Closing the Gender Gap in Climate

Can Electing More Women into Office Narrow the Climate Divide?

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Abstract

In this paper, I seek to identify what the causal effect of electing higher proportions of women in national legislatures in African and Arab nations is on yearly CO_2 per capita emissions. I utilize a two-stage difference-in-difference approach on 64 different countries throughout Africa and the Middle-East. To eliminate the potential endogeneity in using the proportion of women in national legislature as a treatment on per capita CO_2 emissions, I use years since women were granted suffrage in each respective country as an instrumental variable. One nominal economic paper has shown evidence that increased women in parliaments is more than likely causally related to stricter climate policies (Mavisakalyan & Tarverdi, 2019). However, this paper fails to take into account how carbon emissions and women in government have evolved over time. In this paper, I expand on this research by showing directly how carbon emissions

have changed over time to elicit a causal effect between changes in gender compositions at the national legislature level. When the average proportion of women in national legislatures in African and Arab countries increase by 1%, I find a significant effect that total yearly (gigatons of) CO_2 emissions per capita decrease by 0.273 (-0.306, -0.240) on average.

Introduction

The current climate and economic research today overwhelmingly supports the thesis that women—especially in developing countries—are disproportionately affected by the negative externalities of climate change. During periods of increased droughts, which are exacerbated due to the effects of climate change, women often make long trips to meet agricultural, hygiene, and family needs. This has been linked to increased violence against women (UNDP, 2019). Increased risks of poverty, food insecurities have also been shown to increase as a result of increased rates of natural disasters and variable weather patterns. One article reports that, "When a family is faced with the impact of the climate crisis, girls' education is one of the first things families drop" (Medlicott, 2021). Another meta-analysis consisting of 53 studies report that two-thirds of those studies find that women are more often the victims of death or injury in the case of extreme weather events (Seller, 2016). If women are, on average, affected disproportionately affected by climate externalities, then one would expect that, on average, rational policymakers who are women would enact policies concerning climate policy at higher and potentially at more effective rates (pertaining to protecting the needs of women) relative to other compositions of policymakers with lower proportions of women, assuming policymakers who are women have the means and power to do so.

The outline of this paper will proceed as follows: I will first introduce the data I collected for this study and assess its reliability; I will then discuss the identification strategy I use to identify used to estimate the causal effect. I will then finish by discussing the results of my econometric analysis and its implications and then conclude.

Data

The data used for this analysis was strategically gathered using publicly available sources (either by ethical web-scraping methods or direct download). The Inter-Parliamentary Union (IPU) maintains a consistent archive of records that monitors the number of women elected into the national legislature for each country. Not every government maintains the same structure of government, however, the IPU distinguishes between women elected into the lower-house and upper-house levels of government throughout multiple time periods. I take the average proportion of women holding office in the national legislature overall in a given year by taking a weighted average of the two houses. I obtained emissions data, including total yearly (gigatons of) CO_2 emissions per capita per country from the Emissions Database for Global Atmospheric Research (EDGAR) as maintained by the European Commission. I obtained country-year-level specific covariates, such as population and GDP per capita, from the

World Bank. I obtained the data base linking each country's year they passed some legal right of suffrage from "A Lexical Index of Electoral Democracy" on the Harvard Dataverse (Skaaning et. al, 2015). I checked country-specific anomalies using data from the Pew Research Center (Schaeffer, 2020). Finally, I used U.S. Department of State's regional definitions to define which countries I define as either African or Arab.

By adjusting for time, country, and other specific endogenous effects on the proportion of women in government in a country (c) at year (t), I hope to show that, effectively, when proportions of women in government increase, CO_2 emissions decrease. However, one potential problem with data is that I fail to account for country-specific government intricacies that could prevent that could prevent women or a legislature from passing effective climate policy–such as failing to account for government corruption or other unique anomalies (wars between countries, coups, etc, that prevent a government from acting efficiently)–at every year. I assume that on average, these factors in the idiosyncratic term of my model are 0 for any given country at year t.

