economies

MDPI

*Article*

# Linear and Nonlinear Relationships Between Energy Consumption, Renewable Energy, Greenhouse Gas Emissions, and Economic Growth in Saudi Arabia: A Combined ARDL and Threshold Approach

**Besma Hamdi [1,*], Afifa Ferhi [2], Awatef Louhichi [3] and Olfa Gammoudi [1]**

[1] Management Information Systems Department, Applied College, University of Ha'il, Hail City P.O. Box 2440, Saudi Arabia; o.gammoudi@uoh.edu.sa
[2] Department of Quantitative Methods, Faculty of Economics and Management, University of Sfax, Sfax City P.O. Box 3018, Tunisia; ferhi_afifa1983@yahoo.fr
[3] Department of Economics and Finance, College of Business Administration, University of Ha'il, Hail City P.O. Box 2440, Saudi Arabia; a.louhichi@uoh.edu.sa
* Correspondence: hamdibes2014@gmail.com

**Abstract**

The shift toward renewable energy is a priority for many countries seeking to mitigate climate change, support sustainable economic growth, and enhance energy security. This study examines the relationships between fossil fuel energy consumption, renewable energy, $CO_2$ emissions, $CH_4$ emissions, and economic growth in Saudi Arabia over 1990–2022, using an Autoregressive Distributed Lag (ARDL) model and a Threshold Regression (TR) model. The bounds test indicates a long-run relationship among the variables, while the error-correction term suggests that approximately 52.7% of short-run deviations from long-run equilibrium are corrected within one year. Renewable energy is positively and significantly associated with economic growth in both the short and long run. However, the threshold analysis provides insufficient statistical evidence of a nonlinear relationship: the estimated split at 0.0425 yields a wild-bootstrap *p*-value of 0.943, while the levels specification yields a *p*-value of 0.483. Accordingly, the estimated split is not established as a statistically significant threshold. $CO_2$ emissions are positively associated with economic growth in both renewable-energy regimes, whereas $CH_4$ emissions show a positive and significant long-run association but negative short-run dynamics, with no statistically significant regime-specific effects. Overall, the findings support a positive association between renewable energy and economic growth in the linear model but do not establish a significant threshold effect. These results inform the discussion of energy-transition policies in Saudi Arabia.

**Keywords:** renewable energy; fossil fuel energy consumption; greenhouse gas emissions; economic growth; ARDL bounds testing; threshold regression; Saudi Arabia

Check for updates

Academic Editor: George E. Halkos

Received: 2 October 2026
Revised: 18 October 2026
Accepted: 26 October 2026
Published: 3 November 2026

**Copyright:** © 2026 by the authors. Licensee MDPI, Basel, Switzerland. This article is an open access article distributed under the terms and conditions of the Creative Commons Attribution (CC BY) license.

## 1. Introduction

Over the past few decades, the relationship between energy consumption and economic growth has garnered significant attention in the economic and environmental literature (Kraft & Kraft, 1978; Waheed et al., 2019; González-Álvarez & Montañés, 2023; Zou & Chau, 2023). Energy consumption is a key component of economic development, particularly in resource-rich and transition economies, where it closely tracks industrialization and

economic expansion. However, this dynamic is accompanied by increasing environmental externalities, including greenhouse gas emissions such as carbon dioxide ($CO_2$) and methane ($CH_4$), which contribute to climate change and broader environmental degradation (IPCC, 2022). In this context, several empirical studies have highlighted a trade-off between economic growth and environmental quality, often analyzed through the Environmental Kuznets Curve (EKC) hypothesis (Grossman & Krueger, 1995), which posits that environmental degradation increases during the early stages of development, then declines as income rises. In this context, the transition to renewable energy has emerged as an important pathway for reconciling economic growth with environmental sustainability (C. Chen et al., 2020; Anser et al., 2021; Addis & Cheng, 2023). Compared with conventional fossil fuels, renewable energy generally has lower greenhouse gas emissions and may promote long-term economic growth through technological advancement and improved energy efficiency (IPCC, 2022; IEA, 2023). Although capital accumulation, labour, and technological progress are central to conventional explanations of economic growth, energy consumption is also an important dimension to consider, particularly in energy-intensive economies such as Saudi Arabia.

However, the effects of renewable energy and energy consumption on economic growth remain ambiguous and potentially nonlinear (Wang & Wang, 2020; Ivanovski et al., 2021; Baz et al., 2021). Several empirical studies highlight that the energy–growth relationship is highly dependent on the level of economic development, human capital, technological capabilities, and institutional quality in each country (Destek & Sinha, 2020). Accordingly, some empirical studies have examined whether the energy–growth relationship is asymmetric or varies across threshold regimes, although the findings differ across countries and model specifications. Furthermore, most empirical studies have focused on carbon dioxide ($CO_2$), paying less attention to methane ($CH_4$), despite its importance as a greenhouse gas (IPCC, 2022). This study considers methane emissions from industrial processes alongside $CO_2$ emissions to examine their relationships with economic growth and energy-related variables. However, the selected $CH_4$ indicator captures industrial-process emissions rather than total methane emissions or emissions specifically from the energy sector. Recent empirical studies highlight the use of dynamic and non-linear econometric models to examine the multidimensional interactions among economic growth, energy consumption, and environmental degradation, thereby offering deeper insights into sustainability issues (Youssef et al., 2016; Alqaralleh, 2021; Alshehry & Belloumi, 2024).

Although research on the energy–growth–environment relationship is expanding, empirical findings remain inconsistent in three respects. First, the relationships involving methane emissions receive comparatively less attention than those involving $CO_2$ emissions. Second, many studies rely on linear specifications that may not capture nonlinear or regime-dependent relationships. Third, evidence remains limited for oil-dependent economies undergoing structural transformation, particularly Saudi Arabia under Vision 2030.

Empirical evidence on the relationship between renewable energy and economic growth remains mixed. While some studies report positive contributions through efficiency gains and technological advances, others suggest that the relationship varies with economic development, energy structures, and institutional conditions. These differences motivate further country-specific investigation without presuming that a nonlinear relationship necessarily exists.

In this context, Saudi Arabia offers an instructive case study. As one of the world's major oil producers and currently implementing structural reforms through Vision 2030, the Kingdom is balancing economic diversification with ambitious climate commitments, as illustrated by the Saudi Green Initiative, which aims to reduce carbon emissions and accelerate the expansion of renewable energy.

Against this background, this study asks whether the effect of renewable energy on economic growth in Saudi Arabia is linear or varies according to the level of renewable energy development. To address this question, the study examines the relationships between economic growth, energy consumption, renewable energy, $CO_2$ emissions, and $CH_4$ emissions within a unified empirical framework. First, the Autoregressive Distributed Lag (ARDL) approach is employed to examine the short- and long-run relationships among the variables within a linear framework. Second, threshold regression is employed to test whether the effect of renewable energy on economic growth changes across different regimes defined by the level of renewable energy.

Our study contributes to the literature on environmental and energy economics in several ways, building on several previous studies (Apergis & Payne, 2010; Shahbaz et al., 2013; Destek & Sinha, 2020; Mujtaba et al., 2022; G. Liu et al., 2023). For oil-dependent economies such as Saudi Arabia, the energy transition raises important questions about the role of renewable energy in supporting economic growth while addressing environmental concerns. This study examines whether the relationship between renewable energy consumption and economic growth in Saudi Arabia varies across regimes defined by renewable energy consumption. By selecting renewable energy consumption as the threshold variable, the analysis tests this possibility empirically without presuming that a regime-dependent relationship exists.

From an econometric perspective, the study combines the ARDL approach with threshold regression to examine the energy–growth relationship from complementary perspectives. The ARDL model captures short- and long-run relationships among the variables, while threshold regression tests whether the relationship between renewable energy consumption and economic growth varies across regimes (Bekhet et al., 2017; Waheed et al., 2019; Addo et al., 2021).

The remainder of the paper is organized as follows: Section 2 provides a review of the relevant literature, Section 3 outlines the methodological framework and the dataset used in the analysis, Section 4 reports and discusses the empirical results, Section 5 discusses the results, and Section 6 concludes and sets out the policy implications and strategic recommendations for Saudi Arabia's economic development and energy transition agenda.

## 2. Literature Review

Meadows et al. (2004) emphasize the important role of energy, particularly coal and oil, in the development of modern industrial economies. This role has made energy an important issue in economic thought, although it was not until the 1970s that energy-related issues became more prominent in the discipline. This shift can be explained by two main contextual factors: growing environmental concerns and the increasing depletion of fossil fuel resources, which emerged as critical global issues. One of the most influential contributions reflecting this shift was the Club of Rome's 1972 report, *The Limits to Growth*. Based on system dynamics simulations, the report highlighted the inherent limits to sustained economic expansion, emphasizing that continued growth could become unsustainable in the face of finite natural resources and escalating environmental degradation. In the 1970s, contributions by Solow (1974) and Stiglitz (1974) advanced the theoretical discussion of economic growth, exhaustible natural resources, and the implications of resource constraints for long-term sustainability. These seminal works renewed interest in neoclassical natural resource economics during the 1970s by formally integrating resource constraints into growth theory. The oil crises of 1973 and 1979 further intensified academic and policy interest in the economic implications of energy dependence. Nevertheless, growing awareness of the economic and environmental risks of heavy reliance on fossil fuels—such as energy market volatility, greenhouse gas emissions, and ecological degradation—has

reinforced renewable energy's role as a pivotal pathway to sustainable development (Ozcan & Ozturk, 2019). Given its ability to supply energy with zero or minimal greenhouse gas and air pollutant emissions, renewable energy has emerged as a promising substitute for fossil fuels, attracting growing interest from academics, practitioners, and policymakers alike (Amri, 2018; Belaïd & Zrelli, 2019; Dong et al., 2021). However, the potential contribution of renewable energy to economic growth cannot be assumed to be uniform across countries or levels of renewable energy use. This makes the relationship between renewable energy consumption and economic growth an important empirical question, particularly in economies that remain heavily dependent on fossil fuels.

### *2.1. The Link Between Renewable Energy and Economic Growth*

The empirical literature on the relationship between renewable energy consumption and economic growth remains inconclusive, with findings varying across countries, periods, and empirical specifications.

A first group of studies highlights a positive relationship between renewable energy and economic growth. Xuan (2025b) examined the relationship between renewable energy consumption and economic growth in Germany over the period 1990–2023 using the autoregressive distributed lag (ARDL) model. The results indicate a positive and statistically significant long-run relationship between renewable energy consumption and economic growth. Using the System Generalized Method of Moments, Wen et al. (2025) examined the relationship between green finance, renewable energy investment, economic growth, and environmental sustainability across 22 European Union member states over the period 2000–2022. Their findings indicate a positive relationship between renewable energy investment and economic growth. Using data for European countries over the period 1970–2019, Guliyev and Tatoğlu (2023) examined the relationship between renewable energy and economic growth, highlighting the importance of considering energy efficiency when assessing this relationship. Similarly, F. Liu et al. (2025) found a U-shaped relationship between renewable energy and economic growth across 30 provinces in China from 2001 to 2022. In this context, Song et al. (2025) examined the nonlinear effects of renewable energy consumption on economic growth across 31 Chinese provinces from 2000 to 2022. Their findings suggest that the relationship between renewable energy consumption and economic growth varies across threshold regimes, with a negative association below the estimated threshold and a positive association above it.

Other studies examine the dynamic and causal relationship between renewable energy and economic growth. Using data from 35 OECD countries, Raifu et al. (2025) investigated the causal relationship between renewable energy and economic growth and reported bidirectional causality when labor and capital were incorporated into the analysis. Their findings suggest that the relationship should be examined within a broader economic framework that accounts for other production factors.

Evidence from Saudi Arabia and the Gulf Cooperation Council (GCC) provides a relevant regional perspective on the energy–growth–environment nexus. AlNemer et al. (2023) examined the relationships between renewable and non-renewable energy consumption, $CO_2$ emissions, and economic growth in Saudi Arabia using spectral Granger causality and wavelet analysis. Their findings indicated a long-run causal relationship between non-renewable energy consumption and economic growth, while renewable energy consumption was associated with economic growth across different frequencies and with lower $CO_2$ emissions. The results also revealed variations in these relationships across time-frequency scales. At the GCC level, Bekhet et al. (2017) investigated the dynamic relationships among carbon emissions, financial development, economic growth, and energy consumption using an ARDL model. They reported long-run relationships and causal

linkages among these variables in most GCC countries. For Saudi Arabia, their results indicated unidirectional causality from carbon emissions to energy use.

