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Private provision of public goods: Do individual climate protection efforts depend on perceptions of climate policy?

Abstract

This paper extends the economic literature on the private provision of public goods by examining the relevance of perceptions of climate policy to voluntary contributions to the public good of climate protection. Based on an analytical model which allows for perceptions of climate policy such as justification of international climate policy, procedural trust and procedural justice to affect voluntary climate protection activities, we examined data from representative surveys among citizens in the USA and Germany. Our microeconometric analysis confirmed the prediction that the perceived justification of international climate policy is positively related to voluntary contributions to climate protection in both countries. We also found empirical support (mainly for the USA) that higher perceived procedural justice lowers citizens' propensity to adopt climate protection activities. In contrast, we found no support that higher perceived procedural trust reduces citizens' propensity to adopt such measures. In a broad interpretation, our empirical results imply that individuals' perceptions about the process of providing public goods should also be considered when analyzing the factors explaining voluntary individual contribution to public goods.

1. Introduction

Standard economic theory predicts that individuals have virtually no incentive to voluntarily contribute to the provision of public goods (e.g., Holländer, 1990). Nonetheless, the findings from numerous free-rider experiments and stated preferences studies suggest that individuals do not only act in their own self-interest (e.g., Andreoni, 1988; Blanco et al., 2012; Kerr et al., 2012). Social preferences like prestige, respect, reputation, the contempt of others (e.g., Olson, 1965; Becker, 1974; Banerjee and Shogren, 2012), preferences for fairness, and inequity aversion (e.g., Fehr and Schmidt, 1999; Dannenberg et al., 2012), or feelings of warm glow (e.g., Andreoni, 1990) may motivate individuals to voluntarily contribute to the provision of a public good. Social preferences and feelings of warm glow also affect individual contributions to climate protection (e.g., Ahlheim and Schneider, 2002; Kotchen and Moore, 2008; Allcott, 2011; Araghi et al., 2014; Blasch and Farsi, 2014) – arguably the most prominent global public good.

In this paper, we study whether perceptions and beliefs associated with the process of the provision of a public good also influence the willingness for voluntary contributions to public goods. Specifically, we explore whether voluntary climate protection efforts depend on beliefs in the justification of international climate policy as well as on perceptions of procedural trust and procedural justice in international climate policy.

This research question is related to previous studies on the private provision of public goods in various ways. First, external interventions may enhance or decrease intrinsic motivation to adopt pro-social behavior (e.g., Frey, 1994; d'Adda, 2011). Extensive literature on motivation crowding has demonstrated that external circumstances like monetary incentives or institutional settings have the potential to change preferences or the perceived moral obligation. They can lead to crowding-in or crowding-out effects, i.e. enhance or reduce the motivation to voluntarily contribute to public goods (e.g, Frey and Oberholzer-Gee, 1997; Bohnet et al., 2001; Frey et al., 2001; Brekke et al., 2003; Nyborg and Rege, 2003). These findings imply that perceptions and beliefs associated with the public good providing process may influence private contributions positively or negatively. Additionally, Frey et al. (2001) showed that individuals derive procedural utility from the political process itself, while Nyborg and Rege (2003) and Bó et al. (2010) found that the presentation of a policy as well as the political institution influence moral motivation as much as the policy instrument itself.

Second, the belief or trust that the public good providing process leads to a socially optimal provision of the public good may motivate private contributions. Experimental studies on

'conditional cooperation' have found that people are more willing to contribute to charities and public goods if they observe, believe, or are informed that others are willing to do the same (e.g., Fischbacher et al., 2001; Kocher et al., 2008; Alpizar et al., 2008; Herrmann and Thöni, 2009; Khadjavi and Lange, 2013). Similarly, subjects who contribute and while becoming aware that others do not tend to lower their contributions (e.g., Fehr and Gächter, 2000). Conformity, social norms, or reciprocity have been found to be reasons to motivate this behavior (e.g., Frey and Meier, 2004; Nielsen et al., 2014). Thus, motivation of individuals to voluntarily contribute to public goods may depend on the perceived trustworthiness of their counterpart and beliefs about actual behavior and motives of others (e.g., Rabin, 1993; Berg et al., 1995; Fischbacher and Gächter, 2010).

While the existing literature provides evidence that perceptions of the public good providing process may indirectly affect individuals' contributions to a public good, such a direct link has not been established yet. This paper attempts to take a first step towards closing this gap by analyzing the impact of individuals' perceptions of international climate policy on their voluntary contributions to climate protection. We first derived testable predictions from a simple analytical model, in which a representative individual's utility function also includes perceptions of climate policy such as justification, procedural trust, and procedural justice. We then econometrically analyze the impact of these perceptions on the willingness to adopt five domestic climate protection activities. Our microeconometric analyses are based on unique data from representative surveys conducted simultaneously among citizens in the USA and Germany. Both countries are large emitters with strong political clout and are therefore considered to play a key role in future international climate policy. Thus, our analysis also allows for a comparison across countries.

The remainder of this paper is structured as follows: Section 2 offers a brief background on the state of the United Nations climate negotiations and also discusses justification, procedural trust and justice in international climate policy. Section 3 presents the analytical approach and derives the theoretical predictions for our empirical analyses. Section 4 describes the survey and our econometric approach. Results are presented in Section 5. The concluding Section 6 discusses our main findings and points out future research needs.

