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A Measure of Comovement for Economic Indicators: Theory and Empirics

by

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Abstract

This paper proposes a measure of dynamic comovement between (possibly many) time series and names it cohesion. The measure is computed on the basis of the estimation of the cospectrum and it is appropriate for processes which are costationary in first differences. In the bivariate case, the measure is a function of the correlation, the drifts and the variances and it is defined at each frequency. The multivariate measure is a corresponding weighted average. We show that it relates in a simple way to cointegration.

Cohesion is useful to study problems of business cycle synchronization, to investigate long-run dynamic properties of multiple time series such as convergence of output per-capita, to identify dynamic clusters.

We fully describes the theoretical properties of the measure and provide two empirical illustrations. The first is on comovement of the output growth of 450 manufacturing sectors in the USA. The second is on comovement of output growth between US states and European regions.

Keywords: Business cycle, Sectoral comovements, Convergence .

JEL Classification: E3, C1

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1 Introduction [1]

There are few empirical relations which have the status of "stylized facts" in economics. One of them is that macroeconomic aggregates comove. This observation has been the source of speculation of economic theory since its birth. In modern theories of the business cycle people have asked whether comovements can be explained by large aggregate shocks, monetary or real, or whether an explanation should be found in non-linear propagation mechanisms. Every macroeconomic textbook starts by a statement on comovements between aggregates. However, paradoxically, this is one of the facts that is least well documented and on what there is more confusion of meaning and terminology. "Comovement" is a lose term, possibly describing different phenomena and, consequently, with many different interpretations. What are really the stylized facts and what should macroeconomics be trying to explain? Appropriate measures of comovement between time series processes should be developed to provide a meaningful answer to this question.

Such measure should also be useful for the study of the pattern of comovements between different disaggregate economic variables, such as sectoral and regional output or individual consumption since the analysis of disaggregate behavior can be informative about the aggregate. In recent literature, for example, the conjecture has been made that persistent aggregate fluctuations can be explained by micro shocks propagating locally through input-output relations, spillovers and local interaction (e.g. Long and Plosser, 1983, Cooper and Haltiwanger, 1990, Shea, 1994, Horvath, 1995). If this is the case, observed aggregate persistence may be caused by local rather than aggregate shocks. In general, any story based on local spillovers and complementarities should produce dynamic clusters in the relevant cross-section, defined by higher degree of comovements within clusters than between clusters. To test for this characteristics we should once again define a measure of comovement.

The informal discussion on comovements usually refers to something close to a notion of correlation. However, the traditional way with which the time series literature has dealt with measurement of comovements is based on a notion of rank reduction (see Ahn and Reinsel, 1988) which has a different meaning. In this category belongs the idea of cointegration (Engle and Granger, 1987): two processes are cointegrated if the spectral density at frequency zero has rank one, codependence (Gourieroux and Peaucelle, 1992), which refers to linear

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combinations of correlated processes which are of lower autoregressive order than others, common features (Engle and Kozicki, 1993), which are linear combinations which are unpredictable with respect to past information and common cycles (Vahid and Engle 1993) which are defined as common features in first differences for processes which are cointegrated. This class of measures presents several problems. First, high cross-correlation neither implies nor it is implied by cointegration, common cycles or common features and identical dynamics neither implies nor is implied by them (Quah, 1993 and Forni and Reichlin, 1997). Second, these measures are binary. For example, two processes are either cointegrated or not, but we can't establish different degrees of association. Finally, in order to establish rank reduction, we need to estimate the parameters of a VAR which may be problematic when the number of time series is large. For all these reasons, while the notion of rank reduction is certainly interesting to characterize some aspects of the dynamic properties of multivariate time series, it is not the appropriate one for the study of comovements.

In this paper we develop a measure of comovement which is closer to the notion of correlation and is based on the estimation of cross-spectra. Our measure, however, unlike correlation, takes also in consideration the drifts and the variances. This is because variables in levels may comove weakly even if their first differences are strongly correlated, owing to large differences in both deterministic and stochastic trends. The comovement index can be decomposed by frequency and frequency band and can then be used to study business cycle as well as long-run questions. Moreover, it can be generalized in such a way as to provide a summary measure of the degree of comovement within a group of variables or between two groups of variables. The latter index, named "cohesion", can be used for instance to see how strongly the outputs of a set of sectors or regions comove, both in the long-run and at the cyclical frequencies, and compare results with other sets of sectors or regions.

To illustrate our proposed measure and to provide further motivation we implement two empirical applications. In the first one, the objective is to study the "local interaction hypothesis". We study the pattern of cohesion of sectoral output for 450 manufacturing sectors in the USA since 1958 (4-digit disaggregation level) and we show how the size and shape of cohesion between different groups of sectors convey information on the nature of shocks and propagation mechanisms.

In the second, we study output per capita of US states and European nations. We first evaluate differences in overall cohesion of countries within Europe and of states within the US. Then we evaluate the dynamic shape of cohesion to establish convergence and business cycle synchronization.

The paper is organized as follows. We first describe the properties of our comovement and cohesion indeces. We then compare it with other comovement measures used in the literature. In Section five we estimate it for the two empirical examples, compute bootstrap confidence intervals and illustrate its use in dynamic clustering. We end with summary and conclusions.

2 Contemporaneous comovement

Let us consider two stochastic processes, say x_{1t} and x_{2t}. We are looking for an index c measuring how strongly x_{1t} and x_{2t} "move together". Our task will be greatly simplified by assuming that the first differences Δx_{1t} and Δx_{2t} are jointly covariance-stationary, an assumption which seems reasonable for many economic variables. In this way we can construct our measure as a function of the first and the second moments of Δx_{1t} and Δx_{2t}, i.e. their covariance σ_{12}, their standard deviations σ_1 and σ_2, and means μ_1 and μ_2.

A quite natural requirement is that c is a positive function of the correlation coefficient $\rho = \sigma_{12}/\sigma_1\sigma_2$ and is normalized so that its maximum value is 1. Indeed, one could be tempted to set simply $c = \rho$. This choice however entails a concept of comovement which is weaker than that we have in mind.

To clarify this point, assume that Δx_{1t} and Δx_{2t} are perfectly correlated, so that $\Delta x_{1t} = \beta + \gamma \Delta x_{2t}$. Integrating both sides we get $x_{1t} = \alpha + \beta t + \gamma x_{2t}$, where α is some constant. Note that, owing to the deterministic trend βt, the two variables are not cointegrated. This case is illustrated in Figure 1a for $\beta = 1$ and $\gamma = 4$: it is clear that the behaviour of x_{1t} and x_{2t} does not correspond to the intuitive meaning of comovement.

