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TESTING FOR AN ENVIRONMENTAL KUZNETS CURVE IN LATIN-AMERICAN COUNTRIES

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Abstract

This paper presents empirical estimates of Environmental Kuznets Curves for a panel of Latin-American countries over the period 1975-1998. It uses a new econometric technique that allows for more flexible assumptions in a panel data framework with a large time dimension. Unlike most previous studies we test for slope heterogeneity of the income coefficient in the search of a common empirical relation between carbon dioxide emissions and income. Our results point to the existence of some heterogeneity among countries, but with specific patterns for those sharing certain characteristics.

I. Introduction

According to some researchers, there is an inverse U-shaped pattern between pollution and economic growth (see the survey presented by Stern, 1996). This regularity implies that pollution increases with income until a “turning point” in which pollution begins to decrease while income is still rising. Because of its similarity to the relationship between income inequality and the level of income (Kuznets, 1955), the inverted U-shaped curve that relates pollution with income is called the Environmental Kuznets Curve (EKC).

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Since the first EKC study, presented in 1991 by Grossman and Krueger, the research in this field has grown very fast but empirical evidence does not support the EKC hypothesis in a general way. Results are strongly dependent on the pollutant indicators, the functional form, the econometric procedure and the explanatory variables included in the regressions, the time period considered and the countries included in the sample.

Over the last two decades there have also been extensive theoretical research on the EKC hypothesis. Technological progress, structural changes, enforcement of environmental regulations are some of the explanatory factors researchers have pointed out to compensate the negative impact that a larger scale of economic activity produces on the environment. Bulte and van Soest (2001) show an alternative explanation to EKC relative to environmental degradation in developing countries based on imperfect markets for factors and commodities[1].

Research on the EKC hypothesis is far from being an academic entertainment since the existence or absence of such a curve has important policy implications. If the EKC were a generalized phenomenon, environmental degradation will automatically fall in the long run as income becomes sufficiently high. However, if the proposition does not hold, public intervention would be necessary to curb pollution and make sustainable development a reality[2].

The Earth Summit held in Rio de Janeiro in 1992 and the 1997 Kyoto Summit have called the international attention upon environment, particularly upon the heating of the planet as a consequence of the greenhouse effect.

Since the main gas producing the greenhouse effect is carbon dioxide, we focus on this pollutant to investigate the existence of an EKC in Latin American countries. Previous research dealing with this topic is quite scarce (see Table 1). Studies concerning developing countries have mainly focused on sulphur emissions (de Bruyn *et al*., 1997; Panayotou, 1997; Kaufman *et al*., 1998) and also on deforestation (Koop and Tole, 1999; Bhattarai and Hammig, 2001; Halkos and Tsionas, 2001[3]). However, there is an important debate about the role that should be played by developing countries in curbing CO_2 emissions. The Kyoto Protocol contains a specific commitment taken by industrialised and transition economies (the so-called Annex B countries)[4] to reduce their emissions over the period 2008-2012 down to the level attained in 1990, but no commitment exists for developing countries.

In Latin America and the Caribbean total CO_2 emissions generated by the energy sector have been steadily raising since 1970 as shown by the CEPAL (2002). When emissions per unit of GDP are considered, a raising trend is observed, but when emission intensity and per capita income are considered, the path is not clear (pages 290 and 291, *op. cit*.). Anderson and Cavendish (2001) present a dynamic simulation model to develop scenarios for SO_2, and CO_2 abatement in developing countries. Their results show that, without a pollution abatement policy, in a developing country with an initial per capita income of $ 2,500 and growth rate of 4%, emissions will rise to over five times today levels in the present century.

The main aim of this paper is to investigate the relationship between economic growth and CO_2 emissions in Latin American countries in order to compare the observed patterns with those followed by Annex B countries.

We use a fresh methodology, the pooled mean group estimator, based on Pesaran *et al*. (1999) that allows for slope heterogeneity in the short run imposing restrictions only in the long run and testing for their validity. To our knowledge, this methodology has only been applied to EKCs for sulphur emissions in Perman and Stern (1999) and to income growth equations in Bassanini and Scarpeta (2001). There are 19 countries[5] in the sample (only Guatemala is excluded due to missing data), including some Caribbean countries. We consider the evolution of income and CO_2 emissions from 1975 to 1998.

The paper is structured as follows. Section II summarises previous research in CO_2 EKCs. Section III presents the econometric approach, Section IV shows the empirical findings and, finally, Section V concludes.

II. Exploring the Evidence

Since the early 90's a number of authors have estimated EKCs for various indicators of environmental degradation. In this section we briefly review, in a chronological order, the studies which are more relevant for our analysis. Table 1 summarises some results of previous EKC studies on CO_2 emissions conducted along the last decade.

Shafik and Bandyopadhyay (1992) estimated EKCs for ten different environmental indicators from 1961 until 1986. They found that sulphur oxides conform to the EKC hypothesis with a turning point at $ 5,000 per capita income, but they did not find evidence for carbon emissions per capita which increase unambiguously with rising income.

Holtz-Eakin and Selden (1995) formulated a quadratic EKC function for carbon dioxide emissions, with an estimated turning point at $ 35,428 per capita income, and a log-quadratic function, which showed a very high turning point ($ 8 million). Although they concluded that there is a diminishing marginal propensity to emit CO_2 as economies develop. They predicted that emissions growth will continue because output and population will grow more rapidly in lower-income nations that have the highest marginal propensity to emit. Agras (1995) found a turning point for SO_2 at $ 6.654 per capita income but for CO_2 he questions the existence of an EKC. However, Sengupta (1996) found evidence for this gas as well as Schmalensee *et al.* (1998).

Moomaw and Unruh (1997) analysed per capita CO_2 emissions in a set of countries over the period 1950 to 1992. They found a great heterogeneity among them. OECD member states showed a discontinuous transition in which the CO_2/GDP relation changed from a strong positive covariance to a negative or weakly correlated relation. Another subset of countries, dominated by centrally planned economies and some developing countries, showed a positive correlation. Finally,