2. Climate policy background

Within the United Nations' climate negotiations, countries have recognized that the global mean temperature must not rise by more than 2°C above the pre-industrial level in order to

limit the dangerous impacts of anthropogenic climate change to acceptable levels (UNFCCC, 2009). However, the voluntary pledges countries made to the Copenhagen Accord in 2009 as well as decisions reached at subsequent climate summits are unlikely to be consistent with a path towards reaching the 2°C target (e.g., Höhne et al., 2012).Moreover, progress on a post Kyoto global climate agreement, which is scheduled to come into effect after 2020, has been slow. Industrialized countries (particularly the USA) fear that greenhouse gas emission targets may negatively affect the competitiveness of their economy (e.g., Pauwelyn, 2007) while emerging and developing countries (e.g. China) fear that emission targets will inhibit their future economic growth ("cap on development") (e.g., Banerjee, 2012). The widening gap between actual and required climate protection efforts raises doubts about the effectiveness of current international climate policy in general, thereby undermining its justification even for those not questioning the existence of climate change. Since decisions at the international level are executed at the domestic level, perceptions of justification may also influence the individual willingness for contributions to climate protection. The findings by Oberholzer-Gee et al. (1997), for instance, suggest that a higher social acceptability of international climate agreements may lead to a higher propensity to take on financial burdens associated with national implementations of international climate policy.

International climate policy is further characterized by a lack of procedural trust and procedural justice. Procedural trust is generally defined as the confidence in a structure or process (e.g., Furlong, 2005) but may also reflect a belief in sufficiently high social preferences of the counterpart instead of selfish motives (e.g., Sliwka, 2007). In climate policy, the free-rider problem in particular has challenged procedural trust. For example, the USA never ratified the Kyoto Protocol and in 2011 Canada formally withdrew from it. In 2012, Japan reneged on its voluntary pledge made under the Copenhagen Accord. Likewise, China for a long time refused to have its greenhouse gas emissions monitored by others and current dispute revolves around procedures to verify the pledges by third parties. In the absence of a supranational authority, however, procedural trust is particularly conducive to enforcing the outcomes of international climate negotiations. The belief in the process of international climate policy may lead to binding agreements even if parties distrust each other. The conceptual study by Pittel and Rübbelke (2013) pointed out that increased trust in the process of international climate policy and the perceived intentions of other countries may raise the probability of cooperative behavior in climate protection. Several studies also suggested that attitudes towards policy instruments are influenced by the trust in politicians or governments (e.g., Torgler and García-Valiñas, 2007; Jagers and Hammar, 2009; Jagers et al., 2010).

Unlike distributive justice (e.g., Lange et al., 2007, 2010; Dannenberg et al., 2010; Carlsson et al., 2013), procedural justice has gained little attention in the climate policy literature (e.g., Okereke, 2010; Schleich et al., 2014). The concept of procedural justice traces back to Lind and Tyler (1988) who find that individuals are often as concerned about the justice of the process as they are about the outcome itself. If individuals perceive the process as lacking fairness and transparency, they are reluctant to consider the outcomes as legitimate and acceptable. In the context of international climate policy, procedural justice requires adequate representation of all countries at the climate negotiations, in particular of those countries that are most vulnerable to climate change (e.g., Klinsky and Dowlatabadi, 2009; Okereke, 2010).

3. Analytical approach

We present a simple analytical model which allows perceptions of climate policy such as justification, procedural trust, and procedural justice to affect voluntary climate protection activities. The following equation describes the utility of a representative individual i with quasi-linear preferences defined over a numeraire good and a public good (i.e. climate protection in our case):

$$u_i = \omega_i - C_i(g_i) + \alpha B_i(\beta G_{-i} + g_i) \quad (1)$$

Here, ω_i is the fixed income by individual i, g_i reflects her voluntary contributions to the public good climate protection, and G_{-i} is the perceived contribution of the n-1 representative individuals in the other countries. $C_i(.)$ reflects the costs individual i faces when contributing to climate protection. Marginal costs are positive and increase for higher individual contributions to climate protection, i.e. $C_i'(.) > 0$ and $C_i''(.) > 0$. $B_i(.)$ captures the benefits that i enjoys from climate protection. Marginal benefits are positive and decrease in climate protection, i.e. $B_i'(.) > 0$ and $B_i''(.) < 0$.

By maximizing her utility, the individual may account for perceptions of international climate policy for two types of reasons. First, she may have little faith in the usefulness of international climate negotiations per se since she does not believe that protecting the climate is a public good problem. This perceived lack of justification of climate policy is assumed to be captured by the parameter α ($0 \leq \alpha \leq 1$) which discounts the perceived benefits from climate protection. Second, an individual may discount the contribution by representative individuals from other countries because she distrusts other countries' intentions and motives or disapproves of the procedures at international climate conferences. This perceived lack of proce-

dural trust and perceived lack of procedural justice is assumed to be captured by the parameter β ($0 \leq \beta \leq 1$).

The representative individual chooses her contribution to the public good g_i to maximize her utility. Assuming that she takes the contributions of the representative individuals of other countries as given leads to the First Order Condition for a Nash Equilibrium:

$$C'(.) = \alpha B'(.) \quad (2)$$

Totally differentiating Equation (2) and further assuming that the representative individuals are identical in all n countries, i.e. $G_{-i} = (n-1)g_i$, yields:

$$\frac{dg_i}{d\alpha} = \frac{B'(.)}{C''(.) - \alpha[\beta(n-1)+1]B''(.)} > 0 \quad (3)$$

$$\frac{dg_i}{d\beta} = \frac{\alpha(n-1)g_i B''(.)}{C''(.) - \alpha[\beta(n-1)+1]B''(.)} < 0 \quad (4)$$

Note that the numerator is positive by the Second Order Condition.

