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Gender and climate change:
Do female parliamentarians make difference?

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ABSTRACT: This paper investigates whether female political representation in national parliaments influences climate change policy outcomes. Based on data from a large sample of countries, we demonstrate that female representation leads countries to adopt more stringent climate change policies. We exploit a combination of full and partial identification approaches to suggest that this relationship is likely to be causal. Moreover, we show that through its effect on the stringency of climate change policies, the representation of females in parliament results in lower carbon dioxide emissions. Female political representation may be an underutilized tool for addressing climate change.

JEL classification: D70, J16, Q54, Q58.

Keywords: gender, political representation, climate change, environmental policy.

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1. Introduction

Climate change is a serious threat, and it demands prompt policy response (Stern, 2008). Political commitment to address the issue is critical, yet sources of large differences in such commitments across countries are not fully uncovered. This paper focuses on the relevance of political identity, and in particular, the gender of politicians. We ask whether representation of females in political decision-making contributes to climate change policy action around the world.

That politician's gender identity may have implications for policy outcomes has been established in the literature. Studies have shown that female political representation contributes to significant changes in domestic and international policies including higher spending on health (Bhalotra and Clots-Figueras, 2014; Mavisakalyan, 2014) and education (Svaleryd, 2009; Clots-Figueras, 2012), more laws and expenditures relevant to female needs (Chattopadhyay and Duflo, 2004; Clots-Figueras, 2011), higher disbursements of foreign aid (Hicks et al., 2015, 2016). Furthermore, female politicians have been associated with outcomes such as better quality of institutions (Dollar et al., 2001; Swamy et al., 2001) and higher rates of economic growth (Jayasuriya and Burke, 2013). While the existing studies on the determinants of climate change policies across countries have highlighted the relevance of countries' formal and informal institutions (Fredriksson and Neumayer, 2013; Fredriksson and Wollscheid, 2015; Ang and Fredriksson, 2017; Mavisakalyan et al., 2017), to the best of our knowledge the role of politicians' identity has not received dedicated attention in this literature.

Gender differences in attitudes towards climate change identified in the general public suggest that females have greater awareness and concern about climate change than males do (McCright, 2010; McCright and Dunlap, 2011). Based on theories of gender socialisation, these differences can be linked to differences in values and social expectations conferred through socialization whereby cooperation and carefulness - values relevant for climate change action - are more emphasized in females than in males (Gilligan, 1982; Beutel and Marini, 1995). Gender differences in climate change concern can also be linked to differences in social roles performed in the society with production of climate change seen to be more closely linked to activities performed by males than females (Spitzner, 2009). Furthermore, consequences of climate change are likely to be gender-differentiated as well, with females more disproportionately bearing the costs of climate change due to their 'gendered labour and care roles and social status' (Seager et al., 2016, p.13).

In 'citizen candidates' models, in the absence of complete political commitment, politicians implement policies consistent with their preferences (Osborne and Slivinski, 1996; Besley and Coate, 1997). However, it is unclear whether we should expect to observe gender differences in preferences of politicians similar to those observed in the general public. Indeed, it is possible that females who pursue leadership roles in a predominantly male environment are similar to males (Adams and Funk, 2012). Consistent with this proposition, Sundstrom and McCright (2014) do not find robust evidence for gender differences in environmental concerns among Swedish parliamentarians although such gender differences are observed in the general public in Sweden. In the context of the US, however, Fredriksson and Wang (2011) find that female parliamentarians in the House of Representatives have more pro-environmental views compared to their male counterparts. Building on these observations, our study evaluates the implications of female representation in politics for climate change policies adopted by countries around the world.

From an econometric perspective, we are confronted with a problem of omitted variables: adoption of climate change policies and election of females to parliament may both be the product of underlying characteristics of countries we do not observe. We employ two key strategies to ascertain that our estimates are not driven by confounding factors. First, we use an instrumental variable based on countries' history of female political empowerment and estimate a 2SLS model. Second, we use a partial identification approach proposed by Oster (2016). The main idea of this approach is to study how large would the amount of selection on unobservables need to be, relative to the amount of selection on observables, to explain away the entire casual effect of female representation.

Based on a large sample of countries, we document a robust positive association between female representation in a country's parliament and the stringency of its climate change policies. The impact we identify is statistically significant and economically meaningful. As an illustration, our 2SLS estimation results imply that lifting the female representation in Bahrain, a country where females comprised just over 2% of parliamentarians in the study period, to the level of Denmark, a country with over 37% female representation, could lead to nearly 7-fold increase in the stringency of the country's climate change policies (in practice Denmark's climate change policies are over 8 times as stringent as Bahrain's). In extended results, we further demonstrate that through its effects on the stringency of climate change policies, female parliamentary representation results in lower carbon dioxide emissions.