TABLE 1

CO_2 EKC STUDIES IN A CHRONOLOGICAL ORDER

Autors	Turning Points	PPP	Additional Variables	Data source for CO_2	Time period	Estimation technique	Functional form	EKC	Countries
Shafik and Bandyopadhyay (1992)	$7Million	Yes	Yes (Market premium, dollar index)	Marland (1989)	1961-86	Fixed Effects, Random Effects	Linear, Quadratic and Cubic (logs)	No	118-153
Holtz-Eakin and Selden (1995)	$35,428(level) $8 Mill. (logs)	Yes ($1,986)	No	ORNL[b]	1951-86	Two ways Fixed Effects	Quadratic (levels and logs)	Yes	108
Tucker (1995)	Decreasing over time		No	WRI (1994)	1971-91	Yearly Cross-sectional analysis. First Differ.	Quadratic	In 11 years	137
Sengupta (1996)	$8,740	Yes ($1,985)	No	ORNL[b]		Fixed Effects	Quadratic	Yes	16 Developed + Developing
Cole, Rayner and Bates (1997)	$25,100(levels) $62,700 (logs)	No	Yes (Trade, pop.d., tech)	Marland *et al*. (1994)	1960-92	Generalized Least Squares	Linear, Quadratic (levels and logs)	Yes	7 World Regions
Moomaw and Unruh (1997)	$12,813	($1,985)	No	World Bank (1992)	1950-92	Fixed Effects	Structural Transition Model, Cubic form	N shaped	16 Developed
Roberts and Grimes (1997)	$8,000-$10,000	Yes	No	ORNL[b]	1962-91	Cross-section analysis	Quadratic	Yes, after the 70s	Developed + Developing
Schmalensee, Stoker and Judson (1998)	Within sample	Yes ($1,985)	No	ORNL[b]	1950-1990	Two ways Fixed Effects	Spline Function	Yes	141
Agras and Chapman (1999)	$13,630	Yes ($1,997)	Yes (Price, trade var.)	IEA[a] and ORNL[b]	1971-89	Autorregressive-Distributed Lag with Fixed Effects	Quadratic (logs)	No	34
Galeotti and Lanza (1999)	$15,073-$21,757	Yes ($1,990)	No	IEA[a]	1971-96	Least Squares Dummy Variable	Non linear Gamma and Weibull	Yes	110
Panayotou, Peterson and Sachs (2000)	$29,732-$40,906 (1950-1990)	Yes	Yes (Trade, K, pop. d.)	CDIAC [c] (1997)	1870-1994	Feasible Generalized Least Squares	Quadratic	Yes for Developed	17 Developed
Heerink *et al*. (2001)	$68,871	Yes	Yes (Inequality)	Marland (1989)	1985	Generalised Method of Moments	Quadratic (logs)	Yes	118-153
Roca *et al*. (2001)	Y^2 non sign.	No ($1,986)	Yes (Energy prices)	IEA[a]	1973-96	Time series, cointegration	Linear, Squared and Cubic (logs)	No	Spain
Baiocchi and di Falco (2001)	Y^2 non sign.	Yes	No	World Resources Institute		Nonparametric method	Local polynomial	No	160
Bengochea *et al*. (2001)	$24,427-$73,170	Yes ($1,993)	No	OECD Environmental Data	1980-95	Fixed Effects, Random Effects, Instrumental Variables, First Differ.	Linear, quadratic, cubic	For some countries	EU
Dijkgraaf and Vollebergh (2001)	$20,647	No ($1,990)	No	OECD 2000 IEA (1991)[a]	1960-97	Fixed Effects, Seemly Unrelated Regression	Linear, quadratic, cubic. Slope heterogeneity	Yes 5 rich countries	24-OECD
Martínez-Zarzoso *et al*. (2002)	$4,914-$18,364	Yes ($1,993)	No	World Development Indicators 2001	1975-98	Pooled Mean Group	Linear, quadratic, cubic. Slope heterogeneity	N shaped	22-OECD

Notes: a: International Energy Agency: Greenhouse Gas Emissions: The Energy Dimension (Paris, OECD, 1991), b: Oak Ridge National Laboratory, c: Carbon Dioxide Information Analysis Centre, e: World Resources Institute.

a third group described chaotic behaviour between CO_2 emissions and income. Thus, the authors conclude that neither the quadratic nor the cubic functional formulation of the EKC hypothesis can provide a reliable indication of future behaviour. Moreover, even in cases where transition is observed (an EKC exists), decreasing of CO_2 emissions does not appear to correlate with specific income levels but with specific points in time in response to exogenous shocks influencing these economies as the petroleum crisis in 70s. In the same line, Roberts and Grimes (1997), in their work about carbon dioxide emissions intensity, argued that the inverted U curve reached statistical significance only after 1970. However, they consider that this fact is not the result of countries passing through stages of development, but of efficiency improvements in a small number of wealthy nations combined with worse performance in poor and middle-income countries.

De Bruyn *et al.* (1998) considered three pollutants (CO_2, NO_x, and SO_2) in four countries (the Netherlands, the UK, the USA and the former West Germany). The authors found that emissions were positively correlated with income but it is possible to abate them because of technological progress and structural change. Their conclusion is that emissions behaviour corresponding to EKC hypothesis could be found, but estimations from panel data does not seem to be able to capture all the dynamic processes involved, thus one can not accept as a general rule that economic growth improves environmental quality.

Agras and Chapman (1999) test two EKCs using both per capita energy consumption and per capita CO_2 emissions to represent environmental degradation in 34 countries during the period 1971-1989. They observe that energy per capita consumption decreases as gasoline prices increase, so they do not find significant evidence for the existence of an EKC within the range of current incomes for energy in the presence of price and trade variables. Despite the fact that they report a \$ 13,630 turning point for CO_2 emissions, they conclude that, since energy use at all income levels is price elastic, rising levels of GDP in the long run can increase energy use and, therefore, Governments need to undertake policies now to start to reduce levels of pollution and CO_2 emissions.

Galeotti and Lanza (1999) considered 110 countries, 30 belonging to Annex B and 80 coming from non Annex B. Despite their analysis conform the EKC hypothesis, they forecast that future global emission between 2000 and 2020 will rise as a result of the faster growth rate exhibited by developing countries. Panayotou *et al*. (2000) also found a CO_2 EKC for 17 developed countries.

Heerink *et al*. (2001) analyse the relationship between environmental degradation, income and income inequality in a set of countries. They use several environmental indicators and the EKC hypothesis holds for some of them. As far as CO_2 emissions are concerned, the relationship with income is positive and non linear and income inequality is found to have a significant negative impact on the level of emissions.

In a recent paper, Roca *et al*. (2001) analyse trends of annual emission flux of six atmospheric pollutants in Spain. Only SO_2 emissions can support the EKC hypothesis. Looking at the evolution of per capita CO_2 emissions, their work confirms the Moomaw and Unruh's theory. They find three stages corresponding to the adjustment to the new situation of sharp increases in energy prices: strong emissions growth until the end of the seventies, a subsequent relative emissions stabilization during the 1980s and, finally, an increase in emission in the last decade. Baiocchi and di Falco (2001) neither find an EKC for carbon dioxide.

Bengochea *et al*. (2001) find evidence only for a few countries in the European Union while for a 22-Annex B sample. On the other hand, Martínez-Zarzoso *et al.* (2002) conclude that a cubic specification conforms better with the data.

Dijkgraaf and Vollebergh (2001) use panel data methodology to analyse CO_2 emissions in 24 OCDE countries during the period 1960-1997. The EKC hypothesis is only confirmed for eleven countries. The authors reject model specifications that feature homogeneity assumptions across countries. They calculate turning points with panel and time-series estimates. The unweighted mean turning point in the panel general model is \$ 20,647. When time-series estimates are used, turning points vary among countries from \$ 12,505 to \$ 31,407.

III. The Econometric Approach

We estimate a standard EKC equation on the basis of annual data using pooled cross-country time series. We assume that the long-run EKC function is

$$\begin{aligned}&\ln co_{it} = \alpha_{0i} + \alpha_{1i} \ln yh_{it} + \alpha_{2i}\left(\ln yh_{it}\right)^2 + \alpha_{3i}\left(\ln yh_{it}\right)^3 + \beta_{1i} t + \mu_{it}\\ &i = 1,\ 2,\ \ldots N, \qquad t = 1,\ 2,\ \ldots,\ T.\end{aligned} \tag{1}$$

where co_{it} = CO_2 emissions per capita, yh_{it} = GDP per capita in dollars at 1993 PPP, α_{0i} is a country-specific intercept, t is a time trend and μ_{it} is an error term. The two additional terms are the natural log of GDP per capita squared and cubed. The cubed term has only been included in a few of the models tested in the literature for CO_2 emissions (Shafik and Bandyopadhyay, 1992; Moomaw and Unruh, 1997; Roca *et al*, 2001; Bengochea *et al*., 2001; Dijkgraaf and Vollebergh, 2001). Equation (1) will be tested with and without this term for comparative purposes.

