RESEARCH ARTICLE

The dynamic role of institutional quality, renewable and non-renewable energy on the ecological footprint of OECD countries: do institutions and renewables function as leverage points for environmental sustainability?

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Abstract

This article examines for the first time the impact of disaggregated energy sources and institutional quality on the ecological footprint (EF) of 29 OECD countries, by explaining how the diversification in countries' energy mix and their institutional performance are associated with sustainable environmental performance. We use panel data from 1984 to 2016 and we apply second-generation techniques to arrange the critical issues of cross-sectional dependence and heterogeneity. The applied cointegration tests expose a long-run equilibrium relationship that associates renewable/non-renewable energy consumption, economic growth, institutional quality, and the EF of OECD countries. The robust cross-sectional augmented distributed lag (CS-DL) estimator shows that economic growth and the adoption of non-renewable energies are detrimental to the environment, while the operational quality of institutions adds to ecological sustainability. Concurrently, the negative effect of renewables on EF does not seem to cause a significant beneficial impact on the environment. Moreover, there is evidence that non-renewable energy and institutional quality have a bidirectional causal association with EF. Also, a weak unidirectional causal effect is running from the EF to renewables consumption. The study further demonstrates the inefficient integration of renewable energy forms in OECD countries and the concomitant essential role of institutions on environmental sustainability by providing relevant policy orientations.

Keywords Ecological footprint, · Renewable energy, · Institutional quality, · Cointegration · Panel data analysis, · Cross-sectional dependence

Introduction

Environmental sustainability constitutes a fundamental component of the process of sustainable development, along with the economic and social sustainability. The concept of environmental sustainability was first developed by Goodland

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(1995), who defines it as the ability to ensure human well-being through the maintenance of natural capital. This conceptualization was further formulated in the early twenty-first century, when OECD environmental strategy was adopted by OECD countries (OECD 2001). The strategy introduced four criteria for a sustainable environment. These criteria are the regeneration of natural resources, the substitutability capacity of fossil energy sources, the assimilation of waste, and the avoidance of irreversible adverse effects on ecosystems.

The operationalization of the former criteria for environmental sustainability is imperative for policymakers in order to eliminate climate change, global warming, pollution, and depletion of resources. Since economic growth is directly associated with environmental survival (Sharma 2011), joint policy programs tend to integrate environmental policies into the economic ones in the policy formulation arena (Jordan and Lenschow 2008). In this bidimensional policy framework, this

study intends to probe the concomitant role of institutional quality and dissociated energy forms on the ecological footprint (EF) of OECD countries.

In the energy economics literature, the use of energy is a central determinant in fostering environmental efficiency (Charfeddine and Mrabet 2017). Efficient and clean energy use contributes to efficient and environmentally friendly exploitation of resources and effective environmental management (Nathaniel et al. 2020). Renewable energy (RE) exerts a significant role in national energy policies as it provides secure energy, expands countries' energy autonomy, increases employment, and combats environmental problems (IRENA and REN21 2018; Liu et al. 2020a, 2020b). Although globally the portion of renewables[1] on the total supply of primary energy was 12.9% in 2008, most of the predictions assume a 17% share in 2030 and a 27% share in 2050[2] (IEO 2019). The electricity sector faces the largest and fastest increase in the use of renewables since the share of RE that covered needs in electrical power reached 24% in 2016. This was more than a double growth from 2015 and the fastest one since 1990 (IEO 2019). It is estimated that due to technological innovations and government RE policies, in 2050, renewables will account for nearly half of the world's total electricity production.

On the other hand, the dominant role of fossil fuels in the majority of developing and developed countries and the increasing public awareness about ecological sustainability intensify the implementation of effective energy policies (Fudge et al. 2008). Fossil energy is considered the main culprit of anthropogenic emissions and the acceleration of clean energy policies is vital for the innovation of the energy system (Edenhofer et al. 2011). However, the level of renewables penetration in countries' energy mix is strongly correlated with the governmental support policies (IRENA and REN21 2018). The high cost of renewable technology implementation may hinder this penetration for the sake of short-run fast economic growth, as a result of a policy level myopia (Sharif et al. 2020b).

The implementation of effective renewable energy policies on the route for a sustainable environment presumes a robust institutional framework (Panayotou 1997; Bhattarai and Hammig 2001). The administrative and institutional settings enable and support environmental policies, not only at the policymaking level but also at the implementation and monitoring level (Lovei and Weiss 1998). Furthermore, effective institutions may influence the rate of renewables penetration in the countries' energy mix (Bhattacharya et al. 2017), thus causing an indirect effect on the ecological protection. Iacobuta and Gagea (2010) argue that the main obstacle to sustainable development is not insufficient resources, but

inadequate institutional operation. Ostrom (1998) states that institutional quality reflects the governments' structure and effectiveness which is formulated through their designed policies and takes place in the environmental regulatory framework. Institutions could be categorized into formal which refer to the economic, political and legal institutions, and informal institutions implying societal norms and constraints, and cultural-cognitive procedures (Dasgupta and De Cian 2016). Although both types formulate the social allocation of incentives, assets, and resources (United Nations 2016), this study adopts the "formal institutions" concept in an extended form that captures the overlapping role of governance.

Effective institutions are instrumental in achieving higher economic growth by minimizing the environmental cost (Panayotou 1997). Governments' capacity to formulate and execute policies and regulations that rule the free market, strengthens the standard of contract performance, security of property rights, rule of law, and political power impartiality (Canh et al. 2019). Governmental stability is crucial for the qualitative performance of institutions and thus for productive environmental management. Political instability can have a negative influence on the rule of law as a result of inefficient governmental performance (Abid 2016). The well-established legal and regulatory framework creates inclusive institutions, enforces a sense of trust to the public, and builds institutional transparency and accountability (Ritzen and Woolcock 2000). Impartial institutions support a strong rule of law accompanied by sound legislation, which has the potential to control the companies' compliance with the respective environmental standards (Lau et al. 2014; Hassan et al. 2020). On the other hand, corruption may be proved to impede environmental quality indirectly. Corrupted governmental officials could promote lax environmental laws and regulations and impede the deployment of RE policies (Fredriksson et al. 2004; Wang et al. 2018). The limited stringency in environmental legislation increases the intensity of energy use, but the existence of governmental stability could regulate the scale of energy consumption (Damania et al. 2003). Besides, institutional corruption and bureaucratic barriers hinder the development and deployment of RE policies by governmental authorities (IRENA and REN21 2018).

Taking the aforementioned into account, this study is motivated by the strong public awareness about environmental issues and the challenge of formulating effective environmental and energy policies under a robust institutional environment. In a holistic approach to environmental sustainability, we intend to explore the dynamic impacts of RE consumption and institutional quality on the EF of 29 OECD countries. The study adds to the relevant literature in several ways.

Firstly, according to our investigation, this study is the first one that engages institutional measures in the energy-environmental framework of OECD countries. Examining the impact of disaggregated energy sources on their

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[1] Mostly bioenergy, direct solar energy, geothermal energy, hydropower, ocean energy, and wind energy.
[2] In the most optimistic scenarios, these shares of renewables will reach 43% in 2030 and almost 77% in 2050.

environment, we highlight the instrumental role of institutions. On that account, we employ a well-defined institutional index that takes into account the fundamental dimensions of institutional performance. Secondly, unlike previous studies, our analysis employs extended institutional data coverage for the time period 1984–2016. Thirdly, we utilize the EF as a proxy for environmental performance in OECD countries. This indicator represents not only the atmospheric pollution, but also the overall ecosystem's damage in a heuristic way. EF encompasses the concept of ecological sustainability and reflects the human demand on nature compared to the biocapacity required for the eco-friendly operation of an economic system. Finally, most of the studies in the relevant literature use first-generation econometric techniques that do not consider the issues of cross-sectional dependency, heterogeneity, endogeneity, and the misspecification of the employed panel data dynamic models. In this work, we mainly employ second-generation tests to address methodological flaws related to these issues. We examine the long-run effects of institutional quality and disaggregated energy in the EF via the recently developed Cross-sectional Augmented Distributed Lag (CS-DL) estimator (Chudik et al. 2016) which produces robust and reliable results. We further apply the Dumitrescu and Hurlin (2012) test to detect causality effects in the panel set. To our knowledge, this study employs for the first time these novel methodological approaches to find out how the institutional environment and the disaggregated energy consumption affect the environmental performance of OECD economies.

Our findings suggest that the institutional quality in OECD countries has a healing environmental effect, while RE consumption seems to produce an insignificant effect on EF. Furthermore, the impact of non-RE consumption and economic growth on the EF is evidenced to be positive and significant. The results are in line with the reports of IRENA and REN21 (2018) and IEA, IRENA, UNSD, WB, WHO (2019) regarding the incomplete integration of renewables in the OECD countries' energy-mix system, and the need for further relevant policy implementation.

The remaining paper has the following structure: The "Environmental sustainability and OECD countries facts and perspectives" section presents some environmental facts and perspectives in the OECD countries. The "The ecological footprint concept" section refers to the concept of ecological footprint, and the "Literature review" section reviews the relevant literature. The "Data and methodology" section describes the data, the model specification, and the methodological approach of the study. The "Empirical findings and discussion" section exhibits and comments on the empirical findings, and finally, the "Conclusions and policy orientations" section provides conclusions along with policy suggestions.

Environmental sustainability and OECD countries facts and perspectives

The increasing environmental threats and the possible irreversible consequences to human existence have facilitated the international community's awareness about a sustainable environment. Global warming and climate change cause devastating effects on the environment, such as ecosystem collapse, glacier retreat and changes in sea level, biodiversity loss, extreme weather phenomena, desertification, and human health threats. In 2015, the CO_2 concentration in the atmosphere was 400.1 ppm (parts per million), while in 2017 it reached 405.5 ppm, which is a 146% increase above the pre-industrial levels (United Nations 2019). Compared with the pre-industrial era, it is estimated that human activities are responsible for a 0.8 to 1.2°C increase in global warming, with the potential to rise above the threshold of 1.5°C in 2030–2052 (IPCC 2018). Anthropogenic emissions have irreversible consequences in the ecosystem's sustainability, putting the biodiversity in jeopardy (United Nations 2019). The continuous decline in environmental quality and the further direct negative consequences to the social and economic dimensions of human life increased countries' environmental awareness. This was recently demonstrated by the ratification of the Paris Agreement, in May 2019. The participants of this agreement committed to deploying nationally determined contributions (NDCs) to eliminate climate change (United Nations 2019).

While many developing countries lay less emphasis on the adoption of renewable energies due to technology and funding constraints, most developed countries have increasingly focused on clean energy policies. Until the late 1980s and during the process of their environmental policy integration (EPI), most OECD countries gradually adopted national and sectoral strategies and environmental projects. In this process, they introduced specialized institutions responsible for the monitoring of sustainable development targets (Jordan and Lenschow 2008). In 2011, OECD countries validated the Green Growth Strategy (GGS) in order to formulate a common policy framework for the promotion of efficient energy targets and environmental sustainable pathways (OECD 2012). Although energy use in non-OECD countries is projected to increase by approximately 70% until 2050, the said increase is estimated at only 15% in OECD countries. Concurrently, until 2050, the energy-induced emissions in OECD countries are predicted to decline by almost 0.2% every year, presenting a 14% decline in comparison with the 2005 levels. (IEO 2019).