More recent studies have further examined nonlinear relationships between renewable energy consumption and economic growth in Saudi Arabia. Mighri and AlSaggaf (2023) used NARDL-based approaches to investigate asymmetric relationships between renewable energy consumption and economic growth, highlighting the importance of distinguishing between positive and negative changes in renewable energy consumption. Hamdi et al. (2026) subsequently examined renewable energy, industrialization, entrepreneurial activity, and economic growth using threshold and quantile regression approaches. Their findings showed that renewable energy was not statistically significant in either threshold regime, indicating that the identification of a threshold does not necessarily imply a statistically significant relationship between renewable energy and economic growth in each regime.

Collectively, these studies provide complementary evidence on the energy–growth–environment nexus in Saudi Arabia and the GCC, but they employ different indicators and econometric approaches. The present study complements this literature by examining the relationships between energy consumption, renewable energy, greenhouse gas emissions, and economic growth within an ARDL framework, while testing whether the relationship between renewable energy consumption and economic growth varies across threshold regimes.

More directly related to the present study, Ferhi and Helali (2024b) used panel threshold and smooth transition regression models for OECD countries to examine whether the relationship between renewable energy and economic growth varied across regimes. Ferhi and Helali (2024a) further examined the environmental and welfare implications of renewable energy.

Taken together, the literature reports mixed evidence regarding the relationship between renewable energy consumption and economic growth. Differences in findings may reflect country-specific conditions, energy-system characteristics, and the econometric approaches employed. The Saudi-specific evidence discussed above further highlights the importance of testing whether the relationship varies across regimes rather than presuming that a threshold exists.

### *2.2. The Link Between Economic Growth and Greenhouse Gas Emissions*

The empirical literature on the relationship between economic growth and greenhouse gas emissions reports mixed findings across countries, time periods, and types of emissions.

A first group of studies highlights a positive relationship between economic growth and environmental degradation. Kukla-Gryz (2009) examined the relationship between economic growth, international trade, and air pollution in 86 low- and lower-middle-income countries. The findings indicate that economic growth and trade can contribute to increased air pollution, although the pathways differ across income groups: changes in the structure of economic activity are particularly relevant in developing countries, whereas more direct scale-related effects are observed in developed countries. Similarly, Xuan (2025a) examined the dynamic relationship between carbon dioxide ($CO_2$) emissions and economic growth in Australia over the period 1990–2023 using the autoregressive distributed lag (ARDL) approach. The results indicate a positive relationship between GDP and $CO_2$ emissions. Consistent with this evidence, Q. Chen and Zhang (2025) report positive and statistically significant relationships among economic growth, energy consumption, and $CO_2$ emissions.

Evidence from Saudi Arabia further highlights the importance of considering the composition of economic activity when assessing the growth–emissions relationship. Tlili et al. (2025) investigated the relationship between economic growth and $CO_2$ emissions in Saudi Arabia over the period 1970–2020 using ARDL and NARDL models. Their findings

suggested that oil-related economic growth was associated with higher $CO_2$ emissions in the long run, whereas non-oil economic activity was associated with lower emissions. These results suggest that the relationship between economic growth and environmental degradation may depend on the structure of economic activity and the econometric approach employed.

The relationship between methane ($CH_4$) emissions and economic growth has also been examined in several empirical studies, although the findings vary across countries and contexts. Using data for selected oil-producing African countries, Yusuf et al. (2020) stated a positive long-run relationship between economic growth and methane emissions, together with evidence of a nonlinear relationship. Similarly, Djoukouo (2021) reported evidence consistent with the Environmental Kuznets Curve (EKC) hypothesis for methane emissions and economic growth in Central African countries, alongside bidirectional Granger causality between the two variables. More recently, Destiartono and Firmansyah (2024) examined the relationship between economic growth and methane emissions in Indonesia, also considering the role of renewable energy in this relationship. Taken together, these studies suggest that the relationship between economic growth and methane emissions may vary across countries and contexts. In the present study, methane emissions are considered separately from $CO_2$ emissions; however, the interpretation of the results depends on the specific methane measure used.

A second strand of the literature examines whether the relationship between economic growth and environmental degradation is nonlinear, including the possibility of an inverted U-shaped relationship. Li and Haneklaus (2021) examined the relationships among renewable energy, economic growth, fossil fuel consumption, urbanization, and $CO_2$ emissions in China over the period 1990–2020. Their findings indicate an inverted U-shaped relationship between GDP per capita and $CO_2$ emissions in both the short and long run. Similarly, S. Jiang et al. (2022) reported an inverted U-shaped relationship between economic growth and air pollution, suggesting that the relationship between economic activity and environmental degradation may change as economic growth progresses.

A third group of studies examines the direction of causality and the extent to which the relationships between economic growth and emissions vary across countries. Mensah et al. (2019) found evidence of unidirectional causality from carbon emissions to economic growth in non-oil-exporting countries in both the short and long run, whereas bidirectional causality was reported in the long run for oil-exporting countries. Raza et al. (2025) examined the relationships among $CO_2$ emissions, economic growth, and energy consumption in BRICS countries over the period 2009–2022. Their findings indicated a long-run relationship among these variables, although the abstract does not establish a specific direction of causality.

Finally, Oprea et al. (2025) found a negative long-run relationship between GDP per capita and $CO_2$ emissions in Iceland. They also identified an inverted U-shaped relationship between renewable energy consumption and $CO_2$ emissions, indicating that this relationship varies with the level of renewable energy consumption.

On the whole, the literature suggests that the relationships between economic growth and greenhouse gas emissions vary across countries, emission measures, and economic structures. Some studies document nonlinear relationships or evidence of causality, while others highlight differences associated with national energy systems and the composition of economic activity. These findings highlight the importance of examining environmental indicators separately and avoiding the assumption that similar relationships hold across countries. Accordingly, the present study examines the relationships between economic growth and environmental indicators in Saudi Arabia and tests whether the relationship between renewable energy consumption and economic growth varies across threshold regimes, without presuming the existence of a threshold in advance.

### *2.3. The Link Between Economic Growth and Fossil Fuel Energy*

The empirical literature on the relationship between economic growth and fossil fuel energy consumption highlights heterogeneous and sometimes contradictory findings across countries and regions.

A first group of studies reports a positive association between fossil fuel energy consumption and economic growth. Mensah et al. (2019) examined the dynamic relationships among economic growth, fossil fuel energy consumption, carbon emissions, and oil prices in 22 African countries from 1990 to 2015, distinguishing between oil-exporting and non-oil-exporting countries. Using the pooled mean group (PMG) panel ARDL approach, they found bidirectional causality between fossil fuel energy consumption and economic growth in both the short and long run across the country groups. They also reported bidirectional causality between fossil fuel energy consumption and carbon emissions. However, the relationship between carbon emissions and economic growth differed across the two groups: carbon emissions Granger-caused economic growth in both the short and long run in non-oil-exporting countries, whereas bidirectional causality was found only in the long run among oil-exporting countries. In addition, oil prices exhibited unidirectional causality with economic growth, fossil fuel energy consumption, and carbon emissions across the country groups. These findings highlight the interdependence of energy use, economic performance, emissions, and oil prices, while indicating that the direction of these relationships may vary with countries' oil-exporting status. Similarly, Rahman and Velayutham (2020), analyzing five South Asian countries from 1990 to 2014, found that both renewable and non-renewable energy consumption were positively associated with economic growth. Ivanovski et al. (2021) reported a positive and statistically significant relationship between non-renewable energy consumption and economic growth in both OECD and non-OECD countries. Likewise, Z. Jiang et al. (2022) found that fossil fuel energy consumption was associated with higher economic growth in the East Asia–Pacific region, where positive shocks to fossil fuel consumption were associated with higher growth. AlNemer et al. (2023) also reported a significant long-run causal relationship between fossil fuel energy consumption and economic growth in Saudi Arabia. In the same context, Alao et al. (2025) found that economic growth in West Africa remained highly dependent on fossil fuel consumption. Finally, Thach (2025) reported positive relationships between both renewable and non-renewable energy consumption and economic growth, although the estimated association was stronger for renewable energy.

A second strand of the literature highlights broader and more complex relationships between energy consumption, economic growth, and environmental outcomes. For instance, Usman et al. (2021), examining 15 high-emitting countries over the period 1990–2017, found that renewable and non-renewable energy consumption, together with financial development, were associated with higher economic growth while also reducing environmental pressure.

Overall, the empirical evidence indicates that the relationship between energy consumption and economic growth varies across countries and energy systems. The findings of Mensah et al. (2019), in particular, illustrate that these relationships may also depend on oil-exporting status and the interactions among energy consumption, carbon emissions, and oil prices. Such differences support examining the Saudi Arabian energy–growth relationship in its specific economic and energy context, while treating the possibility of nonlinear relationships as an empirical question rather than an assumption.

Taken together, the reviewed literature motivates testing the relationship between renewable energy consumption and economic growth and assessing whether this relationship varies across threshold regimes in Saudi Arabia. Accordingly, the study examines the following hypotheses:

**H1.** *Renewable energy consumption is significantly associated with economic growth in Saudi Arabia.*

**H2.** *The association between renewable energy consumption and economic growth differs across regimes defined by the level of renewable energy.*

## 3. Econometric Methodology

### *3.1. Model Specification*

#### 3.1.1. ARDL Approach

To examine the associations between the fossil fuel share, renewable energy, $CO_2$, and $CH_4$ emissions and economic growth in Saudi Arabia, this study employs the Autoregressive Distributed Lag (ARDL) model, following Pesaran and Shin (1995) and Pesaran et al. (2001).

The approach suits the present system for two reasons. It remains valid whatever the mix of I(0) and I(1) regressors, provided none is I(2), and here, every series is I(1) (Section 4.2), with none integrated of order two, satisfying the integration-order requirement for the bounds-testing approach. It also performs well with small samples (Narayan, 2005), an important property given 33 annual observations. Following Pesaran et al. (2001), the ARDL model is reparametrized as an unrestricted error-correction model (UECM), which jointly estimates the short-run dynamics and the long-run relationship. With LGDP denoting the logarithm of real GDP, the model is:

$$\begin{aligned}\Delta LGDP_t = \alpha_0 + \sum_{i=1}^{P} \beta_i \Delta LGDP_{t-i} + \sum_{i=0}^{q1} \gamma_i \Delta EC_{t-i} + \sum_{i=0}^{q2} \delta_i \Delta RE_{t-i} + \sum_{i=0}^{q3} \emptyset_i \Delta CO2_{t-i} + \sum_{i=0}^{q4} \theta_i \Delta CH4_{t-i} \\ + \lambda_1 LGDP_{t-1} + \lambda_2 EC_{t-1} + \lambda_3 RE_{t-1} + \lambda_4 CO2_{t-1} + \lambda_5 CH4_{t-1} + \varepsilon_t\end{aligned} \quad (1)$$

where

$\Delta$ denotes the first difference operator;

$P$, $q1$, $q2$, $q3$, and $q4$ denote the optimal lag lengths of the dependent and explanatory variables;

$\lambda_1, \lambda_2, \lambda_3, \lambda_4$, and $\lambda_5$ represent the long-run coefficients related to the variables involved in the model;

$\varepsilon_t$ represents the stochastic error term.

The bounds test is a single-equation procedure: it treats the fossil fuel share, renewable energy, and the two emission series as long-run forcing variables for real GDP (Pesaran et al., 2001). This assumption is strong in a hydrocarbon economy, where emissions are largely a by-product of output and where long-run feedback from growth to emissions has been reported for oil exporters (Mensah et al., 2019). It is therefore tested in two ways (Appendix A). First, Granger non-causality is tested in both directions using the Toda and Yamamoto (1995) procedure, which is valid regardless of the integration and cointegration properties of the series. Second, long-run forcing is tested directly as weak exogeneity with respect to the long-run relationship (Johansen, 1992): the lagged equilibrium error must not enter the equation of a forcing variable. This test is run both in single-equation marginal models and in a Johansen (1991) vector error-correction model in which all five series are endogenous; the latter also gives a system estimate of the cointegrating vector that does not rest on the forcing assumption. In all cases, the long-run coefficients describe the level relationship that links the five series, normalized to LGDP; they are not causal effects of emissions or renewable energy on output, which is why the $CO_2$ coefficient is interpreted as a scale effect.

