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Does Financial Crisis Really Affect French Economic Growth? Fresh Insights from Nonlinear Approach

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Abstract ;

In this paper, we attempt to analyze the financial crisis impact on the French economic growth in France. For this end, we use the industrial production index, the consumer price index, the money market rate and the transactions volume during the period from 2000 to 2015. We also employ to nonlinear approach to study such effect and identify the existence of a nonlinear long-term relationship between variables. The empirical results clearly show that the linear and non-linear coefficients of the shifted residuals for a single period have negative and positive signs. Therefore, French economic growth has a non-linear dynamic towards its fundamental value.

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Introduction

Over the past two decades, the global economy has experienced major financial crises (e.g. monetary crisis, banking crisis). As a result, and in most cases, such dramatic phenomena result in substantial adverse economic and social effects (contraction of real activity, bankruptcy of some companies, unemployment increase and poverty aggravation). For instance, the advent of crises in countries of South America and South-East Asia during the last half of the 1990s led to spectacular falls in monetary and stock market values.

With the occurrence of the 2007-2009financial crisis, deep recessions are considerably felt by many developed countries and reflected by drops in employment rates and sharp rises in budget deficits (Bozio et al., 2015). Such crisis has particularly struck the European Union at a time when it was experiencing a period of prolonged growth (Matei and Calapod, 2015). Indeed, in 2006, the growth of the Eurozone economy was characterized by an increase of 2.7% and the employment rate growth was 1.5% with an increase of 2 millions of jobs. As a matter of fact, in December 2006, the unemployment rate recordeda significant decrease, with the lowest level in the last 15 years (7.5%). From academic standpoint, several researchers have notably focused on studying the impact of financial crisis on the stock markets in the developed and developing countries (e.g. Wong, 2014; Collingio and Frenkel, 2019). For instance, Bordo (2008) showed that many stock markets have adversely influenced by the 2007-2009 financial crisis. Aloui et al. (2011) analyzed the financial crisis based on a multivariate copula approach. More specially, they examine the interdependence of four emerging markets (China, Russia, India and Brazil) with the US market over the period 2004-2009. They find substantial dynamic dependence between the US and emerging markets. Wong (2014) analyzed the interaction between the US and seven East Asian stock markets during the financial crisis. The empirical results show that such crisis has strengthened the linkages between the US and East Asian stock markets. More recently, Collingio and Frenkel (2019) examine the response of national and Euro-wide equity indices as well as of government bond yields. They display that financial market participants tend to react more highly to monetary and policy after crisis. They also find that cross-country differences in terms of responsiveness of government bond yields seem to correlate with average national unemployment rates and inflation rates. Nikkinen et al. (2019) test for the effect of the financial crisis on internal linkages within the European frontier stock markets (Croatia, Estonia, Romania, Slovakia and Slovenia) and external linkages between such markets and developed stock markets (US, UK and Germany). They prove that both long- and short-run external relationships were reinforced during the crisis. However, internal linkages seem to be weak, expect for Croatia and Slovenia.

Other researchers have instead focused on the effect of the financial crisis on the economic sector given losses recorded by financial markets and main economic aggregates such as GDP growth and investment (e.g. Reinhart and Rogoff, 2009; Cerra and Saxena, 2005; Ksantinia and Boujelbene, 2014). For example, Cecchetti et al. (2009) reported that only 20% of crisis has a permanent effect on the level of GDP. Yuan et al. (2010) analyzed the impact of financial crisis on energy on consumption and economic growth in China. They clearly display that the decrease of exports produced by such crisis implies a decrease of 7.33% and 9.21% in GDP and energy consumption, respectively. Operaa and Bilan (2015) point out the effects of the crisis on the revenues, expenditures, budget balance and debt in Romania. Makin (2019) assesses the G20's macroeconomic policy response to the crisis and its public debt legacy for advanced economies.

This study lies to the aforementioned literature and attempts to investigate the impact of the financial crisis on the GDP growth and investment in France over the period 2000-2005. Although several researchers point out the significant impact of such crisis on economic sector, there still exists the question whether the financial crisis temporarily or permanently affects the macro-financial indicators. The size and duration of these adverse effects significantly varied among countries (Bozio et al., 2015). So, this study offers new insights into this issue from the French case. For this end, we use nonlinear approach which allows us to assess not only the impact of the financial crisis but also the short- and long-term dynamics after crisis. Such approach therefore allows to use the classic method based on the dummy variable (0,1) to better apprehend the effects of financial crisis.

