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Let's get real: a factor analytical approach to dissaggregated business cycle dynamics

by

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ABSTRACT

This paper develops a method for analysing the dynamics of large cross-sections based on a factor analytic model. We use "law of large numbers" arguments to show that the number of common factors can be determined by a principal components method, the economy-wide shocks can be identified by means of simple structural VAR techniques and the unobserved factor model can be estimated by applying OLS equation by equation. We distinguish between a technological and a non-technological shock. Identification is obtained by minimizing the negative realizations of the technology shock. Empirical results on 4-digit industrial output and productivity for the US economy from 1958 to 1986 show that: (1) at least two economy-wide shocks, both having a long-run effect on sectoral output, are needed to explain the common dynamics; (2) although the technological shock accounts for at least 50 % of the aggregate dynamics of output, it cannot by itself explain dynamics at business cycle frequencies; (3) sector-specific shocks explain the main bulk of total variance but generate mainly high frequency dynamics; (4) both the technological and the non technological component of output show a peak for *positive* sectoral comovements of output at business cycle frequencies; (5) technological shocks are strongly correlated with the growth rates of the investment in machinery and equipment sectors and their inputs.

JEL Classification: C51, E32, O30.
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1. Introduction[1]

Many interesting questions about cyclical fluctuations and economic growth can be answered to only by studying the dynamic behavior of sectoral variables. When data contain information on time for a large cross-section of sectors, traditional econometric techniques used in the macroeconomic literature such as Vector Autoregressive (VAR) and Vector Autoregressive Moving Average (VARMA) models are not appropriate since they require the estimation of too many parameters. This is why new methods which allow for the reduction of the parameter space need to be developed.

The objective of this paper is both methodological and descriptive.

At the methodological level we develop a simple framework for the dynamic analysis of large cross-sections. The basic model is a dynamic factor analytic model as in Sargent and Sims (1977). The sectoral variables are decomposed into two unobservable components: a common component, driven by macroeconomic shocks, and a purely sectoral component. When the cross-section is large, simple large numbers arguments can be used to show that, due to orthogonality, the sectoral idiosyncratic component dies out on average relative to the common component (Chamberlain 1983, Granger 1987 and Forni and Lippi 1995). Here we exploit this result in order to develop a new estimation procedure. More specifically, we show that the number of common factors can be determined by a principal components method, the economy-wide shocks can be identified by means of simple structural VAR techniques and the unobserved factor model can be estimated by applying OLS equation by equation. This is a great simplification with respect to existing methods (see for example Quah and Sargent 1994).

An additional contribution of the paper is the identification of the common factors. We distinguish between a technological and a non-technological shock. Identification is obtained by imposing a "quasi-positivity" constraint. More precisely, the technological shock process is defined as the shock for which the absolute sum of the negative realizations are minimized. By using this criterion we are taking the view that technological shocks, in general, take the form of technical improvements and, in this case, must be positive. However, they may exceptionally be negative since they include special events such as oil shocks or institutional changes affecting the organization of production.

At the descriptive level we characterize the nature of fluctuations of output and productivity in US manufacturing by analysing the dynamics of

[1] We would like to thank John Shea for providing the data. Thanks for helpful comments are due to Renato Flores, Carlo Giannini, Christian Gourieroux, Clive Granger, Wolfang Haerdle, Alan Kirman, Marco Lippi, Enrique Sentana, Marc Watson, Michael Woodford, two anonymous referees and the participants at the ECARE-CEPR conference on empirical macroeconomics.

450 sectors (4-digit classification) from 1958 to 1986. We ask the following questions. First, how many shocks are common to all sectors? The answer to this question would provide an empirical justification for the choice of the stochastic dimension in aggregate models of fluctuations. Are business cycle models driven by only one shock a good characterization of aggregate behaviour or else, do we need to work with multi-shocks models? Second, we quantify the relative importance of macro and sector-specific dynamics. Several papers in the literature have addressed this issue (Lilien 1982 and, more recently, Davis and Haltiwanger 1992 and 1994 and Horvath and Verbugge 1996, amongst others). We go beyond reporting variance ratios, by analysing separately the whole dynamic profile of the common and idiosyncratic elements. We are then able to answer precisely to the question of whether purely sector-specific shocks generate cyclical fluctuations as claimed, for example, by Long and Plosser (1983) and by the literature on strategic complementarities (e.g. Cooper and Haltiwanger 1990 and Shea 1994). Moreover, we analyse in detail the propagation mechanism of economy-wide technological shocks by looking not only at the contribution of technology to the total variation of output and productivity, but also at whether real shocks are capable of generating a cycle as manifested by positive sectoral comovements at business cycle frequencies. Finally, we explore whether the rate of growth of output is associated with technological innovation in some industries. If the answer is affirmative, this would imply that, as suggested by De Long and Summers (1991, and 1992) there are key sectors whose technological progress affects the overall economic growth through strong positive externalities.

2. The model

Let us begin by assuming a countable infinity of sectors $i = 1, \ldots, \infty$. We specify a dynamic factor analytic model, as for instance in Sargent and Sims (1977), Geweke and Singleton (1981) and, more recently, Quah and Sargent (1994). More precisely, we assume that we have m variables of interest and that for each sector i the m-vector $y_t^i = (y_{1t}^i, y_{2t}^i, \cdots, y_{mt}^i)'$ can be written as

$$y_t^i = A^i(L)u_t + \epsilon_t^i, \tag{1}$$

where

$$\epsilon_t^i = (\epsilon_{1t}^i, \epsilon_{2t}^i, \cdots, \epsilon_{mt}^i)'$$

is a vector of sector-specific factors - the idiosyncratic components - possibly autocorrelated but mutually orthogonal at all leads and lags, with variances bounded above by the reals σ_h with $h = 1, \cdots, m$;

$$u_t^i = (u_{1t}, u_{2t}, \cdots, u_{qt})'$$

is a vector of q unit variance white noises, the common shocks, identical for all sectors and variables, mutually orthogonal and orthogonal to ϵ_t^i for all i; $A^i(L)$ is a $m \times q$ matrix of rational functions in the lag operator L. We call $A^i(L)u_t$ "the common component". All the variables are in deviation from the mean, wide-sense stationary and linearly regular, with rational spectral density matrix.

The above model is used to estimate the dynamics of the rate of growth of output and labor productivity ($m = 2$) for $n = 450$ manufacturing sectors of the US economy from 1958 to 1986 (for a more precise description of the data and the data sources see Appendix 2).

The pure factor analytic model (1) implies that sectoral variables are driven by shocks which are either common to our n sectors or purely sectoral at the 4-digit level. Both types of shocks are allowed to generate heterogenous dynamics across sectors, but autoregressive linkages and intermediate-size shocks which are common to subsets of sectors are ruled out. This could be seen as an excessive simplification since the former should capture dynamic input-output relations and the latter reveal strategic complementarities within clusters of sectors. On the other hand, if these effects were empirically significant, model (1) would fail specification tests. In particular, the orthogonality condition on the idiosyncratic components would be violated. As shown by the orthogonality test on the estimated idiosyncratic components (see Appendix 1.B), the latters are nearly orthogonal so that we can safely conclude that model (1) captures the essential empirical dynamic features of our data.[2]

A static version of the same framework has been proposed in the financial literature to model systematic and idiosyncratic risk (see for example Chamberlain 1983). In macroeconomics unobserved component models have been extensively used to estimate permanent and transitory dynamic components (see Harvey 1989 for a discussion of permanent transitory decompositions in the dynamic factor analytic framework and Stock and Watson 1988 for a different approach). Our framework differs insofar as both unobserved components are allowed to have permanent and transitory dynamics. Harvey's model can be seen as a particular case of the dynamic factor model (1) since in both his model and model (1) the components are mutually orthogonal. On the other hand, our model should be distinguished from the common trend representation proposed by Stock and Watson (1988) where the two components are driven by the same vector of shocks.

