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Federal policies and local economies: Europe and the US

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Abstract

This paper establishes stylized facts on regional output fluctuations in Europe and the US. Moreover, it proposes a measure of the potential output target of the future European central bank, estimates the potential variance stabilization of a fiscal federation and constructs a regional map of the potential beneficiaries of monetary and fiscal federal policies. The econometric model is an extention of the dynamic factor model à la Sargent and Sims (1977. In: Sims, C.A. (Ed.), New Methods in Business Research. Federal Reserve Bank of Minneapolis) where we introduce 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 (1996. Empirical Economics, Long-Run Economic Growth (special issue) 21 (1996) 27–42. Review of Economic Studies 65 (1998) 453–473) to propose an estimation method which exploits the large cross-sectional dimension of our data set. Our analysis shows that (i) Europe has a level of integration similar to that of the US and that national shocks are not a sizeable source of fluctuations: around 75% of output variance is explained by global and purely local dynamics; (ii) Europe, unlike the US, has no traditional business cycle; (iii) the core of the most integrated regions in Europe does not have national boundaries; (iv) the future European Central Bank has a potential stabilization target of about 18% of total output fluctuations; (v) a fiscal federation, if implemented, could have a smoothing effect on output in

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addition to what done by national fiscal policy, which accounts also for about 18% of total output fluctuations.

JEL classification: C51; E32; O30

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1. Introduction

This paper is about the potential output stabilization effects of European federal policies – the established monetary union and a virtual fiscal federation. We measure potential costs and benefits from the point of view of the regions of the Union rather than its nations and we distinguish between stabilization of income over time and over the cross-section of regions.

Our work is motivated by two observations.

First, intertemporal stabilization policies in a monetary union aim at an aggregate cycle, but, if the degree of 'commonality' of regional fluctuations with the aggregate cycle is low, the European Central Bank will not be very effective at reducing regional fluctuations. This is an observation which was at the center of the debate on the optimality of the EMU as a currency area. Here we propose a simple way to measure the potential output stabilization target of the new central bank and the cost of the loss in flexibility of federal monetary policies as compared with the old national system.

Second, fiscal policy at a federal level could do more than intertemporal stabilization of common fluctuations since it can act as a provider of insurance: taxes and transfers, by reallocating income across regions, have automatic stabilization effects. Since only idiosyncratic risk can be insured, national fiscal systems can only smooth region specific shocks whereas in a European wide federation even national shocks can be regarded as idiosyncratic. Whether a European federation would be more effective than a nation in providing insurance depends on how large is the national component of output volatility. Here we propose an estimate for this quantity.

To produce these numbers we develop a model which is suitable to analyze the entire dynamic profile of regional correlation of output. The model we propose is an extension of the traditional dynamic factor model à la Sargent and Sims. We build on Forni and Reichlin (1996,1998a) and Forni et al. (1999) to propose a simple method of estimation for panels of time series. In each region, total output growth is seen as generated by shocks which are purely specific to that local economy, shocks affecting all areas in the same nation and Europe-wide shocks. Correspondingly, GDP growth is decomposed into a local, a national, and a European component.

Our estimates of potential regional targets of common European monetary policy and fiscal insurance will give, as a byproduct, the geographical distribution of potential winners from these policies. This distribution will allow us to ask the political economy question of whether policy makers, in the future, are likely to face national constituencies or coalitions of regions belonging to different countries.

The paper is organized as follows. We first describe the model, the estimation procedure and the data. We then report empirical results and discuss policy implications. Next come a methodological section and summary and conclusions.

2. The model

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

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

for $j = 1, \ldots, J$ and $i = 1, \ldots, I^j$. E_t^{ij}, N_t^{ij} and $\mathscr{L}_t^{ij}$ are the European component, the national component and the local component, respectively. $a^{ij}(L)$, $b^{ij}(L)$ and $c^{ij}(L)$ are 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 intermediate-level factor, common for regions belonging to the same country but orthogonal across countries.[1]

Some comments may be useful in order to avoid possible misunderstandings about the model.

First, the local shock of a region cannot affect other regions and the national shock of a nation cannot affect other nations. Therefore, if a shock originated in a large country, e.g. a large policy shock in Germany, has non-negligible effects on other countries, it must be considered as a European shock, not as a German national shock. Similarly, a shock should be labeled as national or local depending on its effects rather than on its origin.

Second, the European 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 to allow for possible heterogeneity of local

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[1] Note 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 jth factor.

economic structures. Different regions may be affected by the same shock with different time delays or even with opposite signs. In such cases, the correlation across regions of the European components may be low, or even negative. In the same way, 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.

Finally, 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 (1996, 1998a), 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

$$y_t^i = N_t^i + \mathscr{L}_t^i = b^i n_t + c^i l_t^i, \quad i = 1, \dots, I.$$

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 one, owing to the orthogonality of the local shocks. 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 $\bar{y}_t$. 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 percentage of the variance of the average explained by the local shocks? The problem of finding the optimal regressor is particularly relevant when the number of cross-sectional observations is not very large, as is the case with the European part of our data set (see Section 4).

In Appendix A 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 $\mathscr{L}_t^i$ 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 and can therefore be used for diagnostic checking.

Since the matrix Σ cannot be estimated directly from the data, we followed a two-stage procedure. In the first stage we assumed Σ being equal to the diagonal matrix having the same main diagonal as Γ, multiplied by a scalar. The implied eigenvector was 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.[2]

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 B. 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 jth national aggregate.

With this procedure we estimate, for each region, the three components – European, national and local. All the variances and spectral densities described in Section 5 are computed from the estimated components.

It should be observed that both the estimation procedure proposed in Forni and Reichlin (1996, 1998a) and the modified version used here raise problems concerning consistency of the estimates. Appendix C discusses consistency and other methodological issues, such as the extension of model (1) to the case of more common factors, and provides further references.