In my sample, the United Arab Emirates (UAE) and Saudia Arabia are excluded in my sample because they don't give full suffrage to women: UAE held its first national elections in 2006, but only a select amount of men and women were eligible to participate (Schaeffer, 2020). Saudi Arabia doesn't hold national elections. Women are, however, enfranchised on the local level government. This leaves us with a total of 64 different countries in the sample. Using all the data which I was able to find data for the number of women in national legislature for a

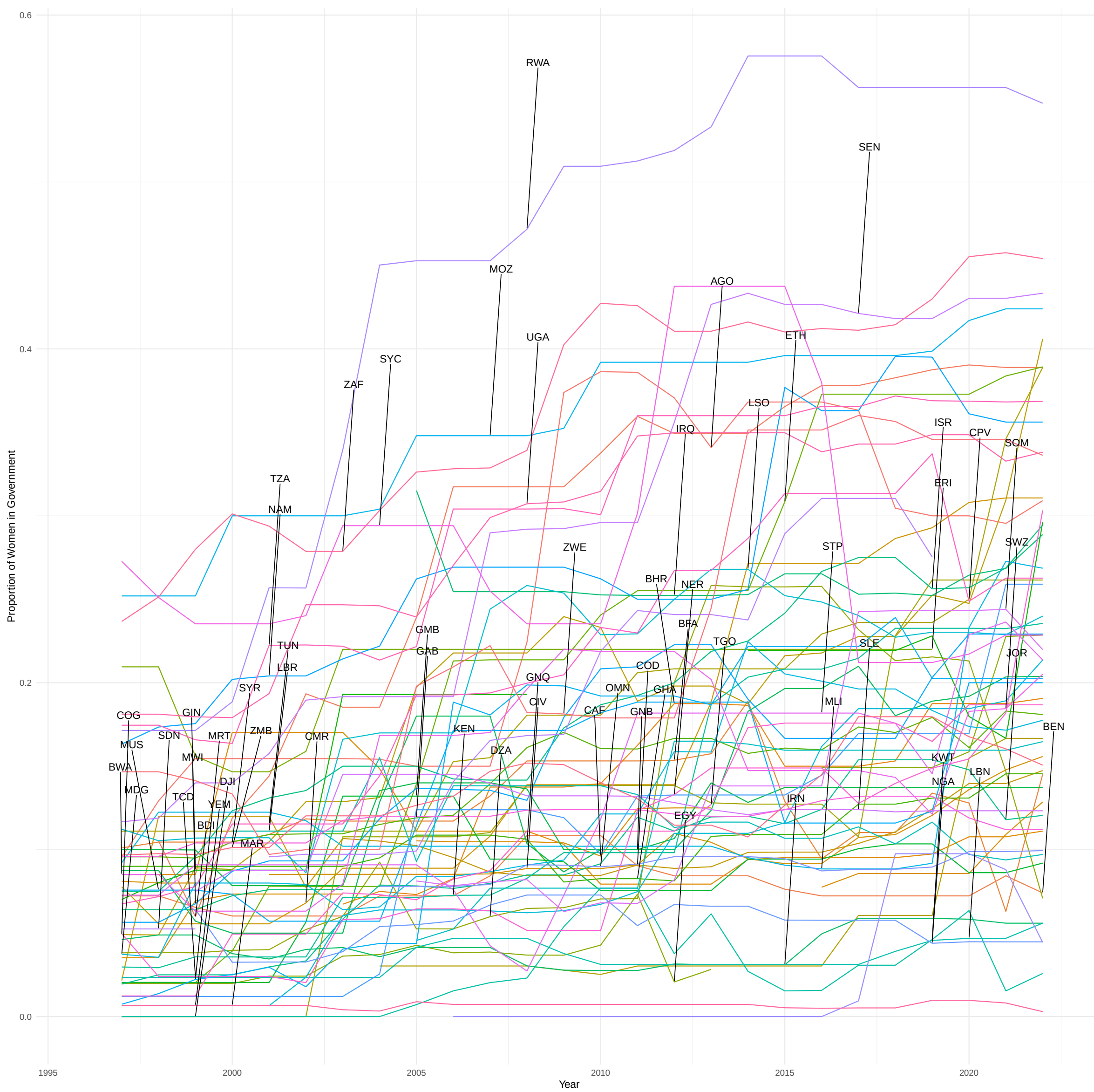

Figure 1 – Average Proportion of Women in National Legislatures Over Time

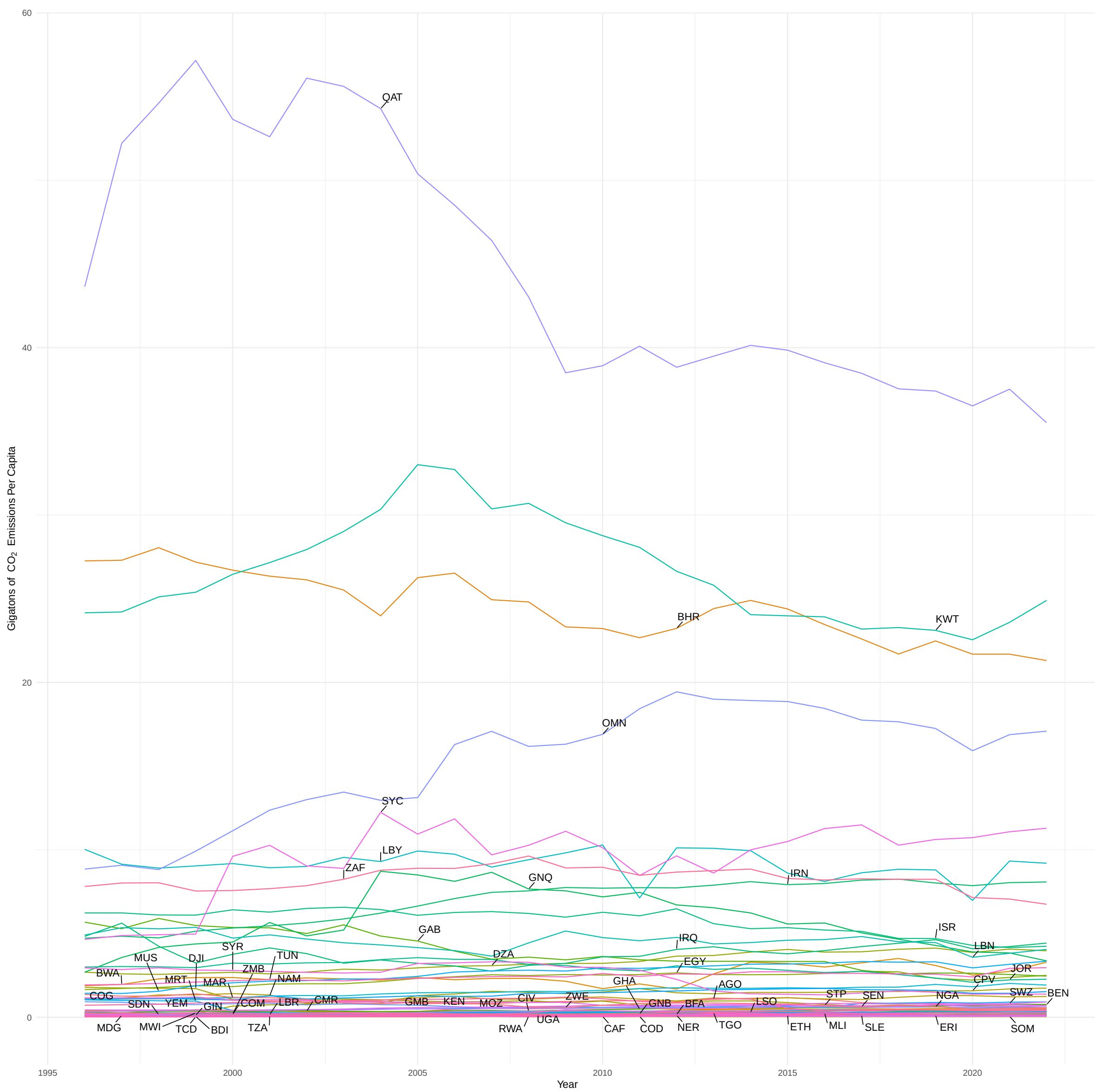

Figure 2 – Per Capita Carbon Emissions Over Time

given country using data from the IPU, I was left with data spanning 1997-2022 when merged with CO_2 emissions data from EDGAR.

