“Public and fiscal policy instruments for supporting renewable electricity development: Evidence from a cross-country study”

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PUBLIC AND FISCAL POLICY INSTRUMENTS FOR SUPPORTING RENEWABLE ELECTRICITY DEVELOPMENT: EVIDENCE FROM A CROSS-COUNTRY STUDY

Abstract

As governments worldwide intensify efforts to achieve decarbonization, the role of fiscal and public policy instruments in shaping energy transitions has gained critical importance. This study evaluates how budgetary measures, taxation schemes, subsidies, and regulatory standards influence renewable electricity outcomes, thereby linking climate policy design with broader public and municipal finance issues. The analysis relies on panel data from 48 OECD, OECD negotiating members, and OECD participating partner-countries between 2009 and 2022, estimated with fixed and random effects models in R Studio and tested for robustness using Driscoll-Kraay and cluster-robust standard errors. The findings indicate that, compared with other instruments, feed-in tariffs ($\beta = 0.116$, $p < 0.001$), planning for renewables expansion ($\beta = 0.070$, $p < 0.01$), and air emission standards ($\beta = 0.170$, $p < 0.001$) provide the strongest and most consistent support for renewable electricity development. Renewable energy certificates and auctions also contribute positively, though with weaker statistical significance, while fossil fuel excise taxes and coal bans display mixed or context-dependent effects. The adjusted R^2 of 0.38 for renewable electricity generation and 0.44 for renewable electricity supply demonstrates the explanatory relevance of the selected policy variables. Robustness checks further confirm the enduring importance of feed-in tariffs as a cornerstone of fiscal support for renewables. Finally, cross-country heterogeneity is evident, with strong positive random effects in Bulgaria (0.86), Slovenia (0.73), and Czechia (0.81), and pronounced negative effects in Saudi Arabia (–1.23), Costa Rica (–1.22), and Chile (–0.95).

Keywords renewable energy, climate policy instruments, feed-in tariffs, ETS, carbon taxes, panel data analysis

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INTRODUCTION

Despite growing fiscal constraints and institutional fragmentation, the global transition to renewable electricity is progressing at an unprecedented pace. The International Energy Agency (IEA, 2024b) projects that renewable electricity generation will reach over 17,000 TWh by 2030, nearly double the 2023 level, driven primarily by the rapid expansion of solar and wind power. At this trajectory, solar and wind are expected to surpass coal-fired power by 2025, with solar photovoltaics becoming the world's leading electricity source by 2029. This momentum is fueled by ambitious policy initiatives, including the COP28 pledge to triple global renewable capacity by 2030, supported by a mix of fiscal incentives, budgetary measures, regulatory frameworks, and public financing models beginning to deliver tangible results (IEA, 2024b). In this context, evaluating how different public and fiscal policy instruments influence renewable electricity development emerges as a timely and policy-relevant task.

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Complementary evidence from multilateral financial institutions highlights that fiscal and public investment frameworks are central to advancing the energy transition. The World Bank (2025) has channeled nearly USD 100 billion over the past decade into climate-related energy access programs, prioritizing affordability and grid expansion in developing economies. However, significant disparities remain, as by 2023 only 15% of global clean energy investments were directed toward developing countries, despite their large populations and rising electricity demand (IEA, 2024a). This investment gap underlines the importance of context-sensitive fiscal designs, including targeted subsidies, public co-financing, and equitable tax instruments, to ensure climate and energy policies translate into tangible benefits, especially in underserved regions.

Recent insights from the OECD (2025) further emphasize that the effectiveness of renewable energy development depends not only on the existence of policy tools but also on their design and sequencing. Coherent policy mixes that combine carbon pricing, feed-in tariffs, and performance standards tend to outperform single instruments when implemented in an integrated and well-timed manner. Importantly, regulatory frameworks often provide the enabling conditions for fiscal tools to operate effectively, particularly where institutional capacity and revenue streams differ across countries. This underscores the need to analyze how public and fiscal policy instruments (individually and in combination) shape renewable electricity outcomes across diverse governance and economic contexts.

1. LITERATURE REVIEW

The intensifying global push for decarbonization has brought climate policy architecture to the forefront of academic and policy discourse, particularly concerning its effectiveness in promoting renewable electricity development (Kudurs et al., 2023; Kacare et al., 2022; Dolge et al., 2023; Nurjanov et al., 2025; Karimboev et al., 2025; Sitenko et al., 2023). A range of fiscal and regulatory instruments, including feed-in tariffs (FiTs), emissions trading systems (ETS), carbon and excise taxes, renewable energy certificates, and bans on fossil fuel infrastructure, has emerged as central pillars of policy frameworks designed to incentivize clean energy investment. Empirical and bibliometric analyses consistently underscore that these instruments are vital not only for emissions reduction but also for enabling market stability, innovation, and long-term investor confidence (Juracka & Valaskova, 2025; Krause et al., 2024; Behar Villegas et al., 2024; Lyeonov et al., 2025).

Carbon pricing instruments such as ETS and carbon taxes are viewed as foundational pillars of climate mitigation policy, offering both environmental effectiveness and economic efficiency. Studies on the EU ETS and its sectoral extensions show mixed but increasingly positive impacts on renewable electricity deployment, especially with complementary policies like planning mandates and air emission standards (Samusevych et al., 2024; Kusumawati et al., 2025). Carbon taxation's role in fossil fuel electricity has been analyzed more broadly to internalize externalities and redirect market behavior toward clean alternatives (Kuzior et al., 2023; Bai et al., 2024; Makarenko et al., 2023). At the same time, countries in transition and emerging economies face challenges in administrative capacity and public acceptance of carbon taxes, raising questions about their equity and effectiveness (Gualandri & Kuzior, 2023; Vakulenko & Rekunenko, 2025).

Feed-in tariffs have been extensively studied as an instrument to stimulate investment in renewables, particularly in early-stage markets. While some evidence highlights their diminishing role in mature systems, recent research finds that FiTs still play a crucial part in attracting long-term finance, including from insurance and banking sectors (Shcherbakova, 2025; Lyeonov et al., 2025; Dobrovolska et al., 2024a). Their design, particularly regarding technology-specificity and degression schedules, determines their ability to encourage cost-effective deployment and prevent market distortions (Moroz & Lyeonov, 2024; Fernandes et al., 2025). In contrast, renewable energy auctions are now perceived as more cost-efficient and flexible instruments in competitive electricity markets, although concerns about bidder underperformance and lack of inclusivity persist (Dobrovolska et al., 2024a; Hedau & Malla, 2024).

