Business cycle convergence in EMU: A first look at the second moment*

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Abstract

We propose the analysis of the dynamics of the standard deviation of business cycles across euro area countries in order to evaluate the patterns of cyclical convergence in the European Monetary Union (EMU) for the period 1977-2002. We identify significant business cycle divergence taking place in the mid-eighties, followed by a persistent convergence period spanning most of the nineties. This convergent episode finishes roughly with the birth of the European Monetary Union. Furthermore, we use a simple autoregressive model with structural breaks in order to characterize the different phases of business cycle synchronization in terms of persistence of convergence/divergence and level of synchronization. We link the systematic trend towards synchronization of business cycles on the run-up to EMU with the process of fiscal homogeneization across euro area economies.

Keywords: Business cycles, business cycle convergence, European Monetary Union.

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1 Introduction

One of the most important debates related with the European integration process since the end of the decade of the eighties has dealt with the creation of the European Economic and Monetary Union (EMU). Two issues in relation to membership in EMU have been emphasized. Firstly, whether all members which want to and demonstrate to be able to join the currency area could be a member of the arrangement or whether, on the contrary, only a core of countries in Europe should participate in EMU. The second issue has dealt with the Maastricht convergence and stability criteria and their importance in support of the maintenance and performance of the currency area. Both issues are still important and their shadow underlies the European enlargement process. This is especially true for the latter since, under the Accession Treaty and due to the no opt-out provision stipulated in the Copenhagen European Council in 2002, all the new member states go straight into the Stage Three of EMU and have to take on the Maastricht convergence criteria. Recently, the relevance of the convergence criteria has been placed in doubt in practise by the breaks of the Stability and Growth Pact (SGP) criteria that took place in some European countries since the first years of the new century. At the same time, several authors (see Kopits and Székely, 2003; for example) argue about the convenience of relaxing these convergence criteria for the new members in order to allow faster real convergence and the adoption of the *acquis communaitaire*.

This discussion has to be assessed within the framework of the Optimum Currency Area (OCA) Theory. This branch of literature had its formal birth with Mundell's (1961) article and is often divided in two main branches, the first one centered in the optimum geographic domain of a currency area and the definition of criteria through which this optimality can be defined, and the second one, centered in the analysis of the trade-off between the costs and benefits from adopting a single currency. Since Mundell's (1961) proposition that a region, defined as area within which there is factor mobility, but shows factor immobility with other regions, is an OCA, a vast body of literature has developed. Early on, McKinnon (1963) extended the criteria by considering optimality in terms of openness and size of the economy. Together with factor mobility, openness and size of the economy, a large number of OCA criteria have been suggested in the literature (see for example Tavlas, 1993, or Mongelli, 2002). Most of these criteria can be summarized under the consideration that a region which shows a high degree of business cycle homogeneity is an OCA. Therefore, most of the empirical research in this area has concentrated in judging the suitability of potential or existing currency area members on the basis of these prerequisites. The current literature on OCA criteria has been highly influenced by the work of Frankel and Rose (1998), who put forward the "endogeneity of OCA" hypothesis by which the structure and relations of the economies that join a currency area are likely to change dramatically as a result of the effective participation of the currency area. As a corollary, a country would be able to satisfy the criteria for participation in a currency area better *ex post* (that is, after joining the currency area) than *ex ante* and thus, the suitability to form a currency area cannot be analyzed on the basis of these prerequisites without considering the possible endogeneity of the criteria. Frankel and Rose considered the endogeneity between two of the OCA criteria, trade integration and business cycle coherence, but other potential endogeneities have been considered (see De Grauwe and Mongelli, 2005, for a survey). The underlying rationale for the endogeneity defined by Frankel and Rose is theoretically ambiguous. The increase in trade after joining a monetary union may induce more synchronization in business cycles if intra-industry trade predominate over inter-industry trade (see European Commission, 1990) or, on the other hand, may induce business cycles to become more idiosyncratic if countries become more specialized as a result of the prevalence of inter-industry trade over the rest of effects (see Krugman, 1991). Frankel and Rose (1998) find evidence that this relationship is empirically unambiguous and that international trade integration is positively related with more synchronized business cycles (see also Fidrmuc, 2004, for evidence using intra-industry trade indicators).

