What drives Health Care Expenditure since 1950 in France?[1]

A time-series study with structural breaks and non linearity approaches

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Preliminary paper - Do not quote

This study is preliminary and must be improved. It is only at this stage a working paper for discussion at Third joint CES-HESG meeting. The text should be read by a native English speaker.

Introduction

The emergence of State's Debt and challenging phase of the Welfare State conduct analysts to better understand the dynamics of social spending. Particularly, the preservation of population health is a concern shared by all OECD countries. Health Care Expenditure (HCE) have steadily risen in OECD countries and have therefore attracted a great attention in the political and scientific debate. The share of GDP devoted to HCE significantly varies across countries and on average reaches 8.8% in 2008. France dedicates to health one of the highest share of GDP in the world (11.2 % behind USA with a proportion of 16.2%).

France suffers from a lack of macroeconomic studies able to analyze the determinants of HCE cause of the weakness of available data. In particular, it seems necessary to understand how various health system reforms and economic crisis have affected HCE growth in France, taking into account non-stationarity and cointegration (co-movement over the long run) properties between time series. Indeed, a recent empirical literature has investigated these classical problems by controlling structural breaks (co-movement over the short run) and cross-section dependence. For example, Jewell *et al.* (2003) and Carrion-i- Silvestre (2005) conclude that HCE and GDP are stationary around at least one break. In this connection, Narayan and Narayan (2008) advice to extend work on identifying turning points in GDP and HCE time series.

Our goal, for the first time with French data, is to explain the role of GDP on HCE taking into account for structural breaks and non linearity on the long-run economic relationship between HCE and GDP controlling by price effect, innovation proxy and Medical density. We use 1950-2009 annual data on HCE (called CSBM, Medical Care Consumption) from French Health Data Bases 2011 from Irdes (Institute for Research and Information in Health Economics). In first section, we present a short review of literature and specificities of French Health Care System. Thus, we investigate empirical issues by cointegration analysis (section 2). Into the section 3, we take into account for structural breaks. The section 4 deals with relationships between HCE and GDP with non linearity approach. Finally, we propose a provisional finding.

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Section 1. Brief overview of the literature and specificities of French case

HCE, particularly in OECD countries, have provided a huge and varied empirical literature. Many studies investigated the HCE determinants, particularly income. French case has not benefited from this field of research particularly because of data problems.

1.1. HCE Determinants

The literature shows that the macroeconomic determinants of HCE concern to income effect (Newhouse, 1977), Technology (Newhouse, 1992), Demographic and epidemiological changes (Culyer, A.J., 1988; Zweifel *et al.*, 1999), Price effect (Baumol, 1967), Induced demand (Evans, 1974; Fuchs, 1978) and Institutional factors (Health system reforms, physician remuneration, out of pocket (Hitiris, T., Posnett, J., 1992)). But these determinants have not received equal attention in international studies.

The most robust determinant of HCE is GDP. Following the seminal work of Newhouse (1977), most studies investigate the long-run economic relationship between HCE and income in a panel approach (McCoskey and Selden, 1998; Roberts, 1999; Gerdtham and Lothgren, 2000; Okunade and Karakus, 2001; Jewell et al., 2003; Carrion-i-Silvestre, 2005; Dregen and Reimers, 2005; Freeman, 2003; Wang and Rettenmeier, 2006; Chou, 2007). Some studies examine the stationarity and cointegration properties of HCE (MacDonald and Hopkins, 2002; McCoskey and Selden, 1998; Hitiris, 1997; Jewell et al., 2003; Narayan, 2006). Newhouse showed income elasticity significantly greater than unity (1.35), GDP per capita in explaining 92% of variations in HCE. Considering the short time dimension of the HCE variable, most studies have conducted non stationary panel tests since the middle of the 2000's to improve the power of the stationarity and cointegration tests. Recent works insist especially on the cross-section dependence which appears as an important characteristic of health data (Jewell *et al.*, 2003; Freeman, 2003; Carrion-i-Silvestre, 2005; Wang and Rettenmaier, 2006; Chou, 2007). For example, using a panel of 20 OECD countries during 34 years, Baltagi and Moscone (2010) show that health care is a necessity rather than a luxury, after controlling for cross-country dependence and unobserved heterogeneity. Moscone and Tosetti (2010), using a panel of 49 US States followed over 25 years, study the same relation but the cross-section dependence is incorporated in models through an approximate multifactor structure (Bai and Ng, 2004). More precisely, they adopted a bootstrap method called the Continuous-Path Block Bootstrap (CBB) proposed by Paparoditis and Politis (2001, 2003), and recently used by Fachin (2007) to test for panel co-integration in the presence of cross-section dependence. They find elasticity comprised between 0.35 and 0.46. For USA, this result is confirmed by Freeman (2003) with elasticity equals to 0.82 (with dynamic OLS approach). Herwatz and Theilen (2010) nuance the debate about the nature of health care good. They show that, on average, health care is a necessity good in presence of relatively young population but seems to become a luxury good in aging economies.

Since Newhouse (1992), technology is also considered as an important driver of HCE. However, it appears very difficult to use an appropriate proxy for changes in medical care technology. That's why several proxies are used as surgical apparatus (Baker and Wheeler, 2000; Weil, 1995), health care R&D spending (Okunade and Murthy, 2002), life expectancy and infant mortality (Dregen and Reimers, 2005) or a time index (Gerdtham and Lothgren, 2000).

Moreover, what is the contribution of demographic and epidemiological changes to HCE growth? Results are not so obvious even if the simple correlation between HCE and age appears clear. The explanatory power of age has been discussed by Zweifel *et al.*, (1999) and age doesn't constitute a stable explanatory variable for HCE, since morbidity and mortality are changing over time. Then health status and "time to death" arise determinant to explain HCE (Yang, Norton *et al.*, 2003, Seshamini and Gray, 2004).

Moreover since Baumol theory (1967), we know that health sector's productivity is low relative to other sectors causing inflation in health spending. However, there is no empirical consensus on the effect of real prices on HCE. Hartwig (2008) and Okunade *et al.* (2004) find a positive and statistically significant effect, and Gerdtham *et al.* (1992) and Murthy and Ukpolo (1994) report an insignificant effect. This appears difficult to study this determinant particularly into countries which insure a high level of public expenditures.

The seminal paper of Arrow (1963) emphasized that imperfect information of the health care consumer is recognized as a key feature of market failure. Induced demand then defines the physician's ability to provide a medical service delivery different from the patient would choose if he benefited from the same information (Rice, 1983). This assumption is plausible when physician are paid to fee for service which is the case in France.

