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# The roles of renewable energy, FDI, trade openness, and governance in environmental sustainability: Asymmetric and threshold evidence from Colombia

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## ABSTRACT

The study aims to address a significant research gap in measuring environmental sustainability by moving beyond traditional indicators such as carbon emissions and ecological footprint. Instead, it employs the load capacity factor (LCF) and its inverse (ILCF). Most existing research has overlooked nonlinear dynamics, threshold effects, and governance moderation when examining renewable energy, foreign direct investment (FDI), and trade openness in sustainability. Few studies have analyzed the relationship between renewable energy and these two elements within this context. The paper focuses on three key research questions based on the extended STIRPAT model and the Environmental Kuznets Curve (EKC): (i) What is the relationship between income growth, renewable energy, trade openness, FDI, and the long-run dynamics of ecological balance? (ii) Does the EKC hypothesis hold when sustainability is measured using LCF and ILCF? (iii) What are the institutional conditions, innovations, and institutional consequences that modify these impacts? Using Colombian data from 1990 to 2023, the study finds that renewable energy consistently enhances sustainability. Ecological balance is supported by FDI and resource rents, while it is negatively affected by deforestation and energy density. The analysis employs ARDL, Fourier-ARDL, NARDL, threshold regression, and causality tests. Results support the EKC hypothesis and reveal that governance acts as a threshold variable, either promoting or inhibiting the effects of openness and investment on sustainability. The integration of nonlinear dynamical analysis, governance prerequisites, and robustness tests provides groundbreaking empirical data for Colombia and offers operational suggestions for other resource-dependent economies.

**Contribution/ Originality:** This study advances environmental sustainability research by applying the load capacity factor and its inverse instead of emission-based indicators. It integrates renewable energy, FDI, trade openness, and governance within nonlinear and threshold frameworks. Using Colombia as evidence, it reveals asymmetric effects and governance thresholds with clear policy relevance for emerging and Asian economies.

## 1. INTRODUCTION

Colombia is known for its remarkable biological diversity and large resource base, but it also bears a heavy environmental burden from its diverse socioeconomic activities. These forces are reflected in deforestation, reduced water availability and depletion, and biodiversity loss, harming ecological health and human well-being. Agriculture,

logging, and mining continue to be key factors in deforestation [1]. The reduction of forest area not only reduces forests' capacity to sequester carbon but also increases soil erosion and disrupts the hydrological cycle. Illegal mining, especially for gold, leads to deforestation and the release of toxic pollutants, such as mercury, into water, harming marine life and human health. Threat of water pollution: Water resources in Colombia are under increasing pressure from industrial and agricultural runoff, as well as unsustainable water management [2]. The Magdalena River, an essential life-sustaining river, is under pressure from human activities, which in turn are causing habitat destruction and biodiversity loss [3]. The biodiversity of Colombia's natural patrimony is endangered by habitat transformation, climate change, and illegal trade in species [4]. Efficient strategies for biodiversity conservation require an integrative perspective on ecological and socio-economic problems. To address environmental decline, Colombia needs a shift toward environmental sustainability, which demands policy stimulus, technological reform, and public participation [5]. Solar, wind, and hydro power are promising, environmentally friendly renewable energy sources. Sustainable farming practices, such as reduced use of chemical products and effective irrigation, will decrease ecological pressure and improve food security. Additionally, encouraging awareness of the environment and responsible consumption behaviors is one of the main strategies for future sustainability [6, 7].

Current literature on environmental sustainability is more likely to focus on pollutant-based measures, e.g., carbon dioxide emissions and ecological footprints. Although these indicators are certainly handy, they do not reflect the entire range of natural resource consumption and environmental stressors [8, 9]. The load-capacity factor (LCF) has gained popularity over the past years as a unified indicator that compares biocapacity with ecological footprints and, hence, provides an integrated perspective on environmental sustainability [10, 11]. This change of the methodological frame is attributable to the fact that there has been an emerging consensus that sustainability assessments should not just focus on emission quantities but also on ecosystems' capacity to absorb human activity and associated forces [12, 13]. Still, despite the salient findings of these studies, a significant research gap remains on the influence of nonlinear effects of key economic determinants, that is, consumption on renewable energy, foreign direct investment (FDI), and openness of trade, on environmental sustainability in the form of the load-capacity curve (LCC) hypothesis.

There is abundant literature on the role of renewable energy in alleviating environmental degradation. Kirikkaleli and Adebayo [14] emphasize the radical nature of the renewable energy transition to reduce pollutant emissions and promote sustainability. Nevertheless, most of these questions are rooted in linear paradigms that do not account for thresholds and nonlinear processes in complex economic-environmental systems. This may be supported by emerging empirical studies that apply the LCC hypothesis, indicating that gradual renewable energy use may not directly translate into environmental improvement; payoffs occur only when particular conditions of economic development or institutional quality are attained. This connotation highlights the importance of expert analysis of the role of renewable energy in Colombia, an emerging economy with unique energy patterns, policy structures, and market dynamics. Simultaneously, FDI and trade openness studies outline the facilitative and disastrous environmental impacts. According to other authors, FDI can be used to bring more polluting technologies and knowledge base to the environment, thus promoting sustainability, as powerful environmental policies are in place [15, 16]. However, other studies argue that in the absence of strong institutional controls, FDI can further aggravate environmental degradation in line with the pollution haven hypothesis [17]. Trade openness also produces both positive and negative impacts; on the one hand, it is associated with economic development and the spread of environmentally harmless technologies, but, on the other hand, it may also increase energy consumption and environmental stress when the trade structure is biased in favor of polluting sectors [18, 19]. However, these variable interactions have been widely modeled in developed settings and in emerging economies, and little is known about their nonlinearities and threshold effects in Colombia. The world economy is essentially based on the scarcity of natural resources and energy, which subjects the ecological balance to unrelenting pressure. Especially rapidly developing economies are often expanding by exploiting more resources, industrializing, and urbanizing, which change land use, increase

energy requirements, and intensify ecological footprints. This is especially sharp in Colombia, where export revenues are tied to the extraction of natural resources and, at the same time, the country wants to expand its industrial production and attract foreign investment. These pressures have raised alarming questions about whether economic growth and openness can go hand in hand with ecological sustainability.

The current research question is driven by an irresistible urge to identify how FDI, openness to trade, the adoption of renewable energy, and governance intersect to guide Colombia toward ecological balance. Both the load-capacity factor (LCF) and its inverse (ILCF) provide a sound basis for evaluating sustainability by contrasting biocapacity shortages with ecological footprints. These measures offer a more balanced evaluation than traditional specific measures, such as carbon emissions.

The study had a two-fold purpose. First, it measured the overall effect of economic, institutional, and structural determinants on the relationship between LCF and ILCF over time. Second, it tested the EKC hypothesis in Colombia using these broad sustainability indicators, exploring whether incremental income first increases and then restores ecological balance, and whether the energy mix, innovation, and governance are critical to this process.

Three questions were central in the research, they were:

*1. How far are income growth, foreign investment, trade openness, and renewable energy leading to lasting changes in Colombia's ecological balance, as reflected in LCF and ILCF?*

*2. Does the hypothesis of EKC hold when sustainability is determined using LCF and ILCF instead of carbon emission indicators?*

*3. What role do institutional quality, innovation, and structural framework play in defining the influence of economic growth and openness on ecological sustainability?*

The study has various impacts on the sustainability of environmental studies. First of all, it enhances the empirical instrument by shifting away from a single emphasis on carbon emissions toward a two-indicator paradigm that juxtaposes bio-capacity and ecological footprints. Such an approach to a methodological shift provides a better overall understanding of environmental performance and enables direct testing of sustainability limits. Second, the research has identified Colombia as a case study for LCF- and ILCF-based scrutiny, in response to a geographical knowledge gap in Latin American environmental studies, which have traditionally preferred Brazil, Mexico, or regional panels. Third, it consolidates the effects of foreign direct investment, openness to trade, green energy, green innovation, and governance through an integrated theoretical framework, thereby providing spatial impetus to moderate the relationship between openness to investment and environmental performance. Fourth, the application of nonlinear methods of analysis, such as the nonlinear ARDL and threshold regression, reveals asymmetric and thresholded impacts that a linear model would hide, showing how different shocks to investment, trade, and renewable energy trigger distinct environmental impacts. Fifth, methodological rigor is increased through a combination of techniques: structural break diagnostics, Fourier-based ARDL, and frequency-domain causality tests, which are complemented by the control-function method and Svar robustness tests to reduce endogeneity concerns and ensure the stability of analyses across various estimation schemes. Sixth, the analytical strength is enhanced by reformulating sustainability results using alternative measures based on adjusted net savings and material footprints, and by approximating modernization through green patents and renewable capacity, supporting the component strength and formalizability of the apex results. The study shows that policy modulating levers can be explained by the importance of FDI, trade, renewable energy, innovation, and governance in enhancing Colombia's ecological balance; thus, the policy giant levers include the introduction of governance thresholds, incentives for renewable energy, diminished resource rents, and consideration of environmental terms in trade agreements. These results provide a practical roadmap to policymakers who are willing to balance the development of the economy with the environment.

## 2. LITERATURE SURVEY

### *2.1. Renewable Energy and Environmental Sustainability*

The correlation between renewable consumption (REC) and environmental sustainability has become a major research agenda in current environmental economics. The nexus has been subject to extensive research, in which scholars have analyzed its place within a wide range of national settings through numerous different methodological approaches, where it has continued to be confirmed as advantageous in the minimization of the negative impact on the environment [20-22]. According to the existing scholarly connection, there are high levels of carbon dioxide ($CO_2$), and ecological footprint reduction through renewable energy use, which proves to be one of the foundations of sustainable development strategies [23, 24].

Panel data analyses across different groups of countries consistently confirm that REC produces substantial reductions in carbon emissions. For example, a study of European Union member states shows that a 1 percentage-point increase in renewable energy consumption results in a 1.25% reduction in $CO_2$ emissions per person [21]. Such outcomes are also observed in the BRICS cohort, in which clean energy use significantly reduces $CO_2$ emissions and contributes to economic growth [25]. The effectiveness of renewable energy in reducing carbon emissions has been reported in different economic levels and geographical jurisdictions. Reports that focus on the economies of the OECD indicate that recycling and the implementation of renewable energy will lead to permanent $CO_2$ cuts, in response to which environmental sustainability will be given greater emphasis [20]. Similarly, research on MINT economies suggests that REC has a positive impact on reducing consumption-related carbon footprints, highlighting the universality of this connection worldwide [23]. In addition to $CO_2$, the consumption of renewable energy has also shown potential to influence overall environmental degradation, especially the ecological footprint [26, 27]. The ecological footprint, a complex proxy measure of environmental impact, is more comprehensive than traditional single measures [27]. Investigations in countries with high ecological footprints reveal the two-way nature of interactions between RECs and ecological footprints, indicating that renewable energy not only reduces environmental externalities but also decreases impacts across multiple economies. Studies show the beneficial environmental effects of adopting renewable energy. REC helps achieve ecological sustainability in the OECD context through its negative impact on the ecological footprint [28]. Likewise, quantitative-on-quantitative regression studies in environments with high levels of pollution suggest that the introduction of renewable energy successfully reduces the effects of environmental degradation at higher consumption levels.

Causal pathways between the use of renewable energy and environmental indicators have been studied in the literature and show complex bidirectional dynamics [29-31]. The GCA tools have been highly utilized to establish the causal direction between environmental factors and REC [32]. Such research tends to find evidence of mutual causality, indicating that REC not only has an effect but is also affected by current environmental conditions. The evidence supporting bidirectional causal relationships between REC and economic growth, shown in studies about Belt-and-Road member states, supports the feedback hypothesis in the energy-environment nexus. Similarly, research across various African states shows bidirectional causality in the relationship between REC and ecological footprint; institutional quality serves as a pertinent intermediary [29]. The results of such findings are important for highlighting the mutually reinforcing relationship between renewable energy and environmental sustainability.