Table 1: Summary Statistics of Key Variables

Variable	Mean	SD	Min	Max
CO2 Emissions Per Capita	3.224	7.345	0.021	57.16
Proportion of Women in Gvt	0.155	0.108	0	0.575
GDP	46603801502	88805073403.5	75951136	644035510272
Population	19253149	27937416	77319	218541216
GDP per Capita	4646.585	10138.79	99.757	98041.36
Suffrage	1968.109	13.363	1947	2006

Identification

To estimate the causal effect that electing more women in national legislatures in African and Arab countries has on reducing per capita CO_2 emissions, I employ a two-stage difference-in-difference identification strategy.

Let Y_{ct} represent the total (gigatons of) carbon emissions per capita for a country (c) in a year (t). I seek to identify consistently and interpret the causal effect of δ, the effect of higher proportion of women in national legislatures in African and Arab countries on national CO_2 emis-

Table 2: Summary Statistics by Country

Country	N	CO2 Emissions Per Capita	Proportion of Women in Gvt	GDP	Population	GDP per Capita	Suffrage
AGO	26	0.92681	0.25617	61136510425	23819407.27	2395.2409	1975
BDI	26	0.04481	0.29223	1858691481	9013842.27	195.7666	1962
BEN	26	0.48000	0.07758	9427907948	9513753.81	931.5446	1960
BFA	26	0.14869	0.12574	9879179362	16222351.12	565.0930	1960
BHR	26	24.41692	0.14652	23866885356	1110339.23	19985.1394	2002
BWA	26	2.51423	0.10956	11679307166	2093686.46	5383.7271	1966
CAF	26	0.05631	0.09069	1717825115	4516393.23	371.8470	1960
CIV	26	0.41704	0.09852	37241947845	21228611.50	1671.4381	1960
CMR	26	0.37169	0.16573	26614285194	20110188.96	1264.4503	1960
COD	26	0.04538	0.10226	26137691195	67798246.61	349.2091	1967
COG	26	1.16296	0.19112	9890831710	4342198.92	2129.2625	1960
COM	26	0.22754	0.04802	839948182	658580.65	1232.7882	1975
CPV	26	1.22450	0.18505	1483184743	517811.46	2782.5743	1975
DJI	26	0.98342	0.12020	1568798898	908429.23	1606.6542	1977
DZA	26	3.33192	0.12144	133413556539	36256149.08	3572.0745	1962
EGY	26	2.32654	0.07382	210728516057	87759979.69	2261.6092	1956
ERI	26	0.15304	0.20727	1120484041	3025058.73	399.8786	1993
ETH	26	0.10427	0.21588	43426199788	89604476.46	424.0938	1955
GAB	26	3.89038	0.14328	12441574754	1731276.23	6892.6506	1960
GHA	26	0.48408	0.10488	35446770393	25480258.61	1257.3404	1957
GIN	26	0.20485	0.17269	7558680162	10412517.15	678.6265	1958
GMB	26	0.23292	0.08174	1276629190	1948392.35	641.2549	1965
GNB	26	0.18654	0.11840	874252813	1580784.61	520.2188	1974
GNQ	26	5.84538	0.13341	10680310300	1098129.11	9046.2942	1968
IRN	26	7.04769	0.04200	323334435131	75438778.31	4182.7934	1979
IRQ	26	3.84692	0.20797	128507863985	32491105.15	3634.0622	1958
ISR	26	5.72923	0.18622	250706702966	7614703.85	31445.1946	1948
JOR	26	2.93385	0.09033	25772832748	7476728.62	3171.5507	1974
KEN	26	0.30900	0.12401	49487400802	40920573.00	1097.4821	1963
KWT	26	26.68577	0.02475	102805583557	3032646.92	32629.9300	2006
LBN	26	4.53423	0.03430	32455205151	5123114.12	6143.9463	1952
LBR	26	0.20858	0.11397	2268958422	3925143.88	522.0986	1947
LBY	26	9.11577	0.12542	52387031513	5970720.88	8670.7073	1963
LSO	26	0.23662	0.19993	1813006912	2071184.42	868.2581	1966
MAR	26	1.56115	0.09184	91087307855	32367924.31	2710.8600	1962
MDG	26	0.10985	0.12502	9465270164	21695233.31	419.6738	1960
MLI	26	0.15212	0.12074	10274167660	15662439.38	608.4462	1960
MOZ	26	0.17431	0.35431	11612120970	23558217.61	478.6773	1975
MRT	26	0.67550	0.14266	5071728438	3476739.54	1375.6409	1960
MUS	26	2.70000	0.13275	9303361634	1234473.96	7466.9955	1968
MWI	26	0.27958	0.15362	7514886577	14841171.27	485.1673	1964
NAM	26	1.39731	0.27434	9018993290	2113443.23	4143.8100	1990
NER	26	0.08123	0.10870	7682887769	17080118.61	414.6091	1960
NGA	26	0.66092	0.05481	313024619441	161517413.54	1842.3328	1977
OMN	26	15.47154	0.08991	56775810363	3281428.62	16145.8623	2003
QAT	26	44.70962	0.02556	108516614223	1637385.77	57861.2484	2003
RWA	26	0.08800	0.44140	5978302715	10348245.77	536.6373	1962
SDN	26	0.32985	0.18022	47869298018	34068352.38	1223.7120	1964
SEN	26	0.53638	0.29679	15447790277	12665052.88	1161.7314	1960
SLE	26	0.12796	0.12040	2694746567	6391405.35	394.4146	1961
SOM	26	0.06238	0.14398	8216255642	12121495.23	537.6863	1964
STP	26	0.61331	0.13043	255006484	179928.11	1295.4893	1975
SWZ	26	0.88365	0.14917	3443873221	1098973.65	3091.8267	1968
SYC	26	9.72346	0.28266	1080544320	89129.92	11943.1457	1976
SYR	26	2.33077	0.11750	84109722132	19402662.58	4348.6179	1953
TCD	26	0.07246	0.09761	8092429967	12009809.12	629.2804	1960
TGO	26	0.26342	0.10986	4458357376	6566135.35	635.4826	1960
TUN	26	2.44115	0.20983	37604341839	10925330.88	3401.2543	1959
TZA	26	0.16458	0.29519	34798787505	46143525.54	716.0034	1963
UGA	26	0.10385	0.28969	21152773534	32832378.15	589.9307	1962
YEM	26	0.81108	0.00665	22320277039	24799160.46	860.6911	1990
ZAF	26	8.27500	0.36477	308320233630	52097664.77	5820.1135	1994
ZMB	26	0.28300	0.13306	15927797878	13906848.08	1061.2813	1964
ZWE	26	0.83592	0.22631	13726241713	13273904.50	985.4527	1979

