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Credit Market Development and Economic Growth an Empirical Analysis for Greece

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Abstract: Problem statement: This study investigated the relationship between credit market development and economic growth for Greece for the period 1979-2007 using a Vector Error Correction Model (VECM). Questions were raised whether economic growth spurs credit market development taking into account the negative effect of inflation rate on credit market development. This study aimed to investigate the short-run and the long-run relationship between bank lending, gross domestic product and inflation rate applying the Johansen cointegration analysis. **Approach:** To achieve this objective classical and panel unit root tests were carried out for all time series data in their levels and their first differences. Johansen cointegration analysis was applied to examine whether the variables are cointegrated of the same order taking into account the maximum eigenvalues and trace statistics tests. Finally, a vector error correction model was selected to investigate the long-run relationship between economic growth and credit market development. **Results:** A short-run increase of economic growth per 1% induces an increase of bank lending 2.2%, while an increase of inflation rate per 1% induces a relative decrease of bank lending per 5.6% in Greece. The estimated coefficient of error correction term is statistically significant and has a negative sign, which confirms that there is not any problem in the long-run equilibrium between the examined variables. **Conclusion:** The empirical results indicated that there is a long-run relationship between economic growth and credit market development for Greece.

Key words: Credit market, economic growth, panel unit roots, vector error correction model

INTRODUCTION

The relationship between economic growth and financial development has been an extensive subject of empirical research. The question is whether banks or stock markets proceed or follow economic growth unless there is a complementary relationship between them. The main objective of this study was to investigate the relationship between economic growth and credit market development taking into account the effect of inflation rate on credit market development. Economic growth favors credit market development at times of low inflation rates. This study tries to confirm this hypothesis examining a model of banking system in which bank lending is dependent on gross domestic product and consumer price index.

The literature on financial liberalization encourages free competition among banks as the way forward to achieve economic growth. However, it has largely overlooked the possibility that endogenous constraints in the credit market, such as imperfect information, could be a significant obstacle to efficient credit allocation even when assuming that banks are free from interest rate ceilings. Stiglitz and Weiss (1981) were the first to consider the importance of banks in allocating credit efficiently, particularly to new and innovative investments. The expected return of the borrowers is an increasing function of the riskiness of their projects, the higher the risk the higher the return. This fact would discourage less risky investments from taking place, although they could be more productive (selection effect). Safe borrowers, which deal with banks only, will be left with no other choice.

King and Levine (1993) use different measures of bank development for several countries and find that banking sector development can spur economic growth in the long run. Levine (2002) emphasizes the critical importance of the banking system in economic growth and highlight circumstances when banks can actively spur innovation and future growth by identifying and funding productive investments. It is argued that banks can finance development more effectively than markets in developing economies and in case of state-owned banks, market failures can be overcome and allocation of savings can be undertaken strategically (Gerschenkron, 1962). Those banks that are unhampered

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by regulatory restrictions can exploit economies of scale and scope in information gathering and processing (Levine, 2002). Banks can ease distortions emanating from asymmetric information through forming long-run relationships with firms and, through monitoring, contain moral hazard. As a result, bank based arrangements can produce better improvement in resource allocation and corporate governance than market-based institutions (Bhide, 1993).

The effect of inflation on financial development is much more complicated. A rise of initially low inflation may not lead to detrimental consequences for financial activity, whereas a rise in the rate of inflation that is initially high may substantially depress activity on financial markets and entail reduction in financial depth.

If this hypothesis is true, then there is an inflation threshold in relationship between financial depth and inflation and this threshold can be regarded as an optimum rate of inflation with respect to financial development and therefore be a target for monetary authorities.

The model hypothesis predicts that economic growth facilitates credit market development taking into account the negative effect of inflation rate on credit market development and economic growth.

This study has two objectives:

- To examine the stationarity tests of the examined variables estimating classical and panel unit roots tests
- To examine the long-run relationship among economic growth, inflation rate and credit market development using Johansen co-integration analysis taking into account classical and panel unit root tests

The remainder of the study proceeds as follows: Initially the data and the specification of the multivariate VAR model are described. For this purpose stationarity test and Johansen co-integration analysis are examined taking into account the estimation of vector error correction model. Finally, the empirical results are presented analytically and some discussion issues resulted from this empirical study are developed shortly, while the final conclusions are summarized relatively.