A stronger requirement, implying cointegration, would be

$$x_{1t} = \alpha + \gamma x_{2t}. \tag{1}$$

But even this appears unsatisfactory in many cases, depending on the nature of the variables x_{1t} and x_{2t}. For instance, think of x_{1t} and x_{2t} as being the per-capita incomes of two different countries. Equation (1) implies that the difference $x_{1t} - x_{2t}$ is equal to $(\gamma - 1)x_{2t}$ (up to a constant term). Hence if the variables are I(1) then $x_{1t} - x_{2t}$ is also I(1), unless $\gamma = 1$. This entails that the probability of x_{1t} and x_{2t} being closer than any given distance goes to zero as t goes to infinity, even if they are equal in $t = 0$. Clearly this is not what we mean when saying that the per-capita incomes of two countries move perfectly together. If we do not want $x_{1t} - x_{2t}$ to be I(1) we have to impose both $\beta = 0$ and $\gamma = 1$, i.e. $x_{1t} = x_{2t}$ up to a constant term. Put another way, we want a measure c such that $c = 1$ if and only if, in addition to $\rho = 1$, we have $\mu_1 = \mu_2$ and $\sigma_1 = \sigma_2$ (implying that $\Delta x_{1t} = \Delta x_{2t}$).

Per-capita incomes of different countries is only a paradigm for a wide range of important economic variables. For instance $j = 1, 2$ may represent a region or a sector and x_{jt} may be the unemployment rate, the inflation rate, or the interest rate. Clearly if x_{1t} and x_{2t} are aggregate savings of, say, US and Belgium, not taken in logs, it would make little sense to require $x_{1t} = x_{2t}$ up to a constant term. In this case condition (1) seems more appropriate. But if savings are normalized by dividing by population or are taken in logs, as it is usually done in empirical work, this requirement is perfectly suitable. The same holds for variables like consumption, investment, exports and so on.

Summing up, we are looking for a statistic c which is an increasing function of ρ, but depends also on σ_1, σ_2, μ_1 and μ_2, and reaches its maximum when $\rho = 1$, $\mu_1 = \mu_2$ and $\sigma_1 = \sigma_2$.

Additional requirements are needed to consider the case of negative comovements, i.e. the case in which x_{1t} and x_{2t} move in opposite ways. To take account of this case, we allow for negative values of c and take -1 as the minimum value. Moreover, we require that $c \to 0$ as $|\mu_1 - \mu_2| \to \infty$ or $\sigma_j \to \infty$. This entails that, if we start from a negative c, a growing $|\mu_1 - \mu_2|$ (or σ_j) will eventually raise c, rather than reducing it, reflecting the fact that the negative relation becomes weaker. Our motivation is that the means and the variances should affect only the size of c, not its sign, which is determined by ρ: a large difference between μ_1 and μ_2 simply means that x_{1t} and x_{2t} have very different drifts and hence comove weakly, not that they comove negatively. The same holds if σ_j approaches infinity while ρ, the means and the other standard deviation are held constant. This is illustrated in Figure 1b, where we show two series which are perfectly correlated, but do not comove due to a large difference in their variances.

In sum, we have that c is positive, negative or zero if and only if ρ is positive, negative or zero. Moreover, if the sign of ρ changes, the sign of c changes, but $|c|$ does not, so that $c(x_{1t}, x_{2t}) = -c(x_{1t}, -x_{2t} + 2\mu_2 t)$.

An implication is that $c = -1$ if and only if $\rho = -1$, $\sigma_1 = \sigma_2$ and $\mu_1 = \mu_2$, i.e. if and only if $x_{1t} - \mu_1 t = -(x_{2t} - \mu_2 t)$ up to a constant term and $\mu_1 = \mu_2$; in words, if the average of the two series equals their common deterministic trend. Figure 1c illustrates two variables with perfect negative comovement.

Notice that, if $\mu_1 = \mu_2$ and $\sigma_1 = \sigma_2$, we are requiring $c = \rho$ both in the case $\rho = 1$ and in the case $\rho = -1$. We are requiring $c = \rho$ also when $\rho = 0$. Indeed, the only motivation for our departure from the correlation coefficient ρ is that we want to take into account the differences in the variances and the means. If $\mu_1 = \mu_2$ and $\sigma_1 = \sigma_2$, it seems sensible to assume that $c = \rho$ for any value of ρ, not only in the cases $\rho = 1$, $\rho = -1$ and $\rho = 0$.

The above discussion justifies the following list of properties for the measure of comovement c.

A.1 $c = f(\rho, \sigma_1, \sigma_2, \mu_1, \mu_2)$ where f is continuous and takes on values in $[-1, 1]$.

A.2 (a) c is increasing in ρ; (b) $\rho > 0$ implies $c > 0$; (c) c is antisymmetric in ρ, i.e. $f(\rho, \sigma_1, \sigma_2, \mu_1, \mu_2) = -f(-\rho, \sigma_1, \sigma_2, \mu_1, \mu_2)$.
A.3 $|c|$ is strictly decreasing in $|\mu_1 - \mu_2|$; (b) $\lim_{|\mu_j| \to \infty} c = 0$, $j = 1, 2$; $\lim_{\sigma_j \to \infty} c = 0$, $j = 1, 2$; (c) if $\mu_1 = \mu_2$ and $\sigma_1 = \sigma_2$, then $c = \rho$.

Obvious additional requirements are:

A.4 c commutes with respect to x_{1t} and x_{2t}, i.e. $f(\rho, \sigma_1, \sigma_2, \mu_1, \mu_2) = f(\rho, \sigma_2, \sigma_1, \mu_2, \mu_1)$.

A.5 c is invariant with respect to multiplication of x_{1t} and x_{2t} by a positive constant a, i.e. $f(\rho, \sigma_1, \sigma_2, \mu_1, \mu_2) = f(\rho, a\sigma_1, a\sigma_2, a\mu_1, a\mu_2)$.

We are now ready to introduce our measure of contemporaneous comovement. Our starting point is the opposite of the squared distance between Δx_{1t} and Δx_{2t},

$$-E(\Delta x_{1t} - \Delta x_{2t})^2 = -\sigma_1^2 - \sigma_2^2 + 2\sigma_{12} - (\mu_1 - \mu_2)^2. \tag{2}$$

The above quantity increases when distances reduces (as we want) and has a maximal value of zero. In order to normalize (2) we suppose for a moment that Δx_{1t} and Δx_{2t} are uncorrelated, in which case $E(\Delta x_{1t} - \Delta x_{2t})^2$ would simplify to

