Investigating factors affecting global environmental sustainability:
evidence from nonlinear ARDL bounds test

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Abstract
The concept of environmental sustainability formed the basis of the Paris Agreement and the United Nations conference in Rio de Janeiro. Empirically, without environmental sustainability, everything else could fall apart or be aimless. This study investigates factors affecting global environmental sustainability spanning 1966Q1 to 2019Q4. However, there are many micro-/macroeconomic factors engendering the environment, and the absence of robust clarity on whether factors such as economic growth, urbanization, trade openness, and energy consumption matter for global environmental sustainability remains a global academic dilemma in the economics literature. This paper utilized the unrestricted nonlinear autoregressive distributed lag (NARDL) bounds test techniques to model their relationship. Furthermore, the study adopted fully modified ordinary least squares (FMOLS), dynamic ordinary least squares (DOLS), and canonical cointegrating regression (CCR) methods to test the research hypothesis, catering to the problem of endogeneity and serial correlation. Up-to-date of this study, no empirical study has examined the nexus of these variables within the global framework. The outcomes suggested that (i) NARDL bounds test of cointegration confirmed evidence of long-run and short-run relationships among the variables; (ii) long-run asymmetric relationship was affirmed among the variables; and (iii) DOLS, FMOLS, and CCR models demonstrate that economic growth, energy consumption, and trade openness are positively significantly correlated with environmental sustainability except for economic growth which shows negative and insignificant correlation. These findings validate the protracted argument in literature that these estimated variables are significant for global environmental sustainability. This study recommends that environmental policymakers integrate global economic incentives with favorable regulatory changes for achieving the goals of a global sustainable environment in the long-run equilibrium.

Keywords CO₂ emissions · Environmental sustainability · Global context · NARDL model

Introduction
Environmental sustainability (CO₂ emissions) has been debated several times on corporate agenda around the world. It is concerned with intergenerational decisions making, targeting social welfare or happiness, treating natural resources as if they have intrinsic rights and value, and preventing the consequences of artificial global warming have been highlighted in recent debates. Since the eighteenth century, that is, the beginning of industrial activities, literally hundreds of regional and bilateral agreements and more than 10 global environmental treaties have been negotiated and signed to ensure environmental quality (Hunter 2018). For instance, the nonbinding international environmental treaty of 1992, the legally binding framework of the 1997 Kyoto Protocol, the formation of the UNFCCC, the 2015 Paris Agreement, and followed by the IPCC2 set achieving energy security and the long-term goal of sustaining global warming at the 1.5 °C range (Wang et al. 2022). However, despite general progress made in environmental quality, the main factors affecting the natural environment remain a global dilemma, and it has emerged as one of the significant challenges in the economics literature (Wang et al. 2022). Moreover, the past three decades show that the global human tolls of environmental impacts...
remain profoundly worrying, and it requires robust global policy actions. Weltgesundheitsorganisation, World Health Organization, and European Centre for Environment (2021) highlight that the global community is suffering from serious social and environmental problems, including health diseases, carbon emissions, and climate change. Several theories and empirical evidence have suggested that human resource demands have damaged environmental support systems (Hunter 2018; Malthus 1965). Similarly, studies by Ricardo and Mill assumed scarcity of resources, while the laws of thermodynamics noted that a complete separation between the economy and the environment is impossible (Söllner 1997). Economic activity will always use a certain amount of resources; thus, higher productivity could gradually affect the environment. On the other hand, theories by Hartwick-Solow (Gutés 1996), Kuznets (1955), and Daly’s (1973) criticized the above views. For example, Hartwick-Solow’s study focused on advancing global technologies and substitutions effects that reduce the creation of harmful residues. Another essential reference point could be the material balance hypothesis (Dykes et al. 2020).

Apart from the theories mentioned earlier, empirical studies by Nassani et al. (2021) and Yildiz et al. (2017) have argued that the lack of safeguarding the natural environment and its species of animals is the most crucial factor affecting the ecosystems. Nevertheless, a lack of robust clarity on whether other factors such as economic growth, globalization, energy consumption, livestock productivity, urbanization, and trade openness matter for global environmental sustainability remains as an academic dilemma in the economics literature. The prevention of environmental degradation is an important issue in the global context. Economists and environmental policymakers in recent years have renewed their interests in investigating factors affecting environmental sustainability from a diverse perspective. Moreover, such an investigation could provide useful information to those responsible for formulating global policy to tackle climate change. Unfortunately, up to this present study, no paper has investigated whether factors such as trade openness, urbanization, energy consumption, and economic growth matter for environmental sustainability in context. To close such a knowledge gap in the literature, this paper analyzes factors affecting global environmental sustainability spanning 1966Q1 to 2019Q4.

From the global perspective, environmental challenges are further exacerbated by inconsistent weather crises caused by global warming. Global efforts to alleviate extreme environmental challenges remain a global dilemma; the recent year’s acceleration rate is unprecedented. For example, Australia and the USA recently experienced one of the most devastating bushfires ever recorded. Sovacool, Griffiths, Kim, and Bazilian’s (2021) study noted that global CO2 emissions rebounded by 4.9% due to fossil fuels and cement consumption. A long-term trend of the global CO2 emissions forecast of 2022 is 36.4% higher. The most recent analysis shows that the widespread usage of fossil fuels raised the global CO2 emissions to 32.8% from 1990 to 2021 (Jiang et al. 2022). On a country percentage basis, China and India surpassed their 2019 estimated emission peaks in 2021. India led the upward surge in emissions; the Global Carbon Project also noted that emission growth of 10.7% dropped to 1.4% (2.4 billion mt) 2020 in China. 2021 Chinese emissions have grown from 5.5 to 11.1%. Nevertheless, during the COVID-19 pandemic, emissions estimated value in the US and the EU fell around 11%, while the rest world fell 7% between 2019 and 2021. Table 1 summarizes fossil CO2 emissions by region/country in 2020 and 2021.

In addition to raising CO2 emissions, the world population threatens environmental sustainability. Intensive urbanization could harm the environment by depleting the soil and marine ecosystems. Taylor (2022) expressed that more than 500 land species of animals are on the brink of extinction and may likely be lost within 20 years. Also, in 1950, the world’s urban population was 751 million, has rapidly grown to 4.5 billion in the past few years, and is projected to expand to about 6.68 billion in 2050 (Taylor 2022). The Caribbean and Latin America (81.5%), Europe (75%), America Northern (82.75%), and Oceania (68%) are the most urbanized regions in the world. In Asia, the level of urbanization is now approximating 52%. Africa has the fastest population growth rate of 3.56%, followed by Asia (1.98%). However, recent statistics show that 44% of the

| Table 1 Fossil CO2 emissions in 2020 and 2021 |
|-----------------------------------------------|
| Region/Country | 2020 emissions (billion tonnes/yr) | 2020 growth (present) | 2021 growth (present) | 2021 emissions (billion tonnes/yr) |
| China          | 10.7 | 1.40% | 5.50% | 11.1 |
| USA            | 14.7 | -10.60% | 11.60% | 5.1 |
| India          | 2.4 | -7.30% | 12.60% | 2.7 |
| EU27           | 2.6 | -10.90% | 11.60% | 2.8 |
| All others (incl.IAS*) | 14.4 | -7.00% | 2.90% | 14.8 |
| World (incl. IAS*) | 34.8 | -5.40% | 7.90% | 36.4 |

Data source: Friendlingstein (2022)
populations of Africa live in urban areas; unfortunately, African towns remain mostly rural (Taylor 2022).

