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ABSTRACT
This paper tests for a sulfur Kuznets curve by examining the sulfur emissions per capita-GDP per capita relationship individually, for 25 OECD countries over 1950-2005 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. The predominant post-1950 income-sulfur emissions relationshipみthe case for 24 of the 25 countries studiedみwas either (i) inverted-Vs, where the emissions-income relationship became negative, or (ii) decoupling, where income no longer affected emissions in a long-run, statistically significant way. Seventeen of the 21 transitional breaks uncovered occurred over 1965-1978; hence, in concert with previous work, we conclude that shared timing among countries is important in income-environment transitions.

Keywords: Sulfur emissions; Environmental Kuznets curve; OECD countries; nonlinear flexible form; multiple endogenous breaks; income-emission relationships.

JEL classifications: C18, C22, C50, O44, Q53, Q56.

1. Introduction

Sulfur dioxide causes adverse local health effects and is one of the precursors of acid rain; hence, because people demand better environmental conditions as income rises, sulfur dioxide is the type of pollutant for which one might expect to find an inverted U-shaped relationship with incomeみa so-called Environmental Kuznets Curve (EKC). Indeed, the earliest EKC papers considered either sulfur emissions or concentrations, and those papers often estimated within-sample income turning points (Grossman and Krueger 1991; Shafik 1994; Selden and Song 1994). However, more recent work (Stern and Common 2001; Perman and Stern 2003; Wagner 2008; Vollebergh et al. 2009; and Stern 2010) has called into question those earlier results by rejecting an inverted-U relationship between sulfur emissions and income, despite that Stern (2006) noted a decline in the global trend of sulfur emissions.

This paper contributes to the sulfur Kuznets curve literature by examining the income-sulfur emissions relationship individually for 25 OECD countries using endogenous breaks modeling. In doing so, we address or avoid several important issues that have historically plagued the EKC literature (we believe this is the only paper to address/avoid all these issues simultaneously). Specifically, our approach (i) fully addresses the integration and cointegration properties of the data using recent, sophisticated econometric techniques (Stern 2004); (ii) avoids the pitfall of performing a nonlinear transformation on integrated income (Wagner 2008); and (iii) focuses on the time-series data of single countries, and thus, addresses the crucial question of a specific countryゃevolution of its income-environment relationship (as recommended by Stern et al., 1996). More to that last point, de Bruyn et al. (1998, p.173) argued that むthe EKC, as estimated from panel data does not capture dynamic processes well enough to justify the claim that economic growth is de-linked from environmental pressure in individual countries.め Moreover, regime change (modeled via an endogenous break) is arguably a more realistic model of emissions-income transitions than smooth, continuous curvature. Importantly, our paper is one of a recent strand in the literature to shed the restrictive polynomial model in favor of a flexible functional form approach. Additionally, by analyzing countries individually, we avoid the issues of cross-sectional dependence (Wagner 2008) and heterogeneity (Dijkgraff and Vollenbergh 2005).

2. Literature review

Besides producing contrary results, the EKC literature has been criticized along several lines (e.g., Stern 2004; Wagner 2008; Vollebergh et al., 2009; Carson 2010). Hence, in this brief review we discuss what we surmise is the state of the art in EKC/sulfur emissions-income modeling.[1] A few sulfur EKC papers have avoided the restrictive polynomial regression by applying a semi-parametric regression estimator (e.g., Millimet et al., 2003; Bertinelli and Strobl 2005). While this approach represents an improvement in terms of flexible functional forms, given its semi-parametric nature, the resulting estimates cannot be subjected to standard statistical tests (e.g., t-tests).[2] Also, semi-parametric methods, like parametric ones, must address nonstationarity in the regressors (and developing semi- and nonparametric methods are a nascent area of econometric research). Thus, Bertinelli and Strobl (2005) ultimately analyzed a first difference model, which (i) has a different interpretation than a model estimated in levels; and (ii) in their case, produced results that were much less definitiveみmonotonicity was not demonstrated clearly.[3]

Wagner (2008) performed defactored panel regressions (97 countries over 1950-2000) to address cross-sectional dependence and to avoid a nonlinear transformation of an integrated