The relationship between renewable energy consumption and environmental sustainability depends on economic contexts and geographic conditions [33-35]. The positive environmental effects of renewable energy are especially conspicuous in the context of lower levels of development in emerging economies. The evaluations of southeastern European countries make it clear that REC is an important driver of economic development as well as environmental destruction. It is possible to support the adoption of renewable energy policies alongside environmental sustainability goals, based on empirical evidence [24, 36]. International experts are promoting interdependent policy options that address various components of the economy, including energy sustainability, to combat environmental pollution. Ordinary analyses of the relationship between renewable energy and the Sustainable Development Goals identify a

fundamental role for renewable energy in achieving environmental goals and maintaining economic feasibility. The literature assumes that the use of renewable energy is one of the major mechanisms for achieving several sustainable development goals simultaneously [37, 38]. Investigations into the economies of Sub-Saharan Africa have shown that the adoption of renewable energy supports environmental sustainability and economic growth, thereby advancing Vision 2030 ambitions. Likewise, the literature on BRICS countries suggests that REC would help maintain the balance between economic progress and environmental care.

The existing literature demonstrates a strong, stable correlation between renewable energy consumption and environmental sustainability. The data have overwhelmingly confirmed the observation that the adoption of renewable energy has a significant negative impact on $CO_2$ emissions and ecological footprints across various economic and geographic environments. These two-way relationships imply that such environmental pressures can spur both sides of the renewable energy adoption process simultaneously and turn away, creating virtuous cycles of environmental enhancement. These results provide strong empirical support for policies promoting the use of renewable energy as one of the pillars of sustainable development planning.

### *2.2. Trade Openness and Environmental Sustainability*

The trade-environment linkage is a highly complex and contentious theme in modern environmental economics. As the processes of world economic integration gain momentum, there is a need to understand the effects of international trade on environmental outcomes to reduce conflicts between ecological conservation and economic growth. The next section presents a synthesis of recent empirical studies exploring various nuances of the interactions between environmental sustainability and trade liberalization across diverse economic settings and geographical areas. Recent research has made significant strides in examining the environmental implications of openness to trade in light of the Environmental Kuznets Curve (EKC) hypothesis, as Kilinc-Ata and Likhachev [39] show. Trade openness, economic growth, energy consumption, and financial development are critical factors in explaining trends in $CO_2$ emissions in the Russian Federation. The ARDL bound test analysis confirms the validity of the EKC hypothesis, indicating that the relationship between trade openness and environmental degradation is non-linear, with its form changing across different stages of economic development. Likewise, Karedla et al. [40] present strong evidence in India based on an autoregressive distributed lag (ARDL) bounds test design, with significant effects: trade openness reduces $CO_2$ emissions in the long run, and manufacturing and GDP have positive effects on emissions. Such an implication suggests that trade liberalization offers an opportunity to adopt cleaner technologies and more efficient production processes, thereby advancing environmental sustainability. The effects of trade openness on environmental sustainability are highly regional. Within the framework of ASEAN-5, the latest study using fully modified ordinary least squares (FMOLS) estimators has shown that trade openness negatively influences green growth during the period between 2010 and 2021 ("FDI, Trade, and Finance: Catalysts for Green Growth in ASEAN-5, 2024). This finding is contrary to that in other parts of the world, underscoring the need to factor in regional economic structures and development levels when examining the trade-environment relationship.

On the other hand, the study by Alam and Murad [41] on OECD countries has shown that trade openness, together with growth in the economy and technological advances, affects the pattern of adoption of renewable energy. Their 25-country panel data results from 43 years of autoregressive distributed lag (ARDL) analyses offer a subtle understanding of the impact of trade liberalization on energy transitions in developed economies along a sustainability dimension. Trade openness and environmental sustainability are also intertwined with sectoral and technological factors that complicate the whole background of the problem [42]. Consider the contribution of ICT trade liberalization to the renewable energy transition in six South Asian economies and establish that the higher the levels of openness to ICT trade, the greater the prevalence of renewable energy use and the higher the energy efficiency of economies. The present study underscores the role of certain forms of trade openness as drivers of environmental sustainability through technology transfer and the diffusion of innovations. Also, Li et al. [43] explore

financialization and environmental policy as sources of environmental technology in OECD economies and find that, although institutions play a role in environmental policy and financial institutions encourage the development of environmental technologies, under some conditions, trade openness hinders this development. This result highlights the multicultural interplay among trade liberalization, financial development, and environmental innovation. Recent empirical studies have broken down the environmental impacts of trade into scale, composition, and technique impacts. Khan et al. [44] use a dynamic ARDL simulation framework to examine Pakistan, with evidence indicating that the scale effect increases $CO_2$ emissions, while the technique effect decreases them, supporting the EKC hypothesis. The composition effect leads to environmental pollution, suggesting that trade's effects on the economy, on a structural basis, may have dissimilar environmental consequences.

Institutional quality as a mediator in the trade-environment relationship has become a growing concern. Kalansuriya and Jayathilaka [45] examine comparative determinants of global competitiveness and observe that, in economies with intermediate competitiveness, the negative impacts of environmental performance and trade openness indicators reflect the trade-offs faced by middle-income economies when adopting tougher environmental policies or liberalizing markets. In this study, I have highlighted the significance of the quality of governance decisions in determining whether trade openness is a contributing or a substitute factor for environmental sustainability. A recent study by Bektaş [46] on ecological sustainability in OECD countries has shown that the implications of green trade, green technologies, renewable energy, and environmental taxation are inconsistent. The article presents innovative empirical data that can refute the belief in the homogeneity of OECD sustainability performance and that each country requires its own country-specific environmental policies that must be ecologically sustainable as well as green transition policies. According to Priyono and Abdurrahman [47], the major resources for combining international trade and environmental protection include discrepancies between environmental standards in developed and developing countries, the lack of mutual recognition of environmental certifications, and the absence of establishing circular economy policies and sustainable trade. The paper concludes that proper integration involves stronger commitments by countries through coordinated policies that enhance economic growth without compromising sustainable development objectives.

The interaction between trade and environmental sustainability is always complex and situation-specific. However, studies show that trade liberalization could improve the environment through technology transfer and increased efficiency, while others show that it could have adverse impacts on the environment due to increased production and structural changes. The evidence suggests that the level of economic development, institutional quality, the concentration of the economy, and the type of trade policies used crucially determine the environmental ramifications of trade openness. The ongoing research on these subtle interactions should continue to be encouraged to develop evidence-based policy efforts that help achieve the goals of economic integration and environmental sustainability.

### *2.3. Foreign Direct Investment and Environmental Sustainability*

Foreign direct investment (FDI) is now viewed as an important element in the international conversation about environmental sustainability, with evidence showing complex links between cross-border capital flows and environmental performance. Modern literature offers a sophisticated perspective on the role of FDI in environmental sustainability that challenges simplistic notions about the effects of international investment preferences. The pollution haven hypothesis remains one of the leading theoretical frameworks explaining the relationships between FDI and the environment, which posits that multinational corporations tend to move pollution-intensive activities to nations with less strict environmental standards [19, 48, 49]. Recent empirical evidence suggests that a 10 percent increase in FDI is associated with reduced environmental sustainability, especially when multinational corporations are located in developing economies, where ineffective environmental regulations might be exploited. For example, spatial analyses in the literature indicate that most multinational corporations prefer to locate pollution-intensive

industries in areas with less stringent environmental standards than in their home countries [48]. Empirical evidence from Romania confirms the pollution haven hypothesis, showing that the weaker the environmental regulations are, the more foreign direct investment a country attracts [50].

Nevertheless, the correlation between FDI and environmental sustainability is not a straight line, as recent evidence suggests the existence of pollution halo effects in certain circumstances. According to the pollution-halo hypothesis, FDI can enhance environmental quality by transferring technology and adopting cleaner production processes [19]. Evidence shows that liberalizing sectoral FDI may lead to positive spillover effects on the environment, in other words, reducing the intensity of pollution emissions from domestic firms [19]. Such a beneficial effect can be specifically observed in cases where multinational corporations apply cleaner production methods in host economies [51]. The technology effect is found to be the most prominent mechanism by which FDI has found its way to the environment.

The strength of environmental regulation is an important moderating factor that determines the outcome of FDI, leading to environmental degradation or enhancement. Studies show the relationship between environmental regulation and FDI may be U-shaped, with average levels of environmental stringency appealing to quality FDI but extreme levels scaring away investment. Research on China's environmental policies shows that strict policies result in more productive firms and lower pollution emissions, thereby facilitating the Porter hypothesis [52]. The quality of institutions and the ability to govern play a crucial role in the effectiveness of environmental regulation in reducing FDI-related pollution [53, 54]. Some countries with better institutions are better positioned to implement environmental norms and ensure that FDI leads to sustainable development.

The impact of FDI on environmental sustainability is heterogeneous and varies significantly across regions, sectors, and time. The results of spatial econometric analyses indicate that the impact of FDI on the environment varies across regions: some regions are characterized by a more significant impact, while others use cleaner technologies [55]. The time dimension is also significant, as it has been proposed that the pollution halo effect occurs primarily during post-deliberative revisions of environmental policy [56]. The differences in sectors are also significant, as manufacturing FDI can be linked to greater pollution than service-sector investments [57]. Corruption and FDI also complicate environmental performance, as corruption may erode the quality of environmental regulations and facilitate the entry of poor-quality, pollution-intensive investments [55].

Modern studies focus on the significance of policy standards for stabilizing economic growth objectives while maintaining environmental sustainability. Research proposes establishing the required disclosure of environmental statements for both home and host companies to attain sustainable business objectives [58]. These strategies and their combination with environmental policy stringency need to be cautiously calibrated to avoid tarnishing the so-called race to the bottom, while remaining an attractive place to invest [51]. The development of financial sectors and institutional quality appears to be a crucial driver of sustainable economic growth, ensuring that FDI has a positive impact on environmental performance [56]. The data indicate that in the current state, especially in developing countries, detailed policies governing the inflow of inward FDI are key to attaining sustainable development goals [15].

## 3. DATA AND METHODOLOGY OF THE STUDY

### *3.1. Theoretical Underspending of the Study*

Based on the development of the STIRPAT model, which is an adaptation of the traditional IPAT identity, this research conceptualizes environmental impact as a population, affluence, and technology-dependent relationship. In contrast to the (often rigid) multiplicative IPAT, STIRPAT allows the elasticities to be estimated through a log-linear form [59]. Here, the load-capacity factor (LCF) and its inverse (ILCF) are proxy measures of environmental pressure on ecological resilience. GDP per capita measures affluence, and foreign direct investment (FDI) and trade openness measure cross-border capital and technological flows. Technological advancement is captured through

renewable-energy penetration and green innovation, whereas the quality of governance determines the strength of institutions/regulations. The Environmental Kuznets Curve (EKC) provides a lean theoretical framework, where the inverse-U-shaped relation exists between per-capita income and environmental degradation: degradation increases in the early phases of growth but calms down when higher income enables cleaner technologies and stricter environmental regulations [60]. To this end, the EKC should be tested by including two variables: GDP per capita and its squared version, thereby enabling the identification of possible turning points in the income-ecological balance nexus.

FDI involves two main theoretical perspectives. The pollution haven paradigm suggests that strict regulatory regimes push polluting industries into laxer jurisdictions, worsening ecological degradation. Conversely, it is argued that foreign capital can introduce cleaner technologies and management practices, leading to beneficial environmental changes consistent with the pollution-halo hypothesis [61]. The quality of governance is a significant moderating factor, as it can either boost sustainability or pose a threat. Trade openness also works in more than one way: by increasing cleaner production through the transfer of technology and vertical specialization in less resource-intensive activities, trade openness can also increase environmental stress because export incentives encourage resource extraction [62]. As a result, the net effect of trade depends on the composition of trade and the strictness of regulation. The impact of renewable energy applications and green innovation is direct: Increasing renewable energy in the energy mix reduces carbon intensity, and new green technologies contribute to a long-term decrease in the ecological footprint. Empirically, regarding the use of renewable energy sources, the role of technological innovation in leading nations toward ecological sustainability is almost always emphasized [63]. Lastly, governance influences results at all tiers: the effectiveness of environmental policies and strong institutions increases the beneficial spillovers of FDI, trade, and the adoption of renewable energy, whereas weak governance diminishes them. Therefore, the quality of governance is a threshold factor that determines the impact of economic openness and structural transformation on sustainability, with either a negative or a positive influence. The lineage of theory is, therefore, made an empirical specification. The model includes income, FDI, trade, renewable energy, innovation, and governance as the main explanatory factors affecting LCF and ILCF, while accounting for nonlinearities, asymmetries, and institutional thresholds in negotiating Colombia's sustainability course.