The methodology proposed here to estimate the model exploits an important property of factor models. Due to orthogonality, when aggregating across a large number of sectors the idiosyncratic component vanishes rel-

[2] The issue of AR linkages and intermediate shocks is analysed in more details in Forni and Reichlin (1996b), both at the theoretical and the empirical level.

atively to the common component.

To clarify what we mean, let us introduce for each variable h a sequence of real numbers ω_h^i, $i = 1, \ldots \infty$, such that we can find positive reals L_h and U_h fulfilling

$$L_h \leq \omega_h^i \leq U_h.$$

Now consider a strictly increasing sequence of positive integers i_k, $k = 1, \ldots, \infty$ and let $D_n = \{i_1, \ldots, i_n\}$. The variance of the aggregate idiosyncratic component

$$\bar{\epsilon}_{ht}^n = \frac{\sum_{i \in D_n} \omega_h^i \epsilon_{ht}^i}{\sum_{i \in D_n} \omega_h^i}$$

is bounded above by $n^{-1}(U_h^2 \sigma_h / L_h^2)$. Hence $\lim_{n \to \infty} \text{var}(\bar{\epsilon}_{ht}^n) = 0$.

On the other hand, the common components $y_{ht}^i - \epsilon_{ht}^i$ are not mutually orthogonal, so that, in general, their average will not vanish asymptotically. A positive lower bound for all but a finite number of cross-covariances between the common components is a sufficient (but not necessary) condition for this to be true.[3] It follows that for n large the weighted average

$$\bar{y}_{ht}^n = \frac{\sum_{i \in D_n} \omega_h^i y_{ht}^i}{\sum_{i \in D_n} \omega_h^i}$$

is approximately equal to $B_h^n(L) u_t$, where

$$B_h^n(L) = \sum_{i \in D_n} \omega_h^i A_h^i(L) / \sum_{i \in D_n} \omega_h^i$$

and $A_h^i(L)$ is the h-th row of the matrix $A^i(L)$. In other words, as stated in the following Proposition, the percentage of the total variance explained by the common component is close to unity.

Proposition 1. *As* $n \to \infty$, $\text{var}\left(B_h^n(L) u_t\right) / \text{var}(\bar{y}_{ht}^n) \to 1$.

There are two implications of Proposition 1. First, when the cross section is large, we can use sectoral averages to identify the dimension of the common shocks. Second, both the common shocks and the factor model can be identified and estimated by q cross-sectional averages, where q is the dimension of the common shock u_t.

In the next two Sections we will discuss these two implications in detail.

[3] Necessary conditions for the same model analysed here are given in Forni and Lippi (1995). Chamberlain (1983) provides necessary and sufficient conditions for the static version of the model where, however, the elements of the idiosyncratic component are not restricted to be mutually orthogonal.

3. Identification of the number of common shocks

Let us consider a data set concerning n sectors. Now take a partition consisting of s subsets $G_1, G_2, \cdots, G_s$, call $n_1, n_2, \ldots, n_s$ the number of elements in these sets and define the ms vector of aggregates:

$$Z_t = \begin{pmatrix} \sum_{i \in G_1} y_t^i / n_1 \\ \sum_{i \in G_2} y_t^i / n_2 \\ \vdots \\ \sum_{i \in G_s} y_t^i / n_s \end{pmatrix} \qquad (2).$$

Proposition 1 implies that, if $n_1, \cdots, n_s$ are large, the idiosyncratic components are negligible so that Z_t has approximately a (possibly infinite) moving average representation driven by u_t, say $C(L)u_t$. Hence, if $C(L)$ has maximum rank q, the spectral density of Z_t, $f_Z(\lambda) = C(e^{-i\lambda})C(e^{i\lambda})'$, will have reduced rank, equal to q, almost everywhere in the interval $[0, \pi)$.

Unfortunately, no standard tests for the rank of a spectral density matrix are available. Moreover, in the present context a further difficulty arises. As long as $n_1, \cdots, n_s$ are finite, the idiosyncratic component does not disappear completely and the smallest $sm-q$ eigenvalues of $f_Z(\lambda)$ are not exactly zero, which makes the rigorous definition of a null hypothesis problematic. For these reasons, as an alternative to a formal test, we propose the following 4-step procedure.

STEP 1 Select randomly l different partitions of the sectors in the data set and compute the corresponding vectors $Z_t^j, \quad j = 1, \cdots, l$.

STEP 2 For each j, compute the spectral density of Z_t^j, and decompose it in the following way:

$$f_Z(\lambda) = P(\lambda)D(\lambda)\overline{P(\lambda)}'$$

where $D(\lambda)$ is a diagonal matrix with eigenvalues

$$[\mu_1(\lambda), \cdots, \mu_{ms}(\lambda)]$$

on the principal diagonal and

$$\text{rank} D(\lambda) = \text{rank} f_Z(\lambda).$$

The latent roots $\mu_1(\lambda), \cdots, \mu_{ms}(\lambda)$ are the spectra of the dynamic principal components of Z_t (see Brillinger 1981).

STEP 3 Order the $\mu_k(\lambda)$'s in such a way that $\int_0^\pi \mu_k(\lambda)d\lambda > \int_0^\pi \mu_{k+1}(\lambda)d\lambda$ and compute the ratio:

$$R_r^2 = \frac{\int_0^\pi \sum_{k=1}^r \mu_k(\lambda)d\lambda}{\int_0^\pi \sum_{k=1}^{ms} \mu_k(\lambda)d\lambda} \qquad (3)$$

for $r = 1, \cdots, ms$. R_r^2 gives us the percentage of the trace of the covariance matrix of Z_t^j accounted for by the first r principal components.

STEP 4 Set $q = r$ if $R_{r-1}^2 < .95$ and $R_r^2 > .95$ for all the l experiments.

In the empirical application of this paper we proceeded as follows. We reordered sectors by extracting randomly without replacement natural numbers from 1 to 450 to form the sequence i_k, $k = 1, \ldots, 450$. Then we partitioned the sectors in three groups of 150 sectors each by taking $G_1 = \{i_1, \ldots, i_{150}\}, \ldots, G_3 = \{i_{301}, \ldots, i_{450}\}$. We repeated the experiment 50 times to get the vectors Z_t^j, $j = 1, \ldots, 50$. Since we have two variables we have six aggregates forming the vector Z_t^j.

Notice that, as stated by Proposition 1, we could have constructed Z_t by taking weighted averages rather than simple averages. The weighted procedure is more appropriate when treating data sets with a smaller cross-sectional dimension, since weights can be chosen so as to minimize the expected variance ratio between the idiosyncratic and the common component (see Forni and Reichlin 1996a for details). As illustrated by our diagnostic later on, the data set analysed here is sufficiently large so that there is no need for this complication.

