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CO_2 emissions, GDP and trade: a panel cointegration approach

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ABSTRACT

This paper examines the relationships among per capita CO_2 emissions, per capita GDP and international trade based on panel data spanning the period 1960–2008 for 150 countries. A distinction is also made between OECD and non-OECD countries to capture the differences of this relationship between developed and developing economies. We apply panel unit root and cointegration tests and estimate a panel error correction model. The results from the error correction model suggest that there are long-term relationships between the variables for the whole sample and for non-OECD countries. Finally, Granger causality tests show that there is bidirectional short-term causality between per capita GDP and international trade for the whole sample and between per capita GDP and CO_2 emissions for OECD countries.

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1. Introduction

The relationships between economic growth (measured by increases in real GDP per capita) and pollution, as well as between economic growth and international trade, have been analyzed extensively during the last two decades. Also, as countries around the world continue to grow and develop there is increasing interest in elucidating more comprehensively the dynamic relationships among these variables. The purpose of this paper is to estimate the long-term relationships between per capita CO_2 emissions, per capita GDP and international trade, and to examine short-term causal relationships among these variables. To meet these objectives, we have analyzed a comprehensive panel data set, and two subsets of the data, using the econometric techniques of cointegration and error correction.

There are two well-established lines of research in the literature on this topic. The first originates from studies on environmental economics and is based on joint analysis of GDP and pollution. Much of this work has focused on testing the environmental Kuznets curve (EKC) hypothesis, according to which there is an inverted U-shaped relationship between pollution and GDP. The EKC hypothesis was proposed and tested in a seminal paper by Grossman and Krueger (1993). Stern (2004) and Dinda and Coondoo (2006), among others, have reviewed the literature on economic growth and environmental pollution in considerable detail. These reviews demonstrate that no single relationship fits all pollutants for all places and times. However, the existence of an EKC-type relationship has important policy implications. Specifically, policies that stimulate growth (e.g. trade liberalization, economic restructuring, etc.) may reduce environmental pollution in the long run.

The second line of research originates from studies on international economics and is primarily focused on the relationships between international trade on one hand and pollution and GDP growth on the other. Several authors have investigated whether international trade leads to increased pollution as a consequence of increased production or income (e.g. Copeland & Taylor 1994; Frankel & Romer 1999; Rodriguez & Rodrik 1999; Frankel & Rose 2002). These studies indicate that international trade can affect the environment, even if the empirical relationships between trade, GDP and different types of pollution are not clear-cut. For instance, openness to trade can have positive or negative effects on the environment (Grossman & Krueger 1993), because the overall effect is due to the combined impact of changes in industrial composition, increasing GDP and increasing demand for environmental quality. Furthermore, there is an extensive body of literature on the relationship between economic growth and international trade (see, for example, the surveys by Edwards 1998; Giles & Williams 2000a, 2000b; and Lewer & Van den Berg 2003). Much of this research deals with the link between exports and GDP by testing the export-led growth and growth-led export hypotheses. Different studies have yielded substantially divergent results, making it difficult to draw unambiguous conclusions. More recent studies have addressed the potential

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simultaneity of increases in pollution, GDP (or national income) and international trade rather than assuming (possibly erroneously) that trade and GDP are exogenous determinants of pollution (see Antweiler et al. 2001; Frankel & Rose 2005; Managi 2006; Managi et al. 2009). Frankel and Rose (2005) used an instrumental variables technique to test for a causal relationship between international trade and environmental pollution by analyzing cross-country data for 1990. The central focus of their work was the effect of trade on the environment for a given level of GDP per capita. They derived three equations: one for GDP, one for environmental pollution (specifically sulfur emissions) and one for trade. They also examined the endogeneity of trade openness, which was included as an explanatory variable in both the GDP and environmental quality equations, by introducing a gravity model of bilateral trade as a research instrument. The three derived equations were then used to test the validity of a proposed causal relationship between international trade and environmental pollution. Their results show that trade reduces sulfur dioxide emissions.

Some of the studies from both lines of research have focused on the relationship between GDP and the environment (e.g. the EKC) or between GDP and trade, while other authors such as Frankel and Rose (2005), Managi (2006) and Managi et al. (2009) have studied the relationship among CO_2 emissions, GDP growth and international trade using a single unified model to explicitly describe the endogeneity of GDP and trade.

Why is it interesting to study the relationship among GDP, international trade and CO_2 pollution? Much attention has been paid to global environmental problems: in particular the relationship between CO_2 emissions and trade liberalization policies. The debate focuses on two different but related issues (Huang & Labys 2002). The first, following the agenda of the Kyoto Protocol, is the rising trend in carbon emissions. One of the most important challenges for environmental policy in the near future will be to reduce these emissions. Hence, an understanding of the relationships between CO_2 emissions, GDP and international trade is essential for formulating effective public policy. The second major issue is trade openness, which probably promotes GDP growth but may also increase pollution. The ongoing globalization of the world's economy is increasing the volume of international trade, and this has further contributed to the growing interest in the relationships between international trade, economic growth and environmental pollution.

We have examined the relationships between CO_2 emissions, GDP and international trade by using three time-series econometric techniques – unit root testing, cointegration and the related Error Correction (EC) model – to analyze a panel data set. One of the key objectives of the present paper is to determine whether the time series for CO_2 emissions, GDP and international trade follow similar temporal trends. In addition, the directions of short-run causality among these three variables are examined. The analyzed data set consists of a data panel covering 150 countries for the 1960–2008 period. Separate estimates are presented for all countries, OECD countries and non-OECD countries. The sample was split into these two groups of countries because most developing countries, which are heavily represented among the non-OECD economies, are not signatories of the Kyoto Protocol. Consequently, the relationships that this paper attempts to capture are likely to differ substantially between developed and developing economies. Moreover, in recent decades many poor countries have experienced rapid economic development after adopting liberal economic policies (Akyüz & Gore 2001).