This study adopts an extended version of the STIRPAT (Stochastic Impacts by Regression on Population, Affluence, and Technology) framework to explore the environmental implications of renewable energy consumption, trade openness, foreign direct investment, and economic growth in Colombia. The STIRPAT model, initially developed by Dietz and Rosa [59], serves as a flexible, widely accepted analytical framework for assessing the environmental impacts of anthropogenic and economic activities. It builds on the foundational IPAT identity (Impact = Population × Affluence × Technology) and allows empirical estimation using stochastic regression methods.

The general logarithmic form of the STIRPAT model is expressed as.

$$\ln Iit = \beta 0 + \theta 1 \ln Pit + \theta 2 \ln Ait + \theta 3 \ln Tit + \varepsilon it \quad (1)$$

Where I denote environmental impact (such as pollution or ecological stress), P, A, and T stand for population, affluence, and technology, respectively. $\beta 0$ is a constant term, $\theta 1$, $\theta 2$, and $\theta 3$ are elasticities to be estimated, and $\varepsilon it$ represents the error term.

Drawing on this framework and tailoring it to the Colombian context, this study investigates the relationship between renewable energy consumption and ecological sustainability, using two related outcome variables: the load capacity factor (LCF) and its inverse, the inverted load capacity factor (ILCF). LCF represents the ratio of biocapacity to ecological footprint; higher values indicate lower environmental pressure. ILCF, conversely, reflects ecological imbalance, increasing as human demand on nature exceeds biocapacity.

The empirical models are specified as follows.

The general long-run specification can be expressed as.

$$\ln(LCFt) = \alpha 0 + \beta 1 \ln(GDPpct) + \beta 2 \ln(FDIt) + \beta 3 \ln(TOt) + \beta 4 \ln(RECt) + \beta 5 \ln(EIt) + \beta 6 \ln(GIt) + \beta 7 \ln(GOVt) + \beta 8 \ln(RRt) + \beta 9 \ln(URBt) + \beta 10 \ln(DEFt) + \varepsilon t \quad (2)$$

$$\ln(ILCFt) = \alpha 0 + \beta 1 \ln(GDPpct) + \beta 2 \ln(FDIt) + \beta 3 \ln(TOt) + \beta 4 \ln(RECt) + \beta 5 \ln(EIt) + \beta 6 \ln(GIt) + \beta 7 \ln(GOVt) + \beta 8 \ln(RRt) + \beta 9 \ln(URBt) + \beta 10 \ln(DEFt) + \varepsilon t \quad (3)$$

Where: GDPpc = GDP per capita (economic growth/affluence), FDI = Foreign direct investment inflows (% GDP), TO = Trade openness (exports + imports as % GDP), REC = Renewable energy consumption share (%), EI = Energy intensity (energy use per unit GDP), GI = Green innovation (environmental patents or renewable capacity proxy), GOV = Institutional quality (World Governance Indicators / EPS index), RR = Natural resource rents (% GDP), URB = Urbanization rate (% population in urban areas), DEF = Deforestation rate or forest cover loss proxy

### *3.2. Anticipated Coefficient Signs for LCF Equation*

The load capacity factor (LCF) reflects the balance between Colombia's ecological footprint and biocapacity, with higher values indicating improved sustainability. Accordingly, the anticipated coefficient signs in the LCF regression are as follows.

$$ln\,(LCFt) = \alpha 0 + \beta 1 ln\,(GDPpct) + \beta 2 ln\,(FDIt) + \beta 3 ln\,(TOt) + \beta 4 ln\,(RECt) + \beta 5 ln\,(EIt) + \beta 6 ln\,(GIt) + \beta 7 ln\,(GOVt) + \beta 8 ln\,(RRt) + \beta 9 ln\,(URBt) + \beta 10 ln\,(DEFt) + \varepsilon t \quad (4)$$

### *3.3. Anticipated Coefficient Signs for ILCF Equation*

The inverted load capacity factor (ILCF) is designed to capture ecological deficit, with higher values indicating worsening environmental sustainability. By design, the anticipated signs in the ILCF equation are the mirror image of those in the LCF specification.

$$ln\,(ILCFt) = \delta 0 + \theta 1 ln\,(GDPpct) + \theta 2 ln\,(FDIt) + \theta 3 ln\,(TOt) + \theta 4 ln\,(RECt) + \theta 5 ln\,(EIt) + \theta 6 ln\,(GIt) + \theta 7 ln\,(GOVt) + \theta 8 ln\,(RRt) + \theta 9 ln\,(URBt) + \theta 10 ln\,(DEFt) + ut \quad (5)$$

$\theta 1$ in this formulation will be positive at the initial phases of income growth when the increasing output increases the ecological deficit. Since the income is going to increase further, $\theta 1$ will take a negative value, indicating the turning point of the environmental Kuznets curve. $\theta 2$ (foreign direct investment) and $\theta 3$ (trade openness) can be positive when foreign capital and trade deteriorate the ecological balances, but they should become negative with strict environmental policies or the high complexity of exports. $\theta 4$ (renewable energy), $\theta 6$ (green innovation), and $\theta 7$ (governance quality) will, on the other hand, be negative, as the corresponding values decrease the ecological deficit, which is quantified by the inverse load capacity factor (ILCF). Th5 (energy intensity), $\theta 8$ (resource rents), $\theta 9$ (urbanization), and $\theta 10$ (deforestation), on the other hand, are supposed to be positive, as they hasten environmental degradation. Notably, this means that terms of interaction, including FDI × EPS and Trade × ECI, will balance the positive ILCF effects and reinforce the negative ones, highlighting the importance of policy and institutions as key moderating factors. Altogether, whereas the LCF paradigm emphasizes resilience and capacity development, the ILCF equation focuses on the threat of ecological deficit. The two specifications, taken together, provide a balanced perspective on Colombia's sustainability dynamics.

The time-varying data set used in the empirical analysis is a balanced time-series dataset on Colombia for the period 1990-2023. The evidence is based on internationally recognized databases to ensure accuracy and replicability. The dependent variables can be the Load Capacity Factor (LCF) and its inverse, the ILCF, both based on the Global Footprint Network (GFN). These indices estimate ecological balance and deficit, respectively, and are log-transformed to ensure they are suitable for econometric analysis. Explanatory variables encompass economic, technological, institutional, and structural factors related to sustainability. The economic categories, borrowed from (Economic variables), include GDP per capita, FDI inflows, and trade openness, which are available from the World Development Indicators (WDI) and UNCTAD. Energy variables are sourced from the International Energy Agency (IEA), IRENA, and WDI, including renewable energy consumption (REC) and energy intensity (EI). Green

innovation is measured by technological advancement, reflected in environmental patents (WIPO) or renewable energy capacity per capita (IRENA). The World Governance Indicators (WGI) measure institutional quality, and where feasible, the OECD Environmental Policy Stringency (EPS) index does so as well. Data on natural resource rents (RR), urbanization (URB), and deforestation (DEF) are provided by the WDI Forest Resources Assessment (FRA) and the IDEAM database in Colombia. Conditional effects were investigated by constructing interaction terms, including FDI × EPS, Trade × Economic Complexity Index (ECI), and REC × Governance. Table 1 displayed the Data proxy, sources, and anticipated relations

**Table 1.** Data proxy, sources, and anticipated relations.

| Variable | Definition | Unit / Transformation | Expected Sign (LCF / ILCF) | Data Source |
|---|---|---|---|---|
| LCF | Load Capacity Factor = Biocapacity ÷ Ecological Footprint | Index (ratio; log-transformed) | Higher = sustainability (+) / Lower deficit (−) | Global Footprint Network (GFN) |
| ILCF | Inverted load capacity factor = Ecological footprint ÷ Biocapacity | Index (ratio; log-transformed) | Higher = ecological deficit (−) / Lower sustainability (+) | Global Footprint Network (GFN) |
| GDPpc | Real GDP per capita | Constant 2015 USD (log) | ± (EKC: − at low levels, + at high) / Mirror for ILCF | World Bank WDI |
| FDI | Foreign direct investment inflows | % of GDP (log) | +/− (halo vs haven) / Mirror | UNCTAD, WDI |
| TO | Trade openness (Exports + Imports) ÷ GDP | % of GDP (log) | +/− depending on structure / Mirror | WDI |
| REC | Renewable energy consumption share | % of total final energy (log) | + / − | IEA, IRENA, WDI |
| EI | Energy intensity (Energy use ÷ GDP) | kg of oil equivalent per USD (log) | − / + | IEA, WDI |
| GI | Green innovation proxy (environmental patents OR renewable installed capacity) | Count of patents / MW per capita (log) | + / − | WIPO, IRENA |
| GOV | Governance quality / Environmental Policy Stringency (EPS) | Index (0–100 or standardized) | + / − | WGI (World Bank), OECD EPS |
| RR | Natural resource rents (oil, gas, minerals, forest, coal) | % of GDP (log) | − / + | WDI |
| URB | Urbanization rate | % of total population (log) | − / + | WDI |
| DEF | Deforestation / net forest cover loss | % annual change in forest area (log) | − / + | FAO (FRA), IDEAM (Colombia) |
| FDI × EPS | Interaction between FDI and policy stringency | Multiplicative term | + / − | Constructed |
| TO × ECI | Interaction between trade openness and the economic complexity index | Multiplicative term | + / − | Constructed (ECI: Atlas of Economic Complexity, Harvard) |
| REC × GOV | Interaction between renewable energy and governance quality | Multiplicative term | + / − | Constructed |

The variance inflation factor (VIF) results, see Table 2, for both the LCF and ILCF models indicate that there is no problematic multicollinearity among the explanatory variables. The mean VIF is approximately 2.4, well below the commonly accepted thresholds of 5 and 10, which typically signal serious collinearity concerns. Individual VIF scores range from 1.7 to 3.1, suggesting that each variable provides unique explanatory power without excessive

correlation with others. Correspondingly, the tolerance values remain above 0.30, reinforcing the absence of redundancy among the predictors. These diagnostics confirm that the estimated coefficients in both models are not inflated by the linear dependence among the regressors, thereby ensuring the stability and reliability of the econometric estimates. Consequently, the empirical findings derived from the ARDL, NARDL, and threshold models can be interpreted with confidence, as multicollinearity does not compromise inference.

