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˙National Policies and Local Economies:
Europe and the US

by

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ABSTRACT

This paper studies the synchronization of output fluctuations in European regions and US counties. We extend the two component dynamic factor model a' la Sargent and Sims (1977) by introducing an intermediate-level shock, which is common to all regions (counties) in each country (state), but it is not common to Europe (US) as a whole. We build on Forni and Reichlin (1995, 1996) to propose a simple method of estimation which is based on Law of Large Numbers results and exploits the large cross-sectional dimension of our data set. Our empirical findings show that Europe has a level of integration similar to that of the US. In general, we find that the national dimension in Europe is not very important: around 75 % of output variance is explained by global and purely local dynamics. Similar numbers are found for US counties and US states. The study of the dynamic profile of the components, however, shows that Europe, unlike the US, has no traditional business cycle. Shocks are very persistent and the bulk of the variance is in the long-run. We also find a European core of regions with a particular high level of integration. The core, however, is not defined by a set of nations, but by regions belonging to different countries.

1. Introduction[1]

The question of how far we are from an integrated Europe is, to a certain extent, the same as asking how much regional GDP fluctuations comove. In each region, total output growth can be seen as generated by shocks which are purely specific to that local economy, such as local government or area-specific national measures, shocks affecting all areas in the same nation, such as monetary policies non-coordinated across countries, and Europe-wide shocks such as, for example, those measures determined at the level of the European Community. We can then decompose GDP growth into three components—local, national and international. The relative variance of these three components will be determined by the regional economic structure and the nature of policy.

Europe-wide shocks, as well as the other shocks, have propagation mechanisms which, in principle, may be heterogeneous, because of the heterogeneity of local economic structures. Different regions may be affected by the same shock with different time delays or even with opposite signs. In the first case, the correlations across regions of the international components will be low, indicating asynchronous cycles. In the second case, the same correlations will be negative, which is a manifestation of asymmetric shocks.

If policies and structures were homogeneus, we would observe three characteristics. First, the international component of output growth would have a large variance relative to the total on average. Second, this would be true for all regions. Third, the international components of each region would be strongly positively correlated across regions. The first characteristic will tell us how large is the potential target of a global macro policy, while the second and the third will give us information on how regionally diversified should macro policy be.

This paper analyses these issues empirically by using European regional data and provides a comparison with the results obtained for US counties and states. The model we propose is an extension of the traditional dynamic factor model (see e.g. Sargent and Sims, 1977). We build on Forni and Reichlin (1995, 1996) to propose a simple method of estimation which is based on Law of Large Numbers results and exploits the large cross-sectional dimension of our data set.

The main advantage of our method with respect to the econometric strategies proposed elsewhere is that we can analyse a data set with high ge-

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ographical disaggregation while retaining both sophisticated dynamic modelling and parameter heterogeneity. We are not relying on arbitrary averaging either over time or across cross-sectional units, and we can do this without sacrifying simplicity.

Notice that both dynamics and cross-sectional heterogeneity are relevant information for policy purposes. The dynamic aspect because comovements at business cycle frequencies have different policy implications than comovements in the long-run. Do the variances of output growth and their cross-sectional comovements peak at business cycle frequencies, i.e. at the frequencies likely to be targeted by traditional macro-policies or else do they peak at low frequencies, potential targets of structural policies? Heterogeneity is relevant because, to understand how effective a common European policy could potentially be, we need to have information on synchronicity of regional cycles and asymmetry of shocks.

The questions which we would be asking are the following. How large is the European component of regional output fluctuations on average and is it as large as in the US? What is the dynamic shape of the European-wide cycle and does it differ from the US case? Are the European components strongly positively correlated across regions? Is there a European core of local areas which are "more European" than others? Is the core made of regions belonging to the same nation or is it independent from the political dimension? How important is the national dimension with respect to the purely local one?

2. The model

Let us denote with y_t^{ij} the growth rate of output for the i-th region of nation j, expressed in deviation from the mean. We assume

$$y_t^{ij} = E_t^{ij} + N_t^{ij} + L_t^{ij} = a^{ij}(L)e_t + b^{ij}(L)n_t^j + c^{ij}(L)l_t^{ij}, \qquad (1)$$

for $j = 1, \ldots, J$ and $i = 1, \ldots, I^j$. E_t^{ij}, N_t^{ij} and L_t^{ij} are the European component, the national component and the local component respectively. The functions $a^{ij}(L)$, $b^{ij}(L)$ and $c^{ij}(L)$ are rational functions in the lag operator L. The European shock e_t, the national shocks n_t^j and the local shocks l_t^{ij} are unobserved unit-variance white noises, mutually uncorrelated at all leads and lags.

The difference of this model with respect to the traditional dynamic factor model or index model (see Sargent and Sims 1977, Geweke 1977) is that the factor n_t^j is neither common nor idiosyncratic. It is an

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intermediate-level factor, common for regions belonging to the same country but orthogonal across countries[2].

Notice that the international components E_t^{ij}, as well as the national components of regions belonging to the same country, though driven by the same shock, have heterogeneous response functions, so that they may be both positively and negatively correlated. We do not assume restrictions on the long-run effects $a^{ij}(1)$, $b^{ij}(1)$ and $c^{ij}(1)$, so that all shocks are permanent in general, but may be transitory for particular regions.