sions per capita. Research from Mirziyoyeva & Salahodjaev (2023) suggests that women in government promote economic growth. This could pose potential problems of reverse causality: We need to be certain that any effect we see on women representation causes a reduction (or increase) in carbon emissions. If carbon emissions are highly correlated with economic growth and industrialization then an OLS regression on Y_{ct} to estimate δ would almost certainly be biased. To first eliminate the potential for reverse causality–the supposition that perhaps lower carbon emissions causes a higher proportion of women to become elected during the same year; or rather, perhaps lower carbon emissions is correlated with other variables (such as economic and human development) that are linked to higher proportions of women becoming elected–I lag the treatment variable of interest, W_{ct}, the average proportion of women in national legislature for country c during year t one year. This implies that any change in Y_{ct} will inherently come *after* a change in W_{ct}. Additionally, I introduce a matrix of country-year-specific covariates, X_{ct}, to adjust for additional confounding. This matrix includes a country's GDP per capita over time. This is also lagged one year to control for the fact that the effects on these covariates will necessarily imply that for any increase/decrease in Y_{ct}, the change will thus be associated with the previous year's economic growth for the respective country c.

In the two-stage design, I further suspect that there exists some endogeneity in W_{ct} s.t.

$$\exists \quad c_0, t_0 \quad s.t. \quad \mathbb{E}[W_{c_0 t_0 - 1} \epsilon_{c_0 t_0}] \neq 0$$

Both CO_2 emissions and women representation in government are highly correlated with industrialization and development. While X_{ct} can eliminate some of the endogeneity of W_{ct} by including GDP per capita, these covariates alone are not enough to explain a country's development. The rate at which a country develops, both industrially and politically, in regards to the countries in this study may be highly related to which European power originally occupied the territory and thus, how the power left the state when occupation was ceased, including the political institution(s) established. Endogeneity may also arise in what natural resources are available to the country at year t. Countries with a greater propensity to use carbon-intensive resources will thus have, on average, higher CO_2 emissions. If this also impacts women disproportionately such that women in these countries are more compelled to run for office and enact effective climate policies, then the estimand for δ will be biased in the negative direction.

