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Evidence from panel unit root and cointegration tests that the Environmental Kuznets Curve does not exist*

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The Environmental Kuznets Curve (EKC) hypothesis – an inverted U-shape relation between various indicators of environmental degradation and income per capita – has become one of the 'stylised facts' of environmental and resource economics. This is despite considerable criticism on both theoretical and empirical grounds. Cointegration analysis can be used to test the validity of such stylised facts when the data involved contain stochastic trends. In the present paper, we use cointegration analysis to test the EKC hypothesis using a panel dataset of sulfur emissions and GDP data for 74 countries over a span of 31 years. We find that the data is stochastically trending in the time-series dimension. Given this, and interpreting the EKC as a long run equilibrium relationship, support for the hypothesis requires that an appropriate model cointegrates and that sulfur emissions are a concave function of income. Individual and panel cointegration tests cast doubt on the general applicability of the hypothesised relationship. Even when we find cointegration, many of the relationships for individual countries are not concave. The results show that the EKC is a problematic concept, at least in the case of sulfur emissions.

1. Introduction

The Environmental Kuznets Curve (EKC) hypothesis – an inverted U-shape relation between various indicators of environmental degradation and income per capita – has become one of the 'stylised facts' of environmental and resource economics (e.g., Stokey 1998). This is despite considerable criticism on both theoretical and empirical grounds (e.g., Stern *et al.* 1996; Ekins 1997; Ansuategi *et al.* 1998; Stern 1998; Stern and Common

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2001). The EKC has been interpreted by many as indicating that no effort should be made to adopt environmental policies in developing countries – when those countries become rich the current environmental problems will be addressed by policy changes adopted at that later time (e.g., Beckerman 1992). As a corollary it is implied that developing countries are 'too poor to be green' and that little in the way of environmental clean-up activity is being conducted in developing countries. These views are challenged by recent evidence that, in fact, pollution problems are being addressed and remedied in developing economies (e.g., Dasgupta *et al.* 2002). In addition to the data and case studies provided by Dasgupta *et al.* (2002), Stern (2002) and Stern and Common (2001) show that for sulfur – widely believed to show the inverted U-shape relation between emission and income per capita – emissions in fact rise with increasing income at all levels of income, but that there are strong time effects reducing emissions in all countries across all income levels. In our opinion, this new evidence supersedes the debate about whether some pollutants show an inverted U-shape curve and others – for example carbon dioxide and 'new toxics' (Dasgupta *et al.* 2002) – a monotonic relationship. All pollutants show a monotonic relation with income, but over time pollution has been reduced at all income levels, *ceteris paribus*. Similarly the debate about whether the downward sloping portion of the EKC is an illusion resulting from the movement of polluting industries to offshore locations is also now moot. This phenomenon might lower the income elasticity of pollution in developed economies relative to developing economies, but it does not seem sufficient to make it negative. The true form of the emissions–income relationship is a mix of two of the scenarios proposed by Dasgupta *et al.* (2002) illustrated in figure 1. The overall shape is that of their 'new toxics' EKC – a monotonic increase of emissions in income. But over time this curve shifts down. This is analogous to their 'revised EKC' scenario, which is intended to indicate that over time the conventional EKC curve shifts down.

Cointegration analysis can be used to test the validity of supposed stylised facts, such as the EKC, when the data are time series that are integrated. Classical regression analysis assumes that all the variables involved are stationary. A covariance stationary variable has a constant mean and variance, while strict stationarity implies that all aspects of the distribution are identical in any sample taken from the data. Integrated variables are one class of non-stationary variable.[1] The simplest example of an integrated variable is a

[1] Variables with a deterministic time trend are also non-stationary. For these variables the deterministic time trend should be removed before carrying out regression analysis. However, removing a deterministic time trend from integrated variables does not render them stationary and differencing must be applied.

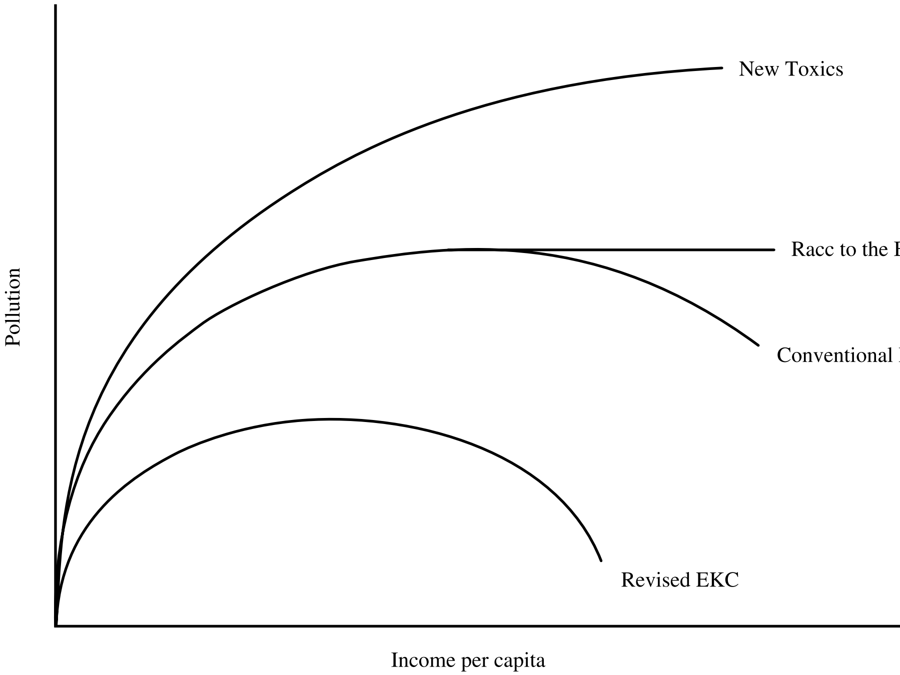

Figure 1 Environmental Kuznets Curve for different scenarios (Source: Dasgupta *et al.* 2002).

random walk. More complex processes include random walks with noise and integrated random walks. These variables are often also referred to as unit root, or stochastically trending variables in the published econometric literature. Some integrated variables can be made stationary after first differencing, in which case they are termed integrated of order one or I(1) variables. A particular historical realisation of an integrated process is called a stochastic trend.

