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 23 OECD countries over 1950-2010 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 15 of 23 countries studied, the uncovered emission-income relationship was either (i) decouplingð where income no longer affected emissions in a statistically significant way, or (ii) saturationð where the emissions elasticity of income is declining, less than proportional, but still positive. For only four countries did the emissions-income relationship become negativeð 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.

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 relationshipð also called an Environmental Kuznets Curve (EKC)ð is one of the most popular questions in environmental economics (e.g., see reviews by Dinda 2004 and 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 emissionsð 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 between emissions and income (or EKC) means that the income elasticity of emissions is negative for countries in the highest income segment. If the income elasticity of emissions declines with income but remains positive (a phenomenon determined in Liddle 2013; Liddle 2014), 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). 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 literatures 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*ð i.e., a saturation effect or S-curve;[2] or (ii) that at high levels of GDP per capita the relationship with emissions is insignificantð 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. A reduced-form, linear model that allows for multiple endogenously determined breaks is used to address the econometric and modeling issues mentioned aboveð the linear model with multiple

[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] 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.

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 countryŝ emissions-GDP relationship separately. Lastly, by focusing on the time-series data of single countries, we address the crucial question of a specific countryŝ evolution of its income-environment relationship (as recommended by Stern et al., 1996 and de Bruyn et al., 1998).

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

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.

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 Mommaw 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 (2012) analyzed the CO_2 per capita and GDP per capita relationship for Spain over 1857-2007 using a cointegration model with endogenous breaks. They found (over three regimes) a declining but always positive income elasticity.

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 ñpiecesò (or income groupings) to be determined exogenously, and those income groupings are the same for all countries (in a panel analysis); hence, that approach 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 (including 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 ñcountry-specific characteristics such as technological progress, structural evolution, or external shocksò (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.[3]

[3] 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).

4. Data and Methods

4.1. Data

We analyze the CO_2 emissions per capita and real GDP per capita relationship for 23 advanced/OECD countries.[4] 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.[5] 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 during those two periods, all countries display a substantial break in GDP per capita around the Great Depression (e.g., 1930-1939). Yet, allowing for endogenous breaks involves an information trade-off; indeed, Harvey et al. (2013), Kejriwal and Perron (2010), and Kejriwal and Lopez (2013) recommend allowing for a maximum of two structural breaks (and considered over 100 time observations). But if we 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 1950ð an era of substantial economic growth and development for the countries consideredð is exactly the time in which we might expect to observe emissions-GDP transitions. Therefore, we restrict our sample to 1950-2010, 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 the Penn World Tables (Heston et al., 2012). (Both series are transformed into natural logs.)

Figures 1 and 2

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

[5] Because their GDP per capita data does not begin until 1950, the series for Hungary and Poland are not included in Figures 1 and 2.

4.2 Unit root tests with endogenous breaks

There are several unit-root tests that allow for structural breaks. Kejriwal and Perron (2010) is a sequential test that first considers one break versus no breaks, and then if one break is found, considers two breaks versus one, and so on. Carrion-i-Silvestre et al. (2009) allow for structural breaks in both the null and the alternative hypotheses, but assume all breaks are of the same magnitude. However, that homogeneity of break magnitudes assumption was challenged by Harvey et al. (2013), who developed a test that allows for breaks of different sizes. This paper adopts the more flexible Harvey et al. procedure in testing for unit-roots and breaks. Their procedure (HLT) 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 $\mathbf{DT}_{t}(\tau)$ is a vector of indicator variables, $1(t > [\tau T])(t - [\tau T])$, T is the sample size, $\tau_{0} = [\tau_{0,1,}$é $, \tau_{0,m}]^{ô}$, is a set of sample fractions, *m* is the maximum number of unknown breaks, $\gamma =$ $(\gamma_{1,}$é $,\gamma_m)$ are parameters associated with breaks, and u_t is a mean zero stochastic error process. A trend break in series y_t occurs at time $[\tau_{0,i}T]$ when γ_i Í 0 (i=$_{1,}$é $_{,m}$), 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 τ_L, τ_U are trimming fractions. The test statistic is $MDF_m = \inf DF^{GLS}(\tau)$, where $DF^{GLS}(\tau)$ is the standard t-ratio associated with ϕ in the fitted ADF equation: $\Delta u_t = \phi u_{t-1} + \sum \psi_j \Delta u_{t-j} + e_t$. 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.