1.2. What about French studies?

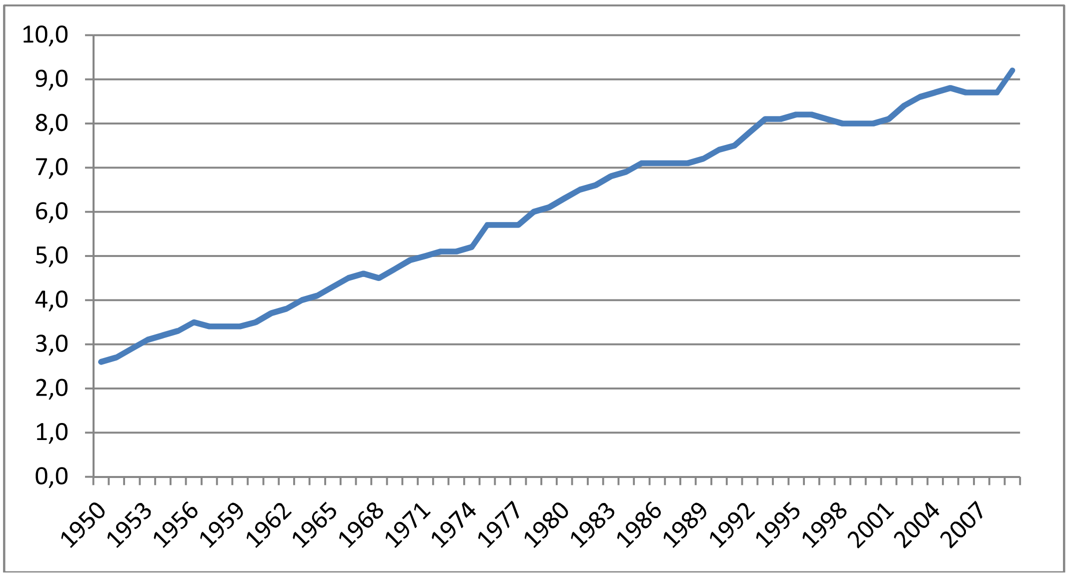
Firstly, we can notice that France is not retained in the sample in a large majority of comparative studies of OECD Countries panel. Only French authors have attempted to integrate France in their comparative studies (Mahieu, 2000; Bac et Cornilleau, 2002, Bac, 2004). The reason is simple, there are no long time series on French National Health Expenditure index comparable to others countries. Moreover, the only studies on French data are characterized by small sample and not using all the appropriate econometric methods to take into account the stationarity and cointegration properties of HCE. By the way, these studies ignore the presence of structural breaks and cross-section dependence.

However, we can draw some lessons from the few published studies. L'Horty *et al.* (1997) work on the 1960-1995 period and try to explain HCE per capita with Error Correction Models (ECM). After controlling price effect and with introducing linear trend (proxy of technology effect), they find that health is a necessity good. They conclude that increases in income would explain 50% of HCE growth since 1960. Missegue and Pereira (2005) and Mahieu (2000) corroborate this result (respectively 0.93 and 0.76). Concerning the HCE determinants, Azizi and Pereira (2005) estimated that between 1970 and 1979 population ageing was annually responsible for 0.82% of spending growth with 0.65 points due to population increase and 0.17 points due to age structure changes. Dormont *et al.* (2006) found, between 1992 and 2000, that demographic changes were responsible of 0.92% annually. Their results show that the impact of changes in practice (including generation and technical progress effects) is 3.8 times higher than demography. As remind of Albouy *et al.* (2009), several studies attempt to estimate an elasticity of the amount of care to the relative price of health. The results suggest an elasticity of -0.6 to -1.0 (Murillo, 1993; L'Horty *et al.*, 1997). The causal interpretation of the volume-price elasticity of care is not trivial. A demand effect (higher prices result in lower demand for care) is arguable. In France, the large insurance cover and the third-party payment mechanism make patients relatively insensitive to the price changes. The existence of a "compensation mechanism" of the supply side cannot be excluded. A decrease of prices can result in an increase in the volume, if the physicians want to achieve an income and induce an extra care demand.

1.3. The French Health Care System

The Public French health care system covers more 99% of the population, without age or other socio-economic conditions. The level of reimbursement of HCE varies significantly depending on expenditure item. More than 90% of spending in hospital care (45% of HCE) are supported by Social Security while ambulatory care (including physician visits) are reimbursed up to around 65%. In addition, 93% of the population has additional coverage. The French health care system is driven by regulatory mechanism; health care volume and price are strongly controlled thereby influencing the rate and structure of growth of HCE. Since Juppé Law in 1996, a National celling for health insurance expenditure is set as a tool of macroeconomic regulation through the Social Security Funding Act. Upstream the number of physician is regulated by a quota (*numerus clausus*) for entry into second year of medical studies. Health prices meet also specific control ways depending on the nature of health care. We can illustrate it with two examples. Visiting physician prices depend on the sector. For 75% of GPs depending on sector 1, visiting prices are fixed (visiting GP costs 23 Euros). In the sector 2, established in 1980 and nearly closed in 1990, GP's fees are free. In addition, the drug prices are set by agreement between Pharmaceutical industry and Economic Committee for Health Products (CEPS in French). They reflect the actual benefit and improvement to the therapeutic arsenal. These specificities can play a role on HCE but also relationship between HCE and GDP. Since 1950 in France, the HCE growth has continuously increased. Over the past 20 years, the pace has however, significantly, decelerated. The share of HCE in GDP, which rose 3.3% in 1955 to 7.1% in 1985, then increased to 8.7% in 2008. Since the early 80's, the increase in HCE was driven by volume, even if "peak" of price increases were recorded (around 3%) in the years 1988-1989, as well as 2002-2003, in line with the increase in payroll-related hospital 35 hours and significant revaluations of medical fees.

Chart 1: Share of HCE in GDP – 1950-2009



Section 2: A linear cointegration analysis of the cointegration HCE/GDP elasticity

At first, we study the elasticity between GDP per capita and HCE per capita on the basis of previous studies in a linear approach.

A simple model is as follows:

$$\ln HCE = \alpha_0 + \alpha_1 \ln GDP + Z_t + u_t \quad (1)$$

Where α_0 is a constant term, α_1 represents the elasticity and Z_t a vector of control variables. This vector varies according to the selected model (see table 4), it can incorporate a linear trend, the population over 65, relative health price and other determinants like medical density, Social security coverage and pharmaceutical research (table 1).

Table 1: Database

Variables	Definition	Availability	Mean	Min	Max
Health Care Expenditure	Medical Care Consumption (deflated € Million – Base 2000)	1950-2009	61 141.47	4 898.76	156 350.10
Gross Domestic Product	Gross Domestic Product (deflated € Million – Base 2000)	1950-2009	779 643.60	220 492	1 423 562
Health Price	Deflated relative health price (Base 100 - 1950)	1950-2005	99.39	88.56	111.93
Population over 65	65+ / total population (%)	1950-2010	13.65 %	11.40 %	16.80 %
Medical Density	Medical density per 100 000 people (Office based practitioners)	1961-2009	148.59	67.50	196.79
Social Security	Share of HCE reimbursed by Social Security (%)	1950-2009	70.89 %	50.98 %	80.00 %
Pharmaceutical research	Global expenditure in pharmaceutical research (deflated € Million - Base 2000)	1965-2007	1 997.35	160.84	4 917.39

Source: Eco-Santé, Institute for Research and Information in Health Economics

2.1 Testing linear cointegration relationship

In a first step, we compute usual unit root tests (Augmented Dickey Fuller, DF-GLS and KPSS) to investigate integration order of GDP and HCE time series. If both series are integrated at the same order, then we have to test the possibility of cointegration between HCE and GDP. We assume to exclude other variables of the cointegration vector (see table 2).

