

Can Rising Incomes Support Sustainable Well-Being? No Evidence for a “Double Dividend” in the United Kingdom

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Abstract

The double dividend hypothesis suggests that acting in more pro-environmental ways will not only benefit the environment but also carries a subjective well-being premium for those who do so, hence paying off doubly. We test the well-being payoff part of this hypothesis using a two-way fixed effects estimator with high-quality panel data for the United Kingdom during the period 2009 to 2019. Examining the link between pro-environmental behavior and subjective well-being (measured as life satisfaction), we find no evidence for even a small relationship between the two, despite a large sample size. In addition, we examine a precondition for the double dividend narrative, namely the relationship between rising incomes and pro-environmental behavior. Here we find that, in the United Kingdom, increasing incomes translate into slightly less pro-environmental behavior for the period in question. Both findings cast some doubt on the practicability of a double dividend narrative: improving material conditions may lead to a deterioration in pro-environmental behaviors, but even if one could transition to more pro-environmental behaviors, a double dividend may not materialize.

Keywords

income, pro-environmental behavior, subjective well-being, sustainability, United Kingdom

Introduction

Can economic growth simultaneously enhance subjective well-being and sustainability? As incomes rise, societies increasingly question not only whether additional income translates into higher well-being, but also whether these income gains come at the expense of the natural environment and, consequently, the well-being of future generations. In this way, debates about material progress are directly linked to concerns about non-material well-being and long-term sustainability.

Based on research that provides evidence for an association between pro-environmental behavior and subjective well-being (for overviews and meta-analytic evidence, see, e.g., Kasser, 2017; Krumm, 2024; Zawadzki et al., 2020), researchers have put forward a “double dividend” policy narrative, that is, pro-environmental behavior both benefits the environment and the individual (Herziger et al., 2020; Jackson, 2005). This means that if rising incomes lead to more pro-environmental behaviors, societies would reap a double dividend and “sustainability transitions” (loosely understood as acting more pro-environmentally) would be facilitated, as individuals stand to gain from

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acting pro-environmentally. On the other hand, if rising incomes were detrimental for the environment, encouraging pro-environmental behavior may require policymakers to persuade individuals to reduce harmful consumption or incur material sacrifices to reduce their negative impact on the environment (e.g., through less travel, fewer cars, smaller housing etc.). Here, the double dividend narrative could at least promise that these material sacrifices would yield a psychological payoff in the currency of subjective well-being. More broadly, the double dividend argument illustrates how subjective well-being is increasingly used to evaluate policy trade-offs that conventional economic indicators may not capture. Establishing whether pro-environmental behavior actually produces a well-being payoff is therefore important not only for environmental policy, but also for the credibility of well-being-based policy appraisal.

The double dividend narrative thus rests on two empirical propositions. First, the well-being dividend itself depends on whether pro-environmental behavior improves subjective well-being. Second, a broader green-growth version of the narrative also becomes more plausible when rising incomes facilitate, rather than undermine, pro-environmental behavior. Regarding the first relationship, even though the above-mentioned meta-analysis by Zawadzki et al. (2020) indicates an average positive association between pro-environmental behavior and subjective well-being, the authors caution that this link may not be interpreted causally. Indeed, most studies included in the meta-analysis suffer from limited causal credibility, as results typically rely on observational data from cross-sections and are analyzed with standard regression techniques that are often insufficiently controlled or lack control variables altogether (see, more extensively, Binder et al., 2025). The few analyses that employ more credible identification strategies suggest that any (causal) effect between pro-environmental behavior and subjective well-being is likely to be absent, or substantially smaller, and/or in the opposite direction than the very large coefficients reported by Zawadzki et al. (2020).

Regarding the second relationship, that is, between income and pro-environmental behavior, this may act as a precondition for the success of the double dividend narrative by facilitating a transition to more pro-environmental behaviors. In other words, how persuasive the double dividend narrative will be to policymakers and citizens alike might depend on whether rising incomes lead to more or less pro-environmental behavior. But here, too, the evidence for this relationship is rather mixed: on the individual level, some studies have shown that higher incomes can go hand-in-hand with more pro-environmental behavior (e.g., Grimmer et al., 2016; Milfont & Markowitz, 2016; Pleeing et al., 2021), but others report that people with higher incomes tend to have a larger ecological footprint and are less inclined to engage in pro-environmental behavior (e.g., Berthold et al., 2023; Chancel, 2022; Wiedmann et al., 2020).

Against this background, our study jointly examines both relationships by using a high-quality and detailed panel data set from a large country (the United Kingdom) experiencing income gains over the sample horizon. More concretely, we analyze an index of 11 pro-environmental behaviors measured across 3 waves (2009/2010, 2012/2013, 2018/2019) of the Understanding Society household panel survey (for more information, see <https://www.understandingsociety.ac.uk/>). This allows us, on the methodological side, to employ a two-way fixed effects estimator which differences out all time-invariant confounds. In addition, we aim for a more systematic selection of (time-variant) control variables via directed acyclic graphs, something that is not common in the literatures relevant to our study and which allows us to move away from uncontrolled or haphazardly controlled partial correlations and make (incrementally) more plausible causal interpretations of the relationships between our variables of interest, even with only observational data at hand (see, e.g., Bartram et al., 2024; Pearl & MacKenzie, 2018).

Given the mixed empirical evidence found in the literature, we also explore whether this might be driven by heterogeneities in pro-environmental behaviors analyzed: for instance, it might be the case that individuals with higher incomes engage in specific types of pro-environmental behavior such as travel a lot, have large housing and heating costs, etc. Or it might be the case that people favor less costly and more “symbolic” pro-environmental behaviors over more costly but more “impactful” behaviors (see, e.g., Diekmann & Preisendörfer, 2003). In our study, we therefore compare symbolic with impactful pro-environmental behaviors, as well as categorize pro-environmental behaviors as related to energy, mobility and shopping, when analyzing our relationships of interest. Finally, our dataset further allows us to examine heterogeneities in our relationships of interest across gender, age, and education.

Our results, in a nutshell, cast doubt on an important well-being component of the double dividend narrative and its precondition: the relationship between income and pro-environmental behaviors is uniformly negative (but substantively rather small), be it for the overall index of pro-environmental behavior, or the sub-indices related to energy, mobility and shopping. Rising incomes are also negatively related to both symbolic and impactful pro-environmental behaviors (but slightly more so for impactful behaviors). And despite a large sample size, none of these pro-environmental behavior indices exhibit a statistically significant relationship with our measure of subjective well-being (life satisfaction): coefficients are uniformly so small as to round to zero.

The remainder of the paper is structured as follows. Section “Background” reviews the relevant literature and outlines our model. Section “Methodology” describes our methodology, and Section “Data Set and Variables” presents the data and variables. Section “Results” reports our results, followed by a discussion in Section “Discussion.” Section “Conclusion” concludes.

Background

Pro-environmental Behavior and Income

The focus of our analysis is the relationship between income and pro-environmental behavior on the one hand, as well as the relationship between pro-environmental behavior and subjective well-being on the other hand. Pro-environmental behavior can be defined as any action that reduces environmental stress or has a positive environmental impact (Kollmuss & Agyeman, 2002; Schmitt et al., 2018; Steg & Vlek, 2009).¹ Pro-environmental behavior can encompass a wide range of behaviors, including reducing energy and water usage, recycling, using public transport, avoiding flying, purchasing green products, consuming organic or locally grown food, and participating in climate activism (De Matos et al., 2025; Furchheim et al., 2019; Kollmuss & Agyeman, 2002). Different taxonomies of pro-environmental behavior exist such as the distinction between public-sphere actions (e.g., joining environmental protests) and private-sphere actions (e.g., reduced water usage), as well as between regular and occasional actions (Capstick et al., 2022). Factors influencing pro-environmental behavior include environmental knowledge, emotional involvement, values, and situational constraints. The adoption of such behaviors can also be influenced by environmental values, self-image, and identity (Whitmarsh & O’Neill, 2010).

The measurement of pro-environmental behavior can be achieved through self-reported surveys, behavioral observations, and composite indices (Lange et al., 2023; Preisendörfer & Diekmann, 2021). In practice, most studies rely on self-reported surveys that capture a broad spectrum of behaviors, ranging from recycling and energy conservation to environmentally conscious consumption (Franzen & Vogl, 2013).