OECD countries are considered to have played a pioneering role in the architecture of Sustainable Development Goals (SDGs) (Destek and Sinha 2020). In the process of accomplishing SDG 7 which calls for sustainable energy, they have intensified the use of RE sources to mitigate environmental degradation risks (SDG 13). In OECD Europe,

the wind and solar renewables generation has a potential to increase from a 20% share in the renewables mix to a 50%, from 2018 to 2050. Concurrently, the ratio of fossil fuels generation for the same period is estimated to decline from about 38% to 18% of the generation mix. Furthermore, most of the energy policies in OECD countries are related to the growth in electricity demand which is mainly met by renewables. (IEO 2019).

From the first stages of their environmental management, OECD countries created national and local independent institutions and environmental agencies (Lovei and Weiss 1998). In an early study, Scruggs (1999) scrutinized the important role of policy-oriented institutions to the quality of environmental performance in OECD countries. In addition, Fredriksson et al. (2004) highlighted the negative role of corruption in formulating effective and efficient energy policies in these countries. Destek and Sinha (2020) observed that most of the OECD countries do not face an environmentally sustainable economic growth, implying the existence of ineffective institutional performance.

Industrialized countries such as OECD member states are increasingly interested in the production and use of renewables (Saidi and Omri 2020). The quality of institutions is also considered to play a vital role in their sustainable development process (Khalil et al. 2007). Moreover, OECD countries are mostly engaged in the design, adoption, and execution of SDGs 7, 8, 13, 15, and 16 (UN SDGs 2015). These goals urge the need for sustainable growth by supporting the use of clean energies and the improvement of institutional quality and capacity on environmental performance and ecological protection. Considering the lack of literature in approaching both the institutional quality and disaggregated energy effects on the environment of OECD countries, this study attempts to figure out how these important factors influence their EF. In particular, we investigate the concomitant role of institutional quality and RE/non-RE energy utilization in the ecological sustainability of OECD countries.

The ecological footprint concept

Most of the existing studies in the literature use CO_2 emissions or other pollutants as indicators that represent environmental performance (e.g., Shafiei and Salim 2014; Abid 2017; Salman et al. 2019; Liu et al. 2020a, 2020b). However, ambient air pollutants capture the environmental damage only partially, since human activities influence also the land and water of the ecosystem. Furthermore, the impact of the energy-driven economic activity on the environment should be assessed through the overall ecosystem's damage (Destek and Sinha 2020). In this sense, the Ecological Footprint (EF) indicator extends the conventional environmental indicators into a broader perspective of ecological sustainability.

Rees (1992) and Wackernagel and Rees (1996) were the first who conceived and developed the indicator of EF. It is defined as an ecological sustainability measure that encompasses not only the total carbon footprint aspect but also the aspects of cropland, grazing, forestry, and built-up land and fisheries. EF shows how humans, regions, and countries' demand for resources is compared with the ability of nature to regenerate and absorb the waste generated under the existing technological and resource management processes (Wackernagel and Rees 1996). It is a measure that reflects the biologically productive earth capacity required for a human or population to balance the resources' consumption with the respective production and absorb the generated waste (Wackernagel and Monfreda 2004).

Since 1970, the world faces a gradual increase in the EF which shows that human demand exceeds the ecosystem's biocapacity, leading to a substantial ecological deficit. Most of the developed countries face an ecological deficit (Yasin et al. 2020) since the rate of natural regeneration of resources and waste absorption cannot meet the pace of human demand on nature. The EF is an easy monitoring tool for policymakers and it is measured in global land hectares which represent the biologically productive area required for sustainable resource utilization every year (Kitzes and Wackernagel 2009; GFN 2020). In 2016, among the OECD countries, Luxemburg had the highest EF consumption per capita of 12.91 global hectares per capita (gha pc), while Mexico had the lowest one of 2.6 gha pc. Taking into account that for sustainable resource utilization and regeneration of the ecosystem the current biocapacity per person is estimated at 1.7 gha pc (GFN 2020), it is obvious that OECD countries are subjected to an escalated ecological deficit.

Literature review

Environment—renewable energy relationship

In the energy and environmental economics literature, the association between energy consumption and environmental quality has been the topic of intensive research. Most of the relevant studies examine energy consumption at an aggregated form (Menyah and Wolde-Rufael 2010a, 2010b; Arouri et al. 2012; Abul et al. 2019; Abbasi et al. 2020; Ansari et al. 2020; Nawaz et al. 2020) and therefore do not consider the distinct effects of disaggregated energy sources on the environment. Since this study decomposes energy consumption in renewable and non-renewable, we focus on the distinctive effects of these two forms of energy on the environment.

Most of the studies demonstrate that the use of renewables has a substantial contribution to environmental sustainability. Dogan and Seker (2016a) followed an Environmental Kuznets Curve (EKC) approach from 1980 to 2012 in the

European Union (EU). They concluded that RE and trade decrease CO_2 emissions in the EU, while non-renewables exert a respective increasing effect. The same authors (Dogan and Seker 2016b) presented similar findings for the top 23 RE-consuming countries. Lee (2019) investigated the dynamic relationship between RE consumption and CO_2 emissions in the EU from 1961 to 2012. He concluded that RE has been used successfully to mitigate the EU environmental deterioration in both the short and the long term. Paramati et al. (2017) showed the positive contribution of RE on mitigating CO_2 emissions using data from the eleven next generation emerging economies (N-11)[3] in the period 1990–2012. Sharif et al. (2019) also found that RE promoted environmental quality in 74 nations during the period 1990–2015.

In a single-country study, Danish et al. (2017) conducted the Autoregressive Distributed Lags (ARDL) bound test for Pakistan from 1970 to 2012. They concluded that RE consumption is environmentally friendly, as opposed to non-RE, which is considered to be the main culprit for environmental degradation. Besides, they reported a bidirectional causality between the two types of energy consumption and CO_2 emissions. Dogan and Ozturk (2017) also adopted an ARDL approach in an EKC model and confirmed that RE consumption exerts a negative influence on CO_2 emissions in the US. In the case of Malaysia, Bekhet and Othman (2018) used the Vector Error Correction (VECM) Granger causality, the Dynamic Ordinary Least Squares (DOLS), and Fully Modified Least Squares (FMOLS) methodologies. Their findings showed that RE mitigates CO_2 emissions reporting a unidirectional causal effect from the latter to RE use.

Sharif et al. (2020b) investigated the relationship between RE and environmental quality using the Quantile-on-Quantile methodological approach for the 10 mostly polluted countries. Though they concluded to a two-way causality between RE and environmental performance, they provided mixed results. In 15 countries where RE occupies a critical role in their energy mix, Saidi and Omri (2020) showed the positive effect of RE on environmental quality, positing no long-run causality. Nathaniel S.et al. (2020) used data from the MENA[4] countries and applied the Augmented Mean Group (AMG) method. They pointed out that factors such as financial development, urbanization, growth, and fossil fuel energy use degrade the environment and RE does not exert a substantial impact on eliminating this degradation. Similarly, Menyah and Wolde-Rufael 2010a, 2010b found that the adoption of RE does not have a substantial impact on the US environmental quality. Lin and Moubarak (2014) found no causal association between RE consumption and CO_2 emissions in China either

in the short or in the long run. In contrast with the previous studies, Apergis et al. (2010) examining 19 countries of different stages of development found that though nuclear energy consumption mitigates CO_2 emissions, RE consumption has the opposite effect. Furthermore, Boluk and Mert (2014) indicated that both renewable and fossil energy use is harmful to the environmental quality of 16 EU countries. Using a panel fixed effects model, they showed that the effect of renewables on the environment is half per energy unit harmful than the effect of non-renewables. The challenge of investigating environmental sustainability in a broad ecological perspective gave birth to studies that adopted the ecological footprint (EF) variable as a novel environmental indicator. Destek et al. (2018) examined the existence of the EKC using the EF of 15 EU countries. They found that RE and trade openness have a positive environmental impact while non-RE a negative one. Danish et al. (2019) explored how the real GDP, urbanization, natural resources, and RE affect the EF consumption in BRICS[5] countries. They found that RE consumption decreases the EF of the examined countries. In the same line, Wang and Dong (2019) using the AMG estimator concluded an adverse relationship between RE and the EF of 14 Sub-Saharan Africa (SSA) countries. They also reported a long-run unidirectional causal effect from RE use to their EF. In a study for Turkey, Sharif et al. (2020a) used the Quantile ARDL (QARDL) to investigate how RE and non-RE consumption influence its EF. Findings showed that RE influences negatively Turkey's EF confirming a bidirectional causal effect between them.

The promotion of RE forms in OECD countries during the last decades and their endeavor of promoting environmental quality through changes in their energy policies have been demonstrated by some recent studies. For a sample of 25 OECD countries over the period 1980 to 2011, Apergis and Payne (2014) showed that RE reduces per capita CO_2 emissions in the short-run. In a study by Alvarez-Herranz et al. (2017) for 17 OECD countries from 1990 to 2012, findings revealed the important role of RE adoption and energy innovation in mitigating greenhouse gas emissions. Shafiei and Salim (2014) utilized a STIRPAT[6] model for 29 OECD countries over the same period. They confirmed the positive role of non-RE and the negative role of RE on CO_2 emissions in the long-run. Zaghdoudi (2017) probed how oil prices and RE affected the environment in 26 OECD countries from 1990 to 2015 in an EKC framework. Findings showed that low oil prices expanded the use of fossil fuels which harms the environment and scrutinized the significance of promoting RE

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[3] The 11 fastest developing economies: Bangladesh, Egypt, Indonesia, Iran, Korea, Mexico, Nigeria, Pakistan, Philippines, Turkey, and Vietnam.

[4] Middle East and North Africa.

[5] Brazil, Russia, India, China, and South Africa.

[6] STIRPAT (Stochastic Impacts by Regression on Population, Affluence, and Technology) is an extension of the IPAT (impact = population, affluence, technology) environmental accounting equation. It is a stochastic model that predicts non-proportional and non-monotonic functional relationships among the factors that are supposed to be related with the ecosystem's performance.

with a view to eliminate environmental pollution. In a recent study, Destek and Sinha (2020) used second-generation techniques to investigate the role of RE and non-RE consumption, together with the impact of trade openness, on the EF of 24 OECD countries. Investigating the period 1980–2014, the results confirmed the contribution of renewables to promoting environmental sustainability by decreasing the EF of OECD countries.

Environment—institutional quality relationship

While several studies investigate the factors that influence environmental performance, there is little evidence in the literature about the role of institutions in promoting environmental sustainability. Approaching the environmental-social dimensions of sustainable development, Lehtonen (2004) emphasized the important role of institutions at the national level as they have to arrange the appropriate framework for local activities that promote environmentally sustainable development. In a political economics approach, Scruggs (1999) investigated the nature of the relationship that associates the political-economic institutions of 17 OECD countries with their environment, from 1970 to the 1990s. He found that the corporatist institutions[7] are associated with better environmental quality, whereas pluralistic institutions[8] do not have a significant effect on it.

