the world indicator capture fluctuations originating in financial or monetary markets or the stance of fiscal policy? Once again, the answer is mixed. While the correlation with US stock returns, average G-7 government deficit and average growth of M3 is small and insignificant (point estimates -0.08, 0.10 and 0.15 respectively), the world indicator appears to be significantly correlated with the average short term real and nominal interest rate prevailing

³Here and in the following "significance" means that the centered 95 percent band does not include zero.

in the G-7 (point estimates -0.28 and -0.57, respectively).

Besides average nominal interest rates and commodity prices, what else is behind the movements of the world indicator? It appears that the world trade (point estimate 0.50), the average consumption/output ratio (point estimate -0.46) and US personal income (point estimate 0.27) are the variables with the most significant correlations. Interestingly, as in Cochrane (1994) we find that the consumption/output ratio is almost as good as any other variables in explaining common movements in the G-7.

In sum, our world indicator captures a number of influences (trade, commodity prices, monetary policy, spending capacity) that analysts and academics have indicated to be important to understand the dynamics of business cycles in the industrialized world. One explanation for the remarkable stability in the explanatory power of this index may be to the fact that it robustly and flexibly captures different sources of fluctuations by allowing for time variations and adaptive changes in the weights.

There are another couple of other interesting facts which our investigation has discovered. First, there is a strong negative relationship between the average explanatory power of the world indicator for national cycles and volatility of GDP (see figure 6). Such a relationship is at times studied in growth literature to evaluate the desirability of stable growth. In our context, this patterns implies that synchronicity with the rest of the G-7 improves (worsens) as the volatility of domestic GDP fluctuations is reduced (increases). Second, the world indicator has larger explanatory power for employment than output (and larger correlation with the average employment growth than with average GDP growth). This result squares well with those derived by the international RBC literature where employment correlations larger than GDP correlations are interpreted as suggestive of the presence of an important world cycle.

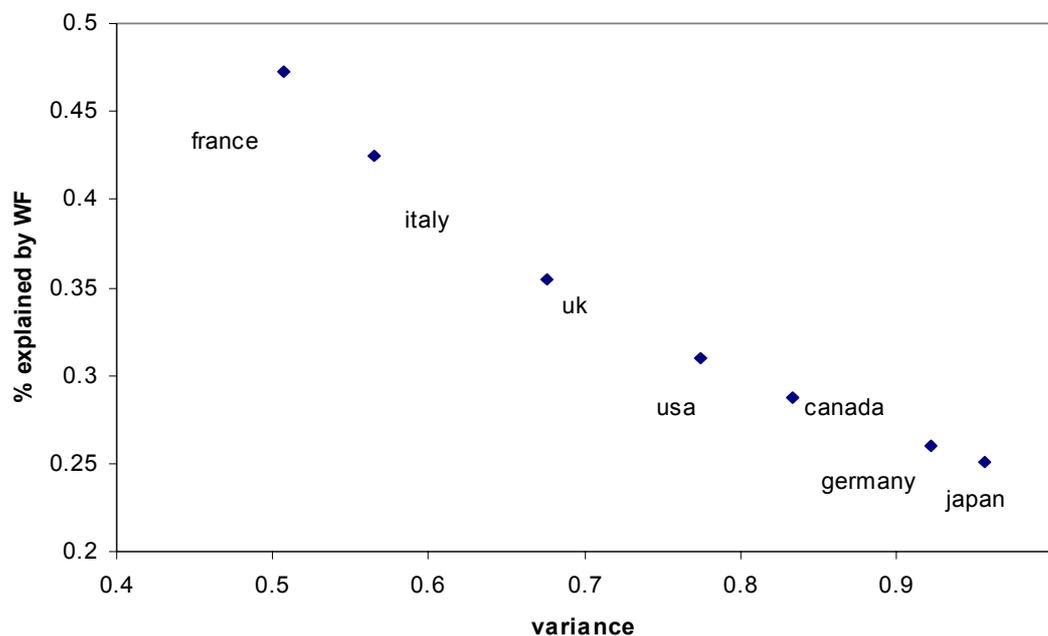


Figure 7. Synchronicity and GDP volatility

We have also attempted to identify the informational content of country specific indicators. We have correlated them with the individual series used in the model (the growth rate of real GDP, industrial production, employment and real retail sales) as well as with other domestic variables like the 3 month interbank nominal interest rate, the nominal yield on 10 year bonds, the real short term and long term interest rates, the growth rate of M3, the general government deficit, the consumption output ratio and the real exchange rate and with a few other international variables (the US personal income, the commodity price index, the US technology disturbances, US stock returns). We found that US personal income is significantly correlated with all the country specific indicators as are short term interest rates (either nominal or real) and money growth. Interestingly, the informational content of national indicators is not related to the stance of fiscal policy in any country nor to the local consumption to GDP ratio, except in Canada.

