# Consistency and variation in the associations between Refugee and environmental attitudes in European mass publics

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## ABSTRACT

We investigated the associations between refugee and environmental attitudes among 36876 respondents from 20 countries included in the European Social Survey Round 8 (2016). Three preregistered hypotheses were supported: (H1) there was a positive association between these attitudes across countries (meta-analytical partial correlation = .16), (H2) anti-immigration party voters held more negative environmental attitudes, and (H3) pro-environmental party voters held more positive refugee attitudes. Against our predictions, the linear association between refugee and environmental attitudes was not moderated by political affiliation (H4) or political engagement (H5). Exploratory analyses further showed that these attitudes were more strongly associated among the young, the more educated, and among the most extreme populist right voters.

## 1. Introduction

Madison Grant was a leading thinker, activist, and household name in 1920s America. As a eugenicist, Grant espoused scientific racism, co-founded the American Eugenics Society, and was vice president of the Immigration Restriction League. Adolf Hitler referred to his work as the "bible" (Burgers, 2011). As a conservationist, he was co-founder of the Save the Redwoods League and the National Parks Association, and is credited with the saving of many different species of animals (Spiro, 2009). For Grant, eugenics was a way of ensuring the survival of the race that had made the United States a prosperous country. Conservation, on the other hand, was a way of preserving the land with which nature-—and natural selection—had endowed them. Even as Grant developed the idea of the blond-haired, blue-eyed Nordics as the "master race", he simultaneously endeavored to preserve the natural beauty of nature (Spiro, 2009).

Refugee migration, and especially the so-called "refugee crisis" in 2015, has fueled the rise of far-right parties in European politics. These parties have successfully leveraged natives' anxieties to mobilize voters and enact more restrictive asylum policies (*e.g.*, for a recent review, see Hangartner, Dinas, Marbach, Matakos, & Xefteris, 2019). Although they share Madison Grant's contention that the state should protect endangered white men (Spiro, 2009), they do not share his belief in the necessity of environmental action. Quite the opposite, at a time in which

most political parties understand or accept climate change as an imminent environmental catastrophe, the nationalist far-right stands out due its denial of climate change, its anthropogenic causes, and its negative consequences (Lockwood, 2018; Schaller & Carius, 2019).

As noted in a recent review, despite the apparent "congruence between right-wing populism and climate skepticism, there is a surprising dearth of academic research that investigates its nature and causes" (Lockwood, 2018, p. 713). We sought to contribute to such research by investigating if this congruence can be consistently found at the level of the individual or whether it depends on the political discourse that the individuals are engaged in. One previous study has looked at the individual-level association between immigrant animosity and lack of environmental concern— pro-refugee and pro-environmental concerns were positively associated in a sample of some 10000 candidates running in the 2017 Finnish municipal elections (Lönnqvist, Ilmarinen, & Sortheix, 2020). This does not, however, automatically mean that the masses would be similarly constrained in their thinking. Political elites are generally thought to know "what—goes—with—what" (*e.g.*, laissez-faire with free enterprise), whereas members of the mass public seldom make coherent connections between different attitudes (*e.g.*, Converse, 1964; Lupton, Myers, & Thornton, 2015; Zaller, 1992). Suggesting that pro-refugee and pro-environmental attitudes may be positively associated even among mass publics, a recent study found across 20 European countries a positive association between pro-immigration

and pro-environmental attitudes (Graça, 2020). However, immigrants (unauthorized and authorized) and refugees are not only definitionally distinct categories, with the former choosing to settle in another country, and the latter being forced to flee, but these categories are also discursively construed and evaluated very differently by natives (Lynn & Lea, 2003; Verkuyten, 2005). We thus sought to investigate whether climate skepticism goes together not only with animosity towards immigrants (Graça, 2020), but also with animosity towards refugees. Moreover, we sought to identify possible individual level moderators of this association, with a special interest in political affiliation and political engagement.

### 1.1. Does engagement with political discourse moderate the association?

To probe the nature of the processes that may contribute to an association between refugee and environmental attitudes, we sought to investigate whether the strength of this association varies in a predictable manner. A long-standing line of work in political psychology emphasizes individual differences in psychological dispositions as the basis for political attitudes and behavior (Jost, Federico, & Napier, 2009). In this view, underlying psychological structures (*e.g.*, need for security), interact with communication by the political elites to form a personal political orientation (Federico & Malka, 2018; Feldman & Johnston, 2014; Jost et al., 2009). Federico and Malka (2018) traced the modern origins of this view in Max Weber's discussion of "elective affinities," a concept Weber used to explain why certain political ideas appeal to certain types of people.

Several underlying dispositions could underlie a positive association between refugee and environmental attitudes. For instance, within the framework provided by Schwartz's Values Theory (1992, pp. 1–65), self-transcendence values (these concern the welfare of others, equality, and tolerance) predict both environmental and refugee attitudes (Davidov, Meuleman, Billiet, & Schmidt, 2008; Schwartz, Sagiv, & Boehnke, 2000). On the other hand, social dominance orientation (SDO), which refers to the preference for social hierarchy and inequality (Sidanius & Pratto, 1999) and right-wing authoritarianism (RWA), referring to valuing the power of perceived authority over others (Altemeyer, 1996), both predict negative attitudes towards the environment and towards refugees (Craig & Richeson, 2014; Graça, 2020; Häkkinen & Akrami, 2014; Milfont et al., 2018; Stanley & Wilson, 2019). Finding a moderately strong and ubiquitous association across different groups of people would support a view that emphasizes the psychological make-up of the individual—environmental and refugee attitudes could appeal to the same underlying psychological dispositions, whatever the nature of those dispositions, resulting in an inherent connection between attitudes towards the environment and towards refugees.

A moderated perspective would, on the other hand, emphasize the role of political discourse in attitude structuring. In this view, the masses follow the political elites' (leaders, politicians, journalists, scientists) cues and opinions, and in doing so learn which attitudes go together (Gabel & Scheve, 2007; Lenz, 2012; Malka & Soto, 2015; Minozzi, Neblo, Esterling, & Lazer, 2015; OBrien & McGarty, 2009; Zaller, 1992). However, not everyone will receive the messages of political discourse to the same degree, and only those who are the most politically engaged will adopt a structure of attitudes similar to that of the political elite (*e. g.*, Converse, 1964; Lupton et al., 2015; Zaller, 1992). The strength of the association between refugee and environmental attitudes would thus depend on the political engagement of the individual.