We will assume that all these variables are I(1) and cointegrated for individual countries, making the error term an I(0) process for all *i*. These assumptions are based on the test results shown in Appendix 3. According to the unit root test results (Table A.1), the null of unit root cannot be rejected for the variables in levels for all the countries, whereas the first differences of the variables are stationary (the null of unit root is rejected). Therefore, the univariate test statistics strongly support the view that income and emissions are I(1) processes.

Concerning cointegration (Table A.2), we find clear evidence in favour of a statistically significant relationship for most countries.

Taking the maximum lag equal to one, the ARDL(1,1,1,1) equation is given by

$$\begin{aligned}\ln co_{it} &= \gamma_i + \delta_{10i} \ln yh_{it} + \delta_{11i} \ln yh_{i,t-1} + \delta_{20i}\left(\ln yh_{it}\right)^2 + \delta_{21i}\left(\ln yh_{i,t-1}\right)^2 \\ &+ \delta_{30i}\left(\ln yh_{it}\right)^3 + \delta_{31i}\left(\ln yh_{i,t-1}\right)^3 + \beta_{10}t + \lambda_i \ln co_{i,t-1} + \eta_{it}\end{aligned} \tag{2}$$

The error correction equation is

$$\begin{aligned}\Delta \ln co_{it} &= \phi_i\left[co_{i,t-1} - \alpha_{0i} - \alpha_{1i} \ln yh_{it} - \alpha_{2i}\left(\ln yh_{it}\right)^2 - \alpha_{3i}\left(\ln yh_{it}\right)^3\right] \\ &- \delta_{11i}\Delta \ln yh_{i,t-1} - \delta_{21i}\Delta\left(\ln yh_{i,t-1}\right)^2 - \delta_{31i}\Delta\left(\ln yh_{i,t-1}\right)^3 + \eta_{it}\end{aligned} \tag{3}$$

were:

$$\alpha_{0i} = \frac{\gamma_i}{1-\lambda_i}, \alpha_{1i} = \frac{\delta_{10i} + \delta_{11i}}{1-\lambda_i}, \alpha_{2i} = \frac{\delta_{20i} + \delta_{21i}}{1-\lambda_i}, \alpha_{3i} = \frac{\alpha_{30i} + \delta_{31i}}{1-\lambda_i},$$
$$\phi_i = -(1-\lambda_i)$$

The empirical analysis of the EKC model specified above involves a system of N*T equations that can be examined in several ways. The approach chosen depends on part on the size of N and T and the quality of data across these two dimensions. The main econometric approaches used in the literature have been based on cross-section regressions and different forms of pooled cross-section time-series regressions.

Authors who have focused their research on a small number of countries, such as those in the OECD area, have often exploited the time dimension of the data (Hilton and Levinson, 1998; Roca *et al*., 2001) or have used pooled cross-country and time-series data (Suri and Chapman, 1998). There are also some examples of EKC regressions based on cross-section data (Tucker, 1995; Roberts and Grimes, 1997).

Most recently, researchers have used techniques based on panel data methodology as in Cole *et al.* (1997), Stern *et al*. (1998), de Bruyn *et al*. (1998) and Dijkgraaf and Vollebergh (2001). The main advantage of these aforementioned techniques for the analysis of EKC equations is that the country-specific effects can be controlled for by using static fixed-effect (SFE) or dynamic fixed-effect estimators (DFE). The SFE or DFE estimators generally impose homogeneity of all slope coefficients, allowing only the intercepts to vary across countries. DFE imposes (N–1)(2k + 2) restrictions on the unrestricted model in equation (3): k

long-run coefficients, *k* short-run coefficients and the convergence coefficient and the common variance. So, the validity of DFE, in particular, depends critically on the assumptions of common estimated parameters, that in turn requires both common income elasticity and common EKC patterns across countries. Since the evolution of CO_2 emissions and income differ across countries, these assumptions are difficult to reconcile with observed emissions patterns across countries. Then, as Pesaran and Smith (1995) pointed out, under slope heterogeneity, estimated coefficients will be affected by an heterogeneity bias.

At the other extreme of the SFE and DFE estimators we find the mean-group approach (MG) that consists of estimating separate regressions for each country and calculating averages of the country-specific coefficients. There are $N(2k + 3)$ parameters to be estimated. Each equation has $2k$ coefficients on the exogenous regressors, an intercept, a coefficient on the lagged dependent variable and a variance. The small-sample downward bias in the coefficient of the lagged dependent variable remains. Furthermore, although this estimator is still consistent, it is likely to be inefficient in small country samples, where any country outlier could badly influence the averages of the country coefficients.

The pooled mean group estimator (PMG) involves both pooling and averaging. It is an intermediate estimator between the DFE and the MG that allows short-run coefficients, the speed of adjustment and error variances to differ across countries but imposes common long-run coefficients. The PMG estimator is specially suited for panels with large T and N. Pesaran *et al.* (1999) show that for T and N greater than 20 the PMG estimator is clearly superior to estimators such as the SFE, DFE and the MG. However, they also use the PMG to estimate energy demand functions for smaller sample sizes: $N = 10$ and $T = 17$, and also in this case the PMG estimator performs better than alternative estimators. The advantages of this method are that it does not impose homogeneity of slopes in the short-run and it allows for dynamics. Thus, we have choose this procedure to estimate the relationship between carbon dioxide emissions and economic growth. Next section summarises the results obtained.

IV. Empirical Results

The 19 countries under study do not exhibit a single behaviour. On the contrary, a great heterogeneity is observed in the scatter diagrams shown in Appendix 2 relating emissions and income, both in logarithms.

Equation (1)-(3) in various forms has been estimated for the sample of Latin-American and Caribbean countries over the period 1975-1998. Several specifications have been tested allowing for a linear, quadratic and cubic form for the income-emissions relationship, each of them with and without a time trend.

First, a common ARDL (1,1,1) was run for all countries. The best specification in terms of diagnostic test was the quadratic form without time trend[6]. Table 2 presents results for an EKC for three alternative pooled estimates MG, PMG and DFE. Results for a linear and cubic specification are available upon request.

TABLE 2

QUADRATIC SPECIFICATION. SELECTION OF THE ESTIMATION METHOD
ONLY *Ly* RESTRICTED IN THE LONG-RUN

Dep. variable: lco	Without time trend. One lag (1,1,1)			
	Mean Group	Pooled Mean Group	Hausman test	Dynamic Fixed Effects[1]
Convergence coefficient	– 0.38**	– 0.28**		– 0.23**
Long run coefficients				
Ly	1.58**	1.48**	0.00	– 2.78**
Ly^2	0.005	0.003		10.24**
Short run coefficients				
Δly	4.23**	1.08		– 2.88*
Δly^2	– 0.21**	0.007		0.24*
Nº of countries	19	19		19
Nº of obs.	380	380		380
Log likelihood	478.46	436.64		– 130.64

Note: [1] t-stat. Calculated using heteroskedasticity consistent standard errors.

