COUNTRY-SPECIFIC ENVIRONMENTAL KUZNETS CURVES: A RANDOM COEFFICIENT APPROACH APPLIED TO HIGH-INCOME COUNTRIES*

CURVA AMBIENTAL DE KUZNETS: UN ENFOQUE DE COEFICIENTES ALEATORIOS APLICADO A PAÍSES DE ALTO INGRESO

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Abstract

Cross-country estimations of the Environmental Kuznets Curve (EKC) to empirically analyze the relationship between income and pollution have generally assumed a common structure for all countries. Since this latter feature is not supported by economic theory, this paper uses the Random Coefficient Model proposed by Swamy (1970) and empirically estimates EKCs for sulfur dioxide with specific turning points from a sample of 73 high and low income countries. A crucial aspect is that there are large differences between the estimated turning points of the EKCs for the different countries in the sample, which points to the relevance of using the approach employed here since assuming a common structure for all countries erroneously hides this relevant empirical feature. Moreover, the analysis of the structure of the EKCs estimated suggests that regulatory processes resembling market mechanisms could induce the empirical emergence of EKCs. Finally, taking into consideration the most recent concerns in the literature, we econometrically checked, on the one hand, for the validity of the usual theoretical assumption of exogeneity of the per capita income variable in the EKC relationship and, on the other hand, for an eventual structural change causing the sign change in the pollution-per capita income relationship of the EKC. The weak exogeneity and the structural break tests employed rendered plausible that income per capita is really the driver variable determining the EKC relationship found.

Key words: *Environmental Kuznets Curve, random coefficients, panel unit roots, weak exogeneity, structural break.*

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* Comments from reviewers were very useful; however, as usual, all remaining errors are the authors' responsibility.

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Resumen

Las estimaciones que utilizan datos de grupos de países para comprobar la existencia de la Curva Ambiental de Kuznets (en inglés, the Environmental Kuznets Curve (EKC)), la que empíricamente analiza la relación entre ingreso y contaminación, en general asumen una estructura común para todos los países. Sin embargo, como este supuesto no se deriva de la teoría económica y contradice la intuición, en este trabajo se emplea el modelo de "Coeficientes Aleatorios" propuesto por Swamy (1970) para estimar empíricamente las EKCs para SO_2 *(dióxido de sulfuro) y los puntos de quiebre específicos a partir de una muestra de 73 países de alto y bajo nivel de ingreso. Un hallazgo crucial es que existen amplias desigualdades entre los puntos de quiebre estimados de las EKCs para los diferentes países en la muestra, lo cual indica la relevancia del enfoque empleado porque el supuesto usual de una estructura común para todos los países erróneamente esconde esta característica empíricamente relevante. Más aún, el análisis de la estructura de las EKCs estimadas sugiere que procesos reguladores que replican mecanismos de mercado pueden inducir la emergencia de EKCs en los países que implementan tales procesos. Finalmente, para hacerse cargo de los cuestionamientos más recientes en la literatura, este trabajo testea econométricamente, por una parte, la validez del supuesto teórico usual de la exogeneidad de la variable de ingreso per cápita en la relación de EKC, y, por otra, la eventualidad de que un cambio estructural sea responsable del cambio de signo en la relación contaminación-ingreso per cápita de la EKC. Los tests de exogeneidad débil y de quiebre estructural empleados indican la plausibilidad de que la variable de ingreso per cápita sea realmente la variable causante que determina la relación de EKC encontrada.*

Palabras clave: *Curva Ambiental de Kuznets, coeficientes aleatorios, test unitario de panel, exogeneidad débil, cambio estructural.*

JEL Classification: *C23; Q53; Q58.*

I. Introduction

Since the early 1990s, the so called environmental Kuznets curve (EKC), an empirical inverted U relationship between pollution and income per capita, has been studied in different contexts due to its eventually promising implications for making economic growth sustainable in the future.

The EKC is an empirical finding showing that the emission levels of a given pollutant to the environment, or its concentration levels in the environment, first rise as income per capita of a country or a city rises along time and then, after reaching a maximum (called 'turning point'), they decline as income per capita continues growing.

Following Grossman and Krueger (1992), who first described the EKC, other authors have been optimistic about the implicit promise contained in the inverted U relationship estimated for some pollutants in different countries or cities (Selden and Song, 1994; World Bank 1992). These authors have the expectation that economic growth would restrain in some way the increase in pollution generally provoked by the increase in production and economic activity. However, other authors are skeptical about this (Panayotou, 1997; Figueroa and Pastén, 2000), since the inverted U relationship exists only for some pollutants, while other contaminants show a monotonically positive relationship with economic activity or income, such as CO_2 and methane which have been pointed out as two of the main green house gases responsible for climate change.

Not surprisingly then, learning more about the EKC, its determinants and the different institutional contexts in which it exists has been an important research challenge in the last fifteen year. As a result, a number of empirical studies have been published on the EKC for different countries and different types of contaminants. However, one crucial drawback of most cross-country studies empirically testing the presence of an EKC is their assumption that the coefficients of the inverted U relationship are the same for every country, implying that the expected shape of the EKC is the same for every country and the predicted turning point in income is also the same for every country.

However, since countries show significant differences in political, social, economic and biophysics factors, one should expect that different countries exhibit different patterns for their relationships between environment (or the level of different pollutants) and income. Therefore, the assumption that the EKC coefficients are constant across countries would be misleading most of the time. To solve this problem, some papers have used a fixed or random effects model allowing for a greater degree of heterogeneity. However, these models assume that the intercepts are different, but they still assume a constant slope for all countries. As a result, and since the EKC is a quadratic relationship, the implied shape and turning point predicted by the EKC are still country invariant.

Furthermore, when one assumes a constant slope for the EKC relationship across countries, testing the presence of an EKC turns out to be a test of an EKC for *every country* in the sample against the null hypothesis of absence of an EKC for *every country* in the sample. However, a more realistic strategy to estimate and capture the underlying empirical relationship is to allow for some countries to display an EKC while others show different and more complex patterns. If this turns out to be the case empirically, then pooling all countries together and testing the EKC for all of them is a biased procedure.

This paper tests the presence of an EKC using time series cross section (TSCS) data and finds the common elements to those countries exhibiting an EKC and, therefore, it allows for the fact that empirically the EKC is expected to be found only for a particular set of the countries included in the analysis. Contrary to the general practice, here we assume estimated coefficients (and the implied turning points) that are country specific. This approach is particularly relevant when the estimating procedure forces outlier countries that are affected by particular mechanisms into a common relationship structure that tends to bias the results.

Moreover, exploiting the time series-cross section characteristic of our data it is possible to explore several dynamic aspects of the EKC such as the order of integration of the series, the causality direction between emission and income and the eventual presence of structural breaks in the time series.

The pollutant studied in this paper is sulfur dioxide (SO_2), which allows for comparing our results with those of several works that have studied this contaminant and tested for the presence of an EKC. The next section reviews the theory and existing empirical results concerning the EKC, particularly for SO_2 emissions. In section three we present the data and the empirical results. Section four presents some robustness checks, and section five shows preliminary evidence that the EKC more likely will arise in regulatory processes resembling market mechanisms. Our conclusions are reported in the final section.

II. Empirical and Theoretical Background

After Grossman and Krueger (1992, 1995), the literature concerning the EKC grew rapidly. Selden and Song (1994) found an EKC for several indicators of urban air pollution emission. Different authors have made estimations for different pollutants and in most cases they found similar relationships when local pollutants with cheap abatement costs were considered. A few examples are sulphur dioxide (Shafik and Bandyopadhyay 1992, Shafik 1994, Cole *et al.*, 1997; Panayotou, 1993, 1997; Kaufmann *et al.*, 1998); carbon dioxide, nitrates, energy, traffic volumes (Holtz-Eakin and Selden, 1995); water quality (Beede and Wheeler, 1992; Hettige *et al.*, 2000). Nevertheless the EKC relationship is not undisputed, especially as a general empirical phenomenon expected to occur with all pollutants. A report of the World Bank (1992) finds monotonically increasing relationships (municipal waste) and monotonically decreasing relationships (lack of clean water, lack of urban sanitation). An entire issue of Environment and Development Economics edited by Barbier (1997) analyzed the EKC. Other recent papers surveying the subject are Stern (1998, 2003), Yandle *et al.* (2004) and Dinda (2004).