Table 2: Usual Unit root tests

Test	Variables	Lags *k*	Stat	Tabulated value (1%)
ADF	lnHCE	0	-0.34	-4.12
	lnGDP	0	-0.80	
KPSS	lnHCE	-	0.25	0.22
	lnGDP	-	0.24	
DF-GLS	lnHCE	0	-3.75	-3.77
	lnGDP	0	-0.31	

Since both series are clearly integrated I(1) considering the table 2, we test the possibility of a cointegrating relationship using the Johansen procedure (1988, 1991). Testing a cointegrating relationship is equivalent to show that the vector of residuals u_t is stationary.

There are two main cointegration approaches to test and then estimate the long-run model (x): single equation approaches and multivariate VAR approaches. The oldest single equation approach is the Engle and Granger 2 step method (1987) which consists in testing the stationarity of OLS regression and then using OLS to obtain a cointegrating vector (or a long-run estimate). The Johansen (1988, 1991) procedure is a multivariate VAR approach. This method allows to study many cointegrating relationships between the series. In this case, Engle and Granger and Johansen methods are equivalent because only two series may be cointegrated (thus, there is only one potential cointegrating vector).

The Johansen methodology (see table 3 results of the Trace and Eigen value tests) exhibits a clear cointegrating relationship between HCE and GDP per capita with or without supplement trend (except when the trend is quadratic specified). All in all that indicates that there is a clear-cut equilibrium relationship between the two series.

Table 3. Cointegration tests in bivariate system

	Eigenvalue	Hypothesis	Alternative Hypothesis	Statistics (Prob)
Trace test	0.361	r=0	r=1	28.44 (0.00)
	0.050	r=1	r=2	02.91 (0.08)
Eigen Value test	0.361	r=0	r=1	25.54 (0.00)
	0.050	r=1	r=2	02.91 (0.09)

The table 3 gives evidence of the existence of a cointegrating relationship when a linear deterministic trend is considered. The other results of Eigen Value and Trace tests (not reproduced here to save place) lead to reject the null hypothesis of no cointegration except in the case of a quadratic trend .Therefore, in the following models, we do not integrate the quadratic trend (contrary to Pereira and Missegue's study). It is quite possible that this trend reflects the extension of the insurance cover of the French population in the first part of the period (1950-1980) followed by deceleration and then explains a big part of the relationship between GDP and HCE.

2.2 Model Estimation

At a first time, we estimate the model (1) by OLS. Indeed, OLS provide superconsistent estimates when the data seem to support the assumption of a single cointegration vector. However, we have to assume that all regressors are exogenous. An estimation method taking into account the possible endogeneity of the regressors and improving the Engle and Granger single equation approach is thus needed.

Table 4: OLS estimates

Variable	(1)	(2)	(3)	(4)	(5)	(5)bis	(6)
Constant	-6.39	-6.21	-5.44	-5.53	-2.34	5.22	2.20
LnGDP	1.70	1.66	1.53	1.52	1.34	0.56	0.88
LnPrice	-1.28	-1.21	-0.90	-0.93	-1.26	-0.93	-1.16
Linear Trend		0.00	0.01	0.00		0.02	-
Pop over 65			-0.06	-0.03	0.01	-0.01	-0.00
Medical Density				0.00	0.00	0.00	0.00
Social Security					0.39	1.77	2.05
Pharmaceutical research							0.00

We consequently performed the DOLS method proposed by Saikkonen (1991) and Stock and Watson (1993) *via* a dynamic OLS (DOLS) regression. Indeed, in small sample, the DOLS estimator is more precise, as it has a smaller mean squared-error than the MLE (see Stock and Watson, 1993).

Determinants of HCE are analyzed from several explanatory models. Since 1950, relationship between HCE and GDP are obvious. According to the specification used, HCE can appear as a luxury good into models 1, 2, 3, 4 (with an elasticity comprised from 1.5 to 2.2) or a necessary good when we introduce Social security (5, 6). The price elasticity of health spending is significantly negative (from -1.9 to -0.5). The decrease in the relative price of health observed in France is mainly due to the drug prices. It contributes to the growth of health spending in particular through a supply effect (regulation of prices may lead to increase volumes). The period 1960-1980 is characterized by strong growth in HCE that can *a priori* be imputed to demand factors but also supply of care (GPs).

The extension of the insurance cover (extension of social security to farmers in 1961, the self-employed in 1966, voluntary insurance in 1967) significantly contributed to the growth of HCE in the 60's and 70's in making sustainable demand for health care. Models 2, 4 and 5 are used to correct the impact of GDP on HCE by introducing a linear trend, which may explain the socioeconomic conditions and innovation. Indeed, we unfortunately do not benefit from long time series (scanners, mammography) and pharmaceutical research coefficient is significant but its intensity is not strong. Nevertheless, the comparison between models 5 and 6 shows that a large part of trend could be explained by Social security.

More surprising is the negative role of population aging on health spending. The macroeconomic literature on this subject shows that the aging effect is very slightly positive on HCE. Our results do not support these works. It is possible that the age structure incorporates contrasting confounding factors. Age and generation effects contribute to increasing health costs, however a better health at each age allows reducing health costs. In addition, the series is marked by a low-related to the First World War that could disrupt the long-term influence of age structure of health spending. We think that it is possible that this series is not linear and can include a break, which can affect coefficients.

Index of medical density can approach the phenomenon of supply induced demand. Finally, we show that the density plays a positive role on HCE, corroborating this assumption.

Table 5: DOLS estimates

Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant	-6.10	-9.43***	-4.84***	-9.15***	-1.07	-0.26
LnGDP	**1.68*** **	**2.16*** **	**1.53*** **	**1.70*** **	**0.99*** **	**0.88*** **
LnPrice	-1.26***	-1.93***	-1.04***	-0.70*	-0.54***	-0.64
Linear Trend		-0.01***	0.01***	-0.01	0.02***	-
Population over 65			-0.04***	-0.01	-0.03***	-0.02
Medical Density				0.00***	0.00***	0.00***
Social Security					0.76***	2.20***
Pharmaceutical research						0.00***

Note: Lags and leads have been selected using the Akaike criterion.

2.3 An unstable linear relationship?

We want to check the stability of the relationship of the GDP elasticity of HCE. We could indeed expect a time varying relationship between HCE and income per capita: after introducing some major structural reforms or undergoing economic crisis, the households are likely to change their consumption behavior. The regulation policy of the French health care system (level of coverage provided by compulsory and complementary health insurance, increase of out of pocket spending via deductibles) could influence the behavior of access to health care and the HCE trend. Moreover, many factors could play a role on HCE growth (in value and volume) : therapeutic decisions favoring the use of medicines; Per case prospective payment; Increase of Generic drugs; modification of relative prices; intensification of the demand for healthcare (cultural or innovation reasons, ageing population).

