The Permanent Component of GNP and Consumption: Results from an Univariate Analysis for Switzerland

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1. INTRODUCTION

Traditional empirical studies of macroeconomic fluctuations did not spend a lot of time to model the permanent or trend component of macroeconomic time series. The Keynesian analysis of the 60's as well as the neoclassical studies of the 70's proceeded on the assumption that the long run behavior of macroeconomic variables can be modeled as deterministic trend functions, which were to reflect the smoothly changing supply conditions. Thus, macroeconomic fluctuations were mainly understood as transitory responses of the economy to demand disturbances. At the beginning of the 80's, this view was challenged. On an empirical level, the very influential paper of NELSON and PLOSSER (1982) argued that the trend behavior of most macroeconomic time series is adequately represented by stochastic and not deterministic trends. This implies that the effect of shocks on the level of macroeconomic variables is to a certain extent permanent and not purely transitory as in the deterministic trend case. Moreover, NELSON and PLOSSER claim that a substantial proportion of the fluctuations of US GNP growth is accounted for by permanent shocks.[1] This finding was considered as empirical support for the concept of real business cycles models as formulated by KYDLAND and PRESCOTT (1982) and LONG and PLOSSER (1983), which model economic fluctuations as equilibrium responses of the economy to real shocks.

The NELSON/PLOSSER paper triggered-off a lot of subsequent research which tried to establish the importance of the permanent or persistent component of the US GNP fluctuations [i.e. WATSON (1986), CAMPBELL/MANKIW (1987a, b), COCHRANE (1988)]. These studies, however, do not unambiguously support the NELSON/PLOSSER findings.

Most recently, the relevance of the results of the work cited above was questioned. The papers of CHRISTIANO and EICHENBAUM (1989) and LIPPI and REICHLIN (1989) are two excellent sources developing and summarizing this critic. Two points are to be mentioned. First, the measures of persistence developed in this literature say nothing about the relative importance of transitory or cyclical and permanent or persistent shocks for macroecon-

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1. Nearly at the same time, the related paper by BEVERIDGE and NELSON (1981), which develops an approach to estimate the permanent component of a time series, appeared.

omic fluctuations. In order to analyse this question, arbitrary and sometimes doubtful identifying restrictions have to be adopted in this literature. Moreover, this work provides us only with rather imprecise estimates of the long run effect of shocks. Second, the central stochastic trend (unit root) assumption adopted in this literature becomes doubtful when it is compared to a framework of a deterministic trend with few breaks as developed by PERRON (1989) and RAPPOPORT and REICHLIN (1989).

The primary aim of this paper is to estimate the effect of permanent shocks to GNP and private consumption in Switzerland critically using the univariate time series methodology applied in the papers cited above. The remaining part of this paper is organized as follows: Sections 2 and 3 briefly discuss the methodology used. The empirical results are contained in Section 4, and Section 5 concludes.

2. UNIVARIATE MEASURES OF PERSISTENCE

Let y_t be the time series of interest and consider the following univariate time series representation of its first difference Δy_t

$$\begin{aligned}\Delta y_t &= \mu + \varepsilon_t + c_1\, \varepsilon_{t-1} + c_2\, \varepsilon_{t-2} \ldots \\ \Delta y_t &= \mu + C(L)\, \varepsilon_t \end{aligned} \tag{1}$$

where $C(L) = 1 + c_1L + c_2L^2 \ldots$ is a stable lag operator polynomial with the properties $\Sigma \mid c_j \mid < \infty$ and $\Sigma\, c_j = C(1) \neq 0$. ε_t is white noise with variance σ^2_ε and μ is a scalar constant term.

$C(1)$ gives the accumulated effect of a ε_t shock on the level of y_t. Thus, it is natural that $C(1)$ is used as a measure of persistence or permanence of the effects of shocks on y_t. The role of $C(1)$ can be highlighted by considering the alternative model of trend stationarity for y_t.

$$y_t = \delta + \mu t + A\,(L)\, \varepsilon_t \tag{2}$$

where $A(L)$ is a stable lag operator polynomial and δ a level constant term. Of course, this model implies that ε_t has no permanent effect on y_t, as in the long run y_t reverts to the deterministic trend $\delta + \mu t$. Taking first differences provides us with

$$\Delta y_t = \mu + (1-L)\, A\,(L)\, \varepsilon_t\,.$$

Of course, this corresponds to (1) when we define $C(L) = (1\text{-}L)A(L)$. However, $C(1)$ is obviously zero in the trend stationarity case, as it should be if we use it as a measure of persistence.