Appropriate methods of inference depend in important ways on whether data are integrated or not. In general, the residual from a regression of integrated variables is also integrated. This violates the assumptions of the classical regression model and the distribution of the regression parameters is highly non-standard. This is a so-called spurious regression (Granger and Newbold 1974). However, if the integrated variables share stochastic trends, and no relevant integrated variables are omitted or irrelevant variables included, the residual will be stationary. In this case, the variables are said to be cointegrated. If additionally there is no serial correlation in the residual the traditional regression inference applies. Hence, cointegration testing is a powerful test of misspecification. Not only can it assess whether the regression can be interpreted in the conventional way but it can also test whether the appropriate variables are included in the model. In the present paper, we use cointegration analysis to test the EKC hypothesis. The results are an additional line of evidence supporting the arguments discussed.

Most EKC analyses use panel datasets (Stern 1998). Only recently has any attention been paid to the time series properties of this data (e.g., Stern 2000, 2002; Stern and Common 2001). All other existing studies of the EKC assume that the data are stationary in the time-series dimension. Given the accumulated evidence that income – the explanatory variable in the EKC – may be an integrated variable, this assumption should at least be questioned.[2] It is well known that a regression between integrated or deterministically trending variables can indicate the existence of a significant relationship between the variables using classical inference when, in fact, the variables are unrelated and the regression is spurious. The same is true for panel data when the data are non-stationary in the time-series dimension. It is, therefore, important to test whether the variables used in EKC studies are integrated; and if they are, to take this non-stationarity into account in subsequent modelling and inference.

In the present paper we systematically test the stationarity assumption and examine its implications for the EKC. Using recently developed tests for unit roots and cointegration in panel data, we find that the data are integrated in the time-series dimension, that there is no single cointegrating relation between emissions and income and income squared in the dataset as a whole, and that although there may be cointegrating relations in some individual countries, only some of these estimated relations support the EKC hypothesis. Even though the environmental Kuznets curve is a non-linear function of income, it can be analysed using linear cointegration methods. The EKC relation is linear in the parameters and therefore can be analysed using linear regression and related methods. Also, although no cointegrating relation should be expected between income and income squared, we test for a relation between emissions and these two variables not for a relation between the two income variables themselves.

The present paper deals only with the relationship between sulfur emissions and income. There are two emissions variables with globally representative, long-term datasets – sulfur and carbon dioxide. Of these, sulfur emissions are widely felt to be the pollutant that most strongly supports the EKC hypothesis (e.g., Grossman and Krueger 1991; Selden and Song 1994; Shafik 1994; Panayotou 1997). Our results regarding a unit root in income mean that unit root and cointegration tests must also be applied in EKC

[2] List (1999) tests for unit roots in the deviations of regional USA pollution emissions from the national average. He finds no unit roots in these deviations, but does not test for a unit root in the national average.

studies concerning other pollutants and environmental impacts, as income is the explanatory variable in all EKC studies.

The present paper is structured as follows. Section 2 briefly surveys the existing published literature in support of and against the existence of a sulfur EKC and discusses the hypotheses that we test later in the present paper. Section 3 describes the data used in the analysis and Section 4 examines its time-series properties. We then test (Section 5) for the existence of cointegrating relationships between the variables of interest. In Section 6 we specify and estimate a dynamic panel model in error correction form, and then test whether the unconstrained individual country estimates converge to a common global cointegrating vector or EKC relation. Section 7 concludes.

2. The EKC for sulfur: evidence and hypotheses

Studies which examine the relationship between sulfur emissions or concentrations and income include Grossman and Krueger (1991), Shukla and Parikh (1992), Panayotou (1993, 1995, 1997), Selden and Song (1994), Shafik (1994), Cole *et al*. (1997), de Bruyn (1997), Vincent (1997), de Bruyn *et al*. (1998), Kaufmann *et al*. (1998), Torras and Boyce (1998), List and Gallet (1999) and Stern and Common (2001). A key parameter in the EKC published literature is the turning point – the level of per capita income at which emissions or impacts stop rising and begin to decline. Stern *et al*. (1996), Stern (1998), and Stern and Common (2001) show that there is a very wide range of turning points in the literature listed above – from approximately US$3000 in 1990 Purchasing Power Parity (PPP) Dollars to over US$100 000. Estimated turning points appear to depend on a number of factors. First, relations estimated with ambient concentrations have a lower turning point. This is because industrial and urban decentralisation and the building of higher chimneys and smokestacks in the course of economic development tend to lower concentrations even if total emissions do not decline. Second, using PPP exchange rates will yield higher estimated turning points than (incorrectly) using ordinary exchange rates. Additional explanatory variables and different functional forms will also have an effect on the results. However, even among a comparable group of studies using emissions and PPP income, there is still a wide range of turning points from approximately $8000 to the maximum levels reported. Stern and Common (2001) find that this variation is related to the range of income levels included in the studies. The Cole *et al*. (1997) sample includes 11 Organization of Economic Cooperation and Development (OECD) countries and has as a turning point of approximately US$8000. Selden and Song (1994) include 22 OECD and eight developing countries and yield a turning point of approximately US$10 000. List and Gallet (1999) include all 50 USA states from 1929 to

1994, which provides a range of incomes from US\$1162 to US\$22 462 in 1987 USA dollars. Their estimated turning point is approximately US\$22 000. The sample used by Stern and Common (2001) covers 73 countries over 31 years and, therefore, is the most comprehensive. The estimated global turning point is the highest found at approximately US\$100 000.