In this study, we make a distinction between pro-environmental behaviors related to mobility, energy, and shopping as well as between “symbolic” and “impactful” pro-environmental behaviors. Symbolic behaviors involve highly observable acts that signal greenness but have small ecological impact (e.g., carrying reusable bags), while impactful behaviors are typically less visible, but have high ecological impact (e.g., reducing meat consumption, avoiding flights). This framing is used in psychology to study green signaling (e.g., Griskevicius et al., 2010) and closely follows the work of Diekmann and Preisendörfer (2003) who distinguished between low-cost behaviors that are often symbolic, easy to adopt, and socially visible and high-cost behaviors that involve a more substantial sacrifice or change. In the context of the relationship between income and pro-environmental behaviors, this distinction may be important, as higher incomes are more likely to be associated with more symbolic behaviors motivated by status signaling rather than with high-cost behaviors.

The relationship between income and pro-environmental behavior is contested. On the one hand, higher income can increase the capacity to adopt costly but sustainable practices, such as purchasing organic food or investing in green technologies, by reducing financial barriers to these choices (Neubert et al., 2022). On the other hand, greater affluence often leads to higher overall consumption, particularly in high-impact domains such as mobility and energy use, thereby offsetting potential environmental gains (Preisendörfer & Diekmann, 2021; Whitmarsh et al., 2017). In addition, higher income is frequently associated with values such as materialism, which are generally less supportive of pro-environmental behaviors (Alzubaidi et al., 2021). Moreover, people with higher incomes may perceive environmental risks as less immediate or personally relevant, which can reduce their engagement in pro-environmental behavior (Berthold et al., 2023; Lo, 2014).

Empirical work underscores this tension: while some studies show that economic resources facilitate environmentally friendly behaviors, others find that rising income is strongly associated with higher carbon emissions (Wiedmann et al., 2020). Longitudinal and cross-national analyses further suggest a dynamic pattern, whereby environmental concern often increases in the early stages of economic development but declines once affluence reaches a level where consumption-driven lifestyles and shifting values become more dominant (Franzen & Vogl, 2013). Moreover, the relationship between income and pro-environmental behavior is potentially not linear. Berthold et al. (2023) find that both higher income and subjective financial scarcity reduce the willingness to engage in pro-environmental behavior, depending on expectations about future resource availability and the perceived effort required in environmental action.

Pro-environmental Behavior and Subjective Well-being

Subjective well-being refers to people's cognitive and affective evaluation of their lives (Diener, 1984) and can be defined as the "degree to which an individual judges the overall quality of their own life-as-a-whole favourably" (Veenhoven, 1984, Chapter 2). People draw on two key sources when evaluating their subjective well-being: their emotional states and their cognitive judgments. This means individuals may assess their subjective well-being based on how they generally feel and at the same time compare their current life to both the ideal and worst imaginable. Though distinct constructs, both have been shown to correlate with each other (Graham, 2016; Schimmack, 2008), especially in Western contexts (Blanchflower & Bryson, 2024a). For our analysis, we focus on life satisfaction as an evaluative measure of subjective well-being.

Regarding the relationship between pro-environmental behavior and subjective well-being, two meta-analyses summarizing the literature find positive and statistically significant correlations between pro-environmental behavior and well-being: Krumm (2024) pegs the relationship at $r = .081$ overall, but distinguishing high-impact versus low-impact behaviors only finds a significant positive relationship for low-impact behaviors ($r = .159$), whereas the meta-analytic correlation coefficient between high-impact pro-environmental behavior and subjective well-being is practically zero ($r = -.003$). Zawadzki et al. (2020), for a slightly larger number of studies aggregated, find an overall correlation of $r = .213$ between pro-environmental behavior and evaluative measures of well-being (such as life satisfaction). Focusing on the latter meta-analysis, this high correlation coefficient reflects only unconditional (zero-order) correlations, and it includes estimates from studies that analyze the reverse causal relationship (from subjective well-being to pro-environmental behavior).² In effect, only 16 (out of 78) studies in the meta-analysis claim they deal expressly with the causal direction from pro-environmental behavior to life satisfaction. Of those, the majority finds positive associations between pro-environmental behavior and life satisfaction (e.g., Laffan, 2020; Schmitt et al., 2018; Welsch & Kühling, 2010). For instance, Schmitt et al. (2018) find a positive and significant relationship between their overall measure of pro-environmental behavior and life satisfaction (standardized coefficient $\beta = .19$, $n = 2,220$ US & Canada, p. 135), but not all associations between their 39 individual pro-environmental behaviors are as large as the overall coefficient and not all are statistically significant (Schmitt et al., 2018, Table 4, p. 136). With other large-scale data, similarly positive relationships have been reported by Welsch and Kühling (2010, for World Values Survey data) or Laffan (2020, for the United Kingdom). Other research finds no evidence for a relationship at all (e.g., Delhey et al., 2026; Suárez-Varela et al., 2014), and some studies even find negative effects (Binder et al., 2020a; Furchheim et al., 2019; Verhofstadt et al., 2016). For example, Binder et al. (2020a) show that, after controlling for a range of factors including green self-image and personality traits, an index of 20 pro-environmental behaviors is negatively associated with life satisfaction in a large sample of Spanish students ($\beta = -.08$, $n = 640$, p. 9). Taken together, these mixed findings are partly explained by important methodological limitations in the existing literature. Much of the research in this area is purely correlational, with analyses applying few or no control variables at all, producing large zero-order correlations (see, e.g., Capstick et al., 2022). However, such large zero-order correlations are consistent with much smaller causal effects and smaller or statistically insignificant coefficients in better controlled multivariate regression analyses (Binder & Blankenberg, 2017; Binder et al., 2025) or experimental research (Dillenseger & Mouminoux, 2026; Prinzing, 2024). Establishing a plausible case for a double dividend therefore requires more than descriptive zero-order correlations; it requires credible evidence consistent with a causal relationship between pro-environmental behavior and subjective well-being. And even though there may be a "robust association" between pro-environmental behavior and subjective well-being, so far, the overall evidence consistent with a causal impact of pro-environmental behavior on subjective well-being in general and life satisfaction in particular, is rather weak (see more extensively Binder et al., 2025).

Methodology

Econometric Strategy

We use a two-way fixed effects ordinary least squares regression model, which accounts for both individual-specific unobserved heterogeneity and time fixed effects. Our dependent variables are pro-environmental behavior (as a sum index adding up behavioral frequencies on a 44-point scale) as well as life satisfaction, which is measured on a 7-point, endpoint-labeled numerical response scale, where values have a natural ordering but do not necessarily presuppose a cardinal interpretation. For the latter, applying an OLS approach has been a dominant model choice in the literature despite ordered choice models being technically more correct. Ease of interpretation of the regression coefficients, but also the ability to apply fixed effects estimators (which do not have comparable fixed effects ordered choice estimators with similar properties) inform this choice.³

Our regression equations thus take the following form:

$$PEB_{it} = \beta INCOME_{it} + \gamma W_{it} + \alpha_i + \lambda_t + \varepsilon_{it} \quad (1)$$

and

$$LS_{it} = \beta PEB_{it} + \gamma Z_{it} + \alpha_i + \lambda_t + \varepsilon_{it} \quad (2)$$

where PEB_{it} is the sum index of pro-environmental behavior, LS_{it} denotes life satisfaction, and W_{it} and Z_{it} are the vectors of control variables to be discussed and defended in the subsequent subsection. α_i represents individual-specific time-invariant fixed effects, λ_t period dummies, and ε_{it} is the idiosyncratic error term, purged of the afore-listed time and person fixed effects. We use heteroscedasticity-robust standard-errors clustered at the individual level to further account for intra-person correlation of otherwise not directly modeled disturbances.

With the two-way fixed effects estimator, we do not compare levels of the dependent variable across individuals but instead examine how the dependent variable changes within individuals as a result of changes in the independent variable. In a standard (one-way) fixed effects (FE OLS) model, individuals serve as their own control (Allison, 2009, p. 1), meaning that only within-individual variation contributes to the estimation of the coefficients, while panel members without such changes do not contribute. Importantly, when individuals are observed over multiple periods before and after a change, the fixed effects coefficient captures the difference in average levels of the outcome associated with that change. With the two-way fixed effects estimator, we also account for potential period effects through time dummies (within estimation with control group), allowing individuals who do not experience a change in the independent variable to serve as a control group for the identification of common time trends. Under the assumptions that (1) there are parallel trends between the treatment and control groups, (2) that all relevant time-varying confounders are controlled for, (3) that past outcomes may not influence treatment selection, and (4) that lagged treatments may not influence current outcomes, the two-way fixed-effects estimator identifies an estimate of the average treatment effect on the treated (ATT; cf. more extensively Allison, 2009; Bruederl, 2010; Hill et al., 2020).