NELSON and PLOSSER analysed this problem with a two-step approach. First, they tested the hypothesis $C(1) \neq 0$ (the unit root hypothesis) using the unit root tests of DICKEY and FULLER (1979, 1981).[2] Second, after accepting $C(1) \neq 0$ (the unit root hypothesis), they fitted ARMA models for Δy_t $[C(L)=\Theta(L)/\Phi(L)]$ and analysed the properties of $C(L)$.[3]

The estimation of long run properties of time series by ARMA models was criticized by COCHRANE (1988) who argues that ARMA models are primarily designed to capture the short run dynamics and are not suited to estimating the long run effects of shocks.

He suggests a non-parametric measure of persistence, namely

$$V^k \equiv \frac{var\,(y_t - y_{t-k})}{k\,\sigma^2_{\Delta y}} = 1 + 2\sum_{j=1}^{k-1} \frac{k-j}{k}\,\rho_j \tag{3}$$

where $\sigma^2_{\Delta y} = Var\,(\Delta y_t)$ and ρ_j are the j'th autocorrelation coefficients of Δy_t. V^k is zero when Δy_t is uncorrelated and it is lower (higher) than 1 when Δy is mainly negatively (positively) autocorrelated. These properties give an intuitive motivation for V^k as a measure of persistence. In addition, it can be shown that there is a close relation to $C(1)$, namely

$$\lim_{k \to \infty} V^k = (C(1))^2 \frac{\sigma^2_{\varepsilon}}{\sigma^2_{\Delta y}} \tag{4}$$

Thus, V^k, the standardized variance of $y_t - y_{t-k}$, tends to zero if the variable is trend stationary $(C(1) = 0)$.

V^k can be estimated in different ways [COCHRANE (1988), CAMPBELL/MANKIW (1987a)]. The simplest approach consists of inserting the sample autocorrelation coefficient in (3).

Finally we should note that time aggregation may have a crucial influence on the value of $C(1)$. As shown by LIPPI and REICHLIN (1989) temporal aggregation may result in a sizeable decrease in $C(1)$ when $C(1)>1$ and an increase in $C(1)$ when $C(1)<1$.

2. The test is essentially based on the regression of $\Delta y_t = \alpha + \beta t + \gamma y_{t-1}$ and considers the hypothesis $\gamma = 0$. The extension of this approach developed by PERRON (1989) uses the broken trend $\alpha_1 + \alpha_2 d_t + \beta_1 + \beta_2 d_t t$ instead of $\alpha + \beta t$, where d_t is the dummy variable for the break.

3. There is the alternative approach adopted by CAMPBELL and MANKIW (1987a), who estimate C(1) by an ARMA model and test C(1)=0 in this framework.

3. PERMANENT AND TRANSITORY COMPONENTS

So far we considered some approaches to estimate the persistence or permanence of the effects of shocks on y_t. Now it is tempting to assume that a $C(1)$ value clearly larger than one shows the relative importance of permanent shocks for the fluctuations of y_t. However, such a conclusion is completely wrong. In order to estimate the relative importance of transitory and permanent movements one has to adopt additional (arbitrary) assumptions. The paper of BEVERIDGE and NELSON (1981) proceeds on the assumption that permanent and transitory movements of y_t are driven by only one independent shock. In addition, the permanent component is modeled as a random walk. Let us define the permanent or trend component as T_t and the transitory or cyclical component as c_t:

$$y_t = T_t + c_t$$

Taking first differences yields

$$\Delta y_t = \Delta T_t + \Delta c_t \tag{5}$$

On the other hand we have the univariate representation (1) of Δy_t. This representation can be written as

$$\Delta y_t = [C(1) + (1-L)\, B\,(L)]\, \varepsilon_t \tag{1a}$$

where the coefficient of $B(L)$ can be determined by using the condition

$$C(L) = C(1) + (1-L)\, B\,(L)$$

and by equating the coefficients of the same power of L.

