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The Environmental Kuznets Curve in a Resource-Dependent Economy: Empirical Evidence from Algeria, 1970–2023

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ABSTRACT

This study examines the Environmental Kuznets Curve hypothesis in Algeria using an extended annual dataset and an autoregressive distributed lag framework (ARDL). The analysis considers economic growth together with energy use, urbanization, foreign direct investment, and inflation, with finite-sample critical values used for the ARDL bounds testing. Structural-break, stability, multicollinearity, and variable-exclusion checks are also conducted, while Toda–Yamamoto causality is used as the primary predictive-causality procedure. The results support the existence of a long-run relationship among the variables but do not confirm the conventional inverted-U Environmental Kuznets Curve. Income is positively associated with CO_2 emissions in the long run, whereas the quadratic income term is statistically insignificant, indicating no reliable evidence of the curvature required for an Environmental Kuznets Curve. Foreign direct investment is negatively associated with emissions, while inflation shows a positive long-run association. The individual effects of energy use and urbanization are less robust because of strong collinearity between the two variables. The causality analysis does not detect significant predictive causality between emissions and the macroeconomic variables under the preferred Toda–Yamamoto causality specification. Overall, the findings suggest that continued income growth alone is unlikely to generate an automatic decline in emissions in Algeria and that environmentl improvement will depend more directly on structural, energy, and environmental policy.

Keywords: *Algeria, Environmental Kuznets Curve, CO_2 emissions, ARDL bounds testing, Toda-Yamamoto causality*

JEL Classification: Q56, Q43, C32, O44, E3

INTRODUCTION

The relationship between environmental degradation and economic development continues to be one of the primary topics in development economics and environmental policies. The EKC hypothesis was put forward for the first time in the working paper written by Grossman and Krueger (1991). It was later theorized by Panayotou (1993), the author who provided the first theoretical explanation for the existence of an inverse U relationship between environmental degradation and income; when income rises, pollution increases until it reaches its peak at a certain level of income, after which it starts decreasing due to better regulation and technological changes. The EKC hypothesis was further examined by Stern (2004), the author who analyzed

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more than 80 studies on the topic and reached the conclusion that EKC is not universal because policies and institutions influence the decrease of pollution in terms of higher income levels. At the same time, the empirical studies regarding CO2 as an indicator of pollution have produced mixed results, especially in resource-dependent countries (Al-Mulali et al., 2015; Apergis & Payne, 2009), which is consistent with the findings by Badeeb, Lean & Shahbaz (2020) regarding the systematic weakening of curvature in resource-based countries as a result of natural resource dependence.

As one of Africa's largest economies and a major natural gas producer, Algeria's economic performance has been strongly influenced by volatility in the hydrocarbon industry. The Algerian economy has experienced several shocks, such as the oil price shock of 1986 and the period of civil strife between 1992 and 1996. Such shocks have affected the trend of greenhouse gas emissions in the country and its ability to implement reforms in environmental policy.

Despite the continued importance of this topic, the Algerian literature remains extensive but limited in terms of the methods used. Most studies in this area use a fairly limited number of variables, and among the Algeria-specific EKC studies reviewed for the present research, we didn't identify a study that includes inflation as an explanatory factor for CO_2 emissions. In addition, the vast majority of studies in this area apply ARDL bounds testing, where estimates are evaluated using the asymptotic critical values of Pesaran et al. (2001). The appropriateness of this approach can be questioned given the relatively small sample size available for Algeria. In turn, Chekouri et al. (2021) differ from these studies by using wavelet coherence and Breitung–Candelon causality rather than ARDL. The combination of ARDL bounds testing with Toda–Yamamoto causality is found in three studies, including the present study, with the former methodology applied using asymptotic critical values over relatively short time periods (1990–2019 and 1975–2014, respectively). It is also worth mentioning that the empirical results in this literature are quite heterogeneous. Some studies confirm the traditional Environmental Kuznets Curve (EKC), while others report EKC turning points above current income levels (Bouznit & Pablo-Romero, 2016; Amri, 2017). Furthermore, some studies find an N-shaped EKC based on the ecological footprint as the environmental indicator (Chaouachi & Balsalobre-Lorente, 2022; Shehzad et al., 2022; Malki et al., 2025).

The current study addresses these gaps directly. The annual dataset spans the period 1970–2023 and, to the best of our knowledge, represents the longest Algeria-specific sample used in the EKC literature reviewed. The study employs the ARDL bounds test of Pesaran et al. (2001) with finite-sample critical values from Kripfganz and Schneider (2020), incorporating seven variables, including inflation. The Toda-Yamamoto modified Wald causality test (Toda & Yamamoto, 1995) is adopted as the primary causal inference tool, with standard Granger causality (Granger, 1969) reported as a secondary robustness check.

This work is structured as follows: introduction; literature review, which provides a general background on research conducted on this subject and the contribution of this study; methodology, including the data used; results together with discussion of the results; and, lastly, conclusion about the major findings.

LITERATURE REVIEW

The Environmental Kuznets Curve: Theoretical Foundations

The Environmental Kuznets Curve (EKC) is defined as a nonlinear relationship between environmental degradation and income that exhibits increased environmental degradation at lower income levels and a change after income surpasses a threshold level, thus following an inverted-U shape. The terminology comes from Kuznets (1955), who observed a nonlinear relationship between income and inequality in the context of industrialization, an idea that was later expanded to the environmental sphere by economists. The use of the theory for explaining

environmental issues became common in the early 1990s when Grossman and Krueger (1991) analyzed the environmental impact of NAFTA (North American Free Trade Agreement) and found evidence of a nonlinear relationship between income and pollution. The processes underlying the EKC have been explained by Panayotou (1993).

Three effects are generally considered when explaining the EKC. The scale effect is related to the increase in output and energy consumption that accompanies economic development, especially at the earlier stages when production is relatively energy intensive. The composition effect refers to changes in the structure of the economy as development progresses, including a shift from energy-intensive activities towards services. The technique effect concerns changes in production methods, particularly the adoption of cleaner technologies as income rises. These effects are discussed in the EKC literature by Panayotou (1993), Grossman and Krueger (1995), and Stern (2004).

Stern (2004), in his review of the EKC literature, questioned the idea that higher income necessarily leads to better environmental outcomes. His assessment shows that the evidence for an inverted-U relationship varies considerably depending on the pollutant under investigation, the countries considered, and the way factors such as trade and institutional quality are incorporated into the analysis. The EKC appears to have stronger empirical support for some local pollutants, particularly sulfur dioxide, because their effects are experienced directly within the countries where they are emitted. This argument is less convincing for carbon dioxide. The effects of CO_2 emissions are global, meaning that the environmental costs are not confined to the country responsible for the emissions. As a result, domestic incentives to reduce them may be weaker. Dinda (2004), Tol (2009), and Mitić et al. (2019) similarly point to substantial variation in the empirical evidence rather than a universal income–environment relationship. More recent country-level studies also illustrate this point. For example, Mitić et al. (2024) apply an ARDL framework to Serbia and examine the EKC hypothesis within a particular national context. Taken together, these studies suggest that economic growth alone cannot be expected to produce environmental improvement. The trajectory of environmental quality is also shaped by policies, institutions, and the structural characteristics of the economy.

Empirical Evidence: Global and MENA (Middle East and North Africa) Studies

The MENA region comprises the countries of the Middle East and North Africa, including Algeria, and is characterized by substantial dependence on hydrocarbon production, energy-intensive economic activity, and widespread fossil-fuel subsidies, all of which are relevant to how the region's income–environment relationship is likely to unfold. Evidence for the Environmental Kuznets Curve varies considerably across regions. Some global studies find that CO2 emissions decline once an economy reaches a sufficiently high-income level, a pattern documented for G7 economies (Koirala et al., 2025). While broad cross-country studies generally support this pattern, findings differ according to model specification and country sample (Farooq et al., 2022), indicating that the outcome depends on local structural factors rather than a single global regularity.

Evidence for the MENA region is considerably more mixed: some studies find support for the EKC hypothesis, while others find little to no evidence of the characteristic inverted-U pattern (Kostakis et al., 2023). This divergence is plausibly linked to the region's reliance on oil and energy-intensive industries; higher energy consumption and inward foreign investment are both associated with higher emissions in economies where hydrocarbons dominate the economic structure (Gorus & Aslan, 2019). This is consistent with Badeeb, Lean and Shahbaz (2020) finding that natural resource dependence systematically weakens the curvature underlying the conventional EKC relationship, suggesting that a country's resource-dependence profile, rather than income growth alone, shapes its environmental trajectory.

The evidence is far from consistent across countries. In the MENA region, the EKC appears to depend on the structure of the economy and, in particular, on the importance of natural resources and energy in economic activity.