Since the three components are mutually orthogonal, the variance of y_t^{ij} can be decomposed into the sum of the variances. The percentage of the total output variance explained by the European component measures the extent to which income fluctuations of region ij are affected by Europe-wide events. Moreover, we can distinguish between long-run and cyclical fluctuations by looking at the spectral density function. Similar considerations hold for the national and the local components.

The average across regions of the variance explained by the European component can thus be considered as a synthetic index of the importance of Europe-wide comovements in local incomes. Such an index cannot be interpreted immediately as measuring the degree of synchronization between regional GDP fluctuations, because of the heterogeneity of the response functions; in order to get a complete picture about synchronization of cycles we have to look at the cross-correlations between the European components of different areas.

3. Estimation procedure

To estimate the model we use an adapted version of the procedure proposed in Forni and Reichlin (1995, 1996), which is based on the implications of the Law of Large Numbers. To get an intuition of the basic idea, consider only regions belonging to the same country and assume that the European component is zero. Assume also for simplicity only contemporaneous responses to the shocks. Dropping the index for the nation, model (1) becomes

$$y_t^i = N_t^i + L_t^i = b^i n_t + c^i l_t^i, \quad i = 1, \dots, I.$$

Now consider the average $\bar{y}_t = \bar{b} n_t + \sum c^i l_t^i / I$. If I is large, the local component $\sum c^i l_t^i / I$ should be small in variance as compared with the common

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[2] Notice that (1) can be interpreted as a $J + 1$ common factors model with the regional responses restricted in such a way that regions in nation j react only to the j-th factor. This restriction cannot be tested formally, since the unrestricted version of the model cannot be estimated unless there is a large number of time observations.

one, owing to the orthogonality of the local shocks.[3] Hence $\bar{y}_t$ is almost collinear to the national shock. But this means that the unobserved common factor becomes observable, so that we can simply substitute $\bar{y}_t$ for n_t and estimate the model by applying OLS equation by equation (clearly we have an estimation bias, which will be smaller the smaller is the percentage of the variance of $\bar{y}_t$ explained by the local idiosyncratic component).

The same argument holds when considering for instance a weighted average of regions, rather than the simple average $\bar{y}_t$.[4] Hence we have different candidates to use as regressors in OLS estimation. Obviously some of them will be better than others, depending on the percentage of idiosyncratic variance surviving aggregation. Are there optimal weights, i.e. weights minimizing the variance explained by the local components? The problem of finding the optimal regressor is particularly relevant when the number of cross-sectional observations is not very large, as it is the case with the European part of our data set (see Section 4).

In Appendix 1 we show that the optimal regressor, within the class of all linear combinations of the y_t^i's, is obtained by using as coefficients the entries of the eigenvector corresponding to the larger eigenvalue of the matrix $\Sigma^{-1}\Gamma$, where Σ is the covariance matrix of the local components L_t^i and Γ is the covariance matrix of the variables y_t^i. Moreover, the reciprocal of the larger eigenvalue is an estimate of the percentage of the idiosyncratic variance in this linear combination.[5]

Since the matrix Σ cannot be estimated directly from the data, we followed a two-stage procedure. In the first stage we assumed proportionality of Σ and the diagonal matrix having the same main diagonal as Γ. The weights obtained under this hypothesis were then used for a preliminary estimate of the model and the diagonal entries of Σ (the non-diagonal entries were set equal to zero according to the orthogonality assumption). In the second stage this estimate of Σ was used to get the final regressor.[6]

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[3] For a discussion on this point see Granger (1987). In the static case of the example, the Large-Numbers argument can easily be made rigorous by assuming a countable infinity of regions satisfying $b^i > \mu > 0$, $c^i < \nu$, and taking the limit for $I \to \infty$. For a generalization to the dynamic case, see Forni and Lippi (1997), Chapter 1.

[4] A sufficient condition is that weights satisfy upper bounds and positive lower bounds before normalization.

[5] Connors and Korajczyk (1986), building on a result by Chamberlain (1983), propose to use the first k principal components as regressors for the estimation of a static k factor CAPM model. The result in Appendix 1 shows that these regressors are not optimal.

[6] An alternative strategy is to continue iteration until convergence of regressors.

Now let us come back to the general case of the three-level factor model (1). A fully detailed description of our estimation procedure is reported in Appendix 2. Here we give the main lines. First, by averaging across regions belonging to the same country we obtained J national aggregates with no local component. Second, by averaging across these national aggregates we obtained a European aggregate with neither national nor local component. Finally, the model was estimated by regressing region ij on both the European and the j-th national aggregate.

4. The Data

The quality of data on regional output is poorer in Europe than in the US: the sample period is shorter and the level of geographical disaggregation not as fine; moreover, the level of disaggregation at which data are available is not homogenous across different European countries. We tried to cope with this problem by constructing two different data sets.

For the first one we aimed at the longest possible series, by reducing the number of countries and local disaggregation and merging heterogeneous data. The sources are Regio Eurostat and Eurostat, regional yearbook, 1983. We selected from the Regio data set observations on GDP in national currency from 1977 to 1993, for 82 regions.[7] Then we deflated by the national consumer price indexes published by the Eurostat. For the period 1973-77 and for the same regions, we used data on gross value added at market prices in national currency published in the regional yearbook of 1983 and deflated in the same way.

In the second data set, we tried to include as many nations as possible. Data on GDP in national currency are available from Regio for 138 regions (including Greece, Spain and Portugal) from 1980 to 1993.[8] To compute real GDP we deflated by the national consumption price index as before.

Our simulations however indicates that the iterative procedure, while being computationally more cumbersome, do not reduce the idiosyncratic variance in general.

