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RESEARCH PAPER

Exploring the Influence of Female Political Participation on CO_2 and Greenhouse Gas Emissions: Insights from the European Union

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Abstract
This study examines the impact of female political participation on environmental degradation in European Union (EU) member states from 2013 to 2023. Drawing on feminist ecological economics, environmental justice, and inclusive governance theory, the analysis assesses whether greater gender representation in political institutions is associated with reduced CO_2 and greenhouse gas (GHG) emissions. Using panel data for 26 EU countries and applying robust econometric techniques—including fixed effects with Driscoll–Kraay standard errors, quantile regression, and instrumental variable methods—the study finds that higher female political representation significantly reduces per capita emissions, with stronger effects in high-emission contexts. Mediation analysis further reveals that women's participation enhances clean household energy use and strengthens governance, though government effectiveness alone shows a positive link with emissions, reflecting economic expansion. To complement these econometric results, a Random Forest a supervised machine learning analysis is conducted, confirming the importance of female political participation alongside GDP per capita, urbanization, and energy consumption as key predictors of emissions. By providing EU-wide empirical evidence on the gender–environment nexus, the study highlights the transformative potential of inclusive institutions for achieving the goals of the European Green Deal and underscores the importance of integrating gender equality into climate policies.

Keywords Female political participation · Environmental degradation · European Union · Sustainable Governance · Gender and environment nexus · Machine learning

Introduction

In recent decades, environmental degradation has emerged as one of the most urgent concerns, intersecting with aspects of governance, equity and social inclusion. In the context of growing demands for sustainable transitions, the inclusivity of political institutions, especially in terms of gender representation has gained recognition.

The European Union (EU), as a global leader in climate action and gender equality, offers an appropriate setting to

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investigate how female political representation influences environmental performance. Despite policy advancements, women are still underrepresented in the main political institutions across EU member states. Only 32–40% of seats are held by women at local, national and EU levels (Zamfir 2024). Examining how this underrepresentation influences environmental outcomes is important for policymaking.

Empirical studies have shown that female political leadership tends to be associated with a stronger emphasis on environmental protection. For instance, Ergas and York (2012) found that empowering women politically can contribute to better environmental outcomes. Similarly, Mavisakalyan and Tarverdi (2019) showed that greater female parliamentary participation correlates with stronger environmental policies, while Sundström and McCright (2014) noted women's greater environmental concern and stronger pro-environmental values. Yet empirical evidence -especially in a cross-national EU context- remains limited. This motivates this study's exploration of the gender–environment nexus within the EU context.

Even if there is a growing body of research linking political institutions to environmental outcomes, most studies have focused on governance quality, regulatory capacity and political regimes. Democratic institutions characterized by social-liberal features such as participatory processes, social inclusivity and government accountability are associated with stronger environmental regulations and lower pollution levels (Povitkina and Jagers 2022). However, the role of gender representation in political institutions remains understudied in comparative EU contexts. This paper seeks to fill this gap by examining whether greater female political participation is linked to lower levels of CO_2 and greenhouse gas (GHG) emissions across EU countries. It is based on panel data and governance indicators to assess this relationship and its underlying mechanisms.

This study investigates the relationship between gender representation and environmental outcomes within the EU, emphasizing the mediating influence of government effectiveness and household energy consumption in shaping emission trajectories. The following research question can be formulated: *How does female political participation influence per capita CO_2 and GHG emissions across EU member states, and to what extent do government effectiveness and household energy consumption mediate this relationship?* Addressing this question sheds light into the role of inclusive governance in shaping environmental performance.

The theoretical foundation of this study draws on feminist ecological economics, environmental justice, and inclusive governance theory. These perspectives suggest that institutions characterized by gender inclusivity are more responsive and environmentally responsible. Women in political leadership are often more likely to advocate for environmental protection and support inclusive decision-making processes. Based on this theoretical framework, the study proposes two hypotheses:

H_1 *Higher female political participation is associated with lower per capita CO_2 and GHG emissions.*

H_2 *This relationship is mediated by improvements in governance effectiveness and greater adoption of sustainable household energy consumption.*

To enhance analytical clarity, hypothesis H_2 is further examined as two sub-hypotheses:

H_{2a} *The relationship between female political participation and per capita CO_2 and GHG emissions is mediated by improvements in governance effectiveness.*

H_{2b} *The relationship between female political participation and per capita CO_2 and GHG emissions is mediated by greater adoption of sustainable household energy consumption.*

Existing literature supports these expectations. Using panel data from 49 countries, Sanchez-Olmedo et al. (2025) find that a one percent increase in the Global Gender Gap Index is associated with a reduction of 0.14 metric tons of CO_2 emissions per capita. Gonda (2024) applies quantile regression to BRICS economies and shows that the environmental benefits of women's empowerment are stronger at higher quantiles of emissions. Female leadership is most effective in contexts of elevated environmental stress, making it an essential driver for high-emission countries.

These hypotheses posit that greater female political participation fosters robust environmental policies, both directly and indirectly via enhanced institutions and sustainable household energy practices.

This study has four specific objectives:**O1**: To measure the effect of female political participation on per capita CO_2 and GHG emissions in EU member states; **O2**: To evaluate whether government effectiveness mediates this relationship; **O3**: To assess the role of female political representation in shaping household energy consumption patterns, particularly the use of clean versus fossil fuels; and **O4**: To analyze cross-national variation in these relationships within the EU.

Together, these objectives define the analytical scope and empirical contribution of the study. Using panel data from EU member states and applying fixed-effects estimation, Driscoll–Kraay standard errors, quantile regression, and panel causality tests, the study finds that higher female political representation is significantly associated with lower per capita CO_2 and GHG emissions. In addition to conventional econometric methods, this study also incorporates machine learning analysis through Random Forest regressions. This non-parametric approach complements the econometric estimations by capturing complex, non-linear relationships and identifying the relative importance of predictors of emissions. Integrating both econometric and machine learning techniques strengthens the robustness and credibility of the findings.

The originality of the study lies in providing new perspectives of the role of gender-inclusive institutions in shaping environmental performance in a robust, cross-national EU context. While theoretical perspectives and prior empirical studies suggest a link between gender inclusion and environmental outcomes, few studies have explored this relationship within a unified, EU-wide framework. This study addresses this critical research gap by applying a comprehensive panel data methodology to assess whether gender parity in political decision-making enhances environmental quality.

From a policy perspective, the findings highlight the importance of integrating gender equality into climate action strategies, including the European Green Deal. Empowering women politically can serve as a catalyst for more inclusive, equitable and effective environmental governance in the EU.

The paper is structured as follows. Sect. "Literature review" reviews the relevant literature. Section 3 presents the data and methodology. Sect. "Discussions" discusses the empirical results. Sect. "Conclusions and policy implications" discusses the main findings. Section 6 concludes with policy implications.

Literature Review

Female political leadership has the potential to promote environmental sustainability in advanced economies, particularly when women hold top government positions (Wu 2024). However, when examining the effects across different levels of governance in Europe, findings suggest a more nuanced picture: while women's participation at the ministry level is associated with an increase in air pollution, participation at lower levels—such as membership in parliaments and regional assemblies—is linked to reductions in air pollution. This supports the argument that, in male-dominated political environments, women at higher levels may adopt traditionally male policy approaches (Tremblay 1998).

Nevertheless, policymaking has evolved. Contrary to Tremblay's (1998) findings, more recent research shows that women in politics today not only advocate for women's issues but also place greater emphasis on climate action (Salamon 2023). Higher environmental awareness and greater participation in activist movements among women (Koengkan et al. 2024) contribute to the development of more climate-friendly policies. These effects tend to be more pronounced and immediate in higher-income democratic countries (Salamon 2023). In Europe, women's political participation has increased over time, as illustrated in Fig. 1 below.

Figure 1 above illustrates the gender distribution in political participation across three spheres within the European Union: ministries, parliaments, and regional assemblies. As of 2023, while some progress has been made in promoting gender equality in certain regions, men continue to dominate political roles overall. For example, Finland has achieved 55% female representation in ministerial positions and 46% in both parliamentary and regional assemblies. Similarly, Sweden reports 51% female ministers, while Spain shows 46% female ministers and 42% female members of parliament (MPs), as reflected in Figures A, C, and E. In contrast, some countries remain strongly male-dominated. Romania, for instance, has 90% male ministers and 80% male MPs and regional representatives. Cyprus also demonstrates a significant gender imbalance in political representation, as shown in Figures B, D, and F.

Overall, men hold the majority of political positions across the EU, occupying 67% of parliamentary and cabinet roles and 70% of regional assembly seats, compared to just 33% and 30% held by women, respectively. This gender imbalance contrasts with the European Union's stated commitment to participatory governance, as outlined in the Lisbon Treaty (European Commission 2025a), which emphasizes citizen engagement and inclusive policymaking. To address these disparities, the EU has implemented a range of initiatives aimed at promoting gender equality. The current European Commission, for example, has committed to building a "Union of Equality" (European Commission 2025b), advancing this objective through its Gender Equality Strategy 2020–2025 and other related policies and directives. In parallel, the EU has also made strides in environmental policy, establishing a comprehensive framework of laws, action programs, and strategic initiatives. These efforts focus on critical areas such as climate change, biodiversity conservation, pollution reduction, and sustainable resource management — highlighting the EU's broader agenda for inclusive and sustainable governance.

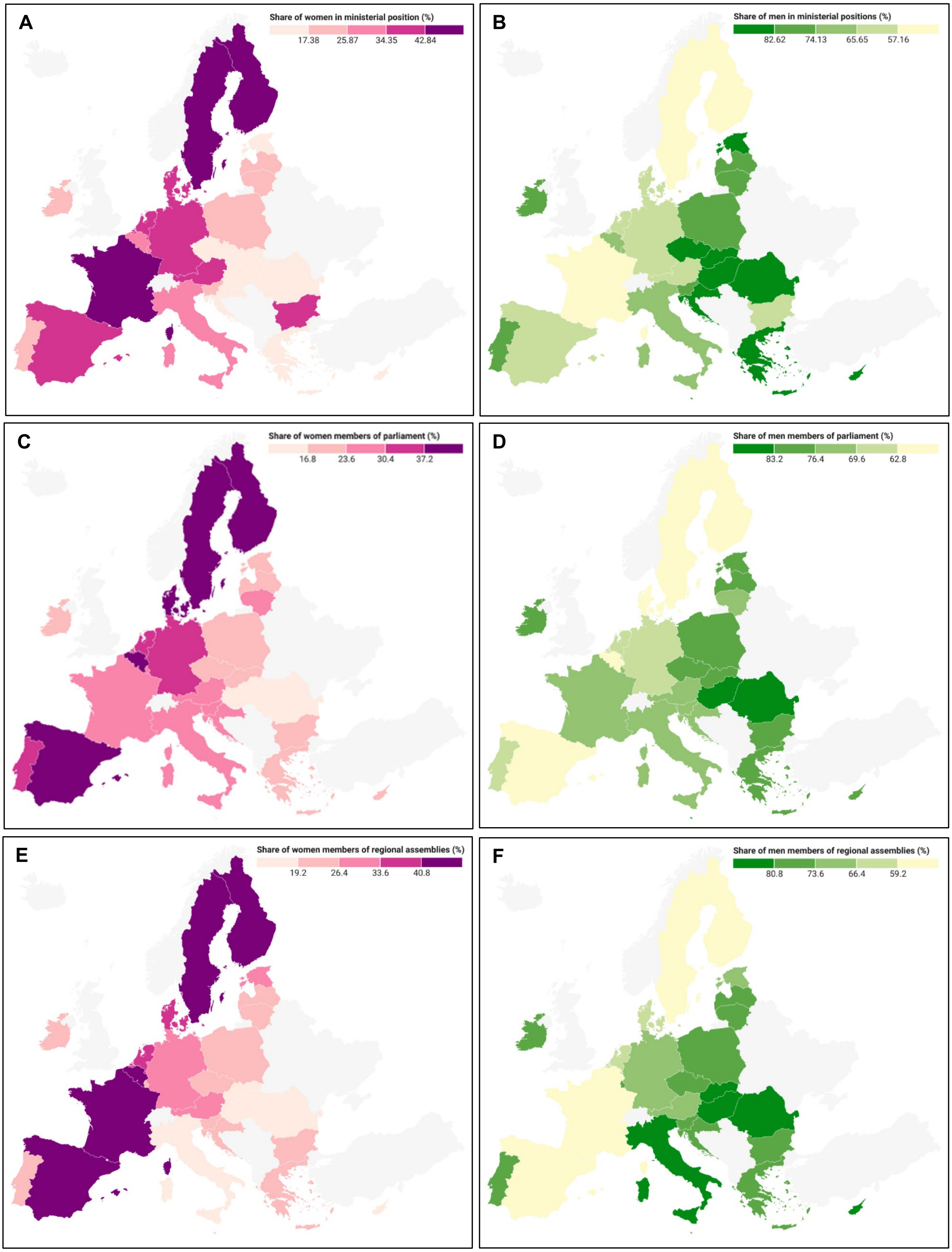
A
Share of women in ministerial position (%)
17.38 25.87 34.35 42.84
B
Share of men in ministerial positions (%)
82.62 74.13 65.65 57.16
C
Share of women members of parliament (%)
16.8 23.6 30.4 37.2
D
Share of men members of parliament (%)
83.2 76.4 69.6 62.8
E
Share of women members of regional assemblies (%)
19.2 26.4 33.6 40.8
F
Share of men members of regional assemblies (%)
80.8 73.6 66.4 59.2

Fig. 1 Gender Distribution in Political Representation (%) in the EU. Notes: This figure presents the gender representation in ministries, parliaments, and regional assemblies across the EU in 2023. Figures **A**, **C**, and **E** represent female political representation, while Figures **B**, **D**, and **F** represent male political representation. The data is based on the European Institute for Gender Equality (EIGE) 2025. This figure was created by the authors

Governance Quality, Institutions, and Environmental Outcomes

Institutional quality—including its dimension of government effectiveness—enables countries to reduce fossil fuel consumption and address urban challenges, partly through mechanisms such as green financing (Sun et al. 2025), thereby contributing to environmental protection (Akpan and Kama 2024). Indeed, effective governance has been demonstrated to bolster environmental protection (Mavragani et al. 2016).In the specific context of the EU, higher institutional quality helps mitigate the negative environmental impacts of tourism and economic growth (Musa et al. 2021). Although institutional quality may also facilitate trends such as increased tourism and economic output—which can have adverse environmental effects—the net impact on environmental performance across the EU-28 remains positive (Usman et al. 2020). It should also be acknowledged that there is potential endogeneity issue and interaction effects (Zhang et al. 2024) between governance and emissions.

Climate and energy policies are influenced not only by the quality of institutions but also by the level of public trust in them. In Europe, political trust significantly shapes the acceptability of regulations (Zannakis et al. 2015). While generalized trust tends to increase support for environmental taxes, it does not necessarily extend to subsidies or bans. In contrast, political trust is more strongly associated with support for taxes than for subsidies or bans (Davidovic and Harring 2020). At the national level, trust in impartial institutions is key to translating concern about climate change into support for climate policies, though trust in partial institutions also plays a role (Kulin and Sevä, 2021). Regarding gendered trust, (Kreutzer 2023) shows that women's institutional trust declines more significantly following natural disasters – as such, there are differences in gender under environmental stress. In the EU countries, men have in general less trust in political support for climate policies (Kulin et al. 2024). As both knowledge and values are essential to the decision-making process, inclusive governance—where decisions are efficient, effective, fair, and morally acceptable—should be the aim (Renn and Schweizer 2009). In this context, including all genders in decision-making is crucial, not only for environmental governance but for decision-making in general.