The remainder of this paper is organised as follows. The next section describes the data used in the study. This is followed by a discussion of our empirical approach and the results of baseline analysis in section 3. In section 4 we test the sensitivity of the results to the choice of control variables and functional form, while section 5 is dedicated to addressing the issue of endogeneity employing instrumental variable and partial identification approaches. We extend the analysis to study the implications of female representation for carbon dioxide emissions in section 6. The final section concludes.

2. Data

We assemble a dataset for a sample of up to 91 countries based on various sources. Table 1 specifies the sources and presents summary statistics for the variables used in the baseline analysis.

Our measure of climate change policies is Climate Laws, Institutions and Measures Index (CLIMI) derived by Steves et al. (2011) based on the 2005-2010 annual national communications to the UNFCCC, and used in other published studies (e.g., Fredriksson and Neumayer, 2013, 2016). CLIMI measures the climate change mitigation policies adopted by countries. It is based on 12 components grouped into four key policy areas, with weights used to reflect the contribution of each of the components and areas to climate change mitigation: (i) international cooperation (0.1) [Kyoto ratification (0.5), Joint Implementation or Clean Development Mechanism host (0.5)]; (ii) domestic climate framework (0.4) [cross-sectoral climate change legislation (0.33), carbon emissions target (0.33), dedicated climate change institution (0.33)]; (iii) significant sectoral fiscal or regulatory measures or targets (0.4) [energy supplies/renewables (0.3), transport (0.13), buildings (0.07), agriculture (0.13), forestry (0.17), industry (0.2)]; and (iv) additional cross-sectoral fiscal or regulatory measures (0.1) [cross-sectoral policy measures (1)]. CLIMI ranges from 0 and 1, with higher values representing stricter policies. The average CLIMI score in the sample is 0.350. Tonga has the lowest CLIMI score in the sample at 0.011, while the UK has the highest score at 0.801.

[Table 1 about here.]

To capture the representation of females in politics we use the number of seats held by female members in single or lower chambers of national parliaments, expressed as a percentage of all occupied seats (Female in Parliament). Since CLIMI is measured based on information collected over the period 2005-2010, this and other time-variant explanatory variables in our analysis are averaged over this time period. The proportion of females in parliament in the sample averages at 19.25% (ranging from 0 to 52.55%). Figure 1 plots CLIMI against Female in Parliament revealing a strong positive relationship: countries with higher proportion of females in parliament have substantially more stringent climate change policies.

[Figure 1 about here.]

To substantiate on this relationship, our baseline analysis controls for several important characteristics of countries as highlighted in previous studies on the determinants of climate change policies. Following Fredriksson and Neumayer (2013, 2016), we control for countries' GDP per capita and openness (defined as imports plus exports divided by GDP). While the demand for climate change policies is expected to increase with GDP per capita, the expected effect of countries' openness on the stringency of climate change policies is ambiguous. As Neumayer (2002) points out, openness may contribute to cooperation on environmental problems, however it may also hamper it if exporting countries' interests are threatened.

The literature has highlighted the role of institutions in determining countries' commitments to address climate change. Countries' political regime is one important factor, with democracies promoting the enactment of environmental regulations more than autocracies do (e.g., Murdoch and Sandler, 1997; Farzin and Bond, 2006). Our analysis includes measures of countries' political regimes based on Polity IV data set (Marshall et al., 2016). The polity score captures the regime authority spectrum on a 21-pont scale ranging from -10 (hereditary monarchy) to +10 (consolidated democracy). We convert this into three regime categories, distinguishing between autocracies (-10 to -6), democracies (+6 to +10), and other regime types (Marshall et al., 2016). In our sample, 60% of the countries are democracies while 12% are autocracies (the remaining 18% have other political regimes).

Legal heritage of countries has been identified as a determinant of a broad range of its legal rules, regulations, and a number of other outcomes (La Porta et al., 2008). There is also evidence that it affects climate change policy outcomes across countries (Fredriksson and Wollscheid, 2015; Ang and Fredriksson, 2017). Legal origins are usually divided into two main groups: English common law and civil law; the later has four sub-strands: French, German, Scandinavian, and Socialist legal traditions (La Porta et al., 1998, 2008). Our analysis includes controls for these legal origins. Around 40% of the countries in the sample have French legal origin. Socialist legal origin underlies almost 31% of the countries, followed by English legal origin shared by nearly 19% of countries. German and Scandinavian legal origin groups are small: each of these groups comprises only 5% of countries in the sample.

3. Baseline results

To evaluate the baseline effect of the proportion of females in parliament on climate change policy of country i, we estimate the following model using ordinary least squares (OLS):

$$CLIMI_i = \alpha X_i + \beta FemParliament_i + \varepsilon_i \text{ for all } i = 1, ..., N. \quad (1)$$

where $FemParliament_i$ is the proportion of females in a country's parliament, X_i is a vector of controls for economic and institutional characteristics of countries defined in section 2, and ε_i is a disturbance term.