Similar to urbanization, economic growth, which is meant to measure living standards, has conflicts with raising environmental problems. However, its challenges are very different in different countries. The USA and Europe face particular issues that look different from those faced in India or China. In 2021, global economic growth decelerated from 5.5 to 4.1%. The continued disruption was caused by COVID-19 and changed due to supply bottlenecks. In addition to global economic growth challenges, widely used energy such as fossil fuels has also brought environmental problems. It is assumed that activities such as transport, building infrastructure, water distribution, and food production consume about 75% of global primary energy, and they emit about 60% of the world’s total greenhouse gases (Cristea et al. 2013). In addition to these factors, trade openness is another environmental issue the world faces globally. According to Guterres (2022), global trade value has reached $28.5 trillion in 2021. Ironically, it may be impossible to globalize trade, utilize energy consumption to industrialize, and increase economic growth per capita without increasing environmental pollutions because most of these factors instigate carbon emissions. Thus, environmentalists believe a trade-off must exist between environmental factors and proxy for environmental sustainability. Understanding factors affecting environmental sustainability in a global context seems significant.

Based on the above discussions, this study makes various contributions to the literature of environmental economics in several ways. First, the earlier studies in this regard focus more on the panels of different countries in describing the vigorous relationship between renewable and nonrenewable energy consumption and environmental degradation (Sharif et al. 2019). To the best of our knowledge, very little and sporadic research exists in the scope of the study (Yildiz et al. 2017); no one has paid attention to factors such as economic growth, urbanization, energy consumption, and trade openness matter for environmental sustainability in a global context. To close this existing gap in the literature, the study broadens knowledge of the relationship among trade openness, urbanization, economic growth, energy consumption, and environmental sustainability (CO2 emissions). Second, unlike previous studies that has used inappropriate proxies such as ecological footprint for environmental sustainability, which most often has led to discrepancies in the reported results (Husnain et al. 2021; Pata 2021), this present study utilized per capita CO2 emission as a proxy variable for environmental sustainability. In recent study, per capita CO2 emission has been demonstrated as a better and more comprehensive environmental variable than an ecological footprint (Pata 2021). Third, also, previous literature has neglected the “unit root test with structural breaks” (Olasunkanmi 2021). The omission of structural breaks in the time series analysis represents a critical research gap. This paper fills the gap in the economics literature by employing unit root tests of “one Endogenous Structural Break and Zivot-Andrews endogenous single break unit root test.” In addition, other unit root tests such as the conventional ADF unit root test and Phillip-Peron (PP) unit root were employed to ensure that the stationarity properties of the series in the estimations are not I(2) or higher.

Fourth, the traditional ARDL bounds test has come under massive criticism for incorrectly specifying the nature of a long-run relationship (Shahbaz et al. 2014; Shin et al. 2014). For instance, a study that investigated nexus between unemployment and economic output in Canada, Japan, and the USA was done by Shin et al. (2014) found strong evidence of a long-run asymmetry relationship between unemployment and output. When the traditional ARDL bounds test was applied on the same subject, the results showed no evidence of a long-run relationship. In this context, utilizing nonlinear autoregressive distributed lag (NARDL) estimator is critical if one must avoid drawing an incorrect inference or making incorrect policy decisions. To close the existing gap in the literature, the present paper adopted the unrestricted NARDL bounds test of the cointegration framework. Like traditional ARDL bounds test, the NARDL model is largely flexible in allowing a mixed order of I(1) and I(0) regressors, which is not applicable for the traditional cointegration tests of Johansen and Juselius (1990). In addition, the NARDL model has the advantage of separating the reactions of the dependent variable (i.e., negative and positive changes) and independent variables by utilizing the Wald test.

Fifth, furthermore, evidence of asymmetry and cointegration relationships among variables does not necessarily suggest model stability, model robustness, and problem of multi-collinearity over the sample period. Hence, to fill the gap, this study utilized residual models of Breusch-Godfrey serial correlation LM test, Breusch-Pagan-Godfrey heteroskedasticity test, cumulative sum chart (CUSUM), and Ramsey regression equation specification error (RESET) test, respectively. Lastly, the lack of endogeneity control, which most often contributes to reverse causality, sample bias, and measurement error, is omitted in the time series analysis and represents a critical research gap. This paper fills the gap in the economics literature by employing dynamic ordinary least squares (DOLS), fully modified ordinary least squares (FMOLS), and canonical cointegrating regression (CCR), respectively. Although there are other robustness check approaches in the literature, our select estimators are relatively the more stable for addressing second-order bias, endogeneity, and serial correlation issues that could exist in cointegrating regressions.

The remaining of this paper is organized to track the following subsections: “Literature review” elaborates the
related literature. “Data sources and model” covers the source of datasets and methodology framework applied in this paper; “Methodology” discusses the empirical formulation, analysis, and findings. Finally, “Empirical results and discussions” presents the conclusion, policy implications, and study limitations, and it exposes some future research directions.

**Literature review**

Changes in climate have turned out to be significantly extreme during the last several decades, a hot concern for argument worldwide and a global happening due to the hazard to sustainable advancement. An environmental problem, especially global warming, is one of the most challenging issues in the modern world’s agenda. CO2 emission has been a leading issue, and all efforts have been put into eliminating it (Godil et al. 2021). Understanding factors affecting environmental sustainability (CO2 emission) is a necessity for environmental policymakers and economists across the globe. It is particularly important for decarbonization, green practices and prudential policy guide that benefits societies. In recent years, the ecological modernization theory has become one of the leading perspectives of decarbonization and the role of green practices in business operations (Khan et al. 2019). The theory argues that modern green technology and innovations (green practices in businesses) enhance environmental performance in terms of reducing the harmful effect on environmental sustainability (Sharif et al. 2020). It can be observed from the trend analysis that CO2 emissions from different sources have been persistently increasing over time. Against this backdrop, this study aimed to analyze factors affecting global environmental sustainability (measures by CO2 emissions) spanning 1966Q1 to 2019Q4. This study pays special attention to the global context because of the dynamic structure of the long-standing global environmental challenges. The current global CO2 emissions represent upwards of 90% Nassani et al. 2021). Recent publications have argued that the lack of preserving ecosystems was the most crucial factor affecting environmental sustainability (Nassani et al. 2021; Yildiz et al. 2017). Furthermore, the laws of thermodynamics (Genovese et al. 2017) noted that a complete separation between the economy and the environment is impossible. Thus, economic activity will always use a certain amount of resources. Theoretically, higher productivity could gradually decline the environment (Gutés 1996). Despite some level of general theoretical consensus, there were criticisms. Hartwick-Solow (Gutés 1996), Kuznets (1955), and Daly (1973) theories criticized the above views. Porter’s hypothesis suggested the mechanisms through which stringent environmental policies could reduce the high consumption of nonrenewable energy (Usman and Balsalobre-Lorente 2022). The objective of this study is to address the question of whether urbanization, economic growth, energy consumption, and trade openness affect global environmental sustainability (CO2 emissions) spanning from 1966Q1 to 2019Q4. To a greater extent, this paper provides answers to the two specific significant questions: (1) what happens to environmental sustainability (CO2 emissions) when urbanization, economic growth, energy consumption, and trade openness increase or decrease? (2) What could be the direction and magnitude of the beta coefficients of environmental sustainability when urbanization, economic growth, energy consumption, and trade openness increase (X+) or decrease (X−)? The empirical evidence of the nexus between economic growth, urbanization, energy consumption, trade openness, and CO2 emissions is detailed in the subsections.