[1] Dinda (2004) and Stern (2004) provide reviews of "first generation" EKC literature.
[2] The resulting traces are effectively analyzed visually—i.e., do they display downward curvature?
[3] Yet, a null hypothesis of linearity could not be rejected.

variable, and so rejected an inverted-U for sulfur emissions; indeed, in most of Wagner's regressions the coefficients for both linear and quadratic income terms are positive and significant. Stern (2010) employed the between estimator on the same data set to address the two issues raised by Wagner, as well as an issue raised by Volleberg et al. (2009)—that time effects are not uniquely identified. Stern (2010) rejected both an inverted-U and a convex monotonic shape for sulfur. The between estimator, while in principle a long-run estimator, effectively ignores unit roots and cointegration possibilities. Neither Wagner nor Stern was interested in the issue of alternative/flexible function forms of the emissions-income relationship.

By considering semi-parametric estimations, Volleberg et al. (2009) did provide some flexibility in functional form; however, they were most interested in separating income from time effects in a panel framework. Using their pairwise estimation procedure (applied to OECD countries only), they found a consistent positive relationship between income and sulfur emissions, but an inverted-U pattern in relationship with time. Volleberg et al. claimed to be unconcerned about either breaks or unit roots in their series. Yet, we do find evidence for both of those concerns, albeit with the advantage of an additional 19 (annual) observations.

Wang (2013) tested for unit roots and cointegration for a panel of OECD countries while allowing the exponent on a nonlinear income term to vary—thereby capturing functional form flexibility. Wang determined that emissions, income, and income raised to the power of 0.1 to 1.9 are all $I(1)$ variables, but income squared is $I(2)$; consequently, estimates involving income and income squared cannot cointegrate and are likely to be spurious. Wang found for an OECD country panel some evidence of EKC-type sulfur-income relationships, but established that emissions, income, and income transformed (although raised to a power less than two) do not cointegrate either, and thus, estimates involving those terms could be spurious as well.

The only recent, single country, time-series based analysis of sulfur emissions we know of is Fosten et al. (2012), who focused on the UK. Fosten et al. uncovered evidence of an inverted-U using nonlinear threshold cointegration and the typical polynomial model (without any apparent adjustments for the nonlinear transformations of integrated income). Therefore, it appears that our approach makes an important contribution to the sulfur emissions-EKC literature, because it (i) analyzes OECD countries individually; (ii) considers both unit roots and cointegration; (iii) at the same time avoids a nonlinear transformation of integrated income; and (iv) still fully allows for flexibility in functional form.

3. Data and Methods

3.1. Data

We analyze the SO_2 emissions per capita and real GDP per capita relationship for 25 advanced/OECD countries.[4] Figures 1 and 2 plot for those countries the long-run (1870-2005) 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 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) recommended allowing for a maximum of two structural breaks (and

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

[5] Because either their GDP per capita data does not begin until 1950 or is not available at all, the series for Hungary, Luxembourg, and Poland are not included in Figures 1 and 2. Emissions data for Ireland and Korea, which are shown in Figure 1, begin in 1924 and 1910, respectively; for all countries, emissions data ends in 2005. GDP per capita data for Korea, shown in Figure 2, begins in 1911.

considered over 100 time observations). But if we restrict our analysis to allow for no more than two endogenous breaks, such breaks likely would/may be calculated to occur before 1950 for most countries. However, the period beginning in 1950みan era of substantial economic growth and development for the countries consideredみis exactly the time in which we might expect to observe emissions-GDP transitions. Therefore, we restrict our sample to 1950-2005, and use SO_2 emissions per capita data from the NASA Socioeconomic Data and Applications Center (Smith et al., 2011) 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

3.2 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, $\tau_{0\,=}\,[\tau_{0,1,}$へ, $\tau_{0,m}]$ゃ, is a set of sample fractions, m is the maximum number of unknown breaks, γ = ($\gamma_{1,}$へ,γ_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 ぜ (i=$_{1,}$へ$_{,m}$), 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.

