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The Nexus of CO2 emissions, energy consumption, economic growth, and trade-openness in WTO countries

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Discussion
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Energy Consumption, Economic Growth,
and Trade-Openness in WTO Countries

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IMPRESSUM

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The nexus of CO₂ emissions, energy consumption, economic growth, and trade-openness in WTO countries

Lars Sorge* and Anne Neumann**

Abstract: This paper analyzes the dynamic relationship between CO₂ emissions, energy consumption, GDP, and trade-openness from 1971 to 2013, based on the Environmental Kuznets Curve (EKC) hypothesis for 70 WTO countries. Using recently developed second-generation panel data methods, the empirical results support the EKC hypothesis for the high-, middle-, and lower-income panels used. Concerning the energy consumption and economic growth nexus, the causality results support the conversion hypothesis for the high-income panel, whereas the neutrality hypothesis holds for the lower- and middle-income panels. Based on the causality results, trade-openness does not positively impact CO₂ emissions, GDP leads CO₂ emissions, and trade-openness causes energy consumption within any income panel. The net effect of economic growth, however, could help to stabilize future CO₂ emissions within any income panel.

Keywords: Environmental Kuznets Curve, CO₂ emissions, energy consumption, economic growth, trade-openness, Granger causality, second-generation panel data methods

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1 Introduction

A substantial amount of greenhouse gas emissions result from the increased use of energy, primarily fossil fuels, as the main driver for production processes and economic growth. Given the global challenge of climate change, initiatives by individual countries are not sufficient to fight global warming. According to the Intergovernmental Panel on Climate Change (2014), continued emissions of carbon dioxide (CO₂) will lead to further surface warming and cause damage to the ecosystem that is likely to be irreversible. Moreover, the report emphasizes that while climate change affects all continents, the world's poorest countries remain the most vulnerable.

Global economic prosperity could provide a solution according to a theoretical concept formulated in 1991: the Environmental Kuznets Curve (EKC) hypothesis postulates that after a certain threshold income level is reached, environmental quality improves over the course of further economic development (Grossman and Krueger, 1991). Research analyzing the causal relationship between energy consumption and economic growth evolved after the oil price shocks in the 1970s and throughout the years prior to the Kyoto Protocol (Payne, 2010). Recently, an additional strand in the energy economics literature combines the energy consumption - economic growth nexus with the EKC hypothesis and incorporates measures of environmental pollutants (Zhang and Cheng, 2009). Expanding upon previously published energy economics studies, this paper uses the empirical results obtained from a dataset of 70 WTO countries clustered into high-, middle-, and lower-income economies from 1971 to 2013 to understand the determinants of CO₂ emissions. We focus on international trade within the EKC framework to formulate appropriate energy and environmental policies for the given income groups.

This paper makes three important contributions to the published literature. First, we incorporate CO₂ emissions as the EKC relationship appears to be valid only for pollutants that do not have long-term effects on a global scale with high abatement costs. Second, we apply recently developed nonstationary panel data methods to improve the quality of the results. Third, since the EKC-framework is routinely criticized for its failure to consider

international trade, we also control for trade-openness. Therefore, the aim of this paper is to examine the dynamic relationship between CO₂ emissions, economic growth, energy consumption, and trade-openness based on the EKC hypothesis for 70 WTO countries clustered into three income-related panels over the longest period possible. The empirical results obtained help to understand the determinants of carbon dioxide emissions with a special focus on international trade within an EKC framework, thus facilitating the development of energy and environmental policies in any given income group.

The remainder of this paper is organized as follows. Section 2 reviews the literature. Section 3 presents the data and explains the empirical strategy. Section 4 reports the empirical results. Section 5 concludes and offers suggestions for future research.

2 Background and Literature Review

According to Zhang and Cheng (2009), the analysis of the relationship between economic growth, energy consumption, and environmental pollutants can be clustered into three research strands in the energy economics literature. The first strand concerns the environmental pollutants and economic growth nexus. The literature in this strand tests the validity of the traditional EKC hypothesis, inspired by the seminal contributions of Grossman and Krueger (1991), Shafik and Bandyopadhyay (1992), Panayotou (1993), and Selden and Song (1994). Over the past two decades several authors tested the validity of the EKC hypothesis, yet the empirical results are still controversial, particularly because of the absence of a single environmental indicator (Dinda, 2004).

The second strand concerns the energy consumption and economic growth nexus. In theory, higher economic development requires more energy consumption and more efficient energy use requires a higher level of economic development (Halicioglu, 2009). The contribution of Kraft and Kraft (1978) was followed by comprehensive research investigating the causality relationship between energy consumption and economic growth. The empirical results conflict because of the different econometric approaches employed and the absence of greenhouse gas emissions measures.

The third strand concerns the dynamic relationship between economic growth, energy consumption, and environmental pollutants. Initiated by Ang (2007) and Soytaş et al. (2007) in a time series data framework, Apergis and Payne (2009) extend the approach of Ang (2007) to a panel data framework.

Our literature review surveys this third strand in light of the validity of the EKC hypothesis, based on extensive reviews of empirical studies analyzing the impact of energy consumption and economic growth on CO₂ emissions by Al-Mulali et al. (2015a), Al-Mulali et al. (2015b), Ozturk and Al-Mulali (2015), and Al-Mulali and Ozturk (2016). We find 32 studies that include the influence of trade as a regressor, but 25 of them comprise single country analyses within a time series framework, and thus are not comparable. Of the remaining seven studies, five do not explore the causal relationship among the variables or are not theoretically or empirically based on the EKC hypothesis. As a result, only two studies are directly comparable in terms of the econometric and theoretical approaches while including the influence of trade in the analysis.

Farhani et al. (2014) investigate the causal relationship between CO₂ emissions, energy consumption, economic growth, and trade-openness for a selection of 10 Middle East and North African (MENA) countries.¹ Energy consumption, GDP, its square, trade-openness, manufacture value added, and modified HDI index (MHDI) explain CO₂ emissions. The results of the panel EKC estimation obtained with both fully modified OLS (FMOLS) and dynamic OLS (DOLS) confirm an inverse U-shaped relationship between CO₂ emissions per capita and per capita real GDP. All estimated coefficients are statistically significant at the 1% level and the coefficient on trade-openness is positive. The estimated turning point of the panel EKC occurs for the FMOLS (DOLS) estimation at an income level of 10.37 (10.41) in logarithms. Farhani et al. (2014) also find bidirectional causality between all variables in both the short- and long-runs.

Kasman and Duman (2015) investigate the relationship between CO₂ emissions, energy consumption, income, and trade-openness in a panel data framework for new and can-

¹Algeria, Bahrain, Egypt, Iran, Jordan, Morocco, Oman, Saudi Arabia, Syria, and Tunisia, 1990 to 2010.

didate European member states from 1992 to 2010.² Energy consumption, GDP, its square, trade-openness, and the share of urban population applying second-generation nonstationary panel data methods that account for cross-sectional dependence explain CO₂ emissions. A first model uses energy consumption, GDP, and GDP² as independent variables and a second model adds trade-openness and share of urban population. The outcome based on the fully modified OLS (FMOLS) estimation technique supports the validity of the EKC hypothesis for both models. The estimated turning point for model 1 (model 2) occurs at 8.22 (8.20) in logarithms. The estimated coefficients on trade-openness as well as urbanization are statistically significant with a positive impact on CO₂ emissions. The estimated turning points of both models do not differ substantially. The results for the panel suggest that unidirectional causality runs from energy consumption, trade-openness, and urbanization to CO₂ emissions in the short-run. Kasman and Duman (2015) find unidirectional causality from GDP to energy consumption as well as from GDP, energy consumption, and urbanization to trade-openness in the short-run. The results provide evidence for short-run unidirectional causality from urbanization to GDP and from urbanization to trade-openness and long-run bidirectional causal relationships between CO₂ emissions, energy consumption, income, and trade-openness.

Based on the above studies, we find only two that include trade-openness as explanatory variables within a panel data framework. Our evaluation of the panel data related literature within the third strand reveals that an analysis of the link between CO₂ emissions, economic growth, energy consumption, and the influence of trade-openness based on the EKC hypothesis was conducted for a panel of MENA countries and newly EU member states for a relatively short time period.

A relatively new area of research that emerged in the 1990s develops the theory of nonstationary panel econometrics, by combining nonstationary time series data with the increased data and power resulting from adding a cross-sectional dimension. This improves the power of statistical inference (Baltagi and Kao, 2001). First-generation panel data

²The 15 new European Union (EU) member states (Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Malta, Poland, Romania, Slovak Republic, and Slovenia) and the three candidate countries (Iceland, Macedonia, and Turkey).

methods applied over the last two decades, however, assume cross-sectional independence, although the presence of cross-sectional dependence can produce misleading inferences (Chudnik and Pesaran, 2015). Only Kasman and Duman (2015) analyze the relationship in question with second-generation panel data methods that account for cross-sectional dependence.

3 Data and Methodology

This section presents the empirical strategy used in the analysis: Section 3.1 describes the utilized panel dataset and a graphical representation of the relevant variables. Section 3.2 outlines the empirical specification of the econometric model associated with the EKC hypothesis. Section 3.3 presents the test used to detect the potential presence of cross-section dependence. Section 3.4 investigates the stationarity properties of the variables with a second-generation panel unit root test. Section 3.5 describes the long-run relationship among the potentially integrated variables determined with a panel cointegration test that accounts for cross-sectional dependence. Section 3.6 describes the estimation technique used to obtain the long-run coefficients of the EKC after establishing the cointegrating relationship. Section 3.7 specifies a panel-based vector error correction model to investigate the causal relationship of the series in the short- and long-runs.