Gender-Inclusive Institutions and Sustainable Development

There is limited research on the institutional effects of women's empowerment on sustainability as a whole. However, evidence from developing countries suggests that collective action—such as cooperatives, self-help groups, and grassroots environmental movements—has improved women's financial stability, social status, and access to education and health insurance (Arya and Shukla 2025). In contrast, the relationship between gender and the environment is less pronounced in developed countries, where greater distance from the sources of food, energy, and water supply weakens the direct connection (Buckingham 2020). In European policy frameworks, social and environmental justice intersect, notably through the European Green Deal. The Just Transition Mechanism incorporates gender-sensitive indicators; the Green Jobs and Skills agenda promotes gender-balanced access to reskilling opportunities in emerging green sectors; and the Mobility and Energy agenda emphasizes the need for gender-sensitive transport and energy planning. The EU's Gender Equality Strategy further underscores the importance of mainstreaming gender across all EU policies, calling for the integration of gender perspectives into green recovery plans and the promotion of equal opportunities in innovation. Environmental degradation has gendered dimensions, often reflected in disproportionate health impacts, increased care burdens, and limited access to decision-making processes and resources. To address these disparities, participation and power must be equalized. Indeed, countries with higher representation of women in positions of power are more likely to adopt ambitious climate policies (Mavisakalyan and Tarverdi 2019).

Household-Level Energy Use and Environmental Degradation

Household decisions significantly shape greenhouse gas (GHG) emissions, particularly in high-income countries, through factors such as residential energy use, transportation choices, food consumption, and the consumption of goods and services. These decisions are influenced by various factors, including income and wealth (Galvin and Sunikka-Blank 2018), as well as gender and age (Räty and Carlsson-Kanyama, 2009), among others. At the institutional level, elements such as energy subsidies, infrastructure, and public investment also influence these choices on a broader scale. However, democracies do not necessarily manage energy consumption more effectively to enhance well-being compared to other forms of governance (Mayer 2017). Environmental degradation is driven not only by industrial and transportation activities but also by household-level consumption. In Europe, between 2008 and 2020, environmental degradation showed improvement, partly due to a shift in household consumption patterns toward services rather than goods (Jankiewicz 2024).

Gaps in the Literature

The current literature on environmental degradation and governance reveals several notable research gaps that warrant further investigation. First, there is limited integration of female political participation into empirical analyses of environmental outcomes, despite growing evidence that gender-inclusive decision-making can influence sustainability agendas. This omission restricts our understanding of how women's representation in political institutions may contribute to shaping environmental policies and practices. Second, existing studies often overlook the complex ways in which gendered political representation interacts with the quality of governance—such as transparency, accountability, and institutional effectiveness—to affect environmental indicators, including carbon emissions. Finally, there is a marked absence of cross-national panel studies, and panel data studies incorporating mediation, that systematically examine the relationship between institutional gender equality and environmental performance across EU member states. Addressing these gaps is essential to developing a more comprehensive and inclusive framework for understanding the political determinants of environmental sustainability.

Theoretical Framework Linkage

The shift from a paradigm emphasizing women's participation and representation—grounded in intergenerational justice, healing, and sustainability—toward a technological and market-based, solution-oriented Anthropocene rhetoric has diminished the perceived importance of environmental policymaking (Ahmed 2006). In response, ecofeminist perspectives, which seek to reunite nature and culture and argue that women are particularly well positioned to ensure a sustainable future (D'Eaubonne, 1978), have gained renewed attention over the past decade. At the same time, inclusive governance—defined both in terms of process (how decisions are made, who is included, and why), and outcomes (how wealth and prosperity are distributed and shared, and why) (OECD 2020)—has become increasingly central to discussions on climate change and sustainability (Annahar et al. 2023). As sustainability cannot be separated from the individuals and institutions shaping it, social training for democracy must occur through participatory processes. This entails preliminary deliberation and ensuring that each participant has an equal formal say in decision-making (Kaufman 2017, p. 192). Such participatory approaches can lead to improved environmental outcomes by fostering cooperation and the formation of intermediary coalitions at the local level (Martinez et al. 2008). However, critics have noted limitations in applying these theories to larger and more complex industrial societies (Warren 1996, p. 242). Nevertheless, when environmental justice is upheld—understood as inclusion in decision-making, equity in environmental laws and values, and recognition in shaping environmental institutions and knowledge (Figueroa 2023, p. 767)—better sustainability outcomes are also more likely. The integration of feminist ecological economics and environmental justice theories enables a multidimensional understanding of governance, bridging the gap between normative arguments for gender equality and quantifiable environmental outcomes. This approach allows us to derive testable hypotheses grounded in both critical theory and institutional practice. Taken together, a conceptual framework can be drawn that integrates feminist ecological economics, inclusive governance theory, and environmental justice and participatory democracy (see Fig. 2 below).

Data and Empirical Results

This section presents the dataset, describes the methodologies used to achieve the research objectives, and analyzes the empirical results.

Data

The empirical analysis examined the impact of female political participation on environmental degradation across the 26 member states of the European Union (EU), namely: Austria, Belgium, Bulgaria, Croatia, Cyprus, Czechia, Denmark, Estonia, Finland, France, Germany, Greece, Hungary, Ireland, Italy, Latvia, Lithuania, Luxembourg, the Netherlands, Poland, Portugal, Romania, Slovakia, Slovenia, Spain, and Sweden, as illustrated in Fig. 3 below.

The 26 EU member states were selected because of their diverse socioeconomic, political, and environmental contexts, which together provide a comprehensive representation of the Union's landscape. Although this focus on EU countries may limit the immediate global generalizability of the findings, it offers a powerful and controlled setting for analysis. The EU presents a distinctive context in which significant socioeconomic diversity exists among member states operating under a common framework of ambitious supranational environmental and gender equality policies. This institutional consistency enables a clearer identification of the relationship between national-level female political participation and environmental outcomes, while

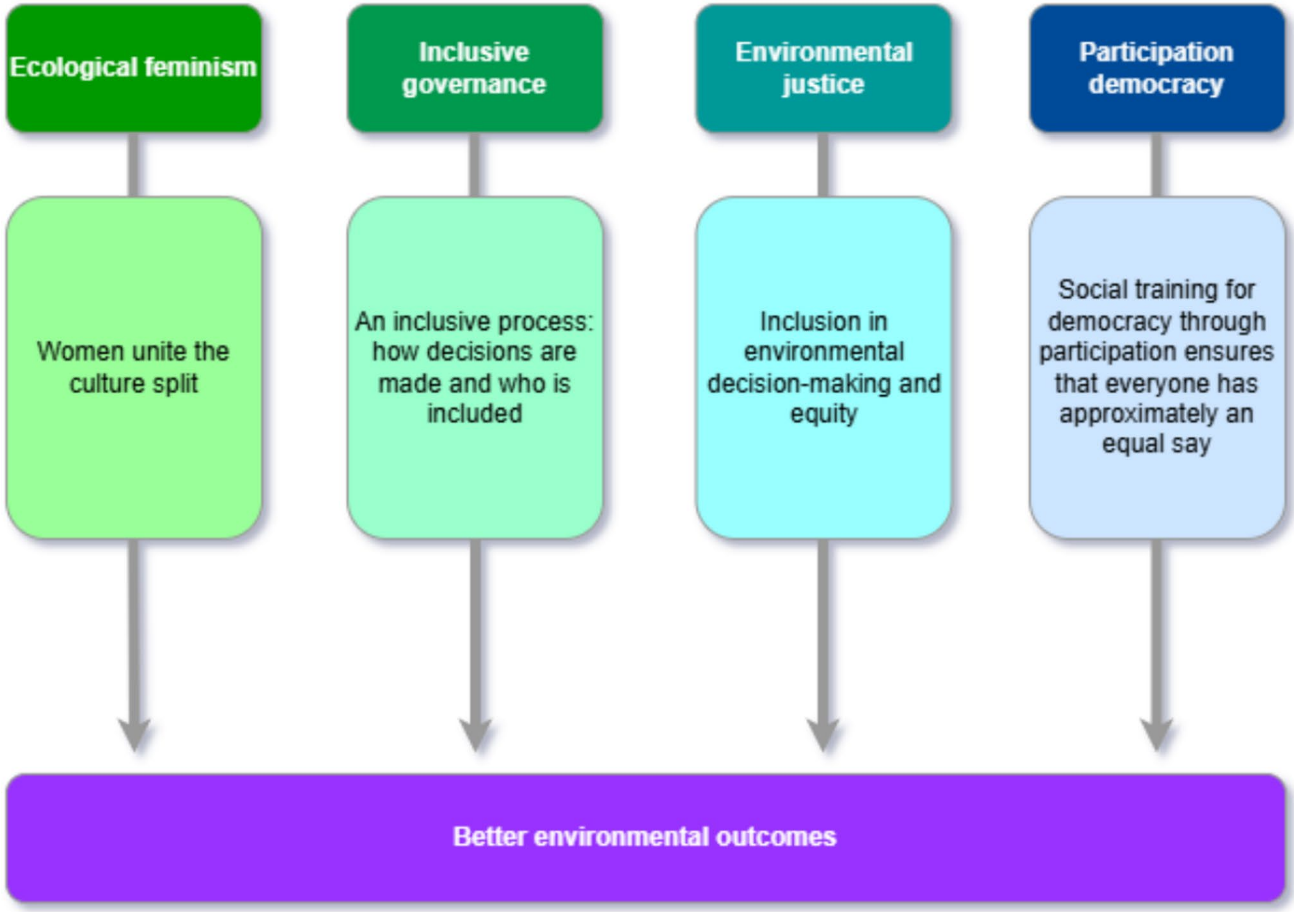

Fig. 2 Conceptual framework linking feminist ecological economics, inclusive governance, and environmental justice to environmental outcomes. The authors created this figure

the internal variation across countries supports robust cross-national comparison within a coherent governance system. Such diversity facilitates a nuanced exploration of the interconnections between gender dynamics, socioeconomic conditions, and environmental degradation. The study spans the period from 2013 to 2023, chosen based on data availability and consistency. Earlier data were excluded due to a lack of standardization, which could have introduced bias into the analysis. Moreover, the country (Malta) was excluded in the analysis due to data regarding the variables CO_2e/cap and GHGe/cap. To strengthen internal validity, the study carefully selects indicators that capture not only emission outputs but also socio-institutional drivers, ensuring a robust empirical framework. Variables were harmonized across time and countries, with rigorous controls for multicollinearity and missing data. Table 1 below presents the key dependent and independent variables used in the study.

To construct the FFP_Index, we apply equal weights to women's representation in ministerial posts (W_min), national parliaments (W_parl), and regional assemblies (W_reg):

$$FFP_{Index} = \frac{W_{min} + W_{parl} + W_{reg}}{3} \tag{1}$$

where, W_{min} is share of women in ministerial positions, W_{parl} is share of women in parliaments, and W_{reg} is share of women in regional assemblies. This approach aligns with conventions in composite index construction, where equal weighting is a common practice to preserve transparency and comparability in the absence of definitive empirical guidance on relative institutional influence (e.g., GEM uses equal weighting across its three components: parliamentary representation, economic decision-making roles, and income share) (Klasen 2006; Betata 2007; Seipel 2010; and Rios et al. 2024). While ministerial roles may wield greater executive power and could merit greater weight, current empirical data do not consistently support a specific differential weighting. Therefore, equal weights provide a theoretically neutral and methodologically transparent baseline. Moreover, all these data was retrieved from European Institute for Gender Equality (EIGE) (2025). The following section presents the empirical results, examining how

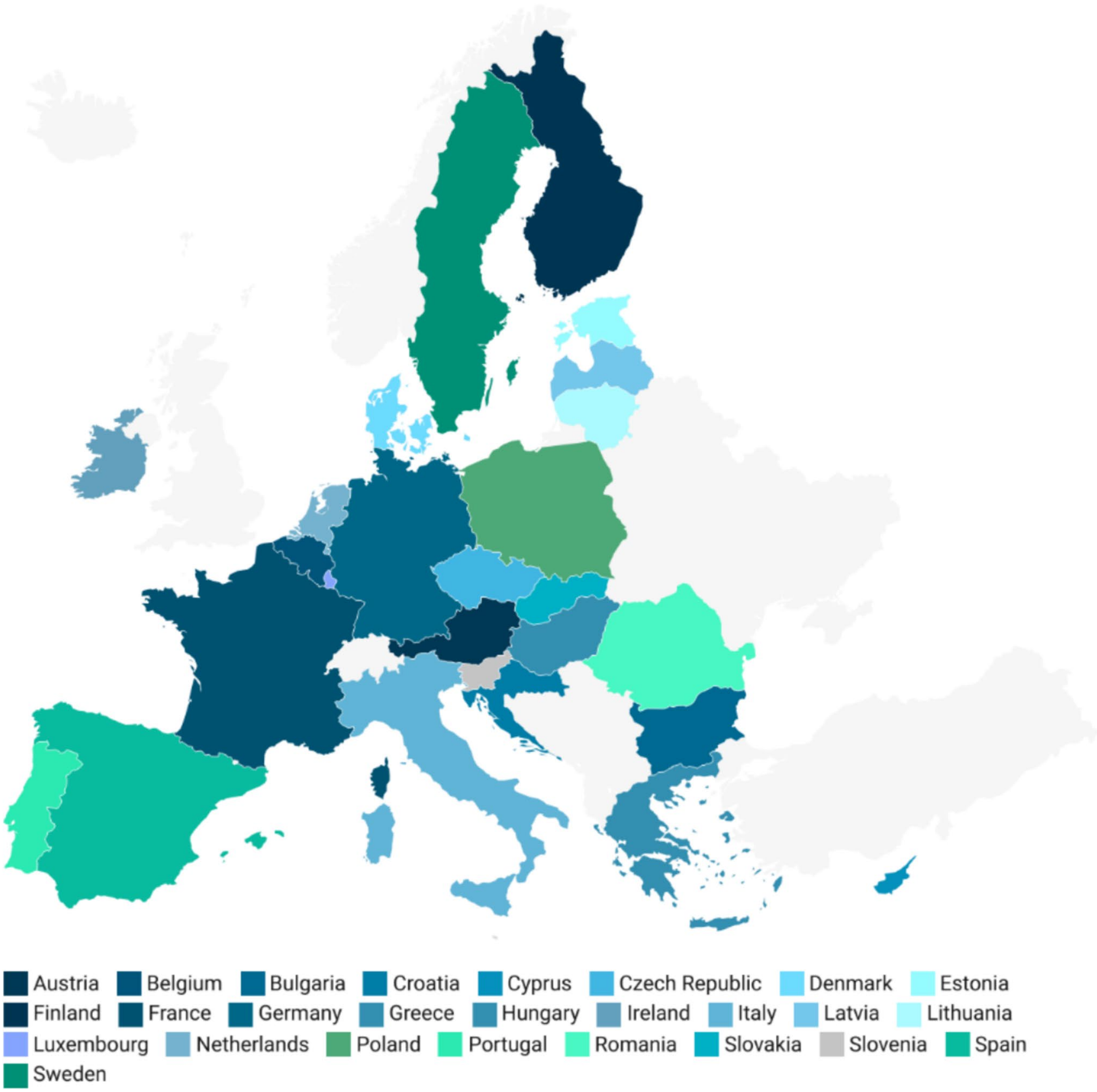

Fig. 3 Graphical representation of the 26 European Union countries. The analysis and visualization were conducted in RStudio (version 2025.05.1). The dataset was structured using the packages dplyr and tidyr, and the world map was generated with rnaturalearth, sf, and ggplot2

the female political participation impact the environmental degradation in the EU.