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[2] An alternative procedure would be to iterate estimation until Σ converges. We tried this procedure, but we got a slightly worst estimation of the percentage of the idiosyncratic variance in the aggregate. We also tried some simulations suggesting that iteration, while being computationally more cumbersome, gives little or no improvement.

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.[3] Then we deflated by the national consumer price indexes published by the Eurostat. For the period 1973–1977 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. Groningen (Netherlands) experienced an exceptionally large (idiosyncratic) income variation during the eighties; we excluded this outlier from the data set when computing average variances and spectra.

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.[4] To compute real GDP we deflated by the national consumption price index as before.

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 Hawaii were excluded because of missing data, the District of Columbia because it contains one region only. In the final data set we have 48 states and 3075 counties. Notice that the geographical disaggregation here is much finer than European NUTS2; a better comparison would have required NUTS3 data, which unfortunately are not available. To cope with this, we distinguished between large, medium and small areas. Large areas are defined as those having income greater than 30 billion of US dollars in 1993; medium areas have more than one billion but less than 30 billion dollars; small

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[3] These are the 11 NUTS1 regions for West Germany (the west landers); the 11 NUTS1 UK regions; 21 NUTS2 French regions (Corsica and the colonies are excluded); the 20 NUTS2 Italian regions; the 9 NUTS2 Belgian regions; the 10 Dutch 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.

[4] 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 (Azores and Madeira are excluded). West Berlin was estimated as explained in the previous note.

areas have less than one billion dollars. Small areas, while being the majority of US areas, are absent in Europe.

5. Empirical results

5.1. The basic facts

Table 1 shows national averages across regions of the percentage of variance explained by the three components for the two European data sets.

For most countries, for both data sets and for the two classes of dimension considered, the common European component explains the bulk of output variance. Exceptions are Greece, Portugal and the UK which show a large nation-specific dynamics, well above 50%.[5] If we exclude the UK, in the first data set the European component is 52.9% for large regions and 46.5% for medium sized regions. Similar figures are obtained for the second data set when Greece, Portugal and the UK are excluded. Note also the enormous variance of Greece and Portugal in the fifth column of Table 1.

In the last column of Table 1 we report the optimal weights computed by our estimation procedure and placed on the national aggregates in order to compute the European common factor. Note that, in both data sets the German aggregate[6] does not have the larger weight, but only the third, coming after both France and the Netherlands. This is most likely due to the fact that Germany leads the other European economies and the phase shift reduces the contemporaneous correlation. Note also that Greece, Portugal and the UK have very small weights.

Table 2 shows the variance decomposition for the US. The comparison with Europe is striking: the average size of the US-wide component is similar to that of the European component of Table 1 when the UK is excluded (first data set) and when Greece, Portugal and the UK are excluded (second data set). Moreover, the US state component seems to be of the same order of magnitude than the national component in Europe.

Overall, these results indicate that pre-monetary union Europe was already very much ‘integrated’, with a small component of output volatility explained by national shocks. Clearly, some caution is suggested by the limitations of the data set already commented in the previous section.

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[5] For the 1973–1993 period, Italy shows an idiosyncratic component which is larger than the European. However, when only data from 1980 are considered, Italy is aligned with the continental model.

[6] We recall that the German aggregate is not the German GDP, but the optimally weighted average of the GDPs of the German regions.

Table 1
Average variance explained by the European component, the national component and the local component

Country	σ_E^2/σ_y^2 $\times 100$	σ_N^2/σ_y^2 $\times 100$	$\sigma_{\mathscr{L}}^2/\sigma_y^2$ $\times 100$	Total	σ_y^2 $\times 1000$	No. of Regions	Optimal weights
1973–1993							
Germany	65.8	21.7	12.5	100	0.83	11	0.21
United Kingdom	13.5	59.2	28.3	100	1.23	11	0.01
France	50.0	12.9	37.1	100	1.15	21	0.23
Italy	33.5	27.2	39.3	100	1.19	20	0.13
Belgium	51.7	24.9	23.5	100	1.04	10	0.14
Netherlands[a]	61.4	6.3	32.3	100	1.77	9	0.29
Large regions	**44.0**	**29.9**	**26.1**	**100**	**1.02**	**47**	—
Medium regions	**44.6**	**17.6**	**37.8**	**100**	**1.41**	**34**	—
Large EMU regions[b]	**52.9**	**20.3**	**26.7**	**100**	**0.99**	**37**	—
Medium EMU Regions[b]	**46.5**	**16.3**	**37.2**	**100**	**1.38**	**33**	—
1980–1993							
Germany	59.5	30.5	10.0	100	0.85	31	0.17
United Kingdom	17.5	74.3	8.2	100	0.99	11	0.02
France	49.9	24.5	25.6	100	1.67	21	0.27
Italy	42.1	34.7	23.3	100	1.57	20	0.16
Belgium	54.5	34.9	10.6	100	1.08	10	0.11
Netherlands[a]	50.8	24.0	25.2	100	2.49	9	0.19
Greece	14.8	62.1	23.1	100	7.12	13	0.00
Spain	47.9	24.8	27.3	100	2.57	18	0.08
Portugal	34.7	51.3	14.0	100	9.80	5	0.01
Large regions	**46.8**	**36.2**	**16.9**	**100**	**1.42**	**69**	—
Medium regions	**31.9**	**45.6**	**22.5**	**100**	**3.32**	**68**	—
Large EMU regions[c]	**52.9**	**29.0**	**18.1**	**100**	**1.18**	**56**	—
Medium EMU regions[c]	**47.9**	**27.2**	**25.0**	**100**	**2.02**	**52**	—

[a] Groningen is excluded.
[b] UK is excluded.
[c] UK, Greece and Portugal are excluded.