Obviously, the random walk trend is defined by $\Delta T_t = C(1)\varepsilon_t$ and the cyclical component is given as $B(L)\varepsilon_t$. The change of the trend component, of course, is the long run effect of the shock to y_t. As shown by BEVERIDGE and NELSON (1981) the level of the trend component can be estimated as

$$T_t = y_t + (\sum_{j=1}^{\infty} c_j)\, \varepsilon_t + (\sum_{j=2}^{\infty} c_j)\, \varepsilon_{t-1} + \dots \tag{6}$$

Thus, the accumulated effects of all past shocks, which are not yet contained in the current level of y_t are added to y_t. It may be helpful to show implications of the

Beveridge-Nelson decomposition within two simple examples. First, let us consider the case of an invertible MA(1) process[4] for Δy_t, i.e. $C(L)=1+\theta L$ ($|\theta|< 1$). It is easily verified that the decomposition is as follows

$$\Delta T_t = (1+\theta)\,\varepsilon_t$$
$$c_t = -\,\theta\varepsilon_t$$

The cyclical component is white noise. When Δy_t is positively (negatively) autocorrelated the cyclical component is perfectly negatively (positively) correlated with the change in the permanent component. The variance of ΔT_t $[(1+\theta)^2\sigma_\varepsilon^2]$ is necessarily larger than the variance of c_t $[\theta^2\sigma_\varepsilon^2]$ and Δy_t $[(1+\theta^2)\,\sigma_\varepsilon^2]$ when $\theta > 0$ holds, i.e. the trend is *more* volatile than the series itself. For the case $\theta < 0$ this is no longer true. In addition the variance of the cycle is larger than that of the change in the trend when $\theta < -0.5$.

Second, we will discuss the AR(1) case $C(L) = \dfrac{1}{1-\phi L}$, ($\phi < 1$). In this case the decomposition is obtained as

$$\Delta T_t = \frac{1}{1-\phi}\,\varepsilon_t$$
$$c_t = -\frac{\phi}{1-\phi}\,\frac{1}{1-\phi L}\,\varepsilon_t$$

Thus, the cyclical component follows the same AR(1) pattern as the series itself. ΔT_t and c_t are perfectly negatively (positively) correlated when ϕ is larger (lower) than zero.

The variance of ΔT_t is

$$\left(\frac{1}{1-\phi}\right)^2\sigma_\varepsilon^2$$

and that of c_t is

$$\left(\frac{\phi^2}{(1-\phi)^2}\right)\frac{1}{1-\phi^2}\,\sigma_\varepsilon^2$$

4. This is the model adopted by NELSON and PLOSSER for annual US GNP.

Of course $Var(c_t)$ is larger than $Var(\Delta T_t)$ when $\frac{\phi^2}{(1-\phi)^2}$ or $|\phi| > \sqrt{1/2}$ holds. Finally $Var(\Delta T_t)$ is again larger (lower) than $Var(\Delta y_t)$ when $\phi > 0$ (< 0) holds.[5]

These two examples show that the Beveridge-Nelson decomposition may have rather strange implications in the frameworks of the often used MA(1) and AR(1) models. This is, of course, the consequence of the arbitrary identifying restrictions (T_t random walk and perfect correlation of ΔT_t and c_t). If we keep the random walk assumption and assume no correlation of ΔT_t and c_t as in WATSON (1986) we would obtain different result. However, this model will not be considered here as it is restrictive in the sense that it implies $C(1) < 1$ [LIPPI/REICHLIN (1989)]. Finally note that if we replace the random walk assumption by a general ARMA model for ΔT_t there is an infinite number of ways to decompose Δy_t in orthogonal trend and cycle components [QUAH (1988), CHRISTIANO/EICHENBAUM (1989)].

4. EMPIRICAL RESULTS

In this section the results obtained by applying the methods outlined above to Swiss GNP and private consumption data are reported. Yearly data covering the period 1949-88 [source: Die Volkswirtschaft, several issues] are used. Quarterly National Account data are available since 1965 in Switzerland. However, a time span of hardly 25 years seems a little bit short to analyse the trend behavior of GNP and consumption. Table 1 summarizes the univariate time series properties of the two series. For the log of GNP and the log of consumption the unit root hypothesis clearly cannot be rejected by the augmented Dickey Fuller test accounting either for one or two lagged differences. The changes of these series are positively autocorrelated. For GNP a MA(1) model seems to be appropriate, whereas an AR(1) model fits the changes in log consumption. It should be mentioned that an AR(1) gives essentially the same result with respect to the long run properties of the series in the latter case. Both rate of change series are positively autocorre- lated and the measure of persistence $\hat{C}(1)$ is, therefore, larger than one. In particular for consumption a shock is magnified by 2.5 in the long run whereas this figure is only 1.5 for GNP. Table 2 reports the Cochrane-measure of persistence V^k. In addition, the standard error of this statistic is estimated following CHRISTIANO and EICHENBAUM (1989) and using a formula given by PRIESTLY (1982). Finally, V^k is used to calculate an implicit value for $\hat{C}(1)$ using a formula developed by MANKIW and CAMPBELL (1987a). Thereby, we used a lag k ranging from 5 to 15 years, the maximum lag number which

