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Stochastic Convergence in Per Capita Carbon Dioxide (CO_2) Emissions: Evidence from OECD Countries

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Abstract

This study analyzes the validity of stochastic convergence hypothesis in relative per capita CO_2 emissions in OECD (Organization for Economic Cooperation and Development) countries for the period 1960-2013. In other words, it is aimed to reveal the nature of shocks to relative per capita CO_2 emissions. As such, divergence holds if shocks are permanent, whereas convergence holds if shocks are temporary. To that aim, the two-break LM (Lagrange multiplier) and three-step RALS-LM (residual augmented least squares Lagrange multiplier) unit root tests are employed. The results mostly provide evidence of convergence in case of two breaks. However, when structural breaks are not taken into consideration, divergence gains empirical validity. From the viewpoint of government policy, these results indicate that energy usage or environmental protection policies of OECD countries have not long-run impacts on the relative per capita emissions series of the sample countries. Concerning the break dates, the first breaks mostly cumulated around the two energy crises period, whereas the second breaks generally occurred in the 1990s.

Keywords: *CO_2 emissions, structural breaks, unit root test, convergence, OECD*

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1. Introduction

Global warming and climate change are serious environmental problems today. The high level of emissions from carbon dioxide (CO_2) and other greenhouse gases is the main contributor to these global problems. The combustion of fossil fuels and other human activities are the main reasons for the increased atmospheric concentration of CO_2. In this scenario, CO_2 contributes to climate change with potentially irreversible negative impacts on the world economy once it is released into the atmosphere as a byproduct of the consuming of fossil fuels (Lee & Lee, 2009).

In this context, industrialized countries have started to arrange important environmental agreements and agendas to decrease and control the atmospheric concentration of greenhouse gases emissions since the early 1990s. For instance, the United Nations Framework Convention on Climate Change (UNFCCC), which was adopted in May 1992 and opened for signatures a month later at the United Nations Conference on Environment and Development in Rio de Janeiro, Brazil, is one of the best-known environmental conventions. The ultimate aim of UNFCCC is to reduce CO_2 emissions across countries to combat global climate change and the greenhouse effect, which are caused by high concentrations of CO_2 in the atmosphere. Additionally, the Kyoto Protocol has extended the UNFCCC. Initially adopted on December 11, 1997, the protocol's major feature is that it sets binding targets for reducing greenhouse gas (GHG) emissions for 37 industrialized countries and the European community. In accordance with the Kyoto Protocol, signatories from developed countries were committed to reducing their combined GHG emissions by at least 5% from their 1990 levels by the period of 2008–2012. Finally, the 2015 United Nations Climate Change Conference, held in Paris, France, was an important arrangement. It was also the 21st yearly session of the Conference of the Parties to the 1992 UNFCCC and the 11th session of the Conference of the Parties to the 1997 Kyoto Protocol.

Based on the aforementioned developments, scholars have started to employ empirical studies on per capita greenhouse emissions to find effective solutions and to develop policy instruments that will increase environmental quality. In this respect, convergence issue in the environmental economics has gained empirical concern among scholars. Convergence holds if countries with low per capita emissions increase their emissions while high per capita emissions countries decrease their emissions (Romero-Avila, 2008). In other words, the equal allocation of emissions to all countries on a per capita basis indicates the existence of convergence (Westerlund & Basher, 2008). In particular, scholars have focused on convergence in CO_2 emissions because CO_2 appears as the main cause of global warming. Furthermore, analyzing convergence in per capita CO_2 emissions (PCE hereafter) is crucial for the following reasons.

First, the detailed explanation and evolution of per capita emissions over time is critical to the development of relevant emissions projection models and sufficient policy responses (Stegman, 2005). The projection models concerning the formulation of emission abatement strategies to combat climate change assume emissions convergence (Romero-Avila, 2008). Thus, convergence is accepted as a key tool for long-run CO_2 emissions projections (Westerlund & Basher, 2008). Furthermore, the stationary relative per capita CO_2 emissions (RPCE) series indicates that it is possible to forecast future movements in RPCE based on its past behavior. CO_2 emissions forecasting is critical for an appropriate climate policy response. Second, there is a close relationship between economic development, environmental protection, and energy consumption. Energy consumption leads to increases in the atmospheric concentration of CO_2 emissions.[1] Therefore, determining whether RPCE series is converging provides information about government action for controlling the atmospheric concentration of emissions. If convergence does hold, the government should not interfere excessively with countries that reach convergence in their RPCE series given that relative CO_2 emissions series only deviate from their mean temporarily (Lee & Chang, 2009). In this case, the government's administrative policy should not be to adopt unnecessary targets. Instead, the government should pay attention to the long-running trends in CO_2 emissions.

Based on the explanations above, we aimed to test the stochastic convergence[2] for the OECD sample, which consists of 28 countries over the period of 1960–2013 by using two-step LM and three-step RALS-LM unit root tests that were developed by Lee et al. (2012) and Meng and Lee (2012), respectively. To our knowledge, there is no study using them to test for the convergence hypothesis in PCE. We contribute to the literature by employing cutting-edge unit root tests. These two tests are more powerful than other endogenous unit root tests. First, endogenous unit root tests depend on the nuisance parameter describing the break under the null, and they assume that breaks are absent under the null to eliminate the dependency on nuisance parameter. However, as proposed by Nunes, Newbold and Kuan (1997), Lee and Strazicich (2003, 2004), and Lee et al. (2012), these endogenous unit root tests may lead to spurious rejections under the null when the magnitude of break is not zero. Therefore, rejection of the null hypothesis in these endogenous tests does not indicate a trend stationary series with break and the possibility of unit root with break still remains. However, the LM tests with trend-breaks applied in this study are invariant to nuisance parameters since a transformation procedure is adopted and thereby, they allow for trend-breaks under the null hypothesis. In

[1] In the literature, especially studies that used the framework of the environmental Kuznets curve (EKC) hypothesis, the relationship between economic development, energy consumption, and CO_2 emissions is stressed. According to the EKC hypothesis, there is an inverted U-shaped relationship between economic development and CO_2 emission levels. However, the EKC issue was left out of this study. See Ozcan (2013) for further details on the EKC hypothesis.

[2] The definitions of stochastic and other convergence types will be provided in Section.

addition, as a first step, the LM tests employed in this study jointly determine whether and where structural breaks occur in the data, whereas existing endogenous unit root tests always find and include the number of breaks that are pre-specified in the model (see Lee et al., 2012). Besides, as suggested by Lee et al. (2012), the proposed two-step LM unit root test accurately identifies breaks and has better size properties.