To eliminate this endogeneity I invoke a two-stage difference-in-difference design, using an instrumental variable for W_{ct-1}. A common instrument in the literature for women parliament has been to use the number of years since the country passed a bill/law of suffrage (Mavisakalyan & Tarverdi, 2019). I use similar instrumental variable design here. Let Z_{ct-1} be the number years since a country c has granted suffrage (measured at year t, and lagged one year to align with W_{ct}).

In the difference-in-difference design, I use 62 country-fixed effects (omitting Angola and Burundi dummies due to colinearity). After lagging W_{ct} one year, I am left with data spanning 1998-2022, thus, including year-fixed effects, I omit 1999. I utilize the following two-stage estimation approach:

$$(1) \quad W_{ct-1} = \pi_0 + \lambda Z_{ct-1} + X'_{ct-1}\Psi + \sum_{k=\text{Benin}}^{Zimbabwe} \mu_k \mathbb{1}(\text{Country}_c = k) + \sum_{j=1999}^{2022} \eta_j \mathbb{1}(\text{Year}_t = j) + v_{ct}$$

$$(2) \quad Y_{ct} = \beta_0 + \delta W_{ct-1} + X'_{ct-1}\Omega + \sum_{k=\text{Benin}}^{Zimbabwe} \beta_k \mathbb{1}(\text{Country}_c = k) + \sum_{j=1999}^{2022} \gamma_j \mathbb{1}(\text{Year}_t = j) + \epsilon_{ct}$$

The validity of years since suffrage as an instrument

Mavisakalyan & Tarverdi, although make use of a cross-sectional data set, to their credit, report that, "an IV estimation can be potentially more convincing in a dataset that varies over time" (161). This is precisely what I employ here. I will expand on the research that uses years since suffrage as an instrument (Grier and Maldonado, 2015; Hicks et al., 2016). Relevance of Z_{ct-1} is present with the idea that women in countries where suffrage is granted will see more involvement in government over time as women's rights expand in politics. Mavisakalyan & Tarverdi note that "a country's history of suffrage should be highly relevant for female exposure to politics" (160).

