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DOES INDUSTRIAL EXPANSION STILL DRIVE INCOME GROWTH? AN ARDL ANALYSIS OF INDUSTRIALISATION, ELECTRICITY CONSUMPTION, AND ECONOMIC GROWTH IN UZBEKISTAN

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Abstract. This study examines whether industrial expansion and electricity consumption are associated with higher per capita income in Uzbekistan over the period 2000–2024. Using World Development Indicators data, per capita gross domestic product (GDP) is modelled as a function of industrial value added growth, electricity consumption per capita, and the labour force participation rate within the Autoregressive Distributed Lag (ARDL) bounds-testing framework, with an unrestricted intercept and trend (Case V). Unit root tests indicate a combination of I(0) and I(1) series, with no I(2) variables. The bounds F-statistic of 4.885 exceeds the asymptotic upper bound at the 10% significance level, while the evidence strengthens to the 5% level (F = 6.760) in a parsimonious specification that excludes the weakly identified labour variable. The error-correction coefficient of −0.404 implies a half-life of deviations of approximately 1.3 years. A one-percentage-point increase in industrial value added growth is associated with an increase of approximately 1.0% in per capita income in the short run and 4.1% in the long run. The long-run coefficient remains positive and statistically significant (0.031–0.056) across four specifications. Electricity consumption and labour force participation exhibit no independent long-run association with income after controlling for industrial performance, while the sign of the electricity coefficient is unstable. Given the sample of 25 annual observations, the long-run estimates should be interpreted as conditional associations rather than causal effects. The findings suggest that energy efficiency and productivity-led industrial upgrading offer a more reliable pathway to income growth than expanding electricity supply alone.

Keywords: Industrialization, electricity consumption, economic growth, ARDL bounds testing, error-correction model, Uzbekistan.

Аннотация. В данном исследовании рассматривается взаимосвязь промышленного роста, потребления электроэнергии и повышения дохода на душу населения в Узбекистане за период 2000–2024 годов. На основе данных World Development Indicators валовой внутренний продукт (ВВП) на душу населения моделируется в зависимости от темпов роста добавленной стоимости в промышленности, потребления электроэнергии на душу населения и уровня участия населения в рабочей силе с использованием модели авторегрессионного распределённого лага (ARDL) и теста границ на коинтеграцию с неограниченной константой и трендом (случай V). Тесты на единичный корень свидетельствуют о наличии рядов, интегрированных порядка I(0) и I(1), при отсутствии переменных порядка I(2). Значение F-статистики теста границ, равное 4,885, превышает асимптотическую верхнюю критическую границу на уровне значимости 10%, тогда как в более компактной спецификации, исключающей слабо идентифицируемую переменную труда, статистическая значимость усиливается до уровня 5% (F = 6,760). Коэффициент корректировки ошибок, равный −0,404, указывает на период полусокращения отклонений от долгосрочного равновесия приблизительно в 1,3 года. Увеличение темпов

роста добавленной стоимости в промышленности на один процентный пункт связано с повышением дохода на душу населения приблизительно на 1,0% в краткосрочном периоде и на 4,1% в долгосрочном периоде. Долгосрочный коэффициент остаётся положительным и статистически значимым (0,031–0,056) во всех четырёх спецификациях модели. Потребление электроэнергии и уровень участия населения в рабочей силе не демонстрируют самостоятельной долгосрочной связи с доходом после учёта показателей промышленного развития, при этом знак коэффициента потребления электроэнергии остаётся нестабильным. Учитывая ограниченный объём выборки, включающей 25 годовых наблюдений, долгосрочные оценки следует интерпретировать как условные статистические взаимосвязи, а не как причинно-следственные эффекты. Полученные результаты свидетельствуют о том, что повышение энергоэффективности и модернизация промышленности на основе роста производительности представляют собой более надёжное направление увеличения доходов населения по сравнению с одним лишь расширением предложения электроэнергии.

Ключевые слова: Индустриализация, потребление электроэнергии, экономический рост, тест границ ARDL, модель коррекции ошибок, Узбекистан.

INTRODUCTION

Industrial development has long occupied a central place in explanations of how low- and middle-income economies raise living standards. Manufacturing is credited with increasing returns to scale, strong backward and forward linkages, and the capacity to absorb labour released from low-productivity agriculture (Szirmai, 2012). These characteristics support the classical view that the pace of industrial expansion determines the pace of aggregate productivity growth. However, the empirical evidence supporting this view has become less conclusive. Rodrik (2016) shows that many developing economies now experience declining industrial employment at income levels far below those at which advanced economies experienced similar transitions, raising the possibility that industrialisation no longer delivers the same growth benefits as in the past.

Energy plays an equally prominent role in this debate. Industrial production is energy-intensive, so manufacturing expansion is generally accompanied by rising electricity demand, and many studies have interpreted the resulting correlation as evidence that energy consumption drives output growth. However, the evidence supporting the energy-led growth hypothesis is mixed: the direction and even the existence of causality vary according to the level of development, the structure of the energy sector, and the estimation method (Apergis & Payne, 2010; Soytas & Sari, 2009). This distinction is important for policy formulation. If energy consumption drives economic growth, investment in electricity generation capacity can itself serve as a growth strategy. If energy demand primarily responds to industrial activity, improving energy efficiency rather than increasing consumption becomes the more appropriate objective. These two approaches have substantially different fiscal and environmental implications.

Uzbekistan provides an informative case for examining these relationships. The country has pursued an industry-oriented development path since the early 2000s, and industry now accounts for a substantial share of economic output (UNCTAD, 2025). Rapid economic growth has been accompanied by a marked increase in electricity demand, while the energy system remains heavily dependent on natural gas (Turakulov et al., 2024). Energy intensity remains among the highest globally, although it has declined steadily, suggesting that the relationship between energy consumption and economic output has evolved over the sample period.

The existing empirical evidence for Uzbekistan requires careful methodological consideration despite the growing number of published studies. Several studies use output measured in current local currency without incorporating a deterministic trend, despite the economy experiencing sustained double-digit inflation. In such regressions, strongly trending explanatory variables may appear statistically significant. Available samples are relatively short, comprising approximately 25 consistent annual observations, which limits both the number of parameters that can be estimated and the statistical power of unit root and cointegration tests. In addition, the interpretation of bounds-test results can be strengthened by explicitly specifying the deterministic case and the source of the critical values, as well as verifying the sign and statistical significance of the error-correction term.

This study addresses these methodological considerations using annual data for 2000–2024. Per capita gross domestic product (GDP) is the dependent variable, while the explanatory variables comprise the growth rate of industrial value added, electricity consumption per capita, and the labour force participation rate. A deterministic trend is included because the level series exhibit a strong common trend; without it, the adjustment coefficient becomes implausibly small, and the long-run elasticities lack meaningful economic interpretation. The study makes three contributions. First, it demonstrates that the long-run association between

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industrial expansion and per capita income remains robust to changes in the set of control variables. Second, it distinguishes the industrial channel from the energy channel and finds that electricity consumption has no independent long-run association with income after controlling for industrial performance, a finding that differs from much of the regional literature. Third, it provides transparent statistical inference by reporting Case V critical values for the appropriate number of regressors, simulated small-sample critical values, and the bounds t-test, enabling readers to assess the strength of the evidence for cointegration.