3.3 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 (but not *I*(1), cointegrated ones). To determine the timing of breaks Bai and Perron (1998, 2003) recommended 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) the $\sup F(l+1|l)$ test, a sequential test for 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) was particularly interested in estimating cointegrating relationships that have changed because of structural breaksみ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,ヘ,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 $\lambda_i = (T_1/T,\wedge,T_k/T)$, i.e., the sample fractions associated with $i=1,\wedge,k$ breaks, $\Omega_{i,j}$ is the long-run variance of u_t for $j=1,\wedge,k$, and $T_1,\wedge,T_k$ 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.

4. Results and discussion

Table 1 presents the results for the HLT (2013) 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 K-AK cointegration test for those countries. However, for Australia, Austria, Spain and Switzerland, 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 Poland and UK). The null hypothesis of cointegration was never rejected.

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 δs in Table 3), there is substantial evidence for sulfur Kuznets curves: for 14 countries (Belgium, Denmark, Finland, France, Hungary, Ireland, Italy, Japan, Luxembourg, Netherlands, Norway, Portugal, Sweden, and USA), a positive significant relationship between income and emissions was followed by a negative significant relationship after a break; whereas, for an additional three countries (Canada, Germany, and Korea), an insignificant relationship between income and emissions was followed by a negative significant one after a break. Poland displays evidence of both of those phenomena (significant positive followed by significant negative relationship and insignificant followed by significant negative relationship), depending on whether one or two breaks are considered. If two breaks are considered, the UK tracks the insignificant followed by significant negative relationship pattern; but if only a single break is allowed, it has a negative relationship throughout, which accelerates after a break. Thus, sulfur Kuznets curves (or inverted Us/Vs) were uncovered for 19 of the 25 OECD countries studied.

This substantial evidence of inverted-Us for sulfur provides an interesting contrast to Liddle and Messinis (2014), who analyzed the income-carbon relationship for OECD countries employing the same methods used here. Liddle and Messinis found evidence of inverted-Us for carbon for only four of the 23 OECD countries they considered. Such a difference in results is not surprising since sulfur is a pollutant with local health and environmental impacts, whereas carbon is a global pollutant with likely future impacts of very uncertain ultimate magnitudes.

Table 3

The sulfur emissions-income relationship in Greece displays decoupling as the relationship between the two series but is no longer statistically significant after the break in

1977. If we judge the four previously mentioned countries for which income and emissions had different orders of integration (Australia, Austria, Spain and Switzerland) as evidencing decoupling, too, then inverted-Vs and decoupling are the dominant post-1950 income-sulfur emissions relationships, i.e., the case for 24 of the 25 OECD countries studied.

New Zealand, by contrast, appears to be an anomaly among OECD countries. New Zealand's emission-income relationship displays a N-shaped pattern, with the relationship shifting from positive to negative after a break in 1978 and then back to positive after a break in 1993. While sulfur emissions in New Zealand do show an uptick in the 1990s as coal used in electricity generation increased, it is worth noting that per capita sulfur emissions in 2005 were nearly half their level in 1973 (the high point for emissions over our 55 year observation window).

Next we consider the timing of the breaks in which the emissions-income relationship first became negative, of which we have estimated 21 such breaks. Ten breaks occurred over 1972-1978—a period that (roughly) includes both oil crises and the intermittent period of high oil prices,[6] while seven breaks occurred in the mid-to late-1960s (a period of heightened environmental awareness/concern in many OECD countries);[7] thus, 17 of 21 transitional breaks occurred over 1965-1978. By contrast, Korea and Portugal, with the two lowest starting (1950) incomes among countries studied, had relatively late breaks (1992); whereas, Germany and Poland had breaks that corresponded (at least roughly) to the fall of the Berlin Wall and end of European communism (1989). Hence, overall there is evidence of the importance of shared

[6] The first oil crisis could be dated 1973–1974, which corresponded to OPEC's embargo; whereas, the second oil crisis, is dated 1979–1981, which corresponded to the fall of the Shah in Iran and the beginning of the Iran–Iraq war.

[7] For example, Rachel Carson published *Silent Spring* in 1962; the Apollo 8 picture of earthrise was taken in 1968, and the first Earth Day was held in 1970; the World Wildlife Fund, the Environmental Defense Fund, the Club of Rome, and Friends of the Earth were established in 1961, 1967, 1968, and 1969, respectively; and Clean Air Acts were passed in US in 1963 and updated in 1967, and passed in UK in 1968.

timing in breaks and emissions-income transitions, and the importance of timing has been emphasized in previous papers (e.g., Stern 2010; Volleberg et al. 2009).