3.1 Data

We build a panel dataset using annual data from the World Bank Development Indicators (WDI) for 70 WTO countries from 1971 to 2013, clustering them into three panels³ based on their level of high, middle and lower incomes. The World Bank clusters economies into low, lower-middle, upper-middle, and high-income groups based on Gross National Income (GNI) per capita and corresponding thresholds established in 1989. These thresholds are updated annually at the beginning of the World Bank's fiscal year,

³Due to the lack of data availability for low-income countries, we include low in the lower-income panel

with an adjustment for inflation. Thus, movements of an economy between the four income groups can result from an increase or decrease of a countries' GNI per capita or due to the adjustment of the annually updated thresholds ⁴. With T=43 in each income panel for every given variable, the high-income panel ($N_{high}=29$) has 1,247 observations, the middle-income panel ($N_{middle}=17$) has 731 observations, and the lower-income panel ($N_{low}=24$) has 1,032 observations. The overall panel ($N=70$) covers WTO countries from the seven regions predefined by the World Bank: East Asia and Pacific ($n=10$); Europe and Central Asia ($n=17$); Latin America and Caribbean ($n=19$); Middle East and North Africa ($n=6$); North America ($n=2$); South Asia ($n=4$); and Sub-Saharan Africa ($n=12$). The high-income panel ($N_{high}=29$) has 16 WTO countries from Europe and Central Asia (55%), five countries from East Asia and Pacific (17%), three countries each from Latin America and Caribbean (10%) and the Middle East and North Africa (10%), and two countries from North America (7%). The middle-income panel ($N_{middle}=17$) has eleven WTO countries from Latin America and Caribbean (65%), three countries from East Asia and Pacific (18%), two countries from Sub-Saharan Africa (11%), and one country from Europe and Central Asia (6%). The lower-income panel ($N_{lower}=24$) has ten WTO countries from Sub-Saharan Africa (42%), five countries from Latin America and Caribbean (21%), four countries from South Asia (17%), three countries each from the Middle East and North Africa (13%), and two from East Asia and Pacific (8%).

Table 1: Summary statistics by panel

Panel	CO ₂		GDP		E		T	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
High-income	9.77	5.81	29,620.85	17,268.91	4,094.38	2,595.74	91.87	79.55
Middle-income	3.15	2.25	5,990.35	3,244.58	1,169.40	622.91	59.49	39.61
Lower-income	0.65	0.50	1,445.07	800.71	444.48	146.62	60.52	24.98

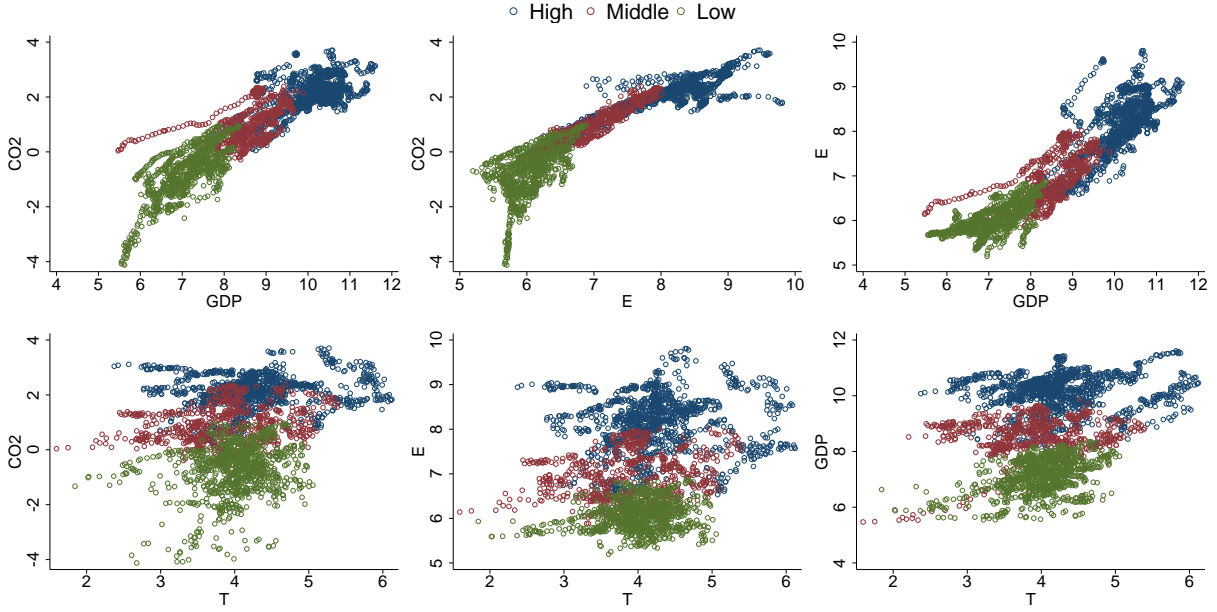
Notes: Mean is the arithmetic mean; standard deviation is denoted by SD; data obtained from the World Bank Development Indicators database (last updated 16 December 2016).

⁴For our panel dataset the following movements occurred over the period from 1987 to 2013: In total, 25 countries moved on average 2.12 times between our three defined income groups over the 27 year period for which the historical classification by income exist. Within our high income panel, movements for 8 economies occurred, whereas 16 countries changed the income group in our middle income panel, and Tunisia changed its income class from the lower-income panel to the middle-income group in 2010.

According to the World Bank's definition (2016), CO₂ emissions are measured in metric tons per capita, per capita real GDP is measured in constant 2010 USD, per capita energy consumption is measured in kg of oil equivalent, and trade-openness is defined as the sum of exports and imports of goods and services measured as a share of gross domestic product. Table 1 shows that the average per capita CO₂ emissions between 1971 and 2013 for the high-income panel are more than 300% (1,500%) higher than the per capita emissions of the middle- (lower-) income panel. Per capita CO₂ emissions for the middle-income panel are almost 500% higher compared to the lower-income panel. The difference in terms of per capita GDP is also remarkable: the average economic wealth expressed as per capita GDP for the high-income panel is nearly 500% and more than 2,000% higher than for the middle- and lower-income panel, respectively. The middle-income panel is on average 400% wealthier than the lower-income panel. As for average per capita energy consumption, the high-income panel emits more than 300% (900%) CO₂ emissions than the middle- (lower-) income panel. Per capita emissions of the middle-income panel are more than twice as high as the lower-income panel. The ratio of trade to GDP for the high-income panel is only about 1.5 times higher than for the middle- and lower-income panels; surprisingly, the trade-openness indicator is almost identical for the latter two panels.

Figure 1 illustrates the relationship between per capita CO₂ emissions, per capita energy consumption, GDP per capita, and trade-openness for each income panel from 1971 to 2013. The top left figure shows a positive relationship between GDP per capita and CO₂ emissions per capita for the overall panel and within every panel. An inverted U-shaped relationship postulated by the EKC hypothesis is not clearly visible for any income panel. The top middle figure shows a positive relationship between energy consumption per capita and CO₂ emissions per capita for every income group. However, the slope for the data points varies substantially between the groups: while for the lower income panel, an increase in energy consumption is accompanied by a very sharp increase in per capita emissions, for both the middle- and high-income panels, an increase in per capita energy consumption results in a less pronounced raise in per capita CO₂ emissions. The top right

Figure 1: Scatter plots of CO₂, GDP, energy consumption, and trade-openness by panel



Notes: All possible variables converted into natural logarithms.

figure reveals the opposite for the relationship between GDP and energy demand. For the high-income panel, an increase in GDP growth results in a very steep increase in demand for energy. In contrast, for the lower income panel an increasing income results in less intensively increasing energy consumption per capita. The three bottom plots of the figure show the influence of trade-openness on CO₂ emissions per capita, energy consumption per capita, and GDP per capita. The bottom right scatter plot is the most interesting; it shows a positive relationship between trade-openness and income in every income group, with the largest positive impact of international trade on GDP for the lower-income panel. The other two scatter plots at the bottom are displayed for completeness, i.e. it is not possible to identify the impact of trade-openness on emissions or energy consumption based on visual inspection.

3.2 Empirical specification

Our panel data framework follows the econometric approach in Apergis and Payne (2009) whilst avoiding the problem of misspecification (Ang, 2007). Based on the EKC hypothesis in its simplest form, we specify the long-run relationship between the relevant

variables in baseline model 1:

$$CO_{2it} = \alpha_i + \beta_{1i}GDP_{it} + \beta_{2i}GDP_{it}^2 + \beta_{3i}E_{it} + \epsilon_{it}, \quad (1)$$

where $i = 1, \dots, N$ denotes each country in the panel, subscript $t = 1, \dots, T$ refers to the time period measured in years, CO_2 is CO_2 emissions per capita, GDP is GDP per capita, GDP^2 is the square of per capita GDP, E is per capita energy consumption, and ϵ_{it} represents the idiosyncratic error term. Converting all possible variables into natural logarithms, we interpret the coefficients on β_1 , β_2 , and β_3 as the proportion of the percentage changes in one variable to the percentage changes in another variable, i.e. the partial long-run elasticity estimates of per capita CO_2 emissions with respect to GDP per capita, squared GDP per capita, and energy consumption per capita, respectively. With the variables transformed into their natural logarithms, their first differences approximate their growth rates. The sign of the coefficient on β_1 is expected to be positive and negative for β_2 if there is an inverted U-shaped relationship as suggested by the EKC hypothesis. Since an increase in energy consumption yields an increase in CO_2 emissions, the expected sign of the coefficient on β_3 is positive for all income panels.⁵

The panel-based literature often omits trade-openness as an explanatory variable, Arrow et al. (1995), Stern et al. (1996), and Cole (2004), however, theoretically justify controlling for international trade in an EKC regression. According to trade theory, developing countries start to specialize in the production of labor and resource-intensive goods, as these countries become comparatively well-endowed in this input, whereas developed and wealthy countries specialize in the production of capital-intensive goods. Thus, the production structure of developing countries is more environmentally harmful compared to wealthy countries, with lesser-developed countries as net exporters of pollution-intensive goods and wealthier countries as net importers. Following the pollution haven hypothesis introduced by Copeland and Taylor (1994), theoretically it implies that present low-income

⁵The existing literature largely ignores the potential multicollinearity among variables. Including the squared GDP per capita term is necessary to test the diminishing effect on CO_2 according to the EKC theory. A major advantage of panel data over both time series and cross-section data is less multicollinearity among the variables (Baltagi, 2013).

countries will be unable to outsource their resource-intensive manufacturing industries to other countries with lower environmental standards. Furthermore, a higher degree of trade-openness as an essential component for sustained economic growth increases the size of the economies of the lower-income countries with the increased demand for energy that in turn increases pollution (Sadorsky, 2012). Therefore, we build a second model containing trade-openness to investigate the impact of international trade on CO₂ emissions within every income group:

$$CO_{2it} = \alpha_i + \beta_{1i}GDP_{it} + \beta_{2i}GDP_{it}^2 + \beta_{3i}E_{it} + \beta_{4i}T_{it} + \epsilon_{it}, \quad (2)$$

where T denotes trade-openness. The sign of β_4 is expected to be negative for the high- and middle-income panels, whereas β_4 is expected to be positive for the lower-income panel.

