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Credit Market Development and Economic Growth: An Empirical Analysis for Ireland

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Abstract:

This study investigated the causal relationship between credit market development and economic growth for Ireland for the period 1978-2007 using a vector error correction model (VECM). The purpose of this study was to investigate the short-run and the long-run relationship between the examined variables applying the Johansen cointegration analysis. For this purpose unit root tests were carried out according to Phillips-Perron (1988) and Kwiatkowski et al (1992), but also taking into account Levin et al (2002) panel unit root test. Finally, a vector error correction model was selected to investigate the long-run relationship between credit market development and economic growth taking into account the inflation rate. The results of Granger causality tests indicated that there is unidirectional causality between credit market development and economic growth for Ireland.

JEL Classification: C12, C13, C32, C52, E51

Keywords: Credit Market, Economic Growth, Panel Unit Roots, Granger causality

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1. 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 precede or follow economic growth unless there is a complementary relationship between them. The main objective of this paper 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 favours credit market development at times of low inflation rates. This paper 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.

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.

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.

Ball and Mankiw (1995) indicate that higher inflation necessarily raises inflation uncertainty. Higher inflation uncertainty increases the riskiness of all credits and therefore even previously ‘high quality borrowers’ get treated as the risky ones. To assure that credits are paid back banks may resort to more severe credit rationing.

Khan et al (2001) argue that besides inflation there might be several other factors affecting financial activity. Among them there are GDP per capita, the degree of openness and the share of public consumption in GDP as a measure of financial repression. Intuitively, the impact of these factors on financial development seems to be straightforward. A rise in GDP per capita and the degree of openness are likely to enlarge financial depth while a rise in financial repression and higher inflation seem to have an opposite result.

The remainder of the paper proceeds as follows: Section 2 describes the specification of the model, develops the Johansen cointegration analysis, analyses the vector error correction models and presents Granger causality tests, while section 3 presents the empirical results. Finally, section 5 provides the conclusions of this paper since a short discussion is summarized in section 4.

2. Data and Methodology

2.1. Data analysis

In this study the method of vector autoregressive model (VAR) is adopted to estimate the effects of economic growth on credit market development through the effect of consumer price index. 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 causal relationships, the following multivariate model is to be estimated

$$BC = f\,(CPI, GDP) \qquad (2.1)$$

where BC are the domestic bank credits to private sector, CPI is the consumer price index, GDP is the gross domestic product.

Following the empirical studies of King and Levine (1993a), Vazakidis (2006), Vazakidis and Adamopoulos (2009b,d; 2010a,b,c), Adamopoulos (2010a) 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 (Katsouli, 2003; Vazakidis and Adamopoulos, 2009a; 2009b, Adamopoulos, 2010b). The data that are used in this analysis are annual covering the period 1978-2007 for Ireland, regarding 2000 as a base year. All time series data are expressed in their levels and are obtained from International Financial Statistics, (International Monetary Fund, 2007).

2.2. Unit root tests:

For univariate time series analysis involving stochastic trends, Phillips-Perron (PP) and Kwiatkowski et al (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-Perron (PP) (1988) test is an extension of the Dickey-Fuller (DF) (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 (Laopodis and Sawhney, 2007) which is very general and can be used in the presence of heteroscedastic and autocorrelated innovations is specified as follows:

$$\ln(1+r) = \alpha + \beta\left(\frac{\mathrm{t - T}}{2}\right) + \delta \ln(1+\mathrm{r_{t-1}}) + \zeta_{\mathrm{t}} \qquad (2.2)$$

for $t=1,2,\ldots,T$ where r_t denotes interest rate at time t, $(t\text{-}T/2)$ is a time trend and T is the 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 $Z(t_\delta)$, $Z(\Phi_2)$, $Z(\Phi_3)$, respectively and their corresponding equations are as follows:

$$Z(t_\delta) = \left(\frac{\sigma_0}{\sigma_{\mathrm{Tl}}}\right) t_\sigma - \left(\frac{\mathrm{T}^3}{3^{1/2} 4\mathrm{D}^{1/2}_{\mathrm{xx}\sigma Tl}}\right)\left(\sigma^2_{\mathrm{Tl}} - \sigma^2_0\right) \qquad (2.2a)$$