As a second unit root test, we employ the RALS-LM test that has improved power with non-normal errors and is fairly robust to some forms of non-linearity (see Meng et al., 2013). However, as proposed in Meng and Lee (2012), the specific form of non-linearity in the real world data is rarely known. In this case, the non-linearity tests are often less powerful than the Dickey- Fuller type linear tests. However, the RALS-LM unit root test is based on the linearized model specification, and thus it will not be subject to this difficulty. Besides, a simple transformation of the data which eliminates dependency on the trend breaks is adopted. Thereby, the same critical values can be used at different break locations. Thanks to this transformation, the nuisance parameter and spurious rejection problems are solved. In addition, the proper number of breaks is determined from the data in the RALS-LM test, whereas other endogenous break unit root tests usually find and include the number of breaks as pre-specified in the model. As sum, by employing the transformed LM and RALS-LM unit root tests, we remove test statistic's dependency on nuisance parameter which many endogenous break unit root tests have.

The remainder of the paper is organized as follows. Section 2 is a literature review. In section 3, we explain our data and methodology. Section 4 provides empirical results. We then conclude the study in the last section.

2. Literature Review

Many studies have tested for convergence in per capita GHG emissions, especially CO_2. However, they do not reach a consensus because they use different country samples, time periods, and methods. The first strand includes studies that have employed conventional univariate and panel unit root tests. For instance, Strazicich and List (2003), using the IPS (Im, Pesaran, & Shin 2003) panel unit root test, tested both the stochastic conditional and β convergence types in 21 industrial countries over the period of 1960–1997. Their results provided evidence of convergence. Applying the DF-GLS unit root test developed by Elliot, Rothenberg and Stock (1996), Aldy (2005) tested the cross-sectional and stochastic convergence hypotheses for 88 countries and a sub-group of 23 OECD countries for the period of 1960–2000. The author obtained mixed support for the stochastic convergence for the OECD sample.

In the second strand, researchers allowed for structural breaks in the frameworks of univariate unit root tests. For instance, List's (1999) seminal paper employed a unit root test that was based on the trend break model of Perron and Vogelsang (1992). In the study, List examined the convergence of per capita emissions of nitrogen oxide and sulfur dioxide for the 10 energy Environmental Protection Agency regions of the United States over the period of 1929–1994. The results indicated that stochastic convergence for both air pollutant emissions held for only two regions. Lanne and Liski (2004) employed an additive outlier model through unit root tests that allow for breaks for 16 OECD countries over the period of 1870–2002. They found that 10 out of 16 series were stationary. In their study, Lee et al. (2008) employed a suite test statistic, proposed by Sen (2003), to study 21 OECD countries within the period of 1960–2000. The authors obtained results favorable to convergence in RPCE series. Chang and Lee (2008) employed LM unit root tests developed by Lee and Strazicich (2003, 2004) to test for convergence in RPCE series among 21 OECD countries for the period of 1960–2000. The results provided evidence of convergence. Along the same line, McKitrick and Strazicich (2006) also used LM unit root tests to test for stationarity in PCE for the world and 121 individual countries for the period of 1950–2000. The stationarity result was generally obtained. Payne et al. (2014) used residual augmented least squares – Lagrange Multiplier (RALS-LM) unit root test with breaks for the District of Columbia from 1900 to 1998. The results supported convergence for each state and the District of Columbia. For 39 African countries, Solarin (2014) employed LM unit root test to analyze the stochastic convergence from 1960 to 2010 and obtained evidence of convergence.

The third strand of studies has applied panel unit root tests allowing for structural breaks and/or cross-sectional dependence when testing for convergence in RPCE. First, Barassi, Cole and Elliott (2008) employed a battery test that allowed for cross-sectional dependence when testing for convergence in PCE among OECD countries for the period of 1950–2002. The authors found that RPCE series did not converge. Lee and Chang (2008) employed the panel SURADF unit root test, developed by Breuer, McNown and Wallace (2001, 2002), to test for convergence in RPCE series among 21 OECD countries from 1960 to 2000. The overall results indicated that convergence was supported for nearly one-third of the countries. Furthermore, Westerlund and Basher (2008) found evidence favorable to the convergence in RPCE series for 16 industrialized and 12 developing countries over the past century by using panel unit root tests that controlled for cross-correlation through a factor model. Westerlund and Basher (2008) also employed the panel unit root tests of Phillips and Sul (2003), Moon and Perron (2004), and Bai and Ng (2004) to test for convergence in RPCE series for 28 developed and developing countries for the period 1870–2002. The authors provided strong support in favor of convergence.

In this strand, there are also some studies that employed panel unit root tests with structural breaks. For instance, Romero-Avila (2008), using the KPSS panel unit root test developed by Carrion-i-Silvestre, Barrio-Castro and Lopez-Bazo (2005), tested the validity of both the stochastic and deterministic convergence hypotheses. The results were generally favorable to the both convergence types in RPCE series among 23 OECD countries for the period of 1960–2002. Lee and Chang (2009) also employed the KPSS panel test to examine the stochastic convergence in RPCE series among 21 OECD countries over the period of 1950–2002. Their results strongly supported the stochastic convergence hypothesis.

The fourth strand of studies used a nonparametric approach to convergence. For instance, Nguyen Van (2005), who employed nonparametric methods to examine the convergence of RPCE series in 100 countries for the period of 1966–1996, reported that industrialized countries exhibited a convergence pattern, although he found little evidence of convergence for the whole sample. Aldy (2006) tested both the sigma (σ) convergence and stochastic convergence in per capita income and CO_2 emissions series for the period of 1960–1999 within the framework of a nonparametric Markow-transition matrix. The results signaled little evidence in favor of convergence in CO_2 emissions series. Using a robust distributional approach and nonparametric distribution tests, Criado and Grether (2011) analyzed convergence in PCE within a panel consisting of 166 countries spanning the period of 1960–2002. The results showed that divergence held before the oil price shocks in the 1970s, whereas some group-specific convergence patterns emerged for the period of 1980–2000. Ezcurra (2007), who examined the spatial distribution of PCE in 87 countries over the period of 1960–1999 using a nonparametric approach, found that there was convergence over these 40 years but that it would not continue indefinitely. Stegman (2005) employed the stochastic kernel estimation of the intra-distributional dynamics of cross-country PCE over time and attained little evidence of absolute convergence across countries.

