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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.

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., 2018), 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

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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
CO_2 emissions per capita	World Bank (2016)	5.830	5.327
Democracy	Marshall et al. (2016)	0.637	0.483
Autocracy	Marshall et al. (2016)	0.121	0.328
Other regime	Marshall et al. (2016)	0.242	0.431
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 $).

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 a 6-fold increase in the stringency of the country's climate change policies (in practice Denmark's climate change policies are around 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 and dependent 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 United Nations Framework Convention on Climate Change (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

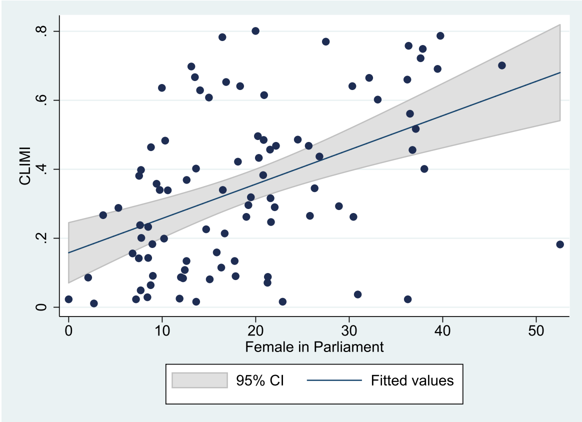

Fig. 1. Relationship between CLIMI and female in parliament.

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 to 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.

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 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%). Fig. 1 plots CLIMI against Female in Parliament revealing a strong positive relationship: countries with higher proportion of females in parliament have considerably more stringent climate change policies.

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. Following Fredriksson and Neumayer (2013), we additionally control for countries' per capita carbon emissions (CO_2 emissions per capita) - a variable that 'reflects the amount at stake for CO_2 emitters and thus their lobbying incentives' (p. 14).[1]

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-point 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, 64% of the countries are democracies while 12% are autocracies (the remaining 24% 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.

[1] The results are not sensitive to dropping this variable from the analysis.

Table 2
Baseline OLS regressions.

Control variables	All (1)	All (2)	All (3)	All (4)	All (5)	No outliers (6)
Female in Parliament (%)	0.010***	0.007***	0.007***	0.005***	0.005***	0.004***
	(0.002)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
GDP per capita (log)[†]		0.125***	0.188***	0.147***	0.138***	0.136***
		(0.015)	(0.017)	(0.021)	(0.023)	(0.024)
Openness		−0.046*	−0.047*	−0.012	−0.012	0.016
		(0.027)	(0.024)	(0.032)	(0.033)	(0.025)
CO_2 emissions per capita			−0.019***	−0.015***	−0.015***	−0.015***
			(0.003)	(0.004)	(0.004)	(0.004)
Democracy				0.166***	0.159***	0.177***
				(0.040)	(0.038)	(0.037)
Autocracy				0.050	0.043	0.074
				(0.058)	(0.058)	(0.056)
English legal origin					−0.056	−0.073
					(0.072)	(0.072)
French legal origin					−0.083	−0.091
					(0.063)	(0.063)
German legal origin					0.071	0.054
					(0.056)	(0.055)
Socialist legal origin					−0.044	−0.075
					(0.067)	(0.066)
Constant	0.161***	−0.920***	−1.393***	−1.150***	−0.999***	−0.978***
	(0.045)	(0.116)	(0.138)	(0.164)	(0.215)	(0.221)
Adjusted R^2	0.203	0.559	0.638	0.701	0.710	0.730
N	91	91	91	91	91	87

Note.— Dependent variable is CLIMI. Columns (1)–(5) report the results based on the full sample. Column (6) 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 $).

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):

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

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 Fig. 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.

Column (3) presents the results with CO_2 emissions per capita added to the list of controls. The coefficient on Female in Parliament is unaffected. CO_2 emissions per capita is negatively correlated with CLIMI. This is consistent with findings in Fredriksson and Neumayer (2013) and suggests that anti-climate change policy lobbying is successfull in places with large CO_2 emissions per capita.

The characteristics of countries' political regime are added as controls in column (4). Reassuringly, the estimated positive significant coefficient on Female in Parliament remains robust to including these controls although it's magnitude is 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 (5) 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 0.05 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).

In column (6) 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 (5) and drop those observations for which

Table 3
Variance Inflation Factors: Baseline regression.

Variables	VIF	1/VIF
Female in Parliament (%)	1.390	0.718
GDP per capita (log)[†]	3.490	0.286
Openness	1.260	0.794
CO_2 emissions per capita	3.490	0.286
Democracy	1.960	0.511
Autocracy	1.730	0.577
English legal origin	5.210	0.153
French legal origin	6.530	0.156
German legal origin	2.070	0.484
Socialist legal origin	6.430	0.156

Note.— [†] PPP (constant 2011 international $).

$|\text{DFbeta}| > 2/\sqrt{N}$ (Belsley et al., 1980). The results are remarkably similar to those reported in column (5) 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.

