IFN Working Paper No. 1332, 2020

Gender and Climate Action

Niklas Elert and Erik Lundin

Research Institute of Industrial Economics
P.O. Box 55665
SE-102 15 Stockholm, Sweden
info@ifn.se
www.ifn.se

Gender and Climate Action

March 1, 2022

Niklas Elert* and Erik Lundin**

Abstract

It is well-known that men and women differ in their views regarding the severity of climate change, but do they also differ in their support for climate policy and in undertaking climate action in their everyday lives? Previous survey evidence on these questions is inconclusive, but we can answer them using unique survey data from the Swedish Environmental Protection Agency (SEPA). Regression analysis confirms that Swedish women believe more strongly than men that climate change will affect Sweden. Further, women engage in more climate-mitigating behavior than men, even conditional on climate beliefs. The association between gender and climate policy support is less robust, and disappears altogether when climate beliefs are controlled for, demonstrating that climate beliefs is the main mechanism explaining the relationship between gender and policy support.

Keywords: Climate change; public opinion; gender; environmental beliefs

JEL-codes: H23; O44; Q54; Q58; J16

* Corresponding author. Research Institute of Industrial Economics (IFN), Box 55665, SE-102 15, Stockholm, Sweden. E-mail: niklas.elert@ifn.se.

** IFN. E-mail: erik.lundin@ifn.se

Acknowledgments: We are grateful for data from the Swedish Environmental Protection Agency, and valuable comments from Carl-Magnus Bjuggren, Emma Heikensten, and Magnus Henrekson. Financial support from Jan Wallanders och Tom Hedelius Stiftelse, the Marianne and Marcus Wallenberg Foundation, and the Swedish Energy Agency is gratefully acknowledged.

1. Introduction

There is considerable support for the argument that gender (i.e., the socially constructed roles, behaviors, activities, and attributes that a society considers appropriate for men and women) affects pro-environmental behaviors via mediating factors such as pro-environmental values and attitudes (Shiva and Mies, 2014; Bell and Braun, 2010; Ergas et al., 2021). At first glance, this also appears to be the case for attitudes and behavior with respect to climate change, increasingly seen as a (literally) man-made problem with negative effects that dis-proportionately harm women (Vilagrasa 2002). Survey research shows a small but consistent gender gap in climate change opinions; in a literature survey examining patterns of survey research from around the world, McCright et al. (2016a) rank gender as the third most consistent predictor of climate change beliefs and concerns after pro-environmentalism and left-wing political orientation. However, while climate perceptions likely influence adaptation processes and potential mitigation measures (Maddison 2006), the extant climate change survey literature only occasionally includes gender as a variable to examine individual-level "support for climate policy and pro-climate behavioral intentions, and [gender's] performance in such models is not consistent" (McCright et al. 2016a, p. 182). Thus, extant survey evidence casts doubt on a gender effect on climate policy support and mitigation behavior, but this result needs further exploration, as gender is here usually included only as a sociodemographic control. This dearth of knowledge on the relationship between gender and climate policy support and behavior is troublesome; after all, political action is likely crucial to meet ambitious climate targets (Ergas et al. 2021; IPCC 2021), and everyday mitigation behavior may also play an important part in reaching such goals (Vandenbergh and Steinemann 2007; Barkenbus 2010; European Commission 2011).

The purpose of this study is to shed further light on the link between climate beliefs, policy support, and mitigation behavior. To our knowledge, the relationship between gender and these three climate outcomes has scarcely been studied at the same time before in the survey literature. This is problematic, because gender differences in climate beliefs may, in and of themselves, explain (a part of the) differences in both policy support and mitigation behavior. Thus, we use survey data to examine three hypotheses, specifically whether systematic gender differences exist in climate beliefs (H1), in support for climate mitigation policies (H2), and in everyday climate mitigation behavior (H3). Moreover, when examining the second and third hypothesis, we take care to examine whether climate beliefs affect these relationships, and whether there is a gender difference in climate policy support and mitigation behavior independent of gender difference in climate beliefs.

To test the hypotheses, we employ survey data from the Swedish Environmental Agency (SEPA), gathered from an online questionnaire in May 2018. The population was chosen to represent the Swedish national average with respect to sex, age, and residential location. Results from OLS and negative binomial regressions corroborate H1, showing that Swedish women hold stronger climate beliefs than men. By contrast, support for H2 is weaker, as the association between gender and support for political interventions to reduce climate change (taxes, subsidies, and information campaigns) disappears when climate beliefs are controlled for. However, even when climate beliefs are controlled for, women engage in more everyday climate mitigation behavior, essentially corroborating H3. The results shed new light on the link between gender, policy support, and mitigation behavior, suggesting that while gender differences in policy support accrue entirely to belief differences, women will engage in more mitigation behavior even compared to men holding similar climate beliefs.

2. Background

2.1 Literature review

Researchers increasingly explore how men and women relate to and understand the environment differently (e.g., Stern et al. 1993; McCright 2010; McKinney and Fulkerson 2015). Studies consistently find that women express stronger concern about environmental issues (e.g., Bord and O'Connor 1997; Davidson and Freudenberg 1996; Stern et al. 1999; Dietz et al. 2002; Xiao and McCright 2015) and engage in more pro-environmental behaviors such as recycling and green consumtion than men (e.g., Hunter et al. 2004; Pisano and Lubell 2017; Xiao and McCright 2014; Li et al. 2019). Women also make up a large majority of leadership and membership in most environmental justice organizations (e.g., Bell 2013; Bell and Braun 2010), and tend, in positions of power, to promote more environmental stewardship (Norgaard and York 2005; McKinney & Fulkerson 2015; Austin & McKinney 2016; Ergas et al. 2021). The ecofeminist tradition argues that these differences result from historic structure, cultural traditions, and social forces positioning women as, e.g., caregivers, subsistence providers, and being generally more altruistic than men

(Shiva and Mies, 2014; Steg 2016; Terry 2009; Warren 1990; Bell and Braun 2010). Women also appear more risk-averse than men (Bord and O'Connor, 1997; Byrnes et al., 1999), while men have more confidence regarding their ability and performance, although such gender differences are most pronounced for tasks judged as masculine (Lundeberg et al. 1994; Barber and O'Dean 2001). This seems to be the case also for pro-environmental behaviors, as they can align with traditional feminine *or* masculine roles (Swim et al. 2019).

Climate change is perhaps the most urgent environmental issue in the 21st century. Climate change perceptions are essential to document because such perceptions influence adaptation processes and potential measures to adapt to or reduce climate change effects (Maddison 2006; Semenza et al. 2008; Debono et al. 2010; Hansen et al. 2012). Yet while survey-based studies document a consistent gender gap in climate change opinions, survey evidence is inconsistent on gender's relationship with policy support and mitigation behavior – two crucial vehicles to overcome climate challenges. This is evident from McCright et al.'s (2016a) examination of survey research from around the world encompassing examine 87 studies, 62 of which use data from the United States. In 26 of the 34 covered studies with a gender focus, women report greater concern about climate change than do men (18 out of 21 US studies find the same gender gap). By contrast, eight out of 16 studies (five out of 12 US studies) examining the relationship between gender and climate policy support report insignificant results (see e.g., Krosnick et al. 2006; O'Connor et al. 2002). Whereas some studies find women to be more supportive of climate policies (see, e.g., Zahran et al. 2006), others report the opposite relationship (see, e.g., Leiserowitz 2007; O'Connor et al. 1999). Likewise, among the seven survey studies examining the relationship between gender and pro-climate behavioral intentions or behavior, five report insignificant results (three out of four US studies) (McCright et al. 2016a).

Part of the explanation for these inconclusive results may be, as McCright et al. (2016a) point out, that gender is included less often in models explaining support for climate policy and pro-climate behavioral intentions. Without exception, moreover, survey studies include gender as a single-item socio-demographic control, without much interpretation or theorizing, and McCright et al. (2016a) argue for the need to fully integrate recent theorizing from gender scholarship while employing more direct measures of gender or gender identity. Another part of the problem is likely that few studies focus on all three relationships simultaneously. Indeed, the only study (McCright et al. 2016b) to examine gender's effect on climate beliefs, support for climate policy, and behavioral intentions at the same time does not focus on these relationships. Covering European Union data, the study finds that women have greater climate concerns and are more ready to alter their behavior than men, but no effect from gender on support for climate change policy. Two additional studies (O'Connor et al. 1999; 2002) examine gender's effect on support for climate policy and behavioral intentions simultaneously, and while neither finds that women are more supportive of climate policy, O'Connor et al. (1999) find a positive gender effect on behavioral intentions: Women are more likely to report a willingness to take voluntary action to mitigate climate change, but "(o)nce there are controls for views toward the government and climate change, and general environmental beliefs, men and older respondents are actually somewhat more likely to vote for government policies to address climate change than are women and the young" (O'Connor et al. 1999, p. 468).

These findings, we believe, underscore why survey researchers should study gender's effects on beliefs, policy support, and everyday behavior simultaneously, to disentangle the direct and indirect effect (working through beliefs) from gender on climate policy support and mitigation behavior. The issue is whether women and men with the same climate beliefs still differ in terms of climate policy support and everyday climate mitigation behavior, or whether belief differences explain the entire gender difference in policy support and mitigation behavior. In fact, among the 8 studies that find an insignificant link between gender and policy support, all but one (Aldy et al. 2012) control, in one way or another, for climate beliefs, but O'Connor et al. (1999) is the only study to include beliefs in a separate regression, making it possible to disentangle the direct and indirect gender effect. Among the five studies finding an insignificant link between gender and behavioral intentions, all but one (Semenza et al. 2008), control for political beliefs in all regressions, making it hard to disentangle the direct and indirect gender effect here as well.