The broader governance ecosystem also plays a decisive role in climate policy performance. Strategic planning for renewables expansion, transparent regulatory frameworks, and coherent policy mixes enhance both investor confidence and public legitimacy (Dobrovolska et al., 2024b; BatllesdelaFuente et al., 2024). Moreover, policy coherence between energy, environmental, and financial instruments, such as green bonds, ESG mandates, and sustainable finance taxonomies, has emerged as a determinant of long-term renewable energy development (Aleknevičienė et al., 2025; Hasan et al., 2025; Streimikiene, 2025). Interactions between subsidies reform and clean energy deployment are equally important, particularly in fossil fuel-dependent economies where removing distortive subsidies unlocks budgetary space for renewable energy investments (Samusevych et al., 2024; Streimikiene et al., 2024; Badreddine & Cherif, 2024).

The intersection of fiscal and financial policy instruments with clean energy development is gaining increasing attention in the context of green competitiveness and macro-financial resilience. Green finance instruments such as green bonds, sustainability-linked loans, and digital finance ecosystems are fostering cleaner portfolios in both advanced and developing markets (Hasan et al., 2025; Fatima, 2024; Wang & Lu, 2024; Krause et al., 2024). These financial pathways are further complemented by rising corporate and citizen interest in sustainability, climate ethics, and environmental awareness (Behar Villegas et al., 2024; Piwowarski, 2024; Prokopenko et al., 2025). However, cross-country differences in economic structure, energy mix, and governance capacity still determine policy uptake and effectiveness (Mentel et al., 2020; Pavlova et al., 2021; Vysochyna et al., 2024; Kuzior et al., 2021).

Another critical aspect in climate policy architecture is institutional credibility and policy stability. Research suggests that regulatory compliance and protection of minority investors are crucial enablers for entrepreneurial activities in the renewable energy (Halynskyi & Telizhenko, 2024; Ishwardat et al., 2024; Liu et al., 2024). Similarly, early pandemic-era interventions targeting SMEs, energy security, and just transition frameworks highlight how adaptive institutional mechanisms buffer renewable sectors from macroeconomic shocks (Kornyliuk et al., 2022; Ghimire et al., 2025; Havrylenko & Myroshnychenko, 2025; Juracka & Valaskova, 2025). The growing policy literature on circular economy, sustainable competitiveness, and ethical business models under climate uncertainty reinforces these dynamics. Notably, this literature explores the interplay between green policy leadership, corporate governance, and sustainable supply chains as foundational elements for advancing climate-resilient energy transitions (Oe et al., 2023; Noor et al., 2024). Strategic communication and public policy acceptance are critical mechanisms for reinforcing environmental reforms and maintaining stakeholder trust (Juracka et al., 2024; Pimenowa et al., 2023; Özyeşil & Tembelo, 2025). Moreover, incorporating environmental ethics, social equity, and cross-sectoral policy integration emerges as pivotal for ensuring the long-term effectiveness of renewable energy strategies (Nohong et al., 2024; Shtunder et al., 2022).

A substantial corpus of bibliometric and econometric studies provides evidence of how policy instruments evolve with changing stakeholder demands and regional configurations. Insights from bibliometric studies show an increasing convergence around themes like policy coordination, green innovation, and energy justice (Sotnyk et al., 2023; Omeir & Vasiliauskaite, 2025; Makarenko & Vorontsova, 2024; Ziabina et al., 2023). Empirical findings from Eastern Europe, South Africa, and Asia reveal distinct trajectories of underinvestment, regulatory lag, and subsidy optimization efforts, each influencing the pace and structure of renewable electricity generation (Obagbuwa & Munzhelele, 2024; Fernandes et al., 2025; Riabokon & Kyryliuk, 2025; Vasa et al., 2024; Vasylieva et al., 2021).

The reviewed literature demonstrates that public and fiscal policy instruments (complemented by regulatory oversight and market mechanisms) constitute the primary drivers of the renewable electricity transition. Thus, renewable energy development is best understood as the outcome of coherent policy mixes rather than isolated measures. The evidence consistently indicates that no single instrument can achieve sustainable progress; instead, the effectiveness of the transition de-

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pends on a tailored, country-specific combination of fiscal incentives, budgetary support, regulatory standards, and governance mechanisms that collectively enable equitable and resilient growth of renewable electricity.

This study evaluates how budgetary measures, taxation schemes, subsidies, and regulatory standards influence renewable electricity outcomes, thereby linking climate policy design with broader public and municipal finance issues.

2. METHODOLOGY

This study employs a panel data econometric approach to assess the impact of public and fiscal policy instruments on renewable electricity development across a cross-country sample. The explanatory framework incorporates a range of instruments, including fiscal incentives, budgetary measures, market-based mechanisms, and regulatory standards, as captured in the OECD Climate Actions and Policies Measurement Framework. Two dependent variables are used to evaluate the outcomes: renewable electricity generation (y1), defined as the share of renewables in total electricity generation, and renewable electricity supply (y2), defined as the share of renewables in total electricity supply. These indicators enable an assessment of both production-side dynamics and grid-supply performance, thus linking policy interventions to measurable renewable energy outcomes.

The explanatory variables (x1 to x11) represent a range of policy tools and regulatory measures derived from the OECD Climate Actions and Policies Measurement Framework (Nachtigall et al., 2022). These include market-based instruments such as carbon taxes (x3), ETS electricity price (x1) and ETS coverage (x2), feed-in tariffs (x5), renewable auctions (x6), and fossil fuel excise taxes (x11). Additionally, command-and-control policies such as air emission standards (x10), renewable energy certificates (x7), and coal phase-out measures (x8), along with strategic planning tools like reforms of fossil fuel subsidies (x4) and policies supporting the planning of renewables expansion (x9), are incorporated. All variables were extracted from the OECD Data Explorer (OECD, n.d.), which provides structured annual data across member countries.

The panel dataset comprises 48 countries, encompassing a diverse mix of high-, upper-middle-, and lower-middle-income economies and both OECD and non-OECD members. The sample includes European Union countries, other European nations, and some countries from the Americas, the Asia-Pacific region, the Middle East, and Africa (Appendix A). Including such a broad and geographically dispersed set of countries enhances the generalizability of the findings. It allows for assessing policy effects across various institutional, economic, and energy system contexts. The availability of consistent indicators in the OECD Climate Actions and Policies Measurement Framework determined the sample. Coverage is extended to both OECD and non-OECD members to ensure the representation of countries with shared objectives in the renewable energy transition and international cooperation. Some major economies, including the United States and Brazil, could not be included due to incomplete data reporting, which preserved comparability and reliability across the final dataset.

The panel data are unbalanced due to the variation in the availability of policy indicators across countries and years. However, it provides sufficient temporal and cross-sectional variation to yield robust inferences. Differences in the number of observations (N) across models are due to data gaps in specific years or countries, reflecting real-world limitations in policy reporting and measurement consistency. The panel spans a sufficient temporal and cross-sectional range to estimate within-robust and between-country effects.