Our analytical model therefore leads to the following predictions:

Prediction 1: The propensity of individuals to voluntarily adopt climate protection activities increases with perceived justification of international climate policy.

Prediction 2: The propensity of individuals to voluntarily adopt climate protection activities decreases with perceived procedural trust in international climate policy.

Prediction 3: The propensity of individuals to voluntarily adopt climate protection activities decreases with perceived procedural justice in international climate policy.

In our econometric analysis, we test these theoretical predictions by examining the relationship between perceptions of international climate policy and the stated willingness of citizens in the USA and Germany to adopt five different climate protection activities.

4. Empirical analysis

4.1 Data and variables

Our data were collected in two representative online surveys of citizens aged 18 and older in the USA and Germany between May and June 2013. Both samples were drawn from the GfK (Gesellschaft für Konsumforschung) Online Panel. In the USA, 1010 respondents and in Germany 1005 respondents completed the self-administered questionnaire. The questions

referred to general assessments of climate change, specific voluntary climate protection activities, assessments of international climate policy and climate negotiations, basic values, as well as socio-demographic and socio-economic information. On average, respondents in the USA took 30.4 minutes to complete the survey and respondents in Germany 31.8 minutes. While we cannot entirely rule out potential self-selection, we have no evidence that our samples are not representative for the underlying populations of U.S. and German adults based on the socio-demographic characteristics age, income, education, marital status and household size.

The respondents in the survey were asked whether they planned to adopt the following climate protection activities: Actions to save energy at home, buy energy-efficient appliances, buy a fuel-efficient car, use or purchase renewable energy, and reduce the consumption of meat and dairy products. Based on the binary structure of the response options, we constructed the five dummy variables *energy savings*, *energy-efficient appliances*, *fuel-efficient car*, *renewable energy*, and *less meat or dairy products* that take the value one if the respondent plans to adopt the respective activity. These variables serve as the dependent variables in our econometric analyses. Table 1 reports the means for all variables across all respondents from the USA and Germany. Accordingly, the percentages range from 42% for *less meat or dairy products* in the USA to 87% for *energy savings* in Germany. For all activities, the percentages are higher for Germany than for the USA.

To elicit their perceptions of international climate policy, the survey asked respondents how strongly they agreed with particular statements on a symmetric scale with five ordered response categories.[1] In order to construct an indicator for perceived justification of international climate policy, we use the citizens' responses to the question "How important do you consider future international agreements are for combating climate change?".[2] On this basis, we construct the dummy variable *perceived justification* for the observations where "rather important" or "very important" is chosen as the answer. To create an indicator for perceived procedural trust in international climate policy we examine to which extent respondents agree with the statement "Commitments made at international climate negotiations will not be kept anyhow". The dummy variable *perceived procedural trust* is set equal to one if the respondent

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[1] Potential problems with this kind of scale are central tendency bias, acquiescence bias, and social desirability bias. These issues were addressed by "don't know/no answer" options to distinguish true neutral from unsure responses, a scale design involving balanced keying, and closed ended and "neutral" wording of the items (for a detailed discussion see also Schleich et al., 2014).

[2]. We differentiated between "very unimportant", "rather unimportant", "neither important nor unimportant", "rather important", and "very important". This question was only posed to the large majority of the respondents in both countries who stated that global climate change is already occurring or will occur in the future. Only these observations enter the econometric analysis.

agreed "very weakly" or "rather weakly".[3] Finally, to build an indicator for perceived procedural justice in international climate policy we rely on the respondents' agreement with the statement "All countries have the same opportunities to represent their interests at international climate conferences".[4] The dummy variable *perceived procedural justice* equals one for the answer categories "rather strongly" or "very strongly.

Table 1 reports the means for these three variables. The percentages for perceived justification are fairly high, for perceived procedural trust are moderate, and for perceived procedural justice are quite low in both countries, but responses differed slightly across countries. Our theoretical approach predicts a positive effect of *perceived justification* and negative effects of *perceived procedural trust* and *perceived procedural justice* on the propensity to adopt climate protection activities.

In addition to these climate policy indicators, our econometric analysis included a wide range of control variables. The first group of control variables refers to the estimated contribution and the estimated financial consequences of the climate protection activities. The dummy variables *effectiveness* takes the value one if the respondent believed the respective activity to contribute "rather a lot" or "a lot" to climate protection[5]. Similarly, *financial advantage* is equal to one, if a respondent believed an activity to provide personal financial advantages.[6] The second group of control variables includes two indicators for environmental preferences, *environmental awareness* and *identification with green politics*. *Environmental awareness* is constructed by adding up the values of six dummy variables, which are based on six statements from the new ecological paradigm (NEP) scale for measuring environmental concern (Dunlap et al., 2000).[7] Thus, *environmental awareness* ranges from zero (lowest level) to six (highest level). If a respondent strongly or rather strongly identified herself with green poli-

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[3] Since these statements are negatively keyed we used these answer categories to reflect weak agreement.

[4] In all cases we differentiated between "very weakly", "rather weakly", "neither weakly nor strongly", "rather strongly", and "very strongly".