3.3 Cross-section dependence test

Cross-section dependence is defined as the contemporaneous correlation among countries after controlling for individual characteristics (Moscone and Tosetti, 2009). The interdependencies across sections in panel data may result from unobserved global shocks and local interactions. First-generation panel data methods commonly used in empirical applications of nonstationary panel econometrics, however, assume cross-section independence. Pesaran (2007) shows that first-generation panel unit root tests have substantial size distortions when the assumption of cross-section independence does not hold, which could lead to an over-rejection of the null hypothesis.

The cross-section dependence (CD) test suggested by Pesaran (2004) is suitable for stationary and unit root dynamic heterogeneous panels with short T and large N under a variety of panel data models. The CD test relies on an average of pairwise correlation coefficients of OLS residuals from the individual regressions in the panel. The CD test

statistic is defined as:

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^N \hat{\rho}_{ij} \right), \quad (3)$$

where $\hat{\rho}_{ij}$ is the estimated pairwise correlation coefficient that includes the estimated residuals $\hat{\epsilon}_{it}$ from a standard panel-data model $y_{it} = \alpha_i + \beta_i X_{it} + \epsilon_{it}$ needed to calculate the CD statistics:

$$\hat{\rho}_{ij} = \hat{\rho}_{ji} = \frac{\sum_{t=1}^T \hat{\epsilon}_{it} \hat{\epsilon}_{jt}}{\left(\sum_{t=1}^T \hat{\epsilon}_{it}^2 \right)^{1/2} \left(\sum_{t=1}^T \hat{\epsilon}_{jt}^2 \right)^{1/2}}. \quad (4)$$

Under the null hypothesis of zero dependence across panel members, the CD test statistics are distributed as standard normal for N sufficiently large. Moreover, the test is robust to the presence of nonstationary processes, parameter heterogeneity, or structural breaks, and it performs well even for small samples (Baltagi, 2013).

3.4 Second-generation panel unit root test

Second-generation panel unit root tests release the assumption that the individual time series in the panel are cross-sectionally independently distributed. Pesaran (2007) proposes a second-generation panel unit root test based on augmenting the ADF regressions with the cross-section average of lagged levels and first differences of the individual series. Thus, the cross-sectional dependence arising from a single common factor is filtered out. The homogeneous null hypothesis states that each country in the panel contains a unit root and is being tested against the heterogeneous alternative that allows to differ across countries. For the more general case of serially correlated disturbances, the cross-sectionally augmented Dickey-Fuller (CADF) regression is given by:

$$\Delta y_{it} = \alpha_i + \rho_i y_{i,t-1} + c_i \bar{y}_{t-1} + \sum_{j=0}^J d_{ij} \Delta \bar{y}_{t-j} + \sum_{j=1}^J \beta_{ij} \Delta y_{i,t-j} + \epsilon_{it}. \quad (5)$$

To account for serial correlation in the residuals, we apply the usual ADF regression for each i , and also include $\bar{y}_{t-1} = N^{-1} \sum_{i=1}^N y_{it-1}$ and $\Delta \bar{y}_t = N^{-1} \sum_{i=1}^N \Delta y_{it}$. We consider a cross-sectionally augmented version Pesaran (2007) of the standardized Im et al. (2003)

test based on the average of individual CADF statistics:

$$CIPS(N,T) = \bar{t} = N^{-1} \sum_{i=1}^N CADF_i = N^{-1} \sum_{i=1}^N t_i(N,T), \quad (6)$$

where $t_i(N,T)$ is the CADF statistic for the i -th individual in the panel which is given by the estimated t -ratio of ρ_i obtained from the above CADF regression. Simulation results indicate that the Pesaran (2007) panel unit root test has satisfactory size and power even for a relatively small number of cross-section units and for small values of the time dimension (Baltagi, 2013).

3.5 Second-generation panel cointegration test

Westerlund (2007) implements four new error correction-based panel cointegration tests to test the null hypothesis of no cointegration. The tests are based on structural rather than residual dynamics which avoids the problem of a common factor restriction. The general idea is to test the null hypothesis by inferring whether or not the error correction term in a conditional error correction model is equal to zero. The rejection of the null hypothesis of no error correction implies a rejection of the null hypothesis of no cointegration. To accommodate for cross-sectional dependence, using bootstrapped robust critical values for the four panel cointegration test statistics prevents statistical misspecification. Therefore, we construct four test statistics by using least square estimates of the following conditional error correction model:

$$\Delta y_{it} = \delta_i^i d_t + \alpha_i (y_{i,t-1} - \beta_i^i X_{i,t-1}) + \sum_{j=1}^{J_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{J_i} \gamma_{ij} \Delta X_{i,t-j} + \epsilon_{it}, \quad (7)$$

where $i = 1, \dots, N$ stands for each unit in the panel, subscript $t = 1, \dots, T$ refers to time, and d_t contains the deterministic components being either fixed effects, or fixed effects and linear time trends. X_{it} describes the set of non-strictly exogenous regressors included in the model. The data-generating process is general enough to allow for unit-specific short-run dynamics for the variables in their first differences and heterogeneous slope

parameters. To estimate the error correction term we reparameterize equation (7) as:

$$\Delta y_{it} = \delta_i^c d_t + \alpha_i y_{i,t-1} + \lambda_i^c X_{i,t-1} + \sum_{j=1}^{J_i} \alpha_{ij} \Delta y_{i,t-j} + \sum_{j=0}^{J_i} \gamma_{ij} \Delta X_{i,t-j} + \epsilon_{it}, \quad (8)$$

where $\lambda_i^c = -\alpha_i \beta_i^c$ and thus the least square estimate of α_i can be used to provide a valid test of the null hypothesis against the alternative hypothesis, and parameter α_i is the error correction or speed of adjustment term determining the amount of time at which the system converges back to the long-run equilibrium relationship $y_{i,t-1} - \beta_i^c X_{i,t-1}$ after a sudden shock.

Next, we group the four test statistics into two categories: panel statistics P_τ and P_α are based on pooling information about the error correction along the cross-sectional dimension of the panel, and group mean statistics G_τ and G_α do not pool information. The null hypothesis is the same for each statistic and the alternative hypothesis differs between the two categories. For the panel statistics, we test the null hypothesis $H_0 : \alpha_i = 0$ for all i against the alternative hypothesis $H_1 : \alpha_i = \alpha < 0$ for all $i = 1, \dots, N$. A rejection of the null hypothesis indicates cointegration of the panel as a whole. For the group mean statistics, we test the null hypothesis $H_0 : \alpha_i = 0$ for all i against the alternative hypothesis $H_1 : \alpha_i < 0$ for all $i = 1, \dots, N$. A rejection of the null hypothesis indicates cointegration for at least one of the cross-sectional members.

3.6 Panel long-run estimation

Pedroni (2001) proposes a dynamic OLS (DOLS) estimator that can be applied to nonstationary cointegrated panels. The DOLS estimation technique adds lags and leads of the differenced explanatory variables into the cointegrating regression in order to control for endogeneity due to the correlation between the regressors and the disturbances, and to obtain consistent and efficient estimators of the long-run relationship. We base the group mean panel DOLS estimator suggested by Pedroni (2001) on the following augmented

regression equation:

$$y_{it} = \alpha_i + \beta_i X_{it} + \sum_{j=-j_i}^{J_i} \gamma_{ij} \Delta X_{i,t-j} + \epsilon_{it}, \quad (9)$$

where $i = 1, \dots, N$ stands for each unit in the panel, subscript $t = 1, \dots, T$ refers to time, $j = 1, \dots, J$ is the number of individual lags and leads in the DOLS regression, and X_{it} describes the set of regressors included in the model. The slope coefficients β_i are permitted to vary across panel members. From equation (8), the group mean panel DOLS estimator $\hat{\beta}_{GM}^*$ averages over the individual specific time series DOLS estimates. Now, we aggregate the individual coefficient estimates and the corresponding t -statistics to obtain the DOLS estimates for the panel as a whole, i.e. the information is pooled along the between-dimension over the entire panel. The between-dimension estimator allows for more flexibility in the presence of heterogeneity of the cointegrating vectors (Pedroni, 2001).