Some different theoretical explanations for the EKC hypothesis have been proposed. Arrow *et al.* (1995) suggested out that the pattern of environmental degradation is the natural progression from a basic agrarian process free of environmental impact to high polluting industrial economies and then to clean economies of services. The explanation by Suri and Chapman (1998) is that advanced economies export intensive pollution processes to less advanced economies. Jones and Manuelli (1995) argue that the externalities associated with intensive pollution processes can be better internalized by advanced collective decision-making institutional frameworks, which may only be implemented in developed countries. John and Pechenino (1994) present an overlapping generation model in which the environment degrades over time unless the society invests in it, which is only possible in advanced countries. Andreoni and Levinson (2001) proposed an explanation based on abatement technologies with increasing returns to scale. According to these authors, high-income individuals demand more consumption and less pollution. When abatement is possible, high income individuals would more easily achieve those demands. Additional analytical models are presented in Andreoni and Levinson (2001) and Mazzani *et al.* (2006).

All the empirical studies mentioned above use one of two general approaches: cross section analysis, or panel data (under fixed or random effects). In both cases the implied models assume that the slope of the corresponding relationship is the same for every country in the sample, while the constant may be different. Since the EKC is a quadratic relationship, the restriction tested must be over the slopes rather than over the constant. Since those slopes are generally assumed the same for every panel, the implied test is that the relationship (up to a constant) is the same for every country. Theoretically, it is possible that, under a given set of conditions, the turning points were statistically the same for every country; however, the underlying restrictions need to be tested rather than assumed a priori.

Only recently a few studies have started to address the slope-heterogeneity issue in EKC analysis. List and Gallet (1999) use panel data on state-level sulfur and nitrogen oxide emissions in the USA in order to test the validness of the "*one size fits all*" traditional approach to the EKC. The authors pointed out that "Parameter estimates suggest, that previous studies, which restrict cross-sections to undergo identical experiences over time, may be presenting biased results"[1]. Therefore, using a seemingly unrelated regression model (SUR) based in Zellner (1962), they found state-level differences in the predicted EKC in American states where such relationship exists. Koop and Tole (1999) using a Random Coefficient model (RC) similar to the model used in this paper, examine the relationship between deforestation and GDP. Using data for 76 developing countries, in the 1961-1992 period, their findings indicate that the previously supported EKC hypothesis does not longer hold under cross-country slope-heterogeneity. However they did not attempt to predict individual country level patterns.

The recent works of Cole (2005) and Markandya *et al.* (2006) share some similarities with our work. Both studies also focus on sulfur dioxide and use the same data base. However, while those studies take an 'all or nothing' approach to test the presence of an EKC relationship, here we allow the data to discriminate countries which present and countries which do not present such a relationship. Contrary to the a priori procedure followed by Markandya *et al.* (2006), by which a set of countries is chosen to test the EKC hypothesis, here an endogenous search procedure is used that allows the data to determine the set of countries in which the EKC hypothesis applies.

Here we use the RC model proposed by Swamy (1970) in order to allow for coefficient heterogeneity between countries. This approach would produce more reliable results than country specific constant models since there are country-specific differences which are function of the underling variables rather than time invariant. Additionally individual slopes for every country are estimated using the Swamy-Mehta estimator (Swamy and Mehta, 1975) and country-specific turning point are predicted where the data supports the EKC hypothesis.

Following Hsiao (2003), the Swamy-Mehta estimator (or shrinkage estimator) for each country could be seen from a Bayesian perspective as the Bayes predictor that determines a Bayesian turning point for country specific EKC. In a Bayesian framework, the parameters of the model are considered *"random variables"* and all probability statements are conditional. In this context the overall coefficients (from a pooled regression) express the prior distribution of the coefficients reflecting the original state of knowledge (or lack of it) about the real distribution of the parameters. With individual coefficients estimated,

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1 List and Gallet (1999), page 409.

the prior distribution of the parameters is revised. Inference can be made once the posterior distribution of the parameters is obtained.

Additionally to the RC model, a few other econometrics models have been proposed in order to address the coefficient heterogeneity issue. Maddala (1971) proposes an error component model similar in spirit to the RC model. In both cases the overall GLS estimator is a weighted average of the within and the between unit estimators. However, the intercountry variance is different in both models since Maddala restricted the randomness only to the constant, while the RC model allows randomness in constant and slope. In this sense RC model is more appropriate to address the slope-heterogeneity issue of concern here. Nevertheless if Maddala's original procedure is extended to allow for slope-randomness, then both, Maddala's GLS estimator and Swamy's GLS estimator, are indistinguishable.

A second approach to study parameter's stability is multilevel (o hierarchic) models. Coefficients coming from a multilevel model could be assumed random or fixed and the procedure is supported by maximum likelihood (ML) estimation. Asymptotically the bayesian ML coefficient and the Swamy's GLS coefficient are equivalent. In fact in the analysis of random coefficient applied to non-linear models, Greene (2001) uses multilevel models rather than Swamy's estimator. Here we use the Swamy's estimator instead of the ML multilevel estimator for convenience, since sometime the likelihood function does not converge for some sub-samples (i.e. low income countries). However, restricted ML was applied yielding coefficients that resembled those based in Swamy's GLS estimator (not reported).

III. Data and Empirical Results

The more general form usually estimated to test for the presence of an EKC is:

(1) $$y_{it} = \alpha_{0i} + \beta_1 x_{it} + \beta_2 x_{it}^2 + \varepsilon_{it}$$

where the subscripts i and t denote a country and time; y is annual pollution emissions divided by population, measured in tons of SO_2 per capita and expressed in natural logarithm; x is the natural logarithm of income per capita. β_1, β_2 are regression coefficients, α_0 is a specific unobservable effect and ε is the error term.

If an EKC exists, $\beta_1 > 0$ and $\beta_2 < 0$ and the estimated turning point is given by:

(2) $$\hat{x} = \exp(-\beta_1 / 2\beta_2)$$

Since in (1) there is a unique slope of the EKC, the model imposes a common structure for all countries. This implies that the expected effect of a given increment in income over the environment is the same for different countries. However, even though this simplifying assumption has been useful to recognize certain stylized facts, the interpretation and policy guidance derived from it may be misleading if in fact the interaction between income and environment is not the same for every country.

To avoid this problem, here we use a Random Coefficient Model, which allows the structure to differ between countries[2], and correspond to:

$$y_{it} = \alpha_{0i} + \beta_{1i} x_{it} + \beta_{2i} x_{it}^{2} + \varepsilon_{it} \tag{3}$$

The model in (3) is more flexible than the model in equation (1) since it allows the coefficients to differ between countries and, therefore, the estimated turning points can also be country specific. The estimated turning points are given by:

$$\hat{x}_i = \exp(-\beta_{1i} / 2\beta_{2i}) \tag{4}$$

Coefficients β_{1i} and β_{2i} are our Swamy-Mehta (Bayesian) parameters discussed before. As shown in Hsiao (2003, p. 170), the estimator $\beta_i = \left[\beta_{1i} \quad \beta_{2i}\right]$ in (3) is a weighted average of the (overall) Swamy estimator $\beta = \left[\beta_1 \quad \beta_2\right]$ and country by country OLS estimators[3]. The Bayesian estimator for each country tends to shrink the country's OLS estimator toward the overall estimator common to all countries. This is relevant in the fixed time series context since each country is obtaining information examining the behavior of others. For countries that belong to a similar group, the expected response should be similar with actual differences given only by a random component.

Nevertheless, even though we are assuming coefficients independently (and randomly) distributed between countries, once a particular coefficient is drawn for a particular country, it remains fixed for that country. Therefore it makes sense to predict individual coefficients and individual turning point where they exist.

If the slope variance between countries is small, a "fixed or random effect model" is indistinguishable from the "random coefficient model". However, if such a variance is large, then there is not slope-homogeneity between countries and the assumption of a common turning point is erroneous. Additionally, the assumption of zero intercountry slope variance will lead to a lower standard deviation for the overall slope when a random or fixed effects model is used in comparison to the RC model, which has strong consequences for statistical inference.

We estimate our model in (3) for 73 countries over the period 1960 to 1990. The data on SO_2 emission levels (in tons per year) is obtained from the ASL and Associates (ASL and Associates, 1997) dataset compiled by the U.S. Department of Energy[4]. The data on PPP-income per capita at constant prices of 1990 and on population are taken from Heston *et al.* (2002)[5].

