Does Switzerland share common business cylces with other European countries?

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1. INTRODUCTION

One salient characteristic of aggregate output fluctuations among countries is their positive correlation. These findings have been well documented by several authors using a variety of methods, for instance CANOVA and DELLAS (1993), BACHUS, KEHOE and KYDLAND (1993), ARTIS and ZHANG (1995), and more recently LUMSDAINE and PRASAD (1997). From a technical point of view, this issue has often been expressed in terms of comovements of international time series data, for instance industrial production on a monthly basis, or real quarterly GDP.

The existence of common components in economic activity across countries has given birth to mainly two kinds of economic research : for a long time, international business cycle research focused on the transmission mechanism of aggregate fluctuations, whereas applied econometrics concentrated on the potential gain in efficiency that may result in a multivariate framework with time series sharing common properties. In the first research area, economic interdependencies of goods or assets markets, the presence of common exogenous disturbances, the relative weight of various shocks (with foreign or domestic origin) were analysed and proposed as a valid explanation of common economic fluctuations. In the second area, a *common feature hypothesis* has been developed by ENGLE and KOZICKI (EK) (1993) for instance, and used to test and illustrate the presence of common components shared between important time series for various countries. Although the EK approach has been quite strongly criticized (CUBADDA, 1999), it nevertheless represents a means for testing the existence of a common synchronous international business cycle (more accurately EK have tested for common *binational* business cycles), in a difference-stationary framework, i. e. by taking the first difference of the time series. KOOPMAN and HARVEY (KH) (1997) have also proposed some interesting tools to account for common components in time series. They suggest the use of multivariate structural time series models to describe cyclical movements of economic time series arising out of a common business cycle, in a trend-stationary framework, i. e. within a decomposition in unobserved components, as trend and cycle for instance. It

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should be stressed that within these two approaches, the degree of synchronization between two cycles is not only measured by the correlation (or the cross-correlation as in ARTIS and ZHANG), but by more specific criteria, that imply much deeper links and the existence of common short term influences and a common source of stochastic disturbances.

Between 1991 and 1996, the Swiss economy experienced a six-year output stagnation period. Such a long slowdown in economic activity raises a lot of questions. The present paper does not directly enter the debate on what generates and propagates international business cycles and does not intend to list all of the potential macroeconomic shocks that accounted for Switzerland's six year stagnation. It concentrates however on two issues, namely:

1. Can we admit that a substantial fraction of short term economic fluctuations in Switzerland is not country-specific but shared with other EU countries?
2. In case there is a strong EU-common component in Swiss aggregate output fluctuations, what kind of interpretation should we give to the six-year stagnation of the Swiss economy between 1991 and 1996?

In addition to testing with which countries Switzerland shares a common business cycle component – an issue that could be of interest in itself – the answers to these two questions have various economic implications. From a policy perpective, if business cycle fluctuations in Switzerland are highly positively correlated with other European countries and moreover *common* to them (using various definitions of *common features*), then the six years' stagnation should be attributed to some kind of temporary slowdown of the medium-long run evolution of output, rather than to six years of a dampening business cycle. Another potential explanation could be an overreaction of the Swiss business cycle during those years, which should be explained. Some comments will be devoted to these important issues later on.

The first section of this paper discusses the definition of business cycle components in time series by focusing on some properties of the official seasonally ajusted quarterly GDP figures released in the OCDE international database (*Main Economic Indicators*). Section two introduces and applies the *serial correlation common feature* test developped by EK to the quarterly GDP data presented in table 1. Section three presents and estimates univariate and multivariate *trend-plus-cycle* structural models, mainly introduced in HARVEY (1989) and recently proposed in KH to take account of potential cyclical components and especially of common business cycle components. Section four comments the main results and presents some concluding remarks.

Our results were obtained using Econometric Views, version 2.0. (QUANTITATIVE microsoftware, 1996), PcFiml version 9.00 (DOORNIK and HENDRY, 1997), and STAMP version 5.0, for the structural time series models (KOOPMAN, HARVEY, DOORNIK and SHEPHARD, 1995).

2. PRESENTATION OF THE DATA AND VARIOUS COMMENTS

Table 1 sets out the various countries selected for our analysis, the sample periods available and the Ljung-Box Q-statistics for the first difference of the log of the seasonally adjusted quarterly GDP series (we use seasonally adjusted series, because for most countries, the raw data are not available for the quarterly real GDP). The Ljung-Box test is a simple global autocorrelation test, which is based here on the first 6 autocorrelations. For Switzerland, we use and adjust the "old" quarterly GDP series available over the period 1965:1/1979:4 (constant prices 1980) and use the "new" one available over the sample period 1980:1/1997:4 (constant prices 1990), in order to get a time series covering the whole sample 1965:1/1997:4. For Germany, we use the West Germany quarterly GDP available over the sample 1965:1/1994:4 for the test of serial correlation common feature and the quarterly GDP for the whole of Germany over the sample 1970:1/1997:4. For this last serie, we take into account explicitely the impact of the reunification with a correction for the level shift that arose in the first quarter of 1991, when estimating the *trend-plus-cycle* structural time series models (in univariate of multivariate contexts). "Eur15" sands out for the aggregated GDP of the Euro zone, that is the GDP of Belgium, Germany, Spain, France, Ireland, Italy, Luxembourg, the Netherlands, Austria, Portugal, Finland, plus the GDP of the United Kingdom, Sweden, Danemark and Grece.

Table 1: Quarterly GDP time series used, samples and Ljung-Box stat. first difference

Time series y_t	sample q	Ljung-Box, Q(6) $(1-B)\log y_t$
1. United States	q = 152; 1960:1/1997:4	25.78**
2. Canada	q = 144; 1962:1/1997:4	22.22**
3. Japan	q = 124; 1967:1/1997:4	59.22**
4. West Germany	q = 150; 1965:1/1994:4	64.29**
5. Germany reunif.	q = 112; 1970:1/1997:4	22.89**
6. France	q = 112; 1970:1/1997:4	18.30**
7. Italy	q = 112; 1970:1/1997:4	51.94**
8. United Kingdom	q = 152; 1960:1/1997:4	9.18
9. Spain	q = 112; 1970:1/1997:4	221.39**
10. Finland	q = 92; 1975:1/1997:4	14.62*
11. Netherland	q = 84; 1977:1/1997:4	4.71
12. Austria	q = 128; 1964:1/1995:4	4.30

Time series y_t	sample q	Ljung-Box, Q(6) $(1-B)\log y_t$
13. Switzerland	q = 132; 1965:1/1997:4	55.99**
14. Eur15	q = 132; 1965:1/1997:4	30.14**

**Significance level 1% *Significance level 5%

Attempts to remove trends and seasonal fluctuations in economic time series have a long history, going back to at least the end of the last century. Since then, one of the best known problems analysed has been the introduction of spurious fluctuations into the trend or smooth series by various methods. Surprisingly enough in spite of critical warnings (CARO, 1996), only few macroeconomists have spent time analysing the estimation techniques of National accounts in general, and of the quarterly series in particular. The latter are very often used in empirical analysis. Nevertheless, as the results of the Ljung-Box tests suggest, although all the retained series represent the same event in all countries (the "seasonally adjusted GDP at contant prices"), it is quite clear that they do not seem to share the same properties. This is even more clearer if we give a look at the estimated correlogram for the first differences of the series (after log transformation) shown in figure 1. According to the global Ljung-Box test for the first six autocorrelations, some series appear to be of the autoregressive class and some do not. If we focus on the autocorrelation of order 1, we come up with two extreme cases : Spain and the United Kingdom. Whereas the autocorrelation of order 1 is very high for the first country, it is not significant for the second. For two other countries, the Netherlands and Austria, according to the global Ljung-Box test, the first 6 autocorrelations are not significantly different from zero. If we perform individual tests for each autocorrelation from order 1 to 6, it appears that the autocorrelation of order 3 seems different from zero only in the case of the United Kingdom and Finland. For Austria and the Netherlands, no individual lag from order 1 to 6 seems significant. Although the seasonal adjustement methods used in each country can hardly account for such differences, it is worth noting that the smoother the series are (for instance corrected for seasonal influences and from irregular fluctuations[1]), the stronger the autocorrelation structure of the first difference will be, and especially the autocorrelation of order 1. In the case of Spain, where the autocorrelation structure is the strongest, the published official GDP quarterly series refers to the trend-cycle component computed with the TRAMO-SEATS program (GOMEZ and MARAVALL, 1996) and can thus be well explained. In the case of the United Kingdom, the Netherlands, Finland and Austria, we can of course advance that a non well defined correlogram may be due to the shortness of the series. We suspect however that the methods used to compute the raw data have something to do with it, and certainly also the seasonal adjustement procedure used. From a more theoretical point of view, we actually never have any prior knowledge of the auto-

1. i. e. the so called *trend-cycle component*, that should not be mixed up with the *trend-cycle decomposition*

correlation structure of the first difference of real quarterly GDP, even if one would not be surprised to find out that there may exist positive correlations from one quarter to the other, due for instance to the enduring nature of various production processes. In that respect, the autocorrelation structure for the first difference of the seasonally adjusted GDP for the US, Canada, Germany, France, Italy, Spain and Switzerland corresponds to what one would expect. For our purpose, we have to keep in mind that although we are analysing the same macroeconomic aggregate for different countries, the properties of the available series are not the same. In our view, this finding should not be only explained by the fact that business cycles have different characteristics in each countries. The way the data are compiled also certainly plays an important role.