The OLS estimates of equation 1 are presented in Table 2. Column (1) presents a parsimonious specification which excludes other controls. Consistent with Figure 1, we estimate a positive highly significant coefficient on Female in Parliament with a magnitude suggesting a 0.10 point increase in CLIMI associated with a 10 unit increase in Female in Parliament.

Next, in column (2), we control for countries' GDP per capita and their openness. Introducing these controls into the regression leaves the sign and statistical significance of the coefficient on Female in Parliament unaffected, although the magnitude of the coefficient is slightly smaller. This is not surprising given the observations on the positive link between countries' economic development and socioeconomic status of females (Duflo, 2012). Countries with higher GDP per capita have more stringent climate change policies as captured by CLIMI. Openness on the other hand, is negatively associated with CLIMI although this estimate is not robust to introducing additional controls in subsequent specifications.

The characteristics of countries' political regime are added as controls in column (3). Reassuringly, the estimated positive significant coefficient on Female in Parliament remains robust to including these controls although it's magnitude is further reduced - an effect that is consistent with findings on positive association between the level of democracy and female parliamentary representation (Paxton et al., 2010). We establish that democracies, relative to other political regimes, have more stringent climate change policies. This is in line with other findings in the literature (e.g., Murdoch and Sandler, 1997; Farzin and Bond, 2006). The effect of autocracies on the stringency of climate change policies is indistinguishable from zero.

Finally, in column (4) we include controls for the legal heritage of countries. We confirm the positive significant coefficient on female parliamentary representation in this specification. A 10 unit increase in the female share of parliament is associated with 0.04 point increase in CLIMI. This effect persists, in spite of the findings in the literature on the positive effect of Scandinavian legal origin on female parliamentary representation (Austen and

Mavisakalyan, 2016). Furthermore, we find that relative to Scandinavian legal origin, countries with English, French and Socialist legal origins have less stringent climate change policies. No differences in the effects of German and Scandinavian legal origins on countries' climate change policies are found.

In column (5) we undertake to evaluate whether the results we find might be sensitive to the presence of influential observations. To identify these, we calculate DFbetas for CLIMI from baseline regression in column (4) and drop those observations for which $|\text{DFbeta}|>2/\sqrt{N}$ (Belsley et al., 1980). The results are remarkably similar to those reported in column (4) thereby confirming that the positive significant association between female parliamentary representation and the stringency of climate change policies we find is not driven by influential observations in the sample.

[Table 2 about here.]

4. Robustness checks

Additional controls. We have established a positive and statistically significant relationship between female representation in parliament and the stringency of climate change policies across countries. Yet, the possibility that female representation simply acts as a marker of unobserved characteristics of countries cannot be ruled out at this stage of the analysis. To mitigate this possibility, we control for additional variables that could be correlated with the unexplained component of CLIMI. The results of this exercise are summarised in Table 3. To allow for comparisons, in column (1) we report the estimation results of the baseline model (these are identical to those presented in column (4) of Table 2).

Average years of schooling in the population is one variable potentially correlated with the unexplained component of CLIMI. Research has documented that the prevalence of gender discriminatory attitudes around work decreases with education (e.g., Mavisakalyan, 2015). Meanwhile, education is also positively associated with pro-environmental behaviour (e.g., Franzen and Meyer, 2010; Mavisakalyan et al., 2017). Indeed, we estimate a positive significant coefficient on average years of schooling (sourced from Barro and Lee (2013)), however it leaves the estimated coefficient on Female in Parliament unchanged (column (2)).

Our baseline regressions control for important institutional features of countries. However there may be other characteristics of countries' institutional environment potentially correlated with both Female in Parliament and CLIMI. Most crucially perhaps, Female in Parliament may be simply picking up the effect of left-wing orientation in political power, which has been linked with pro-environmental outcomes (e.g., Neumayer, 2003, 2004). To address this possibility, we control for the political orientation of the national leader's party available from Keefer (2012).[1] Countries' historical experience of democracy is another important factor that has been linked with outcomes related to both gender equality (Beer, 2009) as well as climate change policies (Fredriksson and Neumayer, 2013). We therefore additionally control for a measure of democratic capital stock accumulated in years 1800-2010 that comes from Fredriksson and Neumayer (2013). The results of the model that controls for these additional institutional variables are presented in column (3). While the estimated coefficient on Left government is insignificant, Stock of democracy is positively significantly correlated with CLIMI. The estimate on Female in Parliament is remarkably similar to that from the baseline model.