The fifth strand includes studies that employed nonlinear unit root tests to test for the stochastic convergence. Yavuz and Yilanci (2012) analyzed convergence in RPCE for G7 countries by employing the threshold autoregressive panel unit root test spanning the period of 1960–2005. They found that convergence held only during the first regime, while divergence was confirmed in the second regime. Camarero, Mendoza and Ordoñez (2011) examined convergence in RPCE series among 22 OECD countries for the period of 1870–2006, using the unit root test developed by Kapetianos et al. (2003). The authors' results indicated that there was no robust convergence. Panopoulou and Pantelidis (2009) examined CO_2 emissions convergence among 128 countries within the period of 1960–2003 using Phillips and Sul's (2007) methodology, which is based on a nonlinear time-varying factor model. Their results indicated that all 128 countries' RPCE series converged in the sample's early period. Furthermore, in the recent years of the sample, there were two convergence clubs. Anorou and Emmanuel (2014) used sequential panel selection method in 15 African countries from1971 to 2011. The results indicated that 11 countries converged, whereas 4 countries diverged. Presno, Landoja and Fernández (2015) employed nonlinear stationarity analysis for the 28 OECD countries over the period of 1901-2009. Their results showed that developed countries were convergent under smooth transitions. Wu et al. (2016) employed the continuous dynamic distribution approach to the panel data of 286 Chinese cities. The results supported convergence for the period of 2002–2011.

In the last strand, there are studies applying other methods than unit root test to analyze the convergence issue. For instance, Nourry (2009) used a pairwise approach for 127 countries to analyze the stochastic convergence from 1950 to 2003. Their results did not provide evidence in favor of convergence. Barassi et al. (2011) examined the convergence of CO_2 emissions within 18 OECD countries over the period of 1870–2004, employing a local whittle estimator of the fractional integration parameter. The results indicated that convergence held for 13 of 18 OECD countries. Oliveira and Vargas Mores (2015) used random and fix effects estimators to examine 118 countries for the period of 1970–2008. They found a strong convergence in large global and regional samples of countries. El-Montasser et al. (2015) examines GHG emissions convergence among the G7 countries, using the pairwise testing technique along with a number of unit root tests proposed by Pesaran (2007) for the period between 1990 and 2011. However, their results do not confirm the hypothesis of convergence. Li and Lin (2013) investigate the convergence in PCE for 110 countries over the period 1971–2008, applying generalized method of moments (GMM) approach. The results manifest an absolute convergence within subsamples grouped by income level, while provide little evidence of absolute convergence in the full sample. Herrerias (2012) analyzes the convergence issue in PCE across the EU-25 countries from 1920 to 2007, using the distribution dynamics approach. It supports the hypothesis of convergence of carbon dioxide emissions across the European countries.

3. Data and Methodology

3.1. Data

The primary data used in this paper for 28 OECD[3] countries were the annual PCE measured in metric ton, which were taken from the World Bank, World Development Indicators (WDI, 2015) database. The country sample and time interval were dictated by data availability.

In designing a precise examination indicator, we followed Aldy (2005), Chang and Lee (2008), Lee and Chang (2008, 2009), Nguyen-Van (2005), Romero-Avila (2008), and Strazicich and List (2003), who used Carlino and Mills' (1993) methodology to test for the stochastic convergence in RPCE. Carlino and Mills (1993) examined the convergence of per capita income for eight regions in the United States by calculating the log of the ratio of per capita income relative to the average per capita income of United States and then conducting a unit root test. Based on this methodology, we calculated a yearly sample average for the 28 OECD countries

[3] These are Austria, Australia, Belgium, Canada, Chile, Denmark, Finland, France, Greece, Hungary, Iceland, Ireland, Israel, Italy, Japan, Luxembourg, Mexico, Netherlands, New Zealand, Norway, Poland, Portugal, Spain, Sweden, Switzerland, Turkey, the United Kingdom, and the United States.

under study and then took the natural logarithm of RPCE for each country *i* as follows:

$$X_{i,t} = ln \frac{CO_{2_{pci,t}}}{average CO_{2_{pct}}} \tag{1}$$

where $CO_{2_{pci,t}}$ is the PCE for country i and average $CO_{2_{pct}}$ is the yearly average value in the sample span. Searching for the unit root in $X_{i,t}$ (RPCE) provides a clear conclusion about whether stochastic convergence holds. Moreover, as stated by Chang and Lee (2008) and Nguyen-Van (2005), some common trends in emissions are avoided thanks to this relative measure of per capita CO_2 emissions.

In general, there are three different types of convergence, stochastic convergence, α convergence, and β convergence, addressed first in the growth literature and then in the environmental literature (Jobert et al., 2010; Panopoulou & Pantelidis, 2009). However, we focused on stochastic convergence in this study. Stochastic convergence indicates that the effects of temporary shocks on RPCE dissipate over time or, likewise, that the time series does not possess a unit root (List, 1999). As such, a unit root in the log of RPCE indicates that a given country's RPCE does not converge stochastically toward the sample average because shocks to RPCE cause permanent effects. The conventional way to test for the stochastic convergence is by employing the unit root tests. In the econometric framework, stochastic convergence holds if the log of RPCE trends stationary. From the economic policy point of view, stationarity implies that the effects of the reduction or the recondition policies of CO_2 emissions are temporary over time, and series will revert to a trend path in the long run. As such, controlling the mean value of trend paths in the long run is a crucial target of all countries rather than creating a transitory reduction in the short run (Chang & Lee, 2008).

Aside from stochastic convergence, αconvergence indicates a reduction in the spread or dispersion of a data set over time, whereas β convergence is supported by the existence of a negative relationship between the growth rate of the variable of interest and its initial level.[4] In addition, the cases of conditional and unconditional convergence can ascertain whether convergence takes place after controlling for country-specific characteristics, which helps determine differences in steady state emission levels. We first plotted the log of RPCE for each country under scrutiny to provide a visual inspection. As seen in Figure 1, there is a gradual narrowing of cross-country disparities in RPCE across OECD countries, indicating a converging pattern.

[4] However, β convergence has been criticized by many authors. For example, Quah (1993) asserted that cross-country regression in β convergence test assumes that all countries have the same convergence rate. Also, Quah (1996, 1997) argued that it is uninformative for a distribution's dynamics because they only capture representative economy dynamics and suggests a dynamic distributional approach to convergence analysis. Furthermore, Romero-Avila (2008) stated that conditional β convergence, as a form of cross-sectional convergence, represents a much weaker notion of convergence than stochastic convergence.