Energy Use, Urbanization, FDI, and Inflation as CO_2 Determinants

Energy Use

Energy use is universally recognized as the most proximate structural driver of CO_2 emissions in the empirical literature, and its inclusion in EKC models is standard. The foundational mechanism is direct: fossil fuel combustion for electricity generation, transportation, and industrial production constitutes the primary emission source in most economies (Apergis & Payne, 2009). Using a panel of six Central American countries, Apergis and Payne (2009) established that energy consumption Granger-causes CO2 emissions in the short run, while bidirectional causality prevails between energy use and emissions in the long run. Their panel cointegration framework confirmed that energy is not merely a correlated control variable but an independent causal driver whose dynamics directly shape the timing and magnitude of any EKC turning point. Beyond the volume of energy use, the composition of the energy mix matters. Stern (2004) argued that the technique effect of the EKC is contingent on economies shifting away from fossil-fuel-intensive production toward cleaner energy sources as income rises; without this shift, rising income cannot generate sufficient decarbonization to close the inverted-U.

In resource-rich economies where domestic energy is heavily subsidized as documented by the IMF across MENA and OPEC countries (Coady et al., 2017), the price signal necessary to incentivize efficiency investment is chronically suppressed, compounding the volume effect with a market-distortion effect that keeps emissions elevated even as income grows. For Algeria specifically, the energy–emissions nexus is structurally entrenched. As a major hydrocarbon producer and exporter, the economy's income dynamics are closely linked to fossil energy use, creating structural collinearity between GDP and energy consumption. Empirical evidence from the MENA region shows that energy consumption remains among the most significant and robust long-run determinants of CO_2 emissions even after controlling for income, urbanization, and trade openness (Al-Mulali & Ozturk, 2015). This collinearity is a central motivation for adopting the ARDL bounds testing approach in this study. Although the ARDL framework accommodates variables with different integration orders, it does not eliminate the identification and precision problems associated with severe multicollinearity; instead, it is able to identify long-run relationships even when the variables themselves are highly persistent.

Urbanization

Insufficient clarity exists regarding the link between urbanization and CO_2 emissions. In essence, urbanization entails the construction of infrastructure that is energy-intensive, and the presence of transportation and industrial activities in dense areas may result in pollution. The link has been shown in studies by Al-Mulali et al. (2013), in which an empirical study found a long-term relationship between urbanization, energy use, and pollution among 20 MENA countries from 1980 to 2009. At the same time, a more complicated relationship may exist, as shown by the inverse-U relationship observed by Martínez-Zarzoso and Maruotti (2011) in the case of 88 developing countries. The inverse-U relationship coincides with the theory of ecological modernization, according to which urbanization, once above a certain level, results in greater efficiency and reduced pollution. However, in the case of the MENA region, the positive relationship appears to prevail. Al-Mulali and Ozturk (2015) found a positive relationship between urbanization and environmental degradation in the MENA region. Further, there is no evidence of an inverse-U relation between urbanization and CO_2 emissions in the MENA region, as reported by Abdallh and Abugamos (2017). Therefore, the expected benefits of urbanization in

a rent-based rather than productivity-based economy such as Algeria may not come about because urbanization is mainly resource-induced.

Foreign Direct Investment

As for the effect of FDI on CO_2 emissions, two theories can be distinguished. First, there is the Pollution Haven Hypothesis (PHH), developed by Copeland and Taylor (1994). It claims that foreign investments go to countries with weak regulations related to environmental protection. As a result, polluting industries can be relocated to these countries, leading to increased levels of emissions. Secondly, the Pollution Halo Hypothesis (PLH) claims that FDI brings advanced technologies, improves the quality of regulation, and results in reduced emission intensity in the host country, which can contribute to the turning point of the EKC (Singhania & Saini, 2021). However, no clear consensus can be identified in the current literature on the topic. Singhania and Saini (2021), using the generalized method of moments (GMM) on a sample of 21 high-emitting countries, find support for the PHH, as they report a positive relationship between FDI and CO_2 emissions, especially for developing countries. On the contrary, Seker et al. (2015), using the ARDL bounds test for Turkey, found a negative long-term effect of FDI on emissions, thus supporting the PLH. Behera and Dash (2017) highlight that the nature of the relationship depends on the level of development, while Saqib et al. (2023) underline the significance of the investment sector composition, regulatory quality, and human capital in host countries. In oil-rich MENA countries, where FDI goes mainly to the energy sector, the PHH seems more applicable.

Inflation

The role played by inflation as a factor affecting the EKC has been discussed very little in the academic literature, despite its relevance to the topic. This research is aimed at addressing this omission. The theory operates on the premise of administered energy pricing, whereby the prices of fuel and electricity in nominal terms are fixed in economies in which the prices of these products are set by government entities. In such cases of energy pricing, sustained inflation causes a decrease in the real cost of energy, regardless of changes in subsidies that are typical of many MENA economies. For example, according to Coady et al. (2017), global fossil fuel subsidies in 2015 stood at around $5.3 trillion, with a considerable portion accounted for by MENA and OPEC countries.

Furthermore, beyond the use of subsidies, the macroeconomic instability arising from inflation may lead to shorter planning horizons among firms and hinder investment in durable capital, such as energy-efficient technologies and other technologies, thereby reducing the technique effect discussed in the Environmental Kuznets Curve (EKC) literature (Stern, 2004). While empirical studies, including that of Shahbaz et al. (2015) for Malaysia, have shown that macroeconomic conditions can significantly affect energy demand, the effect of inflation on the relationship between income, energy, and emissions remains under-researched. Methodologically, ignoring inflation as a variable related to income, energy use, and economic policies may introduce omitted variable bias and affect the estimated coefficients of the EKC drivers, as discussed in the multivariate analysis literature (Field, 2013; Hair et al., 2010).

Studies Specific to Algeria

Algeria has also been part of a number of regional panel studies. El Menyari (2021) looks at a group of four North African economies—Morocco, Algeria, Tunisia, and Egypt—over the period from 1980 to 2014 by employing second-generation panel techniques (CIPS, CADF, Westerlund cointegration, and Dumitrescu–Hurlin causality), and finds that electricity consumption leads to an increase in CO_2 emissions, whereas tourism has a reducing effect on them, with the EKC hypothesis being confirmed for this panel. Touitou (2021) does the same for seven North African

countries using a simple OLS model over the years 2000 to 2018 and finds that GDP per capita shows the effect of income on emissions in five out of the seven countries under review. More recently, Hamdi et al. (2026) applied a Pooled Mean Group ARDL model to five North African countries (Morocco, Algeria, Tunisia, Libya, and Egypt) for the period from 1980 to 2023, testing both the EKC and the pollution haven/halo hypotheses together using greenhouse gas emissions per capita. Notably, the quadratic GDP term in their study is positive rather than negative, and therefore does not support an inverted-U shape, which provides an independent and methodologically separate finding consistent with the absence of a confirmed EKC also identified in the current study, while their quadratic FDI term supports a pollution-halo interpretation. The panel evidence regarding Algeria is therefore mixed, since El Menyari (2021) finds support for an inverted-U pattern but Hamdi et al. (2026) do not, highlighting how sensitive EKC results are to both the composition of the sample and the specification, even within the same wider region. Neither of these two panel studies takes inflation into account as a covariate, and neither provides the same level of detail about Algeria as a focused single-country analysis, because the long-run coefficients are averaged across the diverse economies.

One of the earliest identified Algeria-specific attempts to test the EKC hypothesis is Toumache (2009), who estimates simple OLS regressions of CO_2 emissions per capita on GDP per capita and its square using only 14 annual observations (1993–2006), without unit-root or cointegration testing. The results contradict the sign pattern required for a genuine inverted-U relationship, which is a negative linear and positive quadratic coefficient, the reverse of the EKC prediction, and the quadratic term is statistically insignificant in two of the three specifications reported; the study's own conclusion explicitly calls for a more rigorous analysis using a longer time series. Empirical research on the Environmental Kuznets Curve in Algeria has grown substantially over the past decade, spanning several distinct methodological traditions, summarized below. Most early ARDL-based and VAR-based studies report mixed evidence, ranging from no confirmed turning point to a turning point far beyond Algeria's observed income range to outright rejection of the EKC hypothesis, while others confirm a standard inverted-U relationship (Lacheheb et al., 2015; Bouznit & Pablo-Romero, 2016; Amri, 2017; Belaid & Youssef, 2017; Touitou & Langarita, 2021; Touitou, 2021). A second group of studies uses ecological footprint rather than CO2 emissions as the environmental measure and generally reports an N-shaped rather than inverted-U relationship (Shehzad et al., 2022; Chaouachi & Balsalobre-Lorente, 2022; Malki et al., 2025), with the latter two emphasizing that renewable energy adoption and subsidy reform, not income growth alone, ultimately reduce environmental degradation in Algeria. Chekouri et al. (2021) took a markedly different approach, using wavelet coherence rather than an ARDL framework, finding that the growth-emissions relationship varies across time scales. Most recently, Dekkiche and Mezouri (2022) and Elbadri et al. (2023) applied a Toda-Yamamoto causality framework alongside ARDL bounds testing and a Fourier-Bootstrap ARDL approach, respectively, reaching opposite conclusions on long-run equilibrium.