Consistent with this moderated perspective, some research run in the United States has shown that those who are highly educated (Sidanius & Duffy, 1988) and highly politically engaged (Federico & Schneider, 2007; Sidanius & Duffy, 1988) are the most likely to organize their economic and social attitudes along party lines, whereas those low in political engagement are more likely to adopt a "mixed bag" of attitudes (Feldman, 2013, pp. 591–626; Feldman & Johnston, 2014). Also in the

above referred study on Finnish municipal candidates (Lönnqvist et al., 2020), candidates from parties for which environmental and refugee issues were particularly important showed a stronger connection between attitudes on these issues. That is, the association was particularly strong among candidates of the anti-immigrant and pro-environmental parties. We sought to investigate whether similar patterns could be found among voters.

### 1.2. The present research

We investigated whether refugee attitudes go together with environmental attitudes among the more than 40000 respondents from 23 countries included in the European Social Survey Round 8 (2016). We expected voters of either anti-immigrant or pro-environmental parties, as well as the politically more engaged, to show a stronger association between these attitudes. To characterize political parties in Europe, we relied on the Chapel Hill Expert Survey (CHS; Polk et al., 2017) which estimates party positioning on European integration, ideology and policy issues. We pre-registered the following five hypotheses:

**H1**. There is a positive association between environmental and refugee attitudes among individual respondents across countries in the European Social Survey (2016).

**H2**. Those who voted for pro-environmental parties will report higher pro-refugee attitudes than those who voted for anti-immigration parties.

**H3**. Those who voted for anti-immigration parties will report lower pro-environmental attitudes than those who voted for pro-environmental parties.

**H4**. The strength of the association between environmental and refugee attitudes is stronger among those who voted for pro-environmental or anti-immigration parties as opposed to the voters of other parties.

**H5**. The strength of the association between environment and refugee attitudes is stronger among more politically engaged individuals.

## 2. Method

### 2.1. Preregistration

The hypotheses were pre-registered at https://osf.io/q95xg. Data is available from https://www.europeansocialsurvey.org/data/round-index.html and https://osf.io/mdjtr (combined file).

### 2.2. Participants and procedure

#### 2.2.1. Participants

We started out with 44387 European Social Survey respondents from 23 countries. Because CHS data regarding party ideology was not available for Iceland, Israel, or Russia, these countries were excluded, leaving us with 38520 participants from 20 countries. We further excluded 1644 (4.27%) participants who had missing data on at least one of the variables, leading to a final sample size of 36876 participants. Participants reported whether they, in the previous national elections were eligible to vote, whether they had voted or not, and, if they had voted, for which party they had voted.

#### 2.2.2. Voting groups

Participants reported on their voting behavior in the previous national elections (the years of the elections fell between 2011 and 2016). Altogether 21605 (58.59%) participants had voted for one of 170 political parties. Participants could also indicate that they had not voted (n = 7813, 21.19%) and that they were not eligible to vote (n = 3291, 8.92%). The remaining participants that could not be assigned to a party had "voted blank" (n = 97, 0.26%), responded "don't know" (n = 1206,

3.27%), responded "refuse to answer" (n = 1955, 5.30%), "voted for other party", and "provided invalid vote" (n = 15, 0.04%; only Spanish participants reported providing invalid votes).

Voters were classified into three groups. The participants were categorized as voting "other party" if they reported voting for a party that we had classified as "other" based on the Chapel Hill Survey or if they voted for a party not included in the survey (see 2.2.3 below; n = 894, 2.42%). Participants who reported not being eligible to vote were further categorized as being under-aged (less than 18 (16 in Austria) years old at the time of the election; n = 1813, 4.92%), not having citizenship (n = 1213, 3.29%), or not eligible for other reason (n = 265, 0.72%).

The voting system in Germany gave rise to two votes, of which the one cast for a specific party was used. For Lithuania, there were three voting variables, of which the one cast for political parties (candidate lists) was used. The final voting group categories were: "Did not vote", "Don't know", "Invalid vote", "Not eligible (NE) due to age", "NE citizen", "NE other", "No answer", "Other party", "Pro-environmental party", and "Anti-immigration party".

### 2.2.3. Classification of anti-immigration and pro-environmental parties

The 170 parties that had been voted on were classified as anti-immigration, pro-environmental, or neither (categorized as "other party"). This was primarily done based on the criteria provided by the 2014 Chapel Hill Expert Survey (Polk et al., 2017). CHS 2014 provides ratings by experts (337 political scientists with specialization in political parties) regarding party positioning on various issues, including immigration and environment. We used CHS 2010 or CHS 2017, in this order, to fill gaps (recall that the elections were held 2011 to 2016, and CHS 2014 did not include ratings on all parties involved in these elections). In CHS 2014 (but neither in CHS 2010 nor CHS 2017), experts also listed the three most important issues (e.g., immigration, environment) for each party. Of the participants who reported to have voted for a certain party or coalition, 96.01% voted for a group that could be found in one of the CHS datasets. For each of these parties, scores reflecting their standing on immigration and environment were imported from the CHS data (raw scores ranged from 0 to 10, high scores indicated anti-immigration and anti-environment positioning). When the vote was cast for a coalition consisting of multiple parties, the average of the parties was used. CHS can be found at https://www.chesdata.eu/.

A party was classified as anti-immigrant if it met either of the following two criteria: (1) Immigration was listed as one of the party's three important issue (and scored above 0.5SD from the mean, to exclude pro-immigration parties) in CHS 2014, or (2) it scored 1SD above the mean score on immigration in CHS.

A party was classified as pro-environmental if it met any of the following four criteria: (1) Environment had been listed as one of the party's three important issue (and scored above 0.5SD from the mean, to exclude anti-environmental parties) in CHS 2014, or (2) it scored 1SD above the mean score on environment in CHS, or (3) was listed as belonging to a family of "green" parties in CHS17, or (4) if the party was not rated by the CHS, it belonged to the ecological-green party family in ParlGov.org database for political parties (the party family categories employed for criteria 3 and 4 could not be applied for anti-immigration classification).

The above criteria gave us 33 anti-immigration and 30 pro-environmental parties. The former had been voted on by 4005 participants, and the latter by 2175. Tables S1 and S2 in the Supplementary Materials present the parties classified as anti-immigration or pro-environmental. A list of all parties, along with the number of voters in ESS2016, their size, and their scores on immigration and environment in CHS can be found from https://osf.io/mdjtr/.

## 3. Measures

### 3.1. Attitudes towards refugees

We measured attitudes towards refugees with three items responded to on a 5-point Likert scale (1 = I totally agree, 2 = Agree, 3 = Neither agree nor disagree, 4 = Disagree, 5 = Totally disagree). In response to the prompt "Some people come to this country and apply for refugee status on the grounds that they fear persecution in their own country" participants indicated their agreement with the statements: "The government should be generous in judging people's applications for refugee status," "Most applicants for refugee status aren't in real fear of persecution in their own countries," and "Refugees whose applications are granted should be entitled to bring in their close family members." Responses were reverse coded to indicate higher agreement (pro-refugee attitudes). For Hungarian participants, only the first of the items was available, and the scores for that item were used.

