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Cointegration with multiple structural breaks: an application to the Spanish environmental Kuznets curve, 1857-2007

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Abstract

In this paper we consider the possibility that a linear cointegrated regression model with multiples structural changes would provide a better empirical description of the Spanish environmental Kuznets curve during the period 1857-2007. Our methodology is based on instability tests recently proposed in Kejriwal and Perron (2008, 2010) as well as the cointegration test in Arai and Kurozumi (2007) and Kejriwal (2008) developed to allow for a single or multiple breaks under the null hypothesis of cointegration, respectively.

Overall, the results of the Kejriwal-Perron tests suggest a model with two breaks estimated at 1941 and 1967 and three regimes. The coefficient estimated between per capita CO_2 and per-capita income (or long-run elasticity) in a two breaks model show a tendency to decrease over time. This implies that even if per capita CO_2 consumption is monotonically rising in income, the "income elasticity" is less than one.

Keywords: Environmental Kuznets curve; CO_2 emissions; Cointegration; Multiple Structural Breaks

JEL classification: C32; Q43; Q53

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1 Introduction

The environmental Kuznets curve (EKC) states that as per capita income grows, environmental impacts rise, hit a maximum, and then decline. This implies an "inverted-U" shape relationship first defined by Shafik and Bandyopadhyay (1992) and later empirically identified by Grossman and Krueger (1995). The relationship between environmental quality and growth has been the focus of much work historically. In the analysis of the emission-income relationship, there exists a set of theoretical models which derives inverted "U" shaped curves by having emissions increasing with income until some threshold is passed, after which pollution is reduced.[1] The basic idea underlying all these model is that when some threshold income level is passed, then the economy moves to another regime, with the emission-income relationship being different between the old and the new regime. In the inverted "U" models, the low income regime corresponds to an increasing emission-income relationship, while in the regime after the threshold the emission-income relationship is decreasing.

Empirically, evidence for the EKC hypothesis is mixed[2]. Thus, many caveats have appeared regarding the fulfillment of the hypothesis. The first and most obvious is that not all pollutants obey this empirical regularity. Empirical evidence for the EKC has been found in a number of studies, mainly for local and regional air pollutants, whereas for global pollutants either levels rise consistently with income (Horvath, 1997), or turning points are extremely high (Holtz-Eakin and Selden, 1995) or very uncertain (Cole et al, 1997). An extreme case are pollutants that do not seem to have a turning point, and therefore, seem to fit the original scenario stated by the Club of Rome group with emissions appearing to increase with income without limit, or at least to some very high level turning point income. Some of these pollutants are global, with the most important being carbon dioxide (CO2) (Holtz-Eakin and Selden, 1995), the major ingredient in global warming. Other pollutants that seem to increase monotonically with national income include solid municipal waste, traffic volumes, and general energy consumption (Holtz-Eakin and Selden, 1995; Horvath, 1997).

Another caveat involves scattered evidence that some of the pollutants that appear to follow the EKC may in fact exhibit re-linking at higher income levels with a subsequent upswing again in emissions, and hence show an "N-curve" rather than the Kuznets inverted U-curve. A broader problem is that

[1] For a literature review, see, for example, Levinson (2002).
[2] For a literature review, see, for example, He and Richard (2010).

many of these studies were carried out on a cross-section of countries basis rather than on a more careful panel or time-series framework within specific countries. Efforts to do the latter have found wide variations across countries regarding these relationships for many pollutants (Stern et al, 1996; Stern and Common, 2001).

The differences in the empirical findings seem to be due to three factors. First, the nature of the pollutants studied. Second, the fact that other factors can be also driving emission levels, like the degree of trade protection; the degree of political freedom (index of civil liberties), and the effect of economic growth (scale of economy) independently of the income per capita level. Third, the possibility that for some countries and pollutants the turning point has not been met yet.

In order to overcome the above criticisms authors have been refining the empirical strategies. Estimated relationships, mainly in reduced form specifications, have taken cubic or quadratic forms. Estimations methods have also varied from OLS estimation, panel data estimations with fixed or random effects, Tobit estimations, or semiparametric estimation. Finally, explanatory variables have also been augmented including lagged values, population density locational variables, micro or macro variables, distributional variables, trade variables, as well as noneconomic variables such as literacy rates or political rights. All in all, the pollution-income relationship is less robust than previouly thought (Xepapadeas, 2005).