LITERATURE REVIEW

The argument for the distinctive role of industry can be traced to Kaldor, who maintained that manufacturing benefits from dynamic economies of scale that are less prevalent in other sectors, allowing manufacturing growth to stimulate productivity growth throughout the economy. Szirmai (2012) finds that manufacturing served as an engine of growth in developing economies for much of the twentieth century, although the strength of this relationship has weakened over time and varies across regions. Rodrik (2016) offers a more cautious perspective, documenting premature deindustrialisation, whereby industrial employment shares peak at substantially lower income levels than in previous periods. Taken together, these studies suggest that the relationship between industrialisation and economic growth is an empirical question that depends on specific temporal and geographical conditions.

A parallel body of literature examines the relationship between energy consumption and economic output. Ang (2007) identifies long-run relationships among emissions, energy consumption, and output in France using a vector error-correction framework. Soytas and Sari (2009) report that the causal relationships among energy consumption, emissions, and income in Turkey are weaker than those implied by the energy-led growth hypothesis. Zhang and Cheng (2009) find that causality in China runs from economic output to energy consumption rather than in the opposite direction. Apergis and Payne (2010) identify a positive long-run relationship and bidirectional causality between renewable energy consumption and economic output in OECD countries. Methodologically, much of this literature has adopted the bounds-testing approach developed by Pesaran, Shin, and Smith (2001), which accommodates a combination of I(0) and I(1) regressors and estimates short- and long-run parameters within a single-equation framework. Narayan (2005) demonstrates that asymptotic critical values may underestimate the appropriate critical values in samples containing 30–80 observations, an issue of particular relevance to single-country studies based on 25 annual observations.

Research on Central Asia has expanded considerably. Sobirov, Karimov, and Bekjanov (2023), using panel ARDL methods, find that foreign direct investment, infrastructure, innovation, and trade openness exhibit positive long-run associations with economic development in Commonwealth of Independent States (CIS) countries. Saidmamatov et al. (2023) document substantial cross-country differences in the agriculture–water–energy–environment nexus. Kozhagulov et al. (2025) identify Kazakhstan and Uzbekistan as the principal contributors to regional emissions, reflecting the energy-intensive industrial structures of both economies. Kuldasheva and Ahmad (2025) report positive effects of female labour force participation on economic growth in the region.

For Uzbekistan, Apergis, Kuziboev, Abdullaev, and Rajabov (2023) apply the ARDL approach to data covering 1985–2020 and find that hydropower reduces emissions, while natural gas and oil increase them. Notably, they exclude economic output from the model to address multicollinearity among strongly trending series. Ibadullaev et al. (2024) report that renewable and non-renewable energy consumption support economic development in the long run over the period 1990–2021. Halmuratov et al. (2025) employ a vector error-correction model (VECM) for 1990–2023 and find that energy consumption positively influences per capita output, while trade openness exhibits a negative coefficient. Xolmurotov et al. (2025) and Nurjanov et al. (2025) extend this research to unemployment and exports, while Nuriddinova and Ahmad (2025) examine the environmental Kuznets curve hypothesis for Uzbekistan.

RESEARCH METHODOLOGY

The analysis uses annual data for Uzbekistan over 2000-2024 (25 observations) from the World Development Indicators (World Bank, 2025). The sample starts in 2000 because earlier transition-period series are not comparable. The dependent variable is the logarithm of gross domestic product per capita in current local currency units (lnGDPPC). The main regressor is the annual growth rate of industrial value added in percentage points (IND). IND measures the pace of industrial expansion, not the industry share of GDP, so its coefficient refers to the speed rather than the level of industrialization. The controls are the logarithms of electricity consumption per capita in kWh (lnEPC) and of the labour force participation rate (lnLAB). The renewable share of total energy consumption (REC), available for 23 years, is used only in a robustness check. IND and REC are kept in their original units because they are rates or shares.

Using a nominal dependent variable is a limitation that this paper shares with part of the literature it discusses. Nominal per capita GDP grows about 300-fold over the sample, against roughly threefold growth in real terms. The deterministic trend is included to absorb this common drift: the implied long-run trend of 0.216 log points per year is close to the average nominal growth rate. A linear trend cannot, however, capture changes in the inflation rate, such as the acceleration after the 2017 currency liberalisation, and re-estimation at constant prices is the most important extension of this work.

Unit root properties are checked with the Augmented Dickey-Fuller (Dickey & Fuller, 1979) and Phillips-Perron (Phillips & Perron, 1988) tests with a constant and trend; the bounds approach is valid only if no series is I(2). The long-run relationship of interest is

$$\ln GDPPC_t = \beta_0 + \beta_1 t + \beta_2 \ln EPC_t + \beta_3 \ln LAB_t + \beta_4 IND_t + \varepsilon_t \quad (1)$$

and is tested through the unrestricted conditional error correction model (Pesaran et al., 2001)

$$\Delta y_t = \alpha_0 + \alpha_1 t + \rho\, y_{t-1} + \sum_{j=1}^{k} \theta_j x_{j,t-1} + \sum_{j=1}^{k} \sum_{m=0}^{q_j - 1} \gamma_{jm} \Delta x_{j,t-m} + \varepsilon_t \quad (2)$$

where *y* is lnGDPPC and *x* the vector of regressors. Lag orders are selected by the Akaike information criterion with a maximum of two lags, which the sample size requires; the criterion selects ARDL(1,1,1,1). The null of no level relationship, H0: ρ = θ1 = θ2 = θ3 = 0, is tested with a Wald F-statistic compared with lower (all regressors I(0)) and upper (all regressors I(1)) bounds. Because the intercept and trend enter unrestricted, Case V critical values for k = 3 regressors apply; these are well above the Case III values usually reported. As asymptotic bounds understate small-sample critical values (Narayan, 2005), the paper also reports bounds simulated for T = 24 with the selected lag structure (20,000 replications). A significant F-test is necessary but not sufficient for cointegration (Kremers, Ericsson & Dolado, 1992), so the bounds t-test on ρ is reported as well.

Long-run coefficients are recovered as

$$\hat{\beta}_j^{LR} = -\frac{\hat{\theta}_j}{\hat{\rho}} \quad (3)$$

with delta-method standard errors that account for the covariance between numerator and denominator. Short-run dynamics are described by the error correction model (Engle & Granger, 1987)

$$\Delta y_t = \alpha_0 + \alpha_1 t + \sum_{j=1}^{k} \sum_{m=0}^{q_j - 1} \gamma_{jm} \Delta x_{j,t-m} + \rho\, ECT_{t-1} + \varepsilon_t \quad (4)$$

$$ECT_{t-1} = y_{t-1} - \sum_{j=1}^{k} \hat{\beta}_j^{LR} x_{j,t-1} \quad (5)$$

where a negative and significant ρ indicates convergence and its absolute value gives the share of a deviation corrected within a year.