MATERIALS AND METHODS

Data and specification model: In this study the method of Vector Autoregressive Model (VAR) is applied to estimate the effects of economic growth and inflation rate on credit market development. The use of this methodology predicts the cumulative effects taking into account the dynamic response among credit market development and the other examined variables (Shan, 2005).

In order to test the long-run relationships, the following multivariate model is to be estimated:

$$BC = f\ (CPI, GDP) \qquad (1)$$

Where:
BC = The domestic bank credits to private sector
CPI = The consumer price index
GDP = The gross domestic product

Following the empirical studies of King and Levine (1993), Vazakidis (2006), Vazakidis and Adamopoulos (2009c; 2009d; 2010a; 2010b), the variable of economic growth (GDP) is measured by the rate of change of real GDP, while the credit market development is expressed by the domestic bank credits to private sector (BC) as a percentage of GDP.

This measure has a basic advantage from any other monetary aggregate as a proxy for credit market development. Although it excludes bank credits to the public sector, it represents more accurately the role of financial intermediaries in channeling funds to private market participants (Katos *et al*., 1996; Vazakidis and Adamopoulos, 2009a, 2009b; Adamopoulos, 2010a; 2010b).

The data that are used in this analysis are annual covering the period 1979-2007 for Greece, regarding 2000 as a base year and are obtained from international financial statistics yearbook International Monetary Fund (2007). All time series data are expressed in their levels and Eviews econometric computer software is used for the estimation of the model.

Unit root tests: For univariate time series analysis involving stochastic trends, Phillips-Perron (PP) (1988) and Kwiatkowski *et al*. (1992) (KPSS) unit root tests are calculated for individual series to provide evidence as to whether the variables are integrated. This is followed by a multivariate co-integration analysis.

Phillips and Perron (1988) test is an extension of the Dickey and Fuller (1979) test, which makes the semi-parametric correction for autocorrelation and is more robust in the case of weakly autocorrelation and heteroskedastic regression residuals. According to Choi (1992), the Phillips-Perron test appears to be more powerful than the ADF test for the aggregate data.

Although the Phillips-Perron (PP) test gives different lag profiles for the examined variables (time series) and sometimes in lower levels of significance, the main conclusion is qualitatively the same as reported by the Dickey-Fuller (DF) test. Since the null hypothesis in the Augmented Dickey-Fuller test is that a time series contains a unit root, this hypothesis is accepted unless there is strong evidence against it. However, this approach may have low power against stationary near unit root processes.

Following the studies of Vazakidis and Adamopoulos (2009c, 2010a), the Phillips-Perron unit root test according to Laopodis and Sawhney (2007), the Phillips-Perron unit root test which is very general and can be used in the presence of heteroscedastic and autocorrelated innovations is specified as follows:

$$\ln(1+r) = \alpha + \beta\left(\frac{t\text{-}T}{2}\right) + \delta\ln(1+r_{t-1}) + \zeta_t \quad \text{for } t = 1, 2, \ldots, T \tag{2}$$

Where:

r_t = Denotes interest rate at time
t, (t-T/2) = Time trend
T = Sample size

Equation 2 tests three hypotheses: The first hypothesis is that the series contains a unit root with a drift with a drift and a time trend: H_0^1: δ=1. The second hypothesis is that the series contains a unit root but without a time trend: H_0^2: β=0, δ=1. The third hypothesis is that the series contains a unit root but without a drift or a time trend: H_0^3: a=0, β=0, δ=1. The statistics that are used to test each hypothesis are Z (t_δ), Z (Φ_2), Z (Φ_3), respectively and their corresponding equations are as follows:

$$Z(t_\delta) = \left(\frac{\sigma_0}{\sigma_{Tl}}\right)t_\sigma - \left(\frac{T^3}{3^{1/2}4D_{xx\sigma Tl}^{1/2}}\right)\left(\sigma_{Tl}^2\text{-}\sigma_0^2\right) \tag{2a}$$