[7] These are the 11 NUTS1 regions for West Germany (the west landers); the 11 NUTS1 UK regions; 21 NUTS2 Franch regions (Corsica and the colonies are excluded); the 20 NUTS2 Italian regions; the 9 NUTS2 Belgian regions; the 10 Duch regions obtained by taking NUTS1 for Est-Netherlands and NUTS2 for other areas. West Berlin was estimated for 1992 and 1993 by applying the ratio West Berlin/Berlin of 1991 to the Berlin data.

[8] In this data set the disaggregation level is NUTS2 for all nations but the UK and Est-Netherlands. West Germany: 31 regions; UK: 11 regions; France: 21 regions (Corsica and the colonies are excluded); Italy: 20 regions; Belgium: 9 regions; Netherlands: 10 regions; Greece: 13 regions; Spain: 18 regions; Portugal: 5 regions (Acores and Madeira are excluded). West Berlin was estimated as explained in the previous note.

The source of US data is BEA, Regional Economic Information System. Total personal income by county, from 1969 to 1993, was deflated by the US implicit GDP deflator. Alaska and District of Columbia were excluded. In the final data set we have 48 states and 3170 counties. Notice that the geographical disaggregation here is much finer than European NUTS2; a better comparison between Europe and the US would have required NUTS3 data, which are unfortunately not available.

5. Empirical Results

What follows is a description of our results, organized by five main economic questions.

• How large is the European component of regional output as compared with the other components and with the US component of county output?

Table 1 shows the average across regions of the percentage of variance explained by the three components, for the two European data sets. In both of them the European component is the largest one and accounts for almost 50% of the variance of output growth. United Kingdom and Greece are remarkable exceptions: here the most important component is the national one, whereas the European component is rather small.[9]

When passing from the first to the second and more recent data set, the size of the international component decreases in Germany, but increases in France and, most notably, in Italy. The average figure grows slightly, despite the inclusion of Greece and Portugal. This suggests that the importance of the European component is increasing over time.[10]

Table 2 shows the variance decomposition for US. The comparison with Europe is striking: the average size of the US-wide component (last line) is similar to that of the European component of Table 1. If we exclude Greece and the UK, the degree of economic integration of US counties appears smaller than that of European regions.[11]

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[9] When these countries are excluded, the average becomes about 52% in the first data set and 55% in the second one.

[10] Notice also that the reduction of the European component of Netherland is due to an exceptional event: a very large local shock in one region at the end of the eighties, which has a smaller effect in the longer series.

[11] When evaluating this result we should remind that the geographical partition is finer in the US data set. If US regions were larger, the local component would likely be smaller and the US-wide component would be greater.

Table 1. Percentage of variance explained by the European component, the national component and the local component (national and European averages)

Country	European Component	National Component	Local Component	Total
1973–1993				
Germany	67.2	20.7	12.1	100
United Kingdom	14.0	59.7	26.3	100
France	52.4	12.4	35.2	100
Italy	36.6	27.4	36.0	100
Belgium	51.8	26.9	21.3	100
Netherland	64.0	6.1	29.9	100
Total Average	46.8	24.3	28.9	100
1980–1993				
Germany	61.2	29.6	9.2	100
United Kingdom	17.1	75.5	7.4	100
France	58.5	20.0	21.4	100
Italy	49.5	31.7	18.8	100
Belgium	54.5	35.9	9.6	100
Netherland	55.1	21.5	23.4	100
Greece	14.7	64.1	21.1	100
Spain	50.9	28.5	20.6	100
Portugal	34.0	50.5	15.5	100
Total Average	48.1	35.6	16.2	100

• How large is the national component of output fluctuations in Europe?

From Table 1 we can see that the national component in Europe accounts for only about 25% on average. This indicates that nations count little, not more than states in the US (Table 2). There is, however, some diversity within Europe, with the UK, Greece and Portugal showing a national component above 50 %. Notice also that the average national component in the second data set increases to 35%, partly due to the inclusion of Greece and Portugal. This increase, however, is at the expense of the local component, while the European-wide component stays roughly constant.

• What is the dynamic profile of national output fluctuations in Europe and how does it compare with the dynamic profile of output in US states?

Figure 1 shows the average spectra of the three components for Europe (first data set) and the US. Although, as we have seen, in terms of total variance the two cases are similar, the dynamic profiles of the components differ a lot. The main difference between Europe and the US is that in Europe there is no typical business cycle, neither global, national or regional, while

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Table 2. Percentage of variance explained by the US-wide component, the state component and the local component