Bhattarai and Hammig (2001) showed the critical role of political institutions and governance in the deforestation process of 66 Latin American, African, and Asian countries. They found that institutional and governance improvements reduce the levels of deforestation in Latin America and Africa but increase them in Asia. Culas (2007) presented evidence that in 14 countries of Latin America, Africa, and Asia, improvements in institutions related to secure property rights for forests could eliminate deforestation. He showed that the contract enforceability of government and the efficiency of bureaucracy can promote forestland conservation. Dutt (2009) examined the impact of institutional indicators in an EKC model of 124 countries from 1984 to 2002. Findings revealed the crucial role that political institutions, governance, and socioeconomic conditions play in reducing CO_2 emissions. Examining 28 EU countries, Castiglione et al. (2012) reported the principal role of a strong rule of law in achieving environmental sustainability. Gani (2012) showed that rule of law, political stability, and control of corruption promote the environmental quality of 99 developing countries, while government effectiveness and regulatory quality do not have a significant effect on it. Ibrahim and

Law (2016) applied the system Generalized Method of Moments (SYS-GMM) in a panel of 40 Sub-Saharan countries, identifying the importance of institutional arrangements in environmental performance. They also showed the important institutional influence in the relationship between trade and the environment. Abid (2016) demonstrated that political stability, effective governance, and the reduction of corruption decrease CO_2 emissions in 25 SSA countries, while the well-defined regulatory framework and rule of law increase them.

Liu X. et al. (2020) emphasized the three forms of governance (political, economic, and social-institutional) in the context of the EKC hypothesis for 5 high polluting countries (China, India, Japan, Russia, and the USA). They utilized the DOLS and FMOLS methods providing evidence for the catalytic role of efficient governance on environmental performance. Abid (2017) argued that the level of environmental quality is vitally connected with the nature of economic growth, and institutions play a key role in this process. Using data from 58 MEA[9] and 41 EU countries, he found that institutional quality reduces the environmental degradation of EU countries but has no effect on the environment of the MEA region. Ali S. et al. (2020) examined the impact of institutional efficiency, along with other factors such as urbanization, foreign direct investment (FDI) inflows and trade, on the EF of 47 OIC[10] countries. They observed that except for institutions all the other factors have a positive relationship with the EF.

Environment—energy consumption—institutional quality relationship

The third strand of the literature incorporates both energy utilization and institutional measures in the examination of the effects of economic development on the environment. Tamazian and Rao (2010) used panel data of 24 transition economies and showed that institutional quality promotes environmental protection, while financial liberalization exerts a positive influence in the environment only in the presence of strong institutions. Salman et al. (2019) incorporated the factors of institutional efficiency and energy consumption in the examination of the relationship between economic growth and CO_2 emissions in 3 East Asian countries. Using "law and order" as an institutional indicator, they found that the quality of local institutions is important to lessening environmental degradation in these areas. In a panel of 47 developing countries, Ali H.S. et al. (2019) applied the system GMM estimation which revealed that economic growth, trade openness, energy use, and urbanization increase CO_2 emissions while institutional quality produces the opposite effect.

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[7] Policy concentrated institutions that facilitate the policy implementation among different groups of interest, from regulators to environmental groups.
[8] Institutions that are not strongly policy concentrated and influenced by a small number of interest groups in formulating policies.

[9] Middle Eastern and African.
[10] Organization of Islamic Cooperation.

Bhattacharya et al. ([2017](#)) highlighted the positive role of institutions and RE consumption in the environmental quality of 85 countries by applying the system GMM and FMOLS estimators. Sarkodie and Adams ([2018](#)) used an ARDL model to analyze how renewable, fossil, and nuclear energy along with economic growth, urbanization and institutions affect the environment of South Africa. Findings showed that in the long-term, nuclear energy provokes an insignificant effect on South Africa's environment, while non-RE consumption is harmful to it. On the contrary, RE use and robust political institutions exert a beneficial effect on the environment. Al Mulali and Ozturk ([2015](#)) employed the FMOLS to find how energy consumption and political stability along with other factors, such as urbanization and industrial development, influence the EF of 14 MENA countries. They concluded that although energy consumption and the other factors increase EF consumption, political stability reduces it. Likewise, in an EKC framework, Yasin et al. ([2020](#)) established that political institutions promote the environmental sustainability of 53 developed and 57 developing countries, while energy consumption increases their EF.

In contrast with the abovementioned studies, Azam et al. ([2020](#)) explored the effects of institutional quality on 5 environmental indicators and energy consumption of 66 developing countries using the system GMM method. They concluded that the quality of institutions seems to increase environmental degradation, as well as promotes the use of oil and fossil-based energy sources. In the same line, Hassan et al. ([2020](#)) established a positive relationship between institutional quality, energy consumption, and CO_2 emissions in Pakistan applying an ARDL model. Le and Ozturk ([2020](#)) used the Common Correlated Effects Mean Group (CCE-MG), AMG, and Dynamic CCE (DCCE) estimators to conclude that both institutional quality and energy use raise the environmental pollution in 47 EMDEs[11]. Charfeddine and Mrabet ([2017](#)) employed an EKC model and used the EF as an environmental proxy in 15 MENA countries. Their results, based on DOLS and FMOLS estimators, revealed that energy consumption and institutional quality (represented by political rights and civilian freedom) increase the EF in these regions.

Obviously, the studies that incorporate institutional and energy measures in the identification of potential factors that influence environmental performance are mostly recent. However, as regards the role of institutions and renewables on the environment, a divergence is observed in the relevant empirical findings. This could be explained by the special socioeconomic conditions of the country groups or regions being examined, the differences in the time span, model specifications, and methodologies being implemented as well as the environmental indicator employed. Regarding the OECD countries, the literature, except for the study of Scruggs

([1999](#)), lacks relevant empirical research. Furthermore, although the energy sector is strongly connected with the level of sustainable development, no study establishes a common framework in examining the impact of disaggregated energy utilization and institutional quality on the environment of OECD countries. This study attempts to fill this gap by providing empirical evidence through the employment of recent econometric techniques. Using second-generation unit root tests and cointegration methods is able to account for the critical issues of cross-sectional dependence and heterogeneity.

Data and methodology

Data and model specification

The study provides an insight into the role of institutions in a disaggregated energy-environmental model for OECD countries. We use annual data of 29 OECD countries (Appendix Table 12) over the period 1984–2016, and we focus on a second-generation panel data analysis approach. Unlike the majority of existing studies, and similar to Al Mulali and Ozturk ([2015](#)), Charfeddine and Mrabet ([2017](#)), Destek et al. ([2018](#)), Danish et al. ([2019](#)), Sharif et al. ([2020a](#)), and Ali et al. ([2020](#)), we employ EF as an environmental indicator. As described previously ("[The ecological footprint concept](#)" section), this indicator expresses more broadly and inclusively the level of environmental performance. Moreover, in our model specification, we use the real GDP, RE consumption, non-RE consumption, and institutional quality as independent variables. Thus, in line with Bhattacharya et al. ([2017](#)), we formulate the following model to capture the distinct impact of the disaggregated energy sources and the influence of institutions in the environment of OECD countries during the process of economic growth:

$$EF_{it} = f\left(GDP_{it}, RE_{it}, NRE_{it}, IQ_{it}\right) \tag{1}$$

In Eq. ([1](#)) EF_{it} is the ecological footprint consumption in total global hectares (gha); GDP_{it} is the real GDP in billions of constant 2010 US dollars as a proxy of economic growth; RE_{it} and NRE_{it} are the renewable electricity net consumption[12] and fossil fuels electricity net consumption[13] in billion kilowatt-hours; and IQ_{it} is an institutional quality index created through the Principal Component Analysis (PCA) of 6 basic institutional sub-indices: Socioeconomic Conditions (SEC), Government Stability (GS), Corruption (COR), Law and Order (LO), Democratic Accountability (DA), and Bureaucracy Quality (BA).

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[11] Emerging Market and Developing Economies.

[12] Electricity power from renewables such as wind, solar, geothermal, biomass, and water energy.

[13] Electricity power from non-renewables such as coal, oil, and natural gas.

The institutional quality index constitutes a composite measure of the fundamental institutional dimensions of the OECD countries. These dimensions are reflected by the respective social, constitutional (political and democratic), and administrative-judicial systems (basic institutional quality). The PCA approach produces an index that comes from the weighted factor loadings of originally correlated variables. The i_{th} principal component is the result of the linear combination of m original variables:

$$IQ_i = w_{i1}*X_1 + w_{i2}*X_2 + \dots + w_{im}*X_m \qquad (2)$$

where $w_{i1}^2 + w_{i2}^2 + \dots + w_{im}^2 = 1$, $i = 1, 2, \dots, m$; w_{ij} is the weight of the j variable in the i_{th} component; X_m are the original indicators; and m the number of variables. PCA overcomes the problem of collinearity between individual indicators and forms a composite one that encompasses the most crucial information from the original indicators and expresses the highest possible variation of the original dataset (OECD 2008).

The data for EF_{it} were obtained from the Global Footprint Network (GFN), the data for GDP_{it} from the World Development Indicators (WDI) (2020), and the data for RE_{it} and NRE_{it} from the US Energy Information Administration (EIA) (2020). The data for the 6 original institutional indicators which formed the IQ_{it} index were taken from the International Country Risk Group (ICRG) (Table 1). Table 2 presents some descriptive statistics of the model's variables. The US exhibits the highest EF in 2005, whereas Luxemburg the lowest one in 1984. Also, the US has the largest portion in RE utilization in 2016, while the lowest user of RE is Israel in 1989. It is also observed that the US has the highest level of non-RE consumption in 2007, while Norway is the lowest consumer of fossil fuel energies in 1991. Switzerland seems to have the most qualitative institutional framework in 1984, and Turkey the least qualitative one in 1990. The standard deviation statistics imply a heterogeneous behavior across cross-sectional units. In our model, all variables except for

Table 2 Descriptive statistics

LEFC		LGDP	LRE	LNRE	IQ
Mean		7.990665	11.70831	1.141246	1.587996	−0.208986
Median		7.884581	11.63140	1.383672	1.597432	0.518554
Maximum		9.485326	13.22974	2.807543	3.475996	2.679839
Minimum		6.674594	10.23242	−2.522879	−0.744727	−6.107259
Std.Dev.		0.547829	0.559132	0.950230	0.837727	1.786302

the IQ_{it} are expressed in natural logarithmic form, thus equation (1) can be written in the following log-linear form:

$$
\begin{aligned}
LEF_{it} = {} & \beta_{0i} + \beta_{1i}LGDP_{it} + \beta_{2i}LRE_{it} + \beta_{3i}LNRE_{it} \\
& + \beta_{4i}IQ_{it} + \varepsilon_{it}
\end{aligned}
\qquad (3)
$$

where $i = 1,\dots,N$ the cross-sectional units (countries); $t = 1,\dots T$, the time period (1984-2016); β_{1i}, β_{2i}, β_{3i}, and β_{4i} parameters that represent the coefficient estimates and e_{it} is the error term. The study is based on a panel data analysis that takes into account the heterogeneity and cross-sectional (CS) dependency and focuses on second-generation methodologies and econometric approaches.