This paper is organized as follows. Section 2 presents an overview of studies which examine the effects of financial crisis on the economic growth. Section 3 reports the methodology design and Section 4 reports data and the descriptive statistics. The empirical results are presented in Section 5. Section 6 concludes.

2. *Some Insights from Literature Review on the Financial Crisis-Economic Growth Nexus*

Interestingly enough, the worldwide financial system went through a crisis of unprecedented depth and magnitude since the year 2008. As the crisis deepened and spread, it increasingly affected different markets, actors and countries and led, in some cases,to economic crisis. In particular, the extent of such economic crisis clearly reflects not only the soundness of the government's emergency measures but also the need for further strengthening the framework for better financial stability. In this regard, many governments and central banks have adopted a series of economic policy measures which aim for supporting the banking sector, restoring the well-functioning of markets and curbing the spread and adverse effects of the crisis on the real economy.

The spread of the adverse effects of the financial crisis to the country's economy seems to occur particularly because of the lack of spontaneous mechanisms of regulation, in a snowball effect. As reported in the balance sheet of the OFCE[1], the financial crisis leads to a brutal macroeconomic slowdown. This can be reflected in different ways such as the depreciation of real estate as sets, the increase of corporate defaults (+125% in the United Kingdom, +20% in France) and unemployment via bankruptcy filing. From a statistical standpoint, some indicators such as the industrial production index and unemployment rate support such findings. In terms of variation over one year, the industrial production index decreased by about -18% for Spain, -12% for Germany, -11% for France, -9% for the United Kingdom and Italy and -8% for the United States. As well, based on variation over one year, the unemployment increased by about +5.7 points in Spain (at a level of 14.4%), +2.3 points in the United States (7.6%), +1 point in the United Kingdom (6.1%), +0.4 points for France (8.1%).Most notably, as a whole, it is possible to identify several theoretical mechanisms by which crises affect investment and economic growth. In economic theory, there are two types of crises that affect the real economy either by global demand or supply. The monetary crisis influences the speculative selling situation of the local currency. Such impactled to a significant loss of foreign exchange reserves, higher interest rates and a devaluation of the central currency. Also, the currency crisis weakens global demand and supply by increasing imports and investment costs as well as external debt services. The depreciation of domestic currency and the rise in interest rates are forcing companies to go bankrupt.

From academic standpoint, a very few studies have investigated the effect of the financial crises on the country's economic growth. Aghion et al. (2001) developed a monetary crises model in which the equilibrium corresponding to a crisis situation accompanied by a lower production level. They found that the devaluation led to deterioration in the company's balance sheet and a reduction in borrowing and investment capacity for these companies. Moreover, if one might consider that crises are shocks likely to affect the real exchange rate, it is possible that the latter deviated from its equilibrium level. According to Dollar (1992), there is an optimal exchange rate corresponding to an optimal level of resource allocation.

Hence, any shock that has the effect of moving the real exchange rate away from its optimal level leads to resources misallocation and suboptimal growth. But, the monetary crisis brought about by the devaluation can be beneficial in the long term because it can promote exports and investment and economic growth through external competitiveness improvement. Barro (2001) examined the immediate and long-term effects of the 1997-98 crisis on economic growth performance in East Asia. More explicitly, the author examines the behavior of investment and economic growth of ten countries by diving them into two groups based on the extent to which they were affected by the financial crisis of 1997-98.The first group which is characterized by a strong devaluation of their currency includes South Korea, Indonesia, Malaysia, Philippines and Thailand. The second class, however, comprises countries with a low devaluation of their domestic currency such as China, Hong Kong, Japan, Singapore and Taiwan. The empirical results clearly show that the monetary crisis is associated with a loss of 1.3% in the real GDP growth rate and 0.4% in the investment rate. Therefore, Barro (2001) reports that the banking crisis reduces the per capita economic growth rate to around 0.6% per year and the investment rate in the neighborhood of 0.9%.