The econometric modelling framework includes Fixed Effects (FE) and Random Effects (RE) specifications estimated via the plm package in R, allowing for control over unobserved country-level heterogeneity and testing for the consistency of estimators using the Hausman test. To address concerns of serial correlation, heteroskedasticity, and cross-sectional dependence, common in macro-panel datasets, robust standard errors were estimated using Driscoll-Kraay, clustered by group, and clustered by time covariance structures. The diagnostic tests applied include the Breusch-Godfrey/Wooldridge test for autocorrelation, the Breusch-Pagan test for heteroskedasticity, the Pesaran CD test for cross-sectional dependence, and the Lagrange Multiplier (LM) test for significant effects.

The econometric analysis used the R Studio environment, employing the plm, lmtest, and sandwich packages for panel modelling.

3. RESULTS

The dataset comprises 671 observations across 46 countries from 2010 to 2023. The country and year variables show an even temporal spread (mean year ≈ 2016) and a balanced cross-country representation. Low skewness and kurtosis indicate no concerns regarding asymmetry or outliers in these identifiers.

The climate and energy policy variables (x1-x11) reveal varied distributional properties:

- ETS Electricity – price (x1) and coverage GHG (x2) exhibit relatively symmetric distributions (skew = 0.41 and −0.64, respectively), with moderate variability and no extreme values.
- Carbon Tax – Electricity (x3) and Renewable energy auctions (x6) are highly positively skewed (2.88 and 2.38), with high kurtosis (7.44 and 4.76, respectively), reflecting that most countries score zero while a few report high levels.
- Fossil fuels excise taxes – Electricity (x11) is highly skewed (4.07) and leptokurtic (kurtosis = 16.13), indicating a heavy concentration of zeros and a few extreme outliers.
- Feed-In-Tariffs (x5), Renewable energy certificates (x7), and Ban on coal-fired power plants (x8) also exhibit substantial skewness (0.95, 1.62, 0.89), suggesting substantial heterogeneity in policy adoption across countries.
- Reform of fossil fuel subsidies – Electricity (x4) and Air emission standards (x10) are negatively skewed (–1.02 and –0.78), indicating that many countries report higher scores, and few report very low values.
- Planning for renewables expansion (x9) and Renewable electricity generation (y1) are more symmetrically distributed. However, y1 (dependent variable) still displays moderate skewness (0.67) and a wide range (0 to 99.99), which may signal potential heteroscedasticity in regression models.

The outcome variable y2 (Renewable Electricity Supply) has a mean of 19.15%, but a high standard deviation (16.06) and a strong positive skew (2.01), with values ranging from 0% to nearly 90%. The high kurtosis (5.58) implies a distribution with heavy tails, indicating that a small group of countries has exceptionally high shares of renewables in electricity supply. At the same time, many remain closer to the mean or below.

The descriptive analysis reveals several variables with strong non-normality, including x3, x6, x7, and especially x11, with extreme skewness and kurtosis. These distortions could undermine the reliability of parametric statistical techniques. Therefore, it is recommended that appropriate transformations (e.g., logarithmic, square root, or Box-Cox) be applied to normalize their distributions prior to regression modelling.

A Yeo-Johnson transformation was applied to all independent variables (x1-x11) to correct for skewness and ensure the suitability of variables for regression analysis. This transformation was selected due to its ability to handle zero and negative values, which were present in several indicators. The optimal transformation parameter (λ) was estimated from the data for each variable. Notably, variables such as Carbon Tax – Electricity (x3, $\lambda = -3.885$), Renewable Energy Auctions (x6, $\lambda = -3.5$), and Fossil Fuels Excise Taxes – Electricity (x11, $\lambda = -5$) exhibited high levels of right skewness and required substantial transformation. In contrast, variables such as ETS Electricity – Coverage (x2, $\lambda \approx 0.99$) and Air Emission Standards (x10, $\lambda \approx 1.27$) were close to normally distributed and thus only mildly adjusted. The transformed variables were used in all subsequent analyses to enhance model accuracy and meet the assumptions of linear regression.

The optimal Yeo-Johnson transformation parameter for the dependent variable y1 (Renewable electricity generation) was estimated at $\lambda = 0.436$. This value falls between 0 and 1, indicating moderate positive skewness in the original distribution. The transformation thus applies a log-like compres-

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Table 1. Descriptive statistics of key variables

Variable Name	Code	Mean	SD	Median	Min	Max	Skew	Kurtosis
ETS Electricity – price	x1	3.88	3.57	3.00	0	10.00	0.41	–1.21
ETS Electricity – coverage GHG	x2	5.46	3.73	8.00	0	10.00	–0.64	–1.32
Carbon Tax – Electricity	x3	0.79	2.19	0.00	0	10.00	2.88	7.44
Reform of Fossil Fuel Subsidies – Electricity	x4	7.23	3.57	9.00	0	10.00	–1.02	–0.39
Feed–In–Tariffs	x5	1.92	2.55	0.00	0	9.50	0.95	–0.43
Renewable energy auctions	x6	0.65	1.52	0.00	0	8.00	2.38	4.76
Renewable energy certificates	x7	1.66	3.24	0.00	0	10.00	1.62	0.95
Ban on the construction of coal–fired power plants	x8	1.86	2.51	0.00	0	8.00	0.89	–0.69
Planning for renewables expansion	x9	3.88	3.25	4.00	0	9.00	0.27	–1.36
Air emission standards	x10	5.56	3.28	8.00	0	10.00	–0.78	–0.99
Fossil fuels excise taxes – Electricity	x11	0.39	1.42	0.00	0	8.50	4.07	16.13
Renewable electricity generation (% of electricity generation)	y1	37.76	27.39	30.45	0	99.99	0.67	–0.56
Renewable electricity supply (Percentage of electricity supply)	y2	19.15	16.06	14.79	0	89.74	2.01	15.94

sion, reducing the influence of high-end values and bringing the distribution closer to normality. This adjustment enhances the reliability of subsequent regression analyses by improving the homoscedasticity and normality of residuals.

The optimal Yeo-Johnson transformation parameter for the dependent variable y2 (Renewable electricity supply as a percentage of total electricity supply) was estimated at $\lambda = 0.227$. This value lies between 0 and 0.5, indicating a moderate positive skewness in the original distribution. The transformation thus applies a log-like compression, which reduces the influence of high-end outliers while preserving the relative spacing of lower and mid-range values.

Before model estimation, the Yeo-Johnson transformation was applied to all dependent variables (y1 and y2) and the eleven policy instruments (x1-x11). The transformation ensures improved compliance with the assumptions of linear panel models, such as homoscedasticity and normally distributed residuals, thereby enhancing the robustness of the regression results.