We check the stability of the linear cointegrating relationship previously estimated by performing some stability CUSUM tests (see Brown *et al.*, 1975), Ploberger and Kramer (1992) for OLS regressions and Xiao and Philipps (2002) for the case of cointegrating regressions). The CUSUM test is constructed based on sum of recursive residuals (w_t) as follows:

$$w_t = \frac{e_t = y_t - \hat{y}_t}{S_e} = \frac{y_t - \hat{y}_t}{\sqrt{(1 + x_t^{'}(X_{t-1}^{'} X_{t-1})^{-1} x_t)}}$$

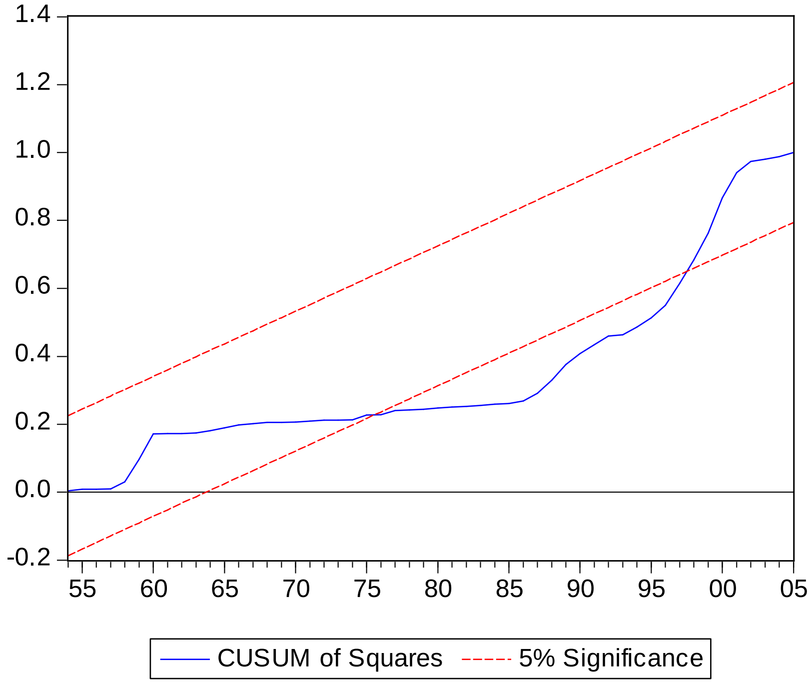
$$t = K+2, K+3, ..., n$$

$$S_t = \frac{n-k}{SCR} \sum_{j=K+2}^{t} w_j$$

$$\left[K, \pm\alpha\sqrt{n-K}\right] et \left[n, \pm 3\alpha\sqrt{n-K}\right]$$

The CUSUM of squares test related to the OLS regression of the model (4) is reported below:

Chart 2: CUSUM of squares



The linear model exhibits clearly non constant coefficients over time and there is evidence of a structural change in the elasticity. The linear model appears clearly unstable, we can probably assume between 1975 and 1995 the existence of a non-linear phenomenon and/or structural change to time. This point is deepened in the both next sections. The instability of the relationship is probably due to non-linearities (either by structural breaks, or by regimes). This relationship seems mispecified because co-integration tests (linear) reject the existence of an equilibrium relationship between HCE and GDP with such a trend. Therefore, OLS or DOLS estimates are completely biased in this case and cause a spurious regression as Yule.

Section 3: Accounting for structural breaks

In this section, we examine the instability of the co-integration relationship outlined in the previous section by considering the possibility of a structural break in the co-integrating relationship or in the deterministic trend. In others words, we consider some form of structural and brutal non linearity in the link between HCE and GDP. Since 1950, several reforms of the French health system were conducted potentially explaining breaks into GDP/HCE relationship. We do not test general form of non linearity at this step because, as raised by Koop and Potter (2001), a linear relationship with breaks could be mistakenly approximated by a non linear model. Controlling break existence is thus a prior to test a more general form of non linearity.

3.1 Unit root tests with endogenous structural breaks

In a first step, we need to reevaluate the properties of our time series by taking account potential structural breaks: are they really non stationary or stationary around some structural breaks?

Since the seminal paper of Perron (1989), it is well known that the usual unit root tests (based on the Dickey Fuller principle) fail to reject the null hypothesis of unit root when structural breaks are not take into account in the procedure. Especially, there is a lack of power of the usual ADF unit root tests when there is a break in the trend. Since structural change may be a characteristic of the HCE dynamics, we examine the stationarity (or the absence of a unit root) of the HCE variable by performing the so called Lee and Strazicich (2003) LM unit root tests. The main advantage of this kind of test is that they are not subject to spurious rejections under the null and allow breaks detection.

In this kind of test, based on the Dickey Fuller principle[3], the break points are detected endogenously form the data *via* a grid search. However, the LM tests from Lee and Strazicich which are based on the seminal work of Schmidt and Phillips (1992) outperforms the others one. Indeed, they are not affected by incorrect placement of breaks in contrast to the other ones: invariance property outlined by Lee and Strazicich (2003). This is very important because the sample finite distribution of the unit root with structural breaks tests is depending on the location of the breaking point. Furthermore, the Lee and Strazicich test is taking account zero, one or two potential breaks in the process of the series.

Moreover, in contrast to other break unit root tests, the Lee and Strazicich (2003) one incorporates structural change both under the null and alternative hypothesis. It is thus not subject to the spurious rejection problem. More precisely, Lee and Strazicich (2003) considered three models in line with the other unit root with breaks tests: a crash model (A), a changing growth (B) and a model allowing both changes (C). In our study, we expect potential two breaks in both level and trend and consequently used the model C. If we observe the both trends (see chart 3 and 4), we can notice a trend changing in HCE between 1975 and 1980 and then in the early 1990s. Regarding the GDP trend, the potential breaks are well known and correspond to the oil crises (early 1970) and the economic crisis of 1993.

[3] See also the tests of Zivot Andrews (1992) and Lumsdaine and Papell (1997). Note that the LM test of Lee and Strazicich (2003) exhibits greater power than the former.