Bordo et al. (2001) analyzed financial crises frequency and their effects on economic growth. Using a sample of industrialized and emerging countries over the period 1880-1997, they distinguished between the periods 1880-1913 (gold standard period), 1919-1939 (interwar period), 1945-1971 (Bretton Woods regime) and 1937-1997 (post-Bretton Woods). In general, crises have been followed by downturns lasting on average 2 to 3 years and costing 5 to 10 per cent of GDP. They also reveal that even if one might controls some features of cycles, recessions with crises seem to be more severe than recessions without them. Such evidence clearly indicates that such pattern reflects causality running from crises to recessions, at least in part, and not simply from recessions to crises. As well, Bordo et al.(2001)report that crises frequency has doubled during the modern period (post-period 1973) and financial globalization has not increased crises severity. Based on a sample of eight Latin American countries (Argentina, Brazil, Chile, Colombia, Ecuador, Mexico, Peru and Venezuela), six Southeast Asian countries (Indonesia, South Korea, Malaysia, Philippines, Singapore and Thailand) and three African countries (South Africa and Egypt), Barro (2007) estimated crises impact using classical growth and investment equations. In such equations, an endogenous variable is the real GDP per capita growth expressed in purchasing power parity while explanatory variables are convergence variable approximated by the lagging real GDP per capita, human capital, physical capital accumulation, human capital, openness rate, banking crisis indicator and twin crisis indicator. The empirical results showed the monetary crisis reduces economic growth rate to around 2% per year and this reduction is persistent for two years in a row. Barro (2007) concluded that the monetary crisis is more robust than the banking crisis which is in the order of 1.5% per year. The crisis has a greater negative impact on monetary and banking crises.Furceri and Mourougane (2009) assess the impact of financial crisis on GDP. They show that, on average, GDP diminished by 1.5% to 2.4% following a crisis and continues over five years. Ksantinia and Boujelbène (2014) propose to analyze the influence of financial crises on GDP growth and investment using twenty-five countries (Argentina, Australia, Austria, Belgium, Brazil, Canada, Chile, China, France, Germany, Hong Kong, India, Indonesia, Italy, Japan, South Korea, Malaysia, Mexico, Netherlands, Singapore, Spain, Switzerland, Thailand, United Kingdom and United States) over the period 1998-2009. Based on a dynamic panel model, they indicate the financial crisis affects significantly and negatively the GDP growth of a country and its level of investment.

3.*Methodology*

In this paper, we use the cointegration theory to examine the impact of financial crisis on economic growth. For this end, we attempt to estimate a model that relates the industrial production index (IPI) with consumer price index (CPI), money market rate (MMR) and stock transaction volume (VT). This model is specified by the following nonlinear model:

$$IPI_t = A\left(CPI_t^{\beta}\right)\exp(MMR_t^{\alpha})(VT_t^{\delta})\exp(\varepsilon_t) \qquad \forall\, 2000:01 \rightarrow 20015:12 \qquad (1)$$

The non-explanatory variables are nested in the error term and the constant (A) is the average effect of the omitted variables. In order to estimate this model, it is necessary to introduce the log-log specification to linearize this model as follows:

$$Log(IPI_t) = Log(A) + \beta Log\left(CPI_t\right) + \alpha(MMR_t) + \delta Log(VT_t) + \varepsilon_t \quad \forall\, 2000:01 \rightarrow 2015:12 \quad (2)$$

In this model, the variables are integrated of order one based the Dickey-Fuller stationarity test (1979-1981). To do this, we use Engle and Granger (1887) double-step method to detect the linear adjustment of French production with the financial crisis.

4. *Data and Descriptive Statistics:*

The data set includes the industrial production index (IPI), the money market rate (MMR), the consumer price index (CPI) and the transaction volume (VT) of the French financial market over the period from 2000 to 2015 with monthly frequencies, a total of 192 observations. Such dataset was provided from the World Bank, the International Monetary Fund, the Paris Stock Exchange and the French Central Bank. The descriptive statistics of the data used in this study is reported in Table 1.