Our analysis is based on the panel EC and cointegration approach recently proposed by Westerlund (2007) to test whether CO_2, GDP and a common measure of international trade are cointegrated: i.e. whether there is a stationary linear combination of the random variables CO_2, GDP and international trade. The heterogeneous panel unit root test developed by Im et al. (2003) is used to check for stationarity. This paper thus fills a gap in the literature by using a dynamic panel error correction model to study the causal linkages among all three variables. In our framework, the per capita GDP, the measure of international trade, and per capita CO_2 emissions are treated as three potentially simultaneous variables, and the issue of short-run causality is addressed through a series of regressions where each variable is regressed against the other two.

The paper contributes to the literature in several ways. First, our work uses a larger data set than previous studies on similar topics that have used a panel approach of any kind. Dinda and Coondoo (2006) used a panel data-based cointegration approach to study incomes and emissions in 83 countries over 30 years, while Managi et al. (2009) used panel data for SO_2 and CO_2 emissions of 88 countries over 27 years, and biological oxygen demand (organic pollutant) emissions of 83 countries over 20 years. Our data set includes 150 countries as a full sample, 30 OECD countries and 120 non-OECD countries, over a period of 48 years. Second, the cointegration approach allows us to address the endogeneity problem that arises from the simultaneous determination of CO_2 emissions, GDP and international trade. This has been one of the most extensively discussed issues in previous publications on trade and the environment (e.g. Frankel & Rose 2005; Managi 2006; Managi et al. 2009). Third, most empirical studies focus on either the relationship between pollution and GDP or that between GDP and international trade. Very few (notable exceptions are the works of Managi 2006; and Managi et al. 2009) are based on panel data, primarily because of the lack of data on pollutant levels over longer periods of time. In contrast, our approach enables us to model the causal relationships of CO_2 emissions, GDP and international trade simultaneously, and examine how these variables change over time in both the short and long run. Fourth, our panel causality tests take into consideration the heterogeneity in the cross-section units and the non-stationary aspects of the panel structure of our data, both of which are neglected in most panel causality studies in this field.

The results from the error correction model suggest that there are long-term relationships between the variables for the whole sample and for non-OECD countries. Finally, Granger causality tests show that there is bidirectional short-term causality between per capita GDP and international trade for the whole sample and between per capita GDP and CO_2 emissions for OECD countries.

The implications of the results could be useful for policymakers to promote policies that integrate simultaneously CO_2 emissions regulations with international trade and GDP growth policies at global level in the long run. Moreover, promoting policies on CO_2 emissions, GDP and trade should be addressed differently between OECD countries and non-OECD countries in short.

The rest of the paper is organized as follows: Section 2 presents the data and the empirical methodology. Section 3 illustrates and discusses the empirical results. Finally, Section 4 presents and discusses our conclusions.

2. Empirical framework

2.1. Data sources and variables

As mentioned in the introduction, the full sample consists of data for 150 countries covering the 1960–2008 period.[1] Separate estimates were prepared for two groups of countries: the OECD nations (30 countries) and the non-OECD nations (120 countries). The basic country-level data – i.e. per capita real GDP (PPP) and information about international trade (exports and imports) – were obtained from the Penn World Table (Mark 7.0 2011). In the analysis below, the per capita GDP is expressed in US$ measured in real 2005 PPP-adjusted dollars and converted to log form, while the indicator of international trade is defined as exports plus imports divided by GDP, that is, the total volume of trade as a proportion of GDP. The corresponding country-level annual data on per capita CO_2 emissions, expressed in metric tons, were obtained from the Tables of National CO_2 Emissions prepared by the Carbon Dioxide Information Analysis Center (CDIAC 2011), Environmental Science Division, Oak Ridge National Laboratory, United States.

The standard summary statistics of our data are available in Appendix A, while the list of countries included in the analysis can be found in Appendix C. As can be seen from Table A1, the mean per capita CO_2 emissions are higher for OECD than for non-OECD countries. In addition, non-OECD countries exhibit the greatest range (distance between Max and Min) in metric tons of CO_2 released per capita. A similar trend is observed for per capita GDP, with the mean being greater for the OECD countries than for the non-OECD countries. With respect to international trade, Table A1 shows that non-OECD countries are more open to trade than OECD countries (as indicated by the ratio of total trade to GDP). It should be noted that the non-OECD sample includes some high- and medium-income countries according to World Bank (2011) classifications.

Hereafter, log values of real GDP per capita[2] are denoted by Y, per capita CO_2 emissions by E and the measure of international trade by T.

2.2. Econometric technique

As indicated in the introduction, this paper examines the relationships among E, Y and T. For these macroeconomic data, growth could occur via deterministic time trend, or it could occur because the annual change in the variable is equal to a constant. In this latter case, the variable is equal to its lagged value plus an intercept and is referred to as having a unit root with drift. E, Y and T are generally growing over time. If variables have different trends processes, they cannot stay in fixed long-run relation to each other, implying that we cannot model the long run, and there is usually no valid base for inference based on standard distributions. If our data appeared to be nonstationary, we may have spurious regression results. Consequently, it is important to investigate if E, Y and T are integrated to determine the appropriate transformations to render them stationary. To address the stationarity properties for Y, E and T, a panel data unit root test of Im, Pesaran and Shin (Im et al. 2003; hereafter the IPS test) based on the well-known Augmented Dickey–Fuller (ADF) procedure is performed to determine whether or not the observed country-specific time series exhibit stochastic trends. IPS test is appropriate for balanced panels and it combines information from the time-series dimension with the cross section dimension, such that fewer time observations are required for the test to have power (less likely to commit a Type II error).[3] IPS test has also been found to have superior test power to analyze long-run relationships in panel data. The null hypothesis of this test means that all series in the panel are non-stationary processes, and the corresponding alternative hypothesis is that some (but not necessarily all) of the individual series in the panel are stationary. Next, cointegration analysis is performed to examine whether the variables are cointegrated (i.e. whether there are stable long-term equilibrium relationships among them). Finally, an Error Correction Model (ECM) is estimated, to test the short-term causality relationships among E, Y and T.

2.2.1. Error correction-based panel cointegration tests

We apply the panel cointegration tests developed by Westerlund (2007) and Persyn and Westerlund (2008). The aim here is to test for the absence of cointegration by determining whether Error Correction exists for individual panel members or for the panel as a whole.