Figure 1: GDP quarterly data, seasonally adjusted, correlogram of the first difference (after log transformation) – estimated with all the data available (see table 1) –

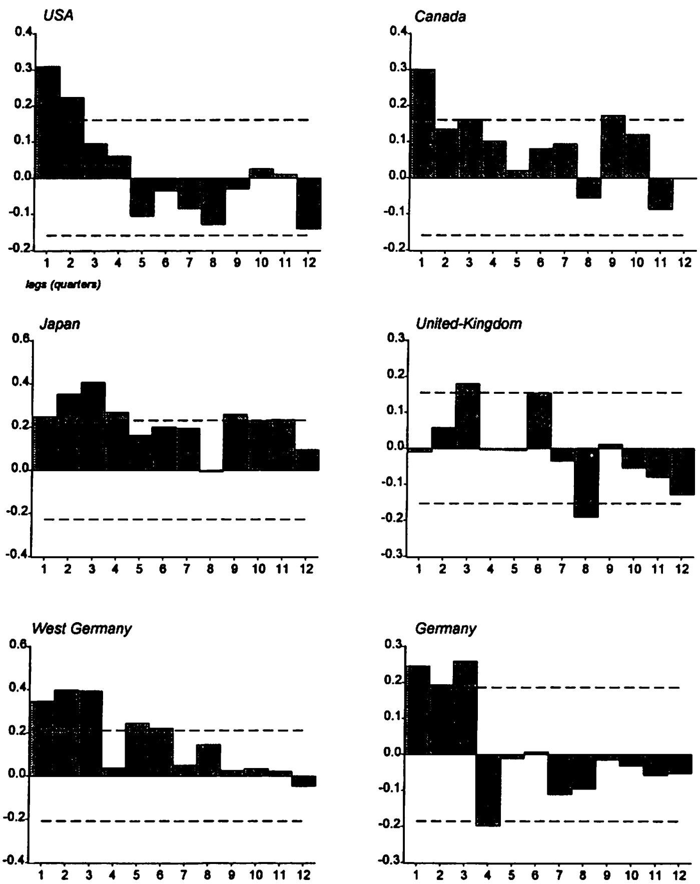
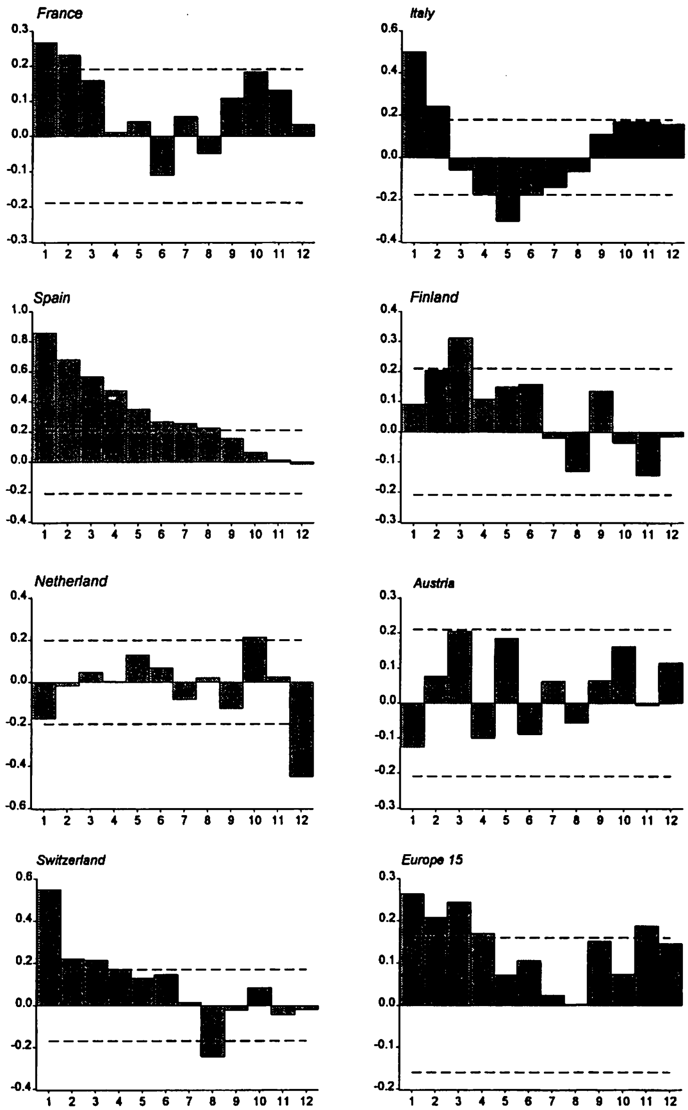


Figure 2: GDP quarterly data, seasonally adjusted, correlogram of the first difference (after log transformation) – estimated with all the data available (see table 1) –



3. INFORMATION ABOUT THE CYCLICAL COMPONENT OF NATIONAL OUTPUT

Different methods can be used to isolate information about the cyclical component of national output. By taking the first differences – assuming that the time series are difference stationary stochastic processes (DS) – or by fitting a deterministic or stochastic trend and isolating a stationary cyclical component, the series are then considered as a representation of trend-stationary (stochastic) processes (TS). These two ways of removing the long-run characteristics that appear in most economic time series have been extensively and intensively discussed in the business cycle and time series literature. WOITEK (1997) is a good recent reference on this topic.

Given the fact that it is more difficult than we generally think to detect the non-stationary origin of a time series, and that EK and KH have proposed statistical tools in the two frameworks (DS and TS) to test for common short term components, we will use both in our study, with the official seasonally adjusted quarterly GDP data for the twelve countries (+ the aggregated GDP for Eur15) presented in table 1. Nevertheless, it is worth noting that recently HECQ (1998) showed that the use of seasonally adjusted data is not without danger when testing for serial common features, especially because the inference conducted on unadjusted observed time series may differ from the one conducted on estimated seasonally adjusted data. The remark of HECQ is even more accurate if we consider the various estimated correlogram of the first difference of the series, presented in figure 1. On the other hand, working with seasonally unadjusted data would suppose first a correct identification of the nature of the seasonality, an issue that is far from being obvious, especially if the data are contaminated by outliers and trend breaks (SMITH and OTERO, 1997). For our purpose, we will concentrate our analysis on the available quarterly seasonally adjusted GDP figures, bearing in mind that the results obtained within only one approach are not sufficient to draw a definitive conclusion. The use of *trend-plus-cycle* bivariate and multivariate decompositions presented in the second part of this paper provides therefore a good alternative and complement when testing for common short term dynamics.

4. SERIAL CORRELATION COMMON FEATURE

4.1. Presentation and some precautionary remarks

Whereas cointegration is an indicator of comovements among non-stationary times series (i. e. long term dependence), the existence of a *serial correlation common feature* (SCCF) in the sense of EK implies synchronous comovements in the stationary components of time series. More precisely, if the hypothesis of SCCF is accepted for two time series (expressed in first differences), it implies that both time series have a certain serial correlation structure (a certain *feature*) and that a linear combination of these two series exists that eliminates all serial correlation (that annihilates the *identified feature*).

The general regression model used by EK is the following:

$$y_t = Ax_t + Bz_t + \varepsilon_t \tag{1}$$

In the case of a bivariate system, we have:

$$\begin{pmatrix} y_{1t} \\ y_{2t} \end{pmatrix} = \begin{pmatrix} a_{11} & a_{12} \\ a_{21} & a_{22} \end{pmatrix} \begin{pmatrix} x_{1t} \\ x_{2t} \end{pmatrix} + \begin{pmatrix} b_{11} & b_{12} \\ b_{21} & b_{22} \end{pmatrix} \begin{pmatrix} z_{1t} \\ z_{2t} \end{pmatrix} + \begin{pmatrix} \varepsilon_{1t} \\ \varepsilon_{2t} \end{pmatrix} \tag{2}$$

To test for the presence of SCCF according to EK, y_t is stationary, x_t contains simply the regression's intercepts, z_t represents one lag of y_t, and ε_t is a bivariate non correlated white noise. In the case $B \neq 0$, the test for SCCF is simply a test for the existence of a linear combinaison of y_{1t} and y_{2t} that rules out the serial correlation identified in B (note that according to EK this is a correlation with only one quarter lagged information). In other words, if the presence of SCCF is accepted, then a scalar δ exists and the combinaison $u_t = y_{1t} - \delta y_{2t}$ is a white noise. EK have proposed a two steps procedure to test for SCCF. The first step is to determine whether the feature (the serial correlation) that could be common is present in each series. The second step consists in testing for a common feature, i. e. the features identified in y_{1t} and y_{2t} could be due to a single component. Two techniques are suggested by EK to test for common features, based on the Lagrange multiplier principle (LM). The first one is based on a iterative grid search for the parameter δ and can be considered as a limited-information maximum likelihood (LIML) procedure and the second one is based on two stage least squares (2SLS). Within the last procedure, the solution to $s(\hat{u}) \equiv \min_{\delta} s(y_{1t} - \delta y_{2t})$ is given by the two stage least squares estimate of δ in the following model:

$$y_{1t} = \delta y_{2t} + a_{11}x_{1t} + e_{1t} \tag{3}$$

where the instrument list is a constant and both variables y_{1t} and y_{2t} lagged with one quarter. The test statistic $s(\hat{u})$ can then be computed as TR^2, where T is the number of observations and R^2 is the traditional measure of fit. These two parameters are taken from the regression of $\hat{e}_{1t}$, the 2SLS residual from regression (3), on a constant, y_{1t-1} and y_{2t-1}. The test statistic is asymptotically distributed as a $\chi^2(s)$ under the null hypothesis (H_0 : common feature), the degree of freedom being the number of instrumental variables in (3), minus the number of right-side endogeneous variables, in our case $s = 1$. The critical value of a $\chi^2(1)$ is 3.84 at the 5% significance level. All the corresponding developments can be found in EK. Other interesting remarks on LM specification tests and the use of TR^2 for the test statistic can also be found in DOORNIK and HENDRY (1997).

A relatively important criticism of this approach says that the concepts of serial correlation and autocorrelation are not clearly defined here (Ericson, 1993). Indeed, according to EK, testing for serial correlation in y_{1t} and y_{2t} is equivalent to testing the hypothesis that y_{1t} is correlated with its first lag *and/or* the first lag of y_{2t} (equivalently for y_{2t}). In equation (2), this can be seen as a test for b_{11} and b_{12} being different from zero for a *feature* in y_{1t}, and b_{21} and b_{22} being different from zero for a *feature* in y_{2t}. EK used GDP seasonally adjusted series for the G7 countries in their article, over the sample 1957:1/

1988:5 (for some series starting with 65:1 and for some others ending at 89:3). For the pair US-UK, EK found that although the presence of a serial correlation feature in both series cannot be rejected, it does not appear that the two countries share a serial correlation common feature. Among all the possible pairs within the G7 countries, the rejections of the serial correlation common feature hypothesis found by EK always involve the UK. Furthermore, one of the main macroeconomic conclusions of EK is that most of the G7 countries share a common business cycle with each other and that UK has its own cyclical pattern. Looking at the various correlograms of the GDP first differences presented in figures 1 and 2, one might have expected such a conclusion only by considering that among the G7, the autocorrelation of order 1 is not different from zero only for UK's first difference GDP[2]. Hence, we could cast some doubt on testing for common short term dynamics, for pairs of countries involving the Netherlands, Austria, Finland and the UK.

4.2. *Cointegration analysis and comments*

Before proceeding to the estimation of the VAR(1) bivariate model with the first differences of the series, as proposed by EK, we need to test for the presence of cointregration. First, we use the two steps procedure of ENGLE and GRANGER (1987). Using the largest sample available, we thus compute several regressions of the form:

$$y_{jt} = \alpha_j + \beta_j t + \sum_{i=1;i\neq j}^{2} \beta_i y_{it} + \varepsilon_{jt} \qquad j = 1,2 \tag{4}$$

where y_{1t} and y_{2t} denote two undifferenced variables, t is a time trend and α_j, β_j and β_i are non time-dependent parameters to be estimated. Then, ADF tests are computed on the series of estimated residuals $\hat{\varepsilon}_{jt}$. If the hypothesis of a unit root in $\hat{\varepsilon}_{jt}$ cannot be rejected, then the series are not cointegrated. For the 13 regressions including Switzerland, table 2 shows the Dickey-Füller t-statistics to test for the hypothesis of a unit root in $\hat{\varepsilon}_{jt}$, using up to 4 lagged values of $\hat{\varepsilon}_{jt}$, an intercept and no time trend. The first row reports the $p - values$ of the Ljung-Box Q-statistics associated with the autocorrelation function of the residuals obtained in the Dickey-Füller regression (6 lags). The critical value mentioned (1% and 5%) have been calculated using the response surface regression of MacKinnon (1991) reproduced in Banerjee, Dolado, Galbraith and HENDRY (1993). According to this first cointegration test, we can assume that there is no evidence against the null hypothesis of no cointegration between any pair of countries involving Switzerland. For the remaining countries taken in pair, we could only reject the hypothesis of a unit root in the estimated residuals for the pair US-Canada, over the sample 1970 : 1/1997 : 4. However, modification of the sample length leads to various results. Thus, even for these two very linked economies, cointegration among the their quarterly GDP series can not be taken as granted.