**Empirical nexus of urbanization and CO2 emissions**

The nexus between urbanization and CO2 emissions has been established and shows contradictory results across the available empirical evidence (Rahman and Alam 2021; Nathaniel and Khan 2020). Among the empirical evidence that examined nexus between urbanization and CO2 emissions includes Kirikkaleli and Sowah (2020); Iheonu et al. (2021); Rahman and Alam (2021); Nathaniel and Khan (2020); and Adedoyin and Zakari (2020). Iheonu et al.’s (2021) study investigated the effect of urbanization on CO2 emissions using sub-Saharan African countries’ datasets. The study employed the panel quantile regression model, and it findings suggest that the urbanization rate increases CO2 emissions across the observed quantiles. Also, bi-directional causality was observed between urbanization and CO2 emissions. Rahman and Alam (2021) study employed the econometric technique of the ARDL bounds test and Toda-Yamamoto Granger causality test to provide evidence on population density, urban population, and CO2 emissions (i.e., measures of environmental pollution) in Bangladesh covering 1973–2014. The outcomes reveal that urban population growth and population density are detrimental to the increase in CO2 emissions in Bangladesh. Similarly, Kirikkaleli and Sowah’s (2020) paper employed the “wavelet coherence technique” to investigate the “co-movement between urbanization and environmental sustainability” (CO2 emissions) in a global context covering 1950 to 2014. Results reveal a significant wavelet relationship between urbanization and environmental sustainability both in the medium and long terms. Also, significant vulnerabilities were identified between urbanization and environmental sustainability at different frequency levels. Finally, Ghazali and Ali (2019) employed the EKC hypothesis based on the STIRPAT method for developing countries covering 1975 to 2018. The results show an inverted U-shaped relationship; the
urban population increases CO₂ emissions. On the contrary, Nathaniel and Khan (2020) and Adedoyin and Zakari (2020) studies have failed to reveal any significant inverted U-shaped relationship among estimated variables.

**Empirical nexus between economic growth and CO₂ emissions**

The recent debates illustrate that the global economic growth upward trend has conflicts with the rising CO₂ emissions. It is observed that growth in GDP per capita is accompanied by growth in CO₂ emission; thus, the upward trend in both GDP per capita and CO₂ emission could mean that GDP per capita is an important driver to the rising CO₂ emissions. Many industries and business leaders believe that we must trade off economic growth and environmental issues (Xiong et al. 2021). Theoretically, we can accomplish both (Kuznets 1955); in fact, economic growth implies increased pollution levels simply due to increased output. Murshed et al. (2021); Nathaniel et al. (2021a); Agboola et al. (2021); Nathaniel et al. (2021b); and Doğan et al. (2022) are recent empirical papers that investigated the association among CO₂ emissions and per capita economic growth.

Murshed et al. (2021) examined the relationship between GDP per capita, ecological footprints, and CO₂ emission in Sri Lanka, Bangladesh, India, and Pakistan. The study used regulatory qualities as a control variable for its cross-sectional panel dependency, slope heterogeneity, and CIPS unit root test. The study reported that regulatory qualities play a significant role in reducing the ecological footprints deficit across the countries. Furthermore, the study argued that economic growth rate and CO₂ emissions have a significant direct relationship. These findings supported the environmental Kuznets curve hypothesis. Furthermore, when renewable energy was employed as another control variable, it authenticated that renewable energy reduced CO₂ emissions. Similar studies by Nathaniel et al. (2021a), and Agboola et al. (2021) supported the Murshed et al. (2021) findings. In contrast, Nathaniel et al. (2021b) and Wan and Yang (2021) studies have failed to identify any negative and significant effects between economic growth and CO₂ emissions. They noted that effective public policy designed such as the 1985 Vienna Convention (i.e., environmental agreement) to protect the ozone layer could help reduce global greenhouse gas emissions.

**Empirical nexus concerning energy consumption and CO₂ emissions**

The impact of energy consumption on CO₂ emission is assumed as one of the major challenges for global development irrespective of its sources. However, energy is needed to power our commercial activities, transport sector, industrial buildings and infrastructure, and water distribution. These activities are the main engines of global economic growth and they cover 75% of a country’s upward trend in GDP per capita. Particularly, renewable energy resources occur in large ecological regions as compared to other energy sources. According to Sharif et al. (2019), global investment in renewable technologies is totaled more than US$214 billion in 35 the year of 2013. Nevertheless, the research on nonrenewable energy (Sharif et al. 2020) shows that the activities consume about 80% of energy and emitted 60% of greenhouse gases, increasing air pollution, and worsening the ecosystem. Nations like India, Russia, and China are among the countries that heavily consume fossil fuels. The need to establish strong policies and standards that could diversify energy consumption has been studied and with contradictory results. Ulucak and Ozcan (2020); Saint Akadiri et al. (2019); Nathaniel and Adeleye (2021); Zhu et al. (2021); Umar et al. (2021); Kirikkaleli and Adebayo (2021); and Sinha et al. (2022) are among the studies that have investigated the link between energy consumption and environmental sustainability (CO₂ emissions).

Ulucak and Ozcan’s (2020) paper utilizes datasets from the OECD countries to examine the relationship between energy consumption, CO₂ emissions, and carbon footprint covering 1980–2016. The study applies an advanced panel data estimation tool of augmented mean group (AMG) estimator. Results reveal that natural resource extraction contributed to increasing CO₂ emissions while renewable energy consumption reduces the negative effects of carbon footprints. When nonrenewable energy was employed as a control variable, it negatively affected the environment. Also, Saint Akadiri et al. (2019) empirical study on energy consumption and CO₂ emissions in South Africa covering 1973 to 2014 provided answers to the question, “what drives what?” The ARDL cointegration employed in this study confirms evidence of long-run cointegration. The Toda-Yamamoto Granger causality test also confirms unidirectional causality running from energy consumption to CO₂ emissions per capita. The ARDL forecast estimator noted that 1% increase in per capita energy consumed could likely lead to 16.7 to 17.2% increase in environmental degradation in both short run and long run. Evidence by Nathaniel et al. (2022) and Umar et al. (2021) have all supported that an increase in nonrenewable energy consumption possibly increases CO₂ emissions. Thus, based on these findings, the green investment could be considered a necessary condition for encouraging clean energy consumption in high pollutant economies.

In contrast, despite the initial negative evidence on the nexus between energy consumption and CO₂ emissions, Kirikkaleli and Adebayo (2021) and Sinha et al. (2022) studies have failed to find any significant negative effect of energy consumption on CO₂ emissions. The shift to renewable energies is expected
to continue. During the COVID-19 lockdown, solar and wind energies became the two most resilient energy sources. Nonrenewable energy output fell between 2019 and 2021, while solar and wind energy experienced an estimated increase of 1.6% in final energy consumption. Sustainability survey in 2021 by Euro-monitor International in Europe noted that 42.3% of professionals interviewed during the survey stated that their company is investing or planning to invest in renewable energies (see Robles 2021). Action to reduce carbon CO2 emissions remains a global priority for companies’ investors and governments across the globe with Chinese companies leading the way.

**Empirical nexus between trade openness and CO2 emissions**

The pollution haven hypothesis assumed that, under the condition of free trade, the industries with intensive pollution could relocate to economies with less environmental regulatory quality. Since the early 1990s, there has been a growing body of empirical studies on the nexus of trade openness, economic growth, and CO2 emissions. The study expressed that trade openness has the direct and indirect structural scale and technical effects, and it plays a major role in determining CO2 emissions (Mishra 2007). A report from UNCTAD noted that the value of global trade reached a record level of $28.5 trillion in 2021 (see Guterres 2022). On a yearly basis, an increase of 25% of global trade occurred in 2019; however, it decreased to 13% as a result of the COVID-19 pandemic, which struck in 2020. Overall, world trade in goods and services significantly declined during the COVID-19 pandemic. Theoretically, the prices of most primary products for the trade are exogenously determined. Thus, trade is affected by a change in the prices of products. Wang and Zhang (2021); Wang and Wang (2021); Dauda et al. (2021); Mutascu (2018); and Tiwari et al. (2022) studies, among others, have examined nexus between trade openness and CO2 emissions and with inclusive results. Dauda et al. (2021) study investigated the potential effect of trade openness on CO2 emissions using panel data from China, Japan, and South Korea covering 1970 to 2019. The study divided the sample into two subsamples, namely before the agreement was signed and after it was signed. The findings indicated that regulatory quality on trade openness positively affects greenhouse gas emissions after signing a trade agreement. In the same vein, Tiwari et al. (2022) utilized a panel dataset from 16 Asian economies covering 1990 to 2019 to investigate the equity market-renewable energy nexus. The novel technique of panel quantile regression proposed by Machado et al. (2018) was employed as an estimator. The result shows that developing the equity market encouraged renewable energy projects of 70% quantile.