The process of monetary integration in Europe has contributed to the development of the academic literature on European business cycle synchronization. Different variables and filtering procedures have been used to measure the business cycle and diverse measures have been used to analyze the coherence or similarity among business cycles. The most widely used measure of business cycle coherence is the correlation coefficient between national cycles. Within this line of study, several findings can be emphasized. Firstly, there exists evidence of homogeneity of business cycles in the European Union (EU) (Agresti and Mojon, 2001, Christodoulakis *et alia*, 1995, Wynne and Koo, 2000) and certain authors claim that we can talk about a European business cycle almost in the same terms that we talk about a US business cycle (Agresti and Mojon, 2001, Wynne and Koo, 2000). Secondly, business cycle correlation in Europe is a relatively recent phenomenon. Artis and Zhang (1997 and 1999) point out that the emergence of a European cycle seems to coincide with the inception of the Exchange Rate Mechanism (ERM) and that it is a specific fact to the group of countries participating in ERM. Inklaar and De Haan (2001), using Artis and Zhang's (1997 and 1999) updated dataset but with different subsamples, found no evidence of a systematic relationship between business cycle homogeneity and monetary integration and pointed out that most ERM cycles are better correlated during the period 1971-1979 than in the period 1979-1987. Finally, other authors locate the convergence period starting in the early nineties (Angeloni and Dedola, 1999, Massmann and Mitchell, 2003).

In this piece of research we analyze the dynamics of business cycle dispersion in Europe for the period 1977-2002. We extract the business cycle from quarterly GDP series for all current members of EMU for the period 1977-2002 using an unobserved components model in the spirit of Harvey (1989) using Kalman filtering methods. As a measure of coherence, the time series of the cross-country standard deviation of business cycles is studied, and significant changes in this measure are assessed using Carree and Klomp's (1997) convergence test. We also analyze the time series properties of our business cycle synchronization measure in order to characterize different systematic periods of convergence/divergence in the cycles of EMU members. To our knowledge, the dynamics of business cycle dispersion has not been exploited until now as an indicator of business cycle coherence.

Our results show a significant period of convergence in business cycles since the end of the seventies to the first years of the eighties. The next decade until the mid-nineties is characterized by a period of business cycle divergence. The last significant convergence period observed since the beginning of the nineties, which seems to finish with the birth of EMU, coincides with the period of fiscal policy homogeneization following the implementation of the Maastricht convergence criteria. Our results provide some evidence on the relationship between fiscal policy homogeneization and business cycle synchronization.

The paper is structured as follows. In section two we present the business cycle extraction method and the basic characteristics of business cycle dispersion in Europe for the period under study. Sections three presents the results of the convergence tests and identifies the different synchronization periods in the sample. In section four, the results are put in context of the homogeneization of fiscal policy following the implementation of the Maastricht Treaty. Section five concludes.

2 Business cycles in EMU

In order to study the convergence of business cycles, an estimate of the cyclical component of the variable of interest (in our case, quarterly real GDP) needs to be obtained. For this purpose, following Harvey (1989) and Harvey and Jaeger (1993), we will decompose the GDP series of each country under study into unobservable trend, cyclical and irregular components. If y_{it} is the (logged) GDP corresponding to country i in period t then

$$y_{it} = \tau_{it} + \phi_{it} + \varepsilon^y_{it}, \;\; \varepsilon^y_{it} \sim \mathbf{NID}(0, \sigma^2_{\varepsilon^y}), \tag{1}$$

where τ_{it} is the trend component, ϕ_{it} is the cyclical component and ε_{it} is the (white noise) irregular component. The trend component, in its most general specification, will be assumed to be a random walk with a drift, where the drift follows a random walk as well, that is,

$$\begin{aligned} \tau_{it} &= \tau_{it-1} + \beta_{t-1} + \varepsilon^\tau_{it}, \;\; \varepsilon^\tau_{it} \sim \mathbf{NID}(0, \sigma^2_{\varepsilon^\tau}), & (2) \\ \beta_{it} &= \beta_{it-1} + \varepsilon^\beta_{it}, \;\; \varepsilon^\beta_{it} \sim \mathbf{NID}(0, \sigma^2_{\varepsilon^\beta}). & (3) \end{aligned}$$