Table 1. Descriptive Statistics of Different Variables

	IPI	CPI	MMR	VT
Mean	4.5517	4.6094	1.8523	6.6751
Median	4.6071	4.61234	1.7766	6.7854
Maximum	4.9111	4.8790	2.3804	7.7481
Minimum	4.0550	4.3226	1.6094	5.2937
Std.Dev	0.2203	0.1468	0.2357	0.5958
Skewness	-0.5001	-0.1382	0.7563	-0.3675
Kurtosis	2.1998	2.1416	2.3265	2.6371
Jarque-Bera	13.1283	6.5058	21.932	5.3757
Probability	0.0014	0.0386	0.0000	0.0680
Observations	192	192	192	192

Source: Prepared by researchers, 2021**.**

From Table 1, the analysis of descriptive statistics reports an information asymmetry of variables under study over the period 2000-2015 on monthly frequency. This asymmetry is detected by the kurtosis coefficient. The Jarque-Bera statistics are significant even at very low levels. Thus, the monthly variables are not normally distributed. Standard deviations are very low for these variables. Hence, it is a good linear fit of each variable relative to its average. Also, the accuracy of these variables is very good because the variance of each variable is very small.To study stationarity, wecomputeneperian logarithm of the industrial production index (LIPI), the logarithm of the consumer price index (CPI), the money market rate (MMR) and the logarithm of transaction volumes of the Paris Stock Exchange. Table 2reports results from both Dickey-Fuller and Phillips-Perron tests.

Table 2. Results from Stationarity Tests

Results from Stationarity Tests

	Dickey-Fuller test (1979-1981)						**Phillips-Perron test (1989)**					
	Lags	Models	In Level		In first differences		Lags	Models	In Level		In first differences	
Variables			T-Statistics	Critical Values	T-Statistics	Critical Values			T-Statistics	Critical Values	T-Statistics	Critical Values
LIPI	3	M3	-2.9201	-3.4337	-11.2910	-3.4339	3	M3	-6.8959	-3.4335		
LCPI	2	M3	-2.3276	-3.4336	-7.9344	-3.4337	3	M1	10.4445	-1.9424	-5.8970	-1.9425
MMR	1	M2	-2.1773	-2.8765	-13.3394	-2.8765	3	M1	-2.20931	-1.9424		
LVT	2	M2	-1.8657	-2.8765	-7.1499	-2.8766	3	M2	-1.84635	-2.8765	-8.6930	-2.8765

Note: M1: Without constant and without trend,M2: With constant and without linear trend and M3: with constant and with linear trend;

From Table 2, these variables contain unit roots given the t-statistics are higher in level than the tabulated values of Mackinnon (1996). These unit roots are gone after a single differentiation. Hence, industrial production, money market rate, consumer price index and transaction volume are integrated in order one. Obtaining unit roots for the money market rate on monthly data was detected with the Dickey-Fuller test (1979). But, we used the Dickey-Fuller-Augmented test to have the existence of unit roots for the logarithm of the industrial production index, the logarithm of the consumer price index and the logarithm of transaction volume within the French stock market. We then use the Phillips-Perron test (1989) to detect absence/presence of unit roots while taking into account the existence of autocorrelation and heteroscedasticity problems for these variables. From Table 2, the results from Philips and Perron test (1988) prove the absence of unit roots for the logarithm of industrial production index (LIPI) and the money market rate (MMR). On the other hand, we detect unit roots presence for the consumer price index (CPI) and the logarithmic transaction volume (LVT) on the Paris Stock Exchange.

5. *Estimation Results and Interpretation:*

We attempt to examine dynamic relationships between the variables under study using from both Granger and Sims causality tests. The results are reported in Table 3.

