# Revisiting carbon Kuznets curves with endogenous breaks modeling: Evidence of decoupling and saturation (but few inverted-Us) for individual OECD countries

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## ABSTRACT

This paper tests for a carbon Kuznets curve (CKC) by examining the carbon emissions per capita-GDP per capita relationship individually, for 21 OECD countries over 1870-2010 using a reduced-form, linear model that allows for multiple endogenously determined breaks. This approach addresses several important econometric and modeling issues, e.g., (i) it is highly flexible and can approximate complicated nonlinear relationships without presuming a priori any particular relationship; (ii) it avoids the nonlinear transformations of potentially nonstationary income. For 10 of 14 countries that were ultimately estimated, the uncovered emission-income relationship was either (i) decoupling—where income no longer affected emissions in a statistically significant way, (ii) saturation—where the emissions elasticity of income is declining, less than proportional, but still positive, or (iii) no transition—where the emissions elasticity of income is (or very near) unity. For only four countries did the emissions-income relationship become negative—i.e., a CKC. In concert with previous work, we conclude that the finding of a CKC is country-specific and that the shared timing among countries is important in income-environment transitions.

Keywords: $CO_2$ emissions; Environmental Kuznets curve; OECD countries; nonlinear flexible form; multiple endogenous breaks; income-emissions elasticities.

JEL classifications: C18, C22, C50, O44, Q43, Q56.

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## 1. Introduction

Whether pollution first rises with income and then falls after some threshold level of income/development is reached—thus forming an inverted U-shaped relationship also called an Environmental Kuznets Curve (EKC)—is one of the most popular questions in environmental economics (e.g., see reviews by Dinda 2004; Stern 2004). Such EKC analyses typically employ panel data and most often focus on emissions per capita. Those emissions are modelled as a quadratic (or sometimes cubic) function of GDP per capita; an EKC between emissions per capita and income is said to exist if the coefficient for GDP per capita is statistically significant and positive, while the coefficient for its square is statistically significant and negative.

One might expect not to find such an inverted-U relationship for carbon dioxide emissions—a global, stock pollutant, whose (uncertain) damages will occur in the future. Yet, several studies have calculated within sample turning points for carbon emissions per capita for either multiple-country panels (e.g., Schmalensee et al. 1998; Agras and Chapman 1999; Martinez-Zarzoso and Bengochea-Morancho 2004; and Galeotti et al. 2006) or for individual countries (e.g., Schmalensee et al. 1998; Dijkgraaf and Vollenbergh 2004; and Azomahou et al. 2006).

It is important to note that an inverted-U relationship (or EKC) between emissions and income means that the income elasticity of emissions is negative for countries in the highest income segment. However, if the income elasticity of emissions declines with income but remains positive (a phenomenon determined in Liddle 2013; Liddle 2015), emissions and income unambiguously have a monotonic relationship, i.e., an EKC is rejected. A declining and less than unity income elasticity suggests that the $CO_2$ intensity (emissions per GDP) follows an inverted-U path (a pattern found for high-income countries in Lindmark 2004).

Not surprisingly, the large EKC literature has generated substantial criticism.[1] Stern (2004) argued that many EKC studies risked spurious findings by ignoring that variables like emissions per capita and GDP per capita are likely nonstationary; later, Muller-Furstenberger and Wagner (2007) argued further that even the EKC studies that did recognize the stationarity properties in the data still risked spurious findings by performing nonlinear (quadratic) transformations of a nonstationary variable (GDP per capita).[2] In addition, Muller-Furstenberger and Wagner (2007) and Wagner (2008) claimed that the studies to date that have employed panel unit root and panel cointegration techniques have relied on methods that incorrectly assume that the cross-sections are independent. Yet, despite that cross-sectional dependence, Dijkgraff and Vollenbergh (2005) rejected panel homogeneity even for OECD countries. Lastly, the polynomial of GDP per capita model (either quadratic or cubic) used in the EKC literature has been criticized for being highly inflexible and for rendering unimportant feasible emissions-GDP relationships for which it cannot test (Lindmark 2004; Liddle 2013). For example, the typical polynomial model does not allow for the possibility (i) that GDP elasticities are significantly different across development levels but still (*forever*) *positive*—i.e., a saturation effect or S-curve;[3] or (ii) that at high levels of GDP per capita the relationship with emissions is insignificant—i.e., a decoupling of the emissions-GDP relationship.

This paper tests for a so-called carbon Kuznets curve (CKC) by examining the $CO_2$ emissions per capita-GDP per capita relationship, individually, for several OECD countries over

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[1]Of course, there are theoretical criticisms of the EKC literature (e.g., Muller-Furstenberger and Wagner 2007; Carson 2010) and additional econometric criticisms (e.g., Stern 2010), which we do not address.

[2] An anonymous referee pointed out that all studies—including Wagner (2008) and this one—divide aggregate GDP and emissions by population, which almost certainly is I(1), and thus, introduce a potential nonlinear transformation.

[3] A quadratic model that produces an out-of-sample turning point implies an S-curve; however, such a model does not allow for the determination of statistically different income elasticities.

the very long-run. A reduced-form, linear model that allows for multiple endogenously determined breaks is used to address the econometric and modeling issues mentioned above—the linear model with multiple breaks is highly flexible and can approximate complicated nonlinear relationships without presuming a priori any particular relationship. Hence, no nonlinear transformations of potentially nonstationary variables are necessary, and the issues of cross-sectional dependence and heterogeneity are avoided/addressed by analyzing each country's emissions-GDP relationship separately. Lastly, by focusing on the time-series data of single countries, we address the crucial question of a specific country's evolution of its income-environment relationship (as recommended by Stern et al. 1996; de Bruyn et al. 1998).

## 2. Previous studies of carbon emissions and breaks (exogenous and endogenous)

While there have been several studies that have employed a structural break methodology, not all have focused on the relationship between carbon emissions and income[4]—a few studies have been concerned only with the behavior of a single series (e.g., carbon emissions). In addition, some studies have focused on single countries.