State	US Component	State Component	Local Component	Total
Maine	57.7	8.8	33.5	100
Massachusetts	54.2	17.6	28.2	100
New Hampshire	55.6	12.0	32.4	100
Rhode Island	52.3	15.5	32.2	100
Vermont	23.3	24.4	52.3	100
Delaware	47.9	44.6	7.6	100
Maryland	55.9	28.9	15.2	100
New Jersey	46.0	22.0	32.0	100
New York	51.7	13.6	34.7	100
Pennsylvania	31.4	34.4	34.2	100
Illinois	38.8	35.7	25.5	100
Indiana	61.8	20.9	17.4	100
Michigan	49.5	33.4	17.1	100
Ohio	26.4	25.1	48.5	100
Wisconsin	39.7	28.2	32.0	100
Iowa	23.9	40.6	35.5	100
Kansas	50.7	28.2	21.1	100
Minnesota	57.6	19.2	23.2	100
Missouri	41.8	47.4	10.8	100
Nebraska	61.7	12.3	26.1	100
North Dakota	54.1	23.3	22.6	100
South Dakota	56.6	16.0	27.5	100
Alabama	54.3	20.7	25.0	100
Arkansas	35.3	32.8	31.9	100
Florida	30.2	35.0	34.8	100
Georgia	31.3	30.4	38.4	100
Kentucky	55.8	35.9	82.4	100
Louisiana	58.2	27.9	14.0	100
Mississippi	32.7	18.5	48.8	100
North Carolina	47.2	29.0	23.8	100
South Carolina	62.1	10.5	27.5	100
Tennessee	28.4	51.5	20.1	100
Virginia	64.5	11.3	24.2	100
West Virginia	16.7	42.9	40.4	100
Arizona	56.6	18.1	25.4	100
New Mexico	66.3	10.1	23.5	100
Oklahoma	53.7	35.4	11.0	100
Texas	59.1	12.5	28.4	100
Colorado	38.1	38.3	23.6	100
Idaho	62.8	9.2	27.9	100
Montana	23.8	25.6	50.6	100
Utah	40.5	16.5	43.0	100
Wyoming	63.4	21.1	15.4	100
California	58.3	10.4	31.3	100
Hawaii	40.1	20.1	39.8	100
Nevada	33.9	34.0	32.1	100
Oregon	60.8	14.3	24.9	100
Washington	20.1	43.7	36.1	100
Total Average	45.5	23.2	31.3	100

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Figure 1. The spectral shape of the three components for Europe and the US

Horizontal axis: frequency; vertical axis: spectral density
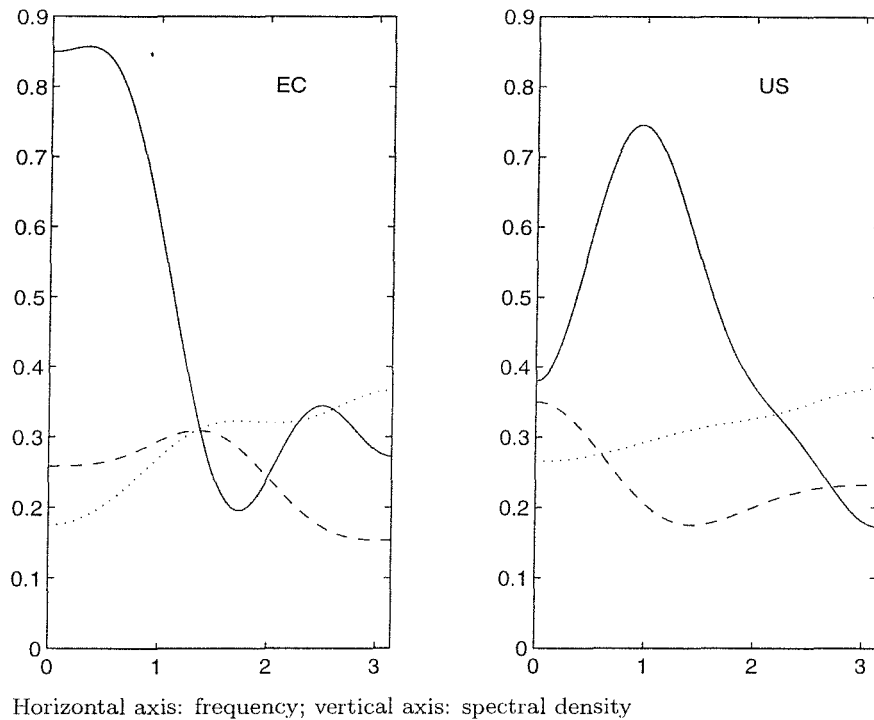
in the US there is a US-wide cycle which peaks at a period of around six years.

This result indicates that Europe-wide shocks are more persistent than US-wide shocks which opens the interesting question of what is the relation between the institutional and market structures and the degree of persistence of shocks. A conjecture is that more flexible markets imply a faster trasmission of shocks throughout the economy so that a larger proportion of the total variance is at high frequencies. In this view, and given results from the previous paragraph, the difference between the US and Europe is not so much in the level of integration, but in the structure of markets. This conjecture is reinforced by Figure 2 which shows results for the six European countries separately. The UK, which is the European country more similar to the US in terms of its market and institutional structure, beside having a large national component, is the only case to show a typical business cycle shape. The other countries confirm the aggregate result of