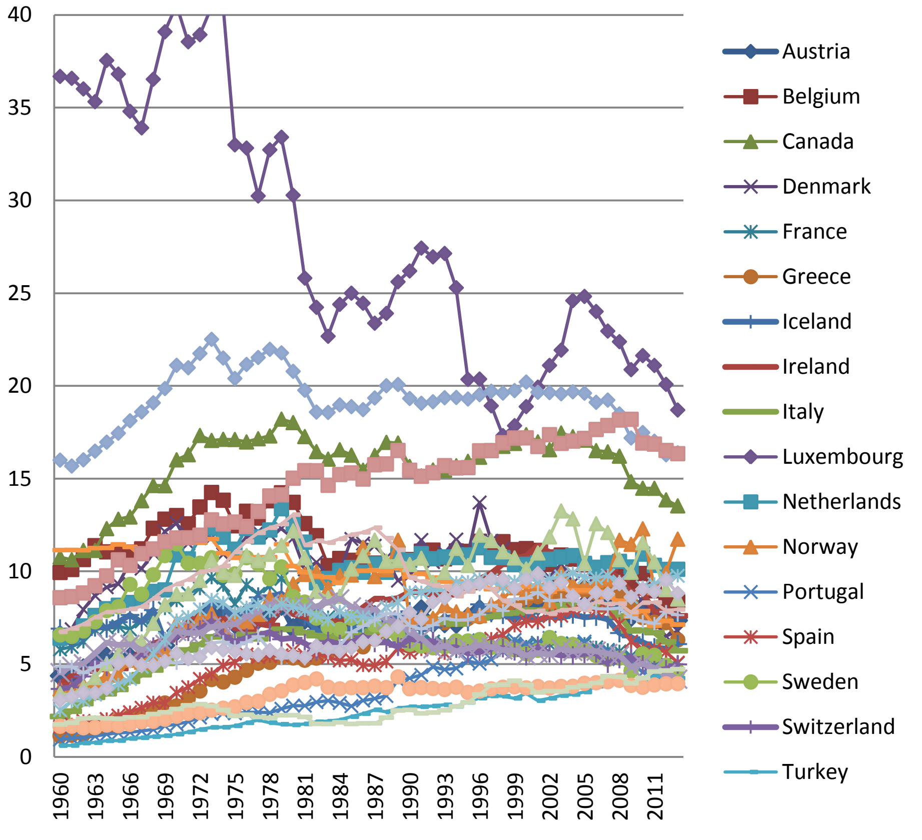

Figure 1. Plot of logarithm of RPCE for 28 OECD countries

However, a visual inspection is not enough to come to a clear-cut conclusion about whether convergence holds. Thus, we need to conduct a formal test using unit root tests. In addition, we decided to allow for breaks in the unit root testing process due to the following reasons. First, our data spanned over 40 years and coincided with important events that created crucial shifts in the trend ways of emissions series such as the first and second oil crises in the 1970s, the Earth Summit in 1992, and the Kyoto Protocol in 1997. In addition, structural breaks can be derived from changes in the degree of environmental-control legislation, changes in the political system, or fluctuations in energy prices. For instance, Romero-Avila (2008) suggested that the 1970 and 1990 amendments to the Clean Air Act, the Council Directive on Limitation of Emissions of Certain Pollutants from Large Combustion Plant, and a final amendment affecting European Union countries were the main policy intervention shocks that may have created significant breaks in the trend of RPCE levels.

3.2. Methodology

We allowed for structural breaks in unit root tests based on the idea of Perron (1989). Perron (1989) stated that if there is a structural break, the power to reject a unit root decreases when the stationary alternative is true and the structural break is ignored. After that, many scholars, such as Zivot and Andrews (1992), Lumsdaine and Papell (1997), Vogelsang and Perron (1998), and Lee and Strazicich (2003, 2004), have developed different unit root tests to account for structural breaks. However, we employed recently developed univariate unit root tests, the two-step LM test and the three-step RALS-LM test with breaks.

We can explain the methodologies of the tests employed based on the studies of Lee et al. (2012), Meng and Lee (2012), and Meng et al. (2013). First, based on the unobserved component 0representation, the following data generating process (DGP) was considered:

$$y_t = \delta^{'Z_t} + e_t , \quad e_t = \beta e_{t-1} + \varepsilon_t \tag{2}$$

where Z_t includes exogenous variables. If $Z_t = [1, t, D_t, DT_t^*]'$, a more general model that allows for both level and trend shifts is obtained. Additional dummy variables could also be included to take multiple breaks into consideration such that:

$$Z_t = [1, t, D_{1t}, \ldots \ldots, D_{Rt}, \; DT_{1t}^*, \ldots \ldots, \; DT_{Rt}^*]', \tag{3}$$

where $D_{it}^* = 1$ for $t \geq T_{Bi} + 1$, $i = 1, \ldots.., R$, and zero otherwise; and $DT_{it}^* = t - T_{Bi}$ for $t \geq T_{Bi} + 1$ and zero otherwise. T_{Bi} represents the break date. Based on the LM (score) principle, the null restriction of β=1 was imposed, and the equation (4) was considered as a first step:

$$\Delta y_t = \delta' \Delta Z_t + u_t \tag{4}$$

where $\delta = [\delta_1, \delta_2, \delta_{3i}', \delta_{4i}']', i = 1, \ldots., R.$ The unit root test statistics were obtained from the following regression:

$$\tilde{S}_t = y_t - \tilde{\psi} - Z_t \tilde{\delta} \tag{5}$$

where $\tilde{S}_t$ represents the de-trended series as

$$\tilde{S}_t = y_t - \tilde{\psi} - Z_t \tilde{\delta} \tag{6}$$

Here, the coefficient $\tilde{\delta}$ was obtained in equation (4) using the first differenced data and $\tilde{\psi} = y_1 - Z_1 \tilde{\delta}.$ In doing so, the dependency on nuisance parameters was removed from the crash model. However, the dependency on nuisance parameters in the trend-break model cannot be removed through this de-trending procedure. The usual LM tests for the trend-break model depend on λ_i^*, which indicates the fraction of sub-samples in each regime so that $\lambda_1^* = T_{B1} / T$, $\lambda_i^* = (T_{Bi} - T_{Bi-1}) / T$, $i = 2, \ldots\ldots, R,$ and $\lambda_{R+1}^* = (T - T_{BR}) / T.$ However, as stated in Lee et al. (2012), the following transformation removes the dependency of the test statistic on the nuisance parameter.[5]

$$\begin{cases} \frac{T}{T_{B1}}\tilde{S}_t, & for\ t \leq T_{B1} \\ \frac{T}{T_{B2}-T_{B1}}\tilde{S}_t & for\ T_{B1} < t < T_{B2} \\ \frac{T}{T-T_{BR}}\tilde{S}_t & for\ T_{BR} < t < T \end{cases} \tag{7}$$

Then, $\tilde{S}_{t-1}$ in equation (5) is replaced with $\tilde{S}^*_{t-1}$ as follows:

$$\Delta y_t = \delta' \Delta Z_t + \varphi \tilde{S}^*_{t-1} + \sum_{j-1}^{k} d_j \Delta \tilde{S}_{t-j} + e_t \tag{8}$$

where $\tilde{\tau}^*_{LM}$ is the t-statistic for $\varphi = 0$. In this case, the unit root test statistic $\tilde{\tau}^*_{LM}$ no longer depends on the nuisance parameter λ^*_i in case of the trend break model. Following the transformation, as the distribution is given as the sum of $R+1$ independent stochastic terms, the asymptotic distribution of $\tilde{\tau}^*_{LM}$ depends only on the number of trend breaks. In case of one trend break ($R=1$), the distribution of $\tilde{\tau}^*_{LM}$ is the same as that of τ_{LM} (untransformed test) using $\lambda = 1/2$, irrespective of the initial location of break(s). Similarly, in the case of two trend-break cases (R=2), the distribution of $\tilde{\tau}^*_{LM}$ is the same as that of untransformed test (τ_{LM}) using $\lambda_1 = 1/3$ and $\lambda_2 = 2/3$. Overall, the same analogy holds for the case of R multiple breaks; the distributions of transformed tests and untransformed tests are the same using $\lambda_i = i/(R+1), i = 1, \dots \dots, R$. Thus, it is not necessary to simulate new critical values at all possible break point combinations. Instead, we only needed critical values that correspond to the number of breaks (R).