5 Conclusions and Directions for future work

This paper studies similarities and convergence of G-7 business cycles using a panel VAR model with cross country interdependencies and time variations. The framework of analysis

is advantageous in several respects. First, the structure is flexible and allows for multiple types of contemporaneous and lagged comovements and for time variations in the correlation structure of cyclical fluctuations across variables and countries. Second, the parsimonious parametrization we use endogenously produces an observable index structure where indicators of world and national cycles are recursively constructed and dynamically span cross country interdependencies. Third, the specification allows us to verify the posterior support for the hypotheses of interest and to analyze what drives the common and the idiosyncratic components of G-7 fluctuations.

We address two interrelated questions. First, we would like to know whether there has been any tendency for G-7 cycles to become more similar in the 1990s or if, on the contrary, they tend to be clustered along geographical, regional or other institutional characteristics. Second, we are curious as to whether there is any evidence that Euro cycles are different from those we observe in the rest of the G-7 or if they have become so in the recent past. In answering these questions we also provide some new evidence on the relative importance of world and country specific cycles, on their evolution over time and over business cycle phases.

Our investigation confirms some of the existing evidence. For example, as in Kose, Otrok and Whiteman (2003) or Lumsdaine and Prasad (2003) we find evidence of a significant world business cycle, despite different empirical techniques and data sets. However, it also provides new insights in the phenomenon. For instance, our results indicate that the common (world) indicator accounts for about 30 percent of the fluctuations in sales, industrial production, output and employment of the seven most industrialized countries, that it captures the more persistent portions of G-7 fluctuations and that it has more information than simple average or principal component measures obtained using G-7 GDPs or IPs. On the other hand, country specific indicators are useful in explaining certain GDP and employment episodes across time, but fail to track cyclical movements in the four variables over the entire sample. Perhaps more interestingly, we find that both world and country specific fluctuations are much more synchronized in contractions than expansions. That is, the uncertainty surrounding estimates of both world and country specific indicators is an order of magnitude smaller in the former than in the latter. Expansions tend to have large idiosyncratic components, both across variables and countries, while declines in economic activity have common timing and similar dynamics, both within and across countries.

Regarding the questions of interest of this paper, we do not find evidence of structural breaks in country indicators in the 1990s. Hence, the often cited idea that national cycles are disappearing finds no support from our analysis. These indicators are as significant in explaining the differential growth rate of GDP across countries in the mid 1990s as they were in the mid 1980s and, if anything, slightly more important. We also find little support for the idea that Euro cycles are different from those of the rest of the world or that a Euro area cycle is emerging in the 1990s. This result should be contrasted with Lumsdaine and Prasad (2003) and Artis et al. (2003) who instead detected the presence of a EU cycle using IP data. Our analysis shows that the Euro signal is much weaker when one considers a broader set of variables and that regional causes have minor explanatory power for G-7 fluctuations throughout the sample.

These set of results taken together imply that movements in the world indicator have been the stable and consistent reason for the commonalities of the fluctuations in the G-7 economies over time and that structural breaks in both the pattern of transmission across countries and in the sources of structural shocks are probably absent.

We document that the indicator captures a variety of influences going from the magnitude of world trade, to the behavior of commodity prices, to the stance of monetary policy in the G-7 and to the spending power of consumers. Interestingly, oil shocks, US technology shocks, world financial or monetary indicators and fiscal policy do not appear to be behind the fluctuations of the world indicator.

