CHALLENGES OF AGEING SOCIETIES

Essays about Health at Older Ages and Pension Attitudes

By

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Author's Declaration

I, the undersigned Réka Branyiczki, candidate for the degree of Doctor of Philosophy at the Central European University Doctoral School of Political Science, Public Policy and International Relations, declare herewith that the present thesis is exclusively my own work, based on my research and only such external information as properly credited in notes and bibliography. I declare that no unidentified and illegitimate use was made of work of others, and no part the thesis infringes on any person's or institution's copyright. I also declare that no part the thesis has been submitted in this form to any other institution of higher education for an academic degree.

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Statement of inclusion of joint work

I confirm that Chapter 1 is based on the paper 'Transition shocks during adulthood and health a few decades later in post-socialist Central and Eastern Europe', published at *BMC Public Health*, which was written in collaboration with Anikó Bíró and that my contribution to the paper equals 50% of the input into the paper. Anikó Bíró analyzed and interpreted the SHARE data regarding the long-lasting associations between shocks related to the transition and health a few decades later. I reviewed the literature, supported the data analysis, and was a major contributor in writing the manuscript. All authors read and approved the final manuscript. Anikó Bíró endorses this statement with her signature below.

Budapest, 31 May 2025

Signature of Anikó Bíró

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Abstract

This paper-based dissertation relates to two main challenges of the welfare states in ageing societies: supporting old-age health and providing adequate old-age benefits in a sustainable and socially legitimate manner. Using individual level data from a harmonized survey covering European countries, and from a conjoint survey experiment embedded in a public opinion survey in Hungary, we explore two main topics. First, we analyze the link between financial difficulties during individuals' active years and health two to three decades later, in older age in Europe. Second, we assess public opinion on fair pensions and the drivers of these preferences, and assess the degree to which these preferences and redistributive considerations are represented in Hungary.

In Chapter 2[1] we assess the long-term health implications of hardship experienced as an adult around the transition of post-socialist Central and Eastern European (CEE) countries and compare them with the implications of unrelated shocks experienced both in CEE and in Western Europe. We use a unique retrospective dataset from the Survey of Health, Ageing and Retirement in Europe (SHARE), which enables us to estimate the implications of stressful periods, financial hardships and job loss occurring around the transition (1987–1993) on both subjective and objective measures of health in 2017.

We find that stressful periods, financial difficulties and job loss around the period of transition are generally associated with worse subjective and objective health at older ages in CEE countries, even after netting out the effect of childhood health and demographic factors. However, the consequences of hardships due to the transition are not specific: health implications of these difficulties seem to be similar to the implications of other shocks possibly unrelated to the transition. The prevalence of stress, financial difficulties, and job loss around the transition contributed to the current health disadvantage in the CEE region. Since similar shocks in the West and in CEE countries before or after the transition had similar health implications, our results draw the attention to the long-lasting impacts of psychosocial stress and financial hardship during adulthood on later health over the life course.

In Chapter 3 we turn to pension preferences and assess pension policy bias and the social legitimacy of earnings-related public pensions in Hungary. While it is well documented that citizens of EU countries strongly support public pensions, little is known about public opinion on pension levels and the factors shaping these pension preferences. Consequently, the size of the pension policy bias and the legitimacy of the system remain unclear. We study attitudes towards pensions using a novel approach: we gather information in a conjoint survey experiment

[1]Chapter 2 is a published co-authored paper: Bíró, A., & Branyiczki, R. (2020). Transition shocks during adulthood and health a few decades later in post-socialist Central and Eastern Europe. *BMC Public Health*, 20, 1-14.

on both perceptions and preferences about pensions for hypothetical pensioners, and compare them to observed pensions in Hungary. Drawing on recent advances in the study of policy bias, we take a distance-based approach and use a straightforward scale, the monthly public pension of old age (in HUF), to estimate both the perceived and the actual pension policy bias. We also assess the relative importance of pensioner characteristics in defining a fair pension to infer about the perceived legitimacy of the current earnings-related pension system. Pensioner attributes include main occupation, years of service, and savings, which are indicative of merit (related to reciprocity) and need (financial) of a pensioner.

We find that voters slightly overestimate monthly amounts of public pensions, except at the median, where the estimate is accurate. Consequently, the perceived and actual policy biases are equal at the median: voters would prefer a substantial increase of approximately 25%. While the perceived gap is largest at the median, the actual bias is more pronounced at the lower end of the pension distribution (the relative difference is 31-36% at the 5^{th} and 10^{th} percentiles). The conjoint experiment indicates that a pensioner's former occupation is the most influential factor in defining a fair pension: the better paid the former occupation, the higher the fair pension. The findings suggest a broad public agreement on the legitimacy of the current earnings-related pension system and the central role of merit, but the substantial pension policy bias points to a deficit in output-legitimacy.

Chapter 4 expands on the study of pension preferences by assessing preferences over the progressivity of public pensions in Hungary, to better understand the lack of redistributive pension policy response to the increasing pension inequality in the country. Understanding mass preferences over the distribution of pensions and pension cuts among pensioners, and the drivers of these redistributive preferences are understudied, nevertheless essential for designing politically feasible austerity measures. We assess preferences over two pension policy outcomes that are contrasting in terms of their salience, in a conjoint survey experiment. First, we ask respondents to define a fair replacement ratio (fair pension as a % of former earnings) for hypothetical pensioners, as a proxy for demand for progressivity. Second, we turn to a salient and unpopular retrenchment measure, a cut to the 13^{th} month extra pension, that the respondents must apply to one of the hypothetical pensioners. We measure the relative importance of several pensioner characteristics indicative of financial vulnerability (such as former occupation, service years, and savings) and compare the responses of financially secure and vulnerable respondents to assess the role of economic self-interest and other-regarding considerations in forming mass preferences.

Public opinion on both pension policies suggests that voters prefer a progressive public pension system. Thus, the lack of redistributive policy response to increasing pension inequality is not explained by mass preferences. There appears to be broad consensus that pension replacement rates should be higher for formerly low-paid pensioners and that pension cuts should primarily affect those with well-paid former occupations. Moreover, we find that financially secure voters are more likely to spare vulnerable pensioners from cuts, indicative

of other-regarding preferences. Our results suggest that strategic pension reforms aiming to boost public support by bundling cuts with expansions should focus on compensating more vulnerable, low-benefit pensioners.

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Contents

List of Figures

List of Tables

Chapter 1

Introduction

The welfare state is hailed as a major and distinct European achievement of modern times, which must react to demographic, social, and economic developments to provide social security and to reach a fair redistribution of life chances among its citizens. Among the challenges, demographic ageing stands out as a major process. The population has been noticeably ageing in Europe since the second half of the 20th century, as life expectancy has been increasing and fertility rates have declined. The process is expected to accelerate further as baby boomers retire, and life expectancy keeps increasing but fertility rates remain low (well below the replacement level of 2.1). The ratio of the old-age population (65+) to the working-age population (aged 20-64) in the European Union (EU) is projected to rise from 36% in 2022 to 59% in 2070 (EC, 2024).

Under current policies in the EU, the total costs of ageing, including public spending on pensions, health care, long-term care and education are projected to increase from 24.4% of the GDP in 2022 to 25.6% of the GDP in 2070, driven by the rising pension expenditures - though there is considerable variation both in the levels and in changes in the levels of public spending across countries (EC, 2024). The bulk of age-related expenditure falls into the category of pensions and health care (11.4% and 6.9% of the GDP in 2022, respectively) (EC, 2024).

A core challenge of welfare states in ageing societies is to remain fiscally sustainable while providing adequate support to preserve their social legitimacy. The dominant response to the fiscal pressure of population ageing has been activation: increasing and prolonging participa-

tion in the labor market, often by curtailing early retirement options and increasing the statutory retirement age. The policy discourse of 'active ageing' fits the wider 'welfare to workfare' trend of the last few decades, where benefit entitlement is conditional on participation, and the goal is to enable people to participate in the labor market and in society instead of being completely dependent on social transfers (Andersen & Larsen, 2024; Dingeldey, 2007). The European Commission jointly with the UNECE (United Nations Economic Commission for Europe) launched the Active Ageing Index (AAI), a composite index of indicators of active ageing performance in EU countries to support development and policy learning (UNECE, 2019; Zaidi & Howse, 2017). Apart from measures of employment and societal participation, there are several indicators of health, reflecting that health is a precondition for active participation, a component of the capacity needed to age actively. There is a clear divide between Northern and Western countries that have a higher AAI score, and Central and Eastern Europe (CEE) that lag behind. The gap is evident in all domains of active ageing, including health (UNECE, 2019).

The observed health gap between CEE and the West is the starting point of Chapter 2. Life expectancy has been increasing across Europe, nevertheless the gap between the lowest and highest life expectancy at birth for men in 2022 was 12 years: 69.4 years in Latvia vs. 81.4 years in Sweden. The gap in 2022 was 8 years for women: 77.9 years in Bulgaria and 85.9 years in Spain (EC, 2025). A more telling indicator for the capacity to age actively is the number of healthy life years at age 65. We see a similar gap of approximately 10 years in the healthy life years at age 65 when comparing Romania and Sweden: 4 years for men and 3.8 for women in Romania and 13.5 years for men and 14.3 for women in Sweden. Previous results from the literature (Laaksonen et al., 2001; Steptoe & Wardle, 2001) suggest that the East-West health gap is partly explained by differences in health behaviors and psychosocial factors, whose health impacts are cumulative over the years. The life course theory of health emphasizes that health inequalities evolve due to socioeconomic circumstances and hardships over the life course (Corna, 2013; S. M. Lynch, 2008; Pearlin et al., 2005). According to life course theory, the lived experience of the transition from planned to market economy in CEE, which resulted in a severe transformational recession (Hodgson Geoffrey, 2006; Kornai, 1994;

Sachs, 1996), is expected to explain some of the East-West health gap.

Empirical tests of the health effect of the transition focused so far on adult mortality and short-term ramifications. Mortality rates rose rapidly in the post-socialist CEE countries (Cornia, 2016), especially in the Baltics, and the immediate and short-term impacts of mass privatization included an increase in alcohol-related deaths, heart disease, and suicide rates (Azarova et al., 2017; King et al., 2009; Scheiring et al., 2019). However, the lack of data impeded the assessment of long-term health implications.

Motivated by the paucity of studies about the long-term health ramifications of economic hardship at the individual level, the main research questions of Chapter 2 are the following.

What are the long-term health implications of hardship experienced as an adult around the transition of post-socialist Central and Eastern European (CEE) countries? And were these hardships during the transition distinct in their health implications?

Understanding the long-term health repercussions of economic hardship during working life may inform welfare state efforts to prevent the accumulation of health risks into old age.

The second part of the dissertation addresses a different challenge faced by welfare states in ageing societies: preserving the social legitimacy of public pensions amid growing fiscal pressures to implement retrenchment. There are two main reasons why understanding welfare attitudes in general, and pension preferences in particular, is relevant. First, social legitimacy, or in simpler terms public support for the public pension system is a precondition for its effective and efficient functioning (Roosma, 2016). Second, we also build on the assumption of the literature on policy responsiveness, which states that welfare attitudes influence welfare policy development because democratically elected representatives are incentivized to conform and respond to public opinion (C. Brooks & Manza, 2006).

Attitudinal studies tend to find that welfare states are supported by the public in advanced democracies of Europe, and the support is especially strong for universal programs, such as public pensions (Deeming, 2018; Ebbinghaus & Naumann, 2020; Svallfors, 2012). Elderly people are consistently seen as the most deserving social category among other needy groups (such as the sick and disabled people, unemployed people, and immigrants) (Van Oorschot, 2006; van Oorschot et al., 2017).

We are aware that voters support public pensions and strongly oppose pension retrenchment, but less academic attention has been paid to the social legitimacy of current public pension systems and how well mass preferences over public pensions are represented. The main question motivating the second part of the dissertation is the following:

How well does public pension policy represent mass preferences over the outcomes and redistributive design of the public pension system in Hungary?

The dissertation takes a multidimensional perspective on welfare attitudes. The theoretical model of multidimensional social legitimacy of welfare states distinguishes four conditions of social legitimacy that may be assessed by measuring support for the different dimensions of the welfare state (Roosma, 2016; van Oorschot et al., 2022). Roosma et al. (2013) describes the welfare state by six dimensions[1]: (1) the goals of the welfare state (the broad goals of providing social security, equality of opportunities or outcomes, and social inclusion), (2) range (areas of life in which the welfare state should redistribute), (3) degree (how much the welfare state should redistribute in each area), (4) the redistribution design (forming each redistributive policy based on questions about who should benefit from it, who should contribute to it, and how much, and why and based on what exact conditions), (5) the implementation process (efficient and effective redistribution), and (6) the outcomes of welfare policies (intended and unintended outcomes).

The four conditions of social legitimacy of welfare states are (1) substantive justice, (2) redistributional justice, (3) procedural justice, and (4) just outcomes (Roosma, 2016). Substantive justice is measured by the support for the above mentioned first three dimensions of the welfare state: support for the goals and for the range and degree of welfare state activities. Redistributional justice is based on the assessment of the redistribution design (most importantly who should receive benefits and who should contribute to the common pocket). Procedural justice refers to the support for the implementation process, based on its efficiency and effectiveness. Finally, the public evaluates the outcomes of welfare policies (i.e. do they succeed in providing social security, reducing inequality, or promoting social inclusion).

We apply the framework of social legitimacy of welfare states to a single welfare policy,

[1] An additional dimension of the welfare mix refers to preferences about the role of state, market, civil society and family in redistribution (Roosma et al., 2013).

pension policy and focus on two areas: redistributional justice and just outcomes. These are aspects that individuals assess mostly based on deservingness considerations (Oorschot, 2000). Measuring public opinion on the outcomes and the redistributional design of a policy is not straight-forward. Instead of relying on attitudinal scales measuring support for the current pension system, we infer about the social legitimacy of public pensions by asking respondents to decide on fair pension policy outcomes in Hungary. We take a distance-based approach introduced by G. Simonovits et al. (2019), which allows us to measure the quality of representation of mass preferences over pensions.

In Chapter 3 we measure the distance between fair, estimated, and actual levels of pensions to have meaningful measures of the scale and variation of the perceived and the actual pension policy biases in Hungary. We also assess whether the earnings-related public pension system is congruent with public opinion to infer about the legitimacy of its redistributional design.

The starting point of Chapter 4 is the lack of pension policy response to recent increases in pension inequality in Hungary, which is surprising under the general predictions of political economy theory. We attempt to better understand this puzzle by assessing public opinion on pension progressivity, proxied by preferences over fair pension replacement rates and targeting a cut in the 13^{th} month extra pension. We also study the role of economic self-interest and other-regarding preferences in forming mass preferences over the pension distribution.

1.1 Methodology

To answer our questions, we use individual level data from a harmonized survey covering European countries, and from a conjoint survey experiment embedded in a public opinion survey in Hungary.

In Chapter 2 we analyze the Survey of Health, Ageing and Retirement in Europe (SHARE), which is a cross-national panel database of micro data on health and socio-economic status of individuals aged 50 or older covering 27 European countries and Israel. We use a unique retrospective dataset from the third and seventh waves of the survey, which include retrospective questions about respondents' life history, such as employment history, periods of stress and fi-

nancial difficulties, and health at younger ages. We estimate multivariate logistic regressions of current health indicators (measured in 2017), with binary measures of hardship during the transition as explanatory variables. We include confounding variables in the multivariate models that are likely to influence both health outcomes at older ages and hardships as an active aged adult. We also added interactions between the individual-level shocks and the region (CEE vs. West) to see whether such shocks have a differential association with later health in different regions.

In Chapter 3 and 4 we assess pension preferences based on a conjoint survey experiment embedded in a public opinion survey in Hungary that was fielded in Spring 2024. Conjoint survey experiments are increasingly popular among political scientists to study multidimensional choices, as they allow researchers to estimate and compare the causal effects of several treatment components (Bansak, Hainmueller, Hopkins, Yamamoto, et al., 2021; Hainmueller et al., 2014), while limiting social desirability bias (Horiuchi et al., 2022).

In the conjoint survey experiment we show hypothetical pensioners (pensioner profiles) with seven attributes in tabular form, including former main occupation, years of service, age, sex, family status, number of children, and savings. As the average marginal component effects (AMCEs) are sensitive to the distribution of attributes (De la Cuesta et al., 2022), we aim for approximating the real-world pensioner attribute distributions. The composition of occupations corresponds to the real-world distribution of occupations, when it comes to socioeconomic classes (such as elite, intellectual, white-collar, blue-collar, unskilled workers) and the sector of the job (private vs. public) to represent the pensioner population in Hungary.

Another feature of the AMCE to keep in mind is that it averages individual preferences both across the distributions of possible pensioners and respondents (Bansak et al., 2023), thus it does not represent a majority preference among respondents (Abramson et al., 2022). It gives us the preferences of an 'average voter' instead of the 'median voter', as it combines the direction and the strength of preferences (Bansak et al., 2023): how strong voters feel about a pensioner characteristic and how many of them feel that way. To have a better general understanding of the pension preferences and their drivers, we also report the marginal mean (MM) of each level of the pensioner attributes next to the AMCE. Also, we follow the advice of

T. J. Leeper et al. (2020) and estimate the difference between marginal means when assessing respondent subgroup differences in forming opinion.

In Chapter 3 the outcome variables of the conjoint experiment are the estimated and the fair amount of monthly pension (in thousand HUF) for each hypothetical pensioner according to voters. We supplement our experimental data with observed levels of pensions as reported by the Central Statistical Office (CSO) in Hungary, to be able to estimate the perceived and the actual pension policy bias.

The dependent variables studied in Chapter 4 are the fair pension replacement rate (%) and the probability of a cut in the 13^{th} month extra pension. Targeting a pension cut is the only forced-choice variable of the conjoint experiment, where respondents must must apply the pension cut to one of the hypothetical pensioners.

1.2 Contributions

We outline our main contributions in each paper of the dissertation. In Chapter 2 we are the first to study long-term health implications of the transition in CEE at the micro level, and compare the shocks of the transition to stress and economic strain at other times and elsewhere in Europe. We are also able to corroborate the predictions of life-course theory of health about the cumulative health impacts of psychosocial stress and economic strain in adulthood.

Our contribution in Chapter 3 is mainly empirical. There is a lack of evidence on the precise magnitude and variation of pension policy bias on a meaningful scale. We offer a novel methodological approach to estimate the perceived and actual pension policy biases throughout the pension distribution. We design a conjoint experiment, where profiles are representative of the recipients of the benefit (pension), and we measure perceptions and preferences on the same scale as the policy outcome (monthly pension in HUF), which allows us to compare the results with observational data and quantify awareness and preference gaps, in other words both the perceived and the actual pension policy bias. Our design could be useful in the policy responsiveness literature, as it allows measuring which groups of voters are better represented by comparing the magnitude of policy bias among them.

Additionally, the conjoint experiment allows us to assess the impact of pensioners' characteristics and former labor market activity on the amount of fair pension they receive from respondents. Preferences over the policy outcome and the redistributional design (public opinion on the conditions for receiving the benefit) are indicative about the 'just outcomes' and 'redistributional justice' dimensions of social legitimacy. Our approach may be applied to other domains of the welfare state, contributing to a vibrant literature on welfare attitudes and welfare legitimacy.

Finally, in Chapter 4 we contribute to the existing literature on pension preferences by studying an often overlooked aspect: preferences over the progressivity of public pensions. Instead of using standard attitudinal scales, we elicit preferences about the pension replacement rate and the targeting of a pension cut in a conjoint experiment in Hungary, offering a direct measure of progressivity that is comparable with pension policy outcomes. We are able to compare the impact of several pensioner characteristics that are indicative of vulnerability in forming preferences. We assess subgroup differences in forming opinion by financial security of the respondents to understand the role of economic self-interest and other-regarding considerations - contributing to a vast literature on the individual-level drivers of redistributive preferences that has been focusing mostly on income distribution.

Chapter 2

Transition shocks during adulthood and health a few decades later in post-socialist Central and Eastern Europe

2.1 Introduction

Health of the population of post-socialist Central and Eastern European (CEE) countries lags behind the European Union average (OECD, 2018)[1]. Previous results from the literature (Laaksonen et al., 2001; Steptoe & Wardle, 2001) suggest that the East-West health gap can partly be explained by differences in health behaviors and psychosocial factors. Health behaviors and psychosocial factors are likely to be related to living conditions during and after the communist era, and to the transition itself. Our aim in this paper is to analyze the link between transition shocks and health 2–3 decades later. Our main research question is whether experienced psychosocial stress as an adult around the transition period had adverse health implications observable at older ages in the life course.

After the fall of the Berlin wall, the dictatorship of the Communist Party came to an end in CEE around 1990, and the Soviet Union dissolved in 1991. The transition implied a dramatic

[1]Chapter 2 is a published co-authored paper: Bíró, A., & Branyiczki, R. (2020). Transition shocks during adulthood and health a few decades later in post-socialist Central and Eastern Europe. *BMC Public Health*, 20, 1-14.

restructuring of the economy and of the social security system. Privatizations took place, prices were no longer fixed, and job security disappeared (Kornai, 2006). Overall, the stress and financial hardship due to job losses, increasing insecurity and rising inequalities affected the health status of the population in CEE (Cornia, 2016; Scheiring et al., 2019).

A severe transformational recession was a common phenomenon in the region (Hodgson Geoffrey, 2006; Kornai, 1994; Sachs, 1996), however, the pace of transformation varied from gradual transition to shock therapy across countries (Balcerowicz, 1994; Godoy & Stiglitz, 2007; Popov, 2000). Bohle and Greskovits (2012) grouped post-socialist capitalist systems into neoliberal (Baltic countries) and embedded neoliberal (Visegrad states) types. The former group combined radical marketization with minimal social protection, while the latter compensated the losers of marketization by more generous welfare states (Bohle & Greskovits, 2012).

So far, studies investigating the East-West health gap and the implications of the transition focused mainly on mortality. Zatonski (2007) documents that adult mortality rates in post-socialist CEE countries started to diverge from Western countries in the 1960s. Also, during the period of transition, adult mortality rose particularly rapidly in the post-socialist CEE countries (Cornia, 2016), especially in the Baltics. Rapid and mass privatization was found to be a significant factor in the declines in life expectancy and in the increase in alcohol-related deaths, heart disease, and suicide rates, also pointing to the role of excess psychosocial stress (Azarova et al., 2017; King et al., 2009; Scheiring et al., 2019).

Less scholarly attention has been paid to the long-term ramifications over the life course, in particular to the link between stressors of the post-socialist transition and later health outcomes at the individual level. Taking a life course perspective has gained prominence in several fields of social science from sociology (Elder Jr et al., 2003; Mayer, 2009) to gerontology (Dannefer, 2003; Ferraro et al., 2009) with the common emphasis on assessing the impact of changes over a long period of lifetime, assuming that early events and impulses affect later life outcomes (Mayer, 2009). Similarly, the life course theory of health focuses on how health inequalities evolve due to socioeconomic circumstances and hardships (Corna, 2013; S. M. Lynch, 2008; Pearlin et al., 2005).

Our paper relates to studies of health trajectories following psychosocial stress and eco-

nomic strain in adulthood (J. W. Lynch et al., 1997; Pearlin et al., 1981). We identify the transition from one-party rule and socialism to democracy and market-economy as an immense source of psychosocial stress due to rapidly increasing job insecurity, inequalities and general uncertainty that may have adverse health implications that unfold over the life course.

Overall, our contribution to the literature is threefold: first, we are among the few (Lazareva, 2020) who study health implications of the transition at the micro level, linking experienced stressors of the transition with both self-rated and more objective health outcomes at the individual level. Second, we take a life-course perspective and assess long-term health implications. Finally, we make use of a harmonized dataset across European countries, which enables us to compare post-socialist countries and to contrast the shocks of the transition to stress and economic strain at other times and elsewhere, unrelated to a system change. Since during the transition both the political and the economic system changed drastically, and not only the economic output or employment dropped within a stable socioeconomic system, it is not obvious whether shocks related to the transition or to an economic downturn are comparable in terms of their health implications.

We exploit individual level life history data to increase our understanding on how difficulties, such as stress, financial hardship and job loss around the transition relates to later population health. We demonstrate the present East-West gap in health outcomes of the population above age 50, and test associations between difficulties around the system change and later health of the individual. We analyze differences in these associations across three groups of CEE countries with different pace of transformation and varying post-socialist market economies: the Visegrad countries, Baltic countries, and Southern CEE countries. This division of the CEE countries is in line with the typology of Bohle and Greskovits (2012) and also with the regions defined by Dingsdale (1999), except for Slovenia and the Visegrad countries that belong to the same group. Since despite the similarities, there are non-negligible economic, political, social and cultural differences among the CEE countries within the three country groups, we also look at the associations between individuals' health and transition-related difficulties on the country level.

2.2 Methods

2.2.1 Data source and country coverage

We analyse the third and seventh waves of the Survey of Health, Ageing and Retirement in Europe (SHARE)[2](Börsch-Supan et al., 2013), which is a cross-national panel database of micro data on health and socio-economic status of individuals aged 50 or older covering 27 European countries and Israel[3]. The third and seventh waves include retrospective questions about respondents' life history, such as employment history, periods of stress and financial difficulties, and health at younger ages. Data were collected in 2009 and 2017 respectively, thus even the youngest cohorts of the sample were already of active age during the times of transition. The seventh wave of SHARE questionnaire contains a retrospective questionnaire for all respondents who did not participate in the third wave, as well as a regular panel questionnaire for all respondents who already answered the retrospective questions in the third wave. Thus, each SHARE respondent who participated in the seventh wave answered the retrospective questions exactly once (either in the third or the seventh wave). We group the countries into postsocialist CEE countries (Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovak Republic, Slovenia) and the rest, labeled as 'West' (Austria, Belgium, Cyprus, Denmark, Finland, France, Germany, Greece, Ireland, Israel, Italy, Luxembourg, Malta, Netherlands, Portugal, Spain, Sweden, Switzerland). We split the German

[2]This paper uses data from SHARE Waves 1–7 (DOIs: https://doi.org/10.6103/SHARE.w1.700, https://doi.org/10.6103/SHARE.w2.700, https://doi.org/10.6103/SHARE.w3.700, https://doi.org/10.6103/SHARE.w4.700, https://doi.org/10.6103/SHARE.w5.700, https://doi.org/10.6103/SHARE.w6.700, https://doi.org/10.6103/SHARE.w7.700); see (Börsch-Supan et al., 2013) for methodological details.

The SHARE data collection has been primarily funded by the European Commission through FP5 (QLK6-CT-2001-00360), FP6 (SHARE-I3: RII-CT-2006-062193, COMPARE: CIT5-CT-2005-028857, SHARELIFE: CIT4-CT-2006-028812), and FP7 (SHARE-PREP: N°211909, SHARE-LEAP: N°227822, SHARE M4: N°261982).

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[3]The two panel databases are: Börsch-Supan A. Survey of Health, Ageing and Retirement in Europe (SHARE) Wave 3 – SHARELIFE. Release version: 7.0.0. SHARE-ERIC. 2019. Dataset. doi: https://doi.org/10.6103/SHARE.w3.700 and Börsch-Supan A. Survey of Health, Ageing and Retirement in Europe (SHARE) Wave 7. Release version: 7.0.0. SHARE-ERIC. 2019. Data set. doi: https://doi.org/10.6103/SHARE.w7.700.

sample according to the place of residence on 1st November 1989 (i.e. before the Berlin wall came down).

We split the post-socialist CEE countries (except for East Germany) further into three groups: the Visegrad countries (V4: Czech Republic, Hungary, Poland, Slovak Republic), Baltic countries (Estonia, Latvia, Lithuania) and Southern countries (Bulgaria, Croatia, Romania, Slovenia).