In addition, we also employed the RALS-LM unit root test, which was developed by Meng and Lee (2012), to utilize the information on non-normal errors and to further improve the power of the transformed LM test statistic $\tilde{\tau}^*_{LM}$. In the RALS procedure, the following term, $\widetilde{w}_t$, was augmented to equation (8).

$$\widetilde{w}_t = h(\hat{e}_t) - \widehat{K} - \hat{e}_t \widehat{D}_2 \tag{9}$$

where $h(\hat{e}_t) = [\hat{e}_t^2, \hat{e}_t^3]'$, $\widehat{K} = \frac{1}{T}\sum_{t-1}^{T} h(\hat{e}_t)$ and $\widehat{D}_2 = \frac{1}{T}\sum_{t-1}^{T} h'(\hat{e}_t)$

The second and third moments of $\hat{e}_t$ were included as $(\hat{e}_t) = [\hat{e}_t^2, \hat{e}_t^3]'$ to capture information of non-normal errors. Then, by letting $\widehat{m}_j = T^{-1}\sum_{t-1}^{T} \hat{e}_t^i$, the augmented term could be defined as such:

$$\widehat{w}_t = [\hat{e}_t^2 - \widehat{m}_2,\ \hat{e}_t^3 - \widehat{m}_3 - 3\widehat{m}_2 \hat{e}_t]'. \tag{10}$$

The augmented terms were obtained from the redundancy condition that implies knowledge of higher moments m_{j+1} are uninformative if $m_{j+1} = \mathrm{j}\sigma^2 m_{j-1}$; augmented terms were obtained from the redundancy condition. The redundancy condition was only satisfied with normal distribution. In case of non-normal

[5] Please see Lee et al. (2012), Meng and Lee (2012), Meng et al. (2013) and Ozcan and Erdogan (2015) for a detailed explanation for procedures and steps of the LM and RALS-LM unit root tests.

distributed error terms, this condition was not satisfied, and therefore, the efficiency may be increased by augmenting equation (8) with $\hat{w}_t$. The transformed RALS-LM test statistic was then obtained in equation (11):

$$\Delta y_t = \delta' \Delta Z_t + \varphi \tilde{S}^*_{t-1} + \sum_{j-1}^{p} d_j \Delta \tilde{S}_{t-j} + \widehat{w'}_t \gamma + u_t \quad (11)$$

The corresponding t-statistic for $\varphi = 0$ is denoted by $\tau^*_{RALS-LM}$. The asymptotic distribution of $\tau^*_{RALS-LM}$ does not depend on the break location parameter. Thus, we did not need to simulate new critical values for all possible break location combinations. The critical values were provided in Meng et al. (2012) for different numbers of observation for R=1,2 and $\rho^2 = 0$ to 1

4. Empirical Results

The model with at most two level and trend breaks was considered in the study. In the first step of the two-step LM unit root test, a maximum structural break number R was defined, and the max F test was applied to identify the break locations and to test the significance of each break with optimal lags. We turned back to the beginning of the first step with break numbers equal to $R-1$, when the null of no trend break is not rejected or when one of the break dummy variables is not significant in case of the rejection of no trend break. This procedure continues until the break number equals to zero, or all the identified break dummy variables were significant. The usual no-break LM unit root test of Schmidt and Phillips (1992) was employed if the break number was equal to zero from the first step; however, the one-break (or R-break) LM unit root test of Amsler and Lee (1995) and Lee and Strazicich (2003) was employed in case of one or more breaks with the break number, location, and the corresponding optimal lags identified in the first step. After that, the LM statistic, $\tilde{\tau}^*_{LM}$, is obtained. Regarding the RALS-LM test, its first two steps were the same as the two-step LM test. In the third step, the higher moment information attained from the second step was used and augmented to the regression of the two-step LM test. RALS-LM test statistic was denoted as $\tau_{RALS-LM}$. The grid search within 0.10–0.90 intervals of the whole sample period was used while searching for the optimal number of breaks.

As a preliminary check, we first employed Augmented Dickey-Fuller (ADF) test of Dickey and Fuller (1981), the LM test of Schmidt and Phillips (1992), and the RALS-LM test without breaks. The results are reported in Table 1.

As seen in Table 1, the unit root null hypothesis was rejected for six countries (Austria, Canada, Denmark, Iceland, Italy, and Switzerland) in the ADF test. For the other 22 countries, RPCE appeared to be stationary. In the case of the LM test, RPCE was stationary for only three countries (Austria, Canada, and Denmark); however, the null of nonstationarity could not be rejected for the remaining 25 OECD countries. Finally, for the RALS-LM test, RPCE is stationary for only six countries (Austria, Belgium, Canada, Denmark, Finland, and Israel), whereas the null of nonstationarity could not be rejected for the remaining 22 OECD countries.