3.7 Panel causality test

For any cointegrating relation an error correction representation exists (Engle and Granger, 1987). Thus, if the variables are integrated of the same order and cointegrated, a vector error correction model (VECM) is appropriate for modeling the system. The representation of a VECM allows us to identify a particular form of causal relationship among the variables for integrated data. The definition of causality based on Granger (1969, 1980) is that a change in one variable occurs before changes in another variable with the underlying assumption that past and present variables cause the future, but that the future cannot cause the present. Following Pesaran et al. (1999), we model the panel VECM equations for the four (model 1) and five (model 2) variable cases with one hypothesized cointegrated relationship:

Model 1:

$$\begin{array}{l}
 (10) \\
 (11) \\
 (12) \\
 (13)
 \end{array}
 \begin{bmatrix}
 \Delta CO_{2it} \\
 \Delta E_{it} \\
 \Delta GDP_{it} \\
 \Delta GDP_{it}^2
 \end{bmatrix}
 =
 \begin{bmatrix}
 \alpha_{1i} \\
 \alpha_{2i} \\
 \alpha_{3i} \\
 \alpha_{4i}
 \end{bmatrix}
 +
 \sum_{j=1}^J
 \begin{bmatrix}
 \beta_{11ij} & \beta_{12ij} & \beta_{13ij} & \beta_{14ij} \\
 \beta_{21ij} & \beta_{22ij} & \beta_{23ij} & \beta_{24ij} \\
 \beta_{31ij} & \beta_{32ij} & \beta_{33ij} & \beta_{34ij} \\
 \beta_{41ij} & \beta_{42ij} & \beta_{43ij} & \beta_{44ij}
 \end{bmatrix}$$

$$\begin{bmatrix}
 \Delta CO_{2i,t-j} \\
 \Delta E_{i,t-j} \\
 \Delta GDP_{i,t-j} \\
 \Delta GDP_{i,t-j}^2
 \end{bmatrix}
 +
 \begin{bmatrix}
 \lambda_{1i} \\
 \lambda_{2i} \\
 \lambda_{3i} \\
 \lambda_{4i}
 \end{bmatrix}
 ECT_{i,t-1}
 +
 \begin{bmatrix}
 \epsilon_{1i} \\
 \epsilon_{2i} \\
 \epsilon_{3i} \\
 \epsilon_{4i}
 \end{bmatrix}$$

Model 2:

$$\begin{array}{l}
 (14) \\
 (15) \\
 (16) \\
 (17) \\
 (18)
 \end{array}
 \begin{bmatrix}
 \Delta CO_{2it} \\
 \Delta E_{it} \\
 \Delta GDP_{it} \\
 \Delta GDP_{it}^2 \\
 \Delta T_{it}
 \end{bmatrix}
 =
 \begin{bmatrix}
 \alpha_{1i} \\
 \alpha_{2i} \\
 \alpha_{3i} \\
 \alpha_{4i} \\
 \alpha_{5i}
 \end{bmatrix}
 +
 \sum_{j=1}^J
 \begin{bmatrix}
 \beta_{11ij} & \beta_{12ij} & \beta_{13ij} & \beta_{14ij} & \beta_{15ij} \\
 \beta_{21ij} & \beta_{22ij} & \beta_{23ij} & \beta_{24ij} & \beta_{25ij} \\
 \beta_{31ij} & \beta_{32ij} & \beta_{33ij} & \beta_{34ij} & \beta_{35ij} \\
 \beta_{41ij} & \beta_{42ij} & \beta_{43ij} & \beta_{44ij} & \beta_{45ij} \\
 \beta_{51ij} & \beta_{52ij} & \beta_{53ij} & \beta_{54ij} & \beta_{55ij}
 \end{bmatrix}$$

$$\begin{bmatrix}
 \Delta CO_{2i,t-j} \\
 \Delta E_{i,t-j} \\
 \Delta GDP_{i,t-j} \\
 \Delta GDP_{i,t-j}^2 \\
 \Delta T_{i,t-j}
 \end{bmatrix}
 +
 \begin{bmatrix}
 \lambda_{1i} \\
 \lambda_{2i} \\
 \lambda_{3i} \\
 \lambda_{4i} \\
 \lambda_{5i}
 \end{bmatrix}
 ECT_{i,t-1}
 +
 \begin{bmatrix}
 \epsilon_{1i} \\
 \epsilon_{2i} \\
 \epsilon_{3i} \\
 \epsilon_{4i} \\
 \epsilon_{5i}
 \end{bmatrix}$$

where $i = 1, \dots, N$ stands for each country in the panel, subscript $t = 1, \dots, T$ refers to time, expression Δ denotes the first difference operator, $j = 1, \dots, J$ is the number of lags, parameter α_{ki} are fixed effects, and ϵ_{kit} is the idiosyncratic error term for the system of equations. The one period lagged error correction term ECT_{t-1} derived from the

normalized long-run cointegrating relationship of the four variables (model 1) and the five variables (model 2) represents the residuals from the following estimated long-run regressions:

Model 1:

$$ECT_{t-1} = CO_{2t-1} - \beta_1 GDP_{t-1} - \beta_2 GDP_{t-1}^2 - \beta_3 E_{t-1}, \quad (19)$$

Model 2:

$$ECT_{t-1} = CO_{2t-1} - \beta_1 GDP_{t-1} - \beta_2 GDP_{t-1}^2 - \beta_3 E_{t-1} - \beta_4 T_{t-1}. \quad (20)$$

The coefficient λ_{ki} on the error correction term for the system of equations determines the speed of adjustment toward the long-run cointegrating relationship and the coefficients on the lagged differenced variables defines the short-run adjustment dynamics. To estimate the short- and long-run parameters and to examine the causal relationship in both models, we use the pooled mean group (PMG) estimator proposed by Pesaran et al. (1999) for panel-based VECM. The PMG estimator allows for individual specific intercepts and also allows the short-run dynamics to differ across countries, whereas the long-run coefficients of the cointegrating relationship are assumed to be homogeneous. The PMG estimator is an intermediate estimator that is based on a combination of pooling and averaging of the coefficients (Pesaran et al., 1999).

We determine two sources of causality by testing the significance on the coefficients in equations (10) to (13) for model 1 and equations (14) to (18) for model 2 from the panel-based VECM, i.e. we test short-run causality by inferring the significance on the coefficients of the lagged dynamic terms and test long-run causality by inferring the significance on the coefficients of the ECT. With respect to short-run causality in equations (10) and (14), causality runs from ΔE to ΔCO_2 if the null hypothesis $H_0 : \beta_{12ij} = 0, \forall_{ij}$ is rejected, whereas causality runs from ΔGDP and ΔGDP^2 to ΔCO_2 if the joint null hypothesis $H_0 : \beta_{13ij} = \beta_{14ij} = 0, \forall_{ij}$ is rejected in both models. Additionally, in equation (14) of model 2, short-run causality runs from ΔT to ΔCO_2 if the null hypothesis $H_0 : \beta_{15ij} = 0, \forall_{ij}$ is rejected in model 2. In equations (11) and (15), causality runs from

ΔCO_2 to ΔE if the null hypothesis $H_0 : \beta_{21ij} = 0, \forall ij$ is rejected, whereas causality runs from ΔGDP and ΔGDP^2 to ΔCO_2 if the joint null hypothesis $H_0 : \beta_{23ij} = \beta_{24ij} = 0, \forall ij$ is rejected. Additionally, short-run causality runs from ΔT to ΔE , if the null hypothesis $H_0 : \beta_{25ij} = 0, \forall ij$ is rejected for equation (15) in model 2. To identify the causality from ΔCO_2 , ΔE or ΔT to economic growth (ΔGDP and ΔGDP^2), we use cross-equation restrictions on equations (12) and (13) for model 1 and equations (16) and (17) for model 2, i.e. short-run causality runs from ΔCO_2 to output growth if the null hypotheses $H_0 : \beta_{31ij} = 0, \forall ij$ and $H_0 : \beta_{41ij} = 0, \forall ij$ are rejected, whereas causality runs from ΔE to economic growth if the null hypotheses $H_0 : \beta_{32ij} = 0, \forall ij$ and $H_0 : \beta_{42ij} = 0, \forall ij$ are rejected. In equations (16) and (17) for model 2, short-run causality runs from ΔT to output growth if the null hypotheses $H_0 : \beta_{35ij} = 0, \forall ij$ and $H_0 : \beta_{45ij} = 0, \forall ij$ are rejected (Apergis and Payne, 2009). In equation (18) of model 2, short-run causality runs from ΔCO_2 to ΔT if the null hypothesis $H_0 : \beta_{51ij} = 0, \forall ij$ is rejected, from ΔE to ΔT if the null hypothesis $H_0 : \beta_{52ij} = 0, \forall ij$ is rejected, and from ΔGDP and ΔGDP^2 to ΔT if the joint null hypothesis $H_0 : \beta_{53ij} = \beta_{54ij} = 0, \forall ij$ is rejected. According to Masih and Masih (1996) and Asafu-Adjaye (2000), short-run causality can be interpreted in the broadest sense as the dependent variable responds only to short-term shocks to the stochastic environment.

We find long-run causality in equations (10) and (11) for model 1 and in equations (14) and (15) for model 2 if the null hypotheses $H_0 : \lambda_{1i} = 0, \forall ij$ and $H_0 : \lambda_{2i} = 0, \forall ij$ are rejected. For the economic growth equations (12) and (13) for model 1 and equations (16) and (17) for model 2, the rejection of both $H_0 : \lambda_{3i} = 0$ and $H_0 : \lambda_{4i} = 0$ indicates long-run causality. Finally, we find long-run causality in trade equation (18) of model 2 if the null hypothesis $H_0 : \lambda_{5i} = 0, \forall ij$ is rejected. If the null hypothesis is rejected, then the dependent variable responds to deviations from the long-run equilibrium, since the coefficient on the speed of adjustment shows how fast the deviations from the long-run equilibrium are eliminated (Pao and Tsai, 2010).

4 Empirical Results

In a first step we test our data for any degree of cross-section dependency. Table 2, which displays the results of the Pesaran (2004) CD test for each variable, shows the average absolute correlation coefficients between the time-series for each country in all income groups. We compute the CD statistic based on the residuals under both a fixed and random effects specification to evaluate the amount of cross-section dependence in the disturbances.

