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Abstract
In this study, building a simple model that incorporates static and dynamic elements, the relationship of financial development and economic growth on the environmental degradation is investigated together with the validation of the Environmental Kuznets Curve (EKC) hypothesis. Our analysis is based on an unbalanced panel data set covering the OECD countries over the period 1970-2014. Our approach strongly accounts for the presence of cross-sectional dependence between the sample variables and utilizes second generation panel unit root tests in order to investigate possible cointegration relationships. The empirical findings do indicate that local (NO_X per capita emissions) and global (CO_2 per capita emissions) pollutants redefine the EKC hypothesis when we account for the presence of financial development indicators. Specifically, in the case of global pollution an N-shape relationship is evident both in static and dynamic framework with a very slow adjustment, whereas a monotonically decreasing relationship is found in the case of local pollutants with a much quicker dynamic adjustment. Lastly, we argue that policy makers and government officials have to cultivate investments in network industries (energy, telecommunications, transportation) by promoting cutting edge research and development financial projects and cost effective mitigation methods.

JEL classifications: C33; G20; Q43; Q53; Q56.

Keywords: Environmental Kuznets Curve; Cross-sectional dependence; Financial development; Panel data.

1. Introduction

During the last years there is a vast body of literature examining the validity of the Environmental Kuznets Curve (EKC) hypothesis using panel data techniques (parametric and semiparametric models) with controversial results (among others, Suarez and Menendez 2015; Apergis *et al.* 2014; Halkos 2013, 2003; Halkos and Tsionas, 2001; Desbordes and Verardi 2012; Cole 2004; Millimet *et al*. 2003; Zaim and Taskin 2000; Holtz-Eakin, and Selden 1995).[1]

At the same time the effects of financial development and economic growth on pollutants' emissions can be decomposed into three different elements: scale, technique and composition effects. More specifically, the greater the financial and economic activity the greater will be the emissions as more inputs are used ("*scale*" effect). However, an increase in the level of economic activity may raise serious environmental concerns thus leading to a reduction in the level of anthropogenic emissions (i.e. greenhouse gases emissions) triggered by the use of cleaner technologies in the production process. The latter justifies the existence of the "*technique*" effect (Grossman and Krueger, 1995). In other words, as income rises it is likely that the demand for cleaner goods also increases. This may induce firms to alter production methods and reduce pollution ("*composition*" effect). As a consequence, the existence of the EKC hypothesis justifies that at lower income levels the scale effect dominates the composition effect, but as income reaches a critical threshold (turning point) the latter effect offsets the former (Halkos, 2013; Jayanthakumaran and Liu, 2012).

Despite the plethora of studies devoted to this topic, existing studies suffer from two shortcomings. First, they assume that the variables or the random disturbances are not correlated across the panel dimension justifying the existence of cross-sectional independence. However, it

[1] The EKC hypothesis implies a non-linear relationship of an inverted '*U*' type between environmental degradation and economic growth.

is common for macro-level data to violate this assumption which will result in low power and size distortions of tests that assume cross-section independence. The latter may arise due to common unobserved effects triggered by changes in environmental legislation in the OECD countries. Therefore, cross section independence is a strong assumption that has to be tested rather than assumed in order to avoid misleading results.

Second, none of the existing studies account for the interdependence between the financial development and economic growth under the presence of local and global pollutants such as NO_x and CO_2 emissions respectively, which in our models acts as a driving force to reveal the validity of the EKC hypothesis.

This study aims to contribute to the financial development-economic growth nexus by building a simple model within a static and dynamic framework to investigate the validity of the Environmental Kuznets Curve (EKC) hypothesis. One of the main novelties of the paper is that it strongly accounts for the presence of cross section dependence along the suggested Pesaran (2004) CD tests. Moreover, it utilizes "*second-generation*" panel unit root tests in order to uncover possible cointegrated relationships an issue that has been overlooked by the existing empirical literature on EKC. The reason for using this kind of unit root testing can be justified by the fact that traditional stationarity tests (known as "*first-generation*" tests) suffer from size distortions and the ignorance of cross section dependence (Apergis, 2016).

The empirical findings do indicate that local (NO_X per capita emissions) and global pollutants (CO_2 per capita emissions) redefine the validity of the EKC hypothesis when we account for the presence of financial development indicators.

The rest of this paper is as follows. Section 2 reviews the empirical literature. Section 3 describes the data and the econometric methodology used in the empirical analysis. Section 4 reports the main empirical findings along with the necessary tests for cross section dependence. Lastly, Section 5 concludes the paper and provides some policy implications.

2. Review of the literature

The literature on EKC hypothesis starts with the pioneering study of Grossman and Krueger (1995) who examined the reduced-form relationship between per capita income and various environmental indicators (i.e. air pollution, river oxygen regime, contamination of river basins) to conclude that environmental quality gradually deteriorates with economic growth (inverted U shape).

The majority of the empirical studies regarding EKC use econometric models (i.e. non linear log models, error correction models, VAR/VECM, etc) dealing with stationarity and cointegration properties where the dependent variable is usually the (per capita) level of pollutants (i.e. CO_2 emissions, NO_X and SO_2 emissions, etc) regressed on different polynomials (powers) of (per capita) GDP, and other covariates including inter alia various efficiency indicators (see for example Halkos 2003, 2013; Halkos and Tzeremes 2009a,b, 2010, 2013; Managi and Kaneko 2006; Shen 2006, 2008; Stern, 2004; Halkos and Tsionas, 2001). The vast majority of these studies consent that there is an inverted U-shape between the level of environmental pollution and economic growth implying the validity of the EKC hypothesis. However, a recent study by Halkos and Polemis (2016) argues that global pollutants such as CO_2 emissions exhibit an N-shape implying that environmental damage starts rising again after a fall to a specific turning point.

On the other hand, relatively few empirical studies adopt a simultaneous equations system to address the impact of economic growth on environmental degradation. Dean (2002) uses a panel simultaneous equations system drawn from a Heckscher-Ohlin model in order to capture certain effects of trade liberalization on the environmental quality (water pollution). His findings suggest that there is a direct negative trade effect on environmental damage which is fully reversed when the income growth is taken into account. In the paper of Jayanthakumaran and Liu (2012) an array of econometric techniques are applied ranging from a quadratic log function specification to a SUR system similar to Dean's approach to provide little support in favor of the EKC hypothesis.

The existing empirical literature on environmental pollution and financial developments is still in its infancy, with controversial results since the researchers acknowledge that emissions may have positive as well as negative effects on financial development (Halkos and Sepetis, 2007; He and Wand, 2102). In the seminal paper of Frankel and Romer (1999) it is argued that during the process of financial development, developing countries will be motivated toward the adoption of cleaner energy technologies, which is a move to reduce the environmental effects. Moreover, they claim that financial development is the driving force for the companies in order to obtain capital and reducing financing costs by adopting environmental friendly techniques. This finding is also evident in the study of Yuxiang and Chen (2010) who argue that promoting financial development policies is a key issue in order to stimulate technological spillovers, which in turns reduce CO_2 emissions and enhance domestic production. Similarly, Cole and Elliot (2005) examine the impact between financial development and environmental degradation as expressed by CO_2 emissions. In their study, they claim that financial tools such as loans, leasing, factoring, treasury bonds, derivatives, allows medium and large scale firms to achieve economies of scale, thereby reducing the use of resources as well as CO_2 emissions. We must stress though that small scale firms are more prone to increase their size as a result of the existence of a strong and robust financial sector and thus raise the amount of carbon dioxide emissions.

On the other hand, there are some studies arguing that financial development creates a negative impact to the environment. More specifically, Zhang (2011) argues that financial development leads to the inefficiency, increasing the toxic releases emissions. Shahbaz (2012) study the impact of financial development on the environmental quality claiming that an organised financial sector attracts Foreign Direct Investments (FDIs), which then stimulate the efficiency of the operation of the stock market and economic activity, leading to a an increasing path of the CO_2 emissions. It is also interesting to mention that there are some studies who found a neutral effect between financial development as expressed by some proxy variables (i.e. treasury bonds, non performing loans, credit risk, stocks, etc) and the level of environmental awareness (Ozturk and Acaravci, 2013).

