Research Square

Preprints are preliminary reports that have not undergone peer review.
They should not be considered conclusive, used to inform clinical practice,
or referenced by the media as validated information.

# A finite pool of agency, not of worry: hardship erodes climate responsibility, but not climate concern, across 30 European countries

**Mudassar Aziz**

mudassar.aziz@psykologi.uio.no

University of Oslo https://orcid.org/0000-0003-3922-9841

**Gulnaz Anjum**

uio

**Article**

**Keywords:**

**Posted Date:** August 20th, 2026

**DOI:** https://doi.org/10.21203/rs.3.rs-10576688/v1

**License:** This work is licensed under a Creative Commons Attribution 4.0 International License.
Read Full License

**Additional Declarations:** There is **NO** Competing Interest.

# Abstract

The people most exposed to climate harm often appear least engaged with climate politics — a pattern widely attributed to a “finite pool of worry”: everyday hardship crowding out concern for distant threats. Pairing the European Social Survey’s health-inequalities module with climate attitudes across 30 countries (N = 50,116), we find the opposite. Chronic illness predicts more climate worry in all 30 countries, and financial strain leaves worry unchanged — an equivalence-tested null. What hardship erodes is agency: poor subjective health and financial strain sharply reduce felt personal responsibility to act, denial doubles with deprivation, and among the very worried the share feeling no personal responsibility rises from 4.1% (living comfortably) to 19.6% (struggling). The deficit is specific: burdened Europeans participate in politics as much or more, yet disown duty-like acts: voting and the climate. This finite pool of agency is pan-European, and it implies enabling, not persuading, Europe’s worried-but-powerless.


# Introduction

Climate change is a health and inequality crisis before it is anything else. Within European countries, heat kills primarily people already living with cardiovascular, respiratory and metabolic disease; poor-quality housing amplifies thermal stress; and low income constrains both protection and recovery[1,2,3]. The same populations have contributed least to the problem: carbon footprints rise steeply with income, within countries as much as between them[4]. Climate-justice scholarship condenses this double asymmetry into a normative demand — vulnerability should confer voice[5] — and public-health scholarship supplies its empirical backdrop, the pervasive social gradient in who gets sick and who stays well[6]. Yet whether Europe’s materially and physically burdened citizens actually voice more climate concern, or perversely less, has never been established at population scale with measurement adequate to the question.

Psychology supplies a famous reason to expect less. In Weber’s formulation, worry draws on a finite pool: as immediate threats — illness, bills, housing — command attention, concern for temporally and spatially distant threats such as climate change is displaced[7,8]. The scarcity literature deepens the mechanism: poverty itself taxes cognitive bandwidth, narrows attention to pressing needs and reshapes preferences[9,10,11], while the post-materialist tradition arrives at the same prediction from sociology — environmental concern as a luxury good that hardship cannot afford[12,13]. If these accounts are right, the justice tension is acute: those who will bear the brunt of climate impacts should report the least worry and the least felt responsibility, and climate advocacy that speaks for them speaks without them.

The direct evidence, however, has never matched the theory’s confidence. Longitudinal work through the COVID-19 pandemic found climate concern in the United Kingdom essentially undisplaced[14]; the most comprehensive multi-method assessment concluded that the finite pool of worry receives at best qualified support, with attention — not worry — behaving like the scarce resource[15]. In wartime Ukraine,

people reporting the severest war impacts expressed more climate concern, not less[16], and Israeli adults with chronic disease reported higher climate worry than healthy controls[17]. Meanwhile, European survey research shows economic insecurity predicting climate denial and scepticism[18,19], and large international studies find that support for climate policy turns as much on perceived fairness and efficacy as on concern[20]. These strands have never been brought into a single test capable of adjudicating between them — partly for want of data measuring hardship properly in the same interview as climate attitudes.

We propose that the conflicting findings resolve once two psychological quantities are separated that the finite-pool debate has run together. Worry is cheap: it requires no resources, competes with nothing, and can attach to any threat a person credits. Felt responsibility to act is not: it presupposes efficacy, control and slack — precisely the resources hardship consumes[9,11,21]. Collective-action research has long shown that efficacy beliefs, more than grievance or emotion, convert concern into engagement[22]. A parallel warning arrives from climate governance: decades of "individualization of responsibility" have assigned mitigation to persons irrespective of their power to act[23,24] — responsibilization that, critics argue, activates agency only where capacity already exists[24]. If hardship depletes the substrate of agency while leaving concern intact, the crowding-out debate has been looking at the wrong margin: the pool that empties is not worry but ownership. We call this the finite pool of agency, and we test it against its rivals with pre-specified, falsifiable signatures.

Four hypotheses organise the analysis. H1 (crowding-out): physical and material burden reduce worry, attribution and responsibility, most strongly for bandwidth-taxing burdens (financial strain, depressive symptoms) and among the multiply burdened. H2 (sensitization): suffering makes systemic risk personally proximate, so health burden increases worry — and, critically, the increase should be largest for climate-sensitive conditions (cardio-circulatory, hypertensive, respiratory, allergic, metabolic disease — the morbidity heat and air quality demonstrably worsen[2]) relative to conditions with no plausible climate pathway; this contrast is our discriminating test between rational risk appraisal and generalized distress. H3 (agency erosion): burden leaves worry intact or higher while reducing felt responsibility, concentrated in subjective global evaluations of one's situation rather than objective morbidity counts, and partially attenuated by bandwidth markers. H4 (justice moderation): whatever gap exists is wider where buffers are weaker — in poorer countries and in health systems generating more unmet need.

Round 11 of the European Social Survey (2023–24) makes the test possible for the first time. Its rotating module on the social determinants of health — the second edition of a module designed by epidemiologists and fielded to clinical standards of measurement[25,26] — records self-rated health, eleven chronic conditions, unmet medical need with reasons, eight housing problems, financial strain, depressive symptoms (CES-D-8) and perceived control, in the same probability-sample interviews as three climate items: worry, attribution to human activity, and felt personal responsibility to reduce climate change (N = 50,116 across 30 countries[27]). Against this measurement base we run every hypothesis through a uniform estimation frame — country-specific weighted regressions pooled by random-effects

meta-analysis with small-sample-robust inference, validated by one-step mixed models — and subject every null to formal equivalence testing. The result is a reversal of the received picture. Hardship does not displace climate concern anywhere in Europe; it displaces climate ownership. The most burdened Europeans are as worried as the comfortable, more likely to deny, no less politically active — and drastically more likely to feel that reducing climate change is not their responsibility.

## Results

Burden is measured in five domains (physical health; healthcare access; housing; material strain; psychological bandwidth) and climate response on three outcomes — worry (1–5), felt personal responsibility (0–10) and human attribution (1–5) — with volunteered denial (0.7% weighted) analysed as a separate margin (Methods). Estimates are country-specific weighted least-squares regressions with demographic controls, pooled by DerSimonian–Laird random effects with Knapp–Hartung inference (k = 30 countries; analysis N = 46,745–48,327 per model); one-step mixed models with random slopes reproduce every headline estimate (Supplementary Tables S4 and S13). Weighted by cumulative burden — the number of domains (of four) in which a respondent falls in the top quartile — 55.2% of Europeans are burdened in at least one domain and 22.3% in two or more (Table 1). Estimates throughout are covariate-adjusted associations; causal interpretation is taken up in the Discussion.