Moomaw and Unruh (1997) took an individual country approach; they tested the stability of a simple linear relationship between $CO_2$ per capita emissions and GDP per capita for a number of developed countries using data spanning 1950-1992, choosing 1973 as the *a priori* break-date, and employing a standard Chow test for structural change. Moomaw and Unruh rejected the null hypothesis of no structural change, typically finding that individual countries switched from a positive to a negative linear relationship between emissions and income at the time of the first oil crisis.

[4] Liddle and Messinis (2015) used the same methods as here but estimated the relationship between *sulfur* emissions and income for individual OECD countries over the period 1950-2005.

Lanne and Liski (2004) examined the $CO_2$ per capita emissions trends over the period 1870-1998 for 16 early industrialized countries using endogenous methods that allowed for multiple structural breaks. In contrast to Moomaw and Unruh (1997), Lanne and Liski rejected the oil price shocks as events causing permanent breaks in the structure and level of emissions; instead, Lanne and Liski found evidence of downturns in increasing $CO_2$ per capita emissions trends occurring early in the 20th century, and evidence of stable declining per capita emissions for only two countries.

Lindmark (2004) focused on the $CO_2$ intensity ($CO_2$/GDP) trends of 46 countries over 1870-1994 and found that most developed countries had declining intensity trends with typically early breaks. However, the income level at those breaks/turning points varied from 5,000 USD to 10,000 USD. Huntington (2005) used a single break procedure to endogenously determine a break in the carbon emissions-GDP relationship for the US over 1870-1998, and similar to Lanne and Liski (2004), found an early break in 1913. Over those two periods (before and after 1913), Huntington estimated a stable income elasticity of 0.9. Both Lindmark and Huntington emphasized the importance of technological advance rather than smooth CKC-type transitions.

Lastly, Esteve and Tamarit (2012a) analyzed the $CO_2$ per capita and GDP per capita relationship for Spain over 1857-2007 using a cointegration model that allowed for two endogenous breaks (similar to the methods used here). They found (over three regimes) a declining but always positive income elasticity. By contrast, Esteve and Tamarit (2012b)—again, focusing on the Spanish case using the same long time-span—employed a threshold cointegration approach with a single break and found evidence of a recent (beginning in 1986), negative relationship between income and carbon emissions. However, it appears that the threshold cointegration method focuses on the determination of a possible asymmetric error correction or adjustment process between the two variables rather than the estimation of elasticities over different regimes (the concern of Esteve and Tamarit 2012a and the present paper).

3. Previous studies considering flexible forms of the carbon emissions-income relationship

More recently several papers have considered methods that introduce more flexible forms than the typical polynomial model. Yet, some of these methods still require the nonlinear transformation of potentially integrated income (e.g., Galeotti et al. 2006; Wang 2013). Papers that do avoid the nonlinear transformation of income while maintaining a fully flexible model form typically fall into two categories. A first group uses fully nonparametric or semi-parametric methods; thus, that group displays plots (with bootstrapped confidence intervals) of the estimated relationship (e.g., Azomahou et al. 2006; Bertinelli and Strobl 2005), rather than includes explicitly determined elasticities with accompanying efficient standard errors (as in parametric estimations). The second group performs linear spline or additive mixed model regressions (e.g., Schmalensee et al. 1998; He and Richard 2010; Zanin and Marra 2012; Liao and Cao 2013).

Yet, both groups have shortcomings/limitations. The piece-wise linear spline model is very data intensive, and so, perhaps most appropriate for panel data. That approach also requires a large number of “pieces” (or income groupings) to be determined exogenously, and those income groupings are the same for all countries (in a panel analysis); hence, the spline model contrasts with the endogenous breaks method employed here, which does not require/force those breaks to occur at any particular income level or time. Furthermore, semi- and nonparametric methods (as well as spline and additive mixed models) must account for nonstationarity (as parametric methods must). But nonparametric methods that fully account for nonstationarity and cointegration are in their infancy (e.g., Chen et al. 2012; Chen et al. 2013), and such methods are certainly not as far along in addressing nonstationarity and cointegration as are parametric methods; hence, the robustness of the previous nonparametric CKC work is not clear.

Besides nonparametric estimations, He and Richard (2010), who analyzed Canada, employed the nonlinear flexible parametric approach of Hamilton (2001). Yet again, the proofs in Hamilton (2001) assumed stationarity. Lastly, the current state of knowledge in the literature seems to be that any determination of delinking or negative relationship between pollution and income likely is a product of “country-specific characteristics such as technological progress, structural evolution, or external shocks” (He and Richard 2010, p. 1084); thus, our proposed approach should be particularly appropriate since it both (i) explicitly estimates regime change (as opposed to the typically smooth transitional estimations of nonparametric models), and (ii) focuses on individual country estimates.[5]

## 4. Data and Methods

### 4.1. Data

We analyze the $CO_2$ emissions per capita and real GDP per capita relationship for 21 advanced/OECD countries.[6] Figures 1 and 2 plot for those countries the long-run (1870-2007) $CO_2$ emissions per capita and GDP per capita series, respectively, in natural logs. The figures clearly indicate why the consideration of breaks is important: for all countries the emissions series display substantial breaks around the two World Wars (e.g, 1914-1921 and 1943-1945); in addition to breaks in $CO_2$ emissions per capita during those two periods, all countries display a substantial break in GDP per capita around the Great Depression (e.g., 1930-1939). Hence, if we

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[5] While most studies using nonparametric methods have considered panels, some nonparametric studies have focused on individual countries (e.g., He and Richard 2010; Zanin and Marra 2012).