### 3.1.1. Attitudes towards the environment

We measured attitudes towards the environment with three items responded to on a 5-point Likert scale (1 = Strongly in favour, 2 = Somewhat in favour, 3 = Neither in favour, nor against laws, 4 = Somewhat against, 5 = Strongly against). In response to the prompt "To what extent are you in favour or against the following policies in [country] to reduce climate change?" participants indicated to what extent they favored "Increasing taxes on fossil fuels, such as oil, gas and coal", "Using public money to subsidise renewable energy such as wind and solar power," and "A law banning the sale of the least energy efficient household appliances." Responses were reverse coded to indicate higher agreement (pro-environmental attitudes).

### 3.1.2. Political engagement

Political engagement was assessed with two items. The first item, "How interested in politics are you?", was responded to on a 4-point Likert scale ranging from 1 (Very interested) to 4 (Not at all interested). This variable was reverse coded to indicate higher engagement. Responses to the second item, "How much time (in minutes) spends consuming news about politics and current affairs (watching, reading or listening) on an average day?" were categorized into four (1 = less than 30 min, 2 = 30–59 min, 3 = 60–119 min, and 4 = 120 min or more). The average of these two variables was used as an indicator of political engagement ($r = 0.32$, $p < .001$).

### 3.1.3. Covariates

As covariates, we included gender, age, years of education, occupation, and place of residence, all of which may be associated with refugee attitudes (Pratto, Stallworth, & Sidanius, 1997). Following the ESS categories, occupation was coded into nine categories ("Armed forces", "Managers", "Professionals", "Technicians and associate professionals", "Clerical support workers", "Service and sales workers", "Skilled agricultural, forestry and fishery workers", "Craft and related trades workers", "Plant and machine operators, and assemblers", "Elementary occupations"). Those who had not reported an occupation were assigned to the occupation categories unemployed (i.e., looking for a job during the last 7 days), retired (i.e., retired during the last 7 days), or "not in paid work" (as coded by the ESS interviewers). Regarding place of residence, cities, towns, and suburbs were categorized as "urban" (0) and others as "rural" (1).

## 4. Equivalence of attitudes across countries

We first sought to investigate the metric equivalence of our scales across countries (it is necessary that the items that constitute the scales are similarly interpreted across countries). Multi-group confirmatory factor analysis was run with the lavaan R package (Rosseel, 2012) to examine measurement equivalence of the attitude scales.

We tested two models. First, each of the three scale items loaded on a latent factor (for which variance was constrained to unity to allow for freely estimated loadings for each item), and these loadings were then constrained to be equal across countries. Following Saris, Satorra, and van der Veld (2009), country-wise modification indices (MI) and expected parameter change (EPC) in the constrained parameters indicated model misspecifications, in this case, loading non-equivalence. We also calculated post hoc power estimates to detect significant misspecification (Saris et al., 2009). The following criteria were used to detect misspecifications:

1. If power to detect was high (at least 0.80), the test for misspecification was significant ($p < .05$), and the absolute standardized EPC was larger than 0.10, the parameter was considered non-equivalent.

2. If power to detect was low (less than 0.80) and the test for misspecification was significant ($p < .05$), the parameter in question was considered non-equivalent.

Measurement within a country was considered as equivalent to measurement in other countries if there were 0-1 non-equivalent loadings, whereas 2–3 non-equivalent loadings was interpreted as suggesting non-equivalence. We pre-registered these criteria along with our decision that if there were 1-3 non-equivalent countries, these would be excluded from the subsequent analyses. However, if there were more than three such countries, the measurement would be considered non-equivalent, and the subsequent analysis would be run with only the first items of each scale.

The equivalence of refugee attitudes was tested with 19 countries because Hungarian participants only responded to the first of the three items assessing these attitudes. The equivalence of environment attitudes was tested for all 20 countries. Participants with missing values (less than 5%) were excluded prior to examining equivalence. Maximum likelihood estimation was used.

Modelling attitudes towards refugees with constrained loadings across countries gave the fit indices CFI = 0.90, TLI = 0.90, RMSEA = 0.11, and SRMR = 0.08. According to the pre-specified criteria, there were six countries in which attitude measurement was non-equivalent to other countries (Belgium, Spain, France, Italy, Poland, and Portugal). Modelling attitudes towards the environment with constrained loadings across countries gave the fit indices CFI = 0.90, TLI = 0.90, RMSEA = 0.08, and SRMR = 0.05. There were six countries in which measurement was non-equivalent to other countries (Switzerland, Czech Republic, Hungary, Ireland, Lithuania, and Slovenia). Because there were more than three countries in which measurement was non-equivalent (the pre-registered decision-rule), the hypotheses were tested with single-item measures for both attitudes.

## 5. Statistical analysis for testing our hypotheses

A multi-level modelling approach with three levels: (1) individuals, (2) voting groups, and (3) countries, was employed for testing all hypotheses, but the specifications of the model differed between hypotheses. We sought to investigate individual level (level-1) association between refugee and environmental attitudes (H1, H4, H5), and whether such association varied across voting groups. For these associations to be untainted by possible level-2 or level-3 associations, we mean-centered the independent variable, environmental attitudes, at these levels. Political engagement was similarly centered. In addition, level-2 associations between voting groups and attitudes were examined (H2 and H3).

### 5.1. Hypothesis 1

The parameter estimate of interest was the fixed effect between environmental and refugee attitudes at the individual level. The effect size was estimated from the proportion of level-1 variance that the fixed effect could explain (whilst controlling for covariates). We also tested whether the fixed effect remained positive when the association was

allowed to vary between level-2 (voting group) and level-3 (country) units.

### 5.1.1. Hypotheses 2 and 3

We expected pro-environmental party voters to report more pro-refugee attitudes (hypothesis 2) and higher pro-environmental attitudes (hypothesis 3) than those who voted for anti-immigration parties. These hypotheses were investigated with level-2 contrasts between voting groups (pro-environmental party, anti-immigrant, other party, various categories of non-voters). The marginal effects of voting pro-environmental vs. anti-immigrant were contrasted, whilst controlling for covariates, and also whilst allowing the voting group-specific effects to vary between countries.

### 5.1.2. Hypothesis 4

To investigate whether the association between environmental and refugee attitudes was stronger among voters of anti-immigration and/or pro-environmental parties, as compared to voters of other parties, the cross-level interaction between refugee attitudes and voting groups was examined, with and without covariates and allowing vs. not allowing for between-country variation. The marginal linear associations contrasted against each other were: i. Voted for an anti-immigration party (1), voted for a party that is not anti-immigration (0), ii. Voted for a pro-environmental party (1), voted for a party that is not pro-environmental (0), iii. Voted for anti-immigration or pro-environmental party (1), voted for a party that is neither anti-immigration nor pro-environmental (0). Voters of other parties served as the reference group in all contrasts.