If only one of the two series is determined to have a unit root, we conclude that the GDP-emissions relationship for that country is already (i.e., as of prior to 1950) described as decoupled, and we do not analyze those series further. Lastly, since all of the series are highly trending, we interpret the rejection of the unit root null as a finding of trend stationary. If both

GDP and carbon emissions are determined to be trend stationary, we estimate the relationship between them using the Bai and Perron (1998, 2003) method of endogenous breaks since that method is robust to trending regressors (but not $I(1)$, cointegrated ones).

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 augments 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 breaksð 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) stability 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,é ,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 $\lambda_i = (T_1/T,$é $,T_k/T)$, i.e., the sample fractions associated with i=1,é ,k breaks, $\Omega_{i,j}$ is the long-run variance of u_t for j=1,é ,k, and T_1,é ,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 Harvey et al. (2013) HLT 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. However, for Austria and Switzerland, the two series are trend stationary; so, we analyze their income-emissions relationships using the Bai Perron (1998; 2003) method (and do not test for cointegration). On the other hand, for Australia, Denmark, Finland, France, and Sweden, the two series are of different order of integration; hence, for those countries, decoupling of income and emissions had (arguably) already occurred, and 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). If the two criteria suggest different, nonzero number of breaks, we consider both possibilities (a case that only occurred for Netherlands and Poland). The null hypothesis of cointegration was never rejected.

Table 2

Table 3 presents the results for the regressions under breaksð for both the nonstationary, cointegrated and trend stationary cases. If we focus on the sign and significance of the income termôs coefficient (the δs in Table 3), by far the most common income-emissions relationship is that of saturationð a statistically significant, declining, but still positive income elasticity; that relationship is the clear case for eight countriesð Austria, Canada, Greece, Ireland, Italy, Japan, Netherlands (when only one break is allowed), and New Zealand. Since carbon emissions are so associated with energy consumption, perhaps a saturation pattern is to be expected. (Saturation in carbonôs income elasticity is the same pattern uncovered in the panel analysis of Liddle 2014.)

Table 3

US displays decoupling of income and emissions beginning in 1970ð as the income elasticity is no longer significant. Similarly, Hungary displays saturation beginning in 1963, followed by decoupling in 1990; so does Netherlands beginning in 1982 for saturation and in 1997 for decoupling, when two breaks are considered for that country. Previously, we mentioned five countries (Australia, Denmark, Finland, France, and Sweden) for which income and emissions had different orders of integration; if we judge those five countries as evidencing decoupling, too, then saturation and decoupling are the primary post-1950 income-emissions relationships, i.e., the case for 15 of the 23 OECD countries studied. Only four countries (Belgium, Germany, Switzerland, and UK) show clear evidence of a carbon Kuznets curveð a significant, negative relationship between income and emissions, and for two of those countries, Belgium and Switzerland, the downturn occurred rather recently (and thus, at a high income level), in 1997 and 1993, respectively.

The income-emissions regimes of four other countries (Norway, Poland, Portugal, and Spain) deserve further discussion. Considering one break (in 1989), Poland displays a CKC; however, when a second break is allowed (in 1999), Polandôs income-emissions relationship takes on an N-shape. That first break in 1989 is associated with the fall of the Berlin Wall and reintegration of East and West Europeð a time of great structural change for Eastern European countries. Hence, for Poland the regime of 1989-1999 was more of a period of structural change/adjustment than a transition period to a less carbon intensive path, and so post-1999, Poland has resumed its rather carbon/energy intensive economic development. For Portugal and

Spain, despite evidence of breaks, both countries have maintained a high, positive, and near proportional relationship between income and emissions (Spain had a period of accelerating emissions relative to income over 1970-1985). Indeed, for both of those countries the sequential method indicated no breaks in their income-emissions relationships; hence, one might surmise that neither Spain nor Portugal have experienced an income-emissions transition or regime change.