Results vary significantly with respect to the estimation method, from MG (the least restrictive, but potentially not efficient) to PMG and to DFE that only allows intercepts to vary across countries. Moving from MG to PMG (imposing only long-run homogeneity to the income variable (*lnyh*) reduces the standard errors and the speed of convergence and reduces the size of the estimated long run parameters. The Hausman test indicates that this restriction (equality of slopes for the income coefficients) is rejected at 1% significance level in both specifications (with and without a linear trend). We also tested for homogeneity in the speed of convergence and short-term dynamics. Moving from PMG to DFE estimators significantly reduces the speed of convergence due to a downward bias in dynamic heterogeneous panel data. Furthermore, the sign and significance of the long-run coefficients change in both specifications. We performed a sensitivity analysis of the PMG results to changes in the lag structure of the dependent and independent variables. Table 3 presents the estimated coefficients when the Schwarz Criterion (SBC) has been used to select the ARDL specifications for each country. The estimated coefficients are very different with respect to the ARDL(1,1,1) specification (Table 2). The estimated speed of convergence is higher for the PMG because the SBC criterion chooses the static model for some countries (with instantaneous adjustment). The maximum likelihood increases for MG estimates and also for PMG when the SBC is applied. Pesaran *et al.* (1999) argued that,

TABLE 3

QUADRATIC SPECIFICATION. SELECTION OF THE ESTIMATION METHOD. ONLY *Ly* RESTRICTED IN THE LONG-RUN
SBC criterion used to chose the lag order

Dep. variable: lco	Without time trend. SCQ Crit			
	Mean Group	Pooled Mean Group	Hausman test	Dynamic Fixed Effects[1]
Convergence coefficient	– 0.41	– 0.25**		
Long run coefficients				
Ly	– 5.937	– 10.513*	0.87	– 3.07**
Ly^2	0.508*	0.789*		0.26*
Short run coefficients				
Δly	10.02	11.49		– 13.82
Δly^2	– 0.509	– 0.58		2.08
Nº of countries	19	19		19
Nº of obs.	380	380		380
Log likelihood	526.70	471.35		– 294.01

Note: [1] t-stat. Calculated using heteroskedasticity consistent standard errors.

because dynamic specification and homogeneity restriction interact in a complex way, what may be the optimal order for country-specific estimates may not be optimal when cross-country homogeneity restrictions are imposed.

Diagnostic tests are reported in Table 4 for the ARDL(1,1,1) and in Table 5 for the SBC criterion used to select lags. There is evidence of serial correlation of the residuals in two countries (Tables 4, 5); functional form misspecification in six countries (Table 4) and in seven countries (Table 5); evidence of non-normality of residuals in only one case (Tables 4, 5); finally, there is evidence of heteroskedasticity in only one case (Table 4).

We run the regression, keeping only those countries whose results do no present any specification problems. Our results indicate that PMG estimates are only slightly different showing similar significance levels. We confirm that PMG seems to be robust to outliers and to the choice of lag order as stated by Pesaran *et al*. (1999).

When a linear form was estimated, we obtained a positive and significant long-run coefficient for income per capita (1.54) and the time trend was also non significant but negatively signed. However, the fit of the individual regressions was, in general, very poor in terms of adjusted R^2 and log likelihood. Moreover, the corresponding diagnostic statistics reported problems for a higher number of countries than when estimating a quadratic EKC. Results for a cubic EKC specification showed estimated coefficients on income per capita and squared income

TABLE 4

GROUP ESTIMATES AND DIAGNOSTIC STATISTICS FOR 19 COUNTRIES.
QUADRATIC FORM WITHOUT TREND. FIXED LAG STRUCTURE (1,1,1)

D.Var. lco	Phi[a]	ly[b]	ly^2[c]	SIGMA[d]	Ch-SC[e]	CH-FF[f]	CH-NO[g]	CH-HE[h]	RBARSQ[i]	LL[j]
Argentin	– 0.190	1.487	– 0.086	0.037	2.14	1.13	0.58	0.03	0.29	42.76
	(0.151)	(0.622)	(0.037)							
Brazil	0.018	1.487	– 0.003	0.047	2.14	1.10	3.29	0.47	0.99	34.76
	(0.043)	(0.622)	(0.084)							
Chile	– 0.523	1.487	– 0.186	0.561	3.47	4.99	0.11	0.04	0.48	– 12.37
	(0.205)	(0.622)	(0.055)							
Colombia	0.043	1.487	– 0.050	0.038	3.75	1.49	0.75	0.00	– 0.17	42.61
	(0.110)	(0.622)	(0.107)							
CostaRic	– 0.301	1.487	– 0.031	0.186	1.29	0.48	0.02	0.02	– 0.19	8.64
	(0.233)	(0.622)	(0.133)							
Dominica	– 0.088	1.487	– 0.077	0.094	3.85	0.60	0.01	0.57	0.54	21.58
	(0.080)	(0.622)	(0.065)							
Dom-Rep	– 0.060	1.487	– 0.067	0.092	3.24	10.09	8.85	0.22	0.98	22.04
	(0.047)	(0.622)	(0.057)							
Ecuador	– 0.324	1.487	0.026	0.389	0.38	21.25	2.10	17.85	0.45	– 5.41
	(0.139)	(0.622)	(0.074)							
Guyana	– 0.105	1.487	0.106	0.112	0.00	18.89	0.23	0.27	0.97	18.23
	(0.069)	(0.622)	(0.071)							
Haiti	– 0.054	1.487	0.069	0.074	0.00	0.74	0.58	0.66	0.98	26.19
	(0.087)	(0.622)	(0.135)							
Honduras	0.024	1.487	0.475	0.094	4.41	15.35	2.78	0.17	0.97	21.58
	(0.044)	(0.622)	(0.792)							
Mexico	– 0.082	1.487	– 0.058	0.067	3.17	2.32	0.02	0.30	0.44	27.95
	(0.123)	(0.622)	(0.127)							
Nicaragu	– 0.491	1.487	– 0.068	0.051	0.00	0.52	3.60	1.45	0.20	36.39
	(0.177)	(0.622)	(0.041)							
Paraguay	– 0.040	1.487	0.097	0.091	0.10	1.57	0.66	0.50	0.99	22.22
	(0.028)	(0.622)	(0.084)							
Panama	– 0.332	1.487	– 0.022	0.076	3.29	5.42	0.47	0.63	1.00	25.62
	(0.078)	(0.622)	(0.038)							
Peru	– 0.703	1.487	0.023	0.246	3.67	4.63	0.09	0.26	0.85	3.33
	(0.267)	(0.622)	(0.039)							
Salvador	– 0.609	1.487	– 0.067	0.026	6.49	1.82	0.55	0.58	0.46	50.14
	(0.162)	(0.622)	(0.045)							
Uruguay	– 0.031	1.487	– 0.030	0.137	0.60	15.92	0.17	0.28	0.98	14.42
	(0.083)	(0.622)	(0.114)							
Venezuel	– 0.295	1.487	0.010	0.049	0.44	0.94	0.53	0.69	0.45	33.96
	(0.104)	(0.622)	(0.038)							