Figure 1 reports the estimated R_r^2 for $r = 1, \cdots, 6$ and for all experiments. The spectra were estimated using a Bartlett window with lag window size equal to seven. For all experiments, the result is that 2 principal components are sufficient to capture more than 95 % of the total variance. From this we conclude that there are two common shocks to our 450 sectors.

The methodology described above can easily be adapted in order to identify the rank of $f_Z(\lambda)$ at a given frequency λ: we have only to reorder the latent roots according to their size at frequency λ and fix $q(\lambda)$ equal to r when the explained variance is greater than 95 % of the total variance. Frequency zero is of particular interest since if the first $p < q$ principal components are sufficient to capture all the variance at frequency zero, then p shocks should be modeled as permanent and $q - p$ as transitory.[4]

The results from this frequency-by-frequency test are shown in Figure 2 which reports the ratios

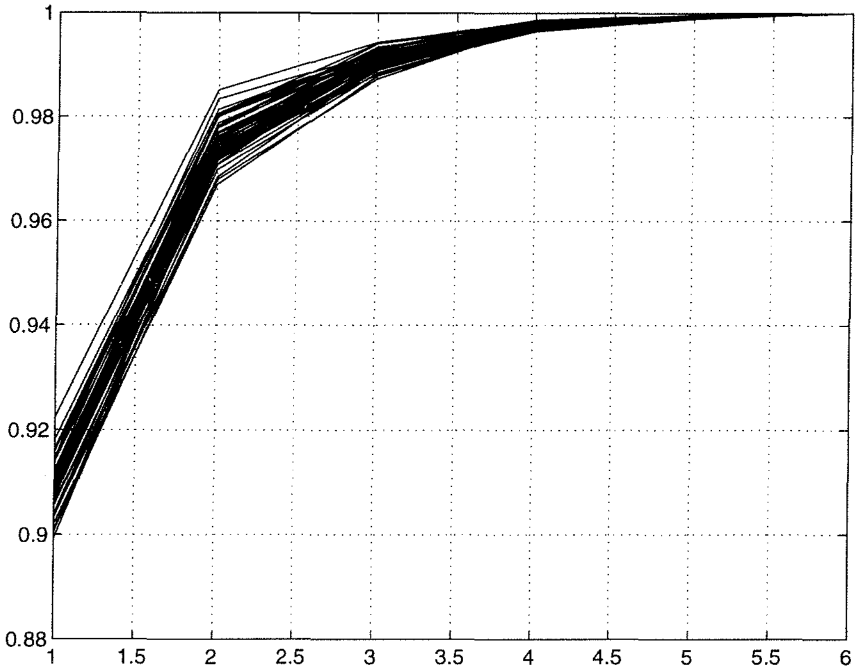
$$\frac{\sum_{k=1}^{2} \mu_k(\lambda)}{\sum_{k=1}^{ms} \mu_k(\lambda)} \quad \text{and} \quad \frac{\mu_1(\lambda)}{\sum_{k=1}^{ms} \mu_k(\lambda)}$$

at each λ for the 50 experiments.[5]

[4] This criterion is a simplified version of the cointegration test proposed by Phillips and Ouliaris (1988). The difference between our criterion and the latter test is that we do not require the construction of confidence bands. Of course, Phillips and Ouliaris's test could also be used; notice however that confidence bands based on asymptotic results are not very reliable when the number of observations is small.

[5] In our case the first two principal components are the same at all frequencies, so that reordering is not needed.

Figure 1: Variance of Z_t explained by the first 6 principal components (l=50 experiments)



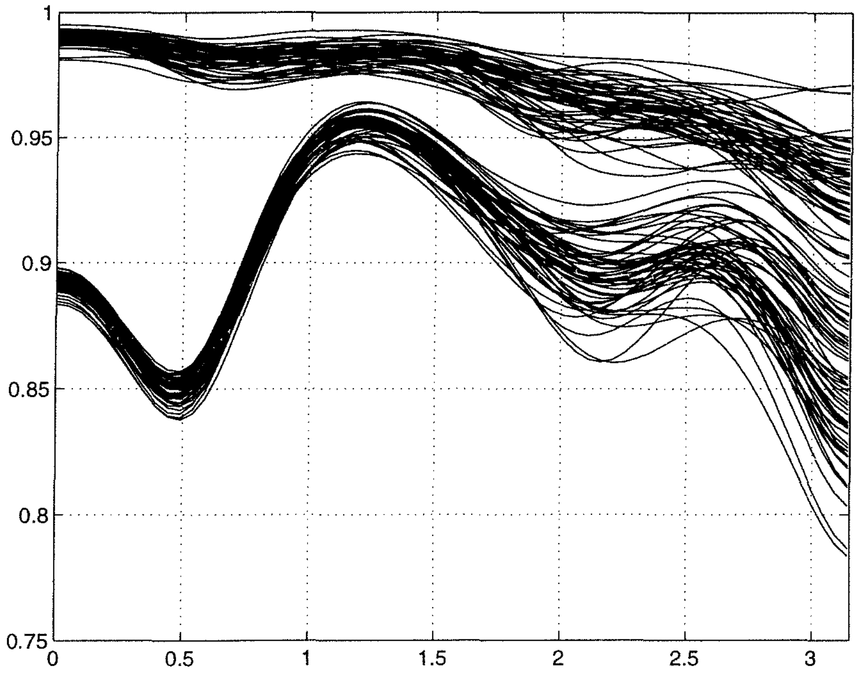
Observe that the variance explained by the first two principal components is similar across frequencies and that results are robust across experiments. Observe also that, while we only need one shock to explain business cycle frequencies, we need at least two to account for low frequency dynamics. This indicates that modelling the two shocks as permanent and transitory, as for instance in Blanchard and Quah (1989), is not appropriate for the US manufacturing sector.

4. Identification and estimation of the common shocks

Another important consequence of Proposition 1 is that we can recover the common shocks by taking any vector of q weighted averages Y_t, and identifying and estimating a VAR or VARMA model for Y_t.[6]

[6] Connor and Korajczyk (1988) have suggested an estimation method, which, like ours, is based on a law of large numbers result. They use a result in Chamberlain and Rothschild (1983) which shows that the common factor tends asymptotically to the principal components of the variables, to estimate the common factor through principal components. Their method, however, is only developed for static models. Moreover, it

Figure 2: Variance of Z_t explained by the first two principal components at different frequencies



Let us ignore the residual idiosyncratic component which is still present in Y_t and assume, for notational simplicity, that Y_t is an exact linear combination of the present and the past of the common shocks. Then we can write:

$$Y_t = \hat{A}(L)\hat{u}_t \tag{4}$$

where $\hat{A}(0)$ is upper triangular, $\det \hat{A}(L)$ does not vanish within the unit circle in the complex plane and $\Sigma_{\hat{u}} = I$. If we limit ourselves to the set of fundamental representations of Y_t[7] , any admissible orthonormal representation of Y_t, that is a representation

$$Y_t = A(L)u_t \tag{5}$$

with $\Sigma_u = I$, is such that

is computationally more burdensome than ours.

[7] As argued by Lippi and Reichlin (1993), in structural VARs, the hypothesis of fundamentalness has no economic justification. We analyse this issue in the context of factor models in large cross-sections in Forni and Reichlin (1996a).

$$u_t = R' \hat{u}_t$$

and

$$A(L) = \hat{A}(L)R,$$

where R is an orthonormal matrix.