Cross-sectional dependency and slope heterogeneity

One of the principal factors that directly influence the adopted methodological strategy in panel data analysis is the issue of errors' CS dependence (De Hoyos and Sarafidis 2006). In the era of globalization, this is a predominant phenomenon in panel data estimations due to socioeconomic linkages among countries or hidden common effects and regional characteristics (Chudik and Pesaran 2013). When CS correlation is present, the conventional estimation methods are criticized for providing inconsistent and unreliable results (Kapetanios et al. 2011). For this reason, our first step is to examine the existence of CS dependency in the residuals of the model

Table 1 Variables description

Variable	Description	Unit of Measurement	Source
EF	Ecological Footprint	Total Global Hectares (gha)	Global Footprint Network (GFN)
GDP	Real GDP	Constant 2010 US $	World Development Indicators (WDI)
RE	Renewable electricity net consumption	Billion kilowatt/hours	US Energy Information Administration(EIA)
NRE	Fossil fuels electricity net consumption	Billion kilowatt/hours	US Energy Information Administration
IQ	Institutional quality Index	PCA composite index of 6 original indicators: SEC, GS, COR, LO, DA and BQ	International Country Risk Guide (ICRG)

Note: *SEC* socioeconomic conditions, *GS* government stability, *COR* corruption, *LO* law and order, *DA* democratic accountability, *BQ* bureaucracy quality

through the Breusch and Pagan (1980) LM test, Pesaran Scaled LM test (2004), Pesaran CD test (2004), and Pesaran bias-adjusted LM test (2008).

Considering a typical panel data model of the following form:

$$y_{it} = a_i + \beta_i X_{it} + \varepsilon_{it} \tag{4}$$

where $i = 1,...,N$ and $t = 1,...$ T express the cross-sections and time escalation respectively; a_i and β_i represent the individual intercepts and slope coefficients; and X_{it} is a $k \times 1$ vector of the independent regressors. The null hypothesis when testing the existence of CS dependence is $Cov(\varepsilon_{it}\varepsilon_{jt}) = 0$ and shows CS independency and the alternative is $Cov(\varepsilon_{it}\varepsilon_{jt}) \neq 0$ and indicates that at least two CS units are correlated with each other. The Breusch and Pagan (1980) LM test which follows the asymptotic χ^2 distribution is appropriate when the dimension of N is fixed and T is large and it is expressed by the following LM statistic:

$$LM = \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} T_{ij}\widehat{\rho}_{ij}^2 \sim \chi^2_{N(N-1)/2} \tag{5}$$

where $\widehat{\rho}_{ij}$ are the values of the pair wise correlation of the errors of Eq. (4) using ordinary least squares (OLS).

Pesaran (2004) introduced a test (the so-called scaled LM test) to overcome the limitation of small N, i.e., for $N \rightarrow \infty$, $T \rightarrow \infty$, which is asymptotically normally distributed and is denoted by the following statistic:

$$LM_S = \sqrt{\frac{1}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \left(T_{ij}\widehat{\rho}_{ij}^2 - 1 \right) \sim N(0, 1) \tag{6}$$

Due to the fact that the LM_S test is subjected to size distortions when T is small and N is large, Pesaran (2004) also suggested the CD test which is based on the average of the $\widehat{\rho}_{ij}$ values. Unlike the previous tests, the CD test addresses the problem of size distortions, it is consistent for fixed T and N values and it performs well even in the case of small T and N (Pesaran 2004). The CD statistic is described as follows:

$$CD = \sqrt{\frac{2T}{N(N-1)}} \left(\sum_{i=1}^{N-1} \sum_{j=i+1}^{N} \widehat{\rho}_{ij} \right) \sim N(0, 1) \tag{7}$$

Nevertheless, the CD test lacks power in cases that the population mean pair-wise correlation, in contrast with the individual ones, is zero. For this reason, Pesaran et al. (2008) introduced the bias-adjusted LM test as follows:

$$LM_{adj} = \sqrt{\frac{2}{N(N-1)}} \sum_{i=1}^{N-1} \sum_{J=i+1}^{N} \frac{(T-k)\widehat{\rho}_{ij}^2 - \mu_{T_{ij}}}{V_{T_{ij}}} \sim N(0, 1) \tag{8}$$

where $\mu_{T_{ij}}$ and $V_{T_{ij}}$ denote the exact mean and variance of $(T-k)\widehat{\rho}_{ij}^2$ respectively.

In accordance with the CS dependence test, another crucial issue that defines the next methodological steps is the issue of slope homogeneity across the individual units of the panel. In most cases the CS homogeneity hypothesis does not correspond to the real world (Urbain and Westerlund 2006). For this reason, we utilize the $\widetilde{\Delta}$ test of Pesaran and Yamagata (2008) which is based on the modified Swamy's (1970) statistic ($\widetilde{S}$), and it is appropriate in panels where both the N and T dimensions could be large in their relative expansion dynamics. In the case of non-normal errors and exogenous regressors, and under the premise that $\sqrt{N}/T^2 \rightarrow 0$ as $(N, T) \rightarrow \infty$, the ($\widetilde{\Delta}$) test follows the standard normal distribution:

$$\widetilde{\Delta} = \sqrt{N} \left(\frac{N^{-1}\widetilde{S} - k}{\sqrt{2k}} \right) \tag{9}$$

where $\widetilde{S}$ is the Swamy statistic[14] and k the explanatory variables. The bias-adjusted $\widetilde{\Delta}_{adj}$ test meliorates the small sample properties of the $\widetilde{\Delta}$ statistic under the premise that errors are normally distributed. This test requires no restrictions on the expansion dynamics of N and T, as $(N, T) \rightarrow \infty$:

$$\widetilde{\Delta}_{adj} = \sqrt{N} \left(\frac{N^{-1}\widetilde{S} - E\left(\widetilde{z}_{it}\right)}{\sqrt{var\left(\widetilde{z}_{it}\right)}} \right) \tag{10}$$

where $E(\widetilde{z}_{it}) = k$ and $var(\widetilde{z}_{it}) = \frac{2k(T-k-1)}{T+1}$ are the mean and variance respectively. The null hypothesis of these tests assumes slope homogeneity ($H_0 : \beta_\iota = \beta$), whereas the alternative one assumes heterogeneity of the CS_s slope coefficients ($H_1 : \beta_\iota \neq \beta_j, i \neq j$) for a non-zero fraction of the pairwise slopes. For dynamic panels, this test produces consistent results when $T \geq N$ (Pesaran and Yamagata 2008), thus it is considered suitable for our study.

Panel unit-root tests

The application of unit root tests is necessary to identify the stationarity properties of the variables. Conventional first-generation tests[15] fail to consider the issues of CS-dependent distribution of the individual units in panels and heterogeneity of slopes. To overcome these problems, Pesaran (2007) suggests the CS augmented Dickey-Fuller (CADF) and CS augmented Im, Pesaran, and Shin (CIPS) unit root tests. These are second-generation tests that produce robust results when panels are subject to CS dependence and heterogeneity.

CADF test extends the standard DF or ADF regressions by adding lagged CS means of the individual units and first

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[14] See Pesaran and Yamagata (2008) for detailed description.
[15] e.g., LLC, Breitung, IPS, Fisher-ADF, Fisher PP

differences. In this way, it filters out the observed CS dependence which is caused by a possible unobserved common factor effect. CADF test builds upon the heterogeneous regression below:

$$\Delta y_{it} = a_i + b_i y_{it-1} + c_i \overline{y}_{it-1} + d_i \Delta \overline{y}_{it} + u_{it} \qquad (11)$$

where a_i is the constant term, $\overline{y}_{i,t-1}$ is the one lag value of the CS mean of $y_{i,t}$ and $\Delta \overline{y}_t$ the first differences of $\overline{y}_t$. The number of lags in Eq. (11) can be modified accordingly in order to filter out the presence of serial correlation in the error term (Pesaran 2007). The CIPS test is estimated by the average of individual CADF statistics of the i_{th} cross-sectional units:

$$CIPS = \left(\frac{1}{N}\right) \sum_{t=1}^{N} CADF_i \qquad (12)$$

The null hypothesis in CADF and CIPS tests assumes a homogenous unit root for all individuals within a panel, whereas the alternative one supports that there is at least one stationary individual in the panel. Due to the detected CS dependency and slope homogeneity in our panel data sample, this study focuses on the aforementioned second-generation tests to avoid wrong inferences.

Panel cointegration tests

Under the premise that the variables are integrated of order one [I(1)], i.e., they are stationary at first differences, the next methodological step is to trace the existence of a cointegration relationship. Pedroni (1999, 2004) panel cointegration tests take into account the CS heterogeneity and detect if there is of a long-run relationship in the examined model. Considering the following equation:

$$y_{it} = a_i + d_i t + b_{1i} x_{1i,t} + b_{2i} x_{2i,t} + \dots + b_{mi} x_{mi,t} + e_{i,t} \qquad (13)$$

t is the number of observations; i the number of CS units; m the number of regressors; and a_i, b_i the intercept and slope coefficients respectively, which can vary across individual units. Provided that all the examined variables are I (1), the null hypothesis assumes that the error term $e_{i,t}$ has a unit root and so there is no cointegration. The alternative hypothesis supports that the error term is stationary implying the existence of cointegration. Pedroni suggests seven statistics that consist of four within-dimension and three between-dimension statistics. The main characteristic of within-dimension statistics is that they consider pooled autoregressive coefficients of the residuals which test them for unit roots. In contrast, between-dimension cointegration statistics are group mean statistics calculated by estimators which consider the average of each unit's individual autoregressive coefficients.

Although the residual-based Pedroni tests are utilized to detect long-run associations in heterogeneous panels, the nature of these tests fails to consider the issue of CS dependence. Westerlund (2007) fills this gap by proposing four panel error-correction based cointegration tests (G_τ, G_a, P_τ, P_α) that overcome any restriction induced by a common factor. According to him, these tests perform greater than the residual-based tests in terms of size accuracy and increase of power. Considering the following error-correction model:

$$\Delta y_{it} = \theta_i' d_t + z_i \left(y_{it-1} - \omega_i' x_{it-1} \right) + \sum_{j=1}^{v_i} z_{ij} \Delta y_{it-j}$$
$$+ \sum_{j=0}^{v_i} \mu_{ij} \Delta x_{it-j} + \varepsilon_{it} \qquad (14)$$

d_t holds the deterministic terms; θ_i is a vector of the relative parameters; and z_i is the error correction term that is tested if is equal to zero or not. The null hypothesis assumes the lack of cointegration when the error correction term is zero, while the alternative one accepts the existence of cointegration. All the test statistics are normally distributed and the G_τ and G_a statistics reveal cointegration to at least one of the cross-sections, while the P_τ and P_α statistics refer to cointegration in the whole panel. In the presence of CS dependence, Westerlund (2007) proposed a bootstrap approach to eliminate inconsistent estimates even when more generic types of CS dependence exist.

Long-run parameter estimates

The investigation of long-run effects in a structural model is of utmost importance as it is joined with the equilibrium state of the model (Chudik et al. 2016). In the presence of CS dependence and heterogeneity in panel data analysis, conventional estimation techniques provide spurious inference (Chudik and Pesaran 2013). The recent dynamic approaches of the CS-ARDL and CS-DL estimators (Chudik et al. 2016) overcome the previous challenges and produce robust long-run estimators. This study focuses on the CS-DL estimator which tends to perform better in small sample sizes where $30 \leq T \leq 50$, in comparison with the CS-ARDL estimator that is consistent only in large samples where $T > 50$ (Chudik et al. 2016).