Stern and Common (2001) argue that this indicates that the EKC model is misspecified. Omitted variables that are correlated with gross domestic product (GDP) result in estimates of the emissions–income relation that are biased and vary according to the sample selected. This is supported by the results of Hausman tests, which find that the random effects model cannot be consistently estimated. Although the fixed effects model can be consistently estimated, the results are conditional on the sample used. Stern and Common (2001) also reject pooling OECD and non-OECD samples to estimate a global EKC model.

Dijkgraaf and Vollebergh (1998) found similar results for carbon emissions in OECD countries. The turning point for an EKC estimated on their entire dataset was within the sample range of incomes. These results fly in the face of the view that there is a low turning point for sulfur emissions while the carbon-income emissions relation is monotonic (Arrow *et al.* 1995). All previous carbon studies had used more globally representative datasets. Dijkgraaf and Vollebergh (1998) also found that results varied widely across individual countries and that pooling the data could be rejected.

The present study is complementary to Stern and Common (2001). While they note that the data may contain stochastic trends, they do not test this. In the present paper, we use recent developments in unit root tests for panel data to address this question. Similarly, they note that EKC regressions may be spurious regressions as evidenced by highly serially correlated residuals, but do not explicitly test this hypothesis. We use recent developments in panel data cointegration tests to test whether the EKC models cointegrate. Finally, Stern and Common (2001) use a classic Chow test to test whether OECD and non-OECD data can be pooled. In the present paper, we test explicitly whether we can impose a single cointegrating vector on all the countries in the sample.

The sulfur EKC hypothesis can be examined within the context of the panel regression model:

$$\ln\left(\frac{M}{P}\right)_{it} = \alpha_i + \chi_t + \delta_i t + \beta_{1,i}\ln\left(\frac{Y}{P}\right)_{it} + \beta_{2,i}\left[\ln\left(\frac{Y}{P}\right)\right]^2_{it} + \varepsilon_{it} \tag{1}$$

where M is emissions, Y is constant price PPP GDP, P denotes a country's population, and t is a deterministic time trend. The variables are observed over a panel of countries ($i = 1 \dots N$) and time periods ($t = 1 \dots T$). We assume that the random disturbances ε_{it} are independent across countries, with variances that may differ among countries. The panel regression model is heterogeneous as the parameters associated with ln (Y/P) and [ln (Y/P)]^2 are allowed to vary from country to country. Additional sources of heterogeneity enter through country-specific effects (α_i), time-specific effects (χ_t) to control for disturbances affecting all countries in the panel at some point in time in a common way, and country-specific linear trends ($\delta_i t$).

The EKC hypothesis is that the EKC has a common form, with $\beta_{1i} > 0$ and $\beta_{2i} < 0$ for all i. The majority of empirical EKC studies using panel data use fixed effects or random effects estimators. Although allowing intercept shifts, the estimators constrain slope parameters to be equal over countries. These homogeneity restrictions imply that the 'turning point' level of income per capita at which emissions per capita begin to decline is identical for all countries. The country and time specific effects merely alter the level of emissions at this turning point. In contrast to the usual practice, we do not impose the homogeneity restrictions a priori, preferring instead to test subsequently whether it is legitimate to impose them.

The EKC model described by equation (1) is a static model. We specify and estimate a dynamic EKC panel model in Section 6.

3. Data

The dataset we use is relatively large in both the N and T dimensions. Estimated sulfur emissions for a broad set of countries are taken from a larger database constructed by ASL and Associates (ASL and Associates 1997; Lefohn *et al.* 1999) for the period 1850–1990. These are bottom-up estimates. Emissions are based on the use of hard coal, brown coal, and petroleum and the extent of metal smelting activity, taking into account estimated sulfur content and estimated sulfur retention or removal from waste streams. This means that the emissions coefficients associated with each fuel and activity, change over time in line with expert opinion. This database has provided, for the first time, estimates of annual sulfur emissions (as opposed to concentrations) at the country level for most of the countries of the world over a century and a half. Income (GDP) is measured in constant price, PPP adjusted, income, and is taken from the Penn World Table. Our panel consists of all countries for which a full set of both GDP and sulfur emissions is available for 1960–1990. There are 74 such countries covering 81 per cent of the world population in 1990. The major region omitted from the sample is the former Soviet Union and some eastern European countries.

Estimated emissions by ASL and Associates (1997) for many developed countries, such as West Germany, Canada, Sweden, and Japan, differ substantially from the better-known OECD estimates. The UK and USA data are similar in both databases. The OECD estimates for the former group of countries tend to overestimate emissions in the early 1970s, and underestimate emissions in the late 1980s and the 1990s relative to the ASL data. The ASL data show much smaller reductions in emissions over time for these countries. In the case of Sweden, emissions rise over time instead of decline. In addition, Vincent (1997) shows a steep decline in Malaysian emissions between 1988 and 1989 because of a switch to gas-fired electricity generation. The ASL data show a big increase in emissions because of increased coal burning at exactly the same point in time. A priori it is not clear whether the OECD or the ASL data is of higher quality. The OECD data is submitted by member governments and may vary widely in quality. The ASL data uses a uniform methodology, but could be poorer than the best individual country estimates in the OECD database. At the global level the estimates are broadly consistent with other earlier globally representative estimates (e.g., Möller 1984; Dignon and Hameed 1989; Hameed and Dignon 1992).

