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The Environmental Kuznets Curve semi-parametrically revisited

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Abstract

This paper re-examines the existence of an Environmental Kuznets Curve across countries using a semi-parametric regression estimator, which places no restrictions on the functional form. In our specification against a linear model and using bootstrapped values, we are unable to reject a linear relationship.

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1. Introduction

Since the seminal paper by Grossman and Krueger (1991), there has been considerable interest in the relationship between economic growth and environmental pollution. Importantly, the authors showed that the link between these follows an inverted U-shaped pattern, now commonly referred to as the Environmental Kuznets Curve (EKC). This finding suggests that lower income regions are ‘too poor to be green,’ and only when these become rich enough will the benefits from a clean environment outweigh its costs. Much of the subsequent literature has focused on estimating the actual turning point and/or investigated whether the shape may depend on the econometric techniques and assumptions employed. One of the main concerns for the latter has been over the appropriate underlying functional form (see

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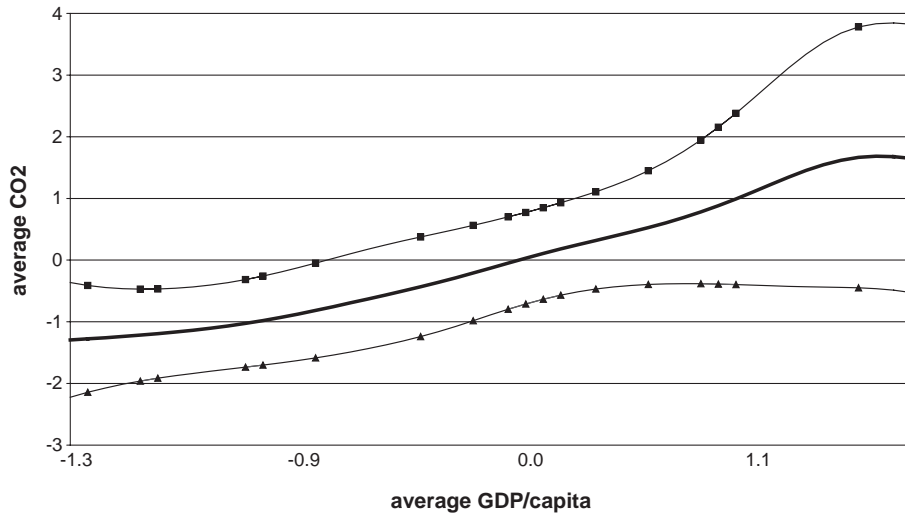


Fig. 1. Sulfur (average 1950–1990).

Dasgupta et al., 2002), where normally researchers allow for the possible non-linearity by introducing higher order terms. However, recently Millimet et al. (2003) have shown with data for US states that such parametric modeling can be rejected in favour of a semi-parametric estimator, which does not impose any a priori restriction on the functional form of the relationship.

In this paper, we employ such a semi-parametric estimator to investigate the existence of the EKC in a cross-country context. So far, results on cross-country studies measuring the relationship between economic growth and pollution have led to rather mixed results concerning the existence of an EKC (see, for instance, Shafik, 1994; Selden and Song, 1994; Grossman and Krueger, 1995; Holtz-Eakin and Selden, 1995; Stern and Common, 2001; Hettige et al., 2000; Harbaugh et al., 2002, among others for recent evidence). However, it must be noted that all of these have imposed relatively restrictive functional forms. Our findings using the semi-parametric estimator suggest that in a cross-country sense, at least in terms of measuring pollution by sulfur and carbon dioxide, the link between environmental pollution and economic growth is actually monotonically increasing for low levels of GDP/capita, and flat thereafter.¹

2. Methodology and data

Most studies examining the EKC have been concerned with estimating the following equation:

$$P_{it} = \alpha + g(Y_{it}) + Z_{it}\delta + v_i + u_{it} \quad (1)$$

where P_{it} is some proxy of environmental degradation (per capita) in country i at time t , Y_{it} is a measure of wealth, usually real per capita GDP at the start of the period, Z_{it} is a vector of variables that controls for other factors, v_i is a unit-specific residual and u_{it} is a disturbance term.