Consider the ECMs described by Equations (1)–(3), in which all variables in levels are assumed to be $I(1)$:

$$\begin{aligned}\Delta E_{i,t} = {} & \alpha_i^E + \lambda_i^E (E_{i,t-1} - \beta_i^E Y_{i,t-1} - \gamma_i^E T_{i,t-1}) \\ & + \sum_{j=1}^{n} \theta_{i,j}^E \Delta E_{i,t-j} + \sum_{j=1}^{p} \varphi_{i,j}^E \Delta T_{+i,t-j} \\ & + \sum_{j=1}^{m} \delta_{i,j}^E \Delta Y_{i,t-j} + u_{i,t} \qquad (1)\end{aligned}$$

$$\begin{aligned}\Delta Y_{i,t} = {} & \alpha_i^Y + \lambda_i^Y (Y_{i,t-1} - \beta_i^Y E_{i,t-1} - \gamma_i^Y T_{i,t-1}) \\ & + \sum_{j=1}^{n} \delta_{i,j}^Y \Delta Y_{i,t-j} + \sum_{j=1}^{m} \theta_{i,j}^Y \Delta E_{i,t-j} \\ & + \sum_{j=1}^{p} \varphi_{i,j}^Y \Delta T_{+i,t-j} + \varepsilon_{i,t} \qquad (2)\end{aligned}$$

$$\begin{aligned}\Delta T_{i,t} = {} & \alpha_i^T + \lambda_i^T (T_{i,t-1} - \beta_i^T Y_{i,t-1} - \gamma_i^T E_{i,t-1}) \\ & + \sum_{j=1}^{p} \varphi_{i,j}^T \Delta T_{i,t-j} + \sum_{j=1}^{m} \delta_{i,j}^T \Delta Y_{i,t-j} \\ & + \sum_{j=1}^{n} \theta_{i,j}^T \Delta E_{+i,t-j} + e_{i,t}. \qquad (3)\end{aligned}$$

Here, the parameters λ_i^k, $k \in \{E, Y, T\}$ are the parameters of the *Error Correction* (EC) term and provide estimates of the speed of error correction towards the long-run equilibrium for country i, while $\varepsilon_{i,t}, u_{i,t}$ and $e_{i,t}$ are white noise random disturbances.

We focus on E and its relation to Y and T; therefore, Equation (1) is our primary equation of interest. Equations (2) and (3) can potentially be ignored if Y and T can be treated as weakly exogenous, and the validity of this assumption can be tested by performing regressions with $\Delta Y_{i,t}$ and $\Delta T_{i,t}$ as dependent variables.

Two different classes of tests can be used to evaluate the null hypothesis of no cointegration and the alternative hypothesis: group-mean tests and panel tests. Westerlund (2007) developed four panel cointegration test statistics (G_a, G_t, P_a and P_t)[4] based on the ECM. The group-mean tests are based on weighted sums of λ_i^k estimated for individual countries, whereas the panel tests are based on an estimate of λ^k for the panel as a whole. These four test statistics are normally distributed. The two tests (G_t,P_t) are computed with the standard errors of λ_i^k estimated in a standard way, while the other statistics (G_a,P_a) are based on Newey and West (1994) standard errors, adjusted for heteroscedasticity and autocorrelations. By applying an ECM in which all variables are assumed to be $I(1)$, the tests proposed by Westerlund (2007) examine whether cointegration is present or not by determining whether error correction is present for individual panel members and for the panel as a whole.

If $\lambda_i^k < 0$, then there is an error correction, which implies that $Y_{i,t}$ and $E_{i,t}$ and $T_{i,t}$ are cointegrated, whereas if $\lambda_i^k = 0$ there is no error correction and thus no cointegration. Thus, the null hypothesis of no cointegration for the group-mean tests (G_a and G_t test statistics) is as follows: $H_0^G : \lambda_i^k = 0$ for all i, which is tested against $H_1^G : \lambda_i^k < 0$ for at least one i. In other words, in the two group-mean-based tests, the alternative hypothesis is that there is cointegration in at least one cross-sectional unit. Therefore, the adjustment coefficient λ_i^k may be heterogeneous across the cross-section units. Rejection of H_0^G should therefore be taken as evidence of cointegration in at least one of the cross-sectional units. The panel tests (P_a and P_t test statistics) instead assume that $\lambda_i^k = \lambda^k$ for all i, so the alternative hypothesis is that adjustment to equilibrium is homogenous across cross-section units. Then, we test $H_0^P : \lambda^k = 0$ against $H_1^P : \lambda^k < 0$. Rejection of H_0^P should therefore be taken as evidence of cointegration for the panel as a whole.

We are mainly interested in the long-run behavior of our model, so the next step is to determine the coefficients of the conditional long-run relationships between E, Y and T when the short-run terms are set to zero. The long-run coefficients can be easily derived from the following long-run equation, obtained from the reduced form of (1) when the terms representing short-run changes are $\Delta E = \Delta T = \Delta Y = 0$, as follows:

$$E_{i,t} = -\frac{\alpha_i^E}{\lambda_i^E} + \beta_i^E Y_{i,t} + \gamma_i^E T_{i,t}$$

Finally, we also test for short-run causality. This implies testing the significance of the coefficients of the lagged difference of the variables (using the Wald restriction test) for Equations (1)–(3). The causality of individual relationships is tested by checking the significance of the t-statistic for the coefficient of the lagged variable, while the joint causality is tested as follows.

We can test the null hypotheses that the other two variables are not sources of short-run causality of E, Y and T by testing whether $H_0 : \varphi^E_{i,j} = \delta^E_{i,j} = 0 \forall i$, $H_0 : \theta^Y_{i,j} = \varphi^Y_{i,j} = 0 \forall i$ and $H_0 : \theta^T_{i,j} = \delta^T_{i,j} = 0 \forall i$ (Equations (1)–(3)), respectively, and if these null hypotheses are rejected, we will have bidirectional causality.

3. Results and discussion

The panel unit root test results for E, Y and T over the full sample are summarized in Table 1. The decision of whether or not to reject the null hypothesis of unit root for the panel as whole is based on the $W_{[t-bar]}$ statistic which is applicable to heterogeneous cross-sectional panels.