2. For UK's first difference GDP, reestimating the correlogram over the sample 1960:1–1989:4 gives approximatively the same picture as in the figure 1.

Table 2: Engle-Granger cointegration test, quarterly real GDP of Switzerland and other 13 countries, bivariate analysis

	USA	Can	Jap	W. Ger	Fra	Ita	UK	Spa	Finl	Nle	Aust	Eur15
Q_6	0.977	0.960	0.984	0.862	0.968	0.948	0.565	0.971	0.93	0.816	0.153	0.326
ADF_t	-2.56	-2.72	-2.59	-2.52	-2.60	-2.64	-2.73	-2.68	-1.93	2.05	2.31	-2.46
critical value 1%	-3.98	-3.98	-3.98	-3.98	-4.00	-4.00	-3.98	-4.00	-4.02	-4.03	-3.98	-3.98
critical value 5%	-3.34	-3.34	-3.34	-3.34	-3.50	-3.50	-3.34	-3.50	-3.37	-3.37	-3.34	-3.34

Second, for two groups of various countries that will be used in a multivariate *trend-plus-cycle* structural model in section 5, that is Switzerland, France, Italie and Germany, and the same group plus the United States, the JOHANSEN test (1988) of cointegration was also performed. The JOHANSEN cointegration test is based on the rank of the matrix Π in the multivariate following regression :

$$\Delta y_t = \Gamma_1 \Delta y_{t-1} + \ldots + \Gamma_{p-1} \Delta y_{t-p+1} + \Pi y_{t-p} + \mu + \varepsilon_t \qquad (5)$$

where y_t is defined as *(CH Fra Ita Ger)'* or *(CH Fra Ita Ger USA)'*, in other terms the level of the real quarterly GDP series of our four/five countries, and Π represents the long-run matrix. With our four (five) dimensional system, each of the matrix of parameters in equation (5) is of order 4×4 (respectively 5×5). Taking into account the sample size (1970:1/1997:4=112 observations), we had to pay attention not to consider too much lags. Setting a value of $p = 4$ was found sufficient to get a relatively low autocorrelations function of the estimated residuals. With Π being the matrix of long-run responses, its rank q determines how many linear combinaisons of the variables are stationary, i. e. how many cointegration relations are present in the data. For the first group of four countries, if $q = n = 4$, all variable in y_t are stationary, whereas a value of $q = 0$ implies that Δy_t is stationary, and the series can be assumed to be noncointegrated. When $0 < q < n$, we can conclude that there are q cointegrated linear combinations within y_t. To determine the rank of Π, i. e. the number of significant eigenvalues in Π, the lambda-max and the trace tests are used, for the sequential testing of $H_0 : r \leq q$ $(q = 3, 2, 1, 0)$. The same procedure was used for our five dimensional VAR model ($H_0 : r \leq q$ $(q = 4, 3, 2, 1, 0)$). More details on this procedure and on the two tests can be found in J.A. DOORNIK and D.F. HENDRY (1997). The results of the JOHANSEN cointegration tests are reported in tables 3 and 4. Again, the second cointegration test presented here indicates that within the four and the five dimensional VAR (with four lags), the various quarterly GDP series cannot be assumed to be cointegrated. For the five dimensional VAR, including the four European countries plus the US, it is true that we can slightly rejet the null hypothesis of $r = 0$, and thus accept that there is at least one cointegration relation. However, computing the JOHANSEN test for the same VAR model with various sample periods and changing the lags length shows that the results are not very robust. In their original article, only by running the ENGLE-GRANGER two steps cointegration test, EK found no evidence against the null hypothesis of no cointegration between any pair of the following countries *Canada, France, Germany, Italy, Japan, UK* and *US*, over the sample 1957:1/1989:4.

Table 3: Johansen cointegration test among quarterly GDP series, VAR model with four countries (Switzerland, France, Italy, Germany) and four lags

	Number of cointegration vectors			
	$r \leq 3$	$r \leq 2$	$r \leq 1$	$r = 0$
Lambda-max test	1.71	6.01	9.70	29.97
critical value 5%	3.80	14.10	21.00	27.10
Trace test	1.71	7.72	17.42	43.39
critical value 5%	3.80	15.40	29.70	47.20

Table 4: Johansen cointegration test among quarterly GDP series, VAR model with five countries (Switzerland, France, Italy, Germany, USA) and four lags

	Number of cointegration vectors				
	$r \leq 4$	$r \leq 3$	$r \leq 2$	$r \leq 1$	$r = 0$
Lambda-max test	0.34	6.18	8.81	26.01	27.45
critical value 5%	3.80	14.10	21.00	27.10	33.50
Trace test	0.35	6.53	15.34	41.39	68.84*
critical value 5%	3.80	15.4	29.70	47.20	68.50

Some futher comments need to be made on the absence of cointegration among quarterly GDP series of various OECD countries (according to our results and to EK). First, it should be kept in mind that the concept of cointegration is closely linked to the idea of "equilibrium". In other words, using quarterly GDP series, if the hypothesis of cointegration had been accepted, we would implicitely have admitted that GDP quarterly series among OECD countries may drift apart in the short run, but that they do not diverge from each other in the long run. However, empirical evidence seems to suggest the opposite. Cyclical and short term fluctuations are highly correlated among quarterly GDP time series for most of the OECD countries, or at least for some groups of countries in the same geographical region. Regarding the long term fluctuations, they may sometimes diverge quite a lot among countries, even within a group where the cyclical fluctuations seem to remain in phase for a long period. Trying to give an interpretation to this last observation without refering to a specific theory of business cycle fluctuations is rather difficult. Furthermore, any comment on the absence of cointegration among quarterly GDP series for various countries also depends on the degree of interdependance assumed between secural growth and short-run fluctuations. If we refer to a simple desequilibrium model and accept that policy errors and external disturbances are responsible for economic fluctuations, while supply side conditions are responsible for the long run evolution of an economy, and more country-specific by nature, we can then find a justification for the absence of cointegration among the GDP time series. Various factors, like the ones driving the quantity and the productivity of human and physical resources, the structure of the local industry, and to a certain extent, the legal, the institutional and the financial frameworks, may indeed play a certain role for the long run growth of an economy (when not for the structure of the short-term fluctuations, but this last issue is still debated in the literature) and seem rather country-specific. If, on the other hand, we accept that systematic instability are the results of endogeneous factors and that they are inherent to the markets functioning, then the interpretation of the absence of cointegration will widely differ. Last but not least, the absence of cointegration among European countries can also be understood in terms of the convergence hypothesis of the economic growth literature, as stated by the IMF in its last World Economic Outlook (October 1999), to explain the observed divergence in economic growth in the EU during the 90s. In this context, countries with low initial levels of output per worker tend to experience higher rate of growth that countries with higher initial conditions. One indirect conclusion of this interpretation is that in some years, when the level of output per worker will be the same among European countries (a form of "catching-up" in the level of activity will once be achieved), then the various quarterly GDP time series in the Euro area will be cointegrated, but this seems not to be the case for the moment. To sum up, even though the cointegration hypothesis of quarterly GDP time series among industrialized countries raises various considerations, according to the results of the two tests presented, and also to the results of EK, we can safely assume that our series are not cointegrated, with a certain reserve for the pair *US-Canada*.

4.3. Application of the EK procedure

Given the results of the preceding sub-sections, the same procedure used by EK was applied to our series. For the twelve countries plus the Eur15 aggregate presented in table 1, we first perform a serial correlation feature test for both series involved in each possible 78 pairs. In our case, we choose a Wald test for corresponding coefficients restrictions in the matrix B from equation (1), i. e. $H_0 : b_{11}$ and $b_{12} = 0$ for a non serial correlation feature in y_{1t}, in the spirit of EK, and $H_0 : b_{21}$ and $b_{22} = 0$ for a non serial correlation feature in y_{2t}. Table 5 presents the p-value of this test statistic. We further test the SCCF hypothesis (*I*nstrumental *V*ariables estimation variant) only with pairs of countries for which the Wald test allows the rejection of the previous hypothesis H_0 at the 5% significance level. Tables 6 and 7 report the main results: $\hat{\delta}$ is the instrumental variable estimate, $t_{\hat{\delta}}$ represents its t statistic, $\chi_2(1)$ is the chi-squared test of overidentifying restrictions (i. e. the common feature test), and $Q(6)$ is the Ljung-Box global test statistic for the first six autocorrelations of the linear combinaison $u_t = y_{1t} - \hat{\delta} y_{2t}$. Keep in mind that if the hypothesis H_0 : common feature is accepted, then u_t should be a white noise. The critical value for the Ljung-Box test at the 5% significance level is 12.6.

Table 5: Wald test for serial correlation in a VAR(1) model, P-value of the test statistic, estimation performed with all available observations

country	USA	Can	Jap	W Ger	Fra	Ita	UK	Spa	Finl	Nle	Aust	CH	Eur15
USA		0.000	0.000	0.000	0.009	0.011	0.000	0.015	0.020	0.021	0.012	0.002	0.000
Can	0.000		0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000
Jap	0.000	0.019		0.000	0.609	0.466	0.004	0.144	0.720	0.984	0.000	0.001	0.000
W Ger	0.000	0.000	0.000		0.004	0.000	0.797	0.000	0.022	0.556	0.000	0.000	0.005
Fra	0.004	0.000	0.004	0.005		0.000	0.000	0.000	0.169	0.391	0.004	0.000	0.000
Ita	0.000	0.000	0.000	0.000	0.000		0.000	0.000	0.000	0.001	0.000	0.000	0.000
UK	0.091	0.588	0.986	0.226	0.467	0.300		0.128	0.211	0.014	0.004	0.126	0.776
Spa	0.000	0.000	0.000	0.000	0.000	0.000	0.000		0.000	0.000	0.000	0.000	0.000
Finl	0.655	0.060	0.696	0.513	0.505	0.197	0.412	0.241		0.776	0.604	0.242	0.658
Nle	0.016	0.009	0.295	0.110	0.251	0.295	0.085	0.096	0.272		0.141	0.162	0.169
Aust	0.319	0.064	0.021	0.006	0.001	0.314	0.001	0.463	0.152	0.000		0.000	0.020
CH	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000	0.000		0.000
Eur15	0.000	0.000	0.000	0.004	0.003	0.000	0.001	0.000	0.094	0.233	0.000	0.000	