To assess the robustness of the results with regard to collinearity, we refer to Variance Inflation Factors (VIF) based on the baseline regression in column (5) of Table 2. The VIFs of the control variables reported in Table 3 do not reveal serious collinearity problems; all variables are well below the rule of thumb VIF value of 10 (Marquaridt, 1970; Neter et al., 1989; Kennedy, 1992).

4. Robustness checks

4.1. 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 4. To allow for comparisons, in column (1) we report the estimation results of the baseline model (these are identical to those presented in column (5) 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., 2018). Indeed, we estimate a positive coefficient on average years of schooling (sourced from Barro and Lee (2013)). It is however statistically insignificant in this particular estimation, while the estimated coefficient on Female in Parliament is 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).[2] 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). The estimated coefficients on both are statistically insignificant. The estimate on Female in Parliament is remarkably similar to that from the baseline model.

Another possibility we consider is that countries with preference for female representation may simply have more policies in general. In the absence of a direct measure for the number of policies in a country, we consider a proxy limited to a certain domain: business regulation environments. Doing Business country rankings sourced from the World Bank (2018) cover 10 areas of business regulations across countries - starting a business, dealing with construction permits, getting electricity, registering property, getting credit, protecting minority investors, paying taxes, trading across borders, enforcing contracts and resolving insolvency - and capture both quantity and quality aspects of the rules. Admittedly, this measure is narrowly focused and does not cover the full range of policies and institutions that comprise the regulatory environment in a country. It is perhaps to some extent for that reason, that we estimate an insignificant coefficient on Doing Business ranking (column 4); however it has a negative sign suggesting that complexity of the regulatory environment in this area is negatively correlated with the stringency of climate change policies. We continue to observe a significant positive coefficient on Female in Parliament, while its magnitude is smaller compared to the baseline estimate.

The significance of history in determining contemporary outcomes may manifest through channels other than those hitherto

[2] 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).

Table 4
OLS regressions with additional controls.

Control variables	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Female in Parliament (%)	0.005***	0.005***	0.004***	0.003**	0.003**	0.003**	0.003*	0.003**
	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
Average years of schooling[†]		0.016	0.026**	0.023**	0.022*	0.021*	0.022**	0.042***
		(0.011)	(0.010)	(0.010)	(0.011)	(0.011)	(0.011)	(0.013)
Left government			−0.005	−0.007	−0.007	−0.024	−0.028	−0.056
			(0.040)	(0.040)	(0.040)	(0.039)	(0.039)	(0.038)
Stock of democracy			0.128	0.113	0.099	0.071	0.082	0.173*
			(0.087)	(0.084)	(0.095)	(0.087)	(0.089)	(0.087)
Doing Business ranking				−0.001	−0.001	−0.000	−0.000	−0.001
				(0.001)	(0.001)	(0.001)	(0.001)	(0.001)
British colony					−0.044	−0.065	−0.057	−0.016
					(0.104)	(0.089)	(0.090)	(0.085)
French colony					−0.036	−0.047	−0.037	−0.029
					(0.048)	(0.046)	(0.045)	(0.051)
Oil & Gas net export per capita (1000 US$)						−0.011	−0.014*	−0.013*
						(0.008)	(0.008)	(0.007)
Latitude of capital						0.234**	0.194	−0.101
						(0.111)	(0.118)	(0.103)
Small island							−0.085*	−0.042
							(0.043)	(0.058)
Climate vulnerability							0.001	−0.001
							(0.004)	(0.004)
Asia								−0.002
								(0.063)
Europe								0.045
								(0.090)
North America								−0.107
								(0.084)
Oceania								−0.142
								(0.095)
South America								−0.193**
								(0.082)
Constant	−0.999***	−0.970***	−0.784***	−0.539*	−0.522*	−0.593**	−0.635**	−0.538
	(0.215)	(0.261)	(0.241)	(0.283)	(0.292)	(0.278)	(0.306)	(0.339)
Baseline controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Adjusted R^2	0.710	0.724	0.747	0.750	0.746	0.761	0.758	0.789
N	91	80	77	77	77	77	77	77

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.

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 (5) are from a model that includes dummies to identify former British and French colonies.[3] 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; Mavisakalyan and Tarverdi, 2018). 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 (6) 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 in this specification. We estimate a positive significance coefficient on Latitude of capital, however it is not robust to subsequent alterations to the list of controls.

Further, in column (7) we control for small island states (source: World Bank (2016)) and climate vulnerability (source: Wheeler (2011)) - variables that have been considered in the literature to account for the demand for climate change policies (Fredriksson and Neumayer, 2013). According to our results, small island states actually have less stringent climate change policies, however this estimate is not robust to subsequently controlling for other geographic characteristics of countries. The estimated coefficient on climate vulnerability is insignificant, while we continue to observe positive significant coefficient on Female in Parliament.

Finally, in the last column of Table 4, we include dummies for continents. After controlling for important socio-economic, institutional and geographic characteristics of countries, the only significant effect we estimate is on the dummy for South America: the climate change policies in the countries of this continent are on average less stringent compared to those in African countries.