$$Z(\Phi_3) = \left(\frac{\sigma_0^2}{\sigma_{Tl}^2}\right)\Phi_3 - \left(\frac{1}{2\sigma_{Tl}^2}\right)\left(\sigma_{Tl}^2 - \sigma_0^2\right)x \left[T(\delta-1) - \left(\frac{T^6}{48D_{xx}}\right)\left(\sigma_{Tl}^2 - \sigma_0^2\right)\right] \tag{2b}$$

$$Z(\Phi_2) = \left(\frac{\sigma_0^2}{\sigma_{Tl}^2}\right)\Phi_2 - \left(\frac{1}{3\sigma_{Tl}^2}\right)\left(\sigma_{Tl}^2 - \sigma_0^2\right)x \left[T(\delta-1) - \left(\frac{T^6}{48D_{xx}}\right)\left(\sigma_{Tl}^2 - \sigma_0^2\right)\right] \tag{2c}$$

Where:

$$\Phi_3 = \frac{T\left(\sigma_0^2 - \left(\bar{r}\text{-}\bar{r}_{t-1}\right)^2 - \sigma^2\right)}{2\sigma^2} \tag{2d}$$

$$\Phi_2 = \frac{T\left(\sigma_0^2 - \sigma^2\right)}{3\sigma^2} \tag{2e}$$

And σ^2 is the OLS residual variance, σ_0^2 is the variance under the particular hypothesis for the standard t-test for δ=1. D_{xx} is the determinant of the (X'X), where X is the T_3 matrix of explanatory variables in Equation 2. Finally, σ_{Tl} is a consistent estimator of the variance of ζ and is computed as follows:

$$\sigma_{Tl}^2 = \sum_{t=1}^{T}\frac{\zeta_t^2}{T} + \frac{\left(2\sum_{t=1}^{1}\sum_{t=s=1}^{T}\left(1-s/(1+1)\right)\zeta_t\zeta_{t-s}\right)}{T} \tag{2f}$$

where, s and l are the lag truncation numbers and s < l. The estimator σ_{Tl} is consistent under general conditions because it allows for effects of serially correlated and heterogeneously distributed innovations. The three statistics are evaluated under various lags (l = 0-12).

Since the null hypothesis in the Augmented Dickey-Fuller test is that a time series contains a unit root, this hypothesis is accepted unless there is strong evidence against it. However, this approach may have low power against stationary near unit root processes. Kwiatkowski *et al*. (1992) present a test where the null hypothesis states that the series is stationary.

The KPSS test complements the Augmented Dickey-Fuller test in that concerns regarding the power of either test can be addressed by comparing the significance of statistics from both tests. A stationary series has significant Augmented Dickey-Fuller statistics and insignificant KPSS. Following the studies of Chang (2002), Adamopoulos (2010b; 2010c), Vazakidis and Adamopoulos (2010b), according to Kwiatkowski *et al*. (1992) the test of KPSS assumes that a time series can be composed into three components, a deterministic time trend, a random walk and a stationary error:

$$y_t = \delta t + r_t + \varepsilon t$$

where, r_t is a random walk $r_t = r_{t-1} + u_t$. The u_t is iid (0, σ_u^2). The stationarity hypothesis implies that $\sigma_u^2 = 0$.

Under the null, y_t, is stationary around a constant (δ=0) or trend-stationary (δ≠0). In practice, one simply runs a regression of y_t over a constant (in the case of level-stationarity) or a constant plus a time trend (in the case of trend-stationary). Using the residuals, e_i, from this regression, one computes the LM statistic:

$$LM = T^{-2}\sum_{t=1}^{T}S_t^2 / S_{\varepsilon t}^2$$

wher,e $S^2_{\varepsilon t}$ is the estimate of variance of ε_t:

$$S_t = \sum_{i=1}^{t} e_i, t = 1, 2, \ldots\ldots T$$

The distribution of LM is non-standard: The test is an upper tail test and limiting values are provided by Kwiatkowski *et al*. (1992) via Monte Carlo simulation. To allow weaker assumptions about the behavior of ε_t, one can rely, following Phillips (1987) on the Newey and West (1987) estimate of the long-run variance of ε_t which is defined as:

$$S^2(l) = T^{-1}\sum_{t=1}^{T} e_i^2 + 2T^{-1}\sum_{s=1}^{l} w(s,l)\sum_{t=s+1}^{T} e_i e_{i-k}$$

where, w(s,l) = 1- s/(l+1). In this case the test becomes:

$$\nu = T^{-2}\sum_{t=1}^{T} S_t^2 / S^2(l)$$

which is the one considered here. Obviously the value of the test will depend upon the choice of the 'lag truncation parameter', l. Here we use the sample autocorrelation function of Δe_t to determine the maximum value of the lag length l.

The KPSS statistic tests for a relative lag-truncation parameter (l), in accordance with the default Bartlett kernel estimation method (since it is unknown how many lagged residuals should be used to construct a consistent estimator of the residual variance), rejects the null hypothesis in the levels of the examined variables for the relative lag-truncation parameter (l) (Katos, 2004).

Besides classical unit roots in this study the methodology of panel units roots tests is examined.

Following the study of Christopoulos and Tsionas (2004) and Levin *et al*. (2002) denoted as LLC panel unit root tests respectively resulted to the same conclusion. They consider the following basic ADF specification:

$$\Delta y_{it} = \alpha y_{it-1} + \sum_{j=1}^{p_i} \beta_{ij}\Delta y_{it-j} + X'_{it}\delta + \varepsilon_{it} \quad (3)$$

where we assume a common α= ρ-1 but allow the lag order for the difference terms, p_i to vary across cross-sections. The null and alternative hypotheses for the tests may be written as: H_0: a=0 but H_1: a<0. In LLC panel unit root test, the null hypothesis is the existence of a unit root, while under the alternative, there is no unit root.

Levin *et al*. (2002) consider the model:

$$y_{it} = \rho_i y_{i,t-1} + z'_{it}\gamma + u_{it} \quad (3a)$$

where, z_{it} are deterministic variables, u_{it} is $iid(0,\sigma^2)$ and P_1=P. They assume that there is a common unit root process so that ρ_t is identical across cross-sections.

The LLC test statistic is a t-statistic on p given by:

$$t_\rho = \frac{(\hat{\rho}-1)\sqrt{\sum_{i=1}^{N}\sum_{t=1}^{T}\tilde{y}_{i,t-1}^2}}{s_e} \quad (3b)$$

Where:

$$\tilde{y}_{it} = y_{it} - \sum_{s=1}^{T} h(t,s) y_{is}, \ \tilde{u}_{it} = u_{it} - \sum_{s=1}^{T} h(t,s) u_{is}$$

$$h(t,s) = z'_t(\sum_{t=1}^{T} z_t z'_t) z_s, \ s_e^2 = (NT)^{-1}\sum_{i=1}^{N}\sum_{t=1}^{T}\tilde{u}_{it}^2$$

And $\hat{\rho}$ is the OLS estimate of p . It can be shown that if there are only fixed effects in the model, then:

$$\sqrt{N}T(\hat{\rho}-1) + 3\sqrt{N} \to N(0, \tfrac{51}{5}) \quad (3c)$$

And if there are fixed effects and a time trend:

$$\sqrt{N}(T(\hat{\rho}-1) + 7.5) \to N(0, \tfrac{2895}{112}) \quad (3d)$$

In *et al*. (1997) denoted as IPS panel unit root tests respectively resulted to the same conclusion. In IPS panel unit root test, the null hypothesis is the existence of a unit root. The IPS statistic is based on averaging individual Dickey-Fuller unit root test (t_i) according to:

$$t_{IPS} = \frac{\sqrt{N}(\bar{t} - E[t_i \mid \rho_i = 0])}{\sqrt{var[t_i \mid \rho_i = 0]}} \to N(0,1) \quad (4)$$

where, $\bar{t} = N^{-1}\sum_{i=1}^{N} t_i$. The moments of $E[t_i \mid \rho_i = 0]$

And $var[t_i \mid \rho_i = 0]$ are obtained by Monte Carlo simulation and are tabulated in IPS (Christopoulos and Tsionas, 2004).