5. The conditions to be exploited are $1 + \phi L + \phi L^2 \ldots = \frac{1}{1-\phi} + (1-L)\,[b_0 + b_1 L + b_2 L^2 \ldots]$ obviously $b_0 = 1 - \frac{1}{1-\phi} = \frac{-\phi}{1-\phi}$ and b_1 is determined according to $\phi = -b_0 + b_1$, which gives $b_1 = \phi - \frac{\phi}{1-\phi} = \frac{-\phi^2}{1-\phi}$. $\phi^2 = -b_1 + b_2$, then implies $b_2 = \phi^2 - \frac{\phi^2}{1-\phi} = -\frac{\phi^3}{1-\phi}$ and in general $b_j = -\frac{\phi}{1-\phi}\phi^j$ holds.

seems reasonable given our sample. This exercise gives similar results as the ARMA models of table 1. However, the estimated standard errors of $\hat{V}^{k}$ show that the estimates of persistence are rather imprecise. This is, of course, not surprising given the sample of 40 years.

The ARMA models of table 1 are now used to decompose the series into a permanent and transitory component according to BEVERIDGE and NELSON. These results are presented graphically in figures 1 and 2. As we know from section 3 an MA(1) model with $\theta > 0$ implies that the trend component is more variable than the series itself. This property is clearly seen in figure 1: The permanent GNP component closely follows the actual time path, as the cyclical component is white noise, but it is more volatile. For consumption things look different. The cyclical component follows an AR scheme and leads to sustained deviations of the actual from the permanent level of the series. Again, the permanent component is more volatile than the actual series. Finally let us briefly consider what happens to our result, if we use a model of a broken trend as an alternative to the difference stationary model. To this end we applied PERRON's unit root test taking into account an oil crisis break in the mid of the seventies (1949-74, 1975-88). The results of this exercise are reported in table 3. For GNP the unit root hypothesis can be rejected at least at the 5% significance level for one and two years lagged differences. For consumption this is only true for the first variant of the test. Thus, taking into account only one structural break in the trend during the 40 years leads to serious doubts about the persistence of shocks found on the difference stationary model. It suggests that only "epochal" shocks like the oil crises have a permanent effect on the level of GNP and consumption in Switzerland, whereas "ordinary" shocks have only transitory effects.

5. CONCLUSION

This paper estimates the persistence of shocks to Swiss GNP and consumption and decomposes these variables into permanent and transitory components according to the method developed by BEVERIDGE and NELSON. The application of the standard Dickey Fuller unit root test indicates that the series are difference stationary and a parametric as well as a non-parametric measure of persistence indicate that the long run impact of shocks to GNP and consumption is larger than its short run effect. In addition, the Beveridge-Nelson decomposition indicates the change in the trend component to be more volatile than the cyclical component. Both findings are consistent with the view that permanent or real shocks are the most important source of business cycle fluctuations. However, these rather clear cut results become ambiguous when we take account of the two following points. First, the existence of a unit root in GNP and consumption becomes doubtful when we account for a break in the middle of the seventies. Second, the Beveridge-Nelson decomposition is based on arbitrary identifying restrictions, namely a random walk permanent component and perfect correlation between transitory and permanent shocks. The first point means that only "epochal" shocks like the first oil shock have permanent effects whereas "ordinary" shocks are only transitory. The second point calls for a multivariate analysis of the relative importance of transitory and permanent shocks.