In contrast, the study by Mutascu (2018) and Wang and Zhang (2021) has reported mixed results on the nexus between trade openness and CO2 emissions. The study employed datasets from 182 upper-middle-income and high-income countries covering 1990 to 2015. The outcomes show that in upper-middle-income countries, trade openness negatively affected CO2 emissions, while high-income countries’ trade openness positively affected carbon emissions during the datasets covered. When lower-middle-income countries were considered for the analysis, the study failed to find any significant impact of trade openness on CO2 emissions. However, when issues concerning the heterogeneous became highlighted in the analysis, trade openness specified positive influences on CO2 emissions in the rich economies but no significant impacts in emerging nations. Mutascu (2018) paper also explores the nexus between trade openness and CO2 emissions using datasets for Bangladesh covering 1975 to 2013. The study reveals that trade liberalization and CO2 emissions have negative coefficients, while energy consumption and CO2 emissions show positive coefficients in the short-run and long-run. The panel study by Wang and Wang (2021) on 104 countries/regions covering 2000 to 2014 supported Mutascu (2018) and Wang and Zhang (2021) studies. In view of these inclusive empirical findings, nevertheless, the empirical literatures have generated considerable inconclusive debates. This paper contributes to the empirical literature by examining factors affecting global environmental sustainability by utilizing the nonlinear ARDL bounds test and its application, as detailed in the next section.

**Data sources and model**

Based on the empirical debates, we use time series quarterly data from the global context in our investigation. The sample period is set as 1966Q4 to 2019Q4 because of data availability issues. The analysis of our datasets is line with the approach developed by Shin et al. (2014) and Pesaran et al. (2001). The data consist of urbanization (measures urban population), economic growth (measures per capita GDP in the constant price of 2016 US dollars), energy consumption (measures in kg of oil equivalent per capita), trade openness (measures world trade), and environmental sustainability (measures metric tons per capita CO2 emissions). Per capita CO2 emissions, per capita GDP growth, and urban population as a share of total population were sourced from the World Bank’s World Development Indicators, while the datasets of global trade as a percentage of GDP and energy consumption were obtained from Our World in Data (OWID). The selection of these datasets is the basis of theories and empirical insights (Saboori et al. 2012; Wang and Wang 2021). The dataset’s selected period was constrained by the lack of relevant data in the more recent years, and it gives a total of 216 observations. All the variables were seasonally adjusted, and they were chosen to capture the
particular characteristics within the global context. The statistical software used in the regressions is EViews 10 (Lucini et al. 2020). Environmental sustainability (i.e., measures by per capita CO$_2$ emissions) is used as a dependent variable (Hite and Seitz 2021; Shin et al. 2014). CO$_2$ emission per capita is calculated on the basis of the burning of fossil fuels, including the manufacture of cement, and emissions from coal, gas, and petroleum. Likewise, energy consumption, urbanization, trade openness, and economic growth were carefully selected as explanatory variables based on prior empirical studies (Wang and Wang 2021; Hite and Seitz 2021). Table 2 codes and summarizes the variables employed in the study, while Table 3 depicts descriptive statistics.

Table 3 describes the basic features of the datasets and the basis of quantitative analysis. It measures the central tendency and variability of the outliers. It provides two types of averages: mathematical and positional averages. The kurtosis values show statistical measures of variables distribution, whereas the skewness presents differentiates extreme values in datasets, and the Jarque-Bera statistic measures goodness-of-fit or normality test. On the whole, the result indicates that there are no outliers in the datasets. The mathematical coding of our baseline equation linear functional form of the variables is as follows:

$$Y = f\left( X_{1t}, X_{2t}, X_{3t}, X_{4t} \right)$$

The essence is to determine whether all regressors are decomposed into their positive (POS) and negative (NEG) shocks. The NARDL bounds test–decomposed functional form is as follows:

$$Y = f\left( X^+_{1t}, X^+_{2t}, X^+_{3t}, X^+_{4t}, X^-_{1t}, X^-_{2t}, X^-_{3t}, X^-_{4t} \right)$$

where $X^+_t$ and $X^-_t$ denote the decomposed variables as well as the partial sums of positive (+) and negative (−) changes in $X_t$. Hence, the linear functional analysis on EViews software is shown as follows:

$$LCO_2P = f(LGDPO, LGDPN, LEC, LECO, LTOPO, LTOPN, LURPO, LURPN)$$

where $LCO_2P$ denotes per capita CO$_2$ emissions and POS indicates positive shock, while NEG presents negative shock (Shin et al. 2014).

### Methodology

#### Pre-estimation data checks — stationarity tests

The study’s empirical analysis entails establishing the stationarity of variables, which helps us avoid spurious regression. Stationarity implies that the variable’s mean zero has a constant variance (Dickey and Fuller 1979). The variable might become stationarity at levels or first difference, etc. This study employs the stationarity test of Zivot-Andrew (1992), Dickey and Fuller (1979), unit root test with one endogenous structural break of Perron (1990), and Phillips and Perron (1988). These unit root tests were observed to confirm that none of the variables utilized in these estimations is I(2) or
higher. The mathematical application for these unit root tests is structured in the following equations:
\[
\Delta x_t = \phi x_{t-1} + \varepsilon_t
\]
where \(\Delta x\) presents changes in the time series variable and \(t\) denotes notation of time, while \(\phi\) is slope coefficient, and \(\varepsilon\) presents distributed error term.

The efficiency of Dickey-Fuller regression is accomplished by augmenting the following mathematical application:
\[
\Delta x_t = \alpha_0 + \gamma t + (\phi - 1)x_{t-1} + \sum_{i=1}^{n} \beta_i \Delta x_{t-p+i} + \varepsilon_t
\]
where \((\phi - 1)\) is a null hypothesis, \(\alpha_0\) would be said to follow a random walk, and \(\Delta\) stands (lagged differences) for the first difference operator. The operator either follows Schwarz Information Criteria, Akaike Information Criteria, etc.

Previous studies have often neglected structural breaks in time series analysis (Perron 1994), which could cause unit biased or false null hypotheses. The mathematical application of the additive outlier (AO) and the innovational outlier (IO\(_1\) and IO\(_2\)) models are shown below:
\[
y_t = \mu + \delta D(T_b) + \beta X_{t-1} + \sum_{i=1}^{k} \alpha_i \Delta x_{t-1} + \varepsilon_t
\]
where \(\delta D(T_b) = 1(t > T_b)(t - T_b)\), and \(\tilde{y}_t = \alpha \tilde{y}_{t-1} + \sum_{i=1}^{k} \alpha_i \Delta \tilde{y}_{t-1} + \varepsilon_t\); the AO model represents gradual change, and AO model represents rapid change. All models are reported with asymptotic critical values (Perron 1994).

**Nonlinear ARDL bounds test of cointegration**

While the traditional ARDL bounds test enables the evaluation of the long-run and short-run interactions of the variables, in this paper, we followed NARDL bounds test of cointegration proposed by Pesaran et al. (1999) and later validated by Pesaran et al. (2001). This approach responds that the dependent variable, \(Y_t\), increases (+) and decreases (−) of each independent variable of \(X_t\) (Shin et al. 2014). The fundamental application the time series model is specified:
\[
Y_t = \beta_0 + \beta_1 X_t + \mu_t
\]
where \(Y_t\) is the target variable, \(X_t\) represents the regressors, and \(\beta_1\) shows the change in \(Y\) per unit change in \(X\). It also captures the direction and magnitude \(Y_t\) reaction to changes in \(X\). In keeping with the two-step framework Engle and Granger (1987) developed, which accounts for the potential structural shift, we reformulate partial sums of \(X_t^+\) or partial sums of \(X_t^-\). Recall that the mathematical application of the long-run asymmetric model is written:
\[
Y_t = \beta_0 + \beta_1^+ X_t^+ + \beta_1^- X_t^- + \mu_t
\]