Causal Models and Control Variable Selection

While our study draws on panel data, it does lack experimental variation. Hence, a causal interpretation of the estimated relationships for our main variables of interest via multivariate regression depends on a convincing and systematic selection of relevant control variables (see assumption (2) above; Bartram et al., 2024; Pearl & Mackenzie, 2018). This requires a departure from the usual reference to “control variables that are typically used in the literature” and from empirically driven procedures to establish which control variables to use. One of the contributions of our article is to place the analysis of the relationship between income, pro-environmental behavior, and subjective well-being on a more solid causal footing. To this end, we utilize a principled approach of selecting control variables based on an explicit causal model developed with the so-called “Directed Acyclic Graphs” methodology.

Directed acyclic graphs are visual tools used in causal inference to map causal relationships and assumptions between variables, helping to identify potential sources of bias like confounding and selection bias, and to determine the appropriate adjustments (i.e., what variables to control for) needed to estimate causal effects from observational data. They use nodes for variables and directed arrows for causal effects, forming a graph that is “acyclic” because it contains no feedback loops (Bartram et al., 2024; Cinelli et al., 2024; Pearl & Mackenzie, 2018; Rohrer, 2018). Careful consideration about which control variables should be included in multivariate regression, and which may introduce bias, can help in making causal claims from observational data more credible. Many different patterns for the relationship between potential third variables and the dependent and independent variables of interest have been identified, with differing needs to control for them or not (Cinelli et al., 2024). After providing arguments for the causal structure of potential confounds in our relationships of interest and drawing a directed acyclic graph, identification of these confounding patterns can be automatized, and programs exist that quite conveniently then identify which of the variables in the graph need to be controlled for (see, for instance, <https://dagitty.net/dags.html>). What determines the credibility of a causal interpretation with this method, is, however, how plausibly the researcher identifies relevant control variables from existing literature and the success in plausibly defending those variables as relevant or not. We proceed to do so in the following subsections.

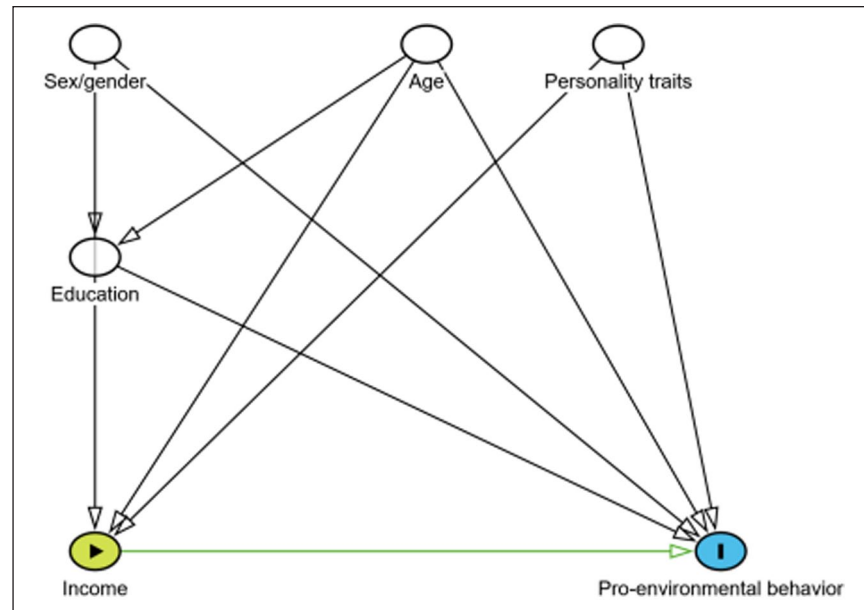


Figure 1. Causal model: income → pro-environmental behavior.

Relevant Control Variables: Income → Pro-environmental Behavior

We start by developing a causal model for our analyses with pro-environmental behavior as dependent variable, which is presented in Figure 1. The literature has established that personality can impact both income and pro-environmental behavior: recent meta-analyses by Alderotti et al. (2023) and Vella (2024) report that evidence from primary studies strongly indicates that higher personal earnings are linked to higher levels of openness, conscientiousness, and extraversion, whereas earnings tend to be significantly lower among individuals scoring higher on agreeableness and neuroticism. These findings hold even when controlling for education and cognitive ability. Research has also connected personality to pro-environmental behavior and attitudes (Gifford & Nilsson, 2014; Milfont, 2021; Soutter & Möttus, 2021), with neuroticism and extraversion emerging as the main traits lacking a consistent association with environmentalism. Due to the links between personality and both our dependent and independent variables, personality traits need to be accounted for as relevant control variable, otherwise our estimates would be biased.

Regarding gender, there is a considerable wage gap between men and women worldwide, although this has been declining over time (Gebrewolde et al., 2025; Matysiak & Cukrowska-Torzewska, 2021; Weichselbaumer & Winter-Ebmer, 2005). In addition, a relationship has been established between gender and pro-environmental behavior, as women tend to engage more in pro-environmental behavior than men and display more pro-environmental attitudes (see Franzen & Vogl, 2013; Gifford & Nilsson, 2014; Preisendörfer & Diekmann, 2021). Studies have attributed this relationship to differences in degree of socialization (Zelezny et al., 2000), value priorities (Dietz et al., 2002), and risk perception (McCright, 2010). Again, since gender is related to both independent and dependent variable, it needs to be included as control variable.

A similar logic applies to age: the relationship between age and income is well-established in the literature and tends to be U-shaped (Mincer, 1974). Income tends to increase with age, reaching a peak in people's late 40s or early 50s (OECD, 2023) and then declines due to retirement transitions or reduced work capacity. Following work on Mincer wage equations (Heckman et al., 2006), this pattern reflects human capital accumulation in the early stages of life, followed by depreciation of skills and lower productivity growth later in life. There is also evidence linking age and pro-environmental behavior: Otto and Kaiser (2014) attribute this relationship to learning: the more exposed people are to environmental information, the more pronounced their pro-environmental behavior. At the same time, younger people may sometimes show more concern but face practical obstacles that limit behavior (Cantillo et al., 2025). Age therefore also needs to be included as a control variable.

Lastly, education is (causally) related to income (see, e.g., Card, 1999; Harmon et al., 2003; Psacharopoulos & Patrinos, 2018). Higher education levels also lead to individuals acting increasingly environmentally friendly (Gifford &

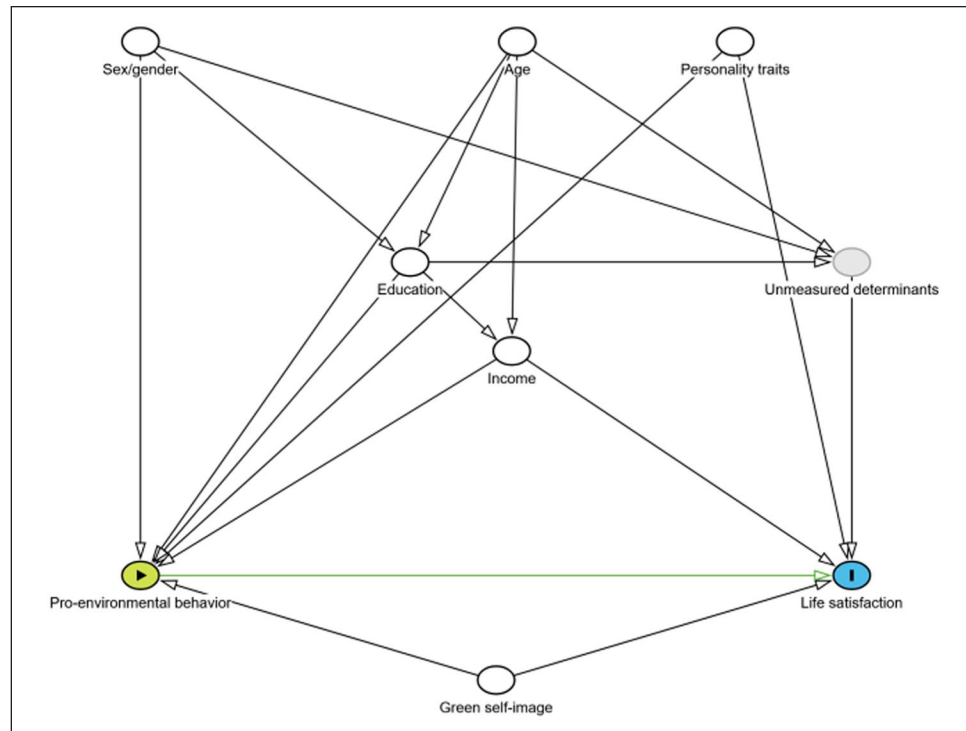


Figure 2. Causal model: pro-environmental behavior → life satisfaction.