We were not able to reject the null unit root hypothesis for the Y and T series when expressed in level form. However, E is stationary without a trend term. When using the first differences, the null of unit roots is strongly rejected at the $P < 0.01$ significance level for all three series, implying that the series are I (1). This finding is confirmed by all tests employed for all three country samples examined: i.e. the full sample and both the OECD and non-OECD subsamples, although the corresponding values are not presented herein.

We proceed by testing whether Y, E and T are cointegrated (see Appendix B for the specifications used in the four cointegration tests). We adopt the Westerlund-based panel cointegration tests using a single lag and lead, $h_i = q_i = 1$. The lead and lag orders were selected based on the minimum AIC (Akaike's Information Criterion). We perform cointegration tests with both a constant and a trend, no constant or trend, and with a constant but no trend. We also consider the robust P-values obtained after bootstrapping using 800 replicates after testing for cross-sectional dependence among residuals.

Results obtained from the model with a constant but no trend suggest that there is no cointegration for Y and T (Table 2; see Table B1 in Appendix B for results from the other cointegration tests). However, as can be seen in Table 2, our results for the whole sample – i.e. from the panel cointegration tests – indicate that there is a long-run cointegrating relationship for E among the series under consideration, based on Equation (1). The P_t and P_a statistics indicate that the null hypothesis of no cointegration for E should be rejected at the $P < 0.01$ level. The other models (neither constant nor trend, and both a constant and trend) also indicate that the null hypothesis of no cointegration for E should be rejected at the $P < 0.01$ level. The robust P-values indicate that the null hypothesis of no cointegration should be rejected at the $P < 0.05$ level for the full sample and the $P < 0.01$ level for non-OECD countries.

As can be seen from the P-values, for the income equation (Y) the null hypothesis of no cointegration cannot be rejected for either the full sample or the OECD sample. However, the P_t and P_a values indicate that the null hypothesis of no cointegration (and hence no stationary equilibrium relationship among the variables) should be rejected at $P < 0.01$ for the non-OECD sample. At the same time, the robust P-values indicate that we can reject the null hypothesis of no cointegration for neither the full sample nor the OECD and non-OECD countries.

Table 1. Im-Pesaran-Shin test for unit root in panels for the full sample.

	Levels		First differences	
	Constant	Constant and trend	Constant	Constant and trend
Variable	$W_{[t-bar]}$ Statistic			
E	−3.237***	−1.033	−24.923***	−19.998***
Y	21.922	10.907	−22.094***	−18.305***
T	6.054	4.588	−33.093***	−30.063***

***Indicates significance at the $P < 0.01$ level.

Table 2. Results of the Westerlund-based panel cointegration tests.

		With constant but no trend											
		Full sample				OECD				Non-OECD			
Model	Test	Value of test	Z-value	P-value	Robust P-value	Value of test	Z-value	P-value	Robust P-value	Value of test	Z-value	P-value	Robust P-value
Y	G_t	−1.706	4.379	1.000	0.998	−1.585	2.676	0.996	0.954	−1.720	3.746	0.996	0.908
	G_a	−5.885	6.326	1.000	0.995	−5.508	3.158	0.999	0.934	−5.966	5.516	1.000	0.946
	P_t	−18.152	2.740	0.997	0.789	−7.420	1.901	0.971	0.745	−64.092	−43.894	0.000	0.858
	P_a	−4.919	2.073	0.981	0.523	−4.263	1.569	0.942	0.614	−18.629	−25.007	0.000	0.759
E	G_t	−2.01	0.329	0.629	0.741	−1.916	0.706	0.760	0.741	−2.032	0.040	0.516	0.002
	G_a	−7.701	2.78	0.997	0.171	−5.147	3.473	1.000	0.984	−8.336	1.377	0.916	0.024
	P_t	−27.016	−5.845	0.000	0.030	−8.020	1.320	0.907	0.735	−24.819	−5.862	0.000	0.000
	P_a	−8.023	−4.728	0.000	0.000	−4.098	1.731	0.958	0.794	−8.401	−4.968	0.000	0.000
T	G_t	−1.693	4.550	1.000	0.999	−0.791	7.400	1.000	1.000	−1.545	5.827	1.000	1.000
	G_a	−5.686	6.715	1.000	0.958	−1.311	6.824	1.000	1.000	−4.695	7.737	1.000	1.000
	P_t	−20.841	0.135	0.554	0.121	−2.225	6.932	1.000	0.980	−17.871	0.867	0.807	0.890
	P_a	−6.051	−0.407	0.342	0.203	−1.262	4.509	1.000	0.966	−5.495	0.726	0.766	0.818

We then used xtwest to test for cointegration, using the AIC to choose the optimal lag and lead lengths for each series and with the Bartlett kernel window width set to $4(T/100)^{2/9} \approx 3$.

For the trade equation (T), there is cointegration across the panel as a whole when the model is estimated without constant and trend terms. However, the addition of either a constant alone or a constant and a trend term makes all of the test statistics nonsignificant for all of the samples. Thus, the null hypothesis of no cointegration in the trade equation cannot be rejected for the model with either a constant or both constant and trend terms.

Because of differences in their construction, 'group-mean' and 'panel' tests can give different results, and the G_a and G_t test statistics do not indicate that the null hypothesis of no cointegration can be rejected, even at $P < 0.10$ (except for E in the non-OECD countries, for which the robust P-values of the G_a and G_t test statistics indicate that the null hypothesis can be rejected at the $P < 0.05$ level).

Caution is required when interpreting the results of our tests for the emission equation. Given the definitions used, one would expect the group-mean tests to reject the null hypothesis more often than the panel tests (because at least one series is cointegrating in the former case, which might not necessarily show up in the latter test), and not the opposite. When analyzing a small data set, such as that used here (T = 48), the results of the two tests should be interpreted carefully.[5] As a consequence, for our data it seems that panel tests are probably more appropriate than group-mean tests.[6]

The economic implication of the existence of cointegration is that there is a stable equilibrium long-run relationship among the variables E, Y and T. Table 2 provides evidence of cointegration in the emissions equation for both the full sample and non-OECD countries. However, the other models suggest that there is no cointegration of Y, except for some evidence of cointegration for Y based on the P-value obtained from the panel tests for non-OECD countries. Thus, results based on the income equation should be interpreted with caution.