Stern (2002) compares estimates of the EKC using a sample of this data for 64 countries for 1973–1990 to estimates for the 1960–1990 period for 74 countries that we present here. The results are very similar for the two samples but clearly estimation precision is reduced in the smaller sample. The results in the current paper are, therefore, not likely an artefact of the choice of sample period.

4. Time-series properties of the data

For each variable, we conduct unit root tests for each individual country as well as for the panel as a whole using panel unit root tests. The panel tests should be more powerful, but it is worth comparing the results to individual time-series tests.

4.1 Individual country unit root tests

The testing procedure followed the search method proposed by Campbell and Perron (1991) and elaborated by Holden and Perman (1994). This employs a sequence of F and t-type tests beginning from an Augmented Dickey–Fuller (ADF) regression including both a deterministic trend and intercept term. The lag lengths in the ADF regressions were chosen separately for each country using the Hall (1991) procedure – a conventional step-down procedure that begins with a preselected maximum lag in the ADF regression, which ensures that the residuals are approximately white noise. The maximum lag length considered was three. We subtracted common (cross-country) time means from the data. This transformation is equivalent to estimating each individual country regression with time dummies whose regression coefficients are common across all countries and are those that would be obtained from a pooled regression over the whole panel.

The results of these tests support the view that each time series is an I(1) process. For both the log of sulfur emissions per capita and the log of income per capita, the unit root null could be rejected in, at most, 10 of the 74 countries (although not the same countries in each case). We found similar results in the case of the squared transformation of log per capita income. This is as expected from the results of Ermini and Granger (1993), who find that for low degree polynomials the I(1) property of the untransformed series is retained although any constant drift term in the original random walks is replaced by a linear deterministic time trend in the transformed series.[3]

4.2 Panel unit root tests

In recent years, frameworks have been developed for implementing unit root tests in panel data (Levin and Lin 1993; Quah 1994; Pedroni 1995; Im *et al.* 1997; Maddala and Wu 1999). The procedures have been surveyed and discussed in Maddala and Kim (1998) and the 1999 Supplement of the Oxford Bulletin of Economics and Statistics. By exploiting more information, panel unit root tests offer the prospect of ameliorating some important weaknesses of existing single time series tests, including low-power and large-size distortions. Exact comparisons of power are questionable, though, because there is an important difference in what is being tested (Maddala and Kim 1998). Panel unit root tests have as the null hypothesis a unit autoregressive root for every country in the panel, whereas an individual series test has as the null a unit root in that series, independently of what might be the case elsewhere.

In the present study, we report two forms of panel unit root test statistic, one similar in spirit to the Levin and Lin (1993) testing framework (hereafter called the 'panel' statistic), and the other based on the group mean t statistic developed by Im *et al.* (1997) (hereafter called the 'group' statistic). Test statistics are reported in table 1, for regressions including and excluding country-specific linear trends. The 'panel' statistics are derived from regressions including time dummies to eliminate common time effects that might

[3] Full details of the unit root tests for each variable and for each country can be downloaded from: http://homepages.strath.ac.uk/~hbs96107/ekc.htm

Table 1 Panel unit root test statistics

Ln (Y/P)		
Panel: Regression without trends	8.93	Do not reject unit root null
Panel: Regression with trends	4.34	Do not reject unit root null
Group: Without trends and without common time dummies	1.41	Do not reject unit root null
Group: Without trends but with time dummies	8.86	Do not reject unit root null
Group: With trends and common time dummies	1.27	Do not reject unit root null
Ln (M/P)		
Panel: Regression without trends	1.23	Do not reject unit root null
Panel: Regression with trends	1.41	Do not reject unit root null
Group: Without trends and without common time dummies	−0.05	Do not reject unit root null
Group: Without trends but with time dummies	−0.23	Do not reject unit root null
Group: With trends and common time dummies	−2.51	Reject unit root null
[ln (Y/P)]^2		
Panel: Regression without trends	9.23	Do not reject unit root null
Panel: Regression with trends	3.97	Do not reject unit root null
Group: Without trends and without common time dummies	2.67	Do not reject unit root null
Group: Without trends but with time dummies	9.27	Do not reject unit root null
Group: With trends and common time dummies	0.91	Do not reject unit root null

otherwise impart cross-country error correlations. We report the ‘group’ statistics with and without these time dummies. Each statistic is constructed to have an asymptotic standard normal distribution.

The statistics point strongly to the presence of a stochastic trend in each of the three series for all countries in the panel. The panel statistics reinforce the findings of the individual country ADF test statistics, suggesting that the widespread failure to reject the null of non-stationarity is not attributable to low power.

4.3 Implications

Given the presence of stochastic trends in the data, conventional measures of significance, such as t and F statistics and R^2 obtained from the least squares estimation of the EKC model in equation (1), cannot be relied upon to distinguish between true long run relationships and spurious regressions. For example, in panel spurious regressions, t and F statistics are divergent, making the probability that a null will be rejected go to one as N increases (see Kao 1997). But even when regressions cointegrate the distributions of the parameters are non-standard if the residual is serially correlated. It follows from this that significance tests in much of the existing EKC published literature are likely to have been conducted with critical values far smaller in absolute value than they should have been. A re-examination of EKC regression results might, in many cases, find that apparently significant relationships are spurious and not significant at all. Our next task is to test for the existence of real (as opposed to spurious) sulfur emissions–income relationships using cointegration analysis.

5. Cointegration analysis

Just as we carried out both individual and panel unit root tests, we employ both individual country and panel cointegration tests.[4]

5.1 Individual country cointegration tests

In the present study, we test for cointegration in the individual countries by carrying out an ADF test on the residuals from a cointegrating regression of the form of equation (1), but rather than pooling the data we estimate a separate regression for each country. The lag length in the ADF regression is chosen to achieve serially uncorrelated residuals.