Following the studies of Christopoulos and Tsionas (2004); Kiran *et al*. (2009) and Breitung (1999) finds that IPS suffers a dramatic loss of power when individual trends are included and the test is sensitive to the specification of deterministic trends.

The Breitung (1999) denoted as (BR) panel unit root test assumes that there is a common unit root process so that ρ_i is identical across cross-sections. Under the null hypothesis, there is a unit root, while under the alternative, there is no unit root. LLC and Breitung examine the same basic ADF specification:

The Breitung panel unit root test differs from LLC in two distinct ways. First, only the autoregressive portion (and not the exogenous components) is removed when constructing the standardized proxies:

$$\Delta\tilde{y}_{it} = \left(\Delta y_{it} - \sum_{j=1}^{p_i} \beta_{ij} \Delta y_{it-j}\right) / s_i \tag{5a}$$

And $$\tilde{y}_{it-1} = \left(y_{it-1} - \sum_{j=1}^{p_i} \beta_{ij} \Delta y_{it-j}\right) / s_i \tag{5b}$$

where, $\tilde{\beta}, \beta$ and s_i are as defined for LLC.

Second, the proxies are transformed and detrended:

$$\Delta y_{it}^* = \sqrt{\frac{(T-t)}{(T-t+1)}}\left(\Delta\tilde{y}_{it} - \frac{\Delta\tilde{y}_{it+1} + \dots + \Delta y_{iT}}{T-t}\right) \tag{5c}$$

And:

$$y_{it}^* = \tilde{y}_{it} - \tilde{y}_{i1} - \frac{t-1}{T-1}\left(\tilde{y}_{iT} - \tilde{y}_{i1}\right) \tag{5d}$$

The persistence parameter is estimated from the pooled proxy equation:

$$\Delta y_{it}^* = a\, y_{it-1}^* + \nu_{it} \tag{5e}$$

Breitung test shows that under the null, the resulting estimator a* is asymptotically distributed as a standard normal. The Breitung panel unit root test requires only a specification of the number of lags used in each cross section ADF regression, p_i and the exogenous regressors. In contrast with LLC, no kernel computations are required.

The econometric software Eviews which is used to conduct the PP and KPSS tests, reports the simulated critical values based on response surfaces. The results of the Phillips and Perron (1988) and Kwiatkowski *et al*. (1992) unit root test and of Levin *et al*. (2002); Im *et al*. (1997) and Breitung (1999) panel unit roots tests for each variable appear in Table 1a-b. If the time series (variables) are non-stationary in their levels, they can be integrated with integration of order 1, when their first differences are stationary.

Johansen co-integration analysis: Since it has been determined that the variables under examination are integrated of order 1, then the cointegrated test is performed. The testing hypothesis is the null of nonco-integration against the alternative that is the existence of co-integration using the Johansen maximum likelihood procedure (Johansen, 1988).

Once a unit root has been confirmed for a data series, the question is whether there exists a long-run equilibrium relationship among variables. According to Granger (1986), a set of variables, Y_t is said to be co-integrated of order (d, b)-denoted CI(d, b)-if Y_t is integrated of order d and there exists a vector, β, such that $\beta' Y_t$ is integrated of order (d-b).

Co-integration tests in this study are conducted using the method developed by Johansen and Juselius (1990). The multivariate co-integration techniques developed by Johansen and Juselius (1990; 1992) using a maximum likelihood estimation procedure allows researchers to estimate simultaneously models involving two or more variables to circumvent the problems associated with the traditional regression methods used in previous studies on this issue. Therefore, the Johansen method applies the maximum likelihood procedure to determine the presence of co-integrated vectors in non-stationary time series.

Following the study of Chang and Caudill (2005); Johansen (1988) and Osterwald-Lenum (1992) propose two test statistics for testing the number of cointegrated vectors (or the rank of Π): The trace (λ_{trace}) and the maximum eigenvalue (λ_{max}) statistics.