In order to examine the relationship between the x_{it} regressors and the $y_{i,t}$ dependent variable, we consider an ARDL (ν, ρ) model specification in a panel data context as follows:

$$y_{i,t} = \sum_{j=1}^{v_{y_i}} a_{i,j} y_{i,t-j} + \sum_{j=0}^{\rho_{x_i}} \beta_{i,j}' x_{i,t-j} + \varepsilon_{i,t}, \qquad (15)$$

where $x_{i,t}$ is a $k \times 1$ vector of regressors and $\varepsilon_{i,t} = \theta_i' f_t + e_{it}$, where f_t is a vector of unobserved common factors; and v_{y_i}, ρ_{x_i} are the optimal lag lengths to ensure that $\varepsilon_{i,t}$ does not exhibit a

serial correlation process across cross-section units. Equation (15) can be expressed as a cointegrating relationship in the following form:

$$y_{i,t} = \varphi_i x_{it} + \gamma_i'(L)\Delta x_{it} + \widetilde{\varepsilon}_{i,t} \qquad (16)$$

where φ_i is the vector of the long-run coefficients. The CS-DL estimator is structured as follows:

$$y_{i,t} = c_{yi} + \varphi_i' x_{it} + \sum_{j=0}^{\rho_{x_i}-1} \mu_{ij}\Delta x_{i,t-j} \sum_{j=0}^{\overline{y}} \pi_{y,ij}\overline{y}_{i,t-j}$$
$$+ \sum_{j=0}^{\overline{x}} \pi_{x,ij}'\overline{x}_{i,t-j} + \varepsilon_{i,t} \qquad (17)$$

where $\overline{x}_{it}$ and $\overline{y}_{it}$ are the respective average values of x_{it} and y_{it}; $\rho_{\overline{x}} = \left[\sqrt[3]{T}\right]$, $\rho = \rho_{\overline{x}}$ and $v_{\overline{y}} = 0$. Except for the main merit of this estimator considering its consistent performance when the sample's time dimension lies between 30 and 50 periods[16], it is also robust in the presence of unit roots in independent variables and/or unobserved factors. Moreover, this estimator is reliable if serial correlation exists in e_{it} and f_t and when a number of hidden common factors exist, and it is consistent when there are possible breaks in e_{it} and dynamic wrong specifications (Chudik et al. 2013).

In addition to the CS-DL estimator, for comparison reasons, this study also employs the group-mean (MG) panel DOLS estimator (Pedroni 2001) which addresses the issues of heterogeneity, endogeneity, and autocorrelation in the detection of long-run relationships. The use of leads and lags which form the parametric nature of DOLS make this estimator perform better as T increases than the nonparametric FMOLS estimator (Wagner and Hlouskova 2010). Concomitantly, DOLS demonstrates greater performance in the case of small samples than the FMOLS estimator (Kao and Chiang 2001). Nevertheless, this estimator is robust only in the presence of CS dependence induced by time factors and is not consistent in the presence of other forms of CS dependency (Pedroni 2001).

Panel causality test

As a final part of the methodological analysis, the study attempts to investigate the existing causal associations in the model applying the Dumitrescu and Hurlin (2012) D-H causality test. This step is critical for implementing relevant policy design. The D-H test is suitable for heterogeneous panels subjected to CS dependence and it is robust in small sample performance (either small T or N). However, it requires all the observed variables to be stationary at the same level. It is

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[16] In this study, T=33.

implemented on the basis of the Granger (1969) non-causality test by averaging the individual Wald statistics of CS units. The D-H model specification is described as follows:

$$y_{i,t} = c_i + \sum_{l=1}^{L} a_i^{(l)} y_{i,t-l} + \sum_{l=1}^{L} b_i^{(l)} x_{i,t-l} + e_{i,t} \qquad (18)$$

It is a fixed coefficient VAR model where x and y are the model's variables examined for N cross-section units under T periods; L is the lag structures; $a_i^{(l)}$ and $b_i^{(l)}$ denote the autoregressive parameters and the coefficient slopes that differ across CS groups. The null Homogeneous Non-Causality hypothesis assumes the absence of causal relationships in any of the CS units. The alternative Heterogeneous Non-causality hypothesis considers the existence of causality at least for a subgroup of them (Dumitrescu and Hurlin 2012). The mean Wald statistic which tests the null hypothesis is:

$$\overline{W}_{N,T} = \frac{1}{N} \sum_{i=1}^{N} W_{i,T} \qquad (19)$$

where $W_{i,T}$ is the individual units' Wald statistics. Dumitrescu and Hurlin (2012) show that as T initially and N sequentially increase, the test statistic converges to a standard normal distribution. In contrast with previous approaches to the D-H test, this study applies the bootstrap estimation procedure of the test, which is appropriate when CS units are not independent.

Empirical findings and discussion

The construction of an institutional quality index (IQ) that reflects the basic institutional dimensions of the examined OECD economies is provided through the PCA analysis displayed in Table 3. All six individual indicators are assigned equal weights in forming the principal components. The eigenvalues reported in the upper part of this table show the maximum eigenvalue 3.187 of the first factor, 1.087 of the second factor, and 0.657, 0.516, 0.302, 0.247 the respective eigenvalues of the third, fourth, fifth, and sixth factors. These values indicate that the first principal component explains 53.13% of the overall standardized variance, while the proportional variations of the second and third factors are 18.13% and 10.96% respectively. The proportions of the fourth, fifth and sixth factors are found lower than 10%. In the lower part of the table, it is clear that only the first principal component's loadings exhibit no negative values, so we use this for the construction of our index. The factor loadings suggest the individual contributions to the overall variance of the first principal component. As can be seen, the highest contributions are that of "bureaucracy quality" (with a weight of

Table 3 Principal component analysis of the IQ index

Eigenvalues

Number	Value	Difference	Proportion	Cumulative Value	Cumulative Proportion	
1	3.187541	2.100007	0.5313	3.187541	0.5313	
2	1.087534	0.429852	0.1813	4.275076	0.7125	
3	0.657682	0.140983	0.1096	4.932758	0.8221	
4	0.516699	0.213920	0.0861	5.449458	0.9082	
5	0.302779	0.055016	0.0505	5.752237	0.9587	
6	0.247763	---	0.0413	6.000000	1.0000	

Eigenvectors (loadings)

Variable	PC 1	PC 2	PC 3	PC 4	PC 5	PC 6
SEC	0.371077	0.479148	−0.098029	−0.746156	0.132667	0.220816
GS	0.245438	0.731557	0.360414	0.518748	−0.043142	−0.061043
COR	0.407791	−0.385895	0.537446	0.051781	0.590619	0.210789
BQ	0.497869	−0.119041	−0.163025	−0.050795	0.031530	−0.841311
LO	0.466704	−0.265509	0.216443	−0.032604	−0.777126	0.244656
DA	0.412410	−0.040753	−0.705847	0.409668	0.163672	0.367997

Note: *SEC* socioeconomic conditions, *GS* government stability, *COR* corruption, *LO* law and order, *DA* democratic accountability, *BQ* bureaucracy quality

0.497), "law and order" (weigh 0.466), "democratic accountability" (weigh 0.412), and corruption (weigh 0.407).

The most critical point that defines the methodological route that is going to be followed in the study is the CS dependency test. Table 4 depicts the results of the four relative tests of Breusch-Pagan (1980) and Pesaran (2004, 2008) mentioned in the previous section. All tests reject the null hypothesis of CS independence which means that unequivocally, in a globalized economic environment, there are common factors that affect OECD countries. In addition, Table 5 shows the findings of the slope heterogeneity test of Pesaran and Yamagata (2008) where the null hypothesis of homogeneity is rejected at a 1% level of significance. The rejection of slope homogeneity reveals a country-specific heterogeneity in the examined panel. This finding implies that the homogeneity restriction in the investigation of the causal relationships between the variables of interest could lead to wrong estimations. The presence of CS correlation and heterogeneity

indicates that the second-generation tests probably provide more consistent inferences than the first-generation ones.

For the identification of unit roots in the variables of our model, we employ the CADF and CIPS tests (Pesaran 2007) which are robust to the dependency and heterogeneity of the CS units. However, for comparison reasons, we both apply the first generation unit-root tests of Im, Pesaran, and Shin (2003) (IPS), and the Fisher-Augmented Dickey-Fuller (ADF) (Maddala and Wu 1999). Table 6 presents the results of the first generation unit root tests and Table 7 displays the respective results of the second-generation tests. The outcome in Table 6 shows a mixed order of integration for the variables of the study, as some of them are stationary at levels either with constant or with constant and trend. However, all variables become stationary at their first differences. In contrast with the results from the first-generation tests, the results displayed in Table 7 demonstrate that all the variables of our model contain unit roots at levels but become stationary at first differences. These findings allow us to move forward with detecting if the

Table 4 Cross-section dependence tests results

Type	Statistic	*p*-value
Breusch-Pagan LM test	5290.631	0.000
Pesaran Scaled LM test	171.4169	0.000
Pesaran CD test	61.28005	0.000
Pesaran bias-adjusted LM test	15.600	0.000

Note: The null hypothesis of cross-sectional independence is rejected at 1% level of significance

Table 5 Slope heterogeneity test results

Type	Statistic	*p*-value
$\widetilde{\Delta}$	28.79	0.000
$\widetilde{\Delta}_{adj}$	31.83	0.000

Note: The null hypothesis of slope homogeneity is rejected at 1% level of significance

Table 6 Unit root tests results (first generation)

Fisher ADF					
Levels			First differences		
Variable	Constant	Constant and trend	Variable	Constant	Constant and trend
LEF	72.318	34.245	DLEF	453.401***	676.318***
	(0.197)	(0.994)		(0.000)	(0.000)
LGDP	85.419**	36.620	DLGDP	286.135***	252.828***
	(0.011)	(0.987)		(0.000)	(0.000)
LRE	74.627*	74.875*	DLRE	631.381***	626.937***
	(0.069)	(0.067)		(0.000)	(0.000)
LNRE	59.992	49.458	DLNRE	484.536***	462.438***
	(0.403)	(0.780)		(0.000)	(0.000)
IQ	70.489	42.435	DIQ	322.083***	263.734***
	(0.125)	(0.937)		(0.000)	(0.000)
IPS					
Levels			First differences		
Variable	Constant	Constant and trend	Variable	Constant	Constant and trend
LEF	−2.118**	3.106	DLEF	−20.169***	−26.502***
	(0.017)	(0.999)		(0.000)	(0.000)
LGDP	−0.702	2.180	DLGDP	−13.490***	−12.571***
	(0.241)	(0.985)		(0.000)	(0.000)
LRE	4.491	−0.646	DLRE	−32.689***	−27.836***
	(1.000)	(0.259)		(0.000)	(0.000)
LNRE	−0.865	3.292	DLNRE	−20.892***	−22.022***
	(0.193)	(0.999)		(0.000)	(0.000)
IQ	−2.305**	1.438	DIQ	−14.468***	−12.487***
	(0.016)	(0.924)		(0.000)	(0.000)

Notes: Lags length selected based on SIC. Parentheses show the respective *p*-values. ***, **,* indicate that the null hypothesis of unit root is rejected at 1%, 5%, 10% level of significance respectively

independent variables of the model have a long-run association with the ecological footprint of OECD countries.