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

Table 5
OLS regressions with alternative controls.

Control variables	(1)	(2)
Female in Parliament (%)	0.005∗∗∗	0.003∗
	(0.001)	(0.001)
Human Development Index	0.970∗∗∗	0.837∗∗
	(0.146)	(0.361)
Openness	0.003	0.003
	(0.037)	(0.033)
CO_2 emissions per capita	−0.010∗∗∗	−0.013∗∗
	(0.004)	(0.005)
Democracy	0.164∗∗∗	0.088∗
	(0.036)	(0.048)
Autocracy	0.073	0.218∗∗∗
	(0.054)	(0.064)
English legal origin	−0.081	−0.084
	(0.073)	(0.114)
French legal origin	−0.065	−0.034
	(0.061)	(0.067)
German legal origin	0.055	−0.064
	(0.052)	(0.062)
Socialist legal origin	−0.095	−0.119
	(0.062)	(0.076)
Average years of schooling, females†		0.019
		(0.013)
Left largest government party		−0.029
		(0.034)
Stock of democracy		0.200∗∗
		(0.086)
Doing Business ranking		−0.000
		(0.001)
British colony		−0.009
		(0.085)
French colony		−0.036
		(0.050)
Fuel, ores and metals export (% of merchandise exports)		−0.002∗∗∗
		(0.001)
Latitude of capital		−0.183
		(0.126)
Small island		0.000
		(0.052)
Climate vulnerability		−0.003
		(0.003)
Asia		−0.092
		(0.072)
Europe		−0.088
		(0.104)
North America		−0.267∗∗
		(0.104)
Oceania		−0.292∗∗
		(0.109)
South America		−0.251∗∗∗
		(0.091)
Constant	−0.427∗∗∗	−0.225
	(0.119)	(0.215)
Baseline controls	Yes	Yes
Adjusted R^2	0.722	0.812
N	91	74

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.

The estimated positive effect of Female in Parliament on CLIMI persists after these characteristics of countries have been accounted for.

As another attempt to test the robustness of our central result on the link between Female in Parliament and CLIMI to the choice of control variables included in the model, we make further alterations to the list of controls. In Table 5, we report the results of the regressions with baseline and additional lists of controls where we replace: (i) GDP per capita with the Human Development Index by the United Nations Development Programme (2018); (ii) average years of schooling in the total population with the average years of schooling of the female population World Bank (2016); (iii) the political (left-wing) orientation of national leader's party with the orientation of the largest government party (Keefer, 2012); and (iv) natural resource export per capita with natural resource export as a share of total merchandise export World Bank (2016).

Table 6
OLS regressions with separate policy areas of CLIMI as dependent variables.

Control variables	Domestic framework (1)	Sectoral measures (2)	Cross-sectoral measures (3)	International cooperation (4)
Female in Parliament (%)	0.002***	0.002***	0.000	0.000
	(0.001)	(0.001)	(0.000)	(0.000)
Baseline controls	Yes	Yes	Yes	Yes
Adjusted R^2	0.605	0.624	0.412	0.074
N	91	91	91	91

Note.— Dependent variables are the separate policy areas of CLIMI. Robust standard errors in parentheses. ***$p < 0.01$, **$p < 0.05$, *$p < 0.1$.

The estimated coefficients on Female in Parliament are identical to those obtained from the models with initial specifications of baseline and additional controls. There is a highly significant positive relationship between a country's Human Development Index and the stringency of its climate change policies. The estimates on the average years of schooling of the female population and the left-wing orientation of the largest government party are insignificant. But we do estimate a significant negative coefficient on the fuel, ores and metals export as a share of merchandise exports: there is a decline in CLIMI with an increase in this share.

4.2. Alternative dependent variables

Our estimates of the relationship between female representation in parliament and climate change policies are remarkably robust to the choice and definition of control variables in the model. Are they also robust to the way we measure the countries' climate change policies? As a first attempt at addressing this question, we consider the four areas of CLIMI - domestic climate framework; sectoral fiscal or regulatory measures; additional cross-sectoral fiscal or regulatory measures; and international cooperation-as separate dependent variables in Table 6. While the first two policy areas each enter with the weights of 0.4 in the construction of the overall index, the contribution of the two later policy areas is limited to the weights of 0.1 each. Consistent with this, we find that the positive significant effect of Female in Parliament on CLIMI might be driven by its effect on domestic framework and sectoral measures - the coefficients on Female in Parliament in the models where the measures for these policy areas are used as the dependent variable are positive and significant. The relationship of Female in Parliament with cross-sectoral measures or with international cooperation is indistinguishable from zero.

In addition to looking at individual policy areas of CLIMI, we consider alternative measures of countries' environmental policies. A number of such measures exist, with data spanning several years which result in unbalanced panel datasets. We study these next, in a panel fixed effects estimation framework which employs Female in Parliament and other time-variant characteristics of countries as regressors.[4]

We start with OECD's Environmental Policy Stringency Index (EPS) - a measure that captures the degree of stringency of 14 environmental policy instruments, primarily related to climate and air pollution (Botta and Kozluk, 2014). The data is available for the period from 1990 to 2012, however it is limited to 27 OECD and 6 BRICS countries.[5] The results presented in the first column of Table 7 confirm the positive significant relationship between Female in Parliament and this measure of environmental policies. Consistent with the baseline results, we also observe a positive significant relationship between a country's per capita GDP and the stringency of its environmental policies, while the effect of per capita CO_2 emissions is negative.