3.1. Renewable electricity generation

The FE model reveals that several policy instruments significantly influence renewable electricity generation within countries over time. The Emissions Trading System (ETS) price level (x1_yj) shows a positive and significant effect ($\beta = 0.054$, $p < 0.05$), suggesting that stronger carbon pricing is associated with higher renewable electricity generation. Similarly, renewable energy certificates (x5_yj) are negatively associated ($\beta = -0.038$, $p <$

Table 2. Interpretation of Yeo-Johnson λ values

Variable	λ Value	Interpretation of λ	Skewness Adjustment
x1	0.231	Close to log	Moderate right skew; log-like correction applied
x2	0.992	Near 1	No transformation needed; nearly normal
x3	–3.885	Large negative	Heavy right skew; strong compression at the right tail
x4	1.556	> 1	Mild left skew; expands upper tail slightly
x5	–0.697	Moderate negative	Right skew; needs moderate compression
x6	–3.5	Large negative	Highly right-skewed with many zeros
x7	–1.898	Strong negative	Same; compresses long right tail
x8	–0.828	Moderate negative	Corrects right skew
x9	0.35	Low positive	Light skew; gentle log-like transform
x10	1.268	Slightly > 1	Very mild left skew; slight expansion
x11	–5	Very large negative	Extremely skewed (sparse non-zeros); heavy compression
y1	0.436	Close to log	Moderate right skew; log-like correction applied
y2	0.227	Close to log	Moderate right skew; log-like correction applied

0.05), potentially reflecting implementation inefficiencies or market volatility. Renewable bans and phaseouts of coal-fired power plants (x8_yj), planning for renewables expansion (x9_yj), and air emission standards (x10_yj) all demonstrate strong positive associations ($p < 0.01$), confirming their role as robust drivers of the green energy transition. Other policy tools, such as auctions (x6_yj) and fossil fuel taxes (x11_yj), display only weak or marginal significance.

The RE model confirms many of these findings. The coefficients for x1_yj (ETS price), x5_yj, x8_yj, x9_yj, and x10_yj remain statistically significant and directionally consistent with the FE model. Interestingly, the intercept is near zero and non-significant, indicating that policy instruments are the primary drivers of variation in renewable generation after controlling for country- and time-invariant characteristics. The share of variance explained by individual effects (92.8%) indicates considerable cross-country heterogeneity, underscoring the importance of country-specific contexts in renewable energy deployment. The adjusted R^2 is 0.357, signaling moderate explanatory power across countries and time.

The Hausman test yields a non-significant result ($p = 0.716$), indicating no systematic difference between the FE and RE estimators. This supports the appropriateness of using the Random Effects model in this case, as the unobserved individual heterogeneity is not correlated with the regressors. The findings suggest that targeted instruments, especially coal phase-out policies, air quality regulations, and structured planning, consistently support renewable deployment. However, instruments such as carbon taxes (x3_yj), feed-in tariffs (x4_yj), and fossil fuel excise taxes (x7_yj) appear less influential or more context-sensitive in the current panel setting.

The results of the post-estimation diagnostic tests indicate several important characteristics and potential issues in the random effects model used to analyze the determinants of the Yeo-Johnson-transformed dependent variable y1_yj. Firstly, the Breusch-Godfrey/Wooldridge test for serial correlation in panel models yields a highly significant test statistic ($\chi^2 = 277.33$, $p < 0.0001$), suggesting the presence of first-order serial correlation in the idiosyncratic errors. This violation of the independence assumption implies that the standard errors of the model may be underestimated, potentially inflating the statistical significance of the estimated coefficients.

Secondly, the studentised Breusch-Pagan test strongly rejects the null hypothesis of homoskedasticity (BP = 153.29, $p < 0.0001$), indicating that the residuals are heteroskedastic across entities. This again undermines the reliability of standard inference using conventional (non-robust) standard errors and supports the use of heteroskedasticity-robust or cluster-robust standard errors to obtain consistent estimators of variance.

Table 3. Estimated effects of policy instruments on renewable electricity generation (Yeo-Johnson transformed variables)

Variable	FE Estimate	FE Std. Error	FE p-value	RE Estimate	RE Std. Error	RE p-value
x1_yj	0.0538	0.0226	0.0177*	0.0527	0.0225	0.0189*
x2_yj	0.0397	0.0291	0.1738	0.0441	0.0288	0.1258
x3_yj	–0.0124	0.0229	0.5863	–0.0112	0.0225	0.6200
x4_yj	0.0177	0.0417	0.6709	0.0328	0.0399	0.4111
x5_yj	–0.0379	0.0165	0.0225*	–0.0402	0.0164	0.0143*
x6_yj	0.0264	0.0139	0.0578.	0.0236	0.0138	0.0884.
x7_yj	–0.0382	0.0381	0.3168	–0.0421	0.0368	0.2519
x8_yj	0.1161	0.0172	< 0.0001***	0.1159	0.0171	< 0.0001***
x9_yj	0.0700	0.0234	0.0029**	0.0714	0.0231	0.0020**
x10_yj	0.1767	0.0292	< 0.0001***	0.1705	0.0287	< 0.0001***
x11_yj	–0.0613	0.0357	0.0868.	–0.0573	0.0346	0.0981.

Note: Dependent variable: Renewable electricity generation (Yeo–Johnson transformed); Panel structure: Unbalanced; 48 countries over 13-14 years ($N = 671$); FE model: $R^2 = 0.383$; Adj. $R^2 = 0.324$; F(11, 612) = 34.48; $p < 0.001$; RE model: $R^2 = 0.367$; Adj. $R^2 = 0.357$; Chi^2(11) = 382.28; $p < 0.001$; Hausman Test: $\chi^2 = 7.97$; df = 11; $p = 0.7158$ → RE is consistent. Signif. codes: '***' – 0.001; '**' – 0.01; '*' – 0.05; '.' – 0.1; 'no symbol' – insignificant.

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Moreover, the Pesaran CD test for cross-sectional dependence reveals a statistically significant result ($z = 2.99$, $p \approx 0.0028$), suggesting that the residuals of different cross-sectional units (countries) are not independent, likely due to unobserved standard shocks, spillover effects, or global policy convergence. This has important implications for the interpretation of the panel estimates and justifies the robust application of estimators to such dependencies, such as Driscoll-Kraay standard errors.

Finally, the Lagrange Multiplier (Breusch-Pagan) test for panel effects yields a vast and significant test statistic ($\chi^2 = 3331.3$, $p < 0.0001$), clearly confirming the presence of unobserved individual (country-specific) effects. This supports the appropriateness of using a panel model over a pooled OLS approach and validates the initial model choice. Taken together, these diagnostic tests underscore the need to apply robust estimation techniques, especially those correcting for heteroskedasticity, autocorrelation, and cross-sectional dependence, to ensure the validity of the inference.

Applying robust standard errors through Driscoll-Kraay and clustered approaches enhances the reliability of inference by addressing issues identified earlier, such as serial correlation, heteroskedasticity, and cross-sectional dependence, in the random effects model for the dependent variable y1_yj. The Driscoll-Kraay standard errors (with maxlag = 2) provide heteroskedasticity- and autocorrelation-consistent inference, robust to cross-sectional dependence. Under this specification, several coefficients are statistically significant at conventional levels: x5_yj ($p < 0.001$), x7_yj ($p \approx 0.019$), x8_yj, x9_yj, and x10_yj (all $p < 0.001$), as well as x11_yj ($p \approx 0.023$). The variables x1_yj, x2_yj, x3_yj, and x6_yj lose statistical significance in this specification, suggesting that autocorrelated or dependent error structures might have driven their earlier significance.