The reason for the selection of this particular period of time is threefold. First, availability of reliable data neglects the possibility of a longer and more up to date time series; second, the period from the sixties to the nineties seems to be the period when most part of the environmental action happened; and third, this

2 The specific model used here follows Swamy (1970) and it is presented in Appendix A.
3 The weights given by the inverse of the variance's estimates.
4 A comprehensive description of this dataset can be found in Stern and Common (2001).
5 Real (PPP converted) GDP per capita (constant prices: chain series).

sample and particularly findings derived from it in Stern and Common (2001) seems to be the benchmark for the ongoing debate regarding the robustness of the EKC hypothesis (Stern 2004).

Stern and Common (2001), using a common structure for all countries, found some support for the EKC hypothesis in OECD but not in Non-OECD countries. Given the slope-homogeneity assumed in their paper these findings can be interpreted as accepting the EKC for *all* OECD countries and rejecting the EKC for *all* Non-OECD countries. As discussed before, in this paper we take a less extreme alternative which allows us the possibility that, within a group of countries, the null (EKC) hypothesis is valid for some countries but not for all of them.

One of our first concerns is whether or not to pool the data given the likely high heterogeneity of countries in this sample. In order to address this issue we follow Bartels (1996) in the sense that rather than deciding to poll or not, there is always the possibility of deciding "*how much to pool*". Bartels proposes to estimate different models allowing for varying degrees of pooling and then to make a decision based on the fitness of the model. At one extreme, it is possible to pool the whole sample with the cost of some biasness if in fact the proposed model applies only to some units and not to all the units in the sample. At the other extreme, it is possible not to pool at all, treating any unit as a specific model with the cost of loss of generality. Even though deciding how much to pool involves judgment, it is still more reasonable than using one of the two extreme alternatives.

In order to compare results, we first use the same sample as in Stern and Common (2001) (S&C). Then we restrict the sample to include only those countries where an overall EKC is found (the OECD case). Finally, we use the random coefficient model in order to predict country-specific turning points.

Table 1 reports our empirical results applying Random Coefficients to the same three samples used in S&C. Column 2 shows results for the world as a whole, while columns 3 and 4 show results for Non-OECD and OECD countries respectively. For the whole world and for the Non-OECD countries the t values are not significant, rejecting the presence of an inverted U relationship for both cases. Contrarily, in the OECD case all the coefficients are significant with a 95% of confidence given support to the presence of an EKC for OECD members. These results are consistent with S&C's results. Additionally, the estimated turning point for OECD countries (US$ 12,776 of 1990) is reasonably close to that found in S&C[6]. Finally, the test of parameter constancy reflects the fact that the null hypothesis of different slopes between countries cannot be rejected by the data, which confirms the presumption that motivated this work. The same test also indicates that there is parameter heterogeneity in the three regions and that the OECD case shows the lowest heterogeneity.

Table 2 displays Swamy-Mehta estimators (Swamy and Mehta, 1975) to predict individual coefficients within the OECD group. For most of the countries in the OECD sample the coefficients have the right sign and are highly significant, giving support to the existence of an EKC at a country level. However, for 9

6 In S&C the turning point for OECD members is US$ 9,239 in the case of fixed effects and US$ 9,181 in the case of random effects.

TABLE 1
RANDOM COEFFICIENT MODEL FOR THREE REGIONS OF THE WORLD

	World	Non-OECD	OECD
Numb. Obs.	2263	1550	713
Numb. Groups	73	50	23
Numb. Periods	31	31	31
Constant	–59.32	–46.32	–101.96*
	(–1.3)	(–0.7)	(–2.15)
β_1	13.39	11.99	20.80*
	(1.1)	(0.72)	(2.0)
β_2	–0.84	–0.88	–1.1*
	(–1.1)	(–0.78)	(–1.9)
Turning Point			12,776
Test. of Par. Const.	1.9e+05	1.6e+05	2,4311.63

t values in parenthesis.
*significant at the 5 % level.

countries in the sample (Denmark, Greece, Ireland, Portugal, Spain, Sweden, Turkey, U.K., and Australia) the EKC hypothesis does not stand.

When the cases in which the EKC does not apply are treated as outliers and eliminated, the reduced sample strongly supports the presence of an EKC using a fixed-random effect model and a random coefficient model (see Table 3). The reduction in the test of parameter constancy vis-à-vis the one in Table 1 indicates that this sub sample is more homogeneous that the whole OECD sample, even though there is still high slope variability between countries. Again, using the Swamy-Mehta estimator it is possible to predict individual coefficients. These coefficients along with the estimated individual turning points are reported in Table 4. The estimated turning point goes from US$ 6,229 (Austria) to US$ 12,865 (USA), with large differences between countries (see column 4).

Nevertheless, in order to test the presence of an EKC in the richest countries of the world a sample of only OECD members is inappropriate since it does not comprise the whole universe of high-income countries. For example, in the 1960-1990 period, Turkey was among the thirty poorest countries in the world. Furthermore, using current GDP data, countries not belonging to the OECD such as Korea, Singapore and Hong Kong can be classified as high income countries.

Therefore, to test the hypothesis that an EKC exists for the richest countries in the world, we use the World Bank's definition of high income countries employed in its World Development Indicators (World Bank, 2004). Columns 2 and 3 in Table 5 show the GDP per capita of the richest countries as a proportion of USA GDP per capita in 1960 and 1990, respectively. Our sample comprises rich countries that by 1990 had incomes per capita greater than 40% of the USA income per capita. Using Swamy and Mehta (1975), individual slopes can be estimated and individual turning points predicted (not reported) for each country. The last column in Table 5 shows that in 17 out of the 28 richest countries

TABLE 2
PREDICTED COEFFICIENTS OECD COUNTRIES

Country	β_{1i}	β_{2i}	Country	β_{1i}	β_{2i}
Canada	16.52*	–0.89*	Luxembourg	85.25*	–4.57*
	(4.31)	(–4.36)		(4.12)	(–4.10)
USA	40.58*	–2.14*	Netherlands	19.32*	–1.10*
	(4.88)	(–0.44)		(2.07)	(–2.14)
Japan	16.04*	–0.89*	Norway	65.63*	–3.59*
	(8.58)	(–8.42)		(9.62)	(–9.63)
Austria	10.53*	–0.60*	Portugal	–27.16	1.72
	(2.62)	(–2.70)		(–10.1)	(10.45)
Belgium	20.43*	–1.14*	Spain	–5.86	0.42
	(2.94)	(2.97)		(–0.85)	(1.05)
Denmark	–15.65	0.89	Sweden	0.37	0.06
	(–1.37)	(1.44)		(0.03)	(0.10)
Finland	25.50*	–1.38*	Switzerland	71.43*	–3.82*
	(5.02)	(–4.94)		(4.18)	(–4.21)
France	51.01*	–2.83*	Turkey	–12.35	0.89
	(5.55)	(–5.61)		(–2.16)	(2.44)
Germany	22.44*	–1.25*	United	–13.09	0.64
	(5.19)	(–5.30)	Kingdom	(–1.09)	(0.98)
Greece	–22.37*	–1.47*	Australia	–7.33	0.40
	(–9.46)	(10.33)		(–1.33)	(1.35)
Ireland	–8.57	0.50	New Zealand	103.86*	–5.71*
	(–1.23)	(1.24)		(4.80)	(–4.85)
Italy	41.76*	–2.33*			
	(9.55)	(–9.56)			

t values in parenthesis. *Significant at the 5% level.
All constants α_{0i} (not reported) are significant at the 5% level.

in the world there exists an EKC. In 11 cases (Hong Kong, Sweden, Australia, Denmark, United Kingdom, Spain, Ireland, Cyprus, Portugal, Greece, and Korea) the evidence does not support the existence of a Kuznets curve.

Moreover, most countries with an income per capita less than 50% of the USA's do not exhibit an EKC (the only exceptions being Israel and Saudi Arabia), and more strikingly, for some of the richest countries in the sample there is no evidence of an EKC. As before, those cases are Sweden, Australia, Denmark and United Kingdom and the newly added, Hong Kong. These results (particularly the case of Hong Kong) are indicative that a deterministic and automatic relationship does not arise simply because a country became rich.