A more accurate picture can be given by analyzing the dynamic profile of the different components. The problem can be analyzed through the frequency-domain representation. Fig. 1 shows the average spectra of the three components for Europe (first data set, UK excluded) and the US (medium and large counties). Although, as we have seen, the variances of the output components are similar, the dynamic profiles are very different. There are two main differences between Europe and the US. First, the European common component is very persistent, whereas the US-wide component exhibits a typical business cycle shape, which peaks at a period of around six years. By looking at low frequencies we see that the total long-run variance is similar in Europe and the

Table 2
Average variance explained by the US-wide component, the state component and the local component

State	σ^2_{US}/σ^2_y $\times 100$	σ^2_S/σ^2_y $\times 100$	$\sigma^2_{\mathscr{L}}/\sigma^2_y$ $\times 100$	Total	σ^2_y $\times 1000$	No. of counties
Connecticut	47.3	44.2	8.5	100	0.86	8
Maine	45.4	22.9	31.7	100	1.63	16
Massachusetts	38.0	47.4	14.6	100	1.05	14
New Hampshire	54.8	36.5	8.7	100	1.53	10
Rhode Island	50.3	35.4	14.3	100	1.21	5
Vermont	62.9	19.7	17.5	100	1.14	14
Delaware	54.9	26.2	18.8	100	1.11	3
Maryland	53.6	18.7	27.8	100	1.43	24
New Jersey	57.0	25.3	17.8	100	0.79	21
New York	45.7	29.2	25.1	100	0.64	62
Pennsylvania	65.5	9.8	24.7	100	0.85	67
Illinois	29.4	45.6	25.0	100	3.43	102
Indiana	50.8	30.0	19.2	100	2.83	92
Michigan	60.8	11.2	27.3	100	1.63	83
Ohio	63.3	11.6	25.1	100	1.36	88
Wisconsin	57.1	15.7	27.3	100	1.30	71
Iowa	43.4	39.6	17.0	100	7.52	99
Kansas	19.2	25.2	55.6	100	15.62	105
Minnesota	44.3	28.1	27.6	100	10.81	87
Missouri	41.2	36.3	22.5	100	4.94	115
Nebraska	27.4	36.7	35.9	100	18.48	93
North Dakota	25.2	55.0	19.8	100	42.99	53
South Dakota	30.3	44.3	25.4	100	18.33	66
Alabama	54.3	9.7	36.0	100	2.12	67
Arkansas	46.7	13.3	40.0	100	5.19	75
Florida	40.0	23.1	36.9	100	2.01	67
Georgia	38.2	18.5	43.3	100	3.38	159
Kentucky	33.9	31.2	34.8	100	2.84	120
Louisiana	22.7	29.8	47.5	100	3.46	64
Mississippi	39.1	20.9	40.0	100	4.74	82
North Carolina	55.2	10.6	34.2	100	2.47	100
South Carolina	51.0	13.4	35.6	100	1.83	46
Tennessee	59.3	10.6	30.0	100	2.35	95
Virginia	51.7	9.9	38.4	100	1.84	105
West Virginia	31.6	32.3	36.1	100	1.83	55
Arizona	44.2	15.0	40.8	100	3.68	14
New Mexico	28.3	16.0	55.7	100	5.57	32
Oklahoma	15.4	29.3	55.3	100	6.31	77
Texas	20.5	10.7	68.8	100	16.55	254
Colorado	21.3	13.4	65.4	100	10.64	63
Idaho	19.6	40.1	39.8	100	5.81	44
Montana	22.5	41.7	35.8	100	15.56	56
Utah	27.7	18.1	54.2	100	3.55	29

Table 2 (*Continued*)

State	σ^2_{US}/σ^2_y × 100	σ^2_S/σ^2_y × 100	$\sigma^2_{\mathscr{L}}/\sigma^2_y$ × 100	Total	σ^2_y × 1000	No. of counties
Wyoming	17.3	35.9	46.8	100	5.78	23
California	45.1	8.8	46.2	100	2.40	58
Nevada	21.8	31.0	47.2	100	5.18	17
Oregon	34.9	21.3	43.9	100	7.57	36
Washington	28.8	14.3	56.9	100	4.28	39
Large counties	**51.3**	**27.1**	**21.5**	**100**	**0.93**	**26**
Medium counties	**54.0**	**20.0**	**26.0**	**100**	**1.25**	**732**
Small counties	**28.6**	**28.4**	**42.9**	**100**	**8.62**	**2312**

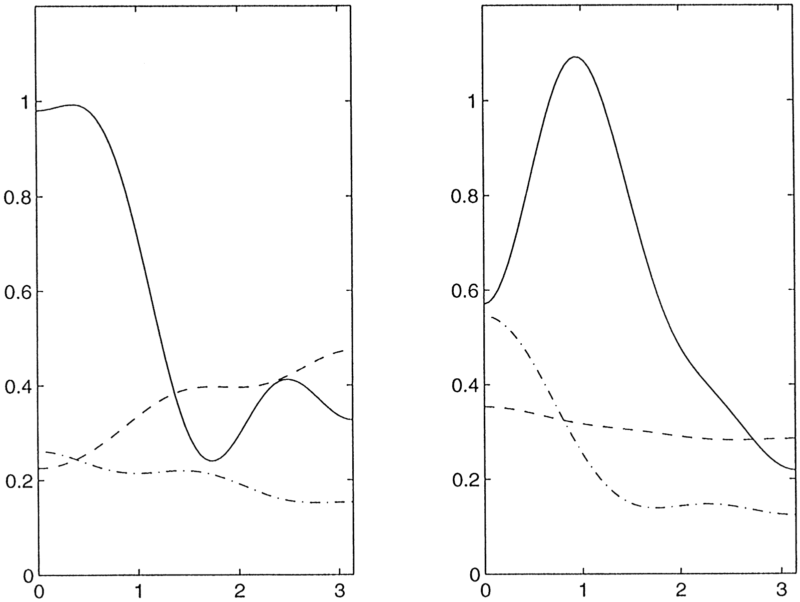

Fig. 1. The spectral shape of the three components for Europe and the US. Horizontal axis: frequency; vertical axis: spectral density. Common component: solid line; National (state) component: dotted and dashed line; local component: dotted line. Europe (first data set, UK excluded) is on the left; US (large and medium counties) is on the right.