The inference landscape shifts when using cluster-robust standard errors by group (country). While the point estimates remain unchanged, standard errors inflate for some variables. Notably, x6_yj becomes significant ($p \approx 0.040$), reaffirming its importance when accounting for intra-country error correlation. x8_yj remains highly significant ($p < 0.001$), while x11_yj regains statistical significance ($p \approx 0.019$). Variables such as x5_yj, though still negative and sizeable in effect, no longer meet the 5% threshold for significance. This highlights how cross-sectional clustering affects the precision of estimates, which is critical in policy-oriented modelling.

The cluster-robust standard errors by time (year) offer yet another robustness check. Here, x5_yj, x7_yj, x8_yj, x9_yj, x10_yj, and x11_yj emerge statistically significant, suggesting these predictors exhibit consistent influence across temporal shocks or macro-level dynamics. The standard errors for time-clustered models are generally smaller for some variables (e.g., x7_yj, x11_yj), possibly reflecting a lower degree of autocorrelation or intertemporal volatility in the data structure than group-level variation. Notably, x7_yj becomes significant at the 1% level in this specification ($p \approx 0.0038$), suggesting its effect is more temporally than cross-sectionally consistent.

The robust estimation results underscore the robust and persistent influence of variables x8_yj, x9_yj, and x10_yj on the outcome across all specifications. Variables like x5_yj and x11_yj show more specification-sensitive significance, while x1_yj, x2_yj, and x3_yj remain statistically insignificant under robust methods. These findings highlight the need to interpret model results within a robustness-aware framework, especially when working with panel data that violate classical OLS assumptions.

The results from the random-effects specification reveal apparent cross-country heterogeneity (Table 5). Countries such as Bulgaria, Czechia, Slovenia, Spain, Greece, and Germany display positive random intercepts, indicating systematically higher dependent variable levels once standard covariates are controlled. These positive effects suggest that structural or institutional conditions in these economies favor higher outcomes beyond what the policy instruments explain. The most significant positive deviations are observed in Bulgaria (0.86), Slovenia (0.73), and Czechia (0.81), all significant at the 5% level given their relatively small standard errors.

Conversely, countries such as Chile, Mexico, Peru, Saudi Arabia, Costa Rica, and Argentina show

Table 4. Robust standard errors from the random effects model for y1_yj

Variable	Estimate	DK SE	DK p-value	Clustered SE (Group)	p-value	Clustered SE (Time)	p-value
(Intercept)	0.0007	0.3238	0.998	0.1340	0.996	0.2444	0.998
x1_yj	0.0527	0.0485	0.278	0.0395	0.182	0.0410	0.199
x2_yj	0.0441	0.0475	0.353	0.0467	0.345	0.0397	0.266
x3_yj	–0.0112	0.0219	0.611	0.0332	0.737	0.0174	0.522
x4_yj	0.0328	0.0255	0.199	0.0687	0.633	0.0264	0.215
x5_yj	–0.0402	0.0094	0.000***	0.0300	0.180	0.0101	0.000***
x6_yj	0.0236	0.0200	0.240	0.0115	0.040*	0.0149	0.115
x7_yj	–0.0421	0.0179	0.019*	0.0262	0.109	0.0145	0.004**
x8_yj	0.1159	0.0230	0.000***	0.0292	0.000***	0.0194	0.000***
x9_yj	0.0714	0.0102	0.000***	0.0413	0.084.	0.0107	0.000***
x10_yj	0.1705	0.0217	0.000***	0.1053	0.106	0.0220	0.000***
x11_yj	–0.0573	0.0251	0.023*	0.0243	0.019*	0.0289	0.048*

Note: Dependent variable: y1_yj (Yeo-Johnson transformed renewable electricity generation); Model: Random Effects (Swamy-Arora); Driscoll-Kraay SEs: robust to heteroskedasticity, serial correlation, and cross-sectional dependence; Clustered SEs: robust to within-country correlation only; SEs: Robust to time-based clustering (year-level shocks). Signif. codes: '***' – 0.001; '**' – 0.01; '*' – 0.05; '.' – 0.1; 'no symbol' – insignificant.

substantially adverse random effects, pointing to lower baseline levels than the pooled benchmark. For instance, Saudi Arabia (–1.23) and Costa Rica (–1.22) are the strongest negative outliers. These adverse effects may reflect structural barriers, institutional inefficiencies, or external constraints that systematically reduce the effectiveness of the policy instruments. The clustering of adverse effects among several Latin American countries (e.g., Argentina, Chile, Colombia, Mexico, and Peru) highlights potential regional patterns worth further investigation.

3.2. Renewable electricity supply

The FE and RE models (Table 6) reveal several important insights into how eleven policy instruments (x1_yj to x11_yj) impact the normalized

Table 5. Country-specific random effects with standard errors for y1

Country	Effect	SE	Country	Effect	SE
Argentina	–0.763	0.1531	Italy	0.509	0.1478
Australia	–0.249	0.1478	Japan	–0.140	0.1478
Austria	0.502	0.1478	Korea	–0.053	0.1478
Belgium	0.060	0.1478	Latvia	0.395	0.1478
Bulgaria	0.859	0.1478	Lithuania	0.303	0.1478
Canada	–0.130	0.1478	Luxembourg	–0.109	0.1478
Chile	–0.949	0.1478	Malta	–0.264	0.1478
China	–0.511	0.1478	Mexico	–0.955	0.1478
Colombia	–0.761	0.1478	The Netherlands	0.500	0.1478
Costa Rica	–1.217	0.1478	New Zealand	–0.671	0.1478
Croatia	0.250	0.1478	Norway	0.216	0.1478
Czechia	0.805	0.1478	Peru	–0.963	0.1478
Denmark	0.340	0.1478	Poland	0.288	0.1478
Estonia	0.661	0.1478	Portugal	0.444	0.1478
Finland	0.369	0.1478	Romania	0.403	0.1478
France	0.481	0.1478	Saudi Arabia	–1.227	0.1478
Germany	0.559	0.1478	Slovakia	0.581	0.1478
Greece	0.599	0.1478	Slovenia	0.735	0.1478
Hungary	0.611	0.1478	South Africa	–0.397	0.1478
Iceland	–0.685	0.1478	Spain	0.637	0.1478
India	–0.596	0.1478	Sweden	0.222	0.1478
Indonesia	–0.624	0.1478	Switzerland	–0.137	0.1478
Ireland	0.465	0.1478	Türkiye	–0.374	0.1478
Israel	–0.499	0.1478	The United Kingdom	0.482	0.1478

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and transformed renewable electricity supply (y2_yj) across 48 countries over 2010–2023. The fixed effects model explains approximately 44.2% of the within-country variation in y2_yj (adjusted R^2 = 0.389), with an overall highly significant F-statistic. The Hausman test fails to reject the null hypothesis (p = 0.987), suggesting the random effects model is consistent and preferable under standard assumptions.

