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A study of CO_2 emissions, output, energy consumption, and trade

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Abstract

This article contributes to the literature by investigating the dynamic relationship between Carbone dioxide (CO_2) emissions, output (GDP), energy consumption, and trade using the bounds testing approach to cointegration and the ARDL methodology for Tunisia over the period 1971-2008. The empirical results reveal the existence of two causal long-run relationships between the variables. In the short-run, there are three unidirectional Granger causality relationships, which run from GDP, squared GDP and energy consumption to CO_2 emissions. To check the stability in the parameter of the selected model, CUSUM and CUSUMSQ were used. The results also provide important policy implications.

Keywords: CO_2 emissions, Energy consumption, ARDL bounds testing approach
JEL Classifications: Q56, Q43, C51

1. Introduction

In the last few decades, pollutant emissions can be considered under the heading of the major worldwide on-going concerns. Worldwide organizations, such as the United Nations (UN) or the World Economic Forum (WEF), have been attempting to reduce the adverse impacts of global warming and climate change on the economy. The Kyoto protocol to the United Nations Framework Convention on Climate Change (UFCCC) is one of the important solutions to reduce greenhouse gases (GHG).

Tunisia, as a developing country, has ratified this protocol in 2003 in order to reduce the continuous increase in GHG emissions over time. Returning to the history, the inventory of GHG for 1994 shows a relatively limited contribution to the greenhouse effect. In fact, the net anthropogenic GHG emissions of Tunisia are 23.4 million tons CO_2 equivalent, which represent 2.66 million tons CO_2 equivalent per capita or 1.8 million tons CO_2 equivalent per thousand US$ of GDP. In gross terms, Tunisian emissions presented 28.9 million tons CO_2 equivalent, meaning 3.3 million tons CO_2 equivalent per capita. The GHG absorption totaled 5.5 million tons of CO_2, meaning 0.6 tons of CO_2 per capita. The analysis of the gross Tunisian emissions by source shows a preponderance of energy in the emission balance. In fact, with 15.3 million tons CO_2 equivalent, energy uses represent more than half the gross national emissions (52.8%), followed by agriculture, which represents with 6 million tons CO_2 equivalent, meaning 20.8% of gross emissions. Next come forests and industrial processes with respectively 3.7 million tons of CO_2 (12.9%) and 2.8 million tons of CO_2 (9.8%) of gross national emissions. Wastes remain a relatively weak source of emissions with only 1 million tons of CO_2, meaning 3.6% of gross Tunisian emissions (Ministry of Environment and Land Planning, 2001).

Between 1994 and 2002, both energy generation and the transport sector have been considered as major contributors to air pollution with 31% and 30%, respectively. With more details, CO_2 emissions account for 92% of the total GHG emissions, while methane (CH_4) emissions account for 7%, and nitrous oxide (N_2O) emissions account for 1%. The transport sector is the top contributor of CO_2 emissions, while GHG emissions of CO_2 from this sector rose from 3.4 million tons to 5.8 million tons, with an annual increase rate of 9%. In the second order, GHG emissions of CO_2 from the energy sector also accounted for 29% in 2002, rising from 23% in 1994 (United Nations Environment Programme, UNEP, 2002, chapter 2). For that, an important program is proposed by the Environment and Management body of the Territory Ministry in collaboration with the National Environment Protection Agency and Tunis International Centre for Environmental Technologies. The objective of this program

consists to preserve the air quality by identification, characterization and reduction of fixed (industrial sector) or mobile (transport sector) pollutions (Chaaban, 2008).

Between 2005 and 2008, the government has been already allowed by clean energy plans to save $1.1 billion in energy bills, relative to initial investments of $200 million in clean energy infrastructure. In December 2009, the government presented the first national Solar Energy Plan and other complementary plans with the objective of increasing the share of energy sources using renewable energy from just under 1% to 4.3% in 2014. To implement this plan, total financial resources have been estimated at $2.5 billion, including $175 million from the National Fund, $530 million from the public sector, $1,660 million from private sector funds, and $24 million from international cooperation, all to be spent by 2016 on 40 renewable energy projects. Approximately 40% of the resources are devoted to the development of energy export infrastructure. Expected to result from the Solar Energy Plan, the energy savings could reach 22% for 2016, with a reduction of 1.3 million tons per year of CO_2 emissions.

Indeed, the Tunisian Solar Programme (PROSOL) is a joint initiative between the Tunisian National Agency for Energy Conservation (*Agence National de Maitrise de l'Energie*, ANME), the Tunisian Company of Electricity and Gas (*Société Tunisienne de l'Electricité et de Gaz*, STEG), the United Nations Environment Programme (UNEP) and the Italian Ministry for the Environment, Land and Sea (IMELS), provides an example of solar thermal market development. This program has helped more than 105 000 Tunisian families get their hot water from the sun, based on loans of over $60 million, a substantial leverage on Prosol's initial $2.5 million cost. Financial and fiscal support combines a capital grant qualifying for a value added tax exemption, customs duty reduction and a bank loan with a reduced interest rate. Repayment of the loan is organized through the regular utility bill of the state electric utility STEG, with local banks receiving support that allows them to finance Solar Water Heating (SWH) projects with reduced interest rates. This arrangement has generated direct financial benefits for the end users, when comparing the size of the monthly installments for a SWH system to the earlier electricity bills. A complementary interest rate subsidy was available during the first two years (2005-2006) of the program, reducing the interest rate of the loan to 0 % to the final end user.

This support was removed in 2007 and annual interest rates for loan repayment have been 6.5%. The government provides a subsidy of 20% of the system cost or $75 per m^2, while customers are expected to finance a minimum of 10% of the purchase and installation costs. Over 50 000 Tunisian families now get their hot water from the sun based on loans amounting

to more than $5 million in 2005 and $7.8 million in 2006. With installed surface of the program reaching 400 000 m², the government has set a more ambitious target of 750 000 m² for the period 2010-2014, a level comparable to much larger countries such as Spain or Italy. As of 2008, PROSOL helped avoid 214 000 tons of cumulative CO_2 emissions. Many jobs have been created, as 42 technology suppliers and more than 1 000 installation companies have sprung up to service the solar market. The tourism and industry sectors are also now involved, with 47 hotels engaged by late 2009, and there are plans to encourage industry to make greater use of the sun's energy (UNEP, 2010).