That leaves perhaps the most curious caseð wealthy, fossil fuel-endowed Norway. Norwayøs income-emissions relationship accelerated in 1970, and while it declined relative to that high elasticity, that relationship still has been more than proportional from 1990ð effectively a U-shaped relationship. Despite its high per capita income and its governmentøs traditional concern for sustainable development (e.g., the UNøs Our Common Future report is also known as the Brundtland Reportð after a Norwegian prime minister), Norway maintainsð particularly for an OECD countryð a relatively energy intensive industry sector. (Indeed, Liddle 2009 determined that Norway was one of six OECD countries for whose industry electricity intensity was converging to a relatively high level with respect to the other 12 OECD countries analyzed.)

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; relatedly, we estimated breaks for Canada in 1969 and 1981. 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.[6] 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 Austria, Canada, and Italy. For Spain Zanin and Marra predicted an oscillating but declining income elasticity (arguably similar to what we uncovered). Esteve and Tamarit (2012) focused on Spain over 1857-2007 using the same methods we do. They calculated a three-regime saturation pattern for Spain with an income elasticity of 0.56 over the final regime (1967-2007). By contrast we calculated a higher, but still less than proportional elasticityð most likely, had we considered a longer time span, a saturation pattern would have been more evident in our results, too (Esteve and Tamarit determined a first break at 1940).

The final way we can analyze the results shown in Table 3 is to focus on the timing of the break dates. For the 18 countries analyzed in Table 3, 29 breaks were identified. Severalð but not allð of the breaks did occur during periods of global/regional shocks (e.g., both Hungary and Poland had a break in 1989, the year the Berlin Wall fell). The first oil crisis could be dated 1973ï 1974;[7] whereas, the second oil crisis, which is dated 1979ï 1981, corresponded to the fall of the Shah in Iran and the beginning of the Iranï Iraq war, and it led to considerably higher prices than the first oil crisis. Indeed, seven breaks occurred during those two (oil-crisis) periodsð eight if we count Greeceôs break in 1982, and two more breaks occurred during the intervening period of high prices (1975-1978), for a total of 10 out of 29 breaks.

Another eight breaks occurred in 1971 or earlier, and the 1960s through the early 1970s (before the first oil crisis) was a period of heightened environmental awareness/concern in many

[6] Zanin and Marra (2012) do not appear to have performed unit root tests, and thus, proceeded to analyse Australia, Denmark, Finland, and France—countries for which we do not estimate an income-emissions relationship since we found their two series to be of different integration orders.

[7] 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.

OECD countries.[8] 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 (Mommaw and Unruh 1997; Volleberg et al., 2009; Stern 2010; and He and Richard 2010) and in energy intensity (Liddle 2012).

6. Conclusions

We used endogenous breaks modeling to examine the carbon emission-income relationship for 23 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 2012 study of Spain). Following several previous studies, the importance of shared timing among countries was uncoveredð in particular the increased interest in the quality of the environment in the 1960s and 1970s in OECD 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 negativeð i.e., a CKC. Indeed, the primary emission-income relationship determined here (i.e., for 15 of 23 countries studied) was either (i) decouplingð where income no longer affected

[8] 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).

emissions in a statistically significant way, or (ii) saturationð where the emissions elasticity of income is declining, less than proportional, but still positive.

This lack of inverted-Us provides an interesting contrast to Liddle and Messinis (2014), who focus on sulfur emissions of OECD countries and employ 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 processôs impact on the carbon emissions-income relationship we can contribute is that those two variables either become less than proportionally, positively related to each other or no longer strongly related to each other at all, as countries reach higher levels of development.