2.2.2 Measures

We assess the current health (as measured in 2017, the 7th wave of the SHARE data) of the 50+ population with several indicators. Self-rated general health is measured on a 5-point Likert scale from excellent to poor, which is a strong predictor of morbidity and mortality (Idler & Benyamini, 1997). From this variable, following the standard approach in the literature (Boerma et al., 2016; Crimmins et al., 2011; Manor et al., 2000; Subramanian et al., 2010), we generate a binary indicator of poor health which equals 1 if the self-rated health is fair or poor, 0 otherwise. Other binary outcome variables indicate whether the respondent suffers from chronic or long-term health problems (long-term illness, henceforth), has a health problem that limits paid work, has certain conditions, such as heart problems, hypertension, diabetes, ulcer, cancer, and chronic lung disease (each condition is assessed by a separate dichotomous variable). Besides reported health conditions, dependent variables include obesity (Body Mass Index 30 or greater) and an indicator of grip strength, which was shown to explain old age disability (Rantanen et al., 1999). Since grip strength, on average, varies by age, gender and the build of the individuals, we create a binary indicator of weak grip strength, which equals one if the grip strength is below the gender, 10-year age group and country specific median of grip strength. For the sake of brevity, we focus on the binary indicators of poor health and long-term illness in the main analysis, as two composite health indicators, with the indicator of long-term illness being more objective. We relegate the results on the other health measures to the Appendix.

To identify shocks around the transition, we look at retrospectively reported periods of stress and financial hardship that started between 1987 and 1993 and at reported end of jobs

between 1987 and 1993 with at least 6 months of gap without employment or immediate retirement afterwards. The latter two indicators measure whether the respondent suffered from economic difficulties, while the stress variable may capture the general burden of uncertainties experienced during the system change as well. These binary measures of hardship are set to zero for those who do not report the analyzed hardship ever (i.e. no stress, no hardship and no end of job with 6 months gap afterwards, except for retirement, respectively). Descriptive statistics are provided in Table 2.1.

Table 2.1: Descriptive Statistics

Variable	West		V4		South		Baltic	
	Mean	SD	Mean	SD	Mean	SD	Mean	SD
Health in 2017:								
poor health	0.393	0.488	0.452	0.498	0.462	0.499	0.642	0.479
long-term illness	0.493	0.500	0.615	0.487	0.494	0.500	0.597	0.490
health limits work	0.213	0.410	0.324	0.468	0.182	0.386	0.318	0.466
any chronic disease	0.529	0.499	0.593	0.491	0.574	0.495	0.606	0.489
hypertension	0.393	0.488	0.480	0.500	0.472	0.499	0.474	0.499
heart problem	0.106	0.307	0.145	0.353	0.103	0.303	0.163	0.370
diabetes	0.130	0.336	0.161	0.368	0.117	0.321	0.100	0.300
ulcer	0.028	0.165	0.052	0.222	0.045	0.208	0.088	0.283
cancer	0.046	0.210	0.044	0.205	0.024	0.153	0.053	0.225
lung disease	0.058	0.234	0.050	0.217	0.045	0.207	0.056	0.231
obese	0.199	0.399	0.299	0.458	0.256	0.437	0.308	0.462
weak grip strength	0.531	0.499	0.517	0.500	0.540	0.498	0.534	0.499
Hardship around transition:								
stressful period	0.137	0.344	0.081	0.272	0.104	0.305	0.123	0.328
financial difficulties	0.056	0.229	0.062	0.241	0.072	0.259	0.120	0.325
job ends with gap after	0.085	0.279	0.109	0.312	0.125	0.330	0.148	0.355
Start of hardship conditional on hardship ever:								
stressful period	0.151	0.358	0.130	0.336	0.161	0.368	0.178	0.383
financial difficulties	0.129	0.336	0.158	0.365	0.155	0.362	0.271	0.445
job ends with gap after	0.242	0.428	0.355	0.479	0.451	0.498	0.314	0.464
Individual characteristics:								
age in 2017	67.396	10.629	65.901	10.069	66.529	10.122	66.624	10.585
female	0.536	0.499	0.559	0.497	0.555	0.497	0.604	0.489
education	0.937	0.685	1.012	0.468	0.940	0.506	1.210	0.553
childhood health	2.192	1.036	2.200	0.981	1.970	0.964	2.616	1.044
hospitalisation during childhood	0.053	0.224	0.063	0.243	0.034	0.183	0.092	0.289
Industry of last job prior 1987:								
agriculture etc.	0.078	0.268	0.186	0.389	0.154	0.361	0.242	0.428
mining	0.013	0.114	0.032	0.177	0.041	0.198	0.010	0.099
manufacturing	0.181	0.385	0.252	0.434	0.305	0.460	0.211	0.408
electricity etc.	0.019	0.137	0.022	0.146	0.023	0.150	0.024	0.153
construction	0.087	0.282	0.077	0.266	0.083	0.276	0.077	0.267
wholesale trade	0.117	0.321	0.081	0.273	0.067	0.251	0.056	0.230
hotels and restaurants	0.033	0.180	0.021	0.143	0.020	0.141	0.022	0.147
transport etc.	0.049	0.216	0.069	0.253	0.080	0.271	0.082	0.274
financial intermediation	0.031	0.172	0.010	0.097	0.006	0.079	0.008	0.091
real estate etc.	0.013	0.114	0.004	0.059	0.001	0.032	0.003	0.057
public admin.	0.093	0.290	0.051	0.220	0.033	0.178	0.033	0.180
education	0.081	0.272	0.071	0.257	0.050	0.217	0.102	0.303
health and social work	0.075	0.264	0.053	0.224	0.034	0.182	0.054	0.227
other community	0.130	0.336	0.072	0.259	0.102	0.303	0.075	0.263
Total number of individuals	43424		12310		10025		8739	

Note: The number of observations vary across variables due to item non-response and due to sample restrictions for the hardship indicators. The indicator of the industry of last job is missing if no working period is reported.

2.2.3 Statistical models

We estimated multivariate logistic regressions of current health indicators, with binary measures of hardship during the transition as explanatory variables. We added the following confounding variables to the multivariate models that are likely to influence both health outcomes

and our explanatory variable (shock indicator): age in 2017, gender, education (categorized as primary, secondary and tertiary, based on the international classification, ISCED-97), the industry code of the last job before the transition, and measures of childhood health (self-evaluated overall childhood health and a dummy for hospitalization during childhood). To account for country specific differences, we included country dummies. We calculated cluster-robust standard errors, clustering on the country level, using the vce (cluster clustvar) option of Stata, as explained by Cameron, Trivedi, et al. (2010), section 3.3.5. All our results are based on weighted data, using calibrated individual weights. Hence, in the weighted sample, smaller countries have smaller weights. Also, with using the calibrated weights, we avoid bias due to unit nonresponse and panel attrition (see Börsch-Supan and Malter (2015) for details).

In the first group of models, we estimated the effects of the shocks around the transition on the subsample of CEE countries for each health outcome and type of hardship, allowing the effects to differ by subgroups of the CEE countries:

$$\Pr(h_{igc} = 1) = \Lambda\left(s_{igc} D_g \alpha_1 + x_{igc} \beta_1 + \gamma_c\right) \tag{2.1}$$

where Λ is the logistic function, h_{igc} is the binary indicator of current health problem of individual i living in country-group g and country c, D_g is a binary indicator of living in country group g, s is the indicator of hardship during transition, x is the set of confounding variables listed above, and γ_c captures the country effects. Our focus is on the exponential of the coefficient vector α_1 (reported in Table 2.2), showing how the odds of a health problem in 2017 relates to having had hardships during transition in a specific country-group. Individuals who never had such hardships (according to the retrospective survey) serve as the comparison group.

To analyze to what extent do the associations between transition related shocks and later health vary within the country groups, we estimate a modified version of Eq. 2.1. Here, we replace D_g with the binary indicators of living in the specific country in the CEE region. Also, we replace s with the binary indicator of experiencing any of the analyzed three shocks around the transition, with individuals who never had such hardships serving as the comparison group. We analyze the three shocks jointly in the country-specific analysis to ensure that we have a sufficient number of observations of transition related shocks in each country.

In the second group of models, we included the Western countries and analyzed a possible interaction between the shocks and the region (CEE versus West) based on the following equation:

$$\Pr(h_{irc} = 1) = \Lambda\left(s_{irc} D_r \alpha_2 + x_{irc} \beta_2 + \varepsilon_c\right) \tag{2.2}$$

where the notation is the same as in Eq.2.1, with country-group specific coefficients replaced with region (r) specific coefficients. The exponential of α_2 (reported in Table 2.3) shows how the odds of a health problem in 2017 relates to having had hardships during transition in CEE or in the West, with individuals who never had such hardships serving as the comparison group.

In the third group of models (Table 2.4), we extended the time period and assessed the impact of shocks in CEE between 1984 and 1996, to see whether difficulties around and probably due to the transition are specific or not:

$$\Pr(h_{ic} = 1) = \Lambda\left(s_{ic} T_{ic} \alpha_3 + x_{ic} \beta_3 + \omega_c\right) \tag{2.3}$$

where the notation is the same as in Eq. 2.1, but instead of estimating country-group specific coefficients, we allow the health implications of hardships to vary with the time period when the difficulties occurred, denoted by T_{ic} (during the transition period versus before or after the transition period).

Finally, we estimated a modified version of Eq. 2.3, where we allow the health implications of hardships to vary with gender, education and age category, restricting the sample again to CEE and considering shocks occurring between 1987 and 1993 (Table 2.5).

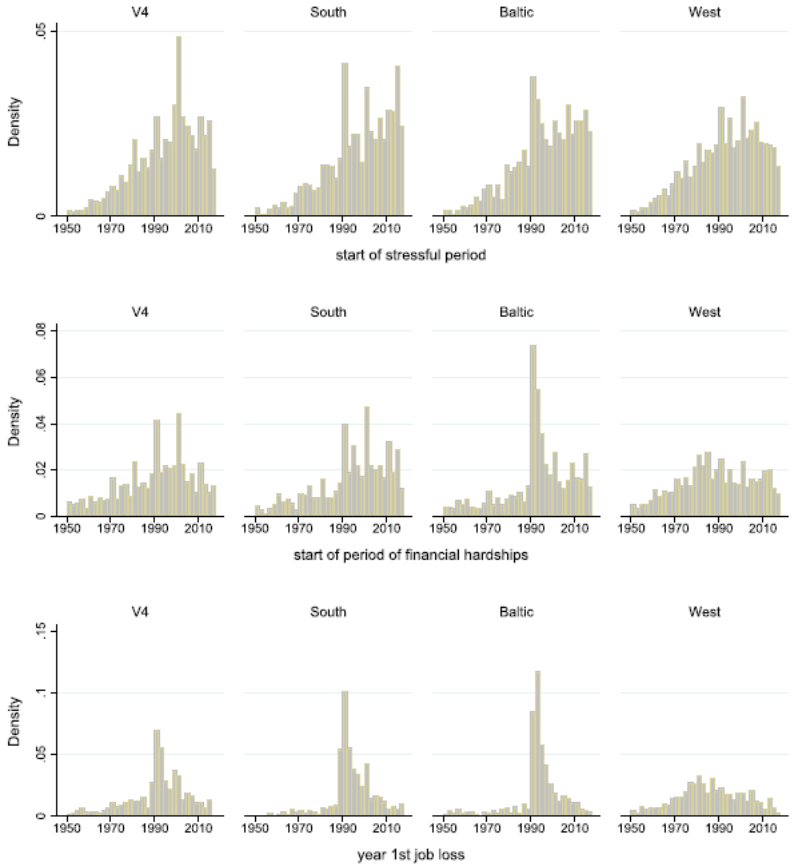
2.3 Results

2.3.1 Shocks around the transition

Looking at the distribution of the start year of periods with difficulties and the year when the first job ends with at least 6 months of gap or retirement afterwards, these have a peak around

the transition (1990) in the postsocialist countries, but not in the West (Figure 2.1 – in the third panel, for illustration purposes, we plot the year of the first job ending, but in the regression analyses we use a binary indicator if any, thus not only the first job ends around the transition).

Figure 2.1: Starting year of stressful periods and financial hardships and end year of 1st job with a 6 months gap or immediate retirement afterwards in CEE and West.



Note: Source: SHARE Waves 3 and 7

The probability that between 1987 and 1993 (i.e. around the transition), someone had a stressful period, a period with financial hardship or an ending of any job with a gap or retirement afterwards is 0.5%point (95% CI: -0.005; 0.014), 3%points (95% CI: 0.018; 0.038) and 17%points (95% CI: 0.158; 0.185) higher in the post-socialist countries than in the West, respectively, among those who experienced the specific shock ever. Overall 19% (95% CI: 0.183; 0.192) of the participants reported at least one of these three shocks around the transition in CEE compared to 15% (95% CI: 0.151; 0.158) in the West. Among those, who experienced at least one of these shocks ever, 31% (95% CI: 0.308; 0.321) had a shock around the transition in the CEE, as opposed to 24% (95% CI: 0.235; 0.244) in the West. Within the CEE group, this fraction is the highest in the Baltic countries with 33% (95% CI: 0.315; 0.338), followed

by the Southern countries with 31% (95% CI: 0.298; 0.321) and the V4 with 27% (95% CI: 0.260; 0.281).

Holding age, gender, education level and the industry of the last job fixed, using West as the comparison country group, and focusing on those who had at least one of the analyzed shocks, respondents from the V4, Baltic and Southern states are 0.5%point (95% CI: -0.018; 0.028), 4.9%points (95% CI: 0.027; 0.071) and 5.3%points (95% CI: 0.029; 0.077) more likely, respectively, to report any of the three analyzed shocks occurring between 1987 and 1993.

Also, among those who stopped working in a job between 1987 and 1993 with a gap or retirement afterward, it is more prevalent in the post-socialist countries that the reason for leaving a job was either being laid off or plant/office shut-down (36% in the West, 95% CI: 0.341; 0.375, versus 62% in CEE, 95% CI: 0.603; 0.634).

2.3.2 Health gap at old age in CEE

Assessing current health of the 50+ population in 2017, we see that self-rated health is on average worse in the CEE countries, but this difference compared to the West is more substantial in the Baltic countries than in the Southern and V4 CEE countries (first part of Table 2.1). Also, most chronic diseases and obesity are more prevalent in CEE, which are known to be influenced by living conditions and lifestyle. However, there are some variations across the CEE country groups, the health status of individuals from the South seems to be on average the most similar to the health of Western people. This is also reflected by the indicators of having any long-term illness and health issues limiting work. On the other hand, based on our data, there is little difference in the prevalence of cancer, chronic lung diseases and weak grip strength between CEE and West. Note, that these statistics are conditional on having survived 20–30 years after the transition. The health differences between the country groups change little if age and gender are controlled for.

2.3.3 Regression results

The first group of models of current health conditions is estimated in the subsample of CEE countries (see Table 2.2 and Table A.1 in the Appendix), with individuals who have never had

the analyzed hardships serving as the comparison group. The results indicate that a stressful period, financial difficulties and job loss around the transition (between 1987 and 1993) are generally associated with worse later health, accounting for observed individual characteristics (demographics and measures of childhood health) and country heterogeneity. The odds of reporting poor health and having a long-term illness increase with reporting shocks around the transition in all three country groups. The results reported in the Appendix indicate that difficulties around the transition are systematically related to higher likelihood of specific chronic diseases and obesity. There are few exceptions, where a negative association is found between health problems and difficulties around transition, such as heart problem and obesity in the South. However, these negative associations are mostly statistically insignificant.

Table 2.2: First Group of Models - Health measures regressed on difficulties occurring between 1987 and 1993 in CEE country groups

Variable	Poor Health		Long-term Illness	
	Coefficient	CI	Coefficient	CI
Stress x V4	1.728***	[1.502–1.988]	2.610***	[2.243–3.037]
Stress x South	2.042***	[1.607–2.596]	2.236***	[2.052–2.438]
Stress x Baltic	1.592**	[1.112–2.280]	1.724***	[1.222–2.434]
Observations	17,452		17,452	
Wald test p-value	0.425		0.034	
Fin. difficulties x V4	1.923***	[1.676–2.663]	2.112***	[1.355–1.665]
Fin. difficulties x South	1.771***	[1.263–2.484]	1.549***	[1.111–2.159]
Fin. difficulties x Baltic	1.175***	[1.111–1.243]	1.522***	[1.208–1.917]
Observations	20,503		20,503	
Wald test p-value	0.007		0.121	
Job ends x V4	1.502***	[1.211–3.054]	1.419***	[1.331–1.514]
Job ends x South	1.599***	[1.430–1.788]	1.343***	[1.222–1.476]
Job ends x Baltic	1.967***	[1.585–2.442]	1.707***	[1.502–1.940]
Observations	20,524		20,525	
Wald test p-value	0.081		0.008	

Note: We control for individual characteristics and country effects. Logit odds ratios are reported. 95% CI displayed in brackets. The Wald test tests the equality of the coefficients of the stress indicators interacted with the country group indicators *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

The associations between the difficulties around the transition and later health are broadly similar across the three country groups of CEE, there is no systematic pattern in the differences in the associations. Figure A.1 in the Appendix shows that difficulties around the transition are associated with worse health in each country of CEE, although these associations are heterogeneous even within the three country groups. Overall, the results suggest that the analyzed relations are stronger in Slovakia within the V4 countries, and in Lithuania within the Baltic countries.

In the second group of models (see Table 2.3 and Table A.2 in the Appendix) we did not

find clear evidence that the health implications of shocks occurring in CEE and the West would be different. We generally see stronger association between the reported difficulties and later health problems in CEE, but Wald tests of the equality of coefficients indicate that most of these differences are statistically insignificant.

Table 2.3: Second Group of Models- Health measures regressed on difficulties occurring between 1987 and 1993 in CEE and West

Variable	Poor Health		Long-term Illness	
	Coefficient	CI	Coefficient	CI
CEE x stress	1.563***	[1.233–1.980]	1.966***	[1.473–2.623]
West x stress	1.526***	[1.114–2.088]	1.506***	[1.343–1.689]
Observations	35,273		35,276	
Wald test p-value	0.826		0.075	
CEE x fin. difficulties	1.773***	[1.408–2.232]	1.588***	[1.243–2.029]
West x fin. difficulties	1.670***	[1.343–2.077]	1.780***	[1.540–2.057]
Observations	43,539		43,541	
Wald test p-value	0.731		0.399	
CEE x job ends	1.502***	[1.368–1.650]	1.238*	[0.979–1.564]
West x job ends	1.270	[0.931–1.733]	1.282***	[1.094–1.501]
Observations	43,344		43,344	
Wald test p-value	0.177		0.707	

Note: We control for individual characteristics and country effects. Logit odds ratios are reported. 95% CI displayed in brackets. The Wald test tests the equality of the coefficients of the stress indicators interacted with the country group indicators *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

While there is a peak in the timing of difficulties around the transition in CEE, difficulties occurring before or after the transition have similar health implications (Table 2.4 and Table A.2 in the Appendix).

Table 2.4: Third Group of Models — Difficulties Around Transition in CEE

Variable	Poor Health		Long-term Illness	
	Coefficient	CI	Coefficient	CI
Transition x stress	1.695***	[1.429–2.010]	2.045***	[1.812–2.306]
Before/after transition x stress	1.562***	[1.203–2.028]	1.940***	[1.378–2.731]
Observations	19,174		19,174	
Transition x fin. difficulties	1.923***	[1.515–2.441]	1.993***	[1.581–2.512]
Before/after transition x fin. difficulties	1.780***	[1.428–2.219]	1.589***	[1.187–2.127]
Observations	21,837		21,837	
Transition x job ends	1.379***	[1.174–1.620]	1.441***	[1.189–1.745]
Before/after transition x job ends	1.403***	[1.147–1.715]	1.156	[0.804–1.660]
Observations	21,773		21,774	

Note: We control for individual characteristics and country effects. Logit odds ratios are reported. 95% CI displayed in brackets. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

We report heterogeneity analysis results in Table 2.5. The heterogeneities in the health implications of difficulties around transition by gender, education or age group are mostly statistically insignificant. The results suggest that the health implications of difficulties around

transition are stronger among men and the younger. The health implications of stress around transition seem to be stronger among the lower educated, whereas the health implications of financial difficulties and job loss around transition are stronger among the more educated.

Table 2.5: Heterogeneity analysis results - Health measures regressed on difficulties occurring between 1987 and 1993 in CEE

Variable	Poor Health		Long-term Illness	
	Coefficient	CI	Coefficient	CI
Stress	2.206***	[1.728–2.817]	2.650***	[1.742–4.032]
Stress × Female	0.701**	[0.512–0.960]	0.814	[0.475–1.394]
Observations	17,452		17,452	
Stress	2.707***	[1.652–4.434]	3.097***	[1.573–6.098]
Stress × Secondary Education	0.655*	[0.416–1.030]	0.787	[0.415–1.493]
Stress × Tertiary Education	0.645	[0.321–1.297]	0.615	[0.269–1.407]
Observations	17,452		17,452	
Stress	1.743***	[1.374–2.212]	2.135***	[1.873–2.433]
Age $<$ 36 in 1990 × Stress	1.092	[0.756–1.577]	1.219*	[0.965–1.540]
Observations	17,452		17,452	
Fin. difficulties	1.910***	[1.494–2.441]	1.874***	[1.315–2.669]
Fin. difficulties × Female	0.874	[0.679–1.125]	0.942	[0.686–1.293]
Observations	20,503		20,503	
Fin. difficulties	1.629	[0.868–3.060]	1.496***	[1.295–1.729]
Fin. difficulties × Secondary Education	1.090	[0.599–1.983]	1.259	[0.902–1.757]
Fin. difficulties × Tertiary Education	1.157	[0.544–2.461]	1.060	[0.649–1.730]
Observations	20,503		20,503	
Fin. difficulties	1.939***	[1.185–3.170]	1.744***	[1.310–2.321]
Age $<$ 36 in 1990 × Fin. difficulties	0.853	[0.517–1.407]	1.066	[0.753–1.507]
Observations	20,503		20,503	
Job ends	1.880***	[1.466–2.410]	1.625***	[1.223–2.159]
Job ends × Female	0.736	[0.465–1.166]	0.787	[0.500–1.237]
Observations	20,524		20,525	
Job ends	1.019	[0.881–1.178]	1.082	[0.866–1.351]
Job ends × Secondary Education	1.672***	[1.407–1.988]	1.383**	[1.070–1.787]
Job ends × Tertiary Education	1.148	[0.861–1.532]	1.017	[0.740–1.398]
Observations	20,524		20,525	
Job ends	1.415***	[1.257–1.593]	1.342**	[1.030–1.749]
Age $<$ 36 in 1990 × Job ends	1.211**	[1.011–1.451]	1.090	[0.694–1.714]
Observations	20,524		20,525	

Note: We control for individual characteristics and country effects. Logit odds ratios are reported. 95% CI displayed in brackets. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

2.4 Discussion

Using retrospective data, we analyzed whether the hardships around the transition contributed to the health gap between post-socialist CEE and western Europe. We showed that the era of post-socialist transition was indeed more often associated with the start of stressful periods, financial difficulties and the termination of a job with a gap or immediate retirement afterwards in CEE than in the West. Within the analyzed CEE countries, difficulties around the transition

were most often reported in the Baltic states; also, respondents from the Baltic states reported on average the poorest health. These results correspond to the macroeconomic evidence that the transition had the most severe economic effect in the Baltic states (Milanovic et al., 1998).

We found evidence that stressful periods, financial difficulties and job loss around the period of transition are mostly associated with worse health at older ages in all groups of CEE countries, even after netting out the effect of childhood health and demographic factors, implying that psychosocial stress as an adult around the transition cumulated over the life course into weaker health at older ages. The associations between the three different stressors (stress, financial difficulties and job loss) and later health are comparable, thus they seem to be similarly important to mitigate during the life course.

While the transition was a drastic restructuring of the political-economic systems absent in the West, we found that major difficulties are negatively related to later health both in CEE and the West, indicating that the consequences of hardships due to the transition are not specific, health implications of these difficulties seem to be similar to the implications of other shocks possibly unrelated to the transition, such as an economic crisis. Nevertheless, the transition implied major difficulties for 19% (95% CI: 0.183; 0.192) of the individuals in CEE, whereas in the West only 15% (95% CI: 0.151; 0.158) experienced hardships during the same period. Also, among those who experienced any of the analyzed difficulties ever, those difficulties were 7.5%points (95% CI 0.060; 0.091) more likely to occur around the transition in the CEE than in the West. Thus, not the transition-specific nature of the difficulties, but the higher fraction of individuals experiencing them around the transition contributed to the current health disadvantage in CEE.

Heterogeneity analysis revealed that the health implications of difficulties around transition were stronger among males and the younger, although most of the gender and age differences were statistically insignificant. The gender differences are in line with the literature which established that the mortality consequences of transition were stronger among males (Azarova et al., 2017; Brainerd, 2001; Cornia, 2016). Stronger health implications of socio-economic strain among those who were hit by the transition at a younger age point to the risk of accumulating health disadvantage over the life course.

Our study is subject to a set of limitations. The results are conditional on having survived around 20 years after transition, thus we could not estimate the mortality effects. If the mortality rate was higher among those whose health was most affected by the hardships around transition (which is likely to be the case) then the negative health implications of the hardships around transition were even stronger than what our results suggest. Health behaviors could not be analyzed due to data limitations. Also, reverse causality is possible from persistent health problems to reporting hardships related to the transition, therefore our results indicate associations rather than causal effects. As an alternative identification, we compared individuals experiencing hardship around the transition with individuals not reporting hardship or facing difficulties only after the transition and these more conservative estimates are in line with the reported results.

2.5 Conclusions

Overall, our results draw the attention to the long-lasting impacts of psychosocial stress and financial hardship during adulthood on later health. Hence, our analysis relates to the literature analyzing the effect of the recent financial crisis and austerity on health (Gili et al., 2013; Karanikolos et al., 2013; Stuckler et al., 2017).

As the different types of stressors have similar associations with later health, policies that have the potential to alleviate more of these interrelated experiences, such as employment protection or activation policies are worth considering. We suggest including alleviating youth unemployment on the policy agenda given that some of the health implications of difficulties around transition were stronger among the younger. However, further research is needed to formulate suitable policy recommendations, both in terms of policy tools and target groups.

In general, at times of economic recession or widespread political and economic restructuring when a large proportion of the population is directly affected by rising insecurity, policy makers should also consider the health implications of their policy responses to mitigate the cumulating health disadvantages over the life-course.

Chapter 3

What is a fair pension? Policy bias and social legitimacy of public pensions

3.1 Introduction

How well does public pension policy represent mass preferences?[1] As the population ages, the majority of advanced democracies face constant pressure to adjust their pension system to remain sustainable and adequate. Governments in representative democracies have a strong incentive to be responsive to public opinion to remain in office. At the same time, it is rather unknown what the public would consider a fair public pension and how well the current public pension system represents mass preferences.

We focus on the representation of mass preferences over the 'basic welfare questions' about 'who should get what and why' regarding public pensions. We compare preferred and observed pension levels, and preferred and observed pension calculation rules, as the alignment of popular deservingness considerations with the outcomes and redistributional design of the pension system is the basis of its social legitimacy (Oorschot, 2000).

We build on two main lines of research: the representation of mass preferences literature, and the welfare attitudes literature. We outline some of the main considerations of both fields that are relevant for our study and then introduce our own approach to assessing the represen-

[1] This paper is an unpublished manuscript.

tation of pension preferences and the social legitimacy of public pensions in Hungary.

3.1.1 Measuring representation of mass preferences

Representation of mass preferences is mostly studied by assessing policy responsiveness: the alignment of public opinion and policy. There is evidence for the relationship between public opinion and policy, when comparing broad measures of positions in opinion and policy, such as ideological orientation or other attitudinal scales about government responsibility (see the overview by Wlezien (2017)). More liberal American states tend to enact more liberal policies (Erikson et al., 1993; Shapiro, 2011). And mass preferences over government responsibilities are reflected in the welfare state effort of advanced democracies (C. Brooks & Manza, 2008).