Recently, some centers and organizations such as International Centre for Trade and Sustainable Development (ICTSD), International Trade Centre (ITC), and World Trade Organization (WTO) conclude that environment and trade are strongly related, and can be mutually supportive. Their essential goal is to propose program on trade, environment and natural resources. This program consists to engage stakeholders and to provide new and innovative thinking on how to make trade policy effectively contribute to sustaining and improving livelihoods in developing countries that depend on natural resources.

Tunisia has been a member of WTO since 29 March 1995. In collaboration with this organization, Tunisian economists have developed a conceptual framework for examining how trade opening can affect environment and pollution. This framework, first applied to study the environmental impact of the North American Free Trade Agreement (NAFTA), separates the impact of trade liberalization into three independent effects: scale, composition and technique. Firstly, the scale's effect indicates the impact on GHG emissions from the increased economic activity resulting from trade. The fundamental conclusion is that trade opening will increase economic activity and hence energy use. Everything else being equal, this increase in the economic activity scale and energy use will lead to higher levels of GHG emissions. Secondly, the composition's effect refers to the way that trade liberalization changes the mix of a country's production towards those products where it has a comparative advantage. The effect on greenhouse gas emissions will depend on the sectors in which a country has comparative advantage. The principal conclusion of this effect will result in less GHG emissions if the expanding sectors are less energy intensive than the contracting sectors. Whether the composition effect results in higher or lower greenhouse gas emissions is therefore difficult to predict in advance. Finally, the technique's effect proves that trade opening can lead to improvements in energy efficiency. This means that the production of goods and services can generate less GHG emissions, the increase in the availability of goods,

services and technologies improves energy efficiency, and the trade can help to meet the challenge of global warming (WTO, 2013).

The remainder of this paper is organized as follow. Section 2 investigates literature review. Section 3 highlights data and methodology. Section 4 presents empirical results. Section 5 states conclusion and policy implications.

2. Literature review

The impact of economic growth, energy consumption and trade openness on the environment has become a dominant question in the recent economic literature. This subject has been mostly examined within the framework of Environmental Kuznets Curve (EKC), in order to suggest the existence of an inverted U-shaped relationship between environmental pollution and per capita income.

Many studies within the context of the EKC are completely devoted to test the validity of the EKC hypothesis for a single country and/or a balanced panel. But in general, these studies can be classified on two categories of empirical researches. The first category consists to study the dynamic relationship between economic growth and environmental pollutants, or between economic growth, energy consumption and environmental pollutants (Selden and Song, 1994; Grossman and Krueger, 1995; Stern et al., 1996; Ekins, 1997; Vincent, 1997; Gale and Mendez, 1998; Stern and Common, 2001; Ang, 2007; Apergis and Payne, 2009; Fodha and Zaghdoud, 2010; Chebbi, 2010), whereas the second category was focused to examine the dynamic relationship between economic growth, energy consumption, environmental pollutants and trade (Suri and Chapman, 1998; Halicioglu, 2009; Jayanthakumaran et al., 2012).

For the first category which not includes trade, Grossman and Krueger (1995), Selden and Song (1994), Vincent (1997), and Gale and Mendez (1998) investigated the impact of economic growth on environmental degradation. But these studies have being criticized by Stern et al. (1996), Ekins, (1997), and Stern and Common, (2001) because the methodological choices of pollutants can significantly influence to change the results. Recently, the literature of the dynamic relationship between pollutant emissions, energy consumption and output has been directly examined by Ang (2007), Apergis and Payne (2009), and Fodha and Zaghdoud (2010). Ang (2007) studied the case of France using cointegration and vector error-correction modeling techniques and his results provide the existence of a robust long-run relationship. The long-run link support that economic growth exerts a causal influence on growth of energy

consumption and growth of pollution. The results also provide that in the short-run there is a unidirectional causality running from growth of energy consumption to economic growth. In other work, Apergis and payne (2009) studied the causal relationship between carbon dioxide emissions, energy consumption, and output for six Central American countries within a panel vector correction model over the and their results indicate that in the long-run the energy consumption has a positive and statistically significant impact on CO_2 emissions but the real output exhibits the EKC hypothesis. For the case of Tunisia, there are two several studies (Fodha and Zaghdoud, 2010; Chebbi, 2010), which investigated this literature. Fodha and Zaghdoud (2010) examined the relationship between economic growth and pollutant emissions (CO_2 and SO_2), and they found that there is a long-run cointegrating relationship between emissions of two pollutants and GDP. They also concluded that an emission reduction policies and more investment in pollution abatement cost will not upset the Tunisian's economic growth. Chebbi's (2010) paper provides some insights into the linkages between energy consumption, carbon emissions and the sectoral components of output growth. For the long-run, the results do not support the neutrality hypothesis between energy consumption and sectoral output growth. For the short-run, the results indicate that the relationships between energy consumption and economic growth, as well as economic growth and environmental pollution are not uniform across sectors (agriculture, industry and services). For that, he suggests to follow a prudent energy and environmental policies, in order to distinguish the differences in the relationship between energy consumption and output growth by sector.