For the detection of a cointegration relationship in the model, we first apply the Pedroni panel cointegration test (1999, 2004) which is applicable for heterogeneous panels. As it is shown in Table 8, almost all the statistics of this test (six out of seven), strongly indicate the existence of cointegration in the panel since the null hypothesis of no cointegration is rejected at the 1% level of significance. Although the residual-based test of Pedroni is pioneering in tracing long-run relationships, it ignores the serious issue of CS dependence. All the Pedroni statistics fail to control the detected cross-sectional dependence in the panel which occurs due to unobserved common factors. Thus, they cannot lead to reliable conclusions about the long-run association between the variables of the examined model. Therefore, we utilize the Westerlund (2007) bootstrap test statistics as described in the preliminary methodological analysis to consider robustly our previous finding of CS dependence. Table 9 reviews the results of this test, where the G_τ, P_τ (at the 5% level of significance), and P_α (at the 10%

level) statistics provide clear evidence for the rejection of the null hypothesis of no cointegration. Since the G_τ and P_τ statistics are considered to be most robust when CS correlations exist (Westerlund 2007), the results of the respective bootstrapped p-values provide strong evidence in favor of cointegration. These outcomes validate the Pedroni tests which evidence a long-run relationship that interrelates RE and non-RE consumption, economic growth, and institutional quality with the steady-state of OECD countries' ecological footprint.

Having established the existence of cointegration in our model, we proceed with the investigation of the long-run coefficients accounting for the issues of heterogeneity and CS dependence. From this perspective, we report the outcomes of the long-run CS-DL estimator in Table 10. The results displayed in this table reveal that the effects of non-RE consumption and real income on the EF are negative and significant. In addition, institutional quality and RE consumption exert a negative effect on EF though the RE effect is not statistically significant. In particular, a 1% increase in non-

Table 7 Unit root tests results (second generation)

CADF

Levels			First differences		
Variable	Constant	Constant and trend	Variable	Constant	Constant and trend
LEF	−1.864	−2.308	DLEF	−3.309 ***	−3.463 ***
LGDP	−1.260	−2.491	DLGDP	−2.947 ***	−3.127 ***
LRE	−1.913	−2.284	DLRE	−3.399 ***	−3.397 ***
LNRE	−1.898	−2.096	DLNRE	−2.789 ***	−3.146 ***
IQ	−1.926	−1.990	DIQ	−4.061 ***	−4.094 ***

CIPS

Levels			First differences		
Variable	Constant	Constant and trend	Variable	Constant	Constant and trend
LEF	−2.007	−2.500	DLEF	−5.146 ***	−5.404 ***
LGDP	−1.817	−2.126	DLGDP	−3.524***	−3.647***
LRE	−1.451	−1.737	DLRE	−3.517***	−4.289***
LNRE	−2.040	−2.540	DLNRE	−5.140 ***	−5.318 ***
IQ	−1.926	−1.990	DIQ	−4.254***	−4.243***

Notes: For CADF lags length selected based on SIC. Critical values at 1% level of significance: with constant −2.30, with constant and trend −2.78. At 5% level: with constant −2.16, with constant and trend −2.65. At 10% level: with constant −2.08, with constant and trend −2.58. For CIPS critical values at 1% level of significance: with constant −2.23, with constant and trend -2.73. At 5% level: with constant −2.11, with constant and trend −2.61. At 10% level: with constant −2.04, with constant and trend −2.54. Lags criterion decision general to particular based on F joint test (max lags 2). *** indicate that the null hypothesis of unit root is rejected at 1% level of significance

RE use increases the EF of the examined OECD countries by 0.22% and a 1% increase in the real GDP increases the EF by 0.87%. Moreover, as institutional quality raises by 1%, EF shows a decrease of 1.2%.

Furthermore, we report the results of the long-run parameters using a first- generation estimator. The DOLS-MG estimator provides similar coefficient signs with the CS-DL estimator and it unveils that all the independent variables have a statistically significant impact on EF. Specifically, this estimator shows that a 1% increase in RE decreases the EF by 0.34%, while a 1% increase in non-RE consumption causes

a rise of 0.21% in the EF. Concurrently, a 1% increase in real GDP increases EF by 0.69%, while a 1% increase of institutional quality lessens EF by 2.6%. Although the findings of both estimators have similarities, this study considers the CS-DL estimation procedure as the benchmark methodological approach. This is because the CS-DL estimator performs better when the time dimension is between 30 and 50 observations. Moreover, unlike the DOLS-MG estimator which is consistent only in time factor-induced CS dependency, the CS-DL estimator captures more general forms of CS dependence.

Table 8 Pedroni cointegration tests results

Within-dimension			Between- dimension	
Statistic	Statistic value (*p*-value)	Weighted Stat. value (*p*-value)	Statistic	Statistic value (*p*-value)
Panel v-Stat.	3.475*** (0.000)	2.078** (0.018)	Group rho-Stat.	−1.061 (0.144)
Panel rho-Stat.	−4.372*** (0.000)	−2.580*** (0.004)	Group PP-Stat.	−9.396*** (0.000)
Panel PP-Stat.	−10.528*** (0.000)	−8.363*** (0.000)	Group ADF-Stat.	−9517*** (0.000)
Panel ADF-Stat.	−10.777*** (0.000)	−8.759*** (0.000)		

Note: Optimal lag length selected based on SIC. ***, **,* indicate that the null hypothesis of no cointegration is rejected at 1%, 5%, 10% level of significance respectively

Table 9 Westerlund Cointegration tests results

Statistic	Statistic value	Robust p-value
G_τ	−2.524**	0.030
G_α	−5.120	0.815
P_τ	−14.286***	0.000
P_α	−5.967*	0.100

Note: The Westerlund bootstrap test statistics are calculated using 200 bootstrap replications. Robust p-values are used in the presence of cross-sectional dependence. ***, **,* indicate that the null hypothesis of no cointegration is rejected at 1%, 5%, 10% level of significance respectively

Based on our findings, we substantiate the studies of Tamazian and Rao Bhaskara (2010), Gani (2012), Al Mulali and Ozturk (2015), Ibrahim and Law (2016), Abid (2016), Bhattacharya et al. (2017), Ali et al. (2019), and Salman et al. (2019) regarding the beneficial role of the institutional framework in the environmental performance. In parallel with these studies, via the use of an inclusive institutional quality index and its impact on the more sophisticated environmental indicator of ecological footprint, we provide similar evidence for most of the OECD countries. Concerning the negative role of non-RE consumption on the environmental quality, our study confirms the conclusions reached in the majority of the relevant studies such as those of Shafiei and Salim (2014), Dogan and Seker (2016 a,b), Danish et al. (2017), Sarkodie and Adams (2018) and Wang and Dong (2019). Nevertheless, RE use does not yet seem to demonstrate a substantial diminishing impact on the EF of OECD countries in the long-run. This finding is close to the results of Menyah and Wolde-Rufael (2010a, b) for the US, Lin and Moubarak (2014) for China, and Nathaniel et al. (2020) for MENA countries.

Table 10 Long-run estimates

	CS-DL mean long-run estimates	DOLS-MG estimates
Dependent Variable: LEF		
Independent Variables	Long-run coefficients (p-value)	Long-run coefficients (p-value)
LRE	−0.022 (0.697)	−0.348*** (0.000)
LNRE	0.221** (0.033)	0.213*** (0.001)
LGDP	0.872*** (0.000)	0.695*** (0.000)
IQ	−0.012** (0.057)	−0.026*** (0.000)

Note: ***, **,* indicate that the null hypothesis is rejected at 1%, 5%, 10% level of significance respectively

The last part of the study investigates the causal relationships between the respective variables in a robust methodological approach, as described by the Dumitrescu and Hurlin (2012) test. Taking into account the CS heterogeneity and dependence in our panel, we employ the bootstrap application of this test that produces consistent estimates. Table 11 depicts the D-H test results. It is shown that there is a unidirectional causal effect running from economic growth to EF and a weak causal effect from EF to the use of RE. Bekhet and Othman (2018) also found a one-way causality from CO_2 to RE, in the case of Malaysia. In line with the studies of Shafiei and Salim (2014) and Zachoudi (2017) for the OECD countries, we detect no causal effect directed from RE to EF. Furthermore, a bidirectional causal relationship seems to exist between non-RE and EF and between institutional quality and EF.

According to the previous findings, the role of RE consumption in the examined OECD countries seems to provoke an insignificant effect on their EF. Nevertheless, the increase in EF may weakly prompt the adoption of RE in OECD economies. That means that the level of renewables' utilization in OECD countries is not high enough to cause a substantial reduction in their EF. Concurrently, non-RE use seems not only to upsurge the EF but also to exert a feedback effect to it. Consequently, the share of non-RE still seems to exercise a dominant effect on the environmental status of the examined countries. On the other hand, institutions have a significant healing effect on the environmental performance of OECD countries. This effect is strengthened by the detected interdependence between the institutional operation and the EF. More precisely, our findings testify that the qualitative operation of institutions constitutes the main factor that contributes to environmental quality in OECD countries. This verdict complements the observations of Destek and Sinha (2020), who argue that the institutional performance in most of the OECD countries is not effective enough to ensure environmental sustainability.

The negative but insignificant effect of renewables in reducing the EF of OECD countries could be mainly ascribed to their differences in their available economic and technological-operational dynamics. There is a lot of diversification in the share of RE consumption on their national energy-mix portfolios. This could be attributed to differences regarding the institutional level, the stage of economic growth, and the phase of the fossil fuels decoupling process. The high cost of renewables and the different phases of sustainable targets materialization of OECD countries provoke an insignificant overall effect of renewables on their EF. Moreover, countries' specific institutional factors such as their legal and regulatory framework and governments' political will in deploying RE policies could explain this finding.

Table 11 D-H panel causality test results

Null hypothesis (H$_0$)	W-bar	Z-bar	p-value	Result
LGDP⇏LEF	4.437***	13.089***	0.005	Causality
LEF⇏LGDP	1.822	3.131	0.140	No causality
LRE⇏LEF	1.989	3.769	0.340	No Causality
LEF⇏LRE	2.381*	5.262*	0.095	Weak Causality
LNRE⇏LEF	3.502***	9.527***	0.000	Causality
LEF⇏LNRE	3.572***	9.795***	0.000	Causality
IQ⇏LEF	16.893***	10.019***	0.000	Causality
LEF⇏IQ	11.533***	3.215***	0.001	Causality

Note: The bootstrap D-H test statistics are calculated using 200 bootstrap replications. Optimal lag length is chosen based on the Bayesian information criterion. ⇏ denotes no causal effect. ***, **,* indicate statistical significance at 1%, 5%, 10% level of significance respectively

Conclusions and policy orientations

This study aims at the simultaneous investigation of the dynamic effects of institutional quality and disaggregated energy consumption on the EF of 29 OECD countries during the period 1984–2016. It is the first attempt to incorporate the institutional dimension in a disaggregated energy-environmental framework for OECD countries. For this reason, we generate a constructive institutional index which includes the most critical institutional parameters. The article also adopts a holistic environmental approach, through the use of the ecological footprint indicator, to extend the concept of environmental performance into the overall ecosystem's performance. Accounting for the detected CS dependence and heterogeneity, though we apply several panel estimation techniques, we base our conclusions on second-generation methodologies, and to the recently developed CS-DL estimator.

The panel cointegration tests reveal the existence of a long-run relationship between the variables in the examined model. The long-run estimates show the stimulative effect of economic growth and non-RE consumption on the EF of OECD countries. Moreover, institutional quality seems to be the most important factor that exerts a positive influence on the environmental performance of these countries. RE consumption has also a negative association with the EF, but it is not proved to be significant enough. In parallel, the D-H causality analysis shows a unidirectional causality from economic growth to the EF and a weak one-way causal effect running from the second to RE consumption. A bidirectional causality is also detected between non-RE and EF and between institutional quality and EF.