Given that we have no prior knowledge of which combination (if any) of deterministic trends and time dummies should be included in the cointegrating regressions, the most complete approach is to entertain all three possibilities. If attention is restricted only to individual equation analysis, then the inclusion of time dummies is not necessary. However, as we will later be examining the panel as a whole, these cross-country effects must be controlled in order to avoid cross-country dependence in the errors. We label the three permutations as case 1 (equation (1) excluding heterogeneous deterministic trends and time dummies), case 2 (equation (1) excluding heterogeneous deterministic trends), and case 3 (equation (1) in full).

Table 2 summarises the ADF cointegration test results. The test statistics provide only weak support for the contention that cointegration is pervasive across individual countries in the panel. In just under one half of all cases (35 out of 74 countries), cointegration between emissions per capita and the first and second powers of income per capita could not be found for any combination of trend and dummies.

Table 3 lists some qualitative results concerning the point estimates from all individual country regressions. The table entries are not restricted to include only the regressions that cointegrate. For the regressions excluding time trends, 42 out of the 74 countries have signs in conformity with the EKC hypothesis. However, over one third of the countries appear to have U shaped (rather than inverted U shaped) emissions–income relationships (if

[4] Full details of results can be obtained online from: http://homepages.strath.ac.uk/~hbs96107/ekc.htm

Table 2 Significant cointegration ADF statistics in individual regressions

Model used	Proportion of ADF t statistics significant at 10% or better
Case 1: No trends and no time dummies	17/74
Case 2: No trends but with time dummies	16/74
Case 3: With trends and time dummies	31/74
Countries with cointegration in all 3 models	7/74
Countries with cointegration in 2 models	11/74
Countries with cointegration in 1 model only	21/74
Countries with cointegration in no model	35/74

ADF, Augmented Dickey–Fuller

Table 3 Individual regressions including time dummies, with and without heterogeneous trends

Model	$\beta_1 > 0$ and $\beta_2 < 0$ (Inverted U shaped emissions–income relationship)	$\beta_1 < 0$ and $\beta_2 > 0$ (U shaped emissions–income relationship)	$\beta_1 > 0$ and $\beta_2 > 0$ or $\beta_1 < 0$ and $\beta_2 < 0$
With time dummies but no trend	42/74	26/74	6/74
With time dummies and trend	34/74	36/74	4/74

indeed any relationship exists at all). The relative proportions are even less favourable to the EKC hypothesis when time trends are included.

5.2 Panel cointegration tests

Like unit root tests, individual (country) cointegration tests suffer from low power. Even if the postulated EKC relationship were generally true, a researcher would accept the non-cointegration hypothesis far more often than should be done. The pooling used in panel cointegration tests can improve power relative to single country based tests. A good, recent survey of the panel cointegration published literature can be found in Banerjee (1999). We employ residuals-based tests of the null of no cointegration, as developed by Pedroni (1999), which are appropriate for heterogeneous panels in which both N and T are of moderately large dimension.

Pedroni's tests are of two types. One set, based on Levin and Lin (1993) and hereafter called 'panel' statistics, pools over the *within* dimension. Numerator and denominator components of the test statistics are summed separately over the N dimension. A second set – in the spirit of Im *et al.* (1997), and hereafter called 'group' statistics – pools over the *between* dimension, obtaining the ratio of numerator to denominator for each country prior to aggregating over the N dimension. In both cases, under the null hypothesis, the variables are not cointegrated for each panel member; the alternative asserts that a cointegrating vector exists for each individual, although this vector may be unique for each individual.[5] However, the alternative hypotheses for the two types of test differ in the following way: The panel statistics test the null that the first order autoregressive coefficient $\rho_i = 1$ for all i against the alternative $\rho_i = \rho < 1$ for all i, whereas the group statistics test the null $\rho_i = 1$ for all i against the alternative $\rho_i < 1$ for all i. In other words, the alternative hypothesis for the panel test asserts that the autoregressive coefficient in the auxiliary ADF regression of the cointegrating residuals is the same for every individual. Maddala and Wu (1999) argue that this alternative is unreasonable and that the group statistics are more appropriate because their alternative hypothesis is less restrictive – simply that the autoregressive coefficient in the auxiliary regression is less than unity, although it may differ for each individual.

The test statistics, shown in table 4, do not yield an unambiguous conclusion about the existence of cointegration over the panel. We report statistics for regressions with both time dummies and linear trends to show the sensitivity of our results to modelling assumptions. Our preferred model is that in the first row, which includes time dummies (to eliminate cross-country common time effects that would otherwise create cross-equation dependence in the error terms), but does not include country-specific deterministic time trends. Regressions without time dummies were not investigated, as that case is of little practical importance given the consensus that time dummies are necessary to validate the conventional estimation assumption of cross-section independence.

For the preferred model, five out of the seven statistics suggest cointegration over the panel as a whole at the 5 per cent level or better. However, the two Rho statistics suggest no cointegration in this specification (or in the other). Inclusion of deterministic trends does little to alter the inference regarding cointegration.

It is important to note that inference is rather sensitive to the choice of the maximum lag length allowed for in the testing procedure (Pedroni

[5] The form of the alternative hypothesis might seem very restrictive and a referee asserted that we were wrong on this point. However, as Banerjee (1999) states, for the two types of panel cointegration test used here:

> 'The first category of tests uses the following specification of null and alternative hypotheses: H_O: $\gamma_i = 1$, for all i, H_A: $\gamma_i = \gamma < 1$ for all i. The second category uses H_O: $\gamma_i = 1$, for all i, H_A: $\gamma_i < 1$ for all i.'

where γ_i (ρ_i in our notation) is the autoregressive coefficient in the i^{th} country ADF regression.