[5] The underlying question is "How much do you believe the following measures contribute to climate protection" with the following five ordered response categories: "very little", "rather little", "neither a little nor a lot", "rather a lot", and "a lot".

[6] The underlying question is "In your opinion, do the following measures provide rather financial advantages (e.g. saving money, financial gains) or rather financial disadvantages (e.g. costs) for you personally" with the following three ordered response categories: "rather financial disadvantages", "neither financial advantages nor disadvantages", and "rather financial advantages".

[7] The underlying six statements are "Humans have the right to modify the natural environment to suit their needs", "Humans are severely abusing the planet", "Plants and animals have the same right to exist as humans", "Nature is strong enough to cope with the impacts of modern industrial nations", "Humans were meant to rule over the rest of nature", and "The balance of nature is very delicate and easily upset" with the five ordered response categories "very weakly", "rather weakly", "neither weakly nor strongly", "rather strongly", and "very strongly". The corresponding dummy variables takes the value one if the respondent agreed to the respective statement rather strongly or very strongly or (in the case of negative keying) rather weakly or very weakly, respectively.

tics, *identification with green politic* equals one.[8] *Effectiveness, financial advantage, environmental awareness* and *identification with green politics* are expected to positively affect the adoption of climate protection activities.

The third group of control variables captures the socio-demographic characteristics of the respondents. *Age* as measured in years and varies between 18 and 85 in the USA and between 18 and 89 in Germany. *Female* is equal to one if the respondent is a woman. The *number of children* varies between zero and eleven in the USA and between zero and five in Germany. *Living together* takes the value one for the respondent's marital status "living with a partner" or "married", and zero otherwise. *High education* equals one if the respondent is qualified to pursue a degree in higher education (i.e. high-school degree in the USA and "Abitur" in Germany). The final group of control variables reflects regional heterogeneity. For the USA, we include the dummy variables *west, midwest, northeast,* and *south* (which is treated as the omitted category in the econometric analysis). For Germany, we add the dummy variable *west* for respondents living in Western Germany. Our econometric approaches comprise single-country models and combined-countries model. For analyses involving observations from both countries, the dummy variable *USA* stands for respondents from the USA.

4.2 Econometric approaches

In our econometric analyses, the dependent variables y_{ij} are dummies which indicated whether a citizen $i = 1,…, n$ plan to adopt a climate protection activity $j = 1,…, 5$. Our first econometric approach involved stacking the data over all activities j. This *aggregate model* allows estimating the determinants of the propensity to adopt one of the five climate protection activities. Therefore, the underlying unobservable latent variable for each citizen i is:

$$y_{ij}^* = \beta_j' \, x_{ij} + \varepsilon_{ij} \quad (5)$$

For each citizen i and for each planned climate protection activity j, the vector x_{ij} comprises a set of explanatory variables with the unknown parameter vector β_j. The dummy variables y_{ij} takes the value one if $y_{ij}^* > 0$. $P(y_{ij} = 1)$ denoted the probability that citizen i plans to adopt the climate protection activity j. We employ binary probit models, i.e. the error terms ε_{ij} are assumed to be normally distributed. To account for unobserved heterogeneity over the different

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[8] The underlying statement was "I identify myself closest with green politics" with the five ordered response categories "very weakly", "rather weakly", "neither weakly nor strongly", "rather strongly", and "very strongly".

activities, we estimated binary random effects probit models.[9] Unobserved heterogeneity is incorporated in the error terms ε_{ij} and is assumed to be uncorrelated with the explanatory variables in x_{ij}. Stacking the data across activities allows us to capture differences in the conditional means of the activity dummies. In all cases, the dummy variable *less meat or dairy products* is treated as the omitted category, i.e. the activity with the smallest percentage in both countries (see Table 1).

In our second econometric approach, we employ binary probit models without random effects to separately estimate the determinants of the planned adoption of the five climate protection activities. Thus, unlike the *aggregate model*, this *single activity model* does not assume the parameter estimates to be identical across the activities. Since the decision to adopt a particular climate protection activity may depend on the choices for the other activities, the use of univariate binary probit models can lead to biased and inconsistent parameter estimations (e.g., Greene, 2012). We therefore used multivariate binary probit models, where the error terms captured possible correlations between the dependent variables.[10]

5. Results

5.1 Aggregate model

Table 2 reports the estimation results for the binary random effects probit models. The first column refers to the combined-countries model, while the second and third columns refer to the single-country models for the USA and for Germany. Statistical tests imply that the hypothesis of no unobserved heterogeneity can be rejected at the 1% significance level, which supports the random effects specification. The results for all combined and single-countries models suggest that the propensity to save energy at home, buy energy-efficient appliances, buy a fuel-efficient car, and use or purchase energy from renewable sources is significantly higher than to reduce the consumption of meat and dairy products.

The estimated parameters for *effectiveness of activity* shows the expected positive sign and are significantly different from zero for the combined-countries model and for Germany but not for the USA alone. Thus, the findings for the combined-countries model are mostly driven by the observations from Germany. As expected, *financial advantages* and *identification with green politics* has a significantly positive impact on the planned adoption of climate protec-

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[9]All maximum likelihood and (in the case of multivariate binary probit models) simulated maximum likelihood estimations were carried out with STATA 12.