In a recent study, Lee *et al.* (2015) investigate the validity of the EKC hypothesis between the level of global pollutants (CO_2 emissions) and financial development for a panel dataset consisting of 25 OECD countries over the period 1971-2007. The authors use panel Fully Modified OLS estimators (FMOLS), rejecting the existence of the EKC for their sample countries. Lastly, Shahbaz *et al.* (2016) re-examine the asymmetric impact of financial development on environmental quality in Pakistan for the period 1985-2014 using quarterly data. Their approach is similar to ours in a sense that they use comprehensive indices of financial development generated by using Bank and Stock market based financial development indicators. They claim that inefficient use of energy negatively affects the level of environmental quality, implying that the adoption of energy efficient technology is of a paramount importance.

3. Data and methodology

The econometric estimation was based on an unbalanced panel of 34 OECD countries covering the period 1970-2014 (n = 34 and T = 44).[2] The latter was dictated by data availability. The environmental variables entering the models (CO_2, and NO_x per capita emissions measured in metric tons of CO_2 equivalent) are obtained by the World Bank (World Development Indicators Database). The banking development indicators were drawn from the World Bank (Financial Development and Structure Database) and were selected following the existing literature (see for example Antzoulatos *et al.* 2011). More specifically, we use the private credit by deposit money banks as a percentage to GDP (CREDIT).

This indicator denotes, the financial resources provided to the private sector by domestic money banks as a share of GDP. Domestic money banks comprise commercial banks and other financial institutions that accept transferable deposits, such as demand deposits. The other indicator (STOCK) is the stock market capitalization to GDP and includes the total value of all listed shares in a stock market as a percentage of GDP. The third financial development indicator (BOND) stands for the corporate bond issuance volume to GDP and denotes the ratio of newly issued corporate bonds by private entities in industries other than finance, holding companies and insurance, divided by GDP in current USD.

Table 1 depicts the main descriptive statistics from the model variables. We must stress that due to the lack of sufficient comparable data, we could not include other banking development indicators such as central bank assets to GDP, credit to bank deposits and non-

[2] These include the following countries: Australia, Austria, Belgium, Canada, Czech Republic, Denmark, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Italy, Japan, Korea, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Slovakia, Spain, Sweden, Switzerland, Turkey, United Kingdom and the United States.

performing loans to gross loans. Lastly, the level of per capita real GDP (in 2005 USD prices) by OECD country is also drawn from the Financial Development and Structure Database.

Table 1: Summary statistics

Variable	**Obs**	**Mean**	**Standard deviation**	**Min**	**Max**
		Dependent variables			
CO_2	1,399	9.083	5.087	1.230	40.59
NO_X	1,530	1.054	0.758	0.150	4.940
		Explanatory variables			
GDP	1,365	24,857	14,708	1,968	87,773
Credit	1,393	64.59	39.04	3.320	262.5
Stock	1,077	49.75	40.87	0.180	250.0
Bond	496	1.845	1.850	0	18.07
FININDEX	1,259	-0.585	0.927	-4.605	1.803

Where appropriate, interpolation was used in the case of missing values while moving average and single and double exponential smoothing techniques were applied to predict the missing values of the variables of interest for recent years of the time period considered. The choice of the appropriate method was determined with the help of measures of accuracy like Mean Absolute Percentage Error (MAPE), Mean Absolute Deviation (MAD), and Mean Squared Deviation (MSD).[3]

[3] The use of these statistics helps us to compare different forecasting fits and smoothing procedures with smaller values indicating a better fitting model.

3.1 The proposed econometric methods

Similarly to other empirical studies (see for example Millimet *et al.* 2003; Zarzoso and Morancho 2004; Apergis 2016), we first estimate separately the following (polynomial) panel data models in a static form[4]:

$$CO_{2it} = \alpha_i + \beta_t + b_0 + b_1 GDP_{it} + b_2 GDP_{it}^2 + b_3 GDP_{it}^3 + c_1 CREDIT_{it} + c_2 STOCK_{it} + c_3 BOND_{it} + e_{it} \quad (1)$$

$$NO_{Xit} = \alpha_i + \beta_t + b_0 + b_1 GDP_{it} + b_2 GDP_{it}^2 + b_3 GDP_{it}^3 + c_1 CREDIT_{it} + c_2 STOCK_{it} + c_3 BOND_{it} + e_{it} \quad (2)$$

i = 1,2,…34 and t = 1,2,…44

where CO_{2it} and NO_{Xit} are the per capita global and local pollutants in country i at time t; α_i and β_t are state and time fixed effects used in order to capture common factors across the cross-section element; GDP_{it} is real GDP per capita (powers) for country i at time t, CREDIT, STOCK and BOND have been defined above. Finally, e_{it} are zero mean i.i.d. errors.

The basic model of unobserved effects may be expressed as

$$y_{it} = X_{it}\beta + d_i + \varepsilon_{it} \qquad t = 1,2,...,T \quad (3)$$

where X_{it} is 1xK. The first method used is the fixed effects (FE) estimator allowing a different intercept for every country and treating the constants as regression parameters. Specifically as it will be shown next in the random effects specification d_i is put into the error term assuming d_i is orthogonal to X_{it} while in the fixed effects is allowed to be randomly correlated with X_{it} with (3) expressed as

$$\hat{\beta}_{FE} = \frac{\sum_{i=1}^{N}\left[(X_{it} - \bar{X}_i)'\left(Y_{it} - \bar{Y}_i\right)\right]}{\sum_{i=1}^{N}\left[(X_{it} - \bar{X}_i)'\left(X_{it} - \bar{X}_i\right)\right]} \quad (4)$$

[4] The degree of the polynomial for each equation has been determined by the maximum number of statistically significant powers. For example in the case of NO_X fourth and higher degree polynomial specifications have the extra powers of GDP to be not statistically significant.

And the feasible Generalized Least Squares (FGLS) is given as

$$\hat{\beta}_{FEGLS} = \frac{\sum_{i=1}^{N}\left[(X_{it}-\bar{X}_i)'\hat{\Omega}^{-1}\left(Y_{it}-\bar{Y}_i\right)\right]}{\sum_{i=1}^{N}\left[(X_{it}-\bar{X}_i)'\hat{\Omega}^{-1}\left(X_{it}-\bar{X}_i\right)\right]} \qquad with\ \ \hat{\Omega} = \frac{\sum\left[(\varepsilon_{it}-\bar{\varepsilon}_i)(\varepsilon_{it}-\bar{\varepsilon}_i)'\right]}{\mathrm{N}} \tag{5}$$

In the FE specification the within transformation for consistent estimators requires T to be large. When we refer to the full set of countries it may be logical to assume that the model is constant. But if the sampled cross-sections are derived from a large population and individual effects are strictly uncorrelated with the regressors it may be suitable to model the individual intercepts as randomly distributed across cross-sections (Greene, 2003).[5]

In the random effects analysis d_i is entered into the error term and (3) becomes

$$y_i = X_i\beta + u_i \tag{6}$$

With $u_i=d_i j_T+\varepsilon_i$ where j_T is Tx1 vector of ones. The unconditional variance of u_i may be expressed as TxT matrix of the form $\Omega = \mathrm{E}(u_i u_i')$ assuming it is positive definite. To apply an FGLS we form a TxT positive definite matrix of the form

$$\hat{\Omega} = \sigma_\varepsilon^2 \mathrm{I}_\mathrm{T} + \sigma_d^2 j_\mathrm{T} j_T' \tag{7}$$

With

$$\hat{\beta}_{RE} = \frac{\left(\sum_{i=1}^{N} X_i'\hat{\Omega}^{-1} y_\iota\right)}{\left(\sum_{i=1}^{N} X_i'\hat{\Omega}^{-1}\mathrm{X}_\iota\right)} \tag{8}$$

In the random effects (RE) individual effects are treated as random, and constants are components of the random disturbances. Both FE and RE are inefficient in the presence of heteroskedasticity (Greene, 2003, Baltagi, 2002) and to tackle heteroskedasticity and various