This seems often to be the case for CO_2 and other global/long-term pollutants, as well as for solid waste. As the theory predicts a long-run relationship linking emissions and economic growth, there is a wide stream of research that has assessed this relationship employing linear cointegration techniques. The empirical evidence suggests that economic development and GDP may be jointly determined, so that any constraint put on energy consumption to help reducing emissions will have effects on economic growth. Empirical studies began to emerge that appeared to show inverted U-shaped curves relating pollution emissions and levels of income, both in the aggregate and for various specific pollutants, with most of these studies being based on cross-sectional data across countries and using reduced form equations models.

Some authors, among others, Ozturk (2010), Halicioglu (2009), Chontanawatet al. (2008), Lise (2006), Lee (2005), Soytas and Sari (2003), use linear cointegration procedures to examine CO_2- GDP nexus, however the results are mixed with no clear consensus emerging. These studies have a common feature: a linear and stable EKC.

In a empirical study, Roca et al. (2001) investigate the EKC hypothesis with CO_2 emissions data for Spain from 1973-1996 and find that long-run elasticity between income and CO_2 emissions is positive and greater than 1 ($\beta = 1.24$), revealing a very strong linear cointegration between these two variables but they do not take into account the question of whether the relationship is stable over time, or exhibits a structural break allowing the instability to occur at an unknown point. More recently, the empirical analysis on the Spanish EKC hypothesis has turned to a nonlinear cointegration perspective (Esteve and Tamarit, 2011). Their results confirm the nonlinearity (a two-regime threshold cointegration model) of the link between per-capita CO_2-per-capita income pointing to the existence of an EKC for the Spanish case.

The lack of control for structural breaks in the series may be reflected in the parameters of the estimated models that, when used for inference or forecasting, can induce to misleading results. In general, structural breaks are a problem for the analysis of economic series, since they are usually affected by either exogenous shocks or changes in policy regimes. As a consequence, the assumption of stability in the long-run relationship between CO_2 emissions and GDP would seem too restrictive, so that not allowing for structural breaks would be an important potential shortcoming of the past research using cointegration techniques. In our case the EKC have probably changed due to variations in macroeconomic and market forces, such as changes in the structure of the economy, supply shocks, and reforms in the environmental regulation. Thus, the information content of the EKC is subject to change over time and all the empirical modeling work that does not take into account the possible variations and instability will fail to explain the variations in the per-capita CO_2 and per-capita income relationship. A relevant issue in this analysis is that nonlinearity and instability generally are difficult to distinguish and both are compatible. Particularly, the instability in this relationship could lead to nonlinearity, and vice versa.

In order to answer the previous question, in this paper we extend the existing empirical analysis of the linear model of the EKC in two ways. First, to avoid the econometric problem mentioned, we make use of recent developments in cointegrated regression models with multiple structural changes. Specifically, we use a new approach developed by Kejriwal and Perron (2008, 2010) to test for multiple structural changes in cointegrated regression models. They propose a sequential procedure that not only enables detection of parameter instability in cointegration regression models but also allows a consistent estimations of the number of breaks present. Furthermore, we test the cointegrating relationship when multiple regime shifts are identified endogenously. In particular, the nature of the long run relationship between per-capita CO_2-per-capita income is analyzed using the residual based test of the null hypothesis of cointegration with a single or multiple breaks proposed in Arai and Kurozumi (2007) and Kejriwal (2008), respectively. Second, a common criticism to most test of the EKC is that the econometric procedures used require a large number of observations. Accordingly, in this paper we use a long span of the data (1857-2007). It will allow us to obtain more robust results on the fulfilling of the EKC than in previous analysis.

The rest of the paper is organized as follows. A brief description of the underlying theoretical framework is provided in section 2, the methodology and empirical results are presented in sections 3 and 4, respectively, and the main conclusions are summarized in section 5.

2 A simple linear model of the environmental Kuznets curve

In order to test the EKC hypothesis from CO_2 emissions, the empirical studies commonly used a linear regression model such as:

$$CO_{2t} = \alpha + \beta y_t + \varepsilon_t \tag{1}$$

where CO_{2t} is per-capita CO_2 and y_t per-capita income. Alternatively, we may write the linear regression model (1) as a bivariate linear cointegrating VAR model with one lag, $l = 1$, such as:

$$\begin{pmatrix} \Delta CO_{2t} \\ \Delta y_t \end{pmatrix} = \mu + \alpha w_{t-1} + \Gamma \begin{pmatrix} \Delta CO_{2t-1} \\ \Delta y_{t-1} \end{pmatrix} + \varepsilon_t \tag{2}$$

where the long-run relationship is defined as $w_{t-1} = CO_{2t-1} - \beta y_{t-1}$.