### 5.1.3. Hypothesis 5

The expected moderating effect of political engagement on the association between refugee and environmental attitudes was investigated by predicting refugee attitudes with the level-1 interaction term between environmental attitudes and political engagement, whilst controlling for covariates. We first tested for the significance of the fixed interaction effect followed by tests in which the interaction was allowed to vary between voting groups and countries, whilst also allowing the main effects to vary.

Null-hypothesis significant tests with type-I error at 0.05 were used for testing the hypotheses. A p-value based on a Satterthwaite approximation of degrees of freedom was used for fixed effects. When testing for model improvement based on the inclusion and exclusion of random effects, or inclusion of multiple fixed effects, a likelihood ratio test (LRT) was used. We employed the lme4 -package (Bates, Mächler, Bolker, & Walker, 2015) and emmeans -package (Lenth, 2019) in R (R Core Team, 2019).

### 5.1.4. Statement regarding statistical power

For each hypothesis, we had sufficient sensitivity to detect even small effects with 80% power and type-I error set at 0.05 (for H1 fixed effect $R^2 = 0.0002$; for H2 and H3 d = 0.07; for H4 and H5 power was approximated with simulations that indicated virtually 1.00 power for cross-level and level-1 interaction effects of 0.05 in standardized magnitude). See https://osf.io/mdjtr/ for more specific power and sensitivity calculations.

## 6. Results

Following our pre-registered decision rule for determining lack of measurement equivalence in the attitude scales, all analyses were conducted with the single-item measures for environmental and refugee attitudes.

## 6.1. Confirmatory analyses

### 6.1.1. *H1: there is a positive association between pro-environmental and pro-refugee attitudes*

Results from the multilevel models are presented in Table 1. For purposes of model estimation, we predicted refugee attitudes with environmental attitudes (the two variables could just as well switch places, but we followed the pre-registered analysis plan). As predicted, the fixed association between refugee and environmental attitudes was positive and statistically significant ($\beta = 0.13$, $p < .001$, 95% CI [0.12, 0.13]). At the level of the individual, environmental attitudes explained 2.00% of the variance in attitudes towards refugees. This association varied between countries ($\sigma = 0.04$) and between voting groups within countries ($\sigma = 0.04$), $\chi^2(4) = 55.58$, $p < .001$. The fixed association remained statistically significant in the random-effects model ($\beta = 0.12$, $p < .001$, 95% CI [0.10, 0.15]). The meta-analytical partial correlation calculated across countries was 0.16 (95% CI [0.13, 0.19], ranging from 0.02 in Poland to .23 in Germany and Norway). Country-specific estimates are presented in Fig. 1. Similar forest plots for voting groups within each country alongside full list of estimates for each country and voting group are presented in the Supplementary Materials https://osf.io/mdjtr/.

### 6.1.2. *H2: Those who voted for pro-environmental parties report higher pro-refugee attitudes than those who voted for anti-immigration parties*

To test hypothesis 2, voting group (now ten categories: pro-environmental, anti-immigration, other party, seven groups of non-voters) was included at the group-level (level-2) in models predicting refugee attitudes. Marginal means for all categories are presented in Table 2. Pro-environmental party voters had positive refugee attitudes (M = 0.41, 95% CI [0.18, 0.65], p < .001) whereas anti-immigration party voters had negative refugee attitudes (M = –0.48, 95% CI [-0.71, -0.25], p < .001). Supporting the hypothesis, the difference

between the groups was statistically significant (difference = 0.90, 95% CI [0.77, 1.02], p < .001, observed standardized mean difference d = 0.86). Pro-environmental party voters were also more positive towards refugees than voters of other parties (difference = 0.36, 95% CI [0.25, 0.46], p < .001, d = 0.34), whereas anti-immigration voters opposed refugees more than voters of other parties (difference = –0.54, 95% CI [-0.63, -0.45], p < .001, d = –0.51). Allowing pro-environmental and anti-immigration level-2 means to vary across countries did not improve the model, $\chi^2(2) = 3.37$, p = .186, indicating that differences between these voting groups did not vary across countries.

### 6.1.3. *H3: Those who voted for anti-immigration parties report lower pro-environmental attitudes than those who voted for pro-environmental parties*

Hypothesis 3 was tested similarly to hypothesis 2, but now the dependent variable was attitudes towards the environment. Marginal means for all categories are presented in Table 2. Pro-environmental party voters had positive (M = 0.48, 95% CI [0.33, 0.63], p < .001) whereas anti-immigration party voters had negative environmental attitudes (M = –0.22, 95% CI [-0.36, -0.08], p = .002). Supporting the hypothesis, the difference between the groups was statistically significant (difference = 0.70, 95% CI [0.59, 0.82], p < .001, d = 0.59). Pro-environmental party voters were also more pro-environmental than voters of other parties (difference = 0.49, 95% CI [0.40, 0.58], p < .001, d = 0.41), whereas anti-immigration voters were less pro-environmental than voters of other parties (difference = –0.22, 95% CI [0.13, 0.30], p < .001, d = –0.18). Allowing means to vary across countries did not improve the model, $\chi^2(2) = 4.62$, p = .099, indicating that differences between these two groups of voters did not vary across countries.

### 6.1.4. *H4: The strength of the association between environmental and refugee attitudes is stronger among those who voted for pro-environmental or anti-immigration parties*

Including the cross-level interaction between attitudes towards the

**Table 1**

Multilevel models examining the association between environmental and refugee attitudes.