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Table 1. HLT (20013) unit root test with breaks, 1950-2010.

	LN GDP per capita		LN CO_2 per capita	
	m=1	**m=2**	**m=1**	**m=2**
Australia	-3.128	-4.280	-3.970*	-5.049*
Austria	-4.245*	-4.902*	-4.567*	-4.959*
Belgium	-2.523	-3.709	-3.208	-3.792
Canada	-2.308	-3.134	-1.989	-3.183
Switzerland	-3.602	-4.639*	-6.010*	-6.287*
Denmark	-2.271	-3.220	-4.182*	-5.165*
Spain	-3.472	-3.544	-1.941	-2.409
Finland	-2.892	-3.074	-4.787*	-5.430*
France	-2.556	-3.251	-3.528	-4.969*
Germany	-1.831	-3.055	-2.736	-4.184
Greece	-2.128	-2.920	-2.526	-3.601
Hungary	-2.310	-3.112	-2.343	-3.095
Ireland	-2.414	-2.595	-3.371	-4.348
Italy	-1.761	-3.596	-2.300	-3.346
Japan	-2.654	-3.054	-2.825	-3.672
Netherlands	-2.721	-3.351	-3.334	-3.979
New Zealand	-2.327	-3.512	-3.448	-4.077
Norway	-2.223	-2.855	-2.104	-3.184
Poland	-2.269	-3.367	-2.564	-3.859
Portugal	-1.998	-2.825	-2.958	-3.832
Sweden	-2.546	-2.848	-2.606	-5.384*
UK	-2.164	-2.815	-3.747	-4.188
USA	-2.879	-3.044	-2.315	-3.037

Notes: * indicates 5% significance level. m=number of breaks.

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, 1950-2010.

	Optimal number of breaks		**K-AK cointegration test**				
	S	**LWZ**	**$V_1(\breve{\lambda})$**	**Date**	**$V_2(\breve{\lambda})$**	**Date 1**	**Date 2**
Belgium	0	2			0.05	1981	1997
Canada	0	2			0.05	1969	1981
Spain	0	2			0.04	1969	1985
Germany	0	2			0.06	1962	1978
Greece	2	2			0.04	1982	1995
Hungary	0	2			0.08	1962	1989
Ireland	1	1	0.079	1970			
Italy	0	1	0.072	1971			
Japan	0	2	0.081	1977			
Netherlands	2	1	0.049	1981	0.03	1981	1996
New Zealand	0	2	0.054	1984			
Norway	2	2			0.12	1969	1989
Poland	1	2	0.054	1989	0.07	1989	1999
Portugal	0	1	0.068	1988			
UK	0	1	0.050	1973			
USA	0	2			0.06	1969	1979

Notes: S=sequential procedure (as described in Kejriwal and Perron 2010). LWZ=Schwarz criterion. The 1% and 5% simulated critical values for $V_1(\breve{\lambda})$ and $V_2(\breve{\lambda})$ are 0.214 and 0.129, and 0.156 and 0.101, respectively. The null hypothesis is cointegrated.

Table 3. Regression estimates with breaks, LN GDP per capita & LN CO_2 per capita, 1950-2010.