Across this considerable and methodologically diverse literature, three limitations recur. First, every study identified omits inflation as a covariate, despite Algeria's documented history of energy subsidy reform and inflationary volatility, both plausibly correlated with income, energy use, and emissions. Second, with the exception of Chekouri et al. (2021), every ARDL-based study compares its bounds-test statistic only against the asymptotic critical values of Pesaran et al. (2001), which can overstate evidence for cointegration at the moderate sample sizes available for Algeria (Kripfganz & Schneider, 2020). Third, of the two studies employing Toda-Yamamoto causality testing (Dekkiche & Mezouri, 2022; Chaouachi & Balsalobre-Lorente, 2022), both rely on samples ending by 2019 at the latest, missing the post-2019 period entirely.

Table 1.Summary of Algeria-Specific EKC Studies

Study	Period	Environmental measures	Method	EKC result
Toumache (2009)	1993-2006	CO_2	OLS (no cointegration testing)	Rejected; wrong-signed, largely insignificant
Lacheheb et al. (2015)	1971-2009	CO_2	ARDL bounds testing	No turning point; unstable results
Bouznit & Pablo-Romero (2016)	1970-2010	CO_2	ARDL bounds testing	Confirmed; turning point beyond observed range
Amri (2017)	1980-2011	CO_2	ARDL with structural breakpoints	Confirmed
Belaid & Youssef (2017)	1980-2017	CO_2	ARDL bounds+ZA unit-root	Rejected; growth detrimental
Touitou & Langarita (2021)	1973-2016	CO2	ARDL bounds testing	Confirmed; standard inverted U
Touitou (2021)	1980-2017	CO_2, SO_2, NO_X	VAR+cointegration	Confirmed; standard inverted U
Shehzad et al. (2022)	1970-2017	Ecological footprint	ARDL	N-shaped
Chaouachi & Balsalobre-Lorente (2022)	1975-2014	Ecological footprint	ARDL+FMOLS/DOLS+Toda-Yamamoto	N-shaped
Chekouri et al. (2021)	1971-2018	CO_2	Wavelet coherence	Relationship varies by time scale
Dekkiche & Mezouri (2022)	1990-2019	CO_2	ARDL bounds +Toda-Yamamoto	Long-run equilibrium confirmed
Elbadri et al. (2023)	1990-2018	CO_2	Fourier-Bootstrap ARDL	Rejected
Bekhti et al. (2024)	1962-2014	CO_2	ARDL bounds testing	Confirmed; standard inverted U
Benatek & Akriche (2024)	1990-2019	CO_2	Engle-Granger cointegration	Rejected
Malki et al. (2025)	1990-2020	Ecological footprint	ARDL bounds testing	N-shaped
Current study	1970-2023	CO_2	ARDL bounds (finite-sample CVs)+GTS+Granger+Toda-Yamamoto	No statistically significant quadratic curvature

Source: Elaborated by the authors.

Contributions

This study contributes to the literature in four respects. First, it provides updated empirical evidence on the income–environment relationship in Algeria using a relatively long annual

sample, thereby reassessing the conventional EKC hypothesis in a resource-dependent economy. Second, the analysis incorporates urbanization, FDI, and inflation, allowing the environmental implications of broader macroeconomic conditions to be examined within a unified long-run framework. Third, the study uses finite-sample critical values from Kripfganz and Schneider (2020), and subjects the principal findings to extensive specification, multicollinearity, stability, structural-break, and variable-exclusion sensitivity analyses, providing a more cautious assessment of the robustness of the estimated income–emissions relationship. Fourth, the study complements the long-run ARDL analysis with predictive-causality evidence, while explicitly distinguishing cointegration, adjustment dynamics, parameter stability, and predictive causality.

DATA AND METHODOLOGY

Data Sources and Variable Definition

All series are drawn from two sources, the World Bank's World Development Indicators (WDI) database (World Bank, 2026) and the GLOBALEconomy database (Global Economy Database, 2026) for Algeria, 1970–2023. Table 2 reports details about each variable.

Table 2. Data Sources, Definitions and Transformations

Variable	Source	Indicator Code	Definition	Unit	Transformation
CO_2 emissions per capita	(Global Economy Database, 2026)	//	Annual CO_2-equivalent emissions per capita from agriculture, energy, waste, and industry, excluding LULUCF.	Metric tons per capita	Natural log
GDP per capita	(World Bank, 2026)	NY.GDP.PCAP.KD	Gross domestic product divided by population, expressed in constant 2015 U.S. dollars to remove the effects of inflation.	Constant 2015 US $	Natural log; mean-centered For estimation
GDP per capita squared	(World Bank, 2026)	Derived from NY.GDP.PCAP.KD	Square of centered ln(GDP per capita), included to test for EKC curvature.		Centered natural log of GDP squared
Energy use per capita	(Global Economy Database, 2026)	//	Use of primary energy before transformation to other end-use fuels.	Kg of oil equivalent per capita	Natural log
Urban population	(Global Economy Database, 2026)	//	Share of population living in areas classified as urban by	% of total population	Natural log

Variable	Source	Indicator Code	Definition	Unit	Transformation
			national statistical offices.		
Foreign direct investment (FDI)	(Global Economy Database, 2026)	//	Net inflows of investment	% of GDP	Level (not logged)
Inflation	(World Bank, 2026)	FP.CPI.TOTL.ZG	Annual percentage change in the cost to the average consumer of a fixed basket of goods and services	Annual %	Level (not logged)

Note: the access date for all variables was June 2025.

Source: Elaborated by the authors.

Descriptive Statistics

Descriptive statistics are an essential step before conducting any statistical or econometric modeling; they provide a comprehensive overview of parameters of the data before any multivariate analysis (Hair et al., 2010). This process helps identify anomalies such as missing observations or outliers, which can affect the accuracy of subsequent estimates (Field, 2013). Table 3 reports the descriptive statistics of the seven variables used in the ARDL time-series analysis. The study covers the period from 1970-2023, giving 54 annual observations, which is reduced to 52 once the lags required to run the model are taken into account.

Table 3. Descriptive Statistics

Variable	Var.used	N	Missing	Mean	Median	SD	Min	Max
ln(CO_2)	Used	54	0	1.098	1.087	0.213	0.358	1.43
ln(GDP) raw	-	54	0	8.250	8.241	0.150	7.804	8.470
ln(GDP) centered	Used	54	0	0.000	-0.009	0.150	-0.446	0.220
ln(GDP)^2	Used	54	0	0.022	0.016	0.031	0.000	0.199
ln(Energy use)	Used	54	0	6.686	6.785	0.503	5.504	7.31
ln(Urban)	Used	54	0	3.886	3.896	0.224	3.418	4.22
FDI	Used	54	0	0.670	0.530	0.609	-0.290	2.71
Inflation	Used	54	0	8.787	6.050	7.445	0.340	31.67

Note: ln(GDP) reported above in both raw and centered form. For estimation, the mean centered ln(GDP) and its squared term were used in all the subsequent tests. The use of centered variables is to reduce multicollinearity between the linear and quadratic income terms. In this paper, we consider ln(GDP)^2 =(ln(GDP))^2.

Source: Prepared by the authors based on Python results.

There is considerable variation in the data over the study period (1970–2023). In particular, ln(CO_2) has a mean of 1.098 (SD = 0.213), whereas the centered ln(GDP) has a mean of 0.000 and an SD of 0.150. Energy consumption has a mean of 6.686 (SD = 0.503), while urbanization has a mean of 3.886 (SD = 0.224). Both variables show a generally increasing trend over the period, although energy consumption varies more. FDI is more variable, with a mean of 0.670 (SD = 0.609)

and values ranging from −0.290% to 2.710% of GDP. Finally, the average inflation rate is 8.787% (SD = 7.445%).