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[5] In our case, although we examine as a specific group the OECD countries, we still consider that these sampled cross-sections are derived from a larger population. As a consequence we rely on the RE model without considering the Hausman test when we account for the inconsistency of the RE estimates.

patterns of correlation between residuals, Generalized Least Squares (GLS) specifications may be used (Grossman and Krueger, 1995). We may also assume that the n disturbance terms at t, ε_t, follow a multinomial normal distribution with zero mean and nxn covariance matrix. The log likelihood function is given by

$$\ln(\beta,\Sigma \Big| n) = \frac{nT}{2}\ln 2\pi - \frac{T}{2}\ln|\Sigma| = \frac{1}{2}\sum_{t=1}^{T} u_t'\Sigma^{-1}u_t \quad (9)$$

With $u_{it} = y_{it} - X_{it}'\beta$ i=1,2,…,n and $\hat{\sigma}_{ij} = \frac{\hat{u}_t'\hat{u}_t}{T}$

The aforementioned analysis was performed in a static framework. With the intention to examine the robustness of our empirical findings permitting for dynamic aspects we use dynamic panel data techniques such as Difference Generalised Method of Moments (DIF-GMM) estimators attributed to Arellano and Bond, (1991) and System Generalised Method of Moments (SYS-GMM) estimators proposed by Arellano and Bover (1995) and Blundell and Bond (1998) respectively. The use of the latter is mainly justified as it improves significantly the estimates' accuracy and enlarges efficiency when the lagged dependent variables are considered as poor instruments as in the first-differenced regressors (Greene, 2003, Baltagi, 2002, Harrington et al, 2014).

In our case and in modelling dynamic effects we have the lagged dependent among the independent variables in the following form:

$$Y_{it} = X_{it}'\beta + \delta Y_{i,t-1} + \alpha_t + u_{it} \qquad i=1,2,\ldots,N\ ,\ t=1,2,\ldots,T \quad (10)$$

where δ being a scalar, X_{it}' a 1xK and β a Kx1 and u_{it} follow a one-way error component model ($u_{it}=\mu_i+v_{it}$); with $\mu_i \sim IID(0,\sigma_\mu^2)$ and $v_{it} \sim IID(0,\sigma_v^2)$ independent both of each other and between them. Then the first difference GMM estimation is given as

$$\hat{\delta}_{GMM} = \frac{\left(\sum_{i=1}^{N} \Delta y_{i,t-1} Z_i\right) W_N \left(\sum_{i=1}^{N} Z_i' \Delta y_i\right)}{\left(\sum_{i=1}^{N} \Delta y_{i,t-1} Z_i\right) W_N \left(\sum_{i=1}^{N} Z_i' \Delta y_{i,t-1}\right)} \tag{11}$$

With the choice of W_N being important with the first-step consistent estimator of d being

$$W_N^* = \frac{1}{\left(\frac{1}{N} \sum_{i=1}^{N} Z_i' \Delta \hat{\varepsilon}_i Z_t\right)} \tag{12}$$

In (3) if X_{it} are predetermined with current and lagged X_{it}s uncorrelated with current term then $E(X_{ij}u_{is})=0$ for s≥t. A combination of strictly exogenous and predetermined X variables may more realistic compared to the two extreme cases with matrix Z_i adjusted according to each case. Arelano and Bover (1995) integrated this approach with the instrumental variables of Hansen and Taylor (1981) with individual series being highly persistent and δ being near to one. In such circumstances FD-GMM may present finite sample biases as the instruments are weak (Verbeek, 2012). Estimators using moment conditions relying on levels and FD refer to system GMM.

Based on the above, the dynamic specifications of all the three models are given by the following reduced form equations:

$$CO_{2it} = \alpha_i + \beta_t + b_0 + \sum_{l=1}^{L} d_{l,i} CO_{2it-l} + \sum_{m=0}^{M} b_1 GDP_{it-m} + b_2 GDP^2{}_{it} + b_3 GDP^3{}_{it} + c_1 CREDIT_{it} + c_2 STOCK_{it} + c_3 BOND_{it} + e_{it} \tag{13}$$

$$NO_{Xit} = \alpha_i + \beta_t + b_0 + \sum_{l=1}^{L} d_{l,i} NO_{Xit-l} + \sum_{m=0}^{M} b_1 GDP_{it-m} + b_2 GDP^2{}_{it} + b_3 GDP^3{}_{it} + c_1 CREDIT_{it} + c_2 STOCK_{it} + c_3 BOND_{it} + e_{it} \tag{14}$$

i = 1,2,…34, t = 1,2,…44 and l is the time lag operator for the dependent variable.

4. Results and discussion

4.1 Cross-Section Dependence

One of the additional complications that arise when dealing with panel data compared to the pure time-series case, is the possibility that the variables or the random disturbances are correlated across the panel dimension. The early literature on unit root and cointegration tests adopted the assumption of no cross-sectional dependence. However, it is common for macro-level data to violate this assumption which will result in low power and size distortions of tests that assume cross-section independence. For example, cross-section dependence in our data may arise due to common unobserved effects due to changes in federal legislation. Therefore, before proceeding to unit root and cointegration tests we test for cross-section dependence. We use the cross-section dependence tests proposed by Pesaran (2004). The tests are based on the estimation of the linear panel model of the form:

$$y_{it} = \alpha_i + \beta_i' x_{it} + u_{it}, \qquad i = 1,...,N; t = 1,...,T \tag{15}$$

where T and N are the time and panel dimensions respectively, α_i the provincial-specific intercept, and x_{it} a $kx1$ vector of regressors, and u_{it} the random disturbance term. The null hypothesis in both tests assumes the existence of cross-section correlation: $Cov(u_{it}, u_{jt}) = 0$ for all t and for all $i \neq j$. This is tested against the alternative hypothesis that $Cov(u_{it}, u_{jt}) \neq 0$ for at least one pair of i and j. The Pesaran (2004) tests are a type of Lagrange-Multiplier test that is based on the errors obtained from estimating Equation 20 by the OLS method.

Based on the above, we carry out the first part of the empirical analysis by examining the presence of cross-section dependence. We use the cross-section dependence tests proposed by Breusch and Pagan (1980) and Pesaran (2004). Both tests strongly reject the null hypothesis of cross-section independence (P-value = 0.000) for all the models, providing evidence of cross-sectional dependence in the data given the statistical significance of the CD statistics (see Table 2). In face of this evidence we proceed to test for unit roots using tests that are robust to cross-section dependence (the so-called "*Second Generation*" tests for unit roots in panel data).

Table 2: Cross-section dependence (Pesaran CD test)

Variable	CD test	P-value	Correlation	Absolute (correlation)
CO_2	11.20^{***}	0.000	0.076	0.546
NO_X	85.13^{***}	0.000	0.536	0.682
GDP	132.95^{***}	0.000	0.942	0.942
GDP^2	131.44^{***}	0.000	0.930	0.930
GDP^3	129.62^{***}	0.000	0.917	0.917
Credit	63.29^{***}	0.000	0.428	0.571
Stock	63.09^{***}	0.000	0.497	0.543
Bond	30.22^{***}	0.000	0.330	0.402
FININDEX	29.59^{***}	0.000	0.241	0.413

Note: Under the null hypothesis of cross-sectional independence the CD statistic is distributed as a two-tailed standard normal. Results are based on the test of Pesaran (2004). The p-values are for a one-sided test based on the normal distribution. Correlation and abs(correlation) are the average (absolute) value of the off-diagonal elements of the cross-sectional correlation matrix of residuals obtained from estimating Eq. (1-3). Significant at ***1%.

4.2 Unit Root and cointegration testing

To examine the stationarity properties of the variables in our models we use the "*second generation*" unit root tests for panel-data. The unit root testing methodology allows for non-linear functions of the I(1) variables, which is really the case here as GDP enters both in levels and in quadratic and cubed form (Apergis, 2016). To the end of the unit root testing, the empirical analysis makes use of the Fisher test as proposed by Maddala and Wu (1999) and developed by Kyung *et al*, (2003). This test explicitly considers cross-sectional dependency in unbalanced panel data set. More specifically, this methodological approach is based on the p-values of individual unit root tests and assumes that all series are non-stationary under the null hypothesis against the alternative that at least one series in the panel is stationary. Unlike the Im-Pesaran-Shin (1997) test, Fisher's test does not require a balanced panel. The results are reported In Table 3 and they support the presence of a unit root across all three variables. In other words, the test results suggest that no variables are integrated of an order greater than one (I-1).[6]

[6] The test was carried out in STATA using the "*xtfisher*" routine.