We consider another policy stringency measure - the Climate Policy Index (CPI) developed by Künkel et al. (2006). This index measures climate policy performance and draws on 'quantitative and qualitative information on national climate policies by sector' for three time periods: 1992, 1997 and 2005. The data is available for 24 countries only. Due to unavailability of data on females in parliaments across countries in 1992, our analysis is restricted to the years 1997 and 2005. The estimated coefficient on Female in Parliament is statistically insignificant, unsurprisingly so, given the sample size (column 2).[6] It should also be noted, however, that this analysis precedes that on CLIMI, and changes in climate policies across a number of countries over time may also explain the differences in the results (Surminski and Williamson, 2012).

Finally, we draw on data from Germanwatch to utilise the climate policy ingredients of their Climate Change Performance Index (CCPI) as our outcome measures. The index evaluates the climate protection performance of 58 countries that are, together, responsible for more than 90 percent of global energy-related CO_2 emissions (Burck et al., 2014). The climate policy components of the overall index are based on assessments by country experts on the extent and quality of each country's commitment to climate policies and regulations. There are particularly sharp differences in the measurement criteria used in the construction of these policy ingredients of CCPI and CLIMI, with the correlation coefficient between the overall policy ingredient and CLIMI of 0.19 (Surminski and Williamson, 2012). Data is available for national and international policies from year 2007 onwards, however our analysis is restricted to the period up to 2014 since no data on per capita CO_2 emissions utilised in the model is available from then on. The results using these policy measures as dependent variables are reported in columns (3) and (4) of Table 7. Female in Parliament is positively significantly correlated with the international policy ingredient of CCPI, however, the estimated coefficient of Female in

[4] To increase variability over time, Polity IV scores rather than dummy variables to distinguish between democracies, autocracies and other regimes are employed in this analysis.

[5] BRICS refers to the association of five major emerging national economies: Brazil, Russia, India, China and South Africa.

[6] Due to lack of variation over the two time periods within the sample of countries considered here, Polity IV score is dropped from the estimation.

Table 7
OLS panel fixed effects regressions with alternative dependent variables.

Control variables	EPS (1)	CPI (2)	CCPI-national policy (3)	CCPI-international policy (4)
Female in Parliament (%)	0.052***	0.009	−0.004	0.005**
	(0.012)	(0.008)	(0.003)	(0.003)
GDP per capita (log)†	1.706***	−0.172	0.306*	0.165
	(0.588)	(0.413)	(0.174)	(0.176)
Openness	0.009	−0.003	0.000	−0.001
	(0.005)	(0.003)	(0.001)	(0.001)
Polity IV score	−0.018		0.015**	0.028***
	(0.019)		(0.006)	(0.005)
CO_2 emissions per capita	−0.205**	0.043	0.004	0.020
	(0.084)	(0.035)	(0.018)	(0.013)
Constant	−15.408***	3.922	−2.679	−1.514
	(5.273)	(4.142)	(1.706)	(1.761)
Adjusted R^2	0.505	−0.011	0.010	0.046
N	526	48	433	433
Countries	33	24	55	55

Note.— Dependent variables are: Environmental Policy Stringency Index (EPS) (OECD, 2016) in column 1; Climate Policy Index (CPI) from Künkel et al. (2006) in column (2), national and international policy components of Germanwatch's Climate Change Performance Index (CCPI) in columns (3) and (4). Robust standard errors in parentheses. ***p < 0.01, **p < 0.05, *p < 0.1. † PPP (constant 2011 international $).

Table 8
Quantile regressions.

Control variables	Quantiles		
	0.25 (1)	0.5 (2)	0.75 (3)
Female in Parliament (%)	0.004	0.005***	0.004*
	(0.003)	(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.

Parliament in the regressions of the national policy ingredient of the index is insignificant.

4.3. Non-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 in Parliament 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:

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

In equation (2), as an arbitrary combination, δ can be 0.25, 0.50, 0.75 quantiles. The relevant results are reported in Table 8. 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.