Table 1

**Weighted sample characteristics by cumulative burden.** Weighted by anweight; N unweighted. Cumulative burden counts domains (of four: health, access, housing, material) in the top quartile of the pooled distribution.

| N (unweighted) | 0 domains | 1 domain | 2 + domains | All |
|---|---|---|---|---|
| | 21,940 | 15,440 | 10,257 | 47,637 |
| Age (years) | 47.66 | 50.35 | 51.75 | 49.4 |
| Female (%) | 47.8 | 53.6 | 58.7 | 52.2 |
| Education (ES-ISCED 1–7) | 4.2 | 3.93 | 3.56 | 3.96 |
| Urban resident (%) | 29.2 | 31.3 | 32.1 | 30.7 |
| Unemployed (%) | 2.5 | 5.5 | 9.3 | 5.1 |
| Self-rated health (1–5, higher=worse) | 1.82 | 2.31 | 2.91 | 2.23 |
| Bad/very bad health (%) | 0.2 | 8.1 | 25.2 | 8.4 |
| Chronic conditions (0–11) | 1.06 | 1.94 | 3.07 | 1.78 |
| Unmet medical need (%) | 0 | 18.6 | 50 | 17.2 |
| Any housing problem (%) | 0 | 27.3 | 61.9 | 22.7 |
| Income difficult/very difficult (%) | 0 | 23.6 | 54.6 | 20.2 |
| CES-D-8 depressive symptoms (0–24) | 4.15 | 5.92 | 8.61 | 5.76 |
| Perceived control (0–10) | 7.66 | 7.21 | 6.54 | 7.24 |
| Climate worry (1–5) | 3.25 | 3.27 | 3.29 | 3.26 |
| Very/extremely worried (%) | 38.7 | 40.9 | 42.4 | 39.9 |
| Felt responsibility (0–10) | 6.31 | 6.19 | 5.94 | 6.16 |
| Low responsibility 0–3 (%) | 13.6 | 15.3 | 19 | 15.7 |
| Human attribution (1–5) | 3.46 | 3.45 | 3.47 | 3.46 |
| Climate denial (%) | 0.5 | 0.7 | 1.1 | 0.7 |

Because unmet need and any housing problem are minority states, their top quartile coincides with reporting the state at all; composition rows partly reflect the definition. Per-country descriptives: Supplementary Table S1.

# Hardship does not displace climate concern (H1)

The crowding-out hypothesis makes an unambiguous prediction: burden should reduce worry, attribution and responsibility, most strongly for bandwidth-taxing burdens and among the multiply burdened. On the

concern margin the data refuse it. Across nine burden measures (Table 2, Fig. 2 left), not one reduces climate worry. Financial strain — the burden scarcity theory nominates first — has a pooled effect of − 0.004 points per standard deviation (95% CI − 0.025 to 0.018, $P = 0.72$): a precise null whose confidence bounds exclude anything beyond ± 2.6% of a worry standard deviation, and which passes a two one-sided tests (TOST) equivalence procedure against bounds of ± 0.05 points ($P_{TOST} = 7.5 \times 10^{-5}$). Poor self-rated health behaves identically (− 0.002 per s.d., $P = 0.80$; $P_{TOST} = 2.7 \times 10^{-7}$), as does low perceived control (− 0.014, $P = 0.28$; $P_{TOST} = 0.004$); the same equivalence holds on the attribution margin, and the financial-strain and self-rated-health nulls remain equivalent under stricter ± 0.03 bounds (Supplementary Table S7). The only burden showing any negative concern gradient is objective low income, and it is small (− 0.028 per s.d., $P = 0.022$) next to its own agency effect (− 0.166, $P = 4.8 \times 10^{-7}$).

Where H1 predicts decline, we find increase. The chronic-condition count raises worry by + 0.065 points per s.d. (95% CI 0.050–0.080, $P = 1.1 \times 10^{-9}$), positive in all 30 countries; unmet medical need (+ 0.025, $P = 0.003$), being hampered by illness (+ 0.023, $P = 0.001$) and depressive symptoms (+ 0.036, $P = 0.001$) follow suit. Nor does accumulation reverse the picture: respondents burdened in two or more domains are more worried than the unburdened (+ 0.074, 95% CI 0.032–0.116, $P = 0.001$) and more likely to endorse human attribution (+ 0.050, $P = 0.035$), while feeling less responsibility (− 0.130, 95% CI − 0.222 to − 0.038, $P = 0.007$) — the dissociation in miniature (Supplementary Table S7). What survives of H1 is its ordering claim, in relative terms only: within the same model the financial-strain coefficient on worry sits 0.082 points below the condition-count coefficient (95% CI − 0.109 to − 0.055, $P = 8.5 \times 10^{-7}$), and depressive symptoms 0.041 below ($P = 0.006$). Bandwidth-taxing burdens do the least worrying — but from a baseline of no crowding-out at all.

Table 2

**Pooled associations of burden with climate worry, felt responsibility and human attribution.** Country-specific WLS (post-stratification weights, HC1 errors) with demographic controls; DerSimonian–Laird random effects with Knapp–Hartung 95% CIs (k = 30). Effects per 1 s.d. of predictor on the outcome's native scale. † null passes TOST equivalence at ± 0.05.

| Burden (per 1 s.d.) | Worry b (95% CI) | $I^2$% | Responsibility b (95% CI) | $I^2$% | Attribution b (95% CI) | $I^2$% |
|---|---|---|---|---|---|---|
| Financial strain | -0.004 (-0.025, 0.018)† | 60 | -0.159 (-0.232, -0.087) | 74 | -0.002 (-0.024, 0.020)† | 72 |
| Poor self-rated health | -0.002 (-0.017, 0.013)† | 37 | -0.182 (-0.242, -0.123) | 70 | 0.005 (-0.015, 0.025)† | 67 |
| Hampered by illness | 0.023 (0.010, 0.037) | 30 | -0.067 (-0.103, -0.031) | 28 | 0.013 (-0.002, 0.027) | 45 |
| Chronic conditions (count) | 0.065 (0.050, 0.080) | 43 | 0.029 (-0.035, 0.094) | 71 | 0.034 (0.017, 0.050) | 56 |
| Unmet medical need | 0.025 (0.009, 0.041) | 53 | 0.036 (0.001, 0.071) | 35 | 0.009 (-0.008, 0.027) | 66 |
| Housing problems (count) | 0.017 (-0.000, 0.033) | 54 | 0.002 (-0.041, 0.045) | 46 | 0.011 (-0.002, 0.025) | 48 |
| Depressive symptoms | 0.036 (0.016, 0.057) | 63 | -0.116 (-0.179, -0.054) | 73 | 0.019 (0.001, 0.036) | 58 |
| Low perceived control | -0.014 (-0.040, 0.012)† | 75 | -0.230 (-0.302, -0.159) | 80 | 0.002 (-0.016, 0.019) | 62 |
| Composite vulnerability | 0.034 (0.016, 0.052) | 55 | -0.068 (-0.122, -0.014) | 64 | 0.017 (-0.001, 0.035) | 64 |

$I^2$ is the percentage of between-country variance attributable to heterogeneity. Sign splits, $\tau^2$, exact P values and one-step replication: Supplementary Tables S4 and S13.

## Health burden raises worry — but not through climate-risk appraisal (H2)

Sensitization gets its generic prediction right and its mechanism wrong. Worry rises with morbidity — from 3.15 at zero conditions to 3.40 at four (weighted means; ordinal-logit OR 1.14 per s.d., $P = 5.4 \times 10^{-43}$; Supplementary Fig. S1) — but the pre-specified discriminating test cuts against a risk-appraisal reading (Fig. 3). If the worried body were reasoning from exposure, worry should load on the conditions climate change demonstrably worsens[2]. It does the opposite. With both counts entered jointly (self-rated health held constant), climate-sensitive conditions contribute + 0.016 points per s.d. ($P = 0.044$), other conditions + 0.072 ($P = 1.5 \times 10^{-11}$); the within-model difference is − 0.056 (95% CI − 0.077 to − 0.034, $P = 1.3 \times 10^{-5}$), negative in 25 of 30 countries, and essentially unchanged by adjustment for depressive symptoms (− 0.052, $P = 2.8 \times 10^{-5}$). The same wrong-signed difference appears for responsibility (−

0.083, $P = 0.039$) and attribution (− 0.034, $P = 9 \times 10^{-4}$), and reclassifying allergies to the comparison set — the one climate-sensitive condition with a positive coefficient — widens the worry difference to − 0.072 ($P = 4.4 \times 10^{-7}$; Supplementary Table S8).

Condition-specific estimates locate the signal precisely (Fig. 3b): worry tracks skin conditions (+ 0.079, $P = 5 \times 10^{-4}$), stomach and digestive complaints (+ 0.069), allergies (+ 0.068), musculoskeletal pain in hand or arm (+ 0.062, $P = 1 \times 10^{-4}$), foot or leg (+ 0.057), severe headaches (+ 0.056) and back or neck (+ 0.040) — the somatic-symptom cluster — while heart or circulation problems (− 0.026), high blood pressure (− 0.006), diabetes (− 0.028), breathing problems (+ 0.016) and current cancer (− 0.019) contribute nothing distinguishable from zero. Living with the diagnoses most threatened by heat brings no extra climate worry; living with everyday somatic distress does. The residual condition–worry association survives CES-D-8 adjustment with only ~ 10% attenuation (+ 0.065 → +0.059), so measured depressive affect does not fully account for it; its distribution across conditions is nonetheless the signature of generalized distress or lived precarity, not of a population consciously connecting its morbidity to a warming climate.