ARDL Model Specification

The empirical analysis employs the ARDL bounds testing approach of Pesaran et al. (2001). The method can be used when the variables are I(0) or I(1), which is relevant for the variables considered in this study. It is also suited to the relatively small sample used here. The ARDL framework allows the short-run dynamics and the long-run equilibrium relationship to be estimated within the same equation. The bounds F-test is used to test for a long-run relationship. Lag selection for the short-run dynamics uses a general-to-specific approach, starting with a general specification allowing up to two lags for each short-run term. Insignificant terms are sequentially removed at the 10% level based on their p-values, while the complete long-run levels relationship with all seven variables at lag 1 is retained. Two short-run terms that do not meet the criterion are nevertheless kept in the final specification, as will be explained in the present section. The model is estimated over 1972–2023 (T = 52). Equation (1) represents the ARDL specification for the EKC model.

$$lnCO2_t = \alpha_0 + \alpha_1 lnCO2_{t-1} + \sum_{j=0}^{2} \beta_{1j}\, lnGDP_{t-j} + \sum_{j=0}^{2} \beta_{2j}\, (lnGDP^2)_{t-j} + \sum_{j=0}^{2} \beta_{3j}\, lnEnergy_{t-j} + \sum_{j=0}^{1} \beta_{4j}\, lnUrban_{t-j} + \sum_{j=0}^{1} \beta_{5j}\, FDI_{t-j} + \sum_{j=0}^{2} \beta_{6j}\, Inflation_{t-j} + \varepsilon_t \quad (1)$$

Where:

- Ln(GDP) and ln(GDP)²: the EKC core: income and its square to test the inverted-U;
- Ln(Energy): energy consumption;
- Ln(Urban): Urbanization;
- FDI: foreign direct investment, included in levels;
- Inflation: at level;
- ε_t: the error term.

The unrestricted error-correction form of the ARDL model is specified as follows:

$$\Delta lnCO2_t = \alpha_0 + \lambda lnCO2_{t-1} + \theta_1 lnGDP_{t-1} + \theta_2 (lnGDP^2)_{t-1} + \theta_3 lnEnergy_{t-1} + \theta_4 lnUrban_{t-1} + \theta_5 FDI_{t-1} + \theta_6 Inflation_{t-1} + \sum_{i=1}^{1} \gamma_i\, \Delta lnCO2_{t-i} + \sum_{j=0}^{1} \delta_{1j}\, \Delta lnGDP_{t-j} + \sum_{j=0}^{1} \delta_{2j}\, \Delta (lnGDP^2)_{t-j} + \sum_{j=0}^{1} \delta_{3j}\, \Delta lnEnergy_{t-j} + \delta_4 \Delta lnUrban_t + \delta_5 \Delta FDI_t + \sum_{j=0}^{1} \delta_{6j}\, \Delta Inflation_{t-j} + \varepsilon_t \quad (2)$$

The existence of a long-run relationship is tested using the bounds testing procedure. The null hypothesis of no cointegration is:

$H_0: \lambda = \theta_1 = \theta_2 = \theta_3 = \theta_4 = \theta_5 = \theta_6 = 0$

Against the alternative:

$$H_1: \lambda \neq 0, \theta_i \neq 0$$

The computed F-statistic is compared against the finite-sample critical values provided by Kripfganz and Schneider (2020), which offer more accurate inference.

Once cointegration is established, the long-run coefficients are derived as:

$$\beta_k^{LR} = -\frac{\theta_k}{\lambda}$$

Where:

- λ**:** coefficient of $CO2_{t-1}$;
- θ_k: coefficients of lagged explanatory variables.

The short-run dynamics and speed of adjustment are captured through the error correction model:

$$\Delta lnCO2_t = \alpha_0 + \sum short\text{-}run\ terms + \phi ECT_{t-1} + \varepsilon_t$$

Where:

- ECT_{t-1}: error correction term derived from long-run equilibrium;
- ϕ: speed of adjustment coefficient (expected to be negative and significant).

Lag selection for the short-run dynamics uses a general-to-specific approach. Starting with a general specification allowing up to two lags for each short-run term, insignificant terms are sequentially removed at the 10% level based on their p-values, while the full long-run relationship with all seven variables at lag 1 is retained. The final specification includes contemporaneous Δln(GDP), the first lag of Δln(GDP)^2, contemporaneous Δln(Energy), and the first lags of ΔFDI and ΔInflation. All other short-run terms, including urbanization, are excluded.

The 10% threshold was used as a guide, not as a rigid rule. Two terms with p-values above 0.10 remain in the final specification (Table 9): the lagged change in ln(GDP)^2 (p = 0.122) and the contemporaneous change in ln(Energy) (p = 0.716).

The squared income term was borderline significant during the reduction procedure (p = 0.094 after dropping Δln(Energy)), and its p-value is 0.122 in the final specification, which includes Δln(Energy). It was retained to preserve both income terms in the short-run specification and allow their joint contribution to be assessed, since the shape of a quadratic relationship depends on both the linear and squared terms (Lind & Mehlum, 2010). Δln(Energy) was kept on theoretical rather than statistical grounds, given the established role of energy consumption as a determinant of CO_2 emissions and environmental degradation (Apergis & Payne, 2009; Al-Mulali & Ozturk, 2015), so its short-run dynamics are retained regardless of their significance. Retaining insignificant short-run terms does not affect the consistency of the estimates, although it may cost some efficiency. In general-to-specific modelling, incorrectly excluding a relevant variable can be more problematic than retaining an irrelevant one (Hendry & Krolzig, 2005).

As a sensitivity check, the model was re-estimated with the 10% rule applied strictly, which removes Δln(Energy). The long-run income results do not change: ln(GDP) stays positive and significant, and ln(GDP)^2 stays insignificant. The long-run coefficients of energy use and urbanization, however, become significant at the 5% level (p = 0.041 and p = 0.035). This is consistent with the strong collinearity between the two variables reported in the robustness analysis. Residual diagnostics and parameter-stability tests for the final specification are reported in Table 10, with the estimation results reported in Tables 8–10.

RESULTS AND DISCUSSIONS

Structural Break Tests

Before conducting the conventional stationarity tests, structural breaks are examined because the 1970–2023 Algerian sample covers major oil-price shocks, civil conflict, economic reforms, and the COVID-19 shock. The Zivot and Andrews (1992) unit-root test is employed to allow for one endogenous structural break, complementing the DF-GLS and ADF tests by evaluating the unit-root hypothesis in the presence of a potential break.

Based on Table 4, the ZA test does not reject the unit-root hypothesis for CO_2, energy use, urbanization, FDI, or inflation at the 5% level. These variables therefore remain non-stationary in levels even when a possible break is allowed for. The outcome is different for ln(GDP) and $\ln(GDP)^2$, for which the unit-root hypothesis is rejected. The estimated break dates are 2002 for ln(GDP) and 1979 for $\ln(GDP)^2$. The former falls around the end of the civil conflict and the subsequent increase in hydrocarbon investment. The latter predates the 1986 oil-price collapse and points to a change in the behavior of the income-squared term at an earlier stage. The first-difference tests, reported separately, confirm the I(1) classification of the non-stationary variables. These break dates are used as diagnostic information and as part of the robustness assessment, but do not introduce break dummies into the baseline ARDL model (Table 4; Figures 2–3).

Table 4. Zivot and Andrews (1992) Structural Break Unit Root Test Results

Variable	ZA Statistic	5% crit. Value	p-value	Lags	Break Year	interpretation
$\ln(CO_2)$	-4.942	-5.073	0.073	0	1986	Unit-root (Borderline)
ln(GDP)	-5.808***	-5.073	0.005	1	2002	Break-stationary
$Ln(GDP)^2$	-7.583***	-5.073	0.001	0	1979	Break-stationary
ln(Energy use)	-4.149	-5.073	0.406	0	1979	Unit-root
ln(Urban)	-4.663	-5.073	0.149	1	1991	Unit-root
FDI	-3.346	-5.073	0.877	0	1996	Unit-root
Inflation	-4.745	-5.073	0.123	0	1997	Unit-root

*Notes: The Zivot–Andrews test allows for one endogenous structural break under the alternative hypothesis. The null hypothesis is a unit root without a structural break. Rejection indicates stationarity around a break. ***, *, and * denote significance at the 1%, 5%, and 10% levels, respectively.*

Source: Prepared by the authors based on Python results.

Second, the Quandt–Andrews (Figure 1) sup-F test detects a statistically significant potential parameter shift in the long-run EKC relationship, with sup-F = 78.14 and an estimated break in 1982. However, because several regressors are I(1) and the conventional critical values assume stationary regressors, this result is interpreted as suggestive evidence of a possible regime change rather than definitive evidence of parameter instability (Hansen, 2000). The subsequent CUSUM tests do not indicate parameter instability.

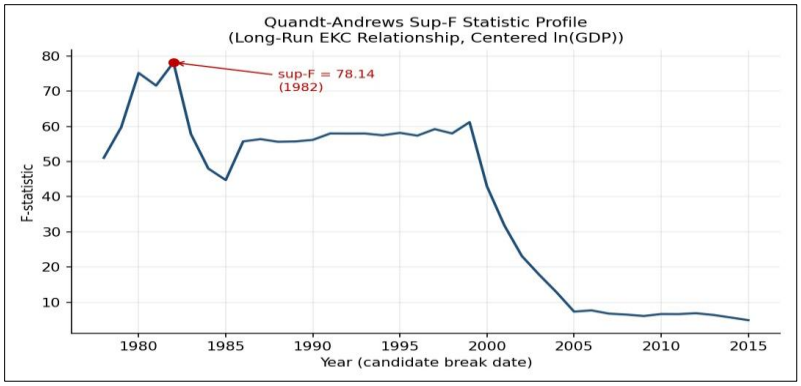

Figure 1. Quandt-Andrews Sup-F Statistic Profile for the Long-Run EKC relationship

Source: Generated by the authors using Python.