Table 6 reports the results obtained from estimations including only the group of the 17 high-income countries exhibiting an EKC. Columns 2, 3 and 4 show the results using a fixed effect model, a random effect model and a random coefficients model, respectively. The three models support the presence of an overall Environmental Kuznets Curve. As expected, the t values are lower for the random coefficient model than for the other two models. The coefficients

TABLE 3
ONLY OECD COUNTRIES WITH EVIDENCE OF AN EKC

	Fixed Eff.	Random Eff.	Rand. Coeff.
Numb. Obs.	434	434	434
Numb. Groups	14	14	14
Numb. Periods	31	31	31
Constant	–73.10*	–73.00*	–204.73*
	(–9.82)	(–9.82)	(–3.51)
β_1	15.34*	15.32*	43.88*
	(9.35)	(9.35)	(3.49)
β_2	–0.85*	–0.85*	–2.40*
	(–9.40)	(–9.40)	(–3.52)
Turning Point	8,322	8,325	9,460
Test of Par. Const.			17,321

t values in parenthesis.
*significant at the 5 % level.

are significant and have the expected signs. Also, the predicted (overall) turning point is similar in the three models. Finally, the test of parameter constancy cannot reject the null hypothesis that the slopes for each country are different, again confirming our initial presumption.

Table 7 shows the estimated coefficients along with the predicted specific turning point for each country in our sample of high-income countries. For all of them, the coefficients have the right sign and are highly significant. The expected turning points go from US$ 6,201 (Netherlands) and US$ 12,863 (USA).

In line with the results presented above we could say that for most countries with an income per capita higher than the fifty percent of US income per capita (i.e. 50 % of the highest income per capita in the sample) an EKC do exist. Contrarily, for most countries with less than 50 % of the highest income in this sample an EKC do not exist. Thus, given the highly skewed distribution of world income distribution, for most of the underdeveloped countries, the predicted turning point is far from being reached[7].

IV. Robustness of the Results

4.1. The order of integration

An arguable shortcoming of the previous analysis is the possibility of a spurious regression between income and emission. If the sample is growing in N (number of countries) rather than in the time dimension this is not a problem. However, a sub sample considering only high-income countries is better characterized

7 Thanks to an anonymous referee for raising this point.

TABLE 4
PREDICTED TURNING POINTS OECD COUNTRIES WITH EVIDENCE OF AN EKC

Country	β_{1i}	β_{2i}	Turning Point (US$/per capita)	Turning Point (% of Austrias' Turning Point)
Canada	16.15	–0.87	10,299	165.3
	(4.11)	(–4.15)		
USA	37.53	–1.98	12,865	206.5
	(4.48)	(–4.50)		
Japan	16.11	–0.89	7,972	128.0
	(8.41)	(–8.25)		
Austria	10.80	–0.62	6,229	100.0
	(2.63)	(–2.71)		
Belgium	20.86	–1.16	7,789	125.0
	(2.97)	(–3.00)		
Finland	25.88	–1.40	9,876	158.5
	(5.00)	(–4.92)		
France	51.25	–2.84	10,114	162.4
	(5.55)	(–5.62)		
Germany	22.48	–1.25	8,001	128.4
	(5.07)	(–5.18)		
Italy	41.66	–2.32	7,681	123.3
	(9.38)	(–9.39)		
Luxembourg	96.07	–5.16	11,134	178.7
	(5.12)	(–5.10)		
Netherlands	19.84	–1.13	6,440	103.4
	(2.12)	(–2.20)		
Norway	66.05	–3.61	9,409	151.1
	(9.53)	(–9.55)		
Switzerland	84.18	–4.50	11,601	186.2
	(5.22)	(–5.26)		
New Zealand	105.42	–5.79	8,925	143.3
	(5.30)	(–5.36)		

t values in parenthesis t values in parenthesis.
All estimates including constants are significant at the 5 % level.

TABLE 5
HIGH-INCOME COUNTRIES

Country	% of GDP of USA 1990	% of GDP of USA 1960	EKC
USA	100	100	YES
Canada	95.1	73.4	YES
Switzerland	91.4	95.1	YES
Luxembourg	90.2	80.1	YES
Norway	82.5	56.7	YES
Hong Kong	82.2	22.7	NO
Sweden	81.8	76.7	NO
Australia	80.0	78.6	NO
Japan	80.0	30.0	YES
Germany	79.4	66.4	YES
Finland	77.9	53.5	YES
Denmark	77.0	68.3	NO
France	77.0	58.8	YES
Belgium	73.3	55.5	YES
United Kingdom	73.2	69.0	NO
Netherlands	72.2	61.4	YES
Austria	70.3	52.0	YES
Italy	69.2	46.1	YES
Singapore	64.9	16.8	YES
New Zealand	63.8	80.4	YES
Spain	53.1	31.6	NO
Israel	51.5	35.1	YES
Ireland	51.4	33.5	NO
Cyprus	46.3	20.6	NO
Portugal	41.1	18.9	NO
Saudi Arabia	40.9	39.3	YES
Greece	37.3	21.2	NO
Korea	37.0	9.1	NO

High income countries in ASL dataset as defined by The World Bank in its World Development Indicators (2004).

TABLE 6
HIGH-INCOME COUNTRIES

	Fixed Eff.	Random Eff.	Rand. Coeff.
Numb. Obs.	527	527	527
Numb. Groups	17	17	17
Numb. Periods	31	31	31
Constant	–86.61*	–86.39*	–182.26*
	(–24.62)	(–24.51)	(–3.66)
β_1	18.38*	18.32*	39.14*
	(23.12)	(23.04)	(3.65)
β_2	–1.02*	–1.02*	–2.15*
	(–22.70)	(–12.51)	(–3.70)
Turning Point	8,183	7,945	8,976
Test of Par. Const.			18,073

t values in parenthesis.
*significant at the 5 % level.

TABLE 7
PREDICTED TURNING POINTS HIGH INCOME COUNTRIES

Country	β_{1i}	β_{2i}	Turning Point (US$/per capita)	Turning Point (% of Netherlands's Turning Point)
Netherlands	19.91 (2.18)	–1.14 (–2.26)	6,201	100.0
Austria	10.86 (2.68)	–0.62 (–2.76)	6,362	102.6
Singapore	27.37 (7.33)	–1.56 (–6.99)	6,454	104.1
Israel	27.49 (3.82)	–1.56 (–3.78)	6,707	108.2
Italy	41.68 (9.52)	–2.33 (–9.53)	7,663	123.6
Belgium	20.94 (3.04)	–1.17 (–3.07)	7,698	124.1
Japan	16.13 (8.52)	–0.90 (–8.36)	7,794	125.7
Germany	22.49 (5.15)	–1.25 (–5.25)	8,071	130.2
Saudi Arabia	5.95 (2.15)	–0.33 (–2.15)	8,227	132.7
France	51.05 (5.66)	–2.83 (–5.72)	8,262	133.2
New Zealand	100.38 (5.52)	–5.52 (–5.48)	8,887	143.3
Norway	65.91 (9.68)	–3.60 (–9.70)	9,454	152.5
Canada	16.18 (4.18)	–0.88 (–4.22)	9,830	158.5
Finland	25.93 (5.09)	–1.41 (–5.00)	9,848	158.8
Luxembourg	93.51 (5.24)	–5.01 (–5.22)	11,089	178.8
Switzerland	82.02 (5.21)	–4.38 (–5.25)	11,649	187.9
USA	37.47 (4.54)	–1.98 (–4.56)	12,863	207.4

t values in parenthesis.
All estimates including constants are significant at the 5 % level.

as growing in the time dimension rather than in the space dimension. Several authors have detected the presence of a unit root in both the pollution and the income indicator (see Stern 2004). Whether or not the results are spurious will depend on the possibility of co-integration between both pollution and income variables. Table 8 presents a set of panel unit root that allows us to determine if such co-integration exists.

In Table 8 a set of panel unit root tests applied to the residuals in the overall random coefficient regression is carried out. The "*Levinlin*" test is based in Levin *et al*. (2002) while the "*Ipshin*" test is based in Im, Pesaran and Shin (2003). The number of lags in the lagged differenced variables was set equal to one[8]. Row one applies the "*levinlin*" test to our sample of countries displaying an EKC. Unit root in the panel is rejected with a 5% of significance. As a matter of comparison, rows two and three in the table show the same statistic for the no-EKC sample and for the sample as a whole respectively. Also in this case the panel unit root hypothesis is rejected at conventional levels of significance. Nevertheless, the *"levinlin"* test is a test of the null hypothesis against the alternative hypothesis of stationarity for "*all*" panel components. A less restrictive test that allows for some non-stationary panel in the alternative hypothesis is given by the "*Ipshin*" test in rows (4)-(6) of Table 8. For countries displaying an EKC, the *Ipshin* tests reject the null hypothesis of panel unit root at the 95% confidence level. Contrarily, as can be seen in rows (5) and (6) the *Ipshin* test cannot reject the null hypothesis of panel unit root in the whole sample, neither can reject it in the Non-OECD sample. We conclude that the residual from the overall EKC regression are panel-stationary and consequently, co-integration in the quadratic equation cannot be rejected by the data rendering our findings no-spurious[9].