Table 3: Fisher panel unit root tests

Variable	Statistic			
	P	Z	L*	Pm
Levels				
CO_2	54.5809 (0.8807)	1.8122 (0.9650)	2.0399 (0.9786)	-1.1507 (0.8751)
NOx	82.5300 (0.1106)	-1.4264* (0.0769)	-1.4060* (0.0807)	1.2459 (0.1064)
GDP	40.1424 (0.9972)	2.3342 (0.9902)	2.3140 (0.9891)	-2.3888 (0.9915)
GDP^2	33.9751 (0.9998)	3.0764 (0.9990)	3.0231 (0.9986)	-2.9176 (0.9982)
GDP^3	30.0662 (1.0000)	3.1197 (0.9991)	3.0910 (0.9988)	-3.2528 (0.9994)
CREDIT	47.5849 (0.9717)	1.7541 (0.9603)	1.8680 (0.9683)	-1.7506 (0.9600)
STOCK	34.7136 (0.9997)	2.8141 (0.9976)	2.7028 (0.9962)	-2.8543 (0.9978)
BOND	59.1352 (0.7698)	1.5266 (0.9366)	1.3536 (0.9111)	-0.7602 (0.7764)
FININDEX	89.283** (0.0445)	-0.1578 (0.4373)	-0.6263 (0.2660)	1.8032 (0.0357)
First Differences				
$\Delta(CO_2)$	274.0547*** (0.0000)	-11.1093*** (0.0000)	-12.4977*** (0.0000)	17.6690*** (0.0000)
Δ(NOx)	330.6820*** (0.0000)	-13.1450*** (0.0000)	-15.4935*** (0.0000)	22.5248*** (0.0000)
Δ(GDP)	190.9629*** (0.0000)	-7.6048*** (0.0000)	-8.2645*** (0.0000)	10.5440*** (0.0000)
$\Delta(GDP)^2$	201.6362*** (0.000)	-7.4326*** (0.000)	-8.6094*** (0.0000)	-11.4592*** (0.0000)
$\Delta(GDP)^3$	201.3489*** (0.000)	-7.2958*** (0.000)	-8.5385*** (0.0000)	-11.4346*** (0.0000)
Δ(CREDIT)	114.5813*** (0.0004)	-3.4850*** (0.0002)	-3.4373*** (0.0004)	3.9943*** (0.0000)
Δ(STOCK)	131.0578*** (0.0000)	-5.1214*** (0.0000)	-5.0516*** (0.0000)	5.4072*** (0.0000)
Δ(BOND)	445.7075*** (0.0000)	-13.6313*** (0.0000)	-20.3759*** (0.0000)	-32.3882*** (0.0000)
Δ(FININDEX)	227.9565*** (0.0000)	-7.0913*** (0.0000)	-9.4297*** (0.0000)	13.7162*** (0.0000)

Note: The number of lags has been set to two according to BIC. The statistics are: P, is the inverse chi-squared statistic, Z is the inverse normal statistic, L^* denotes the inverse logit statistic, while P_m stands for the modified inversed chi-squared statistic. The Augmented Dickey Fuller test is used rather than Phillips-Perron test. The null hypothesis assumes that the variable contains unit root. The numbers in parentheses denote the p-values. Significant at ***1% and *10% respectively.

In order to investigate whether a long-run equilibrium relationship exists among the variables in our three models we implement two cointegration tests proposed by Westerlund (2007) that allow for cross-section dependence and rely on the assumption of weakly exogenous regressors (Demetriades and James, 2011). In general, the tests are an error-correction approach to testing for cointegration that is based on the statistical significance of the error correction term. The intuition behind this approach is that if a long run relationship between the variables in our model, we can write a regression that allows us to estimate the error-correcting terms which reflect the response of the system to random shocks that "pushes" the system towards its long-run equilibrium point. If the error-correction terms are significantly different from zero across sections, then there is evidence in favor of the existence of a long-run relation. The null hypothesis in both tests is that of no cointegration.

The test statistics of the first two tests, denoted G_t, are general enough to allow for individual-specific intercepts and short-run dynamics and is constructed as a weighted average of the estimated error-correcting coefficients across each province in our model. The alternative hypothesis in this test of tests is that at least one section in the panel is cointegrated. The second test assumes that the intercept is the same across sections and tests against the alternative hypothesis that the panel is cointegrated as a whole. The test statistic is denoted by p_t. The results of the tests are presented in the next tables; the critical values were created using a bootstrapping method. The results indicate that the first test rejects the null hypothesis of no cointegration for all three models. However, the second test that restricts the intercept to be the same across all provinces fails to reject the null.[7]

[7] The tests can be carried out in STATA using the "*xtwest*" routine. It should be noted that the results are sensitive to the selection of the lag structure of the model. Persyn and Westerlund (2008) point out that this sensitivity might occur in small datasets.

Table 4: Westerlund ECM panel cointegration tests

Equation	Statistic			
	G_τ	G_α	P_τ	P_α
CO_2 = f (GDP)	-2.603** (0.037)	-11.211 (0.727)	-14.525*** (0.005)	-12.690*** (0.000)
CO_2 = f (GDP)^2	-2.738*** (0.003)	-11.328 (0.691)	-17.195*** (0.000)	-14.914*** (0.000)
CO_2 = f (GDP)^3	-2.738*** (0.003)	-11.328 (0.691)	-17.195*** (0.000)	-14.914*** (0.000)
CO_2 = f (Credit)	-2.798*** (0.001)	-15.772*** (0.000)	-16.684*** (0.000)	-14.608*** (0.000)
CO_2 = f (Stock)	-2.766*** (0.002)	-13.619*** (0.066)	-15.426*** (0.000)	-11.306*** (0.011)
CO_2 = f (Bond)	-2.767*** (0.000)	-12.324*** (0.000)	-15.098*** (0.000)	-12.435*** (0.002)
CO_2 = f (FININDEX)	-2.470*** (0.000)	-7.616 (0.328)	-14.728*** (0.000)	-7.956*** (0.000)
NO_X = f (GDP)	-3.203*** (0.000)	-14.590*** (0.009)	-18.539*** (0.000)	-13.855*** (0.000)
NO_X = f (GDP)^2	-3.261*** (0.000)	-15.506*** (0.001)	-20.211*** (0.000)	-15.910*** (0.000)
NO_X = f (GDP)^3	-3.282*** (0.000)	-16.281*** (0.000)	-21.045*** (0.000)	-16.891*** (0.000)
NO_X = f (Credit)	-3.326*** (0.000)	-17.149*** (0.000)	-18.108*** (0.000)	-16.255*** (0.000)
NO_X = f (Stock)	-3.187*** (0.000)	-17.061*** (0.000)	-16.947*** (0.000)	-14.757*** (0.000)
NO_X = f (Bond)	-3.556*** (0.000)	-17.467*** (0.000)	-17.098*** (0.000)	-15.125*** (0.000)
NO_X = f (FININDEX)	-3.045*** (0.000)	-14.570*** (0.010)	-14.955*** (0.001)	-11.438*** (0.008)

Note: The test regression was fitted with a constant and trend and one lag and lead. The kernel bandwidth was set according to the rule (Demetriades and James, 2011). The null hypothesis assumes that there is no co-integration. The numbers in parentheses denote the p-values. Significant at ***1% and **5% respectively.