As a break-robust check (Table 5), the Gregory–Hansen test is applied under level-shift, level-shift-with-trend, and regime-shift specifications. These specifications allow for a structural change in the intercept and, depending on the model, in the trend and slope coefficients. All three specifications fail to reject the null of no cointegration, with ADF* statistics of –3.784, –4.118, and –4.532, compared with 5% critical values of –4.92, –5.29, and –5.50, respectively. Thus, the Gregory–Hansen test does not provide evidence of cointegration in the GDP–CO_2 relationship when a single endogenous regime shift is allowed. This does not necessarily contradict the multivariate ARDL bounds result because the two procedures test relationships under different model specifications. The Gregory–Hansen test evaluates a GDP–CO_2 relationship with a single endogenous break, whereas the ARDL bounds test evaluates the conditional long-run relationship within the full multivariate model. The Gregory–Hansen test results therefore provide a break-robust rather than a direct test of the multivariate ARDL cointegration relationship. Taken together with the CUSUM results, the Gregory–Hansen test provides additional diagnostic evidence on the stability and specification of the estimated long-run relationship.

Table 5. Gregory & Hansen (1996) Cointegration Test with Regime Shift (Core Income-Emission Relationship)

Model	ADF* Statistic	5% Crit. Value	1% Crit. Value	Break Year	Decision
Level shift (C)	-3.784	-4.92	-5.44	2015	Fail to reject H_0
Level shift +trend(C/T)	-4.118	-5.29	-5.80	2003	Fail to reject H_0
Regime shift (C/S)	-4.532	-5.50	-5.97	1995	Fail to reject H_0

Note: H_0 denotes no cointegration with a regime shift. The ADF statistic is compared with the Gregory–Hansen critical values; Rejection of H_0 indicates cointegration with a structural break.*

Source: Elaborated by the authors based on Python results.

Unit Root Test

Establishing the order of integration of the variables is necessary before applying the ARDL bounds test. The DF-GLS test of Elliott et al. (1996) is used as the main test because it generally performs better than the standard ADF test in small samples. The Schwarz Information Criterion (SIC) is used to select the lag length, with a maximum of five lags. The ADF test is also reported as a check, especially where the DF-GLS results are close to the critical values. All tests include a constant and a linear trend.

The results in Table 6 show that six of the seven variables are non-stationary in levels but become stationary after first differencing. This gives an I(1) classification for ln(CO_2 per capita), ln(GDP per capita), ln(Energy use), ln(Urban population), FDI, and inflation. The result for ln(GDP per capita)^2 is different. Its DF-GLS statistic at level is −4.038 (p = 0.001), leading to rejection of the unit-root hypothesis and its classification as I(0). The borderline results for ln(CO_2 per capita) and ln(Urban population) are also supported by the ADF test.

Table 6. DF-GLS Unit Root Test Results

	Level			First Difference			Final
Variable	**Stat.**	**Result**	**Order**	**Stat.**	**Result**	**Order**	**Order**
ln(CO_2)	−1.188	Non-stat.	I(1)	−3.112†	Stat.	I(0)	I(1)
ln(GDP)	−1.661	Non-stat.	I(1)	−4.141***	Stat.	I(0)	I(1)
ln(GDP)^2	−4.038***	Stat.	I(0)	−5.112***	Stat.	I(0)	I(0)
ln(Energy use)	−0.294	Non-stat.	I(1)	−3.595**	Stat.	I(0)	I(1)
ln(Urban)	−3.758*	Border.	I(1)	−3.322†	Stat.	I(0)	I(1)
FDI	−2.404	Non-stat.	I(1)	−5.624***	Stat.	I(0)	I(1)
Inflation	−2.403	Non-stat.	I(1)	−5.690***	Stat.	I(0)	I(1)

*Notes: ***, **, * denote rejection at 1%, 5%, 10%. † Borderline, confirmed by ADF. All tests include constant and trend. Lag length by SIC, max 5 lags (Elliott et al., 1996). Because unit-root tests with an intercept are invariant to additive shifts, the result for ln(GDP) is unaffected by centering; this invariance does not extend to the squared term, since squaring is a nonlinear transformation.*

Source: Prepared by the authors based on Python results.

The findings from the unit-root test show a mixture of I(0) and I(1) processes but do not support the existence of an I(2) process. This finding fulfills the integration-order assumption of the ARDL approach by Pesaran et al. (2001). In the EKC framework, ln(GDP) is demeaned before forming its squared term. The demeaning process does not alter the order of integration of ln(GDP). The squared term included in the model is categorized as I(0) according to the DF-GLS test results. This classification is based on the formed series and is used here to check whether the variables meet the requirements of the ARDL approach.

Bounds Test for Cointegration

Table 7 shows the ARDL bounds test for a long-run levels relationship. The critical values come from Kripfganz and Schneider (2020), who provide values based on response-surface simulations. These values are chosen here because the sample size is relatively small. The model is estimated under Case III with an intercept and no deterministic trend. Lag selection follows a general-to-specific procedure that keeps all long-run variables and removes short-run terms that are insignificant at the 10% level, with two documented exceptions (see ARDL Model Specification section). The F-statistic is 5.827, which is higher than the 5% I(1) bound of 4.197, while the t-statistic of –5.671 is also lower than its 5% I(1) bound of –4.359. Both statistics therefore lead to rejection of the hypothesis of no long-run levels relationship at the 5% level.

Table 7. ARDL Bounds Test for Cointegration

Test	Stat.	I(0) 10%	I(1) 10%	I(0) 5%	I(1) 5%	Decision
F-statistic (PSS, 2001)	5.827	2.238	3.594	2.671	4.197	Reject H_0 at 5%
t-statistic (PSS, 2001)	−5.671	−2.457	−3.929	−2.814	−4.359	Reject H_0 at 5%

Notes: Critical values from Kripfganz & Schneider (2020), finite-sample (K= 6, T= 52). PSS = Pesaran, Shin & Smith (2001). H_0: no long-run level relationship.

Source: Prepared by the authors based on Python results.

The results provide evidence of a long-run relationship among ln(CO_2 emissions per capita), ln(GDP), ln(GDP)^2, ln(energy use), ln(urbanization), FDI, and inflation in Algeria. The unit-root tests reported above classify the regressors as a mixture of I(0) and I(1), with no I(2) variable identified. The integration-order requirement for the ARDL bounds approach is therefore satisfied. The I(0) classification of the ln(GDP)^2 term is based on the DF-GLS result for the centered series used in the estimation.

Long-Run Coefficients

Table 8. ARDL Long-Run Coefficients and Error Correction Term

Variable	Coeff.	Std. Err.	t-stat	p-value	Sig.
ln(GDP)	1.263	0.198	6.39	0.000	***
ln(GDP)^2	-0.053	0.626	-0.09	0.933	
ln(Energy use)	-0.355	0.227	-1.56	0.126	
ln(Urban)	0.880	0.445	1.98	0.055	*
FDI	-0.108	0.032	-3.33	0.002	***
Inflation	0.006	0.002	2.24	0.031	**
ECT (adj. speed)	-0.593	0.105	-5.67	0.000	***

****, **, * denote significance at 1%, 5%, and 10%. General-to-specific ARDL specification, sample 1972–2023 (T = 52), lag order 1 for all long-run terms. R^2 = 0.658. Critical values from Kripfganz & Schneider (2020).*
Source: Prepared by the authors based on Python results.

Error correction term

The error-correction term ECT was found to be negative and highly significant (−0.593, $p < 0.001$). The size of the coefficient implies that about 59% of the imbalance is corrected within one year. In addition, the half-life of a short-run imbalance is around 0.77 years, which suggests that departures from the long-run path are relatively short-lived.

EKC Hypothesis

The theory of Environmental Kuznets Curves (EKC) developed by Grossman and Krueger (1991, 1995) claims an inverted U-shaped connection between GDP per capita and environmental degradation. This means that in quadratic estimation, the coefficient of ln(GDP) should be positive, and the coefficient of ln(GDP)^2 should be negative.

The long-run estimates give a positive and significant coefficient for ln(GDP) (1.263, $p < 0.001$). The coefficient on ln(GDP)^2 is negative (−0.053), but its p-value is 0.933. Hence, emissions increase with income in the estimated long-run relationship, but there is no statistical evidence of the curvature needed for an inverted-U pattern. On this basis, the EKC hypothesis is not supported.

Since the quadratic term is not statistically significant, the turning-point calculation has little meaning. The results show a relationship between income and emissions, but no statistically significant curvature is found. The structural-break results give some context to this finding. Breaks in income are identified in 1979 and 2002, corresponding to periods of oil price shocks and changes in the Algerian economy.