Notes:
Figures in brackets are the standard errors.
a. Convergence coefficient
b. Estimated coefficient for income per capita
c. Estimated coefficient for income per capita squared
d. Standard error of the regression
e. Godfrey´s test of residual serial correlation
f. Ramsey´s Reset test of functional form
g. Jarque-Bera test of normality of regression residuals
h. Lagrange multiplier test of homoscedasticity
i. Adjusted R^2
j. Log-likelihood

TABLE 5

GROUP ESTIMATES AND DIAGNOSTIC STATISTICS FOR 19 COUNTRIES. QUADRATIC FORM WITHOUT TREND. LAGS (GROUP-SPECIFIC ESTIMATES OF THE LONG-RUN COEFFICIENTS BASED ON ARDL SPECIFICATIONS SELECTED USING THE SCHWARZ CRITERION)

D.Var. lco	Phi[a]	ly[b]	ly^2[c]	SIGMA[d]	Ch-SC[e]	CH-FF[f]	CH-NO[g]	CH-HE[h]	RBARSQ[i]	LL[j]
Argentin	– 0.345	– 10.513	0.595	0.031	0.67	0.03	0.74	0.03	0.51	44.39
	(0.125)	(0.728)	(0.041)							
Brazil	– 0.337	– 10.513	0.632	0.025	0.02	0.00	0.46	0.72	1.00	45.81
	(0.052)	(0.728)	(0.040)							
Chile	– 1.000	– 10.513	0.725	0.477	0.54	7.46	1.39	4.09	0.63	–11.27
	(NA)	(0.728)	(0.057)							
Colombia	– 0.012	– 10.513	0.835	0.031	9.36	0.56	0.49	2.15	0.25	45.49
	(0.017)	(0.728)	(0.096)							
CostaRic	– 0.346	– 10.513	0.819	0.174	1.27	0.76	0.13	0.19	– 0.04	8.53
	(0.214)	(0.728)	(0.079)							
Dominica	– 0.231	– 10.513	0.649	0.086	0.52	0.49	1.18	0.45	0.64	23.14
	(0.096)	(0.728)	(0.047)							
Dom-Repu	– 0.202	– 10.513	0.663	0.080	0.52	9.58	0.68	0.51	0.99	25.19
	(0.066)	(0.728)	(0.042)							
Ecuador	– 0.235	– 10.513	1.065	0.245	3.00	25.05	0.27	0.22	0.79	6.02
	(0.090)	(0.728)	(0.159)							
Guyana	– 0.054	– 10.513	0.920	0.116	0.02	21.59	0.18	0.19	0.97	17.61
	(0.056)	(0.728)	(0.171)							
Haiti	– 0.140	– 10.513	0.685	0.040	3.15	0.53	0.80	0.46	1.00	37.55
	(0.042)	(0.728)	(0.043)							
Honduras	0.024	– 10.513	1.363	0.094	3.95	15.45	3.31	0.17	0.97	21.50
	(0.068)	(0.728)	(1.163)							
Mexico	– 0.299	– 10.513	0.861	0.060	1.52	1.00	9.99	0.10	0.58	29.57
	(0.108)	(0.728)	(0.061)							
Nicaragu	– 0.231	– 10.513	0.712	0.053	0.30	6.75	3.14	0.03	0.11	33.97
	(0.104)	(0.728)	(0.049)							
Paraguay	– 0.040	– 10.513	0.800	0.085	0.51	3.02	1.48	0.38	0.99	23.19
	(0.017)	(0.728)	(0.069)							
Panama	– 0.130	– 10.513	0.701	0.094	4.27	3.71	0.05	0.60	0.99	22.24
	(0.051)	(0.728)	(0.045)							
Peru	– 1.000	– 10.513	0.748	0.210	3.23	3.32	0.30	0.33	0.89	5.48
	(NA)	(0.728)	(0.044)							
Salvador	0.143	– 10.513	0.789	0.028	1.21	0.18	1.18	3.26	0.17	47.61
	(0.050)	(0.728)	(0.053)							
Uruguay	– 0.068	– 10.513	0.701	0.122	0.01	14.99	0.37	0.35	0.98	16.04
	(0.066)	(0.728)	(0.060)							
Venezuel	– 0.231	– 10.513	0.725	0.056	0.59	3.47	0.44	0.69	0.32	29.31
	(0.198)	(0.728)	(0.049)							

Notes:

Figures in brackets are the standard errors.

a. Convergence coefficient

b. Estimated coefficient for income per capita

c. Estimated coefficient for income per capita squared

d. Standard error of the regression

e. Godfrey´s test of residual serial correlation

f. Ramsey´s Reset test of functional form

g. Jarque-Bera test of normality of regression residuals

h. Lagrange multiplier test of homoscedasticity

i. Adjusted R^2

j. Log-likelihood

per capita are both significant with a negative and positive sign respectively. However, the estimated coefficient on cubed income per capita was non-significant and negative signed. Nevertheless, the cubic specification seems to perform better for a number of countries indicating a reduction in the emissions when income is rising but an increase of them when income overpasses a certain level.

Appendix 2 shows scatter diagrams for GDP per capita against CO_2 emissions per capita for the 19 countries considered in our study. We have grouped the graphs in four categories according to the shape of the curves. We observe that 9 countries present a N-shaped curve, 2 countries show a curve with decreasing trend, other 2 countries show a U-shaped curve and 6 countries present an almost lineal up-ward slopping curve. Evidence confirms the need to develop studies that consider the existence of heterogeneity in country-panels and also the appropriateness of single country studies.

V. Conclusions

This paper presents empirical estimates of Environmental Kuznets Curves for a panel of 19 Latin-American and Caribbean countries over the period 1975-1998. A new econometric technique is applied, that allows for more flexible assumptions in a panel data framework with a large time dimension. Unlike most previous studies we test for slope heterogeneity of the income coefficient. A number of functional forms have been tested. A quadratic specification seems to be the more appropriate, although not all the coefficients are significant at conventional levels. Therefore, there is not a clear pattern related to the carbon dioxide emissions path in contrast with Martínez-Zarzoso *et al*. (2002) where a N-shaped EKC was shown for the most part of 22 Kyoto protocol Annex B countries. This finding is in accordance with Perman and Stern (1999)'s study on SO_2 emissions. These authors consider a large sample of countries and conclude that in many cases the cointegrating relation between sulphur emissions and income are not consistent with the EKC hypothesis.

The 19 countries we have studied do not exhibit a single behaviour since a great heterogeneity is observed in the scatter diagrams showing the shape of the relationship between emissions and income. Nevertheless, there is a common fact: emissions have been growing continuously since 1975 to the present time.

Although the levels of emissions in Latin-American and Caribbean countries are still lower than OECD's levels (1.68 and 8.62 annual tons per capita, respectively), similar to Dijkgraaf and Vollebergh (2001), our results might imply that there exists a serious risk that the environmental problem of climate change will not become internalised automatically if countries grow richer. Governments have the responsibility of enacting the commitments reached in the Kyoto protocol.

We are concerned that other explanatory variables like population density, openness to international trade, structural change or variables indicating political reforms could also help to improve the fit of the EKC estimations. We leave these questions open for further research.