In summary, the inconclusive previous empirical evidence in the survey literature suggests that the link between gender, climate concerns, climate policy support, and climate mitigating behavior needs further examination. Notably, the links between these climate constructs in relation to gender needs to be examined, to disentangle between a direct effect from gender on policy support and mitigation behavior and an indirect effect working through climate beliefs. Also, while US numbers concerning the relationship between gender and the three outcomes are roughly on par with findings from other countries, the United States is seen as an outlier in climate perceptions (e.g., Harrison 2010; McCright et al. 2016a). More studies examining these issues in other countries is therefore needed; for example, only one Swedish study covered in McCright et al.'s

(2016a) overview considers gender, finding women to report greater climate change concerns than men (Sundblad et al., 2007).

2.2 Hypothesis development

We formulate three hypotheses to simultaneously examine the relationships between gender, climate beliefs, climate policy support, and everyday climate behavior, and examine the links between these climate constructs. First, following prior survey study data showing the existence of a consistent gender gap in climate beliefs (McCright et al. 2016a), it is reasonable to assume that this relationship holds here as well:

Hypothesis 1: Systematic gender differences exist in climate beliefs, with women believing more strongly than men that climate change is concern.

Not being able to confirm this hypothesis would be surprising. The second relationship of interest, between gender and climate policy support, is less obvious. Granted, cross-country evidence suggests that gender equality and female empowerment at the national level may be crucial to yield better environmental and climate policies (Norgaard and York 2005; Buckingham 2010; Austin & McKinney 2016; McKinney & Fulkerson 2015; Ergas and York 2012; Ergas et al. 2021). Yet, only half of the survey studies examining the relationship between gender and climate policy support find a significant relationship (McCright et al. 2016a). Still, it appears reasonable to believe that the traditional role of women as e.g., caregivers, subsistence providers, and generally more altruistic (Shiva and Mies, 2014; Steg 2016; Terry 2009; Warren 1990; Bell and Braun 2010) also make them more inclined to support climate-policy than men:

Hypothesis 2: Systematic gender differences exist in support for climate mitigation policies, with women being more in favor of such policies than men.

The third relationship of interest is between gender and everyday climate mitigation behavior. As mentioned, the relationship between gender and behavioral intentions has been studied in a survey context before, but five out of seven studies fail to find a significant link. Nevertheless, studies show that women tend to make better ecological and health decisions in their local environment (e.g., Agarwal 1994; Schuler and Hashemi 1994; Hunter et al. 2004; Pisano and Lubell 2017; Xiao and McCright 2014; Li et al. 2019). Moreover, masculinity appears to exacerbate gender disparities in health behaviors, e.g., related to meat consumption (Nakagawa & Hart 2019). Thus, we hypothesize that:

Hypothesis 3: Systematic gender differences exist everyday climate mitigation behavior, with women engaging in more such behavior than men.

As mentioned, we will also focus on how the different climate constructs affect one another, and whether this mediates the gender effect of gender. Specifically, for H2 and H3, we examine whether any gender effect persists even when climate beliefs are controlled for, meaning there is a direct gender effect independent of climate beliefs. The issue is whether women and men with the same climate beliefs still differ in terms of climate policy support and everyday climate behavior, or whether belief differences explain (parts of) their differences in support and behavior.

3. Data

3.1 Survey design

SEPA undertook the survey as an online questionnaire for two weeks in May 2018. The survey population was initially contacted by phone and selected as a representative sample of the Swedish adult population (18 years or older) concerning the reference variables *sex*, *age*, and *residential location* (city, suburban/small town, or rural). The response rate was 52 percent, corresponding to 996 respondents. There is no statistically significant difference in background characteristics between respondents and those who chose not to participate. While respondents were aware the survey would concern societal issues, they did not know beforehand that it would discuss climate change attitudes, limiting concerns about self-selection bias.

Respondents answered climate change questions by indicating their most preferred alternative on a Likert scale. The answers will be used to construct our dependent variables, whereas questions about respondent characteristics constitute our independent variables, together with the reference variables. We refer to SEPA (2018) for a full survey description. Below, we discuss each set of variables.

3.2 Dependent variables

Climate change questions are categorized according to A) Beliefs about climate change in general; B) The type of political interventions that could potentially be employed to mitigate climate change; and C) Which type of voluntary actions the respondent has already undertaken to mitigate climate change. Respondents indicate their level of agreement with each statement, where answers have a natural ordering. Table 1 shows all the questions associated with each category, the way we label them, and the available response categories. Categories A and C have five response categories (associating two options with disagreement), while category B has four (associating one option with disagreement).

Table 1. Overview of all climate-related questions in SEPA survey, coding and labeling

Category A: Beliefs about climate change in general

Question	*Label*
A1. ”Do you think that the climate change is something which, now or in the future, will affect us who live in Sweden?”	A1. Climate change affects Sweden
A2. ”Do you think that we in Sweden can do something to reduce climate change?”	A2. Sweden can reduce climate change
A3. ”Do you think that you can personally do something to reduce climate change?"	A3. The individual can reduce climate change

Questionnaire coding	Empirical re-coding first stage	Empirical re-coding second stage	
		Upper tail	Lower tail
1. No, not at all	0. Non-agreement		1. Strong disagreement
2. No, hardly	0. Non-agreement		0. Weak disagreement
3. Don’t know	0. Non-agreement		
4. Yes, possibly	1. Agreement	0. Weak agreement	
5. Yes, absolutely	1. Agreement	1. Strong agreement	

Category B: The type of political interventions that could potentially be employed to mitigate climate change

Question	*Label*
B1. ”What do you think about the state adding new taxes and fees to goods and services such as petrol, oil, and flights that have major climate impact?”	B1. Tax harmful goods
B2. ”What do you think about the state using tax funds to provide grants or lower taxes when purchasing goods and services which reduce climate-effecting emissions?”	B2. Subsidize pro-environmental goods
B3. ”What do you think about sending more information to households about climate change, in order to make them choose goods and services which reduce climate- effecting emissions?”	B3. Send information

Questionnaire coding	Empirical re-coding first stage	Empirical re-coding second stage	
		Upper tail	Lower tail
1. Not good	0. Non-agreement		N/A
2. Don’t know	0. Non-agreement		N/A
3. Fairly good	1. Agreement	0. Weak agreement	
4. Good	1. Agreement	1. Strong agreement	

Category C: Which type of voluntary actions the respondent has already undertaken to mitigate climate change

Question	*Label*
“Have you done something in your everyday life to reduce your climate impact in the last two years, for example...”	
C1. “... sorting your waste more?”	C1. Recycling
C2. “... reducing your energy use in the home?”	C2. Energy saving

C3. "... reusing things?"		C3. Reusing	
C4. "... changing your choice of holiday destination?"		C4. Holiday destination	
C5. "... eating less meat?""		C5. Meat reduction	
C6. "... changing your choice of daily transport?"		C6. Transport	
Questionnaire coding	Empirical re-coding first stage	Empirical re-coding second stage	
		Upper tail	Lower tail
1. No, not at all	0. Non-agreement		1. Strong disagreement
2. No, hardly	0. Non-agreement		0. Weak disagreement
3. Don't know	0. Non-agreement		
4. Yes, possibly	1. Agreement	0. Weak agreement	
5. Yes, absolutely	1. Agreement	1. Strong agreement	

Note: Climate-related survey questions by category translated from Swedish, how Likert scale responses are coded into ordered responses, as well as how responses are coded into binary variables used when estimating the first and second stage of the question-specific analysis described in Section 4.2.

Table 2 displays summary statistics by gender for the climate questions, together with each gender's level of agreement in percentage points. Agreement is particularly high among the general questions in Category A and always higher for women. Other surveys corroborate this gender difference in Swedish climate change perception, e.g., the European Investment Bank (2018). The level of agreement is more muted for the policy questions in Category B (notably B1, concerning taxing harmful goods), whereas the mitigation behavior questions in category C occupy an intermediate position. Moreover, the mean level of agreement is higher for women for all questions, and the difference is also statistically significant.

Table 2: Summary statistics of the dependent variables, by gender

Category	*Male (n=531)*			*Female (n=463)*			*Difference (n=994)*
A. General questions	Agreement (%)	Mean	Std. Dev.	Agreement (%)	Mean	Std. Dev.	Female mean - Male mean
A1. Climate change affects Sweden	92	4.53	0.88	98	4.8	0.51	0.27***
A2. Sweden can reduce climate change	74	3.88	1.30	89	4.32	0.96	0.44***
A3. The individual can reduce climate change	72	3.8	1.31	87	4.29	0.99	0.49***
Composite index A	N/A	12.20	3.02	N/A	13.41	2.12	1.21***
B. Climate policy questions							
B1. Tax bad goods	45	2.17	1.27	50	2.34	1.25	0.17**
B2. Subsidize good goods	73	2.87	1.14	79	3.11	1.04	0.24***
B3. Send information	89	3.3	0.86	91	3.49	0.76	0.19***
Composite index B	N/A	8.33	2.38	N/A	8.93	2.19	0.60***
C. Individual climate actions							
C1. Recycling	87	4.31	1.07	94	4.58	0.84	0.27***
C2. Energy saving	75	3.76	1.22	78	3.91	1.18	0.15*
C3. Reusing	81	3.96	1.13	88	4.26	0.98	0.30***
C4. Transport	57	3.3	1.44	63	3.51	1.43	0.21**
C5. Meat reduction	43	2.78	1.44	67	3.51	1.45	0.73***
C6. Holiday destination	41	2.79	1.39	53	3.18	1.46	0.39***
Composite index C	N/A	20.91	5.28	N/A	22.96	5.11	2.05***

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

Note: Summary statistics of dependent variables, by gender. Minimum response levels are 1 for each specific question. Max values are 5 for category A and C, and 4 for category B. "Agreement" is the binary variable used in the first stage (the computation of this variable is described in Table 1 above). Standard deviation is the standard deviation of the mean. Significance stars in the last column refer to two-sided t-tests of differences in means. Composite indices are constructed by summing response levels over each category and respondent. In effect, composite indices in category A, B, and C, lies within the range 3-15, 3-12, and 6-30 respectively.