FE and RE models yield similar coefficient estimates and significance patterns, lending robustness to the findings. Specifically, variables x1_yj (ETS price), x8_yj (planning for renewables), x9_yj (air emission standards), and x10_yj (fossil fuel excise taxes) show strong, positive, and statistically significant effects on renewable electricity supply in both models. For instance, x1_yj has an estimated effect of around 0.066 in both models (p < 0.001), indicating that higher carbon pricing in the ETS sector is associated with increased renewable electricity supply. Similarly, x10_yj consistently yields the highest positive coefficient (~ 0.12), suggesting that excise taxes on fossil fuels may be particularly effective policy levers.

Conversely, x3_yj (carbon tax) and x5_yj (FiTs) are negatively associated with y2_yj in both models and statistically significant at the 5% level. This may indicate substitution effects or diminishing returns when these instruments are not well-integrated into broader policy frameworks. Variables such as x2_yj (ETS coverage) and x4_yj (fossil fuel subsidy reform) approach significance in the random effects model but are not consistently strong predictors across both specifications.

These results suggest that carbon pricing mechanisms (x1, x10), supportive regulation and planning (x8, x9), and robust policy environments are central in driving renewable electricity supply. The consistency across FE and RE models, with only marginal differences in coefficient size and significance, strengthens the confidence in these findings. Nonetheless, attention should be paid to potentially collinear or conflicting instruments such as FiTs or standalone carbon taxes (x3, x5), whose effectiveness appears conditional.

The diagnostic tests applied to the random effects model for the dependent variable y2_yj indicate the presence of several econometric issues that must be addressed to ensure the robustness of the estimation results. The Breusch-Godfrey/Wooldridge test for serial correlation yields a highly significant result (χ^2 = 258.72, df = 13, p < 0.0001), providing strong evidence of serial correlation in the idiosyncratic errors. This violates the standard assumption of independently distributed errors and suggests that the model's standard errors may be biased if uncorrected.

Table 6. FE vs RE model for y2_yj

Variable	FE Estimate	FE Std. Error	FE p-value	RE Estimate	RE Std. Error	RE p-value
(Intercept)	–	–	–	–0.00217	0.145828	0.98811
x1_yj	0.066038	0.014448	< 0.0001***	0.065825	0.014383	< 0.0001***
x2_yj	0.031581	0.019038	0.09766.	0.032957	0.018923	0.08157.
x3_yj	–0.03532	0.014752	0.01696*	–0.03451	0.014637	0.01839*
x4_yj	0.031157	0.026519	0.24051	0.035048	0.026037	0.17828
x5_yj	–0.02108	0.010597	0.04712*	–0.02213	0.01055	0.0359*
x6_yj	0.019449	0.008898	0.02921*	0.018603	0.008869	0.03595*
x7_yj	–0.02064	0.024687	0.40354	–0.02193	0.024304	0.36693
x8_yj	0.065826	0.011157	< 0.0001***	0.06564	0.01111	< 0.0001***
x9_yj	0.066561	0.015322	< 0.0001***	0.066515	0.015215	< 0.0001***
x10_yj	0.119069	0.018574	< 0.0001***	0.11698	0.018422	< 0.0001***
x11_yj	0.006876	0.022891	0.76399	0.007803	0.022586	0.72974

Note: Dependent variable: Renewable electricity supply (Yeo–Johnson transformed); Panel structure: Unbalanced; 48 countries over 13–24 years (N = 665); FE model: R^2 = 0.442; Adj. R^2 = 0.389; F(11, 606) = 43.71; p < 0.001; RE model: R^2 = 0.426; Adj. R^2 = 0.416; Chi^2(11) = 483.56; p < 0.001; Hausman Test: χ^2 = 3.23; df = 11; p = 0.9874 → RE is consistent. Signif. codes: '***' – 0.001; '**' – 0.01; '*' – 0.05; '.' – 0.1; 'no symbol' – insignificant.

Similarly, the Breusch-Pagan test for heteroskedasticity is also significant (BP = 217.35, df = 11, $p < 0.0001$), indicating the presence of heteroskedasticity in the residuals. This means that the variance of the error terms is not constant across observations, which again could undermine the efficiency and validity of standard inference procedures unless heteroskedasticity-robust standard errors are used.

In addition, the Pesaran CD test detects strong cross-sectional dependence ($z = 8.6519, p < 0.0001$), suggesting that the residuals of different countries are correlated. This could be due to shared global shocks, policy diffusion, or unobserved common factors that affect renewable electricity supply across countries. Ignoring cross-sectional dependence can lead to downward-biased standard errors and misleading conclusions.

Lastly, the Lagrange Multiplier (Breusch-Pagan) test confirms the relevance of panel data structure ($\chi^2 = 3653.1$, df = 1, $p < 0.0001$), indicating the presence of significant individual effects. This supports panel estimation methods such as random or fixed effects over pooled OLS. Collectively, these results justify the need for robust estimation techniques, such as Driscoll-Kraay or clustered standard errors, which correct for serial correlation, heteroskedasticity, and cross-sectional dependence.

To assess the robustness of the Random Effects model linking policy instruments (x1_yj to x11_yj) to renewable electricity outcomes (y2_yj), three alternative methods for calculating standard errors were employed: Driscoll–Kraay (which accounts for heteroskedasticity, serial correlation, and cross-sectional dependence), cluster-robust SEs by country (group), and cluster-robust SEs by year (time) (Table 7). While the estimated coefficients remain constant across these approaches, as expected since they are derived from the same underlying RE model, their associated standard errors and statistical significance levels vary considerably, indicating that the choice of error correction substantially influences inference.

Several variables emerge as statistically significant under the Driscoll-Kraay correction, which is well-suited for panels with both time-series and cross-sectional dependence. Specifically, x3_yj (Carbon Tax), x8_yj (Ban on coal plants), x9_yj (Planning for renewables expansion), and x10_yj (Air emission standards) all demonstrate strong positive associations with renewable electricity share, suggesting that these policies are effective levers for promoting renewable energy. Notably, x3_yj is negative, indicating that carbon taxes may correlate with short-term reductions in renewable shares, possibly reflecting market distortions or implementation lags. x5_yj (Renewable energy certificates) is borderline significant ($p \approx 0.057$), warranting further scrutiny.

In contrast, clustering by country (group), which controls for within-country autocorrelation and heteroskedasticity, reduces statistical significance across most predictors. Only x8_yj retains significance at the 5% level, while x1_yj (ETS price), x6_yj (Renewable energy auctions), and x9_yj are marginally significant at the 10% threshold. The weakened significance may be due to substantial heterogeneity in policy effectiveness across countries or high intra-group correlation, which inflates standard errors. This clustering strategy is particularly conservative when the number of clusters (countries) is moderate, as in this study.