While the ARDL approach permits the assessment of short-run and long-run linear connections between the variables, it may not adequately capture potential nonlinearities and regime-dependent impacts. Consequently, to determine whether the effect of renewable energy on economic growth varies across different levels of renewable energy expansion, the analysis further employs a nonlinear threshold regression model.

3.1.2. Non-Linear Threshold Regression

Since all variables are I(1), we estimate the baseline threshold regression (TR) in first differences to ensure that the dependent and explanatory variables are stationary, while retaining the threshold variable—lagged renewable energy consumption (RE_{t − 1})—in levels, as its role is solely to define the regimes. A separate levels specification is then estimated as a threshold cointegration model to assess whether the identified regime relationship has long-run content (Balke & Fomby, 1997; Gonzalo & Pitarakis, 2006). The baseline equation is:

$$y_t = \alpha + \beta x_t + \varphi' z_t + \varepsilon_t \quad (2)$$

$y_t$ represents ΔLGDP × 100, the growth rate of real GDP (first difference in LGDP, in %);

$x_t$ represents the regime-dependent regressor, the change in renewable energy consumption (ΔRE);

$z_t$ gathers the first differences in the control variables, ΔEC, $\Delta CO_2$, and $\Delta CH_4$;

$\epsilon_t$ is an error term.

Following Hansen (2000), we estimate a two-regime model in this article and therefore use the following formula:

$$y_t = \begin{cases} \theta_1' w_t + \varepsilon_t, & if\ q_{t-1} \le \gamma \\ \theta_2' w_t + \varepsilon_t, & if\ q_{t-1} > \gamma \end{cases} \quad (3)$$

In Equation (2), φ is the coefficient vector of the controls. In Equation (3), $w_t = (1, x_t, z_t')'$ stacks the intercept, the regime-dependent regressor, and the controls, and $\theta_1$ and $\theta_2$ are the regime-specific coefficient vectors. Equation (4) writes out Equation (3) in full, with superscripts indexing the regime and subscripts indexing the regressor, as in Equation (5). Therefore, with the threshold variable $q_{t-1} = RE_{t-1}$ kept in levels to define the regimes, the regression model is written as follows:

$$\begin{aligned} \Delta LGDP_t = {} & \left(\beta_0^1 + \beta_1^1 \Delta EC_t + \beta_2^1 \Delta RE_t + \beta_3^1 \Delta CO2_t + \beta_4^1 \Delta CH4_t\right) \times I(RE_{t-1} \le \gamma) \\ & + \left(\beta_0^2 + \beta_1^2 \Delta EC_t + \beta_2^2 \Delta RE_t + \beta_3^2 \Delta CO2_t + \beta_4^2 \Delta CH4_t\right) \times I(RE_{t-1} > \gamma) + \mu_t \end{aligned} \quad (4)$$

The SSR-minimizing split reported in Section 4.6 leaves eight observations in the upper regime, so it remains admissible under a stricter rule of eight observations per regime (25% trimming); since a narrower grid cannot raise the sup-F statistic above its full-grid value, the conclusion also holds under this stricter rule. The threshold effect is then tested through the following hypotheses:

$$\begin{cases} H_0 : \beta_i^1 = \beta_i^2 \ \ \forall i = 0, 1, \ldots, 4 \\ H_1 : \beta_i^1 \ne \beta_i^2 \end{cases} \quad (5)$$

The long-run counterpart, estimated in levels as a robustness check, is:

$$\begin{aligned} LGDP_t = {} & (\alpha_1 + b_1' X_t) \times I(RE_{t-1} \le \gamma) + (\alpha_2 + b_2' X_t) \times I(RE_{t-1} > \gamma) \\ & + u_t, \quad X_t = (EC_t, RE_t, CO2_t, CH4_t)', \quad u_t \sim I(0) \end{aligned} \quad (6)$$

Three implementation choices deserve emphasis. First, the dependent variable is ΔLGDP × 100 and all regressors enter as first differences, so both sides of Equation (4) are I(0). Second, the renewable energy consumption never falls over the sample, so the threshold on $RE_{t-1}$ splits the sample into two consecutive sub-periods; because the asymptotic theory of Hansen (2000) assumes a stationary threshold variable, the existence of the threshold is assessed with the sup-F statistic of Hansen (1996) using a fixed-regressor wild bootstrap (1000 replications) rather than asymptotic critical values. The 95% confidence interval for γ is obtained by inverting the likelihood-ratio statistic of Hansen (2000). All coefficients,

the intercept, the switch across regimes, and standard errors are heteroskedasticity-robust. Third, the long-run counterpart of the model, Equation (6), is estimated in levels, and its residual is tested for a unit root against critical values simulated under the null of no cointegration for the same sample size, grid, and trimming, following the logic of Gregory and Hansen (1996). Only a stationary residual would give its regime-specific coefficients, estimated by fully modified OLS (P. C. Phillips & Hansen, 1990), a long-run interpretation. Equation (4) omits the lagged equilibrium error. It therefore asks whether short-run growth dynamics differ across regimes, not whether the speed of adjustment does; the omitted error-correction term, which is stationary, is left in the residual, where it raises the residual variance, lowers the power of the sup-F test, and may bias the regime-specific coefficients. A threshold error-correction model (Balke & Fomby, 1997; Hansen & Seo, 2002) is the natural extension, but with eight observations in the upper regime it would leave almost no degrees of freedom. The long-run side of the threshold question is instead examined through Equation (6).

### *3.2. Data*

This study examines the relationships between energy consumption, renewable energy consumption, $CO_2$ and $CH_4$ emissions, and real output in Saudi Arabia over the period 1990–2022, determined by data availability. We draw real GDP, fossil fuel energy consumption, $CO_2$ emissions, renewable energy consumption, and $CH_4$ emissions from the World Bank's World Development Indicators (WDI). The variables are described in Table 1.

**Table 1.** Description and source of the variables used in the analysis.

| Variable Name | Signs | Description | Source |
|---|---|---|---|
| Real GDP (log) | LGDP | Natural logarithm of real GDP (index 1990 = 100, cumulated from the growth rate of GDP at constant prices) | WDI |
| Renewable energy | RE | Renewable energy consumption (% of total final energy consumption) | WDI |
| $CO_2$ emissions | $CO_2$ | Carbon dioxide ($CO_2$) emissions excluding LULUCF per capita (t $CO_2$e/capita) | WDI |
| Fossil fuel share of energy consumption | EC | Fossil fuel energy consumption (% of total) | WDI |
| $CH_4$ | $CH_4$ | Methane ($CH_4$) emissions from Industrial Processes (Mt $CO_2$e) | WDI |

## 4. Empirical Results

### *4.1. Descriptive Analysis*

Table 2 presents the descriptive statistics. From 1990 to 2022, the logarithm of the real GDP index averaged 5.138, with a standard deviation of 0.348, rising from 4.605 in 1990 to 5.742 in 2022; real output roughly tripled over the period. The fossil fuel share (EC) is high and remarkably stable, with a mean of 99.854% and a standard deviation of only 0.209, confirming that hydrocarbons still dominate the national energy mix. Renewable energy consumption (RE) averages 0.048 and is strongly right-skewed (skewness of 1.822); it is essentially nil until 2009 and then rises rapidly due to the relatively late development

of renewable energy in Saudi Arabia. Any coefficient on renewable energy is therefore identified almost entirely from the years after 2009. Figure 1 plots the series: output rises steadily; renewable energy consumption takes off only after 2010 and jumps in 2018–2019; $CO_2$ per capita peaks in 2015 and declines thereafter; and the fossil fuel share hardly moves, apart from a dip in 2011–2014. $CO_2$ emissions per capita average 13.472, with moderate dispersion, while $CH_4$ emissions average 0.496 and trend broadly upward. The Jarque–Bera statistics indicate normality for LGDP, $CO_2$, and $CH_4$, and non-normality for EC and RE. Overall, the descriptive statistics indicate substantial variation in the variables, while the distributional characteristics of EC and RE reflect the specific structure of Saudi Arabia's energy system.

**Table 2.** Descriptive analysis of variables.

| | LGDP | EC | RE | $CO_2$ | $CH_4$ |
|---|---|---|---|---|---|
| Mean | 5.138 | 99.854 | 0.048 | 13.472 | 0.496 |
| Median | 5.125 | 99.989 | 0.001 | 13.215 | 0.534 |
| Maximum | 5.742 | 99.997 | 0.267 | 17.258 | 0.769 |
| Minimum | 4.605 | 99.204 | 0.000 | 10.710 | 0.044 |
| Std. Dev. | 0.348 | 0.209 | 0.092 | 2.082 | 0.249 |
| Skewness | 0.153 | −1.257 | 1.822 | 0.252 | −0.392 |
| Kurtosis | 1.567 | 3.809 | 4.572 | 1.630 | 1.695 |
| Jarque-Bera | 2.953 | 9.593 | 21.651 | 2.932 | 3.187 |
| Probability | 0.228 | 0.008 | 0.000 | 0.231 | 0.203 |

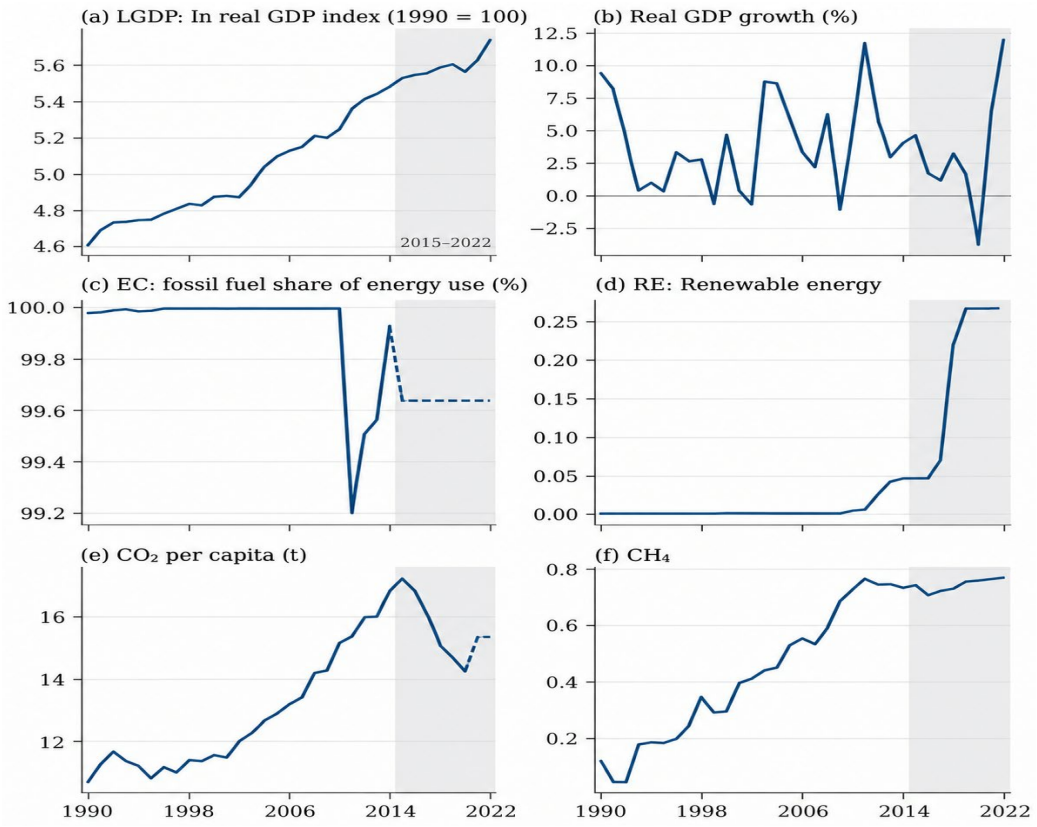



**Figure 1.** Time-series plots of the variables, 1990-2022. Shaded area: 2015-2022, the upper regime of the SSR-minimizing split (Section 4.6). Dashed segments: imputed values (Section 3.2).