The significance of history in determining contemporary outcomes may manifest through channels other than those hitherto considered. History of colonialism is one possible channel, having been linked with gender inequality (Mavisakalyan, 2015) as well as with environmental outcomes of countries (Marchand, 2012). The results reported in column (4) are from a model that includes dummies to identify former British and French colonies.[2] These variables are insignificant, and they do not alter the estimated effect of Female in Parliament on CLIMI.

Next we consider whether our results may be the outcome of omitting important geo-economic characteristics of countries. We first control for the countries' natural resource wealth as measured by the net value of their oil and gas exports taken from Ross and Mahdavi (2015). The literature has documented a negative link between oil and gas wealth and female political participation (Ross, 2008). Resource production, however, also bears important implications for the climate change policies countries implement given the traditionally strong opposition from the oil and gas industry. We additionally control for latitude - a variable that has been linked with Western European influence and quality of institutions (Hall and Jones, 1999) as well as directly capturing important geographic differences across countries (Rodrik et al., 2004). Column (5) reports the results of the regression that controls for these variables. We confirm the positive significant coefficient on Female in Parliament. The estimated coefficient on Oil & Gas net export per capita, while negative, is statistically insignificant as is that on latitude.

[1] The source classifies parties as left if their names reveal them to be communist, socialist, or social democratic or if they are labeled as left-wing (Beck et al., 2001).

[2] These do not completely overlap with English and French legal origins although they are correlated.

Finally, in the last column of Table 2, we include dummies for continents. After controlling for important socio-economic, institutional and geographic characteristics of countries, no independent effect of sharing a continent on the stringency of climate change policies is found. Moreover, the estimated positive effect of Female in Parliament on CLIMI persists after these characteristics of countries have been accounted for.

[Table 3 about here.]

None-parametric analysis. It can be argued that the OLS method imposes some restrictive assumptions on the model specification. Here we relax two of these assumptions; namely assumptions regarding the normal distribution of residuals and the functional form relating the main independent variable to CLIMI.

First, we re-examine the relationship between female parliamentary representation and CLIMI, relaxing the assumption on normally distributed residuals (i.e. ε_i). We do so by estimating the model with quantile regression method. By doing that, we also test the robustness of the results against extreme values of our dependent variable. In quantile regression, the sample is divided into quantiles based on the distribution of dependent variable and therefore the main model (i.e. equation 1) becomes:

$$Q_{CLIMI}\left(\delta \mid X_i, FemParliament_i\right) = \beta_{1,\delta} X_i + \beta_{2,\delta} FemParliament_i + \mu_i \quad \text{for all } i = 1, ..., N. \tag{2}$$

In equation 2, as an arbitrary combination, δ can be 0.25, 0.50, 0.75 quantiles. The relevant results are reported in Table 4. The results corresponding to the first quantile, 0.25, suggest that the effect is not present; however, when we move to the centre of distribution, 0.5 quantile, the effect of female in parliament is significant. Results for the last qunatile, 0.75, show significant effect, although with slightly smaller magnitude.

[Table 4 about here.]

Next, we further explore the relationship between female parliamentary representation and the stringency of climate change policies when some of the functional form assumptions are relaxed. The results we have documented so far are based on the assumption that the function relating $FemParliament_i$ to $CLIMI$ is linear. Relaxing this functional form assumption means the main model in equation 1, now changes to:

$$CLIMI_i = \theta X_i + f\left(FemParliament_i\right) + v_i \text{ for all } i = 1, ..., N. \tag{3}$$

In comparison to equation 1, equation 3 is more flexible as $f(.)$ can accommodate different functional forms. We allow functional form to be different for every point of $FemParliament_i$

and estimate the model using Gaussian Kernel. Figure 2 illustrates the estimates of $f(.)$ and as shown, the effect is always positive and within the confidence intervals. [3]

[Figure 2 about here.]

5. Addressing endogeneity

Employing an instrument. The previous sections have established a robust statistically significant positive relationship between female representation in parliament and the stringency of climate change policies across countries. However, whether this finding can be given a causal interpretation can be questioned because of unobserved heterogeneity: places that are different for a variety of unobserved reasons will differ in their political representation of females as well as in their climate change policy outcomes. In section 4, we made an attempt to mitigate the problem of unobserved heterogeneity by adding proxy variables that could be correlated with the unexplained component of CLIMI, however obviously, unobserved heterogeneity can never be fully accounted for. A conventional way to deal with this problem is to use an instrumental variable, a source of exogenous variation in Female in Parliament, and estimate a 2SLS model. This is what we undertake to do here.