4.3 *Empirical findings*

Using simple OLS to estimate the cointegrating relation will lead to bias in the estimated coefficients unless all of the explanatory variables are strongly exogenous. Furthermore, other OLS estimators that remove the endogeneity bias such as the Fully-Modified OLS (Pedroni, 2000) or the Dynamic OLS (Kao and Chiang, 2000) are inadequate for our data since they assume cross-section independence. As Pesaran and Smith (1995) point out, other traditional methods for estimating pooled models such as the Fixed Effects and the Instrumental Variables estimators proposed by Arellano and Bond (1991) "*can produce very misleading estimates of the average values of the parameters in dynamic panel data models unless the slope coefficients are in fact identical*". Furthermore, the Arellano and Bond (1991) method performs well for $N > T$ which is the case in our data.[8]

As mentioned our intention is to explore the effect of financial development and economic growth on environmental degradation. The stock and bond markets differ in the risk involved in investing in both. Investing in bond markets may be less risky in comparison to stock markets as the former is less volatile. At the same time bond prices fluctuate with changes in market sentiments and in different economic circumstances in a significantly different way and from different factors compared to stocks. Different factors like interest rates and economic motivation policies have an influence on both stocks and bonds with opposite reactions. If stocks are in an increasing trend, investors may move away from bonds and towards the booming stock market while if stock markets stabilize or severe economic problems arise, investors return to the safety of bonds. In our case and in all specifications the signs of credit and bond are negative. Stock has the lowest magnitude in all cases and an opposite effect in comparison to bond.

Table 5 presents the results from the static and dynamic model formulations for the case of the pollutants considered. An N-shape relationship is observed in the static analysis for both pollutants while in the dynamic analysis we still observe an N shape for the global pollutant and a monotonic relation for the local pollutant (see Figures 1 and 2).

[8] The DIF-GMM and SYS-GMM estimators were performed by using the STATA command "*xtabond2*" (Roodman, 2009).

Table 5: Empirical results

	Static results				Dynamic results		
Control variables	**CO_2 Model (MLE)**	**CO_2 Model (GLS)**	**NO_X Model (MLE)**	**NO_X Model (GLS)**	**CO_2 Model$^+$ (DIF-GMM)**	**CO_2 Model (SYS-GMM)**	**NO_X Model$^+$ (DIF-GMM)**
CO_2(-1)	-	-	-	-	0.719*** (0.0509)	0.899*** (0.0176)	-
NOx (-1)	-	-	-	-	-	-	0.314** (0.132)
GDP	0.0004539*** (0.0000697)	0.0004583*** (0.0000703)	0.0000318*** (.0000122)	0.0000349*** (0.0000123)	-0.000512** (0.000244)	0.00018*** (6.51e-05)	-1.23e-05*** (3.70e-06)
GDP^2	-1.06e-08*** (1.71e-09)	-1.05e-08*** (1.73e-09)	-1.47e-09*** (2.95e-10)	-1.45e-09*** (3.02e-10)	1.24e-08* (6.52e-09)	-4.75e-09*** (1.73e-09)	-
GDP^3	7.26e-14*** (1.23e-14)	7.20e-14*** (1.25e-14)	1.07e-14*** (2.12e-15)	1.04e-14*** (2.17e-15)	-8.71e-14* (5.51e-14)	3.47e-14*** (1.23e-14)	-
CREDIT	-0.0094971*** (0.0019115)	-0.0098731 *** (0. 0019263)	-0.0018759*** (0.0003285)	-0.0019601*** (0.0003348)	-0.00639*** (0.00213)	-0.00405** (0.00163)	-0.00292** (0.00139)
STOCK	0.0069834*** (0.0015389)	0.0070643 *** (0.0015694)	0.0009636*** (0.0002638)	0.0009793*** (0.0002715)	0.00248 (0.00215)	0.00370* (0.00191)	0.00118** (0.000565)
BOND	-0.179284*** (0.0303499)	-.1771706*** (0.0309601)	-0.0193849*** (0.0051956)	-0.019343*** (0.0053523)	-	-	-
Constant	4.683581*** (1.0767)	4.492349*** (1.011429)	1.174816*** (0.2150461)	1.077573*** (0.190714)	0.719*** (0.0509)	-0.739 (0.543)	-
				Diagnostics			
Observations	495	495	495	495	1,026	1,060	1,043
Turning points	65,539 31,798	64,144 33,078	79,058 12,531	78,743 14,206	64,593 30,318	64,364 26,894	-
Shape of curve	N- shape	N- shape	N- shape	N shape	Inverted N-shape	N- shape	Monotonically decreasing
LR-test/ R-squared	126.09*** [0.000]	0.227	158.65*** [0.000]	0.320	-	4329.55*** [0.000]	-
F-test	-	-	-	-	105.78*** [0.000]	-	37.75*** [0.000]
AR(1)	-	-	-	-	-3.73*** [0.000]	-3.32*** [0.001]	-1.78 [0.074]
AR(2)	-	-	-	-	1.00 [0.319]	1.41 [0.160]	-0.33 [0.741]
Hansen test	-	-	-	-	29.68 [1.000]	28.49 [1.000]	32.16 [1.000]

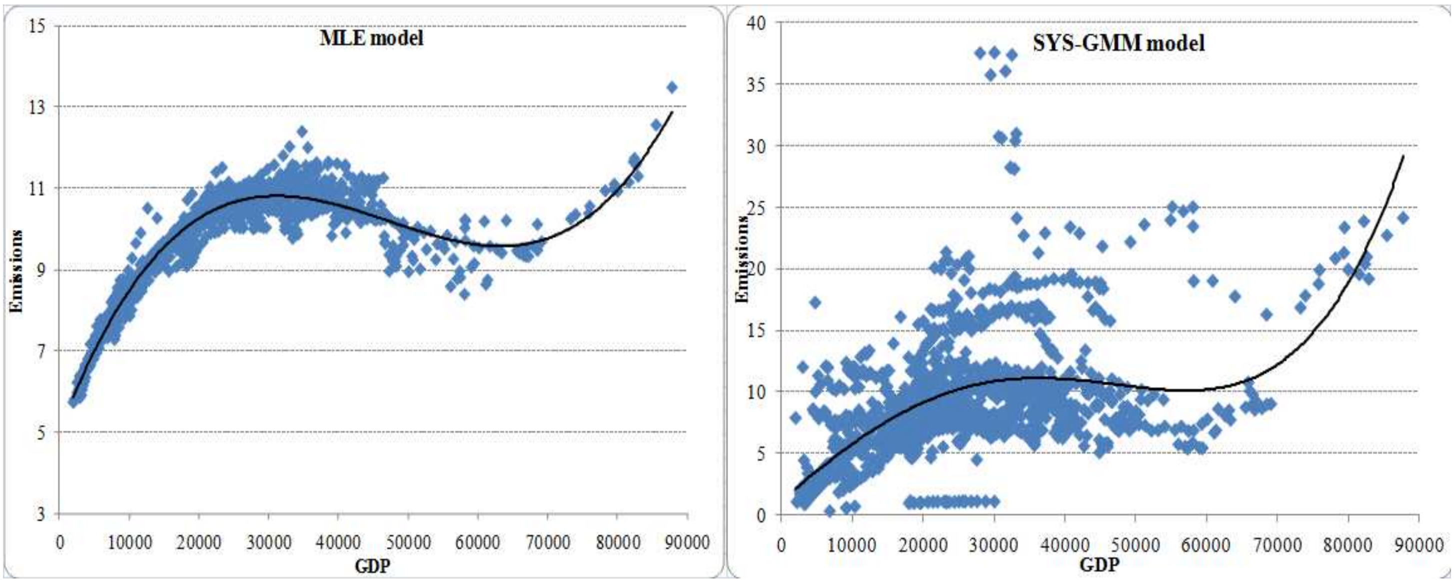
Note: (+) The one step estimators are reported. MLE denotes the GLS maximum likelihood estimator, GLS denotes the random effects estimator, SYS-GMM is the system GMM estimator and DIF-GMM denotes the difference GMM estimator. Robust standard errors are in parentheses. The numbers in square brackets denote the p-values. The choice of the fixed effects estimators (FE) were based on the Hausman test. AR(1) and AR(2) are tests for first and second order serial autocorrelation. LR test denotes the joint statistical significance of all the covariates. Hansen denotes the test of over identifying restrictions of the instruments. Significant at ***1%, **5% and *10% respectively. The estimated peaks and lows are in constant US dollars at 2005 prices.