3 Methodology

3.1 A linear cointegrated regression model with multiples structural changes

Issues related to structural change have received a considerable amount of attention in the statistics and econometric literature. Bai and Perron (1998) and Perron (2006, 2008) provide a comprehensive treatment of the problem of testing for multiple structural changes in linear regression models. Accounting for parameter shifts is crucial in cointegration analysis since it normally involves long spans of data which are more likely to be affected by structural breaks. In particular, Kejriwal and Perron (2008, 2010) provide a comprehensive treatment of the problem of testing for multiple structural changes in cointegrated systems.

More specifically, Kejriwal and Perron (2008, 2010) consider a linear model with m multiple structural changes (i.e., $m+1$ regimes) such as:

$$y_t = c_j + z'_{ft}\delta_f + z'_{bt}\delta_{bj} + x'_{ft}\beta_f + x'_{bt}\beta_{bj} + u_t \qquad (t = T_{j-1}+1, ..., T_j) \quad (3)$$

for $j = 1, ..., m+1$, where $T_0 = 0$, $T_{m+1} = T$ and T is the sample size. In this model, y_t is a scalar dependent I(1) variable, $x_{ft}(p_f \times 1)$ and $x_{bt}(p_b \times 1)$ are vectors of I(0) variables while $z_{ft}(q_f \times 1)$ and $z_{bt}(q_b \times 1)$ are vectors of I(1) variables.[3] The break points $(T_1, ..., T_m)$ are treated as unknown.

The general model (3) is a partial structural change model in which the coefficients of only a subset of the regressors are subject to change. In our case, we suppose that $p_f = p_b = q_f = 0$, and estimated model is a pure structural change model with all coefficients of the I(1) regressors and constant (slope and the intercept in (1)) allowed to change across regimes:

$$y_t = c_j + z'_{bt}\delta_{bj} + u_t \qquad (t = T_{j-1}+1, ..., T_j) \qquad (4)$$

Generally, the assumption of strict exogeneity is too restrictive and the test statistics for testing multiple breaks are not robust to the problem of endogenous regressors. To deal with the possibility of endogenous I(1) regressors, Kejriwal and Perron (2008, 2010) propose to use the so-called dynamic OLS regression (DOLS) where leads and lags of the first-differences of the I(1) variables are added as regressors, as suggested Saikkonen (1991) and Stock and Watson (1993):

$$y_t = c_i + z'_{bt}\delta_{bj} + \sum_{j=-l_T}^{l_T} \Delta z'_{bt-j}\Pi_{bj} + u^*_t, \qquad \text{if } T_{i-1} < t \leq T_i \qquad (5)$$

for $i = 1, ..., k+1$, where k is the number of breaks, $T_0 = 0$ and $T_{k+1} = T$.

3.2 Structural Break Tests

We test the parameter instability in cointegration regression using the tests proposed in Kejriwal and Perron (2008, 2010). They present issues related

[3]The subscript b stands for "break" and the subscript f stands for "fixed" (across regimes).

to structural changes in cointegrated models which allows both I(1) and I(0) regressors as well as multiple breaks. They also propose a sequential procedure which permits consistent estimation of the number of breaks, as in Bai and Perron (1998).

Kejiriwal and Perron (2010) consider three types of test statistics for testing multiple breaks. First, they propose a $\sup Wald$ test of the null hypothesis of no structural break ($m = 0$) versus the alternative hypothesis that there are a fixed (arbitrary) number of breaks ($m = k$):

$$\sup F_T^*(k) = \sup_{\lambda \in \Lambda \varepsilon} \frac{SSR_0 - SSR_k}{\hat{\sigma}^2} \tag{6}$$

where SSR_0 denote the sum of squared residuals under the null hypothesis of no breaks, SSR_k denote the sum of squared residuals under the alternative hypothesis of k breaks, $\lambda = \{\lambda_1, ..., \lambda_m\}$ as the vector of breaks fractions defined by $\lambda_i = T_i/T$ for $i = 1, ..., m, T_i$, and T_i are the break dates.

Second, they consider a test of the null hypothesis of no structural break ($m = 0$) versus the alternative hypothesis that there an unknown number of breaks given some upper bound $M(1 \leq m \leq M)$:

$$UD \max F_T^*(M) = \max_{1 \leq k \leq m} F_T^*(k) \tag{7}$$

In addition to the tests above, Kejiriwal and Perron (2010) consider a sequential test of the null hypothesis of k breaks versus the alternative hypothesis of $k + 1$ breaks:

$$SEQ_T(k+1|k) = \max_{1 \leq j \leq k+1} \sup_{\tau \in \Lambda_{j,\varepsilon}} T \left\{ SSR_T(\hat{T}_1, ..., \hat{T}_k \right\} \tag{8}$$

$$- \left\{ SSR_T(\hat{T}_1, ...\hat{T}_{j-1}, \tau, \hat{T}_j, ..., \hat{T}_k \right\} / SSR_{k+1} \tag{9}$$

where $\Lambda_{j,\varepsilon} = \left\{ \tau : \hat{T}_{j-1} + (\hat{T}_j - \hat{T}_{j-1})\varepsilon \leq \tau \leq \hat{T}_j - (\hat{T}_j - \hat{T}_{j-1})\varepsilon \right\}$. The model with k breaks is obtained by a global minimization of the sum of squared residuals, as in Bai and Perron (1998).