Empirical Results

This subsection presents the results of preliminary tests, panel data diagnostics, model estimation diagnostics, estimation models, post-estimation validation, and robustness check. Estimations were performed using **Stata 17.0** and the **RStudio package (version 2025.05.1)**, with the commands employed listed in the accompanying notes table. Figure 4 shows the methodological framework that this investigation follows.

Preliminary Tests

Before presenting the preliminary tests, it is important to recall that the FFP_Index was constructed as the average of women's representation in ministerial posts, parliaments, and regional assemblies, each assigned equal weight. Equal

Table 1 Variable Descriptions

Variable	Description	Source	Unit	Period	Type
CO_2e/cap	CO_2 emissions per capita (excluding LULUCF)	World Bank Open Data (2025)	Metric tons CO_2e per capita	2013–2023	Dependent variable
GHGe/cap	Total greenhouse gas emissions per capita (excluding LULUCF)	World Bank Open Data (2025)	Metric tons CO_2e per capita	2013–2023	Dependent variable
HFFEC	Household fossil fuel energy consumption	Eurostat Database (2025)	Percentage (%)	2013–2023	Control variable
HCEC	Household clean energy consumption	Eurostat Database (2025)	Percentage (%)	2013–2023	Control variable
URB-POP	Urban population (% of total population)	World Bank Open Data (2025)	Percentage (%)	2013–2023	Control variable
GDP/cap	Gross Domestic Product per capita (constant local currency)	World Bank Open Data (2025)	Euros (€)	2013–2023	Control variable
GovEff	Government effectiveness (perception-based index)	World Bank Open Data (2025)	Index (estimate)	2013–2023	Control variable
FFP_Index	Female political participation (measured as the share of women in ministerial posts, national parliaments, and regional assemblies)	European Institute for Gender Equality (EIGE)	Percentage (%)	2013–2023	Independent variable (core focus)

weighting is a common practice in composite indices where empirical guidance on differential influence is limited, ensuring transparency and comparability across countries (Rios et al. 2024). Preliminary tests, such as those outlined by Fuinhas et al. 2025, Koengkan et al. (2025, 2024) and Wooldridge (2010), were conducted to assess the characteristics of the variables used in this study. Additionally, descriptive statistics for the variables were calculated, as presented in Table 2 below.

Table 2 reports summary statistics for 260 country-year observations. CO_2e/cap averages 6.46 t (SD = 2.64; range = 3.02–18.72 t), and GHGe per capita (GHGe/cap) averages 9.05 t (SD = 3.13; range = 4.78–20.74 t), indicating notable variation in emissions. HFFEC has mean = 3 948.58 (SD = 5 159.01; range = 111.33–27 063.01), while HCEC) is higher on average (mean = 5 682.21; SD = 8 747.30; range = 180.85–38 490.41), reflecting diverse energy mixes across countries. URB-POP averages 72.48% (SD = 12.44; 53.33–98.15%), and GDP/cap shows extreme spread (mean = 224 369.80; SD = 746 435.80; 10 524.87–4 460 942).GovEff clusters around 1.04 (SD = 0.56; – 0.29–2.18), and the variable (FFP_Index) averages 27.24 (SD = 10.32; 0–50.33), highlighting variation in governance quality and women's political representation. A pairwise correlation analysis follows (Jolliffe 2002).

The pairwise correlation analysis (see Table 1A and Fig. 1A in the Appendix) reveals strong associations between CO_2 and total greenhouse gas emissions per capita ($\rho = 0.9338$, $p < 0.01$), as well as between Household Fossil Fuel Energy Consumption (HFFEC) and Household Clean Energy Consumption (HCEC) ($\rho = 0.8676$, $p < 0.01$), suggesting potential collinearity. In contrast, the Female Political Participation Index (FFP_Index) exhibits negligible correlations with both CO_2 ($\rho = 0.0197$) and total GHG emissions (GHGe) ($\rho = -0.0185$), underscoring the need for multivariate analysis to uncover its net effects.

Urban population percentage (URB-POP) shows a positive relationship with emissions ($\rho = 0.3182$ for CO_2 and $\rho = 0.2230$ for GHGe, $p < 0.01$), whereas GDP per capita (GDP/cap) is weakly negatively correlated with emissions ($\rho = -0.1187$ for CO_2, $p < 0.10$; $\rho = -0.1446$ for GHGe, $p < 0.05$), which may reflect emissions decoupling in wealthier economies. Government Effectiveness (GovEff) is positively correlated with emissions ($\rho = 0.3772$ with CO_2; $\rho = 0.3564$ with GHGe, $p < 0.01$), likely indicating higher emissions levels in more developed countries. Strong intercorrelations among URB-POP, GDP/cap, GovEff, and FFP_Index (e.g., GovEff–URB-POP $\rho = 0.5349$, GDP–FFP_Index $\rho = -0.3029$, FFP_Index–GovEff $\rho = 0.5813$; all $p < 0.01$) point to possible multicollinearity, warranting further diagnostic testing.

After conducting the correlation test, multicollinearity was evaluated using the Belsley and Welsch (1980) condition-number approach. Table 2A in the Appendix reports the results of the multicollinearity tests for Panel A and Panel B. The results indicate that multicollinearity is low in both models (mean VIF = 2.61, with individual VIFs below 5).

Normality was assessed using Skewness and Kurtosis tests (D'Agostino and Pearson 1973), with results reported in Table 3A in the Appendix. For seven variables ($n = 260$), the null hypothesis of normality is rejected at the 1% significance level for CO_2e/cap, GHGe/cap, HFFEC, HCEC, GDP/cap, GovEff, and FFP_Index, indicating these series are not normally distributed. Specifically, CO_2e/cap and GHGe/cap display significant departures in both skewness and kurtosis ($p < 0.001$), while HFFEC and HCEC reject normality

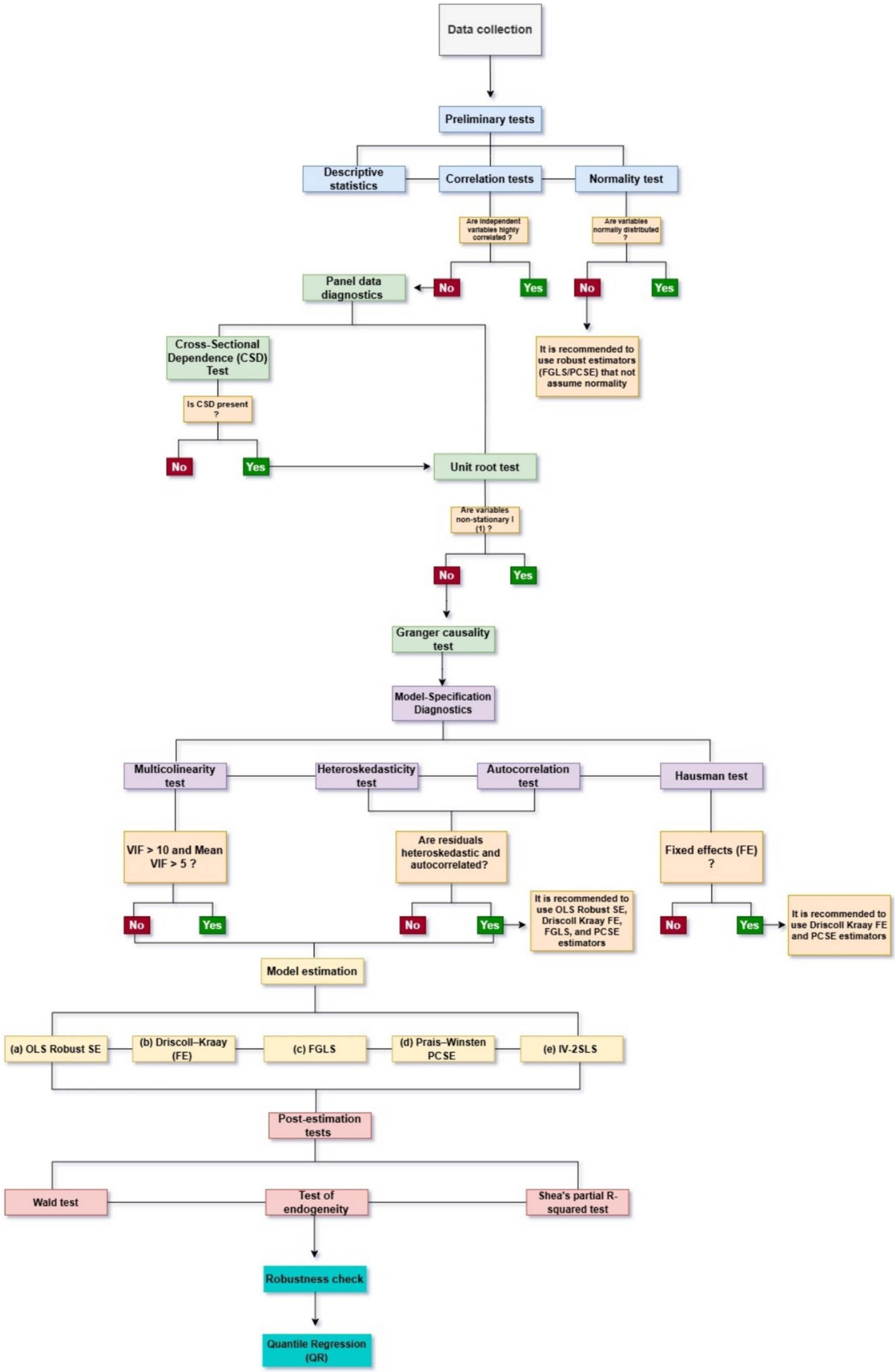

Fig. 4 Methodological framework. This figure was created by the authors

Table 2 Descriptive statistics

Variables	Obs	Mean	Std. Dev	Min	Max
CO_2e/cap	260	6.459788	2.643867	3.018711	18.72295
GHGe/cap	260	9.051421	3.129094	4.780231	20.74058
HFFEC	260	3948.576	5159.009	111.329	27,063.01
HCEC	260	5682.205	8747.304	180.854	38,490.41
URB-POP	260	72.48153	12.43894	53.332	98.153
GDP/cap	260	224,369.8	746,435.8	10,524.87	4,460,942
GovEff	260	1.041373	0.562234	− 0.28694	2.180622
FFP_Index	260	27.24192	10.3156	0	50.33333

Notes: The Stata 17 command *summarize* was used

Table 3 Pesaran (2004) CD test

Variables	CD-test	p-value	Corr	abs(corr)
CO_2e/cap	24.91	0.0000***	0.437	0.619
GHGe/cap	17.67	0.0000***	0.31	0.563
HFFEC	19.24	0.0000***	0.337	0.422
HCEC	32.66	0.0000***	0.573	0.601
URB-POP	36.34	0.0000***	0.637	0.914
GDP/cap	47.06	0.0000***	0.825	0.825
GovEff	4.21	0.0000***	0.074	0.449
FFP_Index	21.12	0.0000***	0.37	0.541

Notes: The Stata 17 command *xtcd* was used; Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$

overall, though individual skewness/kurtosis p-values are not separately reported.

URB-POP does not reject skewness ($p=0.1464$) but shows significant kurtosis ($p<0.001$). GovEff (skewness $p=0.0689$; kurtosis $p=0.0009$) and FFP_Index (skewness $p=0.0671$; kurtosis $p=0.0037$) exhibit non-normality mainly due to excess kurtosis. These findings justify the use of estimation methods robust to non-normal error distributions, such as fixed-effects models with Driscoll–Kraay standard errors or quantile regressions.

The observed violations of normality assumptions across key variables reinforce the appropriateness of applying robust estimation techniques in subsequent analyses, ensuring reliable inference despite the skewed distribution of environmental and socioeconomic indicators. Following these preliminary analyses, panel-data diagnostic tests are essential to verify model assumptions and uphold the validity of the study's results.

Panel Data Diagnostics Tests

Having verified that the variables deviate from normality, this investigation proceeds to test for cross-sectional dependence using Pesaran (2004) CD statistic, which is especially appropriate when the panel comprises more cross-sectional units (N) than time periods (T). This assessment determines whether error terms are correlated across countries, a condition that, if present, could bias standard estimators and render them inefficient. Under the null hypothesis of no cross-sectional dependence, this investigation examines the CD test results presented in Table 3 below.

Table 3 above presents Pesaran (2004) cross-sectional dependence (CD) test for each variable ($n=260$). The null hypothesis of cross-sectional independence is rejected at the 1% level for all eight variables ($p<0.01$), indicating pervasive interdependence among EU country-year observations. Dependence is weakest for the variables (GovEff, corr = 0.074; abs(corr) = 0.449) and household fossil-fuel consumption (HFFEC, corr = 0.337; abs(corr) = 0.422), moderate for CO_2e per capita (corr = 0.437; abs(corr) = 0.619) and total GHG per capita (GHGe/cap, corr = 0.310; abs(corr) = 0.563), and strongest for GDP per capita (corr = 0.825; abs(corr) = 0.825) and urban population share (URB-POP, corr = 0.637; abs(corr) = 0.914). Household clean-energy consumption (HCEC, corr = 0.573; abs(corr) = 0.601) and female political participation (FFP_Index, corr = 0.370; abs(corr) = 0.541) also exhibit substantial cross-sectional correlation. These findings underscore the need to employ panel estimators—such as OLS robust, Driscoll–Kraay fixed effects, FGLS, or PCSE—that account for cross-sectional dependence in subsequent analyses. Moreover, the detection of cross-sectional dependence has important implications for the EU context. It suggests that environmental and governance outcomes in one member state may not be isolated but instead influence, and be influenced by, developments in other countries through mechanisms such as policy spillovers, economic integration, and coordinated EU directives. For instance, ambitious climate or gender equality reforms adopted in one country may diffuse across borders, shaping regional norms and collective strategies. This interconnectedness aligns with our theoretical emphasis on inclusive governance, underscoring that the benefits of greater female political participation may extend beyond national boundaries to foster EU-wide collaboration in environmental policymaking. Given the presence of cross-sectional dependence among the variables, the next step is to assess their stationarity properties using the Levin-Lin-Chu (LLC) unit root test (Levin et al. 2002), which is appropriate for balanced panels and accounts for common time effects across units (see Table 4 below).