Our results are important from a policy point of view for several reasons. First, since variations in economic activity in countries with different institutions, economic structures and/or economic policies are driven by a common causes, markets more than policies appear to be the key to understanding the comovements in economic activity. This also means that policy institutions should probably de-emphasize national and regional cycles and instead focus on the identification of the market(s) or channel(s) that foster cross-country transmission. Second, since national or regional variables drive the national but not the world indicator, policies designed to counteract the tendencies dictated by world conditions may be ineffective. In addition, the presence of significant time variations indicates that reliance on policy actions which have been effective in the past is doomed to failure. Third, since cyclical time variations imply important asymmetries in the shape and the dynamics of international cycles, reliance on linear models in policy analyses may miss important and pervasive features of the data.

There are several interesting questions that our paper has left unanswered. For example, one would like to know more about the decreased posterior variability present in both the world and the national indicators in the late 1990s. What are the causes of this decrease? Is it the size of the shocks which has declined, the synchronization that has increased? If it is the latter, what can we say about the relative importance of contemporaneous vs. lagged transmission? Similarly, one would like to know why there is such a strong negative relationship between international synchronicity and volatility of GDP. Which way does causality go? What factors drive this correlation? Furthermore, since the model tracks reasonably well the four macroeconomic variables used in this study, one may want see whether this ability translates also in useful predictions of the future state of the world economy at various horizons. The results reported in Table 1 are promising. The exercises conducted in Canova and Ciccarelli (2002) suggest that this could be the case using information available up to one or two years in advance, but more evidence is clearly needed. We plan to take up all these questions in future work.

Appendices

A Estimation

A.1 Prior information

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

$$\begin{aligned} p(\Omega^{-1}) &= W(z_1, Q_1) \\ p(\sigma^2) &= IG\left(\frac{\zeta}{2}, \frac{\zeta s^2}{2}\right) \\ p(b_i) &= IG\left(\frac{\varpi_0}{2}, \frac{\delta_0}{2}\right) \\ p(\theta_{t-1} | \mathcal{F}_{t-1}) &= N\left(\bar{\theta}_{t-1|t-1}, \bar{\mathbf{R}}_{t-1|t-1}\right) \end{aligned} \tag{8}$$

where N stands for Normal, W for Wishart and IG for Inverse-gamma distributions, and \mathcal{F}_{t-1} denotes the information available at time $t - 1$. The last conditional distribution also implies the following prior for θ_t

$$p(\theta_t | \mathcal{F}_{t-1}) = N\left(\theta_{t|t-1}^*, \mathbf{R}_{t|t-1}^*\right)$$

where $\theta_{t|t-1}^* = \bar{\theta}_{t-1|t-1}$, and $\mathbf{R}_{t|t-1}^* = \bar{\mathbf{R}}_{t-1|t-1} + B_t$.

The prior distribution on Ω is standard and can be made uninformative by letting $Q_1 \rightarrow 0$. The priors for b_i and σ^2 are general and can be made uninformative letting $\varpi_0, \delta_0 \rightarrow 0$ and $\zeta \rightarrow 0$ respectively. Finally, the prior on θ_t can be derived from normal initial conditions and the Kalman filter.

We collect the hyperparameters of the prior in the vector

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

where $\text{vech}(\cdot)$ denotes the column-wise vectorization of a symmetric matrix. We assume that components of μ are either known or can be estimated in the data, for example, splitting the sample in two pieces, using the first piece (“training” sample) to estimate the μ (by running univariate constant coefficients AR(p) or country by country VAR(p)) and the second to estimate posterior distributions and to conduct inference. The following table reports the values for the hyperparameters used under an informative and a non informative choice:

	ζ	s^2	z_1	Q_1	ϖ_0	δ_0	γ_1	γ_2	θ_0	\bar{R}_0
informative	1.0	$\hat{\sigma}^2$	$N \cdot G + 20$	\hat{Q}_1	10	1	1	0	$\hat{\theta}_0$	I_J
non-informative	0.0	$\hat{\sigma}^2$	$N \cdot G + 1$	\hat{Q}_1	0	0	1	0	$\hat{\theta}_0$	I_J

where $\hat{\sigma}^2$ is calculated averaging the estimated variances of NG AR(p) models, \hat{Q}_1 is the estimated variance-covariance in the time invariant version of (1), $\hat{\theta}_0$ is initialized with a sequential OLS on (1), over the sample 1975-1980, and J is the dimension of θ_t . Notice that the values of the hyperparameters under both the informative and the non informative set-up have been chosen using previous experience (Canova and Ciccarelli, 2002) and information contained in the sample.