Other Long-Run Determinants

Looking at the other variables, inflation, FDI, and urbanization are significant in the long-run relationship with CO_2 emissions. Inflation has a positive coefficient ($\beta = 0.006$, $p = 0.031$), whereas FDI has a negative one ($\beta = -0.108$, $p = 0.002$). Urbanization is also positive and significant at the

10% level (β = 0.880, p = 0.055). Energy consumption, however, is not statistically significant (β = −0.355, p = 0.126). Its coefficient therefore does not provide evidence of an independent long-run effect once the other variables are taken into account. The inflation coefficient also differs between the short and long runs, as shown in Table 9 in the next section.

Short-Run Dynamics

The short-run estimates in Table 9 show that changes in income are associated with CO_2 emissions. The coefficient on D.ln(GDP) is 0.947 and statistically significant (p = 0.006), while the lagged squared-income term is negative but not significant (−0.590, p = 0.122). Although the two income terms are jointly significant (F = 4.185, p = 0.023), this result is mainly driven by the linear income term. The short-run results therefore do not provide clear evidence of nonlinear curvature and are better described as a positive rather than an inverted-U relationship. Urbanization is not included in the final short-run specification because its differenced terms were removed during the general-to-specific reduction for not meeting the 10% retention criterion, although its long-run level coefficient remains significant at the 10% level. The lagged change in FDI is positive and significant (0.060, p = 0.018), while the lagged inflation effect is negative and significant (−0.004, p = 0.011). Energy use has no significant short-run effect (0.259, p = 0.716). Both this term and the lagged squared-income term were retained although their p-values exceed the 10% threshold, for the reasons given in the ARDL Model Specification section.

Table 9. Short-Run ARDL Coefficient Estimates

Variable	Coeff.	Std. Err.	t-stat	p-value	Sig.
D.ln(GDP)	0.947	0.327	2.89	0.006	***
LD.ln(GDP)^2	-0.590	0.373	-1.58	0.122	
D.ln(Energy)	0.259	0.708	0.37	0.716	
LD.FDI	0.060	0.024	2.46	0.018	**
LD.Inflation	-0.004	0.002	-2.67	0.011	**

*Notes: ***, **, * denote significance at 1%, 5%, and 10% levels. D. = first difference; LD. = lagged first difference. The terms LD.ln(GDP)^2 (p=0.122) and D.ln(Energy) (p=0.716) exceed the 10% retention threshold and are retained as explained in the ARDL Model Specifications section.*

Source: Prepared by the authors based on Python results.

Diagnostics and Stability Tests of Residuals

The Breusch–Godfrey test does not reject the null of no serial correlation at the 5% level (p = 0.055), although the result is close to the 5% threshold and becomes significant at 10%. With only 52 observations, this result should therefore be treated with some caution. The restricted White test finds no evidence of heteroskedasticity (p = 0.694), while the full White test could not be estimated given the sample size. The ARCH test also finds no evidence of ARCH effects (p = 0.972), and the Ramsey RESET test does not suggest a functional-form problem (F(2,37) = 0.154, p = 0.858). The residuals are not normally distributed ($\chi^2(2) = 8.196$, p = 0.017). This does not affect coefficient consistency, although it can make inference less reliable in a small sample. Finally, both stability tests remain within their 5% critical values: the recursive CUSUM statistic is 0.700 against 0.948, while the maximum deviation from the CUSUM of Squares test is 0.134 against 0.277.

Table 10. Residual Diagnostic and Parameter Stability Tests

Test	Statistic	p-value	Threshold	Conclusion
Breusch-Godfrey LM (Autocorrelation)	$\chi^2(2) = 5.810$	0.055	> 0.05	Borderline
White-type Test, squares only (Heteroskedasticity)	$\chi^2(24) = 20.058$	0.694	> 0.05	None
Skewness-Kurtosis (Normality)	$\chi^2(2) = 8.196$	0.017	< 0.05	Non-normal
ARCH LM Test	$\chi^2(2) = 0.057$	0.972	> 0.05	No evidence
Ramsey RESET	$F(2,37) = 0.154$	0.858	> 0.05	No evidence
ARDL Recursive CUSUM	Stat. = 0.700	—	CV = 0.948	Stable
CUSUM of Squares	Max dev. = 0.134	—	CV = 0.277	Stable

Notes: CUSUM tests are based on the ARDL specification. "None" indicates no detected problem. The reported White-type test includes squared terms only, and excludes cross-products because the full specification is not estimable with T = 52.

Source: Prepared by the authors based on Python results.

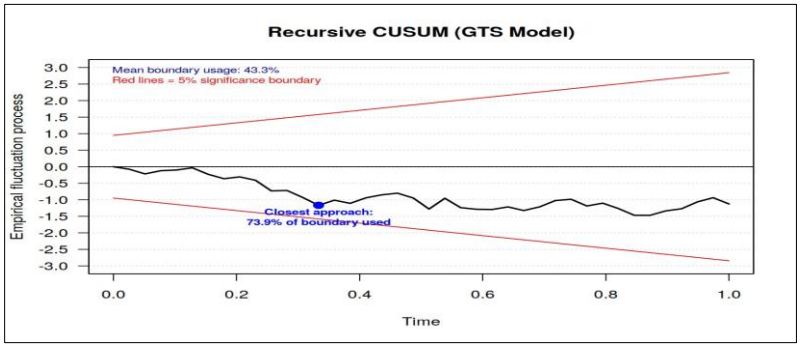

Figure 2. ARDL Recursive CUSUM

Source: Generated by the authors using Python.

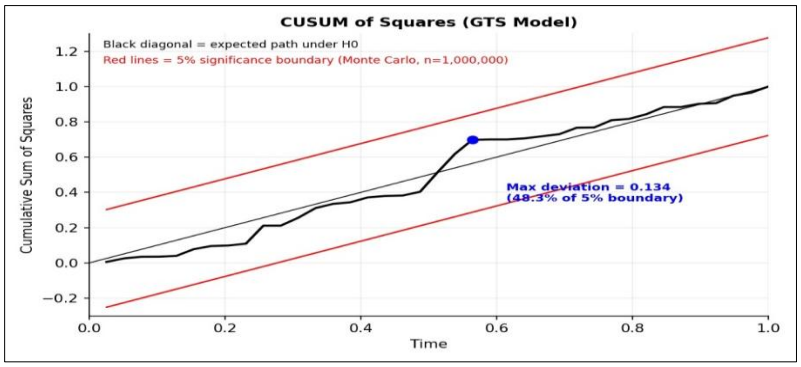

Figure 3. CUSUM of Squares

Source: Generated by the authors using Python.

Short Robustness Checks: Multicollinearity and Model Specification

Two specification issues are addressed. First, ln(GDP) is mean-centered prior to squaring to reduce collinearity between the linear and quadratic income terms. This produces variance inflation factors (VIFs) of 4.13 for ln(GDP) and 1.91 for $\ln(GDP)^2$ (Table 11), substantially reducing collinearity between the two income variables. This specification change does not alter the predicted values, residuals, or significance test of the squared term. Second, ln(Energy) and ln(Urbanization) are strongly correlated (r = 0.966), with VIFs of 42.36 and 44.65, respectively.

Their estimated coefficients may therefore be sensitive to the joint inclusion of both variables in the model.

Table 11. Variance Inflation Factors: Centered Specification

Variable	VIF
ln(GDP)	4.13
ln(GDP)²	1.91
ln(Energy)	42.36
ln(Urban)	44.65
FDI	2.20
Inflation	1.54

Source: Prepared by the authors based on Python results.

Table 12. Sensitivity Analysis (Energy and Urbanization Dropped in Turns)

Variable	Model 1: Full β (p)	Model 2: Energy dropped β (p)	Model 3: Urban dropped β (p)
ln(GDP)	1.263 (0.000)	1.429 (0.000)	1.398 (0.000)
ln(GDP)²	-0.053 (0.933)	0.795 (0.164)	0.655 (0.284)
ln(Energy)	-0.355 (0.126)	*dropped*	0.072 (0.443)
ln(Urbanization)	0.880 (0.055)	0.035 (0.729)	*dropped*
FDI (% GDP)	-0.108 (0.002)	-0.062 (0.014)	-0.065 (0.016)
Inflation (%)	0.006 (0.031)	0.007 (0.014)	0.007 (0.014)
ECT (adj. speed)	-0.593 (0.000)	-0.556 (0.000)	-0.533 (0.000)

Notes: β=Long-run coefficient (delta-method); p-values in parentheses.

Source: Prepared by the authors based on Python results.