[10] The simulated maximum likelihood estimations relied on robust estimations of the standard deviation of the parameter estimates.

tion activities in all models. In contrast, *environmental awareness* only has a significant effect in the combined-countries model.[11] In general, socio-demographic variables only show weak correlations with the planned adoption of climate protection activities. For the combined-countries model we find that women and citizens with high education are significantly more likely to adopt climate protection activities. Higher education also had a significantly positive effect in Germany but not in the USA.

Our main interest is directed at the impact of perceptions of international climate policy. The estimated parameter for *perceived justification* is positive and significantly different from zero for the combined-countries and both single-country models. This finding provides strong support for our first prediction. In addition, *perceived procedural justice* has a significantly negative effect on the planned adoption of climate protection activities in the combined-countries model and in the single-country model for the USA but not for Germany. In general though, these findings confirm our third prediction, at least for the USA.

In contrast, *perceived procedural trust* does not have a significant effect in any model. Hence, we find no empirical support for our second prediction. We also experimented with three additional indicators for perceived procedural trust in international climate policy, derived from the respondents' view on the self-interested use of international climate negotiations by richer or by poorer countries and on the use of international climate negotiations by governments to pacify their citizens. However, none of these three variables was found to have a robust separate effect on the planned adoption of climate protection activities. When all four indicators reflecting procedural trust are included, the null hypothesis that the four parameters were jointly zero cannot be rejected at common significance levels on the basis of Wald and likelihood ratio tests for any of the models.[12]

To assess the robustness of our aggregate model results, we conducted several additional estimations. The results of estimating binary random effects logit models are qualitatively almost identical to those of the binary random effects probit models.[13] While this finding is to be expected given the similar distributions of the error terms in the logit and probit models, it should be noted that both models are based on the potentially restrictive assumption that the

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[11] The insignificant effects in the separate countries models are possibly due to multicollinearity problems with *identification with green politics*.

[12] We also tested whether the negatively keyed statement, which underlies *perceived procedural trust* contributed to the insignificant parameter estimate. For this, we assigned the value of one to a new dummy variable if the respondent agreed "very weakly", "rather weakly", or "neither weakly nor strongly" to the underlying statement. The parameter estimate associated with this new variable is also statistically insignificant.

[13] All results which are not reported for brevity are available upon request from the authors.

unobserved heterogeneity is uncorrelated with the explanatory variables.[14] Although we rejected the hypothesis of no unobserved heterogeneity in the random effects binary probit model at common significance levels, we estimated pooled binary probit models as a further robustness check. Qualitatively, the estimation results for perceptions of justification, procedural trust, and procedural justice are almost identical to the results reported in Table 2.[15]

Our estimation results are very robust when we include additional control variables such as household size[16]. In addition, we also analyzed two further climate protection activities, i.e. reducing car use and reducing the number of flights so that the data were stacked over seven activities. The estimation results for perceptions of justification, procedural trust, and procedural justice with these seven activities are qualitatively very similar to the results presented in Table 1. However, it should be noted that the underlying questions for these two activities were filtered, i.e. only citizens who reported a positive number of kilometers or a positive number of flights could answer these questions so that the number of observations is considerably lower in this case.

As an alternative to the binary probit specification, we also estimated the aggregate model as a count data model, where the dependent variable was the number of activities. Since the dummy variables *effectiveness of activity* and *financial advantage of activity* were measured for a specific activity, they could not be included in these models. To allow for underdispersion, we used generalized poisson models instead of the usual poisson models which assume equidispersion. While these count data models explicitly weighted all five activities equally by simply adding them up, the estimation results for perceptions of justification, procedural trust, and procedural justice were qualitatively almost identical to the results reported in Table 2.

5.2 Single activity models

The estimation results of the multivariate binary probit models are reported in Table 3 for the combined-countries model and in Table 4 and Table 5 for the single-country models. To save

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[14] Applying binary fixed effects logit models is not appropriate in our case since these models can only include explanatory variables which vary across the different climate protection activities, whereas our main interesting indicators for perceptions of international climate policy are invariant.

[15] This pooled approach provided more significant effects for some control variables, however. For example, the coefficients for *environmental awareness*, *living together*, and *high education* were positive and significant, and the coefficient for *age* was negative and significant. Of course, these differences may have been due to erroneously neglecting random effects in the pooled binary probit models.

[16] Our results are also very robust when we include warm glow motives and other factors as explanatory variables in their analysis of CO_2 compensation measures (Lange et al., 2014).

space, we do not report the estimated correlation coefficients in the error terms between the five dependent dummy variables. Since several correlations were significantly different from zero, we prefer the multivariate binary probit model over the univariate binary probit models.[17]

In general, the findings for the single activity models are quite similar to those for the aggregate model, in particular for the effects of climate policy. *Perceived justification* of climate policy is significant for all climate protection activities (except for *less meat or dairy products*) in the combined-countries model and in both single-country models. *Perceived procedural justice* turns out to be significant for all activities (except for *less meat or dairy products*) in the combined-countries model and for all activities (except for *renewable energy* and *less meat or dairy products*) in the single-country model for the USA. For Germany, *perceived procedural justice* is found to be statistically significant for *energy savings* only. Also in line with the findings for the aggregate model, *perceived procedural trust* is not consistently significant in any single activity model. Among the climate protection activities considered in the single activity models, only the consumption of meat and dairy products does not appear to be related to the perception of climate policy, neither among US nor German citizens.