$$Z(\Phi_3) = \left(\frac{\sigma_0^2}{\sigma_{\mathrm{Tl}}^2}\right)\Phi_3 - \left(\frac{1}{2\sigma_{\mathrm{Tl}}^2}\right)\left(\sigma_{\mathrm{Tl}}^2 - \sigma_0^2\right) x \left[T(\delta-1) - \left(\frac{\mathrm{T}^6}{48\mathrm{D}_{\mathrm{xx}}}\right)\left(\sigma_{\mathrm{Tl}}^2 - \sigma_0^2\right)\right] \qquad (2.2b)$$

$$Z(\Phi_2) = \left(\frac{\sigma_0^2}{\sigma_{\mathrm{Tl}}^2}\right)\Phi_2 - \left(\frac{1}{3\sigma_{\mathrm{Tl}}^2}\right)\left(\sigma_{\mathrm{Tl}}^2 - \sigma_0^2\right) x \left[T(\delta-1) - \left(\frac{\mathrm{T}^6}{48\mathrm{D}_{\mathrm{xx}}}\right)\left(\sigma_{\mathrm{Tl}}^2 - \sigma_0^2\right)\right] \qquad (2.2c)$$

where

$$\Phi_3 = \frac{\mathrm{T}\left(\sigma_0^2 - (\bar{\mathrm{r}} - \bar{\mathrm{r}}_{\mathrm{t-1}})^2 - \sigma^2\right)}{2\sigma^2} \qquad (2.2d)$$

$$\Phi_2 = \frac{\mathrm{T}\left(\sigma_0^2 - \sigma^2\right)}{3\sigma^2} \qquad (2.2e)$$

and σ^2 is the OLS residual variance, σ_0^2 is the variance under the particular hypothesis for the standard t-test for $\delta=1$. 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:

$$\sigma_{Tl}^2 = \sum_{t=1}^{T} \frac{\zeta_t^2}{T} + \frac{\left(2\sum_{t=1}^{1} \sum_{t=s=1}^{T} \left(1 - s/(l+1)\right)\zeta_t \zeta_{t-s}\right)}{T} \qquad (2.2f)$$

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 to 12).

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 statistics.

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 $(0, \sigma_u^2)$. The stationarity hypothesis implies that $\sigma_u^2 = 0$. Under the null, y_t, is stationary around a constant ($\delta=0$) or trend-stationary ($\delta \neq 0$). 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 $LM = T^{-2}\sum_{t=1}^{T} S_t^2 / S_{\varepsilon t}^2$ where $S_{\varepsilon t}^2$ is the estimate of variance of $\boldsymbol{\varepsilon_t}$.

$S_t = \sum_{i=1}^{t} e_i$, $t = 1,2,\ldots\ldots T$. 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 behaviour of ε_t, one can rely, following Phillips (1987) on the Newey and West (1987) estimate of the long-run variance of $\boldsymbol{\varepsilon_t}$ which is defined as:

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

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

The Eviews 5.0 (2004) software package which is used to conduct the PP, KPSS, LL tests, reports the simulated critical values based on response surfaces. The results of PP, KPSS tests for each variable appear in Table 1. 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.

2.3. Panel unit roots:

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

$$\Delta y_{it} = \alpha y_{it-1} + \sum_{j=1}^{p_i} \beta_{ij} \Delta y_{it-j} + X'_{it}\delta + \varepsilon_{it} \qquad (2.3a)$$

where we assume a common $\alpha = \rho - 1$ 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 LL panel unit root test, the null hypothesis is the existence of a unit root, while under the alternative, there is no unit root.

Levin, Lin, and Chu (2002) consider the model

$$y_{it} = \rho_i y_{i,t-1} + z'_{it}\gamma + u_{it} \qquad (2.3b)$$

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

The LL test statistic is a $t-$statistic on ρ given by

$$t_\rho = \frac{(\hat{\rho}-1)\sqrt{\sum_{i=1}^{N}\sum_{t=1}^{T}\tilde{y}_{i,t-1}^2}}{s_e} \qquad (2.3c)$$

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 ρ (Christopoulos and Tsionas, 2004). It can be shown that if there are only fixed effects in the model, then

$$\sqrt{N}T(\hat{\rho}-1)+3\sqrt{N} \rightarrow N(0,\tfrac{51}{5}) \qquad (2.3d)$$

and if there are fixed effects and a time trend,

$$\sqrt{N}(T(\hat{\rho}-1)+7.5) \rightarrow N(0,\tfrac{2895}{112}) \qquad (2.3e)$$