Concerning the static specifications in all cases and for both pollutants all explanatory variables are statistically significant and properly signed in all levels of significance. The calculated turning points are all within the sample with the upwards estimated points ranging from 64,144 to 65,539 in the case of CO_2 and 78,743 to 79,058 in the case of NO_x respectively. In the case of CO_2 the downwards estimated turning points range from 31,798 to 31,078 and 12,531 to 14,206 in the case of NO_X respectively. The magnitude of the stock and credit are quite low while the one for the bond is much higher and equal to -0.18 approximately.

Looking in the dynamic model specifications and especially in the DIF-GMM case for CO_2 GDP and its powers are statistically significant and only the financial variable Credit is significant with the extraction of an inverted N shape relationship with turning points within the sample and equal to 30,318 and 64,593 respectively. Similarly, in the case of System-GMM for CO_2 GDP and its powers are statistically significant in all significance levels and now Credit is significant at 5% and Stock at 10% significance levels with the extraction of an N-shape relationship with turning points within the sample ranging from 26,894 to 64,364.

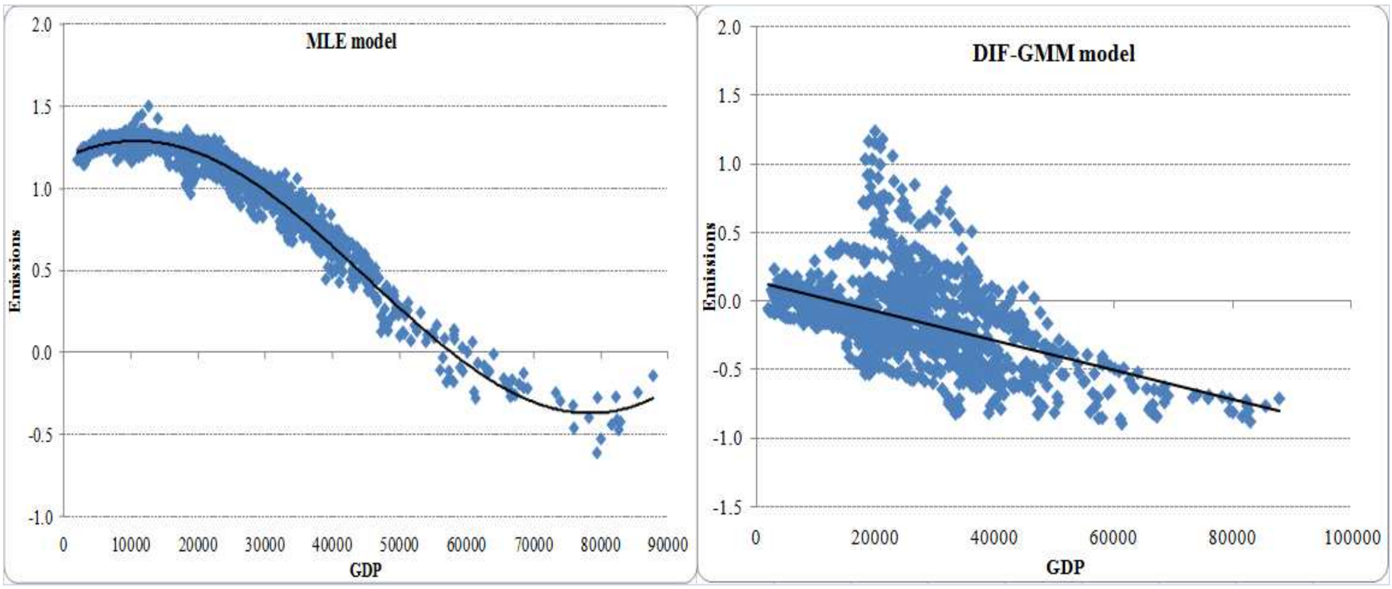
Finally in the case of DIF-GMM for NO_X GDP is statistically significant and Credit and Stock are significant at 5% significance level with a monotonically decreasing relationship of small magnitude. The adjustment coefficients are quite low in the case of CO_2 equal to 0.28 and 0.10 in the cases of difference and system GMM respectively and higher in the case of NO_X (0.69 approximately).

Figure 1: Influence of economic growth on CO_2 emissions



Note: MLE and SYS-GMM denote the (static) GLS maximum likelihood and (dynamic) System GMM model respectively. The graphs were constructed by multiplying GDP, GDP squared, GDP cubed by the sum of the estimated coefficients for current GDP. We normalized by adding to this the value of all the other banking development variables (CREDIT, STOCK and BOND) multiplied by their estimated corresponding coefficients and the constant term.

Figure 2: Influence of economic growth on NO_X emissions



Note: MLE and DIF-GMM denote the (static) GLS maximum likelihood and (dynamic) Difference GMM model respectively.The graphs were constructed by multiplying GDP, GDP squared, GDP cubed by the sum of the estimated coefficients for current GDP. We normalized by adding to this the value of all the other banking development variables (CREDIT, STOCK and BOND) multiplied by their estimated corresponding coefficients and the constant term.

4.4. Sensitivity analysis

In order to check for the validity and robustness of our findings, we re-estimate our (static and dynamic) models of the main determinants of two major pollutants (CO_2 and NO_X emissions) on the basis of a different approximation. More specifically, we used the three banking variables (CREDIT, STOCK and BOND) to form a financial index, denoted as FININDEX.[9] The latter follows closely the spirit of Beck and Levine (2002), Luintel *et al.* (2008) and Antzoulatos *et al.* (2011) who constructed a composite index by taking the log ratio of the sum of stock market and private bond market capitalization divided by private bank credit. The main reason we incorporated the private bond market capitalization in the numerator of the FININDEX is attributed to the fact that the aforementioned variable constitutes a major segment of the financial system in many countries (Beck *et al.*, 2001). An increase (decrease) of this index indicates a development (recession) of capital markets relatively to the development of banks. This may increase (decrease) the level of environmental emissions in a country, which may lead to environmental degradation justifying a call for further actions by policy makers and government officials (i.e. taxes, subsidies, tradable permits, etc).

From the empirical results, it is evident that the main conclusions drawn in the previous section targeted at the shape of the relationship between economic growth and financial development on pollution remain robust (see Table 6). More specifically, regarding the global pollutant (CO_2 emissions), we argue that in the static specifications (see columns 1 and 2) an N-shape form is depicted with very similar turning points ranging from 31,024 to 80,560 US$. The impact of financial development approximated by the FININDEX variable on environmental

[9] Other possible indexes may be market capitalization as a percentage of GDP, turnover ratio or liquid liabilities or domestic credit provided by the banking sector or the value of share trade, each one as share of GDP. According to Tyavambiza and Nyangara (2015) all these measures may provide biased results as they are highly correlated and inappropriate to capture financial sector's development potentials.

degradation as expressed by the level of CO_2 emissions is positive with its magnitude (elasticity) equal to 0.021 in both specifications. This means that if FININDEX shows an increase (decrease) by 100% the level of per capita global pollutant will be increased (decreased) by 2.1%.