$$\sigma_1^2 + \sigma_2^2 + (\mu_1 - \mu_2)^2. \tag{3}$$

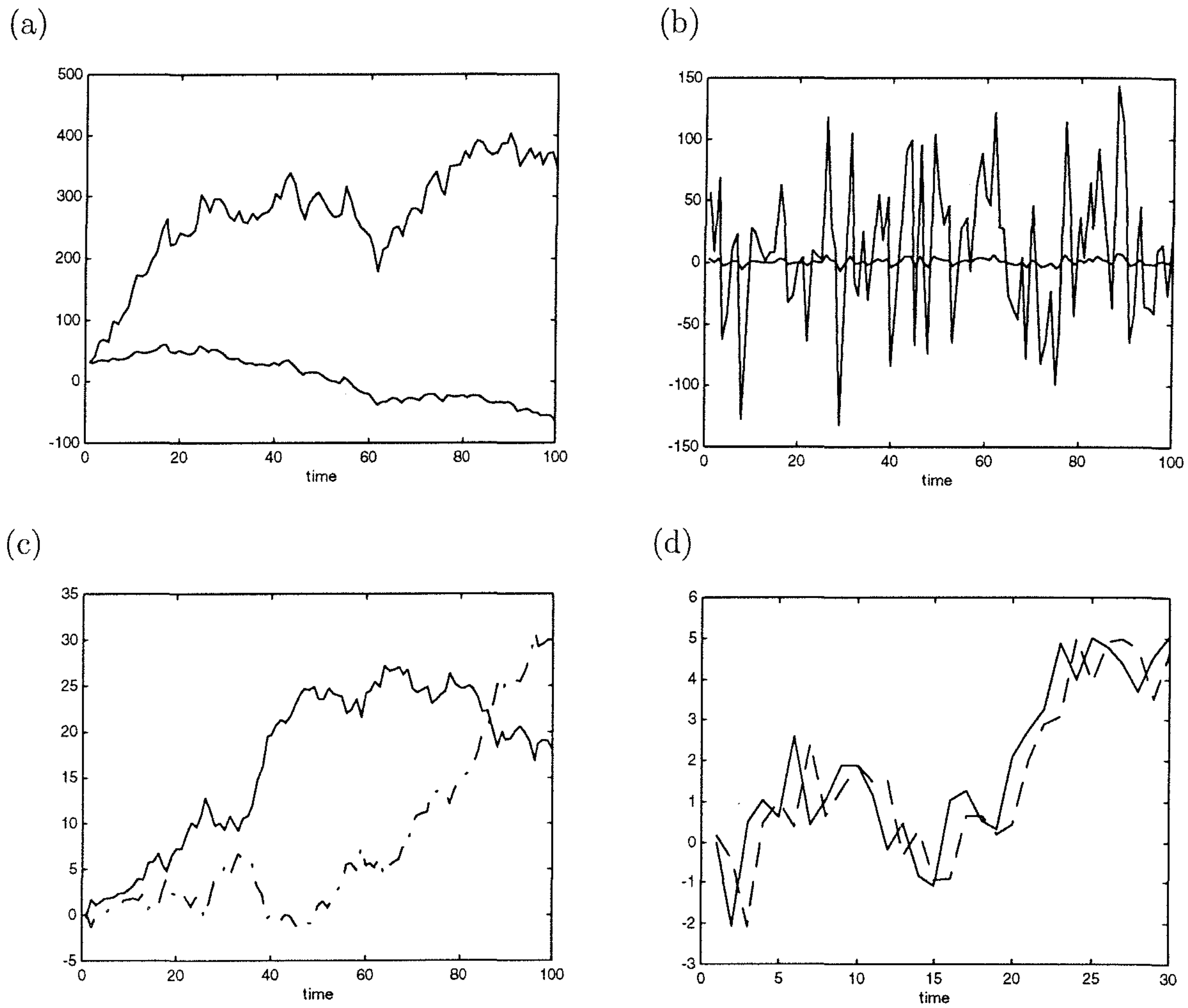

Figure 1: Examples of series with (a) perfect correlation, weak comovement (b) perfect correlation, weak comovement (c) perfect negative comovement (d) no (contemporaneous) correlation, strong comovement

The fraction of (2) and (3) lies in the interval $[-2, 0]$. To satisfy properties A1-A2, we add 1 to this fraction. The resulting index of contemporaneous comovement is then

$$c = \frac{2\sigma_{12}}{\sigma_1^2 + \sigma_2^2 + (\mu_1 - \mu_2)^2} = \rho \frac{2\sigma_1\sigma_2}{\sigma_1^2 + \sigma_2^2 + (\mu_1 - \mu_2)^2} \tag{4}$$

The reader can easily verify that the above function satisfies A.1-A.5, along with

A.6 c is not affected by the addition of equal deterministic trends to x_{1t} and x_{2t}, i.e. $f(\rho, \sigma_1, \sigma_2, \mu_1 + b, \mu_2 + b) = f(\rho, \sigma_1, \sigma_2, \mu_1, \mu_2)$.

The above list of properties is not complete, i.e. A.1-A.6 do not identify uniquely the function (4). To see this, notice that by multiplying the last term at the denominator of (4) by a positive $\omega \neq 1$, a different function is obtained which also satisfies A.1-A.6.

Up to now, we only discussed contemporaneous comovement and abstract from any dynamic considerations. The importance of dynamics is illustrated by Figure 1d, where we show the example: $x_{1t} = x_{1,t-1} + u_t$ and $x_{2t} = x_{1,t-1}$ with u_t white noise. Although the contemporaneous correlation (and also the measure c) between the two series is zero, they clearly comove. The next Section discusses the extension to the dynamic case.

3 Comovement decomposed by frequency and frequency bands

Let us now substitute in formula (4) the static second moments with their dynamic counterparts, that is the spectral densities $\sigma_j^2(\lambda)$, $-\pi \leq \lambda \leq \pi$, for the variances and the co-spectrum $\sigma_{12}(\lambda)$, i.e. the real part of the cross-spectrum, for the covariance. We get the comovement index

$$c(\lambda) = \frac{2\sigma_{12}(\lambda)}{\sigma_1^2(\lambda) + \sigma_2^2(\lambda) + (\mu_1 - \mu_2)^2/(2\pi)}, \tag{5}$$

which measures the comovement of x_{1t} and x_{2t} at frequency λ.[2]

We introduce also a measure of comovement relative to the frequency band $\Lambda = [-\lambda_2, -\lambda_1] \cup [\lambda_1, \lambda_2]$ with $0 \leq \lambda_1 < \lambda_2 \leq \pi$. The latter is given by

$$c(\Lambda) = \frac{2\int_\Lambda \sigma_{12}(\lambda)\, d\lambda}{\int_\Lambda [\sigma_1^2(\lambda) + \sigma_2^2(\lambda)]\, d\lambda + (\mu_1 - \mu_2)^2(\lambda_2 - \lambda_1)/\pi}. \tag{6}$$

[2]Note that, if $\mu_1 = \mu_2$, the denominator can vanish. However, this is not really a problem, since, if Δx_{1t} and Δx_{2t} are ARMA processes, as we assumed, the denominator can vanish only in a finite number of isolated points, where $c(\lambda)$ can be defined by continuity.

Contemporaneous comovement c is obtained by setting $\lambda_1 = 0$ and $\lambda_2 = \pi$, so that Λ is equal to the whole interval $[-\pi, \pi]$.

Notice that $c(\Lambda)$ – and as a special case $c([-\pi, \pi]) = c$ – can be considered as a weighted average of the values taken on by $c(\lambda)$ over Λ, since, denoting by $d(\lambda)$ the denominator of (5), $c(\Lambda)$ is the integral of $c(\lambda)d(\lambda)$ divided by the integral of the weight $d(\lambda)$. In this sense, $c(\lambda)$ decomposes c by frequency. As a consequence, we have $c = 1$ (or $c = -1$) if and only if $c(\lambda) = 1$ (or $c(\lambda) = -1$) almost everywhere on $[-\pi, \pi]$. Similarly, $c(\Lambda)$ decomposes c by frequency bands. Let us consider a partition of $[0, \pi]$ into m intervals $I_1, \dots, I_m$ and set $\Lambda_k = -I_k \cup I_k$, $k = 1, \dots, m$. The reader can easily verify that c is a weighted average of $c(\Lambda_1), \dots, c(\Lambda_m)$.