Passing to the second category which includes trade, few studies have examined the dynamic relationship between pollutant emissions, energy consumption, output and trade. Suri and Chapman (1998) studied the impact of trade policy orientation on the EKC hypothesis, by analyzing the effect of the actual movement between countries of goods that embody pollution has not been considered. The EKC hypothesis was found that while both industrializing and industrialized countries have added to their energy requirements by exporting manufactured goods, the growth has been substantially higher in the former. At the same time, industrialized countries have been able to reduce their energy requirements by importing manufactured goods. Recently, Halicioglu (2009) examined the causal relationship between carbon emissions, energy consumption, income, and trade in the case of Turkey by using the ARDL bounds testing to cointegration procedure. The results from the bounds indicate that there are two forms of long-run equilibrium. The first form is that carbon emissions are explained by energy consumption, income and trade, and the second form is

that carbon emissions, energy consumption, and trade are determinants of income. From the first form which respects the EKC hypothesis, the empirical results suggest that income is the most significant variables followed by energy consumption and trade. In other work, Jayanthakumaran et al. (2012) compare China and India during the annual period by using the bounds testing approach to cointegration and the ARDL methodology to test the long-run and short-run relationships between CO_2 emissions, economic growth, energy consumption, trade, and endogenously determined structural breaks. The results show that the CO_2 emissions in China were influenced by income, structural changes and energy consumption. A similar causal relationship cannot be established for India with regard to structural changes and CO_2 emissions because the informal economy of India is much larger than China's informal economy. Jayanthakumaran et al. (2012) also conclude that India possesses an extraordinarily large number of micro-enterprises that are low energy consumers and not competitive enough to reach international markets.

3. Data and methodology

3.1. Data and model specification

This paper follows the methodology of Halicioglu (2009) and Jayanthakumaran et al. (2012) into a single multivariate framework. With respect to this methodology, the log quadratic EKC equation used to examine the relationship between carbon dioxide emissions (C), real GDP (Y), squared real GDP (Y^2), energy consumption (E) and trade openness (T) in Tunisia is given by the following equation:

$$C_t = \alpha_0 + \alpha_1.Y_t + \alpha_2.Y_t^2 + \alpha_3.E_t + \alpha_4.T_t + \varepsilon_t \quad (1)$$

where t, α_0 and ε denote the time, the fixed country effect and the white noise stochastic disturbance term, respectively. C is per capita CO_2 emissions in metric tons, Y indicates per capita real GDP in constant 2005 US dollars, Y^2 is the square of per capita real GDP, E indicates per capita energy consumption in kg of oil equivalent, and T stands for trade openness (that is equal to the ratio of exports plus imports to GDP). All variables presented in Variables in Eq. (1) are in natural logarithms. Annual data from 1971 to 2008 are obtained from the World Bank Development Indicators (Fig 1).

The parameters α_1, α_2, α_3, and α_4 are the long-run elasticities of CO_2 emissions with respect to real GDP, squared real GDP, energy consumption and trade openness, respectively.

As for the expected signs in equation (1), one would expect that the sign of α_1 expected to be positive whereas a negative sign is expected for α_2 for the EKC hypothesis to be true. The sign α_3 is expected to be positive because more energy consumption can increase the scale of an economy and stimulate CO_2 emissions. The expected sign of α_4 is mixed depending on the level of economic development stage of a country. For the case of developed countries, this sign is expected to be negative as they cease to produce certain pollution intensive goods and begin to import these from other countries with less restrictive environmental protection laws. But for the case of developing countries, this sign expectation is reversed as they tend to have dirty industries with heavy share of pollutants (Grossman and Krueger, 1995). It means also that an increase in trade openness will increase pollution due to a comparative advantage in dirty production under weaker environmental regulations (Jayanthakumaran, 2012).

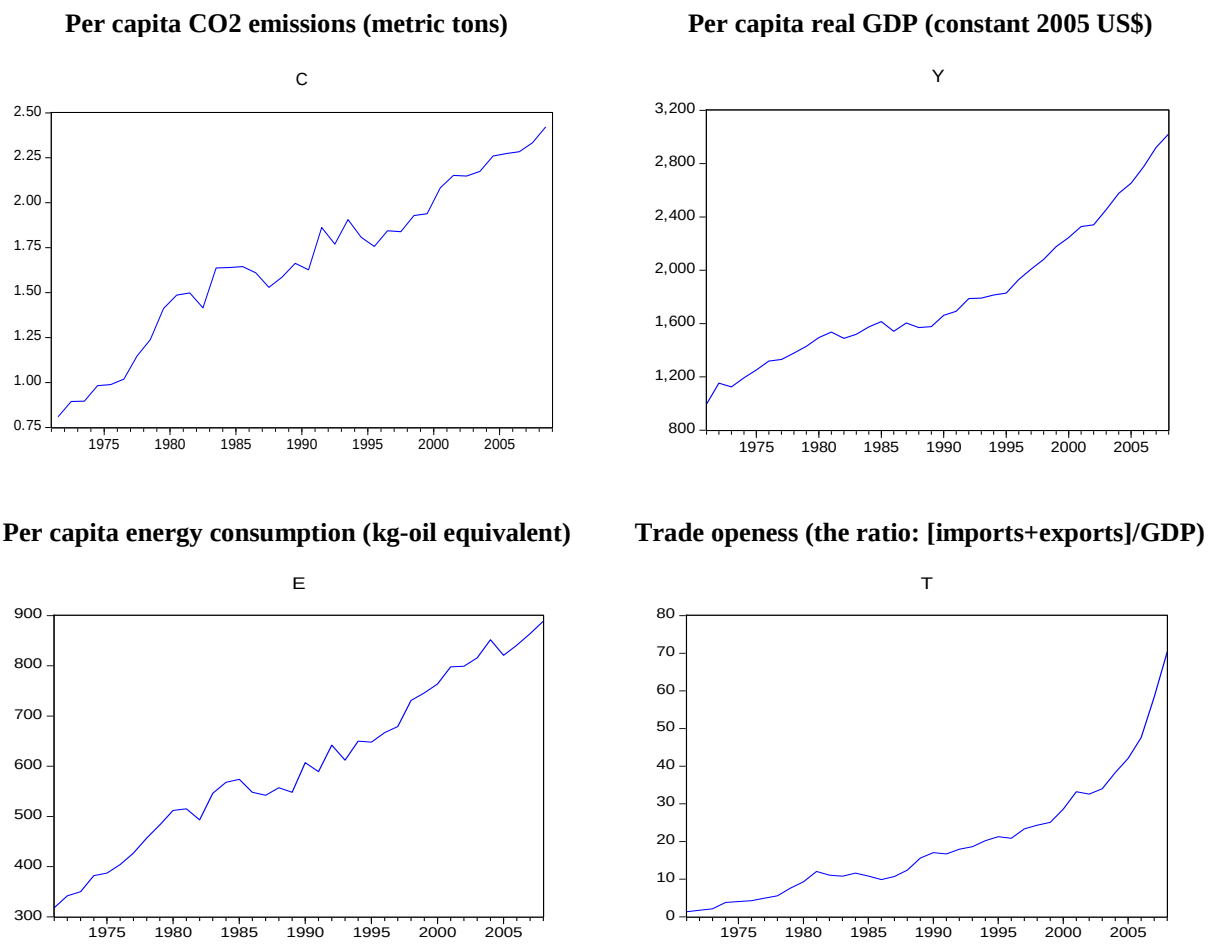


Fig 1. *Plots of per capita CO_2 emissions, per capita real GDP, per capita energy consumption and trade*
Source: compiled from World Bank, Tunisia