The Likelihood Ratio statistic (LR) for the trace test (λ_{trace}) as suggested by Johansen (1988) is:

$$\lambda_{trace}(r) = -T\sum_{i=r+1}^{p} \ln(1-\hat{\lambda}_i) \tag{6}$$

Where:

$\hat{\lambda}_i$ = The largest estimated value of ith characteristic root (eigenvalue) obtained from the estimated Π matrix
r = 0, 1, 2, …p-1
T = The number of usable observations

The λ_{trace} statistic tests the null hypothesis that the number of distinct characteristic roots is less than or equal to r, (where r is 0, 1, or 2) against the general alternative.

Table 1a: PP, KPSS unit root tests

In levels					
		PP_ test stat		KPSS test stat	
Variables	$Z(\Phi_3)$	$Z(\Phi_2)$	$Z(t_\delta)$	h_c	h_t
BC	0.87(k=0)	-1.65(k=0)	-2.06 (k=3)	0.92(l=0)	0.26(l=0)
CPI	-1.28(k=0)	-0.70(k=0)	-3.26 (k=0)	2.65(l=0)	0.15* (l=0)
GDP	13.45(k=0)	0.25(k=1)	-1.99 (k=2)	2.68(l=0)	0.65(l=0)
In 1st differences					
ΔBC	-4.38(k=0)	-4.48(k=2)	-4.40(k=0)	0.08(l=0)	0.09(l=0)
ΔCPI	-4.98 (k=0)	-5.30(k=4)	-5.16(k=0)	0.08(l=0)	0.07(l=0)
ΔGDP	-0.26(k=3)	-1.84(k=3)	-3.98(k=0)	0.57(l=0)	0.14 (l=2)

$Z(\Phi_3)$, $Z(\Phi_2)$, $Z(t_\delta)$, are the PP statistics, h_c and h_t are the KPSS statistics, k, l= bandwidth lengths: Newey-West using Bartlett kernel. The critical values at 1%, 5% and 10% are -2.64, -1.95, -1.61, for $Z(\Phi_3)$, -3.67, -2.96, -2.62 for $Z(\Phi_2)$ and for -4.30, -3.57, -3.22 for $Z(t_\delta)$, respectively. The critical values at 1%, 5% and 10% are 0.73, 0.46 and 0.34 for h_c and 0.21, 0.14 and 0.11 for h_t respectively (Kwiatkowski *et al.* (1992) Table 1).Indicate that those values are not consistent with relative hypotheses at the 1%, 5% and 10% levels of significance relatively

Table 1b: IPS, LLC, BR panel unit root tests

In levels				
	LLC test stat		IPS test stat	BR test stat
Variables	LLC_C	LLC_T	IPS_C	BR_T
BC	-0.17	-0.25	-1.59	0.07
GDP	0.17	0.08	2.63	0.02
CPI	-0.04	-0.52	-0.70	0.03
In 1st differences				
ΔBC	-0.87	-0.88	-1.89	0.10
ΔGDP	-0.30	-0.78	0.71	0.03
ΔCPI	-0.94	-0.93	-1.67	0.04

Notes: LLC is the Levin, Lin and Chu t-test and IPS is the Im, Pesaran and Shin t-test test for unit root test in the model. The critical values for LLC_C test are 5.33 and -4.45 including only constant in levels and first differences respectively. The critical values for LLC_T test are 2.14 and -3.86 including constant and trend in levels and first differences respectively. The critical values for IPS_c test are 2.75 and 2.09 including only constant in levels and first differences respectively. The critical values for BR_T test are 4.85 and -2.93 including constant and trend in levels and first differences respectively

Table 2: Johansen Co-integration tests(BC, GDP, CPI)

Johansen test statistics				
	Critical value		Critical value	
Testing hypothesis	λ_{trace}	5%	λ_{max}	5 (%)
None*	54.55	42.91	30.94	25.82
At most 1	23.60	25.87	16.58	19.38
At most 2	7.01	12.51	7.01	12.51

Trace test and maximum eigenvalue tests indicate 1 co-integrating eqn (s) at the 0.05 level; *: Denotes rejection of he hypothesis at the 0.05 level; **: MacKinnon *et al.* (1999) p-values

In this statistic λ_{trace} will be small when the values of the characteristic roots are closer to zero (and its value will be large in relation to the values of the characteristic roots which are further from zero).