Residuals are tested for serial correlation (Breusch-Godfrey LM, Ljung-Box), heteroskedasticity (Breusch-Pagan), normality (Jarque-Bera, Shapiro-Wilk) and functional form (Ramsey RESET). Parameter stability is assessed with CUSUM and CUSUM of squares statistics. Robustness is examined in two ways: re-estimation with Newey-West (1987) HAC standard errors (Bartlett kernel, two lags), and re-estimation under alternative control sets, recomputing the bounds test, the error correction coefficient and the long-run industrial coefficient each time.

Artificial intelligence tools were used for approximately 10% of the manuscript preparation process, exclusively for language editing, grammatical corrections, and technical improvements to the text. The research methodology, data analysis, interpretation of results, and scientific conclusions were developed independently by the authors. The authors retain full responsibility for the accuracy, originality, and scientific integrity of the manuscript.

ANALYSIS AND RESULTS

Table 1 presents the descriptive statistics. Nominal GDP per capita increased from approximately 131 thousand to 40.0 million soums, representing a 300-fold increase reflecting real economic growth, sustained

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inflation, and the 2017 currency reform. Industrial growth averaged 7.29%, with a standard deviation of 4.58 percentage points. Its pronounced positive skewness (2.95) and excess kurtosis (12.04) reflect the exceptional industrial expansion recorded in 2010. Electricity consumption per capita ranged from 1,473 to 2,119 kWh, while the labour force participation rate remained between 58.9% and 62.9%. The limited variation in lnLAB reduces the precision with which its coefficient can be estimated (Table 1).

Table 1
Descriptive statistics, Uzbekistan, 2000-2024

Variable	Obs	Mean	Std. Dev.	**Min**	**Max**	**Skew.**	**Kurt.**
GDPPC (LCU)	25	9,381,094	11,580,790	131,343	40,002,740	n.r.	n.r.
IND (%)	25	7.287	4.578	1.778	26.069	2.951	12.043
LAB (%)	25	60.420	1.202	58.861	62.879	0.955	0.083
EPC (kWh)	25	1,656.124	192.352	1,473.393	2,118.901	n.r.	n.r.
REC (%)	23	1.213	0.327	0.700	1.700	0.065	-1.440
lnGDPPC	25	15.001	1.734	11.786	17.504	-0.284	-1.120
lnEPC	25	7.406	0.109	7.295	7.659	1.268	0.663
lnLAB	25	4.101	0.020	4.075	4.141	0.955	0.083

Note: GDPPC is GDP per capita in current local currency units; IND the annual growth rate of industrial value added; LAB the labour force participation rate; EPC electricity consumption per capita; REC the renewable share of total energy consumption. Kurtosis is in excess form. n.r.: moments not reported for untransformed level series; see the log-transformed rows used in estimation. Source: authors' calculations based on World Bank (2025).

Figure 1 illustrates the dynamics of key macroeconomic, industrial, and energy indicators in Uzbekistan over the period 2000–2024. Panel (a) shows a sustained increase in nominal GDP per capita, particularly after 2017, reflecting the combined effects of economic growth and changes in nominal values. Panel (b) indicates that industrial value added growth fluctuated considerably throughout the period, reaching an exceptional peak of approximately 26% in 2010 before returning to more moderate growth rates. Despite these fluctuations, industrial growth remained predominantly positive (Figure 1).

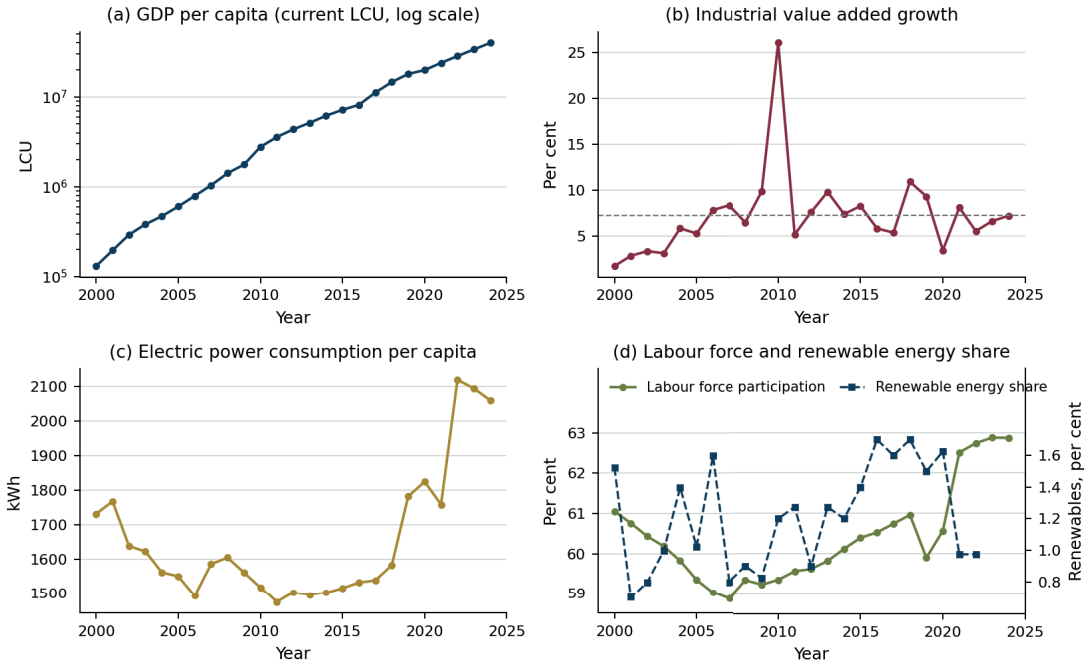


Figure 1. Macroeconomic, industrial and energy variables in Uzbekistan, 2000-2024[1]

1 Note: Panel (a) uses a logarithmic scale. Panel (b) shows the annual growth rate of industrial value added, with the sample mean as

Table 2 and Figure 2 show that per capita income is positively correlated with labour force participation (0.537) and electricity consumption (0.406). Its correlation with industrial growth is positive but weak (0.236) and statistically insignificant, as expected when a growth rate is compared with a level variable. A particularly important finding is the strong positive correlation (0.818) between electricity consumption and labour force participation. Both variables exhibit upward trends, making it difficult to estimate their individual effects precisely with only 25 observations (Table 2).

Table 2
Pearson correlation matrix

Variables	(1)	(2)	(3)	(4)	(5)
(1) lnGDPPC	1.000				
(2) lnEPC	0.406**	1.000			
(3) lnLAB	0.537***	0.818***	1.000		
(4) IND	0.236	-0.242	-0.236	1.000	
(5) REC	0.360*	-0.153	0.080	0.045	1.000

Note: *, ** and *** denote significance at the 10%, 5% and 1% levels. n = 25 (n = 23 for REC).