| Fixed effects | Model 1 | | | Model 2 | | | Model 3 | | |
|---|---|---|---|---|---|---|---|---|---|
| | Est | 95% CI | p | Est | 95% CI | p | Est | 95% CI | p |
| Intercept | 0.06 | [-0.22, 0.35] | .663 | 0.07 | [-0.21, 0.36] | .611 | 0.08 | [-0.21, 0.36] | .598 |
| Age | 0.00 | [0.00, 0.00] | .001 | 0.00 | [0.00, 0.00] | <.001 | 0.00 | [0.00, 0.00] | <.001 |
| Sex | 0.06 | [0.03, 0.08] | <.001 | 0.06 | [0.03, 0.08] | <.001 | 0.06 | [0.03, 0.08] | <.001 |
| Education | 0.02 | [0.01, 0.02] | <.001 | 0.01 | [0.01, 0.02] | <.001 | 0.01 | [0.01, 0.02] | <.001 |
| Residence | −0.08 | [-0.10, −0.05] | <.001 | −0.06 | [-0.08, −0.04] | <.001 | −0.06 | [-0.08, −0.04] | <.001 |
| Occup1 | −0.04 | [-0.21, 0.14] | .687 | −0.04 | [-0.21, 0.13] | .646 | −0.04 | [-0.22, 0.13] | .632 |
| Occup2 | −0.08 | [-0.25, 0.10] | .398 | −0.07 | [-0.24, 0.11] | .446 | −0.07 | [-0.24, 0.10] | .439 |
| Occup3 | 0.01 | [-0.16, 0.19] | .892 | 0.01 | [-0.16, 0.19] | .871 | 0.01 | [-0.16, 0.18] | .900 |
| Occup4 | 0.00 | [-0.18, 0.18] | .998 | −0.02 | [-0.19, 0.16] | .859 | −0.02 | [-0.19, 0.15] | .827 |
| Occup5 | 0.16 | [-0.02, 0.34] | .088 | 0.14 | [-0.04, 0.32] | .123 | 0.14 | [-0.04, 0.32] | .127 |
| Occup6 | −0.07 | [-0.24, 0.11] | .442 | −0.06 | [-0.24, 0.11] | .474 | −0.07 | [-0.24, 0.11] | .460 |
| Occup7 | 0.10 | [-0.08, 0.27] | .273 | 0.07 | [-0.10, 0.24] | .407 | 0.07 | [-0.10, 0.24] | .432 |
| Occup8 | −0.03 | [-0.22, 0.17] | .782 | −0.03 | [-0.23, 0.16] | .734 | −0.04 | [-0.23, 0.16] | .716 |
| Occup9 | −0.04 | [-0.21, 0.14] | .691 | −0.04 | [-0.21, 0.13] | .668 | −0.04 | [-0.21, 0.13] | .657 |
| Occup10 | −0.05 | [-0.23, 0.14] | .601 | −0.05 | [-0.24, 0.13] | .564 | −0.06 | [-0.24, 0.12] | .529 |
| Occup11 | −0.03 | [-0.20, 0.15] | .774 | −0.03 | [-0.20, 0.14] | .718 | −0.03 | [-0.20, 0.14] | .705 |
| Occup12 | −0.03 | [-0.24, 0.19] | .816 | −0.02 | [-0.23, 0.19] | .852 | −0.03 | [-0.24, 0.18] | .812 |
| Env lvl-1 | | | | 0.13 | [0.12, 0.13] | <.001 | 0.12 | [0.10, 0.15] | <.001 |
| Random effects by voting group (SD) | | | | | | | | | |
| Intercept | 0.30 | [0.27, 0.34] | | 0.31 | [0.28, 0.34] | | 0.31 | [0.28, 0.34] | |
| Env | | | | | | | 0.04 | [0.02, 0.06] | |
| ρ Intercept, Env | | | | | | | 0.13 | [–0.19, 0.45] | |
| Random effects by country (SD) | | | | | | | | | |
| Intercept | 0.50 | [0.37, 0.71] | | 0.50 | [0.37, 0.71] | | 0.50 | [0.37, 0.71] | |
| Env | | | | | | | 0.04 | [0.02, 0.06] | |
| ρ Intercept, Env | | | | | | | −0.05 | [–0.53, 0.46] | |
| Residual | 1.04 | [1.03, 1.05] | | 1.03 | [1.02, 1.04] | | 1.03 | [1.02, 1.04] | |
| −2 × logLikelihood | 105003.80 | | | 104295.00 | | | 104239.40 | | |

Note. Occupation groups: 1. Clerical support workers, 2. Craft and related trades workers, 3. Elementary occupations, 4. Managers, 5. Other: Not in paid work, 6. Plant and machine operators, and assemblers, 7. Professionals, 8. Retired, 9. Service and sales workers, 10. Skilled agricultural, forestry and fishery workers, 11. Technicians and associate professionals, 12. Unemployed. Env = Environment attitudes. ρ = random effect correlation.
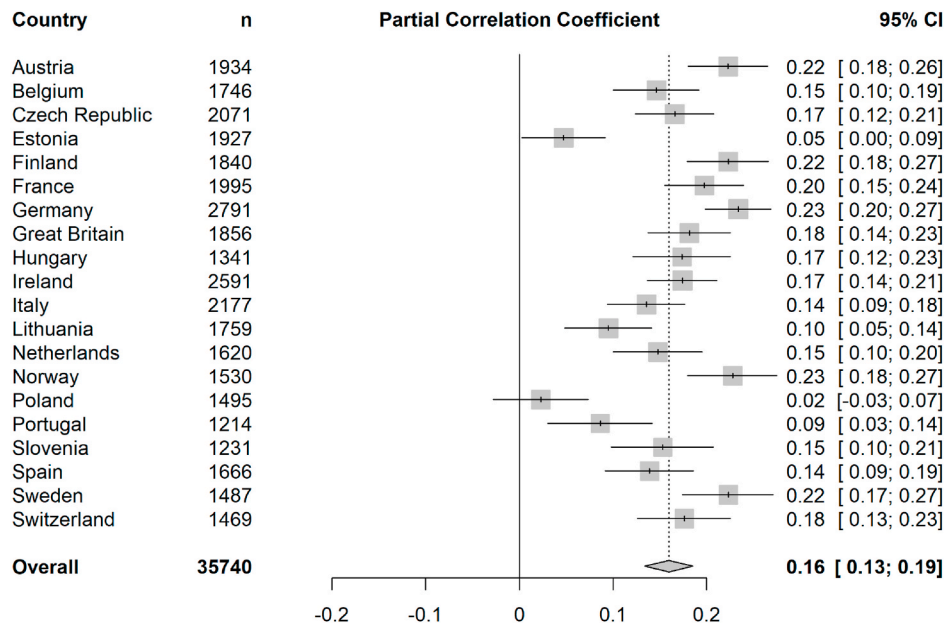
**Fig. 1.** Forest plot of country-specific estimates and 95% confidence intervals for partial correlation between refugee and environmental attitudes while controlling for age, gender, education, and place of residence (urban vs. rural).

**Table 2**
Marginal means and bivariate association for voting groups.

| Voting group | Refugee attitudes M | CI | p | Environmental attitudes M | CI | p | Association between attitudes β | CI | P |
|---|---|---|---|---|---|---|---|---|---|
| Anti-immigration | −0.48 | [-0.71, −0.25] | <.001 | −0.22 | [-0.36, −0.08] | .002 | 0.11 | [0.07, 0.15] | <.001 |
| Did not vote | −0.03 | [-0.26, 0.20] | .781 | −0.12 | [-0.27, 0.02] | .086 | 0.11 | [0.08, 0.14] | <.001 |
| Don't know | −0.06 | [-0.30, 0.18] | .630 | −0.15 | [-0.30, 0.01] | .066 | 0.14 | [0.08, 0.20] | <.001 |
| Invalid vote | 0.02 | [-0.71, 0.74] | .963 | −0.84 | [-1.56, −0.12] | .023 | 0.17 | [-0.59, 0.93] | .664 |
| NE age | 0.25 | [0.01, 0.49] | .040 | 0.17 | [0.01, 0.33] | .034 | 0.15 | [0.10, 0.20] | <.001 |
| NE citizen | 0.37 | [0.13, 0.61] | .003 | 0.02 | [-0.15, 0.18] | .850 | 0.07 | [0.01, 0.13] | .016 |
| NE other | 0.22 | [-0.05, 0.50] | .109 | 0.07 | [-0.14, 0.29] | .489 | 0.14 | [0.04, 0.25] | .008 |
| No answer | 0.08 | [-0.68, 0.84] | .839 | −0.02 | [-0.79, 0.75] | .961 | 0.34 | [-0.11, 0.79] | .141 |
| Other party | 0.06 | [-0.16, 0.27] | .596 | −0.01 | [-0.13, 0.12] | .911 | 0.13 | [0.10, 0.15] | <.001 |
| Pro-environmental | 0.41 | [0.18, 0.65] | <.001 | 0.48 | [0.33, 0.63] | <.001 | 0.14 | [0.09, 0.18] | <.001 |