¹ This result echoes the recent skepticism raised by Stern (2004) over the existence of an EKC internationally in his review of the literature.

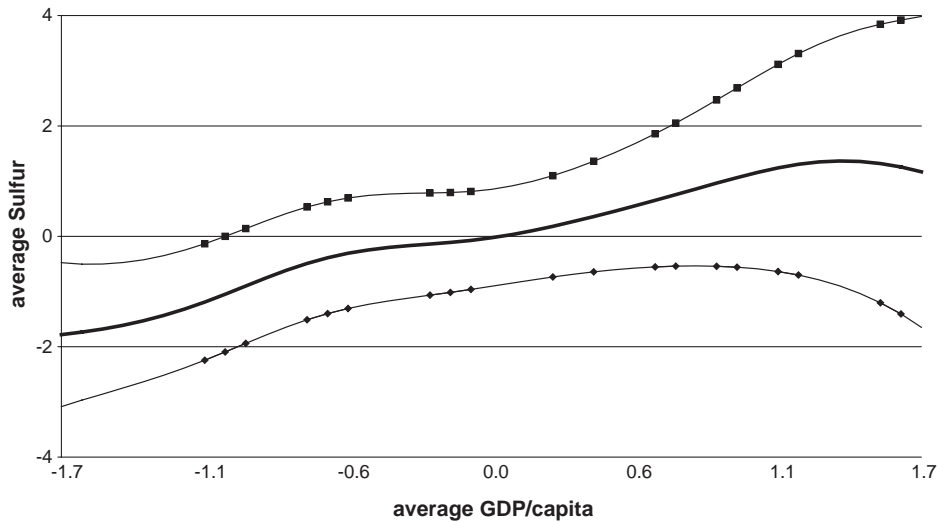


Fig. 2. Carbon dioxide (average 1950–1990).

In order to allow for the possible non-linearity of $g(Y)$, most analyses have simply included a second- and third-order polynomial of Y . As in Millimet et al. (2003), we instead implement Robinson's (1988) semi-parametric Kernel regression estimator (see Blundell and Duncan, 1998, for details and a helpful discussion of the implementation of this method). Accordingly, if we allow $g()$ in (1) to be a smooth and continuous, possibly non-linear, function of Y and assume that the other control variables captured by the vector Z have a linear effect on P , then the estimation of $g(Y)$ can be made by

$$\hat{g}(Y) = \hat{m}_P(Y) - \hat{\delta} \hat{m}_Z(Y) \quad (2)$$

where $\hat{m}_P(Y)$ and $\hat{m}_Z(Y)$ are the (non-parametric) Nadaraya-Watson estimates (Nadarya, 1964; Watson, 1964) of $E(P/Y)$ and $E(Z/Y)$, such that, for a given continuous, bounded, and real shape function, $K_h()$ integrating to one with a smoothing parameter h , $\hat{m}_P(Y)$ (and similarly $\hat{m}_Z(Y)$) is defined as

$$\hat{m}_P(Y) = n^{-1} \frac{\sum_{i=1}^n K_h(y - Y_i) P_i}{\sum_{i=1}^n K_h(y - Y_i)} \quad (3)$$

and $\hat{\delta}$ is the OLS estimator of

$$P - \hat{m}_P(Y) = \delta(Z - \hat{m}_Z(Y)) + \varepsilon \quad (4)$$

The appeal of the estimator (2) lies in its very flexible approach to non-linearity by allowing the relationship between P and Y to vary over all values of Y after purging the effects of other explanatory variables. Specifically, this technique entails first purging the effect of the other factors Z from the relationship between P and Y and then estimating the regression function of P on Y at a particular point by locally fitting constants to the data via weighted least squares, where those observations closer to the chosen point have more influence on the regression estimate than those further away, as determined by the choice of h .²

² For all estimations, we use a Gaussian kernel for K_h and the optimal smoothing parameter suggested by Fox (1990).