Even though we uncovered substantial evidence of sulfur Kuznets curves for individual OECD countries, we do not believe that our findings contradict previous panel analyses that rejected such a relationship (e.g., Wagner 2008; Stern 2010). Indeed, we are analyzing an important, but very unique group of highly developed countries (among the countries considered, perhaps, only Korea achieved high development status during our study period). There appears no reason to believe that effective institutions that improve environmental quality without hampering economic growth are inevitable, nor is there any evidence that economic growth itself is inevitable (e.g., Easterly 2001; Pritchett 1997; Rodrik 1999). Ultimately, the study of transitions presents a conundrum: while we agree with Stern et al. (1996) and de Bruyn et al. (1998) on the importance of examining income-environment relationships individually, one must consider some type of panel-level analysis to support/reject the hypothesis that a particular income-environment transition might be applied generally. In other words, individual country analyses demonstrate what income-environment transitions are achievable but not which such transitions can/should be considered part of the development process.

5. Conclusions

We used endogenous breaks modeling to examine the sulfur emission-income relationship for 25 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. We believe that these three issues have not been addressed simultaneously before in the analysis of sulfur emission-income relationships.

We determined that 24 of 25 countries studied currently have either a negative relationship between income and sulfur emissions per capita or a decoupled/disassociated relationship between income and sulfur emissions so that the two variables are no longer significantly related. However, we interpret this highly consistent finding among OECD countries as evidence that EKC-type environment-development transitions are possible, rather than as evidence that such transitions are at all inevitable. Indeed, echoing previous work, we uncovered a prevalence of shared timing in regards to breaks/transitions as well as somewhat similar levels of development.

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Table 1. HLT (2013) unit root tests with breaks, 1950-2005.

	LN GDP per capita		LN SO_2 per capita	
	m=1	m=2	m=1	m=2
Australia	-3.087	-4.318	-3.976*	-4.487
Austria	-4.155*	-4.965*	-2.246	-2.873
Belgium	-2.520	-3.788	-2.530	-3.027
Canada	-2.471	-2.966	-2.579	-2.823
Switzerland	-3.967*	-4.783*	-2.478	-3.470
Denmark	-2.972	-4.355	-2.430	-4.172
Spain	-3.558	-3.961	-3.473	-4.778*
Finland	-3.006	-3.203	-2.798	-3.893
France	-3.213	-3.830	-2.515	-3.266
Germany	-2.442	-2.739	-2.881	-3.140
Greece	-1.943	-3.296	-2.641	-3.150
Hungary	-2.058	-3.531	-2.320	-3.459
Ireland	-2.637	-2.884	-2.370	-3.794
Italy	-2.576	-4.492	-2.118	-3.671
Japan	-2.796	-3.178	-2.140	-3.903
Korea	-2.193	-3.553	-1.801	-3.311
Luxembourg	-2.838	-3.696	-2.823	-3.756
Netherlands	-2.622	-3.747	-2.775	-2.917
New Zealand	-2.168	-3.610	-1.588	-3.367
Norway	-2.379	-2.971	-1.945	-3.819
Poland	-2.639	-3.048	-2.620	-3.442
Portugal	-2.551	-3.046	-3.15	-4.046
Sweden	-3.041	-3.529	-2.006	-2.910
UK	-2.887	-2.928	-2.181	-3.408
USA	-2.993	-3.751	-2.667	-3.528

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 SO_2 per capita, 1950-2005.