[6] Those countries are Australia, Austria, Belgium, Canada, Denmark, Finland, France, Germany, Greece, Ireland, Italy, Japan, Netherlands, New Zealand, Norway, Portugal, Spain, Sweden, Switzerland, UK, and USA.

restrict our analysis to allow for no more than two endogenous breaks, such breaks likely would be calculated to occur before 1950 for most countries. However, the period beginning in 1950—an era of substantial economic growth and development for the countries considered—is exactly the time in which we might expect to observe emissions-GDP transitions. Therefore, we allow for three endogenous breaks over 1870-2010,[7] and use $CO_2$ emissions per capita data from the Carbon Dioxide Information Analysis Center (Boden et al. 2013) and real GDP per capita data from Angus Maddison (http://www.ggdc.net/). (Both series are transformed into natural logs.)

Figures 1 and 2

4.2 Unit root tests with endogenous breaks/nonlinearities

Again, this paper is about estimating a potentially nonlinear relationship between income and carbon emissions; it is not about determining the order of integration of those series. Interestingly, there is a literature that seeks to determine those integration properties, and that literature has produced highly mixed/conflicting results; for example, see Romero-Avila (2008); Lee and Lee (2009), regarding carbon emissions per capita, and Ben-David et al. (2003); Gaffeo et al. (2005), regarding GDP per capita. While we have no designs to contribute to the debate on the true integration properties of carbon emission per capita and GDP per capita, it is necessary to perform unit root tests before one can employ the estimation methods that we determined are most appropriate. Furthermore, given the nature of the series we consider, we believe it important that unit root tests account for endogenous breaks of potentially different magnitudes, as well as other potential forms of nonlinearity.

In order to avoid spurious rejections when an unknown break is present under the null hypothesis, several unit root test procedures that allow for endogenous breaks have been

[7] The series begin in 1892, 1924 and 1878 for Greece, Ireland, and New Zealand, respectively.

developed, among them are: (1) an ADF-type unit root test that models breaks as innovation outliers by Narayan and Popp (2010), NP, and (2) the quasi-GLS detrending approach that treats breaks as additive outliers, as in Carrion-i-Silvestre et al. (2009), CKP. Both NP and CKP allow for structural breaks in both the null and the alternative hypotheses, but assume—in the case of two breaks—that both breaks are of the same magnitude. However, Harvey et al. (2013), HLT, challenged that homogeneity of break magnitudes assumption, and developed a test (based on the CKP procedure) that allows for breaks of different magnitudes. Note that while the HLT (2013) test only considers trend shifts, it is based on GLS detrending, and thus, is asymptotically robust to level breaks (or "slowly evolving trends"), in the same way the Elliott et al. (1996) test is. Because in finite samples, large level breaks could have an impact, we consider also the NP test, which allows for two structural breaks in both the level and the slope of the trending series.

The HLT test examines a time series, $y_t$:

$$y_t = \mu + \beta t + \gamma' \mathrm{DT}_t(\tau_0) + \mathrm{u}_t, \qquad \mathrm{t} = 1, \ldots, \mathrm{T} \tag{1}$$

where $\mathbf{DT}_{t}(\tau_0)$ is a vector of indicator variables, $1(t > [\tau_0 T])(t - [\tau_0 T])$, T is the sample size, $\tau_0 = [\tau_{0,1}, \ldots, \tau_{0,m}]$ is a set of sample fractions, *m* is the maximum number of unknown breaks, $\gamma = (\gamma_1, \ldots, \gamma_m)$ are parameters associated with breaks, and $u_t = \rho_T u_{t-1} + \varepsilon_t$ is a mean zero stochastic error process. A trend break in series $y_t$ occurs at time $[\tau_{0,i}T]$ when $\gamma_i \neq 0$ $(i = 1, \ldots, m)$, and it is assumed that the break fractions $\tau_{0,i} \in \Lambda$ for all i where $\Lambda = [\tau_L, \tau_U]$, $0 < \tau_L < \tau_U < 1$ and $\tau_L, \tau_U$ are trimming fractions. The unit-root null hypothesis $H_0$: $\rho_T = 1$ against the alternative $H_1$: $1 - c/T$, $c > 0$ and the test statistic is $MDF_m = \inf DF^{GLS}(\tau)$, where $DF^{GLS}(\tau)$ is the standard t-ratio associated with $\phi$ in the fitted ADF equation: $\Delta u_t = \phi u_{t-1} + \sum \psi_j \Delta u_{t-j} + e_t$ obtained from a local GLS regression of $y_{\bar{\rho}} = [y_1, (y_2 - \bar{\rho} y_1), \ldots, (y_T - \bar{\rho} y_{T-1})]'$ on $Z_{\bar{\rho},\tau} = [z_1, (z_2 - \bar{\rho} z_1), \ldots, (z_T - \bar{\rho} z_{T-1})]'$ where

$z_\tau := [1, \tau, DT_\tau(\tau)']'$ with $\bar{\rho} := 1 - \bar{c}/T$ for some $\bar{c} > 0$. Harvey et al. (2013) reiterate the Kejriwal and Perron (2010) point that *m* must be determined in relation to the sample size to avoid power and/or size issues.

Linear (i.e., autoregressive) unit root tests suffer from size distortion in the presence of GARCH or non-stationary volatility. In the spirit of Cavaliere and Taylor (2009), Su et al. (2014), SCR, overcome such size effects by means of a bootstrap method that replicates in the resampled sample data the heteroskedasticity in the original data. The method is robust to heteroskedasticity and achieves the size-corrected power of the usual unit root tests. More precisely, in the data generating integrated process of $y_t = y_{t-1} + \varepsilon_t$ with heteroskedastic errors: $\varepsilon_t = \omega_t \sigma_t$ where t=1,…,T and $\sigma_t$ is iid N(0,1) and $\omega_t$ takes various forms that include a maximum of two breaks. The procedure assesses the volatility of variable $y_t$ by estimating the variance of the series, η(s) by $\hat{\eta}(s) = \dfrac{\sum_{t=1}^{sT} \Delta\hat{y}_t^2 + (\mathrm{sT} - [\mathrm{sT}])\Delta\hat{y}_{[sT]+1}^2}{\sum_{t=1}^{T} \Delta\hat{y}_t^2}$ where $\Delta\hat{y}_t$ is the first difference of the detrended series, $\hat{\eta}(s)$ is a consistent estimator for $\eta(s)$.