Table 3. Results from Granger and Sims Causality Tests

	Granger Causality Test				Sims Causality Sense			
	Explanatory Variables	Lags	Fisher Statistics	Significance	Explanatory Variables	Lags	Fisher Statistics	Significance
LIPI	LPCI	(1,2)	33.5586	0.0000 cause	LCPI	(-1,2)	0.1071	0.7437 Does Not Cause
	MMR	(1,1)	27.5570	0.0000 cause	MMR	(-1,1)	0.1371	0.7115 Does Not Cause
	LVT	(1,1)	20.7593	0.0000 Cause	LVT	-(1,1)	1.6170	0.2050 Does Not Cause
LCPI	LIPI	(1,1)	0.4632	0.4969 Does Not Cause	LIPI	(-1,1)	6.4993	0.0116 Cause
	MMR	(1,1)	4.1718	0.0424 Cause	MMR	(-1,1)	1.0222	0.3132 Does Not Cause
	LVT	(1,1)	0.3067	0.5803 Does Not Cause	LVT	(-1,1)	0.1189	0.7306 Does Not Cause
MMR	LIPI	(1,1)	3.1530	0.0774 Cause	LIPI	(-1,1)	0.0136	0.9070 Does Not Cause
	LIPC	(1,1)	1.7836	0.1833 Does Not Cause	LCPI	(-1,1)	0.1094	0.7411 Does Not Cause
	LVT	(1,1)	0.0330	0.8559 Does Not Cause	LVT	(-1,1)	0.2858	0.8559 Does Not Cause
LVT	LIPI	(1,1)	0.2845	0.5943 Does Not Cause	LIPI	(-1,1)	0.0042	0.5943 Does Not Cause
	LIPC	(1,1)	0.4548	0.5008 Does Not Cause	LCPI	(-1,1)	0.6752	0.4122 Does Not Cause
	MMR	(1,2)	3.4700	0.0331 Cause	MMR	(-1,2)	0.0797	0.7779 Cause

Source: Prepared by researchers, 2021.

As reported in Table 3,there is a dynamic Granger (1969)causality between the consumer price index (CPI), the money market rate (MMR), the transaction volume (VT) and the industrial production index (IPI); a strong dynamic dependence between the monetary sphere specified by these monetary variables and the real sphere approximated by LIPI in France is well-documented. On the other hand, there is no dynamic relationship between industrial production index and inflation. This reflects that French national production is achieved by real and non-monetary indicators. Local inflation does not affect transaction volumes on the Paris stock exchange. Transaction volume and industrial production affect the money market rate as evidenced by Granger (1969). On the other hand, inflation does not cause in a Granger sense interest rate; there is not a strong sensitivity between interest rate and inflation rate in the short term. Real and monetary spheres Dichotomy is evidenced by the lack of causality sense between industrial production index (IPI) and trading volumes (VT) of the Paris stock exchange. The lack of bilateral causality sense is checked in Granger (1969)sense between inflation and transaction volumes. On the other hand, the monetary interest rate is more sensitive to transaction volume. According to Sims causality test (1972), consumer price index (CPI) causes industrial production. Transaction volume causes money market ratein Sims (1972)sense whereas the other variables do not. So, there are no advanced dynamic relationships between these variables.

Table 4shows the empirical results of the long-term relationship estimation of the basic model based on the ordinary least squares (OLS) method.

Table4. Results from the Ordinary Least Squares (OLS) Estimation

Variable	Coefficient	Std. Dev	T-Statistic	Significance
LCPI	0.9415	0.1215	7.7428	0.0000
LVT	0.0366	0.0199	1.8383	0.0676
MMR	-0.2362	0.0527	-4.4746	0.0000
Constant	0.4045	0.5575	0.7256	0.4690
R^2=0.9157, DW= 1.1510, F=681.4110 (0,0000)				
Residual stationarity Test				
			Residual (LIPI)	
Optimal number of delays			3	
Test Nature			ADF	
The residual follows a random walk with constant without linear trend				
T-Statistics			-3.8487 (-1.9425)	

Source: Prepared by researchers, 2021.

From Table 4, the model coefficients seem to be significant. Industrial production is isoelastic at inflation rate; the increase in economic growth is sensitive to the increase in the money supply. The financial crisis effect on economic growth is very small given that industrial production is less-elastic to transaction volume. The French Economic growth is achieved by savings because any increase in domestic production leads to a reduction in the money market rate (MMR). Examining further the target stationarity of industrial production index can help to confirm such relationship. Indeed, the residual of the aforementioned relationship seems to be stationary at the difference-free level given that the t-statistic is less than the critical value of Mackinnon (1996). This leads us to analyze the linear adjustment of the industrial production index with respect to its equilibrium value within an error correction model (ECM). The ECM combines the short-run equilibrium (where all the variables of the model are stationary by the first differences) and the long-run equilibrium (where the variables are stationary by the linear combination provided by the adjustment speed which has a negative and significant sign).The estimation of the ECM model is performed by the ordinary least squares method.