Chart 3: LnHCE – 1950-2009

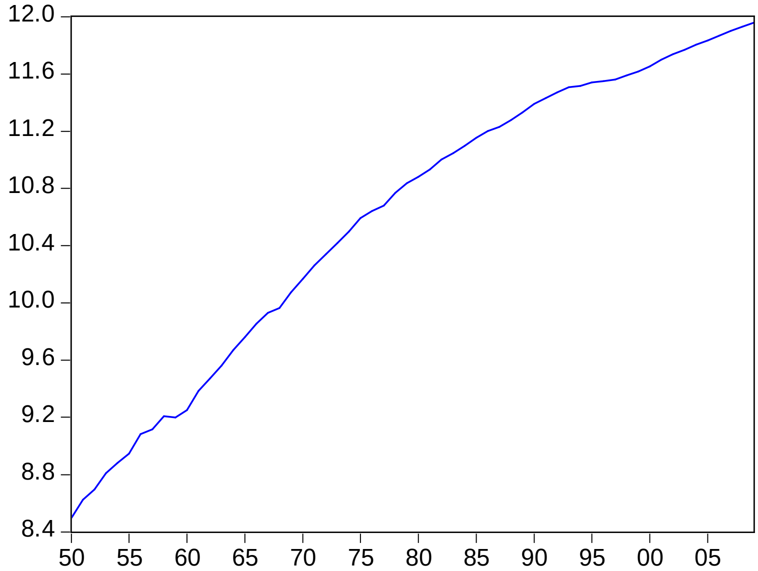
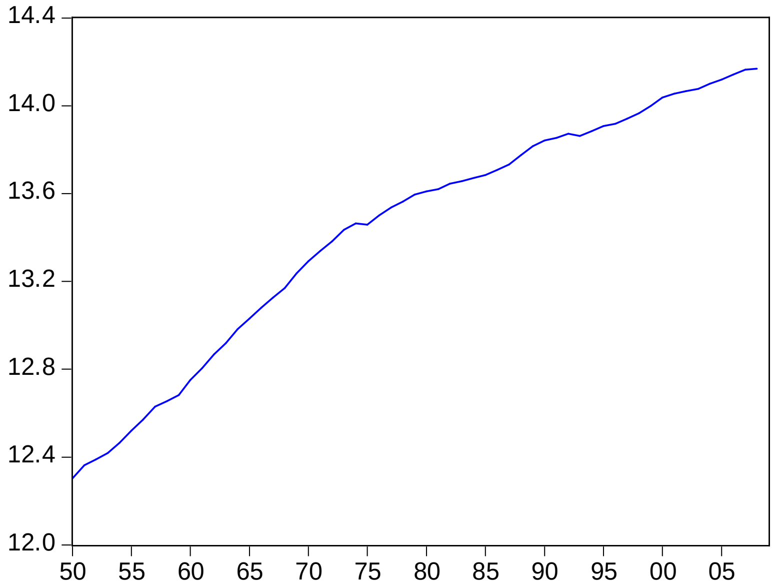


Chart 4: LnGDP – 1950-2009



Following the seminal paper of the authors, the so called LM unit root test is obtained from the following DGP:

$$y_t = d'Z_t + e_t \quad t = 1, \ldots, T \qquad \text{(x)}$$
$$e_t = \beta e_{t-1} + \varepsilon_t$$

where Z_t is a vector of exogenous variables and ε_t is nid.

In the model C with two breaks in level and trend we used here, the vector of exogenous variables Z_t is given as $Z_t = \left[1, t, D_{1t}, D_{2t}, DT_{1t}^*, DT_{2t}^*\right]'$ where $D_{jt} = 1 \text{ for } t \geq T_{Bj}+1, j=1,2, \text{and } 0 \text{ otherwise}$ and where $DT_{jt} = t - T_{Bj} \text{ for } t \geq T_{Bj}+1, j=1,2$ and 0 otherwise. T_{Bj} denotes the time period when the break occurs.

It is important to note that the DGP includes breaks under the null ($\beta = 1$) and under the alternative ($\beta < 1$) hypothesis in contrast to other endogenous break tests as Zivot Andrews (1992) and Lumsdaine and Papell (1997). Indeed, Nunes et al. (1997) and Lee and Strazicich (2001) demonstrate that assuming no break under the null hypothesis causes the test statistic to diverge and lead to significant rejections of the unit root null hypothesis. Thus, we have:

Null hypothesis: $y_t = \mu_0 + d_1 B_{1t} + d_2 B_{2t} + d_3 D_{1t} + d_4 D_{2t} + y_{t-1} + \upsilon_{1t}$

Alternative hypothesis: $y_t = \mu_1 + \gamma_t + d_1 B_{1t} + d_2 B_{2t} + d_3 D_{1t} + d_4 D_{2t} + (1-\alpha) y_{t-1} + \upsilon_{2t}$

Finally, the two breaks LM statistic is generated from the following regression:

$$\Delta y_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-i} + u_t \quad t = 1, \ldots, T \qquad \text{(x)}$$

In the equation (x), $\tilde{S}_t$ is a detrended series defined as $\tilde{S}_t = y_t - \tilde{\psi}_x - Z_t \tilde{\delta},\ t = 2, \ldots, T$ with $\tilde{\psi}_x = y_1 - Z_1 \hat{\delta}$. $\tilde{\delta}$ are coefficients in the regression of Δy_t on ΔZ_t and y_1 and Z_1 are the first observations of y_t and Z_t. Usually, testing the unit root is equivalent to test $\phi = 0$ in equation (X):

$$\tilde{\rho} = T\tilde{\phi}$$
$$LM_\tau = \inf_\lambda \tilde{\tau}(\lambda)$$

In addition, note that lagged terms may be included to correct serial autocorrelation as in the usual Augmented Dickey Fuller. We use this specification and the equation (x) is then rewritten as:

$$\Delta y_t = \delta' \Delta Z_t + \phi \tilde{S}_{t-i} + \sum_{i=1}^{p} \gamma_i \Delta \tilde{S}_{t-i} + u_t \quad t = 1, \ldots, T \qquad \text{(x)}$$

To endogenously determine the location of breaks, the following grid search is used :

$$LM_\rho = \inf_\lambda \tilde{\rho}(\lambda)$$
$$LM_\tau = \inf_\lambda \overline{\tau}(\lambda)$$

The table 6 displays the Lee and Strazicich unit root tests. Allowing for breaks in the unit root tests provide significant evidence in favor of segmented trend stationarity for the lnGDP serie. We choose to introduce two structural breaks. Concerning to both series, the test of Lee and Strazicich shows the non-stationary (for lnHCE : -5.82 and for lnGDP -4,75 and – 5,54 according to lag). In HCE serie, two significant breaks are identified: a break level in 1976 and a trend break in 1994 (both are negative and show a decrease in HCE growth). 1976 corresponds to a drastic austerity budget by Raymond Barre government in 1976 which seeks to limit HCE and 1994 marks an important break in the pace of HCE growth in France due to economic recession in 1993 with a negative GDP of -0.9%. Concerning to lnGDP serie, we test one and six lags (better properties to control autocorrelation) and we identify 1968 break (date of main social crisis in France) and the first oil crisis. These results corroborate the non-stationary of series when taking into account two breaks.

Table 6: Lee and Strazicich Unit root tests with two structural breaks

Variables	k	St-1	Stat	B1	B2	DT1	DT2	$T\hat{B}$
lnHCE	5	-0,80	- 5.82	-0.05 (-2.35)	0.02 (1.14)	0.00 (0.33)	-0.07 (-6.51)	1976, 1994
lnGDP	1	-0.49	-4.75	-0.01 (-0.92)	0.04 (2.99)	0.02 (2.22)	-0.03 (-5.33)	1967, 1975
	6	-0.85	- 5,54	0.01 (0.68)	0.02 (1.48)	0.02 (2.60)	-0.03 (-6.41)	1968, 1981

Lecture: k is the number of lags to correct autocorrelation concerns and $T\hat{B}$ consists in estimated structural breaks. The significance of the LM test statistic is realized via comparison with the critical values of the table 3 of Lee and Strazicich (2003) and the significance of the derived breaks points is established using t-statistic at the 5% level of significance.