3.2. ARDL bounds testing approach

Several econometric procedures were employed to investigate the environmental pollution functions, in the last two decades. With regards to univariate cointegration approaches, there are several examples including Engle and Granger (1987), Fully Modified Ordinary Least Squares (FMOLS) procedures initially suggested by Philips and Hansen (1990), and Dynamic Ordinary Least Squares (DOLS) procedures initially suggested by Saikkonen (1991) and Stock and Watson (1993). There are also many examples of multivariate cointegration procedures including Johansen (1988), Johansen and Juselius (1990), and Johansen's (1996) maximum likelihood technique.

A recent single cointegration approach, known as Auto-Regressive Distributed Lag (ARDL) of Pesaran et al. (2001), has become the most widely used approach by researchers. This cointegration approach, also known as bounds testing, has certain econometric advantages in comparison to other single cointegration procedures. It solves the endogeneity problems and the inability to test hypotheses on the estimated coefficients in the long-run associated with the Engle-Granger method (Engle and Granger, 1987); it studies the long- and short-run estimated parameters in the model in question; it applies the ARDL approach to testing for the existence of a long-run relationship between the variables purely I(0), purely I(1), or fractionally integrated; and the final advantage is that the small sample properties of the bounds testing approach are far superior to th/at of multivariate cointegration (Narayan, 2005).

The following unrestricted error correction regressions of Eq. (1) were estimated with CO_2 as the dependent variable for Tunisia:

$$\begin{aligned}\Delta C_t = \beta_0 + \sum_{i=1}^{p} \beta_1.\Delta C_{t-i} + \sum_{j=0}^{q} \beta_2.\Delta Y_{t-j} + \sum_{k=0}^{r} \beta_3.\Delta Y_{t-k}^2 + \sum_{l=0}^{S} \beta_4.\Delta E_{t-l} + \sum_{m=0}^{w} \beta_5.\Delta T_{t-m} \\ + \beta_6.C_{t-1} + \beta_7.Y_{t-1} + \beta_8.Y_{t-1}^2 + \beta_9.E_{t-1} + \beta_{10}.T_{t-1} + \mu_t\end{aligned} \quad (2)$$

The bounds testing procedure is the first stage of the ARDL cointegration method. Accordingly, a joint significance test, that implies no cointegration hypothesis ($H_0 : \beta_i = 0;\ \forall$ i= 6, 7, 8, 9, 10) against the alternative hypothesis ($H_1 : \beta_i \neq 0;\ \forall$ i= 6, 7, 8, 9, 10), should be performed for Eq. (2).

The bounds testing approach to cointegration requires carrying out the F-test on the selected ARDL models including appropriate lag lengths of selection criterion such as Akaike Information Criterion (AIC) and Schwarz Information Criterion (SIC). In this stage, it is

imposed maximum 2 lags on the level of variables and then selected the optimum lag numbers.

At the second stage, it is also possible to perform for the selected ARDL representation, a general error correction model (ECM) of Eq. (2) formulated as follows:

$$\Delta C_t = \phi_0 + \sum_{i=1}^{p} \theta_1.\Delta C_{t-i} + \sum_{j=0}^{q} \theta_2.\Delta Y_{t-j} + \sum_{k=0}^{r} \theta_3.\Delta Y_{t-k}^2 + \sum_{l=0}^{S} \theta_4.\Delta E_{t-l} + \sum_{m=0}^{w} \theta_5.\Delta T_{t-m} + \lambda.ECT_{t-1} + \xi_t \quad (3)$$

where Δ denotes first differences term; λ is the error correction parameter, ECT_{t-1} is the residuals that are obtained from the estimated cointegration model of Eq. (1), and ξ_t is the disturbance term assumed to be uncorrelated with zero means.

We definite the lagged residuals estimated in Eq. (4) as the error correction term (ECT_{t-1}) then we estimate the parameters related to the short-run model:

$$ECT_{t-1} = C_{t-1} - \hat{\alpha}_1.Y_{t-1} - \hat{\alpha}_2 Y_{t-1}^2 - \hat{\alpha}_3.E_{t-1} - \hat{\alpha}_4.T_{t-1} \quad (4)$$

The ARDL bounds test of cointegration is complemented by Johansen and Juselius's (1990) maximum likelihood to provide a sensitivity check on the results.

A brief reminder of the Johansen and Juselius's (1990) multivariate cointegration methodology is illustrated below:

$$X_t = A + \sum_{i=1}^{k} \Gamma_i.X_{t-i} + \varepsilon_t \quad (5)$$

where $X_t = (C_t,\ Y_t,\ Y_t^2,\ E_t,\ T_t)$ represents a vector of endogenous I(1) variables, A is a vector of constant terms, Γ represents coefficient matrix , k denotes the lag length, and ε_t is the residual matrix.

All variables in Eq. (5) are considered to be potentially endogenous. The cointegrating rank can be found via the trace and the maximal eigenvalue tests. The lag length of the unrestricted vector autoregressive (VAR) structure in Eq. (5) is based on the minimization of Final

Prediction Error (FPE), Akaike (AIC), Schwarz (SIC) and Hannan-Quinn (HQ) criteria, and the maximization of the Likelihood Ratio (LR) criterion.