Consistent with the findings for the aggregate model, *financial advantages* of the climate protection activities exhibits significantly positive effects on the planned adoption of climate protection activities in the single activity models for all activities in the combined-countries model and in the single-country model for the USA as well as for most activities in the single-country model for Germany. In contrast, the impact of *effectiveness* and of *identification with green politics* appears to be more heterogeneous across activities. Both variables seem to matter primarily for the planned adoption of *renewable energy* and *less meat and dairy products*.

The findings for the single activity models further suggest that the effect of socio-economic characteristics on the adoption of climate protection measures vary by activity and country. Younger citizens are found to be more likely to adopt *renewable energy* (in all single activity models) and *energy savings* for the single activity model for Germany. For all other activities and models, *age* was not found to be statistically significant. Similarly, women appeared to be more likely to consume *less meat* and *dairy products* in all single activity models and to en-

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[17] In the multivariate probit models, the number of respondents is slightly lower than in the random effects binary probit analysis since only observations of those citizens with complete information about all dependent and explanatory variables can be included. In order to test the robustness of these results, we also analyzed univariate binary probit models for each activity. The corresponding estimation results were qualitatively almost identical to those discussed in Section 5.2.

gage in *energy saving* activities in the singly-country model for the USA. For the adoption of other activities and models, our gender variable is not statistically significant. The *number of children* also fails to be statistically significant for any activity or model. In contrast, and in line with the findings for the aggregate model, *living together* is positively related to most climate protection activities in the combined-countries model and in the single activity model for Germany but not for the USA. Finally, we find that *high education* increases the likelihood of adopting a *fuel-efficient car* in the combined-countries model and for the singly-country model for the USA but not for Germany. Similarly, *high education* is positively related to the adoption of *renewable energy* in the combined-countries model and for the singly-country model for Germany but not for the USA.

6. Conclusions

A substantial body of theoretical and empirical economic studies, focusing on social objectives or warm glow motives, has identified factors that help explain the private provision of public goods. We extend this literature by examining the relevance of perceptions of international climate policy to voluntary contributions to the global public good climate protection. Specifically, we explored whether beliefs in the justification of international climate policy as well as perceptions of procedural trust and procedural justice in international climate policy affected individuals' voluntary climate protection efforts. Based on an analytical model which allows perceptions of climate policy to affect voluntary climate protection activities, we derived three predictions. The findings from our microeconometric analysis of representative surveys among citizens in the USA and Germany suggest that *perceived justification* of climate policy increases the propensity to adopt climate protection measures (prediction 1) in both countries. Also, a higher *perceived procedural justice* appears to be related to a lower propensity to adopt climate protection measures (prediction 2), in particular in the USA. Third, we found no empirical support that higher *perceived procedural trust* reduces the propensity to adopt climate protection measures (prediction 3). These findings are robust to a wide range of alternative specifications, including aggregate and single activity models.

Our findings are closely related to the literature on motivation crowding: Our results imply that voluntary climate protection activities are motivated by the belief in the justification of international climate policy which leads to a crowding-in effect. In contrast, a perceived lack of procedural justice seems to be compensated by a higher propensity to adopt additional climate protection activities thus leading to a crowding-out effect.

In a broad interpretation, our empirical results imply that individuals' perceptions about the process of providing public goods should also be considered when analyzing the factors explaining the voluntary contribution to public goods. Future research could therefore explore this relationship for the private provision of public goods other than climate protection, such as voluntary donations for social or ethical purposes. Future research could also allow for a richer set of items capturing the various facets of climate policy than that included in our study to better understand the relationship between the perception of international climate policy and the adoption of climate protection activities.

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Appendix

Table 1: Number of respondents and mean for all variables

Variables	Both countries		USA		Germany	
	Number of respondents	Mean	Number of respondents	Mean	Number of respondents	Mean
Planned climate protection activities						
Energy savings	1,938	0.84	965	0.81	973	0.87
Energy-efficient appliances	1,921	0.81	952	0.78	969	0.84
Fuel-efficient car	1,844	0.69	915	0.67	929	0.71
Renewable energy	1,832	0.56	890	0.50	942	0.62
Less meat or dairy products	1,903	0.46	939	0.42	964	0.50
Perceived justification	1,569	0.80	698	0.73	871	0.86
Perceived procedural trust	1,776	0.10	855	0.13	921	0.07
Perceived procedural justice	1,758	0.42	858	0.49	900	0.36
Effectiveness of activity						
Energy savings	1,888	0.61	924	0.61	964	0.61
Energy-efficient appliances	1,892	0.62	926	0.63	966	0.61
Fuel-efficient car	1,874	0.62	918	0.61	956	0.63
Renewable energy	1,824	0.64	875	0.60	949	0.67
Less meat or dairy products	1,795	0.30	847	0.25	948	0.35
Financial advantage of activity						
Energy savings	1,875	0.79	919	0.76	956	0.81
Energy-efficient appliances	1,870	0.67	914	0.73	956	0.62
Fuel-efficient car	1,789	0.63	877	0.66	912	0.61
Renewable energy	1,692	0.39	813	0.50	879	0.29
Less meat or dairy products	1,730	0.38	833	0.39	897	0.37
Environmental awareness	1,833	3.58	905	3.07	928	4.08
Identification with green politics	1,845	0.26	907	0.21	938	0.30
Age	2,015	44.83	1,010	48.51	1,005	41.13
Female	2,015	0.51	1,010	0.53	1,005	0.49
Number of children	2,015	1.14	1,010	1.32	1,005	0.95
Living together	2,008	0.62	1,006	0.62	1,002	0.63
High education	2,006	0.61	1,006	0.68	1,000	0.55
USA	2,015	0,50				
West			1,010	0.22	1,005	0.79
Midwest			1,010	0.23		
Northeast			1,010	0.20		
South			1,010	0.35		