## What hardship consumes is agency (H3)

The responsibility margin inverts the concern margin (Fig. 2, Table 2). Poor self-rated health reduces felt responsibility by − 0.182 points per s.d. (95% CI − 0.242 to − 0.123, $P = 7.3 \times 10^{-7}$; negative in 27 of 30 countries), financial strain by − 0.159 (95% CI − 0.232 to − 0.087, $P = 1.1 \times 10^{-4}$; 25/30), low perceived control by − 0.230 (95% CI − 0.302 to − 0.159, $P = 3.1 \times 10^{-7}$) and depressive symptoms by − 0.116 ($P = 7 \times 10^{-4}$). In categorical terms, respondents finding it very difficult on present income sit − 0.516 points below the comfortable ($P = 0.006$) and those in bad or very bad health − 0.553 below the healthy ($P = 2.2 \times 10^{-9}$), while the corresponding worry contrasts are null (+ 0.033, $P = 0.54$; +0.023, $P = 0.37$). Formally, for every one of the six burden measures the standardized slope on responsibility is more negative than the standardized slope on worry (cluster-bootstrap outcome contrasts, DL/KH-pooled: from − 0.054 s.d., 95% CI − 0.077 to − 0.031, $P = 4.0 \times 10^{-5}$ for financial strain, to − 0.080 s.d., $P = 5.0 \times 10^{-7}$ for depressive symptoms; for the condition count the contrast is negative in 29 of 30 countries, $P = 1.6 \times 10^{-7}$; Supplementary Fig. S2 and Table S9).

Two further signatures identify the deficit as agency rather than concern. First, it is carried by subjective, global evaluations of one's situation, not objective morbidity: holding the condition count constant, the self-rated-health coefficient on responsibility exceeds the count coefficient by − 0.343 points (95% CI − 0.442 to − 0.244, $P = 8.0 \times 10^{-8}$; negative in 27/30 countries) — indeed, conditional on feeling well, each additional diagnosis predicts slightly more felt responsibility (+ 0.105, $P = 0.001$). Second, the strain–responsibility association attenuates by 33.6% (− 0.159 → −0.106, $P = 0.004$) once bandwidth markers (depressive symptoms, perceived control) enter the model, and the health–responsibility association by 26.6% (− 0.182 → −0.134, $P = 2.5 \times 10^{-5}$) — partial mediation consistent with hardship consuming the psychological substrate of responsibility, with perceived control the strongest single correlate in the fully adjusted model (− 0.192, $P = 2.7 \times 10^{-6}$; Supplementary Table S6). Caregiving provides the mirror image:

informal carers feel more climate responsibility (+ 0.099, $P = 0.001$) and more worry (+ 0.037, $P = 2 \times 10^{-4}$) — responsibility for others and for the climate travel together, underscoring that the deficit among the burdened is capacity, not character.

Conditioning on concern makes the justice reading unavoidable (Fig. 1e, f). Among the 17,726 respondents who are very or extremely worried, the share who nonetheless feel essentially no personal responsibility rises from 4.1% (± 0.9) among those living comfortably, through 6.3% and 10.9%, to 19.6% (± 3.7) among those finding it very difficult on present income — a factor of 4.8 — while mean responsibility among the worried falls from 7.66 to 6.10; the health gradient mirrors it (5.2% → 11.8%). The pattern is not compositional: within the worried subsample the adjusted strain effect on responsibility remains − 0.123 per s.d. ($P = 0.005$; negative in 23 of 30 countries), though the categorical very-difficult contrast is estimated more noisily there (− 0.387, $P = 0.080$; Supplementary Table S16). One in five of Europe's worried poor has, in effect, ceded ownership of the response to a threat they fully register.

## Deprivation predicts exit from the climate consensus

Outright denial is rare (0.69% weighted) but socially patterned, extending to Round 11 what Round 8 showed for scepticism[18]: one s.d. of financial strain raises the odds of volunteering that climate change is not happening by 46% (OR 1.456, 95% CI 1.107–1.915, $P = 0.007$; pooled logit with country fixed effects and clustered errors), and the composite vulnerability index by 37% (OR 1.371, 1.044–1.800, $P = 0.023$). Poor self-rated health does not (OR 1.049, 0.902–1.221, $P = 0.53$): denial is a deprivation phenomenon, not a sickness phenomenon (Supplementary Table S10). Because the responsibility item is not asked of deniers, all agency gradients above are conservative lower bounds.

## The deficit is climate ownership, not political voice

If hardship simply drove people out of public life, the climate-agency gradient would be an instance of a general phenomenon and the climate framing would be decoration. The placebo test refuses this reading. On a five-act index of non-electoral participation (signing petitions, boycotting products, demonstrating, contacting politicians, posting political content online), financial strain has no effect (+ 0.004 per s.d., $P = 0.72$), poor self-rated health a slightly positive one (+ 0.028, $P = 0.024$), and the composite vulnerability index a substantially positive one (+ 0.121, $P = 2.8 \times 10^{-11}$): the multiply burdened are more politically active than the comfortable, not less — a surplus carried by the grievance-linked domains, unmet medical need (+ 0.104 per s.d., positive in 30 of 30 countries) and housing problems (+ 0.065, $P = 1.7 \times 10^{-8}$), with financial strain contributing nothing (Supplementary Table S15). Formally, the standardized responsibility slope is more negative than the participation slope for both strain (difference − 0.069, 95% CI − 0.097 to − 0.041, $P = 1.9 \times 10^{-5}$) and health (− 0.097, 95% CI − 0.120 to − 0.073, $P = 2.7 \times 10^{-9}$), negative in 27 of 30 countries each. The one civic act that does mirror climate responsibility is voting — strain − 0.060 s.d. ($P = 9.2 \times 10^{-6}$), health − 0.052 ($P = 7.4 \times 10^{-4}$), vulnerability − 0.048 ($P = 2.5 \times$

$10^{-4}$) — the archetypal duty-based act. What hardship erodes, on this evidence, is felt obligation and ownership — of the vote, of the climate — not the capacity for political action (Supplementary Fig. S6 and Table S15).

# The agency gap is pan-European (H4)

Justice moderation predicts wider gaps where buffers are weaker. We find no reliable evidence of this. The strain–responsibility slope varies across countries ($\tau^2 = 0.023$, $I^2 = 74\%$; range − 0.570 in Latvia to + 0.448 in Hungary, the lone significant positive outlier, whose removal strengthens the pooled estimate to − 0.177), but that variation is not organised by national affluence (meta-regression on log GDP per capita: $\gamma = +0.033$ per s.d., KH $P = 0.39$, $R^2$meta = 2.4%), by health-system strain (country unmet-need rate: $\gamma = -0.004$, $P = 0.92$) or by aggregate deprivation (country mean strain: $\gamma = -0.005$, $P = 0.90$) (Fig. 4; Supplementary Fig. S5). The same nulls hold for the health–responsibility, vulnerability–responsibility and condition–worry slopes, and one-step cross-level interactions reproduce every null (all $P \geq 0.31$; Supplementary Table S11). Two marginal signals — a slightly weaker worry–vulnerability association where national strain is high ($\gamma = -0.016$, $P = 0.056$, $R^2$meta = 20%) and a one-step worry–strain × GDP interaction (− 0.020, $P = 0.054$) contradicted by its own two-step counterpart ($\gamma = -0.002$, $P = 0.88$) — do not survive coherent interpretation and are flagged as exploratory. Descriptively the agency gradient runs steeper in post-communist Europe (− 0.247 vs − 0.105; difference $P = 0.068$), and within countries it is steeper for women (interaction − 0.068, $P = 0.004$) and older respondents (− 0.068 per age s.d., $P = 1.3 \times 10^{-8}$; Supplementary Table S12). The central fact, however, is uniformity: burden erodes climate agency in rich welfare states and poorer systems alike — this is not a deficit that national affluence has, on this evidence, bought away.