We further explore the relationship between Female in Parliament and the stringency of climate change policies across countries when some of the functional form assumptions are relaxed. The results we have documented so far are based on the assumption of a linear function relating $FemParliament_i$ to $CLIMI$. Relaxing this functional form assumption means the main model in equation (1), now changes to:

$$CLIMI_i = \theta X_i + f\,(FemParliament_i) + \upsilon_i \text{ for all } i = 1, \ldots, N. \quad (3)$$

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. Fig. 2 illustrates the estimates

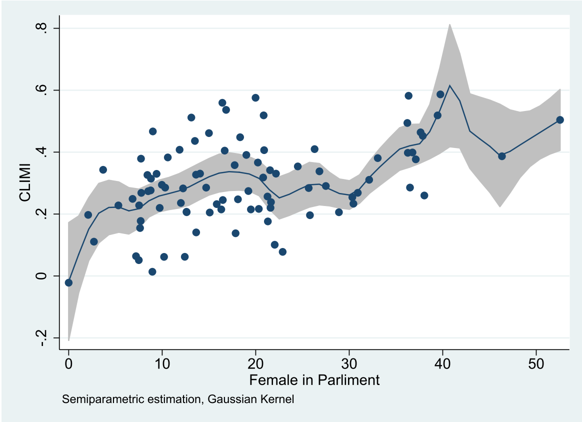

Fig. 2. Semiparametric regression.

of $f(.)$ and as shown, the effect is always positive and within the confidence intervals. [7]

5. Addressing endogeneity

5.1. Employing an instrument

The previous sections have established a rather 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 9, 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 (5) 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.

The 2SLS estimates of equation (1) are presented in columns (4) and (5) of Table 9. Female in Parliament is treated as endogenous and modeled as:

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

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.[8] In defense of our identification strategy, the statistics for the weak identification test (Kleibergen and Paap (2006) F statistics) provides no evidence that the instrument we use is weak or irrelevant.

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 0.12 in CLIMI. This estimate implies that the 35.5 unit difference in female representation between two typical

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

[8] We additionally explored the possible non-linearity in the relationship between Years since suffrage and Female in Parliament by including a squared-term of Years since Suffrage in the estimation of equation (4). The estimated coefficient on this term was statistically insignificant. The results are available on request.

Table 9
2SLS regression.

Control variables	OLS			2SLS	
	(1)	(2)	(3)	1st*Stage* (4)	2nd*Stage* (5)
Female in Parliament (%)	0.005***		0.005***		0.012***
	(0.001)		(0.001)		(0.004)
Years since suffrage		0.002***	0.001	0.190***	
		(0.001)	(0.001)	(0.053)	
GDP per capita (log)[†]	0.138***	0.116***	0.127***	−2.666	0.147***
	(0.023)	(0.024)	(0.025)	(1.796)	(0.026)
Openness	−0.012	0.004	−0.006	2.433	−0.025
	(0.033)	(0.041)	(0.034)	(2.327)	(0.024)
CO_2 emissions per capita	−0.015***	−0.013***	−0.015***	0.439	−0.018***
	(0.004)	(0.005)	(0.004)	(0.324)	(0.005)
Democracy	0.159***	0.170***	0.151***	3.796	0.122***
	(0.038)	(0.038)	(0.037)	(2.486)	(0.044)
Autocracy	0.043	0.095	0.073	5.034	0.035
	(0.058)	(0.063)	(0.061)	(3.675)	(0.058)
English legal origin	−0.056	−0.132**	−0.033	−20.871***	0.125
	(0.072)	(0.062)	(0.069)	(3.855)	(0.132)
French legal origin	−0.083	−0.133**	−0.062	−14.616***	0.048
	(0.063)	(0.057)	(0.060)	(3.398)	(0.096)
German legal origin	0.071	0.042	0.102*	−12.298***	0.196*
	(0.056)	(0.052)	(0.061)	(4.628)	(0.105)
Socialist legal origin	−0.044	−0.170***	−0.059	−24.178***	0.124
	(0.067)	(0.054)	(0.066)	(3.290)	(0.112)
Constant	−0.999***	−0.837***	−1.018***	39.259**	−1.316***
	(0.215)	(0.202)	(0.209)	(16.840)	(0.292)
Adjusted R^2	0.710	0.705	0.738		0.649
Partial R^2				0.101	
Kleibergen and Paap (2006) F statistics[††]				12.72	
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..[†] PPP (constant 2011 international $). [††] The H_o is that the first stage equation is weakly identified.

Table 10
2SLS panel fixed effects regression.

Control variables	OLS	2SLS	
	(1)	1st*Stage* (2)	2nd*Stage* (3)
Female in Parliament (%)	0.052***		0.159***(0.033)
	(0.012)		
Years since suffrage		0.745***	
		(0.113)	
Baseline controls	Yes	Yes	Yes
Adjusted R^2	0.505		0.225
Partial R^2		0.409	
Kleibergen and Paap (2006) F statistics[†]		43.56	
Countries	33	30	30

Note.- Dependent variable is Environmental Policy Stringency Index (EPS) (OECD, 2016). The list of baseline controls is identical to those included in Table 7. Robust standard errors in parentheses. ***p < 0.01, **p < 0.05, *p < 0.1. [†] The H_o is that the first stage equation is weakly identified.

countries with high and low representation, Bahrain and Denmark, would translate into 6-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.

An IV estimation can be potentially more convincing in a dataset that varies over time. Since this is not feasible for CLIMI, a cross-sectional measure, we employ EPS, the measure of environmental policy stringency from OECD (2016) introduced earlier, in a 2SLS panel fixed effects regression. The results are reported in Table 10. The baseline OLS results are reported in column (1) (these are identical to those reported in column (1) of Table 7). We confirm the significant explanatory power of Years since suffrage in explaining Female in Parliament in the first stage results in column (2). The statistically significant relationship between Female in Parliament and EPS persists in the second stage.