3.2 Cointegration test with endogenous structural breaks

Since we outlined some breaks in our univariate time series, we also test the existence of possible breaks in the co-integrating relationship that is in the GDP/HCE elasticity. We use the Carrion-i-silvestre and Sanso (2006) co-integration test[4]. We thus test for the null hypothesis of co-integration against the alternative of no co-integration in the presence of a potential break under both hypotheses. The test is derived for a known and an unknown break with exogenous or endogenous regressors. This Lagrange Multiplier test is very interesting to challenge linear results leading to reject the co-integration and to outline some breaks in the co-integrating relationship. The main advantage of this test over the other in the literature (see for instance Johansen *et al.* (2000)) is that it allows breaks in the slope parameters of the co-integrating vector and not only in the constant and time trend terms. Under the null, the co-integrating vector may shift to one long run regime into another one.

The model considered by the authors is a multivariate extension of Kwiatkowski *et al.* (1992) where deterministic and/or stochastic trends might change at a point of time (TB). The data generating process considered is of the following form[5]:

$$y_t = \alpha_t + \xi_t + x_t^{'}\beta_1 + \varepsilon_t \quad (1)$$
$$x_t = x_{t-1} + \varsigma_t \quad (2)$$
$$\alpha_t = f(t) + \alpha_{t-1} + \eta_t \quad (3)$$

$\eta_t \sim iid(0,\sigma_\eta^2)$, x is vector of k I(1) process (lnHCE and lnGDP here), α_t is a constant and $f(t)$, the heart of the model, is a function collecting deterministic and/or stochastic components.

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[4] We used the gauss code kindly provided by Carrion-i-Silvestre.
[5] This part draws heavily on the work of Carrion-i-Silvestre and Sanso (2006).

The definition of the function f(t) leads to the consideration of six different specifications (An to E). In the way of Perron (1989, 1990), the models An to C affect the deterministic trend component:

$$\begin{aligned}
&An : \xi = 0,\ f(t) = \theta D(T_b)_t \\
&A : \xi \neq 0,\ f(t) = \theta D(T_b)_t \\
&B : \xi \neq 0,\ f(t) = \gamma DU_t \\
&C : \xi \neq 0,\ f(t) = \theta D(T_b)_t + \gamma DU_t
\end{aligned}$$

where $D(T_b)_t = 1 \text{ for } t = T_b + 1 \text{ and } 0 \text{ otherwise}$, $DU_t = 1 \text{ for } t > T_b \text{ and } 0 \text{ otherwise}$ with $T_b = \lambda T$ indicating the estimated date of break ($0 < \lambda < 1$).

This test is based on the KPSS stationarity test (1992) for which under the null hypothesis, the variance of the autoregressive process (equation (3)) is null: $\sigma_\eta^2 = 0$. Under the alternative $\sigma_\eta^2 > 0$. Consequently, under the null hypothesis, the model given by (1), (2) and (3) can be rewritten as:

$$\begin{aligned}
&y_t = g_i(t) + x_t^{'}\beta + \varepsilon_t \ (4) \\
&\text{with } i = \{A, An, B, C\}
\end{aligned}$$

These four models An to C account for structural breaks in the long run deterministic relationship but the co-integrating vector remains unchanged. In the model An, we consider a level shift without time trend, in the model A, we consider a trend and a break in level, the model B captures a change in the slope of the time trend but not in the level and finally, the model C captures level and slope shifts. In the D and E models, the specification allows for a structural break that not only shifts the deterministic component but also changes the co-integrating vector. Thus, in some situations, practitioners would be interested in modelling a co-integration relationship that at a point in time might has shifted from one long-run path to another one (see Carrion and Sanso, 2006). Consequently, a dummy now affects the co-integrating vector (indeed see that $x_t^{'}\beta$ is affected by $DU(t)$ in the model (5)). The differences between D and E are that the model D does not consider any time trend.

$$\begin{aligned}
&D : \xi = 0,\ f(t) = \theta D(T_b)_t + x_t^{'}\beta_2 D(T_b)_t \\
&E : \xi \neq 0,\ f(t) = \theta D(T_b)_t + \gamma DU_t + x_t^{'}\beta_2 D(T_b)_t
\end{aligned}$$

$$\begin{aligned}
&y_t = g_i(t) + x_t^{'}\beta_1 + x_t^{'}\beta_2 DU(t) + \varepsilon_t \ (5) \\
&i = \{D, E\} \\
&g_D(t) = \alpha + \theta DU(t) \\
&g_E(t) = \alpha + \theta DU(t) + \gamma DT^*(t)
\end{aligned}$$

Carrion-i-Silvestre and Sanso (2006) distinguish between models with strictly exogenous regressors and models with non strictly exogenous regressors. In the second case, which is our preferred case, the asymptotic theory does no longer hold. Indeed, the estimation of the vector of cointegration is not efficient with endogenous regressors and Dynamic OLS estimators (see Saikkonen (1991) and Stock and Watson (1993) are needed. Thus, OLS estimation of (4) and (5) are substituted by DOLS estimation of the following equations (note the leads and lages terms in difference in both equations in line with the Stock and Watson (1993)):

$$y_t = g_i(t) + x_t^{'}\beta + \sum_{j=-k}^{k} \Delta x_t^{'}\gamma_j + \varepsilon_t \text{ (6) if } i = \{An, A, B, C\}$$

$$y_t = g_i(t) + x_t^{'}\beta_1 + x_t^{'}\beta_2 DU_t + \sum_{j=-k}^{k} \Delta x_t^{'}\gamma_j + \varepsilon_t \text{ (7) if } i = \{D, E\}$$

After getting the estimated residuals denoted as $\hat{e}_{i,t}$, the test statistic is computed as:

$$SC_i^{+}(\lambda) = T^{-2}\hat{w}_1^{-2}\sum_{t=1}^{T}(S_{i,t}^{+})^2$$

where $\hat{w}_1^2$ is a consistent estimation of the long run variance of ε_t using $\hat{e}_{it}$ and $S_{i,t}^{+} = \sum_{j=1}^{t}\hat{e}_{i,t}$.

Finally, the date break is estimated by minimizing the sequence of squared errors or equivalently by minimizing the sequence of the Bayesian Information Criterion (BIC) obtained when computing the test for all possible structural breaks (see Carrion-i-Silvestre and Sanso for more details):

$$\hat{T}_b = \arg\min_{\lambda \in \wedge}\left[SSR(\lambda)\right]$$

where $SSR(\lambda)$ denotes the sum of squared residuals and λ is a closed subset of the interval (0,1) redefined as $\left[2/T, (T-1)/T\right]$ to minimise the lost of information.

The table 7 outlined the results obtained by performing the Carrion and Sanso's test for model An to E. The results of the model An are not very interesting in our case since the absence of time trend but we give its results for comparison purposes. C and E models are our preferred models.