4.2 Weak exogeneity

A second potential problem of the data used here is the possibility that the right hand side variables in the cointegration relationship in (3) are not exogenous and therefore yield biased results. There are two classes of tests to testing for weak homogeneity in dynamic regression models. The first approach, called Durbin-Wu-Hausman test (Durbin, 1954; Wu, 1973; Hausman, 1978), tests the orthogonality between innovations and the conditioning regressors. The second approach, which is the one used here, corresponds to a more recently developed test consisting in the design of a error correction model in which weak exogeneity is tested through the error correcting behavior (Johansen, 1992).

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8 The results are invariant to a reasonable number of lags chosen

9 Two other alternative approaches in order to test for co-integration were implemented. First the residuals in country by country regressions were tested for stationarity using the critical values reported by Phillips and Ouliaris (1990); however the size of each time series (31 observations) apparently has not enough power against the null in each case. Second, country-specific co-integration vectors were estimated for those countries displaying EKCs and the Johansen test of number of co-integration vectors implemented (Johansen 1988, 1991), as is mentioned in section 4.2; the countries' cointegration relationships support the presence of one cointegration vector for all countries with the only exception of New Zealand.

TABLE 8
PANEL UNIT ROOT TEST
SAMPLE 17 COUNTRIES 1960–1990

$$\Delta\varepsilon_{it} = \alpha + \beta\varepsilon_{it-1} + \sum_{j=1}^{\infty}\gamma_j \Delta\varepsilon_{it-j}$$

	Statistic	5% *cv*	Sample
(1) Levinlin	–10.446 (1)	–7.245	EKC
(2) Levinlin	–9.155 (1)	–4.599	No-EKC
(3) Levinlin	–12.258(1)	–7.4168	Whole
(4) Ipshin	–2.296 (1)	–1.850	EKC
(5) Ipshin	–1.457 (1)	–1.670	No-EKC
(6) Ipshin	–1.543 (1)	–1.670	Whole

Levinlin: Levin-Lin-Chu (Levin *et al.*, 2002).
Ipshin: Im-Pesaran-Shin Test (Im *et al.*, 2003).
Number of lags in brackets.
cv: Critical value.

A vector error correction model of order p (VECM(p)) is designed as:

$$(5) \qquad \Delta Y_{it} = \Pi_i Y_{it-1} + \Gamma_{i1}\Delta Y_{it-1} + \dots\dots\dots + \Gamma_{ip-1}\Delta Y_{it-p+1} + U_{it}$$

where Y_{it} is a 3x1 vector of country *i* time series with the variables appearing in the EKC cointegration relationship (i.e. emission per capita, income, and income squared, all of them expressed in logs). Γ_{ij} with $j = 1,........p$ are 3x3 coefficient matrices and U_{it} is a normally distributed 3x1 white noise process, with zero mean and covariance matrix Σ_{ui}.

In cointegrated models Π_i has reduced rank $r = r(\Pi) < 3$ and can be decomposed as $\Pi_i = \alpha_i \beta_i'$ where α_i and β_i are 3xr matrices containing the loading (or feedback) coefficients and the cointegrating vector respectively. Therefore, in this context, testing for weak exogeneity is equivalent to test for zero restrictions on the α_i matrix.

In order to fit a VECM as in (5) several previous steeps are necessary. Firs it is necessary to check that the country-specific EKC in (3) are effectively cointegration relationships. As was seen in the previous section, this seems to be the case since in general our panel cointegration tests rejects the absence of co-integration. Additionally, since the Engle-Granger procedure used to estimate the presence of cointegration in the EKC assumes the presence of no more than one cointegrating vector (see Engle and Granger, 1987), the Johansen's method (Johansen, 1988,1991) is used here to test the hypothesis that only one cointegrating vector exists in our country specific EKC. In general, the countries' cointegrating relationships support the presence of only one cointegration vector with the only exception of New Zealand that supports zero cointegrating relationships (results not reported). Secondly, in order to implement the Engel and Granger procedure, the residuals for every country specific EKC are estimated and recorded. Therefore,

$$EC_{it-1} = \varepsilon_{it-1} = y_{it-1} - \alpha_{0i} - \beta_{1i}x_{it-1} - \beta_{2i}x_{it-1}{}^2 \tag{6}$$

and next the loading factors α_i are estimated. Table 9 reports the results of the weak exogeneity test along with its t-values. As can be seen from columns 3 and 4 in the table, for most of the countries (fifteen out of seventeen countries) the null hypothesis of weak exogeneity cannot be rejected for the level variables (i.e. income and income squared). Only in the cases of Switzerland and New Zealand the null hypothesis of weak exogeneity is rejected at 95% confidence level for both variables income and income squared. Contrarily, in the case of emissions per capita, as can be seen in column 2 of the same table, the null hypothesis of weak exogeneity is rejected in most cases (fourteen out of seventeen cases). In addition, as can be observed in the same column, two of the tree cases where weak exogeneity is not rejected (France and Luxembourg) are borderline cases.

Summing up, there is strong evidence that in our sample of 17 rich countries exhibiting an EKC relation, income and income squared are (weakly) exogenous variables while emissions per capita is endogenous. Therefore, it makes sense to put emissions at the left hand side and the levels variables at the right hand side of our country-specific EKCs. Finally, what this weak exogeneity test implies is that any deviation from the long run relationship provokes an adjustment in emissions (not in income) toward the EKC. In general this adjustment is pretty fast with an average of 50% of the innovation returning to its steady state at the end of one year.

4.3. Common shocks or income driven EKC?

A final potential problem we address here is the possibility that the reductions in emission of the EKCs empirically corroborated here are induced by a structural change taking place at some specific time rather than by a particular threshold in income as it is assumed by the EKC hypothesis. As an example, column 2 in table 10 shows the actual date of realization of the country-specific turning points per country as predicted in the present paper. The table clearly reflects that, even though all countries in the sample have different income-turning points, the date at which they reach their corresponding turning points is pretty much the same (i.e. 14 out of 17 countries reach their turning point within the five year window from 1968 to 1973). Therefore, it is possible that the conditional process of per capita emissions could have been subject to a structural break in the vicinity of the end of the sixties and beginning of the seventies rather than be driven by a particular income per capita- turning point, which would render our income-pollution relationship miss-specified. In order to test for a possible structural break in a country by country basis, we use here Hansen (2000) test for threshold regressions. The Hansen test in the present context tests for the possibility of a structural break in the conditional realization of per capita emissions such that the estimated income-relation is no longer given by equation (3) but by the following:

$$y_{it} = \alpha_{0i} + \beta_{1i}x_{it} + \beta_{2i}x_{it}{}^2 + \varepsilon_{it} \qquad t \le t^* \tag{7}$$

$$y_{it} = \alpha_{0i} + (\beta_{1i} + \gamma_{1i})x_{it} + (\beta_{2i} + \gamma_{2i})x_{it}{}^2 + \varepsilon_{it} \qquad t > t^* \tag{7'}$$