Nilsson, 2014; Hansmann et al., 2020) but also may increase people's environmental impact (if only through increased income; cf. Preisendörfer & Diekmann, 2021, p. 141). Due to its relationship with both independent and dependent variable, education is a necessary control variable.

Based on the above discussion, our analyses need to control for age, education, gender, and personality traits. The corresponding causal model as exemplified by a directed acyclic graph is shown in Figure 1. Note that we can control for personality traits and gender only indirectly via the fixed effects estimator which differences out time-invariant confounds (this means we need to assume that personality traits do not change [much] over time, something which can be contested, see, e.g., Boyce et al., 2013). In addition, the United Kingdom Household Longitudinal Study (UKHLS) data set allows us to control for time and regional fixed effects, hence allowing us to take into account otherwise unmeasured confounds that vary by time and/or region (for instance rural areas might offer individuals different possibilities for pro-environmental behavior and at the same time also impact their incomes).

Relevant Control Variables: Pro-environmental Behavior → Life Satisfaction

For our model that estimates the impact of pro-environmental behavior on life satisfaction, we draw on the causal model and directed acyclic graph developed in Binder et al. (2025), which is depicted in Figure 2. The set of control variables identified and defended there includes green self-image, income, personality traits, gender, age, and education, the inclusion of which we will now briefly justify with regard to existing research on their relationship to our variables of interest.

Green self-image, or the degree to which individuals view themselves as environmentally conscious (Gatersleben et al., 2002) has been both linked to pro-environmental behavior and subjective well-being. The dominant interpretation in this literature is that green self-image drives pro-environmental behavior, but causal conclusions cannot yet be drawn because many studies often lack adequate controls (Udall et al., 2021). A smaller body of research has explored connections between green self-image and subjective well-being. Binder and Blankenberg (2017) showed for the United Kingdom that green lifestyle is positively related to life satisfaction, while similar findings were found using European survey data (Welsch & Kuehling, 2018) and data on Spanish students (Binder et al., 2020a). Although clear causal evidence for a green self-image–subjective well-being link remains limited, most of the literature assumes the

direction of causality flows from green self-image to both pro-environmental behavior and subjective well-being. Hence, we consider green self-image an important control variable.

In the previous section, we have defended the relationship that exists between pro-environmental behavior and income and personality traits. Both also are related to subjective well-being: income influences people's subjective well-being both directly and indirectly, through its effects on health, living conditions, and access to opportunities. While debates remain about the precise nature of this relationship (Easterlin & O'Connor, 2022), the evidence strongly points to a logarithmic association between income and subjective well-being (Jebb et al., 2018; Kahneman & Deaton, 2010; Killingsworth et al., 2023). Recent research, including natural experiments, reinforces the view that this link is causal rather than purely correlational (e.g., Lindqvist et al., 2020).

Personality traits are also consistently found to be robust predictors of subjective well-being (e.g., DeNeve & Cooper, 1998; Schimmack et al., 2008). A recent meta-analytic regression by Busseri and Erb (2024) showed that higher life satisfaction was uniquely predicted by higher extraversion, agreeableness, emotional stability (i.e., lower neuroticism), and conscientiousness, while openness had no effect.

Whereas income and personality can be considered direct determinants of subjective well-being, variables such as gender, age, and education affect pro-environmental behavior, but show only indirect links to well-being. For instance, although some studies highlight gender differences in subjective well-being (e.g., Blanchflower & Bryson, 2024b; Stevenson & Wolfers, 2009), a comprehensive meta-analysis finds no systematic direct association (Batz-Barbarich et al., 2018). Gender-based disparities in wellbeing that appear in certain countries may instead stem from broader structural inequalities, including income gaps (Bartram, 2022). Age presents a similarly complex case: evidence is mixed regarding whether its association with well-being follows a U-shape, is linear, or is negligible (López Ulloa et al., 2013). More fundamentally, any apparent effect of age is likely mediated by factors such as health, education, and income rather than reflecting a direct causal role (Bartram, 2024). Education, too, appears to affect well-being primarily through indirect pathways. While research on direct education–well-being effects yields mixed results (Clark, 2018), strong evidence indicates that education enhances well-being indirectly via its influence on income (Card, 1999) and health (Conti et al., 2010).

Data set and Variables

Our analysis is based on the United Kingdom Household Longitudinal Study (UKHLS, www.understandingsociety.org.uk), the revamped continuation of the British Household Panel Study (Knies, 2015).⁴ The UKHLS was launched in 2009, and the first wave of interviews was conducted in 2009/2010. The first wave included about 40,000 households in the United Kingdom, covering about 100,000 individuals aged 16 years and older. These households were selected in the beginning via multistage random sampling (random sample of postcodes in the United Kingdom, and random sample of addresses within these postcodes). Given its longitudinal design, the same individuals were followed over time and re-interviewed in subsequent waves. The dataset covers all aspects of individuals' lives (well-being, employment status, health etc.). Important for our purposes, the UKHLS contains subjective assessments of perceived lifestyle, attitudes toward climate change, and environmentally friendly behavior, as well as 11 self-reported variables on the frequency of pro-environmental behavior (e.g., saving water, using public transport or recycling). Our main variables of interest are income and subjective well-being (measured as life satisfaction) as well as the set of variables related to pro-environmental behavior (see Table 1).

Measuring Pro-environmental Behavior

The UKHLS includes 11 questions on environmental behaviors, which were administered in only 3 survey waves: 2009/2010, 2012/2013, and 2018/2019. Respondents were instructed as follows:

“Now a few questions about the environment. Please look at this card and tell me how often you personally do each of the following things.” The behaviors are:

1. Leave the TV on standby overnight
2. Switch off lights in rooms that are not being used
3. Keep the tap running while brushing your teeth
4. Put on more clothes when feeling cold rather than turning on or up the heating

5. Decide not to buy something because it has too much packaging
6. Buy recycled paper products such as toilet paper or tissues
7. Take your own shopping bag when shopping
8. Use public transport (e.g., bus, train) instead of traveling by car
9. Walk or cycle for short journeys of less than 2 or 3 miles
10. Car share with others who need to make a similar journey
11. Take fewer flights when possible

Responses were recorded on a 5-point Likert scale: 1 = “always,” 2 = “very often,” 3 = “quite often,” 4 = “not very often,” and 5 = “never,” plus an additional category of “not applicable, cannot do this.” For some items, the scale was reverse coded in the questionnaire. For consistency, we recoded all items to a scale ranging from 0 to 4, where higher values represent more frequent pro-environmental behavior and 0 indicates that the respondent never performs the behavior. “Not applicable” responses were recoded as missing, ensuring that the mean values in Table 1 reflect the average behavior among respondents for whom the activity is applicable. Participation rates vary substantially across behaviors, with “taking fewer flights” being relatively uncommon and showing a high proportion of missing values.

To capture overall engagement, we created an index of pro-environmental behavior by summing the recoded scores for all 11 items. This “sum index of pro-environmental behavior” ranges from 0 to 44, with the maximum value attained when a respondent reports “always” engaging in every behavior. This index should be interpreted as a measure of individuals’ general propensity to engage in a broad range of pro-environmental behaviors rather than as a direct measure of environmental impact. The 11 behaviors differ in their ecological significance, meaning that identical changes in the index do not necessarily correspond to identical reductions in environmental pressure.