Table 4 above presents the results of Levin–Lin–Chu (LLC) panel unit-root tests for eight variables at lags 0 and 1, with and without a time trend. The null hypothesis of non-stationarity is rejected at the 1% level for most variables, indicating stationarity in levels for CO_2e/cap, GHGe/cap, HFFEC, HCEC, GDP/cap, GovEff, and FFP_Index. CO_2e/cap and GHGe/cap show partial non-stationarity at lag 1 without a trend, while HFFEC, HCEC, GDP/cap, GovEff, and FFP_Index are consistently stationary. The variable URB-POP is stationary at lag 0 but non-stationary at lag 1, suggesting a unit root. Moreover, While the LLC tests

Table 4 Levin-Lin-Chu unit-root test (LLC)

Variable	Lags	Adjustment t (without trend)	p-value	Adjustment t (with trend)	p-value
CO_2e/cap	0	– 2.9776	0.0015***	– 8.6194	0.0000***
CO_2e/cap	1	– 1.0089	0.1565	– 9.1157	0.0000***
GHGe/cap	0	– 2.3366	0.0097***	– 5.9568	0.0000***
GHGe/cap	1	– 1.3590	0.0871*	– 4.2764	0.0000***
HFFEC	0	– 10.4876	0.0000***	– 13.4078	0.0000***
HFFEC	1	– 3.3440	0.0000***	– 7.6344	0.0000***
HCEC	0	– 5.7976	0.0000***	– 15.6569	0.0000***
HCEC	1	– 3.5971	0.0000***	– 3.9008	0.0000***
URB-POP	0	– 15.5581	0.0000***	– 5.4387	0.0000***
URB-POP	1	1.4924	0.9322	0.1462	0.5581
GDP/cap	0	– 2.5722	0.0051***	– 11.7068	0.0000***
GDP/cap	1	– 1.6732	0.0471**	– 7.3292	0.0000***
GovEff	0	– 3.4610	0.0003***	– 8.3898	0.0000***
GovEff	1	– 4.0074	0.0000***	– 9.7157	0.0000***
FFP_Index	0	– 3.0806	0.0010***	– 10.0883	0.0000***
FFP_Index	1	– 6.4612	0.0000***	– 15.3357	0.0000***

Notes: The Stata 17 command *xtunitroot llc* and *xtunitroot llc,* with lag lengths 0 and 1 were used; Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$

confirm that most variables are stationary in levels, this study note that the URB-POP variable exhibits partial non-stationarity at lag 1, despite being stationary at lag 0. This pattern is not uncommon in demographic variables, which often evolve slowly over time. Although, the main estimations remain robust, this suggests that differencing URB-POP could be explored in future robustness checks to further validate causal inferences. Given the consistency of results across multiple estimators and the confirmation of stationarity at lag 0, this study consider this issue non-critical for the present analysis but acknowledge its relevance for extended applications. With stationarity confirmed, the analysis proceeds to Granger causality testing using the Dumitrescu and Hurlin (2012) approach to examine how female political participation influences environmental degradation across EU member states.

Two models are estimated: **Panel A** with CO_2e per capita and **Panel B** with GHGe per capita as dependent variables. Two models are estimated to ensure a comprehensive analysis of environmental degradation. While CO_2 emissions are a major contributor to climate change, they do not capture the full scope of greenhouse gases, which also include methane (CH_4), nitrous oxide (N_2O), and fluorinated gases. Estimating both models allows the study to test the robustness and consistency of results across distinct yet related measures of environmental impact. This dual approach helps assess whether the influence of female political participation and other explanatory variables is specific to CO_2 emissions or generalises to broader GHGe dynamics, thus strengthening the validity and policy relevance of the findings. Moreover, in conducting the Granger causality analysis, this study employed a fixed lag structure (lag 0) given the relatively short time dimension of the panel ($T=11$). While information-criterion–based lag selection (e.g., AIC or BIC) is often recommended in longer panels, applying such methods here would risk overfitting, reduce degrees of freedom, and weaken inference, as exposed by Bun and Windmeijer (2010) and Andrews and Lu (2001). This investigation choice therefore reflects the constraints of the dataset while ensuring reliable estimation. This investigation acknowledge this as a methodological limitation, but note that the consistency of results across both **Panel A** (CO_2e/cap) and **Panel B** (GHGe/cap) supports the robustness of the identified causal relationships. Therefore, Table 5 below shows the results from Granger causality test.

Table 5 presents the Granger-causality results. In **Panel A**, HFFEC Granger-causes CO_2e/cap, but not vice versa, while HCEC and URB-POP exhibit bidirectional causality with CO_2e/cap. GDP/cap and FFP_Index have a unidirectional influence on CO_2e/cap, whereas GovEff shows two-way causality. In **Panel B**, GHGe/cap is weakly predicted by HFFEC but strongly predicts it in return. HCEC, URB-POP, and GovEff all display mutual causality with GHGe/cap, while GDP/cap and FFP_Index again show unidirectional effects. These findings indicate complex interdependencies between socioeconomic variables and emissions. Figure 5 below summarizes the results from Granger causality test.

Following these findings, the next section presents model-specification diagnostics based on two panel-data tests.

Model Specification Diagnostics

Before estimating the models, this conduct specification diagnostics. Heteroskedasticity is tested via the Breusch-Pagan (1979) and Cook-Weisberg (1983) tests, both under the null hypothesis of constant variance. Serial correlation is examined using Wooldridge's (2010) test, where the null assumes no autocorrelation. Addressing, heteroskedasticity, and autocorrelation is essential for obtaining reliable estimators. Table 4A in the Appendix, reports the results of these diagnostics for Panel A and Panel B. Therefore, the results shows that each model suffers from significant heteroskedasticity (Panel A: $\chi^2(1)=148.06$, $p<0.01$; Panel B: $\chi^2(1)=85.81$, $p<0.01$) and autocorrelation (Panel A: $F(1, 25)=93.216$, $p<0.01$; Panel B: $F(1, 25)=100.829$, $p<0.01$). Consequently, both the CO_2e/cap and GHGe/

Table 5 Granger Causality test

Null hypothesis	Lags	Dumitrescu and Hurlin (2012)		
		Z-bar	p-value	Directions
Panel A				
HFFEC → CO_2e/cap	0	5.6164	0.0000***	✅ Causes
CO_2e/cap → HFFEC	0	1.0005	0.3171	❌ Does not cause
HCEC → CO_2e/cap	0	5.1889	0.0000***	✅ Causes
CO_2e/cap → HCEC	0	3.6193	0.0003***	✅ Causes
URB-POP → CO_2e/cap	0	7.5357	0.0000***	✅ Causes
CO_2e/cap → URB-POP	0	4.8667	0.0000***	✅ Causes
GDP/cap → CO_2e/cap	0	9.2845	0.0000***	✅ Causes
CO_2e/cap → GDP/cap	0	0.5048	0.6137	❌ Does not cause
GovEff→ CO_2e/cap	0	2.1012	0.0356**	✅ Causes
CO_2e/cap → GovEff	0	9.2581	0.0000***	✅ Causes
FFP_Index → CO_2e/cap	0	6.6504	0.0000***	✅ Causes
CO_2e/cap → FFP_Index	0	1.9119	0.0559*	✅ Causes
Panel B				
HFFEC → GHGe/cap	0	1.8891	0.0589*	✅ Causes
GHGe/cap → HFFEC	0	4.3718	0.0000***	✅ Causes
HCEC → GHGe/cap	0	2.3947	0.0166**	✅ Causes
GHGe/cap → HCEC	0	2.7483	0.0060***	✅ Causes
URB-POP → GHGe/cap	0	9.2507	0.0000***	✅ Causes
GHGe/cap → URB-POP	0	6.4178	0.0000***	✅ Causes
GDP/cap → GHGe/cap	0	10.4578	0.0000***	✅ Causes
GHGe/cap → GDP/cap	0	0.0060	0.9952	❌ Does not cause
GovEff→ GHGe/cap	0	8.3271	0.0000***	✅ Causes
GHGe/cap → GovEff	0	8.4095	0.0000***	✅ Causes
FFP_Index → GHGe/cap	0	7.4576	0.0000***	✅ Causes
GHGe/cap → FFP_Index	0	1.0851	0.2779	❌ Does not cause

Notes: The Stata 17 commands *xtgcause* was used; The null hypothesis is that X does not Granger-cause Y. All models used lag length 0 unless otherwise stated. Significance levels: *** p<0.01, ** p<0.05, * p<0.1.

cap regressions employ robust estimation methods, namely OLS with robust SE, Driscoll–Kraay FE, FGLS, Prais–Winsten PCSE, and IV-2SLS—to correct for heteroskedasticity and serial correlation.

Following the heteroskedasticity/autocorrelation tests, it is necessary to perform the Hausman test to determine the appropriate panel data specification—fixed effects (FE) or random effects (RE)—for Panels A and B. Table 5A in the Appendix presents the results of the Hausman test. These results were used to determine the appropriate panel data specification—fixed effects (FE) or random effects (RE)—for both models. In Panel A (CO_2e/cap as the dependent variable), the test yields a $chi^2(4)=78.70$ ($p<0.01$), and in Panel B (GHGe/cap as the dependent variable), $chi^2(5)=63.42$ ($p<0.01$). In both cases, the null hypothesis of no systematic difference between FE and RE is rejected at the 1% level, confirming that the fixed effects estimator is preferred. This result justifies the use of FE-based robust methods such as Driscoll–Kraay and Prais–Winsten PCSE in the main analysis to account for unobserved heterogeneity across EU member states.

Model Estimation

To address heteroskedasticity, cross-sectional dependence (CSD), autocorrelation, and potential endogeneity in a large-N, small-T panel data context, this study employs five estimation techniques: (a) Ordinary Least Squares (OLS) with robust standard errors, (b) Driscoll–Kraay standard errors with fixed effects, (c) Feasible Generalized Least Squares (FGLS), (d) Prais–Winsten regression with Panel-Corrected Standard Errors (PCSE), and (e) Instrumental Variables Two-Stage Least Squares (IV-2SLS). These methods were chosen to ensure reliable inference under common violations of classical OLS assumptions.

a) OLS with robust standard errors

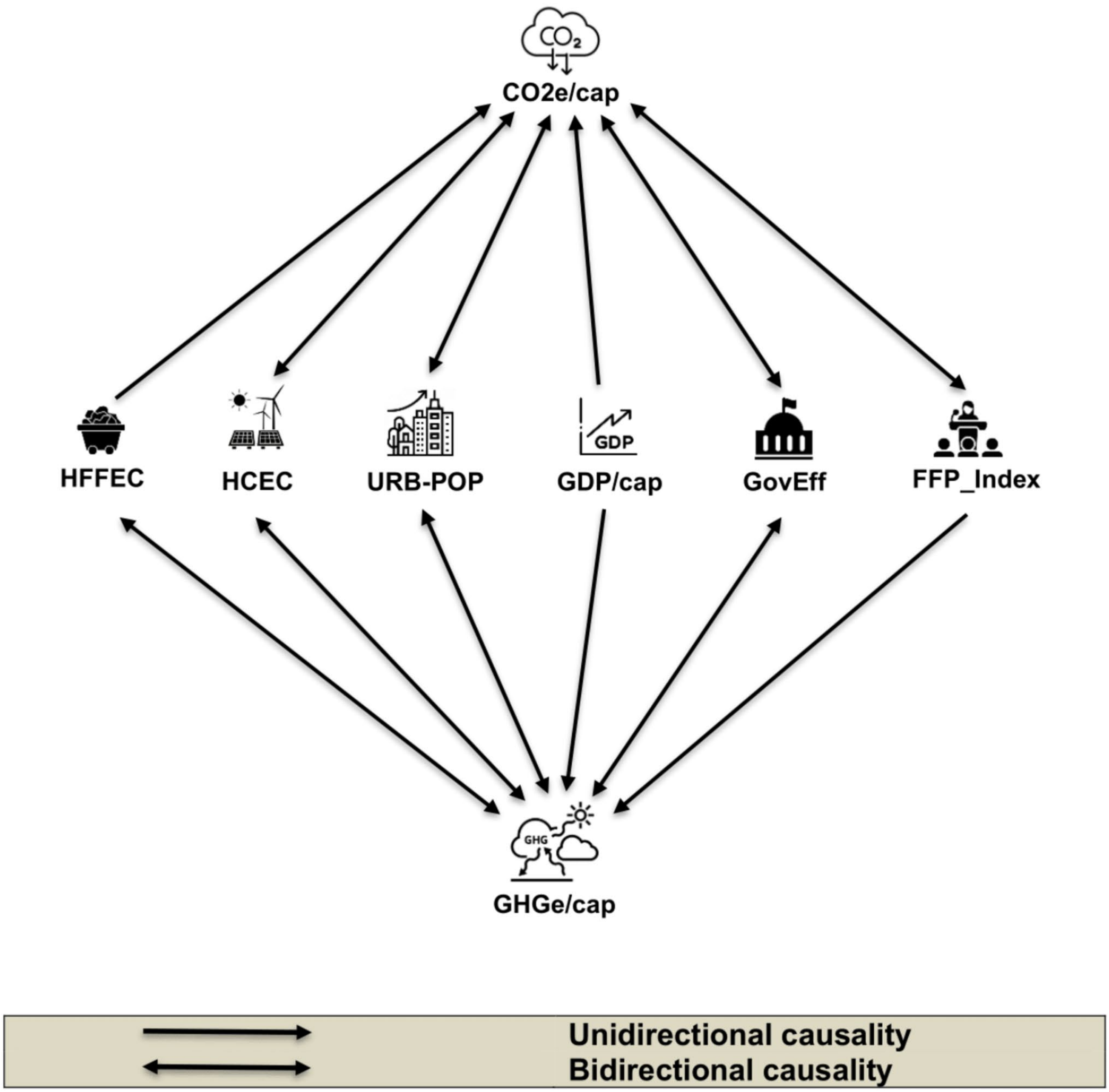

Fig. 5 Summary of Unidirectional and Bidirectional Causalities Based on Granger Causality Test. This figure was created by the authors

Ordinary Least Squares (OLS) with robust standard errors offer a simple yet effective approach to correcting heteroskedasticity, thereby enhancing the reliability of coefficient estimates (Hayes and Cai 2007). The OLS model is typically expressed as:

$$Z_t = W_t\beta + \varepsilon_t \tag{2}$$

where, Z_t is the dependent variable vector for observation$t$, W_t is the independent variable matrix for observation t, β is the vector of coefficients, and ε_t is the error term. The OLS minimizes the sum of squared residuals (ε_t) to estimateβ. Indeed, when using robust standard errors, the heteroskedasticity-consistent covariance matrix is estimated as:

$$\sum\nolimits^{\wedge}_{robust} = (W'W)^1 \, \hat{u}'\hat{u} \, (W'W)^{-1} \tag{3}$$

where $\hat{u}$ are the residuals from the OLS regression. This correction adjusts for heteroskedasticity (non-constant variance of the error term).