Figure 2. The spectral shape of the three components for six European countries
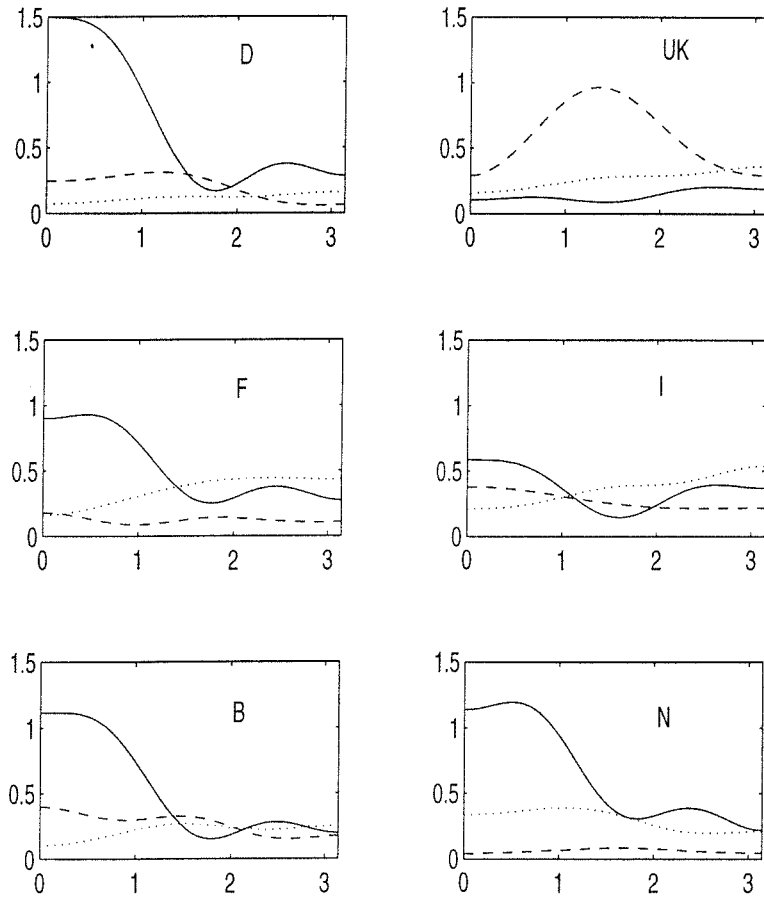
Horizontal axis: frequency; vertical axis: spectral density

a large European wide component with most of the variance concentrated at low frequency.

Figure 3 shows average spectra for six large US states and confirms what seen in the aggregate for the US, i.e. that the typical business cycle shape is a characteristic of the US-wide component. In other words, the business cycle is generated by national shocks.

- How synchronized are European fluctuations?

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Figure 3. The spectral shape of the three components for six US states
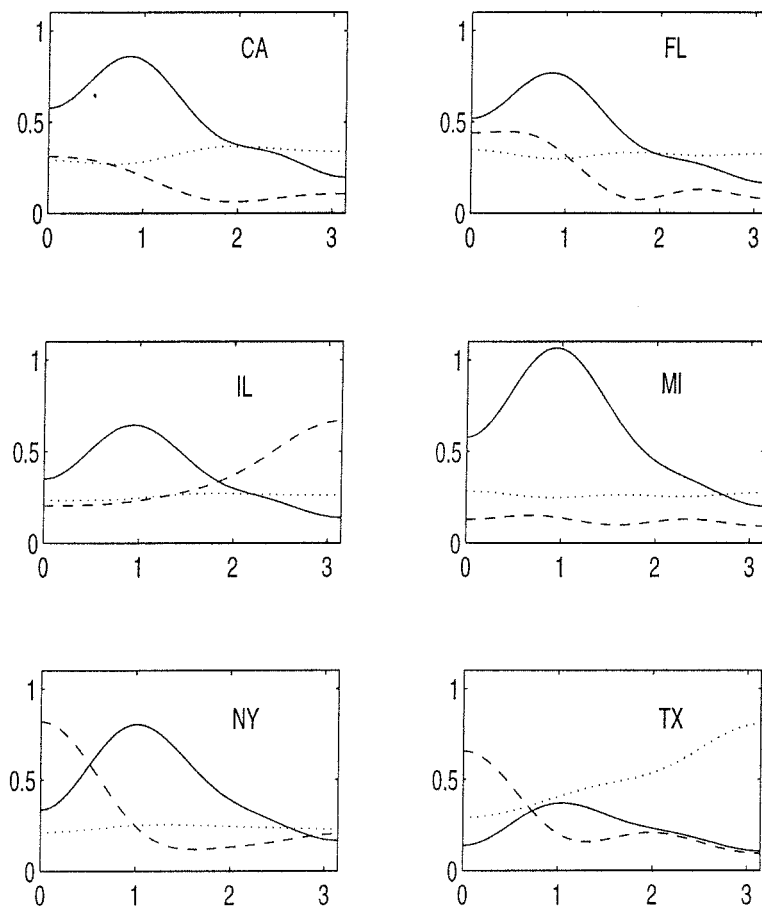
Horizontal axis: frequency; vertical axis: spectral density

To complete information on synchronization we now want to know whether Europe-wide fluctuations are positively correlated across regions.

To this end, we have computed the matrix of contemporaneous correlations of the European components of all regions for the six countries of the first data set (Figure 4). Color intensity describes the correlation value and it ranges between white (correlation equal to plus one) and black (correlation equal to minus one). The gray shade in the upper triangle corresponds to the case of zero correlation. The overwhelming prevalence of white cells shows that there is a lot of large positive correlation amongst