So in an attempt to account for several possibilities: (i) multiple, endogenous breaks in trend of different magnitudes; (ii) multiple, endogenous breaks in both level and trend; and (iii) nonlinear nonstationary volatility, we consider the HTL, NP, and SCR unit root tests. If two of the three tests suggest stationarity, we judge that series to be stationary. If as a result of that judgment, the carbon and income series are of different orders of integration for a country, we do not analyze that country's series further.

4.3 Optimal timing of breaks and cointegration tests and estimation with endogenous breaks

Bai and Perron (1998, 2003) developed a method that allows for multiple endogenous structural breaks in stationary, trending regressors. To determine the timing of breaks Bai and

Perron (1998, 2003) recommend focusing on two statistics: (i) the sup$F_T(k)$ test for the null hypothesis of no structural break against the alternative of a fixed number of $k$ breaks; and (ii) sup$F(l+1|l)$ test, which is a sequential test of the null hypothesis of $l$ break(s) against the alternative of $l + 1$ breaks. The sup$F_T(k)$ test determines whether at least one break is present; if that test indicates the presence of at least one break, then the number of breaks, $m$, is revealed by the sequential examinations of the second set of tests, so that sup$F(l+1|l)$ are insignificant for $l >= m$. The Bai and Perron method determines the break points by a global minimization of the sum of squared residuals. The procedure concludes in favor of a model with $(l+1)$ breaks if the overall minimal value of the sum of squared residuals (over all segments where an additional break is included) is sufficiently smaller than the sum of squared residuals from the $l$ break model (Bai and Perron 1998).

Kejriwal and Perron (2010) updated the Bai and Perron sequential method of endogenous breaks timing to be valid for $I(1)$ cointegrated regressors. Kejriwal (2008) further modified the residual based test of the null hypothesis of cointegration with structural breaks proposed in Arai and Kurozumi (2005) to incorporate multiple breaks under the null hypothesis (K-AK test). Kejriwal (2008) also augmented the cointegrating equation with leads and lags of the first differences of the $I(1)$ regressors to address potential endogeneity. Since Kejriwal (2008) is particularly interested in estimating cointegrating relationships that have changed because of structural breaks (as are we), Kejriwal chose cointegration as the null hypothesis and used the Kejriwal and Perron (2010) sequential instability test along with a modified Schwarz criterion (LWZ) to first ensure the existence of breaks.

Yet, the Kejriwal and Perron (2010) instability test may reject the null of coefficient stability when the regression is a spurious one, i.e., not cointegrated; hence, the Kejriwal (2008) cointegration test with multiple breaks is used to confirm the presence of cointegration, i.e., reject the possibility of a spurious relationship. That test considers the relation

$$y_t = c_i + z_t^{'}\delta_i + \sum_{j=-p}^{p} \Delta z_{t-j}^{'}\Pi_j + \mathrm{u}_t, \qquad if \;\; T_{t-1} < t < T_t \tag{2}$$

for i=1,…,k+1, where k is the number of breaks, $z_t$ is a vector of I(1) variables, $T_0 = 0$, $T_{k+t} = T$, and the third term on the right-hand-side of the equation includes p number of lags and leads of the first difference of the regressors to account for the potential of endogeneity. The resulting test statistic is defined as:

$$V_k(\lambda) = \mathrm{T}^{-2} \sum_{t=1}^{T} S_t(\lambda)^2 / \Omega_{i,j} \tag{3}$$

where $\lambda_i = (T_1/T,…,T_k/T)$, i.e., the sample fractions associated with i=1,…,k breaks, $\Omega_{i,j}$ is the long-run variance of $u_t$ for j=1,…,k, and $T_1,…,T_k$ are recovered from dynamic programming, as in Bai and Perron (2003).

Since the cointegration test is a confirmatory test, for each cross-section, only the number and timing of breaks determined by the sequential procedure and information criteria are considered in the cointegration test. If cointegration is confirmed, the different regimes are estimated similarly by assuming the previously determined number and timing of breaks.

## 5. Results and discussion

Table 1 presents the results for the various unit root tests. Those test results suggest that for most countries the two series are *I*(1); thus, we proceed to the Kejriwal and Perron (2010) stability test and the Kejriwal (2008) K-AK cointegration test for those countries. On the other hand, for Finland, Germany, Greece, Japan, Portugal, and Sweden, the two series are of different order of integration; hence, we do no further analysis on them.

Table 1

Again, to determine the number and timing of breaks, we consider two information/decision criteria, i.e., the sequential method of Kejriwal and Perron (2010) and the LWZ criterion. If the sequential method did not determine a break, we went with the number of breaks determined by the LWZ (as in Kejriwal 2008). Table 2 displays the timing of the breaks and the cointegration results; only for Austria was the null hypothesis of cointegration rejected. Thus, we proceed to the estimation with breaks for the other 14 OECD countries.

Table 2

Table 3 presents the results for the regressions under breaks. If we focus on the sign and significance of the income term's coefficient (the $\delta$s in Table 3), the most common income-emissions relationship is that of saturation—a statistically significant, declining (or at least significantly less than unity), but still positive income elasticity; that relationship is the clear case for five countries—Australia, Canada, Ireland, Italy, and Norway. Since carbon emissions are so associated with energy consumption, perhaps a saturation pattern is to be expected. (Saturation in carbon's income elasticity is the same pattern uncovered in the panel analysis of Liddle 2015.)