This specification of the trend component nests several interesting cases. It should be noticed that if $\sigma^2_{\varepsilon^\tau} > 0$ and $\sigma^2_{\varepsilon^\beta} > 0$, this component induces an $\mathsf{I}(2)$ trend on y_{it}. On the other hand, if $\sigma^2_{\varepsilon^\tau} > 0$ and $\sigma^2_{\varepsilon^\beta} = 0$, τ_{it} is a random walk trend with drift. The case $\sigma^2_{\varepsilon^\tau} = 0$ and $\sigma^2_{\varepsilon^\beta} > 0$ defines a smoothly changing trend,[1] and $\sigma^2_{\varepsilon^\tau} = 0$ and $\sigma^2_{\varepsilon^\beta} = 0$ implies a deterministic linear trend.

The cyclical component is assumed to follow a damped stochastic sine-cosine wave, specified as

$$\begin{bmatrix} \phi_t \\ \phi^*_t \end{bmatrix} = \rho \begin{bmatrix} \cos\lambda & \sin\lambda \\ -\sin\lambda & \cos\lambda \end{bmatrix} \begin{bmatrix} \phi_{t-1} \\ \phi^*_{t-1} \end{bmatrix} + \begin{bmatrix} \theta_t \\ \theta^*_t \end{bmatrix}, \;\; \begin{bmatrix} \theta_t \\ \theta^*_t \end{bmatrix} \sim \mathbf{NID}(\mathbf{0}, \mathbf{\Sigma}_\theta), \tag{4}$$

for $\rho \in [0,1]$, $\lambda \in (0,\pi)$ and $\mathbf{\Sigma}_\theta = \text{diag}(\sigma^2_\theta, \sigma^2_\theta)$, so the disturbances of the cyclical component are assumed independent and of equal variance. It can be easily shown that the specification given by (4) implies that the cycle follows an ARMA(2,1) process, and that the constraints on the parameter space given above restricts the roots of the lag polynomial to lie on the region of the parameter space that leads to pseudo-cyclical behaviour in ϕ_t.

[1]The Hodrick-Prescott trend (Hodrick and Prescott, 1997) appears as a special case of the decomposition of a series into a smooth trend and an irregular component for specific values of $\sigma^2_{\varepsilon^y}/\sigma^2_{\varepsilon^\tau}$ (see Harvey, 1989).

The model specified by (1)-(4) can be written in state space form in a straightforward manner and estimated using maximum likelihood methods via the Kalman filter and the prediction error decomposition. Once the estimates of the parameters in (1)-(4) are obtained, the cyclical component can be recovered as the smoothed estimate of ϕ_{it}, $\hat{\phi}_{it}$, which is given by $\mathsf{E}(\phi_{it}|\{y_{it}\}_{t=1}^{T})$.

The unobserved components model given by (1)-(4) was estimated for the GDP data corresponding to the twelve European countries that currently form EMU. Figure 1 presents the smoothed cyclical components of the quarterly GDP series corresponding to each one of the EMU countries.[2] Figure 2 plots the time series of the cross-country standard deviation of the cyclical component, together with the weighted standard deviation using purchasing power parity adjusted GDP levels in 2000 (relative to the EMU total) as weights, so as to weight down deviations from countries that amount to a small proportion of total production in the aggregate euro area.

The overall dynamics of the weighted and unweighted dispersion measures present similar dynamic patterns, although the difference in the level of the standard deviation indicates that quantitatively most of the lack of business cycle synchronization stems from countries of relatively small size in the euro area. After a period of cyclical convergence from the end of the seventies to the beginning of the eighties, a persistent business cycle divergence trend takes place in the second half of the eighties, which is reversed in the first years of the decade of the nineties. The convergent pattern in the nineties culminates with a global minimum in the dispersion measure for the period under study, which takes place in the first quarter of 1998. The inception of EMU in 1999 is followed by a further reversion in the trend of the dispersion of European business cycles. By the end of the sample (end 2002), the dispersion of business cycles in EU-15 has risen approximately to the levels observed in the mid-nineties.

In the following section we will analyze the statistical significance of the changes in business cycle dispersion across EMU economies for different horizons, and assess the issue of the existence of a structural break in the dynamics of business cycle dispersion across euro area countries.