The main income result remains stable across the three specifications. The coefficient on ln(GDP) ranges from 1.263 to 1.429 and remains significant at the 1% level. FDI, inflation, and the error-correction term also retain their signs and remain statistically significant. By contrast, the urbanization coefficient falls from 0.880 (p = 0.055) to 0.035 (p = 0.729) when Energy is removed, while the Energy coefficient also changes considerably when Urbanization is excluded. These changes are consistent with the strong collinearity between the two variables and mean that their individual effects should be interpreted cautiously. The GDP^2 coefficient is insignificant in all three specifications, although its sign changes when either Energy or Urbanization is excluded. The evidence therefore does not support a statistically significant nonlinear income effect.

Overall, the main conclusions are not materially altered by excluding Energy or Urbanization. The positive and significant long-run association between income and CO_2 emissions remains, as does the absence of statistically significant curvature. FDI and inflation also retain their respective negative and positive long-run associations. The individual coefficients for Energy and Urbanization, however, should be treated more cautiously because of their strong collinearity.

Causality Tests

Standard Granger Causality (VAR(1))

As a preliminary causality analysis, a VAR(1) model in first differences is estimated for the seven variables. BIC and HQIC support one lag, while AIC selects four; given the small sample and the number of parameters, the more parsimonious VAR(1) is retained (Table 13). Wald chi-squared tests are used to assess Granger causality. The centered ln(GDP) and its square are included as separate endogenous variables, with their coefficients tested jointly when income is

the causal or affected variable. The same joint treatment is used in the Toda-Yamamoto test for comparability. The remaining variables are tested individually. Results are reported in Table 15.

Table 13. Lag-Order Selection Criteria for the Underlying VAR System

Lag order (p)	N (obs)	Params/eq	AIC	BIC	HQIC
1	47	8	-42.848	-40.643*	-42.018*
2	47	15	-42.802	-38.669	-41.246
3	47	22	-42.540	-36.478	-40.259
4	47	29	-43.776†	-35.785	-40.769

*Notes: * denotes the minimum BIC and HQIC values across lag orders p = 1–4, while † denotes the minimum AIC value at p = 4. Although AIC selects p = 4, p = 1 is adopted based on the more parsimonious BIC and HQIC criteria and the limited sample size. Candidate lag orders are restricted to p = 1–4 because each additional lag rapidly increases the number of parameters to be estimated. To ensure comparability, all four specifications are evaluated using a common sample of N = 47 observations. The selected p = 1 model is then re-estimated over its full available sample, yielding T = 52 observations for the period 1972–2023 (Table 15).*

Source: Prepared by the authors based on Python results.

The adopted p = 1 specification, discussed above, is used for both the standard Granger and Toda-Yamamoto causality analyses reported below. The Toda-Yamamoto system augments p = 1 with dmax = 1 additional lag (LA-VAR(2)), yielding 15 parameters per equation against 52 observations, that is, 37 effective degrees of freedom per equation, a ratio that reduces (though does not eliminate) the power concerns discussed further below. Residuals diagnostics for both models are reported in Table 14.

Table 14. Stability and Residuals Diagnostics

Diagnostic	Statistic	p-value	Conclusion
Panel A. Lag-augmented VAR(2) in levels (Toda–Yamamoto; preferred specification)			
VAR stability (max eigenvalue modulus)	0.979	—	Stable (close to unit circle)
Multivariate Portmanteau (h=4), asymptotic	Q = 200.863	< 0.0001	Rejected †
Multivariate Portmanteau (h=4), bootstrap (2,000 reps)	Q = 200.863	0.62	Not rejected
Multivariate normality (Jarque-Bera type)	146.519	< 0.0001	Non-normal
ARCH effects, per equation (range across 7 equations)	0.04–5.05	0.025–0.840	Present in 1 of 7 equations
Panel B. VAR(1) in first differences (supplementary standard Granger tests)			
VAR stability (max eigenvalue modulus)	0.752	—	Stable
Multivariate Portmanteau (h=4), asymptotic	Q = 173.991	0.064	Not rejected †
Multivariate Portmanteau (h=4), bootstrap (2,000 reps)	Q = 173.991	0.894	Not rejected
Multivariate normality (Jarque-Bera type)	99.058	< 0.0001	Non-normal
ARCH effects, per equation (range across 7 equations)	0.23–16.40	< 0.0001–0.629	Present in 4 of 7 equations

Notes: Panel A is the lag-augmented VAR(2) in levels (p = 1, dmax = 1) underlying the Toda–Yamamoto tests in Table 16; Panel B is the VAR(1) in first differences underlying the standard Granger tests in Table 15. Both models use T = 52 (1972–2023) and seven endogenous variables, with centered ln(GDP) and its square entered as separate variables. † The asymptotic Portmanteau test relies on a χ^2 approximation with $k^2(h - p)$ degrees of freedom (98 in Panel A, 147 in Panel B), which is unreliable when the number of estimated parameters is large relative to the sample size; the bootstrap version is therefore the more informative test here. It generates

2,000 artificial samples from the estimated model by resampling the residuals in moving blocks of four consecutive years (matching the four lags tested), which preserves the short-run dependence between neighboring observations; the bootstrap p-value is the share of these samples whose Portmanteau statistic is at least as large as the one observed. Here k=7 is the number of variables, h=4 the number of residual lags tested and p the VAR lag order. The multivariate normality test has 2k=14 degrees of freedom in both panels. ARCH effects are tested equation by equation with a first-order ARCH LM test at the 5% level.

Source: Prepared by the authors based on Python results.

Table 14 reports residual diagnostics for the two models behind the causality results. Panel A covers the lag-augmented VAR(2) in levels, which underlies the preferred Toda–Yamamoto tests (Table 16). Panel B covers the VAR(1) in first differences, which underlies the supplementary Granger tests (Table 15).

For the LA-VAR(2), the largest eigenvalue modulus of the companion matrix is 0.979. That is below one but close to it, as we would expect from a levels system containing I(1) variables, and the Toda–Yamamoto procedure does not require the levels VAR to be stationary. The asymptotic multivariate Portmanteau test rejects the null of no residual autocorrelation ($Q = 200.86$; $df = 98$; $p < 0.001$). Its moving-block bootstrap version does not ($p = 0.62$), and it is the more dependable of the two here, given a sample of only 52 observations and a large number of estimated parameters. Residual normality is rejected (JB = 146.52; df = 14; $p < 0.001$), and ARCH effects appear in one of the seven equations (inflation).

The differenced VAR(1) is stable, with a maximum eigenvalue modulus of 0.752. Neither the Portmanteau test ($Q = 173.99$; $df = 147$; $p = 0.064$) nor its bootstrap version ($p = 0.894$) points to significant residual autocorrelation. Residual normality is rejected here as well (JB = 99.06; df = 14; $p < 0.001$), and ARCH effects appear in four of the seven equations.

Table 15. Standard Granger Causality Test Results (VAR (1))

From (Cause)	To (Effect)	χ^2	df	p-value	Direction
Income [x + x^2]	CO_2 emissions	1.29	2	0.524	No predictive causality
Energy use	CO_2 emissions	0.15	1	0.694	No predictive causality
Urbanization	CO_2 emissions	2.81	1	0.094	Predictive causality, 10%
FDI	CO_2 emissions	0.00	1	0.965	No predictive causality
Inflation	CO_2 emissions	0.66	1	0.417	No predictive causality
CO_2 emissions	Income [x + x^2]	7.12	2	0.028	Predictive causality, 5%
CO_2 emissions	Energy use	3.78	1	0.052	Predictive causality, 10%
CO_2 emissions	Urban population	4.50	1	0.034	Predictive causality, 5%
CO_2 emissions	FDI	9.20	1	0.002	Predictive causality, 1%
CO_2 emissions	Inflation	2.10	1	0.148	No predictive causality

Notes: VAR(1) estimation based on the differenced series was carried out for all seven variables to make the series stationary. The sample period is from 1972 to 2023 (T = 52). Wald χ^2 statistics were used to test the significance of the lag-1 coefficients. "Income" refers to the centered ln(GDP) series, which is x = ln(GDP) centered, along with its squared value x^2. The two income terms were tested jointly (2 degrees of freedom), while the other variables were tested individually (1 degree of freedom), as in the Toda–Yamamoto procedure shown in Table 16. As differencing leads to the loss of long-run information from the levels series, the conclusions drawn here can be viewed as a short-run robustness check

Source: Prepared by the authors based on Python results.

Overall, the diagnostics support the use of the LA-VAR(2). The system is stable (largest eigenvalue modulus 0.979, inside the unit circle), and the bootstrap Portmanteau test, which is better suited to a sample of 52 observations than the asymptotic version, finds no residual autocorrelation ($p = 0.62$). ARCH effects appear in only one of the seven equations (inflation). Non-

normality of the residuals is common in macroeconomic samples of this size, and the Toda–Yamamoto Wald tests remain valid under it asymptotically. The Toda–Yamamoto results are therefore taken as the main causality evidence, and the Granger results from the differenced VAR(1) serve as a consistency check.