However, our knowledge of the representation of mass preferences over actual policies is still limited. The core challenge of the literature is finding a common metric, which allows us to compare public opinion with the actual policy. Matsusaka et al. (2010) suggest comparing majority preferences over dichotomous issues (i.e. allow or prohibit capital punishment) to measure congruence with the prevailing policy. Lax and Phillips (2012) also assess congruence (whether the policy matches majority opinion) by assessing support for several policies in a few issue areas and compare the majority preference with the enacted policy. As opposed to the favorable accounts of representation based on responsiveness, these studies find that policies are often incongruent with majority opinion.

Notwithstanding the value of congruence, it is not directly applicable to a broad range of policy outcomes measured on a continuum, such as the level of taxes, minimum wages, transfers, etc. G. Simonovits et al. (2019) introduce a novel measure of representation: policy bias, which is the distance between the average opinion and the enacted policy. This distance-based approach quantifies the difference between mass preferences and policy outcomes on a meaningful scale (by using the same metric to measure preferences and policy outcomes), which allows for studying continuous policy outcomes, and for comparing the degree of representation both across countries, states, policy areas, and voters.

An assumption behind quantifying policy bias on a meaningful scale (where the metrics of the preferences and the policy outcomes correspond to each other) is that voters have mean-

ingful preferences over the policy outcome that are not detached from reality. It is a common concern in the literature that voters do not have meaningful preferences for particular amounts or levels of policy, hence the dominance of questions about relative preferences (Wlezien, 2004, 2017). Although the concern is valid, it is worthwhile to test it empirically for the sake of opening up the field for novel policies and theories about the drivers of representation quality. G. Simonovits et al. (2019) find that respondents had meaningful preferences over the level of minimum wage on a numerical scale (dollars).

Another remaining challenge is to assess the quality of representation in the case of policies whose outcome may not be summarized by a single measure, i.e. a single amount, but may be rather captured by a distribution of outcomes, as it is often the case for welfare policies with varying levels of benefits for different individuals based on the eligibility and calculation rules (the redistributional design of i.e. public pensions, maternity benefits, family allowances, income taxes and tax allowances, etc.). The preferred benefit level may be close to the actual level for one recipient subgroup and far for another subgroup. As a consequence, there is a lack of evidence on the precise magnitude and variation of policy bias on a meaningful scale for several welfare policies.

We offer a novel empirical approach to study policy bias in the case of complex policy outcomes that have a distribution based on the prevailing rules defining who is affected (subject of the policy outcome: *who gets it*) and by how much (level of the policy outcome: *what is given*). Relying on a conjoint survey experiment with hypothetical subjects, representative of the actual subjects of the policy, and with open ended questions about both the estimated and the fair level of policy outcome for the hypothetical subjects allows us to estimate perceived and actual policy bias and their variation across subjects. We apply our approach to assess the magnitude and variation of pension policy bias in Hungary by comparing the distribution of observed, estimated and fair levels of pensions.

3.1.2 Pension preferences and social legitimacy of public pensions

Attitudinal studies tend to find that welfare states and policies have an overall support from the citizens of advanced capitalist economies in Europe, and the support is especially strong

for encompassing programs, such as public pension (Ebbinghaus & Naumann, 2020; Svallfors, 2012), as in European countries the elderly people are consistently seen as the most deserving social category among other needy groups (such as the sick and disabled people, unemployed people, and immigrants) (Van Oorschot, 2006; van Oorschot et al., 2017). Accordingly, there is a general aversion to pension retrenchment, influenced both by economic self-interest and normative views about the role of the state and about solidarity (Boeri et al., 2002; Häusermann et al., 2019; Jaime-Castillo, 2013). As valuable the findings of the pension preferences literature using attitudinal scales are, we lack measurements of the fair levels of pensions to estimate the precise magnitude of pension policy bias.

We are aware that voters support public pensions and strongly oppose pension retrenchment, but less academic attention has been paid to the social legitimacy of the current public pension system in terms of its perceived fairness. In the social legitimacy of welfare states framework Roosma (2016) (building on Rothstein (1998)) distinguishes four conditions for welfare states to be socially legitimate: substantive justice (the goals are just), redistributional justice (who deserves benefits and who should contribute), procedural justice (efficient and effective implementation of policies), and just outcomes (evaluation of welfare policy outcomes). In the pension preferences literature two dimensions of the framework are addressed: redistributional justice (who deserves benefits and who should contribute) and just outcomes (evaluation of welfare policy outcomes: pensions).

Redistributional justice evaluations of public pensions focus mostly on intergenerational fairness. Intergenerational equity would entail that contributions paid and benefits received are spread fairly across successive generations. Pressure on intergenerational equity is especially strong in rapidly ageing societies with a pay-as-you-go (PAYG) system, which redistributes from the active-aged to the elderly population. Sabbagh and Vanhuysse (2014) find that voters in Germany and Israel see the pension system as rather unjust, especially the younger population, based on opinions (measured on an attitudinal scale) about whether the amount of pension pensioners receive and the contributions of current workers towards pensions are just and fair.

The justice evaluation of the outcomes (pensions) involves a decision about the preferred level of intragenerational equity among pensioners, which is often overlooked. There is a

considerable gender gap in pensions (Bettio et al., 2013) and varying life expectancy also contributes to the inequity of lifetime pensions (Crystal et al., 2017; Sanchez-Romero et al., 2020; A. Simonovits & Lackó, 2023), yet public opinion about an equitable pension system and pension distribution is rather unknown. A notable exception is the study of Castillo et al. (2019), which assesses the drivers of a just pension amount of the fully privatized pension system in Chile, and shows that the public accepts high levels of inequality among pensioners, based on their individual achievement.

Intragenerational equity among pensioners may be assessed based on horizontal or vertical equity in a pension system (Clements et al., 2014). Horizontal equity requires in a PAYG system that individuals with similar contributions receive similar pensions (providing actuarially fair pensions where the rates of return on contributions are uniform), whereas vertical equity requires defining pensions based on pensioners' needs (providing adequate pensions for low-wage earners). Consequently, there is a need for redistribution between pensioners with different means to achieve vertical equity, while horizontal equity does not require redistribution.

Horizontal and vertical equity considerations are closely linked to the two main goals public pensions pursue: (1) replacing labor income to smooth consumption over the lifecycle, and (2) alleviating old-age poverty, so that the basic needs of the older population are met. Prioritizing between the two is often necessary due to fiscal constraints, thus there is often a trade-off between horizontal and vertical equity, which comes down to the dominance of merit (individual contributions) vs. need (financial vulnerability) when defining a fair pension.

We lack knowledge about how strong the correlation should be between contributions (earnings during the active years) and benefits (pensions received during the inactive years) in a public pension system according to the public. Consequently, it remains a question how the government should balance between considering merit (replacing earnings) and need (protecting against poverty) of pensioners when reforming the pension system to ensure its social legitimacy. This question pinpoints to deservingness considerations that are part of the assessment of redistributional justice. Oorschot (2000) outlines five deservingness criteria often referred to as the 'CARIN-model'. The five criteria that underlie people's preferences for supporting

specific groups are control (responsibility for neediness), need (level of need), identity (being in-group), attitude (attitude towards support, i.e. gratefulness), and reciprocity (earned support). The preferred role of merit and need in defining pensions is indicative of the relative importance of two different deservingness criteria: need (level of neediness) and reciprocity (how much of the support is earned by paying contributions) (Oorschot, 2000).

We fill this gap by evaluating the congruence of public opinion on the relative importance of pensioners' merit (former earnings) in defining the monthly amount of public pension with the current earnings-related public pension system in Hungary. Instead of relying on descriptive surveys with attitudinal scales that usually fall short of identifying causal relationships and estimating meaningful magnitudes from a policy perspective, we use a conjoint survey experiment, which measures the role of pensioner characteristics in setting fair levels of monthly pensions for hypothetical pensioners. We include attributes of hypothetical pensioners that are relevant based on the prevailing rules of defining the level of monthly pensions (i.e. service years, former occupation). Measuring the relative importance of the attributes when defining a fair level of pension allows us to assess congruence between preferences over deservingness and prevailing pension laws, indicative of the social legitimacy of the policy in terms of its redistributional design.

3.1.3 Our conceptual and empirical approach

To evaluate the representation of mass preferences over public pensions, we consider two dimensions of a socially legitimate welfare policy. First, the outcome of the policy: the level of monthly pension. Second, the rules of defining that policy outcome (elements of the redistributional design): the eligibility rules and calculation of the monthly pension (who benefits, how much and on what conditions). Consequently, we conceptualize the quality of representation in the case of pension policy based on (1) the distance between individual preferences and policy outcome (output-legitimacy of pension levels), and (2) on the congruence between how individuals define fair pension levels and the prevailing pension laws (indicative of redistributional justice). The distance and congruence between public opinion and pension policy are indicative of the social legitimacy of the outcomes and redistributional design of the public pension

policy.

First, we estimate the precise magnitude and variation of pension policy bias on a meaningful scale (HUF). Estimating the magnitude and variation of pension policy bias enables analyses about whose preferences are represented and which groups of voters are better represented (in line with the question raised by Erikson (2015)), for example to test claims about an emerging gray power. In terms of policy, the magnitude of the bias is an indication of the extent voters are dissatisfied with the status quo, which may also be indicative of the urgency of a reform or adjustment of the pension system. The variation of the bias over the pension distribution may suggest which pensioners (low-, middle- or high-benefit) should benefit from a reform and which pensioner should bear the cost of retrenchment to be in line with mass preferences.

Second, we assess whether the drivers of fair pension levels are congruent with the prevailing legislation, which is key to make inferences about the social legitimacy of the redistributional design of the public pension system. Our results could support governments in the balancing act between replacing labor income (which cements inequality due to earnings inequality) and protecting the elderly from poverty. Studying the principles that guide public opinion about a fair public pension (merit/reciprocity vs. need) offers insights into the dominant deservingness criteria, and into what kinds of retrenchment could be justified.

We study public opinion on pensions in a novel setting, where we measure both perceptions and preferences in a conjoint survey experiment about pensions for hypothetical pensioners and combine them with observational data. Instead of standard attitudinal survey items about pensions (measured usually on a Likert-scale), we elicit perceived and preferred fair pensions on a straightforward scale, the monthly old-age pension (in HUF), which is comparable with observed pensions in Hungary. Hungary has an earnings-related public pension system financed by the active population, which is the dominant model in continental Europe. The hypothetical pensioner profiles of the conjoint experiment were designed to represent the Hungarian pensioner population, which allows us to outline the perceived pension distribution and the distribution that the public considers fair. As a result, we are able to estimate the perceived and the actual pension policy bias at several points of the pension distribution by comparing the estimated, the fair and the observed pension levels both at the bottom, the median and the

upper part of the distributions.

Furthermore, we assess the drivers of these pension preferences in the conjoint experiment and quantify their effects on the same meaningful scale (HUF). We focus on the relative importance of pensioner characteristics that are indicative of merit and need, including the former occupation, service years, and savings of hypothetical pensioners. We also study whether self-interest and ideological stance of the respondents influence preferences and whether they mediate the role of merit and need when defining a fair pension.

Our results suggest that respondents would increase current pensions to reach a fair level: the actual policy bias is 55 thousand HUF (144 EUR) at the median, which translates into a substantial preferred raise of 27%[2]. Respondents are aware of pension levels throughout the distribution, but somewhat overestimate them, especially at the lower and upper end of the distribution. As a result, the actual policy bias tends to be larger than the perceived one. The former occupation of pensioners is the main driver of defining their fair pension, indicating the dominant role of merit in fairness considerations about public pensions - in line with the prevailing earnings-related pension system. Overall, the Hungarian pension policy is congruent with a preference for an earnings-related system, however observed pension levels fall short of the preferred levels, pointing to a substantial pension policy bias, and a deficit in output-legitimacy.

Our contribution is mainly empirical. We offer a novel methodological approach to assess the precise magnitude and variation of welfare policy bias. Our approach enables assessments of the quality of representation of welfare preferences (in our case pension preferences), (1) by designing a conjoint experiment, where profiles are representative of the recipients of the benefit (pension), and (2) by measuring perceptions and preferences on the same scale as the policy outcome (monthly pension in HUF), which allows us to compare the results with observational data and quantify awareness and preference gaps, both the perceived and actual policy bias. We also investigate the drivers of a fair pension in terms of the pensioners' characteristics and former labor market activity, which is indicative about the social legitimacy of the earnings-related PAYG system.

[2]This would be a substantial raise compared to the usual rates of annual pension increases that are indexed to inflation. Pensions were increased by 6% in 2024.

The paper is structured as follows. First, we describe the Hungarian pension system in a European context. Then we turn to the design of the conjoint experiment and introduce the data and measures we use. In the results section we start by outlining the perceived and actual pension policy bias, namely the gaps between estimated, fair, and observed pensions. Then we assess the drivers of pension preferences based on the conjoint experiment and see whether they are in line with the pension rules of the current system. Finally, we discuss our results.

3.2 The Hungarian pension system in a European context

Hungary is following the Bismarckian model, which is the dominant pension regime in continental Europe, and which was designed with the main aim of replacing labor-income. Again, benefits are earnings-related and are financed on a pay-as-you-go (PAYG) basis, from the contributions of the active population[3]. Due to the fiscal pressure population aging puts on PAYG systems, European countries have been diversifying pension sources, albeit to a varying degree. Alternative pension sources include statutory funded schemes, occupational schemes, and voluntary personal pension schemes. Eastern European countries pioneered in adding mandatory privately funded schemes to the pension system during the late 1990s. By the end of the 2000s, the majority of CEE (Central and Eastern European) countries adopted a 'three-pillar' model as pension privatization diffused in the region (S. M. Brooks, 2005), building on the suggestions of the World Bank (WorldBank, 1994), where the first pillar refers to the public PAYG scheme, the second to mandatory private accounts and the third to voluntary private accounts. However, a wave of reform reversals followed the economic crisis of 2008, and the weight of the mandatory private pillar was reduced significantly, either by nationalizing it (the extreme example of Hungary) or by reducing or suspending contributions to these schemes (Drahokoupil et al., 2014; Naczyk & Domonkos, 2016).

As a result, the PAYG schemes remained dominant in the region, mostly with a defined benefit (Bulgaria, Czech Republic, Estonia, Hungary, Lithuania), some with a point system that takes into account the relative contribution performance of the pensioner when defining

[3]Nordic countries, the UK and Ireland have Beveridgean pension systems that are tax-funded with flat-rate, targeted benefits with the aim of alleviating old-age poverty.

a pension benefit (Croatia, Romania, Slovak Republic) and a few with notional defined contributions (Latvia and Poland) (Commission et al., 2024; Drahokoupil et al., 2014). Overall, balancing between pension adequacy and sustainability is a common struggle among the EU11 countries[4], as the relative income poverty of the elderly and the pensions-to-GDP ratio are both increasing in the region (Domonkos & Simonovits, 2017).

The Hungarian population - similarly to the population of other EU countries - is ageing rapidly due to fewer children born and rising life expectancy, thus the pressure on the pay-as-you-go pension scheme is increasing. It is expected that there will be only 1.8 Hungarians aged 20-64 for each person aged above 65 by 2070, whereas the same ratio was 4.2 in 1995 (OECD, 2024). The government is reluctant to touch the pension system, despite the increasing fiscal pressure and the commitment to the European Commission (as part of Hungary's recovery and resilience plan) to reform the pension system by March 2025 in order to improve its long-term sustainability while preserving adequacy.

3.2.1 Rules of the Hungarian pension system

The statutory retirement age is 65 years for both men and women since 2012 and the minimum service period is 20 years. (A partial pension is provided above 15 years of work.) However, women can retire regardless of their age if they have an eligibility period of 40 years (based on the so-called 'Women-40' policy); there are no other early retirement schemes since their restriction in 2012.

Pension benefits are calculated based on the number of service years and the average of wages. Since January 2012, pensions are indexed to inflation (to the general consumer price index, CPI). All pensioners receive an extra 13th month's pension benefit, since its reintroduction in 2021. There is a minimum pension of HUF 28 500 (EUR 71) per month, though it has not changed since 2008, increasing the gap between the minimum and average pension. As of January 2024 the average old-age pension was HUF 230 940 (EUR 577) and the median was HUF 204 510 (EUR 511) (CSO, 2024a).

[4]The EU11 countries include Bulgaria, Croatia, Czech Republic, Estonia, Hungary, Latvia, Lithuania, Poland, Romania, Slovakia, Slovenia.

3.3 Research design

We conduct a conjoint survey experiment to assess the estimated and the fair amount of monthly pension (HUF) according to respondents, for hypothetical pensioners representative of the Hungarian pensioner population. First, we compare the results with observed pensions to estimate the perceived and actual pension policy bias. Then, we estimate the sensitivities of assigned pension levels to particular pensioner attributes, which allows us to measure awareness about how public pensions are defined and to compare how merit and needs considerations weigh in when deciding about a fair pension.

Conjoint survey experiments are increasingly popular among political scientists to study multidimensional choices, as they allow researchers to estimate and compare the causal effects of several treatment components (Bansak, Hainmueller, Hopkins, Yamamoto, et al., 2021; Hainmueller et al., 2014). In our case, respondents face two pensioners (1st and 2nd) side by side, with seven attributes in tabular form, including main occupation (indicating merit), years of service (indicating merit) and savings (indicating need). The occupations listed in the conjoint experiment correspond to the real-world distribution of occupations, when it comes to socioeconomic classes (such as elite, intellectual, white-collar, blue-collar, unskilled workers) and the sector of the job (private vs. public) to represent the pensioner population in Hungary (see Tables A.6 and A.7 in the Appendix).

The other attributes of the pensioners that may influence the assigned pension are age (indicative of ageism), sex (to check for a potential gender bias[5]), family status (pointing to needs as single households are more vulnerable on average), and number of children (children may need support or could support their inactive parents and their number may influence sympathy for the pensioner[6]). We list the pensioner attributes and their levels in Table 3.1.

The order of the pensioners' characteristics in the conjoint table varies randomly between respondents[7], but it is constant within respondents to avoid confusion. The levels of each

[5]Checking for gender bias in defining fair pensions is particularly interesting in the Hungarian context to see whether the differential pension regulations due to the Women-40 scheme have any public opinion basis.

[6]The idea of rewarding parents with a higher pension based on the number of children is part of the wider 'family-friendly' discourse of the government in Hungary. It is interesting to see whether the public builds such considerations into pension preferences.

[7]See the balance tests and other technical details of the conjoint experiment in Figure A.4

Table 3.1: Pensioner attributes and their levels

Pensioner attribute	Levels	Theory to test
Age	60 / 65 / 66 / 67 / 68 / 69 / 70 / 71 / 72 / 73 / 74 / 75 / 76 / 77 / 78 / 79 / 80 / 85	Aversion to early retirement, ageism
Years of service	20 / 30 / 40 years	Merit
Sex	Male / Female	Gender bias
Savings	Enough for 1 month / Enough for half a year / Enough for more than half a year	Need
Family status	Single / Married	Need
Number of children	0 / 1 / 2 / 3	Need
Main occupation during active years	Hospital CEO / CEO at a private company / Architect / Highschool teacher / Shop assistant / Hairdresser / Nurse / Tailor / Machine operator / Tram driver / Baker / Cleaner / Deliverer / Doorkeeper / Kitchen assistant	Merit and need

Note: The table show the levels of each pensioner attribute that could appear in the conjoint experiment. The final column outlines the theoretical reasons for including each attribute.

attribute appear randomly, but we restrict some of the combinations of pensioner characteristics to drop highly unrealistic pensioner profiles. A former nurse cannot be male, a former machine operator cannot be female, a former hospital CEO and CEO at a private firm cannot have low savings (enough only for a month), and a former cleaner, deliverer, doorkeeper and kitchen assistant cannot have high savings (enough for more than half a year). For an example of a conjoint table with hypothetical pensioners see Table A.5.

Respondents answer a few questions about the pensioner profiles. One of our innovations is to use open-ended questions about pension levels without any anchor (such as former earnings or current pension level) in the conjoint, to avoid priming effects. A further novelty in our design is to elicit both perceptions and preferences. Respondents define either the estimated or the fair amount of monthly pension (in '000 HUF) for each pensioner profile. The estimated amount is their estimate of the actual monthly pension (perception), and the fair amount is the pension level for the pensioner they consider fair (preference) based on the pensioners' attributes. Only one of the questions (estimate or fair pension) appear randomly to avoid confusion and possible priming effects. (See Figure A.2 for an overview of the structure of the conjoint experiment and A.3 for the formulation of the questions.) Finally, a third innovation is to combine the results of the survey experiment with the observed pension statistics. Mea-

suring awareness, preferences, and policy outcomes on the same scale allows us to estimate the gap between estimates, preferences, and observed outcomes. The monthly amount of public pension (in '000 HUF) is one of the most salient pension policy outcomes for voters, and there are statistics available about observed pensions.

In the first part of the study, we assess actual and perceived representation following a distance-based approach (see G. Simonovits et al. (2019) for a similar approach), where we measure how well a policy outcome reflects preferences directly (in '000 HUF). Another advantage of combining representative pensioner profiles with a direct measure of policy outcome in the conjoint experiment is that the responses are indicative of the perceived and fair distribution of pensions according to voters. We pool all the responses about estimated pensions to draw the perceived pension distribution, and we pool all the responses about fair pensions to derive the fair pension distribution. Consequently, we are able to estimate the actual and perceived policy bias at different points of the estimated, fair, and observed pension distributions, among low-, median-, and high-benefit pensioners as well. We also estimate the perceived policy bias for each occupation that appears in the experiment to see which pensioners are under- or overpaid according to the public, based on the former occupation of the pensioners.

We measure pension bias in absolute terms (the gap in '000 HUF) and also in relative terms (in %). Absolute perception bias is the difference between the estimated and the actual pension, relative perception bias is the ratio of the estimated and the actual pension. Similarly, perceived policy bias in absolute terms is the difference between fair and estimated pensions, and in relative terms it is the ratio of fair and estimated pensions. Finally, actual policy bias in absolute terms is the difference between fair and observed pensions and in relative terms it is the ratio of fair and observed pensions.

In the second part of the study, we estimate the so-called AMCE (average marginal component effect) of each feature on the estimated and fair pension separately, which measures the average relative sensitivity of respondents to specific pensioner attributes. We compare the role of former occupation (reflecting merit), service years (merit) and savings (indicative of need), together with a few other pensioner characteristics that may influence estimated and fair pensions, such as age of the pensioner, sex, family status, and number of children. As the AMCE

of a pensioner attribute is sensitive to the other listed attributes of the pensioner, we aim for including all the relevant characteristics of a pensioner without overwhelming the respondent with too much information.

We also measure individual-level differences in preferences to offer insights into the mechanisms of preference formation. We compare younger and middle-aged adults (between age 18 and 59) and the elderly (60 or above 60 years old). We also explore AMCEs conditional on educational level (primary vs. secondary and higher education) and ideological stance (left-wing vs. right-wing). Differences in preferences across the subgroups based on age point to the role of material self-interest in forming welfare attitudes, while the differences between left and right leaning respondents point to the role of ideology and values. Differences in pension attitudes across more and less educated respondents may point to self-interested considerations (assigning higher pension to similarly educated pensioners) and to a general correlation between education and redistributive preferences.

In general, we estimate average marginal component effects (AMCEs) of pensioner attributes on the outcome variables (estimated and fair pension), which are estimated via OLS regressions with dummy variables for the levels of pensioner attributes (Hainmueller et al., 2014). The AMCE averages individual preferences both across the distributions of possible pensioners and respondents (Bansak et al., 2023), thus it does not represent a majority preference among respondents (Abramson et al., 2022). It gives us the preferences of an 'average voter' instead of the 'median voter', as it combines the direction and the strength of preferences (Bansak et al., 2023)), which makes the estimate suitable as an input into policy considerations, where the relative importance of attributes matter also next to majoritarian preferences (how strong respondents feel about a characteristic and how many of them feel that way). To have a better general understanding of the pension preferences and their drivers, we also report the marginal mean of each level of the pensioner attributes next to the AMCE.

For estimating subgroup differences, we estimate the difference between marginal means of the particular levels of pensioner attributes among different respondents (grouped by age, education and ideology), together with an omnibus F-test, where in a nested model comparison, the restricted model without interactions and the model with interactions are compared (as

advised by T. J. Leeper et al. (2020)). For the analysis we use the software called cregg by T. Leeper (2018) available in the Comprehensive R Archive Network and coefficient plots are based on examples provided by T. J. Leeper et al. (2020). In all of the models, we use the weights provided by the survey agency and we estimate robust SEs clustered at the respondent level.

3.4 Data and measurement

Data were collected during Spring 2024 in Hungary by a survey agency, called TÁRKI Social Research Institute. TÁRKI drew a probability sample, representing the adult population of Hungary. They applied random selection sampling and collected data by Computer-assisted Personal Interviewing (CAPI), where interviewers asked respondents in person and recorded the answers on a computer or tablet.

Overall, we have a probability sample of 1072 respondents, representative of the Hungarian adult population. Respondents saw 3 conjoint tables, each with 2 hypothetical pensioners, so we have a total of 6372 hypothetical pensioners to analyze.

A question before the conjoint experiment elicited opinions about pension adequacy. Retired respondents had to indicate the degree to which their pension can maintain their living standards they enjoyed during their active years on a scale from zero (not at all) to ten (fully). Active respondents had to state their expectation about the adequacy of their future pension on the same scale.

In the conjoint experiment one of the following questions was assigned to respondents randomly. One was about the estimated pension: the monthly public pension level (net, in '000 HUF) the respondent thinks each of the hypothetical pensioners would get in reality. The other was about the 'fair pension', the amount of monthly public pension (net, in '000 HUF) the respondent would consider fair for each of the hypothetical pensioners. As a result we end up with 2450 hypothetical pensioners with an estimated pension and 2815 hypothetical pensioners with a fair pension[8] (see Table A.4 in the Appendix). We complement the responses

[8]With our sample sizes our models are powered (power equals 0.8 and significance level is set to 0.05) to detect an effect size of 0.01 in the case of both outcome variables.

with observed amounts of monthly net old-age pensions that come from the Central Statistical Office.

The variables we use when analyzing subgroup differences in preferences include age groups (adults and elderly, defined as below age 60 and age 60 and above), education level (grouped into primary vs. secondary and higher education), and ideological stance (defined as left- vs. right-wing, scoring 0-5 vs. 6-10 on an attitudinal scale from 0 to 10, where 0 stands for left-wing and 10 for right-wing.)

3.5 Results

We present our results in two main sections. First, we assess pension policy bias. We start with descriptive results on attitudes towards public pensions. Then we describe public opinion on pension adequacy and then compare the estimated, fair and observed pension levels to calculate the perceived and actual policy bias. We also present the estimated and fair pensions by the former occupation of the hypothetical pensioners and show the association between occupational prestige and fair pensions. Second, we assess the drivers of public opinion on pensions to infer about the social legitimacy of the earnings-related system. We report the effect of pensioner characteristics on estimated and fair pensions as well.