Table 1. Results of ADF, LM and RALS-LM tests without break

Country	ADF		LM	RALS-LM			Critical Values		
	τ_{ADF}	$\hat{k}$	τ_{LM}	$\tau_{RALS-LM}$	$\hat{\rho}^2$	$\hat{k}$	1%	5%	10%
Austria	-4.993[a]	0	-5.027[a]	-4.760[a]	0.997	0	-3.691	-3.087	-2.790
Australia	-2.103	0	-1.829	-2.204	0.957	0	-3.674	-3.066	-2.768
Belgium	-2.706	0	-2.680	-3.240[b]	0.829	0	-3.608	-3.002	-2.699
Canada	-3.659[b]	0	-3.497[b]	-3.552[a]	0.711	0	-3.535	-2.937	-2.635
Chile	-1.441	1	-1.582	-1.861	0.983	1	-3.685	-3.079	-2.782
Denmark	-4.201[a]	0	-2.882[c]	-2.978[b]	0.907	0	-3.653	-3.041	-2.741
Finland	-2.806	0	-1.677	-2.841[c]	0.701	0	-3.530	-2.932	-2.630
France	-3.040	0	-2.010	-2.111	0.946	0	-3.670	-3.061	-2.762
Greece	-1.532	3	-1.267	-0.816	0.914	0	-3.656	-3.044	-2.745
Hungary	-2.268	2	-2.009	-1.323	0.968	2	-3.679	-3.072	-2.774
Iceland	-4.259[a]	0	-1.716	-2.474	0.672	1	-3.514	-2.911	-2.607
Ireland	-1.853	0	-2.287	-1.857	0.960	0	-3.675	-3.068	-2.770
Israel	-2.158	0	-2.208	-2.746[c]	0.465	0	-3.362	-2.747	-2.425
Italy	-4.027[b]	0	-0.349	-0.033	1.022	1	-3.692	-3.088	-2.791
Japan	-2.751	2	-1.182	-0.927	0.947	2	-3.670	-3.061	-2.763
Luxembourg	-1.925	1	-2.233	-2.256	0.824	1	-3.605	-2.999	-2.696
Mexico	-1.181	0	-1.227	-1.367	0.994	0	-3.690	-3.085	-2.788
Netherlands	-2.882	0	-1.826	-1.336	0.658	0	-3.507	-2.900	-2.595
New Zealand	-2.587	0	-1.360	-1.993	0.905	0	-3.653	-3.040	-2.740
Norway	-1.333	1	-1.125	-1.063	1.044	1	-3.692	-3.088	-2.791
Poland	-1.292	0	-1.243	-0.955	0.876	1	-3.636	-3.025	-2.724
Portugal	-0.660	0	-1.675	-1.602	1.019	3	-3.692	-3.088	-2.791
Spain	-0.656	0	-1.469	-0.909	0.732	0	-3.548	-2.949	-2.646
Sweden	-2.664	0	-1.121	-2.202	0.642	1	-3.499	-2.888	-2.582
Switzerland	-4.439[a]	0	-2.007	-1.459	0.998	1	-3.692	-3.087	-2.790
Turkey	-2.744	0	-2.387	-2.355	0.923	0	-3.660	-3.049	-2.750
UK	-2.902	4	-1.343	-1.165	0.716	4	-3.539	-2.940	-2.638
United States	-1.846	4	-0.642	-0.521	0.776	4	-3.575	-2.973	-2.670

Notes: $\hat{k}$ is the optimal number of lagged first-differenced terms and we didn't report $\hat{k}$ values again because they are same both for the LM and RALS-LM tests. τ_{ADF} indicates the augmented Dickey-Fuller test statistic, whereas τ_{LM} and $\tau_{RALS-LM}$ are the test statistics of the LM and RALS-LM tests, respectively. The critical values for the ADF test are -4.159, -3.501, and -3.183 at 1%, 5%, and 10% significance levels, respectively. The critical values for the LM test with no break are -3.693, -3.088, and -2.792 at 1%, 5% and 10% levels, respectively. Also, the critical values for the RALS-LM test are tabulated in the last three columns in Table 1. The critical values were obtained for 50 observation numbers for all tests through 100000 iterations. The optimal number of lags $\hat{k}$ is chosen using a general to specific method with the maximum lags equal to four. [a], [b], and [c] indicate significance at 1%, 5% and 10% levels, respectively.

Overall, the unit root tests without breaks generally provided evidence of divergence in RPCE series. However, because allowing for breaks in unit root testing procedure may change the results, we employed the LM and RALS-LM tests with breaks. The results of the tests with two breaks are provided in Table 2. As displayed in Table 2, in the case of the LM test, the null of nonstationarity was rejected for 20 OECD countries, whereas only eight countries (Iceland, Italy, Luxemburg, Mexico, Poland, Spain, Turkey, and the UK) have nonstationary RPCE series. Besides, in the case of the RALS-LM test, the null of nonstationarity could be rejected for 22 countries. Therefore, it could be suggested that allowing for breaks changed the results from divergence to convergence.

Table 2. Results of LM and RALS-LM unit root tests with two breaks

Country	LM	RALS-LM					Critical Values		
	τ_{LM}^{*}	$\tau_{RALS-LM}^{*}$	$\hat{\rho}^2$	$\hat{T}_B$		$\hat{k}$	1%	5%	10%
Austria	-6.204^{a}	-6.204^{a}	1.017	1979	1993	4	-5.005	-4.370	-4.058
Australia	-4.986^{a}	-4.986^{a}	0.992	1978	1989	4	-4.955	-4.321	-4.011
Belgium	-4.062^{c}	-4.062^{c}	0.925	1981	1984	0	-4.777	-4.161	-3.853
Canada	-6.053^{a}	-6.053^{a}	0.727	1975	1996	0	-4.617	-4.004	-3.695
Chile	-4.647^{b}	-4.647^{b}	0.796	1986	2002	1	-4.987	-4.353	-4.041
Denmark	-6.445^{a}	-6.445^{a}	1.044	1989	1997	3	-4.891	-4.259	-3.952
Finland	-6.312^{a}	-6.312^{a}	0.826	1979	2003	1	-4.605	-3.991	-3.682
France	-5.348^{a}	-5.348^{a}	0.724	1982	1988	0	-4.941	-4.307	-3.998
Greece	-5.734^{a}	-5.734^{a}	0.627	1970	1991	4	-4.900	-4.268	-3.960
Hungary	-4.180^{c}	-4.180^{c}	0.702	1988	1992	1	-4.969	-4.335	-4.024
Iceland	-3.737	-3.737^{c}	1.041	1967	1970	2	-4.564	-3.948	-3.638
Ireland	-5.570^{a}	-5.570^{a}	1.014	1973	2000	0	-4.958	-4.324	-4.014
Israel	-4.408^{b}	-4.408^{a}	0.871	1967	1990	4	-4.224	-3.608	-3.288
Italy	-2.699	-2.699	1.032	1967	2000	0	-5.008	-4.373	-4.061
Japan	-4.303^{c}	-4.303^{c}	1.025	1965^{n}	1970	0	-4.942	-4.308	-3.999
Luxembourg	-3.663	-3.663	0.909	1993	1998	4	-4.770	-4.154	-3.847
Mexico	-3.796	-3.796	0.946	1972	1983	0	-5.001	-4.366	-4.054
Netherlands	-6.108^{a}	-6.108^{a}	0.882	1980	1983	0	-4.545	-3.927	-3.617
New Zealand	-4.380^{b}	-4.380^{b}	0.927	1978	1993	3	-4.889	-4.257	-3.950
Norway	-5.004^{b}	-5.004^{b}	0.785	1988	1998^{n}	3	-5.008	-4.373	-4.061
Poland	-2.670	-2.670	0.936	1987	1991	0	-4.846	-4.220	-3.913
Portugal	-5.551^{a}	-5.551^{a}	0.971	1984	2000	4	-5.008	-4.373	-4.061
Spain	-3.051	-3.051	0.715	1968^{n}	1981	3	-4.644	-4.032	-3.723
Sweden	-5.568^{a}	-5.568^{a}	0.701	1976	1995^{n}	4	-4.522	-3.902	-3.592
Switzerland	-5.180^{a}	-5.180^{a}	0.802	1972	1983	0	-5.006	-4.371	-4.059
Turkey	-3.853	-3.853	0.976	1999	2004	3	-4.912	-4.280	-3.971
UK	-3.714	-3.714^{c}	0.745	1975	1989	3	-4.624	-4.011	-3.702
United States	-5.820^{a}	-5.820^{a}	0.991	1978^{n}	1996	2	-4.702	-4.091	-3.783