3 Business cycle convergence and divergence patterns in the euro area

3.1 Testing for business cycle convergence/divergence

A first assessment of the patterns of convergence of business cycles across EMU economies can be done by studying the changes in dispersion plotted in Figure 2. The question that needs to be answered concerns whether the dynamics of the standard deviation of the cyclical component of GDP lead to significant changes in the level of dispersion. Lichtenberg (1994) and Carree and Klomp (1997) tackle the issue of testing for convergence, defined as a reduction of the standard deviation of the variable of interest across economic units. In order to test for the significance of changes in the standard deviation of European business

[2]Detailed results of the estimated models can be obtained from the authors upon request.

cycles, we computed Carree and Klomp's (1997) T_2 test statistic, given by

$$T_{2,t,\tau} = (N - 2.5)\log[1 + 0.25(\hat{\sigma}_t^2 - \hat{\sigma}_{t+\tau}^2)^2/(\hat{\sigma}_t^2\hat{\sigma}_{t+\tau}^2 - \hat{\sigma}_{t,t+\tau}^2)], \tag{5}$$

where $\hat{\sigma}_t$ is the cross-country standard deviation of $\hat{\phi}_{it}$ and $\hat{\sigma}_{t,t+\tau}$ is the covariance between $\hat{\phi}_{it}$ and $\hat{\phi}_{it+\tau}$. Under the null hypothesis of no change in the standard deviation between period t and period $t+\tau$, T_2 is $\chi^2(1)$ distributed, and can thus be used to test for significant changes in dispersion.

$T_{2,t,\tau}$ was calculated for our sample using different potential convergence/divergence horizons ranging from two years ($\tau = 8$) to eight years ($\tau = 32$). Figure 3 presents the changes in the standard deviation of EU-15 business cycles that appeared significant at the 5% significance level for the horizons corresponding to two, four, six and eight years. That is, the variable which is plotted in Figure 3 is defined as

$$c_t = (\hat{\sigma}_t - \hat{\sigma}_{t+\tau})I[T_{2,t,\tau} > \chi^2_{0.95}(1)], \tag{6}$$

where τ is alternatively equal to 8, 16, 24 and 32 quarters, $\chi^2_{0.95}(1)$ is the 95th percentile of the $\chi^2(1)$ distribution and $I[\cdot]$ is the indicator function, taking value one if the argument is true and zero otherwise.

Figure 3 indicates that the medium run dynamics shown in Figure 2 actually led to significant changes in the dispersion of business cycles in EMU for the period under study. In particular, a long period of sizeable and significant convergence took place in the nineties and finished with the inception of the monetary union in 1999. The short run divergence which appears at the end of the sample appears significant when considering dispersion changes in horizons of two, four and six years. This result can be considered as evidence against the endogeneity of optimum currency area criteria (see Frankel and Rose, 1998), although it should be interpreted with care given the short size of the post-EMU sample.

3.2 Synchronization regimes on the way to EMU

In this section we will analyze the time series properties of our business cycle dispersion measure in order to identify systematic periods where different degrees of business cycle synchronization take place. Different unit root tests give contradicting results on the degree of integration of the standard deviation of business cycles in EMU. While a simple Dickey-Fuller test cannot reject the null hypothesis of a unit root at any sensible significance level (the DF test statistic takes a value of -2.17, while the 10% critical value is -2.58), the KPSS test cannot reject stationarity of the series (the KPSS test statistic equals 0.24, while the 10% critical value is at 0.34). We will consider the series to be represented by an autoregressive process potentially subject to breaks in the intercept and the autoregressive parameter. This leads to the possible existence of different regimes of business cycle synchronization. Setting an autoregressive lag of one, which appears sufficient to account for the autocorrelation present in the data, the specification we are considering is the following,

$$s_t = \sum_{j=1}^{R}(\alpha_{0,j} + \alpha_{1,j}s_{t-1})I(T_{j-1} \leq t < T_j) + \varepsilon_t, \tag{7}$$

where ε_t is a white noise disturbance, R is the number of regimes considered (therefore $R-1$ is the number of breaks in the parameters of the process), T_0 is the time index of the first observation and T_R is the time index of the last observation. Table 1 presents the results of the estimation of (7) for the cases $R = 1, 2, 3$ and 4, together with the sup-F test for the null of the model without breaks against each one of the models with breaks. The breaks are estimated in each case by choosing the values $T_1, ..., T_{R-1}$ that globally minimize of the sum of squared residuals, that is,