Figure 2 presents the correlation matrix of the transformed variables used in the econometric analysis. The results indicate that per capita GDP (lnGDPPC) is positively correlated with labour force participation (lnLAB), electricity consumption per capita (lnEPC), industrial value added growth (IND), and renewable energy consumption (REC), with correlation coefficients of 0.54, 0.41, 0.24, and 0.36, respectively. The strongest positive correlation is observed between electricity consumption per capita and labour force participation (0.82), suggesting that these variables exhibit similar patterns over the observation period. This relatively high correlation may introduce multicollinearity concerns when estimating their individual effects within the same model (Figure 2).

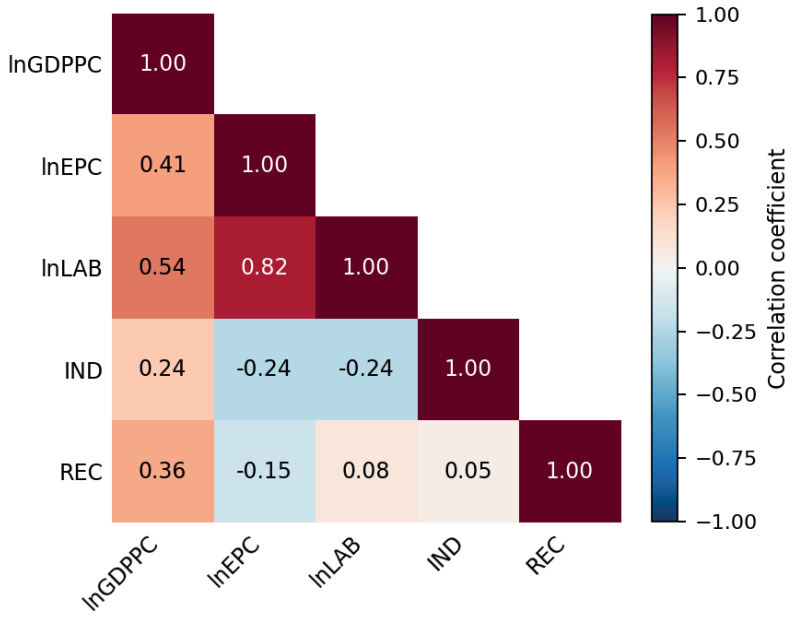


Figure 2. Correlation heat map of the transformed variables[2]

Table 3 presents the results of the unit root tests, indicating a combination of integration orders. Industrial growth is stationary at levels according to both tests. REC is stationary at levels under the Phillips–Perron (PP) test but not under the Augmented Dickey–Fuller (ADF) test, a discrepancy commonly observed for bounded series in small samples. The variables lnGDPPC, lnEPC, and lnLAB are non-stationary at levels

a dotted line. Source: authors' calculations based on World Bank (2025).

2 Note: Coefficients computed on annual observations, 2000-2024 (n = 25; n = 23 for REC).

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but become stationary after first differencing. The only ambiguity concerns lnEPC, for which the ADF test on the first difference yields a p-value of 0.511, whereas the Phillips–Perron test decisively rejects the unit root null hypothesis. Given the limited statistical power of the ADF test in small samples, lnEPC is treated as I(1). None of the variables is integrated of order two, I(2), supporting the applicability of the ARDL bounds-testing approach (Table 3).

Table 3
Unit root tests (ADF and Phillips-Perron, MacKinnon p-values)

Variable	ADF level	ADF 1st diff.	PP level	PP 1st diff.	Order
lnGDPPC	0.8993	0.0259**	0.7207	0.0002***	I(1)
lnEPC	0.9966	0.5112	0.9543	0.0000***	I(1)
lnLAB	0.5329	0.0244**	0.7068	0.0002***	I(1)
IND	0.0000***	0.0000***	0.0054***	0.0000***	I(0)
REC	0.3384	0.0000***	0.0073***	0.0000***	I(0)/I(1)

Note: Constant and linear trend in all test regressions; ADF lag length by AIC. *, ** and *** denote significance at the 10%, 5% and 1% levels.

Table 4 presents the ARDL(1,1,1,1) estimation results. The coefficient of the lagged dependent variable (0.596) indicates substantial but stable persistence, while the deterministic trend coefficient (0.087) is statistically significant at the 5% level, supporting its inclusion in the model. Industrial growth is the only explanatory variable with a statistically significant contemporaneous effect. A one-percentage-point increase in industrial growth is associated with an immediate increase of approximately 1.0% in per capita income (p = 0.002). Its first lag (0.006) is marginally significant, suggesting that the association extends beyond the current year. Neither electricity consumption nor labour force participation exhibits a statistically significant effect. For electricity consumption, this finding suggests that year-to-year changes in electricity use provide no additional information about per capita income after controlling for industrial performance. Regarding labour force participation, the variable fluctuates within a narrow range of approximately four percentage points; therefore, its statistical insignificance may reflect limited identifying variation rather than the absence of a labour-related channel. The high R-squared value of 0.999 is largely attributable to the lagged dependent variable and the deterministic trend and should not be interpreted as evidence of strong explanatory performance on its own (Table 4).

Table 4
Short-run ARDL estimates (dependent variable: lnGDPPC)

Variable	Coefficient	Std. Error	t-Statistic	**Prob.**
lnGDPPC(t-1)	0.5958	0.1464	4.070	0.0010***
lnEPC	-0.1501	0.2878	-0.521	0.6096
lnEPC(t-1)	-0.2573	0.2740	-0.940	0.3620
lnLAB	1.4758	1.8199	0.811	0.4301
lnLAB(t-1)	-0.3937	1.6700	-0.236	0.8170
IND	0.0104	0.0028	3.757	0.0019***
IND(t-1)	0.0061	0.0031	1.997	0.0640*
Trend	0.0873	0.0369	2.363	0.0321**
Constant	3.5352	7.1752	0.493	0.6294
R^2 / Adjusted R^2	0.9993 / 0.9989			
F-statistic (prob.)	2681.39 (0.000)			
Durbin-Watson	1.759			

Note: ARDL(1,1,1,1) selected by AIC, unrestricted intercept and trend. Observations after adjustment: 24. *, ** and *** denote significance at the 10%, 5% and 1% levels.

Table 5 presents the estimated long-run coefficients. Industrial growth has a coefficient of 0.0408 (standard error = 0.0106, p < 0.001), indicating that a sustained one-percentage-point increase in industrial growth is

associated with an approximately 4.1% higher level of per capita income in long-run equilibrium. Since IND represents a growth rate and lnGDPPC is expressed in logarithmic levels, this coefficient captures how the level of per capita income co-moves with the pace of industrial expansion around the deterministic trend, rather than the effect of a one-off change in the level of industrialisation. Therefore, it should not be interpreted as a causal growth elasticity. Neither of the other long-run coefficients is statistically significant. The electricity consumption coefficient is negative (−1.008, standard error = 0.655), but it is not statistically distinguishable from zero, and its sign is unstable across specifications. Consequently, interpreting this result as evidence of energy inefficiency would not be justified. The data do not identify an independent long-run association between electricity consumption and per capita income after controlling for industrial performance. Labour force participation has a relatively large but imprecisely estimated coefficient (2.678, standard error = 3.910), again reflecting limited identifying variation (Table 5).