3.3. Granger causality test

A panel vector error correction model (VECM) is estimated to perform Granger-causality test (Pesaran et al., 1999). This panel followed by the two steps of Engle and Granger (1987) is employed to investigate the long-run and short-run dynamic relationships. The first step estimates the long-run parameters in equation (Eq. 1) in order to obtain the residuals corresponding to the deviation from equilibrium. The second step estimates the parameters related to the short-run adjustment. The resulting equations are used in conjunction with panel Granger causality testing:

$$\begin{pmatrix} \Delta C_t \\ \Delta Y_t \\ \Delta Y_t^2 \\ \Delta E_t \\ \Delta T_t \end{pmatrix} = \begin{pmatrix} \phi_1 \\ \phi_2 \\ \phi_3 \\ \phi_4 \\ \phi_5 \end{pmatrix} + \sum_{k=1}^{m} \begin{pmatrix} \theta_{1,1,k} & \theta_{1,2,k} & \theta_{1,3,k} & \theta_{1,4,k} & \theta_{1,5,k} \\ \theta_{2,1,k} & \theta_{2,2,k} & \theta_{2,3,k} & \theta_{2,4,k} & \theta_{2,5,k} \\ \theta_{3,1,k} & \theta_{3,2,k} & \theta_{3,3,k} & \theta_{3,4,k} & \theta_{3,5,k} \\ \theta_{4,1,k} & \theta_{4,2,k} & \theta_{4,3,k} & \theta_{4,4,k} & \theta_{4,5,k} \\ \theta_{5,1,k} & \theta_{5,2,k} & \theta_{5,3,k} & \theta_{5,4,k} & \theta_{5,5,k} \end{pmatrix} . \begin{pmatrix} \Delta C_{t-k} \\ \Delta Y_{t-k} \\ \Delta Y_{t-k}^2 \\ \Delta E_{t-k} \\ \Delta T_{t-k} \end{pmatrix} + \begin{pmatrix} \lambda_1 \\ \lambda_2 \\ \lambda_3 \\ \lambda_4 \\ \lambda_5 \end{pmatrix} . ECT_{t-1} + \begin{pmatrix} u_{1,t} \\ u_{2,t} \\ u_{3,t} \\ u_{4,t} \\ u_{5,t} \end{pmatrix} \quad (6)$$

where ϕ_j (j=1,2,3,4,5) represents the fixed country effect; k (k=1,…,m) is the optimal lag length determined by the Schwarz information Criterion (SC), ECT_{t-1} is the estimated lagged error correction term derived from the long-run relationship presented in Eq. (1) and estimated via Eq. (4), λ_j (j=1,2,3,4,5) is the adjustment coefficient, and $u_{j,t}$ (j=1,2,3,4,5) is the disturbance term assumed to be uncorrelated with zero means.

Opposite to Eq. (3), all error-correction vectors in Eq. (6) are estimated with the same lag structure ($p = q = r = s = w = m$) that is determined in unrestricted VAR framework.

3.4. Instability tests

Bahmani-Oskooee and Chomsisengphet (2002) suggested that the existence of a cointegration derived from Eq. (2) does not necessarily imply that the estimated coefficients are stable. Generally, the instability of coefficients of regression equations are tested by means of Chow (1960), Brown et al. (1975) Hansen (1992), and Hansen and Johansen (1999) tests. The Chow (1960) test reveals a priori knowledge of structural breaks in the estimation period. In Hansen (1992) and Hansen and Johansen (1999) procedures, instability tests require

I(1) variables and they show the long-run parameter constancy without incorporating the short-run dynamics of a model into the testing-as discussed in Bahmani-Oskooee and Chomsisengphet (2002). Hence, instability tests of Brown et al. (1975), which are also known as cumulative sum (CUSUM) and cumulative sum of squares (CUSUMSQ) tests based on the recursive regression residuals, may be incorporated the short-run dynamics to the long-run through residuals. The CUSUM and CUSUMSQ statistics are updated recursively and plotted against the break points of the model. Provided that the plots of these statistics fall inside the critical bounds of 5% significance, one assumes that the coefficients of a given regression are stable. These tests are explained with more details in Farhani (2012a).

4. Empirical results

4.1. Unit root tests

The time series properties of the variables in Eq. (1) are implemented through two unit root tests: Augmented-Dickey-Fuller (ADF) of Dickey and Fuller (1979, 1981) and Phillips-Perron (PP) of Phillips and Perron (1988). The results displayed in Table 1 shows that all series in Eq. (1) appear to contain a unit root in their levels but stationary in their first differences. Hence, we conclude that they are integrated at order one, I(1), at the 5% level of significance.

Table 1
Unit root tests

		C	*Y*	*Y²*	*E*	*T*
ADF	Level	-2.334041	-0.440151	-0.084117	-1.928492	-2.449106
		(0.1671)	(0.8916)	(0.9439)	(0.3161)	(0.1359)
	First difference	-7.311233	-9.030495	-8.556520	-8.259719	-4.442504
		(0.0000)**	(0.0000)**	(0.0000)**	(0.0003)**	(0.0011)**
PP	Level	-2.793961	-0.418565	-0.037330	-2.775218	-2.312435
		(0.0689)	(0.8955)	(0.9489)	(0.0716)	(0.1735)
	First difference	-7.210741	-8.652443	-8.150293	-8.099610	-4.424094
		(0.0000)**	(0.0000)**	(0.0000)**	(0.0000)**	(0.0012)**

ADF and PP examine the null hypothesis of non-stationary.
For ADF, EViews 7 software is used to select the optimal lag length. The optimal lag length stands for the lag level that maximizes the Akaike Information Criteria (AIC).
For PP, Barlett Kernel is used as the spectral estimation method. The bandwidth is selected using the Newey–West method.
** denotes statistical significance at the 5% level.