Table 3

Belgium, Netherlands, and US display decoupling of income and emissions after a break—as the income elasticity is no longer significant. Only four countries (Denmark, France, Switzerland, and UK) show clear evidence of a carbon Kuznets curve—a significant, negative relationship between income and emissions. Two countries—New Zealand and Spain—display no transition, since their income elasticity is indistinguishable from unity after their final break.

Next, we compare our results to the few recent papers that also use flexible form approaches and focus on individual country estimations. He and Richard (2010) found a similar saturation-type relationship for Canada and emphasized the importance of the oil shocks of the

1970s; however, we estimated much earlier breaks for Canada (1893 and 1918). Fosten et al. (2012), using different methods (nonlinear threshold cointegration without adjustment/concern for the nonlinear transformation of integrated income), also determined an inverted-U for the UK. Zanin and Marra (2012) considered several of the same countries we initially consider. They also found an inverted-U for Switzerland, and predicted (in Figure 2 of that paper) similarly declining income elasticities (but not inverted-Us) for Canada and Italy. For Spain Zanin and Marra predicted an oscillating but declining income elasticity. Esteve and Tamarit (2012a) focused on Spain over 1857-2007 using the same methods we do (but allowing for a maximum of two breaks, as opposed to three). They calculated a three-regime saturation pattern for Spain (as opposed to a four regime pattern here) with an income elasticity of 0.56 over the final regime (1967-2007). We calculated a similarly declining elasticity over time; however, we calculated a final elasticity of unity over 1939-2010.

The final way we can analyze the results shown in Tables 2 and 3 is to focus on the timing of the break dates. There were 37 breaks identified in Table 2. Six breaks occurred during World War I (1914-1918) and roughly 10 breaks (9 if Belgium's 1948 break is not counted) occurred during the Great Depression through World War II (1929-1945). Such a high number of breaks during these periods is not surprising given Figures 1-2. If we focus on the 12 countries analyzed in Table 3 that exhibited a clear regime change (i.e., excluding New Zealand and Spain), and if we consider their final break as most important, then 10 of these 12 most important breaks occurred between 1968-1981 (Table 2 displays the timing of the breaks).

The first oil crisis could be dated 1973–1974;[8] whereas, the second oil crisis, which is dated 1979–1981, corresponded to the fall of the Shah in Iran and the beginning of the Iran–Iraq

[8]OPEC announced on October 15, 1973, their embargo, which would precipitate the first oil crisis that led to a price spike later in 1974; but oil prices already had begun to increase earlier in 1973.

war, and it led to considerably higher prices than the first oil crisis. If we count UK's break in 1972 (since there should be a confidence interval around the break-dates), then six of those 12 most important breaks occurred during the two oil-crisis periods.

Four more of those 12 most important breaks occurred during 1968-1969, and the 1960s through the early 1970s (before the first oil crisis) was a period of heightened environmental awareness/concern in many OECD countries.[9] (The two other breaks occurred in 1918 and 1949.) That shared timing or external shocks have played an important role in apparent inverted-U transitions is a conclusion of previous work on the EKC (Moomaw and Unruh 1997; Volleberg et al. 2009; Stern 2010; He and Richard 2010) and in energy intensity (Liddle 2012).

## 6. Conclusions

We used endogenous breaks modeling to examine the carbon emission-income relationship ultimately for 14 OECD countries. We recommend this approach for studying potential nonlinear relationships because: (i) it does not impose a functional form a priori; (ii) it estimates elasticities for different regimes that are robust to nonstationarity and cointegration; and (iii) it avoids a nonlinear transformation of integrated income. These three issues rarely have been addressed simultaneously in the EKC/CKC literature, and perhaps, never previously addressed in the analysis of several countries (we know only of the Esteve and Tamarit 2012a study of Spain and the Liddle and Messinis 2015 sulfur emissions study). Following several previous studies, the importance of shared timing among countries was uncovered—in particular the increased interest in the quality of the environment in the 1960s and 1970s in OECD

[9] For example, the first Earth Day was held in 1970, and the first United Nations Conference on the Human Environment was held in 1972, which led directly to the creation of several government environmental agencies and the UN Environment Program. Several nongovernmental environmental organizations were established during this period, too, like the World Wildlife Fund in 1961, the Environmental Defense Fund in 1967, Natural Resources Defense Council in 1970, and Greenpeace in 1971. Clean Air Acts were passed in Canada, New Zealand, and UK in 1970, 1972, and 1968, respectively. Lastly, several OECD countries implemented energy/fuel and/or vehicle taxes prior to 1973 (see the OECD/EEA economic instruments database, http://www2.oecd.org/ecoinst/queries).

countries and the oil crises/price spikes of the 1970s and early 1980s. However, it is important to note that for only four countries did the emissions-income relationship become negative—i.e., a CKC. Indeed, the primary emission-income relationships determined here (i.e., for 10 of 14 countries estimated) were either (i) decoupling—where income no longer affected emissions in a statistically significant way, (ii) saturation—where the emissions elasticity of income is declining, less than proportional, but still positive, or (iii) no transition—where the emissions elasticity of income is (or very near) unity.

This lack of inverted-Us provides an interesting contrast to Liddle and Messinis (2015), who focused on sulfur emissions of OECD countries and employed the same methods used here. Liddle and Messinis found inverted-Us for 19 of the 25 OECD countries they studied. Since sulfur has local health and environmental impacts, such a contrast to the income-emissions relationship for the global pollutant, carbon, is not surprising.

Hence, as others have argued (e.g., He and Richard 2010), the finding of a CKC is country-specific. The only generalization about the development process's impact on the carbon emissions-income relationship we can contribute is that, as countries reach higher levels of development, those two variables become either less than proportionally, positively related to each other or no longer strongly related to each other at all.