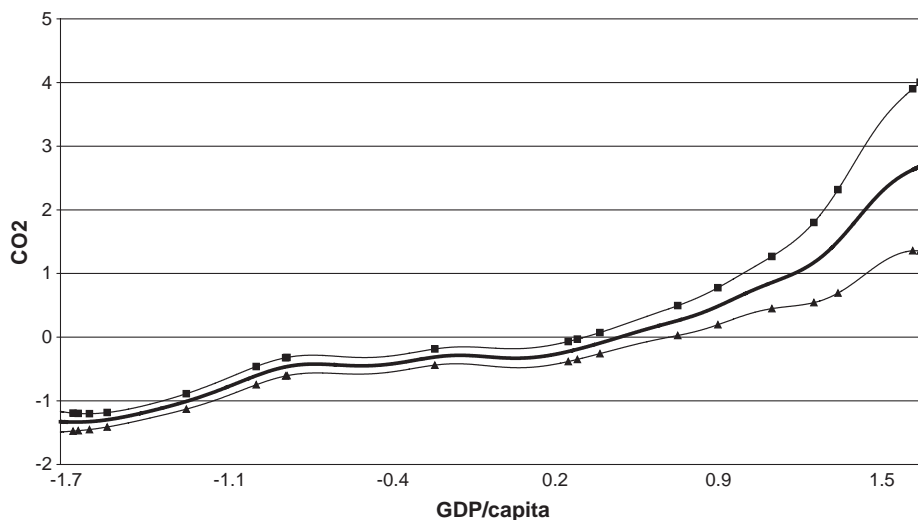


Fig. 3. Sulfur (1950–1990).

One should note that, given its semi-parametric nature, the estimate of $\hat{g}(Y)$ cannot be subjected to the kind of standard statistical tests (such as an F -test or a t -test) of parametric regressions. However, it is possible to calculate upper and lower point-wise confidence intervals, as suggested by Haerdle (1990). Specifically, we calculated bands at the 1st and 99th percentiles along the range of initial income and at every fifth percentile in between. Choosing points according to the distribution of observations also allows one to gauge how the density of the sample affects the approximation bias, since these are inversely related. It should also be noted that the use of the semi-parametric estimator just described leads to a problem of non-identification of an unrestricted intercept term, which leads to a scaling issue when

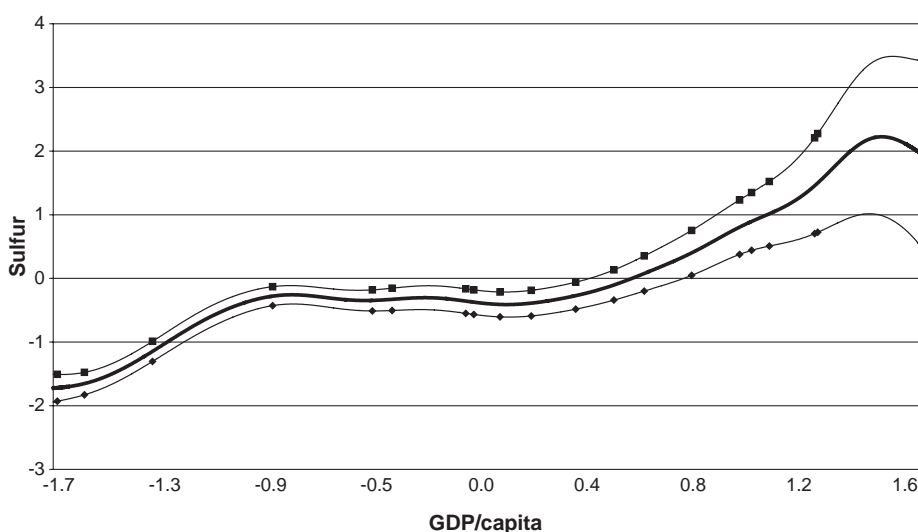


Fig. 4. Carbon dioxide (1950–1990).

Table 1
Im, Pesaran and Shin (IPS, 2003) test for panel unit roots

	Level	First difference
	(1)	(2)
GDP/capita	– 1.202 (0.989)	– 5.266 (0.000)
Sulfur	– 1.026 (1.000)	– 5.196 (0.000)
Carbon dioxide	0.515 (1.000)	– 4.207 (0.000)

(1) Reported values are the t -bar statistics (Im, Pesaran and Shin (IPS, 2003)); (2) p -values in parentheses.

comparing our semiparametric results with any parametric alternative. As in Millimet et al. (2003), we deal with this issue by standardising our data (relative to the full sample).