Notes

1 These authors focus on natural capital as an input in production. Assuming that households affect the environment through their production and consumption decisions, they apply a dynamic model to analyse the environmental impact of such decisions. The authors conclude that the EKC hypothesis is supported or reversed depending on the indicator used to represent environmental pressure. Second, in a partial equilibrium context, the EKC only holds when the household faces an imperfect set of markets.

2 Even if the EKC holds, for different reasons, such avoiding safety and health risks, public intervention is still desirable to reduce faster environmental degradation.

3 These authors also study carbon dioxide emissions.

4 Austria, Belgium, Canada, Denmark, Finland, France, Germany, Iceland, Ireland, Italy, Japan, Luxembourg, Netherlands, Norway, New Zealand, Portugal, Sweden, Spain, Switzerland, United Kingdom and United States.

5 Argentina, Brazil, Chile, Colombia, Costa Rica, Dominica, Dominican-Republic, Ecuador, Guyana, Haiti, Honduras, Mexico, Nicaragua, Paraguay, Panama, Peru, El Salvador, Uruguay and Venezuela.

6 The time trend was not statistically significant.

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APPENDIX 1

DATA SOURCES

2000 World Development Indicators CD-ROM, The World Bank.

Series: CO_2 emissions, industrial (metric tons per capita)

Carbon dioxide emissions from industrial processes are those stemming from the burning of fossil fuels and the manufacture of cement. They include contributions to the carbon dioxide flux from solid fuels, liquid fuels, gas fuels, and gas flaring. Carbon dioxide (CO_2) emissions, largely a by-product of energy production and use account for the largest share of greenhouse gases, which are associated with global warming.

GDP per capita based on purchasing power parity (PPP) (current international $)

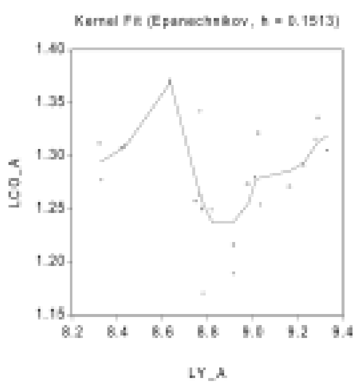
GDP PPP is gross domestic product converted to international dollars using purchasing power parity rates. The data are based on a 1993 reference year. An international dollar has the same purchasing power over GDP as the U.S. dollar in the United States.

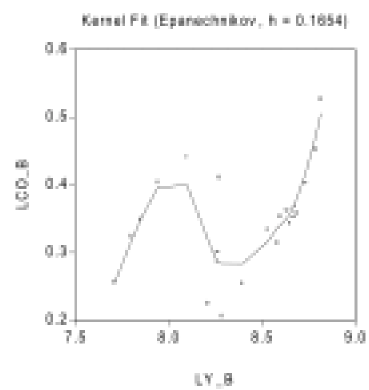
APPENDIX 2

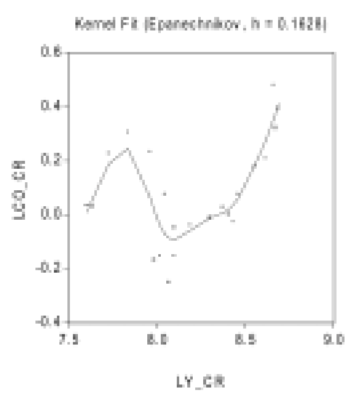
SCATTER DIAGRAMS: GDP AND CO_2 EMISSIONS PER CAPITA IN 19 COUNTRIES

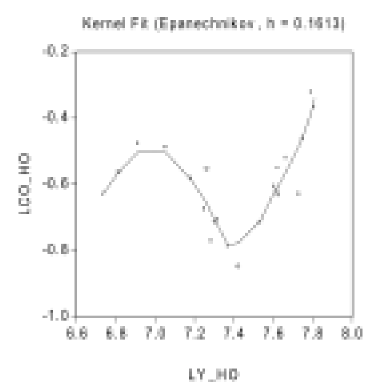
A: Argentina	D: Dominica	HO: Honduras	PE: Peru
B: Brazil	DR: Dominican-Republic	MEX: Mexico	SA: El Salvador
CH: Chile	EC: Ecuador	NIC: Nicaragua	U: Uruguay
COL: Colombia	GUY: Guyana	P: Panama	VE: Venezuela
CR: Costa Rica	HAI: Haiti	PA: Paraguay	

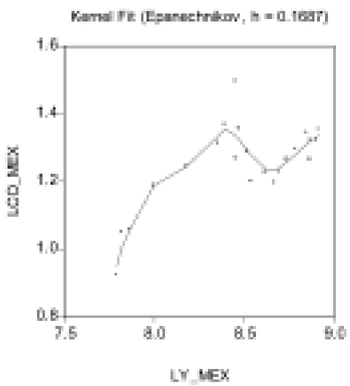
a) Countries with N-shaped EKCs:

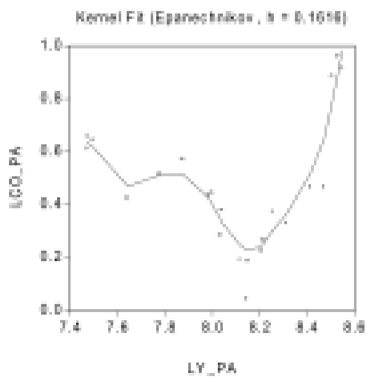









Kernel Fit (Epanechnikov , h = 0.1687)
LCD_MEX
LY_MEX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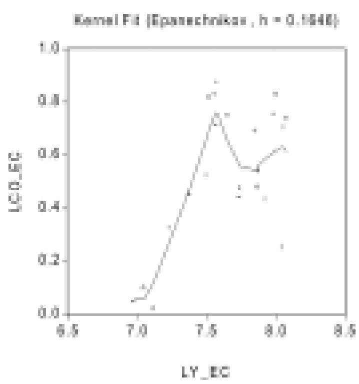
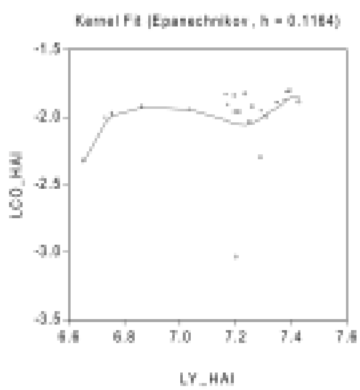
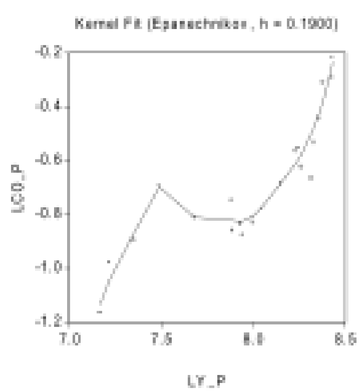
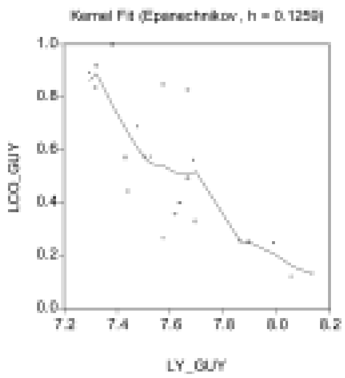
LCO_PA
LY_PA