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Table 1. Unit Root Tests with Breaks or Nonlinearities, 1870-2010

| | HLT (2013) | | SCR (2014) | | NP (2010) | |
|---|---|---|---|---|---|---|
| | **GDP** | **$CO_2$** | **GDP** | **$CO_2$** | **GDP** | **$CO_2$** |
| **Australia** | -3.563 | -3.514 | -1.326 | -2.595 | -5.526* | -3.429 |
| **Austria** | -4.227 | -4.033 | -1.845 | -2.670 | -5.719* | -5.485* |
| **Belgium** | -3.658 | -4.207 | -1.486 | -3.541* | -1.235 | -4.597 |
| **Canada** | -3.956 | -3.801 | -3.381# | -2.758 | -3.400 | -2.732 |
| **Denmark** | -4.167 | -3.387 | -2.359 | -1.034 | -2.066 | -1.901 |
| **Spain** | -2.777 | -2.861 | -1.190 | -2.623 | -1.006 | -1.201 |
| **Finland** | -3.797 | -4.330 | -2.227 | -3.485* | -3.161 | -5.295* |
| **France** | -3.697 | -3.427 | -2.134 | -2.238 | -3.023 | -4.803 |
| **Germany** | -3.857 | -4.589* | -2.330 | -2.945# | -4.659 | -6.139* |
| **Greece** | -5.146* | -2.748 | -1.294 | -2.377 | -5.526* | -3.429 |
| **Ireland** | -3.339 | -3.856 | -2.528 | -1.667 | -5.526* | -3.429 |
| **Italy** | -3.886 | -3.199 | -1.941 | -2.535 | -3.218 | -3.950 |
| **Japan** | -4.886* | -3.490 | -2.366 | -1.638 | -9.369* | -2.073 |
| **Netherlands** | -3.747 | -3.496 | -1.903 | -2.956 | -40.897* | -6.046* |
| **New Zealand** | -3.612 | -3.114 | -2.041 | -3.416 | -5.526* | -3.429 |
| **Norway** | -4.329 | -2.414 | -1.900 | -4.043* | -1.928 | -1.663 |
| **Portugal** | -3.808 | -4.369 | -1.892 | -6.672* | -0.996 | -10.412* |
| **Sweden** | -3.976 | -4.670* | -2.142 | -3.227* | -1.687 | -2.421 |
| **UK** | -3.129 | -3.574 | -2.078 | -3.882 | -2.170 | -4.259 |
| **Switzerland** | -3.369 | -2.808 | -2.785 | -2.165 | -3.027 | -5.435* |
| **USA** | -3.360 | -4.534 | -2.116 | -2.421 | -3.248 | -3.356 |

Notes: * and # indicate 5% and 10% significance levels, respectively. For the HLT test a maximum of two breaks are allowed, and the 5% and 10% critical values are -3.57 and -4.30, respectively. The SCR test is the DF t-test statistic with bootstrapped critical values. The NP test allows for two breaks in levels and the slope of the time trend, and the 5% critical value is -4.937 respectively. Due to missing observations for either GDP or Co2, data for Greece, Ireland and New Zealand begins in 1892, 1924 and 1878 respectively.

Table 2. Optimal number and timing of breaks and K-AK cointegration test with breaks, LN GDP per capita and LN $CO_2$ per capita, 1870-2010.

| | **Optimal # Breaks** | | **K-AK Cointegration Test** | | | | | | | | |
|---|---|---|---|---|---|---|---|---|---|---|---|
| | S | LWZ | $V_1$ | Date | $V_2$ | Date 1st | Date 2nd | $V_3$ | Date 1st | Date 2nd | Date 3rd |
| **Australia** | 0 | 3 | | | | | | 0.06 | 1892 | 1922 | 1949 |
| **Austria** | 1 | 2 | | | 0.10* | 1915 | 1935 | | | | |
| **Belgium** | 0 | 3 | | | | | | 0.07 | 1913 | 1948 | 1981 |
| **Canada** | 0 | 2 | | | 0.06 | 1893 | 1918 | | | | |
| **Denmark** | 0 | 2 | | | 0.02 | 1899 | 1968 | | | | |
| **Spain** | 0 | 3 | | | | | | 0.07 | 1892 | 1919 | 1939 |
| **France** | 0 | 2 | | | 0.08 | 1945 | 1980 | | | | |
| **Ireland** | 0 | 2 | | | 0.01 | 1940 | 1973 | | | | |
| **Italy** | 0 | 3 | | | | | | 0.03 | 1894 | 1942 | 1968 |
| **Netherlands** | 0 | 3 | | | | | | 0.03 | 1916 | 1938 | 1981 |
| **New Zealand** | 0 | 2 | | | 0.06 | 1918 | 1949 | | | | |
| **Norway** | 0 | 3 | | | | | | 0.03 | 1899 | 1939 | 1969 |
| **Switzerland** | 0 | 2 | | | 0.08 | 1940 | 1968 | | | | |
| **UK** | 0 | 2 | | | 0.04 | 1924 | 1972 | | | | |
| **USA** | 1 | 3 | | | | | | 0.04 | 1909 | 1929 | 1980 |

Notes: S=sequential procedure (as described in Kejriwal and Perron 2010). LWZ=Schwarz criterion. * indicates the 5% significance level; where the simulated critical values for $V_2(\hat{\lambda})$ and $V_3(\hat{\lambda})$ are 0.093 and 0.144, respectively. For Ireland and New Zealand the 5% simulated critical values for $V_2(\hat{\lambda})$ are 0.138 and 0.152, respectively. The null hypothesis is cointegration.