Table 6: Serial correlation common feature test in bivariate VAR(1) models, IV estimation

country	Statistic	USA	Can	Jap	W Ger	Frau	Ita	Spa	Aust	CH	Eur15
USA	$\hat{\delta}$		0.58	0.34	-0.02	1.19	0.46	0.22	$\emptyset$	0.18	1.26
	$t_{\hat{\delta}}$		4.05	1.24	-0.07	2.28	2.37	1.44	$\emptyset$	1.19	3.12
	$\chi_2(1)$		0.24	7.22	10.79	1.17	2.56	5.98	$\emptyset$	10.88	0.52
	$Q(6)$		5.19	14.16	21.55	9.25	6.79	10.76	$\emptyset$	15.63	3.56
Can				0.50	0.65	1.37	0.57	0.31	$\emptyset$	0.49	1.39
	$t_{\hat{\delta}}$			1.72	2.43	3.19	3.16	1.91	$\emptyset$	3.27	3.73
	$\chi_2(1)$			9.21	4.51	0.50	4.11	11.22	$\emptyset$	6.67	0.05
	$Q(6)$			11.66	13.23	7.25	8.91	15.51	$\emptyset$	25.22	7.96

country	Statistic	USA	Can	Jap	W Ger	Frau	Ita	Spa	Aust	CH	Eur15
Jap	$\hat{\delta}$				0.81	$\emptyset$	$\emptyset$	$\emptyset$	0.72	0.54	1.71
	$t_{\hat{\delta}}$				3.51	$\emptyset$	$\emptyset$	$\emptyset$	1.62	3.38	3.63
	$\chi_2(1)$				0.31	$\emptyset$	$\emptyset$	$\emptyset$	4.84	3.93	2.70
	$Q(6)$				17.33	$\emptyset$	$\emptyset$	$\emptyset$	6.56	11.05	11.76
W Ger	$\hat{\delta}$					1.18	0.59	0.55	0.34	0.62	0.84
	$t_{\hat{\delta}}$					3.18	3.43	3.75	1.14	3.61	7.14
	$\chi_2(1)$					2.61	2.52	4.16	18.25	7.93	10.73
	$Q(6)$					23.71	39.07	31.19	41.83	35.38	53.42
Fra	$\hat{\delta}$						0.63	0.49	0.23	0.47	1.48
	$t_{\hat{\delta}}$						4.89	4.70	1.15	3.88	4.74
	$\chi_2(1)$						0.12	0.27	11.01	2.21	0.17
	$Q(6)$						2.99	6.67	7.77	6.72	5.58

Table 7: Serial correlation common feature test in VAR(1) bivariate models, IV estimation

Country	Statistic	Spa	Aust	CH	Eur15
Ita	$\hat{\delta}$	0.53	0.82	0.74	0.91
	$t_{\hat{\delta}}$	3.66	3.48	5.02	5.11
	$\chi_2(1)$	19.46	9.19	8.27	0.27
	$Q(6)$	35.42	4.15	12.30	6.46
Spa	$\hat{\delta}$		1.04	0.64	2.04
	$t_{\hat{\delta}}$		3.46	4.90	5.23
	$\chi_2(1)$		13.52	44.19	1.16
	$Q(6)$		13.24	85.84	5.99
Aust	$\hat{\delta}$			0.41	-0.03
	$t_{\hat{\delta}}$			2.57	-0.06
	$\chi_2(1)$			10.43	7.57
	$Q(6)$			25.56	12.82
CH	$\hat{\delta}$				1.83
	$t_{\hat{\delta}}$				5.34
	$\chi_2(1)$				0.47
	$Q(6)$				2.66

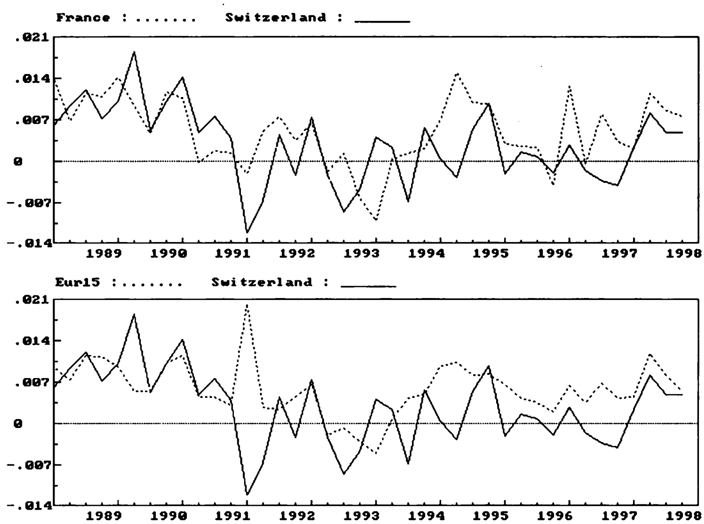
According to the results of table 5, we can retain 39 pairs to test for SCCF. As expected, the pairs including the countries for which the Ljung-Box global test for the first six autocorrelations of the first difference of GDP in table 1 ($Q(6)$ statistic) were the lowest, are also the pairs for which the serial correlation feature is more often rejected in the bivariate VAR(1) model. To simplify, we eliminate the United Kingdom, Finland and the Netherlands in the SCCF test, even if the results of the Wald test are not so clear-cut for some pairs including these countries.

Within the 39 pairs retained, we accept the H_0 hypothesis of SCCF for 17 pairs, ($\chi^2(1) \leq$ the critical value of 3.84). For almost all of these 17 pairs, the obtained autocorrelation structure of the linear combinaison $u_t = y_{1t} - \hat{\delta} y_{2t}$ is not significant according to the global Ljung-Box test (the first 6 autocorrelations). For some pairs including West Germany however, when we perform individual tests for various u_t autocorrelations from order 1 to 6, although the first order autocorrelation is not significant, some autocorrelations with longer lags are slightly different from zero. A very interesting case is the pair France-Spain, where u_t does not display any significant autocorrelation, despite the strong autocorrelation structure of the first difference of Spain's GDP.

In comparison with the results obtained by EK using data ending in 1989, the hypothesis of SCCF can be less often accepted in our application. For instance, we do not accept the SCCF hypothesis for the pairs US-Germany, US-Japan, Canada-Germany and Canada-Japan. However, we do accept the SCCF hypothesis between US and France or Italy, Canada and France, and also between Japan and some European countries. For almost all the European countries taken in pairs, we accept the SCCF hypothesis, but the results obtained for European pairs including West Germany are very often ambiguous. This ambiguity is mostly due to the presence of significant autocorrelations in the linear combination $u_t = y_{1t} - \hat{\delta} y_{2t}$, even if according to the SCCF test we accept the H_0 hypothesis of a common feature. At a more aggregate level, it is interesting to note that we accept the SCCF hypothesis for the pairs US-Eur15, Canada-Eur15, as well as for Japan-Eur15. Regarding Switzerland, we accept that the Swiss quarterly GDP shares short term common feature with France and Eur15. With Italy, West Germany, Japan and Canada, the results are very close to the confidence interval for accepting the test, but at the 5% level significance, we reject the H_0 hypothesis of common feature. For the pair Switzerland-West Germany, the linear combination u_t displays a strong autocorrelation at lag 4, but the first order autocorrelation is no more significant.

Regarding Switzerland, it is interesting to note that despite the six years' stagnation between 1991 and 1996, the quarterly short term GDP fluctuations remained in phase with those of France and Eur15 according to the SCCF test and as shown in figure 3 (the pic in the Eur15 GDP series visible for the first quarter 1991 is due the Germany's reunification effect). Both graphs is figure 3 can be considered as an empirical illustration of the SCCF hypothesis.

Figure 3: First differences of quarterly GDP (after log transformation), Switzerland-France, Switzerland-Eur15 – empirical illustration of the SCCF hypothesis –



5. UNIVARIATE AND MULTIVARIATE TREND-PLUS-CYCLE MODELS

As proposed by HARVEY (1989), HARVEY and JAEGER (1993), KOOPMAN and HARVEY (1997), the stylized facts of macroeconomic time series can be well represented by fitting univariate structural time series models (STM). One of the main characteristics of STM is that they are built explicitly in terms of unobserved components that can be directly interpreted. When used in multivariate frameworks, structural time series models may become behavioural models, FERNANDEZ and HARVEY (1990), KH (1997). For instance, in various macroeconomic time series – in our case the various quarterly GDP time series presented in table 1 – displaying slowly changing trends and potentially short-term (business cycle) fluctuations, multivariate time series models may represent a good means to determine whether some series share unobserved components, i. e. whether the identified unobserved components in a bivariate or multivariate framework are driven by a *common* source of stochastic disturbances.

For quarterly seasonally adjusted economic data, with slowly changing trend and short-term (business cycle) fluctuations, or for yearly series, one candidate model is the so-called *trend-plus-cycle* model. The specification is the following:

$$y_t = \mu_t + \psi_t + \varepsilon_t \tag{6}$$

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t \tag{7}$$

$$\beta_t = \beta_{t-1} + \zeta_t \tag{8}$$

with y_t being the observed series, μ_t the (unobserved) trend, ψ_t the (unobserved) cycle, and ε_t is the residual irregular (unobserved) component. The trend is defined as a local linear trend, with μ_t the level of the trend and β_t the slope (or the first difference of the trend). The cycle component is generated as:

$$\begin{bmatrix} \psi_t \\ \psi_t^* \end{bmatrix} = \rho \begin{bmatrix} c\cos\lambda_c & \sin\lambda_c \\ -\sin\lambda_c & \cos\lambda_c \end{bmatrix} \begin{bmatrix} \psi_{t-1} \\ \psi_{t-1}^* \end{bmatrix} + \begin{bmatrix} \xi_t \\ \xi_t^* \end{bmatrix} \tag{9}$$

with ρ being the damping factor of the cycle ($0 \leq \rho \leq 1$), λ_c the cycle frequency expressed in radians ($0 \leq \lambda_c \leq \pi$), the cycle period can be expressed as $2\pi/\lambda_c$. As can be seen, all the previous unobserved components are, by default, stochastic, with η_t, ζ_t, ξ_t and ξ_t^* being normal white-noise disturbances with variances $\sigma^2_{\eta_t}$, $\sigma^2_{\zeta_t}$, $\sigma^2_{\xi_t}$ (ξ_t and ξ_t^* are two uncorrelated white noise processes sharing the same variance[3]), as is the residual irregular component ε_t (with variance $\sigma^2_{\varepsilon_t}$). In this context, the *strength* of an unobserved component depends on its stochastic disturbances variance. Once the model (6) to (9) has been put in a state space form, the Kalman filter can be used to compute the likelihood function and to estimate by maximization the hyperparameters values ($\sigma^2_{\eta_t}$, $\sigma^2_{\zeta_t}$, $\sigma^2_{\xi_t}$, $\sigma^2_{\varepsilon_t}$, ρ and λ_c), the unobserved trend, cycle and irregular components. As the inital state vector of the state space representation for the *trend-plus-cycle* model has an unspecified distribution (due to some non-stationary unobserved components), attention should be paid to the initialization of the Kalman filter. More details on this topic may be found in KOOPMAN, HARVEY, DOORNIK and SHEPARD (1995), GOMEZ and MARAVALL (1993), and in RONCALLI (1996).