	Optimal number of Breaks		**K-AK Cointegration Test**				
	S	**LWZ**	**$V_1(\bar{\lambda})$**	**Date**	**$V_2(\bar{\lambda})$**	**Date 1**	**Date 2**
Belgium	0	2			0.05	1969	1982
Canada	0	2			0.04	1972	1992
Denmark	0	2			0.03	1976	1993
Finland	0	2			0.05	1976	1991
France	0	2			0.04	1975	1988
Germany	1	1	0.09	1989			
Greece	0	2			0.05	1967	1977
Hungary	0	2			0.10	1965	1988
Ireland	1	1	0.05	1967			
Italy	2	2			0.06	1975	1988
Japan	0	2			0.05	1969	1981
Korea	2	2			0.08	1975	1992
Luxembourg	0	2			0.06	1978	1993
Netherlands	0	2			0.04	1967	1981
New Zealand	0	2			0.04	1978	1993
Norway	0	2			0.05	1969	1981
Poland	1	2	0.09	1989	0.04	1976	1989
Portugal	0	1	0.04	1992			
Sweden	0	2			0.05	1973	1989
UK	1	2	0.09	1988	0.05	1972	1988
USA	0	2			0.04	1969	1986

Notes: S=sequential procedure (as described in Kejriwal and Perron 2010). LWZ=Schwarz criterion. The 1% and 5% simulated critical values for $V_1(\bar{\lambda})$ and $V_2(\bar{\lambda})$ 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 SO_2 per capita, 1950-2005.

		Regime 1		Regime 2		Regime 3	
	Breaks	c_1	δ_1	c_2	δ_2	c_3	δ_3
Belgium	1969	-12.682**	1.118**	13.102**	-1.563**	34.075**	-3.713**
	1982	(1.320)	(0.145)	(3.366)	(0.338)	(1.523)	(0.149)
Canada	1972	-1.109	-0.043	18.315**	-2.007**	4.802**	-0.719**
	1992	(0.870)	(0.092)	(1.330)	(0.131)	(1.942)	(0.188)
Denmark	1976	-11.188**	0.889**	28.676**	-3.144**	113.52**	-11.412**
	1993	(1.106)	(0.115)	(3.547)	(0.352)	(8.220)	(0.793)
Finland	1976	-10.674**	0.897**	23.305**	-2.602**	8.484**	-1.223**
	1991	(0.751)	(0.081)	(2.662)	(0.268)	(2.914)	(0.288)
France	1975	-8.933**	0.628**	64.816**	-6.812**	49.233**	-5.231**
	1988	(0.595)	(0.063)	(3.558)	(0.356)	(3.338)	(0.328)
Germany	1989	-0.877	-0.133	141.3**	-14.13**		
		(1.091)	(0.106)	(5.884)	(0.572)		
Greece	1967	-19.935**	1.794**	-21.92**	1.979**	-5.432**	0.247
	1977	(0.662)	(0.079)	(1.925)	(0.199)	(1.634)	(0.168)
Hungary	1965	-10.952**	1.058**	3.567**	-0.586**	18.887**	-2.292**
	1988	(1.572)	(0.183)	(1.142)	(0.123)	(3.399)	(0.361)
Ireland	1967	-20.071**	1.871**	4.993**	-0.843**		
		(2.099)	(0.236)	(0.555)	(0.059)		
Italy	1975	-21.167**	1.926**	20.041**	-2.338**	58.538**	-6.136**
	1988	(0.791)	(0.081)	(3.006)	(0.304)	(4.383)	(0.430)
Japan	1969	-11.82**	0.957**	44.108**	-4.929**	3.523*	-0.823**
	1981	(0.455)	(0.055)	(2.492)	(0.255)	(1.464)	(0.143)
Korea	1975	-24.119**	2.542**	-5.014**	0.202	21.243**	-2.544**
	1992	(1.362)	(0.183)	(1.396)	(0.160)	(5.367)	(0.549)
Luxembourg	1978	-27.359**	2.491**	6.522*	-0.898**	28.129**	-2.969**
	1993	(2.206)	(0.222)	(3.020)	(0.292)	(6.750)	(0.616)
Netherlands	1967	-18.814**	1.686**	27.794**	-3.103**	35.070**	-3.867**
	1981	(1.591)	(0.168)	(3.051)	(0.304)	(1.322)	(0.129)
New Zealand	1978	-4.744**	0.136#	20.475**	-2.486**	-14.175**	0.996**
	1993	(0.716)	(0.074)	(2.597)	(0.264)	(2.468)	(0.246)
Norway	1969	-8.812**	0.631**	19.858**	-2.269**	23.413**	-2.666**
	1981	(1.342)	(0.143)	(2.122)	(0.211)	(1.138)	(0.108)
Poland	1989	-10.491**	0.919**	14.567**	-1.909**		
		(0.238)	(0.027)	(0.738)	(0.081)		
	1976	-10.194**	0.881**	2.357	-0.506#	16.336**	-2.102**
	1989	(0.270)	(0.032)	(2.637)	(0.293)	(0.674)	(0.074)
Portugal	1992	-12.799**	0.979**	3.109**	-0.673#		
		(0.352)	(0.038)	(3.703)	(0.378)		
Sweden	1973	-5.557**	0.335#	61.184**	-6.399**	30.428**	-3.453**
	1989	(1.567)	(0.161)	(4.986)	(0.498)	(5.226)	(0.513)
UK	1988	4.280**	-0.684**	34.661**	-3.729**		
		(0.694)	(0.074)	(2.056)	(0.203)		
	1972	-3.626**	0.162	5.758**	-0.839**	34.879**	-3.751**
	1988	(0.999)	(0.107)	(1.538)	(0.159)	(1.282)	(0.127)
USA	1969	-5.246**	0.325**	17.322**	-1.932**	17.915**	-1.976**
	1986	(0.828)	(0.086)	(0.954)	(0.095)	(0.876)	(0.084)