Table 4 Panel cointegration test statistics

	Panel statistics				Group statistics		
	V	rho	PP	ADF	rho	PP	ADF
Time dummies included but not trends	2.37‡	–0.02	–2.16†	–1.87†	1.06	–2.83‡	–4.44‡
Both trends and time dummies included	1.23	–0.004	–3.90‡	–5.33‡	1.68	–3.80‡	–8.84‡

V, non-parametric variance ratio statistic; rho, non-parametric test statistic analogous to the Phillips and Perron (PP) rho statistic; PP, non-parametric statistic analogous to the PP *t* statistic; ADF, parametric statistic analogous to the augmented Dickey-Fuller statistic. All statistics distributed as standard normal as *T* and *N* grow large. Rejection of the null of no cointegration is one-sided and involves: variance ratio, large positive values imply cointegration (at 5% significance, reject null of no cointegration if V > 1.645); other six, large negative values imply cointegration (at 5% significance, reject null of no cointegration if statistic < –1.645). †null of no cointegration is rejected at the 5% level; ‡null of no cointegration is rejected at the 1% level

1997). While there are well-accepted routines for choosing lag lengths in single regressions, there is no robust equivalent when dealing with panel estimation. We chose a maximum lag truncation of three using a general-to-specific pretesting procedure.

Subject to all these qualifications, there is reasonable support for the hypothesis that there is a cointegrating relationship between emissions per capita and first and second powers of income per capita for each country in the panel. This suggests that the acceptance of the non-cointegration null in the individual country cointegration tests could be because of low power. But it is important to note that although the alternative hypothesis does imply that there is cointegration in every country it does not imply that there is a single cointegrating vector as is implicitly assumed in most EKC studies. In fact, there is a very large variability of parameter estimates in the individual country functions embodied in equation (1). For a substantial proportion of the countries in our sample, the implied cointegrating relationships are U-shape or monotonic.

6. Dynamic EKC models and testing for a common long-run vector

Explanations given for the existence of an EKC relationship imply that processes of adjustment to the long-run equilibrium are likely to be slow. The static EKC model specified in equation (1) precludes any modelling of such adjustments, and is likely to be statistically misspecified through omission of relevant dynamics. Where the data are integrated of order one in the time series dimension, we can obtain consistent (although possibly highly biased), but inefficient estimates of the long-run parameters from static regressions (see Banerjee *et al.* 1993). More information, and possibly improved estimator efficiency, should be attainable by estimating a dynamic model. A dynamic

EKC panel model has not been estimated before, though de Bruyn *et al.* (1998) estimate dynamic models for a number of countries with time series data.

We estimate an error correction model of the form:

$$\Delta\ln\left(\frac{M}{P}\right)_{it} = \alpha_i\left\{\ln\left(\frac{M}{P}\right) - \beta_{1,i}\ln\left(\frac{Y}{P}\right)_{it} - \beta_{2,i}\left[\ln\left(\frac{Y}{P}\right)\right]^2_{it}\right\}$$

$$+\sum_{j=1}^{p-1}\chi_{ij}\Delta\ln\left(\frac{M}{P}\right)_{it-j} + \sum_{j=0}^{q-1}\delta_{ij}\Delta\ln\left(\frac{Y}{P}\right)_{it-j} + \sum_{j=0}^{r-1}\gamma_{ij}\Delta\left[\ln\left(\frac{Y}{P}\right)\right]^2_{it-j}$$

$$+\mu_i + \eta_t + \varepsilon_{it} \tag{2}$$

Lag lengths p, q and r were selected separately for each country from a maximum of four using the Akaike information criterion. In addition to an efficiency gain, there are further advantages from estimating a dynamic (rather than static) model. First, the dynamic model not only yields information about long run relationships, but also estimates of short run dynamics, and the speed of adjustment to equilibrium. Second, as shown by Pesaran *et al.* (1996), the error-correction form of the autoregressive distributed lag model is robust for both I(0) and I(1) variables, provided that long-run relationships do exist among the variables. Furthermore, where a cointegrating relation does exist between ln (M/P), ln (Y/P) and ln $(Y/P)^2$ for each country, all terms in equation (2) are stationary variables. Hence, classical estimation and inference procedures can be used in the context of this specification, and its statistical adequacy can be assessed using conventional diagnostic statistics. Finally, and perhaps most importantly in the present context, the dynamic framework adopted here is a convenient general framework within which we can test some (usually untested) restrictions.

Equation (2) imposes no restrictions across countries, and so can be estimated efficiently on a country-by-country basis – this is essentially what de Bruyn *et al.* (1998) do.[6] In all but a few cases, test statistics (for absence of residual serial correlation and heteroscedasticity, normality and correct functional form) showed no evidence of equation misspecification.

Looking at the long-run relations in equation (2), we find results that are quite similar to those reported for the static EKC model. Specifically, only 47 out of the 74 countries had the 'correct' concave shape for their emissions–income relationship. We are also able to draw inferences about cointegration

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[6] Detailed regression results at the individual country level – including a full set of diagnostic statistics – can be found online at: http://homepages.strath.ac.uk/~hbs96107/ekc.htm

of the EKC relationships at the individual country level, using the estimated error correction coefficients, which also provide information about the speed of adjustment of emissions to income. For cointegration in country i we require that the i^{th} country's error correction coefficient, α_i, is statistically significant within the interval $\{-1 \leq \alpha_i < 0\}$. This approach is simple and effective in a dynamic single equation context. As the dependent variable and all other explanatory variables are stationary, the error correction coefficient can only be non-zero if the long-run relation is itself stationary. Pesaran *et al*. (2001) develop sets of critical values for testing the null of no long run relationship ($\alpha_i = 0$) against the alternative of a long run relationship $\{-1 \leq \alpha_i < 0\}$ for either I(0) or I(1) regressors.[7] Using this approach, we find that although the coefficient is correctly signed in every country except one (Yugoslavia), it is only statistically significant at 10 per cent or better in 16 out of the 74 cases. Moreover, among these 16 cases, four do not have a concave income–emission relationship.