Table 2: Maximum likelihood estimates (z-statistics) of the parameters in binary random effects probit models, determinants of the general propensity to adopt one of the five climate protection activities

Explanatory variables	Combined-countries model	Single-country model: USA	Single-country model: Germany
Planned climate protection activities			
Energy savings	1.24***	1.27***	1.25***
	(0.08)	(0.12)	(0.10)
Energy-efficient appliances	1.12***	1.12***	1.15***
	(0.07)	(0.11)	(0.09)
Fuel-efficient car	0.62***	0.78***	0.54***
	(0.07)	(0.11)	(0.09)
Renewable energy	0.28***	0.30***	0.26***
	(0.06)	(0.10)	(0.08)
Perceived justification	0.34***	0.35**	0.36***
	(0.09)	(0.14)	(0.13)
Perceived procedural trust	0.00	-0.04	0.09
	(0.11)	(0.16)	(0.16)
Perceived procedural justice	-0.26***	-0.35***	-0.14
	(0.07)	(0.12)	(0.09)
Effectiveness of activity	0.34***	0.10	0.46***
	(0.06)	(0.10)	(0.07)
Financial advantage of activity	0.41***	0.64***	0.26***
	(0.05)	(0.09)	(0.07)
Environmental awareness	0.05**	0.06*	0.04
	(0.02)	(0.03)	(0.03)
Identification with green politics	0.33***	0.25*	0.37***
	(0.08)	(0.14)	(0.09)
Age	-0.00	-0.00	-0.01
	(0.00)	(0.00)	(0.00)
Female	0.14*	0.12	0.14*
	(0.07)	(0.11)	(0.09)
Number of children	0.04	0.03	0.04
	(0.03)	(0.04)	(0.04)
Living together	0.13*	0.09	0.17*
	(0.07)	(0.12)	(0.09)
High education	0.16**	0.14	0.16*
	(0.07)	(0.12)	(0.09)
Northeast		-0.17	
		(0.15)	
Midwest		-0.11	
		(0.15)	
West		-0.28*	0.23**
		(0.15)	(0.10)
USA	-0.06		
	(0.08)		
Constant	-0.71***	-0.57**	-0.92***
	(0.17)	(0.27)	(0.23)
Number of observations	6,195	2,643	3,552
Number of respondents	1,315	568	747

* (**, ***) means that the appropriate parameter or effect is different from zero at the 10% (5%, 1%) significance level.

21

Table 3: Simulated maximum likelihood estimates (z-statistics) of the parameters in the multivariate binary probit model, determinants of the adoption of five climate protection activities separately, combined-countries model for the USA and Germany

Explanatory variables	Energy savings	Energy-efficient appliances	Fuel-efficient car	Renewable energy	Less meat or dairy products
Perceived justification	0.50***	0.47***	0.48***	0.38***	0.11
	(0.13)	(0.12)	(0.11)	(0.11)	(0.11)
Perceived procedural trust	-0.03	0.10	0.02	0.33**	0.06
	(0.16)	(0.16)	(0.14)	(0.14)	(0.14)
Perceived procedural justice	-0.54***	-0.23**	-0.26***	-0.23***	-0.07
	(0.10)	(0.10)	(0.09)	(0.09)	(0.09)
Effectiveness of activity	0.16*	-0.14	0.10	0.28***	0.65***
	(0.10)	(0.10)	(0.09)	(0.09)	(0.09)
Financial advantage of activity	0.33***	0.15*	0.29***	0.18**	0.30***
	(0.10)	(0.09)	(0.08)	(0.09)	(0.08)
Environmental awareness	0.07**	0.05*	0.04	0.01	0.04
	(0.03)	(0.03)	(0.03)	(0.03)	(0.02)
Identification with green politics	-0.14	0.10	0.03	0.44***	0.40***
	(0.11)	(0.11)	(0.10)	(0.10)	(0.09)
Age	-0.00	-0.00	0.00	-0.01***	-0.00
	(0.00)	(0.00)	(0.00)	(0.00)	(0.00)
Female	0.20*	0.09	-0.01	-0.01	0.21**
	(0.10)	(0.10)	(0.09)	(0.09)	(0.08)
Number of children	0.01	0.04	-0.00	0.05	0.00
	(0.04)	(0.05)	(0.04)	(0.04)	(0.04)
Living together	0.20*	0.24**	0.18*	0.18*	0.06
	(0.11)	(0.10)	(0.09)	(0.09)	(0.09)
High education	0.16	0.10	0.24**	0.21**	-0.02
	(0.10)	(0.10)	(0.09)	(0.09)	(0.09)
USA	0.14	-0.02	0.13	0.00	0.06
	(0.11)	(0.10)	(0.10)	(0.10)	(0.09)
Constant	0.23	0.32	-0.43**	0.05	-0.56***
	(0.23)	(0.22)	(0.21)	(0.21)	(0.21)
Number of observations			1,052		

* (**, ***) means that the appropriate parameter or effect is different from zero at the 10% (5%, 1%) significance level.