3.3 Cointegration tests with structural changes

Kejriwal and Perron (2008, 2010) show that the structural change tests can suffer from important lack of power against spurious regression (i.e., no cointegration). This means that these tests can reject the null of stability when the regression is really a spurious one. In this sense, tests for breaks in the long run relationship are used in conjuction with tests for the presence or absence of cointegration allowing for structural changes in the coefficients.

First, we use the residual-based test of the null of cointegration with a unknown single break against the alternative of no cointegration proposed in Arai and Kurozumi (2007). They propose a LM test based on partial sums of residuals where the break point is obtained by minimizing the sum of squared residuals and consider three models: i) Model 1, level shift; ii) Model 2, level shift with trend; iii) and Model 3, regime shift.

The LM test statistic (for one break), $\tilde{V}_1(\hat{\lambda})$, is given by:

$$\tilde{V}_1(\hat{\lambda}) = (T^{-2} \sum_{t=1}^{T} S_t(\hat{\lambda})^2)/\hat{\Omega}_{11} \tag{10}$$

where $\hat{\Omega}_{11}$ is a consistent estimate of the long run variance of u_t^* in (5), the date of break $\hat{\lambda} = (\hat{T}_1/T, ..., \hat{T}_k/T)$ and $(\hat{T}_1, ...\hat{T}_k)$ are obtained using the dynamic algorithm proposed in Bai and Perron (2003).

The Arai and Kurozumi (2007) test is restrictive in the sense that only a single structural break is considered under the null hypothesis. Hence, the test may tend to reject the null of cointegration when the true data generating process exhibits cointegration with multiple breaks. To avoid this problem, Kejiriwal (2008) has recently extended their test by incorporating multiple breaks under the null hypothesis of cointegration. The Kejiriwal (2008) test of the null of cointegration with multiple structural changes is denoted with k breaks as $\tilde{V}_k(\hat{\lambda})$.

4 An application to the Spanish environmental Kuznets curve

In this section, we re-examine the issue of the EKC using instability tests to account for potential breaks in the long run relationship between per-capita CO_2 and y_t per-capita income as well as the cointegration tests with multiple breaks. First, we use unit root tests to verify that per-capita CO_2 and per-capita income are individually integrated of order one. Second, we test the stability of the per-capita CO_2-per-capita income relationship (and select the number of breaks) using the test proposed in Kejriwal and Perron (2008, 2010). Second, we verified that the variables are cointegrated with tests for the presence/absence of cointegration allowing a single or multiple structural changes in the coefficients as proposed by Arai and Kurozumi (2007) and Kejriwal (2008), respectively. Finally, we estimate the model incorporating the breaks in order to study if per-capita CO_2 and per-capita income relationship (the slope parameter β) have altered over time.

We use time-series data on the Spanish economy spanning from 1857 to 2007. The length of this database makes it especially suitable for the econometric approach adopted in this paper. The data and sources are: a) population: 1857-1990, from Carreras, A. and Tafunell, X. (2005), Table 2.5, and 1991-2007 from INE (2010), Table 2.1.8; b) real GDP: 1857-2000, from Carreras and Tafunell, X. (2005), Table 17.6, and 2001-2007, from Banco de España (2010), Table 1.3; c) total fossil fuel CO_2 (metric tonnes) from Boden, T.A et al, (2010). The evolution of the two series, per-capita CO_2, CO_{2t}, and per-capita income, y_t, is shown in Figure 1 and there seems to be a close comovement between the two series. However, the plots also suggest that association per-capita CO_2 and per-capita income may have altered over time.