We exploit the electoral experience of females in society, as proxied by the years since female suffrage was granted, as an instrument to identify the effect of Female in Parliament. Sourced from UN Women (2011), this instrument has been used in other identification approaches applied to different contexts (Grier and Maldonado, 2015; Hicks et al., 2016). The intuition in doing so is simple: a country's history of suffrage should be highly relevant for female exposure to politics, however it is unlikely that it directly affects its contemporary policy outcomes. As an informal way of testing this, in column (2) of Table 5, we report the results of a regression of Years since suffrage on CLIMI which excludes Female in Parliament as a regressor (column (1) presents the OLS estimates of the effect of Female in Parliament on CLIMI; these are identical to those presented in column (4) of Table 2.). The coefficient on Years since suffrage is positive and significant, however it turns insignificant once Female in Parliament is controlled for in column (3) suggesting that the effect of our instrument on the dependent variable is likely to operate through its effect on the endogenous variable.

[3] In the estimation of the relationship for Figure 2 the baseline control variables (i.e. X_i) are considered in their linear form. Results are available upon request.

The 2SLS estimates of equation 1 are presented in columns (4) and (5) of Table 5. Female in Parliament is treated as endogenous and modeled as:

$$FemParliament_i = \gamma X_i + \zeta Suffrage_i + \omega_i \text{ for all } i = 1, ..., N. \tag{4}$$

where $Suffrage_i$ is the number of years since female suffrage was granted in a country, X_i is a vector of controls for economic and institutional characteristics of countries defined in section 2, and ω_i is a disturbance term.

The first stage estimation results reported in column (4) further support the validity of our identification strategy: we estimate a highly significant positive association between the years since suffrage and female parliamentary representation. Column (5) presents the corresponding results from the 2nd stage of the 2SLS model. Not only we confirm the positive significant coefficient on Female in Parliament, the estimated magnitude is larger: a 10 unit increase in Female in Parliament is associated with an increase of 0.10 in CLIMI. This estimate implies that the 35.5 unit difference in female representation between two typical countries with high and low representation, Bahrain and Denmark, would translate into 7-fold difference in the stringency of climate change policies. This is a relatively precise estimate, given around 8-fold difference in the actual stringency of climate change policies between the two countries. In comparison, corresponding OLS estimate from the most extensive parametric model suggests a tangible (over 2-fold) effect of female parliamentary representation on the stringency of climate change policies however it is much smaller than the actual gap in the stringency of climate change policies between the two countries.

[Table 5 about here.]

Partial identification. Our instrumental variable strategy, while convincing at face value, 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. Naturally, it is not possible to control for all possible variables that might be correlated with the years since suffrage and CLIMI. Here we take an alternative strategy to assess the potential impact of unobserved heterogeneity: we use a partial identification approach proposed by Oster (2016) to evaluate how large would the amount of selection on unobservables need to be, relative to the amount of selection on observables, to explain away the entire causal effect of Female in Parliament on CLIMI. To do that, we evaluate the bias-adjusted coefficient derived by Oster (2016):

$$\beta^* \approx \tilde{\beta} - \delta[\dot{\beta} - \tilde{\beta}] \frac{R_{max} - \tilde{R}}{\tilde{R} - \dot{R}} \tag{5}$$

where $\dot{\beta}$ and $\dot{R}$ are the coefficient and the R-squared from estimating the equation 1 where $\alpha = 0$ and $\tilde{\beta}$ and $\tilde{R}$ are the coefficient and R-squared from the regression where $\alpha \neq 0$, i.e. other explanatory variables for CLIMI, in addition to Female in Parliament, are included.

δ denotes the relative importance of observable relative to unobservable variables in generating bias; R_{max} is the R-squared from a hypothetical regression of CLIMI on all observable and unobservable variables. Both of these measures are unknown. Hence, Oster (2016) proposes a bounding approach: the estimated effect of Female in parliament should range from $\tilde{\beta}$ to β^* estimated under an assumption of $\delta = 1$, and given values of $R_{max} \in [\tilde{R}, 1]$. We apply two assumptions on the value of R_{max}: (i) $R_{max} = 0.90$, i.e. that the measurement error in CLIMI accounts for 10% of the variation therein; (ii) $R_{max} = min\{1.3\tilde{R}, 1\}$ - the rule of thumb proposed by Oster (2016).

The results of this analysis are presented in Table 6. In columns (2) and (4) we report the coefficient bounds $[\tilde{\beta}, \beta^*]$ for models with baseline and comprehensive lists of controls (corresponding to those reported in column (4) of Table 2 and column (6) of Table 3). $\tilde{\beta}$ comes from the specifications controlling for all baseline/comprehensive observables. β^* is evaluated using equation (5) by setting $\delta = 1$ and applying the two assumptions on the value of R_{max}. The identified sets, $[\tilde{\beta}, \beta^*]$ exclude zeros in all cases - a finding that suggests that at least some of the estimated effect of Female in Parliament might be causal. Furthermore, in columns (1) and (3) we observe that in all cases $\delta > 1$, i.e. that the unobservables would have to be more important than the observables in explaining CLIMI; this finding provides further support to the validity of our results.

[Table 6 about here.]