A further consideration is that our results are somewhat mixed, especially the robust P-values. For both the full sample and non-OECD countries, only the panel tests suggest there are long-run relationships among E, Y and T. When we account for cross-sectional dependence using the bootstrap approach, we get somewhat different results. For both the full sample and non-OECD countries, cointegration is still confirmed by the panel tests; however, for non-OECD countries the group-mean tests also indicate that the no cointegration hypothesis should be rejected.

Overall, the primary model used in this study suggests that there are long-run relationships among E, Y and T for the whole sample as a panel and for non-OECD countries, both as a panel and as individual panel members.

3.1. Error correction model estimates

Given the evidence of panel cointegration, the long-run relationships among E, Y and T can be further estimated by applying the estimator of Westerlund (2007). Therefore, we estimate Equations (1)–(3) of the ECM, reparameterized based on panel data. Table 3 reports the findings for the three specifications for comprehensiveness, although our focus is on E (Equation (1)).

We approach the interpretation of the regression results presented in Table 3 from the point of view of short-run fluctuations around a long-run equilibrium relationship. In Table 4 we report the results for the long-run relationships of E, Y and T, while Table 5

Table 3. Results of the ECM estimates.

	Full sample			OECD			Non-OECD		
Regressors	ΔY	ΔE	ΔT	ΔY	ΔE	ΔT	ΔY	ΔE	ΔT
Constant	0.249***	−0.309***	−16.44***	0.310***	−0.548**	−8.921**	0.465***	−0.410***	−8.113*
	(12.76)	(−3.51)	(−5.68)	(9.44)	(−2.69)	(−2.92)	(17.86)	(−4.17)	(−2.44)
$Y_{(t-1)}$	−0.030***	0.053***	2.922***	−0.032***	0.075**	0.975**	−0.060***	0.062***	2.126***
	(−12.06)	(4.71)	(7.92)	(−8.46)	(3.24)	(2.80)	(−17.34)	(4.77)	(4.86)
$E_{(t-1)}$	0.003	−0.130***	−0.138	0.0002	−0.055***	−0.320*	0.008***	−0.142***	0.153
	(1.88)	(−20.89)	(−0.63)	(0.15)	(−6.24)	(−2.36)	(3.73)	(−19.49)	(0.58)
$T_{(t-1)}$	0.021***	0.002	−0.093***	0.039***	−0.173***	0.022***	0.028***	0.019	−0.104***
	(5.57)	(0.15)	(−17.68)	(5.64)	(−4.23)	(3.57)	(5.38)	(0.99)	(−17.16)
$\Delta Y_{(t-1)}$	0.123***	0.087	7.135***	0.230***	0.191	−7.008**	−0.141***	0.060	−0.056
	(10.45)	(1.64)	(4.11)	(9.06)	(1.25)	(−3.10)	(−12.64)	(1.45)	(−0.04)
$\Delta T_{(t-1)}$	0.020*	0.017	−0.088***	0.077*	−0.323	0.104***	−0.011	0.016	−0.098***
	(2.40)	(0.47)	(−7.31)	(2.32)	(−1.66)	(3.58)	(−1.00)	(0.29)	(−7.31)
$\Delta E_{(t-1)}$	0.002	−0.057***	0.120	0.006	0.009	0.598	0.002	−0.058***	−0.017
	(0.89)	(−4.72)	(0.30)	(1.19)	(0.34)	(1.46)	(0.54)	(−4.25)	(−0.04)
ΔY		0.058	10.690***		0.134	9.164***		0.057	11.10***
		(1.09)	(6.07)		(0.86)	(3.90)		(0.95)	(5.58)
ΔE	0.008**		0.162	0.014**		−0.690	0.009*		0.283
	(2.94)		(0.40)	(3.08)		(−1.71)	(2.45)		(0.61)
ΔT	0.043***	−0.002		−0.024	0.083		0.042***	0.006	
	(5.28)	(−0.07)		(−0.76)	(0.46)		(3.81)	(0.15)	
N	6898	6898	6898	1380	1380	1380	5520	5520	5520

Values in parentheses are t-values. Significance levels: *, $P < 0.05$; **, $P < 0.01$; ***, $P < 0.001$. Lag and lead lengths both 1.
The 'xtwest' Stata command was applied for the ECM-based panel cointegration test (Persyn & Westerlund 2008).

Table 4. Estimated long-run ECM coefficients.

	FULL SAMPLE			OECD			NON-OECD		
Variable	α_i^k	β_i^k	γ_i^k	α_i^k	β_i^k	γ_i^k	α_i^k	β_i^k	γ_i^k
Y	8.30	0.10	0.70	9.70	0.0075	1.20	7.70	0.13	0.46
E	2.40	0.40	0.10	9.90	1.40	3.10	2.90	0.40	0.10
T	176.8	31.4	1.50	405.50	44.3	14.50	78	20.4	1.40

It is difficult to test the significance of $\alpha_i^k, \beta_i^k, \gamma_i^k$ because the variances for these coefficients may not be available, so we did not estimate their standard errors. Y is the log of GDP.

Table 5. Results of the short-run causality tests.

Causality test	Null hypothesis	Full sample	OECD	Non-OECD
$\Delta Y + \Delta T \rightarrow \Delta E$	$\varphi_{i,j}^E = \delta_{i,j}^E = 0$	7.81**	26.54***	2.66
$\Delta Y \rightarrow \Delta E$	$\delta_{i,j}^E = 0$	0.142** (2.77)	0.671*** (4.78)	0.0235 (0.60)
$\Delta T \rightarrow \Delta E$	$\varphi_{i,j}^E = 0$	0.0273 (0.76)	−0.446** (-2.61)	0.0639 (1.57)
$\Delta E + \Delta T \rightarrow \Delta Y$	$\theta_{i,j}^Y = \varphi_{i,j}^Y = 0$	15.51***	42.01***	7.04
$\Delta E \rightarrow \Delta Y$	$\theta_{i,j}^Y = 0$	0.00490 (1.80)	0.0237*** (4.76)	0.00372 (0.99)
$\Delta T \rightarrow \Delta Y$	$\varphi_{i,j}^Y = 0$	0.0286*** (3.51)	0.122*** (3.77)	0.0271* (2.46)
$\Delta E + \Delta Y \rightarrow \Delta T$	$\theta_{i,j}^T = \delta_{i,j}^T = 0$	30.16***	0.55	12.49***
$\Delta E \rightarrow \Delta T$	$\theta_{i,j}^T = 0$	−0.000743 (-0.19)	0.00244 (0.57)	−0.0000737 (-0.02)
$\Delta Y \rightarrow \Delta T$	$\delta_{i,j}^T = 0$	0.0934*** (5.49)	−0.0140 (-0.62)	0.0488*** (3.53)