Figures A1-A3 in the appendix display histograms of responses to each question in A-C, by gender. The share of women who "absolutely" agree always exceeds the share of men, but no similar pattern is apparent for respondents who "possibly" agree. At the distribution's lower end, the shares of men who ("possibly" or "absolutely") disagree always exceed those of women for all questions. The share of respondents indicating "Don´t know" is low across all questions for both men and women. While we use these to construct of our dependent variables, the variable coding hinges on the method and will therefore be described in section 4.1.

3.3 Independent variables

Table 3 displays summary statistics for the independent variables. The key independent variable is operationalized as the self-reported answer to the question "What is your gender?"[1] The alternatives were *Male*; *Female*; *Other*; and *Do not want to disclose.* Technically, gender is different from the study population's reference variable, which is the binary variable *sex*. However, only two respondents (0.2 percent) indicated *Other*, and no respondent indicated *Do not want to disclose.* Because our concern is statistical inference, we drop these two observations and simply use the binary variable *female*, where unity indicates the respondent identifies as a female. We expect the relationship between this variable and the (three sets of) dependent variables to be positive, meaning female respondents hold stronger climate beliefs (H1), show greater support for climate-mitigation policies (H2), and undertake more everyday climate-mitigating behavior (H3).

Several of the control variables vary significantly with gender. Especially, it worth noting that the share of male respondents that regularly *drives car* (88 percent) exceeds the corresponding figure for females (80 percent) by 0.07.

[1] The Swedish word for gender is *könsidentitet*, and the word for sex is *kön*.

Table 3. Summary statistics of the independent variables

	Male (n=531)		*Female (n=463)*		*Difference (n=994)*
	Mean	Std. Dev.	Mean	Std. Dev.	Female mean - Male mean
Education	2.34	0.64	2.44	0.62	0.10*
Age	53.34	17.42	48.47	17.39	- 4.87***
Drives car	0.88	0.33	0.80	0.40	-0.08**
Student	0.05	0.21	0.08	0.28	0.04*
Unemployed	0.02	0.14	0.03	0.16	0.01
Retired	0.33	0.47	0.25	0.43	-0.08**
Pop > 200 K	0.24	0.43	0.27	0.44	0.03
Pop > 15 K	0.67	0.47	0.68	0.47	0.00
N	463		531		994

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

Note: Summary statistics of independent variables, by gender. Significance stars in the last column refer to two-sided t-tests of differences in means.

The other self-reported characteristics are included as control variables. *Education* describes the respondent's highest completed level of education on a scale from 1-3 where 1 equals compulsory school, 2 upper-secondary school, and 3 a university degree (or equivalents). Generally, more educated people show greater climate concerns, but political beliefs seem to moderate this link in highly developed countries, with educated conservatives being less concerned (Czarnek et al. 2021). Moreover, female education has been posited as powerful tool to combat climate change (Kwauk and Braga 2017; Ergas et al. 2021). We thus expect a positive association with climate attitudes and behaviors.

Age is measured in years (18-84). Whereas O'Connor et al. (1999) find that men and older respondents are more likely to vote for climate-mitigation policy than women and the young, once other factors are controlled for, recent research appears to confirm that younger people care more about climate change than older people (Milfont et al. 2021). Thus, we expect the association between age and climate attitudes and behaviors to be negative.

Drives car is a binary variable, where unity indicates that the respondent drives regularly, which 88 percent of male respondents and 80 percent of female respondents do. While car users' attitudes to climate change differs considerably, with some groups being more amenable to changing their behavior than others (Ali et al. 2018), we expect the relationship between this variable and climate change beliefs and behaviors to be, by and large, negative.

Working, Unemployed, Student and Retired are binary, mutually exclusive indicators of the current labor market status of the respondent. As our baseline category, *working* is not explicitly included in the model. Recent evidence suggests that climate skepticism increases with negative economic shocks and that the effects are concentrated among individuals in the labor force (Meyer 2021), meaning the indicators may reveal a positive association with climate change beliefs and behaviors relative to the base indicator.

Pop>200 K and *Pop>15 K* are binary variables indicating whether respondents live in a city with more than 200 000 or 15 000 residents, respectively. A respondent belonging to the first category is by default also included in the second category, with the base alternative being locations with less than 15 000 residents. In the United States, there is an urban/rural divide on climate change, "with rural voters being more skeptical of both the science of and governmental response to climate change" (Bonnie et al., 2020, p. 3). However, in Sweden there is no documented such relationship, making the expected association between population size and climate beliefs or policies ambiguous.

4. Method

4. 1 Composite index analysis by category

When analyzing Likert scale responses with agreement levels that may be intercorrelated for a respondent, a common approach is to compute composite indices reflecting the respondent's mean level of agreement. Then, the association between the composite index and respondent characteristics can be analyzed directly, as done in the context of gender and climate perceptions by e.g. Zahran et al. (2006), Dietz et al. (2007), McCright et al. (2013), McCright et al. (2016b). Because questions within each category A, B, and C are likely intercorrelated, we begin the empirical analysis by constructing composite indices for each category of questions, using Cronbach's alpha to judge their level of intercorrelation. We then estimate the gender effect according to the following model:

$$y_i = \alpha + \beta_0 female_i + \boldsymbol{X_i \rho} + \varepsilon_i \qquad (1),$$

where y_i is the composite index for respondent i associated with a certain category, and α is a constant. β_0 is the coefficient of interest, estimating the conditional association between the binary variable $female$ and the dependent variable. Finally, $\boldsymbol{X_i}$ is a vector of control variables with its associated coefficient vector $\boldsymbol{\rho}$, and ε_i is the error term, possibly subject to heteroscedasticity. While the composite indices can be considered as ordinal variables, the large set of possible values make it possible to treat them as continuous and estimate (1) using OLS. Because we standardize each index, coefficients can be interpreted as the standard deviation change in the dependent variable following a unit increase in the independent variable.

As mentioned, respondents with greater general climate beliefs may also be more likely to support climate policy measures and engage in everyday climate-mitigating behavior. The issue is whether women and men with the same climate beliefs still differ in terms of climate policy support and everyday climate behavior, or whether belief differences explain all their differences in support and behavior. Therefore, we also undertake specifications for the second and third composite index where the first composite index is included as an independent variable. The identifying assumption is no reverse causality, i.e., support for climate policies and everyday behavior should not shape general climate beliefs. Since the data show that women have stronger general beliefs, the specification allows us to disentangle the *direct effect* of gender, as opposed to the *total effect* unconditional on general climate beliefs in the first specifications. The first index can here be thought of as a *mediator* variable, explaining part of the mechanism through which women score higher on the latter two indices. This approach follows e.g. McCright et al. (2014), who disentangle direct and indirect effects using GMM within a Structural Equation Model (SEM) framework. However, when both the mediator and dependent variables are continuous as in our case, OLS produces results identical to GMM (Rijnhart et al. 2017).

Cronbach's alpha is around 0.8 for both the general climate questions index (A) and the everyday behavior index (C), but only 0.53 for the climate policy index (B), making it a less suitable metric for this category. Moreover, the composite index analysis produces easily interpreted results but obfuscates whether any observable effect is driven by variation in the lower, middle, or upper part of the distribution. To find out, e.g., whether women report higher agreement than men given that subjects are either at least moderately positive or negative to a certain statement, or whether effects are mainly driven by variation around the middle of the distribution, we also estimate the gender effect separately for each survey question.

4.2 Binomial logit analysis by individual questions

An alternative strategy for estimating (1) is the ordered logistic regression. This estimator relies on the proportional odds assumption, meaning that that dependent variables' slope coefficients are the same across response levels. However, Brant tests reject this null hypothesis for most of our survey questions. While a multinomial framework is a potential alternative, a fully specified multinomial choice model would produce a large set of coefficients associated with the relationship between a chosen base response level and all other response levels. This analysis will not reveal anything about the probability of moving between any pair of response levels that does not include the base level. Therefore, we employ another estimation strategy, which accounts for the natural response ordering by splitting the analysis into parts and should be easier to interpret.

In the first stage, we create binary dependent variables by coding as "agreement" (with the value 1) responses that "absolutely" or "possibly" agree with a question in categories A, B, or C (levels

4-5 for A and C; 3-4 for B). The remaining responses are coded as “non-agreement” (with the value 0). We then estimate equation (1) using a conventional binomial logistic regression.

In the second stage, we perform separate analyses for the upper and lower tails of the distribution. First, we restrict the sample to responses that “absolutely” or “possibly” agree with each question (levels 4-5 for A and C; 3-4 for B), to examine if the level of agreement varies with gender, given that the respondents are at least moderately positive. Second, we restrict the sample to responses that “absolutely” or “possibly” disagree with each question (levels 1-2 for A and C). This test cannot be performed for questions in category B, since there is only one level of disagreement (Table 1).

While multicollinearity across the independent variables may result in biased coefficients, this concern seems limited as the highest correlation between any two independent variables is 0.4 (between *pop>200 K* and *pop>15 K* due to the mechanic link between the two). In all estimations, we also account for possible heteroscedasticity using Huber-White standard errors. And while an inspection of histograms in Figures A1-A3 suggest that outliers affecting the results will be unlikely, we effectively remove any such variation by grouping the dependent variables into binary categories in each stage of the analysis.

5. Results

5.1 Composite index analysis by category A, B, and C

Table 4 displays OLS results from the analysis where the dependent variables are standardized composite indices. The total gender effect in specifications (I) is precisely estimated for each index and varies from 0.23 (climate policy) to 0.40 (general climate questions). The interpretation is that the mean response level is 0.23 – 0.40 standard deviations higher for women, conditional on the demographic covariates. When we include the general beliefs index as an independent variable (II), the *direct* gender effect on the climate policy index becomes economically and statistically insignificant, suggesting the effect is entirely driven by women's stronger climate beliefs. In other words, men and women with similar climate beliefs will support climate policy to a similar degree, conditional on the control variables. However, the policy support index' comparatively low Cronbach's alpha suggests a composite analysis may not be entirely appropriate. Further, the *direct* gender effect on everyday mitigation behavior in specification (II) is about half the size of the *total* effect (falling from 0.38 to 0.22), but the coefficient is still precisely estimated. This suggests that climate belief differences explain about half of the gender difference in everyday mitigation behavior.