Table 6: Sensitivity analysis results

	Static results				Dynamic results		
Control variables	**(1) CO_2 Model (MLE)**	**(2) CO_2 Model (GLS)**	**(3) NO_X Model (MLE)**	**(4) NO_X Model (GLS)**	**(5) CO_2 Model$^+$ (DIF-GMM)**	**(6) CO_2 Model (SYS-GMM)**	**(7) NO_X Model (DIF-GMM)**
CO_2(-1)	-	-	-	-	0.781*** (0.0378)	0.786*** (0.0385)	-
NOx (-1)	-	-	-	-	-	-	0.491*** (0.0670)
GDP	0.000529*** (3.91e-05)	0.000530*** (3.93e-05)	1.62e-05** (6.38e-06)	1.72e-05*** (6.40e-06)	-0.000460** (0.000191)	-0.000455** (0.000196)	-1.58e-05*** (5.12e-06)
GDP^2	-1.18e-08*** (9.71e-10)	-1.18e-08*** (9.77e-10)	-1.18e-09*** (1.58e-10)	-1.20e-09*** (1.59e-10)	1.05e-08** (4.98e-09)	1.04e-08** (5.11e-09)	-
GDP^3	7.05e-14*** (7.24e-15)	7.07e-14*** (7.30e-15)	1.07e-14*** (1.18e-15)	1.08e-14*** (1.19e-15)	-7.07e-14* (4.13e-14)	-7.04e-14* (4.22e-14)	-
FININDEX	0.194*** (0.0522)	0.193*** (0.0527)	0.0219** (0.00850)	0.0213** (0.00857)	0.208** (0.0811)	0.202** (0.0807)	0.0195* (0.0105)
Constant	3.512*** (0.834)	3.485*** (0.745)	1.215*** (0.157)	1.203*** (0.138)	-	-	-
Diagnostics							
Observations	986	986	988	988	949	949	954
Turning points	31,024 80,560	31,214 80,054	7,663 65,857	8,039 66,035	66,296 32,714	65,688 32,797	-
Shape of curve	N shape	N shape	N shape	N shape	Inverted N shape	Inverted N shape	Monotonically decreasing
LR-test/ R-squared	280.53*** [0.000]	0.252	439.99*** [0.000]	0.376	-	-	
F-test/Wald test	-	-	-	-	148.15*** [0.000]	682.94*** [0.000]	647.90*** [0.000]
AR(1)	-	-	-	-	-3.65***[0.000]	-3.40 [0.001]	-1.33* [0.098]
AR(2)	-	-	-	-	-0.77 [0.439]	-0.71 [0.476]	-0.34 [0.731]
Hansen test	-	-	-	-	32.70 [1.000]	32.70 [1.000]	30.36*** [0.000]

Note: $^{(+)}$ The one step estimators are reported. MLE denotes the GLS maximum likelihood estimator, GLS denotes the random effects estimator, SYS-GMM is the system GMM estimator and DIF-GMM denotes the difference GMM estimator. Robust standard errors are in parentheses. The numbers in square brackets denote the p-values. AR(1) and AR(2) are tests for first and second order serial autocorrelation. LR and Wald tests denote the joint statistical significance of all the covariates. Hansen denotes the test of over identifying restrictions of the instruments. Significant at ***1%, **5% and *10% respectively. The estimated peaks and lows are in constant US dollars at 2005 prices.

The same form of N-shaped is also evident in the case of local pollutant (NO_X emissions). Similarly, if financial and capital markets in the OECD countries show an increase (decrease) by about 100% the level of production and the subsequent NO_X concentrations emitted in the atmosphere will show a small increase (decrease) equal to 2% approximately.

Table 6 also presents the dynamic results obtained by this estimation process. It is worthwhile to mention that nearly all estimates are statistically significant with the appropriate sign. All underlying estimated equations pass a battery of diagnostic tests. Specifically, the instrument rank is greater than the number of estimated coefficients, while the reported Hansen test indicates that the instrument list satisfies the orthogonallity conditions in all of the three specifications, since the null hypothesis that the over-identifying restrictions are valid cannot be rejected. Regarding the shape of the relationship in each of the three specifications, it is interesting to mention that the CO_2 models reveal a stable inverted N shape relationship in contrast to the NO_X model where a monotonically linear decreasing approximation is evident (negative estimate of the GDP).

The most prominent outcome is the derivation of the long-run effect of the FININDEX. As suggested by Polemis (2016) one of the main reasons of estimating a dynamic model is to capture short-run and long-run effects. From the empirical findings, it is evident that, the long run effect shows a different pattern on each of the two major pollutants. For example, in the case of the CO_2 models (see Table 6 columns 5-7) this effect is almost five times greater than the short run effect of the first period based on Models (5) and (6) respectively, denoting that the level of CO_2 emissions will increase substantially in the long run (0.208 and 0.202 compared to 0.984 and 0.982 respectively). In contrast, this result is reversed in the case of NO_X, where the long run effect of financial development is also positive but is almost twenty five times greater in its magnitude than the short run effect (0.0195 compared to 0.500). This means that financial integration causes some transient disruption to the level of NO_X emissions only in the short-run, whilst the opposite holds in the long run. The different long-run response rate between CO_2 and NO_X emissions (0.982-0.984 for CO_2 and 0.500 for NO_X) may be attributed to the geographical boundaries of the scrutinized pollutant (global vs local pollutant). In other words, CO_2 as one of the main greenhouse gases responsible for the "*global warming*" is a major global pollutant affecting all of the OECD countries it is more likely to demand faster response to financial evolution than NO_X emissions acting at a local level (country or regions). Finally, it is worth mentioning that, the long-run effect of the FININDEX on the level of CO_2 emissions is almost two times greater than in the case of local pollutant (NO_X emissions).

5. Conclusions and policy implications

In our analysis the relationship and the synergies between financial development and economic growth and environmental damage in the form of global (CO_2 emissions) and local (NO_X emissions) pollutants both in static and dynamic formulations was investigated. To explore these possible relationships we have considered the financial aspects both individually as well as an index. From our empirical findings it is clear that the shape of the relationship between growth and financial development on environmental degradation remain robust.

An N-shape relationship is observed in the static analysis for both pollutants while in the dynamic analysis we have an N-shape for the global pollutant but a monotonic relation in the case of the local pollutant. The calculated turning points are in all cases within the sample. The adjustment coefficients are very low for CO_2 emissions ranging from 0.10-0.28 and very high in the case of NO_X emissions and equal to almost 0.7. In all specifications the signs of credit and bond are negative with stock having the smallest magnitude and an opposite influence compared to bond. The effect of financial development using as proxy the constructed financial index (FININDEX) on pollution in the form of CO_2 emissions is positive with elasticity equal to 0.021.

The long-run effect presents a different picture for the two pollutants under consideration. For CO_2 the effect is five times larger compared to the short-run effect of the first period, showing that LR the CO_2 emissions will substantially change. On the contrary, for NO_X emissions LR effect of financial development is again positive but much higher (almost twenty five times greater in magnitude than the SR effect). This implies that financial integration may create transient disruption to NO_X emissions only SR, with the opposite being the case in LR. This in turn shows that financial development in the banking sector is the cause of CO_2 and NO_X emissions with the financial development to facilitate the installation of new more advanced and more cost-effective and energy efficient abatement methods as financial assistance may be obtained at lower cost[10].

Finally, allocated financial resources have to ensure that credit will facilitate firms but not in the cost of environmental degradation. Policy makers and government officials have to stimulate investments in productive sectors like the energy sector and more likely to promote the use of Renewable Energy Sources (RES). This can be accompanied by more financial resources for Research and Development (R&D) and more cost effective mitigation methods.

[10] The important factors for differences among countries in mitigation costs are discussed in Halkos (2010).

References

Antzoulatos A., Apergis N., and Tsoumas C. (2011). Financial Structure And Industrial Structure. *Bulletin of Economic Research*, **63(2)**: 109-139.

Antweiler W., Copeland B. and Taylor M.S. (2001). Is free trade good for the environment? *The American Economic Review,* **91(4):** 877–908.

Apergis N. (2016). Environmental Kuznets curves: New evidence on both panel and country-level CO_2 emissions. *Energy Economics,* **54**: 263–271.

Apergis N., Christou C. and Gupta R. (2014). Are there Environmental Kuznets Curves for US State-Level CO_2 Emissions?, Working Papers **201474**, University of Pretoria, Department of Economics.

Arellano M. and Bond S. (1991). Some Tests of Specification for Panel Data: Monte Carlo Evidence and an Application to Employment Equations. *Review of Economic Studies*, **58(2)**: 277-297.

Arellano M. and Bover O. (1995). Another look at the instrumental variable estimation of error component models. *Journal of Econometrics*, **68**: 29–51.

Baltagi B.H. (ed.) (2002). Recent Developments in the Econometrics of Panel Data, Edward Elgar Publishing, volume 0, number 2701, 10.

Beck T., Demirguc-Kunt A. and Levine R. (2000). A new database on financial development and structure. *World Bank Economic Review*, **14**: 597–605.

Beck T. and Levine R. (2002). Industry growth and capital allocation: does having a market or bank-based system matter?, *Journal of Financial Economics*, **64**: 147–80.