Values in parentheses are t-values. Significance levels: *, $P < 0.05$; **, $P < 0.01$; ***, $P < 0.001$.
For the co-joint test, we used the Wald-test (χ^2).

presents results of the test of the short-run causality relationships. In Table 3, all of the estimated adjustment parameters (i.e. the coefficients of the EC term) are statistically significant and have the *expected* negative sign, except those for the OECD countries when T is taken as the dependent variable. This result is consistent with the findings reported by Dinda and Coondoo (2006), of negative coefficients for Africa, Central America, America as whole, Eastern Europe, Europe as a whole and the world.

In the equation for E, we find that λ^E is negative for each of the three country groups. This implies that if $E_{t-1} > \beta^E Y_{t-1} + \gamma^E T_{t-1}$, the EC term induces a negative change in E back towards the long-run equilibrium. We obtain larger absolute values for the non-OECD countries (0.142) and the full sample (0.130) than for the OECD countries (0.055). This implies that a much longer time will be required for equilibrium to be restored following any deviation from the long-run equilibrium in the OECD countries than in the non-OECD countries. Therefore, this empirical evidence suggests structural divergences between the OECD and non-OECD countries in the speed of adjustment towards long-run equilibrium.

The estimated long-run ECM coefficients are presented in Table 4.

According to our results, for the full sample and non-OECD economies, a 1% increase in Y will increase E by 0.4 metric tons, which represents the long-term effect of Y on E over future periods. The increase of Y will cause deviations from the long-run equilibrium, causing E to be too high. E will then decrease to correct this disequilibrium, with the deviation decreasing by 13% (λ_i^E) in each subsequent time period for the full sample and 14% for non-OECD economies. That is, E will decrease by an average of 0.4 metric tons in response, with the decrease occurring over successive future measurement intervals at a rate of 13% for the full sample and 14% for non-OECD economies. A one-unit increase in T will increase E by 0.10 metric tons in both cases. To reestablish equilibrium E will then decrease by 0.10 metric tons over successive future measurement intervals at a rate of 13% for the full sample and 14% for non-OECD economies per interval. For OECD countries, an increase of 1% in Y will increase E by 1.4 metric tons, while a one-unit increase in T will increase E by 3.1 metric tons. The return to equilibrium will occur at a rate of 5.5% per time interval.

The results of the short-run causality tests are presented in Table 5, where the direction of causal relationships is indicated by (→) for unidirectional causal relationships. According to our results, the relationship between Y and E exhibits bidirectional causality for OECD countries: i.e. a change in Y will affect E and a change in E will similarly affect Y. There is also a bidirectional relationship between T and Y for the full sample, implying that a change in Y will affect T and vice versa. For non-OECD countries, E and Y are causally related to T, and there are unidirectional causal relationships from Y to T.

The main findings can be summarized as follows. There is strong bidirectional short-run causality between CO_2 and GDP for OECD countries. This is consistent with expectations, since the OECD experienced a significant increase in CO_2 emissions that was especially notable in certain countries over the studied period. Furthermore, the higher the country's GDP (and income), the greater the amount of CO_2 that is likely to be released via production and/or consumption.

Dinda and Coondoo (2006) also found cointegrating relationships between CO_2 and GDP for Eastern and Western Europe, Central America, Africa, Japan and Oceania. In addition, they found evidence for their panel as a whole that strongly points to the existence of bidirectional causality.

Finally, there is bidirectional causality between international trade and GDP for the full sample and unidirectional causality between the same variables for OECD countries. For non-OECD countries, there are no direct effects of GDP and trade on emissions. This implies that neither GDP growth nor international trade seems to have any significant short-run effect on CO_2 emissions for non-OECD countries.

4. Conclusions

In this paper, we analyzed cointegration and short-run causal relationships between per capita CO_2 emissions, per capita GDP and international trade based on a cross-country panel data set covering 150 countries during the 1960–2008 period. Our estimates are based on the full sample of countries, as well as on two separate subsamples, comprising OECD and non-OECD countries, respectively.

Using the unit root test procedure, we found that all three series (the logarithm of the per capita GDP, per capita CO_2 emissions and trade measure) follow $I(1)$processes. These findings were then used to apply ECM-based panel cointegration tests (Westerlund 2007). The robust *P*-values obtained from the panel tests indicated that the null hypothesis of no cointegration should be rejected for both the full sample and the non-OECD countries, while group-mean tests indicated that the null hypothesis of no cointegration can be rejected only for non-OECD countries. This suggests that per capita CO_2 emissions, per capita GDP and the measure of international trade are cointegrated. Consequently, there are long-run equilibrium relationships among these three variables for both the full sample and the non-OECD sample. Our results are consistent with previous findings; Dinda and Coondoo (2006) found a cointegration relationship between CO_2 emissions and GDP for 88 countries between 1960 and 1990, while Al-Mulali (2011) found a long-run relationship between CO_2 emissions and GDP for MENA countries.[7]

The possible existence of causal relationships among per capita CO_2 emissions, per capita GDP and international trade has also been tested. The results suggest that there are short-run bidirectional causal relationships between per capita GDP and trade, together with a causal relationship between CO_2 emissions plus GDP and trade, for the full sample. These findings suggest that economic policies should address growth, international trade and environmental pollution simultaneously.

Differences in the direction of causality have been detected between the two subsamples considered. In the OECD sample, our results suggest there is bidirectional causality between per capita GDP and CO_2 emissions. This implies that policymakers should consider CO_2 emissions and economic growth simultaneously. Our results are partially consistent with those of Coondoo and Dinda (2002), who found a unidirectional causal relationship between CO_2 emissions and GDP for developed country groups in North America and Western Europe and a unidirectional causal relationship from GDP to CO_2 emissions for country groups of Central and South America, Oceania and Japan.