The results are very similar in the B and C models leading to the conclusion that the null of cointegration with a break in 1989 in the slope of the time trend cannot be rejected. Very similar results are also derived when we based the test on the model D and E although the break date is now 1985 instead 1989. D and E results are particularly of interest because structural breaks in the deterministic part and also in the co-integrating vector are allowed. The results of the table 7 show that the co-integrating vector has shifted in 1985, we can see this rupture on the chart 1. Whatever the specification of the test we used, the break dates are located at the middle and at the end of the 1980's. The beginning of the 80s was marked by a measure likely to increase health spending, the creation of sector 2 for GP (fees are free). Between 1980 and 1985, the difference in growth rates between GDP and health spending is 4 points (HCE: +5.5% GDP: +1.5%). Given this weak economic growth, the 80' starting 1982 are marked by numerous measures to control HCE, especially among hospital expenditures (main responsible for HCE). Thus, the annual rate of hospital spending growth declined from 18% in 1982 to 7% in 1985. Bérégovoy Plan (1982) initiates these measures with the creation of hospital deductibles, increase of out of pocket for certain drugs and no revaluation of sickness benefits. Plan Dufoix act (1985) reduced the reimbursement of 379 drugs. In 1987, Seguin Plan decided the extension and revision of the list of 25 diseases that can benefit from an exemption from out of pocket payment and the increase in hospital deductibles.

Table 7: Carrion and Sanso co-integration test

Model	Break test	Break date
An	0.051 0.061	1987 1984
A	0.047 0.048	1985 1984
B	0.043 0.043	1989 1989
C	0.037 0.037	1989 1989
D	0.045 0.045	1985 1985
E	0.042 0.042	1985 1985

Note : Null hypothesis : cointegration with break at unkown time. Endogenous regressors are considered (weak exogenous results are unavailable upon request). Critical values are from the table 3 of Carrion-i-Silvestre and Sansó's (2006).

To illustrate the existence of a break in the middle of 80's, we compute again the DOLS regressions outlined in the previous section by adding a dummy in 1985. Whatever the specification used, the dummy is always strongly significant corroborating our assumption. By comparing models 6 (table 5 *versus* table 8) with and without dummy 1985, we show that the introduction of dummy seems determinant on Social Security effect. We think that the role of Social security on HCE was very strong only during the first period 1950-1985 (see appendix).

Table 8: DOLS estimates with dummy in 1985

Variable	(1)	(2)	(3)	(4)	(5)	(6)
Constant	-10.58	-7.67***	-13.02***	-13.14***	-1.97	-9.74***
LnGDP	**1.60*****	**1.66***	**2.22***	**1.98***	**1.03***	**1.56***
LnPrice		-0.87***	-1.29***	-0.62	-0.46***	0.11
Linear Trend			-0.02***	-0.01***	0.01***	-
Pop over 65			0.91	0.59	-0.03***	-3.99***
Medical Density				0.00***	0.00***	0.00***
Social Security					0.73***	-0.79
Pharmaceutical research						0.00*
Dummy 1985	0.08***	0.05***	0.07***	0.04***	0.01***	0.04***

Note: Lags and leads have been selected using the Akaike criterion.

We can conclude that a long-term equilibrium exists between GDP and HCE with a co-integration relationship. We highlight a structural break into the cointegration vector, linked to concentration of health policies at the middle of 80's. Nevertheless, after 1985, nature of cointegration changed, we have now to identify the paths of changing which can be linked to specific variable as induced demand, innovation, ageing population or Social security. In the next section, we try to analyse the possibility of non linearity model.

Section 4. Structural breaks or smooth non linearity?

In the previous section, we tested linear co-integration and co-integration with breaks ie the hypothesis that the link between GDP and HCE might be instable due to the existence of a break. However, a dynamics with structural break implies durable and "abrupt" changes without possible way-back. The true relationship can transit from one regime to another in both directions and might be completely non linear.

4.1 Testing for non linearity

There is a recent literature dealing with non linearity econometric models. One major direction focused on modelling and testing non linear adjustment in deviations from linear long run equilibrium: see Balke and Fomby (1997), Hansen and Seo (2002)… The other direction consider that the equilibrium relationship itself might be non linear. In other words, equilibrium among our interest variables is depending on the state of the health system as represented by one and several transition variables. This approach is more convincing in this context.

We test the possibility that the relationship linking GDP and HCE (both variables being I(1)) undergoes regimes switches. Indeed, if the assumption of linearity is invalid, a re-examination of the HE elasticity is needed. To this aim, the Choi and Saikkonen (2004) cointegration smooth transition regression is assumed (CSTR). The major interest of this approach is to identify the transition or threshold variable to capture the non linearity of the long run relationship between consumption and expenditures. The general methodology consists in identifying a transition value of an explanatory variable (exogeneous to the model or lagged endogeneous) to deal with the dependence of the parameters (here elasticities coefficients) to the dynamics of the "health environment". Consequently, the long-run equilibrium relationship might change smoothly depending of the transition (or threshold) variable that is depending on where the covariates x_t are located relative to the threshold parameter *c*.

Following to the review of literature (section 1), we consider four candidates:

1. ***Population over 65***. We know the positive link between age and health status. But the role of ageing on the GDP/HCE couple is more complex to appreciate. The increase in the elderly causes a deep epidemiological transformation. New neuro-psychiatric disorders appear requiring an appropriate care supply. This variable may introduce changes in the relationship GDP and HCE even if the acceleration of the aging population does that from 2005 with a strong potential impact on spending.

2. ***Medical density***. This indicator of Health Care supply can be a good proxy of demand induced. Indeed, an increased physician density lowers the average rate of personal profit, and may lead the physician to create an unnecessary demand.

3. ***Pharmaceutical research.*** Technological progress corresponds to medical innovations that affect both small appliances (hearing aids, for example) that the heavy equipment (scanners,

Medical imagin). It also refers to the development of medical techniques to improve the quality of life of chronic patients (including widespread use of home dialysis, transplants...).

4. ***Social Security.*** As shown the extension of coverage is able to explain a large part of HCE.

Following the approach recently developed by Choi and Saikkonnen (2004), we test linearity against non linearity of the STR form. The non linear model is given by[6]:

$$\ln HCE_t - \ln GDP_t = \beta_1^{'} x_t + \beta_2^{'} x_t g(z_t, \gamma, c) + \sum_{j=-K}^{K} \pi_j^{'} \Delta x_{t-j} + u_t \qquad (8)$$
$$t = K+1, \ldots, T-K$$

$$\text{with } g = 1 + \exp\left[-\gamma(z_t - c)\right]^{-1}$$

where the function g is a logistic function bounded between 0 to 1, c is a threshold (or location) parameter and γ denotes the smoothness i.e. the slope of the change. The last term of the equation (8) allows to resolve the serial correlation between regressors and errors terms by adding K leads and lags.

Thus, in line with Choi and Saikkonen (2004), we test for linearity in the equation (8) by assuming the null hypothesis: $H_0 : \gamma = 0$ or $\beta_2 = 0$. However, conventional hypothesis testing is complicated because the cointegrating STR model contains unidentified nuisance parameters under the null that is the transition value c and the slope parameter γ. Hence, a possible solution is to employ the first-order (T1) and the third-order (T2) in order to replace the transition function g. The Choi and Saikkonen statistic follows a Chi Square distribution under the null with p degrees of freedom where p is the number of covariates related to the transition function.