Table 4: OLS results by category when dependent variables are standardized composite indices

	A. General Climate Composite Index	*B. Climate Policy Composite Index*		*C. Individual Climate Composite Index*	
	(I)	(I)	(II)	(I)	(II)
Female	0.40***	0.23***	0.024	0.38***	0.22***
	(0.062)	(0.064)	(0.057)	(0.065)	(0.060)
Education	0.11*	-0.0030	-0.058	-0.030	-0.074
	(0.049)	(0.052)	(0.045)	(0.051)	(0.047)
Age	-0.0060*	-0.00054	0.0026	0.00033	0.0028
	(0.0029)	(0.0030)	(0.0026)	(0.0030)	(0.0026)
Drives car	-0.12	-0.25**	-0.19*	-0.12	-0.073
	(0.081)	(0.091)	(0.080)	(0.093)	(0.088)
Student	0.10	-0.084	-0.14	0.11	0.070
	(0.12)	(0.16)	(0.14)	(0.15)	(0.14)
Unemployed	0.10	-0.036	-0.088	0.25	0.21
	(0.23)	(0.20)	(0.19)	(0.21)	(0.19)
Retired	0.15	-0.079	-0.16	0.16	0.098
	(0.11)	(0.11)	(0.092)	(0.11)	(0.094)
Pop > 200 K	0.15*	0.15	0.073	0.23**	0.17*
	(0.077)	(0.084)	(0.072)	(0.083)	(0.077)
Pop > 15 K	-0.096	-0.061	-0.011	-0.17*	-0.13
	(0.074)	(0.073)	(0.063)	(0.072)	(0.066)
General Climate Composite Index			0.52***		0.41***
			(0.028)		(0.033)
N	994	994	994	994	994
Cronbach's alpha	0.81	0.53	0.53	0.78	0.78

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

Note: Results from OLS regressions. Dependent variables are (standardized) indices that are computed for each category by summing individual responses over all questions in that category. Heteroscedasticity robust standard errors in parentheses.

With regard to the control variables, *drives car* has a precisely estimated (i.e. $p < 0.01$) negative effect on the climate policy index (-0.25). The interpretation is that subjects who regularly drive a car score on average 0.25 standard deviations lower than other subjects. *Pop > 200 K* (i.e., citydwellers) instead has a precisely estimated positive effect on the individual action index (0.23). Potential mechanisms for these results are discussed further in section 5.3 and 5.4 below. However, the *direct* effects in specifications (II) are both lower and less precisely estimated than the *total* effects in (I).

5.2 Binomial logit analysis: General climate change questions (category A)

Table 5 displays results from estimations of the first stage where the dependent variables concern general questions about climate change in Category A. Thus, dependent variables are binary responses to three questions: whether “climate change is something which, now or in the future, will affect us who live in Sweden” (A1), whether “we in Sweden can do something to reduce climate change” (A2), and whether the individual “personally (can) do something to reduce climate change” (A3). Coefficients are expressed as odds ratios. With respect to the variable *female*, the odds ratio is the odds of agreement conditional on the subject being female, divided by the odds of agreement conditional on the subject being male. An odds ratio exceeding unity implies that the odds of agreement increase when the subject is female. The bottom rows of the table also present the non-transformed coefficients and marginal probabilities for the female variable, evaluated at the means of each control variable.

Table 5. Odds ratios of agreeing with general climate questions from logistic regressions, first stage.

	A1. Climate change affects Sweden		A2. Sweden can reduce climate change		A3. The individual can reduce climate change.	
	(I)	(II)	(I)	(II)	(I)	(II)
Female	4.76***	4.55***	2.72***	2.59***	2.77***	2.60***
	(1.87)	(1.85)	(0.48)	(0.47)	(0.47)	(0.46)
Education		1.19		1.12		1.10
		(0.28)		(0.15)		(0.15)
Age		0.98		0.99		0.99
		(0.013)		(0.0074)		(0.0077)
Drives car		0.82		0.81		0.83
		(0.45)		(0.22)		(0.21)
Student		2.33		1.22		1.15
		(2.79)		(0.54)		(0.49)
Unemployed		1.09		0.78		1.08
		(1.17)		(0.42)		(0.62)
Retired		2.59		1.67		1.28
		(1.27)		(0.45)		(0.34)
Pop > 200 K		0.74		1.54		1.43
		(0.30)		(0.34)		(0.30)
Pop > 15 K		1.13		0.75		0.65*
		(0.39)		(0.14)		(0.12)
Female coefficient	1.56	1.51	1.00	0.95	1.02	0.96
Female margin	0.058	0.053	0.15	0.14	0.16	0.15
N	994	994	994	994	994	994

* p < 0.05, ** p < 0.01, *** p < 0.001

Note: Results from binary logit regressions, first stage. Dependent variables are binary "agreement"-responses, constructed as described in Table 1. Coefficients are odds ratios. Heteroscedasticity robust standard errors in parentheses. For the *female* variable, raw coefficients and marginal effects (evaluated at the mean of each independent variable) are displayed in the bottom rows.

All odds ratios associated with *female* exceed unity and are economically *and* statistically significant for all three statements. Statement A1, whether climate change affects Sweden, yields the highest odds ratio. The odds of agreeing to the statement increases by a factor of 4.55 when the subject is female, conditional on the control variables. However, the unconditional probability of agreement is also very high for males, at 0.92 compared to 0.98 for women. Therefore, the high odds ratio can largely be explained by the highly non-linear relationship between probabilities and odds as probabilities approach unity. The corresponding odds ratios of the gender variable associated with the statements that Sweden (A2) and the individual can do something about climate change (A3) are 2.59 and 2.60, respectively, demonstrating a gender effect here as well. The table's bottom rows also present the marginal probability associated with the *female* coefficient (evaluated at the means of each control variable), and the corresponding non-transformed coefficient. The probability of agreeing that climate change is a concern (A1) increases by 0.053 when the subject is female, but for the statements that Sweden (A2) and the individual can do something about climate change (A3), these effects are as large as 0.14 and 0.15, respectively. No control variable coefficients are statistically significant in any specification.

Table 6 shows results associated with the upper (I) and lower (II) tail of the second stage test. For the upper tail, i.e., agreement responses, the gender effect is only statistically significant for the statement that climate change is a concern (A1). The odds ratio is 1.67, corresponding to a marginal effect of 0.079. Since most subjects indicated a level of agreement of at least 4, there is little loss in statistical power in these tests. The only statistically significant control variable coefficient is for education, which exceeds unity in each specification. Lower tail tests reveal no detectable gender effect in any specification, and control variables are also imprecisely estimated throughout, though the sample is very small.

Table 6. Odds ratios of agreeing with climate policy questions from logistic regressions, first stage.

	B1. Tax harmful goods		B2. Subsidize pro-climate goods		B3. Send information	
	(I)	(II)	(I)	(II)	(I)	(II)
Female	1.25	1.13	1.37*	1.31	1.29	1.48
	(0.16)	(0.15)	(0.20)	(0.20)	(0.28)	(0.33)
Education		1.30*		0.87		0.70
		(0.14)		(0.11)		(0.13)
Age		1.00		1.00		1.01
		(0.0062)		(0.0070)		(0.010)
Drives car		0.48***		0.61*		1.44
		(0.094)		(0.15)		(0.41)
Student		1.01		0.65		0.47
		(0.31)		(0.22)		(0.19)
Unemployed		0.57		1.09		1.05
		(0.28)		(0.60)		(0.82)
Retired		0.92		0.70		0.87
		(0.20)		(0.17)		(0.35)
Pop > 200 K		1.53*		1.11		0.73
		(0.26)		(0.21)		(0.20)
Pop > 15 K		0.98		0.74		1.22
		(0.15)		(0.13)		(0.32)
Female coefficient	0.22	0.12	0.31	0.27	0.25	0.39
Female margin	0.055	0.031	0.058	0.049	0.023	0.033
N	994	994	994	994	994	994

* p < 0.05, ** p < 0.01, *** p < 0.001

Note: Results from binary logit regressions, first stage. Dependent variables are binary "agreement"-responses, constructed as described in Table 1. Coefficients are odds ratios. Heteroscedasticity robust standard errors in parentheses. For the female variable, raw coefficients and marginal effects (evaluated at the mean of each independent variable) are displayed in the bottom rows.

To summarize, results for category A largely support Hypothesis 1, suggesting women are believe more strongly than men that climate change is concern, but also more confident that Sweden and the individual can do something about it.

5.3 Binomial logit analysis: Climate policy questions (category B)

Table 7 displays results from the first stage when the dependent variables concern questions about climate policy support in Category B, specifically taxing harmful goods (B1), subsidizing pro-environmental goods (B2), and information campaigns (B3). While Table 1 revealed that women exhibited an unconditionally more positive attitude towards each policy measure, we here find no statistically significant gender effect in any specification, even if all point estimates exceed unity. When we include the composite belief index as an additional control (not shown), all point estimates remain insignificant but fall *below* unity, suggesting that to the extent there are gender differences in support for individual policy measures, they are driven by climate beliefs. Among the control variables, it is worth noting the (statistically and economically significant) negative association between *drives car* and support for taxing harmful goods (B1) and subsidizing pro-environmental goods (B2). These results make intuitive sense based on self-interested economic grounds, as the policies will likely increase the relative price of fuel for cars compared to other means of transportation.

Table 7. Odds ratios of self-reported measures taken to mitigate climate change from logistic regressions, first stage.