3.5.1 Assessing pension policy bias

Public opinion on pension adequacy

Most pensioners (55%) think that their pension is far from enough to maintain the living standards they enjoyed during their active years on the labor market. The corresponding share is 63% among active respondents (non-pensioners), and 20% of them answered that their future pension would not be enough at all to maintain their living standards enjoyed during their active years[9]. Hence, active respondents are even more pessimistic about the adequacy of their future

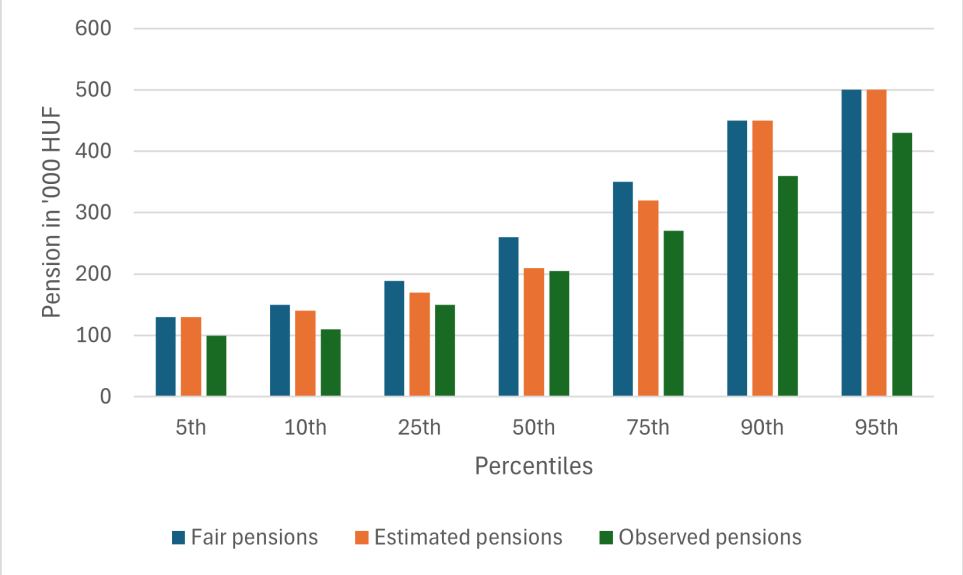
[9]Both pensioner and active respondents were asked about the degree to which their pension is or will be enough to maintain the living standards they enjoyed or enjoy on a scale from 0 to 10, where 0 means that the pension is not adequate at all to maintain living standards during active years and 10 means that the pension is totally adequate to maintain living standards. We considered a score 4 or under as the pension being far from enough to maintain living standards. See Figure A.5 in the Appendix for the distribution of responses.

pension, which is in line with a recent study on pension expectations of the youth in Hungary (Vaskövi, 2024).

Estimated, fair and observed levels of pensions

This section maps the estimated (perceived), the fair (considered as fair) and the observed pensions in Hungary, to assess perception bias (the gap between estimated and observed pensions), policy bias (the deviation between fair and observed pensions), and finally the perceived policy bias (the difference between fair and estimated pensions). We measure each bias at several percentiles of the pension distributions and for different occupations of the hypothetical pensioners.

Figure 3.1: Fair, estimated and observed pensions ('000 HUF) at several percentiles of each distribution



Note: The figure shows the amounts of monthly public pension ('000 HUF) that respondents consider fair next to the estimates of the pension and the observed pension at the 5th, 10th, 25th, 50th, 75th, 90th and 95th percentile of each pension distribution. Data source: observed pensions come from the Central Statistical Office, January 2024 (CSO, 2024b), fair and estimated pensions come from the conjoint experiment, own data collection.

The distribution of responses suggests that most respondents could provide meaningful answers about pensions on a numerical scale, even without an anchor (without i.e. the average pension or a pension amount corresponding to the pensioner profile). Figure 3.1 shows that the median of the pension estimates is almost identical to the actual median pension. Apart from the median, the public tends to somewhat overestimate pensions. The gap between estimated and actual pensions is highest at the lower and upper end of the distribution. Table

Table 3.2: Relative and absolute gaps between the estimated, the fair and the observed pensions at several percentiles of the pension distribution

	Perception bias		**Perceived policy bias**		**Actual policy bias**	
	Rel.	Abs. ('000 HUF)	Rel.	Abs. ('000 HUF)	Rel.	Abs. ('000 HUF)
p5	1.31	31	1.00	0	1.31	31
p10	1.27	30	1.07	10	1.36	40
p25	1.13	20	1.11	19	1.26	39
p50	1.02	5	1.24	50	1.27	55
p75	1.19	50	1.09	30	1.30	80
p90	1.25	90	1.00	0	1.25	90
p95	1.16	70	1.00	0	1.16	70

Note: The table shows the relative bias and the absolute differences ('000 HUF) (columns Rel. and Abs., respectively). Perception bias in relative terms is defined as the ratio of the estimated pensions and the actual pensions at several percentiles of the pension distributions. Estimated minus actual pension gives the absolute difference. Similarly, the perceived policy bias is calculated as the ratio of the preferred (fair) and the estimated pensions. The actual policy bias is the ratio of the preferred (fair) and actual pensions. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: actual pensions come from the Central Statistical Office, January 2024 (CSO, 2024b), perceptions and preferences come from own data collection.

3.2 lists perception biases in a relative term, as the estimated pension over the actual one. The estimated pension at the 5th percentile is around 30% higher than the observed amount at the 5th percentile, and the bias is similar at the 10th and 90th percentiles. Overall, we see a high level of awareness about pension levels, especially at the middle of the distribution, which is a common pattern among primary, secondary and higher educated respondents (see Figure A.6 in the Appendix). Pensioner and non-pensioner respondents gave similar estimates as well, except for the highest benefits, where pensioners gave higher estimates. The estimate at the 95th percentile was close to 20% higher among pensioners compared to non-pensioners (see Figure A.7 in the Appendix).

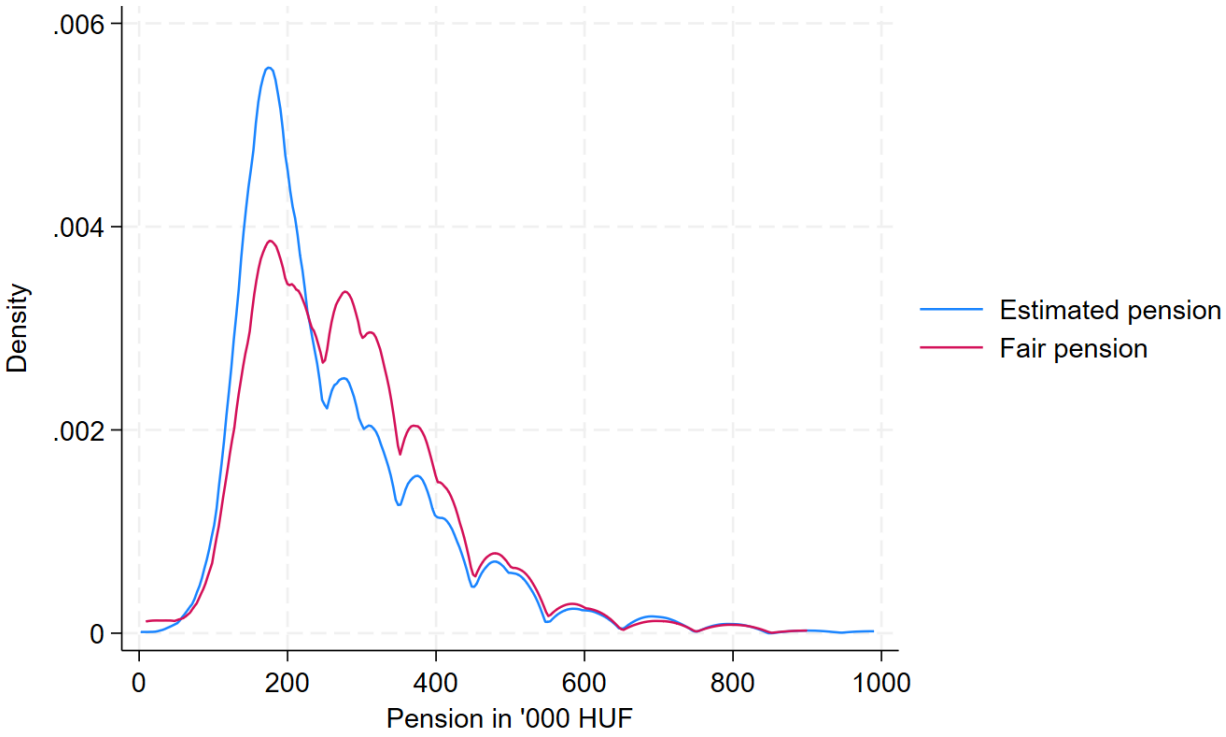
When it comes to fair pensions, we see that the fair amounts are consistently higher than the actual observed pensions throughout the distribution, but we see a gap between fair and estimated pensions only at the middle of the distribution (between 25th and 75th percentiles) (see Figure 3.1). Accordingly, the perceived policy bias is largest at the median, the 50 thousand HUF (EUR 131) difference corresponds to a 24% gap between fair and estimated pensions (see Table 3.2).

Looking at respondent subgroups, we find minimal heterogeneity among respondents based on their education level and pensioner status. The 5th and 25th percentiles of fair pensions are slightly higher among the primary educated (6-8% higher than among secondary and higher educated), whereas the upper end of the distribution, the 95th percentile is slightly larger among

higher educated respondents (4% higher than the amount among primary educated), otherwise the percentiles are almost identical among the subgroups (see Figure A.9 and A.8 in the Appendix), so self-interested considerations do not seem to dominate fairness considerations. The same conclusion applies when comparing fair pensions given by pensioners and non-pensioners, who are active on the labor market and are at the contributing side of the PAYG system. Both groups gave similar amounts that are consistently higher than the observed pensions. We found slightly higher amounts at the lower half and the middle of the fair pension distribution among pensioners (8-12% higher amounts than among non-pensioners), and similar amounts at the upper end of the distribution (see Figure A.10 and Table A.9 in the Appendix). Nevertheless, the mode of fair pensions is lower, around 180 thousand HUF (EUR 471) among the active respondents and 280 thousand HUF (EUR 733) among pensioners (see Figure A.11 in the Appendix).

On average, respondents seem content with pensions at the lower and upper end of the distribution, as their preferred fair pension levels equal their perceptions, and it is at the middle of the distribution where participants would increase pensions the most. We find that low pensions below 120 thousand HUF (EUR 314) and high pensions above 450 thousand HUF (1178 EUR) appear in similar proportions as fair and estimated pensions, and we see a divergence within the range of 120-450 thousand HUF (EUR 314-1178), where fair pensions tend to be higher compared to estimated pensions that peak around 180 thousand HUF (EUR 471) (see Figure 3.2).

Figure 3.2: Distribution of fair and estimated pensions (’000 HUF)



Note: The figure shows the Kernel density estimates for the fair and estimated amounts of monthly public pension (’000 HUF). The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: conjoint experiment, own data collection.

Overall, as respondents tend to overestimate pensions, they underestimate the actual pension policy bias. The actual policy bias is higher than the perceived bias throughout the distribution, especially at the lower end (in the case of relative biases). The fair amount of pension is 36% higher than the actual pension at the 10th percentile. The relative measures indicate that the actual bias is between 16-36% compared to the 0-24% perceived bias. The measures correspond to a preferred raise of 0-50 thousand HUF (EUR 0-131) based on perceived pensions and a raise of 31-90 thousand HUF (EUR 81-236) based on actual pensions. The results indicate that participants may be dissatisfied the most with pensions in the middle of the distribution, however, they would be even more concerned with the lower pensions if they had accurate information about the actual pensions.

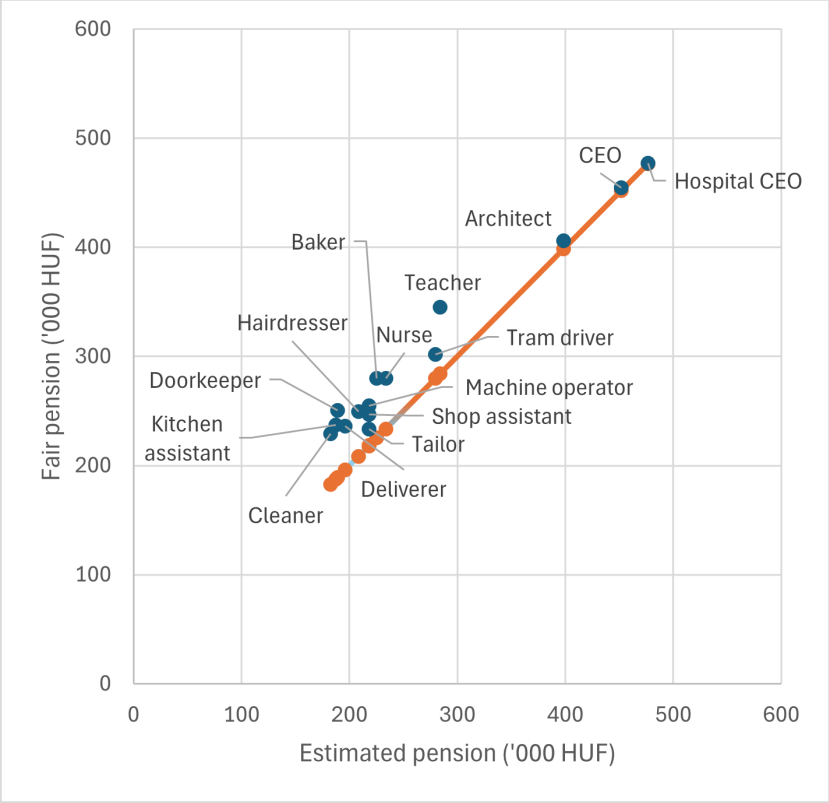
Estimated and fair pensions for different occupations

In this section, we map the estimated and fair pensions for each occupation that appeared in the conjoint experiment. Figure 3.3 shows the average estimates and the average fair amounts for

the occupations, where the gap between the two amounts correspond to the perceived pension bias. The 45 degree line corresponds to a scenario where the perceived bias is zero as perceived and fair pensions are equal on average. Each dot represents an occupation. In general, the higher the occupation on the job ladder, the higher the pensions that respondents estimate and consider fair, and the smaller the perceived bias.

Average estimated and fair pensions of CEOs and architects are close to equal, but average preferred pensions are consistently higher than the average estimates for all other occupations that are lower on the job ladder. The perceived bias is highest for high school teachers and doorkeepers, where the absolute difference is around 60 thousand HUF (EUR 157). Compared to the estimated pensions, respondents would prefer an approximately 20-30% raise for low-skilled occupations (such as bakers, cleaners, deliverers, doorkeepers, and kitchen assistants). Overall, respondents seem to be content with giving much higher pensions to previously high-earning and high-contributing pensioners but would prefer a smaller gap between the pension of a top- and low-earner by increasing the pensions of low-earners.

Figure 3.3: Perceived pension policy bias for several occupations (’000 HUF)



Note: The figure shows the amounts of monthly public pension (’000 HUF) that respondents consider fair on average for each occupation, together with their average estimate of the actual pensions. The closer the dots to the 45-degree line, the smaller the gap between preferred and perceived pensions. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: own data collection.

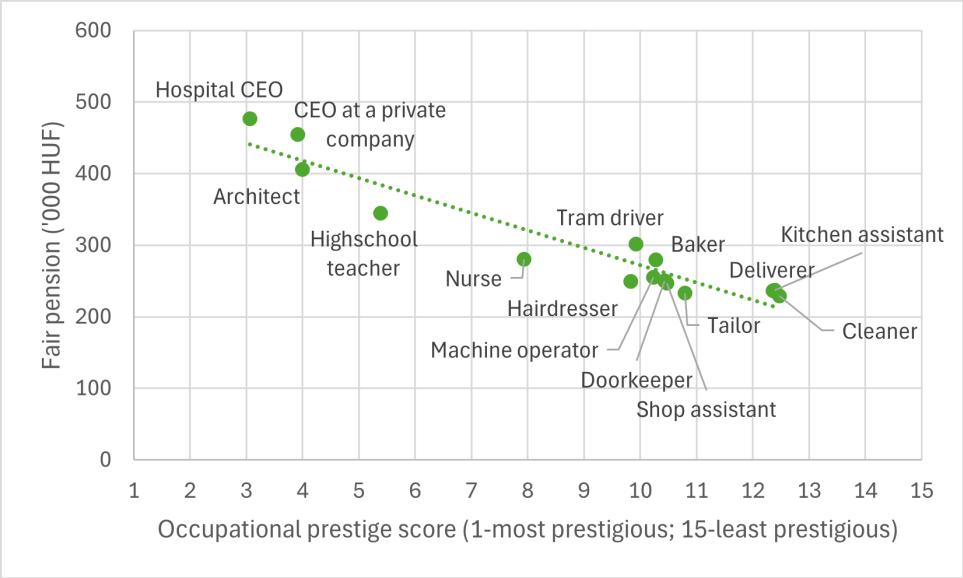
To better understand how voters define a fair amount of public pension to hypothetical pensioners with different former occupations, we compare public opinion about occupation prestige with pension preferences. Occupational prestige is a sociological concept that captures a job's worthiness according to the public. The latest Hungarian occupational prestige scores come from a data collection by the Hungarian Central Statistical Office in 2016 (CSO, 2016)[10]. The prestige scores range from 1 to 15, where the lower the score, the more prestigious the occupation. Figure 3.4 shows the relationship between occupational prestige and the average public opinion about a fair pension for each occupation in the conjoint experiment.

We see a strong correlation between the two: the more prestigious the occupation, the higher the amount of pension respondents consider fair[11]. Most occupations are close to the trend line, but the gap is larger for hospital CEOs and CEOs at a private company, who get a higher average fair pension than their prestige score would predict and for high-school teachers and nurses, who on the other hand get a lower average fair pension than their prestige score would predict. The results indicate that respondents reward occupations that are highly valued in society with higher public pensions, but the above-mentioned 'outlier' occupations suggest that other factors play a role apart from social worthiness. A plausible factor is earnings. The lower fair pension than the corresponding amount based on occupational prestige of teachers and nurses may reflect their notoriously low earnings in Hungary, and consequently their low contributions, while the well-paid CEOs may be outliers due to their 'extremely' high earnings.

[10]The so-called Microcensus survey included 176 occupations that cover the most frequent occupations in society. The respondents (more than 43 thousand) received 15 occupations randomly chosen from the total of 176 and had to rank them based on the prestige of the occupation. The first occupation on the list receives 1 point and the last 15 points. So, on the 1-15 scale, the lower the score, the more prestigious the occupation. In 2016, the most prestigious occupation was hospital CEO; university rector was second and surgeon was third in the ranking. The least prestigious occupation was street prostitute, just after street sweeper (CSO, 2016).

[11]We also see a strong, but slightly weaker correlation between the occupational prestige scores and the average estimated pensions, see Figure A.12 in the Appendix

Figure 3.4: Prestige score and fair pension for the occupations



Note: The figure shows the occupational prestige score (from 1-most prestigious to 15-least prestigious) and the amounts of monthly public pension ('000 HUF) that respondents consider fair on average for each occupation. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: occupational prestige scores are from the the Central Statistical Office, 2016 (CSO, 2016) and fair pensions come from own data collection.

3.5.2 Assessing the legitimacy of earnings-related public pensions

Drivers of the estimated and fair pensions

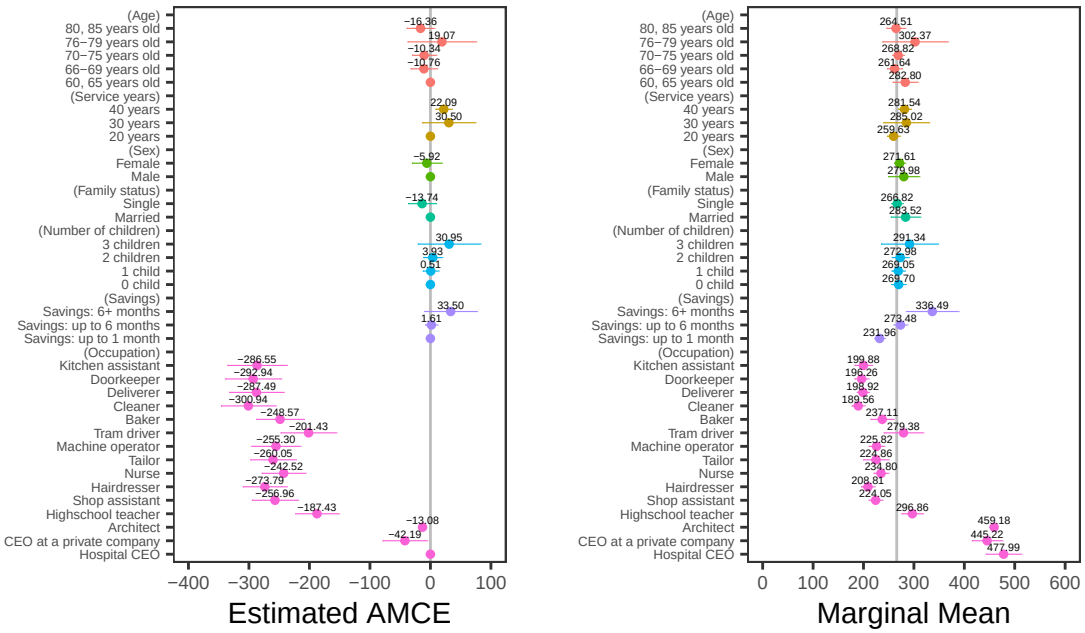
We assess the impact of pensioner attributes on the estimated and the fair amount of monthly pensions. The drivers of estimated pensions reflect perceptions about how public pensions are defined, and the drivers of fair pensions reflect preferences about how public pensions should be defined. In particular, opinions about fair pensions indicate which pensioner attributes would voters prefer to take into consideration. We also compare the drivers of pension preferences among three socio-economic subgroups defined by age (adults and elderly), education level (primary vs. secondary and higher), and ideology (left-wing vs. right-wing voters) of respondents, respectively.

Figure 3.5 shows the average marginal component effect (AMCE) of each pensioner characteristic on the estimated monthly pension ('000 HUF) on the left side and the marginal mean of the attributes on the right side. Both estimates are reported along their cluster-robust 95% confidence intervals. In the case of AMCEs, the reference category is depicted by a dot on the vertical zero line. The marginal mean describes the mean fair pension ('000 HUF) for pensioner

profiles with a particular feature level, all else equal. The vertical line shows the overall mean, which is 266 thousand HUF (approximately 696 EUR), 15% higher than the actual average monthly pension[12].

In terms of drivers, occupations strongly influence the estimates: the higher the former earnings, the higher the estimates, thus respondents understand well that the public pension is earnings-related. The largest gap is between hospital CEOs and cleaners: a cleaner is expected to receive 300 thousand HUF (around 765 EUR) less than a hospital CEO. Participants are also aware that the length of service years matters, although they seem to underestimate its role in increasing pensions. The AMCE is only significant in the case of 40 years of service (compared to the baseline of 20 years, which is the minimum for eligibility). None of the other pensioner characteristics are taken into consideration when estimating the public pension (no other AMCEs are significant), which is generally in line with the current regulations of the system.

Figure 3.5: The effects of pensioner attributes on the estimated pension ('000 HUF)



Note: Figure shows the average marginal component effects and marginal means of pensioner attributes. The dependent variable is the estimated amount of pension ('000 HUF) for each pensioner. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Point estimates are reported above the dots. The dots on the zero line denote the reference category for each attribute, when estimating AMCEs. The vertical line on the right panel indicates the overall mean of estimated pensions. Weights are applied. (The average exchange rate in January 2024: 1 EUR = 382 HUF.)

Again, occupation is the main driver of the assigned fair pension levels. The higher the

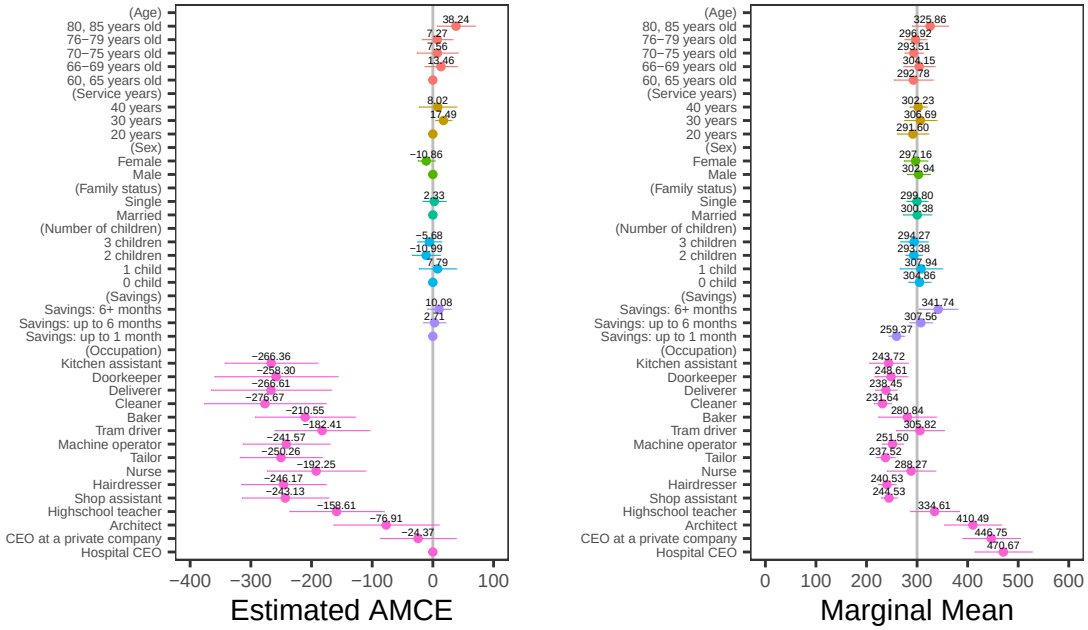
[12]In May 2024 the average monthly public pension was 230 940 HUF (EUR 605) in January 2024 in Hungary according to the Central Statistical Office (CSO, 2024b).

prestige and salary of an occupation, the higher the pension respondents consider fair: the fair pension of a former cleaner is close to 280 thousand HUF (EUR 733) lower than that of a former hospital CEO (see Figure 3.6). The mean of fair pensions is 300 thousand HUF (approximately 765 EUR), which is around 70% of the average and around 80% of the median net earnings and 129% of the actual average public pension in 2024[13]. The marginal means of fair pension by pensioner characteristics vary most across occupations: it ranges from 231 thousand HUF (EUR 605) for a cleaner up to 470 thousand HUF (EUR 1230) for a hospital CEO.

Other pensioner characteristics matter much less when defining a fair pension. Pensioners with higher savings tend to get a higher fair pension, however, when it comes to the effect of savings on the assigned fair pension, we see that respondents are not sensitive to pensioners' savings, as the AMCEs of saving levels are not significantly different from zero. As for service years, plus 10 years of active labor market participation compared to the baseline 20 years increases the fair pension only by 17 thousand HUF and the effect of 40 years of work compared to 20 is not even significant. Compared to the younger pensioners (60 and 65 years old) only the very old (80 and 85 years old) receive a significantly higher fair pension: the increase is 38 thousand HUF (EUR 99) on average. Sex, family status and the number of children are not a decisive factor of public opinion about fair pensions.

[13]In May 2024 average net earnings were 448 700 HUF (EUR 1175), median net earnings were 363 900 HUF (EUR 953) (both include tax benefits) and the average monthly public pension was 230 940 HUF (EUR 605) in January 2024 in Hungary according to the Central Statistical Office (CSO, 2024b).

Figure 3.6: The effects of pensioner attributes on the assigned fair pension (’000 HUF)



Note: Figure shows the average marginal component effects and marginal means of pensioner attributes. The dependent variable is the amount of pension (’000 HUF) the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Point estimates are reported above the dots. The dots on the zero line denote the reference category for each attribute, when estimating AMCEs. The vertical line on the right panel indicates the overall mean of fair pensions. Weights are applied. (The average exchange rate in January 2024: 1 EUR = 382 HUF.)

Comparing the drivers of estimated and fair pensions, we see a similar pattern: estimates and preferences are most sensitive to the former occupation of the pensioner. Respondents consider the earnings-related system fair, although they prefer a somewhat smaller effect of former occupation on pension levels. The AMCE of being a cleaner was the largest for both dependent variables, but the effect was a 300 thousand (EUR 785) estimated malus for cleaners compared to hospital CEOs, compared to the fair difference of 277 thousand HUF (EUR 725). In line with the AMCEs, the range of marginal means for occupations is smaller in the case of fair pensions: 232-471 thousand HUF (EUR 607-1233) compared to the range of 190-478 thousand HUF (EUR 497-1251) in the case of estimated pensions, indicating a preference for more vertical equity in the pension system, mostly by uplifting low-benefit pensioners.