To account for this heterogeneity, we complement the aggregate index with analyses distinguishing more specific behaviors. First, we distinguished between “symbolic” pro-environmental behaviors (not having TV on standby (1), not leaving tap water running while brushing teeth (2), turning off lights in unused rooms (3), not buying something due to excessive packaging (5), using recycled paper (6), and bringing own shopping bags (7)) and “impactful” pro-environmental behaviors (dressing warmer instead heating (4), using public transport (8), walking instead of using car (9), car sharing (10), and taking fewer flights (11)).⁵ We also use an alternative grouping, separating pro-environmental behaviors related to energy (items 1, 2, 3, and 4), mobility (items 8, 9, 10, and 11), and shopping (items 5, 6, and 7). These sub-indices were created as well by simply summing up the respective behavior frequencies. In our analysis, we later focus in the estimation sample only on individuals who reported all 11 behaviors as applicable (and not missing), so as to have a comparable sample of individuals (if a person rates only 1 behavior as applicable, they would score very low in the index otherwise).

Measuring Subjective Well-being

Subjective well-being is assessed using the question on life satisfaction: “Please tick the number that best reflects how dissatisfied or satisfied you are with the following aspects of your current situation . . . Your life overall.” Responses are given on a 7-point Likert scale, ranging from 1 (“completely dissatisfied”) to 7 (“completely satisfied”). Despite their brevity, such single-item measures have been shown to be valid and reliable (e.g., Krueger & Schkade, 2008; Lucas, 2018).

Measuring Income

For income, we use information about equivalized household net income and include a dummy variable that accounts for reporting zero income (although this happens so rarely in our sample that the dummy variable percentage rounds to zero in the descriptive statistics table). These dummy variables help with the fact that there is a discontinuity in income at zero, that is, people having no income are drastically different from those having some income (e.g., being unemployed, or retired or students, etc.).

Measuring Control Variables

In addition to the main analysis variables, we include personal and socio-demographic controls as defended above: age and age² (divided by 100), dummy variables for educational attainment (“1st degree,” “other higher degree,” “A-level

Table 1. Summary Statistics for the Estimation Sample.

Variable	Mean (%)	sd
Life satisfaction	5.24	1.41
Sum index PEB	20.03	6.28
Sum index PEB (impact)	6.19	3.09
Sum index PEB (symbolic)	13.04	4.19
Sum index PEB (shopping)	5.08	2.62
Sum index PEB (mobility)	4.62	3.22
Sum index PEB (energy)	10.33	3.25
Pro-Env. Behavior 1: TV	2.26	1.78
Pro-Env. Behavior 2: Lights	3.36	0.97
Pro-Env. Behavior 3: Water	2.34	1.64
Pro-Env. Behavior 4: Heating	2.37	1.26
Pro-Env. Behavior 5: Packaging	0.79	1.00
Pro-Env. Behavior 6: Recycled paper	1.48	1.29
Pro-Env. Behavior 7: Shopping bags	2.81	1.44
Pro-Env. Behavior 8: Public transport	1.19	1.27
Pro-Env. Behavior 9: Short journeys	1.75	1.36
Pro-Env. Behavior 10: Car share	0.88	1.16
Pro-Env. Behavior 11: Fewer flights	0.79	1.26
Green self-image (GSI)	1.68	0.85
log(Equiv.net household income)	7.45	0.87
Dummy: No net household income	0.00	
Education: Degree	0.28	
Education: Other higher degree	0.13	
Education: A-level etc.	0.21	
Education: GCSE etc.	0.20	
Education: Other qualification	0.08	
Age	47.18	17.15
Age (squared term)	25.20	16.77
year dummy (2012/2013)	0.36	
year dummy (2018/2019)	0.32	
Female (y/n)	0.54	
Region: North East	0.04	
Region: North West	0.11	
Region: Yorkshire and the Humber	0.08	
Region: East Midlands	0.07	
Region: West Midlands	0.08	
Region: East of England	0.09	
Region: London	0.12	
Region: South East	0.13	
Region: South West	0.09	
Region: Wales	0.06	
Region: Scotland	0.09	
Region: Northern Ireland	0.06	
Observations	57,355	

etc.,” “GCSE etc.,” “other qualification,” with “none” as the base category), and region and year dummies. Gender and personality traits are implicitly controlled for through the two-way fixed effects estimator, which differences out all time-invariant confounders.

For the life satisfaction analysis, we also control for green self-image (Binder et al., 2025), measured by the UKHLS question: “Which of these would you say best describes your current lifestyle?” Responses range from 1 (“I don’t really do anything environmentally-friendly”) to 5 (“I’m environmentally-friendly in everything I do”), with 1 as the base category. Given its broad scope and relatively low correlation with specific pro-environmental behaviors, we

Table 2. Two-way Fixed Effects Estimation on the Relationship Between Income and Pro-environmental Behavior.

Dependent variable: pro-environmental behavior indices						
	(1)	(2)	(3)	(4)	(5)	(6)
	All	Symbolic	Impact	Energy	Mobility	Shopping
House. net income equiv. (log)	-0.31*** (-4.30)	-0.11* (-2.39)	-0.18*** (-4.83)	-0.08+ (-1.94)	-0.14*** (-3.58)	-0.09** (-3.04)
No income reported (y/n)	-1.94* (-2.09)	-1.03 (-1.61)	-1.14* (-2.39)	-1.03* (-1.97)	-0.42 (-0.82)	-0.49 (-1.18)
Control variables: education, age	Yes	Yes	Yes	Yes	Yes	Yes
Region and year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	57,355	57,355	57,355	57,355	57,355	57,355
F-statistic	35.36	26.93	20.35	12.02	39.68	102.08
Degrees of freedom	40,438	40,438	40,438	40,438	40,438	40,438
Adjusted R-squared	.04	.04	.03	.02	.05	.13

Note. t statistics in parentheses.

+ $p < .10$. * $p < .05$. ** $p < .01$. *** $p < .001$.

interpret this variable as a proxy for green self-image rather than a direct measure of actual pro-environmental behaviors. Zero-order correlations for all control variables (see Table A1 in Appendix A1) are well below conventional thresholds for multicollinearity concerns.

Results

Income and Pro-environmental Behavior

Do increases in income affect pro-environmental behavior? Table 2 presents the results of this analysis. With respect to income changes, we find that increases in (log) equivalized household net income are associated with decreases in pro-environmental behavior. The effect sizes are rather small: for instance, a doubling of income reduces the overall pro-environmental behavior index (ranging from 0 to 44) by $b = -0.31^{***}$, which corresponds to roughly 5% of the index's standard deviation. The coefficients are also negative when distinguishing between symbolic pro-environmental behavior ($b = -0.11^*$) and impactful pro-environmental behavior ($b = -0.18^*$). Because the two indices use different scales, direct comparison of coefficients is misleading. When expressed in standard deviations (SDs), a one-unit increase in log income corresponds to a 6% SD decrease in the impactful sub-index, compared to a 3% SD decrease in symbolic behavior. Although both effects are small, this suggests that income changes have a weaker influence on symbolic behavior than on impactful behavior.

When decomposing the overall pro-environmental behavior index further into energy, mobility, and shopping behavior, coefficients remain consistently negative: energy ($b = -0.08^+$, 2% of an SD, marginally significant), mobility ($b = -0.14^{***}$, 4% of an SD), and shopping ($b = -0.09^{**}$, 3% of an SD).⁶ Among these, the mobility component exhibits the largest effect, though even here the association is not particularly large. To sum up: while consistently and statistically significantly negative, all effects are on the order of 5% of an SD of the dependent variable and should be classified as substantively small to tiny effects.

Pro-environmental Behavior and Subjective Well-being

In Table 3, we regress life satisfaction on our different pro-environmental behavior indices. The results consistently indicate a lack of evidence for a relationship between pro-environmental behavior and life satisfaction in our sample.

Table 3. Two-way Fixed Effects Estimation on the Relationship Between Pro-environmental Behavior and Subjective Well-being.

Dependent variable: subjective well-being (life satisfaction)			
	(1)	(2)	(3)
Sum index PEB	−0.00 (−0.39)		
Sum index PEB (symbolic)		−0.00 (−0.77)	
Sum index PEB (impact)		0.00 (0.31)	
Sum index PEB (shopping)			−0.00 (−0.31)
Sum index PEB (mobility)			0.01 (1.32)
Sum index PEB (energy)			−0.01 + (−1.71)
Control variables: green self-image, income, education, age	Yes	Yes	Yes
Region and year fixed effects	Yes	Yes	Yes
Observations	57,355	57,355	57,355
F-statistic	11.99	11.51	11.17
Degrees of freedom	40,438	40,438	40,438
Adjusted R-squared	.02	.02	.02

Note. *t* statistics in parentheses.

+*p* < .10. **p* < .05. ***p* < .01. ****p* < .001.