## Robustness

The dissociation is invariant to specification. It reproduces with objective income deciles in place of felt strain (responsibility − 0.166 per s.d., $P = 4.8 \times 10^{-7}$; worry − 0.028, $P = 0.022$), under adjustment for left–right ideology (strain → responsibility − 0.141, $P = 2.5 \times 10^{-4}$), with current cancer added to the condition count (worry + 0.064, $P = 4 \times 10^{-10}$), without survey weights, in ordinal models, when restricted to respondents aged 18 and over (strain → responsibility − 0.152, $P = 1.3 \times 10^{-4}$; conditions → worry + 0.065, $P = 1.5 \times 10^{-9}$), and under leave-one-country-out perturbation (strain → responsibility range − 0.146 to − 0.177; conditions → worry 0.061 to 0.068). Splitting unmet need by reason shows the worry association is carried by system frustration (waiting lists and unavailable appointments: +0.023, $P = 0.003$), not affordability (could not pay: +0.005, $P = 0.44$). Permanent disability reduces responsibility (− 0.040, $P = 0.006$) but not worry; childhood financial hardship leaves a marginal life-course echo on responsibility (− 0.062, $P = 0.054$) and none on worry (Supplementary Table S13).

## Discussion

Across 30 countries and every operationalisation we could construct from Europe's best health-inequality measurement, the finite pool of worry fails on its own terms. Hardship does not displace climate concern: chronic illness raises worry everywhere we look, financial strain leaves it equivalence-tested flat, and the multiply burdened are more worried than the comfortable. What empties under hardship is a different pool — felt responsibility to act. Poor subjective health, financial strain, low perceived control and depressive symptoms each strip a quarter to half a point from a 0–10 responsibility scale that barely moves with objective diagnoses; denial concentrates among the deprived; and among the very worried, feeling no responsibility at all is almost five times more common at the bottom of the income-comfort gradient than at the top. The pattern is pan-European, robust to specification, and specific: the same people participate in politics at equal or higher rates. Europe's burdened are not disengaged; they are dispossessed of climate ownership.

The dissociation reconciles findings that have accumulated against crowding-out without quite replacing it[14,15,16,17]. Worry, we suggest, was never the scarce quantity: it is informationally cheap, emotionally non-rival and compatible with any budget of attention that includes the evening news. Agency is expensive. Feeling personally responsible for a planetary outcome presupposes efficacy beliefs[21,22], perceived control and cognitive slack — the resources that poverty and illness are known to consume[9,10,11]. Our mechanism evidence is correlational but convergent: the agency deficit tracks subjective global evaluations rather than diagnosis counts, attenuates by a quarter to a third under bandwidth markers, and is strongest for perceived control. Meanwhile the worry surplus among the chronically ill carries the signature of generalized distress rather than risk appraisal — it loads on somatic-symptom conditions, not on the cardio-respiratory and metabolic diseases heat actually kills[2] — which is why we read it as proximity of suffering, not climate literacy of the sick.

The voting parallel sharpens the interpretation. Hardship leaves protest, petitioning and boycotting intact — among the multiply burdened, participation is higher, consistent with grievance-based mobilisation[22] — while depressing precisely the two acts framed as duties of ordinary citizenship: voting and personal climate responsibility. A finite pool of agency is therefore not a finite pool of voice. It is a selective withdrawal from internalised obligation, concentrated where obligation is hardest to afford. This distinction matters theoretically, because it relocates the climate-justice problem from motivation to moral economy: the burdened have not stopped caring or acting; they have stopped accepting that this one is on them.

An alternative reading deserves to be stated with equal force, because our data cannot fully exclude it: low felt responsibility among the deprived may be an accurate distributive judgment rather than a psychological depletion. Europe's poor emit least[4], and fairness perceptions are among the strongest predictors of climate-policy attitudes[20]. "Not my responsibility" may mean "I cannot" — or "I have done least, and owe least". Three observations keep both readings live rather than settling the matter. The attenuation by depressive symptoms and perceived control, the concentration among those who feel ill and out of control rather than those with objective diagnoses, and the parallel erosion of voting — an act

with no fair-share logic — all point beyond pure distributive reasoning. But the fair-share account predicts the same income gradient we observe, and nothing in a cross-sectional survey adjudicates cleanly. We regard the two as complementary faces of the same justice fact — one psychological, one normative — and flag their separation (via efficacy measures, responsibility-attribution items and framing experiments) as the sharpest next question this finding opens[28].

The policy implication survives either reading, which is what makes it usable. If the deficit is depletion, then income security, healthcare access and housing quality are climate-agency policy: they restore the slack that ownership requires. If the deficit is a fairness judgment, then climate policy that visibly shields the burdened — and assigns costs where emissions sit — earns back the ownership it currently forfeits[20]. Either way, the dissociation is the population-scale face of responsibilization without empowerment: individualized climate responsibility, detached from power and capacity, is declined precisely by those with least of either[24], and communication that couples personal action to collective, structural framings is better matched to this constituency than appeals to private virtue[29]. Under neither reading is the standard communication instinct — raising concern among the "unengaged" poor — aimed at the right deficit: their concern is already at parity, and among the worried it is agency that is missing. The pan-European uniformity of the gradient adds a sobering rider: no national welfare model in our thirty has closed the gap, so it will not close as a by-product of affluence. And the denial margin carries its own warning: the small minority who exit the climate consensus altogether are recruited from the financially strained, the constituency for whom climate politics most easily reads as an elite project[18,19].

Limitations bound the claims. The design is cross-sectional: we estimate associations, and although reverse causation from climate attitudes to chronic illness or financial strain is implausible, dynamic selection (for example, disengagement and hardship sharing unmeasured causes such as trait negative affectivity or institutional distrust) is not fully excluded — our CES-D-8 and control adjustments are proxies, not guarantees. Felt responsibility is a single item that blends responsibility attribution with efficacy; we have treated it explicitly as felt agency, buttressed by the voting parallel and the control gradient; multi-item validation against the Round 8 efficacy battery is a direct extension of this design. Subjective burden measures can carry reporting styles; the objective-count contrasts guard against this only partially; the gradients replicate within both post-communist and western Europe (Supplementary Table S12), though formal measurement-invariance testing of the burden indices across cultural regions remains an important robustness frontier. The responsibility item is not asked of deniers, making our agency gradients conservative; macro moderation is powered only for k = 30; and Europe is not the world — the vulnerability–voice gap may take different forms where climate impacts are existential rather than prospective, and where the discipline's evidence base remains thinnest[30,31] — an epistemic skew documented for climate psychology at large[32]. Finally, we measure felt responsibility, not behaviour; linking the agency gap to consumption, adaptation and policy support is the natural extension[20].

For two decades the finite pool of worry has told a story in which the poor and the sick cannot afford to care about the climate. In the largest test yet possible, Europeans most burdened by illness and hardship

care at least as much as everyone else — they worry more, deny more, vote less, and above all they no longer feel that the response belongs to them. The pool that runs dry is agency. Refilling it is not a communication task but a distributive one, and it names a constituency climate politics has been mislabelling: not the indifferent, but the worried-but-powerless.

# Reporting summary

Further information on research design is available in the Nature Portfolio Reporting Summary linked to this article.

## Methods

# Data and participants

We analyse the European Social Survey Round 11 integrated file, edition 4.2[27] (fieldwork 2023–24): face-to-face interviews with probability samples of the resident population aged 15 + in 30 countries (Austria, Belgium, Bulgaria, Croatia, Cyprus, Estonia, Finland, France, Germany, Greece, Hungary, Iceland, Ireland, Israel, Italy, Latvia, Lithuania, Montenegro, the Netherlands, Norway, Poland, Portugal, Serbia, Slovakia, Slovenia, Spain, Sweden, Switzerland, Ukraine and the United Kingdom), N = 50,116. The ESS is conducted under the ESS ERIC Research Ethics Committee's framework and complies with national ethical review in participating countries; the present study is a secondary analysis of fully anonymised public-use data and required no additional approval. All variable names, wordings and codings below were verified item-by-item against the edition 4.2 codebook. Every substantive measure has ≥ 98% item coverage except the household income decile (79.2%); complete-case analysis samples are 46,745–48,327 (93–96% of the full sample) per model (Supplementary Table S2).

# Climate outcomes

Worry: "How worried are you about climate change?" (wrclmch; 1 not at all – 5 extremely worried). Felt personal responsibility: "To what extent do you feel a personal responsibility to try to reduce climate change?" (ccrdprs; 0 not at all – 10 a great deal). Attribution: whether climate change is caused by natural processes, human activity, or both (ccnthum; 1 entirely natural – 5 entirely human). Respondents may volunteer at the attribution item that they do not think climate change is happening (code 55); this defines the denial indicator (0.69% weighted) and, because the responsibility item is filtered on not volunteering denial, makes all responsibility gradients conservative lower bounds for burdens that raise denial. Outcomes are analysed on their native scales; pooled standard deviations (worry 0.965; responsibility 2.688) are used wherever standardization across outcomes is required.