Subgroup analysis

We group respondents by age, education level, and ideological stance to compare the role of pensioner attributes among them when defining a fair pension. We focus on the three main pensioner attributes of interest that are indicative about merit and need: occupation, service

years, and savings. The differences in pension preferences by age and education level of the respondent are suggestive of self-interested considerations, while estimated differences by ideology would point to the role of values. In the subgroup analysis we group the occupations into three categories (low-, middle- and high-skilled) to increase statistical power[14].

Overall, subgroup differences are mostly insignificant. Looking at preferences about the fair amount of monthly pension, we find that the elderly prefer higher pensions in the case of most of the attributes. However, differences across age groups are not significant, which may be partly due to the smaller sample size of elderly respondents. Preferences about fair pensions are similar between respondents with different educational levels. However, participants with higher education assign an approx. 30 thousand HUF (EUR 79) lower fair pension to low-skilled pensioners, pointing to the role of self-interest (see Figure A.13 in the Appendix). We find no clear differences among left-wing and right-wing respondents in the relative importance of pensioner attributes (see Figure A.14 in the Appendix). None of the omnibus F test comparisons point to significant subgroup differences either.

3.6 Discussion and concluding remarks

In this paper we offer a novel approach to evaluate the representation of mass preferences over public pensions. We follow a distance-based approach and compare perceptions, preferences and policy outcomes on a meaningful scale (the monthly pension amount in '000 HUF in Hungary). We design a conjoint survey experiment with pensioner profiles that are representative of their population to compare the distributions of estimated and fair pensions with observed pensions, and to assess the impact of pensioner characteristics on estimated and fair pensions. We quantified the perceived and actual pension policy bias, and assessed whether pensions should be defined based on merit or need according to the public to assess how well mass pension preferences are represented, and whether voters deem the current earnings-related system socially legitimate.

[14]Low-skilled occupations include cleaners, deliverers, doorkeepers, kitchen assistants. Middle-skilled occupations are: baker, tram driver, machine operator, tailor, nurse, hairdresser, shop assistant. High-skilled occupations are: high-school teacher, architect, CEO at a private company, Hospital CEO. The marginal means by subgroups for all occupations are reported in the Appendix, see Figures A.15, A.16

We find that respondents are aware of the monthly amount of public pensions and they are also aware how pensions are defined, at least broadly. They understand well that the current public pension system is earnings-related but assign a low weight to the length of service years, which may be due to the complex procedure of calculating one's pension (involving i.e. valorization). The high level of awareness about pension amounts is in stark contrast to the low level of fiscal awareness that is routinely found in previous studies (Csontos et al., 1998) and to the concerns regarding meaningful policy outcome preferences on a numerical scale (Wlezien, 2017).

Respondents perceive the largest pension policy bias at the median, while the actual relative difference is largest at the lower end of the distribution. Participants tend to somewhat overestimate pensions, except for the median pension, where perception bias is close to zero. This may reflect the more frequent media coverage of the median pension compared to other points of the distribution. Consequently, the perceived and actual pension bias is equal at the median: respondents would prefer a substantial increase around 25% to reach a fair amount of monthly pension in the middle of the distribution. This is the largest perceived bias throughout the distribution, whereas respondents seem content with low and high pensions (perceived pension levels equal fair levels below the 10^{th} and above the 90^{th} percentiles). However, the actual policy bias is largest at the lower end of the distribution: the relative difference is 31-36% at the 5^{th} and 10^{th} percentiles. If participants were aware of the actual amounts of pensions throughout the distribution, they would prefer a higher pension at all points, but especially at the lower tail.

Fair pensions by occupation show that respondents accept the earnings-related system and do not prefer a flat-rate system. However, they would prefer a weaker correlation between earnings and pensions based on the comparison of the average estimated and the average fair pensions for different occupations. Average fair pensions for low-earners tend to be 20-30% higher than average estimated pensions, while the mean estimated and fair pensions are equal for high-earners. The results suggest that respondents think that pensioners with low-skilled former occupations are underpaid, but they do not think that high-earners are overpaid when it comes to public pensions. Participants seem to prefer smaller inequality among pensioners

(less vertical inequity), but not necessarily by redistributing benefits from high-benefit to low-benefit pensioners. Fair pensions are strongly associated with occupational prestige: the more prestigious the occupation, the higher the assigned fair pension on average, suggesting that merit (reciprocity) is a dominant driver of pension preferences.

Similarly, when assessing the drivers of a fair pension, the results of the conjoint experiment show that the former main occupation of the pensioner has the highest relative importance, compared to other pensioner characteristics (such as age, sex, savings, family status and number of children). No indicators of vulnerability, such as low savings and being single, affected responses. Surprisingly, years of service did not impact the decision about a fair pension either, despite indicating merit (in terms of contributions that were paid during these years). Respondents assign a fair pension regardless of the length of the career, and it is the occupation which guides their decision. A potential reason behind the dominance of occupation is that it provides rich implicit information about pensioners: their approximate earnings during active years, their education level and, accordingly, their social class and rank in society (both in terms of material and intellectual resources and their 'value added' to society), which are all indicative about how prestigious the occupation is. We find that more prestigious occupations increase the fair amount of monthly pension, as higher merit is rewarded by a higher pension. Our results indicate that among the deservingness criterias of Oorschot (2000) reciprocity (degree of receiving an earned support) is the dominant one.

The finding that the drivers of estimated and fair pensions are similar suggests that respondents broadly perceive the public pension system as socially legitimate when it comes to the rules of defining the pension for an eligible person (a part of the redistributional design). While the drivers of estimated and fair pensions are in line with each other, their magnitude differs, as the effect of a low-paid occupation is smaller on the fair pension than on the estimated one. This points to a preference for stronger solidarity towards the more vulnerable pensioners than perceived in the current system.

The drivers of pension preferences seem to be similar across participants, we find no significant differences across age groups, respondents with different education levels and different ideological stance. Accordingly, we find no evidence of inter-generational cleavages when

defining pensions and self-interest in general does not seem to be a dominant factor either based on the insignificant role of age and education level of the respondents.

Overall, we see a substantial pension policy bias: respondents prefer higher pensions than the actual amounts throughout the distribution but especially for low-benefit pensioners. The median fair pension is 55 thousand HUF or 27% higher than the observed median. The actual relative pension policy bias is largest at the 10^{th} percentile: 36%, pointing to weak output-legitimacy. Nevertheless, the perceived pension policy bias tends to be lower, except at the median.

We find that the current earnings-related public pension system is socially legitimate. We see a broad consensus among participants that pensioners with a more prestigious and better paid former occupation paying higher contributions during their active years should receive higher public pensions. Accordingly, public opinion suggests that pensions should rather smooth consumption over the life-cycle, in other words, merit and reciprocity matter.

The contradictory conclusions about the quality of representation within one welfare policy domain (public pension policy) based on the two aspects we assess: policy bias (measured by the difference between fair and observed levels of pensions) and alignment of preferred and observed redistribution rules, points to the complex nature of policy responsiveness and to the multidimensionality of social legitimacy. Our findings are in line with the multidimensional attitudes patterns towards the welfare state found in Eastern and Southern Europe: the public supports the goals and role of the government while it is critical of the efficiency and policy outcomes of the welfare state (Roosma et al., 2013).

Our study has three main limitations to keep in mind. First, the estimated and fair pensions were defined on the basis of 15 occupations that are meant to be representative of the pensioner population, still the actual pensions of pensioners with these former occupations may not cover all the percentiles of the observed pension distribution. Thus, some of the perception bias and actual policy bias may be due to the selection of occupations. For example, it may be the case that the observed pensions of former hospital CEOs are between the 96-98^{th} percentiles in the actual pension distribution, whereas the estimates for the same occupation cover a wider range of percentiles from 94^{th} to 100^{th} in the distribution of pension estimates. It may have been

clearer to compare the responses with observed pensions corresponding to each occupation, but such data were not available.

Second, our study collected opinion about the pensions of those who are included in the pension system, so we do not measure public opinion on supporting the most vulnerable segment of the elderly population who are excluded from the system. However, in Hungary 92% of the population above age 65 receives an old-age pension (Eurostat, 2024). Nevertheless, we lack information on the attitudes towards supporting the remaining 8% of the elderly population.

Finally, we measured attitudes towards a public pay-as-you-go pension system in a country where other sources of pensions, such as occupational or private pension, are minimal. Our findings are relevant for continental Europe with Bismarckian pension regimes; nevertheless, a stronger presence of alternative pension sources may influence public opinion about the role of the public PAYG pension scheme.

Our final limitation leads us to the potential extensions of our research. Applying our approach to measure attitudes on pensions in other countries could shed light on how general the preference for a merit-based pension is across different pension schemes and welfare regimes. Comparative studies could also measure the association between institutional settings and the magnitude of policy bias to learn about the drivers of policy responsiveness.

A further extension of our study would be to use conjoint experiments with open-ended questions to elicit public opinion on the distribution and eligibility rules of other welfare policy outcomes, such as unemployment benefit, family benefits, or even income tax rates. Such experiments would allow for across-policy comparisons of responsiveness and would give rich information on the drivers of public opinion on the main questions of welfare regimes about who should get and how much from the public pocket. We believe that similar conjoint experiments could benefit welfare attitude studies by testing theories about social legitimacy and deservingness in various policy fields and contexts.

Chapter 4

Preferences over the progressivity of public pensions

4.1 Introduction

Pension inequality has been increasing over the last two decades in Hungary[1]. The minimum public pension of 28 thousand HUF (EUR 71) is in stark contrast with the highest public pensions that are estimated to be over 2-4 million HUF (EUR 5-10 thousand). The gap is still enormous when comparing these 'record-level' pensions with the median pension (around 205 thousand HUF, EUR 510 in 2024). Despite increasing pension inequality, there has been no pension policy response to increase redistribution among pensioners, no pension reforms with progressive distribution effects. The recent reintroduction of the 13th month extra pension in 2021, without a cap maximizing the extra benefit for pensioners above a certain level of monthly pension, even deepens pension inequality. This is puzzling, based on the general prediction of political economy theory, which asserts that increasing inequality increases demand for redistribution, which would be reflected in more redistributive policies in democracies. We assess preferences over the progressivity of public pensions in Hungary to see whether they could serve as an explanation for the seemingly absent policy response to increasing pension inequality. A better understanding of mass pension preferences may also support formulating

[1]This paper is an unpublished manuscript.

pension austerity measures that are politically feasible.

There are ample studies about the akin puzzle of rising income inequality coupled with stagnant or even falling redistribution - a trend in the United States and advanced democracies in recent decades that received much attention - that focus on understanding redistributive preferences. The individual-level drivers of redistributive attitudes may be categorized into two main groups: self-interest and other-regarding preferences (altruism, fairness considerations). Several studies of income redistribution preferences have shown that economic self-interest is a prominent factor (Alesina & Giuliano, 2011; Iversen & Soskice, 2001; Margalit, 2013), mediated by the social beliefs about the sources of income (effort vs luck and connections) (Alesina & Angeletos, 2005; Mijs, 2021), mobility perceptions (Alesina et al., 2018; Benabou & Ok, 2001), the perceived level of inequality in society (Choi, 2019; Dimick et al., 2016), and relative position in society (Alt & Iversen, 2017; Kuziemko et al., 2014; Lupu & Pontusson, 2011). At the same time, empirical studies routinely find that other-regarding concerns matter also (Ballard-Rosa et al., 2017; Cavaillé & Trump, 2015; Corneo & Grüner, 2002), which are often rooted in ideological orientation and cultural-historical experiences of a country (Alesina & Giuliano, 2011; Blekesaune & Quadagno, 2003; Sears et al., 1980), and the degree of identification with poorer social groups (Dimick et al., 2018). A recent experimental study reinforced that both considerations play a role: self-interest is coupled with a prosocial concern for the poor when forming preferences over several economic and welfare policies (Bechtel & Liesch, 2020).

Less scholarly attention has been paid to public opinion on pension distribution. Most studies on pension preferences are motivated by the search for socially acceptable pension retrenchment options. Public opinion remains central as a driver of welfare policies in the literature on welfare politics (C. Brooks and Manza, 2008; Pierson, 1996), and pension cuts are notoriously unpopular among voters (Bansak, Bechtel, et al., 2021; Boeri et al., 2002; Häusermann et al., 2019; Jensen, 2012; Pierson, 1996). Accordingly, their political feasibility is a major concern due to the pressure posed by the aging population of advanced democracies.

The seminal paper of Pierson, 1996 asserts that the politics of welfare state retrenchment is distinct and policymakers have a strong incentive to lower the visibility of the reforms by mak-

ing it hard for voters to detect their effects and to trace responsibility for them, and accordingly, reforms tend to be incremental modifications of existing policies. In a similar vein, Tepe and Vanhuysse, 2009 argue that radical pension generosity cutbacks may be delayed due to risks of short-term vote losses, and governments may attempt to 'muddle through' by incremental medium-size cuts.

Another retrenchment strategy, already described by Pierson (1996), is to compensate politically crucial groups with concentrated interests for the lost benefits. A recent strand of studies argue that compensating potential opponents of a reform is an effective way to overcome inertia (Bonoli, 2000; Häusermann et al., 2019; Lindvall, 2017; Natali & Rhodes, 2004). As policies are complex and multidimensional, there is scope to bundle reforms strategically, where a careful combination of benefit restrictions and targeted expansions may secure majoritarian support, even for austerity (Bansak, Bechtel, et al., 2021; Boeri et al., 2001), and even for public pension retrenchment (Häusermann et al., 2019).

To design a politically feasible pension reform that successfully targets restrictions and expansions, one must be aware of voters' redistributive attitudes and their drivers. Understanding redistributive preferences regarding the pension system is complex. In a pay-as-you-go (PAYG) system there is redistribution from the active to the inactive population, and there is room for redistribution among the pensioners as well (for example by defining minimum or maximum benefits). Most studies on pension preferences focus on the former, assessing intergenerational conflicts over public spending. Findings are mixed about the correlation between age or ageing and pension policy attitudes: some find a significant relationship between the two (Busemeyer et al., 2009), however others report weak or no relationship (J. Lynch & Myrskylä, 2009; Sørensen, 2013). The correlation or lack of correlation between age and attitudes is often interpreted as the result of or as the lack of self-interested considerations.

Another approach to studying the role of economic self-interest in forming pension preferences is to measure public opinion among voters on particular pension reforms, such as increasing the statutory retirement age, decreasing benefits, and raising contributions by the active population. Boeri et al. (2002) reports that the economic situation matters along with ideology when forming opinion about pension reforms, and they interpret the apparent short-sightedness

as selfishness of current generations. Jaime-Castillo (2013) find that solidarity principles influence preferences over pension reforms as well: conservative respondents are more in favor of increasing contributions and oppose postponing retirement, whereas liberal individuals oppose increasing contributions and prefer extending the retirement age.

Overall, most of the literature on pension attitudes focuses on the potential intergenerational cleavages due to the varying economic interests of the active and inactvie population (at least on the short run), based on the assumption that self-interest is the dominant driver of pension preferences. However, there is a gap in the literature when it comes to preferences over redistribution and inequality among pensioners, which may provide explanations for the lack of pension policy response to rising pension inequality in Hungary. Furthermore, preferences over the progressivity of public pensions are relevant as they could be a crucial input when forming socially acceptable and politically feasible pension reforms, especially when targeting austerity among pensioners.

We focus on two main research questions. First, what is the public opinion over pension progressivity? Second, what drives mass preferences over the public pension distribution: economic self-interest or other-regarding preferences?

In this study we measure preferences over the progressivity of public pensions by assessing public opinion on two different pension policy outcomes in Hungary. The term progressivity is mostly used in the taxation literature. Taxation or a tax may be progressive or regressive based on the incidence of taxation distributed across the population. A tax is progressive if it increases as the taxable amount increases, in other words if richer individuals pay a higher proportion of their income/wealth than poorer individuals. Defining progressivity of a public pension system is complex as it should incorporate both the contributions paid and the benefits received over the life-course. We rely on the definition of Biggs et al. (2009), who defines pension progressivity as the degree to which benefits are higher relative to lifetime payroll contributions for lower contributors than for higher contributors. It follows that a more progressive pension system redistributes more from higher to lower contributors.

We run a conjoint survey experiment embedded in a public opinion survey, where respondents see two hypothetical pensioners with several characteristics that are indicative of the

amount of public pension they receive and their financial vulnerability, such as former occupation, number of service years, and level of savings. Respondents decide about two pension policy outcomes based on the pensioner profiles: they have to define the fair pension replacement rate (fair pension as a % of the last salary before retirement) for each hypothetical pensioner, and they have to apply a cut to the 13th month extra pension to one of the hypothetical pensioners.

Pension replacement rates are usually not visible, nor politicized as they depend on complicated pension calculations that are difficult to follow for the majority of voters. As a consequence, a change in pension replacement rates or any differentiation of them across pensioners would be relatively easy to obfuscate and ideal for implementing incremental policy changes. We consider the public opinion on the fair pension replacement rates across pensioners as a direct measure of preferences over public pension progressivity. Differences compared to a uniform distribution of fair replacement rates across pensioners point to a preference for a progressive or degressive pension system: higher replacement rates for formerly low-paid pensioners with lower payroll contributions than for high-earners or vice versa.

The other pension policy we study is salient and highly politicized: the provision or cut of the 13th month extra pension. Such a visible pension cut would most probably be widely unpopular, as it activates loss aversion strongly. The forced choice of choosing one of the hypothetical pensioners to bear the pension cut based on their financial vulnerability is also indicative of the preferences for progressive public pensions: higher probabilities of a pension cut among better-off pensioners than among more vulnerable pensioners reflects pro-poor considerations and a preference for pension policies with a progressive effect on the distribution of public pensions.

Building on the literature on redistributive preferences, we study the role of two main potential drivers of pension preferences: economic self-interest and other-regarding preferences. We group respondents by their socio-economic status and compare the responses of financially secure and vulnerable voters and assess how they distribute pensions and pension cuts among hypothetical pensioners with different socio-economic backgrounds.

We find that respondents prefer a progressive public pension system, based on both pol-

icy outcomes we studied: replacement rates are higher for pensioners with low-paid former occupations and pension cuts are targeted towards better-off pensioners. Pro-poor considerations are strong among the financially vulnerable and secure respondents as well, and they are even stronger among the latter group, suggesting that other-regarding considerations are an important driver of pension preferences. Our results refute the argument that the lack of public support for more progressive pension policies may explain why Hungarian pension policy has not responded to rising pension inequality. They rather suggest that mass preferences over pension progressivity are not represented by the government based on the current public pension system.

We contribute to the existing literature on pension preferences by studying an often overlooked aspect of public pensions, its distribution among pensioners in a survey experimental setting. Our approach has several advantages. First, instead of measuring pension preferences on attitudinal scales, we put respondents in a 'policy-maker position' in the sense that they decide on two straight-forward pension policy outcomes for hypothetical pensioners (though without explicit budget constraints). Instead of overwhelming respondents with complex pension reform packages with unforeseeable consequences on the distribution of pensions, they define fair replacement rates and target a pension cut, which are more directly related to preferences over pension progressivity. Second, we are able to identify causal relationships based on the conjoint survey experiment and compare the role of several pensioner characteristics that are indicative of vulnerability at the same time. Assessing preferences over the progressivity of public pensions is important to better understand the lack of progressive pension policy response to rising pension inequality.

The paper is structured as follows. First, we introduce the Hungarian pension system, focusing on replacement rates and the 13^{th} month extra pension. Then we describe the design of the conjoint experiment, the data, and the variables. In the results section, we assess the drivers of defining a fair replacement rate and of targeting a pension cut, and compare the drivers among financially secure and vulnerable respondents. In closing, we discuss the results and their implications.

4.2 The Hungarian context

We study public pension preferences in Hungary. The Hungarian pension system follows a social insurance model, or the Bismarckian model, which is the dominant pension regime in continental Europe. It is a mandatory earnings-related public pension system with a defined benefit, which is calculated based on the average of wages and the number of service years. Voluntary private pension provision is minimal. The system is financed on a pay-as-you-go basis, where pensions are financed by the working population. Pension benefits are calculated based on the average of wages and the number of service years. The pensionable age is 65 for both women and men with a minimum service period of 20 years. Nevertheless, women may retire before age 65 in case their service period reaches 40 years. Pensions are indexed to inflation.

The at-risk-of-poverty (AROP) rate of those aged 65+ indicates the relative income position of pensioners, as it shows the share of the 65+ population with an income that is lower than the 60% of the national median equivalised income after social transfers (Eurostat, 2025). The AROP rate has been increasing steadily since 2015 jumping from 4.6% (among the lowest values in the EU) to 15% (close to the EU average of 17%) in 2020, as pensions did not follow the wage growth in the period, given that they are indexed to inflation. Nevertheless, the severe material deprivation (an absolute measure of poverty as opposed to AROP) declined during the same period (from 15.7% in 2015 to 9% in 2020) (Commission et al., 2024).

Inequality among pensioners has increased in the last decade. An indicative measure of inequality among pensioners, the relative standard deviation of pensions increased from 0.41 in 2013 to 0.48 in 2022 among men, and for women, it rose from 0.32 to 0.39 in the same period (A. Simonovits, 2023). The very low minimum pension (HUF 28 500 which equals EUR 71) that has been unchanged since 2008, the elimination of the cap of earnings that feed into the pension calculation (dropping the contribution cap) in 2013, and dropping indexation to wages and indexing pensions only to inflation since 2012 coupled with the intense increase of earnings in 2016-2020 all deepened inequity among pensioners over the last decade. The considerable

longevity gap in Hungary[2], which refers to the difference in life expectancies of strata with high and low socioeconomic status, also reduces the redistribution towards low-benefit pensioners in the pension system (A. Simonovits & Lackó, 2023). When it comes to inequality between genders, in Hungary women above age 65+ received a pension that was on average 10% lower than that of men in 2019 (most recent measure), which is much lower than the EU-average gender pension gap of 29% (Eurostat, 2021).

Population aging is a common challenge in the European Union (EU) and Hungary is no exception. The fiscal pressure on the pay-as-you-go pension system is increasing: public pension spending in terms of the GDP is expected to increase by 4.3% by 2070, while contributions remain constant as a share of the GDP (OECD, 2024). As a consequence, there is a need to adjust the pension system. The government committed itself to reform the pension system by March 2025 to remain sustainable and to provide adequate pensions, as part of Hungary's recovery and resilience plan submitted to the European Commission. However, it refuses to take into account the suggestions of experts summarized in a recent OECD (The Organization for Economic Cooperation and Development) report prepared at its own request (OECD, 2024). The main policy recommendations fall into two categories: tightening eligibility conditions and adjusting benefit levels. In the latter category, there is a suggestion to adjust the 13th month's pension benefit by adding an indexed ceiling (OECD, 2024). A. Simonovits and Lackó (2023) also suggests as part of a rational pension reform package to level the 13th month pension at the average to reduce pension inequalities or even set a common level below the average or make the extra benefit degressive to also make the system more sustainable.

4.2.1 Pension replacement rates

The pension replacement rate is an indicator about the relative living standards of pensioners. There are two common definitions of the pension replacement rate. One compares the net pension with the net earnings of the pensioner before retirement (pre-retirement earnings may refer to the amount coming from final active year or to the average earnings throughout the career). The other compares the average net pension with the average net earnings in the

[2]See (Bıró et al., 2021) for an estimate of life expectancy inequality at age 45 during 1991 - 2016

country.

The net pension replacement rate according to the OECD estimate, which compares the net pension to the net pre-retirement earnings[3] for full-time workers is 78.8% for men and 73.7% for women, which is around 11 percentage points higher than the average of the 22 EU countries in the OECD (OECD, 2024).

The rates are much lower when comparing the average net pension to the average net earnings, as reported by the Central Statistical Office (CSO, 2025). They were around 65% during 2007-2017 and then dropped to between 51% and 59% during 2017-2023. However, as Oblath and Simonovits (2024) show, the official earnings statistics overestimate earnings during 2010-2020, leading to an underestimated ratio of average pension benefits to the average net wages. Calculating with revised wages (relying on the national accounts), the ratio is higher, around 70% instead of the official 55% in 2021 (Oblath & Simonovits, 2024).

Overall, the pension replacement rates suggest that the Hungarian pension system performs well in maintaining living standards after retirement. Nevertheless, the low level of minimum pension that has not changed from HUF 28 500 (EUR 71) since 2008 and the limited non-contributory benefits raise concerns about the most vulnerable segment of the elderly population, whose share may increase due to the increasingly diverse employment forms on the labor market (OECD, 2024).

The values of pension replacement rates are not salient, in part because of their complex definitions. Political communication mostly abandons the measure and opts for simpler outcomes and messages about public pensions. Accordingly, most voters are unaware of the exact values of pension replacement rates, and public opinion on their fair levels is also unknown.

4.2.2 The 13th month pension

Several EU countries provide additional pension, which is often called 13th month pension or a bonus often transferred before holidays, such as Christmas or Easter. The rules of the

[3]The net pension replacement rate is defined as the individual net pension divided by the net pre-retirement earnings. The replacement rate is calculated for men and women workers separately (to reflect the shorter average career length of women driven by the Women-40 scheme) with an average wage, with an entry into the labor market in 2022 at the age of 22 and a constant working career until claiming a pension without penalty (OECD, 2024).

additional pension vary across countries. The extra pension may be means-tested and targeted to pensioners with a low benefit or it may be capped (i.e. Belgium, Bulgaria, Croatia, Poland, Slovakia, Slovenia), whereas some countries transfer a 13th or even 14th month pension to all without means-testing (i.e. Austria, Hungary, Italy, Portugal) (Commission et al., 2024).

In Hungary, all pensioners receive an extra 13th month pension benefit. It was first introduced gradually between 2003 and 2006 as part of a larger fiscal expansion. However, as a result of the 2008 global financial and economic recession, the Hungarian government introduced an austerity package, which included the abolishment of the 13th month pension in 2009. The withdrawal cut yearly pension payments by about 10% and, consequently, it was widely unpopular. Another government reintroduced the 13th month pension in 2021 as part of the expansionary response to the COVID-19 crisis. Instead of the originally planned gradual reintroduction spanning over four years, the government reinstated the full amount of the 13th month pension already in February 2022, before the national elections of 2022 (Commission et al., 2024). All pensioners are eligible to receive the full amount of their monthly pension as 13th month pension, regardless of their pension level and the economic performance of the country.

The 13th month pension is salient as it is a flagship policy of the ruling party, FIDESZ, which they promise to 'defend' on billboards all over the country. Voters strongly support the policy. A survey run by a government-friendly organization asked a rather provocative question about whether the 13th month benefit should persist in its current form, despite the suggestion of EU experts to reduce it, and found that 74% of the adult respondents say yes (Intezet, 2025).

4.3 Design

We design a conjoint survey experiment to measure the relative importance of both pensioner and respondent characteristics in targeting a pension cut and in setting a fair replacement rate for each hypothetical pensioner to infer about preferences over pension progressivity. In the case of the pension cut, respondents have to choose between two pensioners and cut back the 13th month benefit of one of the pensioners. The general public opinion on the impact of

particular pensioner attributes shows the role of pro-poor considerations in pension attitudes. We measure preference differences among respondents with low and high socioeconomic backgrounds to assess the role of self-interest and other-regarding considerations in forming pension preferences.

The conjoint survey experiment is a useful tool for studying multidimensional choices, as it includes several treatment components, so it can estimate causal relationships between several independent variables and the outcome variable at the same time (Hainmueller et al., 2014). As a result, the effects are comparable.