The coefficients are so close to zero that they do not have any substantive meaning, and none are statistically significant at conventional levels. Note that while the zero-order correlation Table A1 shows a small but statistically significant zero-order correlation of $r = .04^{***}$ between our pro-environmental behavior sum index and life satisfaction, this relationship is no longer observed in a more rigorously controlled multivariate framework, which also accounts for a range of unobserved, time-invariant confounds that are difficult to control for in cross-sectional models (specifically, e.g., personality traits).

Heterogeneity Analysis

The above regression results provide averages and may mask heterogeneity in the relationships of interest. To explore this, we conduct additional sensitivity analyses to examine whether the estimated effects differ by gender, age, education, as well as regarding the broader region of the United Kingdom the individuals live in. To do so, we treat these variables as moderators in the relationship between income and pro-environmental behavior (Table 4), and pro-environmental behavior and life satisfaction (Table 5). This involves interacting our main independent variable with each of the moderators in separate models to assess, for example, whether gender changes the relationship between income and pro-environmental behavior.

For the regressions on the relationship between income and pro-environmental behavior (Table 4), joint *F*-tests for the interaction terms with the main variables are statistically significant, but nearly none of the interaction terms themselves are significant at conventional levels of statistical significance. Most notably, there is a small age interaction suggesting that a negative income coefficient will turn positive after the age of 65, that is, the negative relationship between income and pro-environmental behavior decreases linearly in age.⁷ Joint *F*-tests for the pro-environmental behavior-subjective well-being relationship (Table 5) are not statistically significant with the exception of the regional moderator. For this model a few regions exhibit significant moderation effects compared to the base category of the London region. All other moderation analyses do not yield statistically significant results. Apart from age (in the regressions Income → PEB), this means that our analysis does not provide evidence that gender and education moderate any of our main relationships of interest (and there is inconsistent evidence for the regional variable). While generally, a much higher degree of statistical power is necessary to detect interaction effects as compared to main effects (e.g., Gelman et al., 2020, Sec. 16.4 O’Boyle et al., 2019), our large sample size gives some form of reassurance that these nonsignificant results do not reflect a lack of power.

Table 4. Heterogeneity Analysis: Income → Pro-environmental Behavior.

	(1)	(2)	(3)	(4)
	Gender	Age	Education	Region
Income	−0.28** (−3.28)	−0.65*** (−3.75)	−0.11 (−0.72)	−0.22 (−1.57)
No net household income (y/n)	−1.92* (−2.07)	−2.21* (−2.36)	−1.86* (−2.02)	−1.81 ⁺ (−1.89)
Female	−4.89* (−2.24)			
Female × Income	−0.05 (−0.52)			
Age	−0.10 (−0.69)	−0.22 (−1.60)	−0.09 (−0.67)	−0.10 (−0.68)
Age (squared term)	−0.08** (−2.91)		−0.08** (−2.93)	−0.08** (−2.94)
Income × Age		0.01* (2.28)		
Degree			1.45 (1.13)	
Other higher degree			2.15 (1.46)	
A-level etc			3.41* (2.43)	
GCSE etc			1.70 (1.21)	
Other qualification			2.01 (1.33)	
Degree × Income			−0.16 (−1.05)	
Other higher degree × Income			−0.18 (−0.99)	
A-level etc × Income			−0.34* (−1.97)	
GCSE etc × Income			−0.22 (−1.23)	
Other qualification × Income			−0.21 (−1.10)	
North East				−2.69 (−1.17)
North West				2.23 (1.39)
Yorkshire and the Humber				−1.08 (−0.63)
East Midlands				−0.39 (−0.18)
West Midlands				0.38 (0.17)
East of England				−1.54 (−0.97)
South East				−0.64 (−0.45)
South West				0.64 (0.41)
Wales				0.65 (0.27)
Scotland				−1.64 (−0.99)
Northern Ireland				0.37 (0.19)
North East × Income				0.08 (0.30)
North West × Income				−0.28 (−1.52)
Yorkshire . . . × Income				0.00 (0.01)
East Midlands × Income				−0.14 (−0.53)
West Midlands × Income				−0.18 (−0.66)
East of England × Income				0.07 (0.36)
South East × Income				−0.01 (−0.05)
South West × Income				−0.22 (−1.18)
Wales × Income				−0.20 (−0.72)
Scotland × Income				−0.08 (−0.44)
Northern Ireland × Income				−0.13 (−0.75)
Control variables: education, age	Yes	Yes	Yes	Yes
Region and year fixed effects	Yes	Yes	Yes	Yes
Observations	57,355	57,355	57,355	57,355
F-statistic	32.66	35.20	29.00	23.90
Degrees of freedom	40,438	40,438	40,438	40,438
Adjusted R-squared	.04	.00	.04	.04

Note. t statistics in parentheses.

⁺p < .10. *p < .05. **p < .01. ***p < .001.

Table 5. Heterogeneity Analysis: Pro-environmental Behavior → Life Satisfaction.

	(1)	(2)	(3)	(4)
	Gender	Age	Education	Region
Sum index PEB	-0.00 (-0.18)	-0.01 (-1.56)	0.01 (0.59)	0.01 + (1.70)
Female	1.20 (1.34)			
Female × Sum index PEB	-0.00 (-0.08)			
Age	-0.05 (-1.14)	-0.02 (-0.41)	-0.05 (-1.14)	-0.05 (-1.21)
Age (squared term)	0.03*** (4.86)		0.03*** (4.86)	0.03*** (4.90)
Sum index PEB × Age		0.00 (1.50)		
Degree			-0.05 (-0.18)	
Other higher degree			0.22 (0.79)	
A-level etc			0.04 (0.18)	
GCSE etc			0.01 (0.05)	
Other qualification			-0.29 (-1.11)	
Degree × Sum index PEB			-0.00 (-0.44)	
Other higher degree × Sum index PEB			-0.02 (-1.43)	
A-level etc × Sum index PEB			-0.01 (-0.83)	
GCSE etc × Sum index PEB			-0.01 (-0.67)	
Other qualification × Sum index PEB			0.00 (0.17)	
North East				0.37 (0.94)
North West				0.40 (1.39)
Yorkshire and the Humber				0.36 (1.18)
East Midlands				0.44 (1.56)
West Midlands				0.63* (2.23)
East of England				0.45 + (1.73)
South East				0.19 (0.79)
South West				0.32 (1.19)
Wales				0.68* (2.08)
Scotland				0.09 (0.25)
Northern Ireland				1.55*** (3.90)
North East × Sum index PEB				0.00 (0.25)
North West × Sum index PEB				-0.01 (-0.80)
Yorkshire . . . × Sum index PEB				-0.01 (-1.36)
East Midlands × Sum index PEB				-0.02 (-1.40)
West Midlands × Sum index PEB				-0.02* (-2.16)
East of England × Sum index PEB				-0.02* (-2.04)
South East × Sum index PEB				-0.01 (-1.04)
South West × Sum index PEB				-0.01 (-1.08)
Wales × Sum index PEB				-0.03* (-2.43)
Scotland × Sum index PEB				-0.01 (-1.28)
Northern Ireland × Sum index PEB				-0.02 + (-1.88)
Control variables: green self-image, income, education, age	Yes	Yes	Yes	Yes
Region and year fixed effects	Yes	Yes	Yes	Yes
Observations	57,355	57,355	57,355	57,355
F-statistic	11.13	11.15	10.12	8.65
Degrees of freedom	40,438	40,438	40,438	40,438
Adjusted R-squared	.02	.01	.02	.02

Note. t statistics in parentheses.

+ $p < .10$. * $p < .05$. ** $p < .01$. *** $p < .001$.

Discussion

Our analysis does not provide evidence that engaging more frequently in pro-environmental behavior would enhance subjective well-being, as measured by our life satisfaction measure. This contrasts with the overall sentiment in the field that there is a “robust association” between pro-environmental behavior and subjective well-being (see the meta-analyses by Zawadzki et al., 2020 and Krumm, 2024). But it is worth noting that our analysis is arguably methodologically stronger than the typical study reporting an association: most such studies rely on cross-sectional observational data, often with smaller sample sizes and fewer, often haphazardly selected control variables. There are studies that make up the cited meta-analytic averages that contain no control variables at all. As Binder et al. (2025) note, specifically highly relevant confounds such as personality traits or green self-image are accounted for by few other studies. In this way, there is no contradiction in that zero-order correlations or insufficiently controlled regression analyses will provide statistically significant (positive) correlations (remember our zero-order correlation is $r = .04^{***}$ as well), but such a correlation would ultimately disappear in more rigorously controlled studies (e.g., Verhofstadt et al., 2016) or experiments (Prinzing, 2024). Our findings are in line with earlier research on the topic for the United Kingdom (see, Binder & Blankenberg, 2017; Binder et al., 2020b, with fewer data points) or results from Spain (Binder et al., 2020a) and Germany (Binder et al., 2025) that all account for the aforementioned confounds (personality, green self-image). Taken together, these findings suggest that previous findings of a positive association may be attributable to confounding factors.