These results reinforce the previous individual country static regression results. Most of the supposed EKC relations are spurious – either the variables have no true long-run relation or relevant variables have been omitted. In total, only 16 per cent of countries have a valid long-run relation among the variables which supports the EKC hypothesis. However, the validity of these results is possibly limited because cointegration tests have low power, especially in such small samples (31 time-series observations). However, it is possible to exploit the full pooled sample to test whether there is a single long-run relation common to all countries.

This turns out to be a relatively simple test to implement, even though we have strong reason to believe that the data are non-stationary. The belief that there is a single EKC relationship common to all countries is equivalent to maintaining that there is a unique cointegrating relationship common to all countries. Then, as shown in Pesaran *et al*. (1998), the autoregressive distributed lag panel model equation (2) is stable with roots of the autoregressive parameters lying outside the unit circle and classical hypothesis tests are applicable.

We proceed by stacking the unrestricted model in equation (2) over individual countries. We assume that the disturbances are independently distributed across time and groups, have zero means, positive country-specific variances, and are distributed independently of the regressors. Independence across time can be achieved by a suitable choice of distributed lag lengths. Inter-dependence across groups can be eliminated where slope coefficients

[7] From Pesaran *et al*. (2001) table C2, we find the critical value of the one sided test t-statistic at 95 per cent [90 per cent] confidence, to reject null of no cointegration, to be –3.21 [–3.53], on assumption of I(1) variables.

Table 5 Dynamic Error Correction Model: unrestricted and restricted estimates for full panel

	Unrestricted model (mean group parameter estimator)	Pooled mean group estimator (homogeneous long-run coefficients)	Static fixed effects
Long-run parameter estimates			
ln (Y/P)	16.56 (1.30)	6.272 (17.25)	3.85 (3.67)
in $(Y/P)^2$	−0.89 (−1.30)	−0.326 (−15.22)	−0.17 (−2.70)
Error correction	−0.366 (−12.7)	−0.24 (−8.30)	na
Implied turning point (US$)	10 975	15 063	82 746
lnL	1828.19	1506.08	−1805.23

na, not applicable. *t* ratios in parentheses. *t* ratios for fixed effects based on robust standard errors. Truncation of maximum allowed lags, here set to three, chosen by a general to specific search procedure, beginning from a starting value of six.

are identical across countries (or largely avoided where slopes vary) by expressing all variables as deviations from their respective cross-sectional means, a procedure we have used in our dynamic estimation. Therefore, the unrestricted estimates can be obtained by ordinary least squares. The homogeneity restrictions are imposed as cross-equation restrictions, and the restricted model estimated by maximum likelihood (ML). The validity of the restrictions is tested using a likelihood ratio test.

Table 5 presents the ML results for the whole sample of countries; tables 6 and 7 present results for OECD and non-OECD subsamples. The latter are given for comparability with Stern and Common (2001). The single point estimates in the left-hand column of statistics in tables 5,6 and 7 are what Pesaran *et al.* (1998) call mean group estimates. Each of these is the simple average of the individual country long run and error correction coefficient estimates from the model (2), in which no restrictions are imposed on regression parameters and error variances across countries.[8]

Also listed are the mean group estimates of the turning point. The mean group statistic is a consistent estimate of the mean of the individual parameters, which is a useful statistic when parameter estimates are randomly clustered around a common mean. The formula for the estimated variance of the mean group coefficient estimate is:

$$\hat{V}(\hat{\theta}_{MG}) = \frac{1}{N(N-1)}\sum_{i=1}^{N}(\hat{\theta}_i - \hat{\theta}_{MG})(\hat{\theta}_i - \hat{\theta}_{MG})' \tag{3}$$

[8] Individual country coefficient estimates are available on the additional results web page: http://homepages.strath.ac.uk/~hbs96107/ekc.htm

Table 6 Dynamic Error Correction Model: unrestricted and restricted estimates for OECD countries only

	Unrestricted model (mean group parameter estimator)	Pooled mean group estimator (homogeneous long-run coefficients)	Static fixed effects
Long-run parameter estimates			
ln (Y/P)	19.78 (2.11)	34.59 (13.31)	12.84 (4.73)
in $(Y/P)^2$	−1.02 (−1.91)	−1.85 (−12.65)	−0.71 (−4.62)
Error correction:	−0.300 (−5.57)	−0.163 (−3.16)	na
Implied turning point (US$)	16 254	11 483	8453
lnL	923.22	843.00	−8.21

na, not applicable. *t* ratios in parentheses. *t* ratios for fixed effects based on robust standard errors. Truncation of maximum allowed lags, here set to three, chosen by a general to specific search procedure, beginning from a starting value of six.

Table 7 Dynamic Error Correction Model: unrestricted and restricted estimates for non-OECD countries only

	Unrestricted model (mean group parameter estimator)	Pooled mean group estimator (homogeneous long-run coefficients)	Static fixed effects
Long-run parameter estimates			
ln (Y/P)	−1.56 (−0.23)	5.75 (11.91)	3.50 (2.73)
in $(Y/P)^2$	0.13 (0.31)	−0.28 (−9.95)	−0.15 (−1.89)
Error correction	−0.331 (−11.09)	−0.221 (−7.40)	na
Implied turning point (US$)	403 minimum	28 792	116 618
lnL	922.32	751.56	−1464.01

na, not applicable. *t* ratios in parentheses. *t* ratios for fixed effects based on robust standard errors. Truncation of maximum allowed lags, here set to three, chosen by a general to specific search procedure, beginning from a starting value of six.

where the $\hat{\theta}_i$'s are the individual country long run parameter estimates, and $\hat{\theta}_{MG}$ is the mean group estimator of the long run coefficient. This is a consistent estimator of the variance in question (Pesaran and Smith 1995).