	C1. Recycling		C2. Energy saving		C3. Reusing		C4. Transport		C5. Meat reduction		C6. Holiday destination	
	(I)	(II)	(I)	(II)	(I)	(II)	(I)	(II)	(I)	(II)	(I)	(II)
Female	2.28***	2.54***	1.18	1.36	1.69**	1.83***	1.32*	1.22	2.68***	2.57***	1.65***	1.70***
	(0.53)	(0.61)	(0.18)	(0.22)	(0.30)	(0.34)	(0.17)	(0.17)	(0.35)	(0.35)	(0.21)	(0.22)
Education		0.95		0.96		0.97		1.06		1.22		0.80*
		(0.17)		(0.12)		(0.14)		(0.12)		(0.13)		(0.085)
Age		1.02		1.01		1.01		1.00		1.00		1.00
		(0.011)		(0.0069)		(0.0082)		(0.0062)		(0.0062)		(0.0062)
Drives car		0.81		2.03***		0.99		0.49**		0.79		0.79
		(0.26)		(0.42)		(0.26)		(0.11)		(0.16)		(0.15)
Student		0.73		0.71		0.59		1.78		1.22		0.89
		(0.32)		(0.22)		(0.20)		(0.62)		(0.38)		(0.27)
Unemployed		1.01		1.73		1.90		0.97		1.25		1.90
		(0.81)		(0.99)		(1.48)		(0.46)		(0.59)		(0.86)
Retired		0.95		1.15		1.03		1.35		1.13		1.04
		(0.41)		(0.30)		(0.31)		(0.30)		(0.25)		(0.23)
Pop > 200 K		0.73		1.48*		1.07		1.42*		1.63**		1.20
		(0.20)		(0.29)		(0.24)		(0.25)		(0.29)		(0.20)
Pop > 15 K		0.72		0.53***		0.79		1.22		1.03		0.83
		(0.20)		(0.100)		(0.17)		(0.19)		(0.16)		(0.13)
Female coefficient	0.82	0.93	0.17	0.31	0.53	0.61	0.28	0.20	0.99	0.94	0.50	0.53
Female margin	0.068	0.073	0.030	0.054	0.070	0.079	0.066	0.048	0.25	0.23	0.13	0.13
N	994	994	994	994	994	994	994	994	994	994	994	994

Table 8 displays results from the upper-tail tests (as mentioned, we cannot perform a lower-tail test here). Here, gender has a positive association with support for information campaign (B3), where the female coefficient displays a statistically significant odds ratio of 1.90, corresponding to a marginal probability of 0.15.

Table 8. Odds ratios of agreeing with general climate questions from logistic regressions, second stage.

	A1. Climate change affects Sweden		A2. Sweden can reduce climate change		A3. The individual can reduce climate change.	
	(I) Upper tail	(II) Lower tail	(I) Upper tail	(II) Lower tail	(I) Upper tail	(II) Lower tail
Female	1.67**	0.11	0.96	1.18	1.10	1.46
	(0.29)	(0.13)	(0.14)	(0.59)	(0.17)	(0.70)
Education	1.56***	0.70	1.43**	0.76	1.34*	0.58
	(0.21)	(0.44)	(0.17)	(0.23)	(0.16)	(0.18)
Age	0.99	1.08*	0.99	1.01	0.99	1.01
	(0.0081)	(0.040)	(0.0067)	(0.019)	(0.0070)	(0.015)
Drives car	0.49*	1.89	0.82	0.24	0.90	0.54
	(0.14)	(2.33)	(0.18)	(0.27)	(0.19)	(0.44)
Student	1.21	1	1.43	1	0.64	1.62
	(0.54)	(.)	(0.51)	(.)	(0.21)	(1.74)
Unemployed	2.14	1	2.02	0.86	1.53	0.44
	(1.61)	(.)	(1.23)	(1.04)	(0.86)	(0.51)
Retired	1.00	0.10	1.04	1.42	0.66	2.23
	(0.27)	(0.15)	(0.25)	(0.99)	(0.17)	(1.41)
Pop > 200 K	0.91	2.70	1.23	0.50	1.19	2.10
	(0.20)	(3.54)	(0.23)	(0.27)	(0.23)	(1.33)
Pop > 15 K	1.08	0.40	0.96	1.17	1.12	0.82
	(0.21)	(0.40)	(0.17)	(0.53)	(0.20)	(0.35)
Levels coded as unity	5	2	5	2	5	2
Female coefficient	0.51	-2.17	-0.036	0.17	0.091	0.38
Female margin	0.079	-0.40	-0.0086	0.024	0.022	0.053
N	945	39	803	174	785	191

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

Note: Results from binary logit regressions, second stage. Dependent variables are binary "agreement"-responses, constructed as described in Table 1. Coefficients are odds ratios. Heteroscedasticity robust standard errors in parentheses. For the *female* variable, raw coefficients and marginal effects (evaluated at the mean of each independent variable) are displayed in the bottom rows.

To summarize, negative binomial results related to questions in category B offer little support for Hypothesis 2, suggesting that once other variables are controlled for, notably climate beliefs, women are no more likely than men to favor pro-climate policies than men.

5.4 Binomial logit analysis: Everyday climate behavior (category C)

Table 9 displays first stage results when the dependent variables concern questions about respondents' everyday climate-mitigating behavior in Category C, specifically recycling (C1), energy saving (C2), reusing (C3), transport (C4), meat reduction (C5), and the choice of holiday destination (C6). There is an economically and statistically significant effect with respect to all listed actions except for energy saving and transport changes. A likely reason for the insignificant effects is that men own about twice as many cars and residential properties in Sweden as women (Ownershift 2019), suggesting that men have a greater potential to take measures in these two areas. The statistically significant odds ratios vary little across specifications, and all lie in the range 1.70 (choice of holiday destination) and 2.54 (meat reduction) when including control variables. When we include the climate belief index (not shown), all odds ratios associated with gender shrink in magnitude, but while the four statistically significant effects weaken, they remain significant. This suggests that while differences in climate beliefs explain some of the behavioral differences across genders, a woman holding the same beliefs as a man is still more likely to engage in climate-mitigating behavior.

The marginal probabilities are between 0.073 (recycling) and 0.23 (reducing meat consumption). These gender differences likely reflect the fact that social forces generally position women as e.g., caregivers, subsistence providers, and generally more altruistic than men (Shiva and Mies, 2014; Steg 2016; Terry 2009; Warren 1990; Bell and Braun 2010). The fact that the most pronounced effect is found in relation to meat reduction confirms previous studies attributing meat consumption to masculinity norms (Nakagawa & Hart 2019).

Table 9. Odds ratios of agreeing with climate policy questions from logistic regressions, second stage.

	B1. Tax harmful goods	B2. Subsidize pro-climate goods	B3. Send information
	(I) Upper tail	(I) Upper tail	(I) Upper tail
Female	0.85	1.32	1.90***
	(0.16)	(0.20)	(0.27)
Education	1.06	1.24	0.96
	(0.16)	(0.15)	(0.11)
Age	0.99	0.99*	1.01
	(0.0084)	(0.0068)	(0.0066)
Drives car	0.76	0.99	1.22
	(0.19)	(0.22)	(0.26)
Student	1.27	1.25	0.89
	(0.54)	(0.45)	(0.30)
Unemployed	7.24	0.63	1.34
	(7.55)	(0.32)	(0.69)
Retired	1.01	1.20	1.06
	(0.32)	(0.30)	(0.25)
Pop > 200 K	1.22	0.94	1.58*
	(0.28)	(0.18)	(0.30)
Pop > 15 K	1.08	0.97	0.95
	(0.25)	(0.17)	(0.15)
Levels coded as unity	4	4	4
Female coefficient	-0.17	0.27	0.64
Female margin	-0.042	0.068	0.15
N	469	751	894

* p < 0.05, ** p < 0.01, *** p < 0.001

Note: Results from binary logit regressions, second stage. Dependent variables are binary "agreement"-responses, constructed as described in Table 1. Coefficients are odds ratios. Heteroscedasticity robust standard errors in parentheses. For the female variable, raw coefficients and marginal effects (evaluated at the mean of each independent variable) are displayed in the bottom rows.

Among the statistically significant background variables, the strong negative association between *drives car* and the transport choice (with an odds ratio of 0.49) makes intuitive sense, as such a change may substantially alter the life of someone dependent on the car as a mode of transportation. In terms of residential energy reduction, *drives car* is instead positive, with a comparatively high and statistically significant odds ratio of 2.03. Rather than being causal, this association may reflect the fact that reducing residential energy consumption is easier for house owners, who are also more likely to use a car than others. There is also a statistically significant negative association between pop > 15 K and energy savings (C2), likely reflecting the greater potential for energy savings among people who live in a house in the countryside. Further, there is a statistically significant positive effect of pop > 200 K and meat reduction (C5), which is in line with previous studies demonstrating that there are twice as many vegetarians per capita in cities compared to the countryside (Novus 2018).

Table 10 displays results for the upper and lower tail of the second stage test. For the upper tail, the gender effect is again both economically and statistically significant with respect to all actions except for energy saving and the choice of transport. The statistically significant odds ratios lie within the range 1.52 (holiday destination) and 1.66 (recycling). The corresponding marginal effects are all at around 0.1. Thus, results confirm a gender effect also among subjects that do engage in actions aimed at reducing climate change, with women doing so to a greater extent than men. Again, however, there is no detectable gender effect in any of the specifications for lower tail.

To summarize, results related to questions in category C largely support Hypothesis 3, suggesting women engaging in more everyday climate-mitigating behavior than men, even when climate beliefs are controlled for.[2]

[2] When we instead use as dependent variable a binary variable indicating if the response was "*Don't know*", as opposed to all other response levels, the *female* variable is nearly always insignificant, suggesting there is no gender effect associated with self-reported lack of knowledge regarding any of the statements. The sole exception belongs to Category B, where women are moderately more likely to choose "Don't know" when it comes to support for taxing harmful goods (B1).

Table 10. Odds ratios of self-reported measures taken to mitigate climate change from logistic regressions, second stage.