To begin, we specify the asymmetric regression model:
\[
\Delta y_t = \beta_0 + \sum_{i=1}^{N} \lambda_i \Delta y_{t-i} + \sum_{i=0}^{q} \delta_i \Delta y_{t-i}^+ + \sum_{i=0}^{q} \delta_i^- \Delta y_{t-i}^-
\]
\[
+ \sum_{i=0}^{q} \delta_i \Delta y_{t-i}^+ + \sum_{i=0}^{q} \delta_i^- \Delta y_{t-i}^- + \rho Y_{t-1} + \varphi_1 X_{t-1}^+ + \varphi_2 X_{t-1}^-
\]
where long-run asymmetric effects of \(X_t\) on \(Y\) is calculated as \(L_{M1+} = -\varphi_1^+\) and \(L_{M1-} = -\varphi_1^-\), short-run asymmetric effects of \(X_t\) on \(Y\). If symmetry is rejected, we conclude the impact of \(X\) on \(Y\) is asymmetric. Following the methodology, as developed by Shin et al. (2014), the regression model for this analysis is as follows:
\[
\Delta LCO_{t-1} + \Delta LTOP_{t-1} + \Delta LURP_{t-1} + \Delta EC_{t-1} + \Delta LGDP_{t-1} + \Delta CO_{t-1} = \beta_0 + \sum_{i=0}^{q} \alpha_1 \Delta LCO_{t-1}^+ + \sum_{i=0}^{q} \alpha_2 \Delta LTOP_{t-1}^+ + \sum_{i=0}^{q} \alpha_3 \Delta LURP_{t-1}^+ + \sum_{i=0}^{q} \alpha_4 \Delta LGDP_{t-1}^+ + \sum_{i=0}^{q} \alpha_5 \Delta CO_{t-1}^+
\]
\[
+ \sum_{i=0}^{q} \alpha_1 \Delta LCO_{t-1}^- + \sum_{i=0}^{q} \alpha_2 \Delta LTOP_{t-1}^- + \sum_{i=0}^{q} \alpha_3 \Delta LURP_{t-1}^- + \sum_{i=0}^{q} \alpha_4 \Delta LGDP_{t-1}^- + \sum_{i=0}^{q} \alpha_5 \Delta CO_{t-1}^-
\]
\[
+ \sum_{i=0}^{q} \alpha_1 \Delta LCO_{t-1}^+ + \sum_{i=0}^{q} \alpha_2 \Delta LTOP_{t-1}^+ + \sum_{i=0}^{q} \alpha_3 \Delta LURP_{t-1}^+ + \sum_{i=0}^{q} \alpha_4 \Delta LGDP_{t-1}^+ + \sum_{i=0}^{q} \alpha_5 \Delta CO_{t-1}^+
\]
\[
+ \sum_{i=0}^{q} \alpha_1 \Delta LCO_{t-1}^- + \sum_{i=0}^{q} \alpha_2 \Delta LTOP_{t-1}^- + \sum_{i=0}^{q} \alpha_3 \Delta LURP_{t-1}^- + \sum_{i=0}^{q} \alpha_4 \Delta LGDP_{t-1}^- + \sum_{i=0}^{q} \alpha_5 \Delta CO_{t-1}^-
\]
\[
\Delta Y_t = a_0 + \sum_{i=1}^{q} a_1 \Delta Y_{t-1} + \sum_{i=1}^{q} a_2 \Delta X_{t-1}^+ + \sum_{i=1}^{q} a_3 \Delta X_{t-1}^- + \sum_{i=1}^{q} a_4 \Delta Y_{t-1} + \varphi_1^+ X_{t-1}^+ + \varphi_2^- X_{t-1}^-
\]
where \(\Delta Y_t = \alpha_0 + \sum_{i=1}^{q} \alpha_1 \Delta Y_{t-1} + \sum_{i=1}^{q} \alpha_2 \Delta X_{t-1}^+ + \sum_{i=1}^{q} \alpha_3 \Delta X_{t-1}^-\). Finally, denote short-run estimation and \(\rho Y_{t-1} + \varphi_1 X_{t-1}^+ + \varphi_2 X_{t-1}^-\) indicates long-run estimation. Also, \(X_t^+\) and \(X_t^-\) are the partial sums of positive (+) and negative (−) changes in \(X_t\). Like the ARDL bounds test, the NARDL bounds test determines the long-run relationship between the regressand or \(X_t^+\) and \(X_t^-\) and regressors or \(Y\) (Pesaran et al. 2001).

For simplicity, we employed \(t\)-test of Banerjee et al. (1998) hypothesis and computed \(F\)-statistic and then compared \(F\)-statistic with the Pesaran et al. (2001) asymptotic critical value. If we reject the null hypothesis (no cointegration), meaning that variables are cointegrated in the presence of asymmetry, then the null hypothesis of no cointegration is formulated as follows:
\[
H_0 : \varphi = 0 \quad H_A : \varphi < 0
\]

**NARDL long-run asymmetric coefficients**

To capture NARDL long-run asymmetric coefficients, we divided the negative coefficient of \(X_t^-\) (i.e., \(\varphi^\)\(^-\)) by the coefficient of \(Y_{t-1}\) (i.e., \(\rho\)):
Furthermore, we divided the negative of the coefficient of $X_t^-$ (i.e., $\phi^-$) by the coefficient of $Y_{t-1}$ (i.e., $\rho$):

$$\frac{-\phi^-}{\rho}$$

Test for long-run asymmetry: Wald test

Next, we adopt the general-to-specific procedure to arrive at the final specification of the NARDL model by trimming insignificant lags. Finally, we perform the Wald test of the null hypothesis. If we reject $H_0$, we have both short-run and long-run asymmetry relationships. In other words, the magnitude of the change in $Y$ when $X$ increases are NOT THE SAME as when $X$ decreases:

$$H_0 : \frac{-\phi^+}{\rho} = \frac{-\phi^-}{\rho}$$

$$H_A : \frac{-\phi^+}{\rho} \neq \frac{-\phi^-}{\rho}$$

Asymmetric dynamic multiplier graph

To account for the potential divergence or structural shift from the short run to the long run, we employed the asymmetric dynamic multipliers graph, where $Y_t$ is adjusted to its equilibrium levels following a negative or positive shock in $X_t$ as illustrated in the following:

$$m^+_h = \sum_{k=0}^\infty \frac{\delta Y_{t+k}}{\delta X_t} m^-_k = \sum_{k=0}^\infty \frac{\delta Y_{t+k}}{\delta X_t} \text{ for } h = 0, 1, 2, \ldots$$

Where $h \rightarrow \infty$, then $m^+_t \rightarrow \frac{-\phi^+}{\rho}$ and $m^-_t \rightarrow \frac{-\phi^-}{\rho}$

Residual diagnostic tests

The result of long-run or short-run cointegrating relationship does not automatically infer that models employed in the analysis are stable over the sample period. Hence, we employed residual diagnostics tests of Breusch-Pagan-Godfrey heteroscedasticity test, Breusch-Godfrey serial correlation LM test of $Q$-statistics, and Ramsey RESET test, respectively.

Robustness checks

Taking cognizance of the robustness checks, the present study follows the example of Tang and Bethencourt (2017) to adopt FMOLS, DOLS, and CCR as the robustness check estimators. Although there are other robustness check approaches in the literature, our select estimators are relatively the more stable for addressing second-order bias, endogeneity and serial correlation issues in cointegrating regressions.