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 SO_2 is the dependent variable.

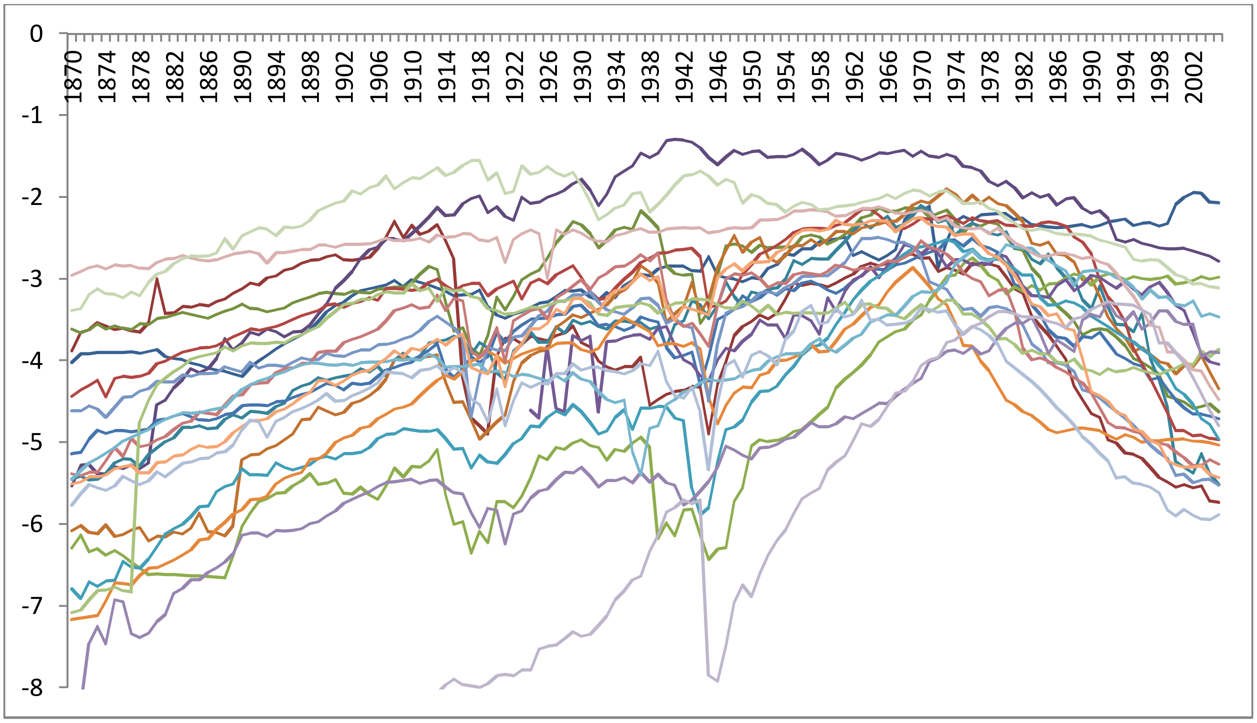

Figure 1. Natural log of SO_2 emissions per capita 1870-2005 for 22 OECD countries. Emissions data from Smith et al. (2011) and population data from Angus Maddison (http://www.ggdc.net/).