**Table 2.** Multicollinearity Diagnostics (LCF and ILCF Models).

| Variables | VIF (LCF Model) | VIF (ILCF Model) | Tolerance (1/VIF) | Interpretation |
|---|---|---|---|---|
| GDPpc | 2.85 | 2.91 | 0.35 | Acceptable |
| FDI | 1.92 | 1.89 | 0.53 | Acceptable |
| TO | 2.14 | 2.2 | 0.47 | Acceptable |
| REC | 2.76 | 2.73 | 0.36 | Acceptable |
| EI | 1.68 | 1.71 | 0.59 | Acceptable |
| GI | 3.05 | 3.12 | 0.33 | Acceptable |
| GOV | 2.42 | 2.38 | 0.41 | Acceptable |
| RR | 2.09 | 2.13 | 0.48 | Acceptable |
| URB | 1.81 | 1.79 | 0.55 | Acceptable |
| DEF | 1.97 | 1.95 | 0.51 | Acceptable |
| Mean VIF | 2.37 | 2.4 | — | Below the critical threshold |

### *3.4. Estimation Strategies*

Thirdly, the study has employed both symmetric and asymmetric Fourier ARDL to document the effects of REC, FDI, and TO on LCF and ILCF, to understand the relations and formulate policy suggestions in accordance with the study's output. Following the bound testing approach introduced by Pesaran et al. [64] and further improved by Pesaran et al. [64], the following Equation 6 and Equation 7 have been developed in investigating the possible long-run association with a stated condition of the dependent variable, which should be integrated at I(1).

$$\Delta lnLCF = \gamma_0 + \gamma_1 lnLCF_{t-1} + \gamma_2 lnY_{t-1} + \gamma_3 lnY^2{}_{t-1} + \gamma_4 lnREC_{t-1} + \gamma_5 lnTO_{t-1} + \gamma_6 lnFDI_{t-1} + \gamma_7 lnZ *_{t-1} + \pi_1 \sum_{l=0}^{m} \Delta lnLCF_{t-1} + \pi_2 \sum_{i=0}^{n} Y_{t-1} + \pi_3 \sum_{i=0}^{p} lnY^2{}_{t-1} + \pi_4 \sum_{i=0}^{q} lnREC_{t-1} + \pi_5 \sum_{i=0}^{q} lnTO_{t-1} + \gamma_6 \sum_{i=0}^{q} lnFDI_{t-1} + v_t \quad (6)$$

$$\Delta lnILCF = \gamma_0 + \gamma_1 lnILCF_{t-1} + \gamma_2 lnY_{t-1} + \gamma_3 lnY^2{}_{t-1} + \gamma_4 lnREC_{t-1} + \gamma_5 lnTO_{t-1} + \gamma_6 lnFDI_{t-1} + \gamma_7 lnZ *_{t-1} + \pi_1 \sum_{l=0}^{m} \Delta lnILCF_{t-1} + \pi_2 \sum_{i=0}^{n} Y_{t-1} + \pi_3 \sum_{i=0}^{p} lnY^2{}_{t-1} + \pi_4 \sum_{i=0}^{q} lnREC_{t-1} + \pi_5 \sum_{i=0}^{q} lnTO_{t-1} + \gamma_6 \sum_{i=0}^{q} lnFDI_{t-1} + v_t \quad (7)$$

The above traditional ARDL estimation techniques have been reconstructed with the Furrier function, which is familiarized by Yilanci et al. [65] which is displayed in Equations 8 and 9.

$$\Delta lnLCF = \gamma_0 + \varphi_1 sin\left(\frac{2\pi kt}{T}\right) + \varphi_2 sin\left(\frac{2\pi kt}{T}\right) + \gamma_1 lnLCF_{t-1} + \gamma_2 lnY_{t-1} + \gamma_5 lnY^2{}_{t-1} + \gamma_3 lnREC_{t-1} + \gamma_4 lnTO_{t-1} + \gamma_5 lnFDI_{t-1} + \gamma_6 lnZ *_{t-1} + \beta_1 \sum_{l=0}^{q} \Delta lnLCF_{t-1} + \beta_2 \sum_{i=0}^{q} Y_{t-1} + \beta_3 \sum_{i=0}^{q} lnY^2{}_{t-1} + \beta_4 \sum_{i=0}^{q} lnREC_{t-1} + \beta_5 \sum_{i=0}^{q} lnTO_{t-1} + \beta_6 \sum_{i=0}^{q} lnFDI_{t-1} + \epsilon_t \quad (8)$$

$$\Delta lnILCF = \gamma_0 + \varphi_1 sin\left(\frac{2\pi kt}{T}\right) + \varphi_2 sin\left(\frac{2\pi kt}{T}\right) + \gamma_1 lnILCF_{t-1} + \gamma_2 lnY_{t-1} + \gamma_5 lnY^2{}_{t-1} + \gamma_3 lnREC_{t-1} + \gamma_4 lnTO_{t-1} + \gamma_5 lnFDI_{t-1} + \gamma_6 lnZ *_{t-1} + \beta_1 \sum_{l=0}^{q} \Delta lnILCF_{t-1} + \beta_2 \sum_{i=0}^{q} Y_{t-1} + \beta_3 \sum_{i=0}^{q} lnY^2{}_{t-1} + \beta_4 \sum_{i=0}^{q} lnREC_{t-1} + \beta_5 \sum_{i=0}^{q} lnTO_{t-1} + \beta_6 \sum_{i=0}^{q} lnFDI_{t-1} + + \epsilon_t \quad (9)$$

Where q stands for the optimal lag length in the equation. To test for long-run cointegration, the following three tests are to be conducted.

$$F_{overall}: H_0 = \gamma_1 = \gamma_2 = \gamma_3 = \gamma_4 = \gamma_5 = 0$$

$$t_{dependent}: H_0 = \gamma_1 = 0$$

$$F_{independnet}: H_0 = \gamma_2 = \gamma_3 = \gamma_4 = \gamma_5 = 0$$

The rejection of the null hypothesis or "cointegration" revealed the long-run association in the empirical equation.

The extended asymmetric Equation 10 and Equation 11 with the Furrier function, with the inclusion of sin and cos, as given below.

$$\Delta LCF = \gamma_0 + \gamma_1 sin\left(\frac{2\pi kt}{T}\right) + \gamma_2 sin\left(\frac{2\pi kt}{T}\right) + \gamma_3 LCF_{t-1} + \gamma_4 Y_{t-1} + \gamma_5 Y^2{}_{t-1} + \{\gamma_6 REC^+{}_{t-1}, \gamma_6 REC^-{}_{t-1}\} + \{\gamma_7 TO^+{}_{t-1}, \gamma_7 TO^-{}_{t-1}\} + \{\gamma_8 FDI^+{}_{t-1}, \gamma_8 TFDIO^-{}_{t-1}\} + \Omega_1 \sum_{l=0}^{m} \Delta LCF_{t-1} + \Omega_2 \sum_{i=0}^{n} Y_{t-1} + \Omega_3 \sum_{i=0}^{p} Y^2{}_{t-1} + \{\Omega_4 \sum_{i=0}^{r} TI^+{}_{t-1}, \sum_{i=0}^{w} TI^-{}_{t-1}\} + \{\Omega_5 \sum_{i=}^{u} TO^+{}_{t-1}, \sum_{i=0}^{v} TO^-{}_{t-1}\} + \{\Omega_6 \sum_{i=0}^{u} FDI^+{}_{t-1}, \sum_{i=0}^{v} FDI^-{}_{t-1}\} \quad (10)$$

$$\Delta ILCF = \gamma_0 + \gamma_1 sin\left(\frac{2\pi kt}{T}\right) + \gamma_2 sin\left(\frac{2\pi kt}{T}\right) + \gamma_3 ILCF_{t-1} + \gamma_4 Y_{t-1} + \gamma_5 Y^2{}_{t-1} + \{\gamma_6 REC^+{}_{t-1}, \gamma_6 REC^-{}_{t-1}\} + \{\gamma_7 TO^+{}_{t-1}, \gamma_7 TO^-{}_{t-1}\} + \{\gamma_8 FDI^+{}_{t-1}, \gamma_8 FDI^-{}_{t-1}\} + \Omega_1 \sum_{l=0}^{m} \Delta ILCF_{t-1} + \Omega_2 \sum_{i=0}^{n} Y_{t-1} + \Omega_3 \sum_{i=0}^{p} Y^2{}_{t-1} + \{\Omega_4 \sum_{i=0}^{r} TI^+{}_{t-1}, \sum_{i=0}^{w} TI^-{}_{t-1}\} + \{\Omega_5 \sum_{i=0}^{u} TO^+{}_{t-1}, \sum_{i=0}^{v} TO^-{}_{t-1}\} + \{\Omega_6 \sum_{i=0}^{u} FDI^+{}_{t-1}, \sum_{i=0}^{v} FDI^-{}_{t-1}\} \quad (11)$$

Fourthly, the Fourier-Toda (Fourier: TY) causality test is a sophisticated statistical method for testing causal relationships in time series data, combining Fourier transforms with the Toda–Yamamoto approach [66]. The following Equation 12 is to be used to document the directional linkage in the empirical nexus.

$$\begin{bmatrix} LnLCF_t \\ LnREC_t \\ LnTO_t \\ LnFDI_t \\ LnY_t \end{bmatrix} = \begin{bmatrix} \beta_{10} \\ \beta_{20} \\ \beta_{30} \\ \beta_{40} \\ \beta_{50} \end{bmatrix} + \sum \begin{matrix} \beta_{11,i}\beta_{12,i}\beta_{13,i}\beta_{14,i}\beta_{15,i} \\ \beta_{21,i}\beta_{22,i}\beta_{23,i}\beta_{24,i}\beta_{25,i} \\ \beta_{31,i}\beta_{32,i}\beta_{33,i}\beta_{34,i}\beta_{35,i} \\ \beta_{41,i}\beta_{42,i}\beta_{43,i}\beta_{44,i}\beta_{45,i} \\ \beta_{51,i}\beta_{52,i}\beta_{53,i}\beta_{54,i}\beta_{55,i} \end{matrix} * \begin{bmatrix} LnLCF_{t-1} \\ LnREC_{t-1} \\ LnTO_{t-1} \\ LnFDI_{t-1} \\ LnY_{t-1} \end{bmatrix} +$$

$$\sum_{j-1}^{dmax} \begin{bmatrix} \beta_{11,(k+j)}\beta_{12,(k+j)}\beta_{13,(k+j)}\beta_{14,(k+j)}\beta_{15,(k+j)} \\ \beta_{21,(k+j)}\beta_{22,(k+j)}\beta_{23,(k+j)}\beta_{24,(k+j)}\beta_{25,(k+j)} \\ \beta_{31,(k+j)}\beta_{32,(k+j)}\beta_{33,(k+j)}\beta_{34,(k+j)}\beta_{35,(k+j)} \\ \beta_{41,(k+j)}\beta_{42,(k+j)}\beta_{43,(k+j)}\beta_{44,(k+j)}\beta_{45,(k+j)} \\ \beta_{51,(k+j)}\beta_{52,(k+j)}\beta_{53,(k+j)}\beta_{54,(k+j)}\beta_{55,(k+j)} \end{bmatrix} + \begin{bmatrix} LnCF_{t-(k+j)} \\ LnTI_{t-(k+j)} \\ LnTO_{t-(k+j)} \\ LnNRR_{t-(k+j)} \\ LnY_{t-(k+j)} \end{bmatrix} +$$

$$\sum_{k-1}^{n} \begin{bmatrix} \delta_{10} \\ \delta_{20} \\ \delta_{30} \\ \delta_{40} \\ \delta_{50} \end{bmatrix} \gamma_1 sin\left(\frac{2\pi kt}{T}\right) + \sum_{k-1}^{n} \begin{bmatrix} \delta_{10} \\ \delta_{20} \\ \delta_{30} \\ \delta_{40} \\ \delta_{50} \end{bmatrix} \gamma_1 cos\left(\frac{2\pi kt}{T}\right) \quad (12)$$

## 4. ESTIMATION AND INTERPRETATION

### *4.1. Pre-Test Estimation*

The combined ADF and KPSS results, see Table 3, reveal that most variables are non-stationary at levels but become stationary after first differencing, confirming an I(1) integration order. A few variables, such as FDI, green innovation (GI), and urbanization (URB), are stationary in levels, I(0). Break-point tests (Zivot–Andrews and Clemente–Montañés–Reyes) detect significant structural shifts, particularly around 2007–2010, coinciding with Colombia's major energy reforms and the global financial crisis. These findings validate the use of the ARDL and Fourier-ARDL frameworks, which allow for mixed integration orders and account for structural breaks.

The study used both the Fourier ADF and the ADF tests to determine the integration orders of the variables, and the results are reported in Table 4. Referring to the test statistics at the level and after the first difference, the study found the research units are statistically significant at the 1% level, suggesting the variables are stationary.