Note: τ_{LM}^{*} and $\tau_{RALS-LM}^{*}$ are the test statistics for the LM and RALS-LM tests, respectively. $\hat{k}$ is the optimal number of lagged first-differenced terms. $\hat{T}_B$ denotes the estimated break points. The test statistics are invariant to the location of trend breaks because transformed tests are implemented. ^{a}, ^{b} and ^{c} denote the significance of the test statistic at 1%, 5% and 10% levels, respectively. As the LM test and RALS-LM test share the same procedure when searching for the break points and the corresponding optimal lags, the break dates were reported only one time to save the space. The optimal number of lags $(\hat{k})$ is chosen using a general to specific method with the maximum lags equal to four. The critical values for the LM test are -5.008, -4.373 and -4.0613 at 1%, 5%, and 10% levels, respectively and were obtained through 100000 iterations for 50 observation. The last three columns provide critical values for the RALS-LM test. ^{n} denotes insignificance of break dummy.

The results of the LM and RALS-LM tests were similar. Furthermore, for Japan, Norway, Spain, Sweden, and the United States, one of the breaks identified for each country was insignificant at the 10% level. Therefore, a one-break unit root test appeared to be more appropriate for these countries. However, we employed additional one-break tests for all countries to examine the impact of including two breaks instead of one; the results are reported in Table 3.

Table 3. Results of LM and RALS-LM unit root tests with one break

Country	LM	RALS-LM				Critical Values		
	τ^*_{LM}	$\tau^*_{RALS-LM}$	$\hat{\rho}^2$	$\hat{T}_B$	$\hat{k}$	1%	5%	10%
Austria	-2.797	-2.605	0.941	2004	2	-4.349	-4.349	-4.349
Australia	-3.480[c]	-3.366	0.994	1979	4	-4.403	-4.403	-4.403
Belgium	-2.907	-3.124	0.901	2000	0	-4.310	-4.310	-4.310
Canada	-4.401[b]	-4.568[a]	0.827	1996	0	-4.231	-4.231	-4.231
Chile	-3.052	-2.859	0.996	1986	1	-4.405	-4.405	-4.405
Denmark	-3.834[b]	-4.641[a]	0.704	1970	0	-4.110	-4.110	-4.110
Finland	-3.385	-3.979[b]	0.937	1980	0	-4.346	-4.346	-4.346
France	-5.343[a]	-6.092[a]	0.700	1983	0	-4.106	-4.106	-4.106
Greece	-2.889	-2.659	0.907	1988	0	-4.315	-4.315	-4.315
Hungary	-3.389	-2.956	0.926	1986	2	-4.335	-4.335	-4.335
Iceland	-4.032[b]	-4.107[b]	0.701	1984	0	-4.108	-4.108	-4.108
Ireland	-3.124	-3.435[c]	0.869	2000	0	-4.275	-4.275	-4.275
Israel	-2.202	-2.346	0.560	1996	0	-3.954	-3.954	-3.954
Italy	-2.464	-2.560	0.972	1967	1	-4.381	-4.381	-4.381
Japan	-3.466	-3.047	1.032	1972	0	-4.409	-4.409	-4.409
Luxembourg	-2.775	-2.271	0.909	1998	4	-4.317	-4.317	-4.317
Mexico	-1.692	-2.215	0.926	1965	0	-4.335	-4.335	-4.335
Netherlands	-5.211[a]	-5.177[a]	0.913	1980	2	-4.321	-4.321	-4.321
New Zealand	-3.772[c]	-3.748[c]	0.989	1983	2	-4.398	-4.398	-4.398
Norway	-4.089[b]	-3.545	0.911	1988	2	-4.320	-4.320	-4.320
Poland	-2.606	-2.218	0.982	1986	2	-4.391	-4.391	-4.391
Portugal	-6.043[a]	-6.276[a]	0.830	1997	3	-4.234	-4.234	-4.234
Spain	-1.903	-2.747	0.507	1981	0	-3.892	-3.892	-3.892
Sweden	-3.170	-3.851[b]	0.712	1978	0	-4.118	-4.118	-4.118
Switzerland	-4.220[b]	-4.071[b]	0.996	1989	0	-4.405	-4.405	-4.405
Turkey	-3.804[b]	-3.997[b]	0.725	1989	2	-4.130	-4.130	-4.130
UK	-1.871	-1.582	0.731	1964	4	-4.136	-4.136	-4.136

Notes: τ^*_{LM} and $\tau^*_{RALS-LM}$ are the test statistics for the LM and RALS-LM test, respectively. k is the optimal number of lagged first-differenced terms. The optimal number of lags is chosen using a general to specific method with the maximum lags equal to four. $\hat{T}_B$ denotes the estimated break points. The test statistics are invariant to the location of trend breaks because transformed tests are implemented. [a], [b] and [c] denote the significance of the test statistic at 1%, 5% and 10% levels, respectively. Since the LM test and RALS-LM test share the same procedure when searching for the break points and the corresponding optimal lags, the break dates were reported only one time to save the space. The critical values for the LM test are -4.409, -3.780 and -3.483 at 1%, 5%, and 10% levels, respectively and obtained through 100000 iterations for 50 observation. The last three columns provide critical values for the RALS-LM test.

As tabulated in Table 3, compared to 20 and 22 rejections for the unit root tests in the case of two breaks, 12 and 13 countries rejected the null hypothesis of a unit root in the cases of one-break LM and RALS-LM tests, respectively. In case of the LM test for Australia, Canada, Denmark, France, Iceland, Netherlands, New Zealand, Norway, Portugal, Switzerland, Turkey, and the United States, convergence in RPCE was confirmed as they had stationary series. In the case of the RALS-LM test, the unit root null hypothesis was rejected, i.e. convergence does hold for Canada, Denmark, Finland, France, Iceland, Ireland, Netherland, New Zealand, Portugal, Sweden, Switzerland, Turkey, and the United States. All of the break dummies were significant at the 10% level or better in one break case.