The multivariate version of the *trend-plus-cycle* model is defined as:

$$y_t = \mu_t + \psi_t + \varepsilon_t, \qquad \varepsilon_t \sim NID(0, \Sigma_\varepsilon) \tag{10}$$

$$\mu_t = \mu_{t-1} + \beta_{t-1} + \eta_t, \qquad \eta_t \sim NID(0, \Sigma_\eta) \tag{11}$$

$$\beta_t = \beta_{t-1} + \zeta_t, \qquad \zeta_t \sim NID(0, \Sigma_\zeta) \tag{12}$$

and the multivariate cycle component ψ_t is generated as :

$$\begin{bmatrix} \psi_t \\ \psi_t^* \end{bmatrix} = \left\{ \rho \begin{bmatrix} \cos\lambda_c & \sin\lambda_c \\ -\sin\lambda_c & \cos\lambda_c \end{bmatrix} \otimes I_N \right\} \begin{bmatrix} \psi_{t-1} \\ \psi_{t-1}^* \end{bmatrix} + \begin{bmatrix} \xi_t \\ \xi_t^* \end{bmatrix} \tag{13}$$

with $E(\xi_t \xi_t') = E(\xi_t^* \xi_t^{*\prime}) = \sum_\xi$ and $E(\xi_t \xi_t^{*\prime}) = 0$. When the vectorial time series y_t contains $N \times 1$ series, ψ_t and ψ_{t-1}^* are then $N \times 1$ vectors. The stochastic disturbances

3. Imposing this constraint was not found to be damaging and brings computational advantages (HARVEY, 1985).

driving the unobserved components in the univariate representation of the model become here multivariate normal disturbances vectors with $N \times N$ covariance matrices. Even if each individual series is modeled as in the univariate case, such a multivariate formulation allows stochastic disturbances for a specific unobserved component to be correlated across two or more time series, which is an intesting property. It should be noticed, that the parameters ρ and λ_c in the multivariate *trend-plus-cycle* model are the same for all series, that is the cycles are *similar* in all the various series, they have the same damping factor and the same frequency (in the digital filtering language we could say the same *phase* but not the same *amplitude*). This may be considered as a constraint imposed on the data on previous information, as a valid hypothesis based on the results obtained after having estimated univariate models by maximum likelihood, or as a result of a likelihood ratio test as presented in HARVEY (1989). In an extreme case, that is when the disturbances driving a particular unobserved component are perfectly correlated across two or more time series, this component is said to be *common*. In our case we will focus on the potential short term business cycle fluctuations that could be common across some GDP quarterly series.

It should be noticed that within this kind of decomposition, the trend component modeled as in (7) and (8) does not necessarily represent the smooth long term evolution of the series that one would normally expect. In contrast with the so called *canonical* decomposition found for instance in MARAVALL and GOMEZ (1995), where all the unobserved components are as stable as possible, and the variance of the irregular component is maximized, within STM all the unobserved components can be contaminated by additional white noise. Without a-priori (economic) information on the stochastic property of the unobserved components, within the STM approach the allocation of noice among the components relies essentially on the maximum likelihood procedure used for identification and estimation. As pointed out by MARAVALL (1985), the choice between an unobserved component contaminated or not by additional white noice depends finally on the objective of the analysis. For our purpose, the fact that the irregular component in (6) will disappear in most of the estimated models is to be considered as a positive criterion for the decomposition. In such a case, the stochastic specification used for the trend and the cycle components is effectively capable of capturing the stochastic property of the data, without having to estimate a residual white noise component.

5.1. Univariate estimates of the trend-plus-cycle model

The results of fitting the univariate *trend-plus-cycle* model (6) to (9) to the quarterly GDP series presented in table 1 are shown in table 8. Beside the standard deviations of the disturbances driving the unobserved components (absolute values are obtained by multiplying by 10^{-4}), the damping factor and the cycle period in quarters, various goodness-of-fit statistics are presented. R_D^2 refers to the coefficient of determination with respect to the first differences; $Q(12)$ is the global Ljung-Box statistic based on the first $p = 12$

residual autocorrelations[4], which is tested against a χ^2_q with q set equal to $p+1$ minus the number of estimated hyperparameters. This number was adapted for Germany, because of the level shift dummy variable used to account for 1991:1, and also in the case of Italy and Switzerland, where intervention variables were used to take account of the break in the trend, respectively for 1974:4 and for 1974:2 and 1974:4[5]. The probability value for a Shenton and Bowman (1975) test for normality and heteroscedasticity is also given ($P.val.N_{SB}$). Note that when the residuals show some departure from normality, this is due to the presence of outliers, and correction for these effects with more dummy variables does not change the overall results of the models (especially $\hat{\rho}$ and $2\pi/\lambda_c$). As pointed out earlier, in structural time series models, the *strength* of an unobserved component depends on its stochastic disturbances variance. For the cycle component, when $\hat{\sigma}_\xi = \hat{\sigma}_{\xi^*}$ is different from zero but small, it is helpful to perform a likelihood ratio test imposing the constraint $\hat{\sigma}_\xi = 0$ (or the constraint $\rho = 0$, as exposed in HARVEY, 1985). The LR statistic is reported in the last column of table 8: when $LR > 2.71$, we reject the $H_0 : \hat{\sigma}_\xi = 0$, at the 10% significance level, respectively when $LR > 3.84$, at the 5% level.

Table 8: Estimate of the univariate trend-plus-cycle model, sample: all available observations

Series	$\hat{\sigma}_\varepsilon$	$\hat{\sigma}_\eta$	$\hat{\sigma}_\zeta$	$\hat{\sigma}_\xi$	$\hat{\rho}$	$2\pi/\lambda_c$	R^2_D	$Q(12)$	$P.val.\ N_{SB}$	LR
USA	0	0	3.82	76.62	0.95	30.1	0.11	12.65	0.0875	7.14
Canada	0	0	4.96	79.83	0.94	39.9	0.10	16.95*	0.5148	3.08
Japan	55.35	0	24.13	14.09	0.98	12.35	0.29	13.97	0.0000	4.47
W. Germ.	22.51	0	11.10	66.41	0.97	30.56	0.22	25.54**	0.0899	3.04
Germ.	0	0	8.05	62.51	0.93	28.84	0.60	15.91*	0.8169	2.95
France	0	55.49	9.37	0	∅	∅	0.16	10.28	0.0011	∅
Italy	0	21.78	14.59	44.04	0.92	12.05	0.32	9.73	0.5431	11.25
Uni. King.	56.31	106.3	0	0	∅	∅	0.01	17.94*	0.0000	∅
Spain	0	0	32.49	0	∅	∅	0.72	32.13**	0.0000	∅
Finland	82.23	0	44.99	0	∅	∅	0.14	12.42	0.3851	∅
Nether.	43.13	80.72	3.15	0	∅	∅	0.03	19.31**	0.0007	∅
Austria	55.77	47.77	9.89	16.01	0.96	13.10	0.09	9.02	0.0679	2.39
Switzer.	0	0	23.99	52.38	0.87	20.00	0.35	38.65**	0.7378	3.81
Eur15	0	56.13	6.49	14.22	0.95	13.88	0.13	7.42	0.0000	1.95

* Significant at 10% level **Significant at 5% level

4. Note that within a structural time series model, the residuals are defined as the one-step-ahead prediction errors and should not be mixed up with the irregular component in (6).
5. These intervention variables are used in the univariate and multivariate decompositions.

Once we have the results in table 8, the AIC or BIC criteriums (not reported here), the value of the LR test, and the autocorrelation structure of the residuals, we are able to establish that a *trend-plus-cycle* model with the specification given in (6) to (9) is a *suitable representation of the quarterly real GDP data* for the US, Canada, Germany (West Germany and whole Germany), Italy and for Switzerland. For West Germany and whole Germany, a significant negative autocorrelation is found in the residuals at lag 4. For Switzerland, the high value of the Ljung-Box statistic stems from a quite strong negative autocorrelation at lag 8 (also found in the "old" official quarterly seasonally adjusted GDP series) that we could not explain. If we take account of a negative $MA(4)$ and $MA(8)$ components in the residuals for Germany and Switzerland, then the Ljung-Box statistic is no more significant for the first 12 autocorrelations. To achieve convergence, we impose $\hat{\sigma}_\eta = 0$ for the Canadian GDP. For the US, Switzerland and Germany, without imposing any constraint, $\hat{\sigma}_\eta = 0$ after convergence, which indicates that the trend estimated is relatively smooth. Moreover, given that $\hat{\sigma}_\varepsilon = 0$ for all these countries, their quarterly GDP data effectively decompose into a trend *plus* a stochastic cycle, without having to take a residual irregular component into account. Hence the stochastic properties of the data seem to be well captured by a smooth stochastic trend and a stochastic cycle component.

5.2. *Bivariate estimates*

The results in table 8 suggest that the damping factor and the cycle period estimated for the US and Canada are very close; the same conclusion can be drawn for Germany and the US and to some extent for Italy and Switzerland (the cycles being stochastic, the estimated value for the cycle period is thus an average). This kind of results suggests that fitting bivariate of multivariate models could be of interest to bring out more characteristics from the data. For the pairs US-Canada, US-Germany and Italy-Switzerland, we fit a bivariate *trend-plus-cycle* model, imposing, as mentioned earlier, the presence of *similar* cycles across the series[6].

The results of fitting a bivariate *trend-plus-cycle* model for the pairs mentioned above are shown in table 9. The estimated standard deviation for the one-step (quarter) prediction error is presented for the univariate and the bivariate models ($\hat{\sigma}_{uni}$ and $\hat{\sigma}_{biv}$). In table 10, the estimated covariance matrices $\hat{\Sigma}_\zeta$ and $\hat{\Sigma}_\xi$ are also presented. More precisely, the covariance matrix is given by the lower triangular part of the $\hat{\Sigma}_{...}$ matrix, the main diagonal contains the variance of the corresponding disturbances term in each series, and the upper triangular part is the corresponding correlation matrix (in our bivariate system, a scalar). We set out the covariance matrices of the slope and the cycle components disturbances.

6. Note that for the European countries, this hypothesis does not seem to be too strong if we look at the industry surveys' results released on a montly basis by Eurostat for instance. A visual inspection of these data would suggest the presence of a similar cycle for the European countries, at least in the industry.