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Figure 4. Contemporaneous correlations of the European components
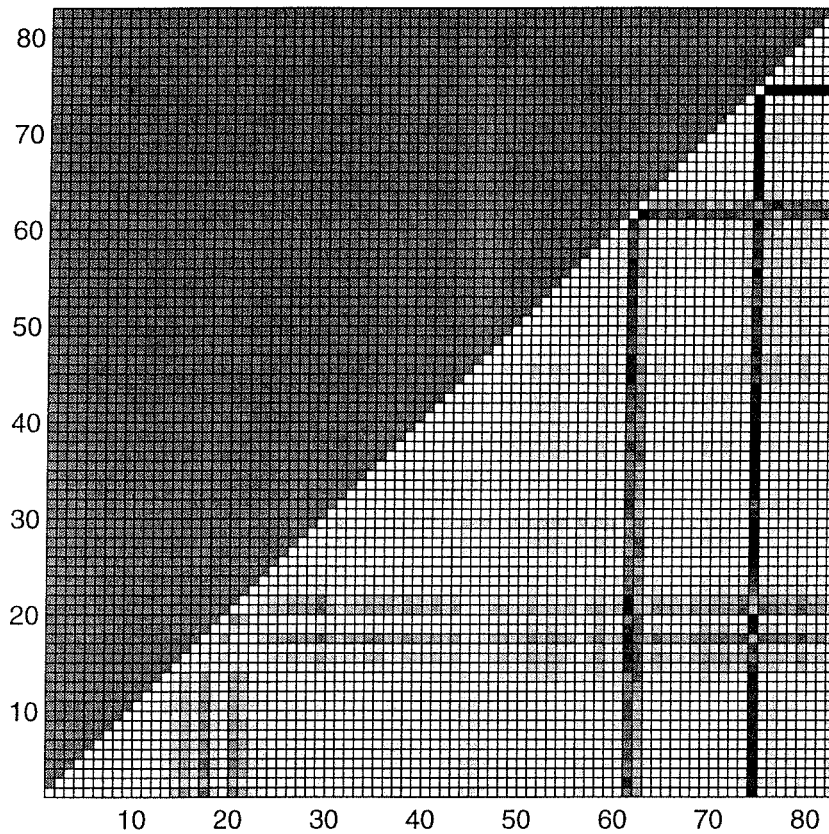
the European components of different regions. This result, coupled with the large average size of the European-wide component, indicates a high degree of synchronization of output fluctuations in Europe.

- What is the European core?

Here we want to ask the question of whether regions which are "more European" (larger relative variance of the European component) belong to a particular geographical area. In order to answer this question we now shift to the second data set, which is the larger one. Figure 5 reports the geographical distribution of variance ratios between European-wide components and total variance.

Light gray indicates a small European component while dark gray indicates a large European component. The Figure shows that a core made by

Figure 5. Percentage of output variance explained by the European component
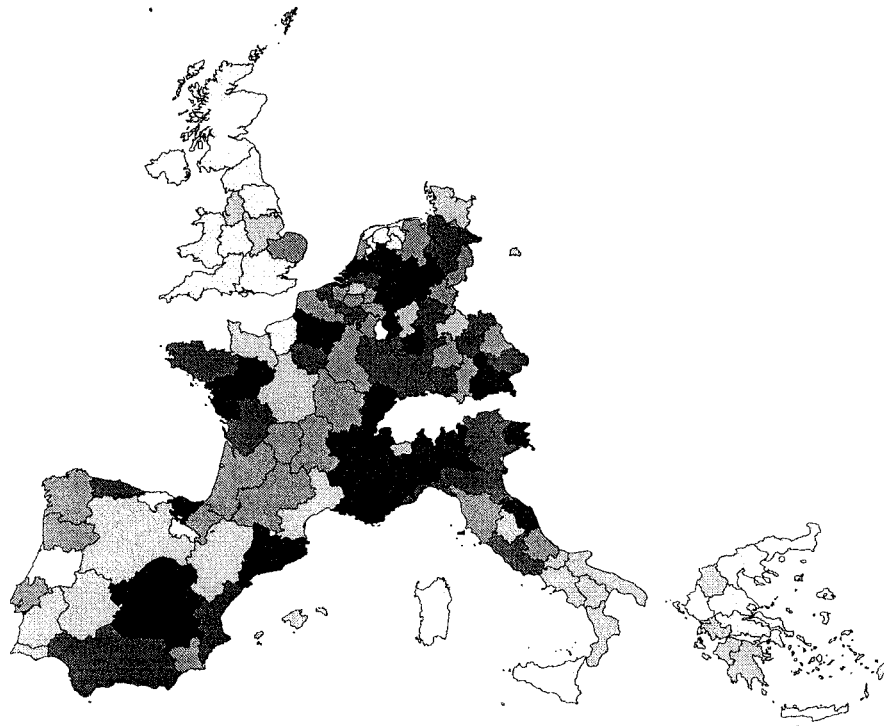
Dark regions have a large European component
The limits for colour changes are 0.23, 0.42, 0.58 and 0.70

the key countries France, Germany and the Benelux does not exist. Dark and light spots are sparse, indicating that almost all countries are partly in and partly out. The only exceptions are Greece and the UK, which are clearly out.

6. Related empirical literature

How do our results compare with the existing empirical literature? Direct comparison of results is not easy since the approaches used to study asym-

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metries of shocks and degree of synchronization of fluctuations are very diverse. Moreover, our study is unique in the literature since it analyses the whole pattern of regional dynamic interaction.

Earlier studies have measured the degree of idiosyncracy of output or employment movements by looking at regional and country dispersion of output growth rates (De Grauwe and Venhaverbeke, 1991), at dispersion of shocks across European nations and US states (Bayoumi and Eichengreen, 1993), at the weight of regional specific variance in total employment variation for European regions and states (Decressin and Fatas, 1995), at country specific versus industry-specific components in output variations (Helg et al., 1995), at the weight of a European-wide common component in regional unemployment variation (Viñals and Jimeno, 1996), at the degree of "dynamic homogeneity" within countries (Fuss, 1996).