Empirical results and discussions

The empirical analysis of the study entails following the standard practice of econometric literature. Our first preliminary estimation step was to utilize descriptive statistics and proceed with four well-known unit root stationarity tests, i.e., conventional ADF unit root test, Phillip-Peron (PP) unit root test, unit root test with one endogenous structural break, and Zivot-Andrews unit root test. The descriptive analysis result indicated no outliers in the datasets (see Table 3), while the unit root tests confirm that none of the variables used in the estimations is at I(2) or higher. To put it differently, all estimated variables, including economic growth, energy consumption, and trade openness, were not integrated at level but only became stationarity at a first difference (i.e., I(1)), except for urbanization, which integrated the order zero (i.e., I(0)) with intercept and trend. Tables 4, 5, 6, and 7 present the summary of unit roots tests observed. In a cursory look at the unit root test with mixed order of I(0) and I(1) regressors, the traditional cointegration tests of Johansen and Juselius (1990) were not applicable for

| Table 4 Summary of conventional adf unit root test result |
|-----------------------------------------------------------|
| **Global data** | **Level** | **First difference** |
| **(1966Q1–2019Q4)** | **Intercept** | **Intercept and trend** | **Intercept** | **Intercept and trend** |
| $LCO_{Pt}$ | 2.453 | 2.371 | 4.941 | 4.985 |
| $LEC_t$ | 3.142 | 3.468 | 4.747*** | 4.752*** |
| $LGDP$ | 4.130* | 4.101* | 4.752*** | 4.752*** |
| $LTOP$ | 1.491 | 1.444 | 3.971** | 4.138** |
| $LURP$ | 0.193 | 2.669 | 3.645*** | 3.785** |

SBIC was selected as lag length criterion; variable was tested first at the level and then at first difference with intercept, and intercept and trend, respectively

*, **, and *** represent 10%, 5%, and 1% significance levels, respectively

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estimations. Instead, the traditional ARDL bounds test and the newly developed NARDL bounds test fulfilled the necessary conditions for cointegration tests. Table 8 presents the results of the traditional ARDL bounds test of cointegration. The outcome of computed $F$-statistics of 7.20% validates evidence of a cointegrating relationship among the variables (i.e., $F$-statistics exceeds the upper bounds of critical values) as shown in Table 8.

It is worthy to note that the traditional ARDL bounds test of cointegration does not account for asymmetric relationships. Hence, we proceeded with the NARDL unrestricted error correction model of cointegration. Table 9 presents the summary results of the NARDL bounds test for cointegration with long-run and short-run diagnostic tests. It comprises four different analyses of long-run, short-run, error correction regression, and cointegrating statistics. The findings indicate that the computed NARDL $F$-statistics of 4.46 > Pesaran upper bound critical value of 3.06 at $\alpha = 10\%$, verifying cointegration’s existence. Furthermore, the error correction term is statistically significant at 1% level, and it has the expected sign and value of -0.031836. The Durbin-Watson statistic of 2.06695, $R^2$-squared value of 0.99359, and adj. $R^2$-squared value of 0.992920 measured that the explanatory variables included in the model have the significant joint powers to explain the dependent variable.

Concerning energy consumption ($LEC$), the result indicates that both POS and NEG shocks in $LEC$ have a POS causal effect on $CO_2P$ emissions in the long run. Similarly, about $LTOP$ and $LURP$, both POS and NEG shocks in $LTOP$ and $LURP$ also have a POS causal effect on $CO_2P$ emissions in the long run. On the contrary, for $GDP$,
Table 8 Summary result of traditional ARDL bounds test of cointegration

| Variables     | ARDL Long Run Bound Test Results |
|---------------|---------------------------------|
| LCO2P(-1)*    | Coefficient | Std. Error | t-Statistic | Prob. |
| LEC(-1)       | 0.013       | 0.002      | 5.522       | 0.000 |
| LGDP**        | -0.006      | 0.003      | -2.199      | 0.029 |
| LTOP(-1)      | 0.001       | 0.001      | 1.424       | 0.156 |
| LURP**        | -0.006      | 0.001      | -0.173      | 0.863 |

ARDL Short-Run Results

| Variables     | Coefficient | Std. Error | t-Statistic | Prob. |
|---------------|-------------|------------|-------------|-------|
| ∆LCO2P(-1)    | 0.215       | 0.069      | 3.141       | 0.002 |
| ∆LEC          | 0.026       | 0.004      | 0.748       | 0.460 |
| ∆LGDP**       | -0.006      | 0.003      | -1.662      | 0.098 |
| ∆LTOP         | 0.012       | 0.003      | 3.770       | 0.000 |
| ∆LTOP(-1)     | -0.007      | 0.003      | -2.159      | 0.032 |

Cointegrating Statistics

| Statistic     | Coefficient | Std. Error | t-Statistic | Prob. |
|---------------|-------------|------------|-------------|-------|
| R-squared     | 0.993       |            |             |       |
| Adj. R-squared| 0.993       |            |             |       |
| Durbin-Watson stat | 2.035 |            |             |       |
| AR1 Lag Length | (2, 2, 0, 2, 0) | (2, 2, 0, 2, 0) |             |       |
| CointEq(-1)*  | -0.010      | 0.002      | -6.897      | 0.000 |
| E-statistic   | 7.205       |            |             |       |
| Asymptotic- n=1000, FPSS | 3.52* |            |             |       |

LCO2P natural log of carbon emissions per capita, LEC natural log of energy consumption, LGDP natural log of economic growth, LTOP natural log of trade openness, LURP natural log of urban population

* *, **, and *** illustrate 10%, 5%, and 1% significant levels, respectively.

Both POS and NEG shock in LGDP have POS and NGE causal effects on CO2P emissions in the long run. The short-run causal effects were also investigated critically. About LEC’s short-run causal effects, we find out that both POS and NEG shock in LEC in the current period have both POS and NEG’s short-run effects on CO2P emissions. However, the POS and NEG terms of LEC in past periods show insignificant effects on CO2P emissions.

Concerning LEC short-run effects, the analysis shows that POS and NEG shock in LEC have both POS and NEG short-run effects on CO2P emissions in the current period. However, in past periods, both POS and NEG terms of LGDP have no significant causal effects on CO2P emissions. In terms of LTOP short-run causal effects, the result shows that both POS and NEG shock in LTOP in the current period has a significant NEG short-run effect on CO2P emissions; however, a POS shock short-run in LTOP in the current period shows insignificant effects. In terms of the past period, the POS or NEG shocks in LTOP have no significant causal effects on CO2P emissions. Finally, in terms of LURP short-run causal effects, the outcome shows that both POS and NEG shock in LURP in the current period have both POS and NEG insignificant short-run effects on CO2P emissions. Furthermore, both POS and NEG shock in LURP have an insignificant short-run effect on CO2P emissions in the past periods. These findings imply that the estimated variables have short-run and long-run causal effects on environmental sustainability within the global framework, as indicated in Table 9.

Given the fact of the cointegration relationship among the variables, we proceeded to estimate long-run asymmetric changes in the independent variables on the dependent variable. The results of long-run asymmetric coefficient calculated values are presented in Table 10. It captures the direction and magnitude of per capita CO2 emissions (CO2P) on the log value of energy consumption (LEC), the log value of economic growth (LGDP), log value of trade openness (LTOP), and urbanization, respectively. The calculated outcomes of the long-run asymmetric coefficient of changes in the log value of energy consumption (LEC) show that per capita CO2 emissions (CO2P) have the positive function of both positive (POS) and negative (NEG) changes in the log value of energy consumption. Hence, a 1% increase in energy consumption will cause per capita CO2 emissions by 26.87%, if we hold other factors constant. To put it differently, a 1% decrease in energy consumption will cause per capita CO2 emissions to decrease by 53.11%, if all other factors are constant. Regarding the long-run asymmetric coefficient of change in the log value of economic growth (LGDP), the analysis shows that per capita CO2 emissions are a positive and negative function of POS and NEG changes in the log value of economic growth (LGDP). Hence, it suggests a 1% increase in global economic growth will cause 12.83% rise in per capita CO2 emissions, keeping other factors constant. In contrast, a relatively 1% decrease in economic growth could cause 26.42%, decrease in per capita CO2 emissions, keeping other factors constant. The outcome for trade openness is in concordance with a previous study by Rafindadi and Ozturk (2016). The results
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Table 9 Summary NARDL bounds test for cointegration with long-run and short-run diagnostic tests