Correspondingly, if the common component in the disaggregated model (1) can be represented as $\hat{A}^i(L)\hat{u}_t$, we have infinitely many representations $A^i(L)u_t$, with $u_t = R'\hat{u}_t$ and $A^i(L) = \hat{A}^i(L)R$. Hence, both the common shocks and the disaggregated factor model are identified by selecting an orthonormal matrix R, in the same way as in the structural VAR literature. An important feature of our estimation procedure is that, since the common shocks are estimated by specifying a VAR or VARMA model for the aggregate variables, we can use the same identification strategies used for structural VAR's to achieve identification in the factor model.

In our two common shocks case, the orthonormal matrix R can be represented as a function of a single rotation parameter, $\theta \in [0, \pi)$:

$$R(\theta) = \begin{pmatrix} sin(\theta) & cos(\theta) \\ -cos(\theta) & sin(\theta) \end{pmatrix}$$

so that identification is reached by selecting a particular value of θ.

Figure 3 reports 15 sets of impulse response functions corresponding to different values of θ and for an estimated VAR(2)[8] . Obviously, for each different rotation we have a different structural model with its implied economic interpretation.

The eighth one, corresponding to $\theta = \pi/2$, is the same as the traditional triangular identification scheme originally proposed by Sims (1980) since it corresponds to $R = I$; the Figure corresponding to $\theta = 2.1$ shows results for the identification scheme proposed by Blanchard and Quah (1989) where one of the shock is restricted to have long-run neutrality on output.

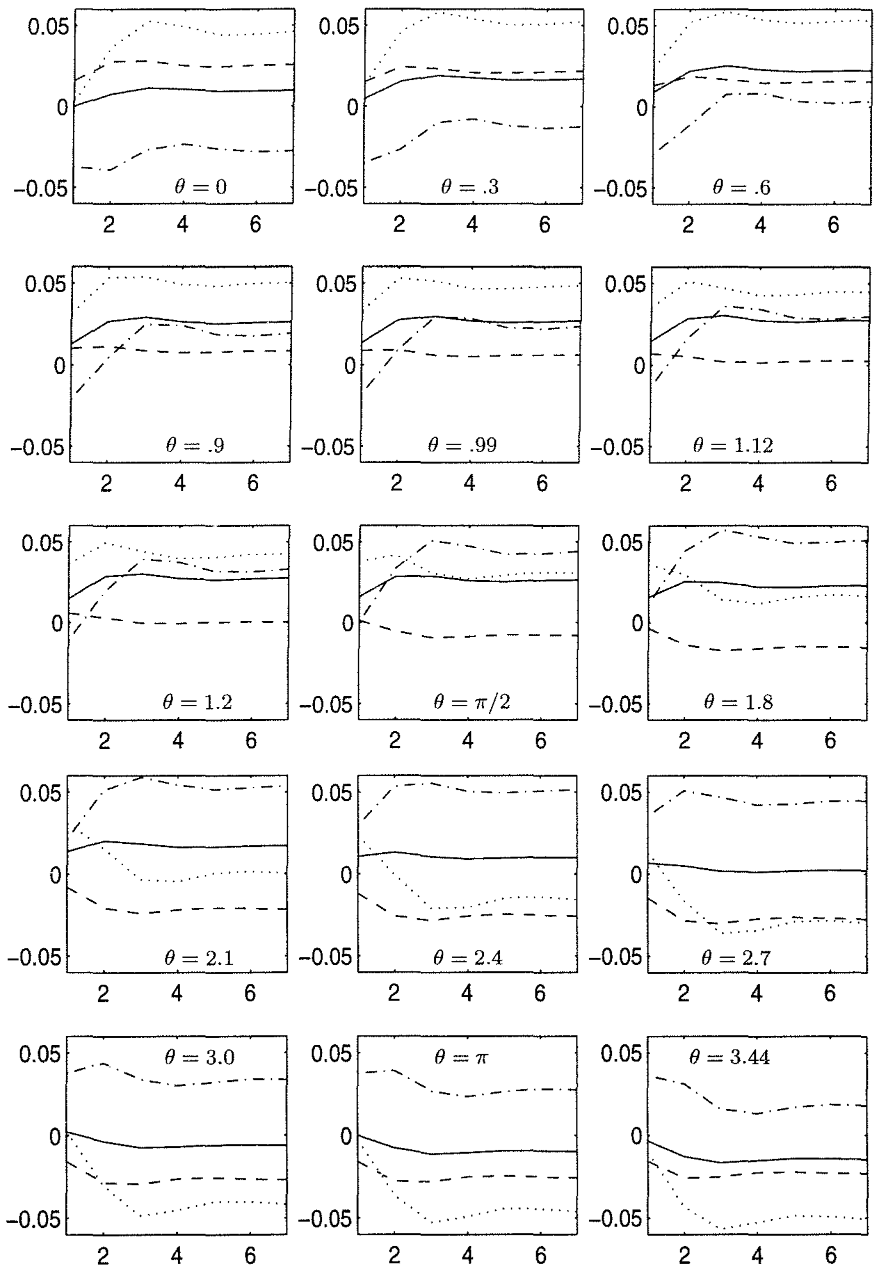
Here we propose to choose the θ for which one of the shocks, labeled technology, has minimum absolute sum of negative values. In the absence of precise theoretical restrictions, this assumption seems less controversial than the common one of long-run demand neutrality and, as we have said, corresponds to the observation that technological shocks are generally positive.

To clarify our identification criterion, let us reintroduce the means of u_t and Y_t explicitely by setting $\tilde{u}_t = u_t + \mu_{\tilde{u}}$ and $\tilde{Y}_t = Y_t + \mu_{\tilde{Y}}$. We then have:

$$\tilde{Y}_t = A(L)\tilde{u}_t = A(1)\mu_{\tilde{u}} + A(L)u_t$$

[8] The lag order has been selected using Akaike information criterion. Standard tests indicate that the levels of the variables in Y_t are not cointegrated.

Figure 3: Impulse response functions for different values of θ



shock u^T on productivity (solid line); shock u^T on output (dotted-dashed line); shock u^{NT} on productivity (dashed line); shock u^{NT} on output (dotted line).

If the levels of the variables in Y_t are not cointegrated as is the case in our data set, $A(1)$ is invertible and $\mu_{\tilde{u}} = A(1)^{-1}\mu_{\tilde{Y}}$. From the choice of θ we can identify u_t and $A(L)$. From $A(L)$ we can then identify $\mu_{\tilde{u}}$ and therefore $\tilde{u}_t$. Now let us call $\tilde{u}_t^T$ the sample realization of the technology shock and $\mathcal{N}$ the set of integers t such that $\tilde{u}_t \leq 0$. Then our identification strategy is to choose θ so as to minimize[9]

$$g = \sum_{t \in \mathcal{N}} \mid \tilde{u}_t^T \mid$$

The technology shock identified in this way is reported in Figure A1 in Appendix 1. Notice that there are three negative realizations in 1974, 1979 and 1981. The first two correpond to the oil shocks.

The impulse response functions are reported in Figure 4.