Table 3. Regression estimates with breaks, LN GDP per capita & LN $CO_2$ per capita, 1870-2010

| | **Regime 1** | | **Regime 2** | | **Regime 3** | | **Regime 4** | |
|---|---|---|---|---|---|---|---|---|
| | $c_1$ | $\delta_1$ | $c_2$ | $\delta_2$ | $c_3$ | $\delta_3$ | $C_4$ | $\delta_4$ |
| **Australia** | -45.081** | 5.212** | -27.832** | 3.294** | -7.117** | 0.837** | -6.564** | 0.818** |
| | (5.099) | (0.610) | (1.863) | (0.222) | (1.727) | (0.199) | (0.514) | (0.054) |
| **Belgium** | -10.968** | 1.414** | -16.916** | 2.105** | -2.607** | 0.412** | 1.340 | -0.027 |
| | (1.092) | (0.134) | (1.415) | (0.169) | (0.489) | (0.054) | (1.253) | (0.128) |
| **Canada** | -44.549** | 5.707** | -16.181** | 2.035** | -3.041** | 0.462** | | |
| | (1.860) | (0.244) | (1.018) | (0.125) | (0.218) | (0.024) | | |
| **Denmark** | -23.780** | 2.907** | -9.934** | 1.171** | 4.220** | -0.317** | | |
| | (1.407) | (0.181) | (0.313) | (0.037) | (0.780) | (0.079) | | |
| **Spain** | -33.332** | 4.169** | -18.431** | 2.220** | -25.675** | 3.102** | -8.877** | 0.996** |
| | (5.170) | (0.699) | (2.540) | (0.337) | (2.350) | (0.303) | (0.207) | (0.023) |
| **France** | -11.354** | 1.396** | -5.450** | 0.665** | 5.930** | -0.549** | | |
| | (0.306) | (0.038) | (0.328) | (0.036) | (1.159) | (0.118) | | |
| **Ireland** | -29.666** | 3.697** | -10.747** | 1.292** | -2.341** | 0.333** | | |
| | (3.758) | (0.471) | (0.449) | (0.054) | (0.288) | (0.031) | | |
| **Italy** | -63.418** | 8.183** | -18.437** | 2.159** | -17.811** | 1.990** | -5.269** | 0.607** |
| | (3.965) | (0.536) | (1.035) | (0.134) | (0.527) | (0.064) | (0.870) | (0.090) |
| **Netherlands** | -22.675** | 2.750** | -18.002** | 2.135** | -8.629** | 1.028** | 0.176 | 0.089 |
| | (1.090) | (0.134) | (1.245) | (0.147) | (0.287) | (0.032) | (1.000) | (0.101) |
| **New Zealand** | -19.419** | 2.292** | -2.309* | 0.241* | -8.199** | 0.922** | | |
| | (0.890) | (0.106) | (0.993) | (0.115) | (0.500) | (0.053) | | |
| **Norway** | -37.393** | 4.864** | -6.678** | 0.796** | -9.904** | 1.113** | -1.380# | 0.220** |
| | (2.871) | (0.388) | (0.771) | (0.098) | (0.850) | (0.098) | (0.807) | (0.082) |
| **Switzerland** | -15.611** | 1.695** | -23.256** | 2.473** | 5.772* | -0.532* | | |
| | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | (0.000) | | |
| **UK** | -3.953** | 0.587** | -2.410** | 0.382** | 3.884** | -0.303** | | |
| | (0.511) | (0.061) | (0.340) | (0.038) | (0.412) | (0.042) | | |
| **USA** | -18.069** | 2.274** | -2.996* | 0.508** | -3.247** | 0.513** | 1.191 | 0.047 |
| | (0.475) | (0.058) | (1.273) | (0.147) | (0.252) | (0.027) | (0.814) | (0.080) |

Notes: #, * and ** indicate 10%, 5% and 1% significance levels of the t-statistic. Standard errors in parentheses. As in Kejriwal (2008), $c_1$, $c_2$, $c_3$, and $c_4$ are the estimates for the constant in regimes 1, 2, 3, and 4, respectively. Likewise, $\delta_1$, $\delta_2$, $\delta_3$, and $\delta_4$ are the coefficient estimates of LN GDP in the four regimes, respectively. The LN $CO_2$ is the dependent variable.