Note. NE = Not eligible to vote. CI = 95% Confidence interval.

environment at level-1 and voting groups at level-2 did not improve the model predicting refugee attitudes, $\chi^2(9) = 7.46$, p = .590. The marginal linear associations for pro-environmental (β = 0.14, 95% CI [0.09, 0.18], p < .001), anti-immigration (β = 0.11, 95% CI [0.07, 0.15], p < .001), and other party voters (β = 0.13, 95% CI [0.10, 0.15], p < .001) were all statistically significant (Table 2) and did not differ from each other (for all pairwise comparisons, p > .325), providing no support for the hypothesis.

*6.1.5. H5. The strength of the association between environmental and refugee attitudes is stronger among more politically engaged individuals*

Political engagement was associated with pro-refugee attitudes (β = 0.06, 95% CI [0.05, 0.08], p < .001). However, the interaction between attitudes towards the environment and political engagement was non-significant (β = 0.01, 95% CI [-0.01, 0.02], p = .301). Allowing main effects and interaction term to vary between voting groups and countries had no influence on the fixed interaction term (β = 0.01, 95% CI [-0.01, 0.03], p = .396) or on the fixed main effect of political engagement (β = 0.06, 95% CI [0.03, 0.09], p < .001).

## 7. Exploratory analyses

### 7.1. Motivation for exploratory analyses

Two of our hypotheses were not supported. Hypothesis 4 was based on one empirical result, which was in the direction opposite to what the authors had predicted (Lönnqvist et al., 2020). In retrospect, lack of support for H4 cannot be considered very surprising. However, the reasoning behind H5 is more difficult to dismiss. One reason H5 was not supported could be poor measurement. Both items we used to assess political engagement referred to "politics", and there may be wide and systematic differences in how different parts of the surveyed population understand the term. For example, older and younger people define politics very differently (e.g., Parry, Moyser, & Day, 1992), and similar differences may exist across other divides (e.g., Eastern and Western Europe). We therefore ran some exploratory analyses employing other variables that could be argued to reflect engagement. Before that how-ever, we tested H5 using the two single items separately rather than aggregating them.

As an altogether different proxy of political engagement, we tested attitude extremity–those who take more extreme positions tend to be more politically engaged and interested (e.g., Abramowitz, 2010; Baldassarri & Gelman, 2008; Baldassarri & Goldberg, 2014; Converse, 1964; DiMaggio, Evans, & Bryson, 1996; Webster & Abramowitz, 2017).

To investigate whether taking a more extreme position was associated with higher issue alignment, we ran a set of exploratory analyses investigating issue alignment across the attitude continuum.

Not only the extreme, but also the young could be expected to be more engaged, given the high importance of environmental and immigration issues in this demographic (van der Brug, 2010). Finally, also education could be important. There is a vast literature showing that education increases political participation (e.g., electoral turnout, civic engagement, political knowledge). To investigate age and education as potential moderator variables, we added their interactions with attitudes into the above-described models. For the sake of consistency, we also entered the other demographic variables as potential moderators.

### 7.1.1. Do demographics matter?

We tested whether the demographic variables that we included as covariates moderated the linear association between attitudes. Random variability by country and voting groups was also estimated. When testing the preregistered hypotheses we employed grand-mean centered covariates. However, we now used group-mean centering (similar to what we did with environmental attitudes) in order to isolate the level-1 effects (e.g., there are age differences between voting groups, and we did not want the results pertaining to age to be confounded by such differences). The covariates were examined one at a time.

**Age.** Age was first divided by 10 in order to estimate the impact of a 10 year change in the independent variable. The interaction term between age and environmental attitudes in the prediction of refugee attitudes was statistically significant ($\beta = -0.01$, 95% CI [-0.01, -0.00], p = .002). Simple slopes indicated that the association between the two attitudes was stronger among the young than among the old (slope at –1SD/+1SD from the grand mean was $\beta = 0.14$, 95% CI [0.12, 0.16], p < .001/$\beta = 0.11$, 95% CI [0.09, 0.13], p < .001).

**Sex.** The interaction term between sex and environmental attitudes was statistically non-significant ($\beta = 0.00$, 95% CI [-0.02, 0.02], p = .913).

**Education.** The interaction term between years of education and environmental attitudes was statistically significant ($\beta = 0.01$, 95% CI [0.00, 0.01], p < .001). Simple slopes indicated that the association was stronger among the more educated (slope at –1SD/+1SD from the grand mean was $\beta = 0.10$, 95% CI [0.08, 0.13], p < .001/$\beta = 0.14$, 95% CI [0.12, 0.17], p < .001).

**Place of residence (urban/rural).** The interaction term between place of residence and environmental attitudes was non-significant ($\beta = 0.00$, 95% CI [-0.02, 0.02], p = .721).

**Occupation group.** Adding the interaction between occupation and environment attitudes improved the model, $\chi^2(12) = 32.84$, p = .001. Marginal effects calculated separately for each occupation category showed that clerical support workers showed a stronger association than did others ($\beta = 0.17$, 95% CI [0.14, 0.21], p < .001; the difference to the mean of other groups was 0.07, 95% CI [0.03, 0.11], p-value adjusted for 13 tests = 0.002). None of the other 12 groups significantly differed from the means of other groups, adjusted p > .098 for all.

### 7.1.2. Is the association stronger among the extreme?

Using an altogether different proxy of political engagement, we examined whether respondents who were more extreme in their environmental attitudes were also more extreme in their refugee attitudes. This was done by adding quadratic and cubic curvature terms to the above models that tested for the linear fixed and random effects of environmental attitudes. We excluded random effect correlations to avoid singular random effect covariance structures and non-converging models. The smallest voting group categories ("invalid vote" (n = 13) and "no answer" (n = 11)) were excluded from all analyses (these would otherwise automatically have been excluded at the stage at which the cubic terms were entered).