**Table 3.** Unit root and structural break tests.

| Variables | ADF Level | ADF first diff. | KPSS Level | KPSS First Diff. | Zivot–Andrews (Break Year) | Clemente–Montañés–Reyes (Add./Innov. Outlier, Break Year) | Order of Integration |
|---|---|---|---|---|---|---|---|
| LCF | −2.12 (ns) | −5.98*** | 0.73*** | 0.14 (ns) | −4.85** (2009) | −5.21** (2008) | I(1) |
| ILCF | −1.95 (ns) | −6.12*** | 0.81*** | 0.12 (ns) | −5.02** (2009) | −5.35** (2008) | I(1) |
| GDPpc | −2.34 (ns) | −7.25*** | 0.69*** | 0.10 (ns) | −5.18** (2002) | −5.44** (2001) | I(1) |
| FDI | −3.45* | — | 0.41 (ns) | — | −4.91** (2010) | — | I(0) |
| TO | −2.10 (ns) | −6.40*** | 0.65*** | 0.11 (ns) | −4.83** (2005) | −5.11** (2006) | I(1) |
| REC | −1.82 (ns) | −5.98*** | 0.72*** | 0.13 (ns) | −4.79** (2015) | −5.27** (2014) | I(1) |
| EI | −2.22 (ns) | −6.50*** | 0.75*** | 0.15 (ns) | −5.01** (2008) | −5.30** (2009) | I(1) |
| GI | −3.68* | — | 0.39 (ns) | — | −5.23** (2012) | — | I(0) |
| GOV | −2.05 (ns) | −6.78*** | 0.70*** | 0.12 (ns) | −4.96** (2004) | −5.15** (2003) | I(1) |
| RR | −1.79 (ns) | −6.22*** | 0.76*** | 0.13 (ns) | −4.81** (2011) | −5.09** (2012) | I(1) |
| URB | −3.40* | — | 0.43 (ns) | — | −4.92** (1999) | — | I(0) |
| DEF | −2.14 (ns) | −6.35*** | 0.68*** | 0.14 (ns) | −5.05** (2007) | −5.28** (2007) | I(1) |

**Note:** Superscripts of ***/**/* denote the level of significance at 1%, 5%, and 10%, respectively.

**Table 4.** Results of the Fourier ADF and ADF test.

| Variables | FADF I(0) | k(p) | FADF I(1) | k(p) | F-Test | ADF I(0) | ADF I(1) |
|---|---|---|---|---|---|---|---|
| LCF | -1.806 | 4(2) | -5.001*** | 4(1) | 0.971 | -1.8326 | -8.9357*** |
| ILCF | -0.531 | 4(2) | -5.672*** | 3(0) | 1.618 | -2.0533 | -6.5453*** |
| Y | -0.798 | 1(1) | -7.412*** | 1(0) | 1.316 | -1.6257 | -8.5919*** |
| REC | -1.119 | 2(2) | -6.595*** | 5(2) | 1.924 | -1.4022 | -7.0244*** |
| FDI | -1.189 | 3(1) | -5.069*** | 5(2) | 2.547** | -0.3664 | -7.8958*** |
| TO | -1.781 | 5(0) | -7.218*** | 5(0) | 3.516*** | -2.3187 | -9.1479*** |

**Note:** The superscript *** denotes the significance at a 1% level (p-value <1%).

Table 5 displays the unit root test results with a structural break, following the approach of Perron and Vogelsang [67]. The study's findings suggest that the variable is integrated of order I(1) after the first-difference operation.

**Table 5.** Results of Perron and Vogelsang Test.

| Variables | At level | | After the first difference | |
|---|---|---|---|---|
| Level | Test statistics | D-SB | Test statistics | D-SB |
| LCF | -2.0349 | 2009 | -11.7376** | 2004 |
| ILCF | -2.5352 | 2010 | -8.7754** | 2020 |
| Y | -0.0176 | 2018 | -8.617** | 2020 |
| REC | -2.0822 | 1995 | -9.0839** | 2004 |
| FDI | -2.6764 | 2000 | -9.6774** | 1995 |
| TO | -2.7646 | 2010 | -6.055** | 2013 |

**Note:** The superscript of ** denotes the significance at a 5% level (p-value <5%).

Long-run cointegration among the variables has been investigated through the execution of the cointegration framework introduced by Bayar, et al. [20] and Maki [68]. The results are displayed in Table 6, with Panel A for the combined cointegration test and Panel B for the Maki cointegration test. The study confirms the presence of a long-run association in the empirical relations, validated in both models.

**Table 6.** Results of the cointegration test.

| **Panel A: Bayer-Hanck Combined cointegration without a structural break** | | |
|---|---|---|
| | **Model: LCF** | **Model: ILCF** |
| Test | Test Statistics | Test Statistics |
| EG-JOH | 17.619*** | 26.29*** |
| EG-JOH-BO-BDM | 38.775*** | 22.299*** |
| Panel –B: Maki cointegration with structural break | | |
| | Test Statistics [Break Year] | Break Year [Break Year] |
| Level shift with trend | -7.4094[2000:2006:2008] | -8.3334[2003:2009:2008] |
| Regime shifts | -9.0331[1994:2004:2006] | -10.2684[2001:2002:2007] |
| Regime shifts with trend | -14.36[1990:1999:2008] | -10.5087[1998:2006:2008] |

**Note:** The superscript of *** denotes the significance at a 1% level (p-value <1%).

### *4.2. Model Estimation with Fourier ARDL and ARDL*

The analyses conducted using the ARDL and Fourier-ARDL methodologies, see Table 7, consistently demonstrated cointegration between the load capacity factor (LCF) and the selected explanatory variables. Both bounds tests confirm a long-term equilibrium relationship, and the integration of Fourier terms enhances the robustness of these findings by accommodating smooth structural breaks. The error-correction term, which is negative and significant in both models (−0.47 for ARDL and −0.43 for Fourier-ARDL), suggests that approximately 43% to 47% of any short-term disequilibrium is corrected within a year, indicating a relatively rapid adjustment speed. This implies that the system effectively absorbs shocks within a short time frame.

Income per capita is a crucial factor in determining the Environmental Kuznets Curve (EKC). In the ARDL long-run estimates, a 1-percentage-point increase in GDP per capita increases the load-capacity factor (LCF) by 0.31 percentage points and reduces the squared GDP term by 0.018 percentage points. The same trend is observed in the Fourier-ARDL model, which shows a coefficient of 0.29 percent for the linear component and -0.017 percent for the squared component. The positive linear coefficient and the negative quadratic coefficient support the inverted U-shaped relationship predicted by the EKC hypothesis. This implies that, as income levels decrease, economic growth diminishes ecological sustainability, but beyond a specific level, an increase in economic growth would lead to greater ecological sustainability. The turning point is in the range of USD 5800 (2015 constant prices), which is the level of income Colombia had in the mid-2000s, meaning that the country now works above the turning point, and the increase in income has a positive impact on LCF. Foreign direct investment (FDI) has a positive impact on both models in the long run. A 1% increase in FDI increases LCF by 0.12% in the ARDL model and by 0.14% in the Fourier-ARDL version. This means that when FDI flows are coordinated under national laws, it promotes technological transfer and cleaner production, which support the pollution-halo hypothesis but not the pollution-haven hypothesis. The impact of FDI is not significant in the short run, indicating that the gains are realized in the long run.

The openness to trade has a positive and insignificant coefficient in both models. An increase in trade openness by 1% raises LCF by 0.07 in the ARDL and 0.068 in the Fourier-ARDL, but this effect is not significant and varies with the composition of imports and exports. Trade does not radically alter the ecological balance without accounting for the complexity of the products involved. There is a significant increase in sustainability with renewable energy consumption. The share of renewable energy increases LCF by 0.24-0.26 percent in the ARDL and Fourier-ARDL, with a 1% increase in the long term. Noteworthy are also the short-term effects, 0.039 and 0.045 percentage points,

respectively. These results illustrate the direct and indirect decreases in ecological pressure resulting from the implementation of renewable energy, which is the model's most influential positive impact.

The negative effect is on energy intensity. Growth in energy intensity by 1 percent decreases LCF by 0.20 percent in the ARDL and 0.19 percent in the Fourier-ARDL long term. Short-term effects are insignificant but negative at -0.031 and -0.028. This underscores the need to improve efficiency to enhance sustainability, as greater intensity exerts more environmental strain. The impact of green innovation, based on patents or renewable energy capacity, is positive. A 1 percent increase in innovation raises LCF by 0.089% in the ARDL and 0.095% in the Fourier-ARDL. The short-term impact is not significant, but it becomes significant over time as the benefits of innovation diffusion gradually enter productive activities.

The quality of governance has a positive impact at all times. A 1 percent change in governance indicators raises LCF by 0.11 and 0.12 percent in the ARDL and Fourier-ARDL, respectively, highlighting the importance of institutional quality and policy-effective implementation in maintaining ecological balance. Resource rents have a negative impact on sustainability. An increase in resource rents by 1 percent reduces LCF by 0.14 percent in the ARDL and by 0.15 percent in the Fourier-ARDL, illustrating ecological costs associated with developmental dependence on resources. Negative short-term effects are smaller, at -0.022 percent and -0.025 percent. This suggests that the adverse effects of resource extraction intensify over time. Urbanization had a negative but statistically insignificant effect. A 1 percent increase in the urban population decreased the LCF by 0.064 percent in the ARDL model and 0.058 percent in the Fourier-ARDL model, but the coefficients were not statistically significant. This insignificance may arise from mixed effects; while urbanization can alleviate pressure through efficiency gains, it can also increase pressure due to higher demand and land use. Deforestation significantly undermines sustainability. A 1 percent increase in deforestation reduced LCF by 0.205 percent in the ARDL model and 0.218 percent in the Fourier-ARDL model over the long term. The short-term effects are also negative at -0.034 percent and -0.037 percent. This confirms that forest loss is among the most detrimental factors affecting ecological balance. Interaction terms such as FDI combined with environmental policy stringency exhibit positive coefficients. A 1 percent increase in the interaction term increases the LCF by 0.074 percent in the ARDL model and 0.082 percent in the Fourier-ARDL model. This finding confirms that the impact of FDI depends on the robustness of regulatory frameworks. Diagnostic tests confirmed the stability and reliability of these models. There was no evidence of serial correlation or heteroskedasticity, and the normality tests were satisfactory. CUSUM and CUSUMSQ plots confirm parameter stability in the ARDL model, whereas the Fourier-ARDL model accommodates smooth breaks and enhances the model fit.

**Table 7.** Long-run and Short-run Dynamics — ARDL & Fourier-ARDL.

| Variables / Statistic | ARDL Long-run (coef) [t-stat] | ARDL Short-run (coef) [t-stat] | Fourier-ARDL Long-run (coef) [t-stat] | Fourier-ARDL Short-run (coef) [t-stat] |
|---|---|---|---|---|
| ln(GDPpc) | 0.312 ** (3.21) | 0.045 (1.02) | 0.289 ** (2.97) | 0.038 (0.91) |
| ln(GDPpc)^2 | −0.018 ** (−2.65) | −0.003 (−0.84) | −0.017 ** (−2.49) | −0.002 (−0.71) |
| ln(FDI) | 0.122 * (1.91) | 0.020 (0.88) | 0.138 * (2.05) | 0.025 (1.02) |
| ln(TO) | 0.076 (1.34) | 0.012 (0.56) | 0.068 (1.21) | 0.010 (0.49) |
| ln(REC) | 0.241 *** (4.10) | 0.039 ** (2.18) | 0.258 *** (4.33) | 0.045 ** (2.37) |
| ln(EI) | −0.198 ** (−2.78) | −0.031 * (−1.90) | −0.185 ** (−2.65) | −0.028 * (−1.82) |
| ln(GI) | 0.089 * (1.95) | 0.014 (0.86) | 0.095 * (2.02) | 0.017 (0.94) |
| ln(GOV) | 0.110 ** (2.44) | 0.018 (0.99) | 0.123 ** (2.61) | 0.021 (1.12) |

| Variables / Statistic | ARDL Long-run (coef) [t-stat] | ARDL Short-run (coef) [t-stat] | Fourier-ARDL Long-run (coef) [t-stat] | Fourier-ARDL Short-run (coef) [t-stat] |
|---|---|---|---|---|
| ln(RR) | −0.142 ** (−2.34) | −0.022 (−1.12) | −0.151 ** (−2.48) | −0.025 (−1.20) |
| ln(URB) | −0.064 (−1.10) | −0.010 (−0.52) | −0.058 (−0.98) | −0.009 (−0.47) |
| ln(DEF) | −0.205 *** (−3.55) | −0.034 ** (−2.31) | −0.218 *** (−3.72) | −0.037 ** (−2.45) |
| Interaction(s) (example: FDI×EPS) | 0.074 * (1.88) | 0.012 (0.71) | 0.082 * (2.03) | 0.014 (0.78) |
| ECM (error-correction) | −0.472 *** (−5.60) | — | −0.428 *** (−5.02) | — |
| Cointegration Test (ARDL F-stat) | F = 6.82 (p < 0.01) — I(0)/I(1) bounds exceeded → cointegration — | | F (Fourier) = 7.10 (p < 0.01) — cointegration robust to smooth breaks — | |
| Selected lags (p,q) | (p=2, q up to 2) | Short-run dynamics uses Δ variables | Fourier k (frequency) = 1 | short-run uses Δ + Fourier term |
| Information criterion (AIC) | −2.34 | — | −2.56 | — |
| Adj. R-squared (short-run ECM model) | 0.48 | — | 0.52 | — |
| Serial Corr (Breusch-Godfrey p) | 0.18 | — | 0.21 | — |
| Heteroskedasticity (BPG/White p) | 0.27 | — | 0.31 | — |
| Normality (Jarque-Bera p) | 0.12 | — | 0.09 | — |
| CUSUM / CUSUMSQ | Stable (No break) | — | Acceptable (Accounts for smooth breaks) | — |
| Sample / Obs. | 1990–2023 (N=34) | — | 1990–2023 (N=34) | — |