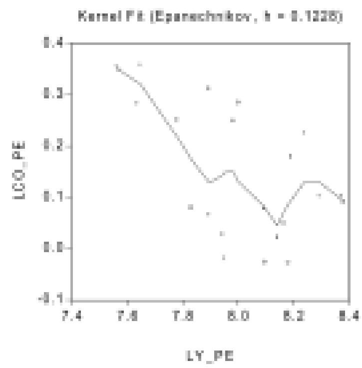

Kernel Fit (Epanechnikov , h = 0.1646)
LCO_EC
LY_EC


Kernel Fit (Epanechnikov , h = 0.1184)
LCO_HAI
LY_HAI

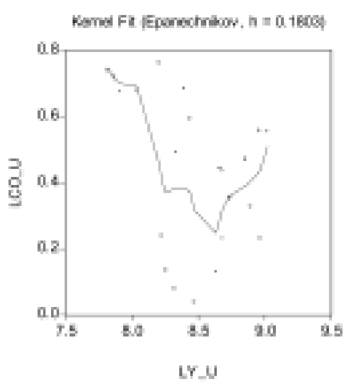

Kernel Fit (Epanechnikov , h = 0.1900)
LCO_P
LY_P

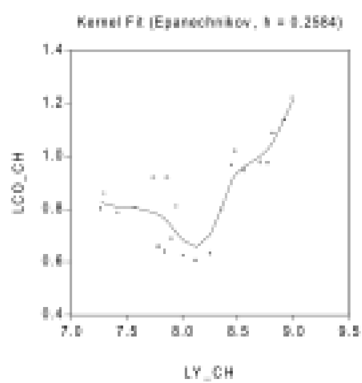
b) Countries with EKCs showing a decreasing trend:



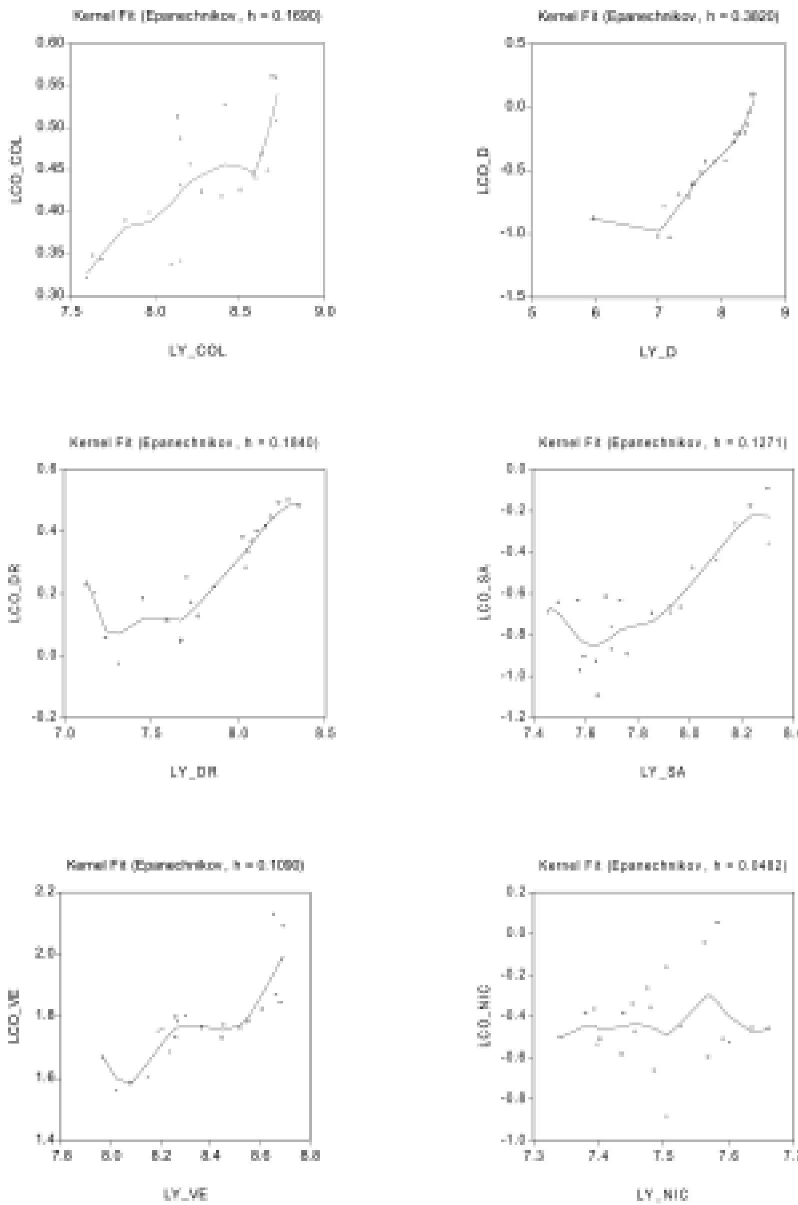


c) Countries with EKCs showing a “U” shaped curve:





d) Countries with EKCs showing increasing trend:



APPENDIX 3

TESTING FOR UNIT ROOTS AND COINTEGRATION

Unit roots

The testing procedure employed here attempt to minimise the problem and distortions caused by the presence of too many or too few deterministic variables. It involves starting with the most general specification of the augmented Dickey-Fuller test (Dickey and Fuller, 1981) and testing downwards. Essentially, moving from Equation 1 to Equation 3:

$$\Delta y_t = \alpha_0 + \gamma y_{t-1} + \alpha_2 t + \sum_{i=1}^{p} B_i \Delta y_{t-1+1} + \varepsilon_t \tag{1}$$

$$\Delta y_t = \alpha_0 + \gamma y_{t-1} + \sum_{i=1}^{p} B_i \Delta y_{t-1+1} + \varepsilon_t \tag{2}$$

$$\Delta y_t = \gamma y_{t-1} + \sum_{i=1}^{p} B_i \Delta y_{t-1+1} + \varepsilon_t \tag{3}$$

In these regressions many lags of the dependent variable are included to ensure that the residuals are serially uncorrelated. Table A.1 summarises our unit root test results. Null of unit root cannot be rejected for the variables in levels for all the countries, whereas the first differences of the variables are stationary (the null of unit root is rejected). Therefore, the univariate test statistics strongly support the view that both variables (income and emissions) are I(1) processes. We do not report tests statistics for $\ln(yh)^2$ since these are virtually the same as those for ln(yh).

The issue of nonstationarity can also be addressed directly within a panel data framework. Recently, a number of researchers (Quah, 1994; Im, Pesaran and Shin, 1996; Pedroni, 1998) have developed new tests which usually are more powerful than those used for single time series. Panel data exploits more information and therefore improves the power. However, since the results from these tests usually reinforce the findings of the individual countries results and we already found that the variables are I(1), we find not need to perform them.