### 4.2. Series Stationarity

Table 3 reports the ADF (Dickey & Fuller, 1979) and Phillips–Perron (P. C. B. Phillips & Perron, 1988) tests in levels, with an intercept and a linear trend, and in first differences with an intercept. A trend is the relevant alternative for the trending series (LGDP, RE, $CO_2$ and $CH_4$) but not for the fossil fuel share, a bounded ratio hovering near 100% (Elder & Kennedy, 2001). On this basis, no series rejects the unit root in levels in the preferred specifications: the preferred ADF statistics are −2.7220 for LGDP, −0.8626 for EC, −1.4249 for RE, −0.9853 for $CO_2$, and −1.3930 for $CH_4$. Two borderline cases deserve mention. For $CO_2$, the ADF test with trend rejects at 5% ($p = 0.0174$), but the Phillips–Perron test does not ($p = 0.8583$), and neither rejects without a trend; for EC, a rejection appears only if a deterministic trend is imposed, which has no economic meaning for a share capped at 100%. After first differencing, every series rejects the unit root at the 5% level or better in both tests. We therefore treat all five variables as I(1), and none as I(2). Even if $CO_2$ or EC were regarded as I(0), the bounds test would remain valid, since it accommodates a mix of I(0) and I(1) regressors (Pesaran et al., 2001).

**Table 3.** Variables' unit root tests.

| | | Level (c) | | Level (c, t) | | First Difference (c) | | |
|---|---|---|---|---|---|---|---|---|
| **Variable** | **Test** | **Stat.** | **Prob.** | **Stat.** | **Prob.** | **Stat.** | **Prob.** | **Order** |
| LGDP | ADF | 0.7438 | 0.9907 | −2.7220 | 0.2271 | −4.4078 *** | 0.0003 | I(1) |
| | PP | 0.6708 | 0.9892 | −1.5260 | 0.8201 | −3.8031 *** | 0.0029 | |
| EC | ADF | −0.8626 | 0.8000 | −3.8255 ** | 0.0153 | −5.4778 *** | 0.0000 | I(1) |
| | PP | −2.5947 * | 0.0941 | −3.6189 ** | 0.0283 | −11.0143 *** | 0.0000 | |
| RE | ADF | −0.2438 | 0.9331 | −1.4249 | 0.8535 | −3.6190 *** | 0.0054 | I(1) |
| | PP | 0.8795 | 0.9928 | −0.4665 | 0.9847 | −3.4515 *** | 0.0093 | |
| $CO_2$ | ADF | −0.9853 | 0.7586 | −3.7837 ** | 0.0174 | −3.3425 ** | 0.0131 | I(1) |
| | PP | −1.0651 | 0.7288 | −1.4090 | 0.8583 | −4.2938 *** | 0.0005 | |
| $CH_4$ | ADF | −1.0659 | 0.7285 | −1.3930 | 0.8630 | −5.9294 *** | 0.0000 | I(1) |
| | PP | −1.1193 | 0.7073 | −1.5088 | 0.8261 | −5.9444 *** | 0.0000 | |

Note: ***, **, and * indicate rejection of the unit-root null at the 1%, 5%, and 10% levels, respectively. Levels are tested with an intercept (c) and with an intercept and a linear trend (c, t); first differences with an intercept. The order of integration is based on (c, t) for LGDP, RE, $CO_2$, and $CH_4$ and on (c) for EC. MacKinnon (1996) probabilities are reported next to each statistic. ADF lag lengths are selected by the AIC (maximum 3); PP tests use the Newey–West bandwidth. LGDP = ln (real GDP index, 1990 = 100).

These tests do not allow for structural breaks, although the sample covers the 1990–1991 Gulf War, the 2008–2009 financial crisis, the 2014–2016 oil price collapse, and the 2020 pandemic contraction. Ignoring breaks biases unit root tests towards non-rejection (Perron, 1989; Zivot & Andrews, 1992), so there is a risk that a break-stationary series is classified as I(1). Such a misclassification would not invalidate the bounds test, which accommodates any mix of I(0) and I(1) regressors provided no series is I(2), and the clear rejections in first differences rule out that case. In addition, the significant error-correction coefficient and the significant long-run coefficients rule out degenerate cases in which the bounds test rejects spurious relationships (McNown et al., 2018).

### 4.3. Selecting the Number of Lags

Table 4 reports the lag-order criteria for the VAR in levels of LGDP, EC, RE, $CO_2$ and $CH_4$. All five criteria (LR, FPE, AIC, SC and HQ) select lag 1. On this basis, the

ARDL orders are chosen using the Akaike criterion, with a maximum of 2 lags, yielding an ARDL (2,1,2,0,2) specification. With 33 annual observations, this parsimony matters: every additional lag costs five degrees of freedom in the VAR and weakens each of the tests that follow. The VAR criteria and the ARDL lag orders answer different questions. In the five-variable VAR, each additional lag adds five parameters to every equation, which explains why all criteria stop at one lag; in the single-equation ARDL, an extra lag on a regressor costs one parameter. Allowing up to two lags follows the recommendation of Pesaran and Shin (1995) to augment the lag order sufficiently to remove residual serial correlation, as confirmed by the Breusch–Godfrey test ($p = 0.316$). The second lags retained for LGDP, RE, and $CH_4$ are significant at the 5% or 10% level, so they carry information that a one-lag cap would push into the residual. The finite-sample bounds in Table 5 are simulated for this lag structure.

**Table 4.** Number of lags.

| Lag | LogL | LR | FPE | AIC | SC | HQ |
|---|---|---|---|---|---|---|
| 0 | 76.0074 | NA | $7.049 \times 10^{-9}$ | −4.5811 | −4.3498 | −4.5057 |
| 1 | 218.0252 | 229.0610 * | $3.803 \times 10^{-12}$ * | −12.1307 * | −10.7429 * | −11.6783 * |
| 2 | 236.1300 | 23.3611 | $6.801 \times 10^{-12}$ | −11.6858 | −9.1416 | −10.8565 |

Note: * indicates the lag order selected by the criterion. VAR in levels of LGDP, EC, RE, $CO_2$, and $CH_4$ with an intercept; common sample.

**Table 5.** Results of the ARDL cointegration test.

| Model | F-Statistic = 5.357 | | | |
|---|---|---|---|---|
| **Significance Level** | **Asymptotic I(0)** | **Asymptotic I(1)** | **Finite-Sample I(0)** | **Finite-Sample I(1)** |
| 10% | 2.45 | 3.52 | 2.71 | 4.06 |
| 5% | 2.86 | 4.01 | 3.25 | 4.89 |
| 2.5% | 3.25 | 4.49 | 3.92 | 5.83 |
| 1% | 3.74 | 5.06 | 5.04 | 7.06 |
| Finite-sample *p*-value (I(1) bound) | | | 0.036 | |

Note: Asymptotic bounds from Pesaran et al. (2001), Table CI(iii), k = 4, case III. Finite-sample bounds are simulated (5000 replications) for T = 31 and the selected ARDL (2, 1, 2, 0, 2) order, with the regressors generated as I(0) (lower bound) and I(1) (upper bound) processes.

### *4.4. Bound Test Results*

Table 5 reports the bounds test of Pesaran et al. (2001). The F-statistic equals 5.357 for k = 4 regressors (case III: unrestricted intercept, no trend). It exceeds the upper bound at the 1% level, so the null of no level relationship is rejected: real GDP, the fossil fuel share, renewable energy, $CO_2$, and $CH_4$ exhibit a long-run level relationship. Because all regressors are I(1), the bounds-test decision is based on the comparison of the F-statistic with the I(0) and I(1) bounds, and here the statistic exceeds the I(1) upper bound. Because asymptotic bounds can be too lenient with 33 observations (Narayan, 2005), we also simulate finite-sample bounds for T = 31 and the selected lag structure: the finite-sample upper bound simulated for T = 31 is exceeded at the 5% level (simulated *p*-value = 0.036). Economically, a level relationship means that real output, the energy mix, renewable energy, and emissions cannot drift apart indefinitely; a long-run relationship ties them together, which justifies the error-correction representation estimated below.

### *4.5. ARDL Short- and Long-Run Evaluation Results*

Table 6 reports the ARDL (2, 1, 2, 0, 2) estimates for LGDP. Statistically, the error-correction term is −0.5273*** ($p = 0.0002$), so that about 52.7% of any deviation from the long-run path is corrected within a year, with a half-life of roughly 0.93 years; a negative, significant coefficient above −1 confirms the bounds test and implies a monotonic return to equilibrium. The long-run coefficients are semi-elasticities of real GDP. Renewable energy has a positive coefficient of 1.2665*** ($p = 0.0000$): an additional 0.1 of renewable energy is associated with real GDP 0.1267 log points higher, that is, about 13.5% ($\exp(0.1267)$. Economically, this magnitude cannot be read as the direct output of renewable energies; rather, it reflects the imprint of the diversification phase during which these energies were built: the years of solar deployment coincide with those of the Vision 2030 investment program; renewable energy acts much as a marker of this structural change as much as a production input. Because renewable energy is close to zero until 2009, this coefficient is identified almost entirely from the years since then, during which output and renewable energy both trended upwards, and it may partly capture this common trend. A Vision 2030 dummy would not separate the two, as it would be almost collinear with the renewable series. $CO_2$ per capita (0.0602***, $p = 0.0004$) and $CH_4$ (0.6603***, $p = 0.0000$) are also positive and significant; one additional tonne of $CO_2$ per capita is associated with output about 6.2% higher ($\exp(0.0602) - 1$), the scale effect of a hydrocarbon economy rather than a return to pollution (AlNemer et al., 2023). The fossil fuel share is insignificant (0.0686, $p = 0.4428$), which is unsurprising for a variable that hardly moves from 99.9%. In the short run, a rise in renewable energy is associated with faster output growth in the same year (0.7570***, $p = 0.0063$). This association is partly offset the following year (−0.3954*, $p = 0.0658$); lagged methane growth carries a negative short-run coefficient (−0.3256**, $p = 0.0129$). Financially, the distinction matters for project appraisal: the long-run semi-elasticity summarizes how output and renewable energy have moved together over the cycle, whereas the short-run terms capture short-term changes associated with renewable energy deployment. These short-run associations may be consistent with installation-related activity, but they do not by themselves identify project expenditures. The model estimates 12 parameters (the intercept, five lagged levels, and six short-run terms) on 31 effective observations, leaving 19 residual degrees of freedom, which explains the gap between the $R^2$ (0.684) and the adjusted $R^2$ (0.500). The delta-method standard errors of the long-run coefficients rest on asymptotic arguments, so *p*-values are taken from the t distribution with 19 degrees of freedom; the long-run t-statistics for RE, $CO_2$, and $CH_4$ (7.98, 4.31, and 5.27) remain well above the 1% critical value of 2.86. In contrast, the lagged short-run renewable term, significant only at the 10% level, is weak evidence. With so few degrees of freedom, the tests also have limited power, so a non-rejection signals an absence of evidence rather than evidence of absence. The diagnostic tests support the specification: the Breusch–Godfrey ($p = 0.316$) and ARCH ($p = 0.332$) tests detect neither serial correlation nor conditional heteroscedasticity, the residuals are normal (Jarque–Bera $p = 0.643$), and the CUSUM and CUSUM of squares statistics remain within their 5% bands. No break dummies are included: the diagnostics show none of the serial correlation, non-normality, or instability that unmodelled breaks would leave in the residuals, and four crisis dummies would reduce the residual degrees of freedom from 19 to 15. Figure 2 plots the CUSUM and CUSUM of squares statistics, both of which remain within their 5% bands over 2004–2022.