The standard Granger tests, estimated in first differences, show little predictive causality toward CO_2 emissions. Income (p = 0.524), energy use (p = 0.694), FDI (p = 0.965), and inflation (p = 0.417) are not significant, while urbanization has weak predictive content at the 10% level (p = 0.094). In the reverse direction, CO_2 emissions do not significantly predict inflation (p = 0.148), and their effect on energy use is marginal (p = 0.052). However, they jointly predict the income block (χ^2 = 7.12, df = 2, p = 0.028), as well as urbanization (p = 0.034) and FDI (p = 0.002). These results reflect short-run predictive content rather than structural causality. The standard Granger and Toda–Yamamoto approaches give broadly similar results toward CO_2 emissions but differ for the reverse effects on income, urbanization, and FDI.

However, the first-differenced VAR focuses on predictive dynamics in changes and does not retain the levels information that may characterize long-run relationships; it also does not test whether income and its square jointly shift the long-run trajectory of emissions in the way the EKC hypothesis requires. The Toda-Yamamoto procedure, described below, instead estimates the VAR in levels while accommodating the integration properties of the variables, and is therefore used as the primary causality specification.

Toda-Yamamoto Modified Wald Causality Test (VAR(2))

The Toda and Yamamoto (1995) modified Wald test procedure is adopted as the primary causality test because it is specifically designed for time series with unknown integration and cointegration properties. The procedure permits Wald testing in the level VAR without requiring pre-testing for cointegration, provided that the maximum integration order is correctly identified (Toda & Yamamoto, 1995); it does not, however, eliminate finite-sample concerns given the modest sample size used here.

The procedure estimates a Lag-Augmented VAR (LA-VAR) in levels with p + dmax lags, where p = 1 (supported by BIC and HQIC, per Table 13) and dmax = 1. The value of dmax reflects the maximum order of integration actually established across all seven variables in Table 6, not an assumption extended from ln(GDP) to its square: centered ln(GDP) and the other five variables are individually confirmed I(1), while its square is separately tested and found I(0) (DF-GLS statistic = −4.038, p = 0.001), so the maximum across the system remains one. The resulting LA-VAR(2) is estimated over 1972–2023 (T = 52). Wald tests are applied to the coefficient of the first p = 1 lag only; the augmented second lag is included for asymptotic validity but excluded from the test restrictions. Results are presented in Table 16. Diagnostics for LA-VAR(2) are reported in Panel A of Table 14.

Table 16. Toda-Yamamoto (1995) Modified Wald Causality Test Results (LA-VAR (2))

From (Cause)	To (Effect)	χ^2	df	p-value	Direction
Income [x + x^2]	CO_2 emissions	1.82	2	0.403	No predictive causality
Energy use	CO_2 emissions	0.08	1	0.784	No predictive causality
Urban population	CO_2 emissions	0.83	1	0.362	No predictive causality
FDI	CO_2 emissions	0.00	1	0.972	No predictive causality
Inflation	CO_2 emissions	0.71	1	0.399	No predictive causality
CO_2 emissions	Income [x + x^2]	0.95	2	0.621	No predictive causality
CO_2 emissions	Energy use	0.18	1	0.672	No predictive causality
CO_2 emissions	Urban population	0.47	1	0.495	No predictive causality
CO_2 emissions	FDI	2.59	1	0.108	No predictive causality
CO_2 emissions	Inflation	0.21	1	0.644	No predictive causality

Notes: LA-VAR(2) estimated in levels with p = 1 and dmax = 1 augmentation lag for I(1) variables (Toda & Yamamoto, 1995). The p = 1 specification is supported by BIC and HQIC (Table 13). "Income" refers to x = ln(GDP) centered and its square, x^2, which are tested jointly (2 df) when income is the causal or affected variable. The joint test avoids dependence on the choice of centering and accounts for the covariance between the coefficient estimates. Wald statistics test the lag-1 coefficients only, excluding the augmented lag-2 term. Sample: 1972–2023 (T = 52).

Source: Prepared by the authors based on Python results.

The Toda–Yamamoto results differ from the first-differenced VAR, especially in the reverse direction. This difference reflects the fact that the two approaches estimate different models: the first uses differences, whereas Toda–Yamamoto uses levels with an additional lag. In the TY results, none of the variables shows significant predictive causality toward CO_2 emissions, either individually (energy use, urbanization, FDI, and inflation) or jointly for income (GDP and GDP^2). This lack of significance does not mean that these variables are unrelated to CO_2 emissions; rather, no predictive relationship is detected at conventional levels in this sample and specification. In the reverse direction, CO_2 emissions also do not significantly predict income, energy use, urbanization, FDI, or inflation.

Comparison of Causality Methods

The two causality tests give similar results for the causal direction of primary interest. Neither test detects predictive causality toward CO_2 emissions at the 5% level from income, energy use, FDI, or inflation, whether tested individually or (for income) jointly. The standard Granger test provides weak evidence for urbanization at the 10% level (p = 0.094), whereas the Toda-Yamamoto test does not (p = 0.362). The two tests disagree, however, on three reverse-direction relationships: CO_2 emissions jointly predicting the income block, and CO_2 emissions predicting urbanization and FDI, all of which are significant under standard Granger but not under Toda-Yamamoto. Because Toda-Yamamoto works directly in levels with lag augmentation, retaining level information that differencing would discard (without itself modeling a long-run equilibrium relationship in the way an ARDL/ECM does), and provides asymptotically valid inference without pre-testing for cointegration, it is adopted as the preferred and primary causality specification in this study, consistent with the maximum order of integration of one among the variables. The standard Granger results are retained as a short-run robustness comparison. Given the limited effective degrees of freedom of the Toda-Yamamoto system documented above, these results are best read as a failure to detect statistically significant predictive causality, rather than as confirmation that Algeria's CO_2 trajectory is structurally exogenous to these macroeconomic variables; a non-rejection in a low-power test does not establish the absence of a true relationship.

CONCLUSION

This paper has presented a comprehensive empirical investigation of the Environmental Kuznets Curve hypothesis for Algeria over 1970–2023 using the ARDL bounds testing approach of Pesaran et al. (2001) with finite-sample critical values from Kripfganz and Schneider (2020). The following principal findings emerge.

First, the ARDL bounds test confirms a long-run relationship among the variables (F = 5.827, t = −5.671). The F-statistic exceeds its upper I(1) bound, while the t-statistic is more negative than its corresponding I(1) critical value at the 5% level. The CUSUM and CUSUM of squares tests reported later do not indicate parameter instability. The ECT coefficient of −0.593 indicates relatively rapid adjustment toward long-run equilibrium, with about 59% of short-run disequilibrium corrected each year.

Second, the results do not confirm the EKC hypothesis in the long run. After ln(GDP) is mean-centered, its squared term is formed to reduce collinearity (Table 11). GDP^2 is not statistically

significant, so there is no evidence of an inverted-U relationship. At the sample-mean level of income, the marginal long-run effect of income on CO_2 emissions is positive and significant ($p < 0.001$). Although a turning point can be calculated from the estimated coefficients, it lies far outside the income range observed in the sample. Given the insignificant GDP^2 term, it cannot be considered an economically meaningful threshold. In the short run, the change in income is positive and significant (D.ln(GDP), $p = 0.006$), while the lagged squared term is not significant ($p = 0.122$). The short-run results therefore also suggest a positive income–emissions relationship without significant curvature.

Third, the long-run coefficient of inflation is positive and statistically significant ($\beta = 0.006$, $p = 0.031$), suggesting a link between macroeconomic instability and CO_2 emissions. FDI is also statistically significant in the long run ($p = 0.002$), indicating that foreign direct investment is another factor associated with long-run CO_2 emissions.

Fourth, the Toda–Yamamoto (1995) test, preferred because the maximum order of integration is one, finds no significant predictive causality between CO_2 emissions and the macroeconomic variables in either direction. The standard Granger test, used as a short-run robustness check, gives the same result for causality toward CO_2 emissions but identifies three reverse-direction relationships that are not found by Toda–Yamamoto: the income block, urbanization, and FDI. This difference reflects the fact that the two tests use different model specifications rather than Toda–Yamamoto being inherently more conservative.

The findings suggest that environmental improvement cannot be expected to emerge automatically from continued income growth, since no economically meaningful income threshold is identified at which emissions would begin to decline on their own. Policy intervention should therefore focus on reducing the carbon intensity of economic activity directly: reforming fossil-fuel subsidies and administered energy prices, accelerating renewable-energy deployment, and improving the sustainability of urban growth. Environmental regulation should be strengthened while foreign direct investment is directed toward cleaner technologies and less carbon-intensive activities, which may reinforce the negative environmental association observed in the estimates; the model does not, however, establish that stronger regulation itself attracts such investment. The findings therefore suggest that changes in the emissions trajectory are unlikely to arise from income growth alone and may require complementary structural and energy-policy interventions.

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