b) Driscoll–Kraay standard errors with fixed effects

The fixed effects model with Driscoll–Kraay standard errors is suitable for panel data with unobserved individual-specific effects that remain constant over time. This method controls for unobserved heterogeneity while adjusting for heteroskedasticity and autocorrelation. Driscoll–Kraay standard errors also account for both temporal and cross-sectional dependence, making them ideal for panel data with these features (Hoechle 2007). The fixed effects model is represented as:

$$Z_{it} = a_i + W_{it}\beta + \varepsilon_{it} \tag{4}$$

where, Z_{it} is the dependent variable for individual i at the time t, a_i is the individual-specific intercept (capturing unobserved heterogeneity), W_{it} is the independent variable matrix for individual i at time$t$, β is the vector of coefficients, and ε_{it} is the error term. The Driscoll-Kraay standard errors correct for both heteroskedasticity and autocorrelation in the panel data model. The covariance matrix of the coefficient estimates is adjusted as:

$$\sum\nolimits_{DK}^{\wedge} = (W'W)^{-1}\left(\sum_{t=1}^{T} \hat{u}_t \hat{u}_t\right)(W'W)^{-1} \tag{5}$$

where $\hat{u}_t$ is the residual for time t and the adjustment accounts for cross-sectional and temporal dependence in the residuals.

c) Feasible Generalized Least Squares (FGLS)

Feasible Generalized Least Squares (FGLS) effectively addresses cross-sectional dependence by correcting for heteroskedasticity and autocorrelation. This method improves the precision of estimates, especially when there is substantial correlation across cross-sectional units (Bai et al. 2021). In the FGLS model, the error term is generally assumed to have the following structure:

$$Z_t = \rho\epsilon_{t-1} + u_t \tag{6}$$

where, Z_t is the error term for time, u_t is the white noise error term, and ρ is the autoregressive parameter. The FGLS estimator modifies the standard OLS approach by estimating a covariance matrix that accounts for both heteroskedasticity and serial correlation. The generalized least squares (GLS) estimator is given by:

$$Z_{GLS} = \left(W\prime\Omega^{-1}W\right)^{-1} W\prime\Omega^{-1}Y \tag{7}$$

where Ω is the error covariance matrix that captures heteroskedasticity and serial correlation. FGLS involves estimating Ω and then using it in the GLS formula.

d) Prais–Winsten regression with Panel-Corrected Standard Errors (PCSE)

Panel-Corrected Standard Errors (PCSE) effectively tackle heteroskedasticity and contemporaneous correlation in panel data, making them well-suited for datasets with both time-series and cross-sectional dimensions. By adjusting the covariance matrix to account for cross-sectional dependence, PCSE provides reliable estimates even when residuals are correlated across panels (Bailey and Katz 2011). The PCSE model follows the same specification as the Pooled OLS model:

$$Z_{it} = a_i + W_{it}\beta + \varepsilon_{it} \tag{8}$$

The key difference is in the adjustment to the covariance matrix, which is corrected for both heteroskedasticity and contemporaneous correlation across panels:

$$\sum\nolimits_{PCSE}^{\wedge} = (W\prime W)^{-1}\hat{\Omega}(W\prime W)^{-1} \tag{9}$$

where $\widehat{\Omega}$ is the panel-corrected covariance matrix that accounts for cross-sectional dependence (contemporaneous correlation) across panels.

e) Instrumental Variables Two-Stage Least Squares (IV-2SLS)

Two-Stage Least Squares (2SLS) is a widely used econometric method for addressing endogeneity in regression models, especially when explanatory variables are correlated with the error term. It is applicable to both cross-sectional and panel data and is crucial for handling issues such as simultaneity, measurement error, or omitted variable bias. The 2SLS estimator operates in two stages to produce consistent parameter estimates (Wooldridge 2010). In the first stage, the potentially endogenous regressors are regressed on all exogenous variables, including instrumental variables, to obtain predicted values:

$$Z_{it} = W_{it\pi} + \mu_{it} \tag{10}$$

where Z_{it} is the endogenous variable, W_{it} includes all exogeneous regressions and instruments, and π represents the first-stage coefficients.

In the second stage, the original equation is estimated using the fitted values from the first stage in place of the endogenous regressors:

$$Z_{it} = \alpha + W_{it}\beta + \varepsilon_{it} \tag{11}$$

This approach ensures consistent estimation of β by removing the correlation between the regressors and the error term. The validity of 2SLS hinges on the relevance and exogeneity of the instruments, which must be strongly correlated with the endogenous variables but uncorrelated with the error term. Moreover, in the IV-2SLS model, endogeneity of the independent variable FFP_Index is addressed using five instruments: HFFEC, HCEC, URB-POP, GDP/cap, and GovEff. These variables are strongly correlated with FFP_Index but not directly with emissions. Instrument validity can be supported by Shea's Partial R^2 values and robust diagnostic tests.

OLS with robust standard errors, Driscoll–Kraay standard errors with fixed effects (FE), FGLS, PCSE, and IV-2SLS address key econometric issues in panel data. OLS with robust errors corrects heteroskedasticity but does not account for autocorrelation or cross-sectional dependence. Driscoll–Kraay standard errors provide robust corrections for heteroskedasticity, autocorrelation, and importantly, cross-sectional dependence, making them especially useful when residuals exhibit both temporal and spatial correlation. FGLS improves efficiency by handling heteroskedasticity, autocorrelation, and dependence but requires correct error specification and may underestimate standard errors when the time dimension (T) is small. PCSE adjusts for heteroskedasticity and contemporaneous correlation, making it suitable for large N and small T panels, though it can be less effective with strong cross-sectional correlation. IV-2SLS addresses endogeneity using instrumental variables but depends on the strength and validity of the instruments. Given the panel structure (N = 26, T = 11) and diagnostic evidence of autocorrelation and cross-sectional dependence, Driscoll–Kraay with fixed effects was selected as the primary model for its comprehensive and robust handling of these issues. Table 6 below presents the estimation results for **Panel A**, where the dependent variable is CO_2e/cap.

Table 6 above presents the estimation results for the dependent variable (CO_2e/cap), using five estimation methods. The Driscoll–Kraay fixed effects model is chosen as the primary estimator due to its robustness in addressing heteroskedasticity, autocorrelation, and cross-sectional dependence. The coefficient for the variable (HFFEC) is positive and highly significant across all models; under the Driscoll–Kraay FE model, it is 0.0002 ($t = 9.31$, $p < 0.01$), indicating that greater fossil fuel use significantly increases CO_2 emissions per capita. Conversely, the variable (HCEC) consistently shows a negative and significant effect, with a coefficient of -0.0001 ($t = -8.06$, $p < 0.01$) in the primary model, demonstrating its role in reducing emissions. The variable (URB-POP) has a positive and significant association (0.0934, $t = 4.00$, $p < 0.01$), reflecting increased environmental pressures from urbanization. The variable (GDP/cap) has a small but significant negative coefficient (− 8.38e-07, $t = -3.51$, $p < 0.01$), suggesting that economic growth may lead to emission reductions through improved efficiency or cleaner technologies. Interestingly, the variable (GovEff) is positively related to emissions (2.0979, $t = 4.86$, $p < 0.01$), which may reflect higher economic activity in countries with stronger governance. Finally, the independent variable (FFP_Index) is negatively associated with emissions (− 0.1379, $t = -3.19$, $p < 0.01$), highlighting the potential of stronger female political involvement in driving environmental improvements. All variables maintain strong statistical significance and show consistent patterns across OLS robust, FGLS, Prais–Winsten PCSE, and IV-2SLS models. Post-estimation tests confirm the joint significance of coefficients and the validity and strength of instruments in the 2SLS approach, reinforcing the robustness of these findings. Overall, the Driscoll–Kraay FE model is preferred in this analysis for its comprehensive handling of key panel data issues. Figure 6 below shows the visualization of the coefficients in **Panel A**.

After completing the estimation for **Panel A**, it is necessary to analyse the results of **Panel B**, where the dependent variable is (GHGe/cap). Table 7 below presents these results.

Table 7 above presents the estimation results for the dependent variable (GHGe/cap), across five estimation methods. The Driscoll–Kraay fixed effects model is selected as the primary estimator due to its robustness in addressing heteroskedasticity, autocorrelation, and cross-sectional dependence. The coefficient for the variable (HFFEC) is positive and highly significant across all models; under the Driscoll–Kraay FE model, it is 2.28e-10 ($t = 10.84$, $p < 0.01$), indicating that increased fossil fuel consumption significantly raises per capita greenhouse gas emissions. Conversely, the variable (HCEC) shows a consistently negative and significant effect, with a coefficient of -1.5e-10 ($t = -10.76$, $p < 0.01$) in the primary model, demonstrating its effectiveness in reducing emissions. The variable (URB-POP) has a positive and significant coefficient of 8.33e-08 ($t = 3.18$, $p < 0.01$), suggesting that higher urbanization is associated with greater environmental pressure and emissions. The variable (GDP/cap) shows a small but statistically significant negative coefficient of -1.1e-12 ($t = -4.02$, $p < 0.01$), indicating that economic growth may contribute to emissions reductions through cleaner technologies or efficiency gains. The variable (GovEff) is positively related to

Table 6 Estimation results

Panel A															
Dependent variable (CO_2e/cap)															
Models															
Variables	(a) OLS robust			(b) Driscoll–Kraay (FE)			(c) FGLS			(d) Prais–Winsten PCSE			(e) IV-2SLS		
HFFEC	0.0002	5.65	***	0.0002	9.31	***	0.0002	6.39	***	0.00012	3.27	***	0.0002	5.06	***
HCEC	-0.0001	-5.16	***	-0.0001	-8.06	***	-0.0001	-7.44	***	-0.0001	-2.24	**	-0.0001	-4.22	***
URB-POP	0.0934	4.81	***	0.0934	4.00	***	0.0961	6.53	***	0.0456	5.28	***	0 .1029	5.53	***
GDP/cap	-8.38e-07	-5.26	***	-8.38e-07	-3.51	***	-1.40e-06	-2.42	**	-3.82e-07	-2.59	**	-9.54e-07	-5.17	***
GovEff	2.0979	6.63	***	2.0979	4.86	***	1.9238	11.27	***	1.5386	4.94	***	2.3386	6.76	***
FFP_Index	-0.1379	-4.13	***	-0.1379	-3.19	***	-0.1272	-18.37	***	-0.0321	-2.60	***	-0.1822	-4.70	***
Constant	1.2664	1.50		1.2664	1.44	***	0.8744	1.04		2.5096	4.27	***	1.4675	1.74	*
Observations	260			260			260			260			234		
Post-estimation tests															
Wald Test for Joint Significance of Coefficients													Tests of endogeneity		
F(6, 253) = 17.93***				F(6, 25) = 170.68***			chi2(6) = 1050.39***			chi2(6) = 144.15***			Robust Score Chi² (6)	7.8008 (0.2531)	
													Robust Regression F(6,221)	1.53121 (0.1690)	
Shea's partial R-squared test															
Variable						Shea's Partial R²							Strength		
HFFEC						0.9345							Strong		
HCEC						0.9405							Strong		
URB-POP						0.9097							Strong		
GDP/cap						0.9260							Strong		
GovEff						0.8767							Strong		
FFP_Index						0.6057							Strong		

Notes: The Stata 17 commands, r*egress*, *xtreg, fe, xtgls, xtpcse*, and *ivreg2* were used for estimations in columns (a)–(e), respectively. Post-estimation tests employed test, *ivendog, estat endogenous*, and *estat firststage*. Instrument strength was assessed using *ivreg2* with the *first option* for Shea's Partial R². Significance levels: *** p<0.01, ** p<0.05, * p<0.1.

emissions, with a coefficient of 2.65e-06 (t = 4.95, p < 0.01), which may reflect increased economic activity and energy use in better-governed countries. Lastly, the independent variable (FFP_Index) is negatively associated with emissions (-1.6e-07, t = -3.19, p < 0.01), highlighting the potential influence of stronger female political involvement on promoting environmental sustainability. All variables show strong statistical significance and consistent signs across OLS robust, FGLS, Prais–Winsten PCSE, and IV-2SLS models. Post-estimation tests confirm the joint significance of coefficients and the strength and validity of instruments used in the 2SLS estimation, reinforcing the robustness of the results. Overall, the Driscoll–Kraay FE model remains the preferred method for its comprehensive treatment of key panel data challenges. The results in Panel B are similar to those in Panel A, which is expected since CO_2 emissions are a major component of total greenhouse gas emissions (GHGe). Figure 7 below shows the visualization of the coefficients in **Panel B**.

Indeed, Fig. 8 below summarises the results of **estimator (b)** found in Tables 6 and 7 above.

Both models consistently show that increased fossil fuel consumption raises emissions, while clean energy use and female political participation contribute to their reduction, reinforcing the robustness of these findings across related environmental indicators.

Machine Learning Analysis

In addition to the econometric estimations, this study conducted a complementary analysis using Random Forest regressions, a supervised machine learning ensemble method. Random Forest constructs multiple decision trees on bootstrap samples of the data and averages their predictions, enhancing predictive accuracy while mitigating overfitting and model variance (Breiman 2001; Biau and Scornet 2016). This method is particularly well-suited for this investigation analysis because it can capture nonlinear relationships and complex interactions among variables, providing

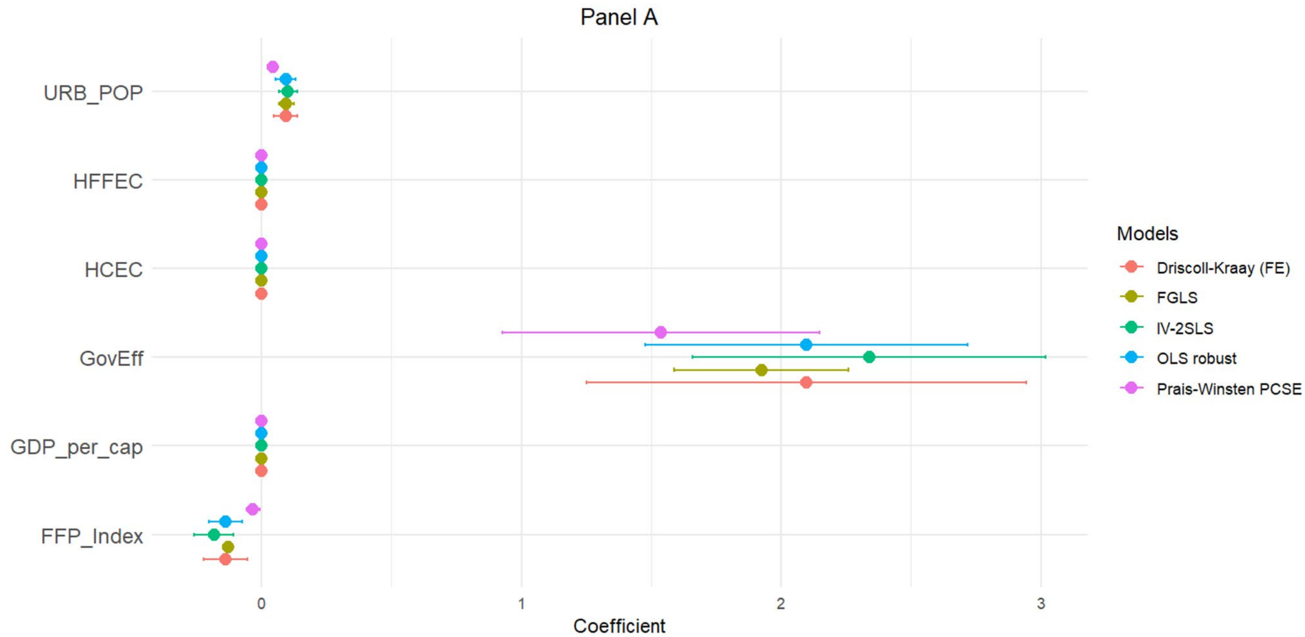

Fig. 6 Visualization of the coefficients in Panel A.The analysis and visualization were conducted in RStudio (version 2025.05.1). The dataset was structured using the packages dplyr and tidyr, and the graphical representation was produced with ggplot2

a robustness check and an alternative, non-parametric perspective on the determinants of CO_2 emissions in EU countries. The model was implemented in RStudio (version 2025.05.1) using the *randomforest* package, with 500 trees (ntree = 500) to ensure stable and reliable predictions, and three variables considered at each split (mtry = 3), following best-practice recommendations for regression tasks. Variable importance was evaluated using two standard metrics: %IncMSE, which measures the increase in prediction error when a variable is randomly permuted, and IncNodePurity, which quantifies the total reduction in node impurity contributed by each variable across all trees. Overall, this Random Forest analysis complements the econometric results by highlighting the relative importance of each explanatory factor, reinforcing the robustness of our findings and offering insights that may be overlooked in parametric models. Table 6A in the Appendix and Fig. 9 below, reports the variable importance measures and model summary for both specifications. The results indicate that the variables GDP/cap and URB-POP are consistently the most influential predictors, followed by HFFEC, HCEC, and GovEff. Importantly, female political participation (FFP_Index) also emerges as a significant contributor, reinforcing the evidence that gender-inclusive institutions play a meaningful role in reducing emissions.