**Table 6.** ARDL estimates: long- and short-run results for LGDP.

| Dependent Variable: LGDP. Selected Model: ARDL(2, 1, 2, 0, 2) | | | | |
|---|---|---|---|---|
| **Variable** | **Coefficient** | **Std. Error** | **t-Statistic** | **Prob.** |
| | Long-run results | | | |
| EC | 0.0686 | 0.0875 | 0.784 | 0.4428 |
| RE | 1.2665 *** | 0.1588 | 7.976 | 0.0000 |
| $CO_2$ | 0.0602 *** | 0.0140 | 4.310 | 0.0004 |
| $CH_4$ | 0.6603 *** | 0.1253 | 5.271 | 0.0000 |
| C | −2.8708 | 8.7986 | −0.326 | 0.7478 |
| | Short-run results | | | |
| D(LGDP(−1)) | 0.3522 ** | 0.1623 | 2.170 | 0.0429 |
| D(EC) | −0.0245 | 0.0334 | −0.732 | 0.4731 |
| D(EC(−1)) | — | — | — | — |
| D(RE) | 0.7570 *** | 0.2464 | 3.072 | 0.0063 |
| D(RE(−1)) | −0.3954 * | 0.2025 | −1.952 | 0.0658 |
| $D(CO_2)$ | — | — | — | — |
| $D(CO_2(-1))$ | — | — | — | — |
| $D(CH_4)$ | 0.0644 | 0.1213 | 0.531 | 0.6018 |
| $D(CH_4(-1))$ | −0.3256 ** | 0.1187 | −2.742 | 0.0129 |
| C | −1.5138 | 4.6243 | −0.327 | 0.7470 |
| CointEq(−1) | −0.5273 *** | 0.1119 | −4.710 | 0.0002 |
| Breusch–Godfrey test | 1.235 | | | 0.316 |
| ARCH test | 0.942 | | | 0.332 |
| Jarque–Bera (prob.) | | | | 0.643 |
| R-squared | 0.684 | | Adjusted R-squared | 0.500 |
| F-statistic | 3.732 *** | | | |
| CUSUM | Stable | | CUSUMSQ | Stable |

Note: ***, **, and * denote significance at the 1%, 5%, and 10% levels, respectively. D (·) is the first-difference operator; CointEq(−1) is the coefficient of LGDP(−1) in the unrestricted error-correction model. Long-run coefficients and their standard errors are obtained by the delta method and read as semi-elasticities of real GDP. "—" marks a lag not retained by the AIC. The Breusch–Godfrey (2 lags) and ARCH (1 lag) rows report the test statistic and its probability.

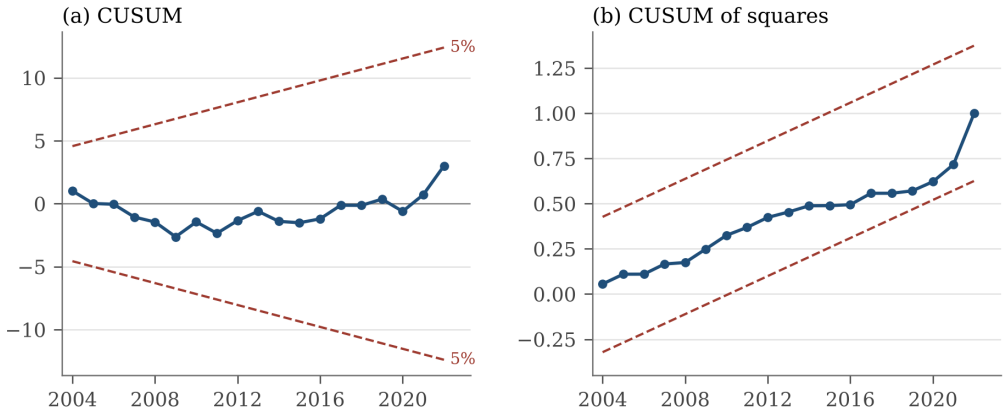



**Figure 2.** CUSUM and CUSUM of squares tests for the ARDL (2, 1, 2, 0, 2) model, recursive residuals 2004–2022. Dashed lines: 5% bands, following Brown et al. (1975) for the CUSUM; for the CUSUM of squares, the 5% critical value (0.374) for 19 recursive residuals is obtained by simulation.

Because the bounds test conditions on the regressors, Table A1 (Appendix A) tests whether they are long-run forcing. The Toda–Yamamoto tests reject Granger non-causality in neither direction between LGDP and any regressor (all $p$-values $\geq 0.24$); with 20 residual degrees of freedom per equation, this reflects the low power of the tests more than an absence of feedback. The weak exogeneity tests are more informative. The lagged equilibrium error does not enter the EC and $CH_4$ equations in the single-equation tests ($p = 0.210$ and 0.214), but it does enter the $CO_2$ equation ($p = 0.036$ in both the single-equation and system tests) and the RE equation ($p = 0.052$ and 0.002). The system loading of EC is also significant, a fragile result for a series that is nearly constant and imputed after 2014. $CO_2$ and renewable energy are therefore not long-run forcing; they respond to deviations from the long-run relationship, in line with the long-run feedback reported for oil exporters by Mensah et al. (2019). The long-run results, however, do not depend on this assumption. The Johansen procedure, in which all five series are endogenous, points to a single cointegrating vector, and its normalized coefficients are close to those in Table 6: 1.158 for RE (against 1.267), 0.079 for $CO_2$ (0.060), and 0.580 for $CH_4$ (0.660); only the fossil fuel share differs, becoming significant (0.188). Output itself adjusts to this vector (loading −0.295, $p = 0.012$), so normalizing the relationship on LGDP is not arbitrary. With a single cointegrating vector, the feedback does not render the ARDL long-run estimates inconsistent, but it reduces the conditional model's efficiency and makes its inference, including the bounds test, less reliable (Johansen, 1992). It also means that the $CO_2$ and RE coefficients are two-way long-run associations, not effects running from these variables to output.

### *4.6. Nonlinear Threshold Regression Model Results*

Table 7 reports the threshold tests. In the main specification, estimated in first differences, the grid search over the observed values of $RE_{t-1}$ (at least six observations per regime) locates the SSR-minimizing split at 0.0425, which separates 1991–2014 (24 observations) from 2015–2022 (8 observations). The split carries no statistical weight: the sup-F statistic is only 1.861, with a bootstrap $p$-value of 0.943, so linearity cannot be rejected, and the 95% confidence interval for the threshold, [0.0000; 0.0425], covers almost the whole admissible grid. The levels check points the same way. Its sup-F statistic (12.814, bootstrap $p = 0.483$) does not reject linearity either, and the ADF statistic of its residual, −5.031, stays above the simulated critical values of −6.688 (5%) and −6.178 (10%) (simulated $p = 0.434$), so there is no evidence of threshold cointegration. Statistically, the regime change that appears when a stationary growth rate is regressed on trended levels does not survive a balanced specification test. The levels threshold, 0.0006, lies at the bottom of the renewable series. It separates 1991–2010 (20 observations), when renewable energy is nil or fixed at this initial value, from 2011–2022 (12 observations), when it becomes measurable, and its 95% likelihood-ratio confidence set covers the whole admissible grid, [0.0000; 0.0425]. Table A2 (Appendix A) reports the complete levels results. In the lower regime, RE takes only three distinct values, so its coefficient (−31.67) rests on two step changes, in 2000 and 2010, and has no economic content. In the upper regime, where renewable energy actually varies, the coefficients on RE (1.47) and $CO_2$ (0.10) are close to the ARDL long-run estimates (1.27 and 0.06). Since the unit root in the residual is not rejected, however, these FMOLS coefficients have no long-run interpretation and their t-statistics cannot be read in the usual way; Equation (6) is informative only as a test of threshold cointegration, which it does not support.

Because renewable energy never declines, the threshold search orders the observations by time, and a threshold on $RE_{t-1}$ is observationally equivalent to a single structural break at an unknown date. The test in Table 7 is therefore, in substance, a test of parameter stability, and the 2014/2015 split cannot be distinguished from a time break associated with the oil

price collapse of 2014 or the launch of Vision 2030 in 2016. Read in this way, the result remains informative: a statistic of 1.861 would not reject stability even if the break date were fixed in advance, so the short-run growth relationship estimated over 2015–2022 does not differ significantly from that of 1991–2014.

**Table 7.** Results of the RE threshold test.

| Specification | Threshold | 95% CI | SSR | sup-F | Bootstrap *p* | Residual ADF |
|---|---|---|---|---|---|---|
| First differences, Equation (4) | 0.0425 | [0.0000; 0.0425] | 215.896 | 1.861 | 0.943 | — |
| Levels, Equation (6) | 0.0006 | [0.0000; 0.0425] | 0.029 | 12.814 | 0.483 | −5.031 [5% c.v. −6.688] |

Note: Threshold variable: $RE_{t-1}$ in levels. Sup-F *p*-values from a fixed-regressor wild bootstrap with 1000 replications (Hansen, 1996); confidence interval by inversion of the likelihood-ratio statistic (Hansen, 2000). Residual ADF: test without deterministic terms on the residual of Equation (6), with critical values simulated under the null of no cointegration (5000 replications) for the same sample size, grid, and trimming (Gregory & Hansen, 1996); 1% critical value = −7.723. For the levels specification, the confidence set is obtained in the same way and is descriptive only, since the theory of Hansen (2000) does not cover I(1) regressors; the complete levels results are reported in Table A2.

Table 8 reports the regime-specific estimates at the SSR-minimizing split, although the threshold itself is not significant. With eight observations and five parameters in the upper regime, only three residual degrees of freedom remain, so these estimates are imprecise. They should be read as a descriptive comparison of the two sub-periods rather than as evidence of regime-specific effects. The dependent variable is the growth rate of real GDP, so the coefficients are read in percentage points. Because ΔLGDP is multiplied by 100, these coefficients are 100 times the corresponding ARDL short-run coefficients, which are in log points: the ARDL coefficient of 0.7570 on D(RE) corresponds to 75.7 percentage points of growth of renewable energy, the same order of magnitude as the regime estimates below. The change in renewable energy has positive but insignificant coefficients in both regimes (106.7101, $p = 0.108$; 27.4406, $p = 0.378$), and their confidence intervals overlap widely: nothing in these estimates supports a sign reversal around a critical scale. The change in $CO_2$ per capita is positive and significant in both regimes (5.0274*** and 4.5269**), corresponding to about 5.03 and 4.53 percentage points of real GDP for an additional unit of $CO_2$ per capita, respectively, which is consistent with the positive association found in the ARDL. The change in the fossil fuel share is strongly negative in the first regime (−9.4318***) but insignificant in the second (2.6898). The coefficient on $CH_4$ is insignificant in both regimes. The model has an $R^2$ of 0.428, and the Jarque–Bera test on its residuals gives $p = 0.055$.

**Table 8.** Threshold regression in first differences: RE regimes on ΔLGDP.

| | Regime 1: $RE_{t-1} \leq 0.0425$ | | | | Regime 2: $RE_{t-1} > 0.0425$ | | | |
|---|---|---|---|---|---|---|---|---|
| **Variable** | **Coef.** | **Std. Err.** | **t** | **P > \|t\|** | **Coef.** | **Std. Err.** | **t** | **P > \|t\|** |
| C | 2.3515 *** | 0.7650 | 3.074 | 0.006 | 3.4877 | 2.2956 | 1.519 | 0.143 |
| ΔEC | −9.4318 *** | 0.7421 | −12.709 | 0.000 | 2.6898 | 8.8791 | 0.303 | 0.765 |
| ΔRE | 106.7101 | 63.7439 | 1.674 | 0.108 | 27.4406 | 30.4717 | 0.901 | 0.378 |
| $\Delta CO_2$ | 5.0274 *** | 1.2405 | 4.053 | 0.001 | 4.5269 ** | 2.1082 | 2.147 | 0.043 |
| $\Delta CH_4$ | −8.0099 | 10.4362 | −0.768 | 0.451 | −14.9173 | 34.9451 | −0.427 | 0.674 |

Note: *** and ** denote significance at the 1% and 5% levels, respectively. The regime-specific estimates are descriptive: the upper regime contains 8 observations for 5 parameters.

## 5. Discussion

The ARDL results indicate a positive association between renewable energy consumption and economic growth in both the short and long run over the 1990–2022 period. In the short run, renewable energy consumption is positively and significantly associated with real GDP, with a coefficient of 0.7570 ($p = 0.0063$), while its lagged change is negative and weakly significant ($-0.3954$, $p = 0.0658$). This pattern is consistent with an investment-related transmission mechanism, whereby renewable energy development may generate demand through project construction, procurement, and related infrastructure activities, followed by a partial reversal as the initial investment impulse dissipates. In the long run, renewable energy has a positive and statistically significant coefficient of 1.2665 ($p = 0.0000$). The long-run coefficient is positive and statistically significant. However, its economic magnitude should be interpreted in light of the measurement scales and functional form of the variables, as well as the limited variation in renewable energy consumption during much of the sample period. The coefficient should not, by itself, be interpreted as evidence of a proportionately large economic effect.