Table 9: Estimates of the bivariate trend-plus-cycle model, US-Canada, US-Germany, Switzerland-Italy; all available observations used

Series	$\hat{\sigma}_\varepsilon$	$\hat{\sigma}_\eta$	$\hat{\sigma}_\zeta$	$\hat{\sigma}_\xi$	ρ	$2\pi/\lambda_c$	R_D^2	$Q(12)$	$P.val.\ N_{SB}$	$\widehat{\sigma}_{uni}$	$\widehat{\sigma}_{biv}$
USA	19.62	20.37	3.23	67.42	0.95	29.13	0.15	13.76	0.0568	85.67	83.66
Canada	19.20	51.35	5.91	55.11	0.95	29.13	0.13	16.22*	0.5028	90.03	87.18
USA	0	35.53	0	70.51	0.94	27.57	0.16	14.69	0.9087	84.39	85.39
Germany	0	46.87	6.88	46.42	0.94	27.57	0.60	16.56*	0.0022	73.71	73.42
Switzerl.	0	53.88	29.61	7.94	0.90	12.24	0.37	21.21**	0.8179	69.55	70.08
Italy	0	0	17.50	40.48	0.90	12.24	0.33	8.25	0.5397	70.70	70.45

* Significant at 10% level ** Significant at 5% level

Table 10: Covariance matrices, trend slope and cycle disturbances

Pairs	$\hat{\Sigma}_\zeta$	$\hat{\Sigma}_\xi$
USA-Canada	$\begin{pmatrix} 1.046 \times 10^{-7} & 1.00 \\ 1.913 \times 10^{-7} & 3.496 \times 10^{-7} \end{pmatrix}$	$\begin{pmatrix} 4.546 \times 10^{-5} & 0.9622 \\ 3.575 \times 10^{-5} & 3.037 \times 10^{-5} \end{pmatrix}$
USA-Germany	$\begin{pmatrix} 5.952 \times 10^{-7} & 1.00 \\ 7.785 \times 10^{-8} & 1.018 \times 10^{-8} \end{pmatrix}$	$\begin{pmatrix} 2.178 \times 10^{-5} & 0.7767 \\ 2.550 \times 10^{-5} & 4.950 \times 10^{-5} \end{pmatrix}$
Switzerl.-Italy	$\begin{pmatrix} 2.452 \times 10^{-6} & 0.6092 \\ 2.824 \times 10^{-6} & 8.766 \times 10^{-6} \end{pmatrix}$	$\begin{pmatrix} 2.260 \times 10^{-5} & 1.00 \\ 3.773 \times 10^{-6} & 6.298 \times 10^{-7} \end{pmatrix}$

From the results in table 9 and 10, we notice that not only the GDP series for US and Canada share a common stochastic trend and thus may be cointegrated, but also that the short term movements (the *business cycle* component with an average period just below 8 years) are driven by highly correlated stochastic disturbances. For the pair US-Germany, there is also a potentially cointegrated trend (a *common* stochastic trend), but the cycles' disturbances seem less correlated. Although in the bivariate decomposition, the trend of the Swiss quarterly GDP becomes more stochastic, the results of the maximum likelihood estimate point towards the presence of common business cycle component between Switzerland and Italy, i. e. that the short term fluctuations for both countries are driven by stochastic disturbances sharing the same origin. In other words, *the stochastic properties of the Swiss and Italian GDP series allow for a bivariate admissible trend plus cycle decomposition, in which the cycle identified for each series is proportional to the other.* Imposing this contraint in the matrix $\hat{\Sigma}_\xi$ for the pair US-Canada and reestimating the model allows to compute a LR test and to accept that these two countries also share a common cycle component. However, the same conclusion cannot be drawn for the pair US-Germany.

The graphs in figure 4 illustrate the slope of the trend and the cycle component obtained within the univariate and the bivariate decompositions, for the pairs US-Canada and Switzerland-Italy. As mentioned earlier, the trend is smoother for US and Canada than the trend obtained for Switerland and Italy, due to the relatively lower estimated standard deviations of the disturbances driving the trend components for the first two countries.

Figure 4: US-Canada, slope of the estimated stochastic trend and cycle components from real quarterly GDP, univariate and bivariate decompositions, sample period: 1962:1/1997:4

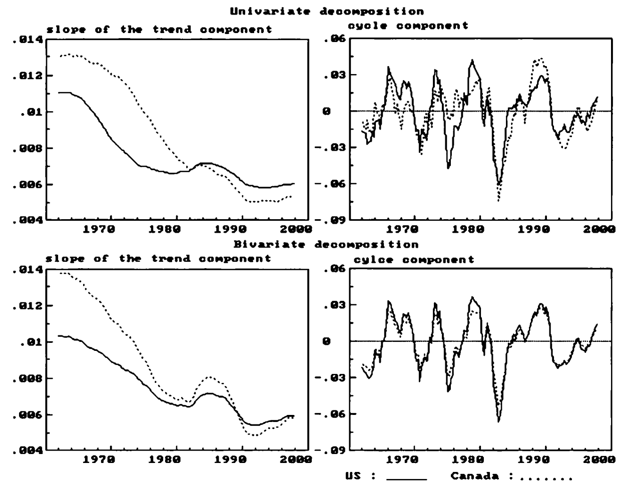
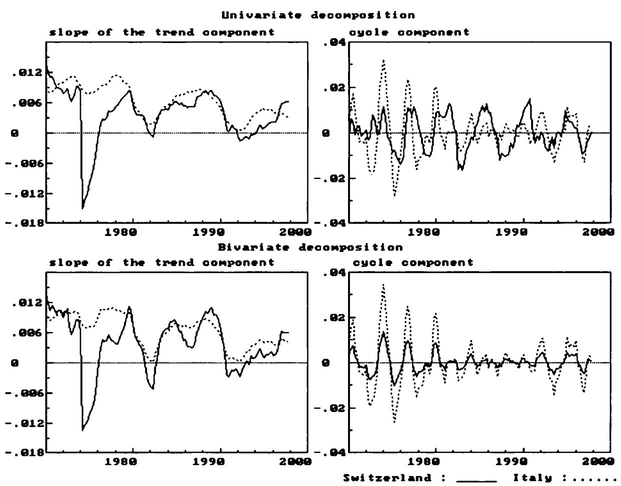


Figure 5: Switzerland-Italy, slope of the estimated stochastic trend and cycle component from real quarterly GDP, univariate and bivariate decompositions, sample period: 1970:1/1997:4



For the pair Switzerland-France, for which the SCCF hypothesis has been accepted, fitting a bivariate *trend-plus-cycle* model can also be of interest. Relying on the results of the previous estimates, and on various goodness-of-fit statistics, the best structural model we could fit for these two countries taken in pair was a model with a trend plus two stochastic cycles (with the constraint $\sigma_{\zeta_t} = 0$). We also used an intervention dummy variable for the break in the trend in 1974:4 (significant for both series). The main results are given in table 11. Again, within this decomposition, we accept that Switzerland and France share common cycle components, although the standardised one-step ahead prediction error is not perfectly a white noise in both cases (i. e. the specification used is not *perfect*). The various graphs in the figure 6 illustrate the decomposition. In this case, the trend does represent the long-run underlying evolution of the quarterly GDP, due to the contraint $\sigma_{\zeta_t} = 0$, and the deviations from the trend are then fully captured by the two stochastic cycles.

Table 11: Estimate of the bivariate trend-plus-two-cycles model (with fixed slope), Switzerland-France; all available observations used

Series	$\hat{\sigma}_\varepsilon$	$\hat{\sigma}_\eta$	$\hat{\sigma}_\xi^{c\_1}$	$\hat{\sigma}_\xi^{c\_2}$	ρ^1/ρ^2	$2\pi/\lambda_{c\_1/c\_2}$	R_D^2	$Q(12)$	$P.val.N_{SB}$
Switzerland	0	14.58	33.72	48.16	0.94/0.99	16.76 / 45.79	0.41	16.07*	0.0514
France	0	52.18	5.90	25.21	0.94/0.99	16.76 / 45.79	0.20	17.36*	0.2972

* Significant at 10% level

Table 12: Covariance matrices, cycle 1 and cycle 2 disturbances

Pairs	$\hat{\Sigma}_{\xi^{c\_1}}$	$\hat{\Sigma}_{\xi^{c\_2}}$
Switzerland-France	$\begin{pmatrix} 10.620 \times 10^{-5} & 1.00 \\ 1.860 \times 10^{-5} & 3.255 \times 10^{-6} \end{pmatrix}$	$\begin{pmatrix} 13.706 \times 10^{-4} & 0.9155 \\ 6.568 \times 10^{-4} & 3.755 \times 10^{-4} \end{pmatrix}$

The graphs in figure 6 present the decomposition of the quarterly GDP for Switzerland and France into a trend and two stochastic cyclical components. It is interesting to note that the first identified and estimated cycle for both countries with an average period just above 4 years has a striking resemblance with the industrial business cycle observable in the industrial business surveys released by Eurostat for France and by the KOF-Insitute for Business Cycle Research (Swiss Federal Institute of Technology) for Switzerland (see figure 7)[7]. The interpretation of the second cycle is not straightforward, but such a cycle can also be found in the German GDP and, to some extent, in the US GDP.

7. For France, the industry survey series refers to the industrial confidence indicator, available on a monthly basis, since 1985:01. For Switzerland, the series refers to an appreciation of the y-o-y variation of industrial production, available on a monthly basis since 1970:01.

Figure 6: Switzerland-France, estimated trend and cycle components from real quarterly GDP, bivariate decomposition with trend and two cycles, sample period: 1970:1/1997:4

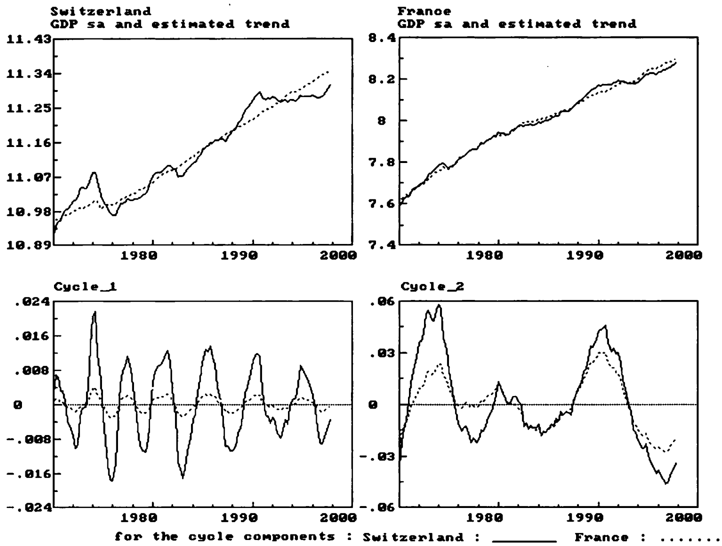
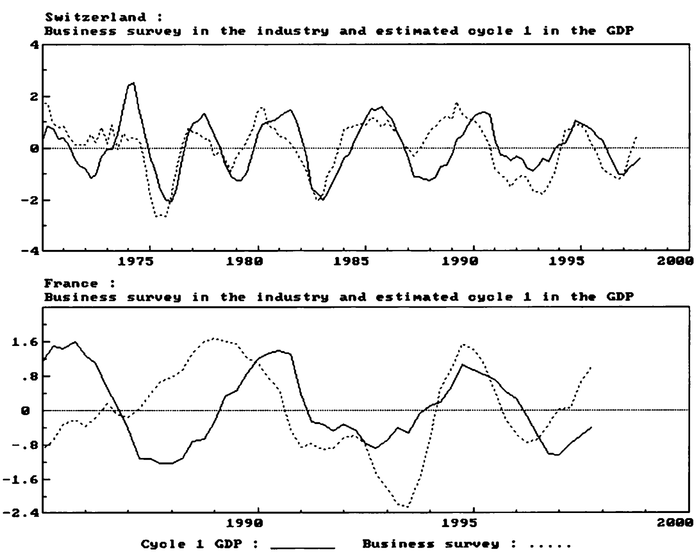


Figure 7: Business surveys data and estimated cycle 1 in the quarterly GDP series



5.3. *Multivariate estimates*

One way of testing the presence of a potential European business cycle is to build a multivariate *trend-plus-cycle* model. Thus, the hypothesis of a similar business cycle among various countries has first to be used, and then, after estimation, the quality of the decomposition and the estimated value in $\hat{\Sigma}_\xi$ can be analyzed. For two groups of countries, that is *(CH Fra Ita Ger)* and *(CH Fra Ita Ger USA)*, named hereafter respectively *group* 1 and *group* 2, we thus estimated a four and a five dimensional *trend-plus-cycle* model. The main results are presented in tables 13 and 14. As regards the correlation of the estimated cyclical components (with a period of about three years), the matrice $\hat{\Sigma}_{cycle}$ presents the correlation among the estimated cycle components.