Table 5
Long-run ARDL coefficients (dependent variable: lnGDPPC)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
lnEPC	-1.0080	0.6545	-1.540	0.1235
lnLAB	2.6775	3.9103	0.685	0.4935
IND	0.0408	0.0106	3.854	0.0001 ***

Note: Long-run coefficients computed as -θj/ρ from the conditional error correction model; delta-method standard errors. *** denotes significance at the 1% level.

Table 6 presents the results of the ARDL bounds tests. The F-statistic of 4.885 exceeds the asymptotic Case V upper critical bound at the 10% significance level (4.432) but remains below the 5% upper critical bound (5.048). When compared with the small-sample critical bounds for T = 24, the statistic falls within the inconclusive region. The t-statistic for the lagged dependent variable (−2.760) does not satisfy the bounds t-test rejection criterion, for which the 10% I(0) critical value is −3.13. Therefore, the evidence supporting a long-run level relationship in the baseline model is moderate rather than conclusive. This evidence becomes stronger when the weakly identified labour variable is excluded. It is also important to note that applying Case III critical values or critical values corresponding to a larger number of regressors would incorrectly suggest rejection of the null hypothesis at the 5% significance level. This methodological distinction helps explain the difficulties in comparing earlier empirical findings (Table 6).

Table 6
ARDL bounds tests for cointegration, ARDL(1,1,1,1), Case V, k = 3

Significance	Asymptotic I(0)	Asymptotic I(1)	T = 24 I(0)	T = 24 I(1)
Panel A. F-test: F = 4.885 (significant at 10%, asymptotic bounds)				
10%	3.474	4.432	4.01	5.35
5%	4.022	5.048	4.96	6.59
1%	5.191	6.345	7.43	9.82
Panel B. t-test on ρ: t = -2.760 (not significant)				
10%	-3.13	-3.84	-3.12	-3.92
5%	-3.41	-4.16	-3.55	-4.39
1%	-3.96	-4.73	-4.45	-5.47

Note: Asymptotic bounds from Pesaran et al. (2001), Case V (unrestricted intercept and trend), k = 3. Small-sample bounds simulated by the authors for T = 24 with the ARDL(1,1,1,1) structure (20,000 replications). Rejection of no cointegration requires the statistic to exceed the I(1) bound (in absolute value for the t-test).

Table 7 presents the error-correction model (ECM) estimation results. The error-correction coefficient is −0.404, indicating that approximately 40% of any deviation from long-run equilibrium is corrected within one year. This corresponds to an adjustment half-life of approximately 1.3 years [ln(0.5)/ln(1 − 0.404)]. When the deterministic trend is excluded, the coefficient decreases in absolute magnitude to −0.059, implying an

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adjustment half-life exceeding 11 years, which is less plausible for an annual macroeconomic relationship. The conventional p-value of 0.015 is reported for completeness but should not be interpreted as a valid test of cointegration for the reasons discussed above. In the short-run specification, ΔIND remains statistically significant at the 1% level, while the electricity consumption and labour force participation terms remain statistically insignificant (Table 7).

Table 7
Error correction model (dependent variable: ΔlnGDPPC)

Variable	Coefficient	Std. Error	t-Statistic	Prob.
ECT(t-1)	-0.4042	0.1464	-2.760	0.0146†
ΔlnEPC	-0.1501	0.2878	-0.521	0.6096
ΔlnLAB	1.4758	1.8199	0.811	0.4301
ΔIND	0.0104	0.0028	3.757	0.0019***
Trend	0.0873	0.0369	2.363	0.0321**
Constant	3.5352	7.1752	0.493	0.6294
Speed of adjustment (%)	40.4			
Half-life (years)	1.34			

Note: † conventional p-value; the valid test of cointegration is the bounds t-test in Table 6. ** and *** denote significance at the 5% and 1% levels.

Table 8 and Figure 3 present the results of re-estimating the model under four alternative sets of control variables: the baseline specification (M1), the specification excluding labour force participation (M2), the specification excluding electricity consumption (M3), and the specification incorporating the renewable energy share (M4, based on 22 observations). Three main findings emerge from this robustness analysis.

First, the error-correction coefficient remains negative across all four specifications, ranging from −0.274 to −0.523. Second, the bounds test rejects the null hypothesis of no cointegration at the 5% significance level in M2 (F = 6.760, exceeding the upper critical bound of 5.817) and at the 10% level in M1 and M3. The results for M4 remain inconclusive, consistent with the reduction in sample size by three observations. Therefore, the strongest evidence of a long-run level relationship is obtained from the specification excluding labour force participation, which exhibits limited variation and a high degree of collinearity with electricity consumption. Third, and most importantly, the long-run industrial growth coefficient remains positive and statistically significant across all four specifications, ranging from 0.031 to 0.056. This stability across alternative model specifications, rather than the statistical significance of any individual estimate, provides stronger support for the observed relationship between industrial growth and per capita income.

In contrast, the electricity consumption coefficient exhibits considerable instability across specifications. It is negative in the baseline model but becomes positive and relatively large in M2 when labour force participation is excluded. Such a reversal in the coefficient's sign following the removal of a highly collinear explanatory variable limits its economic interpretability. Consequently, selecting only the specification that produces the theoretically expected sign would introduce specification-selection bias and could lead to misleading conclusions (Table 8).

Table 8
Robustness of cointegration and the long-run industrial effect

Specification	Obs	F-stat.	I(0) 5%	I(1) 5%	ECT	Long-run IND
M1: EPC + LAB + IND	24	4.885*	4.022	5.048	-0.404**	0.0408***
M2: EPC + IND	24	6.760**	4.883	5.817	-0.438***	0.0350***
M3: LAB + IND	24	5.543*	4.883	5.817	-0.274**	0.0562***
M4: EPC + LAB + IND + REC	22	3.686	3.501	4.574	-0.523**	0.0313***

Note: Case V in all models; asymptotic bounds from Pesaran et al. (2001) for k = 3, 2, 2 and 4 regressors. F-statistic stars refer to the asymptotic I(1) bounds; ECT stars are conventional t-test levels. Standard errors of the long-run IND coefficient: 0.0106, 0.0089, 0.0156 and 0.0091. *, ** and *** denote significance at the 10%, 5% and 1% levels.