Table 4: Simulated maximum likelihood estimates (z-statistics) of the parameters in the multivariate binary probit model, determinants of the adoption of five climate protection activities separately, single-country models for the USA

Explanatory variables	Energy savings	Energy-efficient appliances	Fuel-efficient car	Renewable energy	Less meat or dairy products
Perceived justification	0.50***	0.53***	0.46***	0.37**	0.16
	(0.18)	(0.17)	(0.17)	(0.16)	(0.16)
Perceived procedural trust	-0.25	0.09	-0.02	0.42**	-0.04
	(0.21)	(0.22)	(0.20)	(0.18)	(0.18)
Perceived procedural justice	-0.63***	-0.39***	-0.38**	-0.21	-0.02
	(0.16)	(0.15)	(0.15)	(0.14)	(0.13)
Effectiveness of activity	0.12	-0.60***	-0.05	0.12	0.37**
	(0.15)	(0.16)	(0.15)	(0.14)	(0.15)
Financial advantage of activity	0.27*	0.33**	0.33**	0.38***	0.37***
	(0.15)	(0.14)	(0.13)	(0.12)	(0.13)
Environmental awareness	0.12**	0.15***	0.12**	0.01	0.00
	(0.05)	(0.05)	(0.05)	(0.04)	(0.04)
Identification with green politics	-0.28	0.01	-0.09	0.35**	0.31**
	(0.17)	(0.17)	(0.16)	(0.16)	(0.16)
Age	-0.00	-0.00	0.00	-0.02***	-0.00
	(0.01)	(0.01)	(0.01)	(0.01)	(0.00)
Female	0.28*	0.04	0.04	-0.04	0.24*
	(0.15)	(0.15)	(0.14)	(0.13)	(0.13)
Number of children	0.01	0.05	0.00	0.06	0.01
	(0.06)	(0.06)	(0.06)	(0.05)	(0.05)
Living together	0.10	0.20	0.23	-0.11	-0.06
	(0.16)	(0.15)	(0.14)	(0.14)	(0.13)
High education	-0.13	0.01	0.45***	0.14	0.15
	(0.17)	(0.16)	(0.15)	(0.14)	(0.14)
Northeast	-0.42**	-0.01	-0.10	0.01	-0.16
	(0.20)	(0.21)	(0.19)	(0.18)	(0.17)
Midwest	-0.27	-0.28	0.00	-0.25	-0.16
	(0.21)	(0.19)	(0.19)	(0.17)	(0.17)
West	-0.47**	-0.48***	-0.04	-0.12	-0.06
	(0.20)	(0.18)	(0.18)	(0.17)	(0.16)
Constant	0.81**	0.49	-0.43	0.51*	-0.37
	(0.35)	(0.33)	(0.33)	(0.31)	(0.30)
Number of observations			448		

* (**, ***) means that the appropriate parameter or effect is different from zero at the 10% (5%, 1%) significance level.

Table 5: Simulated maximum likelihood estimates (z-statistics) of the parameters in the multivariate binary probit model, determinants of the adoption of five climate protection activities separately, single-country model for Germany

Explanatory variables	Energy savings	Energy-efficient appliances	Fuel-efficient car	Renewable energy	Less meat or dairy products
Perceived justification	0.60***	0.53***	0.48***	0.38**	0.10
	(0.18)	(0.18)	(0.17)	(0.17)	(0.16)
Perceived procedural trust	0.10	0.06	0.05	0.26	0.26
	(0.27)	(0.23)	(0.22)	(0.22)	(0.21)
Perceived procedural justice	-0.45***	-0.05	-0.17	-0.16	-0.01
	(0.14)	(0.14)	(0.12)	(0.12)	(0.12)
Effectiveness of activity	0.15	0.11	0.21*	0.34***	0.80***
	(0.13)	(0.13)	(0.12)	(0.12)	(0.12)
Financial advantage of activity	0.30**	0.06	0.29**	0.08	0.25**
	(0.14)	(0.13)	(0.11)	(0.12)	(0.12)
Environmental awareness	0.05	0.00	0.01	0.03	0.07**
	(0.04)	(0.04)	(0.03)	(0.03)	(0.03)
Identification with green politics	-0.07	0.18	0.07	0.40***	0.39***
	(0.15)	(0.15)	(0.13)	(0.13)	(0.12)
Age	-0.01**	-0.01	0.00	-0.02***	-0.01
	(0.01)	(0.01)	(0.01)	(0.01)	(0.01)
Female	0.12	0.07	-0.06	0.03	0.20*
	(0.14)	(0.14)	(0.12)	(0.12)	(0.11)
Number of children	0.05	0.07	0.03	0.07	0.02
	(0.07)	(0.07)	(0.06)	(0.06)	(0.06)
Living together	0.24*	0.24*	0.09	0.32***	0.16
	(0.15)	(0.14)	(0.13)	(0.12)	(0.12)
High education	0.30**	0.16	0.09	0.26**	-0.12
	(0.14)	(0.14)	(0.12)	(0.12)	(0.12)
West	0.12	0.03	0.19	0.35***	0.41***
	(0.16)	(0.16)	(0.14)	(0.14)	(0.14)
Constant	0.32	0.41	-0.34	-0.39	-1.01***
	(0.34)	(0.37)	(0.31)	(0.31)	(0.31)
Number of observations			604		

* (**, ***) means that the appropriate parameter or effect is different from zero at the 10% (5%, 1%) significance level.