The data used for our analysis consists of annual information on sulfur emissions taken from the Historical Global Sulfur Emissions database (Lefohn et al., 1999), carbon dioxide emissions taken from World Resource Institute (People and Ecosystems CD-ROM), and GDP per capita figures (real GDP per capita in constant dollars, base year 1985) taken from the World Penn Tables 6.1. As other explanatory variables, Z , we included time- and country-specific dummies.³ Together, this provides us with a total sample size of 3976 observations consisting of 122 countries (of which 95 are LDCs), and 3336 observations for 108 countries (of which 81 are LDCs) for the carbon dioxide sample and the sulfur sample respectively, over the period 1950–1990.⁴

3. Results

In Figs. 1 and 2, we depict the graph for $\hat{g}(\bar{Y})$ and corresponding confidence bands for sulfur and carbon dioxide, respectively, where \bar{Y} is the across-period average of GDP/capita, i.e., we have estimated specification (1) by averaging the two variables of interest (pollution and GDP/capita) over the 40-year period under scrutiny. Accordingly, there is little evidence of a bell-shaped link between either sulfur or carbon dioxide emissions and GDP/capita. Rather, in contrast to an ‘EKC,’ our estimate appears to be decidedly linear, environmental pollution increasing with country wealth for low levels of GDP/capita, and becoming flat thereafter.

Averaging pollution and GDP/capita over the whole period, however, reduces considerably the information contained in inter-annual changes. We thus also show in Figs. 3 and 4 plots of $\hat{g}(\bar{Y})$ for annual sulfur and carbon dioxide emissions. As for our averaged data, there is little support for a bell-shaped link between pollution and GDP/capita., but instead, a linear relationship emerges.⁵ The only exception of linearity is for very high GDP per capita values, but as the distance between the confidence intervals suggests, this portion of both curves is likely to be very poorly estimated because of the lower number of observations around it and the fact that it is near the endpoint.⁶

³ We also experimented with other controls, such as openness and population growth rate. However, these made little qualitatively or quantitative difference and only reduced sample size.

⁴ A list of all the countries can be found in Bertinelli and Strobl (2004).

⁵ One should note that our panel is unbalanced. However, 80% of the countries had more than three quarters of the observations in the carbon dioxide sample (70% for the sulfur sample). Missing data are essentially concentrated in former colonial countries for years before 1960.

⁶ The accuracy of the estimate of $g(Y)$ at Y is positively related to the density of other observations around that point. Furthermore, the approximation bias is larger at the boundaries (see Wand and Jones, 1995).

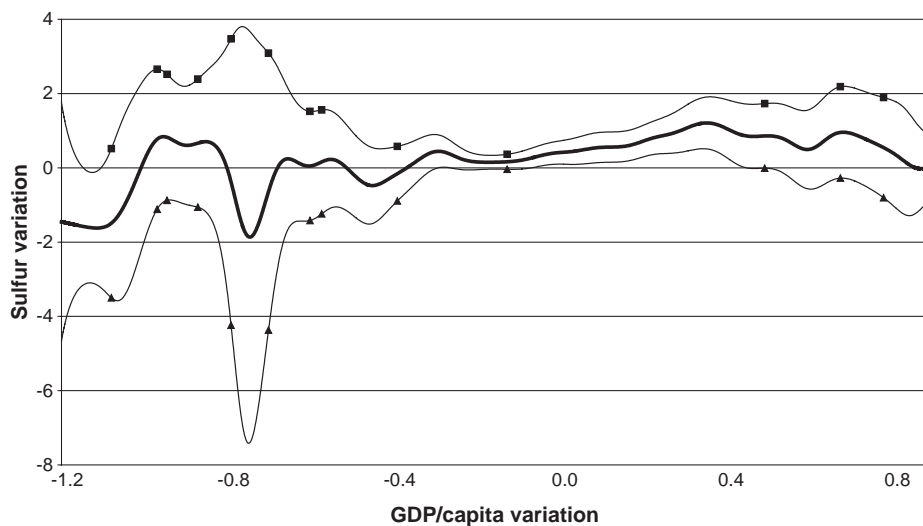


Fig. 5. Sulfur (first-differenced series; 1950–1990).