# Burden measures

Physical health: self-rated general health (health; 1 very good – 5 very bad, analysed so that higher = worse); hampered in daily activities by illness/disability/infirmity/mental problem (hlthhmp; recoded 0 no, 1 to some extent, 2 a lot); and a count (0–11) of chronic conditions experienced in the last 12 months from the CARD 54 checklist — heart or circulation problem; high blood pressure; breathing problems; allergies; back or neck pain; muscular or joint pain in hand or arm; muscular or joint pain in foot or leg; stomach or digestion related; skin condition related; severe headaches; diabetes (hltprhc, hltprhb, hltprbp, hltpral, hltprbn, hltprpa, hltprpf, hltprsd, hltprsc, hltprsh, hltprdi). Checklist batteries flagged refusal/don't-know/no-answer, or with no item marked including 'none of these' (hltprnt), are set missing. Current cancer (hltprca = 1, asked separately as ever/currently/never) is a robustness addition. Healthcare access: unmet medical need — "In the last 12 months, were you ever unable to get a medical consultation or the treatment you needed?" (medtrun; binary), with reasons could-not-pay (medtrnp), waiting list too long (medtrwl) and no appointments available (medtrnaa) defining an affordability vs system-frustration split (waiting list OR no appointments). Housing: a count (0–8) of accommodation problems — mould or rot; damp walls or leaking roof; no indoor flushing toilet; neither bath nor shower; overcrowding; extremely hot or cold; noise; insects or rodents (paccmoro, paccdwlr, pacclift, paccnbsh, paccocrw, paccxhoc, paccnois, paccinro), with the same battery-validity rule. Material strain: feeling about household income (hincfel; 1 living comfortably – 4 very difficult); objective income is the household net decile (hinctnta), reversed for the low-income robustness models. Psychological bandwidth: CES-D-8 depressive symptoms[36] (fltdpr, flteeff, slprl, fltlnl, fltsd, cldgng and reversed wrhpp, enjlf; each 1–4; summed to 0–24 and prorated when exactly one item is missing) and low perceived control (ctrlife, "how much control do you feel you have over your life in general nowadays?", 0–10, reversed). Life-course adversity: severe financial difficulties (fnsdfml) and serious household conflict (cnfpplh) while growing up, both reversed so higher = more adversity. Additional burden markers used in robustness: permanent sickness/disability as main activity (dsbld) and informal caregiving (hlpfmly).

# Derived indices and pre-specified classifications

The composite vulnerability index is the mean of four standardized domain scores — physical health (mean of z-scored poor self-rated health, hampered, condition count), unmet need, housing count, financial strain — computed when at least three domains are observed, then re-standardized. Cumulative burden counts the domains (of four) in which a respondent exceeds the pooled 75th percentile of the domain distribution; because unmet need (17.2% weighted) and any housing problem (22.7%) are minority states, their top quartile coincides with reporting the state at all. The climate-sensitivity classification of conditions was fixed before estimation on epidemiological grounds[2]: climate-sensitive — heart/circulation, high blood pressure, breathing problems, allergies, diabetes (heat- and air-quality-linked morbidity and mortality); other — back/neck, hand/arm, foot/leg, stomach/digestion, skin, severe headaches. Specificity outcomes: non-electoral participation is the count (0–5) of acts in the last 12 months — signed a petition; boycotted products; taken part in a lawful public demonstration; contacted a politician or official; posted or shared political content online (sgnptit, bctprd, pbldmna, contplt, pstplonl)

— computed when ≥ 4 items are valid (pooled s.d. 1.10); voting is participation in the last national election (vote; the ineligible are set missing; pooled s.d. 0.475).

# Statistical analysis

Two-step design. Step 1: within each country, weighted least squares (post-stratification weights, pspwght) with HC1 heteroskedasticity-robust standard errors regress each outcome on the standardized focal burden plus demographic controls: age (z), age² (z), gender, education (ES-ISCED, z), urban residence and current unemployment. Predictors are standardized on the pooled estimation sample; outcomes remain on native scales, so coefficients read as outcome points per 1 s.d. of burden. Countries enter a pooled estimate when the country-specific complete-case sample is ≥ 250 (all 30 qualify in every model reported; ≥150 for the worried-subsample models). Step 2: country coefficients are pooled by DerSimonian–Laird random-effects meta-analysis[33] with Knapp–Hartung adjustment and t(k − 1) reference distribution[34]; we report $\tau^2$, $I^2$, country sign splits and 95% KH confidence intervals. One-step validation: unweighted linear mixed models (REML, L-BFGS) with random intercepts and random slopes for the focal predictor reproduce every headline estimate (Supplementary Table S13c). Mutual-adjustment models enter five (and, extended, seven) burden measures simultaneously.

Equivalence testing. Null concern effects are subjected to two one-sided tests (TOST)[35] against pre-specified bounds of ± 0.05 outcome points per 1 s.d. of burden — approximately 5% of a worry s.d. and smaller than every robustly detected positive effect in the study — using the KH standard error and t(k − 1) distribution; $P_{TOST}$ is the larger one-sided P. As a sensitivity analysis, the financial-strain and self-rated-health nulls were additionally tested against stricter bounds of ± 0.03 (both remain equivalent); the perceived-control null is equivalent only at ± 0.05.

Coefficient contrasts. Within-model differences between two coefficients (for example, climate-sensitive minus other conditions; financial strain minus condition count; self-rated health minus condition count) are computed per country from the joint model with their covariance, then DL/KH-pooled. Cross-outcome contrasts (responsibility vs worry; responsibility vs participation), which involve different outcome scales and the within-person covariance of outcomes, are computed on the same within-country complete-case sample: each outcome is divided by its pooled s.d., both regressions are estimated on identical rows, and the standard error of the within-country difference comes from a nonparametric bootstrap (B = 500 row resamples, weights retained, fixed seed); country differences are then DL/KH-pooled.

Denial. The rare-event denial margin is modelled by pooled logistic regression (GLM binomial) with country fixed effects, normalized survey weights and country-clustered standard errors; estimates are odds ratios per 1 s.d. of burden.

Contextual moderation (H4). Country slopes are meta-regressed (DL weights, KH inference) on three standardized macro moderators: log GDP per capita (PPP, current international $, 2023; World Bank), the survey's own weighted country unmet-need rate, and country mean financial strain. $R^2$meta is the

proportional reduction in $\tau^2$ versus the intercept-only model (floored at 0). One-step validation adds the cross-level interaction to the mixed model. Subgroup pooling contrasts post-communist countries (BG, EE, HR, HU, LT, LV, ME, PL, RS, SI, SK, UA) with the rest via a normal approximation on the pooled estimates; demographic moderation (gender, age) uses one-step interactions.

Robustness battery. (R1) objective income decile in place of felt strain; (R2) left–right ideology adjustment (a sensitivity analysis, as ideology is a plausible mediator); (R3) unmet-need reason split; (R4) condition count including current cancer; (R5) permanent disability and informal caregiving as alternative burden markers; (R6) childhood financial hardship; (R7) leave-one-country-out influence analysis; (R8) ordinal logistic models for the categorical outcomes; plus unweighted re-estimation and adjusted categorical contrasts (Supplementary Table S13). Analysis flags: one one-step model (attribution on self-rated health) and one one-step interaction (responsibility on self-rated health × log GDP) did not converge — their two-step counterparts carry the reported inference; Hungary is a significant positive outlier for the strain–responsibility slope, and leave-one-out ranges are reported; two marginal moderation signals ($P \approx 0.05$–$.06$) are mutually incoherent across estimators and treated as exploratory.

Weights and descriptive inference. Within-country models use pspwght; pooled descriptive quantities use anweight (= pspwght × pweight) with Kish effective-sample-size confidence intervals[37]. All tests are two-sided with exact P values; no multiple-comparison correction is applied; inferential emphasis rests on effect sizes, confidence intervals, cross-country sign consistency and the pre-specified contrast structure rather than isolated significance. All central associations reported in the text nonetheless survive Bonferroni correction across the 27 domain-wise tests of Table 2 ($\alpha = 0.0019$); the smaller secondary associations (unmet medical need → worry; unmet need and composite vulnerability → responsibility) do not clear that threshold and are treated as secondary.