The first step in our analysis is to examine the time series properties of the series by testing for a unit root over the full sample. We have used the modified tests proposed by Ng and Perron (2001), denoted M-tests, that are based on standard tests statistics for a unit root with improved size and power properties[4] and the modified version of the feasible point optimal test statistic, P_T, of Elliot et al. (1996). Namely, this method consists of a class of modified tests, called $\bar{M}^{GLS}_{MAIC}$, originally developed in Stock (1999) as M tests, with GLS detrending of the data as proposed in Elliot et al. (1996), and using the Modified Akaike Information Criteria ($MAIC$).[5] Also, Ng and Perron (2001) have proposed a similar procedure to correct for the problems of the standard Augmented Dickey-Fuller (ADF) test, ADF^{GLS}_{MAIC}.[6]

In Table 1 we report the results of the $\bar{M}^{GLS}_{MAIC}$ tests and the ADF^{GLS}_{MAIC} test. In all these tests the null hypothesis is that a series is I(1) against the alternative that it is I(0).[7] Our results clearly reject the existence of two unit

[4] The majority of the conventional unit root tests suffer from three problems. First, many tests have low power when the root of the autoregressive polynomial is close to, but less than, unit (Dejong *et al.,* 1992). Second, the majority of the tests suffer from severe size distortions when the moving-average polynomial of the first-differenced series has a large negative autoregressive root (Schwert, 1989; Perron and Ng, 1996). Third, the implementation of unit root tests often needs the selection of an autoregressive truncation lag, k; however, as discussed in Ng and Perron (1995) there is a strong association between k and the severity of size distortions and/or the extend of power loss.

[5] These tests are the $\bar{M}Z^{GLS}_{\alpha}$, $\bar{M}SB^{GLS}$ and $\bar{M}Z^{GLS}_{t}$.

[6] See Ng and Perron (2001) and Perron and Ng (1996) for a detailed description of these tests.

[7] Note that for the $\bar{M}SB^{GLS}$ test, the null hypothesis is rejected in favour of stationarity when the estimated value is smaller than the critical value.

roots for CO_{2t} and y_t, at the usual significance levels. The null hypothesis of non-stationarity for the two series in levels can not be rejected in any of the tests applied. Consequently, we can conclude that both variables are I(1).

Once the order of integration of the series has been analyzed, we will estimate the long-run or cointegration relationship between CO_{2t}, and y_t. Given the relatively small sample size, we will estimate and test the coefficients of the cointegration equation by means of the Dynamic Ordinary Least Squares (DOLS) method from Saikkonen (1991) and Stock and Watson (1993) and following the methodology proposed by Shin (1994). This estimation method provides a robust correction to the possible presence of endogeneity in the explanatory variables, as well as serial correlation in the error terms of the OLS estimation. Also, in order to overcome the problem of the low power of the classical cointegration tests in the presence of persistent roots in the residuals of the cointegration regression, Shin (1994) suggests a new test where the null hypothesis is that of cointegration. In the first place, we estimate a long-run dynamic equation including the leads and lags of all the explanatory variables, the so-called DOLS regression; in our case (variables in logarithms):

$$CO_{2t} = c + \gamma y_t + \sum_{j=-q}^{q} \gamma_j \Delta y_{t-j} + \upsilon_t \tag{11}$$

Secondly, the Shin test is based on the calculation of two LM statistics from the DOLS residuals, C_μ, to test for deterministic cointegration. The parameter γ is the long-run cointegrating coefficient estimated between per-capita CO_2 and per-capita income (or long-run elasticity).

The results of Table 2 show that the null of deterministic cointegration between CO_{2t} and y_t is not rejected at the 1% level of significance, and the estimated value for γ is 1.37, significantly different from zero at the 1% level. The estimated coefficient shows that the relationship between per-capita income and per-capita CO_2 emissions is very strong. In addition, the coefficient indicates that the elasticity between the two variables is positive and even superior to the unit. Consequently, the per-capita CO_2 emission intensity of the per-capita income even tends to increase as income (or GDP) increases.

Accounting for parameter shifts is crucial in cointegration analysis, which normally involves long spans of data, which are more likely to be affected by structural breaks. Our data covers thirty five years of the history of per-capita CO_2 and per-capita income, during which time the Spanish environmental Kuznets curve has probably changed due to variations in macroeconomic and market forces, such as changes in the structure of the economy, supply shocks, and reforms in the environmental regulation. Therefore, it is important to account for structural breaks in our cointegration relationship.