	C1. Recycling		C2. Energy saving		C3. Reusing		C4. Transport		C5. Meat reduction		C6. Holiday destination	
	(I) Upper tail	(II) Lower tail	(I) Upper tail	(II) Lower tail	(I) Upper tail	(II) Lower tail	(I) Upper tail	(II) Lower tail	(I) Upper tail	(II) Lower tail	(I) Upper tail	(II) Lower tail
Female	1.66***	0.37	1.28	0.82	1.57**	0.89	1.08	0.83	1.57*	0.73	1.52*	0.76
	(0.26)	(0.21)	(0.19)	(0.28)	(0.22)	(0.41)	(0.19)	(0.20)	(0.29)	(0.15)	(0.31)	(0.15)
Education	0.89	1.20	1.09	0.77	0.99	0.75	1.07	0.85	1.16	0.96	1.04	1.34
	(0.11)	(0.56)	(0.13)	(0.24)	(0.11)	(0.30)	(0.15)	(0.16)	(0.18)	(0.15)	(0.17)	(0.22)
Age	1.01	0.99	0.99	0.98	0.99	0.99	1.00	1.01	0.99	1.00	1.01	0.99
	(0.0072)	(0.022)	(0.0070)	(0.016)	(0.0067)	(0.018)	(0.0079)	(0.011)	(0.0083)	(0.0098)	(0.0097)	(0.0093)
Drives car	1.09	1.23	0.85	0.81	0.96	0.90	0.49**	1.56	0.86	1.06	0.72	0.96
	(0.24)	(1.02)	(0.20)	(0.38)	(0.20)	(0.63)	(0.11)	(0.68)	(0.21)	(0.36)	(0.19)	(0.31)
Student	1.26	0.26	1.46	0.36	1.53	0.41	1.64	0.68	1.56	0.84	1.06	0.90
	(0.43)	(0.44)	(0.56)	(0.26)	(0.52)	(0.40)	(0.59)	(0.53)	(0.59)	(0.46)	(0.51)	(0.46)
Unemployed	0.85	1	1.38	1	0.75	1	0.76	1.96	2.62	3.24	1.97	0.34
	(0.40)	(.)	(0.64)	(.)	(0.34)	(.)	(0.45)	(1.75)	(1.54)	(2.64)	(1.16)	(0.25)
Retired	1.39	2.05	1.69*	3.25	1.09	1.99	0.77	1.51	1.51	1.51	0.93	1.37
	(0.36)	(1.70)	(0.42)	(2.09)	(0.26)	(1.50)	(0.22)	(0.62)	(0.48)	(0.50)	(0.32)	(0.44)
Pop > 200 K	1.10	1.16	0.96	0.78	1.65**	1.04	1.47	1.09	1.35	1.76*	1.14	1.64*
	(0.22)	(0.73)	(0.19)	(0.31)	(0.30)	(0.59)	(0.31)	(0.35)	(0.30)	(0.50)	(0.29)	(0.41)
Pop > 15 K	0.85	0.82	0.79	0.86	0.59**	0.84	1.05	0.70	0.75	1.05	0.44***	0.49**
	(0.15)	(0.49)	(0.14)	(0.37)	(0.098)	(0.44)	(0.22)	(0.18)	(0.17)	(0.23)	(0.10)	(0.11)
Levels coded as unity	5	2	5	2	5	2	5	2	5	2	5	2
Female coefficient	0.51	-0.98	0.25	-0.20	0.45	-0.11	0.077	-0.19	0.45	-0.32	0.42	-0.27
Female margin	0.10	-0.22	0.061	-0.037	0.11	-0.020	0.019	-0.039	0.11	-0.078	0.099	-0.065
N	898	81	759	206	835	133	594	379	534	444	462	478

* p < 0.05, ** p < 0.01, *** p < 0.001

Note: Results from binary logit regressions, second stage. Dependent variables are binary "agreement"-responses, constructed as described in Table 1. Coefficients are odds ratios. Heteroscedasticity robust standard errors in parentheses. For the female variable, raw coefficients and marginal effects (evaluated at the mean of each independent variable) are displayed in the bottom rows.

6. Conclusion

6.1 Discussion of results

The composite index analysis largely supports our three hypotheses, with a few interesting caveats. Notably, the gender effect on climate policy support disappears when we control for general climate beliefs, while the gender effect on behavior weakens substantially. This suggests that women and men with similar climate beliefs will behave fairly similarly in terms of policy support and mitigative behavior. In fact, these results are reminiscent of those of O'Connor et al. (1999), who find that women are more likely to report a willingness to take voluntary action to mitigate climate change, but that the effect disappears when controlling for views toward the government and climate change, and general environmental beliefs.

The binomial results, meanwhile, find strong support for Hypothesis 1, that women are more concerned about climate change than men, and Hypothesis 3, that women engage in more everyday pro-climate behavior. However, the relationship between gender and policy support is insignificant, thus offering little support for Hypothesis 2 except for the stage two results regarding the upper tail of the policy measure "send information" (B3). These findings are largely in line with those of McCright et al. (2016b), the sole previous survey study examining gender's effect on climate beliefs, support for climate policy, and behavioral intentions at the same time (though we examine actual behavior rather than behavioral intentions). McCright et al. (2016b) did not find any association between gender and support for climate change policy either, but a positive association between gender and climate concerns and behavioral intentions.

Methodologically speaking, key strengths of this paper is that we employ two different sets of dependent variables (the composite indices and the dichotomous variables), taking care to use the appropriate model given the characteristics of the dependent variables while also making results comparable to the previous literature. Moreover, the fact that we simultaneously consider gender's effect on three important climate constructs (beliefs, policy support, and everyday behavior) is valuable in that it enables us to distinguish between the effect from gender and climate beliefs on policy support and mitigation behavior. One concern with our data is that all actions are self-reported; as such, they should be interpreted with care (Kormos and Gifford 2014). That said, reporting on past behavior instead of current or future behavior is more likely to reflect actual behavior and could mitigate self-report biases (Gatersleben, 2013; van den Broek et al. 2019). Another limitation is the fact that we have no way to assess our respondents' degree of pro-environmentalism or political orientation, the two most consistent predictors of climate beliefs in the previous literature (McCright et al. 2016a).

6.2 Contributions

This paper adds to the survey literature on climate perceptions in several ways. First, by simultaneously studying three important climate issues that previously have scarcely been studied together. Notably, we are, to our knowledge, the first study to link climate beliefs not only to policy support, but also to everyday climate-mitigation behavior. This focus is important since the evidence in the survey literature is still out concerning whether there is a gender gap in policy support and behavior.

The finding that climate beliefs drive the entire gender difference in climate policy support is perhaps the starkest result of our study, especially when contrasted with the fact that women will engage in more everyday climate-mitigation behavior than men with similar climate beliefs. This nuance was only discernable because we studied these three types of climate outcomes in a consistent framework, and future studies would benefit from the same approach. Policy-wise, these results suggest that better climate knowledge may be a powerful tool to induce greater climate policy support among both genders. That said, our finding that driving a car affects the willingness to support the taxation of harmful goods and alter transport behavior suggests that future studies should also examine how men's and women's willingness to alter mitigation behavior is affected by the costs of doing so (cf. Diederich and Goeschl 2017; Alberini et al. 2018), and whether adverse consequences in terms of a behavioral rebound effect differ by gender (cf. Dorner 2019).

This paper also contributes to the broader literature on gender differences in attitudes and behavior. As mentioned, while the previous literature suggests that some pro-environmental behaviors are considered more masculine than others (Swim et al. 2019), we find that men score lower on all six climate-mitigation behaviors than women, but also that stronger climate beliefs go some ways toward bridging this gap. From a policy perspective, this suggest that better climate knowledge may be a tool to induce greater more mitigation behavior among men. Future studies regarding other areas encompassing both the personal and the political should be on the lookout for similar gender differences, and the extent to which better knowledge can bridge them.

References

Alberini, A., M. Ščasný, & A. Bigano (2018). "Policy-v. individual heterogeneity in the benefits of climate change mitigation: Evidence from a stated-preference survey." *Energy Policy*, *121*, 565–575.

Aldy, J. E., Kotchen, M. J., & Leiserowitz, A. A. (2012). "Willingness to pay and political support for a US national clean energy standard". *Nature Climate Change*, *2*(8), 596–599.

Agarwal, B. (1994). "Gender and command over property: A critical gap in economic analysis and policy in South Asia." *World development*, *22*(10), 1455–1478.

Ali, F., Dissanayake, D., Bell, M. & Farrow, M. (2018). "Investigating car users' attitudes to climate change using multiple correspondence analysis." *Journal of Transport Geography*, *72*, 237-247.

Austin, K. F., & McKinney, L. A. (2016). "Disaster devastation in poor nations: the direct and indirect effects of gender equality, ecological losses, and development." *Social Forces*, *95*(1), 355–380.

Barber, B. M., & T. Odean (2001). "Boys will be boys: Gender, overconfidence, and common stock investment." *The Quarterly Journal of Economics*, 116(1), 261–292.

Barkenbus, J. N. (2010). "Eco-driving: An overlooked climate change initiative." *Energy policy*, *38*(2), 762–769.

Bell, S. E (2013). *Our roots run deep as ironweed: Appalachian women and the fight for environmental justice*. University of Illinois Press.

Bell, S. E., & Braun, Y. A. (2010). "Coal, identity, and the gendering of environmental justice activism in central Appalachia." Gender and Society, 24(6), 794–813.

Bonnie, R., Diamond, E. P., & Rowe, E. (2020). Understanding Rural Attitudes Toward the Environment and Conservation in America. Duke Nicholas Institute.

Bord, R.J. & R.E. O'Connor. (1997). "The Gender Gap in Environmental Attitudes: The Case of Perceived Vulnerability to Risk." *Social Science Quarterly*, 78(4), 830–840.

Brody, S.D., S. Zahran, A. Vedlitz, & H. Grover (2008). "Examining the relationship between physical vulnerability and public perceptions of global climate change in the United States." *Environment and Behavior*, 41, 72–95.

Buckingham, S. (2010). Call in the women. *Nature*, *468*(7323), 502–502.

Byrnes, J. P., D. C. Miller, & W. D. Schafer (1999). "Gender differences in risk taking: a meta-analysis." *Psychological Bulletin*, *125*(3), 367–383.

Czarnek, G., Kossowska, M., & Szwed, P. (2021). "Right-wing ideology reduces the effects of education on climate change beliefs in more developed countries." *Nature Climate Change*, *11*(1), 9-13.

Davidson, D. J. & W. R. Freudenberg (1996). "Gender and environmental risk concerns: A review and analysis of available research." *Environment and Behavior*, 28(3), 302–339.