The short-run positive association is consistent with the investment-driven growth channel documented by Ifa and Guetat (2022), Xuan (2025c), and, for OECD economies, Ferhi and Helali (2024a). The positive long-run association further indicates that renewable energy and real GDP have moved together over the sample period, although the ARDL estimates do not by themselves establish a causal effect. The relatively limited development of renewable energy during much of the sample period and its concentration in the later years should also be considered when interpreting the estimated long-run relationship.

In the long run, $CO_2$ emissions are positively and significantly associated with real GDP, with a coefficient of 0.0602 ($p = 0.0004$). This result is consistent with a scale effect, whereby higher economic activity in a fossil-fuel-intensive economy is associated with higher emissions. The finding is consistent with evidence on oil-exporting economies reported by AlNemer et al. (2023). In the threshold regression, the change in $CO_2$ emissions is also positive and statistically significant in both renewable-energy regimes, with coefficients of 5.0274 ($p = 0.001$) and 4.5269 ($p = 0.043$), respectively. The consistency of the positive $CO_2$ association across the two regimes suggests that the relationship between economic activity and carbon emissions remains evident even as renewable-energy deployment increases. This result should nevertheless be interpreted as an association rather than evidence that emissions themselves generate economic growth. The weak exogeneity tests in Appendix A support this reading: $CO_2$ responds to deviations from the long-run relationship, so the association runs in both directions.

The greenhouse-gas results should be interpreted separately for the two gases because their estimated associations with economic growth differ. $CH_4$ emissions, by contrast, have positive and statistically significant long-run coefficients in the ARDL model (0.6603, $p = 0.0000$), whereas their short-run effect is negative and significant only at one lag ($-0.3256$, $p = 0.0129$). In the threshold regression, however, the coefficients on $\Delta CH_4$ are negative but statistically insignificant in both regimes ($-8.0099$, $p = 0.451$; $-14.9173$, $p = 0.674$). The negative short-run association should therefore be interpreted with caution, and the threshold results do not provide robust evidence of a regime-specific methane effect. These findings should also be interpreted in light of the specific $CH_4$ indicator used, which captures emissions from industrial processes rather than total methane emissions across all sectors.

Given that the share of fossil fuels in Saudi Arabia ranges from approximately 99.2% to 100% across the sample, the variable exhibits very limited variation. Its coefficient is statistically insignificant in the long-run ARDL relationship (0.0686, $p = 0.4428$), while its short-run change is also insignificant ($-0.0245$, $p = 0.4731$). In the threshold regression, $\Delta EC$ is strongly negative and statistically significant in the first regime ($-9.4318$, $p = 0.000$),

but becomes positive and statistically insignificant in the second regime (2.6898, $p = 0.765$). Given the extremely limited variation in the fossil-fuel share and the concentration of the first-regime result around a relatively large movement in the series, these coefficients should be interpreted cautiously. Overall, the results provide limited evidence that marginal movements within an almost saturated fossil-fuel mix are systematically associated with economic growth. They do not, however, indicate how a substantial structural change in the energy mix would affect growth, since such a transition is largely absent from the historical sample.

The threshold regression does not provide statistically significant evidence of nonlinearities in the renewable-energy–growth relationship. In the first-difference specification, the SSR-minimizing split occurs at 0.0425, separating the observations into a lower-renewable-energy regime covering 1991–2014 and a higher-renewable-energy regime covering 2015–2022. However, the sup-F statistic is only 1.861, and the bootstrap $p$-value is 0.943, so the null hypothesis of linearity cannot be rejected. The regime-specific renewable-energy coefficients are positive but statistically insignificant in both regimes (106.7101, $p = 0.108$, and 27.4406, $p = 0.378$). Accordingly, the estimated split should be interpreted as the SSR-minimizing sample division rather than as a statistically established critical threshold. The results do not provide sufficient statistical evidence of a threshold effect or a statistically significant difference in the renewable-energy–growth relationship across the estimated regimes.

The levels specification leads to the same conclusion. Its sup-F statistic of 12.814 has a bootstrap $p$-value of 0.483, providing no statistically significant evidence of threshold nonlinearity. In addition, the residual-based stationarity test does not provide evidence of threshold cointegration under the simulated critical values. Taken together, these findings suggest that the apparent regime differences identified by the threshold search should not be interpreted as evidence of a statistically established nonlinear relationship. The threshold analysis therefore complements the ARDL results by showing that the positive average relationship identified in the linear long-run model is not accompanied by robust evidence of a discrete threshold effect.

Taken together, the results provide support for H1, insofar as renewable energy is positively and statistically significantly associated with real GDP in both the short-run and long-run ARDL specifications. However, the results do not provide sufficient statistical evidence to support H2, as the estimated relationship does not differ significantly across the two regimes. Given the near-monotonic increase in renewable energy consumption, the estimated split also corresponds closely to two consecutive subperiods, 1991–2014 and 2015–2022. This finding differs from evidence of threshold effects reported for Chinese provinces by Song et al. (2025) and F. Liu et al. (2025), as well as from the OECD threshold estimates of Ferhi and Helali (2024b). Such differences may reflect variations in country characteristics, sample periods, renewable energy development, model specifications, and the degree of variation in the underlying data.

For the environmental controls, the threshold model identifies a positive and statistically significant association between changes in $CO_2$ emissions and changes in real GDP in both renewable-energy regimes. In contrast, the coefficients on $CH_4$ are negative but statistically insignificant. These results reinforce the distinction between the two greenhouse gases in the present specification. $CO_2$ appears to co-move positively with economic activity across both regimes, consistent with the scale effect of a fossil-fuel-intensive economy. The methane coefficients, however, do not provide sufficiently robust evidence of a systematic threshold-specific relationship with economic growth. Accordingly, the environmental results should be interpreted primarily as evidence of differing statistical associations rather than as evidence of causal effects.

From a policy perspective, the ARDL results indicate that renewable energy development is positively associated with economic activity in both the short and long run. This finding is relevant to the sequencing of renewable energy development under Vision 2030, particularly regarding grid integration, infrastructure, financing, and complementary technologies. At the same time, the absence of statistically significant threshold effects means that the estimated results do not identify a specific renewable-energy level at which the economic-growth relationship changes. The value of the threshold analysis is therefore primarily diagnostic: it indicates that the apparent regime split found by the grid search is not statistically robust. Policy interpretation should consequently focus on the broader conditions surrounding renewable-energy deployment rather than on a particular numerical threshold.

These findings are broadly consistent with the objectives of Saudi Arabia's Vision 2030, particularly the diversification of economic activity and the development of renewable energy. However, the empirical results should be interpreted as evidence of associations among renewable energy use, emissions, and economic growth rather than as evidence that renewable energy is, by itself, a causal driver of long-term sustainable growth. This distinction is particularly important given the relatively small sample of 33 annual observations and the limited variation in the fossil-fuel share.

Methodologically, the two models are complementary rather than redundant. The ARDL framework captures the dynamic time-series relationship among the variables, including short-run effects and an error-correction mechanism. In contrast, the threshold regression tests whether the renewable-energy–growth relationship differs significantly across regimes. The results from the two approaches are broadly consistent in indicating a positive association between renewable energy and economic activity. At the same time, the threshold analysis does not provide sufficient statistical evidence to conclude that this association changes discretely at a particular level of renewable-energy development.

## 6. Conclusions and Policy Implications

This study investigates whether renewable energy is associated with economic growth in Saudi Arabia and whether this association varies with the level of renewable energy. Focusing on Saudi Arabia's ongoing transition from a hydrocarbon-dependent economy toward a more diversified energy structure, the analysis examines the interactions between economic growth, renewable energy, the fossil-fuel share, $CO_2$ emissions, and $CH_4$ emissions over the period 1990–2022. To capture both long-run dynamics and potential nonlinear effects, the study combines the ARDL bounds-testing approach with a Hansen-type threshold regression, using renewable energy consumption as the threshold variable.

The empirical results indicate that renewable energy is positively associated with economic growth in the linear ARDL framework, while the evidence does not support a statistically significant nonlinear relationship. The ARDL bounds test provides evidence of a long-run relationship among the variables. At the same time, the estimated error-correction coefficient of −0.5273 indicates that approximately 52.7% of short-run deviations from the long-run equilibrium are corrected within one year. In the short run, renewable energy is positively and statistically significantly associated with economic growth, while its long-run coefficient is also positive and statistically significant. These results indicate a positive association between renewable energy and economic growth across both time horizons. However, the magnitude of the long-run coefficient should be interpreted with caution given the limited development of renewable energy during much of the sample period.

The threshold analysis does not provide statistically significant evidence that the relationship between renewable energy and growth changes at any specific level of renewable energy use. The SSR-minimizing split of 0.0425 is not statistically established as a thresh-

old because the wild-bootstrap test does not provide sufficient evidence to reject the null hypothesis of no threshold effect ($p = 0.943$).

The results further reveal distinct associations between greenhouse-gas emissions and economic growth. $CO_2$ emissions are positively and statistically significantly associated with economic growth in the long-run ARDL relationship and remain positively associated with GDP growth in both threshold regimes. This indicates that the relationship between economic activity and carbon emissions remains evident even as renewable-energy use increases. By contrast, $CH_4$ emissions have a positive and statistically significant long-run coefficient, while their lagged short-run coefficient is negative and statistically significant. However, the $CH_4$ coefficients are not statistically significant in either threshold regime. The estimated associations between $CH_4$ emissions and economic growth differ across time horizons, although the threshold estimates do not provide statistically significant evidence of a regime-specific relationship.

### *6.1. Policy Recommendations and Implications*

The findings provide several implications for energy and environmental policy in Saudi Arabia. The positive short- and long-run association between renewable energy and economic growth is consistent with continued development of renewable energy as part of the country's broader economic and energy diversification process. However, the absence of statistically significant threshold effects means that the results do not identify a particular renewable-energy level at which the economic relationship changes. Policy decisions should therefore not be based on the estimated SSR-minimizing value as a specific deployment target.

At the same time, the positive $CO_2$ association indicates that renewable-energy expansion alone has not eliminated the relationship between economic activity and carbon emissions over the study period. In the broader context of Saudi Arabia's energy transition, strengthening grid integration, expanding storage capacity, and continuing implementation of the National Renewable Energy Program may support renewable energy deployment. These measures are policy considerations rather than effects directly estimated by the present model. The methane results also suggest the importance of considering greenhouse gases separately, given the different statistical patterns observed across the short- and long-run specifications.

### *6.2. Limitations and Future Research*

This study has several limitations that should be considered when interpreting the results. The relatively small annual sample limits the precision of regime-specific estimates, while the single-country design restricts the generalizability of the findings beyond Saudi Arabia. The limited variation in the fossil-fuel share and the relatively recent development of renewable energy may also constrain the detection of nonlinear effects. Furthermore, the ARDL specification includes variables that may adjust to deviations from the long-run relationship, limiting a strictly one-directional interpretation of the estimated relationships. Physical capital, labour, and oil-market variables are not included, structural breaks are not modelled explicitly, and the threshold specification does not incorporate an error-correction term. Future research could address these limitations by adopting a production-function specification that includes capital and labour, incorporating oil-market controls, applying unit-root tests robust to structural breaks, and developing a threshold error-correction model. These extensions could be explored when longer or higher-frequency data become available.

**Author Contributions:** Conceptualization, B.H. and A.F.; methodology, B.H.; software, A.L.; validation, B.H., A.F., A.L. and O.G.; formal analysis, B.H.; investigation, A.F.; resources, A.L.; data curation, O.G.; writing—original draft preparation, B.H.; writing—review and editing, B.H.; visualization, A.F.; supervision, O.G.; project administration, B.H.; funding acquisition, B.H. All authors have read and agreed to the published version of the manuscript.

**Funding:** The Deanship of Scientific Research at the University of Ha'il, Saudi Arabia (Project No. RG-23 140) supported this research.

**Institutional Review Board Statement:** Not applicable.

**Informed Consent Statement:** Not applicable.

**Data Availability Statement:** The data used in this study were obtained from the World Bank World Development Indicators (WDI) database and are publicly available. Further details regarding the data and data processing steps are available from the authors upon reasonable request.

**Acknowledgments:** This research has been funded by the Scientific Research Deanship at the University of Ha'il—Saudi Arabia through project number «RG-23 140».