An immediate consequence of definition (5) is

B.1 $c(\lambda) = 0$ for every $\lambda \in [-\pi, \pi]$ if and only if Δx_{1t} and Δx_{2t} are uncorrelated at all leads and lags.

Evaluating $c(\lambda)$ at frequency zero we get an index of long-run comovement between x_{1t} and x_{2t}, which is related to cointegration in the following way.

B.2 If x_{1t} and x_{2t} are difference-stationary, i.e. $\sigma_j^2(0) \neq 0$, $j = 1, 2$, then they are cointegrated with cointegrating vector $(1 \quad -1)$ if and only if $c(0) = 1$.

To see this, call $g(\lambda)$ the spectral density of $\Delta(x_{1t} - x_{2t})$; $x_{1t} - x_{2t}$ is stationary if and only if $g(0) = \sigma_1^2(0) + \sigma_2^2(0) - 2\sigma_{12}(0) = 0$ and $\mu_1 = \mu_2$, i.e. $2\sigma_{12}(0) = \sigma_1^2(0) + \sigma_2^2(0)$ and $\mu_1 = \mu_2$. Since $\sigma_1^2(0) + \sigma_2^2(0) \neq 0$ the above conditions are equivalent to $c(0) = 1$.

Figure 1e plots the comovement index of the two series showed in Fig 1d, which have zero contemporaneous correlation, but are cointegrated with cointegrating vector $(1 \quad -1)$. On average, the index equals zero; however, inspection of all frequency bands reveals that this is the result of large long-run positive comovements canceling out with large short-run negative comovements.

Remark: We can say that two series are cointegrated with arbitrary cointegration vector when the comovement index at frequency zero computed from the standardized series $x_{1t}/\sigma_1(0)$ and $x_{1t}/\sigma_2(0)$ equals one.

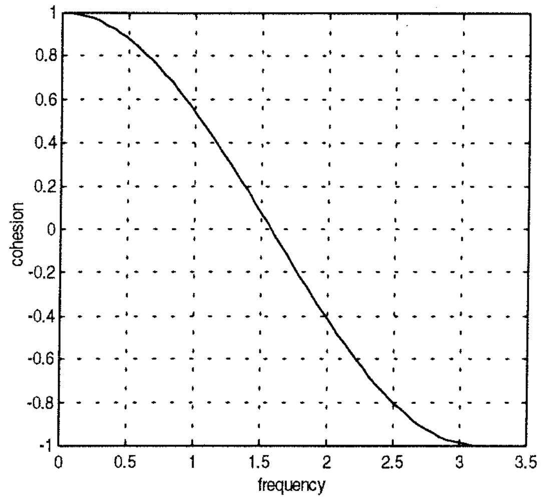

Figure 1e: comovement index of two cointegrated series with zero correlation

4 Cohesion and cross-cohesion

Let us now consider a vector of $n \geq 2$ variables $x_t = (x_{1t} \quad \cdots \quad x_{nt})'$. Moreover, let us attach non-normalized positive weights $w = (w_1 \quad \cdots \quad w_n)'$ to the variables in x_t. Our proposed measure for the internal cohesion of the variables in x_t is motivated in the same manner as in the bivariate case and equals the fraction of the weighted averages formed by the numerator and denominator of the comovement index defined in (4). Contemporaneous cohesion is defined as

$$\mathrm{coh} = \frac{2\sum_{i \neq j} w_i w_j \sigma_{ij}}{\sum_{i \neq j} w_i w_j [\sigma_i^2 + \sigma_j^2 + (\mu_i - \mu_j)^2]}, \tag{7}$$

where for notational simplicity we do not explicit the dependence on w. Note that coh can be seen as a weighted average of the comovement indeces c_{ij}: denoting by d_{ij} the denominator of c_{ij} and setting $v_{ij} = w_i w_j d_{ij}$, we have $\mathrm{coh} = \sum_{i \neq j} c_{ij} v_{ij} / \sum_{i \neq j} v_{ij}$.

At frequency λ, we have

$$\mathrm{coh}(\lambda) = \frac{2\sum_{i \neq j} w_i w_j \sigma_{ij}(\lambda)}{\sum_{i \neq j} w_i w_j [\sigma_i^2(\lambda) + \sigma_j^2(\lambda) + (\mu_i - \mu_j)^2/(2\pi)]}, \tag{8}$$

while the measure of cohesion within the frequency band $\Lambda = [-\lambda_2, -\lambda_1] \cup [\lambda_1, \lambda_2]$ is given by

$$\mathrm{coh}(\Lambda) = \frac{2\sum_{i \neq j} w_i w_j \int_\Lambda \sigma_{ij}(\lambda)\, d\lambda}{\sum_{i \neq j} w_i w_j [\int_\Lambda \sigma_i^2(\lambda) + \sigma_j^2(\lambda)\, d\lambda + (\mu_i - \mu_j)^2(\lambda_2 - \lambda_1)/\pi]}$$

A very simple choice for the weights is $w_i = 1$ for all i. While equal weights may work well in many cases, they appear unsuited when we deal with sectors or regions with very different importance. For instance, if we want to measure the cohesion of the per-capita incomes of the European countries, it is reasonable to give Germany a greater weight than Luxembourg. A natural choice for w_i in this case would be the level of income or population of country i at some t.

We exclude the diagonal terms in the weighted averages (7) and (8) for two reasons. First, it seems reasonable to require that, if the entries in Δx_t are pairwise uncorrelated, then contemporaneous cohesion is zero, and if they are pairwise uncorrelated at all leads and lags, cohesion is zero at all frequencies. These properties hold with the above definition, whereas they would be violated when including the diagonal terms. Second, the inclusion of the diagonal terms would render cohesion dependent of n. As an example, assume $\mu_i = \mu$, $\sigma_i = \sigma$ for all i and $\rho_{ij} = \rho$ for $i \neq j$, and set for simplicity $w_i = 1$ for all i. In this case $c_{ij} = \rho$ for $i \neq j$. According to definition (7) we get coh $= \rho$, while the inclusion of the diagonal terms would give $[\rho(n-1)+1]/n$, so that a group with two uncorrelated variables would have cohesion 0.5 while a group with ten pairwise uncorrelated variables would have cohesion 0.1.

Clearly $\text{coh}(\lambda) \leq 1$ and $\text{coh}(\lambda) = 1$ if and only if all of the variables in x_t comove perfectly at frequency λ. In particular, $\text{coh}(0) = 1$ if and only if the variables in x_t are pairwise cointegrated with cointegrating vector $(1 \quad -1)$, i.e. they have a common trend representation and equal long run responses to the permanent shock. Moreover, $\text{coh}(\lambda) = 1$ at each frequency, i.e. coh $= 1$, if and only if all of the x_{it}'s differ from each other only by a constant term.