| Variables | Coefficient | Std. Error | t-Statistic | Prob. |
|-----------|-------------|------------|-------------|-------|
| **Long-Run Results** | | | | |
| Constant | 0.156 | 0.051 | 3.011 | 0.003 |
| LCO2P(-1)* | -0.031 | 0.014 | -2.223 | 0.027 |
| LEC POS(-1) | 0.008 | 0.004 | 1.949 | 0.053 |
| LEC NEG(-1) | 0.016 | 0.003 | 4.544 | 0.000 |
| LGDP POS(-1) | 0.004 | 0.006 | 0.678 | 0.498 |
| LGDP NEG(-1) | -0.008 | 0.004 | -1.965 | 0.051 |
| LTOP POS(-1) | 0.001 | 0.001 | 1.169 | 0.244 |
| LTOP NEG(-1) | 0.000 | 0.001 | 0.265 | 0.791 |
| LURP POS(-1) | 0.004 | 0.035 | 0.117 | 0.906 |
| LURP NEG(-1) | 0.064 | 0.167 | 0.383 | 0.995 |
| **Short-Run Results** | | | | |
| A (LCO2P(1)) | 0.183 | 0.074 | 2.454 | 0.015 |
| A (LCO2P(2)) | 0.052 | 0.075 | 0.690 | 0.490 |
| A (LCO2P(3)) | -0.073 | 0.070 | -1.044 | 0.297 |
| A (LCO2P) | 0.002 | 0.013 | 0.115 | 0.908 |
| A (LCO2P(-1)) | 0.001 | 0.013 | 0.098 | 0.921 |
| A (LCO2P(-2)) | -0.003 | 0.013 | -0.239 | 0.810 |
| A (LCO2P(-3)) | 0.040 | 0.009 | 4.159 | 0.000 |
| A (LCO2P(-4)) | -0.013 | 0.010 | -1.319 | 0.188 |
| A (LCO2P(-5)) | -0.006 | 0.009 | -0.681 | 0.496 |
| A (LCO2P(-6)) | 0.031 | 0.015 | 2.007 | 0.046 |
| A (LCO2P(-7)) | -0.010 | 0.015 | -0.674 | 0.500 |
| A (LCO2P(-8)) | -0.003 | 0.015 | -0.258 | 0.796 |
| A (LCO2P(-9)) | -0.035 | 0.013 | -2.691 | 0.007 |
| A (LCO2P(-10)) | 0.014 | 0.014 | 0.984 | 0.326 |
| A (LCO2P(-11)) | 0.006 | 0.014 | 0.409 | 0.682 |
| A (LCO2P(-12)) | -0.001 | 0.006 | -0.213 | 0.831 |
| A (LCO2P(-13)) | -0.004 | 0.006 | -0.680 | 0.497 |
| A (LCO2P(-14)) | -0.002 | 0.006 | -0.377 | 0.706 |
| A (LCO2P(-15)) | 0.021 | 0.003 | 3.916 | 0.000 |
| A (LCO2P(-16)) | -0.008 | 0.006 | -1.445 | 0.150 |
| A (LCO2P(-17)) | -0.005 | 0.005 | -0.869 | 0.385 |
| A (LCO2P(-18)) | -0.136 | 0.218 | -0.623 | 0.533 |
| A (LCO2P(-19)) | 0.223 | 0.231 | 0.966 | 0.335 |
| A (LCO2P(-20)) | 0.057 | 0.021 | 2.626 | 0.008 |
| A (LCO2P(-21)) | -0.178 | 0.196 | -0.905 | 0.366 |
| A (LCO2P(-22)) | -0.111 | 0.218 | -0.513 | 0.608 |
| A (LCO2P(-23)) | -0.134 | 0.208 | -0.643 | 0.521 |

POS positive, NEG negative; LCO2P natural log of carbon emissions per capita, LEC natural log of energy consumption, LGDP natural log of economic growth, LTOP natural log of trade openness, LURP natural log of urban population.

*., **, and *** represent 10%, 5%, and 1% significance levels, respectively.

Table 10 presents the stepwise regression; it summarizes parsimonious nonlinear error correction results. All insignificant short-run terms were excluded, although all the long-run terms were retained regardless of whether or not they were significant (Akçay 2019; Shin et al. 2014). The long-run and short-run asymmetry coefficients of economic growth, energy consumption, and trade openness were estimated using Wald test (Akçay 2019). The estimated hypothesis for energy consumption, economic growth, trade openness, and urbanization confirms evidence of a

show long-run asymmetric coefficient for per capita CO2 emissions is a positive function of both POS and NEG changes in trade openness. Hence, 1% increase in trade openness causes CO2 emissions per capita to increase by 3.51%. On the other hand, a 1% decrease in trade openness means that CO2 emissions per capita will decrease by 1.22%, keeping other factors constant. Finally, concerning the long-run asymmetric coefficient of urbanization, the finding noted that a per capita CO2 emission is a positive function of both POS and NEG changes in urbanization. Thus, 1% increase in urbanization cause per capita CO2 emissions to increase by 12.72%, and if urbanization decreases by 1%, per capita, CO2 emissions could decrease by 20.26%, as presented in Table 10.

Table 11 presents the stepwise regression; it summarizes parsimonious nonlinear error correction results. All insignificant short-run terms were excluded, although all the long-run terms were retained regardless of whether or not they were significant (Akçay 2019; Shin et al. 2014). The long-run and short-run asymmetry coefficients of economic growth, energy consumption, and trade openness were estimated using Wald test (Akçay 2019). The estimated hypothesis for energy consumption, economic growth, trade openness, and urbanization confirms evidence of a
long-run asymmetric relationship. However, all variables show no evidence of a short-run asymmetry relationship. These findings are contrary to Shin et al. (2014) paper but in line with the research by Akçay (2019). In terms of the diagnostic statistics, Breusch-Pagan-Godfrey heteroscedasticity test, Breusch-Godfrey serial correlation LM test, $Q$-statistics, and Ramsey RESET test were employed. The Breusch-Pagan-Godfrey heteroscedasticity test indicates that the model is homoscedastic, and no serial correlation was found. The Ramsey RESET Test shows that the estimated models are stable. The observed Durbin-Watson Statistic of 2.214947 validates that the estimated variables have no first-order autocorrelation in the models (Masood et al. 2015). The $R^2$ and adj. $R^2$-squared values of 0.996145 and 0.995487 imply that the variables included in the model correctly explain variations in global environmental sustainability. Table 11 presents the results of the Wald test.

### Table 10: Asymmetric long-run coefficients

| Description of asymmetric Notations | Calculated Value | Coefficients | % of coefficient |
|-------------------------------------|-----------------|--------------|-----------------|
| Long-run coefficient for LEC_POS    | $\varphi_1^P$   | 0.0088553    | 26.87%          |
| Long-run coefficient for LEC_NGE    | $\varphi_1^N$   | -0.016909    | 53.11%          |
| Long-run coefficient for LGDP_POS   | $\varphi_2^P$   | -0.004083    | 12.83%          |
| Long-run coefficient for LGDP_NGE   | $\varphi_2^N$   | -0.0088410   | -26.42%         |
| Long-run coefficient for LTOP_POS   | $\varphi_3^P$   | -0.001119    | 3.51%           |
| Long-run coefficient for LTOP_NGE   | $\varphi_3^N$   | -0.003389    | 1.22%           |
| Long-run coefficient for LURP_POS   | $\varphi_4^P$   | -0.004049    | 12.72%          |
| Long-run coefficient for LURP_NGE   | $\varphi_4^N$   | -0.006449    | 20.25%          |

POS positive, NEG negative, $LCO_2P$ natural log of carbon emissions per capita, $LEC$ natural log of energy consumption, $LGDP$ natural log of economic growth, $LTOP$ natural log of trade openness, $LURP$ natural log of urban population