The findings that supply shocks are less correlated across European countries than across US states (Bayoumi and Eichengreen, 1993), output variance is mostly explained by an industry-specific component than by a country specific one (Helg et al., 1995), employment dynamics at the regional (state) level is more idiosyncratic in Europe than in the US (Decressin and Fatas, 1995) seem to lead to the conclusion that Europe is more heterogenous than the US, both in the sense of a higher degree of idiosyncratic volatility and asymmetry of shocks. Bayoumi and Eichengreen, 1993 and Helg et al, 1995, however, identify a European core formed by France, Germany and the Benelux which correspond more closely to the US case.

Our results partly contradict these findings since Europe emerges as being very close to the US in what concerns the relative weights of the components. Viñals and Jimeno, 1996, found a similar result for unemployment data. On asymmetries, we can confirm the presence of a core, but our analysis of disaggregated dynamics shows that the core is made of regions belonging to different nations. This result, which has obvious policy implications, highlights the importance of a disaggregate analysis.

7. Methodological observations

Why proposing a dynamic factor model instead of more popular options such as VARs or panel data models? VARs are not well suited for modelling large cross-sections of time series, since the number of parameters to estimate is typically too large with respect to the number of available observations over time. In order to estimate a VAR we need a huge number of *a priori* restrictions, which usually economic theory cannot provide. On the other hand, panel data tecniques require two or more observed variables

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with a clearcut *a priori* distinction between 'dependent' and 'independent' variables. By contrast, unobserved factor models can be used also when we have only one variable and can provide a parsimonious representation of the dynamic relations among cross-sectional units. The interest of the methodology proposed in Forni and Reichlin (1995, 1996) and adapted here in order to take into account both intermediate shocks and optimal weights is that it shows that, contrary to the common wisdom (see e.g. Sargent and Quah, 1994, and the related discussion) estimation of dynamic factor models is simpler rather than more difficult when large cross-sections are involved.

As explained in Section 2, the percentage of variance accounted for by the common component is a measure of comovement which emerges naturally from the common-idiosyncratic representation of factor models. This notion of comovement is different from what has been proposed by the literature on common trends (Stock and Watson, 1988), common features (Engle and Kozicky, 1993) and common cycles (Vahid and Engle, 1993). The example below may help to clarify this point. Consider the following very simple dynamic specification of model (1), with only two regions and zero national component:

$$\Delta y_t^1 = ae_t + (1 - bL)l_t^1$$
$$\Delta y_t^2 = ae_t + (1 - bL)l_t^2.$$

The two regions do not have either common trends or common cycles, since the first difference of the linear combination $z_t = y_t^1 - \alpha y_t^2$ is

$$\Delta z_t = (1 - \alpha)ae_t + (1 - bL)(l_t^1 - \alpha l_t^2),$$

which do not have a unit root in the Wold representation (unless $b = 1$ and $\alpha = 1$) and is not serially uncorrelated (unless $b = 0$). Nevertheless, if a is large, the two processes comove strongly according to the measure proposed above. Indeed, their correlation may be arbitrarily close to unity (both at long-run and cyclical frequencies), depending on the size of a.

The example also shows that the concepts of common trends and common cycles are not particularly useful in order to get a measure of comovement between the GDP growth rates of a set of regions or nations. The main reason is that cointegration tests, as well as tests on common features, can only provide a binary result: either the regions comove perfectly or not. For instance, cointegration gives a unambiguous result on the strenght of long-run comovements only if all regions are pairwise cointegrated. This

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condition is rather strong; in practice, it will be never satisfied by a large set of regions, leaving us with no useful indications about the relevant problem.

8. Summary and Conclusion

This paper proposes a method to study synchronization of output fluctuations at different levels of aggregation and compares estimates for European regions and US counties. For all regions we estimate the relative variance of a dynamic component generated by a European-wide shock, its dynamic profile and the pattern of its cross-regional correlations.

We find that, unlike what claimed by previous studies, Europe is as much integrated as the USA. This is measured by the relative variance of a European-wide (US-wide) component with respect to the total variance of output growth. The disaggregated analysis confirms this fact for all countries exept for the UK and Greece. We also find a European core formed of regions where the weight of the European comonent is very large. The core, however, is not determined by the national dimension. In general, we find that the national dimension in Europe is not very important: what matters is the European component and a purely local component. The policy implication we draw from these facts is that a common European policy can target on average about 50% of output variation; if this is complemented by regional policies, we can conclude that around 75% of output could be smoothed without relying on national policies.

The study of the dynamic profile of the components, however, shows that Europe, unlike the US, has no traditional business cycle. Shocks are very persistent and the bulk of the variance is in the long-run. A conjecture, which is supported by the fact that the UK is the only European country with similar dynamic behaviour than the US, is that more flexible markets imply a faster trasmission of shocks throughout the economy and therefore less persistence.

The ensemble of the results indicate that the difference between the US and Europe is not so much in the level of integration, but in the structure of markets. Common European policy designed to target output is potentially effective provided that it aims at the long-run. Macroeconomic policies should be complemented by market reforms.