These findings confirm and extend the econometric results, demonstrating the robustness of the gender–environment nexus across alternative methodological frameworks. Indeed, to confirm the results presented in Tables 6 and 7, this study conducts a robustness check using Quantile Regression (QR). The following subsection presents the results of this robustness analysis.

Robustness Check

Quantile Regression (QR), introduced by Koenker and Bassett (1978), extends traditional mean regression by estimating conditional quantiles of the dependent variable, allowing a more comprehensive analysis of the distributional effects of explanatory variables. Unlike OLS, which focuses solely on the conditional mean, QR captures the heterogeneity of relationships at different points in the outcome distribution, making it particularly suitable when the impact of predictors varies across low, median, and high levels of the dependent variable. In QR, the model estimates the $\tau - th$ conditional quantile (e.g., 25th, 50th, 75th, 95th percentiles) of the response variable Y, given covariates W, as:

$$Q_Y\left(T|W\right) = X\beta'(\tau) \quad (12)$$

where $\beta\left(\tau\right)$ represents the vector of quantile-specific coefficients. The estimation minimizes a weighted sum of absolute residuals, assigning different penalties to over- and under-predictions depending on the quantile τ.

In this study, quantile regression (QR) is used as a robustness check to assess whether the effects of explanatory variables vary across different levels of environmental emissions. Unlike mean-based models (e.g., Driscoll–Kraay FE, FGLS, IV-2SLS), which estimate average effects, QR

Table 7 Estimation results

Panel B															
Dependent variable (GHGe/cap)															
Models															
Variables	(a) OLS robust			(b) Driscoll–Kraay (FE)			(c) FGLS			(d) Prais–Winsten PCSE			(e) IV-2SLS		
HFFEC	2.28E-10	5.67	***	2.28E-10	10.84	***	2.49E-10	6	***	1.62E-10	4.23	***	2.48E-10	5.05	***
HCEC	-1.5E-10	-6.74	***	-1.5E-10	-10.76	***	-1.5E-10	-8.85	***	-1.1E-10	-4.17	***	-1.5E-10	-5.45	***
URB-POP	8.33E-08	4.28	***	8.33E-08	3.18	***	8.1E-08	10.66	***	5.15E-08	6.02	*	9.5E-08	5.29	***
GDP/cap	-1.1E-12	-5.98	***	-1.1E-12	-4.02	***	-2.3E-12	-3.19	***	-5.6E-13	-4.17	***	-1.2E-12	-5.81	***
GovEff	2.65E-06	6.83	***	2.65E-06	4.95	***	2.51E-06	19.4	***	1.9E-06	3.99	***	2.89E-06	6.66	***
FFP_Index	-1.6E-07	-4.21	***	-1.6E-07	-3.19	***	-1.4E-07	-16.67	***	-3.4E-08	-2.55	**	-2.1E-07	-4.76	***
Constant	4.72E-06	5.23	***	4.72E-06	4.62	***	4.67E-06	11.3	***	4.46E-06	6.83	***	4.95E-06	5.43	***
Observations	260			260			260			260			234		

Post-estimation tests					
Wald Test for Joint Significance of Coefficients				Tests of endogeneity	
F(3, 253) = 22.13***	F(6,25) = 70.26***	Wald chi2(3) = 1095.88***	Wald chi2(3) = 121.00***	Robust Score Chi² (6)	7.58418 (0.2702)
				Robust Regression F(3,221)	1.02808 (0.3810)

Shea's partial R-squared test		
Variable	Shea's Partial R²	Strength
HFFEC	0.9345	Strong
HCEC	0.9405	Strong
URB-POP	0.9097	Strong
GDP/cap	0.9260	Strong
GovEff	0.8767	Strong
FFP_Index	0.6057	Strong

Notes: The Stata 17 commands, r*egress*, *xtreg, fe, xtgls, xtpcse*, and *ivreg2* were used for estimations in columns (a)–(e), respectively. Post-estimation tests employed test, *ivendog, estat endogenous*, and *estat firststage*. Instrument strength was assessed using *ivreg2* with the *first option* for Shea's Partial R². Significance levels: *** p<0.01, ** p<0.05, * p<0.1.

captures distributional heterogeneity, providing insight into how relationships may differ in low-, median-, and high-emitting contexts. By estimating effects at $\tau = 0.25$, 0.50, 0.75, and 0.95, QR helps determine whether variables exert consistent or varying influences across the emission spectrum. This approach ensures that results are not limited to the average but are robust throughout the distribution. Moreover, unlike traditional models that assume homogenous effects across observations, quantile regression captures differential impacts along the emissions distribution. This is especially relevant in EU contexts where structural differences among countries may lead to divergent environmental responses to policy inputs. Table 8 below presents the results of this robustness check.

Table 8 above presents quantile regression results for CO_2 emissions per capita (**Panel A**) and GHGe per capita (**Panel B**) across the 25th, 50th, 75th, and 95th quantiles. This method reveals how the effects of explanatory variables vary across the emissions distribution. In both panels, HFFEC has a positive and significant impact on emissions at most quantiles, confirming its role in driving higher emissions. HCEC consistently shows a negative effect, indicating its effectiveness in reducing emissions. URB-POP is positively associated with emissions, especially at higher quantiles, reflecting greater urban pressure in high-emitting contexts. GDP/cap has a negative effect across all quantiles, stronger at higher levels, suggesting that economic growth can support emission reductions through efficiency or cleaner technologies. Government Effectiveness is positively related to emissions, likely due to increased energy use in better-governed economies. The FFP_Index maintains a negative association with emissions, with stronger effects at higher quantiles, reinforcing the environmental benefits of greater female political participation. Overall, these results confirm the robustness of findings from Tables 6 and 7, while highlighting how variable effects differ across the emissions spectrum. Moreover, Fig. 10 below shows the visualization of the coefficients in **Panels A** and **B**.

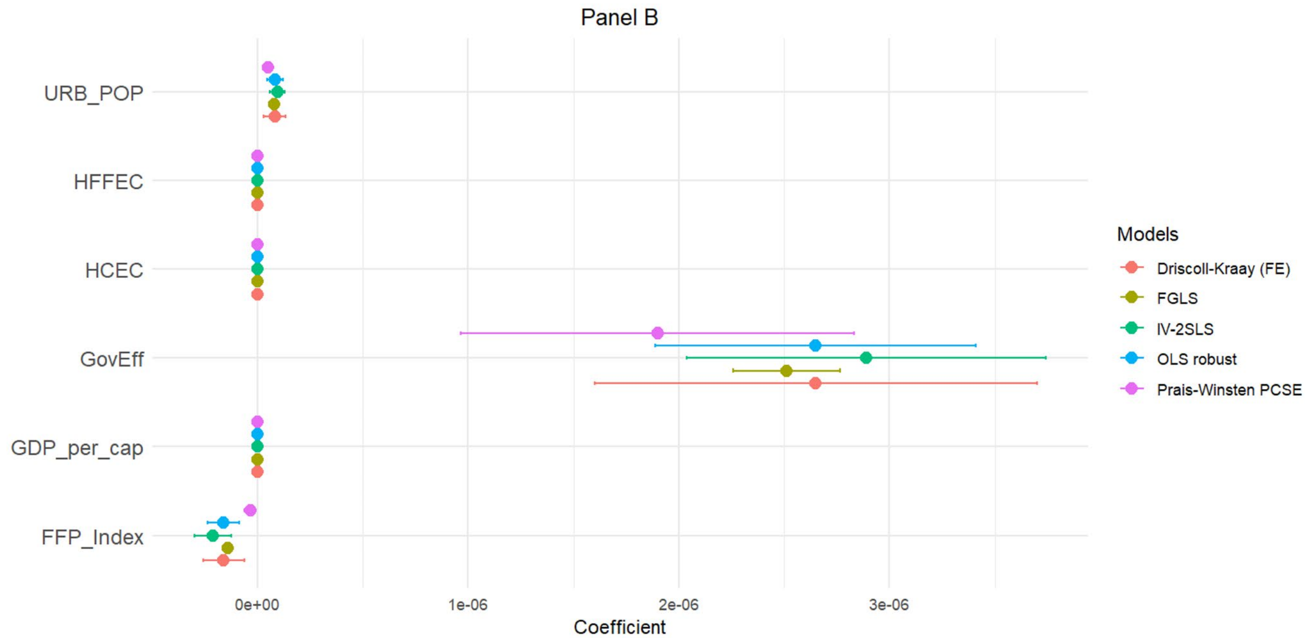

Fig. 7 Visualization of the coefficients in Panel B.The analysis and visualization were conducted in RStudio (version 2025.05.1). The dataset was structured using the packages dplyr and tidyr, and the graphical representation was produced with ggplot2

Discussions

This study investigates how women's political participation, household fossil fuel and clean energy consumption, urbanization, per capita GDP, and government effectiveness affect environmental degradation—measured by per capita CO_2 and total greenhouse gas (GHG) emissions—across European Union (EU) countries. Using both traditional panel data models and panel quantile regression, the analysis reveals not only average effects but also how these relationships vary across different emission levels. While results from both approaches are largely consistent, quantile regression highlights significant heterogeneity across countries, especially those with higher emissions.

One of the most important findings is that household fossil fuel consumption (HFFEC) has a strong and positive association with both CO_2 and total GHG emissions. This effect intensifies in higher-emitting countries, underscoring the urgent need to reduce fossil fuel use, especially at the household level.

In contrast, Household clean energy consumption (HCEC) has a statistically significant negative effect on emissions across all models, and this effect becomes stronger in higher quantiles. This supports the EU's energy transition policies and suggests that investments in clean energy yield the greatest environmental benefits in high-emitting nations.

Per capita GDP (GDP/cap) is found to reduce emissions, with stronger effects at higher levels of emissions. These results align with the Environmental Kuznets Curve (EKC) hypothesis, which posits that after a certain level of economic development, environmental degradation begins to decline. Therefore, economic growth can support sustainability—particularly in wealthier, high-emission countries.

The role of urbanization (URB-POP) is more complex. While it contributes to higher emissions overall—likely due to increased energy demand—this effect is more pronounced in higher-emitting countries. These findings point to the importance of sustainable urban planning and investment in energy-efficient infrastructure, especially in rapidly urbanizing regions.

A particularly striking result concerns women's political participation, measured by the Female Political Participation Index (FFP Index). Across all methods, the index is negatively and significantly associated with CO_2 and GHG emissions. The effect is most pronounced in high-emitting countries, suggesting that women's participation in political decision-making leads to the adoption of more environmentally sustainable policies. This supports theoretical frameworks such as feminist ecological economics and environmental justice, and is consistent with research showing that women tend to prioritize environmental protection, climate action, and social equity (Ergas and York 2012; McCright 2010; Dollar et al. 2001; Mavisakalyan and Tarverdi 2019; Koengkan and Fuinhas 2021).

Unexpectedly, government effectiveness (GovEff) is positively associated with emissions, especially in high-emission countries. Although effective governance is typically assumed to support environmental policy implementation, these findings suggest that it may also facilitate economic

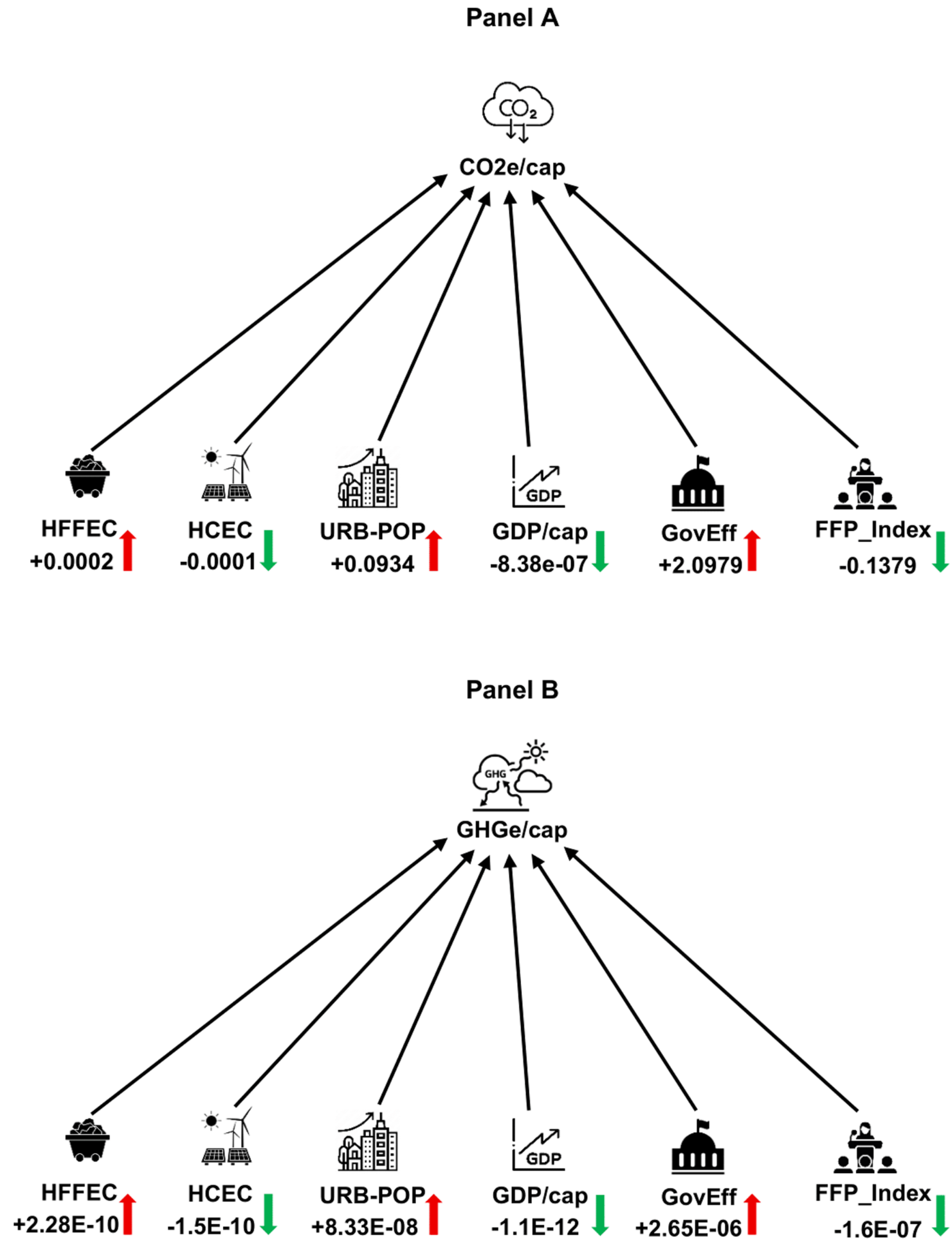

Fig. 8 Summary of the impacts of explanatory variables on the dependent variable based on Estimator, (b) results (Tables 6 and 7). This figure was created by the authors

expansion and energy use, inadvertently increasing emissions. This raises the possibility that combining effective governance with greater female political representation could enhance environmental outcomes. Gender-sensitive governance may thus serve as a complementary tool in achieving sustainability goals.