The fixed quadratic ($\beta = -0.01$, 95% CI [-0.02, -0.00], p = .010) and cubic ($\beta = 0.01$, 95% CI [0.00, 0.01], p = .048) terms were statistically significant. There was also between county variance in the linearity of the association (significant random effects, $\chi^2(2) = 22.14$, p < .001), and when controlling for this, also between voting groups ($\chi^2(2) = 10.00$, p = .007). Inclusion of the random effect terms rendered the fixed quadratic term non-significant ($\beta = -0.01$, 95% CI [-0.02, 0.01], p = .290), but the fixed cubic term remained significant ($\beta = 0.01$, 95% CI [0.00, 0.02], p = .036). Probing the simple slopes for the non-linear association between environment and refugee attitudes indicated that the fixed association was somewhat stronger towards the anti-environment pole (slope at –1SD from the mean: $\beta = 0.15$, 95% CI [0.11, 0.20], p < .001) than around the mean level (slope at the mean: $\beta = 0.10$, 95% CI [0.08, 0.13], p < .001) or towards the pro-environmental pole (slope at +1SD from the mean: $\beta = 0.12$, 95% CI [0.08, 0.16], p < .001). That is, environmental attitudes were more predictive of refugee attitudes among those with the least pro-environmental attitudes (and vice versa). This effect was, however, small in magnitude, and a pairwise comparison of the above point estimates showed that only the contrast between the anti-environment pole and the mean was statistically significant (difference = 0.05, 95% CI [0.01, 0.09], p = .034).

To further examine the above revealed variation in linearity between voting groups, we added the interaction terms of quadratic and cubic environmental attitudes with voting group into the model. Indeed, the shape of the slopes varied between groups ($\chi^2(14) = 24.50$, p = .040), and the marginal coefficients showed that it was the anti-immigration party voters who differed (Table 3). Only the linear coefficient was statistically significant among pro-environmental and other party voters, and these two groups did not differ in terms of linear, quadratic or cubic terms (for all comparisons, p > .169). However, among anti-immigration party voters, the linear coefficient was non-significant ($\beta = 0.05$, 95% CI [-0.01, 0.11], p = .088) whereas the quadratic ($\beta = -0.05$, 95% CI [-0.09, -0.02], p < .001) and cubic coefficients ($\beta = 0.02$, 95% CI [0.00, 0.05], p = .016) were statistically significant. Contrasts showed that anti-immigration voters differed from pro-environmental and other party voters in terms of quadratic coefficients (the differences were 0.07, 95% CI [0.02, 0.11], p = .006 and –0.05, 95% CI [0.02, 0.18], p = .003, respectively), and from pro-environmental voters in terms of linear coefficients (the difference was 0.11, 95% CI [0.01, 0.20], p = .025).

To visualize the above interactions, the associations between environmental and refugee attitudes among anti-immigration, pro-environmental, and other party voters are depicted in Fig. 2. Regarding the anti-immigration group, simple slope analyses showed that the association was strong at the anti-environment pole (slope at –1SD from the grand mean: $\beta = 0.23$, 95% CI [0.15, 0.31], p < .001), and otherwise non-significant (slopes at the grand mean and at +1SD from the grand mean were $\beta = 0.03$, 95% CI [-0.03, 0.10], p = .292 and $\beta = 0.05$, 95% CI [-0.05, 0.14], p = .330, respectively). By contrast, among voters of pro-environmental and other parties, the attitudes were positively and consistently associated throughout the attitude continuum (see also Fig. 2, in which red-dashed lines indicate the linear slopes for each voting group).

The marginal regression coefficients further revealed that the association between attitudes was non-linear also among those who were too young to vote (as indicated by the cubic coefficient: $\beta = 0.04$, 95% CI [0.01, 0.07], p = .009). Simple slopes showed that the association was strong at both poles (slope at –1SD/+1SD from the grand mean: $\beta = 0.19$, 95% CI [0.05, 0.33], p = .008/$\beta = 0.21$, 95% CI [0.11, 0.30], p < .001), but non-significant around the grand mean ($\beta = 0.04$, 95% CI [-0.05, 0.13], p = .344).

### 7.1.3. Political engagement as assessed by single items

We ran the models testing H5 separately for the two political engagement items that we had originally aggregated. The exploratory tests of H5 were consistent with the result of the confirmatory test, according to which the political engagement variable did not moderate the

**Table 3**

Marginal voting group-specific coefficients for environmental attitudes from non-linear models.

| Voting group | Linear | | | Quadratic | | | Cubic | | |
|---|---|---|---|---|---|---|---|---|---|
| | β | CI | p | β | CI | p | B | CI | p |
| Anti-immigration | 0.05 | [-0.01, 0.11] | .088 | −0.05 | [-0.09, −0.02] | .001 | 0.02 | [0.00, 0.05] | .016 |
| Did not vote | 0.10 | [0.05, 0.14] | <.001 | −0.02 | [-0.05, 0.00] | .075 | 0.01 | [-0.01, 0.02] | .400 |
| Don't know | 0.12 | [0.01, 0.22] | .028 | 0.01 | [-0.04, 0.06] | .606 | 0.01 | [-0.03, 0.04] | .681 |
| NE age | 0.05 | [-0.04, 0.14] | .273 | 0.03 | [-0.01, 0.07] | .152 | 0.04 | [0.01, 0.07] | .009 |
| NE citizen | 0.06 | [-0.05, 0.16] | .305 | 0.02 | [-0.03, 0.06] | .479 | 0.01 | [-0.03, 0.04] | .731 |
| NE other | 0.08 | [-0.14, 0.29] | .485 | 0.01 | [-0.08, 0.10] | .850 | 0.02 | [-0.04, 0.08] | .535 |
| Other party | 0.10 | [0.07, 0.14] | <.001 | −0.01 | [-0.02, 0.01] | .518 | 0.01 | [-0.00, 0.02] | .067 |
| Pro-environmental | 0.16 | [0.08, 0.23] | <.001 | 0.01 | [-0.03, 0.05] | .557 | −0.00 | [-0.03, 0.02] | .725 |

Note. NE = Not eligible to vote. CI = 95% Confidence interval.
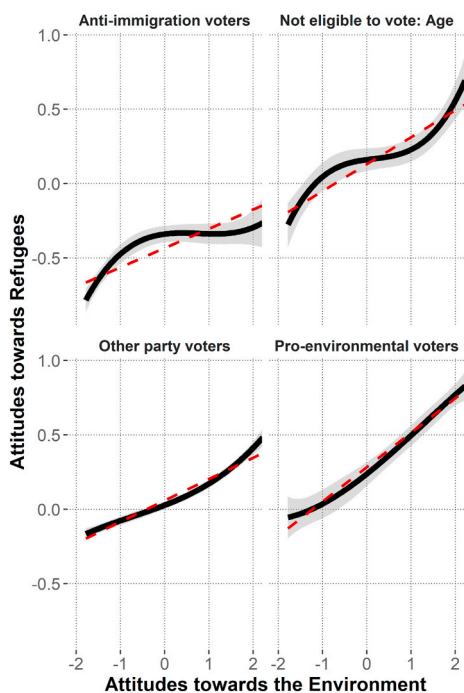


**Fig. 2.** Linear (dashed red-lines) and non-linear (third-degree polynomial, solid black lines) association between environmental and refugee attitudes by voting groups. (For interpretation of the references to colour in this figure legend, the reader is referred to the Web version of this article.)

association between refugee and environmental attitudes. See Supplementary materials for item-specific results.