One issue to be considered with regard to using annual data is the potential time series component in the data. In particular, pollutant emissions and GDP/capita are unlikely to be stationary over time and this is confirmed by the Im et al.'s (2003) t -test statistics for unit roots in heterogeneous panels shown in Table 1. We thus first-differenced our series so that, as confirmed by the test-statistics in the same table, these are difference stationary, although not surprisingly, more volatile plots of the semi-parametric estimates of these, depicted in Figs. 5 and 6, are similarly roughly linear.

Our estimates for both sulfur and carbon dioxide emissions suggest that the relationship between pollution and wealth is in fact linear. To test this more rigorously, we employ the test suggested by

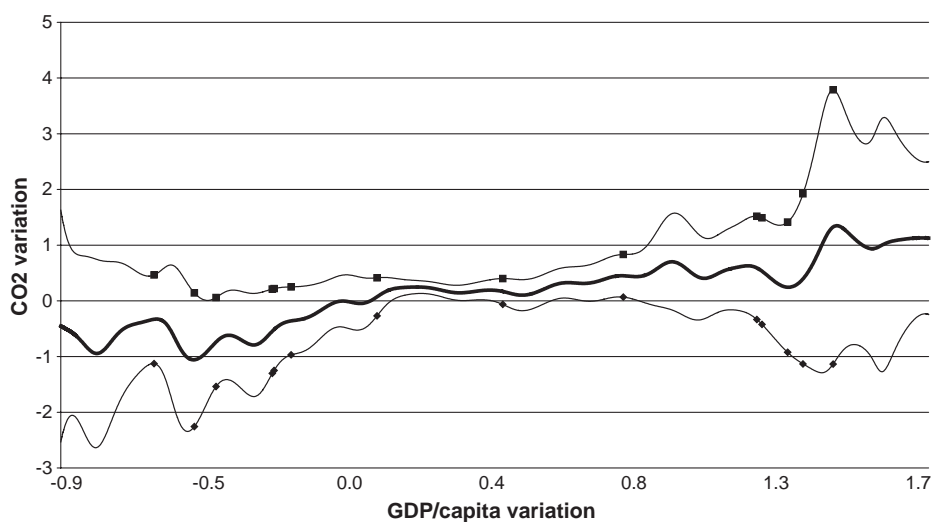


Fig. 6. Carbon dioxide (first-differenced series; 1950–1990).

Table 2
Tests against linearity

	Sulfur			Carbon dioxide		
	Average	Level	First difference	Average	Level	First difference
<i>T</i>	0.918	0.911	0.339	0.505	0.989	0.619
<i>p</i> -values	0.243	1.00	0.66	0.312	1.00	0.58

Ullah (1985), which allows one to test our semi-parametric estimate (SP) against some linear parametric estimate (*P*) of a regression curve by using their residual sum of squares RSS^{SP} and RSS^P , respectively:

$$T = \frac{(RSS^P - RSS^{SP})}{RSS^{SP}} \quad (5)$$

The test is based on the null hypothesis that there is a relationship of interest that is linear. Given that \sqrt{nT} has an analytically intractable distribution under the null hypothesis, we implement the bootstrap procedure suggested by Lee and Ullah (2001) to obtain *p*-values using 100 replications.⁷ The resultant test statistics for both pollutants in cross-country average, levels and first-difference, as well as bootstrapped *p*-values, are reported in Table 2. As can be seen, results suggest that we are in all cases unable to reject the linearity of the relationship between wealth and pollution.⁸

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⁷ This admittedly small number of replications was chosen in consideration of the considerable amount of computational time for each procedure.

⁸ The 5% and 10% critical values are 1.64 and 1.28, respectively; see Ait-Sahalia et al. (2001).

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