$$\{\hat{T}_1, \dots, \hat{T}_{R-1}\} = \arg\min \sum_{t=1}^{T_R} \hat{\varepsilon}(T_1, \dots, T_{R-1})_t^2,$$

where the search for the breaks is done after imposing a minimum of 15% of the full sample to be contained in each regime, in order to avoid spurious results caused by small subsample sizes. The significance level of the sup-F test is obtained in each case by simulating the asymptotic critical values using the method proposed by Bai and Perron (1998).

The results of the sup-F test shown in Table 1 present evidence for the model with three breaks. In the first regime (1977/1-1983/4) the process presents high persistence and convergence to a very low level of dispersion. This is followed by a period of relatively low persistence (an estimate of $\alpha_{1,2}$ of 0.33, as compared to 0.94 in the first regime) and a higher unconditional expectation of s_t, which ends in the third quarter of 1988. The persistence of the dispersion of business cycles, as proxied by the autoregressive parameters $\alpha_{1,3}$ and $\alpha_{1,4}$ is significantly higher than in the second period (0.75 versus 0.33) and does not change significantly between the third and fourth regime. The implied unconditional expectation of the estimated process, however, is much lower in the fourth regime. Figure 4 summarizes the results concerning the long-run unconditional expectation of the autoregressive process with breaks by plotting $\mathbf{E}(s_t|\hat{T}_1, \hat{T}_2, \hat{T}_3, \hat{\alpha}_{0,j}, \hat{\alpha}_{1,j}, j = 1, \dots, 4) = \sum_{j=1}^{4} [\hat{\alpha}_{0,j}/(1 - \hat{\alpha}_{1,j})] I(\hat{T}_{j-1} \leq t < \hat{T}_j)$ together with the original dispersion series. The shaded area in Figure 4 shows the 90% confidence interval corresponding to each one of the estimated breaks.

To sum up, four distinct regimes concerning the level and persistence of business cycle divergence can be found in the euro area for the period under study. A first period spanning the end of the seventies and beginning of the eighties is marked by a systematic and persistent reduction of business cycle dispersion. The following decade (mid-eighties to mid -nineties) is characterized by mean reversion to a significantly higher level of dispersion, and thus to a lower degree of business cycle synchronization in the countries that currently compose EMU. Furthermore, the dynamics of the process became more persistent in the nineties. Finally, the period starting in the mid-nineties is characterized by convergence to a lower level of dispersion.

4 Business cycle synchronization in EMU: the role of fiscal policy

Recently, the importance of the link between fiscal policy and business cycle synchronization has been emphasized by the optimum currency area literature. Darvas *et alia* (2005) find that convergence in fiscal balances is systematically linked to business cycle convergence, and that the relationship exists even when the potential endogeneity of fiscal policy responses is accounted for. Böwer and Guillemineau (2006) show that fiscal policy homogeneization has been one of the robust determinants of business cycle synchronization in EMU and Akin (2006) provides evidence of the importance of similarity in idyosincratic fiscal shocks as a determinant of cyclical convergence in a broader set of countries.