Concerning the break dates, the breaks, as stated by Romero-Avila (2008) and Lee and Chang (2009), are related to the relative emission series; thus, we could not define the breaks for each individual country or related events. Based on the definition of the emissions data, a break in the original series may appear as a break in all 28 individual relative series unless the breaks for individual series are ruled out with the breaks exhibited by the average emission series. Therefore, we abstained from detailed explanations of break dates for each country. However, in general, the first break dates of 20 countries occurred in the period of 1965–1982. In the 1960s, the modern environmental movement began with concerns about air and water pollution such as the Clean Air Acts of 1963 and 1965. The first breaks of Ireland, Iceland, Israel, Italy, and Japan occurred in this period. In addition, two major oil crises from 1970–1982 caused some shocks in RPCE because fossil fuels became the main source of productivity in 1970s due to higher oil prices.

Furthermore, as Lee and Chang (2008) noted, major technological and structural breaks such as the development of nuclear power reduced the demand for oil and led to decreases in CO_2 emissions from the 1970s onward. As such, the first break points of Austria, Australia, Belgium, Finland, France, Netherlands, New Zealand, and the United States occurred during the second oil crisis, whereas the first break dates for Canada, Ireland, Mexico, Sweden, Switzerland, and the UK were related to the first oil crisis.

In particular, the growing international concerns about environmental issues have been discussed among industrial countries since 1988. This environmental consciousness has led developed countries to create many environmental agreements and agendas such as the UNFCCC in 1992, Habitat II (the Second United Nations Conference on Human Settlements) in 1996, and the Kyoto Protocol in 1997. Thanks to these developments, the GHG emission rates, especially CO_2, have started to decrease. In this context, after 1988, 20 OECD countries (Austria, Australia, Canada, Chili, Denmark, Finland, France, Greece, Hungary, Ireland, Israel, Italy, Luxemburg, Norway, Poland, Portugal, Sweden, Turkey, the UK, and the United States) experienced the second break in the RPCE series. The third energy crisis, which was related to the Gulf War in 1990–91, had some influences in RPCE for Greece, Hungary, Israel, and Poland, whereas some OECD countries experienced breaks in RPCE because of the recession of 1981–1983. For instance, the first break dates of Belgium, France, and Mexico and the second break dates of Netherlands, Spain, and Switzerland occurred during the recession period. Furthermore, breaks occurred in six countries (Chile, Finland, Ireland, Italy, Portugal, and Turkey) during the 2000s, a period that included the Western energy crisis and major environmental summits such as the United Nations Millennium Summit in 2000 and the United Nations Earth Summit in 2002.

Overall, we should depend on the results of unit root tests with breaks because all 28 OECD countries analyzed in this study contained one or two breaks, and all breaks identified with one-break tests were significant at the 10% level. Therefore, our empirical findings provided significant support for convergence in RPCE among OECD countries. Allowing for structural breaks in testing for stochastic convergence was crucial given that the results changed from divergence to convergence when breaks were involved. Our results, which were mostly favorable to the stochastic conditional convergence, accorded well with those of Romero-Avila (2008), Chang and Lee (2008), Lee et al. (2008), Lee and Chang (2009), Strazicich and List (2003), and Westerlund and Basher (2007, 2008).

In addition, the results mostly favorable to stationarity have crucial economic and policy implications. The stationary RPCE series indicate that shocks to CO_2 emissions series will have only transitory impacts, implying that RPCE will return to its original equilibrium (to the mean emissions level of OECD) over a short period of time after being hit by a major global structural change or shock. Thus, the long-run projection models for CO_2 emissions could be formulated as an emission abatement strategies to combat climate change. Also, CO_2 emissions forecasting can be used as an appropriate climate policy response because it is possible to forecast future movements in stationary RPCE based on its past behavior. Besides, as relative CO_2 emissions series only temporarily deviate from the sample (OECD) mean for most countries, the governments of OECD countries should not interfere excessively with countries that reach convergence in their RPCE series. In other words, the government's administrative policy should pay attention to the long-run trends in CO_2 emissions instead of adopting unnecessary targets

5. Conclusion and Policy Implications

The issue of convergence in CO_2 emissions is crucial for the aforementioned reasons. We implemented the two-step LM and three-step RALS-LM unit root tests to test for the stochastic conditional convergence for 28 OECD countries within the period of 1960–2013. As a preliminary inspection, we first employed unit root tests without breaks and came to the conclusion that emissions convergence was not generally supported for the OECD countries being considered. We then employed the two-break LM and RALS-LM unit root tests. In case of two breaks, 23 countries had significant break dummy variables, and the stationarity did hold for the 16 and 18 OECD countries in LM and RALS-LM unit root tests, respectively. Furthermore, among the five countries where the one-break unit root test was more appropriate, three countries have stationary series. Based on these results, it can be concluded that stochastic conditional convergence held in most OECD countries under study.

Furthermore, our stationarity results indicated that following a major structural change, RPCE series will revert to sample (OECD) mean over a period of time (Lee & Chang, 2009). In other words, RPCE series for most OECD countries doesn't diverge from mean emissions level of OECD in the long-run. Besides, the following policy implications might be suggested based on these findings. First, as asserted by Romero-Avila (2008), targeting per capita emissions stabilization and reduction is an effective way to combat global climate change and the greenhouse effect. However, stationarity means that the effects of CO_2 emissions reduction or the recondition policy are temporary over time, and series will revert to a trend path in the long run. Thus, controlling the mean value of trend path in the long run is a crucial target of all countries rather than being a transitory reduction in the short run (Chang & Lee, 2008).

Second, the validity of convergence in per capita emissions among OECD countries convinces developing countries, such as China and India, to accept emissions abatement obligations. Also, as suggested by Westerlund and Basher (2008), the fulfillment of certain emission goals and commitments by developed countries is a necessary condition for the application of multilateral climate change agreements like the Kyoto Protocol. Third, stationarity indicates that it is possible to forecast future movements in RPCE series by examining its past behavior. Finally, from the viewpoint of government policy, government-designed energy use or environmental protection policies do not have persistent impacts on the RPCE series of OECD countries. Thus, as stated by Lee and Chang (2009), when CO_2 emissions temporarily deviate from the trend path, the administrative policy of the government should not adopt unnecessary targets because policy actions are not required to return CO_2 emissions to its trend path.

Overall, our results highlight the fact that taking structural breaks in CO_2 emissions level is so crucial in the analysis of convergence issue. Due to shocks in global energy markets, CO_2 emissions series are subject to structural breaks in their mean and/or trend levels. If we ignore those breaks in unit root testing procedure, the divergence hypothesis gains empirical validity for CO2 emissions among OECD countries. However, this result appears to cast shadow over reality. As such, CO2 emissions series converge to the mean emissions level of OECD sample through mean or/and trend breaks in an increasing or a decreasing way.

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