The lower bound of the cohesion index is -1 for $n = 2$, since, if $n = 2$, the cohesion of x_t coincides with the comovement index of x_{1t} and x_{2t} (in dependently of w). For $n > 2$ the lower bound is greater, since of course we cannot have perfect pairwise negative correlation within a group of three variables or more. To illustrate this point, let us consider that contemporaneous cohesion is minimized when $\mu_i = \mu$ and $\sigma_i = \sigma$ for all i. In this case $c_{ij} = \rho_{ij}$ and coh $= (w'Rw - w'w)/\sum_{i\neq j} w_i w_j$, where R is the correlation matrix. This expression cannot be less than $-w'w/\sum_{i\neq j} w_i w_j$, which value depends on the particular choice of the weights. An example is the case of equal weights, with the minimum $1/(n-1)$

tending to zero as n gets larger.

Here is a list of properties for the index of cohesion. Properties C.1 and C.2 are trivial; Property C.3 is an immediate consequence of A.5 and A.6; Property C.4 has been discussed above.

C.1 $\text{coh}(\lambda)$ is independent of the order of the variables in x_t (provided that the weights are reordered in the same way).

C.2 $\text{coh}(\lambda)$ is homogeneous of degree zero in the weight vector w.

C.3 The cohesion indices of x_t and $ax_t + bt\iota_n$, ι_n being the vector of ones, is the same for any pair of scalars a and b.

C.4 (a) $-1 \leq \text{coh}(\lambda) \leq 1$; (b) $\text{coh}(\lambda) = 1$ if and only if $c_{ij}(\lambda) = 1$ for all pairs i, j; (c) if $c_{ij}(\lambda) = 0$ for all pairs i, j, then $\text{coh}(\lambda) = 0$.

Note that, while pairwise correlation implies zero cohesion, the converse is not true: when observing a small cohesion index we cannot distinguish whether it originates from small pairwise comovements or large negative and positive covariances canceling out each others. This problem could be avoided by using the alternative measure

$$\text{coh}^* = \frac{2\sum_{i \neq j} w_i w_j |\sigma_{ij}|}{\sum_{i \neq j} w_i w_j [\sigma_i^2 + \sigma_j^2 + (\mu_i - \mu_j)^2]}.$$

On the other hand, coh^* has the disadvantage that it does not distinguish between negative and positive comovements. If we want to retain both informations we can do this by comparing coh and coh^*.

The cohesion index can be easily generalized to an index measuring the cross-cohesion between the n-vector x_t and the m-vector y_t. For the sake of simplicity, let us assume for the moment that x_t and y_t have no common elements. Let us specify two vectors of weights w_x and w_y. Contemporaneous cross-cohesion of x_t and y_t is given by

$$\text{coh}_{xy} = \frac{2\sum_{i=1}^{n}\sum_{j=1}^{m} w_{xi} w_{yj} \sigma_{ij}}{\sum_{i=1}^{n}\sum_{j=1}^{m} w_{xi} w_{yj} [\sigma_i^2 + \sigma_j^2 + (\mu_i - \mu_j)^2]}.$$

The definition of the corresponding dynamic measures is trivial. Note that if both x_t and y_t are scalars, cross-cohesion reduces to comovement.

In some cases it could be interesting to evaluate the cross-cohesion of two overlapping sets of variables. In this case we should eliminate from the average the self-comovements of the variables in the intersection, as we have done in defining the cohesion index (7). Hence the above definition of cross-cohesion should be generalized in the following way. For notational simplicity, let us reorder the variables in x_t and y_t in such a way that, if x_t and y_t have k variables in common, the latter variables occur in the first k places. The generalized index of cross-cohesion is then

$$\frac{2\sum_{i=1}^{n}\sum_{j=1}^{m} w_{xi}w_{yj}\sigma_{ij} - \sum_{i=1}^{k} w_{xi}w_{yi}\sigma_i^2}{\sum_{i=1}^{n}\sum_{j=1}^{m} w_{xi}w_{yj}[\sigma_i^2 + \sigma_j^2 + (\mu_i - \mu_j)^2] - \sum_{i=1}^{k} w_{xi}w_{yi}\sigma_i^2}. \tag{9}$$

According to this definition, the cross-cohesion of x_t with x_t is equal to the cohesion of x_t.

5 Empirical illustration

5.1 Real output in US manufacturing

This section uses the cohesion measure to show how to proceed empirically to answer one of the most classic questions in macroeconomics. Are the features of aggregate business cycle explained by aggregate shocks or by sectoral shocks with strong "local" linkages? Recent studies (eg Forni and Reichlin, 1996, 1998) have found that aggregate manufacturing output in the US has a pronounced business cycle measured as a peak in the spectral density corresponding to a period of around six years. The question we ask is whether this peak can be explained by sectoral shocks which are not purely idiosyncratic, but that have propagation mechanisms which affect some neighboring sectors.

Whether aggregate business cycle is explained by aggregate or sectoral disturbances is a long-standing question (Lilian, 1982 is one of the earliest and most influential papers). The observation of sectoral comovements, however, is not sufficient to support the sectoral view since sectoral comovements may be generated by macro-shocks. Moreover, since the effect of purely idiosyncratic sectoral shocks should cancel out in the aggregate, for the sectoral explanation to be valid, we should find that the sectoral shocks are not purely idiosyncratic and produce fluctuations in the output of their own sector as well as in that of neighboring sectors [3]

[3]Several mechanisms have been suggested; we can broadly classify them in two categories: input-output linkages (e.g. Long and Plosser, 1983, Horvath, 1995) and linkages through demand for final goods or other complementarities (e.g. Cooper and Haltiwanger, 1990, Shea, 1994).

Our cohesion measure can be used easily to detect whether these "local" effects are present in the data.

Although we are aware that other relevant criteria can be chosen, here we will follow the SIC classification to define distance between sectors. We consider 450 manufacturing sectors (four-digit classification) for the USA over the 1958-86 time period. We then aggregate to obtain the twenty sectors of the two-digit classification. We ask two questions. First, whether output of sectors that belong to the same two-digit classification is more cohesive (within cohesion) than that of randomly selected groups of sectors. Second, whether cohesion is particularly pronounced at business cycle frequencies.

Our empirical strategy starts from the following observations. If sectoral four-digit shocks were purely idiosyncratic, by aggregating from the four-digit to the two-digit level, their effect would disappear owing to the law of large number and observed aggregate behaviour should reflect the effect of aggregate shocks only. If there are strong comovements within the two-digit groups, measured by relatively (with respect to random groups) high cohesion, the law of large number works slowly and the behaviour of the aggregate should then partially reflect the effect of micro shocks. This implies that if (i) the aggregate has a peak at business cycle frequency, (ii) within two-digit cohesion peaks at business cycle frequency and (iii) cohesion within two-digit groups is high relatively to cohesion within randomly selected groups, aggregate business cycle in manufacturing is explained by micro shocks with strong local effects. On the other hand, if (i) holds but cohesion within two-digit is large but flat, then the explanation for the aggregate shape should be based on the effect of aggregate shocks. In this case we should observe that the spectral density of 4-digit sectoral output has a peak at business cycle frequency.