In our conjoint experiment respondents see two pensioners with a few attributes summarized in a table, which include age, sex, family status, the number of children, the main former occupation, years of service, and savings (see Table 4.1). We aim to include all the relevant characteristics of a pensioner that are used for calculating their public pension and that are indicative of their financial vulnerability without overwhelming respondents with too much information. As pensioners receive an earnings-related benefit, their former occupation and the number of service years are the basis of their public pension determination. Better-paid occupations and more years on the labor market translate into a higher amount of public pension. The level of savings pensioners have is a direct measure of their financial vulnerability. The other characteristics, such as age, sex, family status and number of children may only be indicative of vulnerability. Older pensioners tend to have a somewhat lower at-risk-of-poverty rate than younger pensioners: 4.6% among the population aged 76+ compared to the 6.1% among those aged 66+ (OECD, 2024). However, the older elderly may be more frail with higher health-related costs to cover. As for the sex, there is a gender pension gap of 10% between men and women, meaning that the public pension of women is on average 10% lower in Hungary, which is lower than the EU-average of 29% (Eurostat, 2021). Accordingly, the share of older women living in relative income poverty is higher than that of men: 7.1% compared to 4.4%, which is partly due to a higher share of older women living alone as they are more likely to be widowed (OECD, 2024). Number of children may be associated with higher risk of poverty due to the higher costs of raising and supporting children throughout the life-course, however, adult children may also provide support for their elderly parents.

Table 4.1: Pensioner attributes and their levels

Pensioner attribute	Levels
Age	60 / 65 / 66 / 67 / 68 / 69 / 70 / 71 / 72 / 73 / 74 / 75 / 76 / 77 / 78 / 79 / 80 / 85
Sex	Male / Female
Years of service	20 / 30 / 40 years
Main occupation during active years	Hospital CEO / CEO at a private company / Architect / High-school teacher / Shop assistant / Hairdresser / Nurse / Tailor / Machine operator / Tram driver / Baker / Cleaner / Deliverer / Doorkeeper / Kitchen assistant
Savings	Enough for 1 month / Enough for half a year / Enough for more than half a year
Family status	Single / Married
Number of children	0 / 1 / 2 / 3

Note: The table shows the levels of each pensioner attribute that could appear in the conjoint experiment. The final column outlines the theoretical reasons for including each attribute.

Respondents receive three conjoint tables with hypothetical pensioners in a row. The same questions appear after each table: one about the fair replacement rate and the other about targeting a pension cut (See Figure A.17 for the formulation of the questions). The order of the pensioner attributes in the table varies randomly across respondents, but it is constant within respondents to avoid confusion. In order to present realistic pensioner profiles, we restricted a few combinations of characteristics: CEOs cannot have low savings (enough only for a month), and a former cleaner, deliverer, doorkeeper and kitchen assistant can not have savings enough for more than half a year. Moreover, a former machine operator cannot be female and a former nurse cannot be male.

Respondents had to answer a few questions related to the pensioner profiles. First, respondents entered a replacement ratio (the monthly pension as a percentage of the last salary) for each pensioner that they considered fair. Second, respondents faced a forced choice task: they had to choose which pensioner must face a cutback in case the government must cut the extra 13^{th} month pension of at least one of the pensioners. Our main aim with both questions is to give insight into voters' preferences over the progressivity of public pensions. In the case of fair replacement rates, a preference for progressivity would mean that better paid hypothetical pensioners receive lower replacement rates than lower paid ones. As for targeting the tax cut, cutting the extra pension of a better-off hypothetical pensioner rather than that of a lower-paid

one would be in line with preferences for progressivity.

Estimating the so-called AMCE (average marginal component effect) of each feature on the fair replacement rate and on the probability of cutting the pension of a pensioner gives us a measure of average relative sensitivity of respondents to specific pensioner attributes. AMCEs are estimated via OLS regressions with dummy variables for the levels of pensioner attributes (Hainmueller et al., 2014). They do not represent the majority preference but rather the 'average' preference, as they combine the direction and strength of opinions by averaging preferences both across the distributions of possible pensioners and respondents (Bansak et al., 2023). It follows that the AMCE of a pensioner attribute is sensitive to the other attributes, so we aimed to include all the relevant pensioner characteristics without overwhelming the respondent. We also report the marginal mean of each level of the attributes.

We are also interested in differences among respondents in their pension preferences that are indicative of the dominance of self-interested and other-regarding preferences. We look at several indicators of socio-economic status: education level, subjective financial security and perceived social status of the respondents. Assigning higher fair replacement rates to pensioners of similar socio-economic status and cutting the 13^{th} month pension of 'out-group' pensioners points to self-interested considerations.

Moreover, we assess whether the role of pensioner attributes differ among respondent subgroups. We estimate the difference between marginal means of the particular levels of pensioner attributes among respondents with low- and high socio-economic status, together with an omnibus F-test, where we compare the model without interaction with the model with interactions in a nested model comparison, as advised by T. J. Leeper et al. (2020). We use the software called cregg by T. Leeper (2018) available in the Comprehensive R Archive Network and coefficient plots are based on examples provided by T. J. Leeper et al. (2020). In all of our models, we use the weights provided by the survey agency and we estimate robust standard errors (SEs) clustered at the respondent level.

4.4 Data and measurement

We collected the data in Hungary, during Spring 2024. A survey agency, called TÁRKI Social Research Institute drew a probability sample that is representative of the adult population of Hungary. Interviewers made face-to-face interviews, where they asked respondents in person and recorded the answers on a computer or tablet.

We have a sample of 1072 respondents, who saw 3 conjoint tables, each with 2 hypothetical pensioners, so we have 6372 hypothetical pensioners in total. Due to non-response, we have 5668 hypothetical pensioners with an assigned fair replacement rate and 4168 hypothetical pensioners who either suffered a pension cut or were spared from the cut[4]. The lower response rate in the case of targeting a pension cut may reflect the opposition to any pension retrenchment. Targeting a cut may have been so uncomfortable or unacceptable for some of the respondents that they refused to answer that particular question. Non-response is not associated with basic socio-demographic characteristics: the composition of the respondents answering about replacement rates and pension cuts is similar (see Table A.10 in the Appendix).

A question about public pension spending preceded the conjoint experiment. One of two pension reforms appeared randomly: both highlighted the fiscal trade-off between the level of pension spending and taxes. Respondents had to indicate whether they would support increasing pension spending and as a consequence increasing taxes. Alternatively, they had to decide about supporting decreasing taxes and consequently pensions. Both questions were binary with response options yes and no.

There are two outcomes of the conjoint experiment, based on the questions following each conjoint table with two hypothetical pensioners. The first variable, 'fair replacement rate' is about the share (in %) of the last net earnings of the pensioner the respondent considers fair for each pensioner. The fair replacement rate has a value between 0 and 100%. The second outcome, 'pension cutback' is a forced choice variable, which refers to cutting the extra 13^{th} month pension of one of the pensioners. This is a binary outcome variable, which is 1 if the

[4]With our sample sizes our models are powered (power equals 0.8 and significance level is set to 0.05) to detect an effect size of 0.005 for the fair replacement rate estimation and 0.18 for the pension-cut probability. The much smaller detectable effect size is due to the binary outcome variable of receiving a pension-cut.

hypothetical pensioner suffers the cut and 0 if the pensioner was spared from the cut.

The variables we use when analyzing respondent subgroup differences in preferences include education level, subjective financial security and perceived social status. Respondents were grouped based on their education level into primary vs. secondary and higher educated respondents. Subjective financial security is also a binary variable based on a five-point scale. Respondents who feel they live without financial issues and who feel financially secure with conscious financial planning are categorized as living in financial security. Those, who just make ends meet with their monthly income, who struggle financially from month to month, and who live in poverty are part of the subgroup who lack financial security. The perceived social status may be low, middle or high, based on a social ladder scale from 1 (lowest) to 10 (highest), where we considered self-ratings 1-4 to be low, 5-6 to be middle, and 7-10 to be high perceived social status.

4.5 Results

We introduce our results about preference for progressivity of public pensions in two main parts. First, we describe public opinion about fair pension replacement rates. Second, we assess the drivers of targeting a pension cut.

4.5.1 Public opinion on the fair pension replacement rate

This section begins with unpacking overall public preferences over fair pension replacement rates. First, we assess the impact of pensioner attributes on the fair replacement rate, to see whether indicators of vulnerability matter and there is a preference for progressivity. Second, we disentangle overall preferences and compare pension preferences of respondents with different socio-economic backgrounds to see how self-interest and other-regarding preferences weigh in when forming pension attitudes about progressivity. We look at respondent subgroups defined by their subjective financial security (living in financial security vs. lack of financial security), and their perceived social status (low and middle vs. high social status).

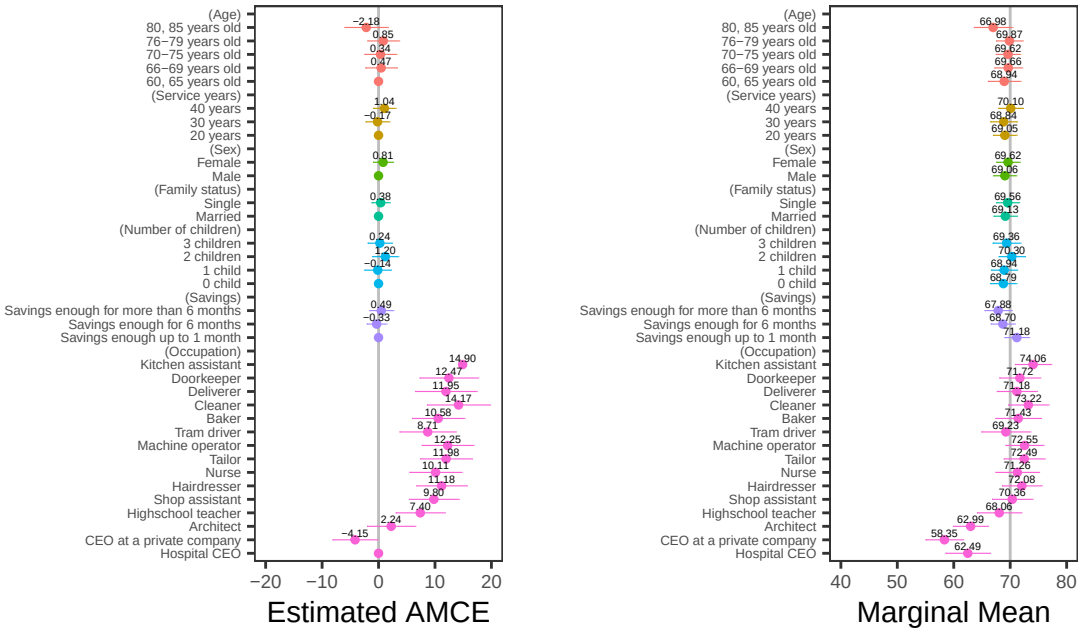
Pensioner attributes' impact on fair replacement rates

We find that the assigned replacement rate is mostly sensitive to the former main occupation of the pensioner. Figure 4.1 shows the average marginal component effect (AMCE) of each pensioner characteristic on the fair replacement rate on the left side and the marginal mean of the attributes on the right side. Both estimates are reported along their cluster-robust 95% confidence intervals. In case of AMCEs, the reference category is depicted by a dot on the vertical zero line. In the case of main former occupations, the reference category is hospital CEO. The AMCEs of occupations indicate that the lower paid the job was, the higher the pension replacement rate that is preferred by the public. Respondents think that the fair pension replacement rate of a former kitchen assistant who was at the lower end of the job ladder should be approximately 15 percentage points higher than the replacement rate of a former hospital CEO, who had a well-paid and prestigious former job. Notably, a CEO at a private company receives a 4 percentage points lower replacement rate than a hospital CEO (which tends to be a public sector job in Hungary). The other attributes, such as age of the pensioner, service years, sex, family status, number of children and savings have muted effects.

Most marginal means line up around 70 percent, the replacement rate that respondents consider fair on average. The marginal mean describes the favored pension replacement rate for pensioner profiles with a particular feature level, all else equal (see the right panel in Figure 4.1). The patterns of marginal means and AMCEs are similar, most of the variation is across occupations: the marginal mean of replacement rates is lowest for a former CEO at a private company (54%) and highest for a former kitchen assistant (74%), reflecting that respondents tend to favor a higher replacement rate for low-paid jobs. The replacement rate is not so sensitive to other characteristics: respondents assign a slightly higher replacement rate for pensioners with less savings, and with more service years, but differences are not significant.

Overall, the substantial differences between hypothetical pensioners with different former occupations in the fair pension replacement rates point to a strong preference for progressivity in public pension assignment. AMCEs of occupations not significantly different from zero would indicate a preference for proportionality, and lower replacement rates for lower-paid occupations would show a preference for a regressive public pension system.

Figure 4.1: The effects of pensioner attributes on the assigned fair pension replacement rate (%)



Note: Figure shows the average marginal component effects and marginal means of pensioner attributes. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Point estimates are reported above the dots. The dots on the zero line denote the reference category for each attribute, when estimating AMCEs. The vertical line on the right panel indicated the overall mean of fair replacement rates. Weights are applied.

Public opinion on replacement rates among respondent subgroups

Turning to the role of respondent characteristics, we find that socio-economic background is associated with fair replacement rates. The mean fair replacement rate is 77% among those respondents who live in financial security, while it is only 61% among those who lack financial security (see Table 4.2). We see a similar relationship between social status and fair replacement rates: the higher the subjective social status, the higher the mean fair replacement rate. The gap between low- and high-status respondents is 9 percentage points on average (68% vs. 77%). However, the level of education of the respondent does not matter: the mean fair replacement rate is around 70% for all three levels of education (primary, secondary and higher education). Similarly, we see no difference between left- and right-wing participants in their preferred pension replacement rates (see Tables A.11 and A.12 in the Appendix).

Table 4.2: Fair pension replacement rate among respondent subgroups(%)

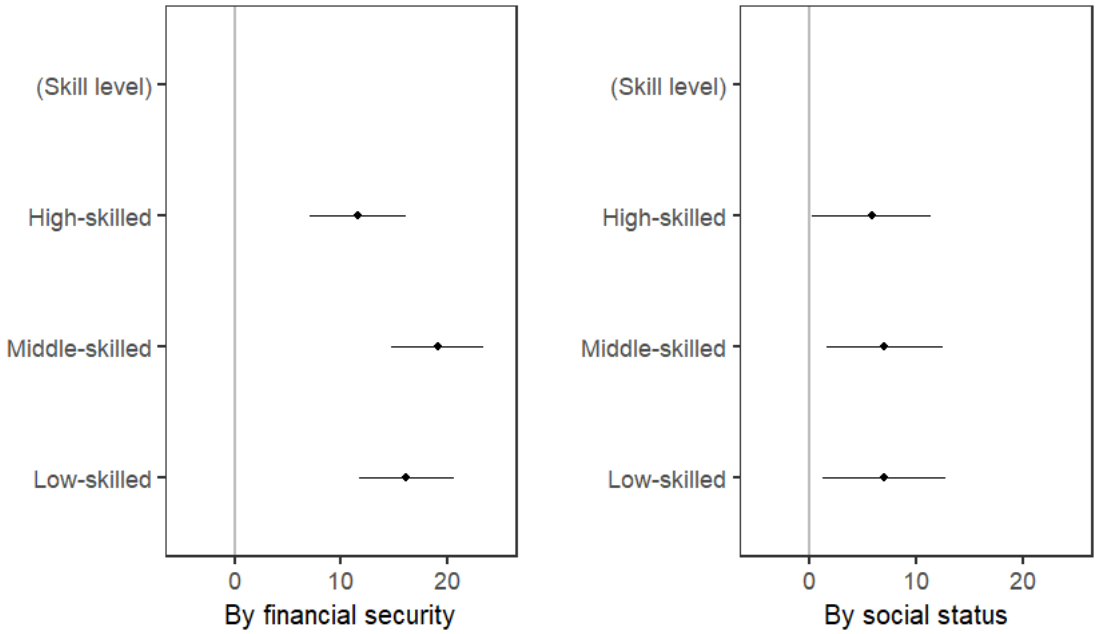
Fair replacement rate (%)	**Financial security**		**Social status**		
	Lack of...	Living in...	Low	Middle	High
Mean	61	77	68	70	77
SD	31	24	29	28	25
Median	70	80	75	75	80
N	2529	3095	2214	2639	692

Note: The table shows the mean and median fair replacement rate by the subjective financial security and subjective social status of the respondents.

When comparing the role of pensioner attributes across respondents with lower- and higher socio-economic status, we find that better-off participants reward all pensioner attributes with a higher fair replacement rate. In Figure 4.2 we report the the pensioner attribute that is the strongest indicator of lifetime payroll contributions and financial vulnerability: former main occupation. We group the occupations into three categories: low-, middle- and high-skilled, to increase statistical power[5]. The differences in marginal means are especially large, between 10 and 20 percentage points, when grouping respondents by their subjective financial security (the omnibus F-tests also suggest subgroup differences both by subjective financial security and social status). Although higher replacement rates benefit high-earners more in absolute terms, the patterns of marginal means do not reflect self-interested responses. A straightforward indication of self-interested considerations would be to see that those living in financial security prefer higher replacement rates only for alike pensioners with high-skilled former occupations. However, we find the opposite: higher socio-economic status participants enter higher replacement rates for all skill levels, but especially for middle- and low-skilled workers, suggesting other-regarding preferences, though confidence intervals overlap.

[5]Low-skilled occupations include cleaners, deliverers, doorkeepers, kitchen assistants. Middle-skilled occupations are: baker, tram driver, machine operator, tailor, nurse, hairdresser, shop assistant. High-skilled occupations are high school teacher, architect, CEO at a private company, Hospital CEO. The marginal means by all occupations are reported in the Appendix, see Figures A.25, A.26.

Figure 4.2: Differences in conditional marginal means of fair replacement rates, by financial security and social status



Note: Figure shows the differences in marginal means for the different skill levels of the former occupation of the pensioner, by the subjective financial security and social status of the respondents. Social status is categorized into lower status and high status (low and middle: 1-6, and high: 7-10 on a scale from 1-10). The dependent variable is fair replacement rate. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Comparing the AMCEs of pensioner attributes among respondents grouped by their subjective financial security yields similar results in terms of demand for a more progressive pension system. Both groups tend to favor a higher replacement rate for pensioners with lower-paid former jobs, but the magnitude is larger among those living in financial security. Most AMCEs of middle-skilled occupations are around 11-12 percentage points among the better-off respondents compared to 7-8 percentage points among the more vulnerable respondents (see Figure A.18 in the Appendix). The point estimates suggest that financially secure respondents prefer a more progressive system: their preferences are more elastic when setting fair replacement rates, favoring the formerly low-earner pensioners.

As opposed to the differences among financially secure and insecure participants, respondent subgroups based on education, age, and ideology attribute similar importance to the pensioner characteristics when defining fair replacement rates. We checked for differences among respondents based on their educational level (as another indicator of socio-economic status), their age (to check for intergenerational cleavages), and ideological stance (to assess its association with pension preferences as it is often related to redistributive preferences). Respondents

with secondary and higher education prefer slightly higher replacement rates than participants with primary education, although most of the differences are insignificant (see Figure A.23 in the Appendix). More educated respondents enter an almost 5 percentage points higher replacement rate for high-skilled pensioners and pensioners with more savings, which is indicative of self-interested considerations, however the point estimates are barely significant. When it comes to age and ideological stance, differences between younger and older, and between left- and right-wing participants are around zero (see Figure A.24 in the Appendix).

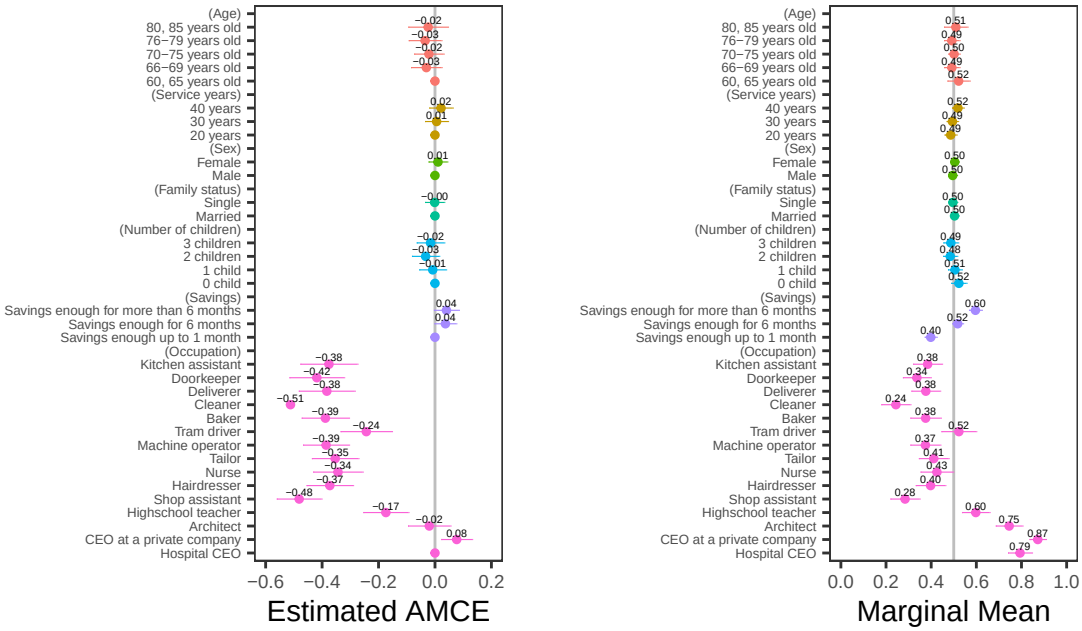
4.5.2 Targeting a pension cut

The public opinion survey also included a question about the trade-off between public pensions and tax levels. The public seems to be resistant to pension spending reforms when the trade-off between pension expenditure and taxes is emphasized, based on a question preceding the conjoint experiment in the survey. Most of the respondents (83%) would not support an increase in pensions that comes with tax increases. A pension cut is even more unpopular (86% of the respondents oppose it), even if it comes with tax cuts, based on a question preceding the conjoint experiment. A strong resistance to any pension cut is also reflected in the lower response rate in the case of the forced choice task of cutting the extra pension benefit (13^{th} month pension) of one of the pensioners.

When assessing the effects of pensioner characteristics on a forced pension cut-back, we find that occupations are driving the targeting of a pension cut - similarly as in the case of fair replacement rates (see Figure 4.3). Being a former cleaner decreases the probability of a pension cut by 51 percentage points compared to being a former hospital CEO. Overall, the lower paid the pensioner was on the labor market, the less likely it was that respondents cut the extra pension of the pensioner. Former CEOs at a private company were the most likely to face a pension cut: they were chosen to face a cut close to 9 times out of 10. The AMCE of being a former CEO at a private company increased the probability of a pension cut by 8 percentage points compared to being a former CEO at a hospital, suggesting that within the same earnings-category participants 'punish' the private sector employee instead of the public sector employee. Among the other pensioner attributes, only pensioners' savings are close to

be significantly different from zero: both savings enough for 6 months and savings enough for more than 6 months increased the probability of cutting the extra pension by 4 percentage points on average, compared to a pensioner with savings enough only up to 1 month. In general, the results indicate that respondents decide on a pension cut primarily based on the financial vulnerability of pensioners.

Figure 4.3: The effects of pensioner attributes on a forced pension cut-back



Note: Figure shows the average marginal component effects and marginal means of pensioner attributes. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Point estimates are reported above the dots. The dots on the zero line denote the reference category for each attribute. Weights are applied.

Respondent subgroup differences in targeting a pension cut

To assess the role of economic self-interest and other-regarding preferences as drivers of the decision about the pension cut, we compare the responses of respondents with lower and higher socio-economic status, measured by subjective financial security and social class.

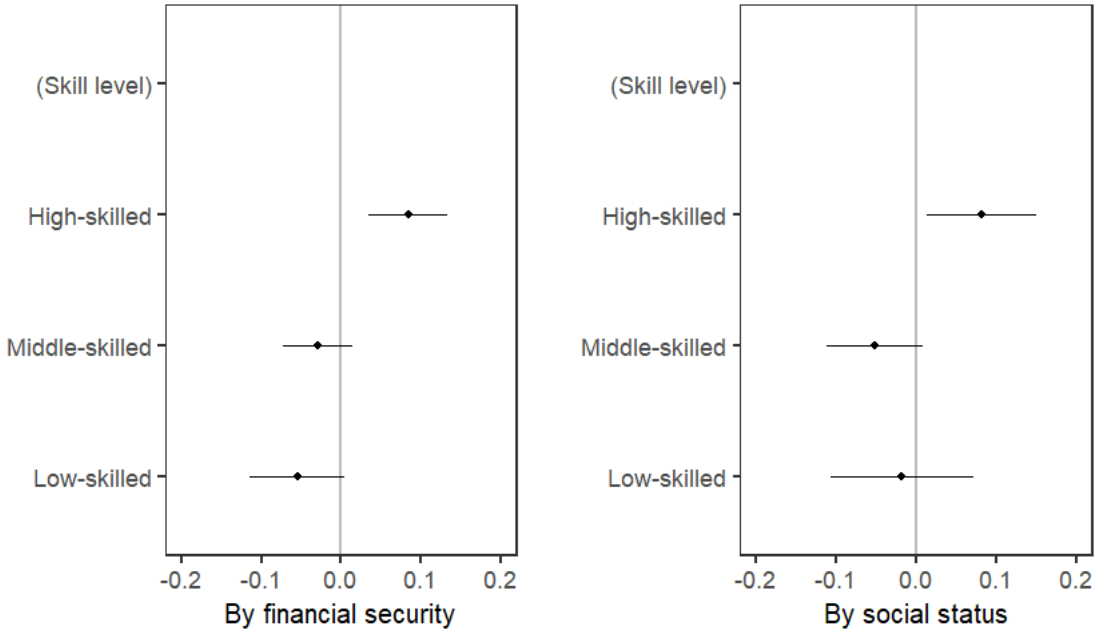
We find evidence for other-regarding concerns among the better-off participants. Respondents living in financial security are more likely to cut the extra 13th month pension of pensioners with supposedly higher pensions due to their high-skilled former occupation, compared to respondents who lack financial security. The other indicators of pensioner vulnerability weigh in similarly among financially secure and insecure individuals.

We see a similar pattern, when comparing low and middle status respondents with high

social status participants: the only significant difference is in the role of high-skilled former occupation, which increases the probability of a pension cut more among high social status respondents. In general, we see that those living in financial security and in higher social classes have a stronger preference for protecting financially vulnerable pensioners from pension retrenchment[6].

Pensioner attributes weigh in similarly among the other respondent subgroups, grouped by their age, their education level, and their ideological stance (see Figures A.30, A.31, A.32, A.33 in the Appendix).

Figure 4.4: Differences in conditional marginal means of pension cut-back probability, by financial security and social status



Note: Figure shows the differences in marginal means for the different skill levels of the former occupation of the pensioner, by the subjective financial security and social status of the respondents. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

We also estimate the AMCEs of pensioner attributes on targeting a pension cut among those living in financial security and those who lack financial security, separately (see Figure A.27 in the Appendix). We see that both groups favor more vulnerable pensioners and cut the extra 13th pensions of pensioners with better-paid former occupations, but the magnitude of the

[6]While we see a significant difference both by financial security and social status in the marginal means of high-skilled occupations, the omnibus F-test suggest only a marginally significant (p leq0.06) overall subgroup difference by financial security and no significant overall difference in targeting the pension cut by social status

effects is larger among financially secure respondents: most point estimates for middle- and low-skilled occupations are between -0.6 and -0.4 among them, whereas the same estimates are between -0.4 and -0.2 among those who lack financial security. The point estimates suggest that financially secure respondents are more likely to spare vulnerable pensioners from a pension cut, reflecting other-regarding preferences.

4.6 Discussion and concluding remarks

In this paper we assessed an overlooked aspect of attitudes towards pensions, which may explain the lack of redistributive policies amid increasing pension inequality: preferences over the progressivity of public pensions, measured in a conjoint survey experiment in Hungary (with an earnings-related PAYG pension system). We show public opinion on two pension policy outcomes that are different in terms of their visibility and the emotions they evoke. The fair replacement rate (fair pension as a % of the last salary of the pensioner) is rather neutral and invisible, thus voters are unaware of the actual rates. Cutting the 13^{th} month extra pension is visible and a highly politicized austerity option in Hungary. In the survey experiment we assess the impact of pensioner characteristics that are indicative of financial vulnerability. Then we study whether the effects of pensioner attributes vary across respondents with different socio-economic backgrounds to see how self-interest and other-regarding concerns weigh in when forming pension preferences.