# Statistics and reproducibility

No statistical method was used to predetermine sample size; all available Round 11 respondents were analysed. The study is observational; no randomization or blinding applies. Exclusions are item-level and reported per model. The 30 countries function as internal replications: for every headline claim we report the number of countries whose independent estimate shares the pooled sign. The full pipeline (Python 3.11; pandas 3.0.2, NumPy 2.4.4, statsmodels 0.14.6, SciPy 1.17.1, Matplotlib 3.10.9, pyarrow 25.0.0) regenerates every number, table and figure from the public microdata; scripts (01–16 plus shared library), per-country estimates and all intermediate results accompany this manuscript and will be archived with a DOI on acceptance.

## Declarations

# Competing interests

The authors declare no competing interests.

# Additional information

Supplementary Information (Supplementary Methods, Tables S1–S16 and Figs S1–S6) accompanies this paper. Correspondence and requests for materials should be addressed to M.A.

# Author contributions

M.A. and G.A. conceived and designed the study. M.A. curated the data and performed the statistical analyses. M.A. and G.A. interpreted the results. M.A. drafted the manuscript; G.A. and M.A. revised it critically. Both authors approved the final version.

# Acknowledgements

This study uses data from the European Social Survey (ESS ERIC). We thank the ESS Core Scientific Team, the Round 11 health-inequalities module team, the national coordinators and the respondents. The ESS bears no responsibility for the analyses or interpretations reported here.

**Data availability**

ESS Round 11 microdata (edition 4.2) are freely available to registered users from Sikt at https://ess.sikt.no (doi:10.21338/ess11e04_2)[27] and are not redistributed here. Country-level moderators (World Bank GDP per capita, PPP 2023) and all derived country-level estimates are included in the replication materials.

**Code availability**

All analysis code (data preparation, estimation, robustness, figures; scripts 01–15 plus shared library) is provided with the submission and will be deposited on GitHub and archived on Zenodo with a DOI on acceptance.

# References

1. IPCC. Climate Change 2022: Impacts, Adaptation and Vulnerability. Contribution of Working Group II to the Sixth Assessment Report of the Intergovernmental Panel on Climate Change (eds Pörtner, H.-O. et al.) (Cambridge Univ. Press, Cambridge, UK and New York, NY, 2022).

2. van Daalen, K. R. et al. The 2024 Europe report of the Lancet Countdown on health and climate change: unprecedented warming demands unprecedented action. Lancet Public Health 9, e495–e522 (2024). doi:10.1016/S2468-2667(24)00055-0.
3. Islam, S. N. & Winkel, J. Climate Change and Social Inequality. DESA Working Paper No. 152 (United Nations, 2017).
4. Chancel, L. Global carbon inequality over 1990–2019. Nat. Sustain. 5, 931–938 (2022).
5. Schlosberg, D. & Collins, L. B. From environmental to climate justice: climate change and the discourse of environmental justice. WIREs Clim. Change 5, 359–374 (2014).
6. Marmot, M. Social determinants of health inequalities. Lancet 365, 1099–1104 (2005).
7. Weber, E. U. Experience-based and description-based perceptions of long-term risk: why global warming does not scare us (yet). Clim. Change 77, 103–120 (2006).
8. Linville, P. W. & Fischer, G. W. Preferences for separating or combining events. J. Pers. Soc. Psychol. 60, 5–23 (1991).
9. Mani, A., Mullainathan, S., Shafir, E. & Zhao, J. Poverty impedes cognitive function. Science 341, 976–980 (2013).
10. Shah, A. K., Mullainathan, S. & Shafir, E. Some consequences of having too little. Science 338, 682–685 (2012).
11. Haushofer, J. & Fehr, E. On the psychology of poverty. Science 344, 862–867 (2014).
12. Inglehart, R. Public support for environmental protection: objective problems and subjective values in 43 societies. PS Polit. Sci. Polit. 28, 57–72 (1995).
13. Franzen, A. & Meyer, R. Environmental attitudes in cross-national perspective: a multilevel analysis of the ISSP 1993 and 2000. Eur. Sociol. Rev. 26, 219–234 (2010).
14. Evensen, D. et al. Effect of “finite pool of worry” and COVID-19 on UK climate change perceptions. Proc. Natl Acad. Sci. USA 118, e2018936118 (2021).
15. Sisco, M. R., Constantino, S. M., Gao, Y., Tavoni, M., Cooperman, A. D., Bosetti, V. & Weber, E. U. Examining evidence for the Finite Pool of Worry and Finite Pool of Attention hypotheses. Glob. Environ. Change 78, 102622 (2023).
16. Błaszczyk, M. & Dyczek, B. Beyond the finite pool of worry: war experiences and climate change concerns in Ukraine. Ambio, doi:10.1007/s13280-025-02291-w (2025).
17. Shinan-Altman, S. & Hamama-Raz, Y. Climate change worry, awareness, risk appraisal, and pro-environmental behaviors: are these factors different for individuals with and without chronic diseases? PLOS ONE 20, e0325836 (2025).
18. Lübke, C. Socioeconomic roots of climate change denial and uncertainty among the European population. Eur. Sociol. Rev. 38, 153–168 (2022).
19. Poortinga, W., Whitmarsh, L., Steg, L., Böhm, G. & Fisher, S. Climate change perceptions and their individual-level determinants: a cross-European analysis. Glob. Environ. Change 55, 25–35 (2019).

20. Dechezleprêtre, A., Fabre, A., Kruse, T., Planterose, B., Sanchez Chico, A. & Stantcheva, S. Fighting climate change: international attitudes toward climate policies. Am. Econ. Rev. 115, 1258–1300 (2025).
21. Bandura, A. Self-efficacy: toward a unifying theory of behavioral change. Psychol. Rev. 84, 191–215 (1977).
22. van Zomeren, M., Postmes, T. & Spears, R. Toward an integrative social identity model of collective action: a quantitative research synthesis of three socio-psychological perspectives. Psychol. Bull. 134, 504–535 (2008).
23. Maniates, M. F. Individualization: plant a tree, buy a bike, save the world? Glob. Environ. Polit. 1, 31–52 (2001).
24. Anjum, G. & Aziz, M. Responsibilized and powerless? Unpacking the individualization of climate mitigation. Energy Res. Soc. Sci. 133, 104585 (2026).
25. Hoven, H., Eikemo, T. A., Backhaus-Hoven, I., Riebler, A., Fitzgerald, R., Martino, S., Huijts, T., Heggebø, K., Vidaurre-Teixidó, P., Bambra, C. & Balaj, M. The second Health Inequalities Module in the European Social Survey (ESS): methodology and research opportunities. Soc. Sci. Med. 380, 118228 (2025).
26. Eikemo, T. A., Bambra, C., Huijts, T. & Fitzgerald, R. The first pan-European sociological health inequalities survey of the general population: the European Social Survey rotating module on the social determinants of health. Eur. Sociol. Rev. 33, 137–153 (2017).
27. European Social Survey ERIC. ESS Round 11 Data, integrated file edition 4.2 (Sikt, 2025). doi:10.21338/ess11e04_2.
28. Norgaard, K. M. Living in Denial: Climate Change, Emotions, and Everyday Life (MIT Press, Cambridge, MA, 2011).
29. Anjum, G. & Aziz, M. Psychological insights and structural solutions: using community frame (C-frame) in climate action and policy response. Clim. Policy, 1–11, doi:10.1080/14693062.2025.2511267 (2025).
30. Anjum, G., Aziz, M. & Nawaz, A. R. Climate anxiety scholarship: a global bibliometric synthesis (2000–2024). J. Anxiety Disord. 118, 103110 (2026).
31. Anjum, G. & Aziz, M. Bibliometric analyses of climate psychology: critical psychology and climate justice perspectives. Front. Psychol. 16, 1520937 (2025).
32. Aziz, M. & Anjum, G. Rethinking knowledge systems in psychology: addressing epistemic hegemony and systemic obstacles in climate change studies. Front. Psychol. 16, 1533802 (2025).
33. DerSimonian, R. & Laird, N. Meta-analysis in clinical trials. Control. Clin. Trials 7, 177–188 (1986).
34. Knapp, G. & Hartung, J. Improved tests for a random effects meta-regression with a single covariate. Stat. Med. 22, 2693–2710 (2003).
35. Lakens, D. Equivalence tests: a practical primer for t tests, correlations, and meta-analyses. Soc. Psychol. Personal. Sci. 8, 355–362 (2017).

36. Radloff, L. S. The CES-D Scale: a self-report depression scale for research in the general population. Appl. Psychol. Meas. 1, 385–401 (1977).