Table 2: Results of Pesaran (2004) CD tests

		Variables in levels					Residuals	
		CO_2	GDP	GDP^2	E	T	\hat{e}^{FE}	\hat{e}^{RE}
High income	abs (corr)	0.51	0.90	0.90	0.67	0.64	0.41	0.40
	CD statistic	11.93 ^a (0.00)	111.29 ^a (0.00)	111.43 ^a (0.00)	69.09 ^a (0.00)	70.64 (0.00)	10.05 ^a (0.00)	9.84 ^a (0.00)
Middle income	abs (corr)	0.49	0.65	0.66	0.60	0.50	0.33	0.34
	CD statistic	24.14 ^a (0.00)	42.30 ^a (0.00)	42.74 ^a (0.00)	21.76 ^a (0.00)	20.84 (0.00)	1.94 ^c (0.06)	1.55 (0.12)
Lower income	abs (corr)	0.55	0.59	0.59	0.55	0.35	0.32	0.32
	CD statistic	29.86 ^a (0.00)	32.30 ^a (0.00)	32.79 ^a (0.00)	27.37 ^a (0.00)	29.61 ^a (0.00)	1.17 (0.24)	1.22 (0.22)

Notes: P-values are in brackets; superscripts *a*, *b*, and *c* represent significance at 1%, 5%, and 10%, respectively; \hat{e}^{FE} and \hat{e}^{RE} are based on a fixed effects and random effects estimation of the following model with all variables in natural logarithms: $CO2_{it} = \alpha_i + \beta_1 GDP_{it} + \beta_2 GDP_{it}^2 + \beta_3 E_{it} + \beta_4 T_{it} + \epsilon_{it}$.

For the high-income panel, the average correlation varies considerably across the variables, from 0.52 in the case of CO_2 to 0.90, or almost perfect correlation, in GDP. The average absolute correlation in the residuals obtained from the fixed effects and random effects specification is 0.41 and 0.40, respectively. The null hypothesis of no cross-sectional dependence can be rejected in all cases for every conventional significance level. For the middle-income panel, the average correlation is distributed more homogeneously across all variables. The null hypothesis can be rejected at the 1% significance level for all variables. Average absolute correlation in the residuals is lower than in the high-income panel. We fail to reject the null hypothesis of no cross-sectional dependence for the CD statistic,

based on the residuals under a random effects specification. For the lower-income panel, there is the least amount of cross-section dependence in the variables, but again the null hypothesis of no cross-sectional dependence can be rejected for each of the variables. The amount of cross-section dependence in the disturbances is not statistically significant. Altogether, the results of the CD tests indicate that the variables are highly dependent across the countries in every income group.

Having established strong evidence in favor of cross-sectional dependencies in the data, it is reasonable to employ a panel unit root and a panel cointegration test, both of which account for cross-section dependence. Table 3 displays the results of the Pesaran (2007) panel unit root tests, which include both an intercept only and an intercept and linear trend. The outcome of the Pesaran (2007) panel unit root test differs among the three income groups.

Table 3: Results of Pesaran (2007) panel unit root tests

	High income		Middle income		Lower income	
	No Trend	Trend	No Trend	Trend	No Trend	Trend
CO_2	-1.848	-2.299	-1.603	-2.180	-1.906	-2.407
GDP	-1.864	-2.076	-1.736	-2.765 ^a	-1.740	-2.713 ^b
GDP^2	-1.858	-2.062	-1.683	-2.717 ^b	-1.700	-2.724 ^b
E	-2.395 ^a	-2.738 ^a	-1.750	-2.373	-1.720	-2.128
T	-2.237 ^a	-2.755 ^a	-1.929	-2.027	-2.058 ^c	-2.273
ΔCO_2	-5.805 ^a	-6.064 ^a	-5.661 ^a	-5.739 ^a	-5.961 ^a	-6.100 ^a
ΔGDP	-4.361 ^a	-4.628 ^a	-4.798 ^a	-4.889 ^a	-4.735 ^a	-5.149 ^a
ΔGDP^2	-4.383 ^a	-4.629 ^a	-4.762 ^a	-4.881 ^a	-4.724 ^a	-5.070 ^a
ΔE	-5.803 ^a	-6.083 ^a	-5.559 ^a	-5.818 ^a	-5.739 ^a	-5.937 ^a
ΔT	-5.312 ^a	-5.395 ^a	-5.192 ^a	-5.196 ^a	-5.902 ^a	-6.080 ^a

Notes: P-values are in brackets; superscripts *a*, *b*, and *c* represent significance at 1%, 5%, and 10%, respectively; critical values are from Pesaran (2007).

For the high-income panel considering both specifications, the null hypothesis is rejected only for the series in levels of per capita energy consumption and trade-openness. For the middle-income panel, the null hypothesis of non-stationarity is rejected only for the series in levels of per capita GDP and its square with a linear trend. For the lower-income panel, the null hypothesis is rejected at the 10% significance level for trade-openness in levels without a linear trend. Similar to the middle-income panel, the null hypothesis of non-stationarity is rejected only for levels of per capita GDP and its square with a

linear trend. The results for the series in first differences indicate that all variables are integrated of order one.⁶

Having established the panel stationarity properties for each series, the next step in the analysis is to apply Westerlund (2007) second-generation panel cointegration tests to test for the existence of a long-run equilibrium relationship for model 1 and model 2. Table 4 shows the results of the four error correction-based panel cointegration tests without any deterministic, with an intercept only, and with an intercept and linear trend specification by income group.⁷

⁶We note that the Pesaran (2007) panel unit root test assumes the existence of one common factor that affects all countries in the same way. This restrictive assumption should be considered when interpreting the results.

⁷We use the Akaike information criterion to determine the optimal lag length within the given limits for each separate time series and use bootstrapped robust critical values to account for the presence of cross-sectional dependence.

Table 4: Results of Westerlund (2007) panel cointegration tests

		No Deterministic		No Trend		Trend		
		Z-Value	P-Value	Z-Value	P-Value	Z-Value	P-Value	
High income	Model 1:							
	G_τ	-2.121 ^c	0.09	-3.025 ^c	0.06	-4.745 ^a	0.01	
	G_α	-0.404 ^b	0.05	-0.438 ^b	0.04	0.817 ^b	0.05	
	P_τ	-4.574 ^a	0.01	-4.884 ^b	0.02	-4.004 ^c	0.09	
	P_α	-4.005 ^a	0.01	-2.732	0.11	0.459	0.30	
	Model 2:							
	G_τ	-2.095 ^b	0.05	-2.040	0.17	-3.687 ^c	0.06	
	G_α	0.911	0.18	0.874 ^c	0.08	2.231	0.14	
	P_τ	-3.859 ^b	0.04	-4.065 ^b	0.05	-3.336	0.12	
	P_α	-2.204 ^c	0.09	-1.372	0.16	1.191	0.42	
	Middle income	Model 1:						
		G_τ	-3.887 ^a	0.00	-4.142 ^b	0.02	-6.603 ^a	0.01
G_α		-1.761 ^a	0.00	-1.611 ^a	0.00	-1.080 ^a	0.00	
P_τ		-5.347 ^a	0.00	-5.542 ^a	0.01	-8.256 ^a	0.01	
P_α		-5.162 ^a	0.00	-4.694 ^c	0.06	-4.318 ^b	0.02	
Model 2:								
G_τ		-3.679 ^a	0.01	-3.331 ^c	0.08	-4.971 ^b	0.05	
G_α		-1.444 ^a	0.00	-0.926 ^b	0.03	0.383 ^b	0.05	
P_τ		-5.933 ^b	0.03	-5.977 ^c	0.09	-7.174 ^b	0.04	
P_α		-5.857 ^a	0.00	-4.528 ^c	0.08	-3.406 ^b	0.04	
Lower income		Model 1:						
		G_τ	-3.377 ^b	0.03	-4.196 ^a	0.01	-4.479	0.16
	G_α	-0.396	0.15	-0.576 ^c	0.06	2.162	0.75	
	P_τ	-5.238 ^a	0.00	-4.229 ^b	0.04	-3.515 ^c	0.10	
	P_α	-5.236 ^a	0.00	-3.320 ^c	0.06	-1.080	0.14	
	Model 2:							
	G_τ	-4.352 ^b	0.04	-5.390 ^c	0.06	-4.156	0.22	
	G_α	-0.611 ^b	0.05	-0.067 ^c	0.07	3.185	0.81	
	P_τ	-5.683 ^b	0.05	-4.741 ^c	0.06	-2.925	0.56	
	P_α	-4.136 ^c	0.06	-2.379	0.11	0.335	0.61	

Notes: Superscripts *a*, *b*, and *c* represent significance at 1%, 5%, and 10%, respectively; number of replications to obtain bootstrapped p-values is set to 100; bandwidth is selected according to the data depending rule $4(T/100)^{2/9} \approx 3$ recommended by Newey and West (1994); Barlett is used as the spectral estimation method.

Again, the results vary across the income panels and are sensitive to the inclusion of a deterministic trend in the cointegrating regression. For the high-income panel with no deterministic included in the cointegrating regression, four out of four (model 1) and three out of four (model 2) tests indicate that the null hypothesis of no error correction, that is no cointegration, is rejected at every conventional significance level. Including a constant leads to a rejection of the null hypothesis in three out of four (model 1) and two out of

four (model 2) tests. For the specification with an intercept and linear trend, three out of four (model 1) and one out of four (model 2) tests show evidence of cointegration. The results for the middle-income panel are not sensitive to the deterministic included in the cointegrating regression. For both model 1 and model 2, for all tests independent of the deterministic specification and test statistic, the null hypothesis of no error correction is rejected. For the lower-income panel with no deterministic included in the cointegrating regression, three out of four (model 1) and four out of four (model 2) tests indicate that the null hypothesis of no error correction is rejected at every conventional significance level. The inclusion of a constant leads to a rejection of the null hypothesis in four out of four (model 1) and three out of four (model 2) tests. For the specification with an intercept and linear trend, the rejection of the null hypothesis is only possible in model 1 for the P_τ statistic at the 10% significance level. The results for model 2 do not indicate cointegration with a linear trend included.