Table5. Results from ECM Estimation

Variables	Coefficient	Std. Dev	T-Statistics	Significance
Constant	0.0107	0.0055	1.9332	0.0547
Δ $LIPI_{t-1}$	-0.2829	0.0708	-3.9946	0.0001
Δ $LCPI_t$	-1.6918	1.2930	-1.3084	0.1924
Δ MMR_t	0.0029	0.0200	0.1482	0.8823
Δ LVT_t	-0.1073	0.0970	-1.1061	0.2701
$Residual_{t-1}$	-0.4240	0.0767	-5.5214	0.0000
R^2=0.363854, DW=2.203663, F=21.04831				

Source: Prepared by researchers, 2021.

From Table 5, and as expected, the adjustment speed has a negative and significant sign. Hence, 42.41% imbalance of the French industrial production index is corrected by the monetary authorities within the financial market in order to avoid the Parisian stock market crisis.

Afterwards, we use the Johansen (1990) multivariate cointegration technique to estimate the long-run relationships between the economic growth and the inflation rate, money market rate and transaction volume. The Johansen (1990)procedure requires fixing the optimal number of autoregressivevector delays based on two information criteria and the likelihood ratio test. The results are shown in Table 6.

Table6. Results from Optimal Number of VAR Delays and Number of Cointegration Relationships

Optimal Number of VAR Delays				
Lags	1	2	3	4
X_{1t} = ($LIPI_t$, $LCPI_t$, MMR_t, LVT_t)				
AIC	-30.8609	-31.1468	-31.1523	-31.1957*
Schwartz	-30.5204	-30.5316*	-30.2604	-30.0251
LR	66.8126 (0.0000)		13.2483 (0.6545)	20.3453 (0.2051)

Number of Cointegration Relationships								
	Test λ_{trace}				**Test λ_{max}**			
X_{1t} = ($LIPI_t$, $LPCI_t$, MMR_t, LVT_t)								
Null hypothesis	r=0	r$\leq$1	r$\leq$2	r $\leq$3	r=0	r=1	r=2	r=3
Alternative hypothesis	r$\geq$1	r$\geq$2	r$\geq$3	r=4	r=1	r=2	r=3	r=4
Statistical value	87 ,4	37,93	15,23	4,31	49,47	22,70	10,91	4,31
Critical value at 5%	53,12	34,91	19,96	9,24	28,14	22,00	15,67	9,24

	LIPI	LCPI	MMR	LBVMT
Vector Co-integrated standardized by LIPI	1	-2,24253	- 0,85157	0,0081

Source: Prepared by researchers, 2021.

From Table 6, the optimal number of delays for the autoregressive vector is equal to 2 according to the Schwartz information criterion and to the likelihood ratio test. Based on the AICcriterion, however, the optimal number of VAR is equal to 4. In this case, we retain 2 as a number of delays given that the likelihood ratio test is more applicable than the two information criteria.To fix the numbers of the cointegration relations, Johansen (1988) proposed the tests of trace and maximum own values. The first test makes it possible to detect the existence of more co-integrations r vectors as for the second; it makes it possible to check the hypothesis of the r+1 vectors presence of the co-integrations. Table 6also shows the trace and maximum own values tests. We find that there are two co-integration relations. From an economic standpoint, there is only one interpretable cointegral relationship. For this end, we use the weak exogeneity test or long-term causality in the Granger (1988)sense to choose a single long-term relationship among these two Co-integrating vectors.

Table 7. Results from Low Exogeneity Test or Granger Long-term Causality Test

X_{1t} = ($LIPI_t$, $LPCI_t$, MMR_t, LVT_t)				
Variables	LIPI	LPCI	MMR	LBVMT
$\chi^{2C}(1)$	6.8535	21.9904	0.2395	1.7146
Significance	0.0088	0.0000	0.6245	0.5186

Source: Prepared by researchers, 2021.

Table 7 reports that industrial production index and consumer prices index are not weakly exogenous, i.e. these two variables are subject to a phenomenon of error correction. On the other hand, the money market rate and the transaction volume are weakly exogenous; the two variables do not contribute to deviation correction of production index towards equilibrium. Hence, the French economic growth is insensitive to the financial crisis as evidenced by the low exogeneity of this crisis.

Error correction vector model estimation is carried out by the maximum likelihood procedure in the co-integrating vector and the adjustment matrixes as shown in Table 8.