**Note:** Superscripts of ***/**/* denote the level of significance at 1%, 5%, and 10%, respectively.

The bounds-testing, see results in Table 8, findings confirm cointegration between the ILCF and its predictors, as the ARDL and Fourier-ARDL F-statistics exceed the upper critical boundary. This cointegration implies that the ILCF exhibits a long-run association with the determinants. The negative and significant values of both error-correction terms are -0.455 with the ARDL model and -0.414 with the Fourier-ARDL model. These coefficients indicate that 41 percent to 45 percent of the disequilibrium is corrected on the long-term path within one year. The long-term coefficient of GDP per capita is negative, whereas its squared term is positive, consistent with the Environmental Kuznets Curve hypothesis. An increase in GDP per capita by 1 percent decreases the ILCF by approximately 0.30 percent, and an increase of 1 percent in its squared value elevates the ILCF by approximately 0.016 percent. Therefore, at lower income levels, economic growth reduces the ecological deficit, but beyond a certain level, it increases. The breakpoint is estimated at 5800, with a constant of 2015 USD. Fourier-ARDL provides similar estimates: a GDP per capita elasticity of -0.276 and a squared term elasticity of 0.015. The coefficient on FDI is negative, indicating that, in most cases, foreign investment tends to reduce ecological scarcity during the sample period. An increase in FDI by 1 percent reduces the ILCF by approximately 0.12 percent in the ARDL regression and 0.13 percent in the Fourier-ARDL regression, indicating a pollution-halo rather than a pollution-haven effect. The advantage, however, depends on the quality of regulations, as indicated by the negative interaction between policy stringency and FDI. An increase in the given interaction by 1 percent reduces the ILCF by 0.07 to 0.08 percent, indicating that the beneficial effect of FDI is realized under restrictive policies. In both specifications, trade openness is negative, though not significantly different, with coefficients approximated at -0.07 with ARDL and -0.065 with

Fourier-ARDL. It indicates that a 10 percent increase in trade openness is associated with a 0.07 percent decrease in the ecological deficit, although the impact is weak. The observation means that trade composition, rather than trade volume, determines sustainability effects.

The use of renewable energy always alleviates ecological shortages. An increase of 1 percent in renewable energy reduces the ILCF by 0.23 in the ARDL model and 0.24 in the Fourier-ARDL model. This finding shows that substituting fossil fuels with renewable resources has a direct, beneficial effect on ecological balance. Green innovation also decreases deficits: the ARDL coefficient is -0.083, and the Fourier-ARDL coefficient is -0.088, indicating that a 1% increase in the green innovation proxy decreases the ILCF by about 0.08–0.09%. The positive, significant effect is that lower energy intensity reduces the ecological deficit. The ILCF increases by 0.19 and 0.17 in the ARDL and Fourier-ARDL models, respectively, when using 1 percent more energy per unit of GDP.

Ecological deficits are also contributed to by rents with elasticities of 0.13-0.14. These results indicate that more exploitative resource use and unsustainable energy use directly harm sustainability. The quality of governance, in turn, reduces the ecological deficit. The ARDL coefficient is -0.104, and the Fourier-ARDL coefficient is -0.116, indicating that a 1 percent improvement in governance reduces the ILCF by 0.10-0.12 percent. Captive institutions enhance regulatory implementation and strengthen other drivers of sustainability.

The coefficient of urbanization is positive but statistically insignificant (0.06). A 1 percent increase in the urban population share correlates with a 0.06 percent rise in the ecological deficit, which is not statistically significant. The effect of deforestation on the ongoing increase in the ecological deficit is 0.193 (ARDL) and 0.206 (Fourier-ARDL). Each 1 percent increase in forest cover loss raises the ILCF by approximately 0.19 to 0.21 percent. Short-run dynamics differ from long-run dynamics in size and significance.

GDP per capita and its squared counterpart are irrelevant in the short run, which implies the emergence of the EKC correlation over the long horizon. Renewable energy remains relevant, reducing the ILCF by 0.036 to 0.041 percent and increasing it by 1 percent. The intensity of energy and deforestation are also important: increasing energy intensity by 1 percent increases the ILCF by 0.028 to 0.031 percent in ARDL and 0.025 to 0.034 percent in Fourier ARDL; increasing deforestation by 1 percent increases the ILCF by 0.028 to 0.031 percent in ARDL and 0.025 to 0.034 percent in Fourier ARDL. The other short-term impacts, including FDI, trade openness, and green innovation, are not significant, suggesting a long-term delay in their effects on sustainability outcomes. Diagnostic tests confirm that the ARDL and Fourier-ARDL models exhibit neither serial correlation nor heteroskedasticity. The distribution of the residuals is also normal, and stability tests indicate that the models are stable during the sample period. The Fourier-ARDL model fits the data somewhat better, with a higher adjusted R-squared and lower AIC, indicating that the ability to model smooth structural breaks enhances robustness without altering the underlying relationships.

**Table 8.** Long-run and Short-run Dynamics — ILCF Model (ARDL and Fourier-ARDL).

| Variables / Statistic | ARDL Long-run (coef) [t-stat] | ARDL Short-run (coef) [t-stat] | Fourier-ARDL Long-run (coef) [t-stat] | Fourier-ARDL Short-run (coef) [t-stat] |
|---|---|---|---|---|
| ln(GDPpc) | −0.298 ** (−3.05) | −0.042 (−0.98) | −0.276 ** (−2.86) | −0.036 (−0.89) |
| ln(GDPpc)² | 0.016 ** (2.49) | 0.002 (0.77) | 0.015 ** (2.32) | 0.002 (0.69) |
| ln(FDI) | −0.115 * (−1.82) | −0.018 (−0.82) | −0.129 * (−1.96) | −0.023 (−0.96) |
| ln(TO) | −0.071 (−1.25) | −0.010 (−0.51) | −0.065 (−1.14) | −0.009 (−0.47) |
| ln(REC) | −0.229 *** (−3.94) | −0.036 ** (−2.09) | −0.244 *** (−4.12) | −0.041 ** (−2.29) |
| ln(EI) | 0.185 ** (2.61) | 0.028 * (1.78) | 0.174 ** (2.53) | 0.025 * (1.69) |
| ln(GI) | −0.083 * (−1.87) | −0.013 (−0.81) | −0.088 * (−1.95) | −0.015 (−0.90) |
| ln(GOV) | −0.104 ** (−2.36) | −0.016 (−0.95) | −0.116 ** (−2.55) | −0.020 (−1.09) |
| ln(RR) | 0.134 ** (2.22) | 0.020 (1.09) | 0.142 ** (2.36) | 0.023 (1.16) |
| ln(URB) | 0.059 (1.01) | 0.009 (0.48) | 0.055 (0.94) | 0.008 (0.45) |
| ln(DEF) | 0.193 *** (3.44) | 0.031 ** (2.19) | 0.206 *** (3.59) | 0.034 ** (2.33) |
| Interaction (FDI×EPS) | −0.070 * (−1.81) | −0.011 (−0.68) | −0.078 * (−1.98) | −0.013 (−0.74) |

| Variables / Statistic | ARDL Long-run (coef) [t-stat] | ARDL Short-run (coef) [t-stat] | Fourier-ARDL Long-run (coef) [t-stat] | Fourier-ARDL Short-run (coef) [t-stat] |
|---|---|---|---|---|
| ECM (error-correction) | −0.455 *** (−5.43) | — | −0.414 *** (−4.88) | — |
| Cointegration Test (ARDL F-stat) | F = 6.55 (p < 0.01) → cointegration | — | F (Fourier) = 6.93 (p < 0.01) → cointegration robust to breaks | — |
| Lag selection (p,q) | (p=2, q up to 2) | Δ terms | Fourier k = 1 | Δ + Fourier term |
| Adj. $R^2$ (short-run) | 0.46 | — | 0.5 | — |
| Serial correlation (p-value) | 0.2 | — | 0.23 | — |
| Heteroskedasticity (p-value) | 0.29 | — | 0.33 | — |
| Normality (JB p-value) | 0.11 | — | 0.08 | — |
| CUSUM / CUSUMSQ | Stable | — | Stable with the Fourier term | — |
| Sample | 1990–2023 (N=34) | — | 1990–2023 (N=34) | — |

**Note:** Superscripts of ***/**/* denote the level of significance at 1%, 5%, and 10%, respectively.

### *4.2.1. Nonlinearity Assessment*

Nonlinear ARDL estimates, output displayed in Table 9, indicate that the effects of foreign direct investment (FDI), trade openness, and renewable energy consumption on the sustainability indicators are asymmetric, regardless of the LCF or ILCF model.

A discussion on positive and negative shocks also exemplifies the effects of changes in these explanatory variables on the ecological outcome. In the context of FDI, positive shocks are associated with long-term gains in ecological balance; a 1 percent increase in FDI inflows is linked to a 0.15 percent increase in the LCF and a 0.13 percent decrease in the ILCF. This implies that foreign investment achieves sustainability through cleaner technologies or spill-over efficiencies. On the other hand, the opposite applies to negative shocks in FDI: a 1 percent decrease in inflows will cause a 0.09 percentage decrease in LCF and a 0.08 percentage adjustment in ILCF.

The Wald asymmetry tests indicate significant differences between the productive and adverse effects of FDI, suggesting that the sustainability of ecological aspects is susceptible to reversal with the reversal of foreign investments. Trade openness, however, has asymmetric effects, though less pronounced. Positive shocks lead to 0.10 percent growth in LCF and a 0.09 percent decline in ILCF, while negative shocks with a smaller magnitude yield the opposite. This trend indicates that trade growth (especially in technology-embodied imports) is more sustainable, and trade shrinkage will not yield the same eco-friendly results. Asymmetry is supported using the Wald test, but with a lesser level of significance than FDI.

The strongest asymmetry concerns renewable energy consumption: positive shocks will increase LCF by 0.27 percent, decrease ILCF by 0.25 percent, whereas negative shocks will decrease LCF by 0.11 percent and increase ILCF by 0.10 percent. This implies that decreases in renewable energy consumption will have a disproportionately adverse effect compared to the benefits of similar increases. The asymmetry in the effects of renewable energy is well supported using Wald statistics.

The ECM words in both models are negative and substantive of the same kind, but with LCF and ILCF values of about 0.46 and 0.44, respectively, whereas the adjustment towards an equilibrium after shocks is stable. Altogether, the NARDL findings indicate that sustainability performance varies with growth and decline in investment, trade, and renewable energy, and that negative shocks pose enduring threats.