Westerlund (2007) shows that the P_τ and G_τ statistics have the smallest size distortions when cross-sectional dependence is present. Against this background, the results indicate that all variables in model 1 and model 2 are panel cointegrated in each of the income panels. Thus, the overall outcome suggests the existence of a long-run equilibrium relationship for both model 1 and model 2 among the integrated variables when no linear trend is included within every income panel. We now turn to estimating the long-term relationship using the DOLS approach. Essentially, we test the validity of the EKC hypothesis and the corresponding partial long-run elasticities. To account for simple structures of cross-sectional dependency, we remove the time trend in the form of $\bar{y}_t = N^{-1} \sum_{i=1}^N y_{it}$. Table 5 displays the results.

All coefficients are statistically significant at the 1% level, which supports the inverted U-shape pattern associated with the EKC hypothesis for both models in any income group. CO₂ emissions per capita increase with GDP per capita up to a turning point and then decrease with higher per capita income for the three income panels. The turning point income varies considerably between the three income groups. Moreover, the coefficient on energy consumption is positive in any income group for both models. Its magnitude

Table 5: Results of Pedroni (2001) DOLS estimates

		<i>CO</i> ₂ is the dependent variable				EKC
		<i>GDP</i>	<i>GDP</i> ²	<i>E</i>	<i>T</i>	Turning point income
High income	Model 1:					
	β	1.75 ^a	-0.09 ^a	0.97 ^a	-	24,584.96 USD
	<i>t</i> -statistic	3.32 (0.00)	-2.87 (0.00)	30.43 (0.00)	-	
	Model 2:					
β	2.97 ^a	-0.14 ^a	0.82 ^a	-0.22 ^a	41,000.15 USD	
<i>t</i> -statistic	3.31 (0.00)	-2.61 (0.01)	31.39 (0.00)	-2.78 (0.01)		
Middle income	Model 1:					
	β	2.09 ^a	-0.09 ^a	0.87 ^a	-	75,083.36 USD
	<i>t</i> -statistic	7.12 (0.00)	-5.82 (0.00)	21.84 (0.00)	-	
	Model 2:					
β	4.26 ^a	-0.22 ^a	0.79 ^a	-0.01 ^a	15,678.07 USD	
<i>t</i> -statistic	6.61 (0.00)	-5.63 (0.00)	20.97 (0.00)	3.17 (0.00)		
Lower income	Model 1:					
	β	3.99 ^a	-0.27 ^a	0.87 ^a	-	1,697.24 USD
	<i>t</i> -statistic	5.25 (0.00)	-4.66 (0.00)	12.72 (0.00)	-	
	Model 2:					
β	4.62 ^a	-0.31 ^a	0.96 ^a	0.05 ^a	1,984.17 USD	
<i>t</i> -statistic	6.34 (0.00)	-5.60 (0.00)	12.23 (0.00)	3.11 (0.00)		

Notes: P-values are in brackets; superscripts *a*, *b*, and *c* represent significance at 1%, 5%, and 10%, respectively; number of lags and leads included in the DOLS regression is set to one.

differs little between models or income groups. The inclusion of trade-openness leads to a higher turning point income, with the exception of the middle-income panel where there is a higher magnitude for model 1. Considering model 2, a higher degree of trade-openness is associated with lower CO₂ emissions per capita only for the high- and middle-income countries, and the reverse is true for the lower-income panel.

We determine the panel elasticity of CO₂ emissions per capita with respect to income per capita by solving $\frac{\partial CO_2}{\partial GDP} = \beta_{GDP} - (2\beta_{GDP^2}) \geq 1$ for β (Pao and Tsai, 2010). For the high-income group, the elasticity obtained with model 1 (model 2) is above unity when GDP is less than 4.33 (7.05) and inelastic if GDP is greater than 4.33 (7.05). For the middle-income panel, the elasticity of CO₂ emissions per capita with respect to income per capita obtained with model 1 (model 2) is above unity when GDP is less than 5.85 (7.39) and inelastic if GDP is greater than 5.85 (7.39).

For the lower-income panel, the elasticity obtained with model 1 (model 2) is above unity when GDP is less than 5.58 (5.95) and inelastic if GDP is greater than 5.58 (5.95).

The results indicate that a 1% increase in per capita energy consumption is associated with an increase in CO₂ emissions per capita for both models in any income group which is not above unity and thus is inelastic. According to Csereklyei et al. (2016), this implies a tendency for decreasing energy intensity in countries that have become richer. The inelastic response of trade-openness on per capita CO₂ emissions in model 2 suggest that a change in trade-openness results only in a weak change in emissions for high-, middle-, and lower-income countries.

The second-generation panel cointegration tests indicates the existence of a long-run relationship between the I(1) variables, although the outcome of the long-run cointegrating regression estimation yields no further insights into the causal relationships between CO₂ emissions, income, energy consumption, and trade-openness. Thus, in a final step we estimate a panel vector error correction model to examine Granger causality relationships between the variables. Tables 6, 7, and 8 display the causality results from the panel VECM based on the PMG estimation for each income group for model 1 and model 2. As both models within any income group produce similar results, we discuss the results obtained via model 2 only. Tables 6, 7, and 8 also display the empirical realization of the Wald chi-squared test statistics for short-run causality and the *t*-statistics for long-run causality. The lag length of the independent variables is set to three based on likelihood ratio tests conducted within any income group.

The results for the high-income panel suggest that unidirectional short-run causality runs from income growth to CO₂ emissions and energy consumption in the short-run; both joint null hypotheses $H_0 : \beta_{13ij} = \beta_{14ij} = 0, \forall ij$ and $H_0 : \beta_{23ij} = \beta_{24ij} = 0, \forall ij$ are rejected for every conventional significance level. Moreover, unidirectional short-run causality runs from trade-openness to energy consumption; the null hypothesis $H_0 : \beta_{25ij} = 0, \forall ij$ is rejected at the 1% significance level. Therefore, in the short-run, changes in economic output have a significant impact on both CO₂ emissions and energy consumption, and a change in trade-openness has a significant consequence on energy consumption. Furthermore, the results suggest bidirectional causality between CO₂ emissions and trade-openness; the null hypotheses $H_0 : \beta_{15ij} = 0, \forall ij$ and $H_0 : \beta_{51ij} = 0, \forall ij$ are rejected for every conventional

Table 6: Panel causality tests high-income panel

Dependent variable	Source of causation (Independent variables)				
	Short-run				Long-run
High income	CO_2	E	GDP & GDP^2	T	ECT
Model 1:					
(10) CO_2	-	2.48 (0.48)	12.61 ^b (0.05)	-	-0.26 ^a (0.00)
(11) E	0.80 (0.85)	-	24.33 ^a (0.00)	-	-0.05 ^a (0.00)
(12) GDP	1.71 (0.63)	1.44 (0.69)	-	-	-0.07 ^b (0.03)
(13) GDP^2	0.76 (0.86)	1.28 (0.73)	-	-	-0.43 ^a (0.00)
Model 2:					
(14) CO_2	-	4.28 (0.23)	12.50 ^b (0.05)	20.36 ^a (0.00)	-0.18 ^a (0.00)
(15) E	3.51 (0.32)	-	13.09 ^b (0.04)	17.12 ^a (0.00)	-0.07 ^a (0.01)
(16) GDP	1.72 (0.63)	3.89 (0.27)	-	13.09 ^a (0.00)	0.00 ^a (0.00)
(17) GDP^2	2.76 (0.43)	7.82 ^b (0.05)	-	14.09 ^a (0.00)	0.94 ^b (0.02)
(18) T	6.48 ^c (0.09)	1.05 (0.79)	35.32 ^a (0.00)	-	0.07 ^a (0.00)

Notes: P-values are in brackets; superscripts *a*, *b*, and *c* represent significance at 1%, 5%, and 10%, respectively.

significance level. Bidirectional causality between income growth and trade-openness is found as the joint null hypotheses $H_0 : \beta_{53ij} = \beta_{54ij} = 0, \forall ij$, $H_0 : \beta_{35ij} = 0, \forall ij$ and $H_0 : \beta_{45ij} = 0, \forall ij$ are rejected on every conventional significance level. The bidirectional causality results suggest that CO₂ emissions and trade-openness as well as income growth and trade-openness impact each other in the short-run. Finally, all five error correction terms are statistically significant, which implies that CO₂ emissions, GDP, its square, energy consumption, and trade-openness are important for the adjustment process to the long-term equilibrium when a shock leads to deviations from the long-term relationship.