Alternatively, the maximum eigenvalue (λ_{max}) statistic as suggested by Johansen is:

$$\lambda_{max}\left(r,\ r+1\right) \ = \ -T\ln(1-\hat{\lambda}_{r+1}) \tag{7}$$

The λ_{max} statistic tests the null hypothesis that the number of r co-integrated vectors is r against the alternative of (r+1) co-integrated vectors. Thus, the null hypothesis r=0 is tested against the alternative that r =1, r=1 against the alternative r=2, r=2 against the alternative r=3 and so forth. If the estimated value of the characteristic root is close to zero, then the λ_{max} will be small.

It is well known that Johansen's co-integration tests are very sensitive to the choice of lag length. Firstly, a VAR model is fitted to the time series data in order to find an appropriate lag structure. The Schwarz (1978) Criterion is used to select the number of lags required in the co-integration test.

The Schwarz Criterion (SC) suggested that the value p=1 is the appropriate specification for the order of VAR model for Greece. Table 2 shows the results from the Johansen co-integration test.

Vector error correction model: Since the variables included in the VAR model are co-integrated, the next step is to specify and estimate a Vector Error Correction Model (VECM) including the error correction term to investigate dynamic behavior of the model. Once the equilibrium conditions are imposed, the VEC model describes how the examined model is adjusting in each time period towards its long-run equilibrium state.

Table 3: Vector error correction model

Sample (adjusted):	1980 2007		
Cointegrating Eq:	CointEq1		
BC (-1)	1.0000		
CPI (-1)	-5.6118		
	(-2.3323)		
GDP (-1)	2.2164		
	(3.3081)		
@TREND (78)	-0.1004		
	-2.6752)		
C	-0.3281		
Error correction:	D (BC)	D (CPI)	D (GDP)
CointEq1	-0.0050	0.0125	0.0558
	(0.0302)	(0.0105)	(0.0094)
	(-0.1669)	(1.1863)	(5.8982)
C	0.0163	-0.0055	0.0495
	(1.1103)	(-1.0703)	(10.7218)
F-stat:	0.0278	1.4073	34.7894
Akaike AIC	-2.1945	-4.3013	-4.5152
Schwarz SC	-2.0993	-4.2062	-4.4201

Since the variables are co-integrated, then in the short run, deviations from this long-run equilibrium will feed back on the changes in the dependent variables in order to force their movements towards the long-run equilibrium state. Hence, the co-integrated vectors from which the error correction terms are derived are each indicating an independent direction where a stable meaningful long-run equilibrium state exists.

The VEC specification forces the long-run behavior of the endogenous variables to converge to their co-integrated relationships, while accommodates short-run dynamics. The dynamic specification of the model allows the deletion of the insignificant variables, while the error correction term is retained (Katos, 2004). The size of the error correction term indicates the speed of adjustment of any disequilibrium towards a long-run equilibrium state (Engle and Granger, 1987). The error-correction model with the computed t-values of the regression coefficients in parentheses is reported in Table 3.

The final form of the Error-Correction Model (ECM) was selected according to the approach suggested by Hendry (Maddala, 1992). The general form of the Vector Error Correction Model (VECM) is the following one:

$$\Delta Y_t = \beta_1 \sum_{i}^{n} \Delta Y_{t-i} + \beta_2 \sum_{i}^{n} \Delta X_{t-i} + \beta_3 \sum_{i}^{n} \Delta Z_{t-i} + \lambda EC_{t-i} + \varepsilon_t \quad (8)$$

Where:

Δ = The first difference operator

EC_{t-1} = The error correction term lagged one period

Λ = The short-run coefficient of the error correction term ($-1<\lambda<0$)

ε_t = The white noise term

RESULTS

The observed t-statistics fail to reject the null hypothesis of the presence of a unit root for all variables in their levels confirming that they are non-stationary at 1, 5 and 10% levels of significance However, the results of the PP, KPSS, LLC IPS and BR tests show that all variables are stationary of the same order when they are transformed into their second differences (Table 1a-1b).