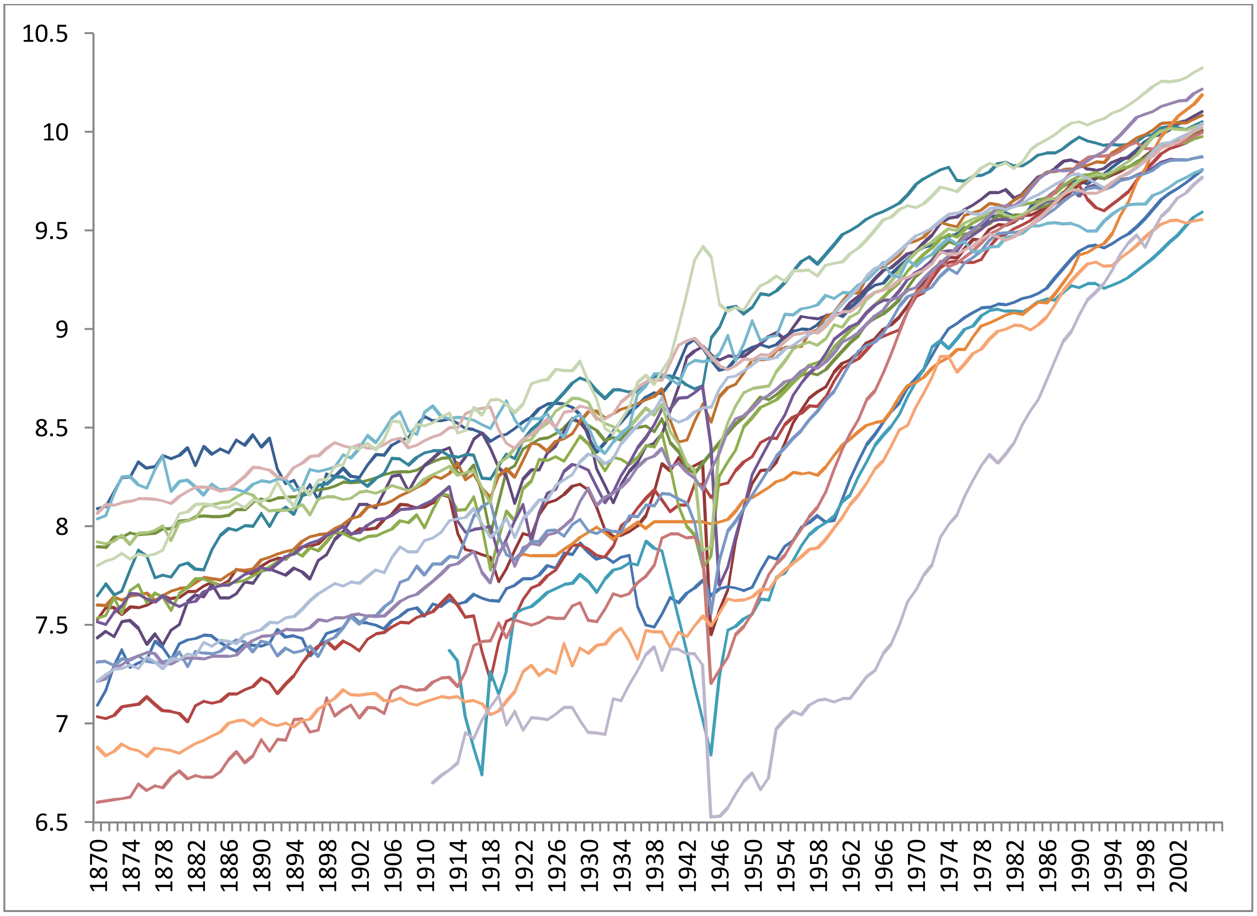

Figure 2. Natural log of real GDP per capita 1870-2005 for 22 OECD countries. Data from Angus Maddison (http://www.ggdc.net/).