Given our identification restrictions, the picture emerging from aggregate estimates is one whereby the common technological shock has a long-run positive effect on both output and productivity, but affects output negatively in the short-run. This suggests that when technological innovations occur, firms reorganize their production process so that in the first year output will grow less than on average. Productivity, however, even in the first year, grows faster than on average because of the immediate impact that the technological innovation has on the demand of labor.

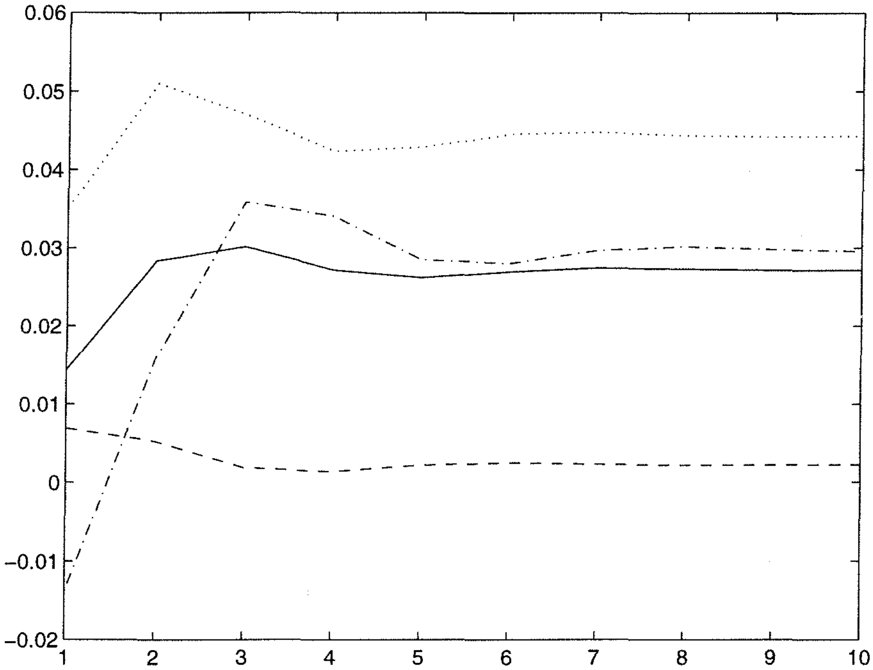
Variance decomposition results indicate that the technological component explains the main bulk of the variance of productivity (87%) and 51% of the variance of output. The result implies that, for aggregate productivity, cyclical fluctuations originating from a common shock are almost all due to technological innovations.

It should be observed that the shape of the impulse of the technological shock on both output and productivity reproduces the S-shape that has been used in the literature to describe slow diffusion of the innovation throughout the economy (e.g. Griliches 1957, Mansfield 1973, Jovanovic and Lach 1989 and 1990). This can be taken as an informal support for our method of identification of the technological shock[10] . Further support

[9] Since the variance of the technological shocks is not affected by rotation, under normality the expected absolute sum of negative values is minimized when the mean of the technology shocks is maximised. In practice, however, maximization of the sample mean of the shock and minimization of g will give different results. In our sample, the former criterion gives $\theta = .99$ (see Figure 3) as against $\theta = 1.12$. One could also consider the minimization of the frequency of negative values of the technology shocks. This criterion, however does not give unique results since frequency is a discrete variable. In our sample the minimum frequency of negative shocks is three and it is reached in the intervals $.66 < \theta < .96$ and $1.12 < \theta < 1.26$.

[10] In the present exercise we obtain the S shape as an empirical result. An alternative strategy would have been to follow Lippi and Reichlin (1994a, 1994b) and identify the technology shock as the shock with an S-shaped impulse by minimizing the distance between the empirical impulse and an S-shaped function.

Figure 4: Impulse response functions - our identification



u_t^{NT} on output (dotted-dashed line), u_t^T on productivity (solid line), u_t^{NT} on productivity (dashed line), u_t^T on output (dotted line).

comes from the correlation coefficient between the shock we have identified as technology and the real interest rate which is positive and highly significant: as predicted by growth theory, a shift in the production function caused by an increase in total factor productivity has a positive effect on the steady state value of the real interest rate.

5. Estimation of the factor sectoral model

Having estimated the common shocks, we can finally estimate the disaggregated model (1). Two alternative strategies can be followed. The first consists in a regression of the sectoral variables directly on the estimated shocks. The second consists in using the aggregates as regressors, i.e. in estimating the model

$$Y_t^i = B^i(L)Y_t + \epsilon_t^i$$

and obtain an estimate for $A^i(L)$ via the relation

$$A^i(L) = B^i(L)A(L).$$

Clearly, the two procedures imply different dynamic specification of (1). We have tried both strategies and obtained similar results. Here we report results only for the latter method with $B(L)$ specified as a polynomial matrix of degree two in L. This method is preferable for both theoretical and practical reasons. First, when the same number of lagged responses are included, it gives a slightly better overall fit, as measured by the ratio of the sum of explained variances to the sum of total variances, for both output and productivity. Secondly, while both methods are affected by an errors in variables problem since in practice the idiosyncratic component does not completely die out in the aggregate, the problem is further aggravated for the second method where the regressors are not the true shocks, but only consistent estimates of the true shocks. Thirdly, the dynamic specification of the former method implies a finite MA structure for the aggregate model, which is inconsistent with our VAR(2) specification.

Notice that in both cases the explanatory variables are the same for all equations so that the model can be estimated consistently, by OLS, equation by equation.

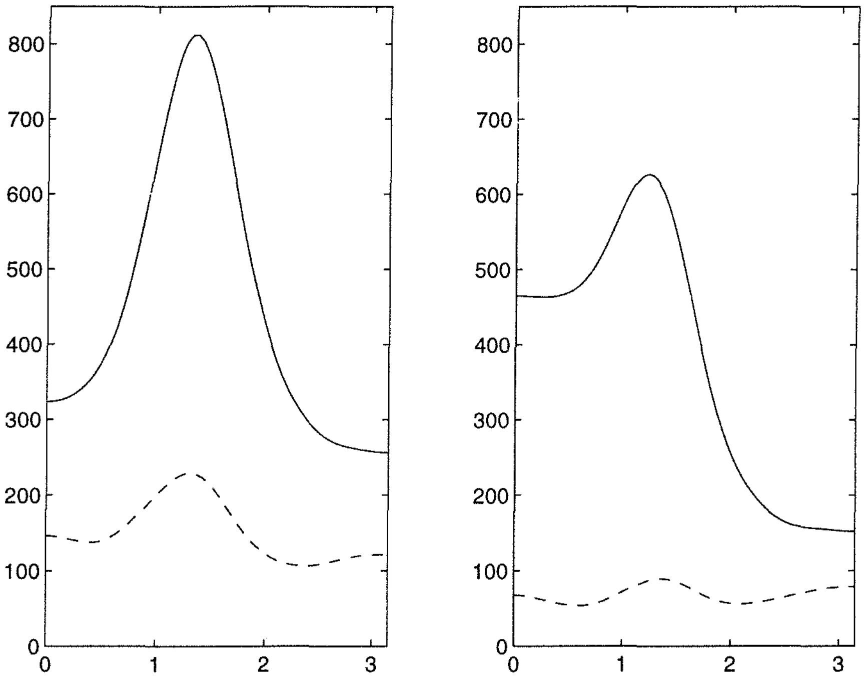
5.1 The relative size and the shape of the common and idiosyncratic components

Let us first assess the relative importance of the common and idiosyncratic components and define an overall measure of fit as the ratio of the sum of the variances of the common components to the sum of the total variances of the variables. This, which is the weighted mean of the sectoral R^2 with weights propotional to the total variances, gives us a percentage of 41% for output and of 29% for productivity. These figures are lower than in previous studies (Horvath and Verbrugge 1996 have estimated the center of the distribution of empirical results from different studies to be 55-60%). Notice, however, that according to our argument, the weight of the common component should decrease with the level of disaggregation and that the 4-digit level of our study is a finer disaggregation level than that on which the cited results are usually based.