In summary, the various causality results for the middle-income panel are as follows: Unidirectional short-run causality runs from income growth to CO₂ emissions as well as trade-openness in the short-run. Therefore, changes in economic output have a significant impact on both CO₂ emissions and trade-openness in the short-run. Unidirectional short-run causality runs from trade-openness to both CO₂ emissions and energy consumption, which implies that changes in trade-openness have a significant impact on both CO₂ emissions and energy consumption in the short-run. Bidirectional causality exists between CO₂ emissions and energy consumption only. Therefore, CO₂ emissions and energy consumption are jointly determined and affected simultaneously in the short-run. For the

Table 7: Panel causality tests middle-income panel

Dependent variable	Source of causation (independent Variables)				
	Short-run				Long-run
Middle income	CO_2	E	GDP & GDP^2	T	ECT
Model 1:					
(10) CO_2	-	7.21 ^c (0.07)	15.94 ^b (0.01)	-	-0.29 ^a (0.00)
(11) E	10.16 ^b (0.02)	-	13.37 ^b (0.04)	-	0.01 ^b (0.04)
(12) GDP	1.34 (0.72)	2.96 (0.40)	-	-	-0.06 (0.36)
(13) GDP^2	1.46 (0.69)	2.75 (0.43)	-	-	-1.08 (0.34)
Model 2:					
(14) CO_2	-	8.75 ^b (0.03)	23.77 ^a (0.00)	6.86 ^c (0.08)	-0.21 ^a (0.00)
(15) E	8.11 ^b (0.04)	-	7.18 (0.30)	21.78 ^a (0.00)	0.00 (0.64)
(16) GDP	1.42 (0.70)	1.28 (0.74)	-	2.52 (0.47)	0.04 ^a (0.00)
(17) GDP^2	2.39 (0.50)	0.62 (0.89)	-	0.97 (0.81)	0.01 ^b (0.02)
(18) T	2.88 (0.41)	0.73 (0.86)	34.82 ^a (0.00)	-	0.21 (0.17)

Notes: P-values are in brackets; superscripts *a*, *b*, and *c* represent significance at 1%, 5%, and 10%, respectively.

middle-income panel considering model 2, the results of the tests for the significance of the one period lagged ECT indicate that only CO_2 emissions, GDP, and its square are important for the adjustment process to the long-term equilibrium when a shock leads to deviations from the long-term relationship.

For the lower-income panel, the results suggest that unidirectional short-run causality runs from income growth to CO_2 emissions as well as trade-openness in the short-run. Unidirectional short-run causality runs also from energy consumption to CO_2 emissions and from trade-openness to energy consumption. Therefore, changes in economic output have a significant impact on both CO_2 emissions and trade-openness, and a shift in the amount of energy consumed causes a change in CO_2 emissions and a change in trade-openness has a significant impact on the scope of energy consumption. Bidirectional causality exists between CO_2 emissions and trade-openness only. Thus, CO_2 emissions and trade-openness are affected simultaneously in the short-run. For the lower-income panel, all five error correction terms are statistically significant, which suggests that CO_2 emissions, GDP, its square, energy consumption, and trade-openness are important for the adjustment process to the long-run equilibrium when a shock leads to deviations from the long-run relationship.

Table 8: Panel causality tests lower-income panel

Dependent variable	Source of causation (independent Variables)				
	Short-run				Long-run
Lower income	CO_2	E	GDP & GDP^2	T	ECT
Model 1:					
(10) CO_2	-	7.63 ^b (0.05)	11.11 ^c (0.09)	-	-0.10 ^c (0.07)
(11) E	1.49 (0.68)	-	3.47 (0.75)	-	-0.01 (0.16)
(12) GDP	1.00 (0.81)	1.08 (0.78)	-	-	0.02 ^c (0.09)
(13) GDP^2	0.75 (0.86)	0.67 (0.88)	-	-	0.36 ^b (0.05)
Model 2:					
(14) CO_2	-	9.20 ^b (0.03)	11.27 ^c (0.08)	10.45 ^b (0.02)	-0.29 ^a (0.00)
(15) E	1.25 (0.74)	-	6.20 (0.40)	6.53 ^c (0.09)	0.02 ^a (0.01)
(16) GDP	1.18 (0.76)	1.74 (0.63)	-	2.80 (0.43)	0.01 ^a (0.00)
(17) GDP^2	1.07 (0.78)	0.65 (0.88)	-	10.64 ^c (0.10)	0.42 ^c (0.07)
(18) T	7.37 ^c (0.06)	5.20 (0.16)	22.05 ^a (0.00)	-	0.00 ^c (0.08)

Notes: P-values are in brackets; superscripts *a*, *b*, and *c* represent significance at 1%, 5%, and 10%, respectively.

5 Conclusions and Policy Implications

This paper investigated the dynamic relationship between CO₂ emissions, income, energy usage, and trade-openness for a panel of 70 WTO countries from 1971 to 2013, clustered into high-, middle-, and lower-income groups, based on the EKC hypothesis and using recently developed panel time series data methods. The CD test (Pesaran, 2004) was used to determine the presence of cross-sectional dependence within the high-, middle-, and lower-income panels. The results indicated that the null hypothesis of no cross-sectional dependence could be rejected in any income group. The results of the second-generation panel unit root test (Pesaran, 2007) accounting for cross-section dependence suggested that the series on CO₂ emissions, GDP, GDP², energy consumption, and trade-openness were integrated of order one. The results of the error correction-based panel cointegration test proposed by Westerlund (2007) supported the existence of a long-run equilibrium relationship for the four and five integrated variables cases in any income group. The DOLS technique (Pedroni, 2001) used to estimate the cointegrating relationship between

CO₂ emissions, energy consumption, economic growth, and trade-openness provided empirical evidence of an inverted U-shape pattern associated with the EKC hypothesis within all panels for models 1 and 2. The estimated turning point income varied between the three income groups. The coefficient on energy consumption was positive in any income group for both models, but a higher degree of trade-openness was associated with lower per capita CO₂ emissions only for the high- and middle-income panels. A panel-based VECM was estimated using the PMG estimator (Pesaran et al., 1999) to identify the short- and long-run Granger causal relationships among the variables. Table 9 summarizes the various short-run causality results and the empirical results of the DOLS estimation for model 2.

Table 9: Summary of the empirical results for model 2

	EKC	Granger short-run causality	
	Turning point income	unidirectional	bidirectional
High income	41,000.15 USD	GDP & GDP ² → CO ₂ GDP & GDP ² → E T → E	T ↔ CO ₂ T ↔ GDP & GDP ²
Middle income	15,678.07 USD	GDP & GDP ² → CO ₂ T → CO ₂ T → E GDP & GDP ² → T	E ↔ CO ₂
Lower income	1,984.17 USD	E → CO ₂ GDP & GDP ² → CO ₂ T → E GDP & GDP ² → T	T ↔ CO ₂

Notes: → and ↔ indicate unidirectional and bidirectional short-run causality, respectively.

The acceptance of the EKC hypothesis for per capita CO₂ emissions within every income panel highlights several important policy implications. The implementation of stronger and international development cooperation could help to prevent that future economic growth is inevitable accompanied by environmental degradation in early stages of economic development. Nevertheless, economic growth helps to undo the damage after a certain turning point income is reached. In other words, policy measures which benefit sustainable economic growth improve environmental quality with respect to emissions (Yandle et al., 2004). The scale effect, the composition of output, the change in energy

inputs, the use of cleaner production technology, and the improvement in energy efficiency due to technological progress are important factors explaining the EKC theoretically (Stern, 2004). As we found empirical evidence of an inverted U-shape pattern, the aforementioned positive impacts on carbon dioxide emissions outweigh the ecologically harmful scale effect for any income panel. According to Panayotou (1997), the quality of policy measures and institutions can speed up the process to reduce environmental degradation at higher incomes. In other words, more effective environmental regulations can reduce the environmental price of economic growth. Similarly, trade liberalization can raise per capita income and thus the demand for environmental protection. The empirical results obtained from estimating the cointegrating relationship suggest that trade-openness has a negative impact for the lower-income panel in terms of CO₂ emissions. Still, the simple empirical existence of the EKC relationship does not guarantee that CO₂ emissions will decrease automatically with economic growth.

The causality analysis results from model 2 also have important policy implications. For the energy consumption and economic growth nexus, the causality analysis reveals that economic growth leads energy consumption within the high-income panel in the short-run. This provides evidence for the conversion hypothesis: policies designed to reduce energy consumption do not disadvantageously affect per capita income as causality is running from economic growth to energy consumption. The neutrality hypothesis is found for both the middle- and lower-income panels in the short-run since there is no causal relationship between income and energy consumption. Therefore, a reduction in energy consumption might not affect economic growth negatively (Payne, 2010). Tables 6, 7, and 8 show that the conversion hypothesis is empirically supported for the high- and middle-income panels considering the results for model 2 without trade-openness. Moreover, for any income panel, the results for model 2 suggest that income growth leads emissions and that trade-openness causes energy consumption in the short-run. The former result is expected because an increase in output stimulates demand for energy which is a major source for CO₂ emissions when not generated by renewable energy sources. The latter result suggests that a higher volume of trade relations increases economic activity within

export-oriented sectors that generally require additional energy (Sadorsky, 2012).

For the high-income panel, trade-openness and GDP are jointly determined, whereas in both the middle- and lower-income panels, GDP leads trade-openness. Concerning the feedback relationship found for the high-income panel, Helpman and Krugman (1985) argue that trade-openness tends to increase when economies of scale are realized due to productivity gains. The rise in exports then enables cost reductions which lead to further productivity gains. Bhagwati (1988) argues that a higher volume of trade-openness generates more income, which leads to more trade, and so on. On the other hand, the growth-led exports hypothesis holds for the middle- and lower-income panels, because economic growth spurs technology and production capabilities that lead to efficiency gains, thus creating a comparative advantage that facilitates exports (Giles and Williams, 2000).

The finding of unidirectional (bidirectional) short-run causality between energy consumption and CO₂ emissions for the middle-income (lower-income) panel against the background of the lack of causality for the high-income panel indicates that middle- and lower-income countries need to reduce fossil fuel energy consumption, or invest in increasing energy efficiency to reduce CO₂ emissions.

The final implication from the causality analysis concerns the impact of trade-openness on the environment. For the high- and lower-income panels, bidirectional short-run causality is identified between trade-openness and CO₂ emissions. In other words, any change in one will affect the other. In fact, trade-openness leads per capita CO₂ emissions within the middle-income panel in the short-run. Based on the causality results, there is no specific sign that trade-openness could positively affect high-, middle-, and lower-income countries in terms of CO₂ emissions.