A.2 Posterior distributions

To calculate the posterior distribution of the unknowns of the model $\psi = (\Omega, b_i, \sigma^2, \{\theta_t\}_t)$, we combine the above priors with the likelihood of the data, which, with the specification adopted, is proportional to

$$L \propto \left(\prod_{t=1}^T \sigma_t \right)^{-NG/2} |\Omega|^{-T/2} \exp \left[-\frac{1}{2} \sum_t (Y_t - X_t \Xi \theta_t)' (\sigma_t \Omega)^{-1} (Y_t - X_t \Xi \theta_t) \right] \quad (9)$$

where $Y^T = (Y_1, \dots, Y_T)$ denotes the data, and $\sigma_t = (1 + \sigma^2 \mathbf{X}_t' \mathbf{X}_t)$. Using Bayes rule, we have $p(\psi | Y^T) = \frac{p(\psi)L(Y^T|\psi)}{p(Y^T)} \propto p(\psi)L(Y^T|\psi)$. Given $p(\psi | Y^T)$, the posterior distribution for the components of ψ , $p(\Omega | Y^T)$, $p(b_i | Y^T)$, $p(\sigma^2 | Y^T)$ and $p(\theta_t | Y^T)$, can then be obtained by integrating out nuisance parameters from $p(\psi | Y^T)$. Once these distributions are obtained, location and dispersion measures for ψ and for any interesting continuous function of them can be obtained.

For the model considered in this paper, it is impossible to compute $p(\psi | Y^T)$ 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 the parameters of interest. However, while the conditional posteriors of Ω, b_i and $(\theta_t)_t$ are known, the conditional posterior distribution of σ^2 is not standard and therefore a Metropolis steps within the Gibbs sampler is needed.

Under regularity conditions (see Geweke(2000)), cycling through these conditional distributions will produce in the limit draws from the joint posterior of interest. From these, marginal distributions can be computed by averaging over draws.

Denoting $\psi_{-\kappa}$ the vector of ψ excluding the parameter κ , the conditional distributions of interest are

$$\begin{aligned}\theta_t | Y^T, \psi_{-\theta_t}, \mathbf{l}_{t-1} &\sim N\left(\bar{\theta}_{t|t}, \bar{\mathbf{R}}_{t|t}\right) \quad t \leq T, \\ \Omega^{-1} | Y^T, \psi_{-\Omega} &\sim W\left(z_1 + T, \left[\frac{\sum_t (Y_t - X_t \Xi \theta_t)(Y_t - X_t \Xi \theta_t)'}{\sigma_t} + Q_1^{-1}\right]^{-1}\right), \\ b_i | Y^T, \psi_{-b_i} &\sim IG\left(\frac{T + \varpi_0}{2}, \frac{\sum_t (\theta_t^i - \theta_{t-1}^i)' (\theta_t^i - \theta_{t-1}^i) + \delta_0}{2\xi_t}\right), \\ \sigma^2 | Y^T, \psi_{-\sigma^2} &\propto L(Y^T | \psi) \times p(\sigma^2),\end{aligned}$$

where $\bar{\theta}_{t|t}$ and $\bar{\mathbf{R}}_{t|t}$ are the one-period-ahead forecasts of θ_t and the variance-covariance matrix of the forecast error, respectively, calculated by the Kalman Filter as:

$$\begin{aligned}\bar{\theta}_{t|t} &= \theta_{t|t-1}^* + \mathbf{R}_{t|t-1}^* (X_t \Xi)' F_t (Y_t - X_t \Xi \theta_{t|t-1}^*) \\ \bar{\mathbf{R}}_{t|t} &= \mathbf{R}_{t|t-1}^* - \mathbf{R}_{t|t-1}^* (X_t \Xi)' F_t (X_t \Xi) \mathbf{R}_{t|t-1}^* \\ F_t &= \left((X_t \Xi) \mathbf{R}_{t|t-1}^* (X_t \Xi)' + \Omega\right)^{-1}.\end{aligned}$$

and θ_t^i is the i^{th} -subvector of θ , with $i = 1, 2$.