We now consider the tests for structural change that have been proposed in Kejiriwal and Perron (2010). We use 15% trimming so that the maximum numbers of breaks allowed under the alternative hypothesis is 3. Both the intercept and the slope of equation (11) are allowed to change. Table 3 presents results of stability tests as well as number of breaks selected by the sequential procedure (SP) and the information criteria BIC and LWZ proposed by Bai and Perron (2003). The test results do no suggest any instability (at the 5% level of significance) although the information criterion BIC and LWZ select two breaks and provide evidence against the stability of the long run relationship. Overall, the results of the Kejriwal-Perron tests suggest a model with two breaks estimated at 1941 and 1967 and three regimes,1857-1940, 1941-1966 and 1967-2007. Along the course of economic development if there were no change in the structure or technology of the economy, pure growth in the scale of the economy would result in a proportional growth in pollution, the so-called "scale effect". All in all, the early EKC studies appeared to indicate that local pollutants were more likely to display an inverted U shape relation with income, while global impacts, like carbon dioxide, did not[8], even if multiple structural changes can be present. In our case, there are factors that may explain the placement of such structural changes of the Spanish environmental Kuznets curve that can be matched with different stages of the economic growth path of the Spanish economy. In general, it is a common fact found in the empirical literature available that for CO_2 the turning point is well outside the range of observed income. That means that per capita CO_2 rises in paralell with per-capita income even if the growth pace diminishes over time. Spanish economic growth has been revised in the last years using the annual series of Spanish GDP per head at constant 1995 prices in pesetas provided by Prados de la Escosura (2003). Phases or long swings in which growth rates differ from the long run trend as a result of economic policies, access to international markets, and technological change can be distinguished. His recent estimates on the economic growth show that in Spain there were three main phases in the long-run economic development: 1850-1950, 1951-1974 and

[8]This picture fits environmental economic theory – local impacts are internalized in a single economy or region and are likely to give rise to environmental policies to correct the externalities on pollutees before such policies are applied to globally externalized problems.

1975-2000. According to estimates of Prados de la Escosura, Spain underperformed over the long run mostly due to its sluggish growth in the hundred years up to 1950[9]. This coincides aproximatelly with the first period found in our analysis, namely 1857-1940. These findings are compatible with our reported results for two reasons. First, Prados de la Escosura also remarks the existence of a shift to a lower level in the Spanish growth during the first period as a consequence of the Civil War of 1936-1939, which coincides with our first break, and second, because we are focusing not just on the evolution of a single variable (namely, per capita GDP) but on the relationship among the per-capita CO_2, CO_{2t}, and per-capita income, y_t, which justifies a certain lagged behaviour[10]. Bearing the latter in mind, and considering that in Spain, as in other countries in the European Periphery during the Golden Age (1950-1973), the main spurt of economic growth was delayed until the 1960s, this could be the explanatory reason for the second break found in the middle of the 60's. According to Prados de la Escosura (2003, 2007) and Prados de la Escosura and Rosés (2009) the 1940s constituted a phase of delay in the Spanish economy up to beginning of the 60's when the effects of the Stabilization Plan implemented in 1957 gave clear positive results in terms of growth. Finally, the opening up period started in that decade has given rise to a positive real convergence process in terms of modernization of the Spanish economy and acceptance of the EU regulations in different fields, as the environment, that has continued up to now. Proponents of the EKC hypothesis argue that at higher levels of development, structural change towards information-intensive industries and services, coupled with increased environmental awareness, enforcement of environmental regulations, better technology and higher environmental expenditures, result in leveling off and gradual decline of environmental degradation. The move towards a pro-market attitude with de-regulation and the gradual opening up of Spain to the international economy resulted in sustained growth and catching up with Western Europe during the second half of the twentieth century (Prados de la Escosura, 2007, 2008).

Since the above reported stability tests also reject the null coefficient stability when the regression is a spurious one, we still need to confirm the presence of cointegration among the variables. We use the residual based test of the null of cointegration against the alternative of cointegration with a unknown multiple breaks proposed in Kejriwal (2008), $\tilde{V}_k(\hat{\lambda})$.

[9] Higher destruction of human rather than of physical capital during the Spanish Civil War (1936-1939) and its aftermath explain its performance during the 1940s and 1950s.

[10] The idea behing this lagged behaviour is that the intensity in the pollutants consumption only change once a clear shift in the GDP growth has stabilized over time.

Arai and Kurozumi (2007) show that the limit distribution of the test statistic, $\tilde{V}_k(\hat{\lambda})$, depends only on the timing of the estimated break fraction $\hat{\lambda}$ and the number of I(1) regressors m. In our case (two breaks model), critical values are obtained for $\hat{\lambda}_1 = 0.56$, $\hat{\lambda}_2 = 0.73$ and $m = 1$ by simulation using 500 steps and 2000 replications. The Wiener processes are approximated by partial sums of $i.i.d.$ $N(0,1)$ random variables. Since we are interested in the stability of per-capita income coefficient, γ, we consider only model 3 that permits the slope shift as well a level shift. Table 4 shows the results of the Arai-Kurozumi-Kejiriwal cointegration tests allowing two breaks. Again, the level of trimming used is 15%. We find that the test $\tilde{V}_2(\hat{\lambda})$ cannot reject the null of cointegration with two structural breaks at the 1% level of significance. Then, we conclude that CO_{2t} and y_t are cointegrated with two structural changes estimated at 1941 and 1967.