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APPENDIX 1

The optimal weighting procedure

Let us focus the attention on a single nation, so that, dropping the index j, model (1) becomes

$$y_t^i = E_t^i + N_t^i + L_t^i = C_t^i + L_t^i,$$

where the common component C_t^i is orthogonal to the local idiosyncratic component L_t^i. Now let us indicate with Σ the covariance matrix of the vector of the idiosyncratic components $L_t = (\, L_t^1 \quad \cdots \quad L_t^I \,)$ and with Γ the covariance matrix of the vector of the variables $Y_t = (\, y_t^1 \quad \cdots \quad y_t^I \,)$. We are looking for the I-dimensional vector w minimizing the ratio $\mathrm{var}(w'L_t)/\mathrm{var}(w'Y_t)$, or, equivalently, the function

$$\log\big(w'\Gamma w\big) - \log\big(w'\Sigma w\big).$$

This function is homogeneous of degree zero in w, so that it reaches an interior maximum in $R^I - 0$ (0 being the null vector) on a ray through the origin satisfying the first order condition

$$\frac{2\Gamma w}{w'\Gamma w} - \frac{2\Sigma w}{w'\Sigma w} = 0.$$

Assuming the invertibility of Σ, this is equivalent to

$$\Sigma^{-1}\Gamma w = \frac{w'\Gamma w}{w'\Sigma w} w.$$

Imposing the latter condition, in turn, is equivalent to imposing

$$\Sigma^{-1}\Gamma w = \lambda w \tag{2}$$

for some scalar λ and some $w \neq 0$. This is because a couple λ, w satisfying (2) must fulfill $\lambda = w'\Gamma w/w'\Sigma w$.

Hence the first order condition is satisfied by, and only by, the eigenvectors of $\Sigma^{-1}\Gamma$; moreover, the eigenvalues λ are the reciprocals of the objective function $\mathrm{var}(w'L_t)/\mathrm{var}(w'Y_t)$ evaluated at the corresponding eigenvectors. It follows that the solution of the above programming problem must be given by the eigenvector corresponding to the maximum latent root of $\Sigma^{-1}\Gamma$.

If Σ is diagonal, as it is assumed here, equation (2) have a simple interpretation. Since

$$\Gamma w = \begin{pmatrix} \mathrm{cov}(y_t^1, w'y_t) \\ \vdots \\ \mathrm{cov}(y_t^I, w'y_t) \end{pmatrix},$$

equation (2) reduces to

$$\frac{\mathrm{cov}(y_t^i, w'y_t)}{\mathrm{var}(L_t^i)} = \lambda w^i, \qquad i = 1, \dots, I$$

i.e. the weight of region i must be larger, the larger is the covariance of region i with the aggregate and the smaller is the variance of the idiosyncratic component.

Notice also that in the particular case of perfectly correlated common components, i.e. $C_t^i = a^i C_t$, w is proportional to $a^i/\mathrm{var}(L_t^i)$, which clearly shows that 'weights' are not necessarily positive.

APPENDIX 2

The estimation procedure

The complete estimation procedure is in five steps.

Step 1 We washed out the local components by computing, for each nation j, the linear combination

$$y_t^j = \sum_{i=1}^{I^j} w^{ij} y_t^{ij},$$

where the coefficients w^{ij} are those minimising the ratio of the variance of the local component over the total, as explained above. Hence

$$y_t^j \approx a^j(L)e_t + b^j(L)n_t^j = E_t^j + N_t^j,$$

where $a^j(L) = \sum_{i=1}^{I^j} w^{ij} a^{ij}(L)$ and $b^j(L) = \sum_{i=1}^{I^j} w^{ij} b^{ij}(L)$.

Step 2 We eliminated the national components by computing the linear combination

$$y_t = \sum_{j=1}^{J} w^j y_t^j,$$

where the w^j's are chosen again to minimize the ratio between the non-common variance to the variance of y_t. Then

$$y_t \approx a(L)e_t,$$

where $a(L) = \sum_{j=1}^{J} w^j a^j(L)$.

Step 3 Assuming equality in the above relation, along with invertibility of $a(L)$, we can write e_t, and therefore E_t^j, as a linear combination of the present and the past of y_t, so that

$$y_t^j = \alpha^j(L)y_t + N_t^j.$$

We estimated the above equations by OLS, with $\alpha^j(L)$ specified as a second order polynomial.[12] These auxiliary regressions are needed in Step 5 in order to disentangle the national and the European component.

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[12] In principle $a(L)$ can be non-invertible toward the past. In order to allow for roots smaller than unity in modulus we have to specify $\alpha^j(L)$ as a bilateral operator (for a discussion on this point see Forni and Reichlin 1996). We tried different specifications for the $\alpha^j(L)$, including both leads and lags of y_t, but we found that a two-lags specification could not be rejected by the F-test.

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Step 4 A similar reasoning leads to the relation

$$y_t^{ij} = \alpha^{ij}(L)y_t + \beta^{ij}(L)y_t^j + L_t^{ij}.$$

We estimated the above regression equations by OLS. Also in this case we found that a second-order specification for both $\alpha^{ij}(L)$ and $\beta^{ij}(L)$ was good. In this way we got an estimate for the local components L_t^{ij}.

Step 5 By substituting for y_t^j in the above relations we see that

$$E_t^{ij} = \alpha^{ij}(L)y_t + \beta^{ij}(L)\alpha^j(L)y_t$$
$$N_t^{ij} = \beta^{ij}(L)N_t^j = y_t^{ij} - E_t^{ij} - L_t^{ij}.$$

This provides estimates for E_t^{ij} and N_t^{ij}.

A complete estimation of the parameters of (1) is beyond our aims. However, estimates for $a^{ij}(L)$ and $b^{ij}(L)$ could in principle be obtained by estimating $a(L)$ (by univariate ARMA modelling of y_t) and $b^j(L)$ (by univariate modelling of the N_t^j's) and using the relations

$$a^{ij}(L) = \left(\alpha^{ij}(L) + \beta^{ij}(L)\alpha^j(L)\right)a(L)$$
$$b^{ij}(L) = \beta^{ij}(L)b^j(L).$$

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