The significant convergence trend observed since the beginning of the nineties and which seems to finish with the birth of EMU coincides with the period of widespread fiscal consolidation among European countries following the convergence lines stated in the Maastricht Treaty. To the extent that differences in the implementation of fiscal policy were responsible for asymmetric shocks in the countries of our sample, the homogenization of fiscal policies on the run-up to EMU may be held partly responsible for the business cycle synchronization trend observed in the nineties. Figure 5 shows the dynamics of the average and dispersion (standard deviation) of budget deficits across EMU countries in the period 1990-2003 (where the data are homogeneous enough as for cross-country comparisons to make sense), together with our measure of business cycle dispersion.[3] The reduction in asymmetric shocks and business cycle dispersion across euro area countries observed in the nineties took place in parallel not only to a sustained reduction of the average budget deficit among EMU countries, but also to a convergence in fiscal balances (in the sense of a sustained reduction in the dispersion of budget deficit figures across euro area economies).[4] Fatás and Mihov (2003a and 2003b) also document a deeper convergence in the conduct of fiscal policy among EMU countries in this period, which they label "coherence" and which implies a reduction in the use of discretionary fiscal measures across euro area countries. To the extent that the observed differences in business cycle in Europe are due mostly to differences in variables that are under control of the government, the process of fiscal coordination would be behind this trend in business cycle synchronization (see also Christodoulakis *et alia*, 1995).
On the other hand, the business cycle divergence observed across EMU economies in the end of the sample coincides with the reversal in the fiscal consolidation and homogeneization trend of the nineties. In particular, while a considerable reduction of budget deficits is observed on average for the period 1997-2000, the standard deviation of fiscal balances across euro area economies increased significantly in these three years (the T_2 test for equality of dispersion in 1997 and 2000 takes a value of 6.3 and thus strongly rejects the null hypothesis of no change in the standard deviation of budget deficits). While assigning a causal relationship between the homogeneization of fiscal positions and the synchronization of business cycles in the period of convergence could be justified by the asymmetric nature of discretionary fiscal policy across countries, the causal direction of the relationship between the change of trend in 1999 both in the dispersion of business cycles and budget deficits is more difficult to disentangle. This is the case since the divergence in fiscal balances across countries may itself be the result of reactions to different asymmetric shocks hitting the euro area economies.

[3]The data for budget deficits are of yearly frequency and is sourced from Eurostat (General government net borrowing/lending as percentage of GDP). Data for Germany starts only in 1991 and for Spain in 1996.

[4]The same applies if cyclical adjusted measures of the budget deficit are used. Furthermore, a convergence in the size of the government, measured by the share of taxes or government expenditure on GDP, also took place in this period (see Fatás and Mihov, 2003a and 2003b).

From a theoretical point of view, the parallel divergence patterns in business cycles and fiscal stance following the birth of EMU can be seen as a result of the interplay between monetary policy in a currency union and national fiscal policies. Using a simple game-theoretical setting where the interaction of monetary, fiscal and wage policy is modelled, Onorante (2004) shows that fiscal activism is always increased by entry in a monetary union. The reason for such a result is that the potential costs of running higher deficits (in terms of higher interest rates) for a country in the monetary union are lower than if monetary policy was independent, since the costs entailed by the increase of interest rates partly fall on other member countries.[5]

5 Conclusions

In this paper, we analyze the dynamics of business cycle dispersion in EMU for the period 1977-2002. We extract business cycles from GDP data using an unobserved components model and analyze the significance of changes in the standard deviation of cycles across EMU countries. Our results show a significant period of convergence in business cycles spanning the end of the seventies and the first years of the eighties, which is followed by by a period of business cycle divergence. A significant convergence period is observed since the beginning of the nineties and finishes with the birth of EMU.

The last strong business cycle convergence period runs in parallel to the synchronization in fiscal policy initiated by the implementation of the Maastricht Treaty. This result sheds a light on the importance of similarity in idyosincratic fiscal shocks as a determinant of cyclical convergence, and highlights the relevance of fiscal policy as a source of asymmetric shocks in the context of OCA theory.

[5]For theoretical models of the interaction between monetary and fiscal policy in monetary unions, see for example Silbert (1992), Levine and Brociner (1994) or Dixit and Lambertini (2001, 2003).