Debono, R., K. Vincenti, & N. Calleja (2010). "Risk communication: climate change as a human-health threat, a survey of public perceptions in Malta." *European Journal of Public Health*, 22(1), 144–149.

Diederich, J., & T. Goeschl (2017). "To mitigate or not to mitigate: The price elasticity of pro-environmental behavior." *Journal of Environmental Economics and Management*, *84*, 209-222.

Dietz, T., L. Kalof, L., & P. C. Stern (2002). "Gender, values, and environmentalism." *Social Science Quarterly*, *83*, 353–364.

Dietz, T., Dan, A., & Shwom, R. (2007). "Support for climate change policy: social psychological and social structural influences." *Rural Sociology*, 72, 185–214.

Dorner, Z. (2019). "A behavioral rebound effect." *Journal of Environmental Economics and Management*, *98*, 102257.

Ergas, C., & York, R. (2012). “Women’s status and carbon dioxide emissions: A quantitative cross-national analysis.” *Social Science Research* 41(4), 965–976.

Ergas, C., P.T. Greiner, J.A. McGee, & M.T. Clement. (2021). “Does Gender Climate Influence Climate Change? The Multidimensionality of Gender Equality and Its Countervailing Effects on the Carbon Intensity of Well-Being.” *Sustainability* 13(7), 3956.

European Commission, 2011. “You Control Climate Change.” http://ec.europa.eu/clima/sites/campaign/index.htm

European Investment Bank (2018). EIB climate survey 1/6: Assessing citizens’ sentiments towards climate change. https://www.eib.org/en/surveys/citizens-climate-change-survey.htm

Gatersleben, B. (2013). “Measuring environmental behaviour.” In: Steg, L., van den Berg, A.E., de Groot, J.I.M. (Eds.), *Environmental Psychology. An Introduction*, pp. 131–140. Chichester, UK: BPS Blackwell.

Hansen, J.I., M. Satoa, & R. Ruedy (2012). “Perception of climate change. Earth, atmospheric and planetary sciences.” *Proceedings of the National Academy of Sciences of the United States of America (PNAS)*, 109(37), E2415–E2423.

Harrison, K. (2010). “The United States as outlier: economic and institutional challenges to US climate policy.” *Global Commons, Domestic Decisions: The Comparative Politics of Climate Change*, 67-103.

Hunter, L. M., A. Hatch, & A. Johnson (2004). “Cross-national gender variation in environmental behaviors.” *Social Science Quarterly*, *85*, 677-694.

IPCC (2021). Climate Change 2021 The Physical Science Basis. Working Group I contribution to the Sixth Assessment Report of the Intergovernmental Panel on Climate Change https://www.ipcc.ch/report/ar6/wg1/downloads/report/IPCC_AR6_WGI_Full_Report.pdf

Kormos, C., & R. Gifford (2014). “The validity of self-report measures of proenvironmental behavior: A meta-analytic review.” *Journal of Environmental Psychology*, *40*, 359–371.

Krosnick, J.A., A.L. Holbrook, L. Lowe, & P.S. Visser (2006). “The origins and consequences of Democratic Citizens’ Policy Agendas: A study of popular concern about global warming.” *Climatic Change*, 77, 7–43.

Kwauk, C., & A. Braga (2017). Three Platforms for Girls’ Education in Climate Strategies. *Washington, DC: Brookings*.

Leiserowitz, A. (2007). “American Opinions on Global Warming.” Yale University, Gallup & Clearvision Institute. http://environment.research.yale.edu/documents/downloads/a-g/AmericansGlobalWarmingReport.pdf

Li, J., J. Zhang, D. Zhang, D., & Q. Ji (2019). ”Does gender inequality affect household green consumption behaviour in China?” *Energy Policy*, *135*, 111071.

Lundeberg, M. A., P. W. Fox, & J. Punccohar (1994). “Highly Confident but Wrong: Gender Differences and Similarities in Confidence Judgments,” *Journal of Educational Psychology*, LXXXVI, 114–121.

Maddison, D. (2006). “The perception of and adaptation to climate change in Africa.” CEEPA Discussion Paper No. 10, Centre for Environmental Economics and Policy in Africa, University of Pretoria.

McCright, A.M. (2010). “The effects of gender on climate change knowledge and concern in the American public.” *Population and Environment*, 32(1), 66−87.

McCright, A.M., R.E. Dunlap, & C. Xiao, (2013). “Perceived scientific agreement and support for government action on climate change in the USA.” *Climatic Change*, 119, 511–518.

McCright, A.M., R.E. Dunlap, C. Xiao, (2014). “Increasing influence of party identification on perceived scientific agreement and support for government action on climate change in the USA, 2006–2012.” *Weather, Climate, and Society*, 6, 194–201.

McCright, A.M., S.T. Marquart-Pyatt, R.I. Shwom, S.R. Brechin, & S. Allen (2016a). “Ideology, capitalism, and climate: Explaining public views about climate change in the United States.” *Energy Research & Social Science*, 21, 180−189.

McCright, A.M., R.E. Dunlap, & S.T. Marquart-Pyatt (2016b). “Political ideology and views about climate change in the European Union.” *Environmental Politics*, 25(2), 338–358.

McKinney, L.A., & G.M. Fulkerson (2015). “Gender equality and climate justice: a cross-national analysis.” *Social Justice Research* 28(3), 293–317.

Meyer, A. G. (2021). Do economic conditions affect climate change beliefs and support for climate action? Evidence from the US in the wake of the Great Recession. *Economic Inquiry*, forthcoming.

Milfont, T. L., Zubielevitch, E., Milojev, P., & Sibley, C. G. (2021). Ten-year panel data confirm generation gap but climate beliefs increase at similar rates across ages. *Nature Communications*, *12*(1), 1-8.

Nakagawa, S., & Hart, C. (2019). “Where’s the beef? How masculinity exacerbates gender disparities in health behaviors.” *Socius, 5*, 2378023119831801.

Norgaard, K., & York, R. (2005). “Gender equality and state environmentalism.” *Gender and Society*, 19(4), 506–522.

Novus (2018). “Report on vegetarianism among Swedish residents”. https://www.djurensratt.se/sites/default/files/2018-06/vegoopinion-maj2018.pdf

O’Connor, R.E., R.J. Bord, B. Yarnal, & N. Wiefek (2002). “Who wants to reduce greenhouse gas emissions?” *Social Science Quarterly*, 83, 1–17.

O’Connor, R.E., R.J. Bord, & A. Fisher (1999). “Risk perceptions, general environmental beliefs, and willingness to address climate change.” *Risk Analysis*, 19(3), 461–471.

Ownershift (2019). “Who owns Sweden? A benchmark on ownership in Sweden.” Policy Report. http://ownershift.se/wp-content/uploads/2019/06/report-eng-web.pdf

Pisano, I., & M. Lubell (2017). “Environmental behavior in cross-national perspective: A multilevel analysis of 30 countries.” *Environment and Behavior*, *49*, 31–58.

Rijnhart, Judith J M & Twisk, Jos & Chin A Paw, Mai & Boer, Michiel & Heymans, Martijn. (2017). “Comparison of methods for the analysis of relatively simple mediation models”. *Contemporary Clinical Trials Communications*, 7, 130-135.

Schuler, S. R., & Hashemi, S. M. (1994). “Credit programs, women's empowerment, and contraceptive use in rural Bangladesh.” *Studies in Family Planning*, 25(2), 65–76.

Semenza, J.C., D.J. Wilson, J. Parra, B.D. Bontempo, M. Hart, D.J. Sailor, & L.A. George (2008). “Public perception of climate change: voluntary mitigation and barriers to behavior change.” *American Journal of Preventive Medicine*, 35, 479–487.

SEPA (2018). “The public’s views on climate 2018: A quantitative survey of the Swedish public’s views on climate solutions.” https://www.naturvardsverket.se/contentassets/6ffad3e6018c47cea06e6402f0eea066/rapport-allmanheten-klimatet-2018.pdf

Shiva, V., Mies, M. (2014). *Ecofeminism*. Zed Books Ltd..

Steg, L. (2016). “Values, norms, and intrinsic motivation to act proenvironmentally.” *Annual Review of Environment and Resources* 41, 277–292.

Stern, P. C., T. Dietz, & L. Kalof (1993). “Value orientations, gender, and environmental concern.” *Environment and Behavior, 25*(3), 322–348.

Stern, P. C., T. Dietz, T. Abel, G. A. Guagnano, & L. Kalof. (1999). “A value-belief-norm theory of support for social movement: The case of environmentalism.” *Human Ecology Review*, 6(2), 81–97.

Sundblad, E. L., Biel, A., & Gärling, T. (2007). ”Cognitive and affective risk judgements related to climate change.” *Journal of Environmental Psychology*, *27*(2), 97–106.

Swim, J. K. A. J. Gillis, & K. J. Hamaty (2019). “Gender Bending and Gender Conformity: The Social Consequences of Engaging in Feminine and Masculine Pro-Environmental Behaviors.” *Sex Roles*, 82, 363–385.

Terry, G. (2009). “No climate justice without gender justice: An overview of the issues.” *Gender and Development*, 17(1), 5–18.

van den Broek, K. L., I. Walker, & C. A. Klöckner (2019). “Drivers of energy saving behaviour: The relative influence of intentional, normative, situational and habitual processes.” *Energy Policy*, *132*, 811–819.

Vandenbergh, M. P., A. C. Steinemann (2007). “The carbon-neutral individual.” New York University Law Review 82, 1673–1745

Villagrasa, D. (2002). “Kyoto Protocol negotiations: reflections on the role of women.” *Gender & Development*, *10*(2), 40-44.

Warren, K. J. (1990). “The promise and power of ecofeminism.” *Environmental Ethics*, *12*(2).