Clustering by time, which accounts for temporal shocks or policy waves affecting all countries simultaneously (e.g., international climate summits or fuel price shocks), again alters the pattern of significance. Variables x3_yj, x8_yj, x9_yj, and x10_yj are highly significant, reaffirming the robustness of their effects over time. In addition, x5_yj and x4_yj (Fossil Fuel Subsidy Reform) show weak significance ($p < 0.1$), indicating that their influence may manifest more uniformly across years than across countries. This suggests that temporal events or coordinated global trends may amplify or moderate the effects of specific policy instruments.

The divergence in results across these three estimation techniques highlights the importance of robust inference in panel econometric studies. While Driscoll-Kraay SEs offer a comprehensive correction, time clustering is particularly informative in revealing the impact of globally synchronized policy actions or external shocks. Overall, the findings suggest that instruments such as bans on

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Table 7. Robust standard errors from random effects model for y1_yj

Variable	Estimate	Driscoll-Kraay SE	Clustered SE (Group)	Clustered SE (Time)
(Intercept)	−0.00217	0.34674	0.136821	0.254616
x1_yj	0.065825	0.054841	0.038603.	0.045777
x2_yj	0.032957	0.065665	0.041177	0.047319
x3_yj	−0.03451	0.016559*	0.051014	0.013611*
x4_yj	0.035048	0.01987.	0.070371	0.020066.
x5_yj	−0.02213	0.011592.	0.030736	0.012194.
x6_yj	0.018603	0.016656	0.010815.	0.013089
x7_yj	−0.02193	0.042053	0.086303	0.036928
x8_yj	0.06564	0.024429**	0.028725*	0.020793**
x9_yj	0.066515	0.008056***	0.036794.	0.008365***
x10_yj	0.11698	0.034134***	0.107156	0.026715***
x11_yj	0.007803	0.015295	0.036987	0.01642

Note: Model – Random Effects, Dependent variable: y2_yj (Yeo–Johnson transformed renewable electricity supply); Driscoll-Kraay SEs: robust to heteroskedasticity, serial correlation, and cross-sectional dependence; Clustered SEs: robust to within-country correlation only; SEs: Robust to time-based clustering (year-level shocks). Signif. codes: '***' – 0.001; '**' – 0.01; '*' – 0.05; '.' – 0.1; 'no symbol' – insignificant.

coal, air emission standards, and renewables' strategic planning consistently positively influence renewable electricity development, regardless of the SE correction method. Meanwhile, other instruments show context-dependent effects that require deeper investigation into their national implementation dynamics and possible lagged impacts.

Estimating country-specific random effects reveals substantial heterogeneity in how nations perform relative to the average baseline once standard policy instruments, such as ETS electricity prices, fossil fuel subsidy reforms, feed-in tariffs, renewable energy auctions, and excise taxes, are controlled for. Countries like Bulgaria, Czechia, Slovenia, Spain, Greece, and Germany show consistently positive effects, signaling structural conditions that enable more effective renewable electricity deployment and supply expansion. These positive intercept shifts suggest that, beyond pol-

Table 8. Country-specific random effects with standard errors for y2

Country	Effect	SE	Country	Effect	SE
Argentina	−0.763	0.1531	Italy	0.509	0.1478
Australia	−0.249	0.1478	Japan	−0.140	0.1478
Austria	0.502	0.1478	Korea	−0.053	0.1478
Belgium	0.060	0.1478	Latvia	0.395	0.1478
Bulgaria	0.859	0.1478	Lithuania	0.303	0.1478
Canada	−0.130	0.1478	Luxembourg	−0.109	0.1478
Chile	−0.949	0.1478	Malta	−0.264	0.1478
China	−0.511	0.1478	Mexico	−0.955	0.1478
Colombia	−0.761	0.1478	The Netherlands	0.500	0.1478
Costa Rica	−1.217	0.1478	New Zealand	−0.671	0.1478
Croatia	0.250	0.1478	Norway	0.216	0.1478
Czechia	0.805	0.1478	Peru	−0.963	0.1478
Denmark	0.340	0.1478	Poland	0.288	0.1478
Estonia	0.661	0.1478	Portugal	0.444	0.1478
Finland	0.369	0.1478	Romania	0.403	0.1478
France	0.481	0.1478	Saudi Arabia	−1.227	0.1478
Germany	0.559	0.1478	Slovakia	0.581	0.1478
Greece	0.599	0.1478	Slovenia	0.735	0.1478
Hungary	0.611	0.1478	South Africa	−0.397	0.1478
Iceland	−0.685	0.1478	Spain	0.637	0.1478
India	−0.596	0.1478	Sweden	0.222	0.1478
Indonesia	−0.624	0.1478	Switzerland	−0.137	0.1478
Ireland	0.465	0.1478	Türkiye	−0.374	0.1478
Israel	−0.499	0.1478	The United Kingdom	0.482	0.1478

icy interventions, institutional quality, technological readiness, and grid integration strategies are creating favorable conditions for renewables. Bulgaria (0.86) and Slovenia (0.73) stand out as strong over-performers.

On the other hand, countries such as Saudi Arabia, Costa Rica, Chile, Mexico, and Peru register substantial adverse effects, indicating lower baseline outcomes compared to the pooled benchmark. Clustering adverse effects in Latin America (Argentina, Chile, Colombia, Mexico, and Peru) highlights potential regional barriers such as fossil-fuel lock-in, regulatory uncertainty, and limited grid flexibility. Similarly, Asian economies such as India and Indonesia display negative intercepts, consistent with ongoing reliance on coal. These results underline that policy instruments alone are insufficient: broader structural and regional contexts critically determine renewable electricity generation and supply trajectories.

4. DISCUSSION

The empirical findings of this study demonstrate that several public and fiscal policy instruments for supporting renewable electricity development, most notably feed-in tariffs, carbon taxation, emissions trading schemes, and fossil fuel subsidy reforms, exert a statistically significant and positive influence on renewable electricity generation and supply. Specifically, the fixed-effects model with cluster-robust standard errors by year confirms the significance of variables such as x1 (ETS electricity price), x2 (ETS coverage GHG), x5 (feed-in tariffs), and x10 (air emission standards), while controlling for unobserved heterogeneity. These results are consistent with the broader academic consensus that a mix of regulatory and market-based instruments is required to accelerate renewable energy deployment.

These outcomes align with findings from prior studies such as Samusevych et al. (2024) and Kusumawati et al. (2025), which showed that emissions trading schemes become more effective when accompanied by planning regulations and complementary standards. The present study further supports the claim that carbon taxes and ETS mechanisms are foundational instruments in the climate policy mix, especially when public acceptance and institutional capacity are sufficient to ensure enforcement and credibility. This echoes concerns of Gualandri and Kuzior (2023) and Vakulenko and Rekunenko (2025), who highlighted limitations in carbon pricing effectiveness within transitional economies due to institutional weaknesses.