	Kejriwal for *I*(1), cointegrated pairs							
		Regime 1		**Regime 2**		**Regime 3**		**Descriptive**
	Breaks	c_1	δ_1	c_2	δ_2	c_3	δ_3	**pattern**
Belgium	1981	-2.493**	0.383**	-0.719	0.179#	11.939**	-1.046**	CKC
	1997	(0.255)	(0.027)	(1.018)	(0.101)	(2.248)	(0.216)	
Canada	1969	-7.566**	0.916**	-0.885	0.243*	-0.810#	0.225**	Saturation
	1981	(0.645)	(0.068)	(1.018)	(0.102)	(0.428)	(0.042)	
Spain	1969	-7.699**	0.803**	-12.774**	1.365**	-7.937**	0.851**	?
	1985	(0.368)	(0.042)	(1.749)	(0.181)	(0.672)	(0.066)	
Germany	1962	-5.339**	0.675**	-0.976**	0.222**	9.266**	-0.800**	CKC
	1978	(0.636)	(0.066)	(0.441)	(0.045)	(0.332)	(0.032)	
Greece	1982	-14.055**	1.482**	-23.894**	2.514**	-3.356**	0.419**	Saturation
	1995	(0.200)	(0.021)	(3.554)	(0.365)	(1.076)	(0.107)	
Hungary	1962	-10.572**	1.241**	-4.918**	0.608**	0.816	-0.039	Saturation,
	1989	(0.979)	(0.114)	(0.389)	(0.042)	(0.730)	(0.077)	decoupling
Ireland	1970	-10.298**	1.164**	-2.451**	0.329**			Saturation
		(0.604)	(0.068)	(0.189)	(0.019)			
Italy	1971	-16.976**	1.830**	-2.351**	0.304**			Saturation
		(0.325)	(0.035)	(0.409)	(0.040)			
Japan	1977	-9.771**	1.095**	-3.284**	0.410**			Saturation
		(0.166)	(0.017)	(0.459)	(0.045)			
Netherlands	1981	-10.743**	1.189**	-0.117	0.111*			Saturation
		(0.339)	(0.034)	(0.513)	(0.050)			
	1981	-10.585**	1.175**	-4.064**	0.501**	1.766	-0.068	Saturation,
	1996	(0.307)	(0.031)	(1.130)	(0.111)	(2.049)	(0.195)	decoupling
New	1984	-5.791**	0.636**	-3.962**	0.472**			Saturation
Zealand		(0.619)	(0.064)	(0.805)	(0.080)			
Norway	1969	9.443**	0.700**	8.274**	2.484**	9.520**	1.370**	?
	1989	(0.026)	(0.085)	(0.196)	(0.257)	(0.117)	(0.142)	
Poland	1989	-6.738**	0.878**	2.606**	-0.190**			CKC
		(0.194)	(0.022)	(0.387)	(0.042)			
	1989	7.663**	1.133**	10.830**	-1.921**	6.894**	3.288**	N
	1999	(0.040)	(0.038)	(0.342)	(0.380)	(0.606)	(0.770)	
Portugal	1988	-10.390**	1.086**	-9.335**	0.993**			**?**
		(0.159)	(0.017)	(0.913)	(0.093)			
UK	1973	-0.582	0.181**	3.319**	-0.233**			CKC
		(0.471)	(0.050)	(0.192)	(0.019)			
USA	1969	-4.100**	0.575**	1.444	0.030	1.457**	0.019	Decoupling
	1979	(0.544)	(0.056)	(1.225)	(0.122)	(0.280)	(0.027)	
	Bai Perron for trend stationary pairs							
		Regime 1		**Regime 2**		**Regime 3**		**Descriptive**
	Breaks	c_1	δ_1	c_2	δ_3	c_3	δ_3	**pattern**
Austria	1980	-7.051**	0.787**	-2.521**	0.318**			Saturation
		(0.219)	(0.023)	(0.508)	(0.049)			
Switzerland	1974	-16.749**	1.687**	0.467	0.006	6.112**	-0.543**	CKC
	1993	(0.417)	(0.042)	(2.125)	(0.205)	(1.715)	(0.164)	

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 are the coefficient estimates for the constant in regimes 1, 2 and 3, respectively. Likewise, δ_1, δ_2, δ_3 are the coefficient estimates of LN GDP in the three regimes, respectively. The LN CO_2 is the dependent variable.