US, so that the uncertainty about the future income level at, say, 10 or 20 years is nearly the same. However – and this is the second difference – while in the US the main bulk of long-run fluctuations is state specific or local, in Europe the long-run variance is mainly common. This implies that European regions have a larger long-run covariance, i.e. larger cospectra near the vertical axis. In other words, European regions have a 'common destiny' in the long-run, more than US counties.

The latter statement needs an important qualification. We are not claiming that European countries have similar income levels, or that they are converging to the same income level. We only observe that persistent shocks are mainly common. Still, both drift and initial levels, which are not analyzed here, could be rather different, leading to permanent gaps, convergence or even divergence.[7]

The analysis disaggregated by country (Fig. 2) shows that the UK is the only European nation which, like the US, has a typical business cycle. The other countries confirm the aggregate result of a large European wide component with most of the variance concentrated at low frequency (although Italy has a lot of high-frequency variation in the idiosyncratic).

Before we can interpret these results on variance decomposition in terms of degree of integration, we must study the degree of synchronization and symmetry of the propagation mechanisms of the common shock. If the response functions of the common shock had different signs, the interpretation of a large common component at low frequency would imply that regions are diverging, exactly the opposite of what concluded so far.

A rough measure of symmetry and synchronization is given by the regional distribution of the correlation coefficients. In Fig. 3 we report the frequency density distribution of a grid of correlation intervals for European regions (first data set) and US medium and large counties. In both cases, we can see that most correlation coefficients are positive and large.[8] This confirms the interpretation of variance decomposition results given above, even if we can notice that European regional dynamics is slightly more asymmetric than in the US case.

Finally, to complete the picture on European integration, 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 shift to the second data set, which is the most complete in terms of number of countries. Fig. 4 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, unlike what found by studies based on national data (see e.g. Bayoumi and Eichengreen, 1993; Helg et al., 1995), a core made by the

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[7] Note also that we are analyzing total income, as opposed to per-capita income. Different dynamics of total income in the US and Europe could be compensated by migrations, leading to similar dynamics of per-capita income.

[8] The negative values are accounted for by Sicily, Sardinia, some UK regions and Groningen, the Dutch outlier.

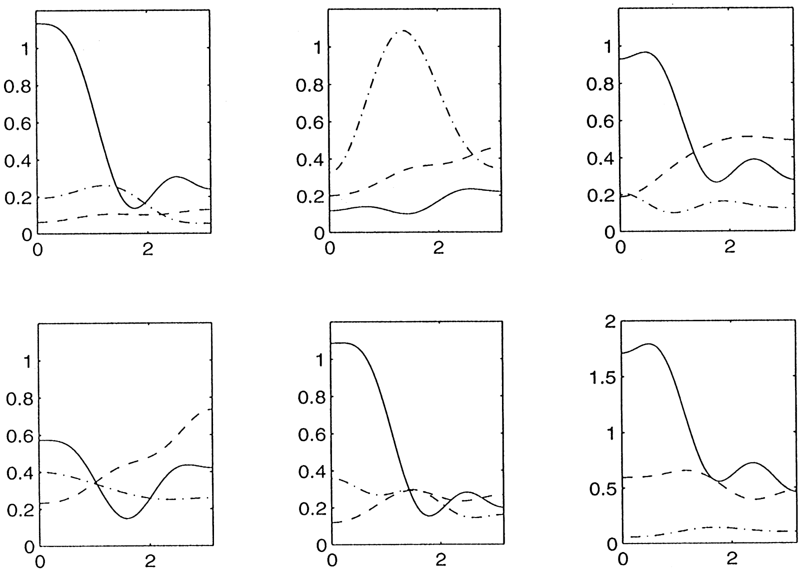

Fig. 2. The spectral shape of the three components for six European countries. Horizontal axis: frequency; vertical axis: spectral density. Common component: solid line; National (state) component: dotted and dashed line; local component: dotted line. From the left to the right: Germany, UK, France; Italy, Belgium, Netherland (Groningen excluded). Netherland has a different scale on the vertical axis.

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 less integrated with the rest of Europe. In general, heterogeneity within nations seems at least as large as heterogeneity across nations.

In summary, European regions are already highly integrated and are expected to 'move together' in the long run more than US counties. The common shock exhibits high persistence in Europe and a typical business-cycle shape in the US. If we exclude Greece and the UK, Europe appears a continent of regions rather than nations, as far as output fluctuations are concerned.

Our estimates on the degree of European integration based on output data are in line with what found by Fatás (1997) for unemployment, Fuss (1997) for income per capita and by Viñals and Jimeno (1996) for employment. They contrast, however, with the earlier literature on the optimality of Europe as a currency area (e.g. Bayoumi and Eichengreen, 1993) which obtained lower estimates. Given the variety of methods used in the literature, results are difficult to compare. One likely explanation for the discrepancies, however, is the higher level of aggregation at which those studies were conducted.