The findings reveal that environmental degradation is shaped not only by economic and energy-related factors but also by institutional and social dynamics—particularly gender and governance. Both panel and quantile regression analyses emphasize the need for targeted policy strategies. For countries facing high environmental pressure, prioritizing gender equality and accelerating the energy transition could produce the greatest environmental benefits.

Enhancing women's involvement in environmental and energy policymaking can be a powerful strategy for sustainable development. Policies that support clean energy access and reduce fossil fuel consumption—especially at the household level—should incorporate gender perspectives. Governance reforms that emphasize transparency, accountability, and inclusiveness can facilitate women's full participation in decision-making processes. Similarly, urban development strategies should be designed with gender sensitivity to better align environmental and social objectives.

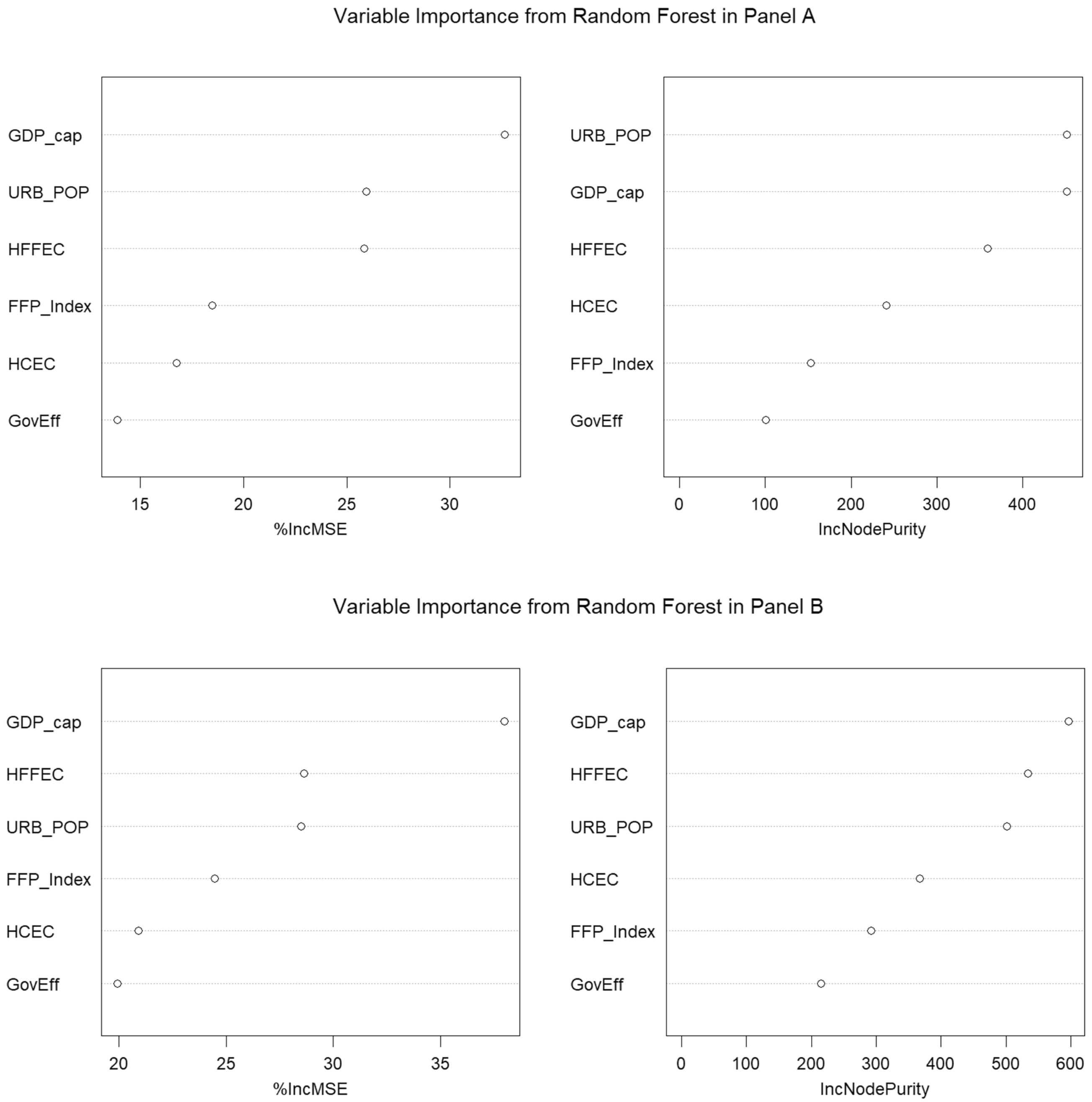

Fig. 9 Variable importance from Random Forest in Panels A and B. The variable importance shown in Panels A and B was obtained using a Random Forest regression model, a supervised machine learning technique, implemented in RStudio with the randomForest package (version 2025.05.1). The variable importance rankings were visualized using the varImpPlot function

This study has several limitations. First, the FFP Index captures women's representation quantitatively but does not reflect qualitative dimensions of political influence, such as leadership in environmental committees, legislative sponsorship, or network centrality. While comparable EU-wide data are unavailable for the study period, future research could integrate such multidimensional measures to strengthen construct validity.

Second, the focus on 26 EU countries provides a controlled institutional setting but limits generalizability. The EU's supranational governance frameworks create homogeneity not present elsewhere, and local or national dynamics may be overlooked.

Third, while household fossil fuel and clean energy consumption are employed as instruments for female political participation, this approach is not without limitations.

Table 8 Estimation results

Panel A												
Dependent variable (CO_2e/cap)												
Variables	Q 0.25			Q.50			Q.75			Q.95		
HFFEC	0.0002	7.15	***	0.0002	7.48	***	0.0002	7.04	***	0.0003	5.28	***
HCEC	− 0.0001	− 4.45	***	− 0.0001	− 5.10	***	− 0.0001	− 4.23	***	− 0.0001	− 8.12	***
URB-POP	0.0735	2.92	***	0.0763	5.54	***	0.0948	3.50	***	0.1495	4.57	***
GDP/cap	− 5.25e-07	− 1.95	**	− 5.03e-07	− 3.55	***	− 7.37e-07	− 3.33	***	− 1.56e-06	− 6.81	***
GovEff	0.911	2.64	***	1.913	4.51	***	2.340	8.00	***	3.652	9.20	***
FFP_Index	-0.104	− 3.33	***	− 0.097	− 4.52	***	− 0.128	− 3.30	***	− 0.265	− 6.43	***
Constant	1.291	1.30		1.392	1.82	*	1.825	1.71	*	2.883	1.00	
Observations	260			260			260			260		
Panel B												
Dependent variable (GHGe/cap)												
HFFEC	3.48e-10	6.45	***	2.78e-10	5.79	***	2.90e-10	7.98	***	9.19e-11	0.68	
HCEC	− 1.64e-10	− 4.57	***	− 1.74e-10	− 5.92	***	− 2.05e-10	− 7.62	***	− 1.21e-10	− 1.88	*
URB-POP	9.32e-08	3.57	***	5.81e-08	3.57	***	1.01e-07	4.80	***	1.02e-07	1.72	*
GDP/cap	− 4.81e-13	− 2.36	**	− 6.73e-13	− 2.18	**	− 1.12e-12	− 6.92	***	− 2.57e-12	− 5.39	***
GovEff	5.58e-07	1.40		2.95e-06	5.89	***	2.98e-06	8.00	***	2.88e-06	5.58	***
FFP_Index	− 7.62e-08	− 2.94	***	− 1.23e-07	− 3.41	***	− 1.91e-07	− 7.13	***	− 3.71e-07	− 5.23	***
Constant	1.73e-06	1.18		5.14e-06	4.85	***	5.42e-06	5.59	***	1.41e-05	2.24	**
Observations	260			260			260			260		

Notes: The Stata 17 command *sqreg* was used to estimate quantile regressions at the 25th, 50th, 75th, and 95th percentiles. Standard errors were computed using bootstrapping with 100 repetitions. Significance levels: *** $p<0.01$, ** $p<0.05$, * $p<0.1$

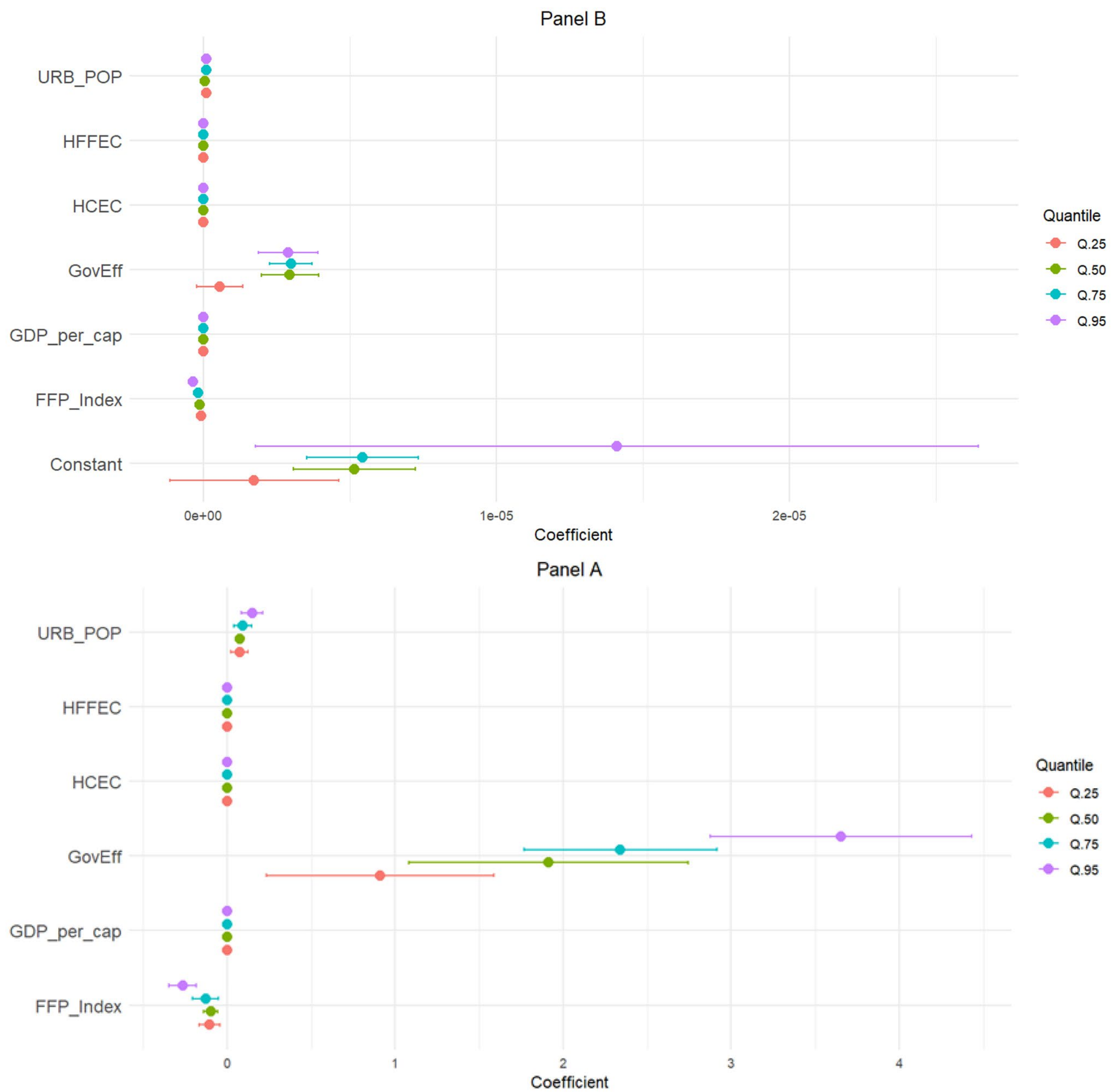

Fig. 10 Visualization of the coefficients in Panels A and B.The analysis and visualization were conducted in RStudio (version 2025.05.1). The dataset was structured using the packages dplyr and tidyr, and the graphical representation was produced with ggplot2

Alternative causal strategies—such as difference-in-differences, regression discontinuity, or synthetic control methods—could be applied in future research with more suitable datasets.

Fourth, the mediation analysis is based on sequential regressions, which are appropriate for small-N, short-T panels but fall short of the rigour offered by structural equation modelling or bootstrapped mediation. Richer and longer datasets would enable such methods to test mediation pathways more formally. Finally, the study period (2013–2023) may be too short to capture long-term structural dynamics. Although chosen for its stability and alignment with key EU strategies, longer horizons and broader geographical coverage could improve understanding of temporal, non-linear, and threshold effects.

In addition, while the findings of this investigation provide important insights into the impact of women's representation on environmental policies in European Union countries, it is important to recognize that this relationship may function differently across political and cultural contexts. For instance, a comprehensive study of BRICS countries (Gonda 2024) found that women's political empowerment, particularly in senior executive positions, significantly reduced carbon emissions. Conversely, other reports suggest that greater marital freedom for women can lead to higher emissions due to increased consumption and energy use. This comparison highlights a dimension that merits consideration, given our study's exclusive focus on the EU context. Omitted variables such as cultural factors, societal environmental awareness, or institutional arrangements may alter both the direction and the strength of the relationship between female representation and environmental outcomes. Future research that accounts for this diversity in an inclusive manner will provide a more holistic understanding.