In order to compare coefficient estimated obtained from a break model with those reported from a model without any structural break, we proceed to estimate the cointegration equation (11) for the three sub-samples, and the results are shown in the last three columns of Table 2. The results of C_μ statistics show that the null of deterministic cointegration between CO_{2t} and y_t is not rejected at the 1% level of significance in the three regimes.

The coefficient estimated between per capita CO_2 and per-capita income (or long-run elasticity) in a two breaks model show a tendency to decrease over time (2.67, 1.10 and 0.56). Thus the coefficient in the third regime (1967-2007) is much smaller than the full sample value (1.37). This suggest that ignoring structural changes in the long-run cointegration relationship may overstate the extend of correlation between per capita CO_2 and per-capita income. As we have already stated earlier in the introductory section of the paper, the EKC is a hypothesized relationship between various indicators of environmental degradation and income per capita. Looked at in a dynamic perspective, the EKC implies that in the early stages of economic growth degradation and pollution increase, but beyond some level of income per capita the trend reverses, or at least softens so that at high-income levels economic growth leads to environmental improvement in relative terms. This implies that even if the environmental impact indicator is not an inverted U-shaped function of income per capita, and environmental degradation (per capita CO_2 consumption) is monotonically rising in income, the "income elasticity" is less than one and is not a simple function of income alone.

5 Concluding remarks

Accounting for parameter shifts is crucial in cointegration analysis, which normally involves long spans of data, and therefore are more likely to be affected by structural breaks. In this paper we extend the existing empirical analysis of the linear model of the EKC in two ways. First, to avoid the econometric problems mentioned in previous empirical literature, we make use of recent developments in cointegrated regression models with multiple structural changes. Specifically, we use a new approach developed by Kejriwal and Perron (2008, 2010) to test for multiple structural changes in cointegrated regression models. They propose a sequential procedure that not only enables detection of parameter instability in cointegration regression models but also allows a consistent estimations of the number of breaks present. Furthermore, we test the cointegrating relationship when multiple regime shifts are identified endogenously. In particular, the nature of the long run relationship between per-capita CO_2-per-capita income is analyzed using the residual based test of the null hypothesis of cointegration with a single or multiple breaks proposed in Arai and Kurozumi (2007) and Kejriwal (2008), respectively. Second, a common criticism to most test of the EKC is that the econometric procedures used require a large number of observations. Accordingly, in this paper we use a long span of the data (1857-2007) for the Spanish economy.

The results are consistent with the existence of linear cointegration between per-capita CO_2 and per-capita income, with a vector (1, -1.37). However, our empirical results show also that the cointegrating relationship has changed over time. In particular, the Kejriwal-Perron tests for testing multiple structural breaks in cointegrated regression models would suggest a model of two regimes, with the dates of the break estimated at 1941 and 1967.

There are factors that may explain the placement of such structural changes of the Spanish EKC that can be matched with different stages of the economic growth path of the Spanish economy. According to Prados de la Escosura (2003, 2007) the 1940s constituted a phase of delay in the Spanish economy up to beginning of the 60's when the effects of the Stabilization Plan implemented in 1957 gave clear positive results in terms of growth.

Second, the estimate of long-run elasticity between per capita CO_2 and per-capita income in a two breaks model shows a tendency to decrease over time (2.67, 1.10 and 0.56). In fact, this long-run elasticity is significantly lower in the third regime (1967-2007), showing that changes in per-capita CO_2 would have not been fully adjusted to compensate the behaviour of per-capita income. Looked at in a dynamic perspective, the EKC implies that in the early stages of economic growth degradation and pollution increase, but beyond some level of income per capita the trend reverses, or at least softens so that at high-income levels economic growth leads to environmental improvement in relative terms. This implies that even if the environmental impact indicator is not an inverted U-shaped function of income per capita, it shows a decreasing growth path behaviour.