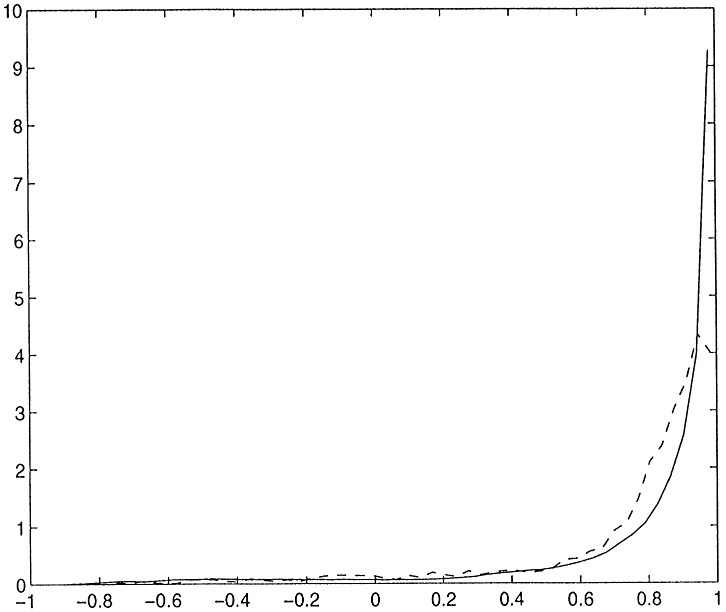

Fig. 3. Density distribution of correlations among the common components. Europe: dashed line; US: solid line.

5.2. *The potential output target of the European central bank*

A central bank can try to smooth output fluctuations by responding to negative exogenous shocks with expansionary measures. In fact, studies on the reaction functions of monetary policy indicate that central banks do react to output gap indicators (Clarida et al. 1998; Gerlach and Schnabel, 1998). On the other hand, there is a large consensus that such expansionary measures can only have temporary effects on output. The long-run response of output to a negative shock cannot be modified substantially: all that a central bank can do is to spread its negative effects over time. This means that monetary policy can lower the spectral density at short- and medium-run cyclical frequencies, but cannot be used as a long-run stabilization device. Hence, if we want to identify the potential target of the European central bank (ECB), we have to disregard long-run variance and focus on cyclical variance.

A further limitation of monetary policy is that it cannot be different for different regions. Reacting to the idiosyncratic shock of a particular region will necessarily produce undesirable effects for other regions. Hence a central bank cannot target idiosyncratic variance. In this respect, there is a basic difference between the ECB and the old system of national central banks. National banks,

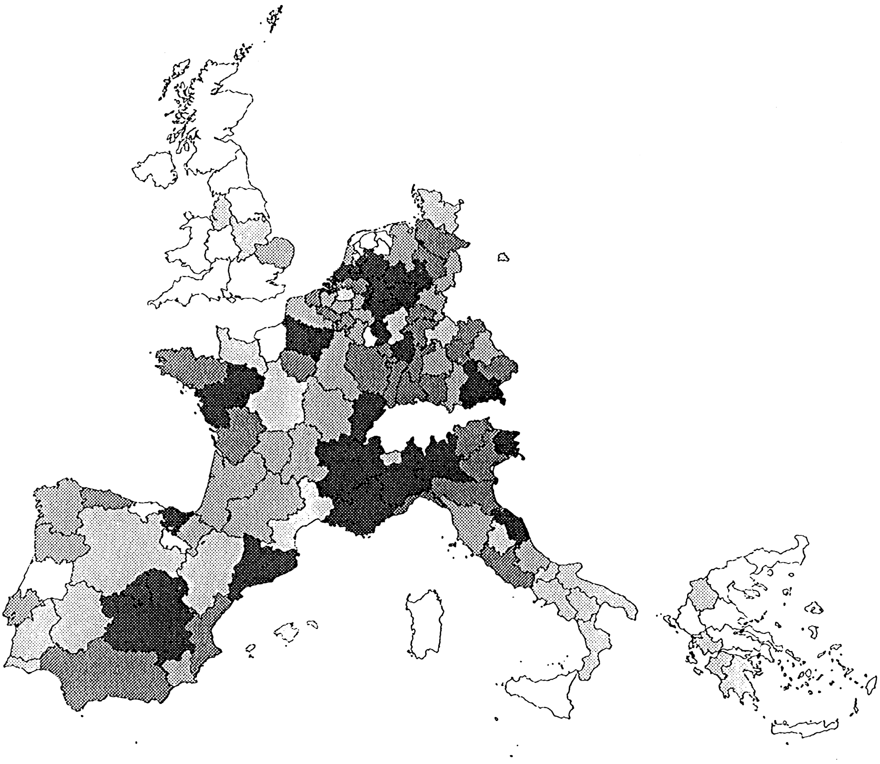

Fig. 4. Percentage of output variance explained by the European component. Dark regions have a large European component. Limits for color changes are 0.23, 0.42, 0.58, 0.70.

by conducting nation-specific policies, can try to stabilize not only the European, but also the national component of regional output. By contrast, the ECB has no power to smooth out the national shocks, so that its potential target is only that part of regional variance that is common to all European regions.

Summing up, the potential smoothing effect of the EMU can be measured by the fraction of total variance which is both cyclical and Europe wide. Moreover, the loss of potential stabilization with respect to the old monetary system can be measured by the percentage of total variance explained by cyclical and national fluctuations.