The results of the LM linearity tests are illustrated in table 9. Two specifications (denoted by (1) and (2)) are distinguished: in the first one, only the cointegrating relationship between HCE and GDP is considered; in the second one, we include the log of the relative price in the long-run relationship (i.e. in the vector of cointegration[7]) to check the robustness of our results. As outlined by the table 9, we can conclude that the evidence supports non linearity when the density is considered but supports linearity when other transition variables are considered. The results concerning Social Security are less clear-cut: the non linearity is supported only when 3 lags and leads are included in the DOLS. All in all, our results are in addition robust to the specification of the long-run relationship.

[6] Five econometric restrictions are needed for the transition variable *g*, see Choi and Saikkonen (2004) for more details.

[7] Therefore, given P_t the log of the relative price, the equation (8) can be rewritten as

$$\ln HCE_t - \ln GDP - \ln P_t = \beta_1^{'} x_t + \beta_2^{'} x_t g(z_t, \gamma, c) + \sum_{j=-K}^{K} \pi_j^{'} \Delta x_{t-j} + u_t.$$

Table 9: LM Linearity tests from Choi and Saikkonen

Transition variable	T1			T2		
	K=1	K=2	K=3	K=1	K=2	K=3
Past Health expenditures	1.20	1.32	1.37	2.17	2.26	2.52
Population over 65	0.03	0.03	0.05	0.72	0.83	1.04
Medical density (1)	5.75**	6.08**	6.66***	6.41**	6.42**	7.07**
Medical density (2)	6.21**	6.48**	10.19***	9.30**	12.01***	15.28***
Pharmaceutical research (1)	0.24	0.39	0.48	3.41	4.57	5.33*
Pharmaceutical research (2)	1.97	3.44	5.05*	5.50	5.07	5.75
Social Security (1)	1.17	1.52	1.83	2.32	1.87	1.83
Social Security (2)	0.23	1.70	5.05*	0.24	2.85	9.80**

Notes: K denotes the leads and lags in the auxiliary regression model. T1 (first order expansion) and T2 (second order expansion) are distributed as asymptotic Chi2 statistic under the null with one and two degrees of freedom respectively (specification (1)) and with two and three degrees of freedom respectively (specification (2)). ***, ** and *: significant at 10%, 5% and 1% level respectively.

4.2 Estimating CSTR model (under construction)

Provisional conclusion

In this preliminary study**,** in a first part (section 2), we conduct a linear analysis after checking the cointegration properties. Results seem robust comparative to other French studies. Health care are considered as a luxury good except after introducing social security coverage and Price effect is negative on HCE. Nevertheless, we show that this first analysis is not perfect cause of the relationships' instability.

Then, we reexamine the bound between GDP and HCE by taking into account structural breaks and possibility of non linear phenomena. We show (section 3) that a structural break changed the cointegration nature at the middle of 80'. We also verify the stability of general elasticity by testing the possibility of nonlinear dynamic due to changes in certain variables.

We can assume that between 1950 and 1985, health expenditures were explained by the increasing of health coverage. However from the mid 80's, the pace of growth in health spending has slowed due to an exogenous shock related to regulation of health expenditures (Bérégovoy Plan, Dufoix Plan). Other variables, such as medical density, should explain the new dynamic GDP/HCE.

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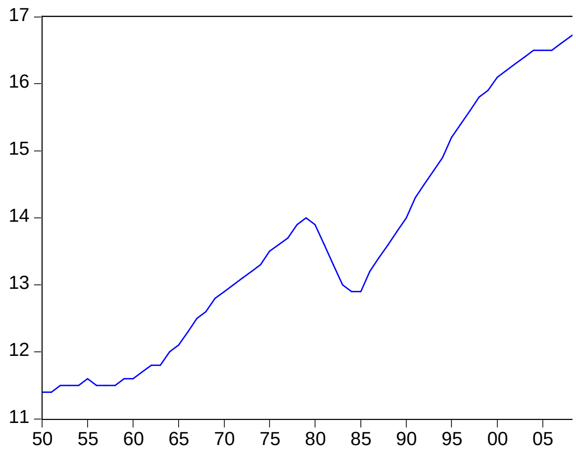
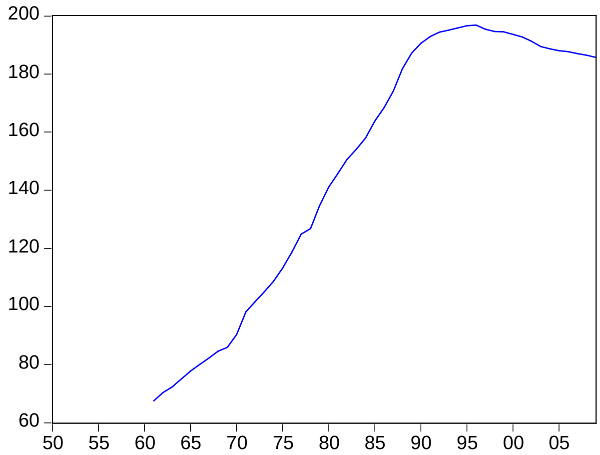
APPENDIX

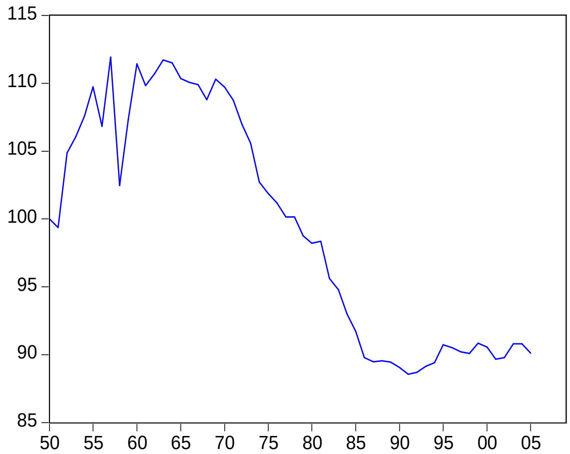
POPULATION OVER 65 (%) / 1950-2009



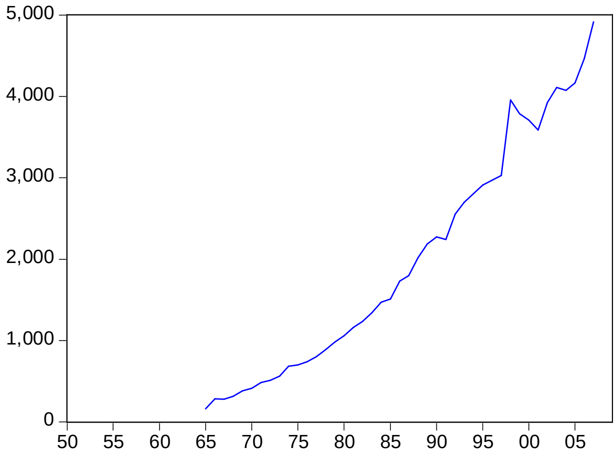
MEDICAL DENSITY / 1961-2009



RELATIVE PRICE OF HEALTH / 1950-2005



PHARMACEUTICAL RESEARCH / 1965-2007



SOCIAL SECURITY (%) / 1950-2009