6. Implications for emissions

Female parliamentary representation appears to be a significant factor in explaining climate change policies across countries. Here we ask whether this finding has implications for actual outcomes: do the policy changes associated with female representation result in lower carbon dioxide emissions? To explore these relationships, we estimate the equation 1 jointly with the following equation determining carbon dioxide emissions, using 3SLS:

$$CO2_i = \eta C_i + \mu D_i + \lambda CLIMI_i + \kappa FemParliament_i + \iota_i \text{ for all } i = 1, ..., N. \quad (6)$$

where $CO2_i$ is the carbon dioxide emissions per capita of a country, C_i is a sub-set of controls of variables included in X_i in equation 1: the measures of countries' legal origins are included in X_i but not in C_i; D_i contains two additional determinants of $CO2$ that are excluded from equation 1: a squared term of GDP per capita (consistent with the Environmental Kuznets Curve hypothesis), and population density; ι_i is a disturbance term.

The results of these estimations are summarised in Table 7. In columns (1) and (2) we set $\kappa = 0$, i.e. we assume that CLIMI is the only mechanism linking Female in Parliament with emissions. In columns (3) and (4) we relax this assumption to allow for other transmission mechanisms. The coefficients of interest are β and λ: is female representation in parliament associated with more stringent climate change policies, and do climate change policies lower carbon dioxide emissions at the same time? Our response to both questions is affirmative. Based on the results reported in columns (1) and (2), a 10 unit increase in female representation, through its effect on the stringency of climate change policies, results in a 0.49 (=0.04*12.153) metric tones decrease of carbon dioxide emissions per capita. The results presented in columns (3) and (4) provide no evidence to suggest that Female in Parliament may affect the emissions directly or through a mechanism other than CLIMI. The significant negative effect of Female in Parliament on carbon dioxide emissions mediated by climate change policy-making remains much the same.

[Table 7 about here.]

7. Conclusion

The lack of political commitment to address climate change around the world warrants an inquiry into underlying sources. In this paper, we have asked whether the lack of female political representation may be one such source. Our results confirm that this is the case: female representation in national parliaments leads to more stringent climate change policies across countries, and by doing so, it results in lower carbon dioxide emissions.

The results of this study have important policy implications. They suggest that manipulation of the gender identity of politicians might yield changes in climate change actions countries are opting for. Moreover, various international campaigns to address climate change may succeed more in places where more females are represented in political power. Various forms of affirmative action to increase female representation in politics have been increasingly introduced by countries in recent years. Our results suggest that these are likely to result in increase in countries' commitments to address climate change.

A number of interesting questions remain to be addressed in future research. If the sources of gender differences in climate change concerns are related to differences in social and economic positions of females and males in the society, will these differences persist as the position of females changes? We established that gender identity of politicians matters for climate change action but other dimensions of identity such as age and ethnicity as well as intersectionality across various dimensions may play a role as well. Exploring the links between various dimensions of politicians' identity and climate change policy-making appears to be a direction with high potential returns to further analysis.

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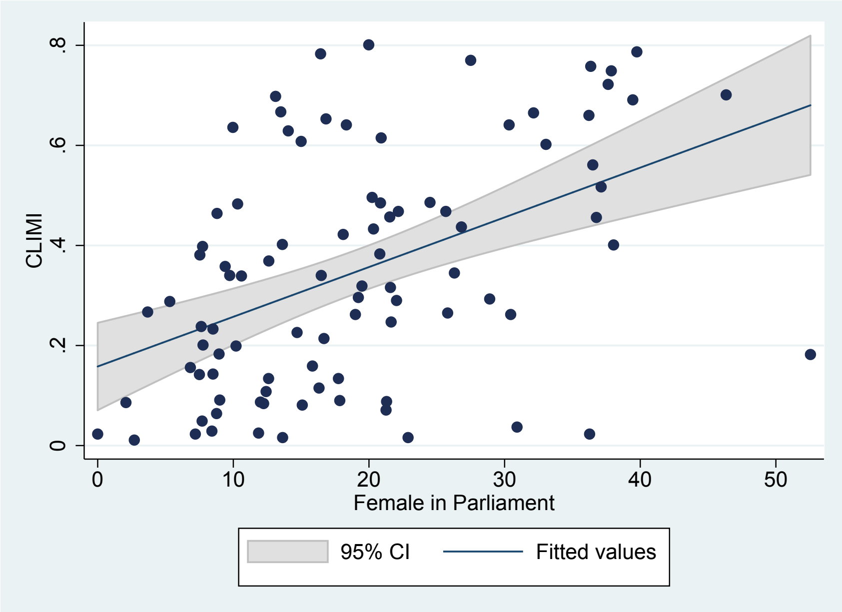