In Table 3 we report variance decompositions at different frequencies for Europe (excluding the UK which has not joined the EMU) and the US. We report average variances for both cycles between five and ten years and cycles of less than five years. Assuming a realistic target to be a cycle of less than five years, we estimate the potential smoothing effect of the common monetary policy to be a variance accounting for 18.3% of the total. Note that this is less

Table 3
Average long-, medium- and short-run variance in Europe and the US

Period	Common comp.	National comp.	Local comp.	Total	Tot. var. × 1000
European regions – UK excluded (70 *regions*)					
> 10 years	17.6	4.5	4.3	26.4	0.31
5 – 10 years	13.5	3.9	5.8	23.2	0.33
< 5 years	18.3	9.7	22.4	50.4	0.54
Total	49.5	18.1	32.5	100.0	1.19
US medium and small counties (758 *counties*)					
> 10 years	12.0	8.4	5.8	26.3	0.29
5 – 10 years	17.7	4.8	5.4	27.9	0.26
< 5 years	24.0	7.2	14.7	45.9	0.56
Total	53.7	20.3	26.0	100.0	1.12

than 50% of the total common component, which is not surprising, since the latter is very persistent.

We also estimate that the loss of output stabilization potential due to the establishment of the ECB is 9.7% of output variance, a number which is slighter higher than in the US.

Notice that, notwithstanding the anti-cyclical monetary policy of the Federal Reserve, the US cyclical variance remains larger than the European one. A counterfactual exercise would be interesting, but it is beyond the scope of this paper.

5.3. The potential insurance role of a fiscal federation

Fiscal policy, like monetary policy, can be used to stabilize aggregate output over time. In addition, however, it has the role of stabilizing local output by acting as an automatic insurance mechanism. Regions which are performing particularly well pay more taxes than regions experiencing temporary slowdowns and receive less benefits.

Abstracting from political economy type of considerations (for a discussion moral hazard problems related to the institution of a fiscal federation, see Persson and Tabellini, 1996), a European-wide insurance mechanism would be more effective than a national one, simply because of size; a fiscal federation can potentially insure both nation specific and purely regional volatility, while a nation can only insure regional variations. Only in the case in which the national component of variance is negligible, a fiscal federation will not have a higher insurable potential than a national insurance system. The size of the national component of output volatility will then tell us how much scope there is for a fiscal federation.

But should we look mainly at long-run fluctuations or short-run fluctuations? While traditional anti-cyclical policies, being based on intertemporal income transfers, can only address short-run volatility, a fiscal federation, by acting through cross-sectional transfers, can in principle reduce both short- and long-run variance. As long as long-run correlation between regional incomes is not perfect, contemporaneous transfers will reduce long-run idiosyncratic risk. From Table 3 we can see that national fluctuations account for about 18% of total variance. Hence the scope for a fiscal federation as an insurance mechanism is not negligible.[9]

5.4. Political economy

From the discussion, we can say that the map of potential losers and winners of a federal fiscal policy is given by the geographical distribution of the variance of the nation-specific component over total variance. Were the heterogeneity to be found within nations and not between nations, we may infer that the degree of consensus against a common fiscal policy will not be different than the consensus towards a national fiscal policy. Fig. 5 reports the map of the nation-specific variance over total variance.

The map shows clearly that Greece and the UK would benefit: more than other countries from fiscal integration. However, once again, if we exclude Greece and the UK, the map of potentially insurable risk is regional, not national. The lighter areas, which represent very low relative national variance, correspond roughly to the dark areas of Fig. 3 of regions with the highest relative common variance (the more integrated regions). These regions belong to Northern Italy, part of Germany and France – the corridor of commercial and productive activities of post-Middle Age Europe. The map suggests that potential opposition to a federation will not be larger than existing regional opposition against national federations.

If we construct the same map for cyclical frequencies only, we obtain a picture of the regions which will suffer most in terms of output stabilization from federal

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[9] In Forni and Reichlin (1998b), by analyzing a simple consumption model, we argue that welfare is related to long-run volatility, rather than volatility per se. The basic intuition is that consumers may do a lot by themselves, through saving decisions, in order to smooth short-run fluctuations, but are impotent with respect to permanent shocks. As a consequence, policy should concentrate on long-run stabilization rather than short cycles. If this argument is correct, we have to focus mainly on the fraction of long-run volatility (rather than total volatility) accounted for by the national component. This fraction however is very similar to the 18% mentioned before (4.5/26.4, see Table 3, first line), so that the main conclusion does not change.

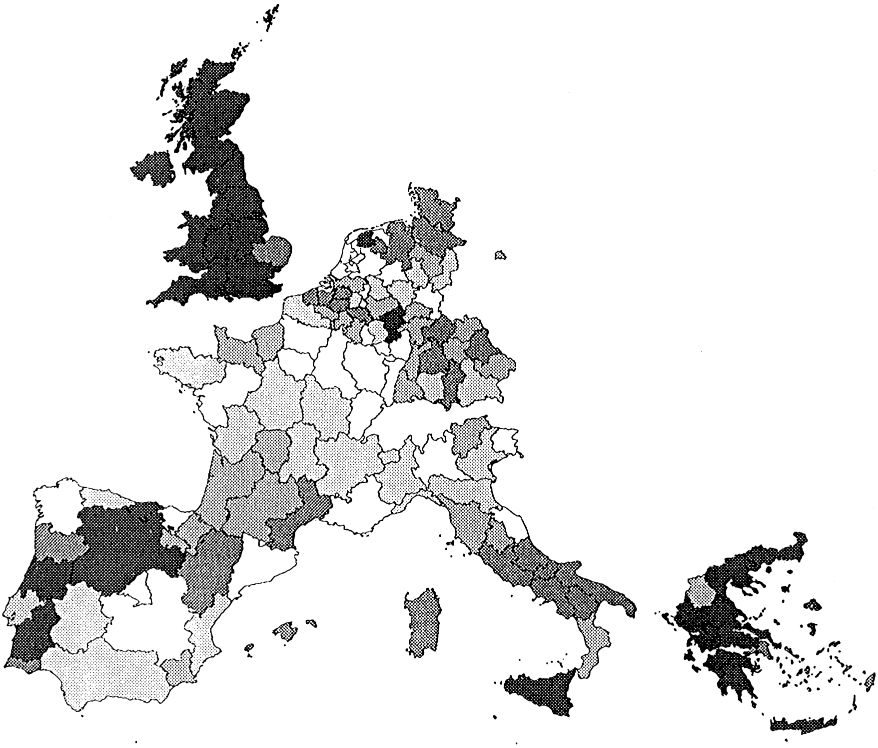

Fig. 5. Variance ratios: national component over total variance. Darker regions have larger variance. Limits for color changes are 0.16, 0.27, 0.37 and 0.56.

monetary policy. This map differs only slightly from Fig. 5 and therefore we do not report it. The conclusion is that, as expected, the more integrated regions will lose less from the establishment of Europe-wide targets. These regions, on the other hand, have less to gain from the establishment of a fiscal federation.