**Conflicts of Interest:** The authors declare no conflicts of interest.

## Appendix A

**Table A1.** Tests of the long-run forcing assumption.

**Panel A. Toda–Yamamoto Granger Non-Causality Tests, VAR(2) in Levels ($n$ = 31)**

| | H0: x does not Granger-cause LGDP | | | H0: LGDP does not Granger-cause x | | |
|---|---|---|---|---|---|---|
| x | Wald $\chi^2(1)$ | $p$-value | Decision | Wald $\chi^2(1)$ | $p$-value | Decision |
| EC | 0.858 | 0.354 | Not rejected | 0.633 | 0.426 | Not rejected |
| RE | 0.310 | 0.578 | Not rejected | 0.860 | 0.354 | Not rejected |
| $CO_2$ | 0.276 | 0.599 | Not rejected | 0.641 | 0.423 | Not rejected |
| $CH_4$ | 0.004 | 0.949 | Not rejected | 1.355 | 0.244 | Not rejected |

**Panel B. Weak Exogeneity (Long-Run Forcing): Loading on the Lagged Equilibrium Error**

| | Single-equation test (n = 31) | | | Johansen VECM (n = 32) | | |
|---|---|---|---|---|---|---|
| Equation | Loading | t | $p$-value | Loading | z | $p$-value |
| ΔLGDP | −0.5258 | −2.850 | 0.009 | −0.2951 | −2.516 | 0.012 |
| ΔEC | 1.3491 | 1.287 | 0.210 | 2.2614 | 4.541 | 0.000 |
| ΔRE | 0.3165 | 2.046 | 0.052 | 0.2765 | 3.110 | 0.002 |
| $\Delta CO_2$ | −6.4001 | −2.223 | 0.036 | −3.4525 | −2.091 | 0.036 |
| $\Delta CH_4$ | 0.3603 | 1.276 | 0.214 | 0.1117 | 0.672 | 0.502 |

**Table A1.** *Cont.*

| Panel C. Johansen Cointegration Tests and Normalized Cointegrating Vector | | | | | | |
|---|---|---|---|---|---|---|
| | | H0: r = 0 | | | H0: r ≤ 1 | |
| Test | Statistic | 5% c.v. | Decision | Statistic | 5% c.v. | Decision |
| Trace | 70.382 | 69.819 | Rejected | 31.231 | 47.854 | Not rejected |
| Maximum eigenvalue | 39.151 | 33.878 | Rejected | 16.410 | 27.586 | Not rejected |
| Long-run coefficient | EC | RE | $CO_2$ | $CH_4$ | | |
| Johansen | 0.1883 | 1.1577 | 0.0788 | 0.5798 | | |
| | (0.0408) | (0.0842) | (0.0072) | (0.0598) | | |
| ARDL (Table 6) | 0.0686 | 1.2665 | 0.0602 | 0.6603 | | |
| | (0.0875) | (0.1588) | (0.0140) | (0.1253) | | |

Note: Panel A: Wald tests on the first lag of the causing variable in a VAR in levels with k + dmax = 1 + 1 lags and an intercept (Toda & Yamamoto, 1995), where k = 1 is the lag order selected in Table 4; 20 residual degrees of freedom per equation. Panel B: single-equation test, Δx regressed on an intercept, the lagged equilibrium error computed from the long-run coefficients in Table 6, and one lag of the first differences in the five variables (24 residual degrees of freedom); Johansen VECM, VAR(1) in levels (Table 4), unrestricted intercept, one cointegrating vector. A significant loading rejects weak exogeneity, that is, long-run forcing (Johansen, 1992). Panel C: asymptotic 5% critical values (Johansen, 1991). With the small-sample correction of Reinsel and Ahn (1992), the statistics for r = 0 become 59.38 (trace) and 33.03 (maximum eigenvalue), the latter significant at 10% (critical value 31.24). Long-run coefficients normalized on LGDP, standard errors in parentheses.

**Table A2.** Levels threshold specification, Equation (6): complete regime-specific results.

| | **Regime 1: $RE_{t-1} \leq 0.0006$ (1991–2010, n = 20)** | | | | **Regime 2: $RE_{t-1} > 0.0006$ (2011–2022, n = 12)** | | | |
|---|---|---|---|---|---|---|---|---|
| **Variable** | **Coef.** | **Std. Err.** | **t** | ***p*** | **Coef.** | **Std. Err.** | **t** | ***p*** |
| C | −283.7575 | 222.6900 | −1.274 | 0.222 | 18.3179 | 6.5530 | 2.795 | 0.027 |
| EC | 2.8746 | 2.2267 | 1.291 | 0.216 | −0.1412 | 0.0657 | −2.149 | 0.069 |
| RE | −31.6653 | 9.7146 | −3.260 | 0.005 | 1.4725 | 0.1284 | 11.467 | 0.000 |
| $CO_2$ | 0.0897 | 0.0125 | 7.203 | 0.000 | 0.1005 | 0.0152 | 6.625 | 0.000 |
| $CH_4$ | 0.4717 | 0.0889 | 5.304 | 0.000 | −0.6509 | 0.4693 | −1.387 | 0.208 |

Note: Dependent variable LGDP; threshold $\gamma = 0.0006$, 95% likelihood-ratio confidence set [0.0000; 0.0425], that is, the whole admissible grid; SSR = 0.029; sup-F = 12.814 (bootstrap $p = 0.483$); residual ADF = −5.031 (simulated 5% critical value −6.688). Because the regimes are consecutive sub-periods, the coefficients are estimated by FMOLS (P. C. Phillips & Hansen, 1990) within each regime, with long-run covariances computed within each regime (Bartlett kernel, bandwidth of 2); *p*-values are from the t-distribution with 15 and 7 degrees of freedom. The residual unit root is not rejected (Table 7), so these coefficients have no long-run interpretation, and their t-statistics are not valid in the usual sense; therefore, no significance levels are indicated. In regime 1, RE takes only three values (0, 0.00061, and 0.00409).

## References

Addis, A. K., & Cheng, S. (2023). The nexus between renewable energy, environmental pollution, and economic growth across BRICS and OECD countries: A comparative empirical study. *Energy Reports*, *10*, 3800–3813. [CrossRef]

Addo, P. M., Manibialoa, C., & McIsaac, F. (2021). Exploring nonlinearity on the $CO_2$ emissions, economic production and energy use nexus: A causal discovery approach. *Energy Reports*, *7*, 6196–6204. [CrossRef]

Alao, R. O., Olanipekun, I. O., Sharimakin, A., & Olanipekun, T. P. (2025). Breaking down the barriers to a green Africa: Detangling the overlapping issues of fossil fuel energy consumption, resource richness, and economic growth. *Resources Policy*, *109*, 105720. [CrossRef]

AlNemer, H. A., Hkiri, B., & Tissaoui, K. (2023). Dynamic impact of renewable and non-renewable energy consumption on $CO_2$ emission and economic growth in Saudi Arabia: Fresh evidence from wavelet coherence analysis. *Renewable Energy*, *209*, 340–356. [CrossRef]

Alqaralleh, H. (2021). On the nexus of $CO_2$ emissions and renewable and nonrenewable energy consumption in Europe: A new insight from panel smooth transition. *Energy & Environment*, *32*(3), 443–457. [CrossRef]

Alshehry, A., & Belloumi, M. (2024). The symmetric and asymmetric impacts of energy consumption and economic growth on environmental sustainability. *Sustainability*, *16*(1), 205. [CrossRef]

Amri, F. (2018). Carbon dioxide emissions, total factor productivity, ICT, trade, financial development, and energy consumption: Testing environmental Kuznets curve hypothesis for Tunisia. *Environmental Science and Pollution Research*, *25*(33), 33691–33701. [CrossRef] [PubMed]

Anser, M. K., Shabbir, M. S., Tabash, M. I., Shah, S. H. A., Ahmad, M., Peng, M. Y.-P., & Lopez, L. B. (2021). Do renewable energy sources improve clean environmental-economic growth ? Empirical investigation from South Asian economies. *Energy Exploration & Exploitation*, *39*(5), 1491–1514. [CrossRef]

Apergis, N., & Payne, J. E. (2010). Renewable energy consumption and economic growth: Evidence from a panel of OECD countries. *Energy Policy*, *38*(1), 656–660. [CrossRef]

Balke, N. S., & Fomby, T. B. (1997). Threshold cointegration. *International Economic Review*, *38*, 627–645. [CrossRef]

Baz, K., Cheng, J., Xu, D., Abbas, K., Ali, I., Ali, H., & Fang, C. (2021). Asymmetric impact of fossil fuel and renewable energy consumption on economic growth: A nonlinear technique. *Energy*, *226*, 120357. [CrossRef]

Bekhet, H. A., Matar, A., & Yasmin, T. (2017). $CO_2$ emissions, energy consumption, economic growth, and financial development in GCC countries: Dynamic simultaneous equation models. *Renewable and Sustainable Energy Reviews*, *70*, 117–132. [CrossRef]

Belaïd, F., & Zrelli, M. H. (2019). Renewable and non-renewable electricity consumption, environmental degradation and economic development: Evidence from Mediterranean countries. *Energy Policy*, *133*, 110929. [CrossRef]

Brown, R. L., Durbin, J., & Evans, J. M. (1975). Techniques for testing the constancy of regression relationships over time. *Journal of the Royal Statistical Society: Series B (Methodological)*, *37*(2), 149–192. [CrossRef]

Chen, C., Pinar, M., & Stengos, T. (2020). Renewable energy consumption and economic growth nexus: Evidence from a threshold model. *Energy Policy*, *139*, 111295. [CrossRef]

Chen, Q., & Zhang, E. (2025). Economic growth, energy consumption and $CO_2$ emissions: A replication. *Energy Economics*, *147*, 108538. [CrossRef]

Destek, M. A., & Sinha, A. (2020). Renewable, non-renewable energy consumption, economic growth, trade openness and ecological footprint: Evidence from organisation for economic Co-operation and development countries. *Journal of Cleaner Production*, *242*, 118537. [CrossRef]

Destiartono, M. E., & Firmansyah. (2024). Does the linkage between GDP, renewable energy, and methane validate the EKC hypothesis? Evidence from Indonesia. *International Journal of Energy Economics and Policy*, *14*(2), 711–720. [CrossRef]

Dickey, D. A., & Fuller, W. A. (1979). Distribution of the estimators for autoregressive time series with a unit root. *Journal of the American Statistical Association*, *74*(366a), 427–431. [CrossRef]

Djoukouo, A. F. D. (2021). Relationship between methane emissions and economic growth in Central Africa countries: Evidence from panel data. *Global Transitions*, *3*, 126–134. [CrossRef]

Dong, F., Qin, C., Zhang, X., Zhao, X., Pan, Y., Gao, Y., Zhu, J., & Li, Y. (2021). Towards carbon neutrality: The impact of renewable energy development on carbon emission efficiency. *International Journal of Environmental Research and Public Health*, *18*(24), 13284. [CrossRef] [PubMed]

Elder, J., & Kennedy, P. E. (2001). Testing for unit roots: What should students be taught? *The Journal of Economic Education*, *32*(2), 137–146. [CrossRef]

Ferhi, A., & Helali, K. (2024a). The impact of renewable energy on the environment and socio-economic welfare: Empirical evidence from OECD countries. *Journal of the Knowledge Economy*, *15*(1), 4776–4799. [CrossRef]

Ferhi, A., & Helali, K. (2024b). The nonlinear nexus between renewable energy consumption and economic growth in OECD countries: Evidence from PTR and PSTR estimates. *Journal of Economic Studies*, *51*(1), 185–201. [CrossRef]

Gonzalo, J., & Pitarakis, J. Y. (2006). Threshold effects in cointegrating relationships. *Oxford Bulletin of Economics and Statistics*, *68*, 813–833. [CrossRef]

González-Álvarez, M. A., & Montañés, A. (2023). $CO_2$ emissions, energy consumption, and economic growth: Determining the stability of the 3E relationship. *Economic Modelling*, *121*, 106195. [CrossRef]

Gregory, A. W., & Hansen, B. E. (1996). Residual-based tests for cointegration in models with regime shifts. *Journal of Econometrics*, *70*(1), 99–126. [CrossRef]

Grossman, G. M., & Krueger, A. B. (1995). Economic growth and the environment. *The Quarterly Journal of Economics*, *110*(2), 353–377. [CrossRef]

Guliyev, H., & Tatoğlu, F. Y. (2023). The relationship between renewable energy and economic growth in European countries: Evidence from panel data model with sharp and smooth changes. *Renewable Energy Focus*, *46*, 185–196. [CrossRef]

Hamdi, B., Ferhi, A., Gammoudi, O., & Louhichi, A. (2026). Unveiling the nonlinear dynamics of renewable energy and economic growth in Saudi Arabia: A combined threshold and quantile regression analysis. *Economies*, *14*(8), 305. [CrossRef]

Hansen, B. E. (1996). Inference when a nuisance parameter is not identified under the null hypothesis. *Econometrica: Journal of the Econometric Society*, *64*, 413–430. [CrossRef]

Hansen, B. E. (2000). Sample splitting and threshold estimation. *Econometrica*, *68*(3), 575–603. [CrossRef]

Hansen, B. E., & Seo, B. (2002). Testing for two-regime threshold cointegration in vector error-correction models. *Journal of Econometrics*, *110*(2), 293–318. [CrossRef]

Ifa, A., & Guetat, I. (2022). Analysing short-run and long-run causality relationship among public spending, renewable energy consumption, non-renewable energy consumption and economic growth: Evidence from eight of South Mediterranean countries. *Energy Exploration & Exploitation*, *40*(2), 554–579. [CrossRef]

Intergovernmental Panel on Climate Change. (2022). Climate change 2022: Mitigation of climate change. In *Contribution of working group III to the sixth assessment report of the IPCC*. Cambridge University Press. [CrossRef]

International Energy Agency. (2023). *World energy outlook 2023*. IEA. Available online: https://www.iea.org/reports/world-energy-outlook-2023 (accessed on 2 August 2026).