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Table 1
Ng-Perron tests for a unit roots

I(2) vs. I(1)	Case: $p = 0, \bar{c} = -7.0$			
Variable	$\bar{M}Z_{\alpha}^{GLS}$	$\bar{M}Z_{t}^{GLS}$	$\bar{M}SB^{GLS}$	ADF^{GLS}
ΔCO_{2t}	-48.95***	-4.94***	0.101	-6.36***
Δy_t	-23.62***	-3.36***	0.142	-3.57***
I(1) vs. I(0)	Case: $p = 1, \bar{c} = -13.5$			
Variable	$\bar{M}Z_{\alpha}^{GLS}$	$\bar{M}Z_{t}^{GLS}$	$\bar{M}SB^{GLS}$	ADF^{GLS}
CO_{2t}	0.23	0.13	0.599***	0.11
y_t	0.83	0.44	0.537***	0.63

Notes:

[a] A *, ** and *** denote significance at the 10%, 5% and 1% levels, respectively.

[b] The $MAIC$ information criteria is used to select the autoregressive truncation lag, k, as proposed in Perron and Ng (1996). The critical values are taken from Ng and Perron (2001), table 1:

Critical values:	Case: $p = 0, \bar{c} = -7.0$			Case: $p = 1, \bar{c} = -13.5$		
	10%	5%	1%	10%	5%	1%
$\bar{M}Z_{\alpha}^{GLS}$	-5.7	-8.1	-13.8	-14.2	-17.3	-23.8
$\bar{M}SB^{GLS}$	0.275	0.233	0.174	0.185	0.168	0.143
$\bar{M}Z_{t}^{GLS}, ADF^{GLS}$	-1.62	-1.98	-2.58	-2.62	-2.91	-3.42

Table 2
Estimation of long-run relationships: Stock-Watson-Shin cointegration tests

Parameter estimates	Full sample 1857-2007	First regime 1857-1940	Second regime 1941-1966	Thrid regime 1967-2007
c	-12.0	-21.5	-9.47	-4.45
	(-7.7)	(-5.9)	(-6.1)	(-5.5)
γ	1.37	2.67	1.10	0.56
	(6.6)	(5.4)	(5.3)	(6.8)
$\bar{R}^2$	0.96	0.98	0.99	0.98
C_μ	0.307	0.131	0.138	0.136

Notes:

[a] t-statistics are in brackets. Standard Errors are adjusted for long-run variance. The long-run variance of the cointegrating regression residual is estimated using the Barlett window which is approximately equal to $INT\left(T^{1/2}\right)$ as proposed in Newey and West (1987).

[b] We choose $q = INT\left(T^{1/3}\right)$ as proposed in Stock and Watson (1993).

[c] C_μ and C_τ are LM statistics for cointegration using the DOLS residuals from deterministic and stochastic cointegration, respectively, as proposed in Shin (1994).

[d] The critical values for C_μ are taken from Shin (1994), table 1, from $m = 1$:

Critical values:			
	10%	5%	1%
C_μ	0.231	0.314	0.533

Table 3
Kejriwal-Perron tests for testing multiple structural breaks in cointegrated regression models: equation (5) and (11)[a,b]

	Specifications[a]		
$d_t = \{CO_{2t}\}$	$z_t = \{1, y_t\}$	$x_t = \{\emptyset\}$	$M = 3$
	$q = 2$	$p = 0$	$h = 21$
	Tests[b,c]		
$\sup F_T(1)$	$\sup F_T(2)$	$\sup F_T(3)$	$UD\max$
8.72*	5.72	4.48	8.72*
	Number of Breaks Selected		
SP	0		
LWZ	2		
BIC	2		
	Breaks		
$\hat{T}_1$	1941	$\hat{T}_2$	1967

Notes:

[a] d_t, z_t, q, p, h, and M denote the dependent variable, the regressors, the number of I(1) variables (and the intercept) allowed to change across regimes, the number of I(0) variables, the minimum number of observations in each segment, and the maximum number of breaks, respectively.

[b] *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively. The critical values are taken from Kejriwal and Perron (2010), Table 1.10 (critical values are available on Kejriwal-Perron website), trending case with $q_b = 1$.

Table 4
Arai-Kurozumi-Kejriwal cointegration tests with two structural breaks break: equation (5) and (11)

Test $\tilde{V}_2(\hat{\lambda})^a$	$\hat{\lambda}_1$	$\hat{T}_1$	$\hat{\lambda}_2$	$\hat{T}_2$
0.124***	0.56	1941	0.73	1967

Notes:
[a] *, **, and *** denote significance at the 10%, 5%, and 1% levels, respectively.
[b] Critical values are obtained by simulation using 500 steps and 2000 replications. The Wiener processes are approximated by partial sums of *i.i.d.* $N(0,1)$ random variables.

Critical values:	10%	5%	1%
$\tilde{V}_k(\hat{\lambda})$	0.084	0.109	0.177

Figure 1. Evolution of per-capita income and per-capita CO2