Overall variance ratios, however, are not sufficiently informative about the role of idiosyncratic shocks for business cycle fluctuations. For this we must look at the distribution across frequencies of the variances of the common and sectoral components. This is captured by the sum of the spectra for the common and the idiosyncratic component (Figure 5).

Notice that, for both variables, while the common component has a typical business cycle shape with a peak corresponding to a period of just over four years, the bulk of the variance for the idiosyncratic component is at the

Figure 5: Sum of the spectra of the common and idiosyncratic components of output (a) and productivity (b)



high frequencies. We should conclude that the business cycle features of the data are mostly explained by economy-wide shocks and that, although the sectoral dynamics is more sizeable than the economy-wide one, it cannot account for cyclical behaviour of output and productivity.

5.2 The impact of technology shocks: dynamic "complements" and "substitutes"

Reallocation effects should not only be captured by the weight of the idiosyncratic component in the total variance, but also by negative comovements of sectoral output and productivity generated by economy-wide shocks. Technology shocks may generate negative comovements because certain industries diminish in importance relative to others (are substituted by others) and demand shocks may have negative effects reflecting changes in the structure of demand produced by an increase in overall income. Positive comovements generated by both type of shocks, on the other hand, may be present at the high and business cycle frequencies because of input-output relations and in the long-run because of complementarities in economic growth.

In order to analyse the weight of the substitution or reallocation effects in the total variability of output and productivity, we need to look at the correlation structure of the impulse response functions associated to the two aggegate shocks.

For simplicity of exposition and "par abus de langage" we call *substitution effects* the negative sectoral comovements generated by aggregate shocks and *complementary effects* the positive sectoral comovements. Of course, these effects do not have anything to do with the entries of a Slutsky matrix.

A measure of complementary and substitution effects can be constructed from the estimates of the spectral density of our panel of sectoral output growth rates and computing the ratio between the sum of the negative values of the co-spectra and the sum of its positive values for different frequencies. This gives us an index of the relative importance of positive covariances amongst sectors relatively to negative covariances. We first calculate, from the estimated coefficients, the implied spectral density matrix of the common components of sectoral output. The real part of the off-diagonal elements of this matrix are the cospectra between the different sectors which give us information about the cross-covariances between sectors at all frequencies. The cospectrum is defined as:

$$s_{ij}(\lambda) = \sum_{k=-\infty}^{\infty} c_{ij}^{k} cos(\lambda k)$$

where i, j are indexes for sectors and c_{ij}^{k} is the covariance at lag k between the common (technological and non) component of output of sector i and sector j. Let us now decompose $s_{ij}(\lambda)$ as

$$s_{ij}(\lambda) = s_{ij}(\lambda)_- + s_{ij}(\lambda)_+$$

where

$$s_{ij}(\lambda)_- = [s_{ij}(\lambda) - \mid s_{ij}(\lambda) \mid]/2$$

and

$$s_{ij}(\lambda)_+ = [\mid s_{ij}(\lambda) \mid + s_{ij}(\lambda)]/2$$

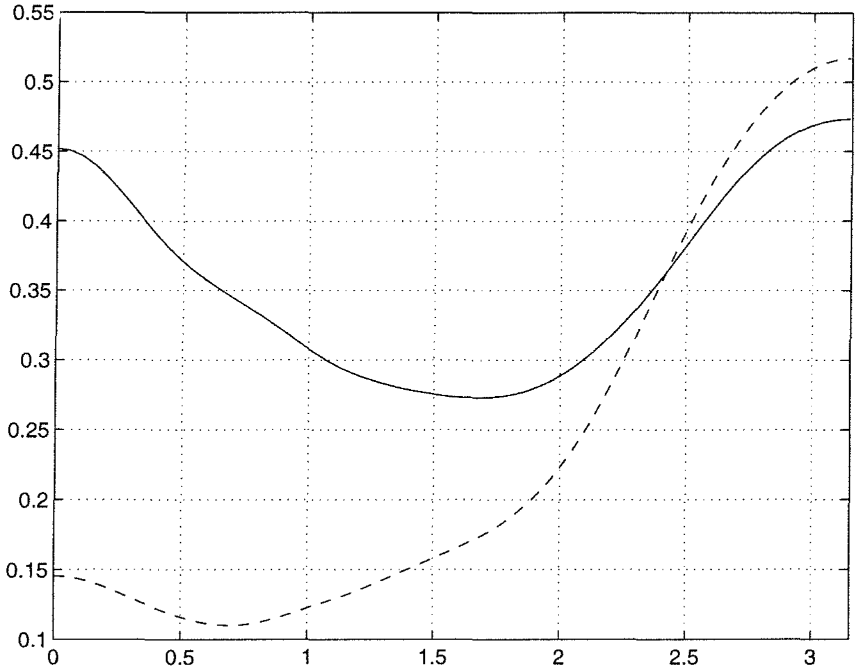
From this we define a measure of the substitution effect of the common shocks as the ratio:

$$S(\lambda) = -\frac{\sum_{i,j} s_{ij}(\lambda)_-}{\sum_{i,j} s_{ij}(\lambda)_+} \tag{6}$$

where the $s_{ij}(\lambda)_-$'s are the negative cospectra while the $s_{ij}(\lambda)_+$'s are the positive cospectra, both at frequency λ. Notice that

$$\sum_{i,j} s_{ij}(\lambda)_+ \sum_{i,j} s_{ij}(\lambda)_+ \geq 0$$

Figure 6: Substitution index: common technology and non technology shock on output



technology shock (solid line), non-technology shock (dashed line)

for any λ, since it is equal to the spectrum of $\sum_i Y_t^i$. It follows that $0 \leq S(\lambda) \leq 1$.

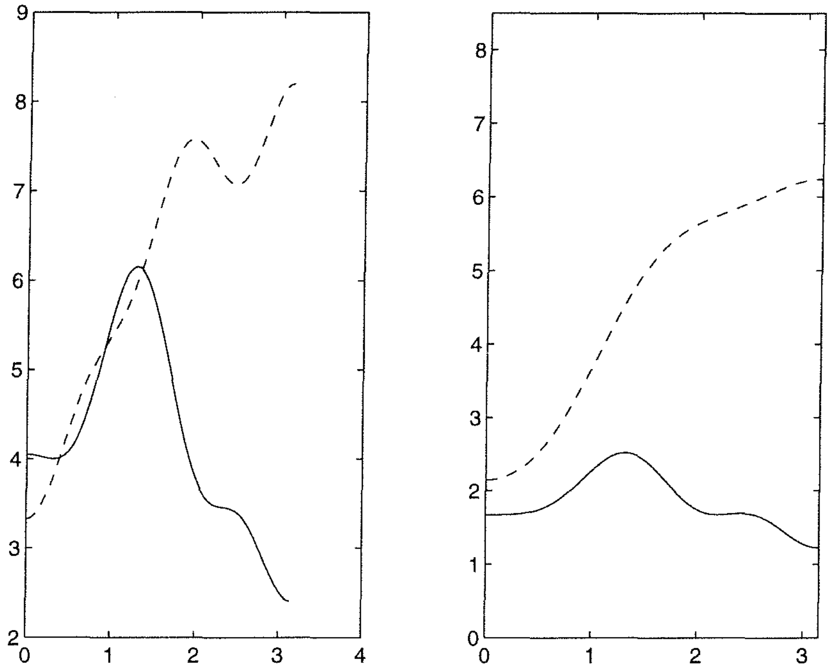
Figure 6 reports the values of $S(\lambda)$ for the technology shock (solid line) and the non-technological shock (dashed line).