2.4. Cointegration test:

Since it has been determined that the variables under examination are integrated of order 1, then the cointegration test is performed. The testing hypothesis is the null of non-cointegration against the alternative that is the existence of cointegration 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 paper 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 and Juselius is

$$\lambda_{trace}\,(r) = -T\sum_{i=r+1}^{p}\ln(1-\hat{\lambda}_i) \qquad (2.4a)$$

where $\hat{\lambda}_i =$ is the largest estimated value of ith characteristic root (eigenvalue) obtained from the estimated Π matrix, r = 0, 1, 2,…….p-1, and T is 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. 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 and Juselius is

$$\lambda_{max}(r, r+1) = -T\ ln(1-\widehat{\lambda}_{r+1}) \qquad (2.4b)$$

The λ_{max} statistic tests the null hypothesis that the number of r cointegrated vectors is r against the alternative of $(r+1)$ cointegrated vectors. Thus, the null hypothesis r=0 is tested against the alternative that r=1, r=1 against the alternative r=2, 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 cointegration 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 Criterion (SC) (1978) and the likelihood ratio (LR) test are used to select the number of lags required in the cointegration test and suggested that the value p=3 is the appropriate specification for the order of VAR model for Ireland. Table 3 presents the results from the Johansen and Juselius (1990) cointegration test.

2.5. Vector error correction model:

Since the variables included in the VAR model are found to be cointegrated, the next step is to specify and estimate a vector error correction model (VECM) including the error correction term to investigate dynamic behaviour 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. The dynamic specification of the model allows the deletion of the insignificant variables, while the error correction term is retained.

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 4. The final form of the vector error-correction model (VECM) was selected according to the *general to specific* methodology suggested by Hendry (Maddala, 1992). The general form of the vector error correction model (VECM) is the following:

$$\Delta BC_t = b_0 + \sum_i^n b_1 \Delta BC_{t-i} + \sum_i^n b_2 \Delta GDP_{t-i} + \sum_i^n b_3 \Delta CPI_{t-i} + \lambda EC_{t-1} + \varepsilon_t \qquad (2.5a)$$

where: Δ is the first difference operator, EC_{t-1} is the error correction term lagged one period, λ is the short-run coefficient of the error correction term ($-1<\lambda<0$), ε_t is the white noise term.

2.6. Granger causality tests:

Granger causality is used for testing the long-run relationship between credit market development and economic growth. The Granger procedure is selected because it consists the more powerful and simpler way of testing causal relationship (Granger, 1986). The following bivariate model is estimated:

$$Y_t = a_{10} + \sum_{j=1}^{k} a_{1j} Y_{t-j} + \sum_{j=1}^{k} \beta_{1j} X_{t-j} + u_t \qquad (2.6a)$$

$$X_t = a_{20} + \sum_{j=1}^{k} a_{2j} X_{t-j} + \sum_{j=1}^{k} \beta_{2j} Y_{t-j} + u_t \qquad (2.6b)$$

where Y_t is the dependent and X_t is the explanatory variable and u_t is the white noise error term in Eq (2.6a), while X_t is the dependent and Y_t is the explanatory variable in Eq (2.6b). The hypotheses in this test may be formed as follows:

H_0: X does not Granger cause Y, i.e. $\{\alpha_{11}, \alpha_{12},\ldots\alpha_{1k}\}=0$, if $F_c <$ critical value of F.
H_a: X does Granger cause Y, i.e. $\{\alpha_{11}, \alpha_{12},\ldots\alpha_{1k}\}\neq 0$, if $F_c >$ critical value of F. (2.6c)

and

H_0: Y does not Granger cause X, i.e. $\{\beta_{21}, \beta_{22},\ldots\beta_{2k}\}=0$, if $F_c <$ critical value of F.
H_a: Y does Granger cause X, i.e. $\{\beta_{21}, \beta_{22},\ldots\beta_{2k}\}\neq 0$, if $F_c >$ critical value of F. (2.6d)

(Seddighi et al, 2000, Katos, 2004, page 1043).

In order to test the above hypotheses the usual Wald F-statistic test is utilised, which has the following form

$$F = \frac{(RSS_R - RSS_U)/q}{RSS_U/(T-2q-1)}$$

where: RSS_U= is the sum of squared residuals from the complete (unrestricted) equation

RSS_R= the sum of squared residuals from the equation under the assumption that a set of variables is redundant, when the restrictions are imposed, (restricted equation)

T = the sample size and q = is the lag length.