But even though our analysis employs a more rigorous approach to selecting and defending control variables, we cannot exclude the possibility that some relevant confounds, or existing confounds were not fully accounted for due to measurement imprecision (see below), which could introduce bias into the estimated relationships and their causal interpretation. One advantage of the two-way fixed effects estimator is, however, that all time-invariant unobserved confounds, such as pro-environmental values, political preferences, or personality traits are automatically removed. Few studies use such an estimator in this research field, which underscores the need for better data to better support evidence-based policymaking.

Yet, even when we use a more robust selection of control variables, longitudinal analyses with fixed effects based on observational data may not fully address the issue of reverse causality.⁸ This limitation is further compounded by the fact that, although we have panel data spanning a long period, the main pro-environmental behavior variables were only measured at three distinct time points (namely 2009/2010, 2012/2013, and 2018/2019). As a result, short-term changes in our variables of interest may be overlooked, which could also explain the lack of evidence we found for the relationship between pro-environmental behavior and life satisfaction. If any life satisfaction gains are transitory, individuals would have adapted to an initially positive payoff of acting pro-environmentally, hence our estimator showing a null effect. If well-being slowly returns to its initial levels, the estimator cannot pick up this short-term well-being gain. Of course, it remains unclear whether a short and transitory well-being gain would be enough to positively reinforce pro-environmental behavior, let alone habitualize it after the potential short-term well-being payoff wears off. Since we do not know about levels of pro-environmental behavior and well-being between the years analyzed, we cannot speculate how quickly a potential short-term gain in well-being wears off.

Nevertheless, we would caution to attribute a difference in findings purely to this difference in estimators, as related work has also shown that with the same data, using cross-sectional OLS only yields a positive association between pro-environmental behavior and life satisfaction when not controlling for green self-image and personality traits (Binder & Blankenberg, 2017, Table 5). In sum, for the UKHLS data set, there does not seem to be evidence, neither in levels nor changes, of any relationship between pro-environmental behavior and life satisfaction.

As regards our findings on income and pro-environmental behavior, similar considerations apply. We have observed that increases in income are associated with, on average, small reductions in the performance of common pro-environmental behaviors over the panel horizon. Here, too, we do not know whether short-term effects differ, but in this case, we have no reason to believe that there should be a short-term positive effect that turns negative after a couple of years and even if this were the case, what matters for the environment is likely not a short-term effect of pro-environmental behavior, but rather longer term (habitualized) behaviors. Again, the coefficients for the relationship between income and pro-environmental behavior are rather small, but the findings raise doubts about whether societal gains in productivity can easily or automatically translate into more pro-environmental behavior. Finally, the environmental implications of economic growth may depend not only on income but also on how productivity gains are allocated. Future research could examine whether shorter working hours, rather than higher consumption, encourage more pro-environmental behavior.

Given the mixed findings in the literature on income and pro-environmental behavior, one might expect there to be heterogeneity regarding different types of pro-environmental behaviors, but our analyses provide little support for this expectation. True, coefficient sizes vary somewhat between the different sub-indices of pro-environmental behavior (i.e., shopping-, energy- and mobility-related behavior), but overall, these differences are small and all in the order of 5% of a SD of the pro-environmental behavior indices for a doubling of income. Large increases in income thus do decrease pro-environmental behavior, but only slightly so. This is also the case when distinguishing symbolic and impactful pro-environmental behaviors. Here, too, and in line with expectations from the literature, impactful behaviors decrease more sharply than symbolic (and less costly) ones, but the effect size is still so small that we should not overinterpret this finding.

In exploratory analysis, we have also investigated how working hours may play into this relationship (see Appendix A2): working time can play a critical role in the relationship between income and pro-environmental behavior, as high-income societies tend to be characterized by long working hours alongside high consumption levels. In this regard, time pressure can foster unsustainable convenience consumption, while reductions in paid working time have been shown to reduce environmental pressures (Antal et al., 2020; Fitzgerald & Schor, 2023). In our exploratory analysis, we find negative relationships between longer working hours and pro-environmental behavior (holding income constant, see Table A2 in the Appendix). While we do not have the space to pursue this issue further in this article, these findings suggest that productivity gains could potentially be translated into shorter working hours rather than higher consumption, thereby encouraging more pro-environmental behavior.

Please note also that what we report in our analysis is the self-reported frequency of behavior, not a more direct measurement of environmental impact. How large the environmental impact of the different behaviors executed is, remains open, but few studies succeed in more objective measurement of the environmental impact of their respondents due to the high cost of eliciting such information reliably (often, this is also inferred from individuals' responses about their lifestyle, not direct measurements of energy used etc.). Drawing on self-reports of pro-environmental behavior is a (necessary) limitation of our study, as it is for much of the literature, and there is ample evidence that such self-reports have limited value in predicting actual behavior (e.g., Blankenberg et al., 2026; Kormos & Gifford, 2014). Because all our variables are self-reported, common-method variance is an additional concern. However, the use of fixed effects estimation, which differences out time-invariant factors such as personality traits, helps to address at least one source of such variance. Respondents may also suffer from imperfect recall or exaggerate their past behaviors due to social desirability bias. Given the prominence of climate change in public discourse, some respondents may have felt pressured to portray themselves as "greener" than they are, even though the survey was large-scale and anonymous.

We also cannot rule out measurement imprecision in some of the variables. For example, we rely on an ad hoc measure of green self-image that has not been validated, as the UKHLS does not offer a better alternative. Similarly, multi-item measures of life satisfaction, which would be more reliable, were not included in the survey. Moreover, the 11 pro-environmental behaviors were selected somewhat ad hoc (cf. Poortinga, 2022). Due to the content of these items, our index of pro-environmental behavior is weighted heavily toward mobility-related behaviors, which are known to decrease with rising income. Other potentially pro-environmental behaviors might show different associations with our variables of interest, but as energy- and mobility-related behaviors account for a large proportion of humankind's environmental impact (Preisendörfer & Dickmann, 2021, p. 136), this focus on the 11 behaviors was not such a bad choice by the data providers.

Overall, our results provide a cautionary note on too optimistic versions of the double dividend narrative. Although the negative relationship between income and pro-environmental behavior is small in our analysis, it suggests that even in a developed economy, further gains in material affluence do not automatically translate into greener lifestyles, with somewhat larger declines observed for more impactful than symbolic behaviors. This is compatible with critiques about affluence-driven lifestyles being detrimental to goals related to sustainability (e.g., Wiedmann et al., 2020). Nor do we find evidence that greater engagement in pro-environmental behavior is associated with higher life satisfaction. Together, these findings contribute to debates on sustainable welfare by suggesting that continued economic growth, if primarily translated into higher consumption, is unlikely by itself to promote either environmental sustainability or individual well-being. Instead, achieving sustainable welfare may require broader behavioral and institutional changes that support both.

A related issue concerns the role of inequality and social comparison. Our analysis focuses on absolute household income, whereas environmentally relevant consumption decisions may also depend on individuals' relative economic position. From this perspective, the modest negative relationship between income and pro-environmental behavior observed in our study may partly reflect broader mechanisms of status competition and positional consumption that are not captured by absolute income alone. Future research could therefore examine whether perceived inequality, or income relative to relevant peer groups, helps explain variation in pro-environmental behavior.

Conclusion

In this article, we have analyzed whether higher income fosters pro-environmental behavior, and whether these pro-environmental behaviors, in turn, enhance subjective well-being, thereby creating a double payoff for both individual and the environment. The existing literature provides mixed evidence for both the influence of income on pro-environmental behavior, as well as for the influence of pro-environmental behavior on subjective well-being. To help clarify these relationships, we drew on rich household panel data from the United Kingdom and adopted a longitudinal perspective. By using directed acyclic graphs, we provided a more principled approach to the selection of control variables in multivariate regression analysis. This approach moves beyond simple correlations toward less biased estimates of a causal effect, even when only relying on observational data. As such, it places causal interpretation on a firmer footing than typical cross-sectional studies that rely on few or ad hoc control variables. Our results indicate that, for the United Kingdom, income increases were associated with less pro-environmental behavior, particularly in resource-intensive domains such as energy and mobility, but the coefficients, while statistically significant, were very small (and a similar relationship was found in an exploratory analysis for working time).