Overall, the results of the conjoint experiment indicate that preferences over replacement rates and pension cuts are most sensitive to the former occupation of the pensioner, compared to other pensioner characteristics (such as age, sex, years of service, savings, family status and number of children). The former occupation provides rich implicit information on the pensioner, such as educational level, social status and social capital, or prestige and influence in society, and it is also the most obvious indicator of former earnings and the amount of public pension in an earnings-related PAYG system. Accordingly, the former occupation provides information about the financial vulnerability of the pensioner.

As a consequence, results suggest that respondents take into consideration the financial

vulnerability of a pensioner when defining fair replacement rates. Lower-paid occupations are assigned a higher replacement rate (keeping in mind that a similar replacement rate translates into higher pensions for former high-earners), which points to preference for a progressive public pension system, where benefits are higher relative to lifetime payroll contributions for lower contributors than for higher contributors. Similarly, pro-poor considerations are strong when cutting the extra 13th month pension of a pensioner. The pension cut is targeted towards pensioners with a well-paid former occupation (and as a consequence a higher monthly pension), who are likely to be less prone to financial difficulties as a result of the pension cut (as opposed to pensioners with a low-paid former occupation).

There seems to be a broad consensus that replacement rates should be higher for formerly low-paid pensioners and that when it comes to retrenchment, pension cuts should rather target pensioners with a well-paid former occupation to protect more vulnerable pensioners from poverty. Participants grouped by their age, ideological stance and education level form their pension preferences similarly: we do not find significant differences among them in the relative importance of pensioner characteristics. While education level, an objective proxy for socio-economic status does not matter, subjective measures of socio-economic status, such as subjective financial security and perceived social status is associated with preferences over pension progressivity. We see that financially secure respondents prefer a more progressive pension system, which goes against the standard economic self-interest based prediction. Their preferred fair replacement rates are more elastic to changes in the perceived pension level of the pensioner. They are also more likely to spare vulnerable pensioners from a pension cut, reflecting other-regarding preferences.

Overall, we find evidence for progressive preferences over pensions among the Hungarian public. Thus, the lack of redistributive policy response to increasing pension inequality is not explained by mass preferences. Instead, our results point to a poor representation of public opinion, which leads us to further questions about the factors that may limit policy responsiveness in the case of public pension policy. Policy response may be absent or slow, due to the effort needed and risks involved in reforming a complex system with long-lasting and potentially unintended consequences. Instead, the government may opt for visible one-off measures

that could ease dissatisfaction and electoral sanctions, such as providing pensioners a capped VAT exemption when buying basic food.

Nevertheless, as tensions will increase over rising pension inequality and over keeping the system both sustainable and adequate, the government must reform or adjust the public pension system eventually. Our results suggest that strategic pension reforms that bundle cuts with potential expansions to boost public support should focus on compensating the more vulnerable, low-benefit pensioners. Retrenchment options that only affect pensioners at the upper end of the pension distribution, i.e. adding a ceiling on the 13th month pension may not be widely unpopular among voters, despite the general aversion towards pension austerity.

Our study has two main limitations to keep in mind. First, there are some caveats about how general the findings about progressive public pension preferences could be. We studied pension preferences in Hungary, where the pension system is dominated by the public PAYG pension scheme and in a context, and where pension inequalities have been increasing during the last decade. The living standards of low-benefit pensioners may condition public preferences over the progressivity of pensions (see the findings of Dimick et al. (2016) about the sensitivity of preferences of the rich to macro-inequality). An extension of our study could be to study preferences over the progressivity of public pensions in other institutional and socio-economic contexts, where there are more pillars of the pension system and where the living standards of low-benefit pensioners are higher.

Second, we interpret the dominant role of former occupation in defining a fair replacement rate and in targeting a pension cut as evidence that voters take into account the level of lifetime payroll contributions and financial vulnerability of pensioners. However, the former occupation of a pensioner may evoke several emotions and thoughts that influence pension preferences, potentially unrelated to contributions or financial vulnerability. A further extension of our study could be to assess the role of former occupations through their impact on the perceived deservingness of pensioners (i.e. based on the deservingness criteria developed by Oorschot (2000)) or on the perceived competence and warmth of pensioners (the dimensions of the stereotype content model (Fiske, 2018)). Such studies could further elaborate on the mechanisms behind pension preference formation.

Chapter 5

Conclusion

This dissertation was motivated by the challenges of welfare states in ageing societies. The first part highlighted the importance of the welfare state in mitigating labor market related risks during working age, such as job loss and financial difficulties, given their negative association with health at older ages. The second part showed evidence that social legitimacy of public pensions is multidimensional: Hungarian voters accept the earnings-related pension system - albeit would prefer a weaker association between earnings and pensions -, nevertheless the substantial pension policy bias points to a deficit in output-legitimacy.

The first paper (Chapter 2) analyzed whether the hardships around the transition contributed to the health gap between post-socialist CEE and western Europe, using a retrospective dataset of SHARE (Survey of Health, Ageing and Retirement in Europe). We found evidence that stressful periods, financial difficulties and job loss around the period of transition are mostly associated with worse health at older ages in all groups of CEE countries, even after netting out the effect of childhood health and demographic factors. Our results suggest that psychosocial stress as an adult around the transition cumulated over the life course into weaker health at older ages, in line with the life-course theory of health.

We also found that similar hardships unrelated to the shocks of the transition period (both in CEE and the West) are also negatively related to later health, indicating that the long-term health implications of hardships due to the transition are not specific. Nevertheless, 19% of the individuals reported major difficulties around the transition in CEE, compared to 15% in the

West during the same period. Thus, not the transition-specific nature of the difficulties, but the higher fraction of individuals experiencing them around the transition contributed to the current health disadvantage in CEE.

Overall, the enduring health consequences of psychosocial stress and financial hardship during adulthood highlight the vital role of the welfare state in offering protection against labor market risks. These stakes are particularly high during periods of economic recession, when large fractions of the population are exposed to heightened insecurity. To prevent the accumulation of health disadvantages over the life course, policy responses to economic recessions should take into account their potential health impacts.

In the second and third paper we turned to mass preferences over public pensions. First, in Chapter 3 we offered a novel approach to evaluate the representation of mass preferences over public pensions by combining a conjoint survey experiment with observed data. The conjoint survey experiment included representative pensioner profiles with several attributes (including occupation, years of service and savings), and we elicited the estimated or the fair pension levels for them. We applied a distance-based measure of policy bias introduced by G. Simonovits et al. (2019) and compared the distribution of estimated, fair, and actual monthly amount of pensions (in '000 HUF) in Hungary, to estimate both the perceived and actual pension policy biases.

We found that voters are aware of the magnitude but overestimate pensions, especially at the tails of the distribution. Consequently, the perceived pension policy bias tends to be lower than the actual bias. The middle of the distribution is an exception. The perceived and actual policy biases are almost the same at the median: voters prefer a substantial increase of approximately 25% (55 thousand HUF corresponding to 144 EUR). Nevertheless, the actual pension policy biases in relative terms are largest at the lower end of the pension distribution (the relative differences are 31-36% at the 5^{th} and 10^{th} percentiles).

If voters were aware of the actual amounts of pensions throughout the distribution, they would prefer a higher pension at all points, but especially at the lower tail. This is in line with our finding that the perceived pension policy bias tends to be largest for pensioners with low-earner occupations, suggesting that voters think that pensioners with low-skilled former occupations are underpaid. Overall, the substantial pension policy biases point to weak output-

legitimacy (Roosma, 2016).

Furthermore, we measured the relative importance of the attributes of hypothetical pensioners when defining a fair level of pension to assess the congruence between preferences over deservingness and prevailing pension laws, indicative of the social legitimacy of the redistributional design of the pension policy. We found that the former main occupation of the pensioner has the highest relative importance, compared to other characteristics of the pensioner. There is a broad consensus among voters that pensioners with a more prestigious and better paid former occupation, paying higher contributions during their active years, should receive higher public pensions. This indicates that among the deservingness criterias of Oorschot (2000) reciprocity (degree of receiving an earned support) is dominant. Consequently, our results suggest that the current earnings-related public pension system is socially legitimate.

Overall, we found a multidimensional attitude pattern toward the public pension system in Hungary. Focusing on two dimensions of social legitimacy (Roosma, 2016), we saw that while the earnings-related nature of the redistributional design of public pensions is socially legitimate, the substantial pension policy bias suggests a poor representation of mass preferences and a deficit in output-legitimacy. Our findings are in line with the finding that the public tends to support the goals and role of the government while it is critical of the efficiency and policy outcomes of the welfare state in in Eastern and Southern Europe (Roosma et al., 2013).

The third paper (Chapter 4) assessed an overlooked aspect of attitudes towards pensions: preferences over the progressivity of public pensions, as a potential explanation for the lack of redistributive policies amid increasing pension inequality in Hungary. We measured public opinion in a conjoint survey experiment with hypothetical pensioners on two pension policy outcomes that are different in terms of their salience: fair pension replacement rate (fair pension as a % of the last salary of the pensioner) and a cut in the 13^{th} month extra pension.

We found that lower-paid occupations are assigned a higher replacement rate, which points to preference for a progressive public pension system, where benefits are higher relative to lifetime payroll contributions for lower contributors than for higher contributors. Similarly, pro-poor considerations are strong when cutting the extra 13^{th} month pension: the pension cut is targeted towards pensioners with a well-paid former occupation.

There is a broad consensus regardless of the age, ideological stance, and educational level of voters that replacement rates should be higher for formerly low-paid pensioners and that pension cuts should target pensioners with a well-paid former occupation. Nevertheless, we found that financially secure respondents prefer a more progressive pension system, as their preferred fair replacement rates are more elastic to changes in the perceived former earnings level of pensioners and they are also more likely to spare vulnerable pensioners from a pension cut. Our results are against the standard economic self-interest based predictions, and reflect other-regarding preferences.

Overall, we found evidence for progressive preferences over pensions among the Hungarian public. Thus, the lack of redistributive policy response to increasing pension inequality is not explained by mass preferences. Public opinion suggests that pension retrenchment should bundle cuts with compensations for the more vulnerable, low-benefit pensioners to boost public support for pension restrictions, in line with the findings of Häusermann et al. (2019).

Finally, we have a few broader conclusions and thoughts regarding studying pension and welfare attitudes. First, we found that voters were able to provide meaningful policy outcome preferences on a numerical scale, despite the common concern in the literature (Wlezien, 2017). We see potential in broadening the range of outcome variables with numerical measures of preferred policy outcomes in the policy representation and welfare attitudes literature, to have a better understanding of the drivers of representation and preference formation.

Second, there is room for a better understanding how the quality of representation of mass preferences over welfare policies and social legitimacy of a welfare policy are related. Another aspect to consider is whether perceived policy outcomes or actual policy outcomes matter more for the social legitimacy of a welfare policy. Similarly, what is more motivating for policy makers: actual or perceived representation of mass preferences?

Third, we did not cover the role of policy feedback, whereas existing or previous welfare policies may also influence welfare attitudes (Pierson, 1993). Policy feedback and responsiveness may co-exist in the policy loop, influencing each other (Meuleman et al., 2020). In our case, the support for an earnings-based public pension system may partially stem from a status-quo bias.

In closing, we reinforced that the reasons for welfare states to provide some protection against labor market risks are manifold, including the relevance of social security in supporting citizens in preserving their health over the life-course. We also found public support for redistribution by the welfare state, based on mass preferences over public pensions. The generational welfare contract (Birnbaum et al., 2017) seems strong insofar as we found no evidence for age-based social cleavages. Instead of economic self-interest, preferences over public pensions are rather shaped by values (regarding deservingness, i.e. the role of merit and need), such as other-regarding considerations. Other-regarding preferences may provide some room for politically feasible pension reforms in ageing societies.

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Appendix A

Appendices

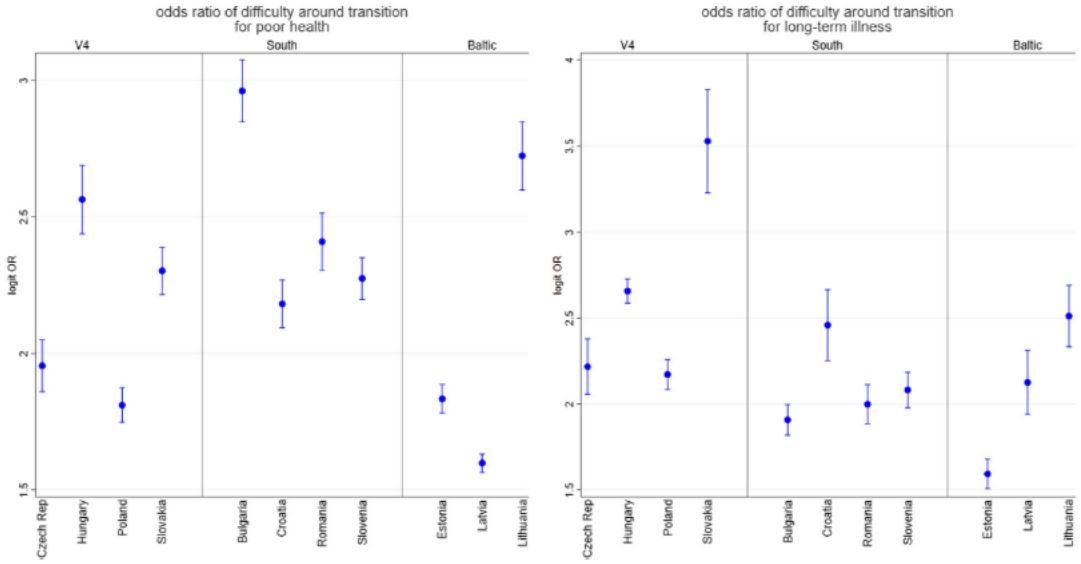
A.1 Appendix - Chapter 1

Table A.1: First group of models – Further health measures regressed on difficulties occurring between 1987 and 1993 in CEE country groups

Variable	Health limits work	Hypertension	Heart problem	Diabetes	Ulcer	Cancer	Lung disease	Obese	Weak grip strength
Stress x V4	1.623***	1.337***	1.346**	1.896***	1.222	1.483	1.699**	1.501***	1.192
	[1.276–2.065]	[1.124–1.591]	[1.055–1.718]	[1.284–2.797]	[0.771–1.938]	[0.920–2.389]	[1.011–2.855]	[1.173–1.919]	[0.878–1.617]
Stress x South	1.916***	1.357***	0.859	0.965	1.965***	1.490**	0.852	1.035	1.101
	[1.812–2.025]	[1.185–1.555]	[0.518–1.422]	[0.792–1.177]	[1.181–3.269]	[1.067–2.080]	[0.486–1.493]	[0.846–1.265]	[0.975–1.243]
Stress x Baltic	1.647***	1.312***	1.195***	1.371	1.500***	1.458	1.307***	1.423***	0.861*
	[1.520–1.784]	[1.177–1.463]	[1.070–1.334]	[0.822–2.287]	[1.304–1.726]	[0.886–2.401]	[1.081–1.582]	[1.198–1.691]	[0.733–1.012]
Obs	14,200	17,393	17,393	17,393	17,393	17,393	17,393	17,087	15,859
Fin. difficulties x V4	1.030	1.269**	1.134**	2.059***	1.038	0.779**	2.714**	1.364***	0.852***
	[0.850–1.247]	[1.002–1.608]	[1.001–1.284]	[1.349–3.142]	[0.704–1.532]	[0.643–0.943]	[1.069–6.895]	[1.262–1.474]	[0.764–0.951]
Fin. difficulties x South	1.423	1.431***	0.802	1.115	2.712***	2.197*	1.350	0.866	1.548***
	[0.768–2.639]	[1.286–1.592]	[0.525–1.225]	[0.840–1.479]	[2.332–3.154]	[0.953–5.067]	[0.865–2.108]	[0.678–1.107]	[1.352–1.772]
Fin. difficulties x Baltic	1.370***	1.231***	1.058	1.340	1.321**	0.990	1.143	1.006	1.238**
	[1.301–1.443]	[1.099–1.380]	[0.923–1.213]	[0.937–1.914]	[1.064–1.639]	[0.878–1.115]	[0.797–1.637]	[0.791–1.279]	[1.040–1.472]
Obs	16,631	20,438	20,438	20,438	20,438	20,438	20,438	20,088	18,680
Job ends x V4	1.856***	1.001	1.471***	1.117	1.173	1.540***	1.154**	0.876	1.309**
	[1.582–2.178]	[0.872–1.148]	[1.410–1.535]	[0.899–1.388]	[0.965–1.427]	[1.264–1.875]	[1.012–1.316]	[0.695–1.105]	[1.023–1.675]
Job ends x South	1.497***	0.954	0.777**	1.092	1.505***	1.518	1.668***	0.816**	1.001
	[1.158–1.935]	[0.898–1.013]	[0.626–0.964]	[0.897–1.328]	[1.306–1.733]	[0.847–2.720]	[1.567–1.776]	[0.694–0.959]	[0.881–1.137]
Job ends x Baltic	1.930***	1.111	1.048	1.001	1.533***	1.173**	1.021	1.138	1.122
	[1.795–2.074]	[0.915–1.350]	[0.945–1.162]	[0.820–1.222]	[1.147–2.049]	[1.029–1.337]	[0.843–1.237]	[0.966–1.340]	[0.955–1.318]
Obs	16,598	20,458	20,458	20,458	20,458	20,458	20,458	20,112	18,721

Note: We control for individual characteristics and country effects. Logit odds ratios are reported. 95% CI displayed in brackets. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Figure A.1: First group of models – Health measures regressed on difficulties occurring between 1987 and 1993 in CEE country groups, with county specific OR.



Note: Logit odds ratios are displayed with 95% CI. The figure displays the OR of reporting any of the analyzed three shocks (stress, financial difficulties, job loss with a gap afterwards)

Table A.2: Second group of models – Further health measures regressed on difficulties occurring between 1987 and 1993 in CEE and West

Variable	Health limits work	Hypertension	Heart problem	Diabetes	Ulcer	Cancer	Lung disease	Obese	Weak grip strength
CEE x stress	1.496***	1.283***	1.388**	1.547***	1.237	1.994***	1.280	1.135	1.101
	[1.256–1.783]	[1.123–1.466]	[1.053–1.830]	[1.207–1.982]	[0.822–1.862]	[1.406–2.828]	[0.893–1.833]	[0.814–1.582]	[0.955–1.270]
West x stress	1.629***	1.079	1.075	0.904	1.248*	1.023	1.064	1.022	0.892
	[1.389–1.912]	[0.956–1.219]	[0.892–1.294]	[0.663–1.233]	[0.990–1.573]	[0.898–1.165]	[0.912–1.242]	[0.671–1.556]	[0.760–1.046]
Obs	27,523	35,197	35,197	35,197	35,197	35,197	35,197	34,306	32,167
CEE x fin. difficulties	1.163	1.291***	1.411	1.459*	1.871**	1.042	2.096***	1.104	0.937
	[0.921–1.470]	[1.141–1.461]	[0.831–2.394]	[0.947–2.247]	[1.074–3.260]	[0.705–1.541]	[1.259–3.490]	[0.882–1.383]	[0.668–1.313]
West x fin. difficulties	1.845***	1.193***	1.024	1.397**	1.485	1.102	1.718***	1.686***	1.220***
	[1.423–2.393]	[1.047–1.360]	[0.673–1.557]	[1.063–1.835]	[0.877–2.515]	[0.625–1.943]	[1.218–2.424]	[1.349–2.107]	[1.068–1.395]
Obs	34,358	43,442	43,442	43,442	43,442	43,442	43,442	42,435	39,954
CEE x job ends	1.531***	1.129	1.255**	1.270***	1.260**	1.152	1.479***	0.911	1.092
	[1.245–1.884]	[0.968–1.316]	[1.053–1.494]	[1.091–1.479]	[1.040–1.526]	[0.789–1.684]	[1.183–1.848]	[0.795–1.043]	[0.849–1.403]
West x job ends	1.302***	0.847	1.042	1.089	1.853***	1.122	1.213*	0.969	1.043
	[1.105–1.535]	[0.634–1.132]	[0.844–1.286]	[0.813–1.460]	[1.352–2.541]	[0.923–1.364]	[0.988–1.489]	[0.823–1.143]	[0.952–1.143]
Obs	34,140	43,246	43,246	43,246	43,246	43,246	43,246	42,273	39,757

Note: We control for individual characteristics and country effects in CEE and West. Logit odds ratios are reported. 95% CI displayed in brackets. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

Table A.3: Third group of models – Further health measures regressed on difficulties occurring between 1987 and 1993 versus 1984–1986 and 1994–1996 in CEE

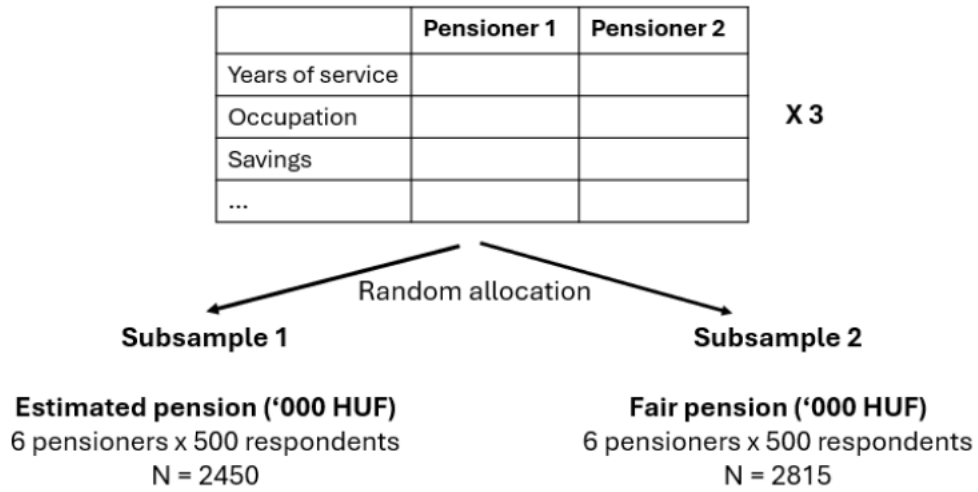
Variable	Health limits work	Hypertension	Heart problem	Diabetes	Ulcer	Cancer	Lung disease	Obese	Weak grip strength
Transition x stress	2.084***	0.993	1.349	1.278**	1.696***	1.246	1.356***	1.146	1.173**
	[1.622–2.677]	[0.600–1.643]	[0.849–2.145]	[1.042–1.566]	[1.335–2.153]	[0.436–3.567]	[1.128–1.630]	[0.925–1.420]	[1.017–1.352]
Before/after trans. x stress	1.565***	1.236***	1.243*	1.408**	1.279	1.733***	1.440**	1.129	1.037
	[1.297–1.889]	[1.060–1.442]	[0.987–1.565]	[1.040–1.907]	[0.860–1.902]	[1.400–2.146]	[1.030–2.014]	[0.793–1.607]	[0.938–1.147]
Obs	15,645	19,108	19,108	19,108	19,108	19,108	19,108	18,786	17,507
Transition x fin. diff.	2.090***	1.070	1.195	1.484***	1.930***	1.496	1.542*	0.963	0.927
	[1.539–2.837]	[0.763–1.500]	[0.776–1.840]	[1.190–1.851]	[1.382–2.695]	[0.745–3.004]	[0.968–2.458]	[0.644–1.439]	[0.667–1.288]
Before/after trans. x fin. diff.	1.226	1.211**	1.307	1.362	1.987**	1.007	2.311***	1.087	1.223**
	[0.954–1.576]	[1.014–1.448]	[0.826–2.067]	[0.811–2.287]	[1.051–3.758]	[0.674–1.505]	[1.352–3.950]	[0.856–1.381]	[1.046–1.430]
Obs	17,770	21,768	21,768	21,768	21,768	21,768	21,768	21,403	19,951
Transition x job ends	1.797***	0.930	1.036	1.393*	1.596***	0.949	1.207	1.053	1.068
	[1.586–2.037]	[0.665–1.300]	[0.609–1.761]	[0.994–1.952]	[1.214–2.097]	[0.657–1.371]	[0.676–2.155]	[0.888–1.250]	[0.821–1.389]
Before/after trans. x job ends	1.578***	1.001	1.156	1.125*	1.356***	1.094	1.517***	0.891*	1.042
	[1.186–2.100]	[0.942–1.063]	[0.876–1.525]	[0.998–1.269]	[1.112–1.654]	[0.689–1.737]	[1.113–2.069]	[0.786–1.010]	[0.955–1.137]
Obs	17,684	21,702	21,702	21,702	21,702	21,702	21,702	21,346	19,913

Note: We control for individual characteristics and country effects. Logit odds ratios are reported. 95% CI displayed in brackets. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$

A.2 Appendix - Chapter 2

A.2.1 Research design and questionnaire

Figure A.2: Structure of the conjoint experiment



Note: The figure shows the structure of the conjoint experiment. Respondents saw three conjoint table in a row and they were randomly allocated into subsample 1 or 2. The dependent variable is the estimated pension (in '000 HUF) in subsample 1 and fair pension (in '000 HUF) in subsample 2.

Figure A.3: Questions of the conjoint experiment

1A. What amount of monthly pension would you consider fair for the 1st and 2nd pensioner? Please enter an amount in HUF.

---------- (thousand HUF) and ---------- (thousand HUF)

1B. What do you think, what is the actual amount of monthly public pension the 1st and 2nd pensioners receive nowadays? Please enter an amount in HUF.

---------- (thousand HUF) and ---------- (thousand HUF)

Note: Respondents were randomly allocated to answer question 1A or 1B.

Figure A.4: Technical details of the conjoint experiment

	Sex	Age	Service years	Family status	Number of children	Savings	Occupations
Respondent characteristics:							
Sex	0.02	-0.07	-0.111*	0.016	-0.077	-0.021	-0.082
	(0.709)	(0.414)	(0.050)	(0.760)	(0.216)	(0.703)	(0.452)
Age	-0.039	-0.022	0.013	-0.005	0.024	-0.021	-0.007
	(0.379)	(0.765)	(0.788)	(0.919)	(0.648)	(0.651)	(0.941)
Education level	-0.02	0.025	-0.011	-0.059	-0.056	0.002	-0.033
	(0.612)	(0.695)	(0.783)	(0.136)	(0.218)	(0.959)	(0.681)
Social status	-0.018	0.015	-0.054	-0.028	-0.019	0.035	0.02
	(0.665)	(0.829)	(0.222)	(0.509)	(0.696)	(0.432)	(0.817)
Intercept	0.067	2.155***	0.795***	0.051	1.178***	0.680***	2.750***
	(0.436)	(0.001)	(0.001)	(0.553)	(0.001)	(0.001)	(0.001)

Notes. Balance tests to verify the random allocation of pensioner attributes across respondents. Linear models regressing pensioner attributes on respondent characteristics. Number of observations: 6204.

Technical details of the conjoint experiment:

Sampling: TARKI applied random selection sampling, and the sample is weighted by considering the common frequencies of four demographic factors – gender, age group, type of settlement and educational level.

Post-stratification weights: Mean: 1.062, minimum:0.34, maximum: 4.24

Period of data collection: from 27 April to 12 May, 2024.

Media salience: The media salience of pensions and pension policy was rather low during the Spring of 2024.

Pre-registration: The conjoint survey experiment was pre-registered at OSF (Open Science Framework), see https://osf.io/q9s4u/overview?view_only=50735c37a8444bc78bf28d4660fcea02

Note: Balance tests and technical details of the conjoint experiment.

Table A.4: Descriptive statistics about the subsamples

Share of...	**Subsample 1**	**Subsample 2**
Female	61%	59%
Aged 60+	36%	38%
Pensioner	29%	31%
Secondary and higher educated	60%	57%
Right-wing	40%	43%

Note: Descriptive statistics about the two subsamples in the conjoint experiment. Subsample S1 received questions about estimated pensions and subsample S2 received questions about fair pensions for the hypothetical pensioners they saw. Allocation of respondents to the subsamples were random.