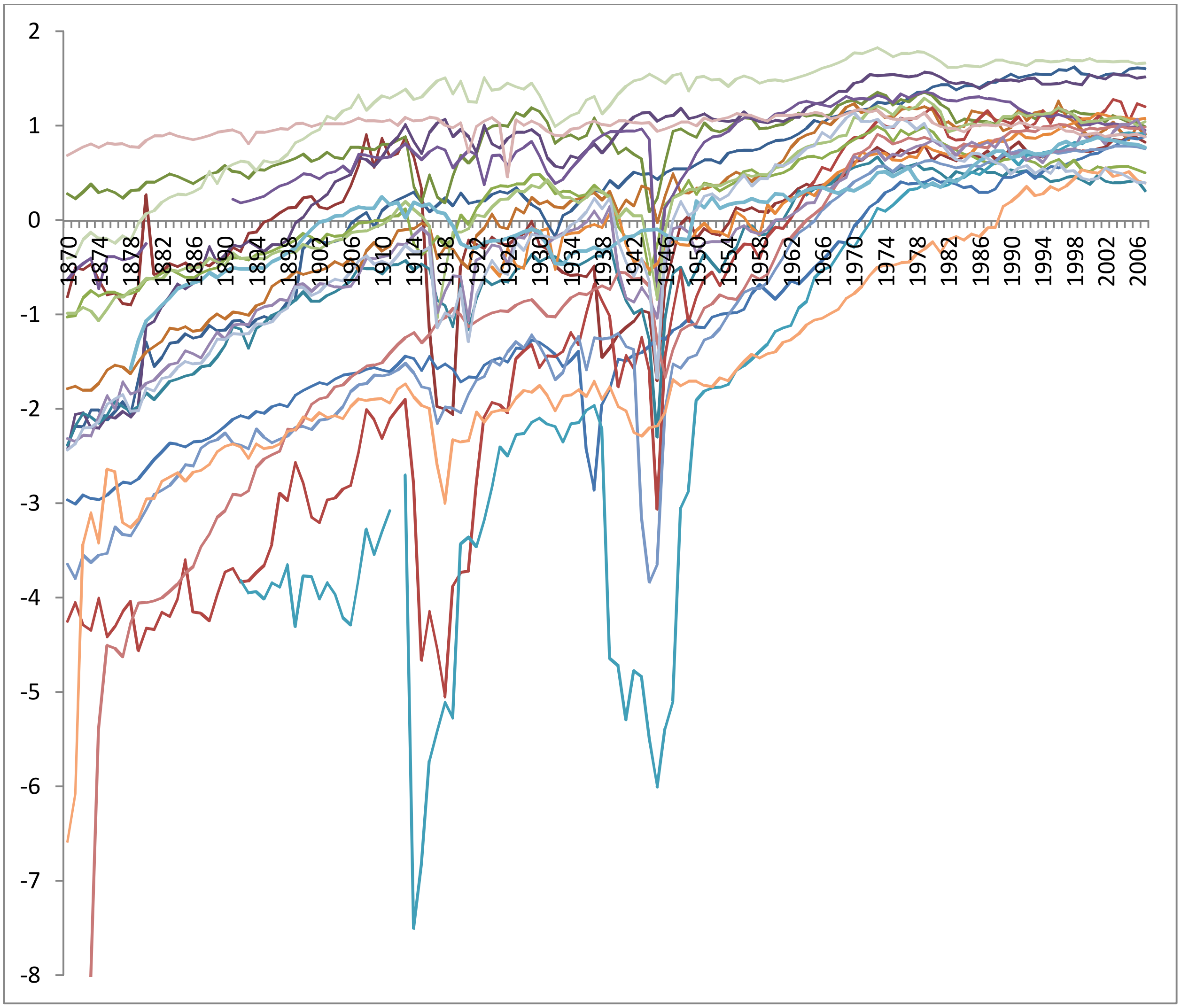

Figure 1. Natural log of CO_2 emissions per capita 1870-2007 for 21 OECD countries. Emissions data from Boden et al. (2013) and population data from Angus Maddison (http://www.ggdc.net/).