## 8. Discussion

We pre-registered five hypotheses, of which the first three were supported (H1). The association between attitudes towards refugees and attitudes towards the environment was positive across large general population samples from 20 European countries (see also Graça, 2020, for similar results pertaining to attitudes towards immigration and attitudes towards the environment). Although the hypothesis was supported, the association was small and varied across countries and across different groups of voters (H2). Pro-environmental party voters were more positive towards refugees than were anti-immigration party voters. Moreover, pro-environmental party voters were more positive towards refugees than were voters of other parties (H3). Anti-immigration party voters were less positive towards the environment than were pro-environmental party voters. Moreover, anti-immigration party voters were less positive towards the environment than were voters of other parties (H4). The strength of the association between environmental and refugee attitudes was not stronger

among those who had voted for pro-environmental or anti-immigration parties (H5), nor among the politically engaged, at least not when engagement was operationalized as self-assessed interest and time spent following politics. However, further exploratory analyses suggested that political engagement may nevertheless play some role in explaining the association between these attitudes.

In retrospect, H4 is easy to dismiss, as our expectations were based on only one prior empirical result that itself was unexpected (Lönnqvist et al., 2020). However, given that H5 was based on a robust literature indicating that political engagement does constrain attitudes, we ran some further exploratory analyses. We found that issue alignment was particularly strong among the educated and among the young. Especially those who were too young to vote and were extreme on one issue tended to be extreme also on the other. Another group that showed strong issue alignment were those anti-immigration voters who took more extreme and negative positions. Among the more moderate anti-immigration voters, environmental and refugee attitudes were virtually uncorrelated.

Supporting the direct association between refugee and environmental attitudes, this association was generally positive across countries and across voting groups. A valid follow-up question would be to ask what, to use Max Weber's term, "elective affinities" (Federico & Malka, 2018), could underlie such an association. However, addressing this question with the type of cross-sectional survey data we employed is not really possible; any explanation would be in danger of tautology. For instance, explaining individuals' refugee and environmental attitudes by pointing, e.g., to their supposedly distinct valuing of equality, would be akin to explaining that "a man fights because of the instinct of pugnacity" (Skinner, 1953, p. 31). And even if the explanation were not strictly tautological, the question of causation would remain unresolved.

Supporting the idea that the association between environmental and refugee attitudes may be moderated by other variables, environmental attitudes explained only two percent of the variance in refugee attitudes, and this association varied between countries and voting groups. Moreover, the attitudes were more strongly associated among the young, the more educated, and among the most extreme populist right voters. All of these groups could be expected to be especially engaged in particularly these issues (Abramowitz, 2010; van der Brug, 2010) or, as in the case of education, in politics in general (Verba, Schlozman, & Brady, 1995).

In sum, our results imply that the psychology of the individual does matter—it could be part of the explanation for the "congruence between right-wing populism and climate skepticism" (Lockwood, 2018, p. 713). However, the association between these attitudes was generally weak, except among those who were politically more engaged. There may thus be very little inherent connection between these attitudes—for ordinary people, what they think of refugees may have very little predictive value with respect to what they think of environmental issues and about the protection of nature. So even if environmental and refugee attitudes do tend to go together, at least in some individuals, the reason for why these attitudes go together may have to be sought for in societal and cultural factors, not underlying psychological dispositions.

Returning to Madison Grant, as the current political scene changes with the rise and fall of new parties and movements, concerns about the survival of the white race may someday again go hand in hand with efforts to preserve nature. Indeed, some European actors on the far right have recently called for merging environmental issues with nationalism, a blend dubbed eco-fascism (Mackay, 2020; Müller & Traufetter, 2019; Onishi, 2019). Whether this leads to a reconfiguration of the association between environmental and refugee attitudes, at least in some countries and among some groups of voters, should be a fascinating question for future research. Another intriguing question is what will happen to the young—will they, with age, adopt more mixed combinations of attitudes, or will they continue to polarize into two groups with opposing views on immigration and on the need for environmental action.

The most obvious limitation of the present study was in the measurement of attitudes and political engagement (see 3.2.). Although we planned to use multi-item attitude scales, the failure of these scales to meet our pre-registered criteria of metric equivalence forced us to use single-items (as we promised to do in the pre-registered plan; moreover, attempts to deviate from this plan, for instance by relaxing our criteria for measurement equivalence, led to negative variance estimates in too many countries). Single items will of course have narrower scope than broader multi-item scales. Problems related to non-equivalence in the measurement of environmental attitudes have recently been noted also in other contexts (Rodríguez-Casallas, Luo, & Geng, 2020). Developing metrically compatible attitude scales should be a high priority for future cross-cultural studies.

Despite the above limitations, the present study allows for the following novel contributions to the literature on the associations between refugee and environmental attitudes: (i) Across European countries, besides being aligned with immigration attitudes (Graca, 2020), environmental attitudes are also aligned with attitudes towards refugees. (ii) Employing expert evaluations of political parties in Europe, we consistently found strong attitude differences between voters of anti-immigration and pro-environmental parties. (iii) Voters of these parties consistently differed from voters of other parties, and did so in opposite directions, indicating that the strongest differences in refugee and environmental attitudes can be found between voters of anti-immigration and pro-environmental parties. Besides these confirmatory results, exploratory analyses showed that refugee and environmental attitudes were more strongly associated among (iv) the young, (v) the more educated, and (vi) the most extreme populist right voters. The association was particularly strong in the last of these three groups, whilst being virtually zero among less extreme anti-immigration voters, suggesting that the merging of environmental issues with nationalism, as currently promoted by some actors of the European far right, may make sense from the perspective of attracting new voters.

## CRediT authorship contribution statement

**Ville-Juhani Ilmarinen:** Conceptualization, Methodology, Software, Formal analysis, Visualization, Data curation, Writing - review & editing, Pre-registration. **Florencia M. Sortheix:** Conceptualization, Methodology, Writing - review & editing, Funding acquisition, Pre-registration. **Jan-Erik Lönnqvist:** Conceptualization, Methodology, Writing - original draft, Writing - review & editing, Funding acquisition, Pre-registration.

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## Appendix A. Supplementary data

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