Table 13: Four dimensional trend-plus-cycle model (group 1); sample 1970:1/1997:4

	$\hat{\sigma}_\varepsilon$	$\hat{\sigma}_\eta$	$\hat{\sigma}_\zeta$	$\hat{\sigma}_\xi$	ρ	$2\pi/\lambda_c$	R_D^2	$Q(12)$	$P.val.\ N_{SB}$	$\hat{\sigma}_{uni}$	$\hat{\sigma}_{multi}$
Switzerland	6.75	44.60	30.07	29.13	0.96	12.75	0.27	40.09*	0.0603	78.07	75.65
France	18.57	38.97	14.82	17.65	0.96	12.75	0.19	7.61	0.3805	59.94	59.06
Italy	4.31	35.20	15.03	40.00	0.96	12.75	0.35	10.93	0.2730	73.28	69.27
Germany	21.84	45.40	18.42	21.46	0.96	12.75	0.64	16.98**	0.8131	73.71	70.35

$\hat{\Sigma}_\zeta$	$\hat{\Sigma}_\xi$	$\hat{\Sigma}_{cycle}$
$\begin{pmatrix} \dots & 0.98 & 0.97 & 0.90 \\ \dots & \dots & 0.98 & 0.92 \\ \dots & \dots & \dots & 0.88 \\ \dots & \dots & \dots & \dots \end{pmatrix}$	$\begin{pmatrix} \dots & 0.79 & 0.89 & 0.65 \\ \dots & \dots & 0.93 & 0.76 \\ \dots & \dots & \dots & 0.90 \\ \dots & \dots & \dots & \dots \end{pmatrix}$	$\begin{pmatrix} \dots & 0.87 & 0.92 & 0.69 \\ \dots & \dots & 0.94 & 0.77 \\ \dots & \dots & \dots & 0.90 \\ \dots & \dots & \dots & \dots \end{pmatrix}$

* Significant at 10% level ** Significant at 5% level

Two findings stand out from the table 13. First, the results of the four dimentional *trend-plus-cycle* STM decomposition seem at least as good as the one obtained with the univariate decomposition. In the case of France, again the incorporation of other countries (other GDP time series) allows the identification of a cyclical component, that does not come out within the univariate decomposition. The estimated one-step-ahead prediction errors (residuals) are not fully white noise for Germany and Switzerland, due to a strong autocorrelation at lag 4 and 8 respectively, as was the case within the univariate decomposition. As in the diagnostic of a traditional VAR model, the vector of estimated residuals should also behave as a white noise vector, i. e. the residuals should not be autocorrelated but should also not be cross-correlated among the four series (or if any correlation are significant, they should be random). In our case, some slightly significant correlations were found among the four series of residuals but not considered as high and frequent enough to reject the decomposition (note that even if 5% of the correlations are significant, the vector of residuals remain a white noise vector).

A second important result that stands out from table 13 is that the presence of an European business cycle can be well established by using a multivariate *trend-plus-cycle* decomposition, as the one propose by the STM approach.The matrices $\hat{\Sigma}_{\xi}$ and $\hat{\Sigma}_{cycle}$ illustrate the very high correlation of the disturbances driving the unobserved business cycle components and also the high correlation of the business cycles among countries. Note that the estimated long-term trends of the series seem also more "cointegrated" (very high values in the matrice $\hat{\Sigma}_{\zeta}$) within this multivariate decomposition, and that the presence of potentially cointegrated trends does not pose here any particular problem. Furthermore, for all the series, the standard deviation of the residuals is also lower within the multivariate STM decomposition ($\widehat{\sigma}_{multi}$) than within the univariate one ($\widehat{\sigma}_{uni}$), as presented in the last column of table 13, which is also a sign that some precision can be gained from the multivariate decomposition.

We then compare the quality of our decomposition with the one resulting from a more popular detrending method, i. e. the Hodrick-Prescott (HP) symmetric linear filter (1997), with a smoothing parameter set to 100 (and not to 1600 as suggested by the two authors). We obtain the results shown in figures 8 and 9. The striking resemblance of the two estimated cyclical components is visible in the graphs presented in figures 8. Figure 9 presents the calculated coherence for the two cyclical components, for each country. From this last figure, we can conclude that the visible resemblance of the two cycles in each graphs of the figure 8 is confirmed by the relatively high value of the coherence for the most important business cycles frequencies. In a way, we can also assess that the multivariate STM decomposition has facilitated the identification of an "optimal" smoothing parameter for the HP decomposition (in our case 100), in order to extract a cyclical component in the quarterly GDP time series analysed.

Figure 8: Quarterly GDP time series, estimated cyclical components, multivariate trend-plus-cycle decomposition and Hodrick Prescott detrended series for comparision (lambda = 100)

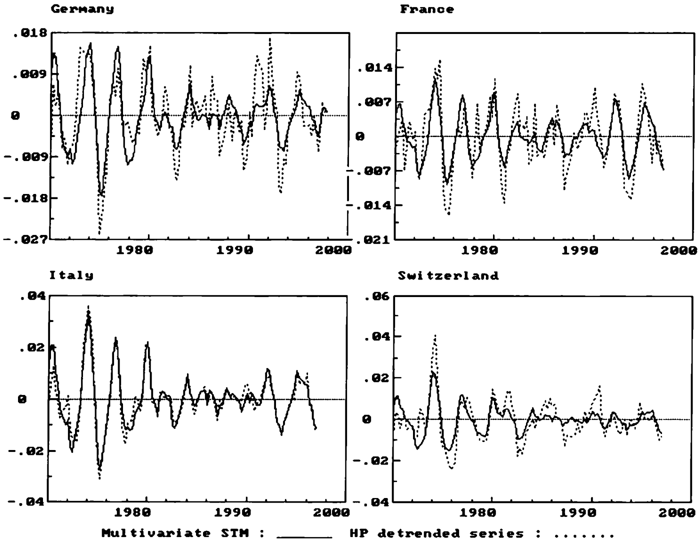
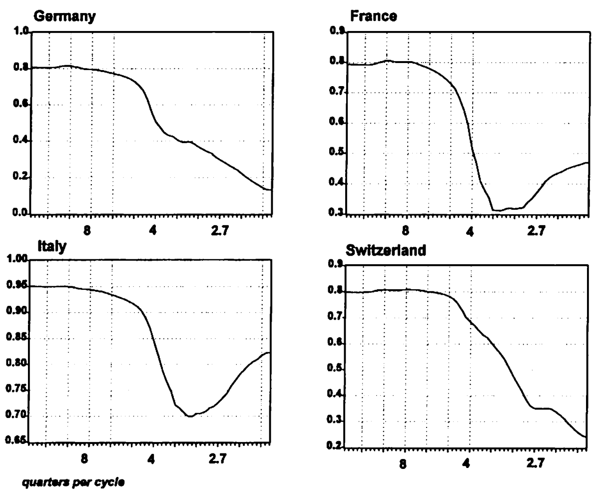


Figure 9: Correlation coefficients at different frequencies (estimated coherence) between cycles STM (multivariate approach) and HP detrended series (univariate decomposition, with lambda = 100)



To consider whether the European business cycle and the US business cycle share some characteristics in common, a second multivariate *trend-plus-cycle* model was estimated for the *group* 2 *(CH Fra Ita Ger USA)*. The main results are presented in table 14.

An examination of the decomposition obtained within the five dimensional structural model reveals that the results are now less convincing. For the US, the multivariate decomposition carried out with other European countries reduces the quality of the decomposition. The estimated one-step-ahead prediction errors (residuals) are more autocorrelated than in the univariate decomposition and the standard deviation of the residuals is also higher. Furthermore, an analysis of the vector of estimated residuals reveals that some autocorrelations seem not random, and more significant cross-correlated values are also obtained. Regarding the correlation among the disturbances driving the estimated cyclical components (matrice $\hat{\Sigma}_\xi$) and the correlation among the same estimated cyclical components (matrice $\hat{\Sigma}_{cycle}$), some very low values confirm that the European countries on one hand, and the US on the other, experience different business cycles. We can thus conclude that the use of structural multivariate time series models allows to assess in a rather easy way whether a group of countries shares common business cycles. This seems to be strongly the case for the four European countries considered *(CH Fra Ita Ger)*, but not for the same countries with the US. Interesting enough is the fact that Switzerland, although not being a former member of the European Monetary System (EMS) or now of the Euro zone, seems to really share a business cycle with ist European neighbours.