For OECD countries, our results suggest that there are also causal relationships from GDP and international trade to per capita CO_2 emissions, and from per capita CO_2 emissions and international trade to per capita GDP. Conversely, for non-OECD countries there are two unidirectional relationships, from per capita GDP to international trade and from per capita CO_2 emissions and per capita GDP to international trade. The absence of causal relationships between per capita CO_2 emissions and per capita GDP in non-OECD countries implies that we do not have clear evidence that GDP affects CO_2 emissions. In contrast, previous studies (e.g. Coondoo & Dinda 2002) have identified a bidirectional relationship between these variables for Asian and African countries in the 1960–1990 period.

We would like to stress that a comprehensive analysis in this field would require a study of relationships among income, trade, emissions and energy, specifying the type of energy used, the structural composition of GDP and available technology, among other factors. However, the empirical framework employed in this study could be used to estimate the short- and long-run elasticities of CO_2 emissions in disaggregated sectors, in order to calibrate the developed models and generate scenarios describing how openness policies might motivate businesses to adopt environmentally friendly and efficient technologies to reduce emissions.

Notes

1. We omitted 16 countries for which we had insufficient historical data on international trade and CO_2 emissions.
2. We take the log of GDP for scale reasons and to facilitate interpretation of the coefficients.
3. Monte Carlo simulation in the IPS study show that the small sample properties of IPS test are superior to those of the best-known Levin and Lin (1993) panel test.
4. $G_t = \frac{1}{N}\sum_{i=1}^{N}\frac{\hat{\lambda}_i^k}{s.e.(\hat{\lambda}_i^k)}$, $G_a = \frac{1}{N}\sum_{i=1}^{N}\frac{T\hat{\lambda}_i^k}{\hat{\lambda}_i^k(1)}$; $\lambda_i^k(1) = \hat{\omega}_{ui}/\hat{\omega}_{xi}$, where $\hat{\omega}_{ui}$ and $\hat{\omega}_{xi}$ are the usual Newey and West (1994) standard error corresponding to the long-run variance estimators $\hat{\omega}_{xi}^2 = \frac{1}{T-1}\sum_{j=-M_i}^{M_i}\left(1-\frac{j}{M+1_i}\right)\sum_{t=j+1}^{T}\Delta x_{it}\,\Delta x_{it-j}$, where M_i is a bandwidth parameter that determines how many covariances to estimate in the kernel. $\hat{\omega}_{ui}$ may be obtained as above using kernel estimation with Δx_{it} replaced by $\hat{v}_{it} = \sum_{j=1}^{m}\hat{\delta}_{i,j}^{E} + \hat{u}_{i,t}$ for (2) and so on for (3) and (4). $P_t = \frac{\hat{\lambda}_k}{s.e.(\hat{\lambda}_k)}$,$P_a = T\hat{\lambda}_k$.

5. The group-mean and panel tests are constructed in different ways and can therefore give different results. They require large N and large T data sets. These tests are also very sensitive to the specific choice of parameters such as lag and lead lengths, and the kernel width.
6. We also estimated the group-mean error-correction model, averaging coefficients of the error-correction equation over all cross-sectional units, together with the implied long-run relationship. However, these results are not reported here because the long-run coefficients were not significant.
7. MENA countries refers to Middle East and North African countries.

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No potential conflict of interest was reported by the author.

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Appendices

Appendix A. Descriptive Statistics

Table A1. Descriptive statistics of variables.

Variable	Unit	Variance	Mean	Std. dev.	Min	Max	Obs.
Full sample							
CO_2 emissions	m.t.	overall	0.931	1.452	0.000	18.390	N = 7350
		between		1.297	0.008	7.807	n = 150
		within		0.660	−4.138	13.362	T = 49
Log (GDP)	$	overall	8.222	1.298	4.522	11.637	N = 7348
		between		1.238	5.353	10.972	n = 150
		within		0.403	5.246	10.963	T-bar = 49
Int. trade	share	overall	0.716	0.520	−0.149	5.866	N = 7350
		between		0.446	0.020	3.096	n = 150
		within		0.270	−0.410	4.078	T = 49
OECD countries							
CO_2 emissions	m.t.	overall	2.392	1.535	0.140	11.050	N = 1470
		between		1.445	0.577	7.807	n = 30
		within		0.579	−0.705	5.635	T = 49
Log(GDP)	$	overall	9.745	0.617	7.498	11.406	N = 1470
		between		0.489	8.669	10.458	n = 30
		within		0.386	8.374	11.025	T = 49
Int. trade	share	overall	0.508	0.410	0.394	3.243	N = 1470
		between		0.355	0.154	2.098	n = 30
		within		0.214	−0.120	1.750	T = 49
Non-OECD countries							
CO_2 emissions	m.t.	overall	0.566	1.175	0.000	18.390	N = 5880
		between		0.963	0.008	5.959	n = 120
		within		0.679	−4.503	12.997	T = 49
Log(GDP)	$	overall	7.840	1.136	1.852	11.637	N = 5880
		between		1.061	5.353	10.972	n = 120
		within		0.417	1.548	10.581	T = 49
Int. Trade	share	overall	0.7628	0.508	1.035	4.432	N = 5880
		between		0.437	2.003	3.096	n = 120
		within		0.262	−36.286	3.670	T = 49

Overall refers to the whole data set. The total variation (around grand mean $\bar{x} = 1/NT \sum_i \sum_t x_{it}$) can be broken down into *within* variation over time for each individual country (around individual mean $\bar{x}_i = 1/NT \sum_t x_{it}$) and *between* variation across countries (for $\bar{x}$ around $\bar{x}_i$). The corresponding breakdown for the variance is *Within* variance: $s_W^2 = \frac{1}{NT-1} \sum_i \sum_t (x_{it} - \bar{x}_i)^2 = \frac{1}{NT-1} \sum_i \sum_t (x_{it} - \bar{x}_i + \bar{x})^2$;
Between variance: $s_B^2 = \frac{1}{N-1} \sum_i (x_i - \bar{x})^2$
Overall variance: $s_O^2 = \frac{1}{NT-1} \sum_i \sum_t (x_{it} - \bar{x})^2$
The second expression for s_w^2 is equivalent to the first, because adding a constant does not change the variance, and it is used at times because $x_{it} - \bar{x}_i + \bar{x}$ is centered on $\bar{x}$, providing a sense of scale, whereas $x_{it} - \bar{x}_i$ is centered on zero.