The picture that emerges is one where technological innovations generate strong negative comovements at low and high frequencies, while they induce positive comovements at business cycle frequencies. The other shock has strong substitution effects in the short run, but generates mainly complementary fluctuations in the long-run.

Figures 7a and 7b report $-\sum s_{ij}(\lambda)_-$ and $\sum s_{ij}(\lambda)_+$ for the technology shock and the non-technology shock.

The Figures illustrate nicely the business cycle features of our data: all the series of the sums of the positive cospectra have peaks at business cycle frequencies, while the series of the negative cospectra are rather flat. Moreover the business cycle is partly real since the technology shock generates positive cospectra at a period of about four years.

Figure 7: Absolute sum of positive (dashed lines) and negative (solid lines) cospectra



technological component (a), non technological component (b)

5.3 Technology, Investment and Growth

What is the mechanism that links technological change and growth?

Some light on the propagation mechanism may come from the identification of the sectors with the strongest correlation between output growth rates and the common technological component. Table 1 describes the 20 sectors with the highest percentage of total output variance accounted for by the technological component.

These core sectors are mainly in the industrial machinery and equipment goods group and in primary and fabricated metals, i.e. they are concentrated in sectors producing investment in capital goods and their inputs. This result is consistent with what noticed by De Long and Summers (1991 and 1992) who found a strong link between equipment investment and output growth for a broad cross-section of nations; they interpreted this as indicating the presence of externalities in the activity of the equipment investment sectors. Our results, as De Long and Summers's, suggest a view of the propagation of technological innovations which is quite different from that suggested by a real business cycle-Solow growth model. In that framework, the technological innovation is identified with total factor productivity and it is purely exogenous. On the contrary, a strong positive correlation between technological innovations and the rate of growth of those key sectors says that since new technology requires new capital goods, it is embodied in capital and it propagates through investment.

Table 1: Sectors with the highest percentage of total variance of output accounted for by the technological component

Sectors	SIC code	R^2
Machine Tool Accessories*	3545	.67
Gray and Ductile Iron Foundries	3321	.66
Air and Gas Compressors*	3563	.65
Ball and Roller Bearings*	3562	.65
Carbon and Graphite Products	3624	.65
Power Transmission Equipment, n.e.c.*	3568	.64
Hardwood Veneer and Plywood	2435	.62
Truck Trailers	3715	.61
Internal Combustion Engines*	3519	.60
Bolts, Nuts, Rivets and Washers	3452	.60
Cement, Hydraulic	3241	.59
Plastic Materials and Resins	2821	.59
Brick and Structural Clay Tile	3251	.58
Iron and Steel Forgings	3462	.58
Machine Tools, Metal Forming Types*	3542	.58
Upholstered Household Furniture	2512	.57
Special Dies, Tools, J'gs & Fixtures*	3544	.57
Blast Furnaces and Steel Mills	3312	.57
Aluminium Die Casting	3363	.55
Sawmills and Planing Mills, General	2421	.55

Starred sectors belong to the broad classification "Industrial Machinery and Equipment".

6. Summary and conclusions

This paper has proposed a methodology for identifying and estimating the contribution of technological innovations in a sample of a large cross-section and time series observations. The data used are output and productivity for 450 manufacturing sectors in the US from 1958 to 1986.

We exploit law of large numbers results to identify the vector of the common shocks by an average quantity. By applying this method and through dynamic principal component analysis we are then able to identify and estimate two common shocks to our data set.

We then identify the technological shocks as those for which the sum of the negative realizations is minimized. This method emphasises the least controversial feature of technological innovations, i.e. that technological innovations are mostly positive.

The ensemble of the empirical results show an interesting picture of the business cycle in manufacturing. First, we found that at least two economy-wide shocks are needed to explain the common dynamics and that, although the technological shock accounts for at least 50% of the aggregate dynamics of output, it cannot by itself explain dynamics at business cycle frequencies. While it is true that technology is an important source of fluctuations, our empirical results do not support the first generation of real business cycle models in which dynamics is driven exclusively by technological innovations. Second, we found that sector-specific shocks explain the main bulk of total variance (60% for output and 70% for productivity). However, sector-specific shocks generate mainly high frequency dynamics so that the idiosyncratic component, unlike the common one, has no recognizable business cycle pattern. This shows that the business cycle is an economy-wide phenomenon and there is no purely sectoral cycle: sectoral technology shocks might be important, but do not generate cycles. Third, we find that a decomposition into a transitory and a permanent component is not an appropriate characterization of dynamics for our data set since rank reduction of the common dynamic component is observed at business cycle frequencies, but not at zero frequency.

A more detailed analysis of the common component which identify separately the behaviour of positive and negative comovements, shows that, as indeed in the NBER definition of the business cycle, the latter is characterized by *positive* sectoral comovements. This is shown by a peak for positive comovements of output at business cycle frequencies in both the technological and non technological component. In the long-run, on the other hand, the technology shock generates a lot of substitution effects (negative comovements), while the other shock has mainly complementary effects (positive comovements).

Finally, we find that technological shocks are strongly correlated with the growth rates of the investment in machinery and equipment sectors and their inputs. This result is consistent with that of De Long and Summers (1991 and 1992), who claim that technology is embodied in the investment in capital goods sectors which then affects growth through strong positive externalities.

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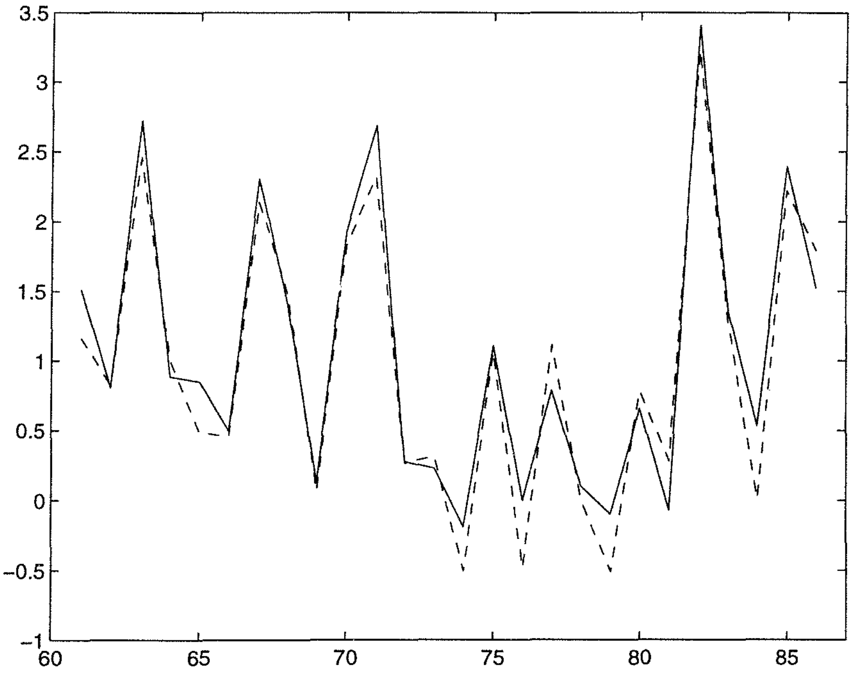
APPENDIX 1