5.5. Diagnostic checking

At this point, some words should be spent about diagnostic checking. As we have already seen, our two European data sets are not very large both in the time and the cross-sectional dimension. Since the estimation procedure is based on the assumption of a large cross-section, one may wander whether in the present case the cross-sectional dimension is large enough to get reliable results.

As shown in Section 3 and Appendix A, the reciprocal of the larger eigenvalue of the matrix $\Sigma^{-1}\Gamma$ is an estimate of the percentage of idiosyncratic variance which is still present in the optimal weighted average that we use as an approximation to the common factors. The numbers we obtained from estimation are quite comforting. For the first data set the larger percentage of local variance in the national averages is 2.4%,[10] while the percentage of non-European variance in the European average is 6.1%. In the second data set, where more countries are present, the latter figure reduces to 4.1%. The percentage of local variance in the national averages is less than 2% for all nations except Portugal, which has 7.8%.

6. Alternative econometric approaches

The questions studied here, as well as several interesting economic problems, cannot be suitably addressed without modeling a large number of time series. An important feature of the dynamic factor model is that it provides a representation of the dynamic relations among many cross-sectional units which is both flexible and parsimonious. Forni and Reichlin (1996, 1998a) show that, contrary to the common wisdom, estimation of the factor model is quite simple when large cross-sections are involved. By contrast, more popular models such as VARs or panel data models fail either in parsimony or in flexibility. In VAR models the number of parameters to estimate is typically too large with respect to the number of available observations over time. On the other hand, panel data techniques, besides requiring a clearcut a priori distinction between 'dependent' and 'independent' variables, entail homogeneity assumptions on the response functions of different units which are severely restrictive when only a few observable explanatory variables are available.

The Panel VARs which have been used to study dynamics for many regions and nations (Blanchard and Katz, 1992; Decressin and Fatás, 1995; Obstfeld and Peri, 1998) while representing an interesting attempt to solve the trade-off, are less flexible than the model employed here. Indeed, the panel VAR model can be seen as a particular case of (1), where the restriction is that regions belonging to the same country have the same propagation mechanism, up to multiplication by a scalar, which must be the same for the national and the European shocks. More precisely, by setting in model (1) $a^{ij}(L) = a^{ij}a^{j}(L)$, $b^{ij}(L) = a^{ij}b^{j}(L)$, and $c^{ij}(L) = c^{ij}c^{j}(L)$ we get

$$y_t^{ij} = a^{ij}f_t^j + c^j(L)l_t^{ij},$$

where $f_t^j = a^j(L)e_t + b^j(L)n_t^j$. By averaging across regions belonging to country j the local components wash out, so that the national average is $y_t^j = a^j f_t^j$, where $a^j = \sum_i a^{ij}/I^j$. Hence

$$y_t^{ij} = \alpha^{ij}y_t^j + c^{ij}c^j(L)l_t^{ij},$$

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[10] Curiously the larger figure is not for the Netherlands, which have only nine regions, but for Italy, which has 20 regions. The smaller percentage is for Germany (0.8%).

where $\alpha^{ij} = a^{ij}/a^j$. The Panel VAR strategy consists in first regressing y_t^{ij} on y_t^j and then performing a restricted VAR on the residuals, separately for each nation j. While homogeneity of the response functions, at least to the European shock, is roughly in line with our empirical findings, the assumption that the coefficient a^{ij} is the same for the national and the European shocks seems rather restrictive. More generally, the panel VAR model above fails to distinguish properly between the national and the European component.

The dynamic factor framework provides a notion of comovements which is very natural. 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 Kozicki, 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 strength of long-run comovements only if all regions are pairwise cointegrated. This 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.[11]

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[11] For a more detailed discussion on cointegration and long-run comovement see Croux et al. (1999).

7. Summary, conclusions and … a caveat

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 Europe is as much integrated as the US. Integration 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 except for the UK and Greece. We also find a European core formed by regions with a large common component. The boundaries of the core, however, are not national so that all nations are partly inside and partly outside the core. 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.

We propose a measure of the potential output stabilization target of the European central bank as the percentage of total variance accounted for by short-run Europe-wide fluctuations. We find the potential target to be about 18% of total variance against a 24% in the US case. The smaller scope for federal monetary policy in Europe is explained by the higher level of persistence of the European common shocks with respect to US wide shocks. We estimate that the loss of stabilization due to the lack of national targeting is 9.7% of the output variance.

Further, we measure the variance stabilization role of a European fiscal federation as the percentage of total variance accounted for by national short-run fluctuations. This is estimated to be around 18% of total variance. Therefore, the ensemble of federal policies for stabilization have a non-negligible potential target of about 40%.