Table 14: Five dimensional trend-plus-cycle model (group 2); sample 1970:1/1997:4

	$\hat{\sigma}_\varepsilon$	$\hat{\sigma}_\eta$	$\hat{\sigma}_\zeta$	$\hat{\sigma}_\xi$	ρ	$2\pi/\lambda_c$	R_D^2	$Q(12)$	$P.val.\ N_{SB}$	$\hat{\sigma}_{uni}$	$\hat{\sigma}_{multi}$
Switzerland	10.84	48.98	33.20	22.19	0.95	12.73	0.24	48.97*	0.0004	78.07	77.22
France	22.34	33.04	14.99	17.88	0.95	12.73	0.20	6.83	0.4399	59.94	58.63
Italy	10.46	31.40	19.30	38.78	0.95	12.73	0.34	12.90	0.2769	73.28	69.41
Germany	15.30	52.88	18.04	21.19	0.95	12.73	0.64	17.55**	0.9915	73.71	69.99
USA	14.53	74.31	24.11	16.96	0.95	12.73	0.07	22.68**	0.0474	85.62	89.79

$\hat{\Sigma}_\zeta$ $\hat{\Sigma}_\xi$ $\hat{\Sigma}_{cycle}$

$$\begin{pmatrix} \cdots & 0.90 & 0.99 & 0.88 & 0.92 \\ \cdots & \cdots & 0.90 & 0.86 & 0.67 \\ \cdots & \cdots & \cdots & 0.88 & 0.92 \\ \cdots & \cdots & \cdots & \cdots & 0.80 \\ \cdots & \cdots & \cdots & \cdots & \cdots \end{pmatrix} \begin{pmatrix} \cdots & 0.75 & 0.95 & 0.69 & 0.66 \\ \cdots & \cdots & 0.92 & 0.71 & 0.04 \\ \cdots & \cdots & \cdots & 0.75 & 0.42 \\ \cdots & \cdots & \cdots & \cdots & 0.06 \\ \cdots & \cdots & \cdots & \cdots & \cdots \end{pmatrix} \begin{pmatrix} \cdots & 0.79 & 0.96 & 0.74 & 0.66 \\ \cdots & \cdots & 0.93 & 0.79 & 0.12 \\ \cdots & \cdots & \cdots & 0.80 & 0.46 \\ \cdots & \cdots & \cdots & \cdots & 0.14 \\ \cdots & \cdots & \cdots & \cdots & \cdots \end{pmatrix}$$

*Significant at 10% level **Significant at 5% level

6. COMMENTS AND CONCLUSION

One of the main conclusions of Arthis and ZHANG (1995) was namely that data rather clearly indicate the emergence of a group-specific European cycle in the ERM period (between 1979 and 1993), somewhat independent of the US cycle. Our findings suggest that the emergence of the European business cycle could also be independent of the ERM, given that even a non member country of the ERM, like Switzerland, has experienced the same short term economic fluctuations. Of course, a reason could simply be that small open economies are forced to import disturbances hitting dominant countries (this is one important conclusion of ARTIS and ZHANG) and to draw a more accurate conclusion, we should have analyzed the business cycle properties of a more important European country not member of the ERM. Returning to the case of Switzerland, its growing linkage during the period 1979-1997 in trade and finance with the ERM countries has certainly been facilitated by the relatively closed links between the Swiss franc and other European currencies, which in turn has strenghthened the presence of an European business cycle in Switzerland during this period. Last but not least, other explanations for the common business cycle hypothesis between Switzerland and other European countries, as the one proposed by the real business cycle theory, should not be ignored. In this context, it is then no more appropriate to consider only the importance of macro-economic policies and external shocks (and the presence of similar reaction functions among various countries) when explaining common business cycle fluctuations. Identical endogeneous shocks and similar interacting movements in profits, investments and credits are also to be considered, when trying to explain the common short-term dynamic among various countries.

The EK serial correlation common feature test and the multivariate structural time series models allow to investigate the quality of comovements of international time series within two particular frameworks. As we pointed out in the introduction, the two approaches refer to different business cycle definitions, and thus can lead to different conclusions. Regarding the Swiss case, it is interesting to note that whatever the definition we use for the short-term movements of aggregated output (first difference or short term stationary deviations from a more or less stochastic trend), we can assert that these short-term fluctuations are shared in common with other EU countries. This result suggests that the *phase* of economic activity fluctuations in Switzerland is driven by forces that are not country-specific but common to other EU countries. However, the same conclusion cannot be drawn for the *amplitude*, which may depend on various factors, that could be more country-specific. Regarding the consequences of various macroeconomic policies on aggregated series, like the real GDP, the analysis could gain in precision by focusing on the amplitude of the fluctuations and not necessarily on the phase. Furthermore, we would like to stress that our findings on Switzerland sharing common business cycle fluctuations with other EU countries do not rule out that it could also experience, from time to time, its own specific fluctuations. Our findings simply suggest that these "idiosyncrasic fluctuations" are not very important.

Regarding the six years' stagnation of the Swiss economy between 1991 and 1996, our analysis lets us draw some conclusions. When we consider that during those years, the Switzerland's main trade partners had to adopt a restrictive economic policy to fulfil the Maastricht criterions and in consequence experienced a period of slow growth, it is hard to believe that Switzerland, which shares a common business cycle component with them, could have experienced something else than slow growth as well. Of course, other typical factors that have already been mentioned by various authors, for instance the appreciation of the real value of the Swiss franc (on a trade-weighted basis) during the years 1993, 1994 and 1995, the slack in domestic demand caused by a sustained weakness of private consumption (LAMBELET, 1996 and NATAL, 1997), and negative fiscal policy impuls, have certainly played an important role. But our findings suggest that the *importance* and *the impact* of the European business cycle component should not be underestimated, when assessing the short term economic fluctuations of Switzerland. Another comment concerns the slowdown of the medium/long term evolution of the GDP quarterly series that can be estimated within an univariate decomposition or within the bivariate decomposition with Italy. Even if after six years of stagnation a time series decomposition is most likely to induce a slowdown of the medium/long run estimated growth path, the acceptance of this result has different economic policy implications. For instance, it fully sustains the current orientation of the economic policy in Switzerland, which in particular attempts to improve the supply side of the internal market. In other words, due to the fact that no other country in the EU experienced such a proctracted stagnation, the six years stagnation may finally be interpreted in two ways. Either the amplitude of the Swiss business cycle overreacted the amplitude of the EU business cycle (an issue that we did not test especially, but that stands out in some of the figures presented), or the six years stagnation are indeed a sign of a slowdown of the underlying medium/long run growth path of the economy. The common business cycle with other European countries, during the years 1991–1997, in addition to restrictive monetary and fiscal policies, represent sufficient factors to explain why the amplitude of the cycle could have overreacted in Switzerland. Regarding a slowdown of the medium/long-term economic growth, to some extent this hypothesis could be explained by the orientation of Swiss structural economic policy (series of reforms after the rejection of the European Economic Space (EEU), deregulation of the sheltered sector), the harsher international competition (due to deregulations in neighbouring countries as well), and to some extent by the limited growth of input factors during this period. All these elements may have had temporary contractionary effects.

Since the signing of the Maastricht Treaty, important progress has been made in the EU countries in reducing inflation and fiscal imbalances and with respect to nominal convergence (interest and exchange rates). Furthermore, the common monetary policy of the Euro area since the beginning of 1999 is expected to reinforce this process and to intensify a real convergence, i. e. the correlation among GDP growth rates in the Euro-Zone (already strong according to the results of ARTIS and ZHANG) and maybe a common business cycle component (already of considerable importance according to our results). During the last six-seven years, the high degree of *nominal* convergence achieved in Europe has certainly been supported and made easier by the existence of a *real* common business cycle component among most of the EU countries (except the UK), as demonstrated by our paper and by other studies. Finally, even if this real short-term common dynamics is likely to be reinforced within the Euro area, we should keep in mind that for many countries, including Switzerland, it has existed for many years despite the presence of various currencies, which suggests that in addition to the monetary transmission mechanisms and the exchange rate regime, there are other strong forces linking economies together.

For the future, one of the main challenge of the European authorities, in order to reinforce the efficiency of the common monetary policy, will be to keep to a minimum business cycle disparities among regions. Given that Euro area monetary policy is no more able to respond to regional imbalances, ideally fiscal flexibility and more structural adjustments are required. But as we know, fiscal adjustments are restricted by the stability and growth pact, and structural adjustments are very slow processes. This situation has led some observers to warn against the potentially catastrophic consequences of imposing a single currency to an ill-suited area. Our main comment, that a real short-term common dynamic does effectively exist in Europe and not only among members of the Euro zone, should represent, however, a good argument for the Euro supporters. Regarding Switzerland, the national authorities could try to take more advantages of this real common short-term dynamics with the rest of Europe, for instance to better adjust the timing of various economic policy measures. During the first part of the 1990s, when important countries in the EU were already entering in a period of slow growth, postponing some fiscal consolidation measures and having a more gradual restrictive monetary policy stance, would have maybe avoided the long stagnation's period of the Swiss economy, but would not have been enough to generate a strong economic growth. From a more general point of view, on the basis of all the various considerations outlined in this paper, analysing the business cycle properties in the Euro area can enhance the understanding and appreciation of the Swiss business cycle. Furthermore, following the business cycle in the neighbouring countries also represents an important prerequisite for the setting and the timing of any important economic policy decision in a small open economy like Switzerland. Even if pretty obvious, this conclusion is worth repeating often.

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SUMMARY

Contemporaneous comovements of aggregated output time series between industrialized countries represent an important characteristic of the international business cycle and have already been well documented. This paper shows that a small open economy like Switzerland shares common short term GDP fluctuations with other European countries, independently of the specific definitions used for the *common component.* The serial correlation common feature (SCCF) hypothesis of ENGLE and KOZICKI (1993) and also the univariate and multivariate structural time series models (HARVEY, 1989) are presented and applied. Sharing common short-term economic fluctuations with other European countries has various economic policy implications for Switzerland. This last issue is commented in the light of the new commom monetary policy framework in the Euro area, since the beginning of 1999.

ZUSAMMENFASSUNG

Gleichlaufende Schwankungen in Zeitreihen aggregierter Daten aus verschiedenen Industrieländern stellen eine wichtige Eigenschaft des internationalen Konjunkturzyklus dar. Sie sind in der Literatur gut dokumentiert. Der vorliegende Artikel zeigt, dass eine kleine offene Volkswirtschaft wie die Schweiz gemeisame kurzfristige BIP-Schwankungen mit anderen europäischen Ländern aufweist, unabhängig von der verwendeten spezifischen Definition der gemeinsamen Komponente. Die Hypothese von ENGLE und KOZICKI (1993) über gemeisame Eigenschaften serieller Korrelation (SCCF) sowie die univariaten und multivariaten strukturellen Zeitreihenmodelle (HARVEY, 1989) werden vorgestellt und angewandt. Aus der Tatsache, dass die Schweiz mit anderen europäischen Ländern gemeisame Konjunkturschwankungen aufweist, lassen sich verschiedene wirtschaftspolitische Schlüsse ziehen. Diese werden im Lichte der gemeinsamen Geldpolitik seit 1999 im Euro-Raum erörtert.

RESUME

Les fluctuations communes des séries temporelles représentant des variables économiques agrégées de différents pays industrialisés représentent une caractéristique du cycle conjoncturel international. L'existence de telles fluctuations a été considérablement documentée dans la littérature. Cet article montre que la Suisse, petite économie ouverte, partage des fluctuations conjoncturelles communes avec d'autres pays européens, indépendamment des définitions qui sont retenues pour qualifier de «communes» de telles fluctuations. L'hypothèse de composante commune de corrélation sérielle de ENGLE et KOZICKI (1993), de même que les modèles univariés et multivariés pour les séries temporelles (HARVEY, 1989) sont présentés et utilisés. Le fait que la Suisse partage avec d'autres pays européens des fluctuations conjoncturelles communes a plusieurs implications pour la gestion des politiques économiques suisses. Ce dernier aspect est commenté sous différents angles, notamment à la lumière de la nouvelle politique monétaire commune introduite dans la zone Euro dès le début de l'année 1999.