Appendix B. Westerlund's ECM-based Panel Cointegration Test

Cointegration is tested according to the following specifications:

$$E_{it} = \mu_i^E + \tau_i^E t + \delta_i^E Y_{it} + \gamma_i^E T_{it} + u_{it} \tag{A1}$$

$$Y_{it} = \alpha_i + \tau_i^Y t + \beta_i^Y E_{it} + \gamma_i^Y T_{it} + \varepsilon_{it} \tag{A2}$$

$$T_{it} = \nu_i + \tau_i^T t + \delta_i^T Y_{it} + \beta_i^T E_{i,t} + e_{it} \tag{A3}$$

Table B1. Results of Westerlund's ECM-based Panel Cointegration Tests.

		No constant no trend									With constant and trend								
		Full sample			OECD			Non-OECD			Full sample			OECD			Non-OECD		
Model	Test	Value of the test	Z-value	P-value	Value of the test	Z-value	P-value	Value of the test	Z-value	P-value	Value of the test	Z-value	P-value	Value of the test	Z-value	P-value	Value of the test	Z-value	P-value
Y	G_t	0.103	17.351	1.000	1.386	14.464	1.000	−0.218	12.165	1.000	−2.571	−0.622	0.267	−2.400	0.837	0.799	−2.627	−1.279	0.101
	G_a	−0.044	12.945	1.000	0.277	6.111	1.000	−0.124	11.418	1.000	−8.879	7.941	1.000	−7.341	4.701	1.000	−9.479	6.206	1.000
	P_t	0.653	9.128	1.000	8.344	10.141	1.000	−1.188	6.831	1.000	−25.231	3.089	0.999	−8.836	4.095	1.000	−111.265	−95.566	0.000
	P_a	0.020	6.294	1.000	0.317	3.147	0.999	−0.044	5.487	1.000	−8.538	3.532	1.000	−5.335	4.175	1.000	−37.829	−44.310	0.000
E	G_t	−1.233	1.738	0.959	−1.138	1.275	0.899	−1.255	1.324	0.907	−2.555	−0.393	0.347	−2.636	−0.702	0.241	−2.532	−0.053	0.479
	G_a	−4.054	3.965	1.000	−2.548	3.281	1.000	−4.427	2.798	0.997	−9.796	6.41	1.000	−6.935	5.004	1.000	−10.507	4.670	1.000
	P_t	−26.994	−11.678	0.000	−6.188	−0.795	0.213	−25.051	−11.128	0.000	−31.694	−4.076	0.000	−13.505	−1.081	0.140	−28.461	−3.771	0.000
	P_a	−6.093	−8.982	0.000	−2.314	0.207	0.582	−6.458	−8.850	0.000	−10.296	0.347	0.636	−8.312	1.762	0.961	−10.448	0.063	0.525
T	G_t	−1.186	2.289	0.989	−0.137	6.506	1.000	−1.245	1.433	0.924	−1.986	7.907	1.000	−1.866	4.315	1.000	−2.337	2.785	0.997
	G_a	−4.042	3.991	1.000	−0.613	5.219	1.000	−3.587	4.481	1.000	−5.385	13.778	1.000	−2.723	8.151	1.000	−7.851	9.659	1.000
	P_t	−20.061	−6.461	0.000	2.795	5.966	1.000	−17.568	−5.496	0.000	−22.468	6.152	1.000	−4.085	9.362	1.000	−24.850	3.511	1.000
	P_a	−4.713	−5.533	0.000	0.860	3.753	1.000	−4.265	−3.947	0.000	−7.332	5.717	1.000	−1.831	7.014	1.000	−8.699	3.241	0.999

Appendix C. List of countries

Afghanistan	Japan*
Albania	Jordan
Algeria	Kenya
Angola	Kiribati
Antigua and Barbuda	Korea, Republic of*
Argentina	Laos
Australia*	Lebanon
Austria*	Liberia
Bahamas	Luxembourg*
Bahrain	Macao
Bangladesh	Madagascar
Barbados	Malawi
Belgium*	Malaysia
Belize	Maldives
Benin	Mali
Bermuda	Malta
Bhutan	Mauritania
Bolivia	Mauritius
Botswana	Mexico*
Brazil	Mongolia
Brunei	Morocco
Bulgaria	Mozambique
Burkina Faso	Nepal
Burundi	Netherlands*
Cambodia	New Zealand*
Cameroon	Nicaragua
Canada*	Niger
Cape Verde	Nigeria
Central African Repub.	Norway*
Chad	Oman
Chile*	Pakistan
China	Panama
Colombia	Papua New Guinea
Comoros	Paraguay
Congo, Dem. Rep. of	Peru
Congo, Republic of	Philippines
Costa Rica	Poland*
Cote d'Ivoire	Portugal*
Cuba	Romania
Cyprus	Rwanda
Denmark*	Samoa
Djibouti	Sao Tome and Principe
Dominica	Senegal
Dominican Republic	Seychelles
Ecuador	Sierra Leone
Egypt	Singapore
El Salvador	Solomon Islands
Equatorial Guinea	Somalia
Ethiopia	South Africa
Fiji	Spain*
Finland*	Sri Lanka
France*	St. Kitts-Nevis
Gabon	St.Vincent & Grenadines
Gambia	Sudan
Germany*	Suriname
Ghana	Swaziland
Greece*	Sweden*
Grenada	Switzerland*
Guatemala	Syria
Guinea	Taiwan
Guinea Bissau	Thailand
Guyana	Togo
Haiti	Tonga
Honduras	Trinidad and Tobago
Hong Kong	Tunisia
Hungary*	Turkey*
Iceland*	Uganda
India	United Kingdom*
Indonesia	United States*
Iran	Uruguay
Iraq	Vanuatu
Ireland*	Venezuela
Israel*	Vietnam
Italy*	Zambia
Ivory Coast	Zimbabwe
Jamaica	

*Indicates OECD countries, the rest are non-OECD countries.