TABLE 9
WEAK EXOGENEITY TESTS

Country	Δy_t		Δx_t		Δx_t^2	
Canada	–0.5	(–2.89)**	–0.12	(–1.44)	–2.36	(–1.48)
USA	–0.17	(–1.27)	–0.09	(–1.24)	–1.65	(–1.26)
Israel	–0.22	(–2.04)**	0.009	(0.28)	0.15	(0.26)
Japan	–0.27	(–2.36)**	–0.08	(–1.31)	–1.46	(–1.32)
Saudi Arabia	–0.84	(–4.03)**	0.005	(0.04)	0.06	(0.02)
Singapore	–0.32	(–2.38)**	0.04	(0.94)	0.55	(0.83)
Austria	–0.39	(–2.17)**	–0.07	(–1.12)	–1.28	(–1.12)
Belgium	–0.43	(–2.70)**	–0.07	(–1.47)	–1.31	(–1.47)
Finland	–0.43	(–2.78)**	–0.02	(–0.32)	–0.32	(–0.30)
France	–0.18	(–1.95)	–0.02	(–1.09)	–0.43	(–1.11)
Germany	–0.40	(–2.78)**	–0.14	(–1.67)	–2.6	(–1.68)
Italy	–0.39	(–2.52)**	–0.09	(–1.39)	–1.57	(–1.42)
Luxembourg	–0.2	(–1.92)	–0.02	(–1.17)	–0.38	(–1.17)
Netherlands	–0.29	(–2.21)**	–0.05	(–1.55)	–0.1	(–1.58)
Norway	–0.49	(–3.36)**	–0.02	(–0.72)	–0.36	(–0.77)
Switzerland	–0.68	(–3.98)**	–0.07	(–2.17)**	–1.3	(–2.17)**
New Zealand	–0.32	(–2.65)**	–0.077	(–2.0)**	–1.43	(–2.00)**

**variable significant at 5% level.

where all variables are defined as before, γ_{1i} and γ_{2i} are shift parameters and t^* is our threshold data-point to be estimated. Following Hansen (2000), the date-point is estimated here as $\hat{t}^* = \underset{t \in \Gamma}{\arg\min}\, \hat{\sigma}^2(t)$, where $\hat{\sigma}^2(t)$ is the estimated variance of residuals OLS regressions for each $t \in \Gamma$, where we have fixed $\Gamma = [0.15, 0.85]$ following the recommendation of Andrews (1990). The estimated t^* for each country in our sample of high income countries are reported in column 3 of Table 10, while the corresponding confidence intervals for the estimated breakpoints are shown in column 4 and are also presented graphically in Table 11. From both tables it is evident that even though some structural shift seems to have happened by the beginning of the eighties, the shift-times are far from being close to the time of occurrence of the turning point and moreover, the country by country confidence intervals are either too wide or with too many countries lying in different interval set.

As a result, we can conclude that even though there are evidence of a structural break in the data, it seems that the estimated data-break do not correspond with the time when according to the EKC the income started to drive emission downward. On the other hand, there is a high dispersion between and within countries regarding the predicted break-points. Finally, even though there is evidence of a structural break in all our country time-series, the existence of such a break do not invalid the presence of an EKC, it only implies a shift in the whole relationship.

V. Further Comments

As it was mentioned before, it seems that the implementation of a regulatory environmental system that highly resembles market mechanisms could explain the presence of an EKC. Barde (1994) pointed out three requirements to properly design instruments that internalize the externalities associated with contamination: first, correction of price distortions in resources and other goods; second, well defined property rights; and third, broad political acceptance of the polluter-pays principle (PPP). These conditions that are hardly met in developing countries, are in general fulfilled in developed (particularly OECD) countries.

Regarding the PPP for example, it was defined and recognized as early as 1972 for OECD countries[10]; while in 1975, the European Community adopted a similar resolution. This is consistent with our econometric results showing that by the end of the sixties and beginning of the seventies, income per capita started to drive per capita emission downward (see column 2 of Table 10). Blackman and Harrington (1998) report that in a 1989 OECD survey on economic instruments (EI) used for environmental protection, at least 14 OECD countries employed between 1 and 20 of such instruments. A total of 151 instruments were in operation, with approximately half of them being charges and a third subsides, with a variety of other instruments such as deposit-refund systems, market creation, and enforcement incentives making up the balance. Thus to get an idea of how often EI are used for environmental regulation in OECD countries, Table 12 summarizes the information of a 1998/1999 survey conducted by the OECD as well as information from the Environmental Protection Agency (EPA). The table is adapted from Barde (1999) and provides a general overview of the use of EI for pollution control in OECD countries, depicting a degree of implementation of EI by these OECD countries that is surely far ahead from the one shown by developing countries.

Intensive use of EI is an indication of policy makers trying to resemble a market mechanism for the environment. However other kinds of policies could help as well, such as the use of cost-benefit analysis in the design of the environmental policies, and the implementation of measures that make political systems more open and transparent. Evidence elsewhere seems to indicate that cost-benefit analysis is more often used in developed than in developing countries. Regarding the openness and transparency of the political system, political databases elsewhere show greater levels of democracy and participation in developed countries in comparison to developing countries.

Additionally, there is some evidence that within the group of developed countries as well, some differences in the degree of internalization of environmental problems do exist. In Table 12, it is shown that 5 out of 24 OECD members did not respond to the survey and additional information which support the presumption that the use of EI was not particularly developed in those cases. Moreover, according to the diversity and the presence of some EI, 4 of the countries that do not present evidence of an EKC in our estimates also do not present strong evidence of using EI (i.e. Ireland, Portugal, Spain and the UK). Nevertheless, a survey performed during the latest 1990s could not reflect the situation for the

[10] Barde (1994), page 3.

TABLE 10
STRUCTURAL BREAK TESTS

Country	Year of occurrence of the turning point	Predicted break $\hat{t}^*$	Confidence interval of the predicted break
Canada	1960-61	1982	1982-1982
USA	1968-69	1981	1981-1982
Israel	1971-72	1976	1976-1976
Japan	1971-72	1982	1982-1982
Saudi Arabia	1971-72	1972	1970-1973
Singapore	1978-79	1965	1965-1970
Austria	1965-66	1983	1979-1984
Belgium	1968-69	1983	1983-1983
Finland	1970-71	1982	1982-1983
France	1968-69	1982	1982-1982
Germany	1967-68	1982	1981-1983
Italy	1971-72	1983	1983-1983
Luxembourg	1972-73	1975	1972-1979
Netherlands	1960-61	1981	1981-1981
Norway	1973-74	1981	1981-1983
Switzerland	1967-68	1975	1975-1977
New Zealand	1968-69	1977	1976-1978

Note: Predicted breaks and confidence intervals estimated using Hansen's methodology (Hansen, 2000).

span of our sample 1960-1990. Addressing this argument, Table 13 displays results of a survey taken during the latest 1980s, and shows at its bottom that the weakest cases of EI implementation correspond to Ireland, Greece, Spain, UK, Turkey and possibly Portugal, a similar group to the one mentioned above that do not show evidence of an EKC in our estimations[11].

All this evidence points to the fact that some of the more obvious variables to be tested as determinants of the EKC hypothesis would not have a clear bias effect that could challenge our conclusions. However, formally testing for this type of possible bias is a pending task which is not attempted here since our interest is to focus on the usual structure of the EKC most used in the literature.

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11 We do not have any consistent database showing the degree of implementation of cost-benefit analysis within the OECD group. However we do have information regarding the degree of political openness within the OECD group in Polity IV database and other databases elsewhere. Per example Polity IV project (Marshal and Jaggers 2003) is a database with political regime characteristic spanning two centuries of information. For the period 1960-1990, every member of the OECD in our database has the highest score in indicators as: Democracy, Political Competition, and Political Participation etc. There are only four exception to this rule; Turkey for the whole period; Spain 1960-1972 (Franco Government until the end of the transition with Adolfo Suarez); Portugal 1960-1975 (final period of the Estado Novo and the military administration until socialist Mario Soares); Greece 1960-1974 (until the end of the regime of the colonels). Not surprisingly, those four countries correspond to countries without strong evidence of an EKC in this paper.

TABLE 11
CONFIDENCE INTERVALS FOR STRUCTURAL BREAKS

Year	65	66	67	68	69	70	71	72	73	74	75	76	77	78	79	80	81	82	83	84	
Canada																					
USA																					
Israel																					
Japan																					
Saudi Arabia																					
Singapore																					
Austria																					
Belgium																					
Finland																					
France																					
Germany																					
Italy																					
Luxembourg																					
Netherlands																					
Norway																					
Switzerland																					
New Zealand																					

Note: Confidence intervals estimated using Hansen's methodology (Hansen, 2000).