Examining this model the following cases can be distinguished

1. If $\{\alpha_{11}, \alpha_{12},\ldots\alpha_{1k}\}\neq 0$ and $\{\beta_{21}, \beta_{22},\ldots\beta_{2k}\}=0$, there exists a unidirectional causality from X to Y, denoted as X→Y
2. If $\{\alpha_{11}, \alpha_{12},\ldots\alpha_{1k}\}=0$ and $\{\beta_{21}, \beta_{22},\ldots\beta_{2k}\}\neq 0$, there exists a unidirectional causality from Y to X, denoted as Y→X
3. If $\{\alpha_{11}, \alpha_{12},\ldots\alpha_{1k}\}\neq 0$ and $\{\beta_{21}, \beta_{22},\ldots\beta_{2k}\}\neq 0$, there exists a bilateral causality between Y and X, denoted as X↔Y

The validity of the test depends on the order of the VAR model and on the stationarity or not of the variables. The results related to the existence of Granger causal relationships among credit market development, economic growth and inflation appear in table 5.

3. Empirical results

Based on Levine and Zervos (1998), Khan et al (2001), Levine (2002), Vazakidis and Adamopoulos, (2009a; 2009b; 2010a) studies the model of credit market development is mainly characterized by the effect of economic growth and inflation. The significance of the empirical results is dependent on the variables under estimation. The results of the PP, and KPSS tests show that the null hypothesis of the presence of a unit root is rejected for all variables when they are transformed into their first differences (Tables 1). The LL unit root test results for both levels and first differences of economic growth, inflation and credit market development are reported in Table 2. The combined results of unit root tests (PP, KPSS, LL) suggested that all variables can be characterized as stationary and integrated of order one, I(1). So, these variables can be cointegrated as well, if there are one or more linear combinations among the variables that are stationary.

Table 1 -Tests of unit roots hypothesis

	Phillips-Perron PP_ test stat			**KPSS LM test stat**	
	$Z(\Phi_3)$	***$Z(\Phi_2)$***	***(t_δ)***	***h_c***	***h_t***
BC	4.07 (k=2)	3.55 (k=1)	1.57 (k=1)	0.61* (l=4)	0.17* (l=4)
CPI	-1.66*** (k=0)	-1.80 (k=0)	-1.70 (k=4)	0.39*,** (l=4)	0.15*(l=3)
GDP	-1.37 (k=1)	-2.81*** (k=2)	-2.88 (k=3)	0.16*,** (l=16)	0.25 (l=0)
ΔBC	-1.94*,** (k=3)	-2.78*,** (k=3)	-3.58*(k=2)	0.45***(l=3)	0.147**,***(l=2)
ΔCPI	-4.30 (k=0)	-4.23 (k=0)	-4.20* (k=0)	0.22 (l=7)	0.16 **,***(l=9)
ΔGDP	-7.43 (k=1)	-7.32 (k=1)	-7.18 (k=1)	0.06 (l=1)	0.05 (l=1)

The critical values for the Phillips-Perron unit root tests are obtained from Dickey-Fuller (1981), $Z(\Phi_3)$, $Z(\Phi_2)$ and $Z(t_\delta)$ are the PP statistics for testing the null hypothesis the series are not I(0) when the residuals are computed from a regression equation without an intercept and time trend, with only an intercept, and with both intercept and time trend, respectively. 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 -4.29, -3.56, -3.21 for $Z(t_\delta)$ respectively.

k= bandwidth length: Newey-West using Bartlett kernel

h_c and h_t are the KPSS statistics for testing the null hypothesis that the series are I(0) when the residuals are computed from a regression equation with only an intercept and intercept and time trend, 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). Since the value of the test will depend upon the choice of the 'lag truncation parameter', l. l= bandwidth length: Newey-West using Bartlett kernel.*, **, *** indicate that those values are not consistent with relative hypotheses at the 1%, 5% and 10% levels of significance relatively.

Table 2 - Panel unit root tests

Countries	**Variables**	**Levels**		**1^{st} Differences**	
		LLC_C	**LLC_T**	**LLC_C**	**LLC_T**
Ireland	BC	0.15967	0.15007	-0.50856	-0.70390
	GDP	-0.44667	-0.29276	-1.27380	-1.27299
	CPI	-0.12402	0.46359	-1.33167	-1.40846

Notes: LLC is the Levin, Lin, and Chu t-test for unit root test in the model.

The critical values for LLC_c test are 3.99, and -6.28 including only constant in levels and first differences respectively.