Therefore, all series that are used for the estimation are non-stationary in their levels, but stationary and integrated of order one I(1), in their first differences. These variables can be co-integrated as well, if there are one or more linear combinations among the variables that are stationary. The results that appear in Table 2 suggest that the number of statistically significant co-integrated vectors for Greece is equal to 1. The process of estimating the rank r is related with the assessment of eigenvalues, which are the following for Greece: $\hat{\lambda}_1 = 0.66$, $\hat{\lambda}_2 = 0.44$, $\hat{\lambda}_3 = 0.22$

For Greece, critical values for the trace statistic defined by Eq. 6 are 42.91 for none co-integrating vectors, 25.87 for at most one vector and 12.51 for at most two vectors at the 0.05 level of significance as reported by MacKinnon *et al*. (1999), while critical values for the maximum eigenvalue test statistic defined by Eq. 7 are 25.82 for none co-integrating vectors, 19.38 for at most one vector and 12.51 for at most two vectors respectively (Table 2).

Then the error-correction model with the computed t-values of the regression coefficients in parentheses is estimated. The dynamic specification of the model allows the deletion of the insignificant variables, while the error correction term is retained by the estimation of the co-intergrated vector. A short-run increase of economic growth per 1% induces an increase of bank credits per 2.2%, while an increase of consumer price index per 1% induces a decrease of bank credits per 5.6% for Greece (Table 3).

The estimated coefficient of EC_{t-1} is statistically significant and has a negative sign, which confirms that there is not any problem in the long-run equilibrium relation between the independent and dependent variables in 5% level of significance, but its relatively value (-0.005) for Greece shows a satisfactory rate of convergence to the equilibrium state per period (Table 3).

DISCUSSION

The model of banking system is mainly characterized by the effect of interest rates, investments and the circulation of money. However, bank development is determined by the size of bank lending directed to private sector at times of low inflation rates leading to higher economic growth rates.

Interest rate is not included in the estimated model of banking system due to the insignificance of estimation results. The significance of the empirical results is dependent on the variables under estimation.

Less empirical studies have concentrated on examining the relationship between economic growth and credit market development taking into account the effect of inflation rate. Most empirical studies examine the relationship between economic growth and stock market development.

The results of this study are agreed with the studies of Levine and Zervos (1998); Khan *et al*. (2001); Levine (2002) and Vazakidis and Adamopoulos (2009a; 2009b; 2010a), Adamopoulos (2010a, 2010b). However, more interest should be focused on the comparative analysis of empirical results for the rest of European Union members-states in future research.

CONCLUSION

This study employs with the relationship between credit market development and economic growth for Greece, using annually data for the period 1979-2007. For univariate time series analysis involving stochastic trends, Phillips and Perron (1988); Kwiatkowski *et al*. (1992) classical unit roots tests and Levin *et al*. (2002); Im *et al*. (1997) and Breitung (1999) panel unit roots tests are calculated for individual series to provide evidence as to whether the variables are stationary and integrated of the same order.

The empirical analysis suggested that the variables that determine credit market development present a unit root. Therefore, all series are stationary and integrated of order one I(1), in their first differences. Since it has been determined that the variables under examination are stationary and integrated of order 1, then the Johansen co-integration analysis is performed taking into account the maximum likelihood procedure. The short run dynamics of the model is studied by analyzing how each variable in a co-integrated system responds or corrects itself to the residual or error from the co-integrating vector. This justifies the use of the term error correction mechanism.

The Error Correction (EC) term, picks up the speed of adjustment of each variable in response to a deviation from the steady state equilibrium. The dynamic specification of the model suggests deletion of the insignificant variables while the error correction term is retained. The VEC specification forces the long-run behavior of the endogenous variables to converge to their co-integrating relationships, while accommodates the short-run dynamics.

A short-run increase of economic growth per 1% leaded to an increase of bank credits per 2.2%, while an increase of consumer price index per 1% leaded to a decrease of bank credits per 5.6% in Greece. Therefore, it can be inferred that economic growth has a positive effect on credit market development taking into account the negative effect of inflation rate on credit market development and economic growth.

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