*, **, and *** illustrate 10%, 5%, and 1% significant levels, respectively

### Table 11: Wald test for long-run and short-run asymmetric and residual diagnostics

| Long-run Coefficient | Long-run effect [+] | Long-run effect [-] |
|----------------------|---------------------|---------------------|
| Exogenous Variables  | Coefficient | t-Stat | P-value | Coefficient | t-Stat | P-value |
| LCO2P                | -0.030     | -1.905 | 0.058   | -----       | -----  | -----   |
| LEC                  | 0.006      | 4.293  | 0.000   | 0.007      | 5.656  | 0.000   |
| LGDP                 | 2.43E      | 1.019  | 0.309   | -4.99E     | -1.483 | 0.139   |
| LTOP                 | 0.004      | 2.347  | 0.019   | -----      | -----  | -----   |
| LURP                 | -0.009     | -1.290 | 0.198   | -----      | -----  | -----   |

Asymmetry Tests:

Test Long-Run Asymmetry

| Variables | WLR F-Stat | P-value | WLR F-Stat | P-value |
|-----------|------------|---------|------------|---------|
| LEC       | 3.588      | 0.059   | 3.198      | 0.075   |
| LGDP      | 0.331      | 0.565   | 16.279     | 0.000   |
| LTOP      | 4.020      | 0.046   | 31.790     | 0.000   |
| LURP      | 4.020      | 0.046   | 31.573     | 0.000   |

Model Diagnostic Test

| Diagnostic Test | Stat | P-value |
|-----------------|------|---------|
| Heteroscedasticity Test: Breusch-Pagan-Godfrey | 1.194 | 0.246 |
| Jarque-Bera Test on Normality (Chi$^2$) value | 59.540 | 0.000 |
| Ramsey RESET Test (F) | 1.446 | 0.230 |
| R-squared       | 0.996 |       |
| Adjusted R-squared | 0.995 |       |
| Durbin-Watson stat | 2.214 |       |

WLR Wald test, $LCO_2P$ natural log of carbon emissions per capita, $LEC$ natural log of energy consumption, $LGDP$ natural log of economic growth, $LTOP$ natural log of trade openness, $LURP$ natural log of urban population

*, **, and *** illustrate 10%, 5%, and 1% significant levels, respectively
Furthermore, another innovation of this study was that we employed the dynamic multiplies to access the dynamicity of our estimated models. Figures 1, 2, 3, 4 and 5 illustrate dynamic multiplies graphs/plots. The result shows evidence of positive and negative dynamic effects among estimated variables. It shows a divergence from the short run to the long run (Brown et al. 1975). The estimated CUSUM value lies at a 5% significant level.
between the two-bonded lines; thus, it confirms the stability of selected models covering from 1966Q1 to 2019Q4.

In terms of robustness checks, Table 12 presents summary results of DOLS, FMOLS, and CCR regression models. The outputs of DOLS and FMOLS models were the same a CCR model in terms of the degree of significance. The log value of energy consumption, trade openness and urbanization indicate a positive coefficient and significant relationship across DOLS, FMOLS, and CCR models, except for the log value of economic growth which has a negative and insignificant relationship. The results obtained from DOLS, FMOLS, and CCR regressions robustly stimulate CO2 emission and align with economics theory as outlined by Genovese et al. (2017). This intuitively implies that the upward trend of energy consumption, economic growth, trade openness, and urbanization does adversely affect CO2 emissions.

| Table 12 Summary of comparison of robustness checks |
|----------------------------------------------------|
| **Regressors** | **DOLS** | **FMOLS** | **CCR** |
| **LEC** | 0.033 | 0.036 | 0.032 |
| | (0.432) | (0.462) | (0.423) |
| **LGDP** | -0.037 | -0.042 | -0.041 |
| | (0.610) | (0.445) | (0.439) |
| **LTOP** | 0.027*** | 0.027*** | 0.027*** |
| | (0.000) | (0.000) | (0.000) |
| **LURP** | 0.503* | 0.489* | 0.492* |
| | (0.092) | (0.067) | (0.070) |
| **Lead** | 1 | | |
| **Lag** | 1 | | |

This table presents the coefficients and associated p-values on the effect of CO2 emissions on energy consumption, economic growth, trade openness, and urbanization; the respective P-values are shown in parentheses.

***, **, and * denote the significance of correlation coefficients at the 1%, 5%, and 10% significant levels, respectively.
Conclusion and policy implication

To sum up, economic activity will always use a certain amount of resources; thus, higher demands for resources have gradually affected the environment. Literally, hundreds of regional and bilateral agreements and more than 10 global environmental treaties have been negotiated and signed to ensure environmental quality (Hunter 2018). However, despite general progress made in ensuring environmental sustainability, the main factors affecting the natural environment have emerged as one of the significant challenges in the economics literature (Wang et al. 2022). Global cooperation is necessary for achieving environmental sustainability (CO2 emissions). Though studies have argued that the lack of preserving natural resources is the most critical factor for environmental sustainability, numerous academic dilemma remains in economics literature due to the absence of robust clarity on whether other factors such as economic growth, urbanization, trade openness, and energy consumption matter for global environmental sustainability. Since these factors and environmental sustainability are integrated of order zero and one, this paper utilized the unrestricted NARDL bounds test technique to model their relationship.

The main contribution of this paper is to add knowledge to the economics literature. Furthermore, following examples from Shin et al. (2014) and Ahmed et al. (2020), this study adopts residual diagnostics tests of Breusch-Pagan-Godfrey heteroscedasticity test, Breusch-Godfrey serial correlation LM test of $Q$-statistics, and Ramsey RESET to access the stability of the estimated models. Taking cognizance of the robustness checks, DOLS, FMOLS, and CCR estimators were observed. These models addressed endogeneity, second-order bias, and serial correlation issues, respectively. Up-to-date of this study, no study has examined the nexus of these variables within the global framework.

Our empirical findings reveal the following. (i) NARDL bounds tests of cointegration confirm evidence of long-run and short-run relationships among the variables. The NARDL bounds test of cointegration further confirmed this evidence explained earlier by Kobbi and Gabsi (2017), supporting Daly’s theory (Daly 1973). (ii) Long-run asymmetric relationship was confirmed among the variables; however, no short-run asymmetric effect was confirmed among the variables. These findings were contrary to empirical studies by Shin et al. (2014) and Akçay (2019). (iii) DOLS, FMOLS, and CCR models demonstrate that economic growth, energy consumption, and trade openness are positively related to environmental sustainability except for economic growth which shows the negative and insignificant relationship. Finally, these findings validate the protracted argument in the literature that these estimated factors are significant for global environmental sustainability. This study is beneficial for global policy actions, and perhaps, it can be used for decisive policy engagements for achieving global environmental quality. In addition, policymakers can prioritize growth-enhancing global reforms that increase awareness of reducing climate crises.

The main limitations of the present study are twofold: Firstly, it included only four global environmental factors, which should have been extended to a wider range of environmental factors to examine environmental sustainability. Undoubtedly, it will be interesting to give more credible conclusions through a multi-national variables comparison. It may be possible to realize some explorations based on sectorial or regional data too. Secondly, only the NARDL bounds test for cointegration tool is implemented, while it is recommended to enlarge the econometric framework by applying, for example, various multivariate cointegration approaches that help to understand the nexus of global environmental factors as previously mentioned. In addition to the NARDL bounds test for cointegration method, several discrete cointegration tools can also be applied. Currently, more plausible ways are considered. Researchers could improve this study for future research direction by incorporating other factors such as livestock productivity index, renewable energy, and globalization by using multivariate wavelet coherence approaches.

Author contribution Dervis Kirikkaleli and James Karmoh Sowah Jr. conducted the investigation and gathered the data. James Karmoh Sowah Jr. wrote the introduction and the literature review, while Dervis Kirikkaleli prepared the methodology and the empirical findings as indicated in this paper. In addition, James Karmoh Sowah Jr. assisted in the explanation of the results. Finally, as the corresponding author, I confirm that the final version of this paper was reviewed and endorsed by all authors.

Data availability The data that support the results of this research are accessible from the World Bank and OWID.

Declarations

Ethical approval We declare that this paper is original, has not been published before, and is not currently being considered for publication by another journal. Therefore, this research does not require ethical authorization or informed consent.

Consent to participate Not applicable

Consent for publication Not applicable

Competing interests The authors declare no competing interests.
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