Table 11
Test of omitted variable bias.

Specification	$\delta_{R_{\max}=0.90}$ (1)	$\left[\widetilde{\beta}, \beta^*_{(R_{\max}=0.90,\delta=1)}\right]$ (2)	$\delta_{R_{\max}=\min\{1.3\widetilde{R},1\}}$ (3)	$\left[\widetilde{\beta}, \beta^*_{(R_{\max}=\min\{1.3\widetilde{R},1\},\delta=1)}\right]$ (4)
Baseline Control	2.07	[0.003, 0.005]	1.725	[0.002, 0.005]
Comprehensive Controls	3.288	[0.002, 0.003]	1.132	[0.0004, 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. $\delta = 1$).

5.2. *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):

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

where $\dot{\beta}$ and $\dot{R}$ are the coefficient and the R-squared from estimating the equation (1) where $\alpha = 0$ and $\widetilde{\beta}$ and $\widetilde{R}$ 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 $\widetilde{\beta}$ to β^* estimated under an assumption of $\delta = 1$, and given values of $R_{\max} \in [\widetilde{R}, 1]$. 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) $R_{\max} = \min\{1.3\widetilde{R}, 1\}$ - the rule of thumb proposed by Oster (2016).

The results of this analysis are presented in Table 11. In columns (2) and (4) we report the coefficient bounds $[\widetilde{\beta}, \beta^*]$ for models with baseline and comprehensive lists of controls (corresponding to those reported in column (5) of Table 2 and column (8) of Table 4). $\widetilde{\beta}$ 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, $[\widetilde{\beta}, \beta^*]$ 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.

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

Table 12
3SLS regressions.

Control variables	Female in Parliament (%) (1)	CLIMI (2)	CO_2 emissions per capita (3)
Female in Parliament (%)		0.018∗∗∗(0.004)	
CLIMI			−13.084∗∗∗(3.660)
Years since suffrage	0.155∗∗ (0.062)		
Other baseline controls[†]	Yes	Yes	Yes
R^2	0.285	0.514	0.781
N	88	88	88

Note.— Dependent variables are: Female in Parliament (%) in column 1; CLIMI in column (2); and CO_2 emissions per capita in column (3). [†] The lists of other controls included in models (1) and (2) are identical to those included in 2SLS estimations reported in columns (4) and (5) of Table 9. The list of other controls in model (3) includes controls for GDP per capite squared-term and population density in addition to the controls reported in column (5) of Table 9 but with the exception of Female in Parliament, COs emissions per capita and dummies for legal origin. Robust standard errors in parentheses. ∗∗∗p < 0.01, ∗∗p < 0.05, ∗p < 0.1. The overall adjusted adjusted R^2 for the system is 0.984. The F-test of joint significance of all excluded variables in the system is 71.61.

result in lower carbon dioxide emissions? To explore these relationships, we estimate the equations (1) and (4) jointly with the following equation determining carbon dioxide emissions, using 3SLS:

$$CO2_i = \eta C_i + \mu D_i + \lambda CLIMI_i + \iota_i \text{ for all } i = 1, \ldots, N. \tag{6}$$

where $CO2_i$ is the per capita carbon emissions of a country, C_i is a sub-set of control variables included in X_i in equation (1). In particular, for identification purposes, the measures of countries' legal origins are included in X_i but not in C_i, i.e. we assume that legal origins might affect $CO2$ emissions only through their effect on countries' environmental policies. Another exclusion restriction from the model is that Female in Parliament does not affect $CO2$ emissions once CLIMI is taken into account. Similar approach has been adopted in other studies that focus on a single potential transmission mechanism at a time (e.g., Sylwester, 2000; Tavares and Wacziarg, 2001; Mavisakalyan, 2011). D_i contains two additional determinants of $CO2$ that are excluded from equations (1) and (4): a squared term of GDP per capita and population density; ι_i is a disturbance term.

The results of estimating equations (1), (4) and (6) simultaneously using 3SLS are summarised in Table 12. The coefficients of interest are β in equation (1) and λ in equation (6): 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 (2) and (3), a 10 unit increase in female representation, through its effect on the stringency of climate change policies, results in 0.24 (=0.018∗13.084) metric tones decrease of carbon dioxide emissions per capita.

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.