Specification Analysis

A. Figure A1 shows the technology shocks derived from the estimation of the VAR(2) (solid line) and the technology shocks derived from the estimation of the same model for the sample of the odd sectors (dashed line). These two processes are almost identical. This result is very comforting for our analysis: first, if two alternative aggregates give us the same estimate of the common technological shock, this justifies our procedure of estimating the common shocks by aggregate quantities; second, the fact that half of the sample produces the same result as in the all sample indicates that there cannot be more than two common shocks.

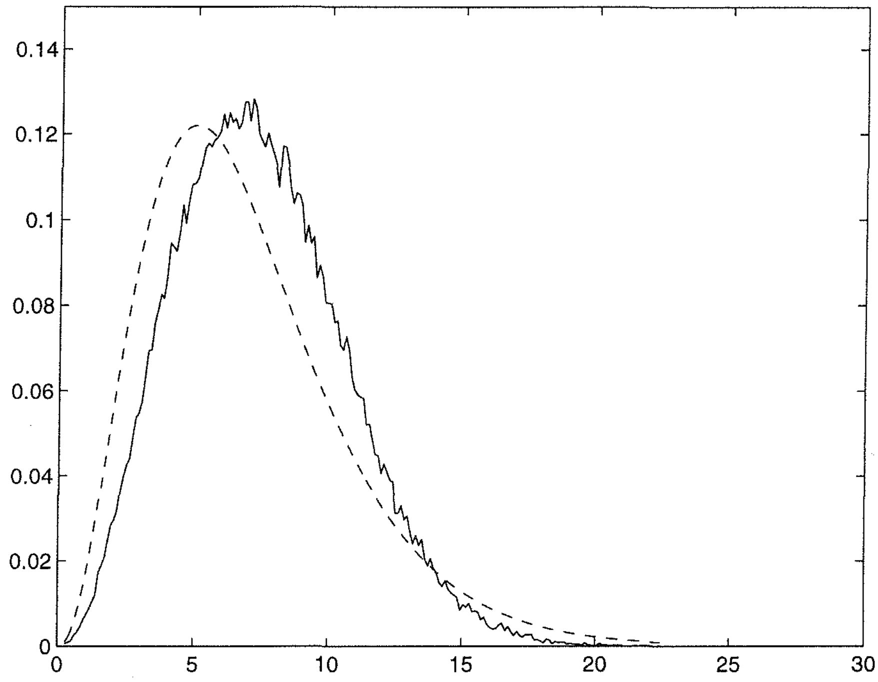
Figure A1. Estimated common technological shock



estimated using the average of all sectors (solid line), estimated using the average of odd sectors only (dashed line).

B. To verify the orthogonality between the sector-specific components we performed a Q test on pre-whitened residuals from the sectoral regressions. For each pair of sectors we computed $Q = T\sum_{k=-3}^{3} r_k^2$, where T is the time dimension of the residuals and r_k^2 denotes the sample cross-correlation of ϵ_{ht}^i and $\epsilon_{h,t-k}^i$. Figure A2 compares the distribution of the Q statistic for the

Figure A2: Distribution of the Q-test statistic for residuals of ouput regressions



idiosyncratic components of sectoral output and the distribution obtained with 450 i.i.d. white noises randomly generated.

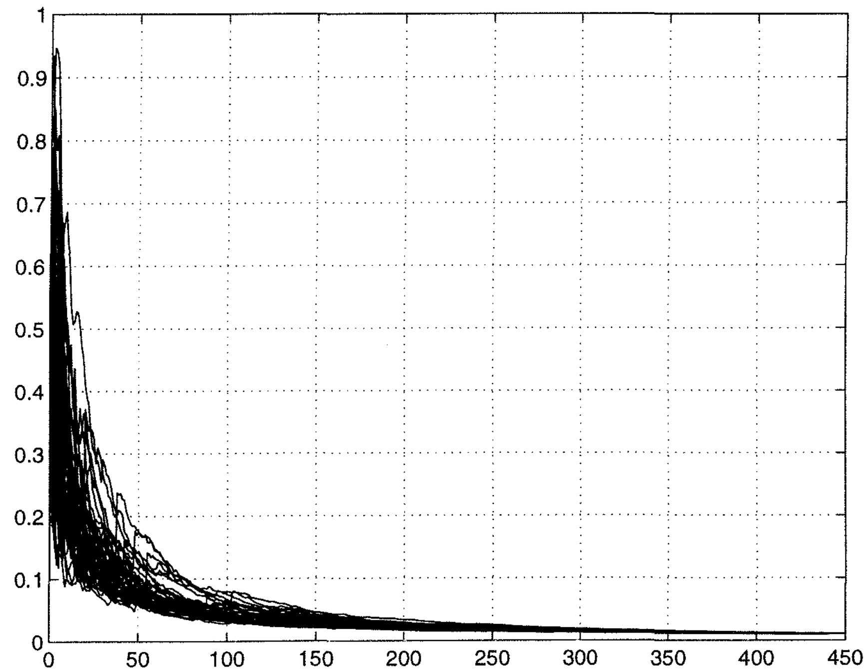
From the comparison we conclude that there is no evidence of large cross-correlations between the estimated idiosyncratic components. Similar results hold for productivity.

C. To verify whether the idiosyncratic component has died out in the aggregate we estimated the ratio of the variance of the aggregate idiosyncratic component to that of the aggregate variable. Call s_h^i the estimated variance of the idiosyncratic component of y_{ht}^i and c_{ht}^i the estimated common component of y_{ht}^i, and $\hat{\sigma}_h$ the sample variance of $\sum_{i=1}^m c_{ht}^i$. Under the orthogonality assumption the above ratio can be estimated by

$$\frac{\sum_{i=1}^n s_h^i}{\hat{\sigma}_h + \sum_{i=1}^n s_h^i}.$$

Results are encouraging since we obtain ratios of .01 for output and .05 for productivity. In order to check how rapidly the variance of the idiosyncratic component goes to zero for increasingly larger aggregates, we performed the following exercise. First, we reordered sectors by extracting randomly without replacement natural numbers from 1 to 450 to form the sequence i_k, $k = 1, \dots, 450$. Second, we computed the above ratio for the sets $\{i_1, \dots, i_n\}$, $n = 1, \dots, 450$. Lastly, we repeated the experiment for 50 different reorderings. Figure A3 illustrates the results for the sample of sectoral output.

Figure A3: Ratios of the variance of the idiosyncratic component to the variance of the sub-aggregates - 50 experiments - output data



APPENDIX 2

Data sources and data treatment

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.

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