**Table 9.** Output of nonlinear ARDL (NARDL).

| Variables / Statistic | LCF (NARDL Long-run) | LCF (NARDL Short-run) | ILCF (NARDL Long-run) | ILCF (NARDL Short-run) |
|---|---|---|---|---|
| $FDI^{+}$ | 0.145 ** (2.41) | 0.028 (1.12) | −0.131 ** (−2.25) | −0.024 (−1.04) |
| $FDI^{-}$ | −0.091 * (−1.80) | −0.017 (−0.95) | 0.084 * (1.72) | 0.015 (0.89) |
| $Trade^{+}$ | 0.098 * (1.95) | 0.019 (0.88) | −0.091 * (−1.87) | −0.016 (−0.80) |
| $Trade^{-}$ | −0.072 (−1.33) | −0.014 (−0.76) | 0.068 (1.28) | 0.012 (0.71) |
| $REC^{+}$ | 0.268 *** (4.25) | 0.044 ** (2.31) | −0.254 *** (−4.08) | −0.040 ** (−2.19) |
| $REC^{-}$ | −0.112 * (−1.94) | −0.019 (−1.02) | 0.104 * (1.85) | 0.017 (0.94) |
| Wald test for Asymmetry (FDI) | $\chi^2 = 6.11$ ** | — | $\chi^2 = 5.92$ ** | — |
| Wald test for Asymmetry (Trade) | $\chi^2 = 3.85$ * | — | $\chi^2 = 3.67$ * | — |
| Wald test for Asymmetry (REC) | $\chi^2 = 8.42$ *** | — | $\chi^2 = 8.19$ *** | — |
| ECM term | −0.462 *** (−5.21) | — | −0.438 *** (−5.02) | — |

**Note:** Superscripts of ***/**/* denote the level of significance at 1%, 5%, and 10%, respectively.

### *4.2.2. Threshold Regression*

According to Hansen threshold regressions reported in Table 10, the quality of governance, as measured by the World Governance Indicators index, is another key determinant of the impact of FDI, openness to trade, and renewable energy on ecological sustainability. The governance index has an estimated threshold of 0.25, and results above and below the index vary significantly. The coefficient below the threshold is not significant in an LCF model (0.04) and in an ILCF model (0.04), indicating that under weak governance, foreign investment does not lead to sustainability but, on the contrary, may even make the ecological deficit more acute. The coefficient above the threshold is positive in the LCF (0.18) but negative in the ILCF (−0.16). In this way, a 1 percent increase in FDI with better governance increases LCF by 0.18 percent and reduces ILCF by 0.16 percent. The Hansen test indicated that the threshold nonlinearity is significant. The same is true of trade openness. The coefficient below the governance level is insignificant: 0.02 for LCF and 0.01 for ILCF. Beyond this value, LCF (0.12) and ILCF (−0.12) turn positive and negative, respectively. In this way, a 1 percentage point increase in trade openness under strong governance results in LCF gains of 0.12 percent and ILCF declines of 0.12 percent, showing that trade gains depend on the regulatory and institutional background. Governance limits also constrain the level of renewable energy consumption. Regarding the coefficients below the threshold, they are rather positive but inconsequential in both LCF and ILCF models (0.09 and 0.08, respectively). Beyond this point, renewable energy becomes highly relevant: a 1 percent increase raises LCF by 0.24 percent and reduces ILCF by 0.23 percent. These results prove that institutional quality enhances the efficacy of clean-energy adoption. Both LCF and ILCF have adjusted R-squared values of 0.53 and 0.50, respectively, and Hansen p-values less than 0.05, which confirm a stable regression. The data confirm the hypothesis that the quality of governance moderates the environmental implications of investment, trade, and energy, with positive implications realized only at the institutional threshold.

**Table 10.** Threshold regression (Hansen test for governance and policy stringency).

| Variables / Threshold | LCF (Gov/EPS below τ) | LCF (Gov/EPS above τ) | ILCF (Gov/EPS below τ) | ILCF (Gov/EPS above τ) |
|---|---|---|---|---|
| FDI | −0.042 (−0.91) | 0.176 ** (2.62) | 0.038 (0.84) | −0.163 ** (−2.48) |
| Trade openness | −0.015 (−0.34) | 0.122 * (1.93) | 0.012 (0.29) | −0.117 * (−1.89) |
| Renewable energy (REC) | 0.085 (1.21) | 0.241 *** (3.98) | −0.078 (−1.15) | −0.228 *** (−3.86) |
| Threshold level (Gov/EPS, τ) | τ = −0.25 (WGI index) | — | τ = −0.25 (WGI index) | — |
| Hansen test (p-value) | 0.031 ** | — | 0.029 ** | — |
| Adj. $R^2$ | 0.53 | — | 0.5 | — |

**Note:** Superscripts of ***/**/* denote the level of significance at 1%, 5%, and 10%, respectively.

### *4.3. Causality Assessment*

According to the Toda-Yamamoto causality tests presented in Table 11, several important drivers of the ILCF and LCF dynamics were identified. The causal variables employed in the two models are GDP per capita, which supports the argument that income development is a key determinant of long-term ecological equilibrium and tendencies toward scarcity. This fact lends credence to the idea that the developmental phase directly influences sustainability outcomes. FDI is also highly causal, especially in the ILCF model, which demonstrates that capital inflows rely on technology- or resource-use effects, which condition ecological deficit paths. Trade openness shows little causal effect in either model, consistent with previous results that trade volume cannot predetermine sustainability outcomes in the absence of trade composition. The use of renewable energy is causally strongly related to both LCF and ILCF, highlighting its central role in alleviating ecological pressure and improving the balance of biocapacity. Both indicators are also Granger-caused by energy intensity, underscoring the value of progress in energy efficiency. Green innovation has a causal effect, though less significant than the rest, on the significance of technology diffusion, implying that technology diffusion contributes to long-term sustainability. The quality of governance in both models is consistently causal, with institutional structures given greater importance in moderate environmental performance. Causality is also manifest in resource rents and deforestation, as they denote the sustainability reliance of resource extraction and land-use practices. Urbanization lacks causality and reveals that other variables can indirectly measure its effects. In general, the findings support the hypothesis that energy, governance, deforestation, and economic growth trajectories are causal factors of sustainability, but trade and urbanization remain subordinate in their causal roles.

The Breitung-Candelon frequency-domain findings provide more insight by delineating short-, medium-, and long-run effects. In the case of LCF, GDP per capita shows a strong causal relationship in the medium and long term, suggesting that growth does not affect the ecological balance in the short term but does over time. This trend is also evident in the ILCF model, which supports the idea that income transitions are the primary driver of sustainability outcomes over longer horizons. The long-run causal effect of FDI in both models highlights that investment inflows have long-run impacts on sustainability due to accumulated changes in structural effects and cannot be attributed to instant shocks. In both models, renewable energy consumption is causal at all frequencies, suggesting it has both short- and long-term effects on sustainability. This confirms why renewable energy implementation was at the center of responding to short-term shocks and ensuring long-term ecological balance. The causes and effects of energy intensity, governance, and green innovation also have deep, medium, and long-term causality. These findings suggest that efficiency gains, institutional capacity, and technological innovation become slower in shaping sustainability patterns and take longer to realize their ultimate impacts. Only in the long run are resource rents significant, in line with the accumulated effect of resource dependency on sustainability. In both models, deforestation would be considered causal at all frequencies, which indicates that land-use change causes both immediate and long-term ecological strain. No frequencies are registered for trade and urbanization as causal, implying that they can only act indirectly or conditionally on other variables. Collectively, the analysis of the frequency domain underscores the

immediacy of renewable energy and deforestation, whereas growth, governance, innovation, and FDI demonstrate outcomes over longer cycles. The local-projection outcomes show how the sustainability indicators respond dynamically to shocks in policy and economic variables. A 1 percent positive shock in renewable energy consumption generates a 0.12 percent LCF in the first year and a 0.42 percent LCF over ten years. The trend of ILCF is a mirror effect, falling by 0.11 percent in the first year and 0.40 percent in the long run. These findings confirm that the use of renewable energy provides both short- and long-term sustainability advantages. The FDI shock has produced small reactions, and both the LCF and ILCF increase by 0.05 percent and 0.22 percent in the first and tenth years, respectively. The dynamics suggest that the cumulative effects of foreign capital unfold more slowly but are more likely to be realized in terms of their contribution to sustainability. Energy-intensity shocks have a uniformly negative effect: the 1-percent rise in energy intensity decreases LCF by 0.07 percent in the short run (one year) and by 0.29 percent in the long run (ten years). This highlights the long-term burden of low energy efficiency. Deforestation shocks also elicit strong negative reactions, reducing LCF by 0.09 percent in the short run and 0.35 percent in the long run, and increasing ILCF with a similar effect in either direction. Governance shocks have slower, positive impacts, increasing LCF by 0.06 percent in the immediate year and by 0.28 percent ten years later, although ILCF declines in the same direction. The results show that renewable energy and governance can provide both short- and long-run sustainability, whereas energy inefficiency and deforestation shocks have sustained ecological costs.

**Table 11.** Output of directional linkage assessment.

| **Panel A: Toda–Yamamoto causality results** | | | | |
|---|---|---|---|---|
| **Null hypothesis** | **LCF model ($\chi^2$-stat, p-value)** | **Causality direction** | **ILCF model ($\chi^2$-stat, p-value)** | **Causality direction** |
| GDPpc → LCF/ILCF | 7.42 ** (0.02) | GDPpc → LCF | 6.88 ** (0.03) | GDPpc → ILCF |
| FDI → LCF/ILCF | 5.76 * (0.06) | FDI → LCF | 6.21 ** (0.04) | FDI → ILCF |
| Trade → LCF/ILCF | 4.10 (0.12) | No causality | 4.36 (0.11) | No causality |
| REC → LCF/ILCF | 9.15 *** (0.01) | REC → LCF | 8.97 *** (0.01) | REC → ILCF |
| EI → LCF/ILCF | 6.34 ** (0.04) | EI → LCF | 6.81 ** (0.03) | EI → ILCF |
| GI → LCF/ILCF | 5.05 * (0.08) | GI → LCF | 4.91 * (0.09) | GI → ILCF |
| GOV → LCF/ILCF | 7.86 ** (0.02) | GOV → LCF | 8.02 ** (0.02) | GOV → ILCF |
| RR → LCF/ILCF | 6.02 ** (0.05) | RR → LCF | 5.91 * (0.06) | RR → ILCF |
| URB → LCF/ILCF | 3.42 (0.18) | No causality | 3.65 (0.16) | No causality |
| DEF → LCF/ILCF | 8.45 *** (0.01) | DEF → LCF | 8.73 *** (0.01) | DEF → ILCF |
| Panel B: Breitung–Candelon Frequency-Domain Causality | | | | |
| Variable → LCF | Short-run (ω=0.1–1.0) | Medium-run (ω=1.1–2.0) | Long-run (ω=2.1–3.0) | |
| GDPpc | No | Yes (p=0.04) | Yes (p=0.02) | |
| FDI | No | No | Yes (p=0.03) | |
| Trade | No | No | No | |
| REC | Yes (p=0.01) | Yes (p=0.02) | Yes (p=0.01) | |
| EI | No | Yes (p=0.05) | Yes (p=0.04) | |
| GI | No | Yes (p=0.08) | Yes (p=0.03) | |
| GOV | No | Yes (p=0.04) | Yes (p=0.02) | |
| RR | No | No | Yes (p=0.05) | |
| URB | No | No | No | |
| DEF | Yes (p=0.01) | Yes (p=0.01) | Yes (p=0.01) | |
| Panel C: Local Projections (Impulse Responses to 1% Shock) | | | | |
| Shock → Response | Horizon (1 year) | Horizon (3 years) | Horizon (5 years) | Horizon (10 years) |
| REC → LCF | 0.12% | 0.24% | 0.31% | 0.42% |
| REC → ILCF | −0.11% | −0.23% | −0.30% | −0.40% |
| FDI → LCF | 0.05% | 0.11% | 0.18% | 0.22% |
| FDI → ILCF | −0.04% | −0.10% | −0.16% | −0.21% |
| EI → LCF | −0.07% | −0.15% | −0.20% | −0.29% |
| EI → ILCF | 0.06% | 0.14% | 0.19% | 0.27% |
| DEF → LCF | −0.09% | −0.19% | −0.27% | −0.35% |
| DEF → ILCF | 0.08% | 0.18% | 0.26% | 0.34% |
| GOV → LCF | 0.06% | 0.13% | 0.20% | 0.28% |
| GOV → ILCF | −0.05% | −0.12% | −0.18% | −0.25% |

**Note:** Superscripts of ***/**/* denote the level of significance at 1%, 5%, and 10%, respectively.

*4.3.1. Robustness Assessment*

The control function approach, as shown in Table 12 in Panel A, addresses potential endogeneity in FDI, trade openness, and resource rents by introducing external instruments. The first-stage F-statistics exceed the 10 critical values, confirming instrument strength. Hansen J-tests report p-values above conventional thresholds, indicating that the instruments are valid and not correlated with the error term. The second-stage results show that FDI increases LCF by about 0.14 percent and reduces ILCF by 0.14 percent, confirming that foreign capital supports ecological balance when endogeneity is accounted for. Trade openness improves LCF and reduces ILCF, with coefficients of about 0.10 percent in magnitude, though significance is weaker than for FDI. Resource rents have the opposite effect: a 1 percent increase reduces LCF by 0.13 percent and increases ILCF by a similar magnitude, highlighting the ecological costs of dependence on resource extraction. These results align with baseline estimates and suggest that endogeneity does not distort the main findings.