Let us conclude with a caveat. This is an empirical paper based on passed data. Is it reasonable to argue about the effect of future policies on the basis of past experience? Will EMU produce a major structural break in the joint behavior of local and national outputs? If yes, our results should be taken with caution. Indeed, existing literature is not unanimous about the size and the direction of future changes. Frankel and Rose (1998) argue that the new system will enlarge intra-EMU trade. Kalemli-Ozcan et al. (1999) suggest that nations will have less incentives to specialize. Such changes in the economic structure of local areas could affect synchronization of national cycles and the size of the common component. However, both trade flows and diversification of local economies are not likely to evolve very fast, so that we should not expect enormous changes in the next few years from this source of variation. This view is shared by Dornbusch et al. (1998), who argue that EMU will mainly change the policy ‘reaction functions’, while leaving the economic structure essentially unaltered. If this is the case, our empirical exercise should provide a reasonable description of the situation that the European central bank will face, with major changes possibly stemming from its policy intervention.

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Appendix A. The optimal weighting procedure

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

$$y_t^i = E_t^i + N_t^i + \mathscr{L}_t^i = C_t^i + \mathscr{L}_t^i,$$

where the common component C_t^i is orthogonal to the local idiosyncratic component $\mathscr{L}_t^i$. Now let us indicate with Σ the covariance matrix of the vector of the idiosyncratic components $\mathscr{L}_t = (\mathscr{L}_t^1 \dots \mathscr{L}_t^I)$ and with Γ the covariance matrix of the vector of the variables $Y_t = (y_t^1 \dots y_t^I)$. We want to find a weighted average such that the 'size' of the local component is minimized with respect to that of the common component. More precisely, we are looking for the I-dimensional vector w minimizing the ratio $\mathrm{var}(w'\mathscr{L}_t)/\mathrm{var}(w'Y_t)$, or, equivalently, the function

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

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

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

for some scalar λ and some $w \neq 0$. This is because a couple λ, w satisfying (A.1) 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 $\mathrm{var}(w'\mathscr{L}_t)/\mathrm{var}(w'Y_t)$ 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 $\Sigma^{-1}\Gamma$. Moreover, the reciprocal of the maximum latent root is an estimate of the percentage of the idiosyncratic variance in the optimal linear combination $w'Y_t$ and can be used for diagnostic checking.

If Σ is diagonal, as it is assumed here, Eq. (A.1) have a simple interpretation. Since

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

Eq. (A.1) reduces to

$$\frac{\mathrm{cov}(y_t^i, w'y_t)}{\mathrm{var}(\mathscr{L}_t^i)} = \lambda w^i, \quad i = 1, \ldots, I,$$

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.

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

Appendix B. 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 minimizing 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.

Step 4: A similar reasoning leads to the relation

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

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 $\mathscr{L}_t^{ij}$.

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

$$E_t^{ij} = \alpha^{ij}(L)y_t + \beta^{ij}(L)\alpha^j(L)y_t,$$

$$N_t^{ij} = \beta^{ij}(L)N_t^j = y_t^{ij} - E_t^{ij} - \mathscr{L}_t^{ij}.$$

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 modeling of y_t) and $b^j(L)$ (by univariate modeling of the N_t^j's) and using the relations

$$a^{ij}(L) = (\alpha^{ij}(L) + \beta^{ij}(L)\alpha^j(L))a(L),$$

$$b^{ij}(L) = \beta^{ij}(L)b^j(L).$$

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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-lag specification could not be rejected by the F-test.

Appendix C. Some issues about methodology

Though the aim of this paper is mainly empirical, the methodology used here raises some issues which deserve further discussion.

While the basic idea behind the estimation procedure is quite simple, deriving the general properties of the estimator is not. As observed in the main text, as long as the number of cross-sectional units I is finite, averaging does not kill completely the idiosyncratic component, so that the factor space can only be obtained with an error, which is similar to a measurement error. As a consequence, the usual properties of OLS estimates no longer hold. For instance, we cannot hope that the estimate of the common and the idiosyncratic components are consistent in T, holding I fixed: convergence may only hold as both I and T go to infinity. Moreover, it is not obvious whether T should go to infinity faster than I, and which rate should be required. Forni and Reichlin (1998a) (FR) show that, under mild assumptions on the parameters, the simple average estimator (see Section 3) is consistent for T, $I \to \infty$, no matter the relative rate. Our Monte Carlo experiments on models with low-order AR and MA common components suggest that the principal component estimator used here largely dominates the simple average estimator for all T and I. However, we do not have theoretical consistency results.

It is worth mentioning here that Forni et al. (1998) (FHLR), Forni and Lippi (1999) and Stock and Watson (1998) (SW) propose dynamic factor models where correlation among the idiosyncratic components is allowed. The estimation procedure used here cannot work in these models. However, the procedures proposed in FHLR and SW are based on the same large numbers intuition and involve principal component estimation (while estimation of Σ is not required). These estimators could be used here, but their performance is worst under simulation for T and I comparable with those of our data set. In both FHLR and SW theoretical consistency result and Monte Carlo experiments are provided.

The model proposed here can be generalized to allow for $q > 1$ common (and/or 'intermediate') factors. In this case q aggregates (together with their leads and lags) are needed in the RHS of the regression equations. These aggregates can be obtained in two ways. First, we can introduce in the model (and the data set) further variables, such as consumption or employment, driven by the same common shocks, and take the first principal component for each variable. Second, we can stick to the one-variable framework and take the first q principal components of the matrix $\Sigma^{-1}\Gamma$. In both cases, if we are interested only in the estimation of the components, the procedure described in Appendix B does not require additional modifications. By contrast, if we are interested in the identification and estimation of the factors themselves and the related impulse-response functions, in the case $q > 1$ further identifying assumptions are needed. A short treatment of this topic can be found in FR. Finally, note that the procedure just outlined works under the assumption that the number of common factors is known in advance. Formal tests for determining the number of common factors in this large cross-section framework are very difficult to construct; however, heuristic procedures are suggested both in FHLR and SW.

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