TABLE 12
GENERAL OVERVIEW OF THE USE OF ECONOMIC INSTRUMENTS FOR POLLUTION CONTROL, 1999

Country	Charges	Country	Charges
Australia	CH, TP, DRS, NCF,PB,S	Japan	CH, LP, S
Austria	CH, DRS, S	Luxembourg	
Belgium	CH	the Netherlands	CH, DRS, S
Canada	TP,	New Zealand	
Denmark	CH, TP, DRS, LP, S	Norway	CH, DRS, NCF, S
Finland	CH, DRS, LP, S	Portugal	
France	CH, TP, S	Spain	
Germany	CH, LP,	Sweden	CH, DRS, NCF, LP, S
Greece	CH, NCF, S	Switzerland	CH, TPS
Iceland	CH, DRS	Turkey	CH, DRS, NCF, LP, S
Ireland		UK	
Italy	CH, DRS,	US	CH, TP, DRS, PB, LP, S

CH: Charges.
TD: Tradable permits.
DRS: Deposit refund system.
NCF: Non compliance fees.
PB: Performance bonds.
LP: Liability payments.
S: Subsides.
Note: Based entirely on questionnaire replies and the EPA report on economic incentives for the US (Adapted from Barde, 1999, page 12).

TABLE 13
ECONOMIC INSTRUMENTS IN OECD COUNTRIES AS 1 JANUARY 1992

Country	Charges on emissions (of which user charges)	Charges on products (of which tax differentia-tion)	Deposit-refund systems	Tradable Permits	Enforcement Incentives
USA	5(2)	6(1)	4	8	2
Sweden	3(2)	11(2)	4		2
Canada	3(2)	7(3)	1	2	2
Denmark	3(2)	10(2)	2		
Finland	3(2)	10(2)	2		
Norway	4(2)	8(2)	3		
Australia	5(2)	1(0)	3	1	2
Germany	5(2)	3(3)	2	1	
Netherlands	5(2)	4(2)	2		
Austria	3(1)	4(2)	3		
Belgium	7(2)	2(2)	1		
Portugal	2(0)	1(1)			
France	5(2)	2(1)			
Switzerland	3(2)	2(2)	1		
Italy	3(2)	2(0)			
Japan	3(1)	1(1)			
Ireland	2(2)	1(1)			
Greece		2(1)			
Spain	3(2)				
UK	1(1)	1(1)			
New Zealand	1(1)				
Turkey					

Note: Adapted from Barde, 1994, page 16.

VI. Conclusions

There is an ongoing debate regarding the robustness of the EKC hypothesis. While Cole (2003) argues in its favor, other authors, particularly Stern (2004) and Perman and Stern (2003) argue against it. Since most of the empirical tests have assumed a common structure for every country in the samples, this *all or nothing* debate is not surprising.

However since in this paper we used for the first time a Bayesian estimator in order to test the EKC hypothesis country by country, our position is less extreme. We argue that while for some countries the EKC hypothesis is robust, for others country it is not.

In order to check for a heterogeneous rather than a common structure, in this study we tested for variable slopes instead of constant slopes to analyze the EKC for SO_2 emissions. Since a test of parameter constancy indicates that the assumption of equal slope variance across countries is not adequate in this particular case, the random coefficient model that allows for coefficient heterogeneity seems more pertinent.

For homogenous developed countries, there is strong evidence of an overall EKC. At country level, for most of the OECD members and for most of the members of the developed world the EKC hypothesis is robust. However a few members of the OECD and a few members of the developed world do not display an EKC.

Using a sample of high-income countries as defined by the World Bank, evidence of an overall EKC is found. At country-specific level, 17 out of 28 countries strongly support the EKC hypothesis, and 11 out of 28 countries do not support the EKC hypothesis.

Based in our econometrics findings together with some additional evidence analyzed, it is possible to conclude that: a) Overall, the EKC hypothesis is robust for developed countries; b) Country-specific EKC can be encountered in most of the OECD and developed countries; c) In a few OECD and developed countries the EKC is not supported by the data; d) Some evidence exist that an EKC is most likely to be found in countries with regulatory processes that resemble market mechanisms; e) An EKC automatically driven by income growth is not inferred by the data since it seems that at least policies and institutions resembling market mechanisms need to be implemented for a country to follow the virtuous path implied by the EKC; f) When (weakly) exogeneity of the income variable in the EKC relationship was tested, the data supported the theoretical presumption that the relationship goes from exogenous income per capita variable to the endogenously determined emissions per capita variable; and, g) For checking whether or not the inverted-U EKC relationship found was a result of an external shock provoking a common structural change the Hansen (2002) test was employed, which empirically showed a high dispersion of the years in which structural changes occurred among the countries displaying an EKC relationship and, therefore rendered plausible that the inverted-U is really determined by the per capita income driver, as it is implied by the theory.

References

Andreoni, J. and Levinson A. (2001), "The Simple Analytics Of The Environmental Kuznets Curve", *Journal of Public Economics,* 80: 269-286.

Andrews, D.W.K. (1990), "Tests for parameter instability and structural change with unknown change point", Discussion Paper 943, Yale University, Cowles Foundation for Research in Economics.

Arrow, K., Bolin B., Constanza R., Dasgupta P., Folke C., Holling C., Jansson B.O., Levin S., Maler K.G., Perring C., and Pimentel D. (1995). "Economic Growth, Carrying Capacity and the Environment", *Science,* 268: 520-521.

ASL and Associates. (1997). "Sulfur Emissions by Country and Year", Report Nº DE96014790, USA Department of Energy, Washington, D.C.

Barbier, E. B. (ed) (1997), Special issue: "The Environmental Kuznets Curve", *Environment and Development Economics* 2, Part 4.

Barde, J. P. (1994). "Economic Instruments in Environmental Policy: Lessons from the OECD Experience and Their Relevance to Developing Economies", Working Paper 92, OECD Development. Center.

Barde, J. P. (1999). "Economic Instruments for Pollution Control and Natural Resources Management in OECD Countries: A Survey", OECD Working Papers on Economic and Environmental Policy Integration.

Bartels, L. (1996). "Pooling Disparate Observations", *American Journal of Political Science Review,* 40: 905-42.

Beede, D. and Wheeler D. (1992). "Measuring and Explaining Cross-Establishment Variation in the Generation and Management of Industrial Waste", World Bank mimeo.

Blackman, A. and Harrington W. (1998). "Using Alternatives Regulatory Instruments to Control Fixed Point Air Pollution in Developing Countries: Lessons from International Experience", Discussion Paper 98-21, Resources for the Future.

Cole, M. A. (2003). "Development, Trade And The Environment: How Robust Is The Environmental Kuznets Curve?" *Environment and Development Economics,* 8: 557-580.

Cole, M. A. (2005). "Re-Examining the Pollution-Income Relationship: A Random Coefficient Approach", *Economics Bulletin,* 14: 1-7.

Cole, M. A., Rayner A.J. and Bates M. (1997). "The Environmental Kuznets Curve: An Empirical Analysis", *Environment and Development Economics,* 2: Nº 4, 401-416.

Dinda, S. (2004). "Environmental Kuznets Curve Hypothesis: A Survey", *Ecological Economics,* 49: 431-455.

Durbin, J. (1954). "Errors in Variables", *Review of the International Statistical Institute*, 22: 22-31.

Engle, R. and Granger, C. (1987). "Cointegration and Error Correction: Representation, Estimation and Testing", *Econometrica*, 55: 251,276.

Figueroa E. and Pastén R. (2000). "Crecimiento y Medio Ambiente: ¿Existe Automatismo en la U Invertida?", In J. M. Villasuso and R. Trejos (eds.). *Comercio e Integración de las Américas*. Banco Interamericano de Desarrollo (BID), Instituto Interamericano de Cooperación para la Agricultura (IICA). pp. 43-49.

Greene, W. (2001). "Fixed and Random Effects in Non Linear Models", Stern School of Business. New York University.

Grossman, G. and Krueger A. (1992). "Environmental Impacts of a North American Free Trade Agreement", Discussion papers in economics. Nº 158, Woodrow Wilson School of Public and International Affairs, Princeton, NJ.

Grossman, G. and Krueger A. (1995). "Economic Growth and the Environment", *Quarterly Journal of Economics,* 110: 2, 352-377.

Hansen, B. E. (2000). "Sample splitting and threshold estimation", *Econometrica*, 68: 575-603.

Hausman, J. (1978). "Specification Test in Econometrics". *Econometrica*, 46: 1251-71.

Heston, A., Summers R. and Aten B. (2002). "Penn World Table Version 6.1", Center for International Comparisons at The University of Pennsylvania (CICUP).

Hettige, H., Mani M. and Wheeler D. (2000). "Industrial Pollution in Economic Development: The Environmental Kuznets Curve Revisited", *Journal of Environmental Economics and Management,* 62: 445-476.