Additionally, this paper affirms the enduring relevance of feed-in tariffs, particularly in attracting investment in early-phase or financially constrained markets. The statistical robustness of FiTs in our panel supports conclusions made by Lyeonov et al. (2025) and Shcherbakova (2025) that FiTs remain instrumental in unlocking long-term financing despite their declining popularity in some mature markets. Conversely, while renewable energy auctions have been promoted as more efficient alternatives, the empirical models in this study found their impact to be weaker and statistically insignificant, which aligns with critiques from Dobrovolska et al. (2024b) and Hedau and Malla (2024) regarding challenges in bidder performance and policy inclusivity.

Including governance-related policy tools, such as air emission standards, planning mandates, and coal plant bans, also yielded important insights. The positive effects of these instruments mirror the literature's emphasis on policy coherence and investor confidence, as discussed by Batlles-delaFuente et al. (2024) and Streimikiene (2025). Moreover, the results reinforce the notion of cumulative policy effects, where combinations of fiscal, regulatory, and planning tools are more effective than isolated measures, supporting frameworks proposed by Aleknevičienė et al. (2025) and Hasan et al. (2025).

The present analysis also highlights that fossil fuel excise taxes and subsidy reforms are crucial drivers of renewable electricity development. These findings echo the arguments made by Samusevych et al. (2024) and Badreddine and Cherif (2024) that removing distortive subsidies promotes cleaner alternatives and generates fiscal space for additional climate investments. Nonetheless, their effectiveness appears to vary across country contexts and is contingent upon vigorous policy enforcement and transparency.

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This study contributes to the expanding body of empirical research emphasizing the necessity of diversified and context-sensitive public and fiscal policy instruments for supporting renewable electricity development. While the impact magnitude of individual instruments varies, the overall evidence supports a strong and positive relationship between coherent public and fiscal policy portfolios and renewable electricity advancement, thus affirming theoretical and policy-oriented conclusions in the existing literature.

Notwithstanding the robustness of the panel data analysis and the application of advanced econometric techniques, several limitations should be acknowledged. First, the dataset is unbalanced, with missing observations for several countries and years, particularly among non-OECD states, which may constrain the generalizability of the results. Second, the public and fiscal policy indicators employed are predominantly categorical and therefore do not fully reflect the intensity, enforcement, or quality of implementation of specific instruments, especially in heterogeneous political and economic environments. Third, the analysis is restricted to policy variables, excluding potentially influential control factors such as energy demand, gross domestic product per capita, technological readiness, and grid infrastructure development, which may mediate or moderate the observed effects. Although robust standard error corrections were applied, the scope for causal inference remains limited, and potential endogeneity between renewable outcomes and policy adoption was not explicitly addressed. Another limitation concerns the sample composition, which includes only OECD members, OECD partner countries, and OECD negotiating members. This restriction reflects the availability of harmonized indicators within the OECD Climate Actions and Policies Measurement Framework. However, it simultaneously reduces the external validity of the findings, as the stringency of policy measures is systematically documented only for these states.

Future research should seek to enhance the dataset by incorporating more granular and continuous measures of public and fiscal policy design, such as tariff levels, auction volumes, and certificate prices, alongside macroeconomic, institutional, and technological covariates. Applying dynamic panel models, including system GMM estimators, or causal inference approaches such as difference-in-differences and synthetic control methods, would strengthen causal identification. Further progress may also be achieved by disaggregating the analysis according to income groups, geographical regions, or clusters of public and fiscal policy instruments, which could illuminate context-specific dynamics and generate tailored recommendations. Finally, with the increasing availability of sub-national data, future investigations should consider intra-country public and fiscal policy heterogeneity, particularly in federal or decentralized governance systems.

CONCLUSION

This study evaluates how budgetary measures, taxation schemes, subsidies, and regulatory standards influence renewable electricity outcomes, thereby linking climate policy design with broader public and municipal finance issues.

The empirical results highlight several statistically significant relationships. For renewable electricity supply (y2), the most consistently positive and robust predictors across specifications were feed-in tariffs (x8), renewable planning policies (x9), and air emission standards (x10), with coefficients ranging from 0.065 to 0.117 and significance levels up to $p < 0.001$. Similarly, for renewable electricity generation (y1), these three variables also remained significant, with x10_yj reaching a coefficient of 0.170 ($p < 0.001$) in the random effects model. Some market-based instruments, such as electricity price levels under ETS (x1), showed moderate positive effects. In contrast, fossil fuel excise taxes (x11) had negative or non-significant impacts, potentially reflecting inconsistent implementation or design flaws. Results remained robust after controlling for heteroskedasticity, autocorrelation, and cross-sectional dependence, especially when applied by Driscoll-Kraay standard errors. The country-specific effects demonstrate that, beyond policy

design, structural and institutional conditions decisively shape renewable electricity outcomes. Countries such as Bulgaria, Slovenia, and Czechia systematically outperform the benchmark, while Saudi Arabia, Costa Rica, and several Latin American economies lag significantly behind. This pattern suggests that supportive institutional environments and regional cooperation must complement coherent policy mixes to achieve meaningful and sustained progress in renewable electricity deployment.

From a policy perspective, these findings suggest that regulatory instruments and structured planning efforts (such as clear renewable expansion strategies and emissions performance standards) may offer more immediate and predictable impacts on renewable electricity outcomes than carbon pricing tools alone. While market mechanisms like carbon taxes and ETS are crucial for internalizing externalities, their effectiveness appears contingent upon coverage thresholds, pricing levels, and accompanying regulatory support. Feed-in tariffs effectively incentivize renewable electricity deployment, especially in contexts where auction mechanisms or certificate schemes are less mature.

Therefore, governments should complement market-based policies with enforceable regulatory standards and clear, long-term planning mandates, especially in emerging or fossil-fuel-reliant economies. Additionally, careful attention must be paid to the policy mix, as overlapping instruments can reinforce or dilute effectiveness. The observed cross-sectional dependence also implies that countries are not acting in isolation; thus, regional coordination and policy harmonization (e.g., through the EU, IEA, or ASEAN) could yield collective gains in accelerating the renewable transition. Lastly, expanding data coverage, improving policy stringency metrics, and integrating technological and grid-related variables could strengthen future analyses in this field.

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APPENDIX A

The sample of countries for analysis:

Argentina, Australia, Austria, Belgium, Bulgaria, Canada, Chile, China, Colombia, Costa Rica, Croatia, Czechia, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, India, Indonesia, Ireland, Israel, Italy, Japan, Korea (South), Latvia, Lithuania, Luxembourg, Malta, Mexico, the Netherlands, New Zealand, Norway, Peru, Poland, Portugal, Romania, Saudi Arabia, Slovakia, Slovenia, South Africa, Spain, Sweden, Switzerland, Türkiye, and the United Kingdom.

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