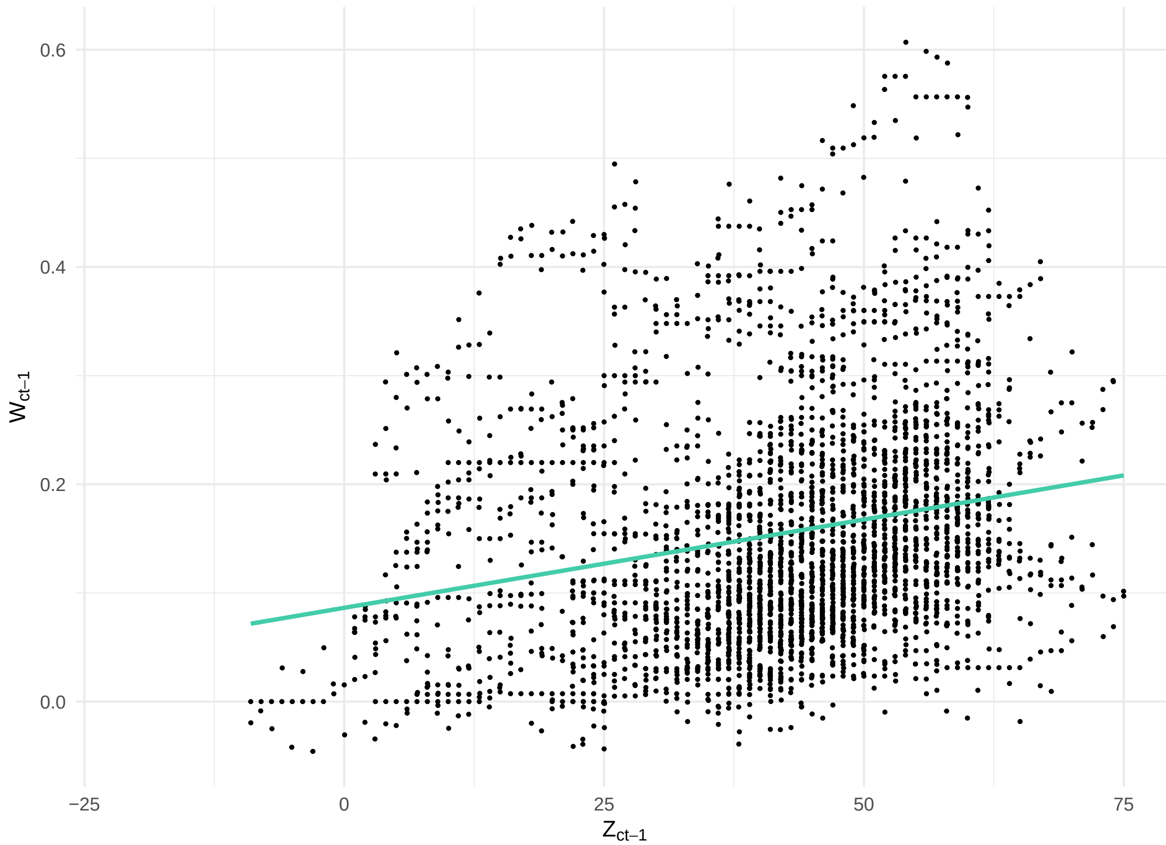

Figure 3 − Relevance of Instrumental Variable

The regression coefficient is significantly positive: 0.00162 (0.00127, 0.00198).

With the difference-in-difference approach, using country and yearly fixed effects, there is reason to believe that $Y_{ct}(W_{ct-1}) \perp Z_{ct-1} | X_{ct-1}, \text{Country}_c, \text{Year}_t$. In other words, regardless of when a country granted suffrage, the potential resulting carbon emissions per capita as function of the proportion of women in the national legislature in a country c are only dependent upon W_{ct-1} after fixed effects are accounted for. However, as Mavisakalyan & Tarverdi note, "[our instrument] would be questionable if there is the possibility that the history of suffrage affects climate change policy outcomes through mechanisms other than the Female in

Parliament." Mathematically, I assume,

$$\mathbb{E}[Z_{ct-1}v_{ct}] = 0 \quad \forall c, t$$

This could be violated if countries whose governments are established after colonial rule have (1) strong dependencies to the formal colonial power in terms of institutional structure and (2) as a function of that institutional structure, relies on carbon-intensive resources to kickstart its economy in post-colonial development. In other words, the instrument will be invalid if there are interactions between time, country, and carbon emissions. Evidently, more research needs to be done to see if this is the case, however, countries that granted suffrage to women immediately after occupation typically took place in the late 1900s, if it was less than random which countries these were, then the economic development–in relation to carbon-intensive resources–is also most likely non-random as well and thus would cause endogeneity in the instrument. In order to assert this, one would have to see a distinct pattern between the time periods of when women were granted suffrage and some other metric associated with carbon-intensive resource use such as oil exports. In this paper, I assume away this endogeneity by claiming that the instrument is as good as random (independent)–there is no distinct correlation between when a country granted suffrage and the types of resources the country typically consumes. Additionally, I assume monotonicity: the longer a country has passed from the year it granted suffrage, the more, on average, women are elected into national legislatures–I assume that there are no countries for which this trend reverses. The monotonicity assumption may be violated if women become less interested in politics over time for certain countries such that their participation decreases and hence the average proportion of women in national legislatures decrease with respect to years since suffrage.

Assumptions for difference-in-differences

In the absence of more (or less) women being elected, the total CO_2 emissions per capita for each country would have followed the same trend over time as the other countries that did not elect more (or less) women in national legislatures at that specific time. Since my treatment variable (W_{ct-1}) is continuous, the parallel trends assumption is hard to verify. However, I assume no other events or factors are influencing CO_2 emissions per capita differently across different countries over time. In other words, I assume no unobserved heterogeneity; my estimate is robust to omitted covariates–this is most likely satisfied if we believe in the exclusion restriction of our instrumental variable.

Following Bertrand et. al (2004), I cluster standard errors by countries, where I assume standard errors across countries are independent from each other. This is most appropriate given that within each country, CO_2 emissions are most likely correlated with each other given that the economic and industrial structure mostly remains consistent within each country from year to year (yet different across country). This assumption could be violated if groups of countries form specific treaties and agreements to reduce CO_2 emissions or make other agreements pertaining to usage of other carbon-intensive resources.