37. Kish, L. Survey Sampling (John Wiley & Sons, New York, 1965).

# Figures

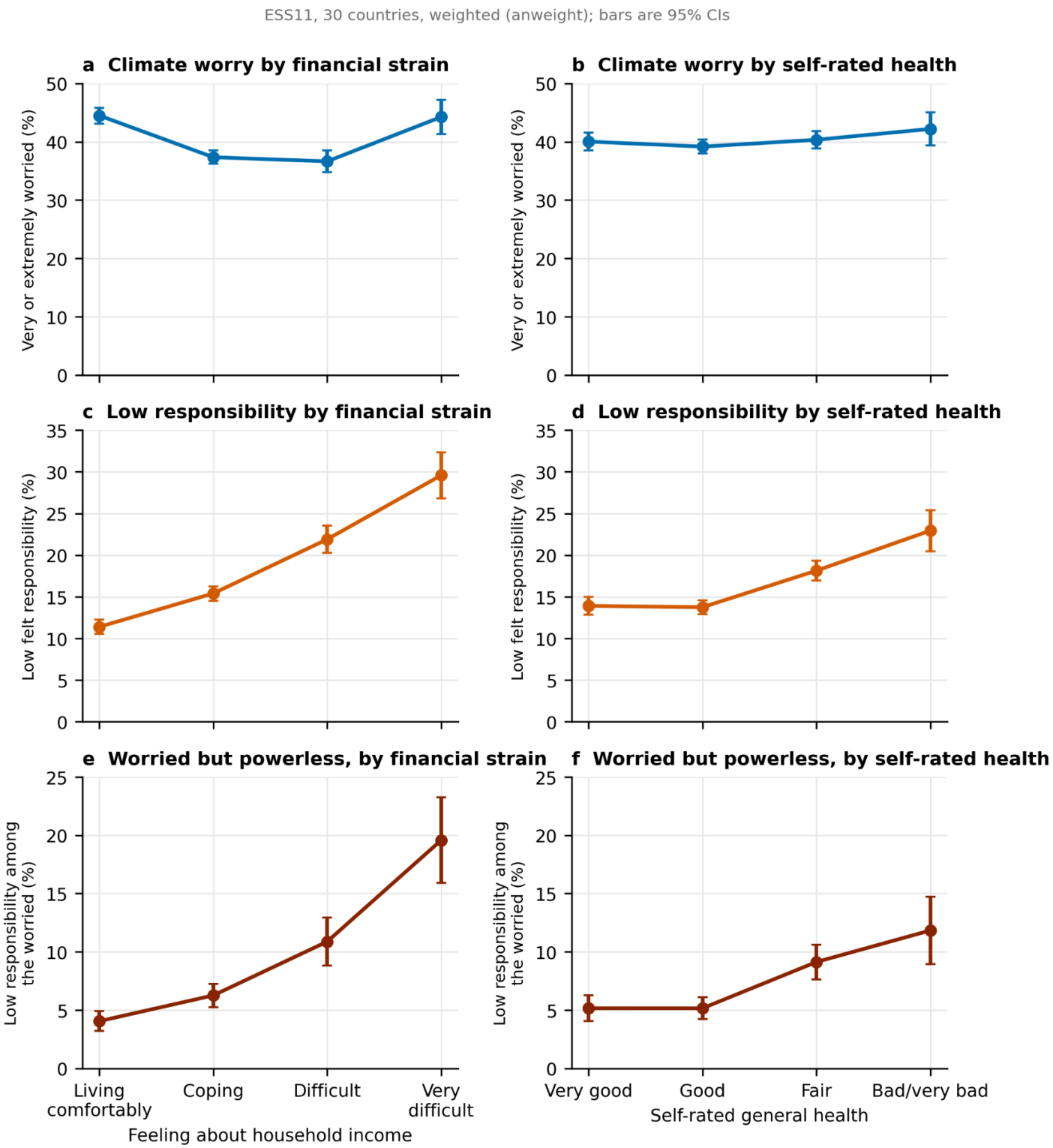



**Figure 1**

**The vulnerability dissociation in the raw data.** Weighted shares (anweight) with 95% CIs (Kish effective-n), ESS11, 30 countries. **a, b**, Share very or extremely worried about climate change by financial strain (a: 44.5% living comfortably vs 44.3% very difficult) and by self-rated health (b: 40.0% → 42.2%). **c, d**, Share reporting essentially no felt responsibility (0–3 of 10) rises from 11.4% to 29.6% across financial strain and from 13.9% to 22.9% across self-rated health. **e, f**, Among the very or extremely worried, the share feeling no personal responsibility rises from 4.1% to 19.6% across financial strain (×4.8) and from 5.2% to 11.8% across self-rated health — Europe's worried-but-powerless.

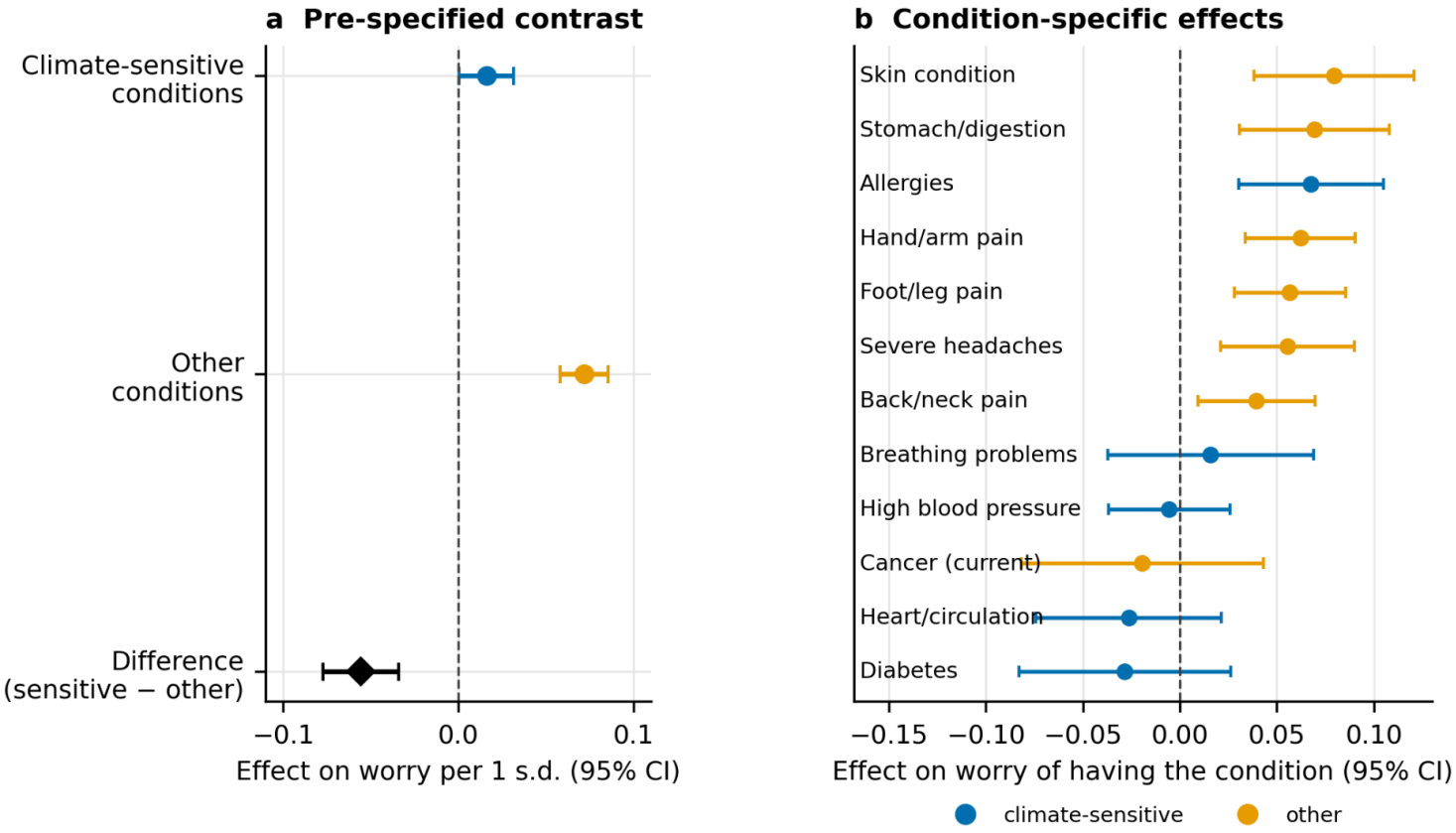
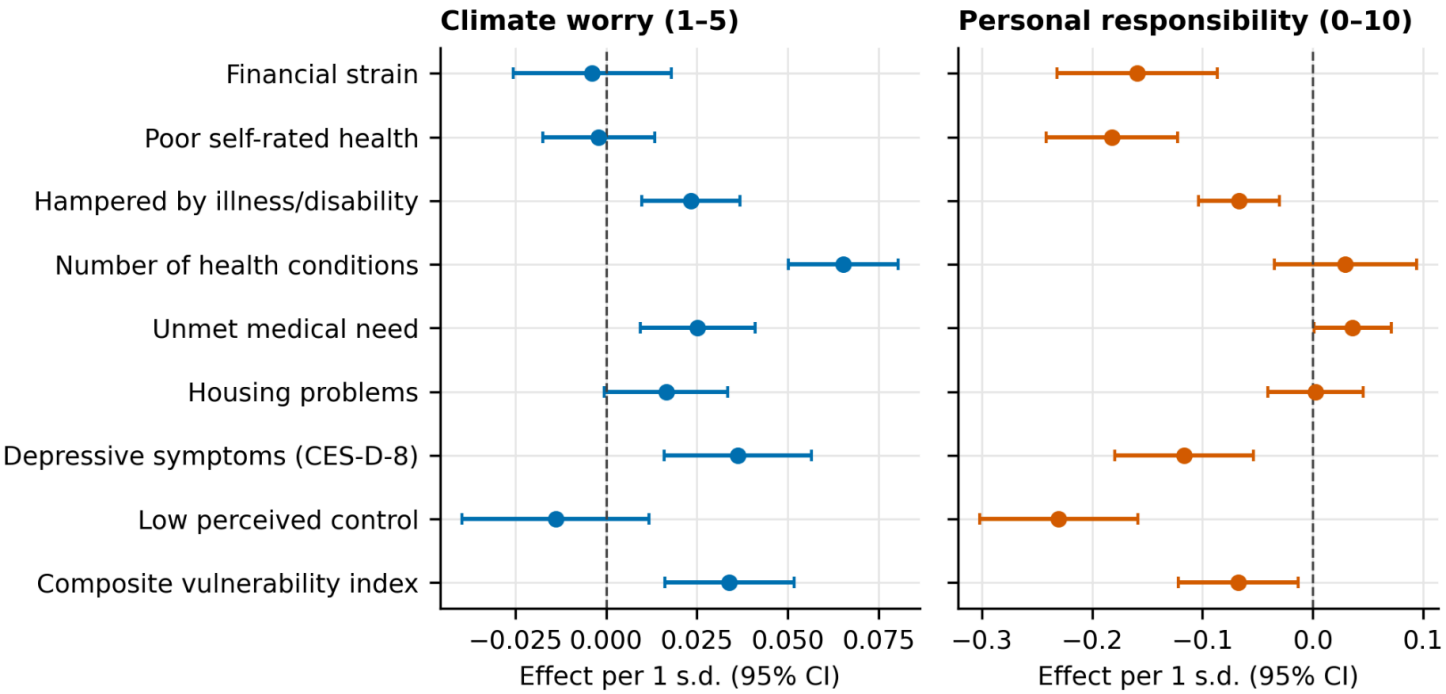