4.2. ARDL cointegration method

Given that C_t, Y_t, Y_t^2, E_t and T_t share common integration properties, Eq. (2) was estimated in two stages. In the first stage of the ARDL procedure, the long-run relationship of Eq. (1) was established in two steps. Firstly, the order of lag, used to estimate later VAR or

VECM models, was obtained from unrestricted VAR by minimized Final Prediction Error (FPE) criterion, Akaike Information Criterion (AIC), Schwarz Information Criterion (SIC) and Hannan-Quinn (HQ) criterion, and by maximized Likelihood Ratio (LR) criterion. The results of this stage are displayed here to conserve the optimal lag which is found to be one, except AIC (Table 2, A). Using this lag, the result of the Johansen maximum eigenvalue test, based on both Johansen (1988) and the modified version suggested by Reinsel and Ahn (1992), points to conclude that there is three cointegrated relationship, at the 5% level of significance (Table 2,B). Secondly, the order of lags (*p*, *q*, *r*, *s*, *w*) on the first differenced variables for Eq. (2) was obtained from the minimization of AIC and SIC criteria (The lag length is used to be two). For AIC, the minimization of this criterion leads to choose the optimal lags: p = 2, q = 2, r = 2, s = 0, w = 1, whereas the minimization of SIC leads to choose: p = 2, q = 0, r = 0, s = 0, w = 1 (Table 2,C). The presence of a diagnostic divergence in the use of these criteria often arrives in the reality (Farhani, 2012b). In this case, it is necessary to understand very well the objective of the study, which consists to control the autocorrelation of the innovations. According to a principle of parsimony (Raykov and Marcoulides, 1999), it is advisable to estimate the model which include the minimum of lags and which presents absence of residuals autocorrelation, too.

Eq. (2) was estimated to obtain the long-run and short-run coefficients for Tunisia using ARDL methodology (Table 3). The results in Table 3 indicate that the coefficient (0.29) is much less when we consider the one-year lag in CO_2 emissions in the equation.

The same table also indicates that per capita real GDP and per capita real GDP squared had a long-run and short-run positive impact on per capita CO_2 emissions. More precisely, a 1% increase in per capita real GDP will increase per capita CO_2 emissions by 4.67% over the long-run, and 6.66% over the short-run, both of which are significant at the 5% level. On the other hand, a 1% increase in per capita GDP squared will lead to decrease CO_2 emissions by 0.28% over the long-run, and 1.76% over the short-run, both of which are significant at the 5% level.

To complement the findings of derive analysis, the elasticity of CO_2 emissions with respect to per capita real income in the long-run is 4.67-0.28Y. The positive sign for per capita real GDP (Y) in natural logarithm and the negative sign for capita real GDP squared (Y^2) in natural logarithm are as expected, supporting the EKC hypothesis that pollution emissions initially increase with income and then decrease after incomes reach a certain level.

Table 2

A. VAR lag order selection criteria

Lag	*LogL*	*LR*	*FPE*	*AIC*	*SIC*	*HQ*
0	210.7386	NA	5.39e-12	-11.75649	-11.53430	-11.67979
1	397.4491	309.4060[a]	5.32e-16[a]	-20.99709	-19.66394[a]	-20.53689[a]
2	422.2375	33.99540	5.91e-16	-20.98500	-18.54088	-20.14129
3	449.1316	29.19937	6.85e-16	-21.09324[a]	-17.53815	-19.86602

B. Johansen maximum eigenvalue test [b]

Hypothesized No. of CE(s)	Eigen-Value	Trace Statistic	5 % Critical Value	Prob.[#]
None**	0.793309	124.5441	69.81889	0.0000
At most 1**	0.635729	67.78905	47.85613	0.0002
At most 2**	0.369583	31.43418	29.79707	0.0321
At most 3*	0.324032	14.82470	15.49471	0.0629
At most 4	0.019985	0.726748	3.841466	0.3939

C. Order of optimal lags (p, q, r, s, w)[c]

Number of lags	AIC	SIC	Number of lags	AIC	SIC
p=1, q=0, r=0, s=0, w=0	- 3.277536	- 2.793683	p=2, q=1, r=0, s=0, w=1	- 3.389592	- 2.767453
p=1, q=1, r=0, s=0, w=0	- 3.222708	- 2.694868	p=2, q=1, r=1, s=1, w=0	- 3.208091	- 2.541513
p=1, q=0, r=1, s=0, w=0	- 3.222287	- 2.694448	p=2, q=1, r=1, s=0, w=1	- 3.336078	- 2.669500
p=1, q=0, r=0, s=1, w=0	- 3.301680	- 2.773841	p=2, q=1, r=1, s=0, w=1	- 3.293765	- 2.582749
p=1, q=0, r=0, s=0, w=1	- 3.352330	- 2.824491	p=2, q=2, r=0, s=0, w=0	- 3.181686	- 2.559546
p=1, q=1, r=1, s=0, w=0	- 3.255082	- 2.683255	p=2, q=2, r=1, s=0, w=0	- 3.172205	- 2.505627
p=1, q=1, r=0, s=1, w=0	- 3.303010	- 2.731183	p=2, q=2, r=0, s=1, w=0	- 3.186938	- 2.520360
p=1, q=1, r=0, s=0, w=1	- 3.307060	- 2.735233	p=2, q=2, r=0, s=0, w=1	- 3.360194	- 2.693616
p=1, q=1, r=1, s=1, w=0	- 3.295252	- 2.679439	p=2, q=2, r=1, s=1, w=0	- 3.154063	- 2.443046
p=1, q=1, r=1, s=0, w=1	- 3.302386	- 2.686573	p=2, q=2, r=1, s=0, w=1	- 3.317933	- 2.606917
p=1, q=1, r=1, s=1, w=1	- 3.322274	- 2.662475	p=2, q=2, r=1, s=1, w=1	- 3.262903	- 2.507449
p=2, q=0, r=0, s=0, w=0	- 3.292301	- 2.759038	p=2, q=2, r=2, s=0, w=0	- 3.281005	- 2.569989
p=2, q=1, r=0, s=0, w=0	- 3.235770	- 2.658070	p=2, q=2, r=2, s=1, w=0	- 3.277436	- 2.521981
p=2, q=0, r=1, s=0, w=0	- 3.235533	- 2.657832	p=2, q=2, r=2, s=0, w=1	**- 3.555412**	- 2.799957
p=2, q=0, r=0, s=1, w=0	- 3.281255	- 2.703554	p=2, q=2, r=2, s=1, w=1	- 3.500187	- 2.700294
p=2, q=0, r=0, s=0, w=1	- 3.444910	**- 2.867210**	p=2, q=2, r=2, s=2, w=0	- 3.228412	- 2.428519
p=2, q=1, r=1, s=0, w=0	- 3.211844	- 2.589704	p=2, q=2, r=2, s=2, w=1	- 3.446689	- 2.602358
p=2, q=1, r=0, s=1, w=0	- 3.243930	- 2.621790	p=2, q=2, r=2, s=2, w=2	- 3.391616	- 2.502846

[a] denotes the optimal lag selection.