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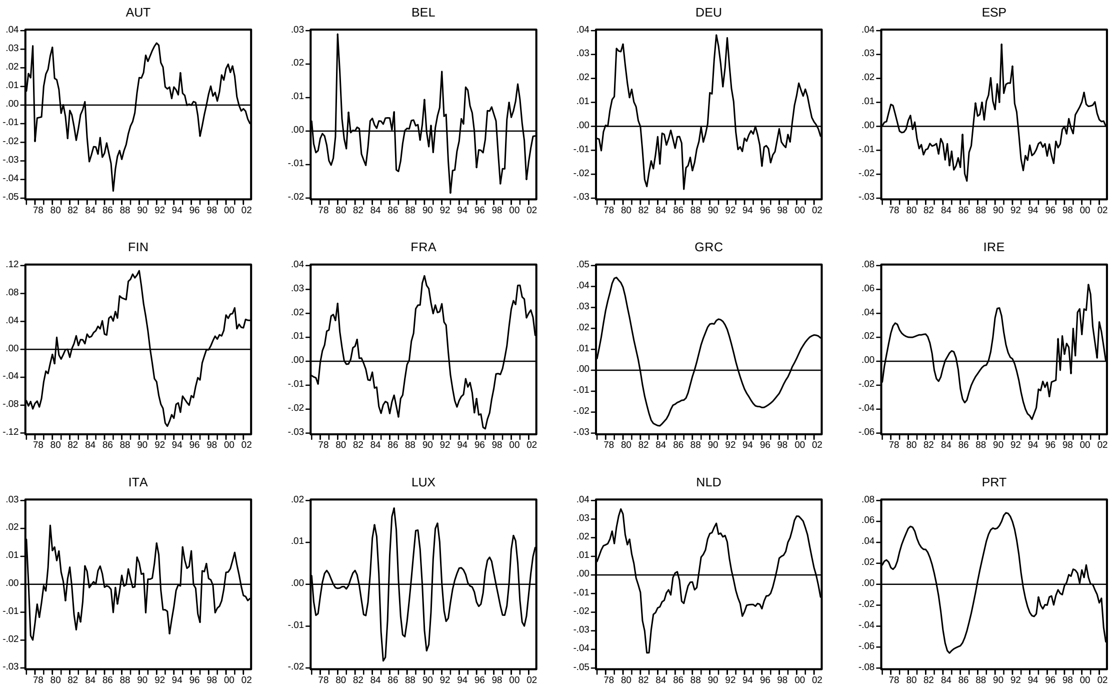