References

Adams, R.B., Funk, P., 2012. Beyond the glass ceiling: does gender matter? Manag. Sci. 58 (2), 219–235.
Ang, J.B., Fredriksson, P.G., 2017. Statehood experience, legal traditions, and climate change policies. Econ. Inq. 55 (3), 1511–1537.
Austen, S., Mavisakalyan, A., 2016. Constitutions and the political agency of women: a cross-country study. Fem. Econ. 22 (1), 183–210.
Barro, R.J., Lee, J.W., 2013. A new data set of educational attainment in the world, 1950-2010. J. Dev. Econ. 104 (Suppl. C), 184–198.
Beck, T., Clarke, G., Groff, A., Keefer, P., Walsh, P., 2001. New tools in comparative political economy: the database of political institutions. World Bank Econ. Rev. 15 (1), 165–176.
Beer, C., 2009. Democracy and gender equality. Stud. Comp. Int. Dev. 44 (3), 212–227.
Belsley, D.A., Kuh, E., Welsch, R.E., 1980. Regression Diagnostics. Wiley and Sons, New York, USA.
Besley, T., Coate, S., 1997. An economic model of representative democracy. Q. J. Econ. 112 (1), 85–114.
Beutel, A.M., Marini, M.M., 1995. Gender and values. Am. Socio. Rev. 60 (3), 436–448.
Bhalotra, S., Clots-Figueras, I., 2014. Health and the political agency of women. Am. Econ. J. Econ. Pol. 6 (2), 164–197.
Botta, E., Kozluk, T., 2014. Measuring Environmental Policy Stringency in Oecd Countries, p. 1177.
Burck, J., Marten, F., Bals, C., 2014. The Climate Change Performance Index Results 2014. Germanwatch and CAN.
Chattopadhyay, R., Duflo, E., 2004. Women as policy makers: evidence from a randomized policy experiment in India. Econometrica 72 (5), 1409–1443.
Clots-Figueras, I., 2011. Women in politics: evidence from the indian states. J. Publ. Econ. 95 (7), 664–690.
Clots-Figueras, I., 2012, 01. Are female leaders good for education? Evidence from India. Am. Econ. J. Appl. Econ. 4 (1), 212–244.
Dollar, D., Fisman, R., Gatti, R., 2001. Are women really the 'fairer' sex? Corruption and women in government. J. Econ. Behav. Organ. 46 (4), 423–429.
Duflo, E., 2012. Women empowerment and economic development. J. Econ. Lit. 50 (4), 1051–1079.
Farzin, Y.H., Bond, C.A., 2006. Democracy and environmental quality. J. Dev. Econ. 81 (1), 213–235.
Franzen, A., Meyer, R., 2010. Environmental attitudes in cross-national perspective: a multilevel analysis of the ISSP 1993 and 2000. Eur. Socio Rev. 26 (2), 219–234.
Fredriksson, P.G., Neumayer, E., 2013. Democracy and climate change policies: is history important? Ecol. Econ. 95, 11–19.
Fredriksson, P.G., Neumayer, E., 2016. Corruption and climate change policies: do the bad old days matter? Environ. Resour. Econ. 63 (2), 451–469.
Fredriksson, P.G., Wang, L., 2011. Sex and environmental policy in the U.S. house of representatives. Econ. Lett. 113 (3), 228–230.
Fredriksson, P.G., Wollscheid, J.R., 2015. Legal origins and climate change policies in former colonies. Environ. Resour. Econ. 62 (2), 309–327.
Gilligan, C., 1982. In a Different Voice. Harvard University Press.
Grier, R., Maldonado, B., 2015. Electoral experience, institutional quality and economic development in Latin America. Oxf. Dev. Stud. 43 (2), 253–280.
Hall, R.E., Jones, C.I., 1999. Why do some countries produce so much more output per worker than others? Q. J. Econ. 114 (1), 83–116.
Hicks, D.L., Hicks, J.H., Maldonado, B., 2015. Are female politicians more responsive to international crises? Appl. Econ. Lett. 22 (6), 493–498.
Hicks, D.L., Hicks, J.H., Maldonado, B., 2016. Women as policy makers and donors: female legislators and foreign aid. Eur. J. Polit. Econ. 41, 46–60.
Jayasuriya, D.S., Burke, P.J., 2013. Female parliamentarians and economic growth: evidence from a large panel. Appl. Econ. Lett. 20 (3), 304–307.