It should also be emphasized that the priorities of women MPs regarding environmental policies are not homogeneous. The literature identifies this heterogeneity along three main axes. First, party affiliation matters: women politicians from left-wing parties tend to advocate stricter environmental regulations, whereas those from right-wing parties may prioritize balancing environmental concerns with economic growth. Second, the type of environmental issue plays a role: climate change, biodiversity loss, and localized pollution often receive different levels of attention. Third, electoral system incentives matter: women's environmental policy initiatives are more visible in proportional representation systems, while this effect is weaker in majoritarian systems. These contextual differences must therefore be considered to fully understand how women's representation translates into environmental policy outcomes.

This study underscores the multifaceted nature of environmental degradation and highlights the transformative potential of women's political participation in promoting sustainability. In high-emission countries, where policy interventions are most urgently needed, enhancing gender-inclusive governance and accelerating the transition to clean energy emerge as mutually reinforcing strategies. These findings offer strong empirical support for integrating gender equality into environmental policymaking and point toward more inclusive, effective, and sustainable solutions for addressing climate change. A key contribution of this study lies in its methodological triangulation. The use of machine learning in parallel with econometric analysis ensures that the findings are not dependent on a single analytical framework. While econometric models offer causal interpretation under specified assumptions, the Random Forest results provide complementary evidence by ranking the importance of explanatory variables in a data-driven manner. Together, these approaches confirm the relevance of female political participation in shaping environmental outcomes across the EU.

Conclusions and Policy Implications

Conclusions

The growing global environmental and climate challenges underscore the urgent need for comprehensive and inclusive strategies to achieve net-zero emissions. This study investigates the impact of female political participation, household energy consumption patterns, urbanization, economic development, and government effectiveness on environmental degradation—specifically, CO_2 and greenhouse gas (GHG) emissions per capita—in 26 European Union countries between 2013 and 2023.To address this, the study employs traditional panel econometric techniques, including five different estimation models, panel quantile regression, and the Dumitrescu and Hurlin (2012) Granger causality test.

These estimation methods were selected based on the characteristics of the panel data to ensure robust and reliable results.The findings reveal that higher levels of female political participation are consistently associated with lower CO_2 and GHG emissions per capita. This outcome supports a growing body of literature suggesting that gender-inclusive governance frameworks are more effective in implementing policies that significantly reduce environmental degradation. Additionally, economic development—as measured by GDP per capita—also contributes to lowering CO_2 and GHG emissions per capita.

Household clean energy consumption is shown to mitigate emissions, whereas fossil fuel energy consumption within households is identified as a key driver of increased emissions. These results underscore the need to accelerate the transition from fossil fuels to renewable energy at the household level. Furthermore, urban population growth and government effectiveness are found to be positively correlated with emissions, suggesting the complexity of balancing development and sustainability goals.

To examine the variability of these effects across the distribution of emissions, the study applies panel quantile regression as a robustness check. The results confirm that the negative relationship between female political participation and emissions holds consistently across the distribution, with particularly strong effects at higher quantiles.

From a theoretical standpoint, these findings align with feminist ecological economics and environmental justice frameworks, both of which emphasize the importance of inclusive and participatory governance in achieving sustainability. By linking female political participation to measurable environmental benefits, the study contributes to the growing literature on the intersection of gender, governance, and environmental performance. It reinforces the argument that women's political involvement at the national level within the EU advances both gender equality and environmental sustainability.

The robustness of these findings is reinforced by their consistency across various estimation techniques, underscoring the reliability of the key conclusions. Ultimately, this study highlights the significant and measurable roles that gender-inclusive institutions, clean household energy use, and economic development play in shaping environmental outcomes in the European Union.

Policy Implications

Based on the empirical findings, this study offers several policy implications for the European Union, emphasizing the need to integrate gender equality and energy transition initiatives into climate strategies to achieve more effective and sustainable results.

Given the robust negative relationship between female political participation and emissions, promoting women's representation should be regarded not only as a matter of social justice but also as a strategic pillar of climate action. Rather than generic calls for parity, targeted interventions are needed to place women in decision-making spaces that directly shape climate outcomes. Institutional entry points such as parliamentary environmental committees, budget and finance committees, and national energy transition task forces should be prioritised, as these are key venues where emissions reduction targets, budget allocations, and regulatory frameworks are debated.

Governments should adopt legislative gender quotas and gender-balanced appointments that explicitly include these strategic committees. To ensure effectiveness, quotas must be carefully designed according to the electoral system, combined with placement mandates that guarantee women are positioned in winnable seats and senior decision-making roles. Implementation challenges may arise in member states with decentralised governance structures or strong party gatekeeping, which can hinder women's entry into key political arenas. To address these challenges, capacity-building initiatives such as leadership training, mentorship programmes, and funding for female candidates should be implemented. These efforts would enhance the readiness of women to engage meaningfully in environmental and energy policy discussions.

The EU should also embed gender equality as a cross-cutting objective in major climate initiatives, including the European Green Deal, REPowerEU, and Fit-for-55 packages. These frameworks should include gender-responsive budgeting to track how financial resources are allocated toward climate and energy programmes that benefit women and vulnerable populations. Concrete metrics for monitoring progress could include: (1) the percentage of women in climate-related committees and regulatory agencies, (2) annual changes in CO_2 and GHG emissions in countries achieving gender parity in parliament compared to those that do not, and (3) the share of public climate investment directed toward inclusive, community-led energy projects.

Household-level interventions remain critical. Given that household fossil fuel use is strongly associated with emissions, EU policies should target women-led and low-income households for subsidies, grants, and infrastructure upgrades that support the adoption of clean energy. These initiatives should be evaluated through metrics such as reductions in household-level emissions and increases in renewable energy penetration in residential sectors.

Finally, effective governance requires the integration of inclusive leadership frameworks. While this study finds a positive association between government effectiveness and emissions—likely reflecting economic expansion—transparent, participatory governance structures that integrate women's voices can enhance the legitimacy and sustainability of environmental policies. Evaluation frameworks should monitor whether increased female participation in governance correlates with measurable improvements in air quality, renewable energy uptake, and progress toward net-zero targets.

In conclusion, mitigating environmental degradation requires harmonised, gender-sensitive strategies that bridge environmental, economic, and social policy domains. Collaboration among ministries of environment, energy, finance, and gender equality should be institutionalised through joint task forces and policy impact assessments. Future research could extend these findings along several complementary lines. At the micro level, investigating mechanisms such as committee dynamics and gender-specific legislative behaviour would shed light on how women's representation translates into concrete environmental policy outcomes. At the institutional level, comparing countries with stronger and weaker environmental frameworks, or interacting measures of government effectiveness with policy stringency, could clarify whether the gender–environment relationship depends on institutional capacity. Further work could also examine the robustness of results under alternative weighting schemes as more refined measures of institutional influence become available.

Crucially, incorporating post-2023 data as they are released would capture the impact of major policy developments, including the continued implementation of the European Green Deal, the REPowerEU plan, and the forthcoming EU Gender Equality Strategy 2026–2030. These initiatives are likely to influence both female political representation and the stringency of environmental policies, potentially reinforcing the nexus identified in this study. Finally, analysing how this relationship evolves amid rising climate ambition, growing energy security concerns, and shifting political dynamics in member states would generate timely insights. A forward-looking research agenda of this kind would help keep empirical evidence aligned with the EU's path towards carbon neutrality and inclusive governance in the coming decade.

Appendix

See Tables 9, 10, 11, 12, 13, 14

See Fig. 11

Table 9 Correlation

Variables	CO_2e/cap	GHGe/cap	HFFEC	HCEC	URB-POP	GDP/cap	GovEff	FFP_Index
CO_2e/cap	1.0000							
GHGe/cap	0.9338***	1.0000						
HFFEC	0.0265	– 0.0455	1.0000					
HCEC	– 0.0209	– 0.113**	0.8676***	1.0000				
URB-POP	0.3182***	0.223***	0.0504	0.1993***	1.0000			
GDP/cap	– 0.1187*	– 0.1446**	– 0.0719	– 0.0583	0.0397	1.0000		
GovEff	0.3772***	0.3564***	0.0399	0.1016	0.5349***	– 0.1234**	1.0000	
FFP_Index	0.0197	– 0.0185	0.2103***	0.2821***	0.557***	– 0.3029***	0.5813***	1.0000

Notes: The Stata 17 command *pwcorr* was used; Significance levels: *** p<0.01, ** p<0.05, * p<0.1

Table 10 Multicollinearity test

Panel A		
Multicollinearity test		
Variables	VIF	1/VIF
	Dependent variable (CO_2e/cap)	
HFFEC	4.34	0.2302
HCEC	4.51	0.2217
URB-POP	1.83	0.5463
GDP/cap	1.18	0.8449
GovEff	1.69	0.5910
FFP_Index	2.10	0.4768
Mean VIF	2.61	
Panel B		
Variables	Dependent variable (GHGe/cap)	
HFFEC	4.34	0.2302
HCEC	4.51	0.2217
URB-POP	1.83	0.5463
GDP/cap	1.18	0.8449
GovEff	1.69	0.5910
FFP_Index	2.10	0.4768
Mean VIF	2.61	

Notes: The Stata 17 commands *vif* was used

Table 11 Skewness/Kurtosis tests

Variables	Obs	Pr (Skewness)	Pr (Kurtosis)	adj chi2(2)	Prob>chi2
CO_2e/cap	260	0.0000	0.0000	73.47	0.0000***
GHGe/cap	260	0.0000	0.0000	57.21	0.0000***
HFFEC	260	0.0000	0.0000	N.A	0.0000***
HCEC	260	0.0000	0.0000	N.A	0.0000***
URB-POP	260	0.1464	0.0000	27.34	0.0000***
GDP/cap	260	0.0000	0.0000	N.A	0.0000***
GovEff	260	0.0689	0.0009	12.43	0.0000***
FFP_Index	260	0.0671	0.0037	10.46	0.0000***

Notes: The Stata 17 command *sktest* was used; Significance levels: *** p<0.01, ** p<0.05, * p<0.1. N.A denotes not available

Table 12 Heteroskedasticity and Autocorrelation tests

Panel A	
Heteroskedasticity and Autocorrelation tests	
Heteroskedasticity test	Autocorrelation test
Chi2(1)=148.06***	F(1, 25)=93.216***
Panel B	
Heteroskedasticity test	Autocorrelation test
Chi2(1)=85.81***	F(1,25)=100.829***

Notes: The Stata 17 commands *xtserial, hettest* were used. Significance levels: *** p<0.01, ** p<0.05, * p<0.1

Table 13 Hausman test

Panel A
chi2(4)=78.70***
Panel B
chi2(5)=63.42***

Notes: The Stata 17 command *hausman* was used. Significance levels: *** p<0.01, ** p<0.05, * p<0.1

Table 14 Variable Importance and model summary from Random Forest Machine Learning Model

Panel A				
Variable Importance				
Feature	%IncMSE	IncNodePurity		
HFFEC	25.83770	359.0686		
HCEC	16.77394	240.9014		
URB-POP	25.96273	451.5597		
GDP/cap	32.64505	451.4038		
GovEff	13.89304	100.5550		
FFP_Index	18.49121	152.9907		
Panel B				
HFFEC	28.64470	534.1565		
HCEC	20.90958	366.5160		
URB-POP	28.48811	500.9983		
GDP/cap	37.95978	596.9429		
GovEff	19.92902	215.1051		
FFP_Index	24.45869	291.7419		
Panels' Summary				
Metric			Value	
			Panels	
			A	B
Type of Random Forest			Regression	Regression
N countries			26	26
Number of Trees			500	500
Variables Tried at Each Split			3	3
Mean Squared Residuals (MSE)			0.4993	0.8037
% Variance Explained			92.83	91.76

Notes: Variable importance and model summary were obtained using a Random Forest regression model, a Machine Learning technique, implemented in RStudio with the RandomForest package (version 2025.05.1). Importance metrics reported are %IncMSE and IncNodePurity. Model parameters: 500 trees (ntree = 500) and 3 variables tried at each split (mtry = 3)

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Data Availability Data is available on request.

Declarations

Conflict of interest No competition of Interest.

Ethical Approval Not applicable.

Consent to Participate Not applicable.

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Declaration of Generative AI in Scientific Writing During the preparation of this work, the author(s) used Grammarly v.1.2.96.1473 to improve the quality of the text and make necessary grammatical and sentence corrections in all text. After using this tool/service, the author(s) reviewed and edited the content as needed and take(s) full responsibility for the final publication.

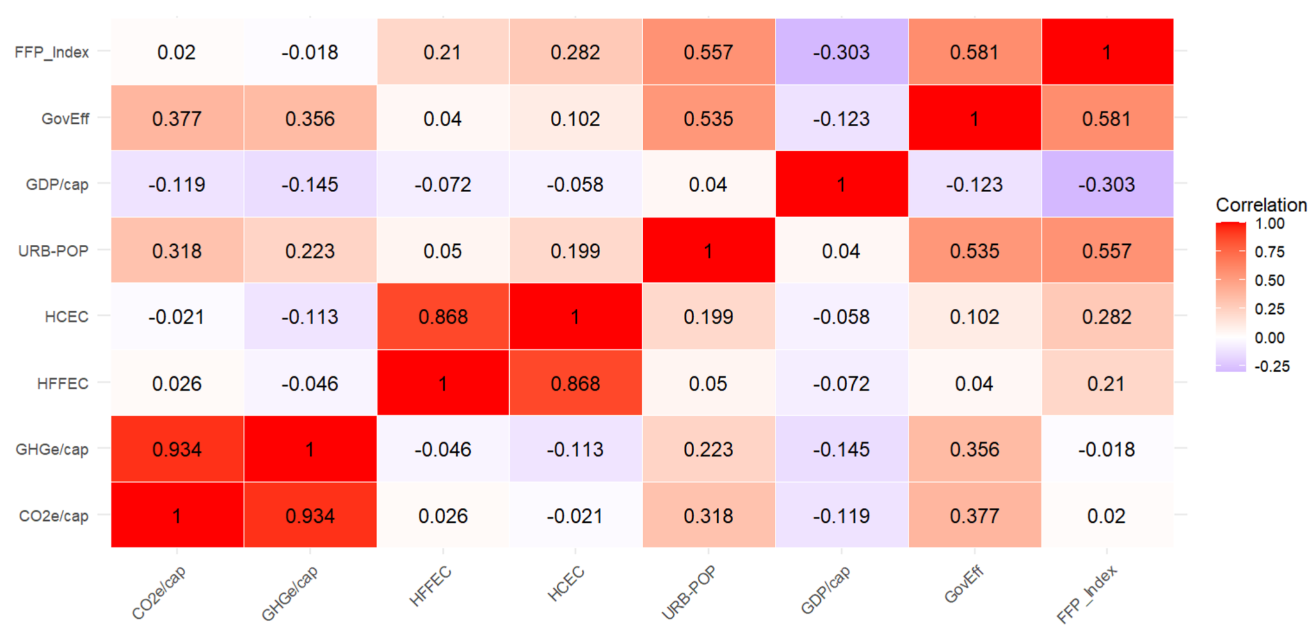

Fig. 11 Correlation matrix of variables. The correlation analysis was conducted using RStudio package (2025.05.1). The correlation matrix was computed with the base functions and visualized using the package corrplot

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