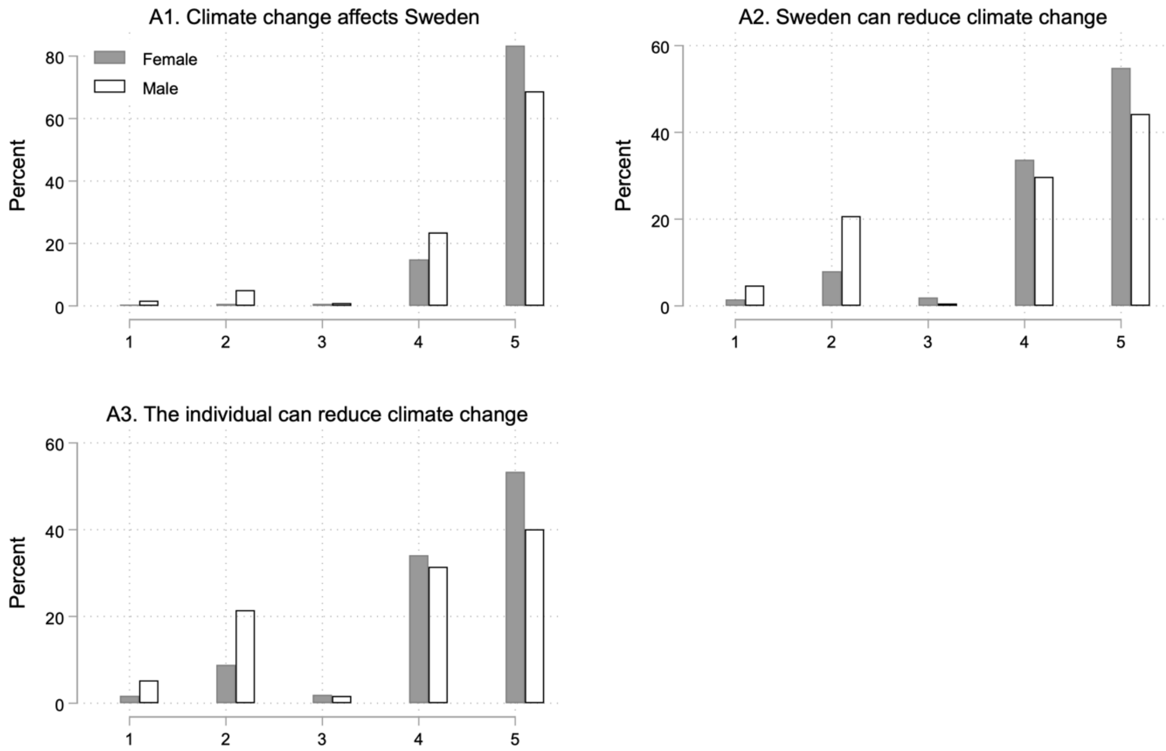
Xiao, C., & A. M. McCright (2014). “A test of the biographical availability argument for gender differences in environmental behaviors.” *Environment and Behavior*, *46*, 241–263.

Xiao, C., & A. M. McCright (2015). “Gender differences in environmental concern: revisiting the institutional trust hypothesis in the USA.” *Environment and Behavior* 47 (1), 17–37.

Zahran, S., S. D. Brody, H. Grover, & A. Vedlitz (2006). “Climate change vulnerability and policy support.” *Society and Natural Resources*, *19*(9), 771–789.

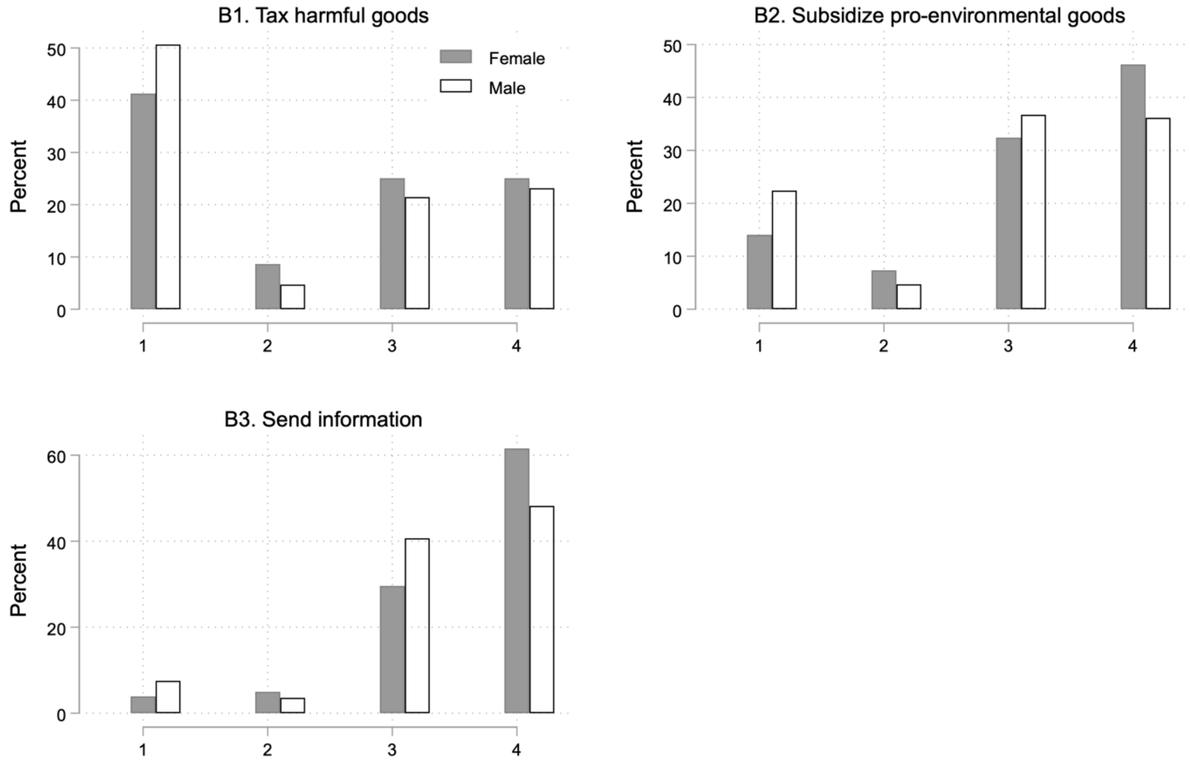
Appendix A

Figure A1. Histograms of responses to general climate questions (category A), by gender.



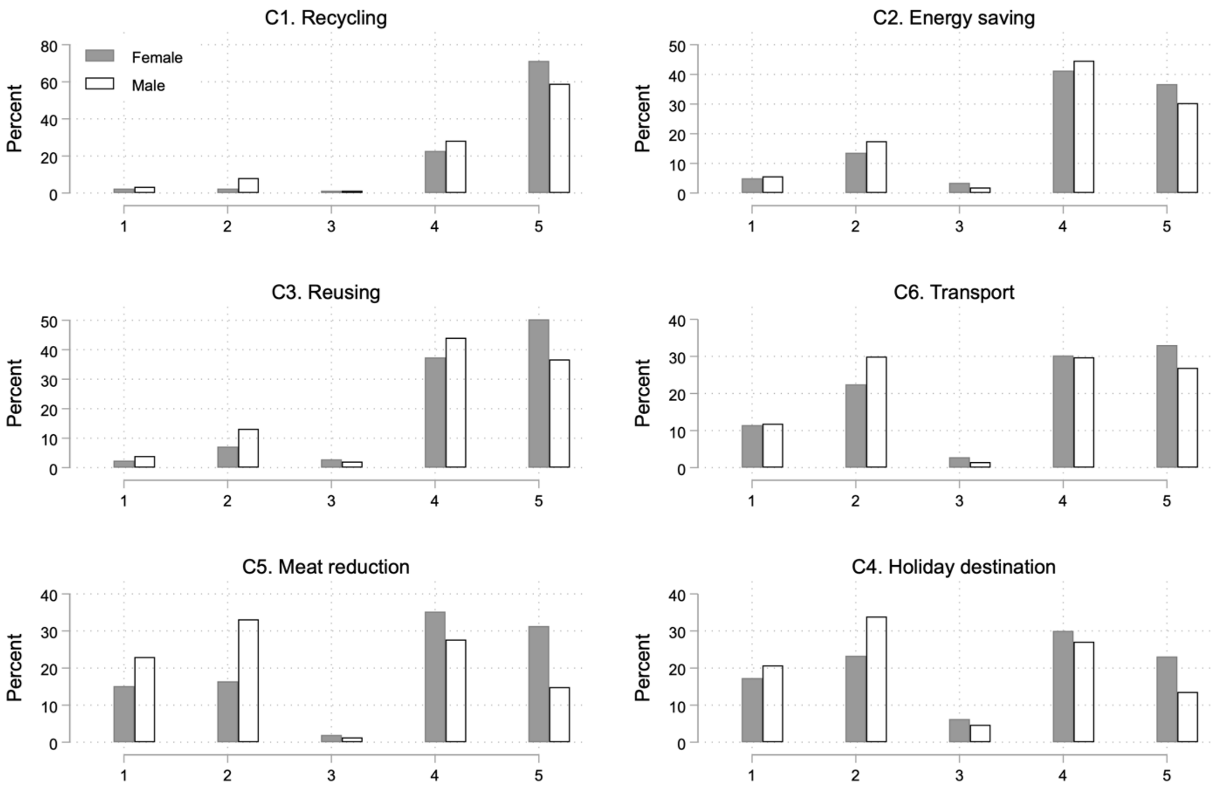
Note: Histograms depicting the percentage of answers by response level for each question in category (A), by gender. Likert scale response levels with increasing levels of agreement. Level 3 corresponds to the response "*Don´t know*".

Figure A2: Histograms of responses to climate policy questions (category B), by gender.



Note: Histograms depicting the percentage of answers by response level for each question in category (B), by gender. Likert scale response levels with increasing levels of agreement. Level 2 corresponds to the response "*Don´t know*".

Figure A3: Histograms of responses to self-reported measures (category C), by gender.



Note: Histograms depicting the percentage of answers by response level for each question in category (C), by gender. Likert scale response levels with increasing levels of agreement. Level 3 corresponds to the response "*Don´t know*".