Figure 2 and 3 show the spectrum of the aggregate and the average of the 450 sectoral spectra. By comparing Figure 2 and 3 we can see that the spectral shape of the aggregate is not reproduced by the shape of the within 2-digit cohesion. While the aggregate shows a pronounced peak corresponding to a cycle of about six years, cohesion seems to be either stronger at low frequencies or constant. These results indicate that the peak in the aggregate spectral shape is produced by aggregate shocks common to all sectors rather than sectoral shocks with local linkages. In the latter case, we should have observed a peak of cohesion corresponding to a six year cycle. This conclusion is reinforced by the comparison between

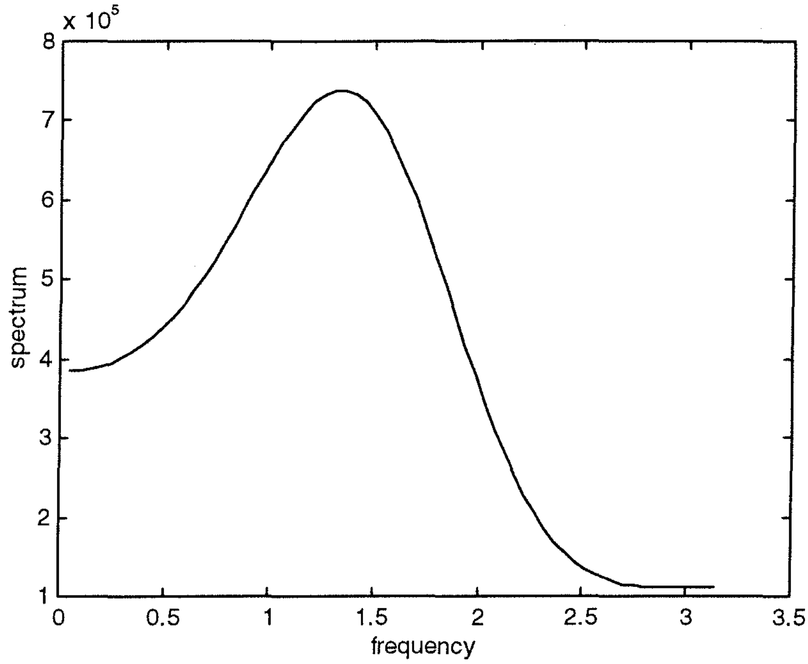

Figure 2: Spectrum of the aggregate output over the 450 manufacturing sectors

Figure 3 and Figure 4: cohesion within two digit does not seem to be significantly higher than cohesion within randomly selected groups.

5.2 Per-capita income: US states and European nations

Here we consider cohesion of per-capita income of 49 US states (within cohesion) from 1962 to 1994 and compare it with cohesion within two groups of European countries: sixteen West-European nations (UK excluded) and the European core composed by France, Germany and the Benelux. We have also computed the cohesion between the European groups and the US. Figure 5 shows that the 16 European countries are less cohesive than US states, especially at business cycle frequency. Moreover, their internal cohesion is very close to the cohesion between the US as a whole and Europe as a whole. In Figure 6 we have also computed bootstrap confidence bands so that we can make more precise statements on whether observed differences are statistically significant. It is interesting to note that, in the long-run, the internal cohesion of Europe is not significantly different from that of the US. Moreover, from Figure 7, we can see that when the European core is considered, European cohesion is very close to US cohesion and much higher than between cohesion, especially at

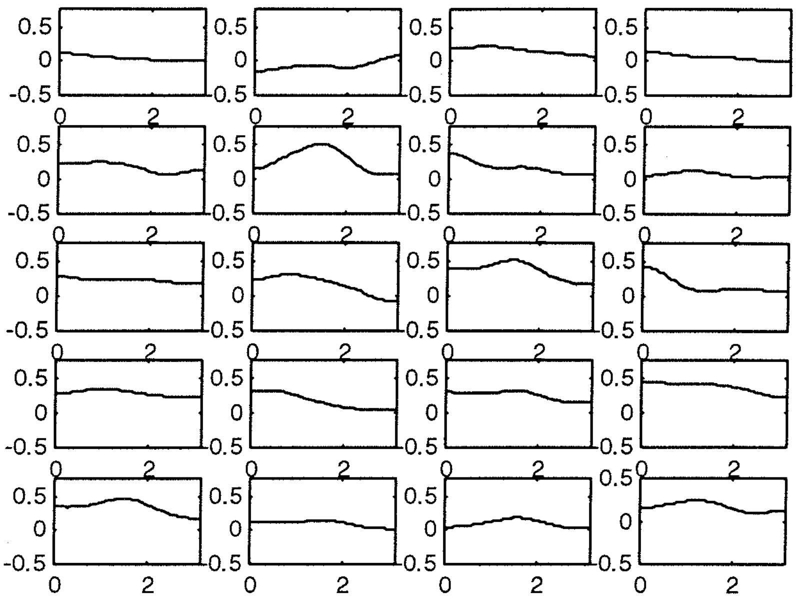

Figure 3: Within Cohesion of 20 groups of manufacturing sectors, 2 digit classification

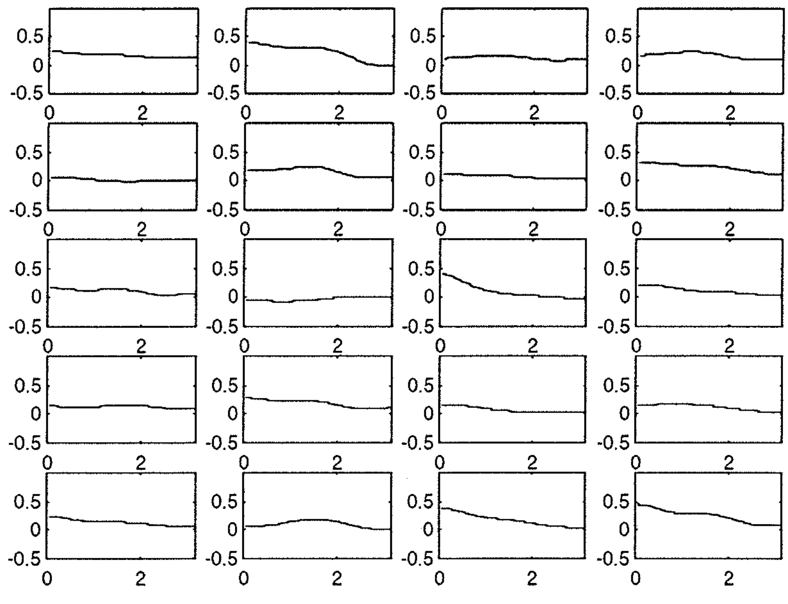

Figure 4: Within Cohesion of 20 groups of manufacturing sectors, random classification