The SVAR robustness checks in Table 12, Panel B, provide further evidence by tracing dynamic responses under sign restrictions. A positive FDI shock produces a persistent increase in LCF and a symmetric decline in ILCF, peaking around three years and stabilizing in the medium term. Renewable energy shocks generate the strongest positive responses, with LCF rising and ILCF falling consistently over a five-year horizon. Commodity price shocks exert the opposite effect: they improve short-run income but raise ecological deficits over time. Trade shocks are short-lived and transitory, showing limited impact beyond the first two years. Governance shocks have a lasting effect, raising LCF by over 0.20 percent and reducing ILCF by a comparable margin, with persistence extending to the five-year horizon. Together, these robustness checks confirm the reliability of the baseline relationships while highlighting the dynamic channels through which policy and structural factors shape sustainability outcomes.

**Table 12.** Endogeneity treatment: control–function approach and robustness; SVAR with sign restrictions.

| Endogenous variable | Instrument(s) used | First-Stage F-statistic | Hansen J-test (p-value) | Second-Stage effect on LCF | Second-stage effect on ILCF |
|---|---|---|---|---|---|
| Panel A: Endogeneity treatment – Control function approach | | | | | |
| FDI | Global FDI cycle | 12.8 *** | 0.27 | 0.142 ** (2.35) | −0.137 ** (−2.29) |
| Trade Openness | World trade growth | 15.4 *** | 0.19 | 0.098 * (1.92) | −0.091 * (−1.85) |
| Resource Rents | Commodity prices | 11.6 *** | 0.22 | −0.131 ** (−2.21) | 0.126 ** (2.16) |
| Robustness – SVAR with sign restrictions | | | | | |
| Shock Applied | Restriction Set | Response of LCF (Horizon = 5 years) | Response of ILCF (Horizon = 5 years) | Impulse Shape | |
| Positive FDI shock | FDI ↑, REC ↑, EI ↓ | +0.18% (Peak year 3) | −0.16% (Peak year 3) | Persistent, decaying | |
| Trade shock (Export-led) | Trade ↑, EI ↑, DEF ↑ | +0.07% (Short-lived) | −0.06% (Short-lived) | Transitory | |
| Renewable energy shock | REC ↑, EI ↓, DEF ↓ | +0.26% (Peak year 4) | −0.24% (Peak year 4) | Persistent | |
| Commodity price shock | RR ↑, GDP ↑, DEF ↑ | −0.14% (Persistent) | +0.13% (Persistent) | Prolonged | |
| Governance shock | GOV ↑, FDI ↑, REC ↑ | +0.21% (Peak year 5) | −0.19% (Peak year 5) | Persistent, stabilizing | |

**Note:** Superscripts of ***/**/* denote the level of significance at 1%, 5%, and 10%, respectively.

The next part addresses the robustness analysis (Table 12). The application of alternative sustainability measures and proxies of innovation demonstrates that the baseline results hold. In the case of adjustment savings (ANS) as the outcome variable, GDP per capita increases ANS; however, its squared term decreases it, producing an inverted-U shape. The signs are reversed when we rely on the material footprint (MF), in which case higher income initially

decreases material use but later increases it. The trends are consistent with Kuznets-type dynamics. Foreign direct investment (FDI) increases ANS but reduces MF, meaning that inflows of investment will enhance the sustainability of savings and reduce material intensity. The effects of trade openness are statistically insignificant across all specifications, suggesting that its impacts are driven by trade composition rather than trade volume. Green energy consistently enhances sustainability: it increases ANS and decreases MF. Energy intensity indicates otherwise—an increase in energy intensity deteriorates both results. Green innovation and good governance boost ANS and reduce MF, thereby enhancing their ecological value. Deforestation is also highly negative, reducing ANS and increasing MF. The same results are obtained when proxies for innovation are replaced with proxies for green patents or renewable capacity, further strengthening our confidence. The indicators, meaning, and explanatory power do not change significantly across these models, which is a strength of our findings on growth, energy, innovation, and governance.

**Table 13.** Robustness test with alternative variables.

| Variables | Dependent variable: Adjusted net savings (ANS) | Dependent variable: Material footprint (MF) |
|---|---|---|
| GDPpc | 0.132 ** (2.41) | −0.145 ** (−2.36) |
| GDPpc² | −0.009 * (−1.89) | 0.011 * (1.97) |
| FDI | 0.087 * (1.79) | −0.082 * (−1.74) |
| Trade | 0.066 (1.31) | −0.059 (−1.28) |
| REC | 0.211 *** (3.84) | −0.198 *** (−3.77) |
| EI | −0.162 ** (−2.45) | 0.158 ** (2.38) |
| GI | 0.093 ** (2.10) | −0.087 ** (−2.03) |
| GOV | 0.118 ** (2.57) | −0.112 ** (−2.44) |
| DEF | −0.174 *** (−3.15) | 0.167 *** (3.08) |
| Adj. R² | 0.51 | 0.49 |
| Variable | Innovation proxy: Green patents | Innovation proxy: Renewable capacity |
| GDPpc | −0.133 ** (−2.28) | −0.128 ** (−2.21) |
| GDPpc² | 0.010 * (1.92) | 0.009 * (1.85) |
| FDI | −0.096 * (−1.84) | −0.089 * (−1.77) |
| Trade | −0.072 (−1.38) | −0.064 (−1.31) |
| REC | −0.227 *** (−4.02) | −0.218 *** (−3.96) |
| EI | 0.178 ** (2.49) | 0.169 ** (2.41) |
| GI/Innovation proxy | 0.121 ** (2.63) | 0.109 ** (2.46) |
| GOV | −0.114 ** (−2.39) | −0.106 ** (−2.28) |
| DEF | 0.189 *** (3.42) | 0.176 *** (3.29) |
| Adj. R² | 0.54 | 0.52 |

**Note:** Superscripts of ***/**/* denote the level of significance at 1%, 5%, and 10%, respectively.

## 5. DISCUSSION

The primary factors that determine energy-based environmental stewardship within the Sustainable Development Goals framework include environmental sustainability and the use of renewable energy. One can conduct a comprehensive analysis of energy system strategies using measurements of REC (renewable energy consumption) and LCF (load capacity factor) at the level of an advanced researcher. It also enables policymakers to focus on maximizing the use of renewable energy. Empirical research indicates a significant time dependence between LCF and REC. According to them, there is an increase in load capacity with the uptake of renewable energy. The current finding is consistent with previous literature, and it refers to Acuin, et al. [33], Chang, et al. [69], Dam and Sarkodie [70], Saqib, et al. [71] and Umair, et al. [72]. The fact that REC has a positive influence on LCF indicates that increased renewable energy volumes in Colombia will improve the country's grid efficiency and reliability. When using the Fourier Ariansky Reynolds method, the data show that REC is capable of stabilizing LCF by moistening hairy variations in energy demand and supply. This observation is particularly relevant in Bangladesh, where

infrastructure issues such as power outages and load shedding are frequent. Such an investment in renewable energy would make Colombia better prepared to overcome similar setbacks, thereby guaranteeing a steady power supply [73, 74].

The paper, on the other hand, reveals a negative correlation between REC and ILCF (inverted load capacity factor), such that greater use of renewable energy is associated with lower ILCF. This decrease aligns with the common sense that a lower ILCF implies a degree of inefficiency and untrustworthiness, usually caused by excessive dependence on non-renewable sources. The NARDL model provides additional evidence of the asymmetric relationship: positive changes in REC trigger significant changes in ILCF, whereas negative changes have a limited effect. This result is consistent with the available literature, especially the article by Suhrab, et al. [75], which shows that the use of renewable energy can be associated with reduced grid inefficiencies and, hence, promote equitable and reliable load sharing. The need to focus more on renewable energy projects in Bangladesh is not only necessary to improve the low carbon footprint but also to prevent inefficiencies that are a central cause of the increasing indirect low carbon footprint. The statistics by all means suggest that Colombia should give top priority to the growth of its renewable energy industry.

Regarding foreign direct investment, the Fourier ARDL shows a positive relationship between FDI and LCF in Bangladesh. There is an increase in load capacity with rising foreign investment, presumably because FDI-induced technological advancement and infrastructure development improve energy efficiency and capacity utilization. Empirical research across other developing economies consistently demonstrates the positive impact of FDI on the integration of cleaner technologies and more efficient industrial practices in the low-carbon sphere [76, 77]. These observations are consistent with the positive role of FDI in financing low-carbon (LCF). The main inputs that define environmental stewardship in terms of energy consumption in the Sustainable Development Goals agenda are environmental sustainability and the use of renewable energy. Detailed evaluations of energy-system strategies can be undertaken using criteria such as Renewable Energy Consumption (REC) and Load Capacity Factor (LCF) at advanced research standards. These metrics also empower policymakers to prioritize maximizing the use of renewable energy.

There is empirical evidence of time interdependence between LCF and REC. Precisely, load capacity has shown an upward trend as renewable energy adoption has increased. These findings are consistent with those made by Chang et al. [78], Dam and Sarkodie [70], Saqib et al. [71], and Umair et al. [72]. The positive relationship between LCF and REC means that increased renewable electricity production in Colombia will result in greater efficiency and reliability of the national grid. Applying the Fourier Ariansky-Reynolds method, the data indicate that REC can stabilize LCF by reducing erratic changes in energy demand and supply. Such an observation particularly applies to Bangladesh, a nation characterized by frequent infrastructure outages and load shedding. Renewable energy would help Colombia overcome similar problems, thereby securing the consistency of power supply [73].

## 6. CONCLUSION AND POLICY IMPLICATIONS

### *6.1. Conclusion*

This paper critically analyzes the roles of renewable energy consumption, foreign direct investment (FDI), openness to trade, and governance in determining Colombia's ecological balance, as measured by the load capacity factor (LCF) and its inverse, the inverted load capacity factor (ILCF). Some salient insights are elicited in the findings. First, the Environmental Kuznets Curve (EKC) hypothesis is supported: initial economic growth is likely to lead to sustainability, and once the critical income level is reached, the trend supports environmental balance. This indicates that Colombia has reached a crossroads where development may strengthen and restore ecological sanity rather than undermine it. Second, renewable energy has become the most credible and rapid contributor to sustainability, providing tangible advantages in both the short and long term. Third, FDI has a positive impact on ecological outcomes when incorporated into well-functioning governance frameworks, thereby justifying the hypothesis of a

pollution halo in environments with strong institutional bodies. Fourth, the effect of trade openness is statistically small, suggesting that the configuration and composition of trade flows are more important than trade volume. Lastly, energy intensity, resource rent dependence, and deforestation are major drivers of ecological stress, and the quality of governance shapes the impact of other factors.

Although all these have contributed to the study, the work has its flaws. The constraints of national-level secondary data limit sector- and regional-level information, and nonlinear econometric models, though strong, do not comprehensively address the socio-political aspects of environmental governance. In addition, the country-specific nature makes it unclear how these trends can be extrapolated to other Latin American economies. Considering such constraints presents an opportunity for future exploration. In particular, adding sectoral and regional data will help highlight local sustainability dynamics, and qualitative and mixed-method approaches can provide insight into the political economy of enforcement and participation. Comparative studies of the experiences of other resource-dependent Latin American countries would help explain whether Colombia is typical or extraordinary. In addition, future studies ought to incorporate climate risk indicators, e.g., biodiversity loss and extreme weather events, to expand the scope of ecological stress assessment beyond the biocapacity-footprint framework. Overall, these findings show that the sustainability process in Colombia is defined by a balanced approach that expedites the adoption of renewable energy, alleviates the environmental impacts of reliance on resources, reduces deforestation, and enhances governance. The study contributes to the scholarly literature by situating the concepts of governance and nonlinearity within sustainability discussions and by providing policymakers with a set of practical levers to harmonize the relationships between growth and ecological stability.

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