Cointegration

It is well known that the test for cointegration using the Engle-Granger methodology suffer from low power compared to the test developed by Johansen. The

TABLE A.1

UNIT ROOT TEST RESULTS

Country	Augmented Dickey-Fuller Statistic (5% critical Value)			
	lco2	Δlco2	ly	Δly
Argentina	−2.73 (–3.66)	−4.71** (–3.67)	−3.21 (−3.64)	−3.84** (–3.65)
Brazil	0.69 (–1.95)	−2.71** (–1.95)	−2.36 (–3.63)	– 4.13** (–3.65)
Chile	– 0.84 (–3.01)	−3.24** (–3.02)	−3.07 (–3.64)	−3.30** (–3.01)
Colombia	−1.60 (–3.01)	– 4.14** (–3.02)	−2.66 (–3.63)	– 4.38** (–3.64)
Costa Rica	−1.65 (–3.01)	−2.78** (–1.96)	−1.72 (–3.00)	−2.91* (–2.65)[a]
Dominica	0.87 (–3.02)	−3.65** (–3.04)	1.36 (–1.95)	−2.65** (–1.96)
Dominican Republic	−0.72 (–3.01)	−4.99** (–3.02)	−2.73 (–3.66)	−3.83** (–3.66)
Ecuador	−2.45 (–3.01)	−4.51** (–3.02)	−2.64 (–3.01)	−2.84* (–2.64)[a]PP
Guyana	−3.43 (–3.65)	−7.18** (–3.67)	−2.86 (–3.63)	−3.29* (–3.26)[a]
Haiti	−1.41 (–3.04)	−3.52** (–3.02)	−1.75 (–3.63)	−3.56* (–3.25)[a]PP
Honduras	−1.23 (–3.01)	−3.93** (–3.02)	−3.25 (–3.63)	−3.85** (–3.02)
Mexico	−2.41 (–3.01)	−2.38** (–1.96)	−3.19 (–3.65)	−3.50* (– 3.27)
Nicaragua	−2.06 (–3.01)	−5.25** (–3.02)	−2.11 (–3.64)	−4.94** (–3.66)
Panama	−0.56 (–3.01)	−2.39** (–1.96)	−2.94 (–3.01)	−3.44* (–3.27)[a]
Paraguay	−0.24 (–3.01)	−3.12** (–3.02)	−3.22 (–3.63)	– 4.29** (–3.64)
Peru	−2.01 (–3.01)	−3.50** (–3.02)	−3.67 (–4.47)***	−2.76** (–1.95)
El Salvador	– 0.93 (–3.01)	−1.81* (–1.62)	−3.03 (–3.63)	−2.37** (–1.95)
Uruguay	−1.90 (–3.01)	−3.85** (–3.02)	−2.36 (–3.62)PP	– 4.11** (–3.64)
Venezuela	−2.60 (–3.69)	−3.27** (–1.96)	−2.42 (–3.63)	−2.41** (–1.95)

Notes: Five percent critical values are in parentheses. [a] indicates ten per cent critical values. Null of unit root is rejected if the test statistic is smaller than the corresponding critical value. **, * indicates rejection at the 5% and 10% level respectively. PP indicates that the Philips–Perron test has been used. Notice that critical values depend on the number of observations and the lag structure of error terms. The number of lags was selected by using a sequential search procedure: one–step reductions of the lag length are made until they can no longer be rejected in testing for the significance of the final included lag using a t test.

set of results presented are therefore obtained using the Johansen methodology: See Johansen (1988) and (1991).

The empirical specifications examined are:

$$\ln co_t = \beta_0 + \beta_1 \ln yh_t + \beta_2 (\ln yh)_t^2 + \varepsilon_t$$

$$\ln co_t = \beta_0 + \eta_t + \beta_1 \ln yh_t + \beta_2 (\ln yh)_t^2 + \varepsilon_t$$

Table A.2 shows the results. It contains the test statistics to establish the number of cointegrating vectors as well as the most stationary vector of coefficients from estimation. We find clear evidence in favour of a statistically significant relationship for most countries. Only in one case, that of Brazil, can the null of no cointegrating vectors be accepted.

Like the unit root test, single individual cointegration tests suffer from low power. This low power may lead to reject cointegration far more often that should be done.

From the panel unit root testing literature discussed above, some authors have also developed panel tests for cointegration. Pedroni (1998) presents a survey of the literature. They are residual-based tests of the null of no cointegration. Their main advantage is the improved power with respect to the single individual cointegration tests. Since we already find cointegrating vectors in all but one of the countries using the traditional methodology, we do not consider necessary to perform these tests.

TABLE A.2

JOHANSEN COINTEGRATION TEST (5% CRITICAL VALUE)

Country	Ly		ly2		No. CE(s)	Eigenvalue	Likel-R.5%
Argentina	0.04	(0.09)	–0.02	(0.01)	0	0.72	38.45**
					1	0.28	12.48
					2	0.26	5.93*
Brazil	–		–		**0**	0.56	30.85
					1	0.34	14.09
					2	0.25	5.88
Chile	0.0000	(0.05)	–0.07	(0.02)	0	0.75	43.27**
					1	0.52	15.30
					2	0.03	0.61
Colombia	–3.08	(0.52)	0.19	(0.03)	0	0.84	60.49**
					1	0.56	23.11
					2	0.29	6.78
Costa Rica	41.14	(4.71)	–2.51	(0.28)	0	0.83	47.19**
					1	0.38	11.17
					2	0.07	1.48
Dominica	–6.32	(1.05)	0.39	(0.06)	0	0.98	120.37**
					1	0.72	30.12*
					2	0.19	4.38
Dominican Republic	5.04	(1.16)	–0.35	(0.07)	0	0.77	49.01**
					1	0.50	19.22
					2	0.23	5.33
Ecuador	–72.67	(39.67)	4.62	(2.50)	0	0.61	38.76*
					1	0.46	19.87
					2	0.30	7.30
Guyana	29.23	(10.76)	–1.81	(0.70)	0	0.63	35.39*
					1	0.52	15.37
					2	0.01	0.32
Haiti	832.4	(1728)	–56.77	(117)	0	0.98	86.46**
					1	0.41	12.10
					2	0.09	1.91
Honduras	34.12	(7.98)	–2.28	(0.52)	0	0.79	45.60**
					1	0.50	15.43*
					2	0.11	2.23
Mexico	21.26	(5.24)	–1.22	(0.30)	0	0.95	69.24**
					1	0.40	10.15
					2	0.01	0.24
Nicaragua	40.11	(81.05)	–2.83	(5.39)	0	0.85	48.60**
					1	0.47	12.25
					2	0.01	0.29
Panama	40.44	(14.47)	–2.52	(0.88)	0	0.78	37.55**
					1	0.34	8.49
					2	0.02	0.49
Paraguay	153.9	(105)	–9.22	(6.29)	0	0.93	68.58**
					1	0.57	17.18*
					2	0.04	0.96
Peru	12.51	(7.59)	–0.77	(0.48)	0	0.69	40.14*
					1	0.48	17.88
					2	0.24	5.35
El Salvador	–322	(–794)	20.54	(50.74)	0	0.77	43.76**
					1	0.38	15.27
					2	0.27	6.09
Uruguay	66.02	(21.80)	–3.79	(1.24)	0	0.86	68.69**
					1	0.67	30.01**
					2	0.37	8.93
Venezuela	2.99	(4.45)	–0.19	(0.26)	0	0.62	39.51*
					1	0.55	20.85*
					2	0.26	5.79

Notes: No. CE(s) denotes the number of cointegrating vectors. Standard errors are in brackets.