The pooled mean group estimates in the central column are derived under the null that the long-run parameters are constant over the panel ($\beta_{1i} = \beta_1$ and $\beta_{2i} = \beta_2$) but permits dynamics, fixed effects and error variances to be heterogeneous over the panel. The restricted (pooled mean group) estimator imposes 146 restrictions on the unrestricted (mean group) model. The corresponding likelihood ratio test statistic is 644, with a probability value indistinguishable from zero. This decisively rejects the hypothesis that the long-run EKC parameters converge to a common cointegrating vector. Inspection of tables 6 and 7 shows that the hypothesis of a common cointegrating vector is also decisively rejected when the sample is restricted to either OECD or non-OECD countries alone.

The final column in tables 5–7 presents statistics from the familiar static fixed effects estimator, which is a heavily restricted special case of equation (2). The model is reparametrised as an autoregressive model in levels and the lag length set to zero. All parameters other than fixed effects are identical over countries and equality of variances is imposed. Dramatic deterioration of the log likelihood decisively rejects the fixed effects model (as it would also the random effects model).

The choice of estimated model has a huge impact on the average turning point. For the full sample of countries, the estimates from the unrestricted model, pooled mean group restricted model, and static fixed effects model are US$10 795, US$15 063 and US$82 746, respectively. As already noted, the unrestricted parameter estimates are very heterogeneous – the calculated mean turning point includes both emissions minima and maxima. For the non-OECD group of countries, the pooled mean group turning point estimate is US$28 792 versus US$116 619 for the fixed effects estimator. Both estimates are out of sample and imply a monotonic emissions–income relation. The turning point implied by the average of the unrestricted estimates is a minimum at US$403. This, too, implies an essentially monotonic relation. Only Tanzania and Myanmar had income lower than this, and even then only for a few years in the 1960s. The estimated turning point for the OECD actually declines as more restrictions are imposed, but, again, the unrestricted estimate is a mix of minima and maxima and so is not particularly meaningful.

The imposition of restrictions also affects the magnitude of the estimated speed of adjustment parameter. The restricted dynamic models (pooled mean group) have slower speeds of adjustment than do the unrestricted models.

7. Conclusions

Empirical work on the EKC using time series or panel data should consider data properties, because appropriate methods of inference depend in important ways on whether data are stationary or non-stationary. If there is no cointegration in a posited regression among non-stationary variables, interpreting the results in the classical way is invalid. Cointegration testing is a powerful test of misspecification.

In the present paper, we investigated the form of the relationship between sulfur emissions and income per capita over a panel of 74 countries. Individual and panel unit root tests suggest strongly that these series are integrated variables. This means that a static EKC regression could be a spurious regression. Even where the regression is not spurious, classical inference is not valid because t and F statistics for significance tests on individual and joint regression parameters have highly non-standard distributions. As income is an integrated variable, unit root and cointegration tests should be applied in EKC studies involving other indices of environmental pressure. If an environmental indicator is integrated, then an EKC regression is potentially spurious, while if it is not integrated it is unlikely to be related to income alone.

Although we were able to reject the non-cointegration null hypothesis for the panel as a whole, this is not sufficient to establish the validity of the EKC hypothesis for sulfur emissions. A large minority of countries have forms of emission–income relationship that are not consistent with the EKC hypothesis – either U shape or monotonically increasing in income.

However, even if we were to accept that there is a cointegrating relationship for each country in the panel we were able, using an error correction model, to reject the restriction that all countries have a common EKC cointegrating vector. This casts serious doubt on the many previous studies for which this is a maintained assumption.

In the present paper, the turning points for the full sample in the first two columns of table 5 seem to indicate a within sample turning point. However, the result in the first column is the mean of turning points where the coefficients were allowed to vary across countries. In some cases the turning point is a minimum, in some cases a maximum, and so seems relatively meaningless in the context of the EKC. Further, as the restriction imposing a single cointegrating vector was rejected the implied turning point in the middle column cannot be seen as valid either. Stern and Common (2001) showed that when the EKC is estimated in first differences, which eliminates the non-cointegration problem and reduces the problem of omitted variables bias, a monotonic relationship between emissions and income is found.

While the environmental Kuznets curve is clearly not a 'stylised fact' and is unlikely to be a useful model, this does not mean that it is not possible to reduce emissions of sulfur. The time effects from an EKC estimated in first differences (Stern and Common 2001) and from an econometric emissions decomposition model (Stern 2002) both show that considerable time related reductions in sulfur emissions have been achieved in countries at many different levels of income. Dasgupta *et al.* (2002) provide data and case studies that illustrate the progress already made in developing countries to reduce pollution. In addition, the income elasticity of emissions is likely to be less than one – but not negative in wealthy countries as suggested by the EKC hypothesis. In slower growing economies, emissions-reducing technological change can overcome the scale effect of rising income per capita on emissions. As a result, substantial reductions in sulfur emissions per capita have been observed in many OECD countries in the last few decades. The true form of the emissions–income relationship is likely to be a mix of two of the scenarios proposed by Dasgupta *et al.* (2002). The overall shape is the shape of their 'new toxics' EKC – a monotonic increase of emissions in income – while the curve shifts down over time in line with their 'revised EKC' scenario.

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