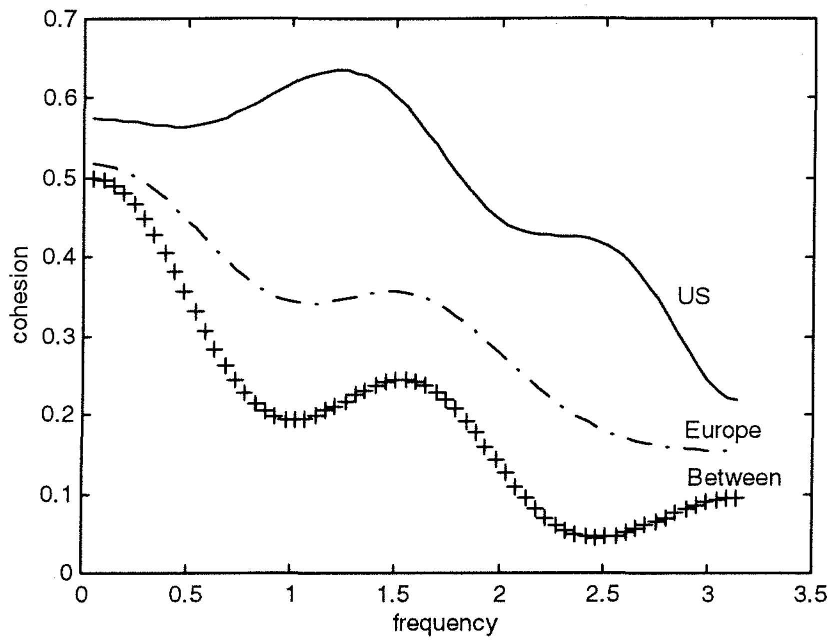

Figure 5: Within and Between Cohesion of the group of 49 US-states and a group of 16 West-European Countries

business cycle frequency.

The proposed measure of cohesion can be used as a measure of distance in computing clusters thereby providing a methodology for dynamic clustering. We will illustrate this by grouping 17 European countries (cfr. Appendix 1) according to the way that their income per capita comove together.

To this purpose, we first need to compute the 17 by 17 matrix of (positive) "dissimilarities", with elements

$$D_{ij}(\Lambda) = 1 - \text{coh}(\Lambda)_{ij} \tag{10}$$

for a given frequency band Λ. We say that countries which strongly comove within the given frequency band, have small "dissimilarities".

To distinguish between short-run and long-run dynamics, we performed calculations for two different frequency ranges, corresponding to cycles of period of ten years or longer (long-run dynamics) and cycles of period less than ten years, representing short-run fluctuations and business cycle dynamics. Of course, this is an arbitrary partition of the spectrum, and

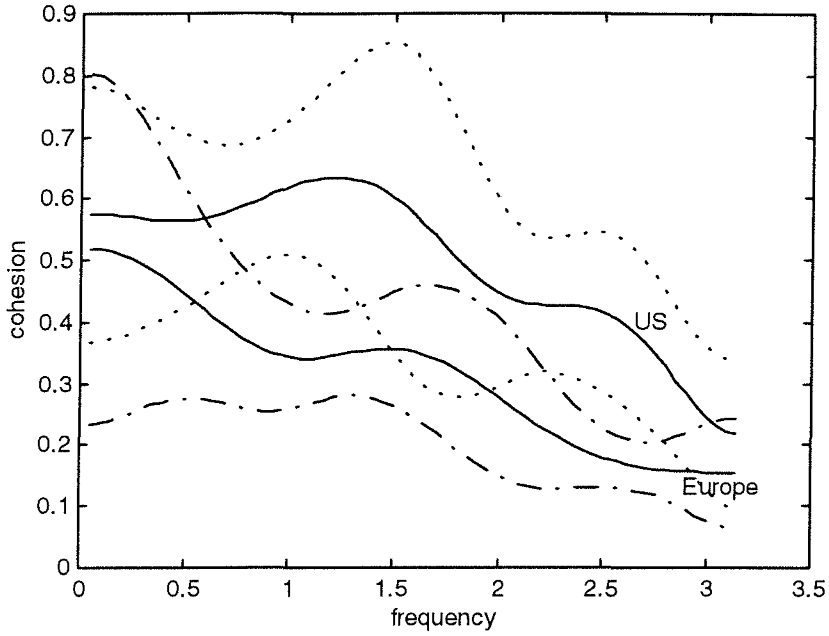

Figure 6: Within Cohesion of GNP/capita for US and Western Europe. Confidence bounds for US are indicated by $\cdots$ and for Europe by $-$.

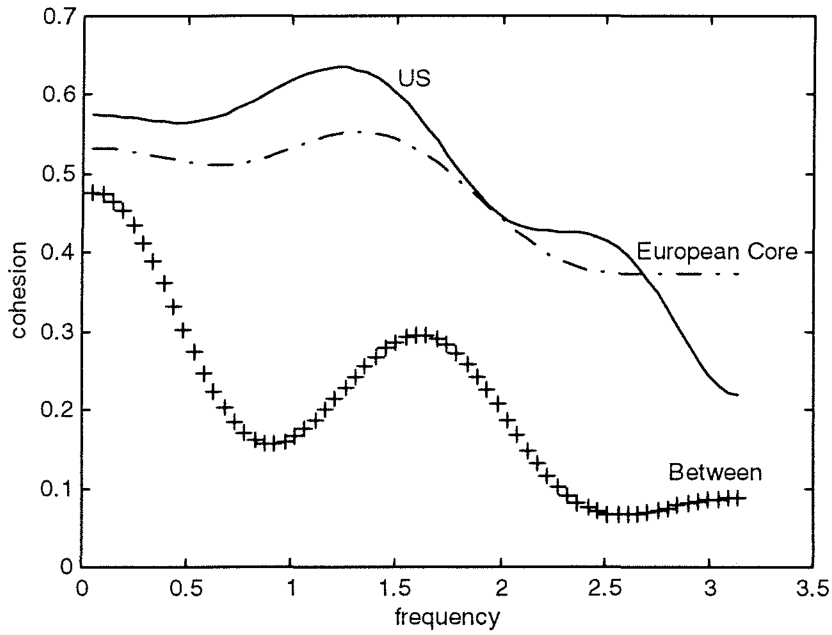

Figure 7: Within and Between Cohesion of the group of 49 US-states and a core group of 5 European Countries