The need for environmental policy orientation in OECD countries is not a new project, since Scruggs (1999) was the first to argue it. In this line, this study supports some important policy implications based on the empirical findings. First of all, as OECD countries are the instigators of the SDGs and the Green Growth Strategy (GGS) plan (OECD 2012), they have to reassess their national energy policy portfolios. Most of these countries are subjected to the world's highest economic growth rates, a fact that makes their need for new energy source exploitation imperative. However, because of their high energy demand, and the broad availability and low cost of fossil fuels energy, it seems that most of their energy needs are still covered by non-renewables. The massive use of fossil fuels inevitably causes a rise in the ecological footprint of these regions.

Therefore, it is evident that OECD countries should give priority to the gradual increase of the renewables share in their national energy-mix, by promoting the substitution of fossil energy from green energy forms. This could be achieved through the establishment of incentive mechanisms for accessible and affordable renewable energies. These incentives could include the implementation of monetary credit channels which can offer lower interest rates on "green enterprises," tax reductions, green projects certifications, and the provision of RE subsidies on households. Furthermore, the cooperation between public and private entities through the development of Public-Private Partnerships (PPPs) could facilitate the adoption and implementation of renewable technologies.

Nevertheless, the transition to RE forms requires regulatory and institutional efficiency. Institutional quality in OECD countries is associated with a well-defined legislative and regulatory framework that protects property rights and ensures transparent economic processes. Moreover, governmental efficiency and effectiveness act as a warranty on the socio-economic environment of these countries, as governments have the dynamics to implement sustainable environmental policies. Governments should enhance their energy strategies and the relevant policy orientation, in order to overcome institutional barriers that hinder the deployment of renewable energies. These barriers are associated with the proper energy market regulation, the RE installed capacity, the investment implementation in research and development (R&D), and the public information and perception towards the utilization of renewables. Energy regulatory bodies could pursue

reforms to reinforce RE, instead of non-RE generation in a level playing field of the energy market. Moreover, institutions should communicate the social dimension of adopting renewables, through an educational approach for the environmental and health benefits.

Finally, OECD governments should strengthen the role of local and national institutions and engage private stakeholders to participate actively in the formulation, monitoring, and evaluation of policies for environmental sustainability. Furthermore, the cooperation among OECD countries (especially in clean technologies applications) is considered critical in the process of materializing their environmental projects in light of their undertaken SDGs. As IEA (2019) argues, the path to a sustainable environment still requires a long way to go, and countries have to undertake a lot of policy measures to integrate renewables into the energy system substantially.

Appendix

Table 12   List of countries

	Countries of the study				
1.	Australia	11.	Hungary	21.	Norway
2.	Austria	12.	Ireland	22.	Poland
3.	Belgium	13.	Israel	23.	Portugal
4.	Canada	14.	Italy	24.	Spain
5.	Chile	15.	Japan	25.	Sweden
6.	Denmark	16.	Korea (South)	26.	Switzerland
7.	Finland	17.	Luxembourg	27.	Turkey
8.	France	18.	Mexico	28.	UK
9.	Germany	19.	Netherlands	29.	US
10.	Greece	20.	New Zealand		

Declarations

Ethics approval and consent to participate   Not applicable.

Consent for publication   Not applicable.

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References

Abbasi MA, Parveen S, Khan S, Kamal MA (2020) Urbanization and energy consumption effects on carbon dioxide emissions: evidence from Asian-8 countries using panel data analysis. *Environ Sci Pollut Res* 27:18029–18043

Abid M (2016) Impact of economic, financial, and institutional factors on CO_2 emissions: evidence from sub-Saharan Africa economies. *Util Policy* 41:85–94

Abid M (2017) Does economic, financial and institutional developments matter for environmental quality? A comparative analysis of EU and MEA countries. *J Environ Manag* 188:183–194

Abul SJ, Satrovic E, Muslija A (2019) The link between energy consumption and economic growth in Gulf Cooperation Council Countries. *Int J Energy Econ Policy* 9(5):38–45

Ali H.S., Zeqiraj, V., Lin, W.L, Law S.H., Yusop Z., Ali Bare U.A., Chin L. (2019) Does quality institutions promote environmental quality? *Environ Sci Pollut Res* 26: 10446–10456

Ali S, Yusop Z, Kaliappan SR, Chin L (2020) Dynamic common correlated effects of trade openness, FDI, and institutional performance on environmental quality: evidence from OIC countries. *Environ Sci Pollut Res* 27:11671–11682

Al Mulali U, Ozturk I (2015) The effect of energy consumption, urbanization, trade openness, industrial output, and the political stability on the environmental degradation in the MENA (Middle East and North African) region. *Energy* 84:382–389

Alvarez-Herranz A, Balsalobre-Lorente D, Shahbaz M, Cantos JS (2017) Energy innovation and renewable energy consumption in the correction of air pollution levels. *Energy Policy* 105:386–397

Ansari MA, Haider S, Khan NA (2020) Environmental Kuznets curve revisited: An analysis using ecological and material footprint. *Ecol Indic* 115:106416

Apergis N, Payne JE (2014) The causal dynamics between renewable energy, real GDP, emissions and oil prices: evidence from OECD countries. *Appl Econ* 46(36):4519–4525

Apergis N, Payne JE, Menyah K, Wolde-Rufael Y (2010) On the causal dynamics between emissions, nuclear energy, renewable energy, and economic growth. *Ecol Econ* 69:2255–2260

Arouri MH, Ben YA, M' Henni H, Rault C (2012) Energy consumption, economic growth and CO_2 emissions in Middle East and North African countries. *Energy Policy* 45:342–349

Azam M, Liu L, Ahmad N (2020) Impact of institutional quality on environment and energy consumption: evidence from developing world. *Environment, Development and Sustainability*

Bekhet HA, Othman NS (2018) The role of renewable energy to validate dynamic interaction between CO_2 emissions and GDP toward sustainable development in Malaysia. *Energy Econ* 72:47–61

Bhattacharya M, Awaworyi CS, Paramati SR (2017) The dynamic impact of renewable energy and institutions on economic output and CO_2 emissions across regions. *Renew Energy* 111:157–167

Bhattarai M, Hammig M (2001) Institutions and the Environmental Kuznets Curve for deforestation: a cross-country analysis for Latin America, Africa and Asia. *World Dev* 29(6):995–1010

Boluk G, Mert M (2014) Fossil & renewable energy consumption, GHGs and economic growth: evidence from a panel of European Union (EU) countries. *Energy* 74:439–446

Breusch TS, Pagan AR (1980) The Lagrange multiplier test and its applications to model specification in econometrics. *Rev Econ Stud* 47(1):239–253

Canh NP, Schinckus C, Thanh SD (2019) Do economic openness and institutional quality influence patents? Evidence from GMM systems estimates. *Int Econ* 157:134–169

Castiglione C, Infante D, Smirnova J (2012) Rule of law and the Environmental Kuznets Curve: evidence for carbon emissions. *International Journal of Sustainable Economics* 4(3):254–269

Charfeddine L, Mrabet Z (2017) The impact of economic development and social–political factors on ecological footprint: a panel data analysis for 15 MENA countries. *Renew Sust Energ Rev* 76:138–154

Chudik A, Pesaran MH (2013) Large panel data models with cross-sectional dependence: A survey. *CAFE Research Paper* 13:15. https://ssrn.com/abstract=2316333 or. https://doi.org/10.2139/ssrn.2316333

Chudik A., K. Mohaddes, Pesaran M. H., Raissi M. (2013) Debt, inflation and growth: Robust estimation of long run effects in dynamic panel data models. *Cambridge Working Papers in Economics* 1350, Faculty of Economics, University of Cambridge

Chudik A, Mohaddes K, Pesaran MH, Raissi M (2016) *Long-run effects in large heterogeneous panel data models with cross-sectionally correlated errors*. In: Hill RC, Gonzalez-Rivera G, Lee T-H (eds) Advances in Econometrics, vol 36. Essays in Honor of Aman Ullah, Emerald Publishing, pp 85–135

Culas RJ (2007) Deforestation and the environmental Kuznets curve: an institutional perspective. *Ecol Econ* 61(2–3):429–437

Damania R, Fredriksson P, List J (2003) Trade liberalization, corruption, and environmental policy formation: Theory and evidence. *J Environ Econ Manag* 46:490–512

Danish ZB, Ulucak R, Khan SU-Din (2019) Determinants of the ecological footprint: Role of renewable energy, natural resources, and urbanization. *Sustain Cities Soc* 54:101996

Danish ZB, Wang B, Wang Z (2017) Role of Renewable Energy and Non-Renewable Energy consumption on EKC: Evidence from Pakistan. *J Clean Prod* 156:855–864

Dasgupta S., De Cian E. (2016) Institutions and the Environment: Existing Evidence and Future Directions. FEEM Working Papers Nota di Lavoro 41.2016, Milan, Italy

De Hoyos RE, Sarafidis V (2006) Testing for cross-sectional dependence in panel-data models. *Stata J* 6(4):482–496

Destek M, Sinha A (2020) Renewable, non-renewable energy consumption, economic growth, trade openness and ecological footprint: Evidence from organization for economic Co-operation and development countries. *J Clean Prod* 242:118537

Destek MA, Ulucak R, Dogan E (2018) Analyzing the environmental Kuznets curve for the EU countries: The role of ecological footprint. *Environ Sci Pollut Res* 25:29387–29396

Dogan E, Ozturk I (2017) The influence of renewable and non-renewable energy consumption and real income on CO_2 emissions in the USA: evidence from structural break tests. *Environ Sci Pollut Res* 24:10846–10854

Dogan E, Seker F (2016a) Determinants of CO_2 emissions in the European Union: The role of renewable and non-renewable energy. *Renew Energy* 94:429–439

Dogan E, Seker F (2016b) The influence of real output, renewable and non-renewable energy, trade and financial development on carbon emissions in the top renewable energy countries. *Renew Sust Energ Rev* 60:1074–1085

Dumitrescu EI, Hurlin C (2012) Testing for granger non-causality in heterogeneous panels. *Econ Model* 29(4):1450–1460

Dutt K (2009) Governance, institutions and the environment-income relationship: a cross-country study. *Environ Dev Sustain* 11(4):705–723

Edenhofer O, Pichs-Madruga R, Sokona Y, Seyboth K, Matschoss P, Kadner S, Zwickel T, Eickemeier P, Hansen G, Schlmer S, von Stechow C (2011) *IPCC Special Report on Renewable Energy Sources and Climate Change Mitigation*. In: Cambridge University Press, Cambridge, United Kingdom. NY, USA, New York

Fredriksson PG, Vollebergh HRJ, Dijkgraaf E (2004) Corruption and energy efficiency in OECD countries: theory and evidence. *J Environ Econ Manag* 47(2):207–231

Fudge S, Hunt L, Jackson T, Mulugetta Y, Peters M (2008) The Political economy of energy regulation in the UK 1945–2007: Paradigms and policy. RESOLVE Working Paper:02–08

Gani A (2012) The Relationship between good governance and carbon dioxide emissions: Evidence from developing economies. *J Econ Dev* 37(1):77–93

GFN (2020) Global Footprint Network. Available at https://data.footprintnetwork.org/#/ . Accessed 01 September 2020