Beck T., Demirguc-Kunt A. and Levine R. (2001). The financial structure database. In: Demirguc-Kunt A. and Levine R. (eds). *Financial Structure and Economic Growth: A Cross-Country Comparison of Banks, Markets, and Development.* Cambridge, MA: MIT: 17–80.

Blundell R. and Bond S. (1998). Initial conditions and moment restrictions in dynamic panel data models. *Journal of Econometrics,* **87**: 115–43.

Blundell R. and Bond S. (2000). GMM estimation with persistent panel data: an application to production functions. *Econometric Reviews,* **19 (3):** 321-340.

Breitung J. and Das S. (2005). Panel unit root tests under cross-sectional dependence. *Statistica Neerlandica*, **59**: 414-433.

Cole M.A. (2004). Trade, the pollution haven hypothesis and the environmental Kuznets curve: Examining the linkages. *Ecological Economics,* **48:** 71–81.

Cole M.A. and Elliot R.J.R. (2005). FDI and the capital intensity of "dirty" sectors: A missing piece of the pollution haven puzzle. *Review of Development Economic Literature*, **42(1):** 7-71.

Dean J.M. (2002). Does trade liberalization harm the environment? A new test. *Canadian Journal of Economics*, **35(4):** 819–842.

Demetriades P. and James G. (2011). Finance and growth in Africa: The broken link. *Economics Letters,* **113**: 263–265.

Frankel J. and Romer D. (1999). Does trade cause growth? *American Economic Review*, **89(3):** 379–399.

Greene W.H. (2003). *Econometric Analysis.* International Edition. 5th Ed. Prentice-Hall.

Grossman G.M. and Krueger A.B. (1995). Economic growth and the environment. *Quarterly Journal of Economics,* **110(2)**: 353–377.

Gutierrez L.H. (2003). The Effect of Endogenous Regulation on Telecommunications Expansion and Efficiency in Latin America. *Journal of Regulatory Economics,* **233**: 257-286.
Halkos G. (2003). Environmental Kuznets Curve for sulfur: evidence using GMM estimation and random coefficient panel data models. *Environment and Development Economics*, **8(4):** 581-601.
Halkos G. (2010). Construction of abatement cost curves: The case of F-gases," MPRA Paper 26532, University Library of Munich, Germany.
Halkos G. (2013). Exploring the economy-environment relationship in the case of sulphur emissions. *Journal of Environmental Planning and Management*, **56(2)**: 159-177.
Halkos G. and Tzeremes N.G (2013). Economic growth and environmental efficiency: Evidence from US regions. *Economics Letters*, **20(1)**: 48-52.
Halkos G. and Tsionas E.G. (2001). Environmental Kuznets curves: Bayesian evidence from switching regime models, *Energy Economics*, **23(2)**: 191-210.
Halkos G. and Sepetis A. (2007). Can capital markets respond to environmental policy of firms? Evidence from Greece. *Ecological Economics*, **63(2-3)**: 578-587.
Halkos G. and Tzeremes N.G. (2009a). Exploring the existence of Kuznets curve in countries' environmental efficiency using DEA window analysis. *Ecological Economics*, **68:** 2168–2176.
Halkos G. and Tzeremes N. (2009b). Economic efficiency and growth in the EU enlargement. *Journal of Policy Modeling*, **31(6):** 847-862.
Halkos G. and Tzeremes N. (2010). Measuring regional economic efficiency: the case of Greek prefectures. *The Annals of Regional Science,* **45(3):** 603-632.
Halkos G. and Polemis M. (2016). The good, the bad and the ugly? Balancing environmental and economic impacts towards efficiency. *MPRA* Paper **72132**, University Library of Munich, Germany.
He J. and Wang H. (2012). Economic structure, development policy and environmental quality: An empirical analysis of environmental Kuznets curves with Chinese municipal data. *Ecological Economics,* **76**: 49–59.
Harrington D.R., Deltas, G., and Khanna M. (2014). Does Pollution Prevention Reduce Toxic Releases? A Dynamic Panel Data Model. *Land Economics*, **90(2)**: 199-221.
Hausman J. and Ros A. (2013). An Econometric Assessment of Telecommunications Prices and Consumer Surplus in Mexico Using Panel Data, *Journal of Regulatory Economics*, **43**: 284–304.
Holtz-Eakin D. and Selden T.M. (1995). Stoking the fires? CO_2 emissions and economic growth. *Journal of Public Economics,* **57:** 85–101.
Kijima M., Nishide K. and Ohyama A. (2010). Economic models for the environmental Kuznets curve: A survey. *Journal of Economic Dynamics and Control,* **34:** 1187–1201.
Kyung S.I., Pesaran M.H. and Shin Y. (2003). Testing for Unit Roots in Heterogeneous Panels. *Journal of Econometrics*, **115:** 53-74.
Lee J.M., Chen K.H and Cho C.H. (2015). The relationship between CO_2 emissions and financial development: Evidence from OECD countries. *The Singapore Economic Review*, **60(5)**: 1550117.
Li Y. and Lyons B. (2012). Market Structure, Regulation and the Speed of Mobile Network Penetration. *International Journal of Industrial Organization*, **30**: 697–707.
Luintel K.B., Khan M., Arestis P. and Theodoridis K. (2008). Financial structure and economic growth. *Journal of Development Economics*, **86(1)**: 181-200.

Maddala, G.S. and Wu, Shaowen. (1999). A Comparative Study of Unit Root Tests With Panel Data and a New Simple Test. *Oxford Bulletin of Economics and Statistics,* **61:** 631-652.
Managi S. and Kaneko S. (2006). Economic growth and the environment in China: An empirical analysis of productivity. *International Journal of Global Environmental Issues,* **6(1)**: 89-133.
Ozturk I. and Acaravci A. (2013). The long-run and causal analysis of energy, growth, openness and financial development on carbon emissions in Turkey. *Energy Economics*, **36(C)**: 262–267.
Pesaran M.H. (2004). General diagnostic tests for cross section dependence in panels. Cambridge Working Papers in Economics **0435**, Faculty of Economics, University of Cambridge.
Polemis M. (2016). New evidence on the impact of structural reforms on electricity sector performance, *Energy Policy*, **92(C)**: 420-431.
Roodman D. (2009). How to do xtabond2: An introduction to difference and system GMM in Stata. *Stata Journal,* **9(1)**: 86-136.
Shahbaz M and Lean H.H. (2012). Does financial development increase energy consumption? The role of industrialization and urbanization in Tunisia. *Energy Policy*, **40:** 473–479.
Shahbaz M. Shahzad, S., Hussain J., Nawaz, A., and A. Shaista(2016). Financial development and environmental quality: The way forward, MPRA Paper **74704**, University Library of Munich, Germany.
Shen J., 2008. Trade liberalization and environmental degradation in China. *Applied Economics* **40(8)**: 997–1004.
Shen J., (2006). A simultaneous estimation of Environmental Kuznets Curve: evidence from China. *China Economic Review,* **17(4)**, 383–394.
Stern D.I., (2004). The rise and fall of the Environmental Kuznets Curve. *World Development* **32(8):** 1419–1439.
Suarez P.R. and Menendez A.J.L (2015). Growing green? Forecasting CO_2 emissions with Environmental Kuznets Curves and Logistic Growth Models. *Environmental Science and Policy,* **54**: 428–437.
Tyavambiza T. and Nyangara D. (2015). Financial and Monetary Reforms and the Finance-Growth Relationship in Zimbabwe. *International Journal of Economics and Financial Issues,* **5(2)**: 590-602.
Westerlund J. (2007). Testing for error correction in panel data. *Oxford Bulletin of Economics and Statistics,* **69:** 709-748.
Verbeek M. (2012). *A guide to modern Economterics.* 4th Ed. Wiley.
Yuxiang K. and Chen Z. (2010). Financial development and environmental performance: Evidence from China. *Environment and Development Economics*, **16**: 93–111.
Zhang Y.J. (2011). The impact of financial development on carbon emissions: An empirical analysis in China. *Energy Policy*, **39**: 2197–2203.