Table 8. Results Estimation from Maximum Likelihood Method

Variables	Standard Co-integrator vectors (matrix β)	Coefficients with error corrections (matrix α)
LIPI	1.000	-0.0332
LPCI	-2.2425	-0.0036
MMR	- 0.8515	0.0031
LBVMT	0.0081	-0.0115

Source: Prepared by researchers, 2021.

The estimated coefficients seem to be significant. Nonetheless, the correction of the deviation from equilibrium is very small (about 3.3244%). This reflects that the French monetary authorities have intervened in a non-punctual manner for correcting economic growth given that a strong real and monetary spheres dichotomy is well-pronounced.

We then study the non-linear dynamics of the financial crisis impact on the French economic growth based on autoregressive models with smooth regime change (STAR). For this, we will refer to the maximum likelihood technique to determine the optimal number of delays of the autoregressive model for the French economic growth (proxied by the industrial product index in logarithmic Neperian). Table 9reports the AR model estimation for this index.

Table9. Model Estimation $\Delta LIPI$

Lags T-Statistics	$\Delta LIPI$
One Lag	-2.7133
Two Lags	1.1333
AR order	AR (1)
Constants	-0.0004
φ_1	-0.0519*
AR variance	0.0044
Ljung-Box Q-Statistics	
Q (6)	9.38 (0.15)
Q (12)	12.75 (0.39)
Q (18)	15.5 (0.63)
Q (24)	29.27 (0.21)

Source: Prepared by researchers, 2021.

From Table 9, we model the first difference of natural logarithm of the industrial production index by an AR(1) model with white noise residuals (as evidenced by the Ljung-Box Q-statistics). The variance of this model is very small, reflecting that the accuracy quality is very good for the AR(1) process. We afterwards examine the nonlinearity of the different components of the natural logarithmic difference of the industrial production index based on Lagrange Multiplier (LM) test (Table 10).

Table 10. Estimation Results from Linearity Test

d	Statistics	Δ LIPI	Δ LPCI	Δ MMR	Δ LVT
d=1	LM	3.2181 (0.0240)	0.9134 (0.4356)	0.0031 (0.9998)	3.1990 (0.0246)
d=2	LM	0.5735 (0.6331)			0.4615 (0.7095)

Source: Prepared by researchers, 2021.

From Table 10, we show that the first difference of the natural logarithms of the industrial production index and the transaction volume are nonlinear for the single delay case given that the LM statistics for these two-first differences are significant. On the other hand, the inflation rate and interest rate are linear as the LM statistics are not statistically significant.

As a last step in the STAR model specification, we choose between the LSTAR and ESTAR processes for the first difference of natural logarithm of the French economic growth. In this respect, Teräs virta (1994) proposed to compare the two auxiliary regressions associated with the linearity tests against these two processes. These two regressions differ only according to the presence of the term related to the LSTAR modelin the regression. The idea behind the Teräsvirta (1994) procedure is to compare the coefficients of the two regressions. If the specified model is ESTAR model, then there is no cube term. For this end, we use a Fisher test sequence. The results of the specifications are presented in Table 11.

Table 11.Estimation Results from Tests for Choosing between LSTAR and ESTAR Models

Variables	Δ LIPI	Δ LVT
d=1	H01 : 0.482 H02 : 8.467* H03 : 0.702	H01 : 8.361* H02 : 1.008 H03 : 0.1807

Source: Prepared by researchers, 2021.

Table 11reports the results of tests for selecting between the ESTAR and LSTARS models. Interestingly enough, the ESTAR models are estimated for the logarithm of the difference industrial production index and the natural logarithm of the first difference transaction volume given that the three Fisher statistics show that there is no cube term for these two primary differences.

The estimation of the ESTAR model parameters is retrieved based on the nonlinear least squares technique. Such procedure is similar to that performed by maximizing the likelihood function with normally distributed errors. As a matter of fact, the smoothing parameter seems to pose a major issue: When it is very high, it hinders the convergence of algorithm. In this regard, Teräsvirta (1994) proposed to standardize this smoothing parameter by dividing it by their standard deviation. The ESTAR model estimation results are shown in Table 12.