Holtz-Eakin, D. and Selden T.M. (1995). "Stocking the fires? CO_2 Emissions and Economic Growth", *Journal of Public Economics,* 57: 85-101.

Hsiao, C. (2003). "Analysis of Panel Data", Econometric Society Monograph 34, 2nd Edition, New York, Cambridge University Press.

Hsiao, C. and Pesaran M.H. (2004). "Random Coefficients Panel Data Model", Working Paper CWPE 0434.

Im, K. S., Pesaran M.H. and Shin Y. (2003). "Testing For Unit Root in Heterogeneous Panels", *Journal of Econometrics,* 115: 53-74.

Johansen, S. (1988). "Statistical Analysis of Cointegration Vectors", *Journal of Economic Dynamics and Control,* 12: 231-254.

Johansen, S. (1991). "Estimation and Hypothesis Testing Of Cointegrated Vectors in Gaussian VAR Models", *Econometrica,* 59: 1551-1580.

Johansen, S. (1992). "Testing Weak Exogeneity and the Order of Cointegration in U.K. Money Demand Data", *Journal of Policy Modeling,* 14: 313-334.

John A., and Pechenino R. (1994). "An Overlapping Generations Model of Growth and the Environment", *The Economic Journal,* 104: 1393-1410.

Jones L. E. and Manuelli R.E. (1995). "A Positive Model of Growth and Pollutions Controls", Working Paper Nº 5205, NBER.

Kaufmann R. K., Davidsdottir B., Garnham S. and Pauly P. (1998). "The Determinants of Atmospheric SO_2 Concentrations: Reconsidering the Environmental Kuznets Curve", *Ecological Economics,* 25: 209-220.

Koop, G. and Tole L. (1999). "Is There an Environmental Kuznets Curve for Deforestation?", *Journal of Development Economics,* 58: 231-244.

Levin, A. Lin C. and Chu J. (2002). "Unit Root in Panel Data: Asymptotic and Finite-Sample Properties", *Journal of Econometrics*, 108: 1-24

List, J. A. and Gallet C.A. (1999). "The Environmental Kuznets Curve: Does One Size Fit All?", *Ecological Economics*, 31: 409-423.

Maddala, G. S. (1971). "The Use of Variance Components Models in Pooling Cross Section and Time Series Data", *Econometrica,* 39: 341-358.

Markandya, A., Golub A. and Pedroso-Galinato S. (2006). "Empirical Analysis of National Income and SO_2 Emissions in Selected European Countries", *Environmental & Resources Economics,* 35: 221-257.

Marshal, M. G. and Jaggers K. (2003). "Polity IV Project: Political Regime Characteristic and Transition 1800-2002", Center for International Development and Conflict Management (CIDCM), University of Maryland, College Park. MD

Mazzanti, M., Musolesi A. and Zoboli R. (2006). "A Bayesian Approach to the Estimation of Environmental Kuznets Curves for CO_2 Emissions", Nota di lavoro 121.2006, Fondazione Enri Enrico Mattei.

Panayotou, T. (1993). "Empirical Test and Policy Analysis of Environmental Degradation at Different Stages of Economic Development", Working Paper WP238, Technology and Employment Programme, International Labour Office, Geneva.

Panayotou, T. (1997). "Demystifying the Environmental Kuznets Curve: Turning a Black Box into a Policy Tool", *Environment and Development Economics,* 2: 465-484.

Perman, R. and Stern D.I. (2003). "Evidence from Panel Unit Root and Cointegration Tests That the Environmental Kuznets Curve Does Not Exist", *Australian Journal of Agricultural and Resources Economics,* 47: 325-347.

Phillips P. C. B. and Ouliaris S. (1990). "Asymptotic properties of residual based tests for cointegration", *Econometrica,* 58: 165-193.

Selden, T. M. and Song D. (1994). "Environmental Quality and Development: Is There a Kuznets Curve For Air Pollution Emissions?", *Journal of Environmental Economics and Management,* 27: 147-162.

Shafik, N. (1994). "Economic Development and Environmental Quality: An Econometric Analysis", *Oxford Economic Papers,* 46: 757-773.

Shafik, N. and Bandyopadhyay S. (1992). "Economic Growth and Environmental Quality: Time Series and Cross Section Evidence", Policy research working paper N° WPS904, World Bank.

Stern, D. I. (1998). "Progress on the Environmental Kuznets Curve?" *Environment and Development Economics,* 3: 173-196.

Stern, D. I. (2003). "The Environmental Kuznets Curve", Paper presented to the International Society for Ecological Economics. June.

Stern, D. I. (2004). "The Rise and Fall of the Environmental Kuznets Curve", World Development Vol. 32, N° 8: 1419-1439.

Stern, D. I. and Common M.S. (2001). "Is there an environmental Kuznets curve for sulfur?", *Journal of Environmental. Economics. and Management,* 41 (2), 162-178.

Suri, V. and Chapman D. (1998). "Economic Growth, Trade and Energy: Implications for the Environmental Kuznets Curve", *Ecological Economics,* 25 (2): 195-208.

Swamy, P. (1970). "Efficient Inference in a Random Coefficient Regression Model", *Econometrica,* 38: 311-23.

Swamy, P. and Mehta J.S. (1975). "Bayesian and Non-Bayesian Analysis of Switching Regressions and of Random Coefficient Regression Models", *Journal of the American Statistical Association,* 70: 593-602.

World Bank (1992). "Development and the Environment", World Development Report 1992, Development and the Environment, Oxford University Press, NY, p. 11.

World Bank (2004). "World Development Indicators", The World Bank Group.

Yandle, B., Bhattarai M. and Vijayaraghavan M. (2004). "Environmental Kuznets Curves: A Review of Findings, Methods, and Policy Implications", PERC. Research Study 02-1.

Zellner, A. (1962). "An efficient Method of Estimating Seemingly Unrelated Regressions and Tests for Aggregation Bias", *Journal of the American Statistical Association,* 57: 348-368.

Wu, D.M. (1973). "Alternative Test of the Independence between Stochastic Regressors and Disturbances", *Econometrica*, 41: 134-139.

APPENDIX A

The Random Coefficient Model

The Swamy's random coefficient model in the context of this paper can be specified as:

$$y_i = X_i B_i + u_i \quad i = 1,2,.........N$$

and

$$B_i = B + w_i$$

where y_i is a vector of T observations of natural logarithm of sulfur dioxide to population for the i th country, X_i is a $T \times 3$ matrix with ones in the first column, the natural logarithm of income per capita in the second column and the square of the natural logarithm of income per capita in the third column. The disturbances u_i and w_i have zero expectation and they have the additional characteristics $E(u_i u_j) = E(w_i w_j) = 0$, $E(u_i u_i) = \sigma_{ii} I$ and $E(w_i w_i) = \Omega$. Under these assumptions, the unbiased estimator of B is given by:

$$\hat{B} = [\alpha_0, \beta_1, \beta_2]' = \left\{ \sum \left[\sigma_{ii} \left(X_i ' X_i \right)^{-1} + \Omega \right]^{-1} \right\}^{-1} \left(\sum \left[\sigma_{ii} \left(X_i ' X_i \right)^{-1} + \Omega \right]^{-1} b_i \right),$$

where b_i are the OLS estimators of B_i where the time series of country i is used. The variance covariance matrix Ω and σ_{ii} are estimated by:

$$\hat{\sigma}_{ii} = \hat{u}_i \hat{u}_i ' / (T-3)$$

$$\hat{\Omega} = \left[\left(\sum b_i b_i ' - \sum b_i \sum b_i ' / N \right) / (N-1) \right]$$

The Swamy-Mehta estimator is the best linear unbiased estimator (BLUE) for B_i and is given by:

$$B_i = \hat{B} + \hat{\Omega} X_i '(X_i \hat{\Omega} X_i ' + \hat{\sigma}_{ii} I)^{-1} (y_i - X_i \hat{B})$$

Finally, the test of parameter constancy can be developed with the statistic proposed by Swamy which follows a chi-square distribution with 2(N – 1) degrees of freedom and is given by:

$$g = \sum \left[\left(b_i - D \right)' X_i ' X_i \left(b_i - D \right) / \hat{\sigma}_{ii} \right], \text{ where } D = \left(\sum \hat{\sigma}_{ii}^{-1} X_i ' X_i \right)^{-1} \sum \hat{\sigma}_{ii}^{-1} X_i ' X_i b_i.$$