Figure 3 illustrates the ARDL bounds-test results and the stability of the long-run industrial growth coefficient across four alternative model specifications. Panel (a) presents the bounds-test F-statistics for the baseline model (M1), the model excluding labour force participation (M2), the model excluding electricity consumption (M3), and the model incorporating the renewable energy share (M4). The corresponding F-statistics are 4.88, 6.76, 5.54, and 3.69, respectively. The horizontal reference lines represent the relevant critical bounds at the 5% and 10% significance levels. The strongest evidence of a long-run level relationship is observed in M2, where the F-statistic exceeds the 5% upper critical bound. M1 and M3 provide evidence of cointegration at the 10% significance level, while the results for M4 remain inconclusive (Figure 3).

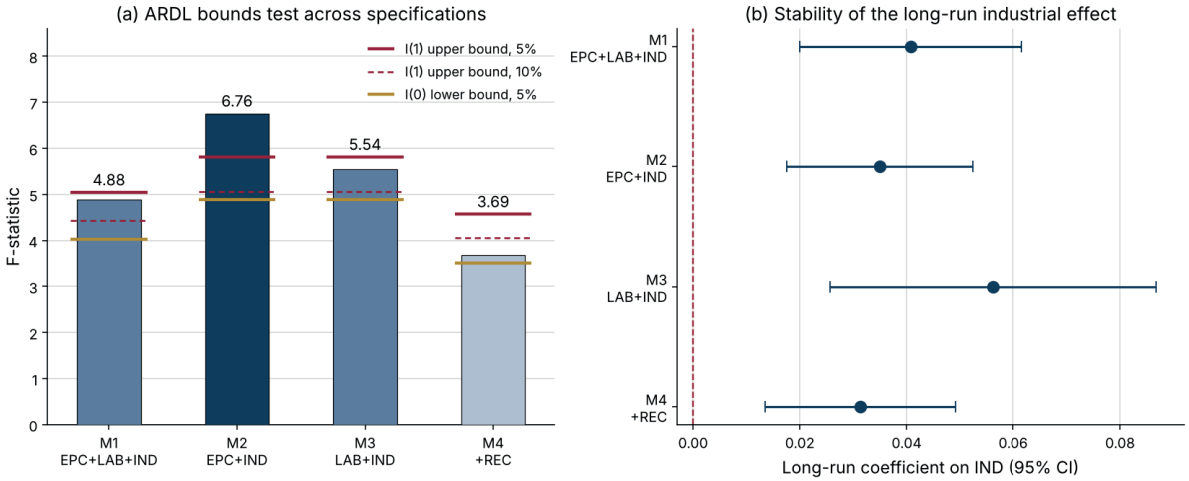


Figure 3. Bounds test statistics and stability of the long-run industrial coefficient[3]

Table 9 presents the diagnostic test results, indicating no evidence of serial correlation (Breusch–Godfrey, p = 0.656; Ljung–Box, p = 0.563), heteroskedasticity (Breusch–Pagan, p = 0.217), or deviations from normality (Jarque–Bera, p = 0.731; Shapiro–Wilk, p = 0.798). The Ramsey RESET test (p = 0.738) provides no evidence of functional-form misspecification (Table 9).

Table 9
Diagnostic and stability tests

Test	Statistic	Prob.	Conclusion
Breusch-Godfrey LM, 2 lags	0.843	0.6562	No serial correlation
Ljung-Box Q(1)	0.335	0.5628	No serial correlation
Breusch-Pagan, $\chi^2(8)$	10.741	0.2168	Homoskedastic
Jarque-Bera	0.626	0.7313	Normal residuals
Shapiro-Wilk	0.975	0.7975	Normal residuals
Ramsey RESET	0.116	0.7381	Correct functional form
Durbin-Watson	1.759	-	No first-order autocorrelation
CUSUM	-	-	Parameters stable
CUSUM of squares	-	-	Variance stable

Note: Graphical evidence is reported in Figures 4 to 7.

3 Note: Panel (a) compares each F-statistic with the asymptotic Case V bounds for the relevant number of regressors; dark bars reject at 5%, mid-blue bars at 10%, light bars are inconclusive. Panel (b) shows the long-run IND coefficient with 95% confidence intervals.

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Figure 4 shows that the fitted values closely track the observed data.

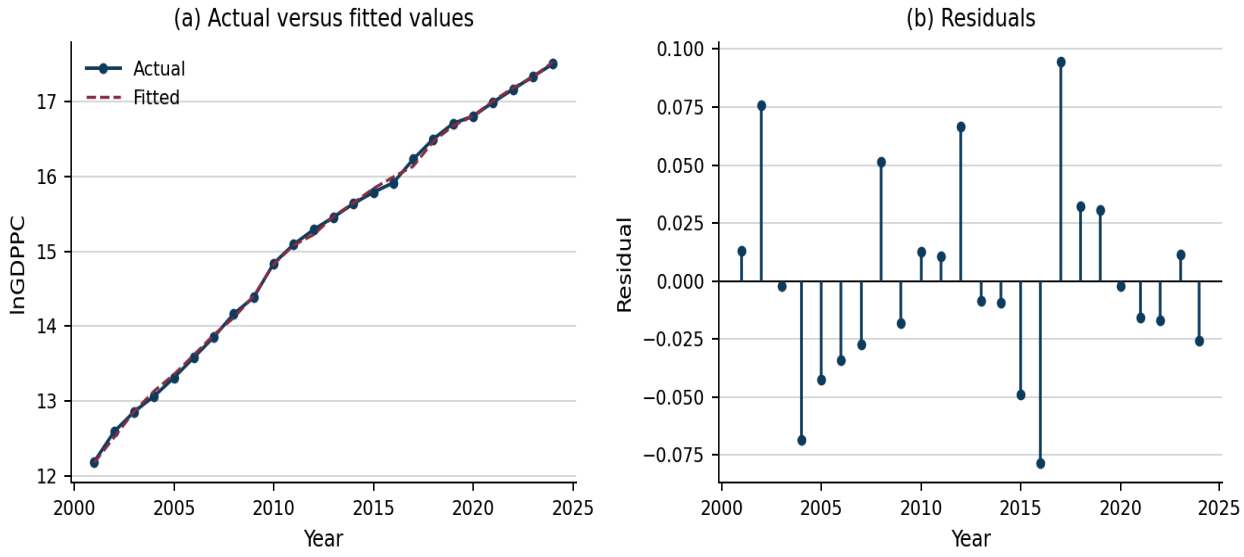


Figure 4. Actual versus fitted values and residuals[4]

The CUSUM statistic remains within the 5% significance bounds, while the maximum deviation of the CUSUM of squares statistic (0.176) is substantially below the critical value (0.340), indicating that the null hypothesis of parameter constancy cannot be rejected (Figure 5).