[b] Trace test indicates 3 cointegrating Eq(s) at the 5% level.

* and ** indicate 10% and 5% level of significance, respectively.

[#] MacKinnon-Haug-Michelis (MHM) of MacKinnon et al. (1999) p-values.

[c] Statistics in bold denote the value of the minimized AIC and SIC.

The statistical significance of the squared per capita real GDP consists to suggest the rules in which GDP raises monotonically with the level of CO_2 emissions. Expected to the EKC hypothesis, the results lead to obtain the support that pollution emissions initially increase with GDP and then decrease after GDP reaches its stabilization point. However, the graphical representation (Fig 2) of per capita CO2 emissions and per capita real GDP does not support the existence of the EKC hypothesis.

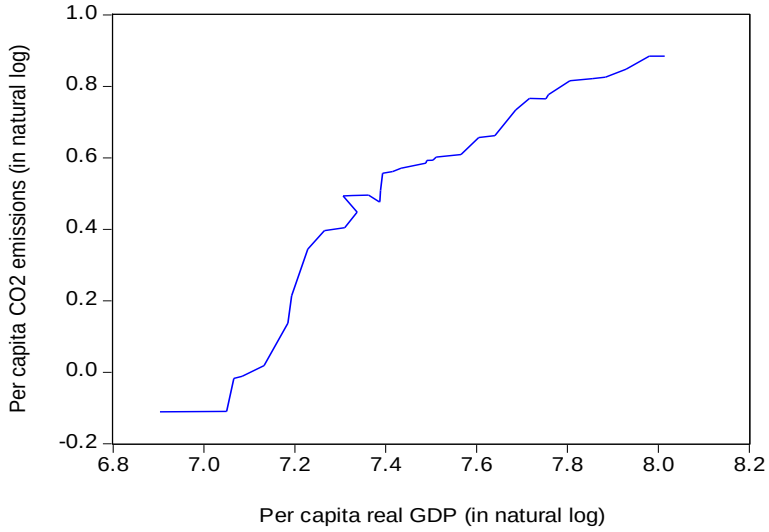


Fig 2. *Plot of CO_2 emissions and real GDP*

The same table also indicates that per capita energy consumption has a long-run and short-run positive impact on per capita CO_2 emissions. The finding results show that a 1% increase in per capita energy consumption will lead to increase CO_2 emissions by 0.45% over the long-run, and 0.7% over the short-run, both of which are significant at the 5% level

The results of the trade indicate a small, but positive, relationship with CO_2 emissions in the short-run only. For more detail, a 1 % increase in trade openness will increase CO_2 emissions by 0.19% in the short-run, which is significant at the 5% level. Our findings are broadly consistent with Selden and Song (1994), Grossman and Krueger (1995), Ang (2007), Halicioglu (2009) and Jayanthakumaran et al. (2012), who have also reported an inverted U-shaped relationship between pollution and output.

The error correction term ECT_{t-1} represents the speed of adjustment of ΔCO_2 to its long-run equilibrium following a shock. The coefficient of –0.55 is significant at the 5% level with a negative sign expected, suggesting that a deviation from the long-run equilibrium level of CO_2 emissions in 1 year is corrected by 55% over the following year. Moreover, a significant error correction confirms the existence of a stable long-run relationship between the regressors (Y_t, Y_t^2, E_t, T_t) and the dependent variable CO_2 emissions (C_t).

Table 3
ECM-ARDL results

A. Dependent variable ΔC_t based on ARDL(2,0,0,0,1) selected lags based on SIC			*B. Cointegrating vector: Dependent variable C_t*		
Regressors	Coefficients	t-ratio	Regressors	Coefficients	t-ratio
ΔC_{t-1}	0.291970	2.28146**	Y_t	4.670921	-2.46434**
ΔY_t	6.663501	-3.67710**	Y_t^2	- 0.278555	2.34331**
ΔY_t^2	-1.718642	3.61816**	E_t	0.445903	-3.97483**
ΔE_t	0.700083	3.45769**	T_t	0.009982	0.39974
ΔT_t	0.186138	-3.03771**	*constant*	-21.70084	
constant	0.015243	1.74773*			
R^2 = 0.73, F = 68.83**, DW = 2.47, SSR = 0.03			ECT_{t-1}	-0.546595	-2.36210**

** and * denote statistical significance at the 5% and 10% levels, respectively.
DW indicates Durbin-Watson statistic. SSR stands for sum squared residuals.

The diagnostic tests indicate that the model passed most of the tests for serial correlation, functional form, normality and heteroscedasticity. The high value of R^2 for ECM-ARDL model shows that the adjustment of ARDL model was extremely good. The F-statistics which measure the joint significance of all regressors in the models are statistically significant at the 5% level, and the Durbin-Watson statistic for the model was greater than two (Table 3).

4.3. Granger causality results

According to the Johansen maximum eigen-value test results revealed in Table 2,B, there exist three cointegrating relationships between C_t, Y_t, Y_t^2, E_t, and T_t. Therefore, the Granger causality test was conducted to Eq. (6) as such that only four long-run relationships were estimated with an error correction term (Table 2,B). Table 4 summarizes the results of the long-run and short-run Granger causality. According to the coefficient on the lagged error-correction term, there exists a long-run relationship among the variables in the form of Eq. (1) as the error-correction term is statistically significant, which also confirms the results of the bounds test. In the long run, real GDP, squared real GDP, energy consumption and trade cause CO_2 emissions. There exists another long-run Granger causality which runs interactively through the error-correction terms from energy CO_2 emissions, real GDP, squared real GDP and trade to the energy consumption.