**Figure 2**

**The pre-specified discriminating test rejects risk-appraisal sensitization. a**, Pooled effects on worry of the climate-sensitive condition count (heart/circulation, high blood pressure, breathing, allergies, diabetes) and the other-condition count (musculoskeletal, digestive, skin, headaches), entered jointly with self-rated health and demographic controls, plus the pooled within-model difference (diamond; DL/KH 95% CI). **b**, Condition-specific effects (each adjusted for the count of other conditions and demographic controls).




Pooled country-specific WLS slopes, DerSimonian–Laird random effects; demographic controls

**Figure 3**

**Nine burden measures, two margins.** Adjusted pooled effects (per 1 s.d.; DL random effects, KH 95% CIs) on climate worry (left) and felt personal responsibility (right). No burden significantly reduces worry; subjective burden — poor self-rated health, financial strain, depressive symptoms, low control — sharply reduces responsibility, while objective condition counts do not.

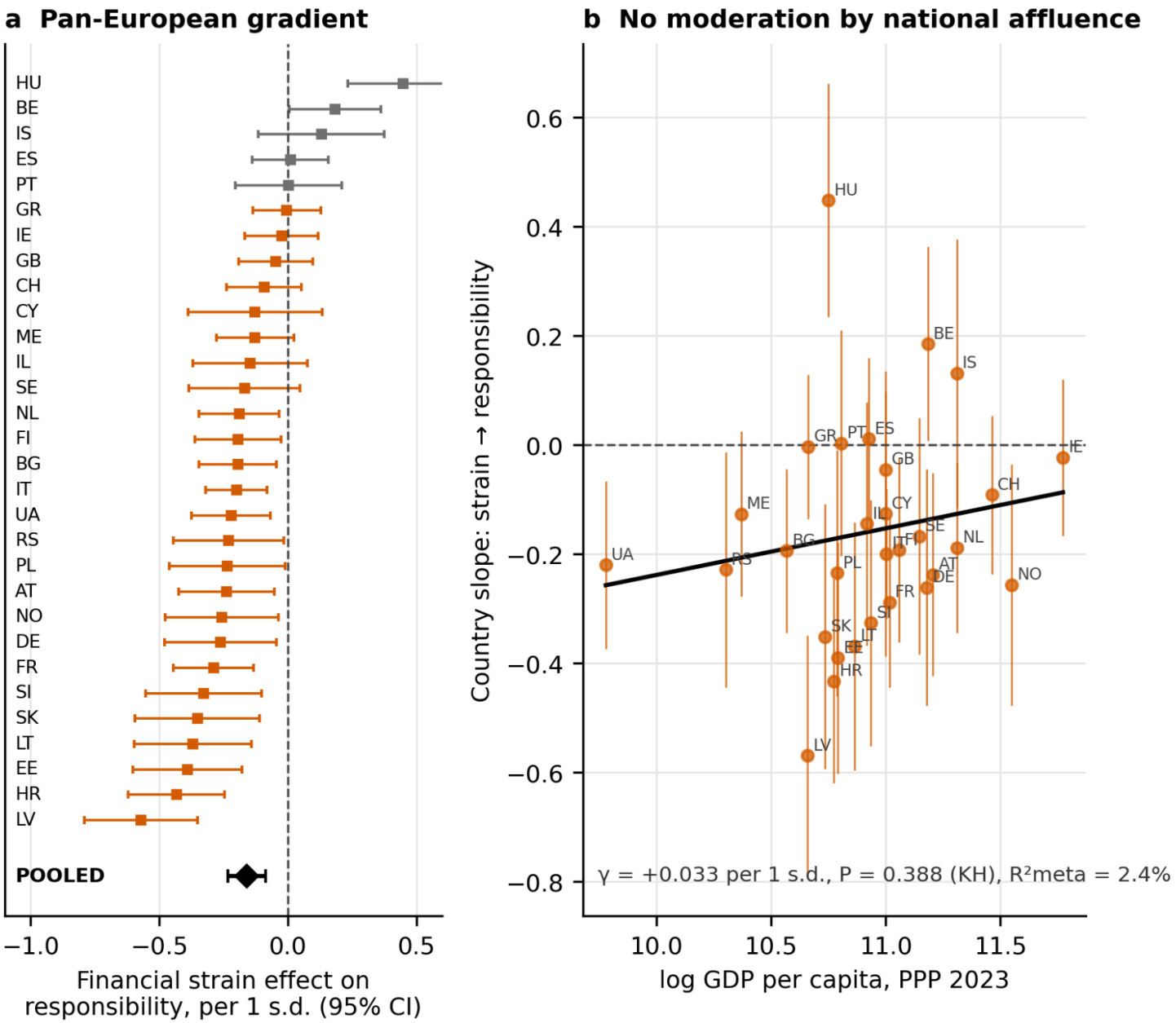



**Figure 4**

**No justice moderation. a**, Country-specific adjusted effects of financial strain (per 1 s.d.) on felt responsibility with the DL pooled estimate (diamond; KH 95% CI); orange, negative point estimates; grey, positive. **b**, Country slopes against log GDP per capita (PPP, 2023) with DL-weighted meta-regression fit.

## Supplementary Files

This is a list of supplementary files associated with this preprint. Click to download.

- ESS11SupplementaryInformation3.docx