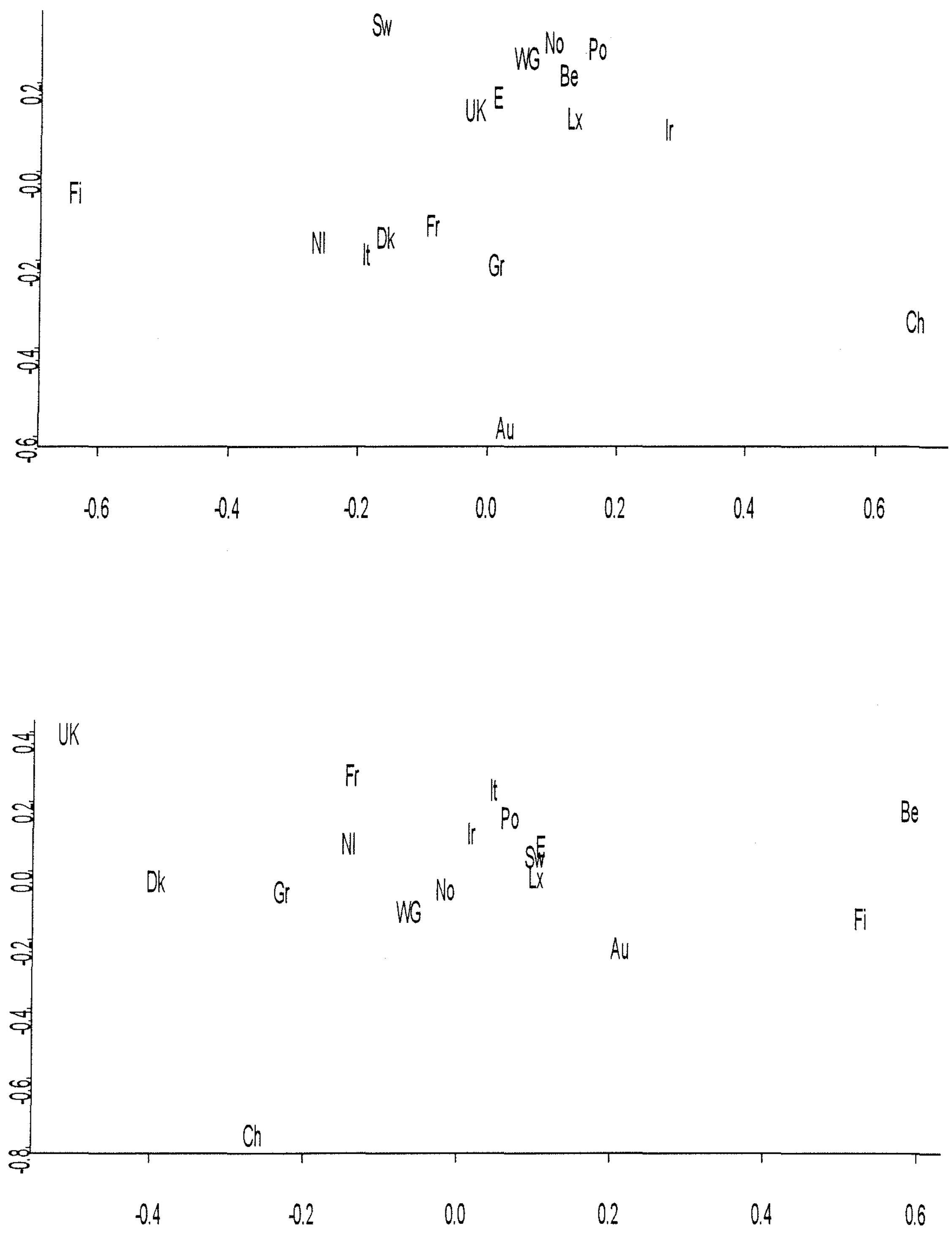

Figure 8: Representation of 17 West-European countries in a plane according to their long run comovements (upper figure) and short run comovement (lower figure)

other partitions of $[-\pi, \pi]$ can be chosen, depending on the empirical application. [4]

Once the dissimilarity matrix has been computed, standard clustering techniques can be applied. We chose to apply the metrical *multidimensional scaling* [5] technique (Cox and Cox, 1994) which represents the 17 countries in a 2 dimensional plane. The Euclidean distances between the points in the plane are now supposed to mimic the dissimilarities computed by (10). Countries which have big dissimilarities, have representations in the plane which are far away from each other. Looking at the objects in the two-dimensional plane, makes it possible to detect possible clusters and outliers. In Figure 8 we illustrate results of the application of multidimensional scaling based on the cohesion indices for long run comovements (Figure 8a), and for business cycle and short run comovements (Figure 8b). No clear grouping emerges and countries seem to have fairly homogenous dynamic behaviour. Not surprisingly, an outlier for both long-run and short-run is Switzerland; the UK, on the other hand, while belonging to the main group in the long-run, has low cohesion with the rest of Europe for cycles of ten years or shorter. This confirms results found in the literature on the basis of different techniques (e.g. Forni and Reichlin, 1997).

6 Summary and conclusion

This paper has proposed a measure of dynamic comovement between (possibly many) time series and named it cohesion. The measure is computed on the basis of the estimation of the cospectrum and it is appropriate for processes which are costationary in first differences. In the bivariate case, the measure is a function of the correlation, the drifts and the variances and it is defined at each frequency. The multivariate measure is the corresponding weighted average. We show that it relates in a simple way to cointegration.

Cohesion is useful to study problems of business cycle synchronization, to investigate long-run dynamic properties of multiple time series such as convergence of output per-capita, to identify dynamic clusters.

We have fully described the theoretical properties of the measure and provided two empirical illustrations. The first is on comovement of the output growth of 450 manufacturing

[4]In practice, the integral appearing in the definition of $\mathrm{coh}(\Lambda)_{ij}$ was computed by the simple trapezium rule for numerical integration.

[5]The procedure cmdscale of the statistical software package Splus was used for this.

sectors in the USA. The second is on comovement of output growth between US states and European regions.

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APPENDIX 1:

Data sources and data treatment

A. *US Manufacturing output*: The data set used is the Annual Survey of Manufacturers (ASM) which is a survey of manufacturing establishments sampled from those responding to the comprehensive Census of Manufacturers. This database contains information for 4-digit manufacturing industries from 1958 through 1986. We have used value added data for output and deflated them by the value of shipments.

Logs of sectoral data on output and productivity were subject to unit root tests. For all data we were not able to reject the null of a unit root (results available on request) at the 5 % level. We then took the differences and removed the mean. The electronic computer sector (SIC 357) was found to have a unit root after being detrended by a segmented trend with change in drift in 1972.

B. *US states income per-capita:* The data are "Per capita personal income (USD)" from REIS, database provided by the Bureau of Economic Analysis, Economics and Statistics Administration of the US department of Commerce.

C. *European nations income per-capita:* The data are GDP at 1990 market prices deflated by the GDP deflator (Mrd USD 1990) and divided by the population. The source is Eurostat.

Countries considered in the large group: Belgium (Be), Denmark (Dk), West Germany (WG), Greece (Gr), Spain (E), France (Fr), Ireland (Ir), Italy (It), Luxembourg (Lx), Netherlands (Nl), Portugal (Po), Switzerland (Ch), Austria (Au), Norway (No), Sweden (Sw), Finland (Fi). Countries considered in the European core: Belgium, West Germany, France, Luxembourg, the Netherlands. For the cluster analysis, the United Kingdom (UK) was added to the large group.

APPENDIX 2:

Bootstrapping technique

Confidence limits around the estimated cohesion coefficients were computed using bootstrap techniques. An overview of methods for bootstrapping time series is given in Berkowitz and Kilian (1996). We applied a standard block bootstrap for the cohesion of per capita income within US and Western-Europe, respectively (cfr. Section 5.2). The number of replicates was 200 and the length of the blocks was chosen to be equal to 8, large enough to retain the cyclical information in the series. The distribution of the bootstrap replicates was relatively close to a normal one, and therefore the confidence limits were computed by the 2 σ-rule. Alternatively, one could bootstrap the $\tanh^{-1}$ transform of cohesion (which can takes values over the whole real line), which made the distribution of the bootstrap replicates closer to a normal, but the obtained confidence limits remained almost the same.

Since cohesion is defined in the frequency domain, it is very appealing to use the spectral approach of Berkowitz and Diebold (1996). By bootstrapping the (multivariate) spectrum instead of the time series itself, we can save computation time, since the spectrum does not need to be recomputed for each bootstrap replicate. Preliminary experiments in this direction were not so successful, probably due to the fact that the time span in the treated examples is rather small. Research on this topic is ongoing.

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