Goodland R (1995) The concept of environmental sustainability. *Annu Rev Ecol Syst* 26:1–24

Granger CW (1969) Investigating causal relations by econometric models and cross spectral methods. *Econometrica* 37(3):424–438

Hassan ST, Danish KSUD, Xia E, Fatima H (2020) Role of institutions in correcting environmental pollution: An empirical investigation. *Sustain Cities Soc* 53:101901

Iacobuta AO, Gagea M (2010) Institutional quality, economic freedom and sustainable development. A comparative analysis of EU countries. *Rev Econ* 3(50):252–260

Ibrahim MH, Law SH (2016) Institutional quality and CO_2 emission–trade relations: evidence from Sub-Saharan Africa. *South African Journal of Economics* 84(2):323–340

IEA, IRENA, UNSD, WB, WHO (2019), Tracking SDG 7: The Energy Progress Report 2019, Washington DC. Available at: https://irena.org/-/media/Files/IRENA/Agency/Publication/2019/May/2019-Tracking-SDG7-Report.pdf . Accessed 14 July 2020

IEO (2019) International Energy Outlook U.S. Energy Information Administration, Office of Energy Analysis, U.S. Department of Energy Washington, DC 20585. Available at: https://www.eia.gov/ieo. Accessed 11 May 2020

Im KS, Pesaran MH, Shin Y (2003) Testing for unit roots in heterogeneous panels. *J Econ* 115(1):53–74

IPCC (2018) Global Warming of 1.5°C. *An IPCC Special Report on the impacts of global warming of 1.5°C above pre-industrial levels and related global greenhouse gas emission pathways, in the context of strengthening the global response to the threat of climate change*. In press. Available at: https://www.ipcc.ch/site/assets/uploads/sites/2/2019/06/SR15_Full_Report_High_Res.pdf . Accessed 14 July 2020

IRENA IEA, REN21 (2018) *Renewable Energy Policies in a Time of Transition*. IRENA, OECD/IEA and REN21

Jordan A and Lenschow A (2008) *Innovation in Environmental Policy? Integrating the Environment for Sustainability*. Edited by: Edward Elgar: Cheltenham, UK

Kao C, Chiang M (2001) On the estimation and inference of a cointegrated regression in panel data. *Adv Econ* 15(1):179–222

Kapetanios G, Pesaran MH, Yamagata T (2011) Panels with nonstationary multifactor error structures. *J Econ* 160(2):326–348

Khalil M, Ellaboudy S, Denzau A (2007) The Institutions and Economic Development in the OECD. *Int Res J Financ Econ* 12:67–79

Kitzes J, Wackernagel M (2009) Answers to common questions in ecological footprint accounting. *Ecol Indic* 9:812–817

Lau LS, Choong CK, Eng YK (2014) Carbon dioxide emission, institutional quality, and economic growth: empirical evidence in Malaysia. *Renew Energy* 68:276–281

Le HP, Ozturk I (2020) The impacts of globalization, financial development, government expenditures, and institutional quality on CO_2 emissions in the presence of environmental Kuznets curve. *Environ Sci Pollut Res* 27:22680–22697

Lee JW (2019) Long-run dynamics of renewable energy consumption on carbon emissions and economic growth in the European Union. *Int J Sust Dev World* 26(1):69–78

Lehtonen M (2004) The environmental–social interface of sustainable development: capabilities, social capital, institutions. *Ecol Econ* 49(2):199–214

Lin B, Moubarak M (2014) Renewable energy consumption – economic growth nexus for China. *Renew Sust Energ Rev* 40:111–117

Liu JL, Ma CQ, Ren YS, Zhao XW (2020a) Do real output and renewable energy consumption affect CO2 emissions? Evidence for selected BRICS countries. *Energies* 13(4):960

Liu X, Latif K, Latif Z, Li N (2020b) Relationship between economic growth and CO2 emissions: does governance matter? *Environ Sci Pollut Res* 27:1–8

Lovei M, Weiss C (1998). *Environmental Management and Institutions in OECD Countries*. World Bank Technical paper no. 391. Pollution Management Series. Washington, DC: The World Bank

Maddala GS, Wu SA (1999) Comparative study of unit root tests with panel data and a new simple test. *Oxf Bull Econ Stat* 108:1–24

Menyah K, Wolde-Rufael Y (2010a) Energy consumption, pollutants emissions and economic growth in South Africa. *Energy Econ* 32(6):1374–1382

Menyah K, Wolde-Rufael Y (2010b) CO_2 emissions, nuclear energy, renewable energy and economic growth in the U.S. *Energy Policy* 38:2911–2915

Nathaniel S, Anyanwu O, Shah M (2020) Renewable energy, urbanization, and ecological footprint in the Middle East and North Africa region. *Environ Sci Pollut Res* 27:14601–14613

Nawaz R, Azam M, Zulfiqar H (2020) Economic Growth, Energy and Environmental Degradation Nexus: A Case Study in Pakistan. *Journal of Economic Impact* 2(1):13–23

OECD (2001) *OECD Environmental Strategy for the First Decade of the 21st Century*. OECD, Paris

OECD (2008) *Handbook on Constructing Composite Indicators: Methodology and User Guide*. OECD Publishing, Paris

OECD (2012) *Green Growth and Developing Countries*. OECD Publishing Paris, Consultation Draft http://www.oecd.org/development/environment-development/50559116.pdf

Ostrom E (1998) A behavioral approach to the rational choice theory of collective action: Presidential address, American political science association, 1997. *Am Polit Sci Rev* 92(1):1–22

Panayotou T (1997) Demystifying the environmental Kuznets curve: turning a black box into a policy tool. *Environ Dev Econ* 2(4):465–484

Paramati SR, Sinha A, Dogan E (2017) The significance of renewable energy use for economic output and environmental protection: evidence from the Next 11 developing economies. *Environ Sci Pollut Res* 24:13546–13560

Pedroni P (1999) Critical values for cointegration tests in heterogeneous panels with multiple regressors. *Oxf Bull Econ Stat* 61:653–670

Pedroni P (2001) Purchasing Power Parity Tests in Cointegrated Panels. *Rev Econ Stat* 83:727–731

Pedroni P (2004) Panel cointegration: asymptotic and finite samples properties of pooled time series tests with an application to the PPP hypothesis. *Econometric Theory* 20:597–625

Pesaran MH (2004) General diagnostic tests for cross section dependence in panels. *Cambridge Working Papers in Economics* No. 435, University of Cambridge, and CESifo Working Paper Series No. 1229

Pesaran MH (2007) A simple panel unit root test in the presence of cross-section dependence. *J Appl Econ* 22(2):265–312

Pesaran MH, Yamagata T (2008) Testing slope homogeneity in large panels. *J Econ* 142(1):50–93

Pesaran MH, Ullah A, Yamagata T (2008) A bias-adjusted LM test of error cross-section independence. *Econ J* 11(1):105–127

Rees WE (1992) Ecological footprints and appropriated carrying capacity: what urban economics leaves out? *Environ Urban* 4(2):121–130

Ritzen J, Woolcock M (2000) *Social cohesion, public policy, and economic growth: implications for countries in transition*. In: Annual Bank Conference on Development Economics, Paris, pp 26–28

Saidi K, Omri A (2020) The impact of renewable energy on carbon emissions and economic growth in 15 major renewable energy-consuming countries. *Environ Res* 186:109567

Salman M, Long X, Dauda L, Mensah CN (2019) The impact of institutional quality on economic growth and carbon emissions: Evidence from Indonesia, South Korea and Thailand. *J Clean Prod* 241:118331

Sarkodie SA, Adams S (2018) Renewable energy, nuclear energy, and environmental pollution: accounting for political institutional quality in South Africa. *Sci Total Environ* 643:1590–1601

Scruggs LA (1999) Institutions and environmental performance in seventeen western democracies. *Br J Polit Sci* 29(1):1–31

Shafiei S, Salim RA (2014) Non-renewable and renewable energy consumption and CO2 emissions in OECD countries: a comparative analysis. *Energy Policy* 66:547–556

Sharif A, Raza SA, Ozturk I, Afshan S (2019) The dynamic relationship of renewable and nonrenewable energy consumption with carbon emission: A global study with the application of heterogeneous panel estimations. *Renew Energy* 133:685–691

Sharif A, Baris-Tuzemen O, Uzuner G, Ozturk I, Sinha A (2020a) Revisiting the role of renewable and non-renewable energy consumption on Turkey's ecological footprint: Evidence from Quantile ARDL approach. *Sustain Cities Soc* 57:102138

Sharif A, Mishra S, Sinha A, Jiao Z, Shahbaz M, Afshan S (2020b) The renewable energy consumption-environmental degradation nexus in Top-10 polluted countries: Fresh insights from quantile-on-quantile regression approach. *Renew Energy* 150:670–690

Sharma SS (2011) Determinants of carbon dioxide emissions: empirical evidence from 69 countries. *Appl Energy* 88(1):376–382

Swamy PA (1970) Efficient inference in a random coefficient regression model. *Econometrica* 38(2):311–323

Tamazian A, Bhaskara RB (2010) Do economic, financial and institutional developments matter for environmental degradation? Evidence from transitional economies. *Energy Econ* 32:137–145

UN SDGs (2015) Transforming our world: *The 2030 Agenda for Sustainable Development*, A/RES/70/1. United Nations, General Assembly, New York

United Nations (2016) Global Sustainable Development Report 2016, Chapter 4: *Inclusive institutions for sustainable development*, Department of Economic and Social Affairs, New York

United Nations (2019) *The Sustainable Development Goals Report:2019* Available at: https://unstats.un.org/sdgs/report/2019/The-Sustainable-Development-Goals-Report-2019.pdf

Urbain JRYJ, Westerlund J (2006) Spurious regression in nonstationary panels with cross-unit cointegration. *METEOR Research Memorandum No* 057

US Energy Information Administration (2020) Available at: https://www.eia.gov/international/data/world . Accessed 01 June 2020

Wackernagel M, Monfreda C (2004) *Ecological footprints and energy*. In: Cleveland CJ (ed) Encyclopedia of Energy, vol 2. Elsevier, Amsterdam, pp 1–11

Wackernagel M, Rees WE (1996) *Our ecological footprint: reducing human impact on the Earth*. New Society Publishers, Gabriola Island, British Columbia, Canada

Wagner M, Hlouskova J (2010) The performance of panel cointegration methods: results from a large scale simulation study. *Econ Rev* 29(2):182–223

Wang J, Dong K (2019) What drives environmental degradation? Evidence from 14 Sub Saharan African countries. *Sci Total Environ* 656:165–173

Wang ZH, Danish ZB, Wang B (2018) The moderating role of corruption between economic growth and CO_2 emissions: Evidence from BRICS economies. *Energy* 148:506–513

Westerlund J (2007) Testing for error correction in panel data. *Oxf Bull Econ Stat* 69:709–748

World Development Indicators (WDI) (2020) World Bank Development Indicators database. Available at https://data.worldbank.org/. Accessed 01 June 2020

Yasin I, Ahmad N, Chaudhary MA (2020) Catechizing the Environmental-Impression of Urbanization, Financial Development, and Political Institutions: A Circumstance of Ecological Footprints in 110 Developed and Less-Developed Countries. *Soc Indic Res* 147:621–649

Zaghdoudi T (2017) Oil prices, renewable energy, CO_2 emissions and economic growth in OECD countries. *Econ Bull* 37:1844–1850