APPENDIX

Table 1:

Time Series Properties of Changes in log GNP and Consumption, Yearly Data 1949–88

	GNP				Consumption			
lag	$t_{\hat{\gamma}}$	$\hat{\rho}$	$\hat{\theta}$	$\hat{C}(1)$	$t_{\hat{\gamma}}$	$\hat{\rho}$	$\hat{\phi}$	$\hat{C}(1)$
1	−1.13	0.39	0.47	1.47	−1.39	0.68	0.88	2.50
		(0.16)	(0.16)			(0.16)	(0.16)	
2	−0.90	0.032			−0.68	0.31	−0.28	
							(0.16)	
3		0.11				0.14		
4		0.15				0.057		
5		0.075				0.079		
6		0.11				0.22		
7		0.11				0.29		
8		0.15				0.27		
9		0.17				0.16		
10		0.15				0.015		
12		−0.073				−0.23		
15		−0.15				−0.21		
Q(5)			0.81				1.02	
Q(15)			3.72				6.33	

$t_{\hat{\gamma}}$: Dickey Fuller Statistic including constant, time and lagged differences

$\hat{\rho}$: Estimated Autocorrelation coefficient

$\hat{\theta},\hat{\phi}$: Estimated MA and IR coefficient

$\hat{C}(1)$: Estimated persistence

Estimated Standard error in parenthesis

Q(k): Box and Pierce statistics (k lags) for the fitted ARMA model

Table 2:

Non-Parametric Measure of Persistence,
Yearly Data 1949-88

	GNP			Consumption		
k	$\hat{V}^k$	se($\hat{V}^k$)	$\hat{C}(1)^k$	$\hat{V}^k$	se($\hat{V}^k$)	$\hat{C}(1)^k$
5	1.41	0.64	1.29	1.83	0.83	1.83
6	1.54	0.75	1.35	2.12	1.04	1.97
7	1.67	0.88	1.40	2.33	1.22	2.07
8	1.79	0.99	1.45	2.50	1.39	2.14
9	1.91	1.12	1.50	2.68	1.57	2.22
10	2.02	1.24	1.54	2.89	1.77	2.30
11	2.14	1.37	1.59	3.12	2.00	2.39
12	2.27	1.52	1.64	3.33	2.22	2.47
13	2.41	1.67	1.69	3.50	2.42	2.54
14	2.53	1.81	1.73	3.63	2.60	2.58
15	2.63	1.94	1.76	3.72	3.75	2.61

$$\hat{V}^k = 1 + 2\sum_{j=1}^{k-1} ((k-j)/k)\,\rho_j$$

$$se(\hat{V}^k) = \hat{V}^k/[0.75[T/(k+1)]]^{1/2}$$

$$\hat{C}(1)^k = [\,\hat{V}^k/(1-\hat{\rho}_1^2)]^{1/2}$$

Table 3:

Perron's Unit Root test
Yearly Data 1949–88

$$\Delta y_t = \gamma y_{t-1} + \alpha_1 + \alpha_2 d_t^{75} + \beta_1 t + \beta_2 t d_t^{75} + \sum_{j=1}^{k} \phi_j \Delta y_{t-j}$$

	GNP		Consumption	
γ	−0.76	−0.72	−0.47	−0.41
	(−5.86)	(−4.42)	(−4.22)	(−3.22)
α_1	7.91	7.51	4.69	4.01
	(5.89)	(4.45)	(4.42)	(3.24)
α_2	0.41	0.37	0.30	0.25
	(0.11)	(2.67)	(3.63)	(2.59)
β_1	0.034	0.032	0.020	0.017
	(5.88)	(4.33)	(4.23)	(3.20)
β_2	−0.019	−0.018	−0.012	−0.011
	(−4.62)	(−3.34)	(−3.87)	(−2.90)
ϕ_1	0.32	0.31	0.55	0.61
	(2.42)	(2.26)	(3.88)	(4.03)
ϕ_2		−0.060		−0.19
		(−0.41)		(−1.10)

"t–values" are given in parenthesis. One sided critical values for the hypothesis $\gamma = 0$ are:
10%: −3.95, 5%: −4.24, 1%: −4.88 (Perron, 1989, Table VI.B)

Figure 1:

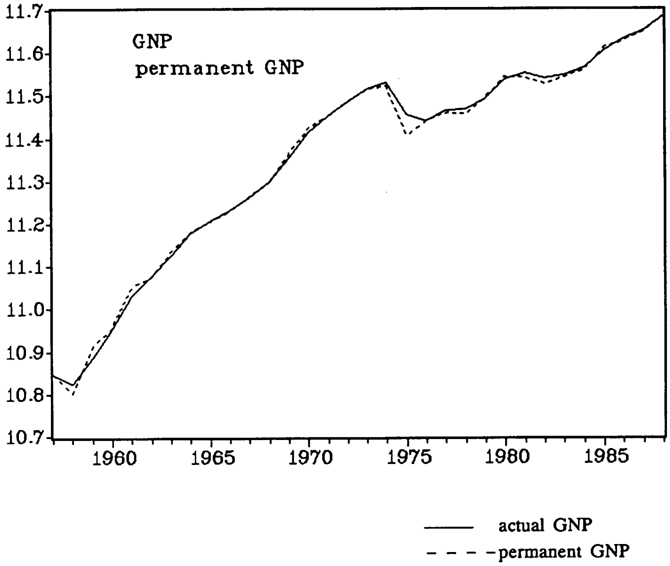
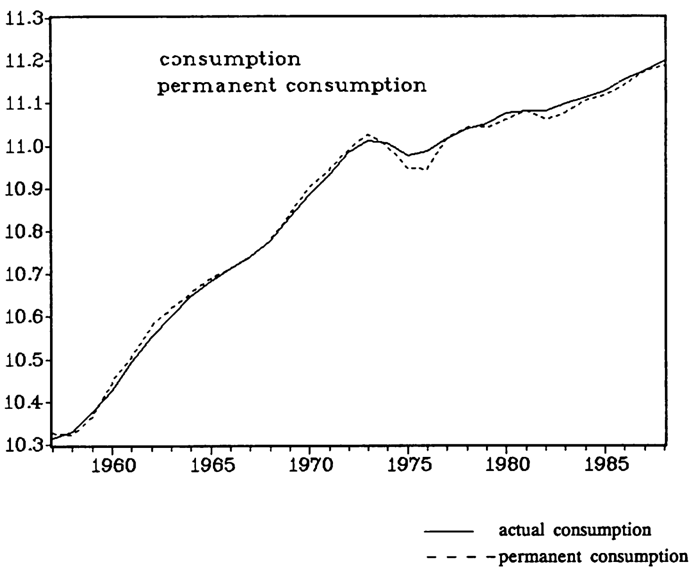


Figure 2:



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SUMMARY

This paper estimates the persistence of shocks to Swiss GNP and consumption, and decomposes these variables into permanent and transitory components according to the method developed by BEVERIDGE and NELSON. The application of the parametric as well as a non-parametric measure of persistence to first differences indicates that the long run impact of shocks to GNP and consumption is larger than its short run effect. In addition, the Beveridge-Nelson decomposition indicates the change in the trend component to be more volatile than the cyclical component. Both findings are consistent with the view that permanent or real shocks are the most important source of business cycle fluctuations. However, these rather clear cut results become ambiguous when we account for a break in the middle of the 70's.

ZUSAMMENFASSUNG

In der vorliegenden Arbeit wurde die Persistenz der Effekte von Schocks auf das reale BSP und den realen Konsum geschätzt, sowie diese beiden Variablen mit der Methode von BEVERIDGE und NELSON in eine transitorische und eine permanente Komponente zerlegt. Die Analyse der ersten Differenz weist darauf hin, dass der langfristige Effekt von Schocks grösser ist als die kurzfristige Auswirkung, und dass die Trendkomponente volatiler ist als die zyklische Komponente. Diese Resultate deuten auf die Relevanz von permanenten (realen) Schocks für die Konjunkturschwankungen hin. Allerdings führt die Berücksichtigung eines Trendbruches in der Mitte der 70er Jahre zu einer Relativierung dieser Ergebnisse.

RESUME

Dans l'article présent, la persistance des shocks auxquels sont soumis le produit national brut réel et la consommation réelle est mesurée et les composantes transitoires et permanentes des deux variables sont estimées d'après la méthode présentée par BEVERIDGE et NELSON. L'analyse des séries différenciées démontre que l'effet à long terme des shocks est plus important que celui mesuré à court terme et que la composante permanente est plus volatile que le composante cyclique. Ces résultats confirment la rélévance de shocks (réels) permanents étant à l'origine des cycles conjoncturels. Pourtant, cette conclusion devient ambigue en prenant compte d'une rupture de trend dans les années septante.