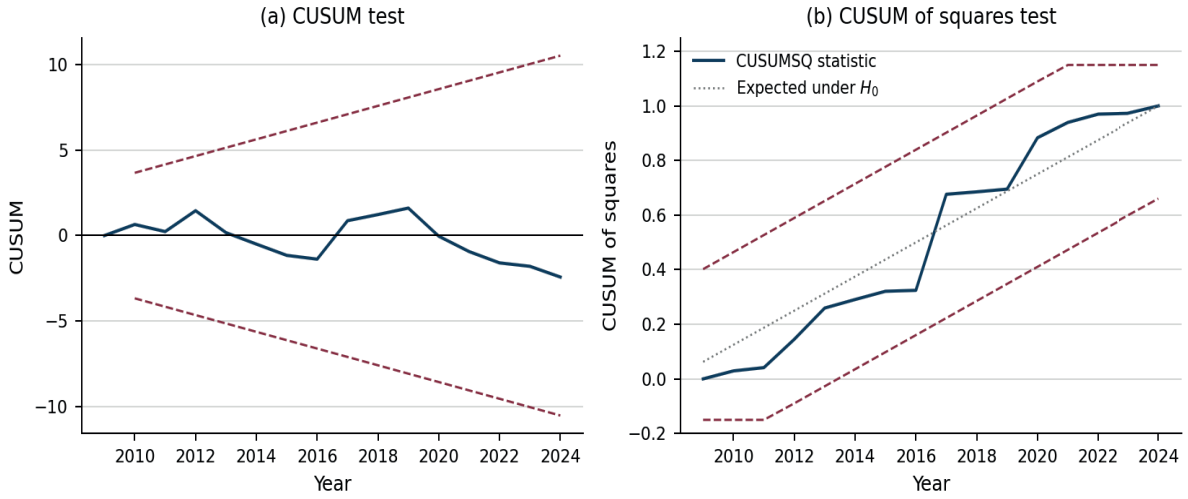


Figure 5. Recursive CUSUM and CUSUM of squares tests

Note: Dashed lines are 5% significance bands; the dotted line in panel (b) is the expected path under parameter constancy, with bands ±1.358/√n.

All residual autocorrelations fall within the confidence bands (Figure 6).

4 Note: ARDL(1,1,1,1) with unrestricted intercept and trend.

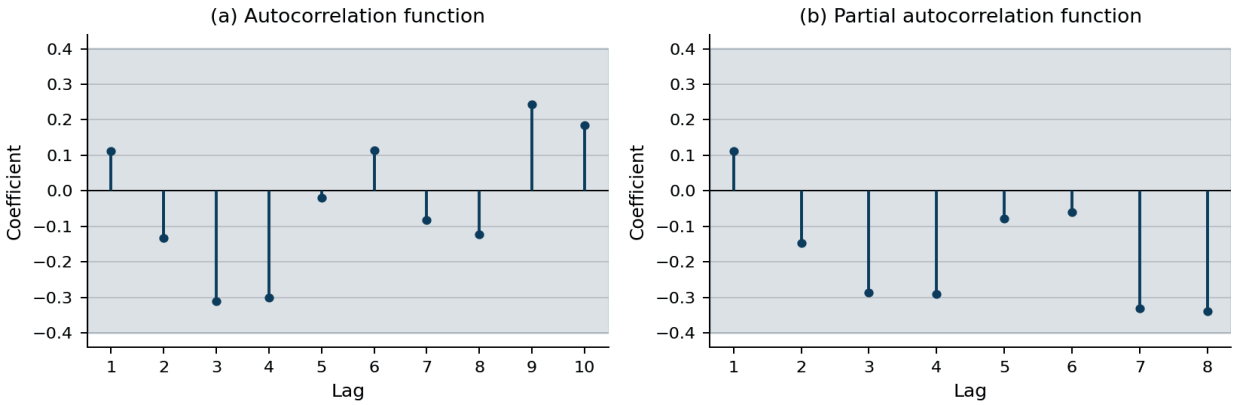


Figure 6. Autocorrelation and partial autocorrelation of residuals[5]

When heteroskedasticity and autocorrelation consistent (HAC) standard errors are applied, the industrial growth coefficients remain statistically significant at the 1% level. However, Newey–West standard errors may be downward biased in samples comprising only 24 observations. Therefore, the results in Table 10 provide additional support for the direction and statistical significance of the estimated industrial growth relationship rather than greater precision in the estimates (Table 10).

Table 10
ARDL estimates with HAC Newey-West standard errors

Variable	Coefficient	Robust S.E.	t-Statistic	Prob.
lnGDPPC(t-1)	0.5958	0.1392	4.282	0.0007***
lnEPC	-0.1501	0.1928	-0.778	0.4484
lnEPC(t-1)	-0.2573	0.1330	-1.935	0.0721*
lnLAB	1.4758	1.0976	1.345	0.1988
lnLAB(t-1)	-0.3937	1.0262	-0.384	0.7067
IND	0.0104	0.0009	11.575	0.0000***
IND(t-1)	0.0061	0.0014	4.355	0.0006***
Trend	0.0873	0.0355	2.458	0.0266**
Constant	3.5352	4.9449	0.715	0.4856

Note: Newey and West (1987) standard errors, Bartlett kernel, bandwidth of two lags. *, ** and *** denote significance at the 10%, 5% and 1% levels.

Figure 7 supports these conclusions. Simulated densities envelop the observed distribution (a), residual smooths are close to flat (b, c), deviations from normality are small (f), and the variance inflation factors for the long-run regressors lie between 1.10 and 2.20 (e). Panel (d) needs a qualification. With 24 observations and nine parameters, leverage is high by construction (maximum 0.928). The largest Cook's distance is 0.766 for 2011, below 1 but above 0.5, and the most influential years are 2010, 2011 and 2017, linked to the industrial surge of 2010 and the 2017 currency reform. The estimates should therefore be read as conditional on the sample period (Figure 7).