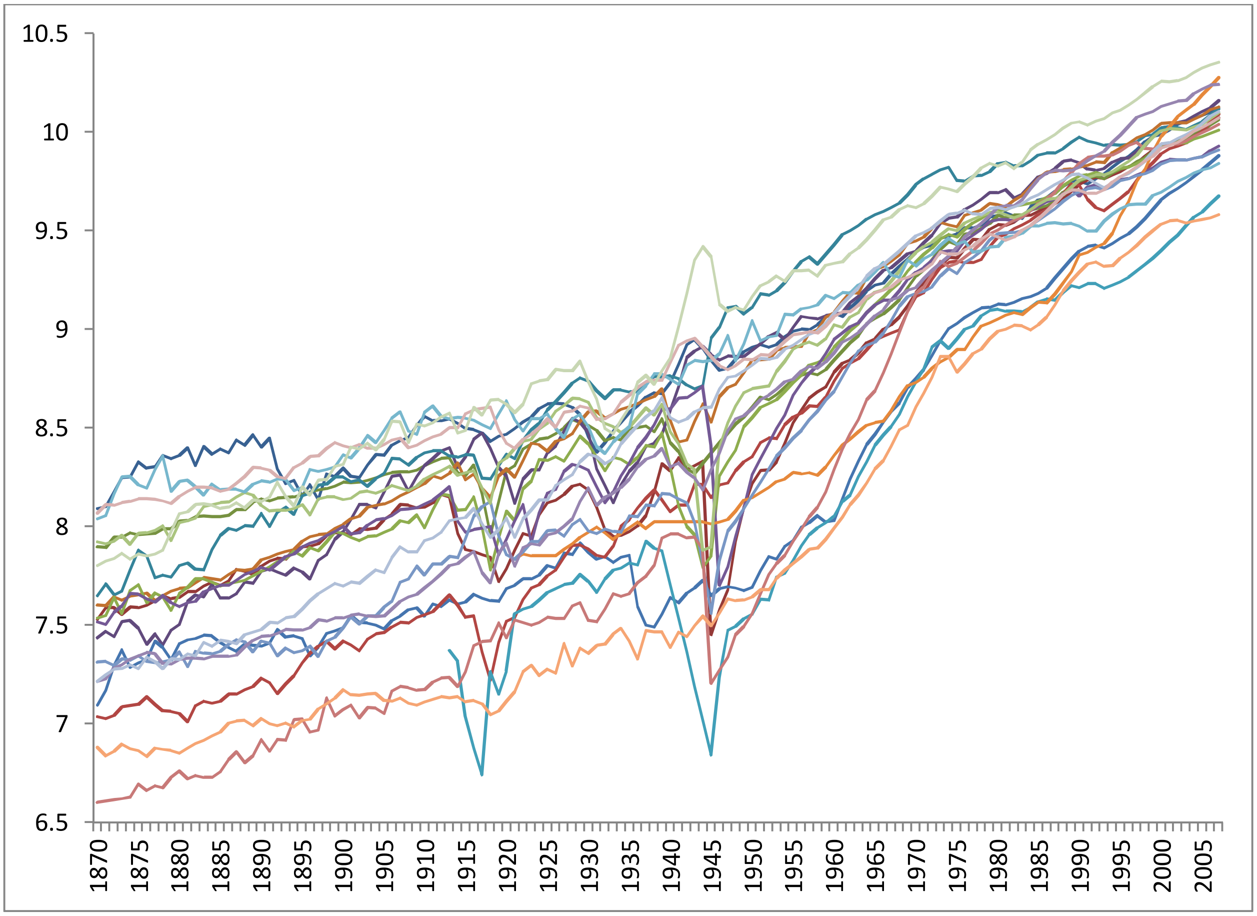

Figure 2. Natural log of real GDP per capita 1870-2007 for 21 OECD countries. Data from Angus Maddison (http://www.ggdc.net/).