Table 3: Regression Results: Two-Stage Difference-in-Difference Regression Results (year and country-fixed effects omitted)

	CO_2 Emissions (Gigatons per Capita)
δ	-27.26**(***)**
	[-30.56,-23.97]
GDP per Capita$_{ct-1}$	0.0000176
	[-0.0000252,0.0000604]
β_0	5.560**(***)**
	[4.982,6.137]
N	1514

95% confidence intervals in brackets

(*)$p < 0.05$, **(**)**$p < 0.01$, **(***)**$p < 0.001$

Results

The difference-in-difference parameter, δ, was estimated using a two-stage design and assumed to be linear. If my assumptions hold, then the economic and statistic implications are relatively straightforward (interpreted on a 1% scale): On average, for every 1% the average proportion of women in national legislatures in African and Arab countries increase by, the total yearly (gigatons of) CO_2 emissions per capita decrease by 0.273 (significant at the 5% level). Interestingly, the effect on GDP per Capita (lagged one year), included in the X_{ct-1} matrix, was found to have an insignificant effect, implying that the effects on GDP per capita growth (as a proxy variable for economic growth) on CO_2 per capita emissions can be almost entirely explained away by the year-level and country-level specific fixed effects.

Economic implications for climate policy

Given that the assumptions hold, the fact that δ yields a significant effect implies that either (1) higher rates of being elected to office in national legislatures does, in fact, reduce yearly CO_2 emissions per capita by virtue empowering women to enact effective climate policies, or perhaps the more peculiar option, (2) the mere increased presence of women in national legislatures somehow either allows those with power to pass climate reform to pressures climate policy reform to take place. (1) Assumes that when women are elected into the national legislature is directly associated to a transfer of power to women to reform policies. We would hope this would be true, however, due to the high amount of corruption even in the high levels of government that has persisted throughout a subset of the countries in this study, one cannot conclude from this study alone that women in African and Arab nations are able to pass more effective climate policy reform (carbon emission reform in this case) when more women are present in the national legislature. Even though this hypothesis could absolutely be true as well, this study can neither confirm nor deny this. Thus, I will consider the possibility of (2): It could be that, coincidentally, when higher concentrations of women are present in national legislatures, the governments, as a whole, are more sensitive and cognizant of climate policy and not necessarily that women are the innate cause of the reduction in carbon emissions. Regardless of which hypothesis holds (or neither), it seems to be the case (with high confidence) that when more women enter the national legislature, we see lower per capita CO_2 emissions.

Plausibility of Heterogeneous Treatment Effects

The estimate for δ is the local average treatment effect of increased proportions on reduced per capita CO_2 emissions for all countries. If the average proportion of women in national legislature affect per capita CO_2 emissions differently on a per country level, then each heterogeneous treatment effect would vary differently from δ (though centered around the true δ). Realistically, heterogeneous effects are most likely present: Countries that have different dependencies on different carbon-intensive resources (such as for Qatar and Kuwait–see Figure 2) and countries that are more prone to corruption and war will have different priorities at the national government level. Thus, even as more women become elected, per capita CO_2 may not decrease depending on the current priorities of the country. A future study could examine the treatment effects for each country. This study assumes away heterogeneous treatment effects for the tradeoff of a larger sample size and a smaller standard error for $\hat{\delta}$.

Conclusion

This paper reaffirms the effect that was outlined by Mavisakalyan & Tarverdi. Utilizing a causal two-stage difference-in-difference econometric framework, As national legislatures in African and Arab nations witness higher concentrations of women we see significantly lower per capita CO_2 emissions. While the exact mechanism through which women's presence influences climate policy remains open for further investigation, these findings hold significant policy implications. By encouraging women to run for office and dismantling barriers to their political participation, governments can potentially achieve environmental benefits alongside broader societal advancements. While this study focused on the average effect across countries, future research can delve deeper into heterogeneous treatment effects. Examining individual countries will allow for a more nuanced understanding of how factors like resource dependence and political structures interact with female representation to influence CO_2 emissions. Furthermore, investigating the specific policies enacted by legislatures with higher female representation can shed light on the mechanisms at play. This study underscores the importance of considering gender equality within climate change mitigation strategies. Understanding how women leadership shapes environmental policy is crucial for developing more effective strategies to achieve a sustainable future.

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