Figure 1: Relationship between CLIMI and Female in Parliament

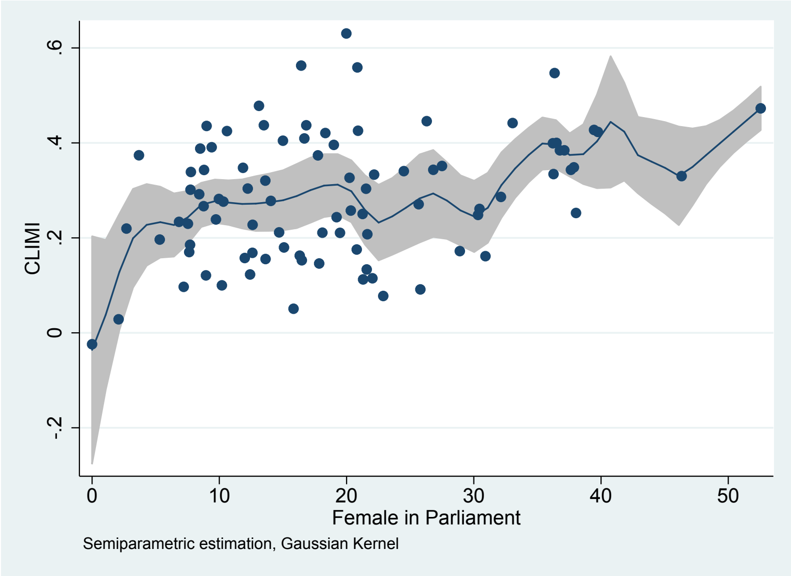

Figure 2: Semi-parametric regression

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Table 1: Summary statistics

Variables	Source	Mean	Std.Dev.
CLIMI	Steves et al. (2011)	0.350	0.235
Female in Parliament (%)	World Bank (2016)	19.250	11.008
GDP per capita (log)†	World Bank (2016)	9.386	1.189
Openness	World Bank (2016)	0.885	0.509
Democracy	Marshall et al. (2016)	0.604	0.492
Autocracy	Marshall et al. (2016)	0.121	0.328
Other regime	Marshall et al. (2016)	0.187	0.392
English legal origin	La Porta et al. (1998)	0.187	0.392
French legal origin	La Porta et al. (1998)	0.396	0.492
German legal origin	La Porta et al. (1998)	0.055	0.229
Scandinavian legal origin	La Porta et al. (1998)	0.055	0.229
Socialist legal origin	La Porta et al. (1998)	0.308	0.464
N		91	

Note.—Variables are defined in section 2. † PPP (constant 2011 international $).

Table 2: Baseline OLS regressions

	All	All	All	All	No outliers
Control variables	(1)	(2)	(3)	(4)	(5)
Female in Parliament (%)	0.010***	0.007***	0.005***	0.004***	0.003***
	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)
GDP per capita (log)†		0.125***	0.095***	0.083***	0.086***
		(0.015)	(0.017)	(0.018)	(0.019)
Openness		-0.046*	0.010	0.021	0.013
		(0.027)	(0.026)	(0.029)	(0.028)
Democracy			0.190***	0.196***	0.195***
			(0.036)	(0.037)	(0.036)
Autocracy			-0.009	0.006	0.024
			(0.061)	(0.063)	(0.066)
English legal origin				-0.176***	-0.157**
				(0.061)	(0.061)
French legal origin				-0.136***	-0.148***
				(0.048)	(0.046)
German legal origin				0.004	-0.005
				(0.038)	(0.036)
Socialist legal origin				-0.129**	-0.139**
				(0.056)	(0.054)
Constant	0.161***	-0.920***	-0.763***	-0.516***	-0.523***
	(0.045)	(0.116)	(0.125)	(0.180)	(0.178)
R^2	0.212	0.574	0.695	0.728	0.744
N	91	91	91	91	88

Note.— Dependent variable is CLIMI. Columns (1)-(4) report the results based on the full sample. Column (5) reports the results in a sample where outliers are removed; these are identified by predicting DFbetas for Female in Parliament (%) from the full sample regression and then dropping those observations for which $|\text{DFbeta}|>2/\sqrt{N}$. Robust standard errors in parentheses.*** p<0.01, ** p<0.05, * p<0.1. † PPP (constant 2011 international $).