Keefer, P., 2012. Database of Political Institutions 2012. Computer file.
Kennedy, P., 1992. A Guide to Econometrics. Blackwell, Oxford.
Kleibergen, F., Paap, R., 2006. Generalized reduced rank tests using the singular value decomposition. J. Econom. 133 (1), 97–126.
Künkel, N., Jacob, K., Busch, P.-O., 2006. Climate Policies - (The Feasibility of) a Statistical Analysis of Their Determinants. Human Dimensions of Global Environmental Change, Berlin.
La Porta, R., Lopez-de Silanes, F., Shleifer, A., Vishny, R.W., 1998. Law and finance. J. Polit. Econ. 106 (6), 1113–1155.
La Porta, R., Lopez-de Silanes, F., Shleifer, A., 2008. The economic consequences of legal origins. J. Econ. Lit. 46 (2), 285–332.
Marchand, S., 2012. Legal origin, colonial origins and deforestation. Econ. Bull. 32 (2), 1653–1670.
Marquaridt, D.W., 1970. Generalized inverses, ridge regression, biased linear estimation, and nonlinear estimation. Technometrics 12 (3), 591–612.
Marshall, M.G., Gurr, T.R., Jaggers, K., 2016. Polity IV Dataset. Computer file.
Mavisakalyan, A., 2011. Immigration, public education spending, and private schooling. South. Econ. J. 78 (2), 397–423.
Mavisakalyan, A., 2014. Women in cabinet and public health spending: evidence across countries. Econ. Govern. 15 (3), 281–304.
Mavisakalyan, A., 2015. Gender in language and gender in employment. Oxf. Dev. Stud. 43 (4), 403–424.
Mavisakalyan, A., Tarverdi, Y., 2018. Oil and women: a re-examination. Energy Econ., https://doi.org/10.1016/j.eneco.2018.01.015.
Mavisakalyan, A., Tarverdi, Y., Weber, C., 2018. Talking in the present, caring for the future: language and environment. J. Comp. Econ.
McCright, A.M., 2010. The effects of gender on climate change knowledge and concern in the American public. Popul. Environ. 32 (1), 66–87.
McCright, A.M., Dunlap, R.E., 2011. Cool dudes: the denial of climate change among conservative white males in the United States. Global Environ. Change 21 (4), 1163–1172.
Murdoch, J.C., Sandler, T., 1997. The voluntary provision of a pure public good: the case of reduced CFC emissions and the Montreal Protocol. J. Publ. Econ. 63 (3), 331–349.
Neter, J., Wasserman, W., Kutner, M. H. l., 1989. Applied Linear Regression Models. Homewood, IL.
Neumayer, E., 2002. Does trade openness promote multilateral environmental cooperation? World Econ. 25 (6), 815–832.
Neumayer, E., 2003. Are left-wing party strength and corporatism good for the environment? evidence from panel analysis of air pollution in OECD countries. Ecol. Econ. 45 (2), 203–220.
Neumayer, E., 2004. The environment, left-wing political orientation and ecological economics. Ecol. Econ. 51 (3), 167–175.
OECD, 2016. Environmental Policy Stringency Index. Computer file.
Osborne, M.J., Slivinski, A., 1996. A model of political competition with citizen-candidates. Q. J. Econ. 111 (1), 65–96.
Oster, E., 2016. Unobservable selection and coefficient stability: theory and evidence. J. Bus. Econ. Stat. 0 (0), 1–18.
Paxton, P., Hughes, M.M., Painter, M.A., 2010. Growth in women's political representation: a longitudinal exploration of democracy, electoral system and gender quotas. Eur. J. Polit. Res. 49 (1), 25–52.
Rodrik, D., Subramanian, A., Trebbi, F., 2004. Institutions rule: the primacy of institutions over geography and integration in economic development. J. Econ. Growth 9 (2), 131–165.
Ross, M.L., 2008. Oil, islam, and women. Am. Polit. Sci. Rev. 102 (1), 107–123.
Ross, M., Mahdavi, P., 2015. Oil and Gas Data, pp. 1932–2014.
Seager, J., Bechtel, J., Bock, S., Dankelman, I., Fordham, M., Gabizon, S., Thuy Trang, N., Perch, L., Qayum, S., Roehr, U., et al., 2016. Global Gender and Environment Outlook. United Nations Environment Programme (UNEP).
Spitzner, M., 2009. How global warming is gendered: a view from the EU. In: Salleh, A. (Ed.), Eco-sufficiency & Global Justice. Pluto Press, pp. 218–229.
Stern, N., 2008. The economics of climate change. Am. Econ. Rev. 98 (2), 1–37.
Steves, F., Treisman, D., Teytelboym, A., 2011. The political economy of climate change policy in the transition region. In: The Low Carbon Transition. EBRD and the Grantham Research Institute on Climate Change and the Environment, LSE.
Sundstrom, A., McCright, A.M., 2014. Gender differences in environmental concern among Swedish citizens and politicians. Environ. Polit. 23 (6), 1082–1095.
Surminski, S., Williamson, A., 2012. Policy Indexes? what Do They Tell Us and what Are Their Applications? the Case of Climate Policy and Business Planning in Emerging Markets. Centre for Climate Change Economics and Policy Working Paper 101.
Svaleryd, H., 2009. Women's representation and public spending. Eur. J. Polit. Econ. 25 (2), 186–198.
Swamy, A., Knack, S., Lee, Y., Azfar, O., 2001. Gender and corruption. J. Dev. Econ. 64 (1), 25–55.
Sylwester, K., 2000. Income inequality, education expenditures, and growth. J. Dev. Econ. 63 (2), 379–398.
Tavares, J., Wacziarg, R., 2001. How democracy affects growth. Eur. Econ. Rev. 45 (8), 1341–1378.
United Nations Development Programme, 2018. Human Development Data (1990-2015). Computer file.
Wheeler, D., 2011. Quantifying Vulnerability to Climate Change: Implications for Adaptation Assistance. Center for Global Development Working Paper 240.
Women, U.N., 2011. Progress of the World's Women: in Pursuit of Justice. UN Women.
World Bank, 2016. World Development Indicators Database. Computer file.
World Bank, 2018. Doing Business. Computer file.