Ivanovski, K., Hailemariam, A., & Smyth, R. (2021). The effect of renewable and non-renewable energy consumption on economic growth: Non-parametric evidence. *Journal of Cleaner Production*, *286*, 124956. [CrossRef]

Jiang, S., Tan, X., Hu, P., Wang, Y., Shi, L., Ma, Z., & Lu, G. (2022). Air pollution and economic growth under local government competition: Evidence from China, 2007–2016. *Journal of Cleaner Production*, *334*, 130231. [CrossRef]

Jiang, Z., Mahmud, A. R., Maneengam, A., Nassani, A. A., Haffar, M., & Cong, P. T. (2022). Non linear effect of Biomass, fossil fuels and renewable energy usage on the economic Growth: Managing sustainable development through energy sector. *Fuel*, *326*, 124943. [CrossRef]

Johansen, S. (1991). Estimation and hypothesis testing of cointegration vectors in Gaussian vector autoregressive models. *Econometrica*, *59*(6), 1551–1580. [CrossRef]

Johansen, S. (1992). Cointegration in partial systems and the efficiency of single-equation analysis. *Journal of Econometrics*, *52*(3), 389–402. [CrossRef]

Kraft, J., & Kraft, A. (1978). On the relationship between energy and GNP. *The Journal of Energy and Development*, *3*, 401–403.

Kukla-Gryz, A. (2009). Economic growth, international trade and air pollution: A decomposition analysis. *Ecological Economics*, *68*(5), 1329–1339. [CrossRef]

Li, B., & Haneklaus, N. (2021). The role of renewable energy, fossil fuel consumption, urbanization and economic growth on $CO_2$ emissions in China. *Energy Reports*, *7*, 783–791. [CrossRef]

Liu, F., Zhou, D., Ding, H., & Wang, Q. (2025). Inhibit or promote: Nonlinear spatial impacts of renewable energy on economic growth in China. *Energy*, *335*, 138108. [CrossRef]

Liu, G., Ofori, C., Ampong, S. A., Appiah-Twum, F., & Alhassan, E. A. (2023). Towards a sustainable environment: Examining the spatial VARIATIONS of renewable energy, environmental pollution, and economic growth in Europe. *Energy Strategy Reviews*, *50*, 101231. [CrossRef]

MacKinnon, J. G. (1996). Numerical distribution functions for unit root and cointegration tests. *Journal of Applied Econometrics*, *11*(6), 601–618. [CrossRef]

McNown, R., Sam, C. Y., & Goh, S. K. (2018). Bootstrapping the autoregressive distributed lag test for cointegration. *Applied Economics*, *50*(13), 1509–1521. [CrossRef]

Meadows, D., Randers, J., & Meadows, D. (2004). *Limits to growth: The 30-year update*. Chelsea Green Publishing.

Mensah, I. A., Sun, M., Gao, C., Omari-Sasu, A. Y., Zhu, D., Ampimah, B. C., & Quarcoo, A. (2019). Analysis on the nexus of economic growth, fossil fuel energy consumption, $CO_2$ emissions and oil price in Africa based on a PMG panel ARDL approach. *Journal of Cleaner Production*, *228*, 161–174. [CrossRef]

Mighri, Z., & AlSaggaf, M. I. (2023). Asymmetric impacts of renewable energy consumption and economic complexity on economic growth in Saudi Arabia: Evidence from the NARDL model. *Environmental Science and Pollution Research*, *30*(3), 7446–7473. [CrossRef] [PubMed]

Mujtaba, A., Jena, P. K., Bekun, F. V., & Sahu, P. K. (2022). Symmetric and asymmetric impact of economic growth, capital formation, renewable and non-renewable energy consumption on environment in OECD countries. *Renewable and Sustainable Energy Reviews*, *160*, 112300. [CrossRef]

Narayan, P. K. (2005). The saving and investment nexus for China: Evidence from cointegration tests. *Applied Economics*, *37*(17), 1979–1990. [CrossRef]

Oprea, S.-V., Bâra, A., & Georgescu, I. A. (2025). Exploring the relationship between $CO_2$ emissions, economic growth, urbanization, renewables and foreign investment in Iceland. *Energy Strategy Reviews*, *61*, 101835. [CrossRef]

Ozcan, B., & Ozturk, I. (2019). Renewable energy consumption-economic growth nexus in emerging countries: A bootstrap panel causality test. *Renewable and Sustainable Energy Reviews*, *104*, 30–37. [CrossRef]

Perron, P. (1989). The great crash, the oil price shock, and the unit root hypothesis. *Econometrica*, *57*(6), 1361–1401. [CrossRef]

Pesaran, M. H., & Shin, Y. (1995). *An autoregressive distributed lag modelling approach to cointegration analysis* (Vol. 9514). Department of Applied Economics, University of Cambridge. Available online: https://EconPapers.repec.org/RePEc:cam:camdae:9514 (accessed on 25 October 2026).

Pesaran, M. H., Shin, Y., & Smith, R. J. (2001). Bounds testing approaches to the analysis of level relationships. *Journal of Applied Econometrics*, *16*(3), 289–326. [CrossRef]

Phillips, P. C., & Hansen, B. E. (1990). Statistical inference in instrumental variables regression with I(1) processes. *The Review of Economic Studies*, *57*(1), 99–125. [CrossRef]

Phillips, P. C. B., & Perron, P. (1988). Testing for a unit root in time series regression. *Biometrika*, *75*(2), 335–346. [CrossRef]

Rahman, M. M., & Velayutham, E. (2020). Renewable and non-renewable energy consumption-economic growth nexus: New evidence from South Asia. *Renewable Energy*, *147*, 399–408. [CrossRef]

Raifu, I. A., Obaniyi, F. A., Nnamani, G., & Salihu, A. A. (2025). Revisiting causal relationship between renewable energy and economic growth in OECD countries: Evidence from a novel JKS's Granger non-causality test. *Renewable Energy*, *244*, 122559. [CrossRef]

Raza, A., Alharthi, M., Shakeel, M., & Khan, M. I. (2025). Nexus between $CO_2$ emissions, economic growth, and energy consumption with special reference to BRICS: A panel data approach. *Energy Nexus*, *20*, 100556. [CrossRef]

Reinsel, G. C., & Ahn, S. K. (1992). Vector autoregressive models with unit roots and reduced rank structure: Estimation, likelihood ratio test, and forecasting. *Journal of Time Series Analysis*, *13*(4), 353–375. [CrossRef]

Shahbaz, M., Tiwari, A. K., & Nasir, M. (2013). The effects of financial development, economic growth, coal consumption and trade openness on $CO_2$ emissions in South Africa. *Energy Policy*, *61*, 1452–1459. [CrossRef]

Solow, R. M. (1974). Intergenerational equity and exhaustible resources. *The Review of Economic Studies*, *41*(5), 29–45. [CrossRef]

Song, D., Hu, Y., Zhang, Q., Wang, S., & Bi, C. (2025). Harnessing AI for renewable energy transition: Threshold effects on China's economic growth. *Sustainable Futures*, *10*, 101349. [CrossRef]

Stiglitz, J. E. (1974). Growth with exhaustible natural resources: Efficient and optimal growth paths. *The Review of Economic Studies*, *41*(5), 123–137. [CrossRef]

Thach, N. N. (2025). Which, renewable or non-renewable energy, is more vital for the economic growth of OECD countries ? A Bayesian hierarchical analysis. *Heliyon*, *11*(3), e42259. [CrossRef] [PubMed]

Tlili, H., Alhamad, S., & Turki, H. (2025). How growth, urbanization, and energy consumption affect $CO_2$ emissions in Saudi Arabia (1970–2020)? An ARDL and NARDL approach to investigate the eco-environmental challenge. *Energy Strategy Reviews*, *59*, 101691. [CrossRef]

Toda, H. Y., & Yamamoto, T. (1995). Statistical inference in vector autoregressions with possibly integrated processes. *Journal of Econometrics*, *66*(1–2), 225–250. [CrossRef]

Usman, M., Makhdum, M. S. A., & Kousar, R. (2021). Does financial inclusion, renewable and non-renewable energy utilization accelerate ecological footprints and economic growth ? Fresh evidence from 15 highest emitting countries. *Sustainable Cities and Society*, *65*, 102590. [CrossRef]

Waheed, R., Sarwar, S., & Wei, C. (2019). The survey of economic growth, energy consumption and carbon emission. *Energy Reports*, *5*, 1103–1115. [CrossRef]

Wang, Q., & Wang, L. (2020). Renewable energy consumption and economic growth in OECD countries: A nonlinear panel data analysis. *Energy*, *207*, 118200. [CrossRef]

Wen, B., He, Y., Jing, X., & Haroon, M. (2025). Advancing renewable energy and green finance for economic growth and ecological resilience. *Energy Strategy Reviews*, *59*, 101747. [CrossRef]

Xuan, V. N. (2025a). An ARDL approach to investigating the relationship between FDI, renewable energy, economic growth, trade openness, and $CO_2$ emissions in Australia. *Results in Engineering*, *27*, 105668. [CrossRef]

Xuan, V. N. (2025b). Nexus of innovation, renewable energy, FDI, trade openness, and economic growth in Germany: New insights from ARDL method. *Renewable Energy*, *247*, 123060. [CrossRef]

Xuan, V. N. (2025c). Toward a sustainable future: Determinants of renewable energy utilisation in Canada. *Energy Reports*, *13*, 1308–1320. [CrossRef]

Youssef, A. B., Hammoudeh, S., & Omri, A. (2016). Simultaneity modeling analysis of the environmental Kuznets curve hypothesis. *Energy Economics*, *60*, 266–274. [CrossRef]

Yusuf, A. M., Abubakar, A. B., & Mamman, S. O. (2020). Relationship between greenhouse gas emission, energy consumption, and economic growth: Evidence from some selected oil-producing African countries. *Environmental Science and Pollution Research*, *27*(13), 15815–15823. [CrossRef] [PubMed]

Zivot, E., & Andrews, D. W. K. (1992). Further evidence on the great crash, the oil-price shock, and the unit-root hypothesis. *Journal of Business & Economic Statistics*, *10*(3), 251–270. [CrossRef]

Zou, G., & Chau, K. W. (2023). Energy consumption, economic growth and environmental sustainability: Evidence from China. *Energy Reports*, *9*, 106–116. [CrossRef]

**Disclaimer/Publisher's Note:** The statements, opinions and data contained in all publications are solely those of the individual author(s) and contributor(s) and not of MDPI and/or the editor(s). MDPI and/or the editor(s) disclaim responsibility for any injury to people or property resulting from any ideas, methods, instructions or products referred to in the content.