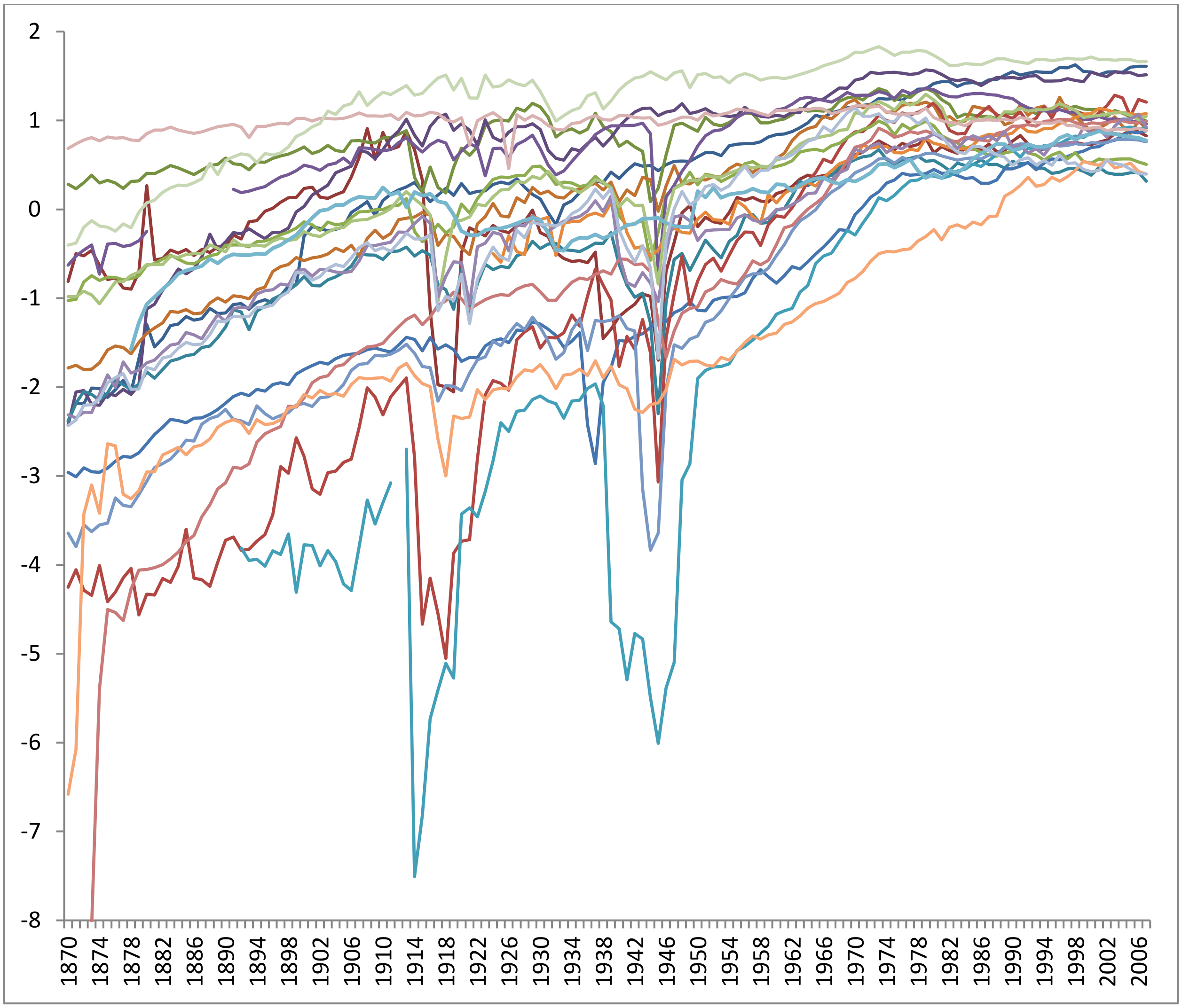


Figure 1. Natural log of $CO_2$ emissions per capita 1870-2007 for 21 OECD countries. Emissions data from Boden et al. (2013) and population data from Angus Maddison (http://www.ggdc.net/).

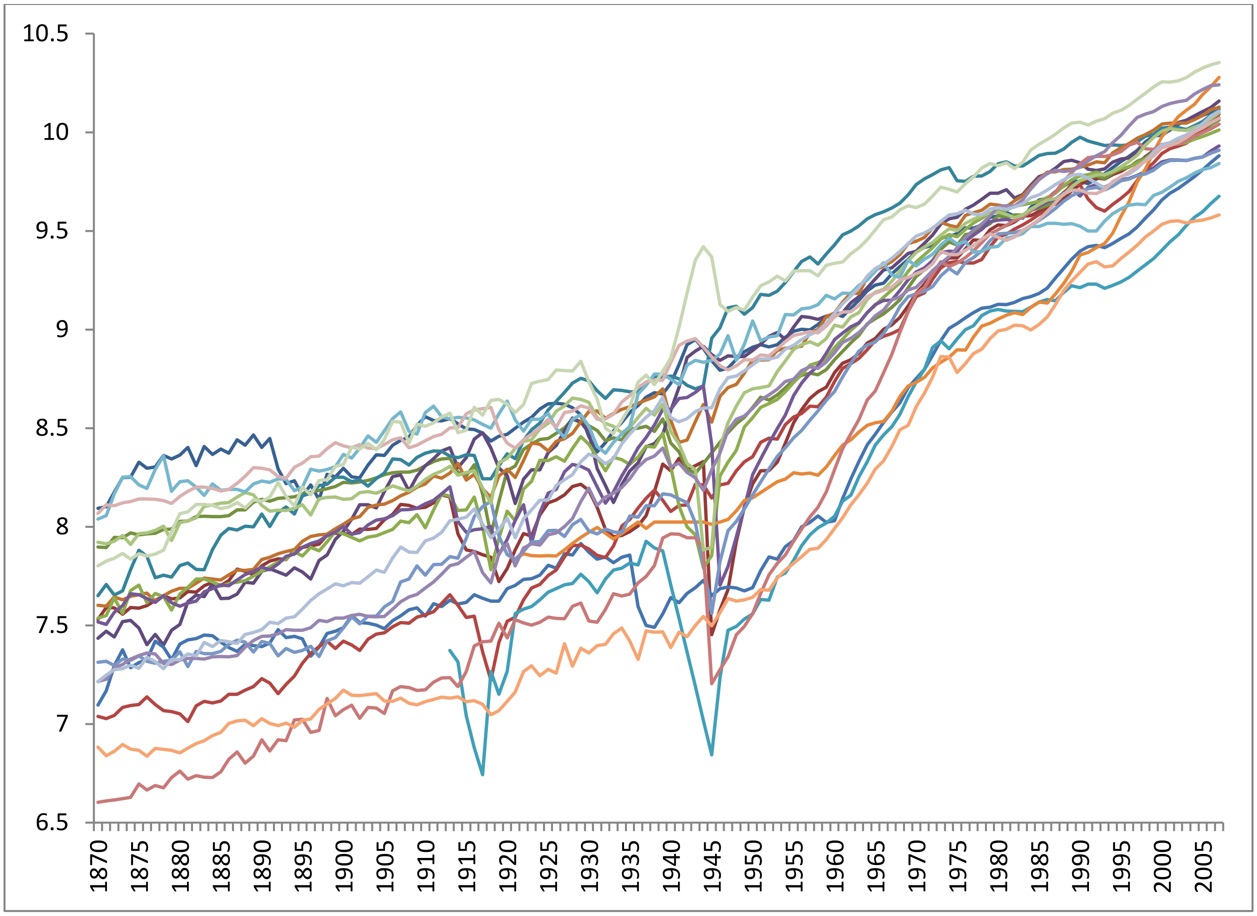


Figure 2. Natural log of real GDP per capita 1870-2007 for 21 OECD countries. Data from Angus Maddison (http://www.ggdc.net/).