Because the analysis in this paper was conducted at an aggregate level, the empirical results may differ across countries, although using a panel data time series approach improves the efficiency of econometric estimates. We suggest that future research should investigate the dynamic relationship between CO₂ emissions, income, energy consumption, and trade-openness in a panel data framework and also examine the linkage in question for each country separately in the panel. A disaggregated analysis considering energy

consumption by source as well as economic growth by sector could provide additional insights into the linkage between output and CO₂ emissions (Payne, 2010). A solution for estimations involving nonlinear transformations of nonstationary variables should be developed for the panel framework (Wagner, 2008).

This paper contributed to the scarce panel based third strand literature by incorporating trade-openness as an explanatory variable to identify its effect on carbon dioxide emissions within any income panel. Identifying an inverted U-shape pattern associated with the EKC hypothesis suggests that the net effect of economic growth should help to stabilize CO₂ emissions in the near future. However, further economic development accompanied by both a stringent climate and energy policy will be critical to prevent the irreversible ecological damage caused by economic activities within any income panel.

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A Appendix

Table 10: Summary statistics for high-income countries

High-income	CO2		GDP		E		T	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Australia	15.59	1.71	38,132.12	9,027.92	5,070.93	541.63	34.86	5.74
Austria	7.68	0.55	35,106.66	8,651.05	3,355.85	479.63	75.06	15.63
Belgium	11.23	1.51	34,015.20	7,631.39	5,015.53	492.59	121.86	21.58
Canada	16.36	1.04	37,835.78	7,339.27	7,622.38	416.82	58.60	11.70
Chile	3.01	0.94	7,807.284	3,345.42	3,345.42	458.26	55.93	12.95
Denmark	10.41	1.55	46,653.63	9,772.76	3,607.66	231.62	74.19	14.23
Finland	10.63	1.14	33,648.98	9,276.78	5,679.14	894.95	62.00	12.02
France	6.90	1.40	32,772.20	6,553.38	3,763.22	388.70	46.01	7.21
Greece	6.74	1.64	21,196.02	4,135.64	2,026.52	513.86	43.54	9.50
Hong Kong	4.80	1.39	19,146.82	8,663.32	1543.77	491.07	266.88	91.66
Iceland	7.38	0.70	31,619.58	7,934.50	9,382.71	3,917.41	73.88	11.36
Ireland	8.70	1.33	30,112.99	14,573.53	2,832.33	504.61	127.08	35.97
Israel	7.57	1.65	22,300.09	5,473.11	2,456.68	471.57	82.89	16.15
Italy	7.05	0.69	29,963.42	6,288.07	2,594.76	351.47	43.57	6.92
Japan	8.74	0.82	33,985.54	8,457.63	3,453.66	507.66	23.43	5.40
Korea, Rep.	6.68	3.27	11,037.89	7,000.27	2,670.63	1,617.45	65.89	16.74
Luxembourg	25.61	6.13	67,986.64	25,989.84	8,972.26	1,476.22	234.28	57.75
Malta	4.99	1.59	13,468.76	5,751.52	1,601.64	522.23	197.96	59.02
Netherlands	10.85	0.83	38,025.06	8,873.78	4,529.63	269.16	110.49	17.49
Norway	9.04	1.43	65,663.49	18,750.10	5,114.55	901.47	72.18	4.21
Portugal	4.15	1.43	16,807.79	4,573.94	1,662.92	627.27	59.15	9.42
Saudi Arabia	14.42	2.33	18,545.67	4,318.89	3,986.66	1597.47	78.11	12.91
Singapore	11.66	3.29	25,643.43	13,130.51	3,807.41	1643.45	344.05	46.44
Spain	5.96	1.06	23,324.78	5,860.99	2,328.50	596.25	42.93	11.24
Sweden	7.05	1.87	39,049.50	8,644.15	5,344.39	400.30	67.47	14.09
Trinidad and Tobago	19.25	8.40	9,996.18	3,640.59	6,878.94	4,319.04	86.95	13.34
United Kingdom	9.56	1.16	29,573.56	7,475.21	3,583.70	233.67	51.33	5.00
United States	19.56	1.40	37,545.68	8,750.57	7,696.15	362.14	20.98	5.04
Uruguay	1.76	0.39	8,039.96	2,073.69	891.03	178.95	42.64	10.19
Panel	9.77	5.81	29,620.85	17,268.91	4,094.38	2,595.74	91.87	79.55

Notes: CO₂ emissions are measured in metric tons per capita; GDP per capita is measured in constant 2010 USD; per capita energy consumption is measured in kg of oil equivalent per capita; trade is the sum of exports and imports of goods and services measured as a share of gross domestic product; mean is the arithmetic mean, and SD denotes the standard deviation; data obtained from the World Bank Development Indicators database (last updated 16 December 2016).

Table 11: Summary statistics for middle-income countries

Middle-income	CO2		GDP		E		T	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Argentina	3.84	0.38	7,970.44	1,229.38	1,566.08	180.27	22.23	10.24
Brazil	1.62	0.32	8,416.59	1,498.34	1,016.20	171.54	20.14	4.50
China	2.91	1.83	1,572.72	1,554.56	943.71	471.88	30.93	17.60
Colombia	1.56	0.14	4,523.36	1,070.94	658.53	43.24	32.74	3.92
Costa Rica	1.26	0.33	5,720.88	1,427.52	665.40	209.16	74.04	11.33
Cuba	2.81	0.44	3,862.67	1,052.40	1,249.57	249.89	54.48	21.34
Dominican Rep.	1.66	0.51	3,322.43	1,145.48	657.19	89.85	63.20	14.90
Ecuador	1.82	0.54	3,851.96	559.63	660.72	152.35	44.45	10.95
Gabon	5.39	2.34	11,095.69	2,090.29	1,545.07	364.96	94.82	13.11
Malaysia	4.21	2.22	5,475.27	2,416.85	1,589.74	788.14	146.18	45.08
Mexico	3.61	0.48	7,611.23	1,091.19	1,351.37	202.98	38.88	15.84
Panama	1.80	0.44	5,227.06	1,614.84	856.82	155.75	130.11	19.24
Peru	1.27	0.27	3,673.60	727.52	550.52	98.15	38.19	9.32
South Africa	8.91	0.91	6,592.99	509.56	2,556.36	250.10	52.67	7.51
Thailand	2.16	1.38	2,848.40	1,452.72	935.68	506.36	83.44	36.04
Turkey	2.76	0.89	6,953.39	2,028.85	1,009.33	284.41	35.54	14.67
Venezuela	5.89	0.73	13,117.24	1,337.35	2,067.50	239.48	49.26	6.59
Panel	3.15	2.25	5,990.35	3,244.58	1,169.40	622.91	59.49	39.61

Notes: CO₂ emissions are measured in metric tons per capita; GDP per capita is measured in constant 2010 USD; per capita energy consumption is measured in kg of oil equivalent per capita; trade is the sum of exports and imports of goods and services measured as a share of gross domestic product; mean is the arithmetic mean, and SD denotes the standard deviation; data obtained from the World Bank Development Indicators database (last updated 16 December 2016).

Table 12: Summary statistics for lower-income countries

Lower-income	CO ₂		GDP		E		T	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Benin	0.23	0.15	643.28	64.49	348.53	27.87	52.78	7.22
Bolivia	1.08	0.36	1,614.26	225.08	463.03	137.12	55.57	12.39
Cameroon	0.28	0.16	1,164.47	204.78	390.00	30.20	45.25	8.02
Congo, Dem. Rep.	0.08	0.05	576.99	282.66	313.82	15.60	46.20	19.99
Congo, Rep.	0.49	0.21	2,567.13	410.87	342.00	73.33	114.14	26.01
Cote d'Ivoire	0.50	0.13	1,573.73	359.46	433.81	68.80	75.52	10.77
Egypt	1.57	0.59	1,678.98	567.03	565.22	215.22	52.61	12.30
El Salvador	0.74	0.29	2,820.35	534.57	608.52	98.30	61.85	10.59
Ghana	0.32	0.07	992.14	215.00	346.57	38.13	55.61	29.99
Guatemala	0.66	0.17	2,425.7 0	256.53	548.74	79.83	48.23	12.41
Honduras	0.73	0.26	1,688.45	240.68	528.65	58.79	88.82	27.63
India	0.80	0.36	688.50	343.28	380.44	96.78	24.04	14.89
Indonesia	1.07	0.53	1,850.45	811.61	570.74	197.52	53.15	10.30
Kenya	0.29	0.05	879.33	68.02	453.35	14.19	58.23	6.78
Morocco	1.08	0.39	1,840.86	586.07	346.77	107.40	57.32	12.07
Nepal	0.08	0.06	401.14	111.52	325.04	26.69	38.56	12.88
Nicaragua	0.68	0.13	1,576.83	424.82	508.62	29.24	66.25	19.11
Nigeria	0.65	0.18	1,660.83	390.06	690.19	52.42	49.34	15.64
Pakistan	0.64	0.22	757.45	198.81	402.94	76.29	33.27	3.51
Philippines	0.79	0.12	1,651.08	265.23	455.81	25.20	67.85	21.71
Senegal	0.45	0.08	909.35	62.08	254.78	26.82	65.79	9.57
Sri Lanka	0.40	0.19	1,541.94	736.96	372.82	73.70	67.91	11.49
Togo	0.24	0.08	541.91	56.53	372.31	58.81	90.88	17.28
Tunisia	1.77	0.50	2,636.59	848.74	644.94	189.15	83.21	15.30
Panel	0.65	0.50	1,445.07	800.71	444.48	146.62	60.52	24.98

Notes: CO₂ emissions are measured in metric tons per capita; GDP per capita is measured in constant 2010 USD; per capita energy consumption is measured in kg of oil equivalent per capita; trade is the sum of exports and imports of goods and services measured as a share of gross domestic product; mean is the arithmetic mean, and SD denotes the standard deviation; data obtained from the World Bank Development Indicators database (last updated 16 December 2016).