The critical values for LLC_T test are 2.30 and -6.35 including constant and trend in levels and first differences respectively,

*, **, *** indicate that those values are not consistent with relative hypotheses at the 1%, 5% and 10% levels of significance relatively.

The number of statistically significant cointegration vectors for Ireland is equal to 1 (Table 3) and is the following:

$$BC_t = -1.36\, CPI_t + 4.05\, GDP_t \qquad (3.1)$$

The cointegration vector of the model of Ireland has rank r<n (n=2). The process of estimating the rank r is related with the assessment of eigenvalues, which are the following for Ireland: $\hat{\lambda}_1 = 0.52$, $\hat{\lambda}_1 = 0.32$, $\hat{\lambda}_1 = 0.065$. The critical values for the trace statistic defined by equation (2.4a) are 24.31 for Ho: r = 0, 12.36 for Ho: $r \leq 1$, 4.16 for Ho: $r \leq 2$ at the significance level 5%, while critical values for the maximum eigenvalue test statistic defined by equation (2.4b) are 17.68 for Ho: r = 0, 11.03 for Ho: $r \leq 1$, 4.16 for Ho: $r \leq 2$ (Table 3).

Table 3 –Johansen and Juselious Cointegration Tests (BC, GDP, CPI)

Country	**Ireland**			
Testing Hypothesis	**Johansen Test Statistics**			
	λ_{trace} 5%	Critical values	λ_{max} 5%	Critical values
None*	29.45	24.31	20.33	17.89
At most 1	9.11	12.53	7.27	11.44
At most 2	1.85	3.84	1.85	3.84

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;

It is obvious from the above cointegrated vector that economic growth has a positive effect on credit market development in the long-run, while inflation has a negative effect on it. According to the signs of the vector cointegration components and based on the basis of economic theory the above relationship can be used as an error correction mechanism in a VAR model for Ireland respectively. The results of the estimated vector error correction model suggested that a short-run increase of economic growth per 1% induces an increase of bank credits per 4.05% in Ireland, and also an increase of consumer price index per 1% induces a decrease of bank credits per 1.36% for Ireland (Table 4). 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 (-0.16) for Ireland shows a satisfactory rate of convergence to the equilibrium state per period (Table 4).

Table 4- Vector Error Correction Model

$$\Delta BC_t = -0.01 + 0.45\ \Delta GDP_{t-1} - 0.39\ \Delta CPI + 0.08\ \Delta BC_{t-3} - 0.16\ u_{t-1}$$

(-0.49)	(1.10)	(-0.3957)	(0.337)	(-2.86)
[0.624]	[0.282]	[0.696]	[0.739]	[0.009]

$R^2 = 0.33$ DW = 1.44

Δ: Denotes the first differences of the variables, R^2 = Coefficient of multiple determinations adjusted for the degrees of freedom (d.f), DW= Durbin-Watson statistic u_t = is the standard error of regression.

In order to proceed to the Granger causality test the number of appropriate time lags was selected in accordance with the VAR model. According to Granger causality tests there is unidirectional causal relationship between economic growth and credit market development with direction from credit market to economic growth for Ireland (Table 5).

Table 5 – Pairwise Granger Causality Tests

Sample 1978-2007

Lags (3)

Null Hypothesis:	Obs	F-Statistic	Probability
CPI does not Granger Cause GDP	30	0.25503	0.8569
GDP does not Granger Cause CPI		0.57617	0.6365
BC does not Granger Cause GDP	30	3.33186	0.0372
GDP does not Granger Cause BC		0.06990	0.9754
BC does not Granger Cause CPI	30	0.37669	0.7706
CPI does not Granger Cause BC		0.25381	0.8578

4. 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 reverse relationship between economic growth and credit market development taking into account the effect of inflation rate. The results of this paper are agreed with the studies of Khan et al (2001) and Levine (2002). However, more interest should be focused on the comparative analysis of empirical results for the rest of European Union members-states in future research.

5. Conclusions

This study employs with the relationship between credit market development and economic growth for Ireland, using annually data for the period 1978-2007. For univariate time series analysis involving stochastic trends, Phillips-Perron (PP) (1988), Kwiatkowski et al (KPSS) (1992) classical unit roots tests, and Levin, Lin, and Chu (LL) (2002) 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 behaviour 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 4.05%, while an increase of consumer price index per 1% leaded to a decrease of bank credits per 1.36% in Ireland.

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. The results of Granger causality tests indicated that there is unidirectional causal relationship between economic growth and credit market development with direction from credit market development to economic growth for Ireland.

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