However, in the short-run Table 4 reveals three unidirectional Granger causality relationships, which run from real GDP, squared real GDP and energy consumption to CO_2 emissions. This result confirms the study of Ang (2007), which suggests that the degradation of the environment does not have a causal impact on economic growth or energy consumption. Instead, expansion of the economy and the energy exert a causal effect on CO_2 emissions. The results call for more environmental protection since environmental pollution

may cause a negative externality to the economic energy through affecting human health and thereby reduce productivity. It means that policies, which impose a higher cost such as environmental taxes, may result in input material substitution by reducing or increasing the pollution levels. The results indicate that the implementation of energy conservation policies has not inversely affected the long-term economic performance of Tunisia. Hence, the results imply that the economy of Tunisia may be less vulnerable to energy shocks, which could adversely affect GDP growth. In the short run, the use of more energy and economic growth is required to pollutant emissions. This problem can be solved by the development of energy conservation strategies.

Table 4
Granger causality test

Dependent variable	*F-statistics* (Probabilities)				ECT_{t-1} [t-stat]
	ΔC_t	ΔY_t -ΔY_t^2	ΔE_t	ΔT_t	
ΔC_t		3.04976**	3.05254**	0.65189	-2.638576**
	#	(0.0456)	(0.0455)	(0.5887)	[-5.16354]
ΔY_t -ΔY_t^2	0.23490		2.10928	0.24202	-2.400441
	(0.8712)	#	(0.1225)	(0.8662)	[-1.34597]
ΔE_t	1.39553	0.99058		1.42514	-1.153546**
	(0.2656)	(0.4121)	#	(0.2571)	[3.51097]
ΔT_t	1.39389	0.09133	0.14058		-0.052913
	(0.2660)	(0.9642)	(0.9348)	#	[-0.13752]

** denotes statistical significance at the 5% level.

4.4. Instability tests results

The SIC based error-correction model of Eq. (2) is selected to implement the CUSUM and CUSUMSQ instability tests. The related graphs of these tests are presented in Fig.3. As can be seen from the graphs of Fig 3, the plots of CUSUM and CUSUMSQ statistics are well within the critical bounds at 5% significance level, implying that all coefficients in the error-correction model are stable. Therefore, the selected CO_2 emissions model can be used for policy decision making purposes, such that the impact of policy changes considering the explanatory variables of CO_2 emissions equation will not cause major distortion in the level of CO_2 emissions, since the parameters in this equation seem to follow a stable pattern during the estimation period (Halicioglu, 2009).

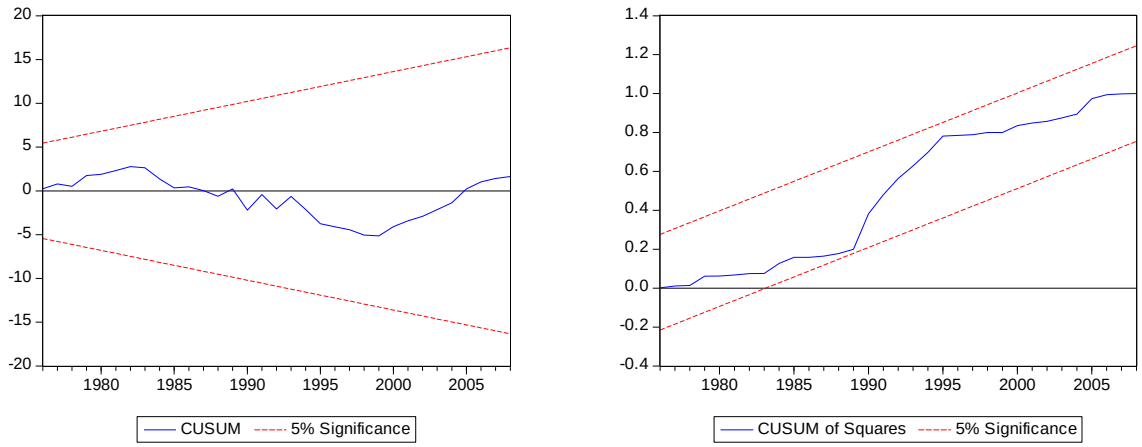


Fig 3. *Plots of CUSUM and CUSUMSQ of recursive residuals*

5. Conclusion and policy implications

This paper consists to analyze the dynamic relationships between CO_2 emissions, real GDP, energy consumption and trade for Tunisia. The long-run relationship in which CO_2 emissions is the dependent variable is used to test the short-run and long-run elasticities of CO_2 emissions with respect to explanatory variables. The long-run elasticity of CO_2 emissions with respect to real GDP is found to be 6.66-3.44Y. Similarly, the elasticity of CO_2 emissions with respect to energy consumption is computed as 0.7. Finally, the long-run elasticity of CO_2 emissions with respect to trade is computed as 0.19. To this extent, the recent study of Chebbi et al. (2011) suggests that the direct effect of trade openness on CO2 emissions is positive both in the short and the long run, but the indirect effect is negative at least in the long run. The overall effect is positive both in the short and long run, highlighting the importance for trade reforms to be accompanied by strong environmental policies.

Granger causality test reveals two long-run relationships. The first indicates that real GDP, squared real GDP, energy consumption and trade cause CO_2 emissions, and the second shows another long-run Granger causality which runs from energy CO_2 emissions, real GDP, squared real GDP and trade to the energy consumption. But in the short-run, Granger causality test reveals three unidirectional relationships, which run from real GDP, squared real GDP and energy consumption to CO_2 emissions. To this extent, the recent study of Fodha and Zaghdoud (2010) investigates the relationship between economic growth and pollutant emissions for Tunisia and they indicate that the unidirectional causality with income causing environmental changes and not vice versa, is implied that an emission reduction policies and more investment in pollution abatement expense will not hurt economic growth. It could be a feasible policy tool for Tunisia to achieve its sustainable growth in the long-run.

The instability tests of the CO_2 emissions equation suggest that policy changes considering the explanatory variables of CO_2 emissions equation will not cause major distortion in the level of CO_2 emissions. The results show that Tunisia should design new environmental policies to reduce environmental degrading.

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