Table 3: OLS regressions with additional controls

Control variables	(1)	(2)	(3)	(4)	(5)	(6)
Female in Parliament (%)	0.004***	0.004***	0.003**	0.003**	0.003*	0.003**
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Average years of schooling††		0.021*	0.027**	0.027*	0.021*	0.037**
		(0.011)	(0.011)	(0.013)	(0.012)	(0.014)
Left government			-0.014	-0.014	-0.036	-0.053
			(0.037)	(0.036)	(0.036)	(0.034)
Stock of democracy			0.161*	0.141	0.062	0.133*
			(0.087)	(0.096)	(0.083)	(0.079)
British colony				-0.058	-0.074	-0.034
				(0.110)	(0.101)	(0.094)
French colony				-0.028	-0.048	-0.008
				(0.055)	(0.049)	(0.055)
Oil & Gas net export per capita (1000 US$)					-0.014	-0.014
					(0.009)	(0.009)
Latitude of capital					0.183	-0.061
					(0.126)	(0.115)
Asia						0.061
						(0.068)
Europe						0.086
						(0.085)
North America						-0.074
						(0.076)
Oceania						-0.071
						(0.105)
South America						-0.112
						(0.081)
Constant	-0.516***	-0.366*	-0.321	-0.310	-0.408*	-0.396*
	(0.180)	(0.208)	(0.203)	(0.202)	(0.211)	(0.224)
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.728	0.738	0.749	0.753	0.800	0.829
N	91	80	77	77	76	76

Note.— Dependent variable is CLIMI. Robust standard errors in parentheses.*** p<0.01, ** p<0.05, * p<0.1. †† Total population of 15 years old or above.

Table 4: Quantile regressions

Control variables	Quantiles		
	0.25	0.5	0.75
	(1)	(2)	(3)
Female in Parliament (%)	0.003	0.005**	0.004**
	(0.002)	(0.002)	(0.002)
Baseline controls	Yes	Yes	Yes
N	91		

Note.—Dependent variable is CLIMI. Bootstrapped (50 replications) standard errors in parentheses.*** $p<0.01$, ** $p<0.05$, * $p<0.1$

Table 5: 2SLS regression

Control variables	OLS			2SLS	
				$1^{st} Stage$	$2^{nd} Stage$
	(1)	(2)	(3)	(4)	(5)
Female in Parliament (%)	0.004***		0.004***		0.012**
	(0.001)		(0.001)		(0.005)
Years since suffrage		0.002**	0.002	0.171***	
		(0.001)	(0.001)	-0.054	
Baseline controls	Yes	Yes	Yes	Yes	Yes
R^2	0.728	0.733	0.753		0.643
N	91	89	88	88	88

Note.—Dependent variable is CLIMI. Robust standard errors in parentheses. *** p<0.01, ** p<0.05, * p<0.1.

Table 6: Test of omitted variable bias

Specification	$\delta_{R_{max}=0.90}$	$\left[\tilde{\beta}, \beta^*_{(R_{max}=0.90,\delta=1)}\right]$	$\delta_{R_{max}=min\{1.3\tilde{R},1\}}$	$\left[\tilde{\beta}, \beta^*_{\left(R_{max}=min\{1.3\tilde{R},1\},\delta=1\right)}\right]$
	(1)	(2)	(3)	(4)
BASELINE CONTROL	1.392	[0.001, 0.003]	1.10	[0.0004, 0.003]
COMPREHENSIVE CONTROLS	2.646	[0.002, 0.003]	1.159	[0.0006, 0.003]

Note.— δ indicates the value of proportional selection of unobservables to observable assuming the maximum value of theoretical R^2 is R_{max}. The coefficient bounds are calculated assuming the unobservables are as important as the observable in explaining the outcome variable (i.e. $\delta = 1$).

Table 7: 3SLS regressions

Control variables	$\kappa = 0$		$\kappa \neq 0$	
	(1)	(2)	(3)	(4)
Female in Parliament (%)	0.004***		0.004***	0.001
	(0.001)		(0.001)	(0.043)
CLIMI		-12.153***		-12.304**
		(3.751)		(6.075)
GDP per capita (log) †	0.083***	-14.543***	0.083***	-14.511***
	(0.014)	(3.039)	(0.014)	(3.242)
GDP per capita (log) squared †		1.051***		1.050***
		(0.172)		(0.176)
Population density (log)		-0.346		-0.343
		(0.226)		(0.272)
English legal origin	-0.193***		-0.193***	
	(0.071)		(0.072)	
French legal origin	-0.134**		-0.133**	
	(0.067)		(0.067)	
German legal origin	-0.009		-0.009	
	(0.080)		(0.080)	
Socialist legal origin	-0.147**		-0.147**	
	(0.071)		(0.071)	
Constant	-0.513***	51.741***	-0.513***	51.535***
	(0.152)	(13.545)	(0.153)	(15.375)
Other baseline controls	Yes	Yes	Yes	Yes
R^2	0.727	0.755	0.726	0.754
N	91	91	91	91

Note.—Dependent variable is CLIMI in columns (1) and (3) and CO_2 emissions per capita, metric tones, in columns (2) and (4). κ refers to the coefficient on Female in Parliament in equation 6. Robust standard errors in parentheses.*** p<0.01, ** p<0.05, * p<0.1. † PPP (constant 2011 international $).

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