5 Note: Shaded areas are 95% confidence bands.

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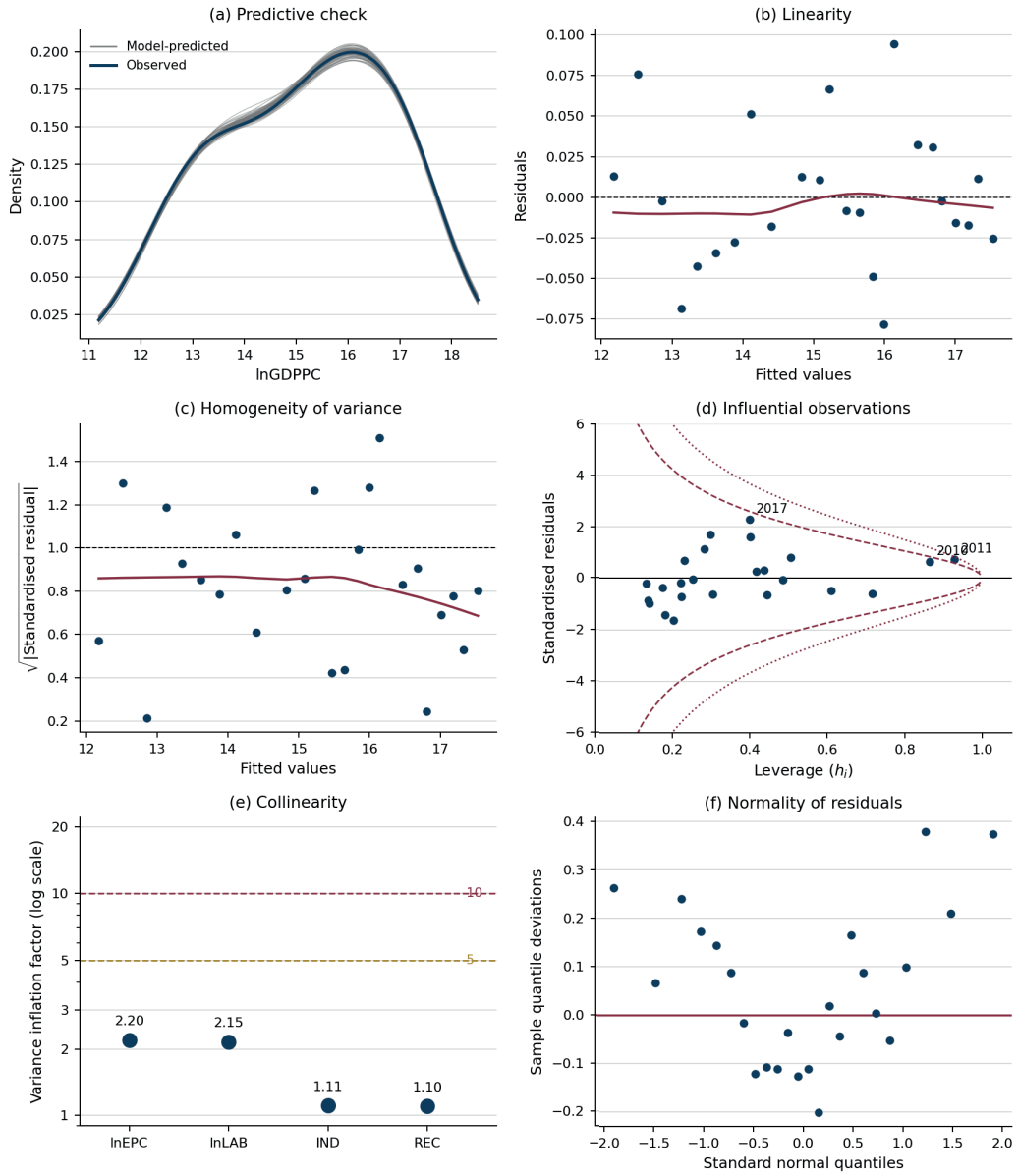


Figure 7. Consolidated model assumption checks[6]

The results support a focused set of conclusions. The pace of industrial expansion is positively and robustly associated with per capita income in Uzbekistan in both the short and long run. A long-run coefficient of approximately 0.041 indicates that a sustained one-percentage-point increase in industrial growth is associated with an approximately 4% higher level of per capita income in long-run equilibrium, with deviations from equilibrium being corrected at a rate of approximately 40% per year. This finding is consistent with the view that manufacturing contributes to structural transformation (Szirmai, 2012) and provides no evidence of premature deindustrialisation (Rodrik, 2016) in Uzbekistan over the period examined.

The absence of an independent long-run relationship between electricity consumption and per capita income does not imply that energy is unimportant for production. Rather, it indicates that electricity consumption provides no additional information about income beyond that already captured by industrial performance. A plausible interpretation is that electricity demand in Uzbekistan is largely derived from industrial activity. This interpretation is consistent with Apergis and Payne (2010) regarding the structural dependence of the energy–growth relationship and with Apergis et al. (2023), who excluded economic output from their model

6 Note: (a) observed density of the dependent variable against 60 simulated densities; (b) residuals and (c) scale-location against fitted values with LOWESS smooths; (d) standardised residuals against leverage with Cook's distance contours at 0.5 and 1.0; (e) variance inflation factors for the long-run regressors; (f) deviations of sample quantiles from normal quantiles.

because of multicollinearity among trending series. However, it differs from the findings of Ibadullaev et al. (2024) and Halmuratov et al. (2025), who report positive effects of energy consumption. Three differences in model specification may explain these contrasting findings. First, models using nominal variables without a deterministic trend may allow strongly trending regressors to capture common trends. Second, excluding the deterministic trend may cause the explanatory variable most closely associated with income to absorb the common trend. Third, omitting industrial performance may allow the electricity consumption coefficient to capture the industrial growth channel. Therefore, the present findings highlight the importance of model specification in interpreting earlier conclusions rather than directly contradicting them.

CONCLUSION AND RECOMMENDATIONS

This study has examined the relationships among industrialisation, electricity consumption, and per capita income in Uzbekistan over the period 2000–2024 using the ARDL bounds-testing approach with an unrestricted trend. Three main findings emerge. First, there is moderate evidence of a long-run relationship: the F-test rejects the null hypothesis of no cointegration at the 10% significance level in the baseline model and at the 5% level when the weakly identified labour variable is excluded. The error-correction coefficient (−0.404) implies an adjustment half-life of approximately 1.3 years. Second, the pace of industrial expansion is positively associated with per capita income in both the short run (approximately 1.0% per percentage point) and the long run (approximately 4.1% per percentage point). The long-run coefficient remains statistically significant and relatively stable in magnitude across all four model specifications. Third, neither electricity consumption nor labour force participation exhibits an independent long-run association with per capita income after controlling for industrial performance, while the sign of the electricity consumption coefficient remains unstable across specifications.

The first policy implication concerns energy. The empirical findings do not provide sufficient evidence that increasing electricity consumption alone raises per capita income. Therefore, expanding electricity generation capacity may be more effective as part of a broader growth strategy rather than as an independent policy instrument. Greater emphasis should be placed on increasing value added per unit of energy consumed through energy efficiency standards for large industrial users, phased tariff reforms that strengthen incentives for energy conservation, and support for the adoption of less energy-intensive technologies. Given Uzbekistan's relatively high energy intensity, substantial opportunities exist for improving energy efficiency.

The second policy implication concerns the composition of industrial growth. Since the pace of industrial expansion, rather than the volume of energy consumed, is associated with per capita income, industrial policy should focus on sustaining growth through productivity improvements rather than capacity expansion alone. Key priorities include increasing the technological content of industrial output, strengthening domestic supplier linkages, and promoting higher value-added production segments instead of relying primarily on the expansion of energy-intensive industries. Increasing value added per unit of energy consumed would contribute to both industrial development and environmental sustainability.

The third policy implication concerns the methodological quality of policy analysis. The differences between the present findings and those of earlier studies arise from model specification rather than differences in the underlying data. Therefore, studies prepared to inform government policy should explicitly identify the deterministic specification and the sources of critical values, report the error-correction coefficient alongside the bounds-test results, and demonstrate how the main estimates respond to reasonable alternative model specifications.

Future research should re-estimate the model using constant-price output, disaggregate industrial value added by industry, and distinguish industrial electricity consumption from residential electricity use to examine the derived-demand interpretation directly. As longer time series become available, future studies should also incorporate structural break tests and tests for asymmetric adjustment.

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