Table 12. Results from STECM ESTAR models Estimation

Results from STECM ESTAR models Estimation

STECM Model		
Estimation of LIPI's long-term relationship		
Variables	Coefficients	Significance
Constant	2.585	0.000
LVT	-1.786	0.000
Stationarity of residuals		
Residuals	Dickey-Fuller	Philips-Perron
T-Statistics	-12.642	-7.214
Critical value 5%	-3.124	-4.732
Linearity test		
Residuals	LM	Significance
d1	2.9888	0.0172
d2	0.7300	0.5353
Linear regime		
Variables	Coefficients	Significance
Constant	0.0012	0.7633
Δ LIPI	-0.3594	0.3667

Δ LVT	-0.1300	0.6573
Residuals-1	-0.2444	0.0832
	Nonlinear regime	
Constant	-0.0012	0.8640
Δ LIPI	0.4482	0.2924
Δ LVT	0.2245	0.5504
Residuals-1	-0.1641	0.3111
γ	0.2312	0.0713
Threshold	0.0166	0.07135
	ESTAR model	
Parameters	Δ LIPI	Δ LVT
	Linear regime	
α_{10}	-0.4721*	0,5711*
α_{11}	-0.3339*	0,3421*
	Nonlinear regime	
α_{20}	4.6714*	-0,0198*
α_{21}	0.9664*	-0,1363*
γ	0.9586*	0,4757*
Seuil	0.0001*	1,6457*
	Ljung-Box Q-Statistics	
Q (6)	11.5 (0.1)	10.17 (0.1)
Q (12)	14.8 (0.2)	21.47 (0.2)
Q (18)	17.5 (0.4)	29.19 (0.3)
Q (24)	31.6 (0.1)	33.21 (0.5)

Source: Prepared by researchers, 2021.

The estimation results show that the parameters of the estimated transitions are statistically significant for the first differences in the French economic growth and Parisian stock market transactions. In particular, the parameters of the linear and nonlinear regimes are significant at 1% level. The transition between the central and regimes is relatively slow. The presence of a nonlinear adjustment with mean reverting remains irreversible for these two first differences as the transition between regimes is smooth.The nonlinear dynamics of French economic growth related to transaction volume is analyzed based on a model with error correction and a smooth regime change (STECM). For this end, we estimate the long-term relationship between economic growth and transaction volume based on the OLS method and we check the stationarity of the residuals level of this relationship. We then test for the nonlinearity of such residuals and use the nonlinear least squares procedure to estimate the STECM model.

The nonlinear dynamics of the financial crisis impact on economic growth is only shown by the static relationship between the natural logarithm of the industrial production index and the trading volume of the Paris Stock Exchange. Such relationship is statistically significant and negative given the French economic growth is stationary in level from Dickey-Fuller (1979)and Philips-Perron (1988) tests. The STECM model is estimated by the Ordinary Least Squares Nonlinear procedure. As well, deterministic linear and nonlinear dynamics are not statistically significant. On the other hand, long-term linear and nonlinear dynamics are statistically significant and negative. Hence, the linear and nonlinear delayed order residuals bring the French economic growth back to a partially stable, long-term situation.

6. *CONCLUSION*

It is noteworthy that the 2007-2009 financial crisis has markedly gained in depth and intensity and revealed challenges for policymakers regarding their financial, monetary and economic policies. In this context, this paper investigates the financial crisis impact on the French economic growth.From an empirical standpoint, we use on a battery of main econometric and statistical tools to analyze the relationship between the industrial production index and the consumer price index, the money market rate and the transactions volume on the Paris Stock Exchange. The dataset is collected from the Central Bank of France, the International Monetary Fund and the World Bank during the period from 2000 to 2015 on monthly frequency. The empirical results show the financial crisis does not affect national production given that the transaction volume is weakly exogenous and does not undergo a phenomenon of deviation rectification of the French economic growth compared to its equilibrium value. Looking deeper into the aforementioned finding, we used the nonlinear cointegration to analyze the existence of a nonlinear long-term relationship between the industrial production and transaction volume. We clearly find that the French economic growth shows a nonlinear dynamic towards its fundamental value. Such finding seems to be interesting to assess the effectiveness of policy actions taken by French government to hamper the effects of financial crisis, ensure economic stabilization and boost the French growth perspectives.

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