We also found no evidence of a meaningful, let alone strongly positive, relationship between pro-environmental behavior and life satisfaction (coefficients were practically zero, none were statistically significant at conventional levels). As these are averages, we have further unpacked potential heterogeneities in our relationships at hand, and redid the analyses focused on pro-environmental behaviors related to energy, mobility, and shopping, as well as symbolic versus impactful pro-environmental behaviors. While the overall patterns were consistent across all types of pro-environmental behavior, the effect sizes differed to some small extent: symbolic pro-environmental behaviors were less affected than impactful pro-environmental behaviors, and mobility- and energy-related pro-environmental behaviors were generally more strongly influenced by income. Our sensitivity analyses also provided no evidence that factors such as gender, or education moderate these relationships. The same applies for unobserved factors that could give rise to regional disparities across the United Kingdom. Only regarding age do we find that the negative relationship between pro-environmental behavior and income decreases in magnitude with age.

In sum, our results may be disappointing for proponents of the “double dividend” narrative, and they challenge policy narratives that economic growth can jointly advance sustainability and well-being. From a public policy perspective, these findings have important implications for policymakers. First, economic growth, as currently structured, may inadvertently hinder the adoption of behaviors that are critical for achieving environmental goals. Strategies aimed at reducing working hours could promote more sustainable behavior without reducing incomes, but additional measures are needed to ensure that these policies also contribute to a contribution to overall well-being. Second, policies should prioritize high-impact actions and combine regulatory, infrastructural, and behavioral measures to foster meaningful environmental change. Given that the effects are consistent across demographic groups, systemic approaches are likely to be more effective than policies targeting specific groups.

Appendix A1

Zero-order Correlations

Table A1. Zero-order Correlations for the Estimation Sample.

	Life satisfaction	Sum index PEB	Green self-image	Log income	Age	Highest education
Life satisfaction	1.00					
Sum index PEB	.04*** (.0000)	1.00				
Green self-image (GSI)	.09*** (.0000)	.30*** (.0000)	1.00			
Log(income, equiv.)	.09*** (.0000)	-.02*** (.0000)	.00 (.9235)	1.00		
Age	.07*** (.0000)	.04*** (.0000)	.20*** (.0000)	.01** (.0044)	1.00	
Highest education level	-.04*** (.0000)	-.10*** (.0000)	-.02*** (.0000)	-.20*** (.0000)	.24*** (.0000)	1.00
Observations	57,355					

Note. *P*-values in parentheses.
 + $p < .10$. * $p < .05$. ** $p < .01$. *** $p < .001$.

Appendix A2

Working Time and Pro-environmental Behavior

In addition to the regressions of pro-environmental behavior on income, we also have explored the relationship between working time and pro-environmental behavior. Working time can play a critical role in the relationship between income and pro-environmental behavior. High-income societies tend to be characterized by long working hours alongside high consumption levels. In this regard, time pressure can foster unsustainable convenience consumption, while reductions in paid working time have been shown to reduce environmental pressures (Antal et al., 2020; Fitzgerald & Schor, 2023). However, the relationship is complex as working hours can influence pro-environmental behaviors both in a direct and indirect way, by shaping income levels and the amount of available free time. Comparative research suggests that societies with shorter average working weeks often report lower per-capita emissions (Knight et al., 2013), although leisure time can have a dual impact. On the one hand, having more free time can help overcome barriers to time-demanding sustainable actions such as repairing, gardening, or choosing slower but environmentally friendly transport options (Wynes et al., 2018). It can also encourage community participation, political engagement (Larson et al., 2015), and, when spent in nature, strengthen environmental values and promote spillover into broader pro-environmental behaviors (Whitburn et al., 2019). On the other hand, leisure may be resource-intensive. One can think here of long-distance travel or motorized sports, tendencies that are especially common when paired with higher incomes (Lenzen et al., 2018). Hence, the environmental implications of less working time may depend less on how much time is available than on the ways individuals choose to spend it.

Regressions in which we replace the income variable with a working hours variable (while holding income constant) are presented in Table A2. Similarly to our income variable, we include dummy variables that account for reporting zero working hours. Our main findings support our findings regarding income and pro-environmental behavior: more working time also translates into fewer pro-environmental behaviors. These findings suggest that channeling productivity gains into shorter working hours could be a strategy to promote greater pro-environmental behavior.

Table A2. Two-way Fixed Effects Model on the Relationship Between Working Hours and Pro-environmental Behavior.

Dependent variable: pro-environmental behavior indices						
	(1)	(2)	(3)	(4)	(5)	(6)
	All	Symbolic	Impact	Energy	Mobility	Shopping
Hours worked (log)	-0.34*** (-3.75)	-0.11 + (-1.84)	-0.21*** (-4.12)	-0.10 + (-1.90)	-0.18*** (-3.30)	-0.07 + (-1.66)
No hours worked (y/n)	-1.16* (-2.31)	-0.38 (-1.11)	-0.75*** (-2.64)	-0.40 (-1.39)	-0.54 + (-1.82)	-0.22 (-0.99)
Income	-0.17* (-2.24)	-0.06 (-1.34)	-0.10* (-2.52)	-0.05 (-1.17)	-0.06 (-1.42)	-0.06* (-1.98)
No net household income (y/n)	-0.87 (-0.92)	-0.66 (-1.03)	-0.51 (-1.06)	-0.80 (-1.51)	0.21 (0.41)	-0.28 (-0.65)
Control variables: education, income, age	Yes	Yes	Yes	Yes	Yes	Yes
Region and year fixed effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	57,355	57,355	57,355	57,355	57,355	57,355
F-statistic	34.64	25.30	21.26	11.39	38.80	94.06
Degrees of freedom	40,438	40,438	40,438	40,438	40,438	40,438
Adjusted R-squared	.04	.04	.03	.02	.05	.13

Note. t statistics in parentheses.
 +p < .10. *p < .05. **p < .01. ***p < .001.

Author Note

Views and opinions expressed are those of the authors only and do not necessarily reflect those of the European Union. Neither the European Union nor the granting authority can be held responsible for them.

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Data Availability Statement

Replication materials are available at: <https://osf.io/tvdgz/>. The data are freely available for researchers at: <https://www.understanding-society.ac.uk/>

Notes

1. Pro-environmental behavior is delineated more narrowly by Berthold et al. (2023) as actions that benefit the natural environment (e.g., recycling) and the avoidance of actions that harm the natural environment (e.g., forgoing air travel).
2. The authors also find statistically significant associations with measures of hedonic and eudaimonic well-being, over all measures, the correlation is even higher at $r = .243$. As argued in more detail in Binder et al. (2025), such a correlation coefficient is implausibly high, as it is comparable in effect size to the meta-analytic effect sizes between unemployment and subjective well-being or losing one's spouse and well-being, two of the most harrowing life events that have been analyzed with regard to their subjective well-being impact.
3. In the case of life satisfaction, it has been shown in countless applications that ordered probit models return similar results to OLS, both because the anchoring of the questions induces some linearity in response styles as well as their technical robustness regarding deviations from the related linearity assumption (for this, see, e.g., Ferrer-i-Carbonell & Frijters, 2004; van Praag, 1991).
4. While we may not share the data set (which is freely available to researchers anyways), we provide replication materials in the form of Stata do- and log-files via: <https://osf.io/tvdgz/>.
5. This classification is mostly along the lines of the high- versus low-impact classification in Krumm (2024).
6. Although the individual sub-indices differ somewhat in effect size, the qualitative pattern of the results remains the same throughout. This suggests that the findings for the aggregate index are not driven by one particular subset of pro-environmental behaviors.
7. A few interaction terms in the income-pro-environmental behavior-analysis are also significant at the 10% level, which is not at all impressive, given the large sample size, and hence does not provide strong empirical evidence.
8. Directed acyclic graphs assume the absence of cycles in causal relationships, which makes it difficult to incorporate reverse causality into this framework. Future research could explore the implementation of multi-period directed acyclic graphs to address this limitation and improve our understanding of causal relationships in panel data.

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