The posterior for σ^2 is simulated assuming a Random Walk Chain Metropolis-Hastings, which, at each iteration l , generates candidates draws according to

$$\left(\sigma^2\right)^* = \left(\sigma^2\right)^{(l-1)} + z$$

where z is assumed normal with mean zero and variance equal to c^2 . The latter is chosen to ensure that the average acceptance probability is in the region 0.2-0.5.⁴

Using posterior draws, the posterior distributions of λ_t and α_t can be estimated using kernel methods and, in turns, the posterior distributions of WLI_t and CLI_t and/or the posterior distribution of the importance of the common components in explaining cyclical fluctuations can be obtained. For example, the posterior mean of WLI can be approximated by $\frac{1}{H} \sum_h \mathcal{W}_t \lambda_t^h$ and a credible 68% interval can be obtained ordering the draws of WLI_t^h for each t .

⁴See e.g. Chib and Greenberg (1995).

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 ergodic (posterior) distribution. We have checked 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 results we present are based on chains with 24000 draws: 600 blocks of 40 draws were made and the last draw for each block is retained after discarding the first 4000. This means that a total of 500 draws is used at each t to conduct posterior inference.

B Specification searches

We have conducted a number of preliminary checks on the nature of the model to examine if all the features included in the specification are really necessary to capture the dynamics of the variables under consideration. In particular, we have examined whether the factorization (3) is exact or not and whether the presence of international interdependencies is necessary to capture the dynamics of the four time series for the G-7 countries. Verification of these hypotheses is important because the specification can be considerably simplified if the factorization is exact and interdependencies are absent. The first hypothesis is examined by checking whether the posterior distribution of σ^2 is more centered around zero than the prior. To formally evaluate the closeness of σ^2 to zero we follow Chib and Greenberg (1995, p. 344) and construct the ratio $\mathcal{S} = \frac{P(\sigma^2 \leq \epsilon | y) P(\sigma^2 > \epsilon | y)}{P(\sigma^2 \leq \epsilon) P(\sigma^2 > \epsilon)}$ for various values of ϵ , where the numerator is computed using posterior draws and the denominator using the inverse gamma prior.

The prior and posterior distributions of σ^2 are presented in figure A.1. The fact that the posterior distribution is shifted to the right with respect to the prior assumption indicates that the factorization is not exact and more is left in the error term. In fact, the ratio \mathcal{S} for $\epsilon = 0.05$ is 0.17 and for $\epsilon = 0.08$ it is 0.36. These small values are evidence against an exact factorization.

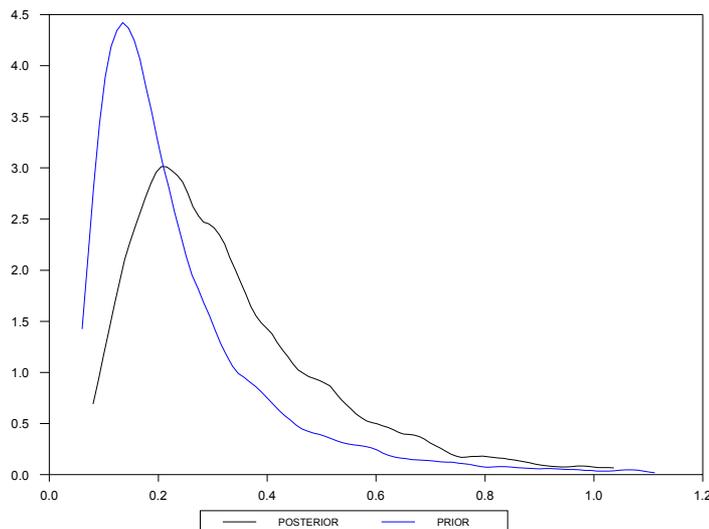


Figure A1. Prior and posterior density of σ^2

The importance of interdependencies is examined by running two different models, one with interdependencies and one without, and using Bayes factors to evaluate their relative approximation to the data. The Bayes factor for a model with and without interdependencies is 2.06 indicating that interdependencies are important to capture the dynamics of G-7 cycles. Intuitively, this means that lagged transmission is as important as contemporaneous comovements in generating common fluctuations in the G-7.

We have also examined the performance of a model with three vectors of indicators (a world, a country specific and a variable specific) against a model with only a world and country specific indicators. The Bayes factor in this case is 1.02, suggesting that the contribution of the third vector of indices is small. In economic terms this means that the dynamics of the four variables of the system are relatively similar, a comforting result since sales, employment, industrial production were chosen because they were suspected to be sufficiently coincident with GDP not only in the US (as the NBER practice indicates) but also in the other G-7 countries.

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