Figure 1: Cyclical component of (log) GDP: EMU countries

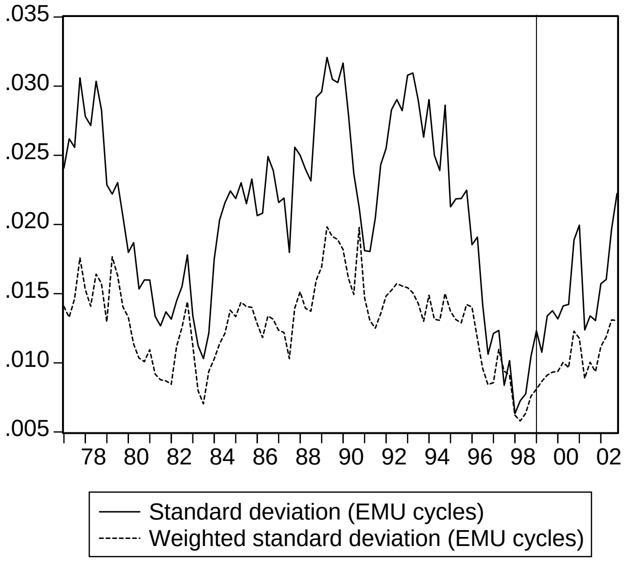

Figure 2: Dispersion of business cycles: EMU countries

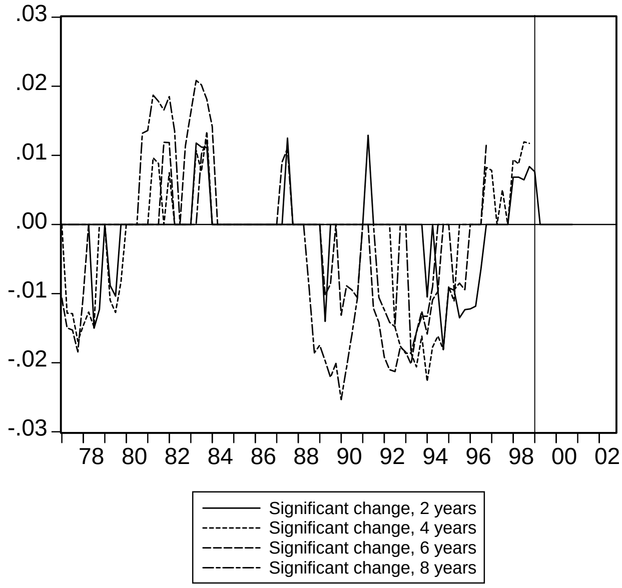

Figure 3: Significant dispersion changes: EMU countries

Table 1: **AR(1) models with structural breaks**

	R =1	R =2	R =3	R =4
$\alpha_{0,1}$	0.002** (0.001)	0.002** (0.001)	0.001 (0.002)	0.001 (0.002)
$\alpha_{1,1}$	0.912*** (0.041)	0.899*** (0.055)	0.942*** (0.080)	0.942*** (0.080)
$\alpha_{0,2}$	-	0.004** (0.001)	0.007*** (0.002)	0.015*** (0.004)
$\alpha_{1,2}$	-	0.752*** (0.092)	0.725*** (0.087)	0.331*** (0.188)
$\alpha_{0,3}$	-	-	0.004** (0.001)	0.007** (0.003)
$\alpha_{1,3}$	-	-	0.752*** (0.089)	0.757*** (0.124)
$\alpha_{0,4}$	-	-	-	0.004** (0.001)
$\alpha_{1,4}$	-	-	-	0.752*** (0.087)
Q(1) test	0.067	0.032	0.005	0.039
Q(4) test	1.580	1.363	1.113	1.892
JB test	0.704	0.442	0.427	0.492
$(T_1, \ldots, T_{R-1})$	-	(1994/4)	(1983/4, 1994/4)	(1983/4, 1988/3, 1994/4)
sup-F test	-	2.888	3.654	4.658**

*(**)[***] stands for significance at the 10%(5%)[1%] level. "Q(z) test" is the Ljung-Box test statistic for autocorrelation up to zth order. "JB test" is the Jarque Bera test statistic for residual normality. The significance level of the sup-F test were computed using the algorithm in Bai and Perron (1998), using 1000 replications with Wiener processes of sample size 500.

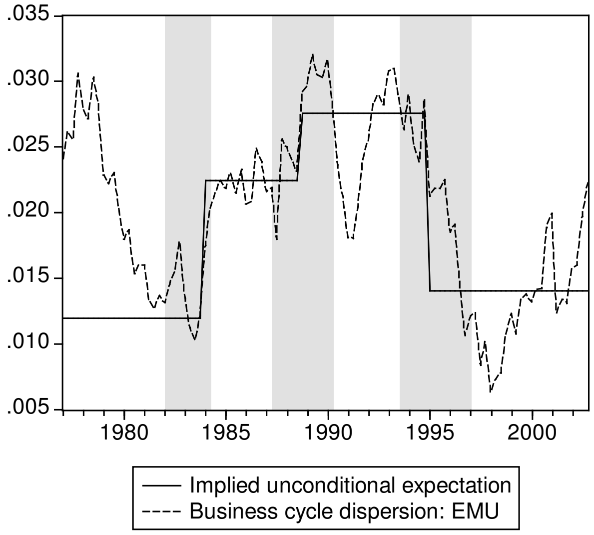

Figure 4: Business cycle dispersion and $\mathbf{E}(s_t|\hat{T}_1, \hat{T}_2, \hat{T}_3, \hat{\alpha}_{0,j}, \hat{\alpha}_{1,j}, j = 1, \ldots, 4)$

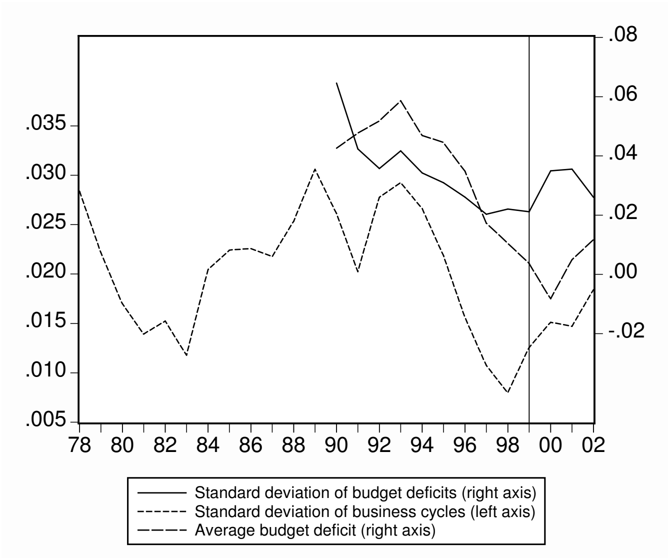

Figure 5: Business cycle synchronization and fiscal policy: EMU countries