Table A.5: Hypothetical pensioners in the coinjoint experiment - an example

	Pensioner 1	**Pensioner 2**
Age	65	75
Years of service	30 years	40 years
Sex	Male	Male
Savings	Enough for half a year	Enough for more than half a year
Family status	Single	Married
Number of children	0	2
Main occupation during active years	Tram driver	Architect

Note: An example of a conjoint table with hypothetical pensioners that respondents received.

Table A.6: Occupations and Associated Characteristics

Occupation	**ISCO-08**	**EGP class**	**Income level**	**Sector**	**Education level**
Hospital CEO	1342	elite	high	public	high
CEO of a private company	1120	elite	high	private	high
Architect	2161	intellectual	high	private	high
High School teacher	2330	intellectual	middle	public	high
Shop assistant	5223	white-collar	middle	private	secondary
Hairdresser	5141	white-collar	middle	private	secondary
Nurse	3221	white-collar	middle	public	secondary
Tailor	7531	blue-collar	middle	private	secondary
Machine operator	8151–8189	blue-collar	middle	private	secondary
Tram driver	8331	blue-collar	middle	public	secondary
Baker	7512	blue-collar	middle	private	secondary
Cleaner	9129	unskilled	low	private	primary
Deliverer	9621	unskilled	low	private	primary
Doorkeeper	5414	unskilled	low	private	primary
Kitchen assistant	9412	unskilled	low	private	primary

Note: The table shows the occupations that appeared randomly in the conjoint experiment and their corresponding ISCO-08 code (International Standard Classification of Occupations), EGP class (based on the Erikson-Goldthorpe–Portocarero classes or Goldthorpe class scheme, as in Kolosi and Pósch, 2014: elite, intellectual, white-collar, blue-collar, unskilled), income level (usual income level: low, middle, high), sector (the usual sector of the occupation in Hungary: public or private), and education level (the usual education level corresponding to the occupation: primary, secondary or high).

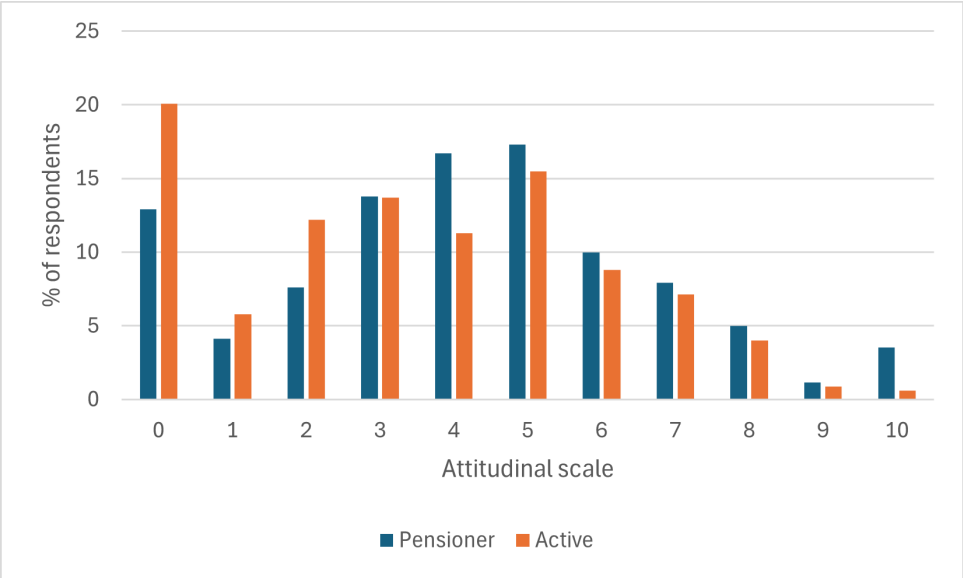
Table A.7: Distribution of occupations in Hungary and in the conjoint experiment

EGP Class	**Share in 2002 in Hungary**	**Number of Occupations in Conjoint Experiment**	**Share of Occupations in Conjoint (%)**
Elite	10.0%	2	13.3%
Intellectual	14.4%	2	13.3%
White-collar worker	14.4%	3	20.0%
Blue-collar worker	30.9%	4	26.7%
Unskilled	30.2%	4	26.7%

Note: The table shows the share of EGP classes (based on the Erikson-Goldthorpe–Portocarero classes or Goldthorpe class scheme, as in Kolosi and Pósch, 2014: elite, intellectual, white-collar, blue-collar, unskilled) in 2002 in Hungary (as measured by Kolosi and Pósch, 2014), and the number and share of occupations that appeared randomly in the conjoint experiment in each social class.

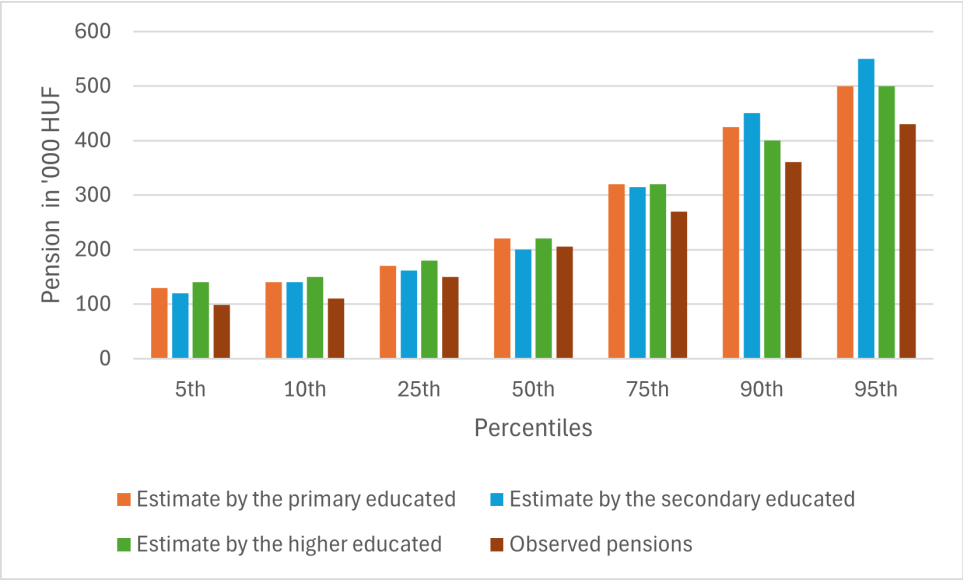
A.2.2 Descriptive results

Figure A.5: Pension is enough to maintain living standards during active years (0 - not at all, 10 - fully)



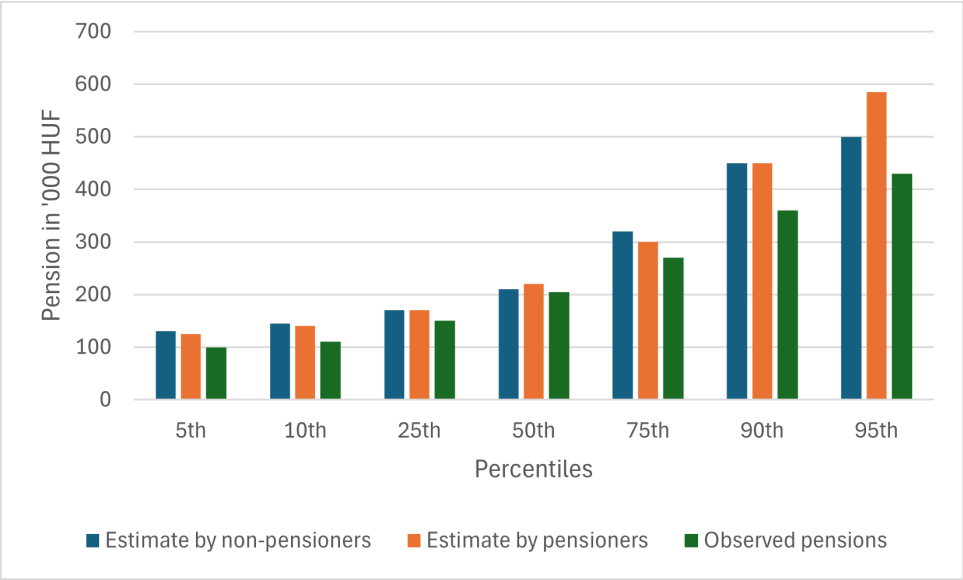
Note: The figure shows the responses of pensioner and active respondents, who were asked about the degree to which their pension is or will be enough to maintain the living standards they enjoyed or enjoy during the active years on a scale from 0 to 10, where 0 means that the pension cannot maintain living standards during active years at all and 10 means that the pension can fully maintain living standards.

Figure A.6: Estimated pensions by education level of the respondent and observed pensions ('000 HUF) at several percentiles of the distribution



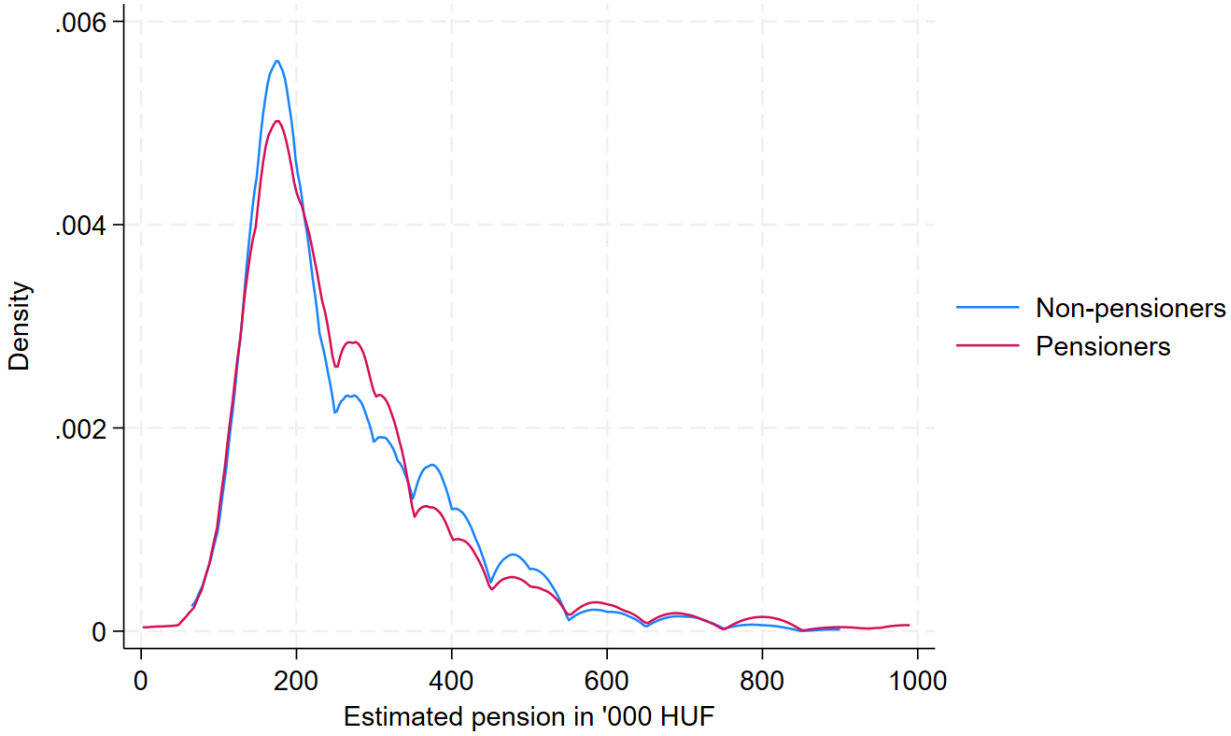
Note: The figure shows the estimated amounts of monthly public pension ('000 HUF) at the 5th, 10th, 25th, 50th, 75th, 90th and 95th percentile of the pension distribution, separately among primary, secondary and higher educated respondents. The final column depicts the observed pension levels. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: observed pensions come from the Central Statistical Office, January 2024 (CSO, 2024b), estimated pensions come from the conjoint experiment, own data collection.

Figure A.7: Estimated pensions among non-pensioners and pensioners and observed pensions (’000 HUF) at several percentiles of the distribution



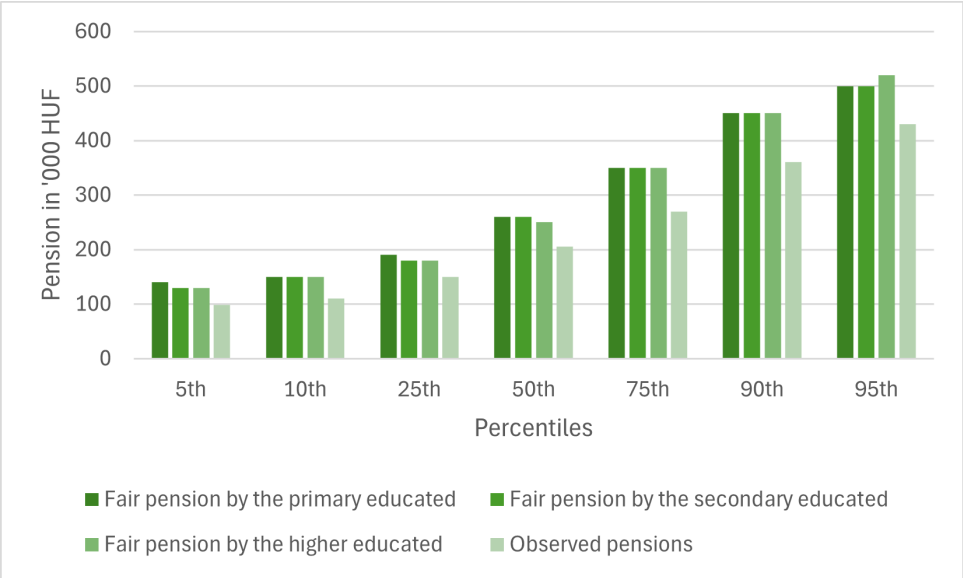
Note: The figure shows the estimated amounts of monthly public pension (’000 HUF) at the 5th, 10th, 25th, 50th, 75th, 90th and 95th percentile of the pension distribution, separately among non-pensioner and pensioner respondents. The final column depicts the observed pension levels. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: observed pensions come from the Central Statistical Office, January 2024 (CSO, 2024b), estimated pensions come from the conjoint experiment, own data collection.

Figure A.8: Distribution of estimated pensions (’000 HUF) among non-pensioners and pensioners



Note: The figure shows the Kernel density estimates for the estimates of monthly public pension (’000 HUF) according to non-pensioners and pensioners. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: conjoint experiment, own data collection.

Figure A.9: Fair pensions by education level of the respondent and observed pensions (’000 HUF) at several percentiles of the distribution



Note: The figure shows the fair amounts of monthly public pension (’000 HUF) at the 5th, 10th, 25th, 50th, 75th, 90th and 95th percentile of the pension distribution, separately among primary, secondary and higher educated respondents. The final column depicts the observed pension levels. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: observed pensions come from the Central Statistical Office, January 2024 (CSO, 2024b), fair pensions come from the conjoint experiment, own data collection.

Figure A.10: Fair pensions among non-pensioners and pensioners and observed pensions (’000 HUF) at several percentiles of the distribution

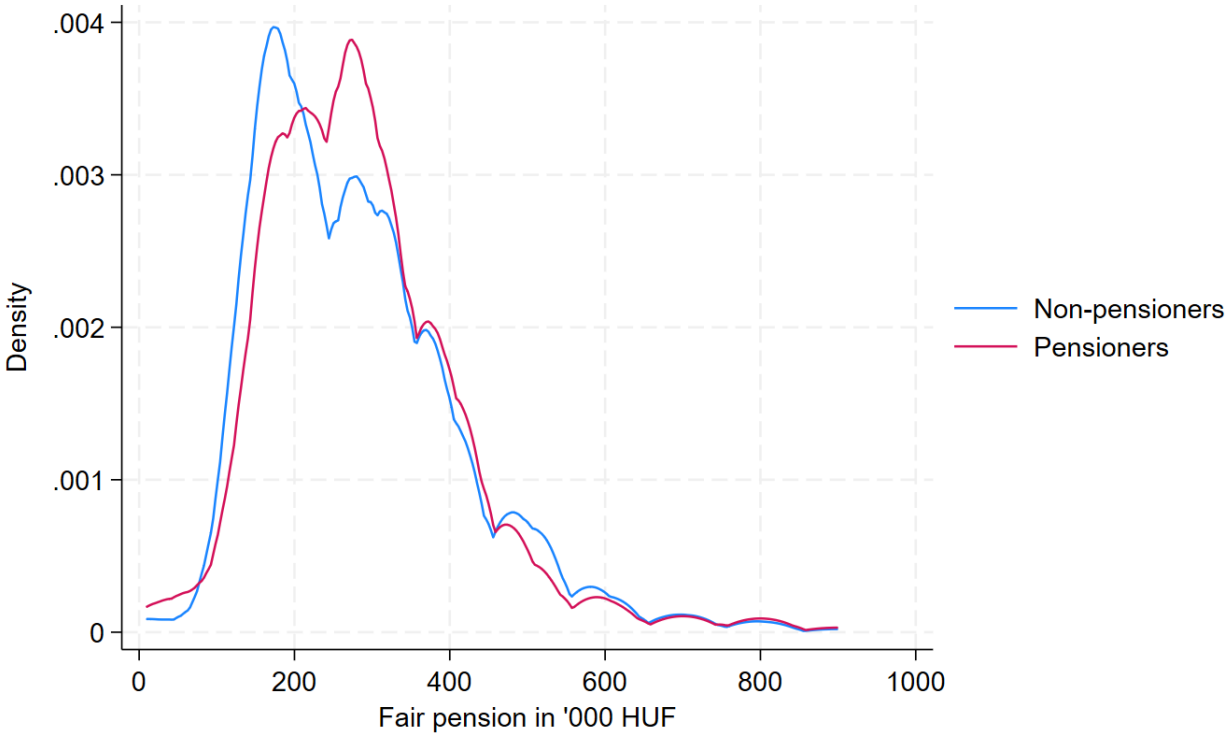
Note: The figure shows the fair amounts of monthly public pension (’000 HUF) at the 5th, 10th, 25th, 50th, 75th, 90th and 95th percentile of the pension distribution, separately among non-pensioner and pensioner respondents. The final column depicts the observed pension levels. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: observed pensions come from the Central Statistical Office, January 2024 (CSO, 2024b), fair pensions come from the conjoint experiment, own data collection.

Table A.8: Relative gaps between the estimated, the fair and the observed pensions among primary, secondary and higher educated respondents

	Primary educated			Secondary educated			Higher educated		
	Perc. bias	Perceived bias	Actual bias	Perc. bias	Perceived bias	Actual bias	Perc. bias	Perceived bias	Actual bias
p5	1.31	1.08	1.41	1.21	1.08	1.31	1.41	0.93	1.31
p10	1.27	1.07	1.36	1.27	1.07	1.36	1.36	1.00	1.36
p25	1.13	1.12	1.27	1.08	1.11	1.20	1.20	1.00	1.20
p50	1.07	1.18	1.27	0.98	1.30	1.27	1.07	1.14	1.22
p75	1.19	1.09	1.30	1.17	1.11	1.30	1.19	1.09	1.30
p90	1.18	1.06	1.25	1.25	1.00	1.25	1.11	1.13	1.25
p95	1.16	1.00	1.16	1.28	0.91	1.16	1.16	1.04	1.21

Note: The table show the relative policy bias among primary, secondary and higher educated respondents. Perception bias in relative terms is defined as the ratio of the estimated pensions and the actual pensions at several percentiles of the pension distributions. The perceived policy bias is calculated as the ratio of the fair and the estimated pensions. The actual policy bias is the ratio of the fair and actual pensions. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: observed pensions come from the Central Statistical Office, January 2024 (CSO, 2024b), perceptions and preferences come from own data collection.

Figure A.11: Distribution of fair pensions ('000 HUF) among non-pensioners and pensioners



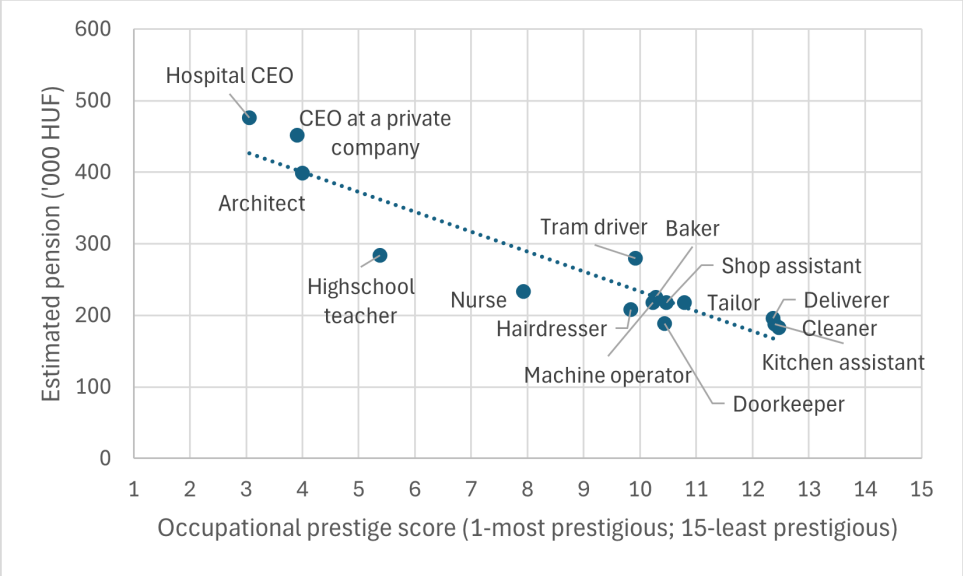
Note: The figure shows the Kernel density estimates for the fair amounts of monthly public pension ('000 HUF) according to non-pensioners and pensioners. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: conjoint experiment, own data collection.

Table A.9: Relative gaps between the estimated, the fair and the observed pensions among primary, secondary and higher educated respondents

	Non-pensioners			Pensioners		
	Perc. bias	Perceived bias	Actual bias	Perc. bias	Perceived bias	Actual bias
p5	1.31	1.00	1.31	1.26	1.12	1.41
p10	1.32	1.03	1.36	1.27	1.07	1.36
p25	1.13	1.06	1.20	1.13	1.18	1.33
p50	1.02	1.19	1.22	1.07	1.27	1.37
p75	1.19	1.09	1.30	1.11	1.17	1.30
p90	1.25	1.00	1.25	1.25	1.00	1.25
p95	1.16	1.00	1.16	1.36	0.85	1.16

Note: The table show the relative policy bias among non-pensioner and pensioner respondents. Perception bias in relative terms is defined as the ratio of the estimated pensions and the actual pensions at several percentiles of the pension distributions. The perceived policy bias is calculated as the ratio of the fair and the estimated pensions. The actual policy bias is the ratio of the fair and actual pensions. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: observed pensions come from the Central Statistical Office, January 2024 (CSO, 2024b), perceptions and preferences come from own data collection.

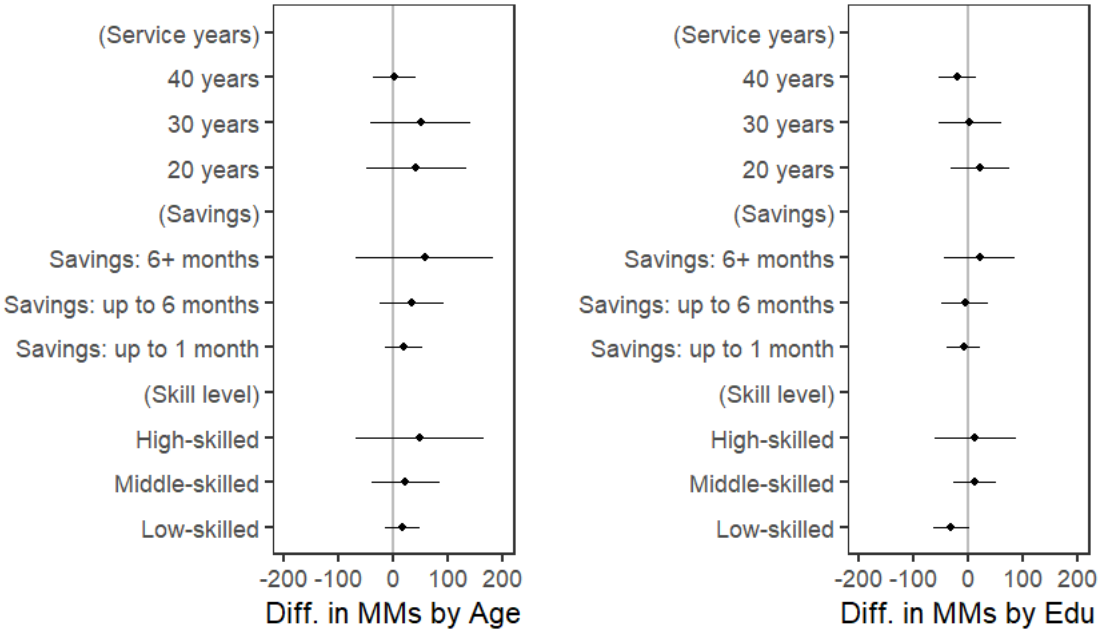
Figure A.12: Prestige score and estimated pension for the occupations



Note: The figure shows the occupational prestige score (from 1-most prestigious to 15-least prestigious) and the amounts of monthly public pension ('000 HUF) that respondents estimate for each occupation. The average exchange rate in January 2024: 1 EUR = 382 HUF. Data source: occupational prestige scores are from the the Central Statistical Office, 2016 (CSO, 2016) and fair pensions come from own data collection.

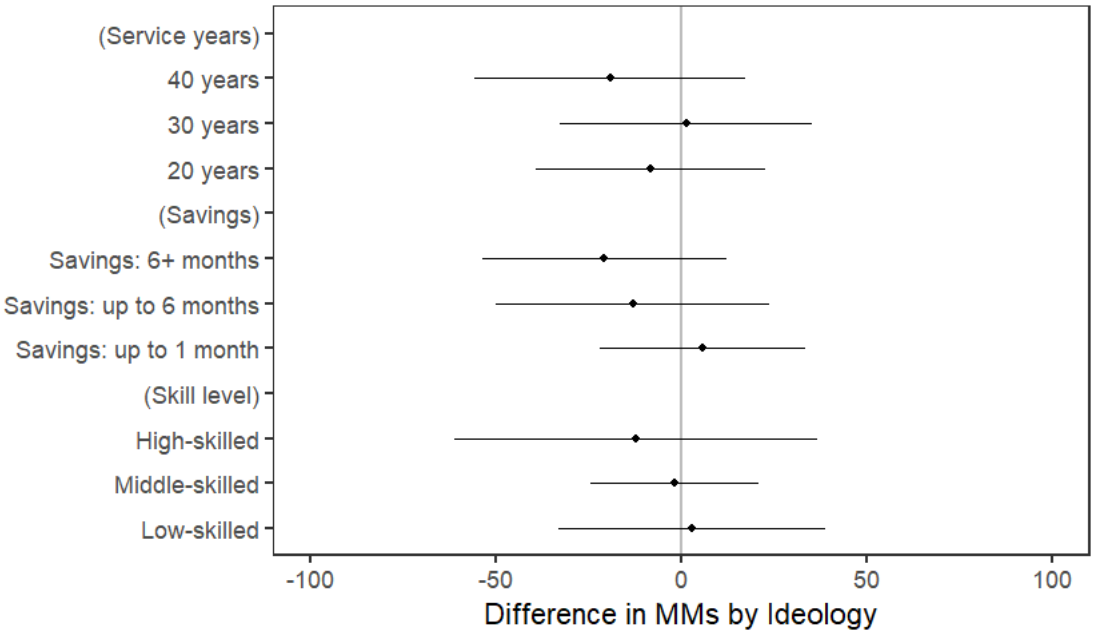
A.2.3 Results of the conjoint experiment

Figure A.13: Differences in conditional marginal means of fair pensions (’000 HUF), by age and education



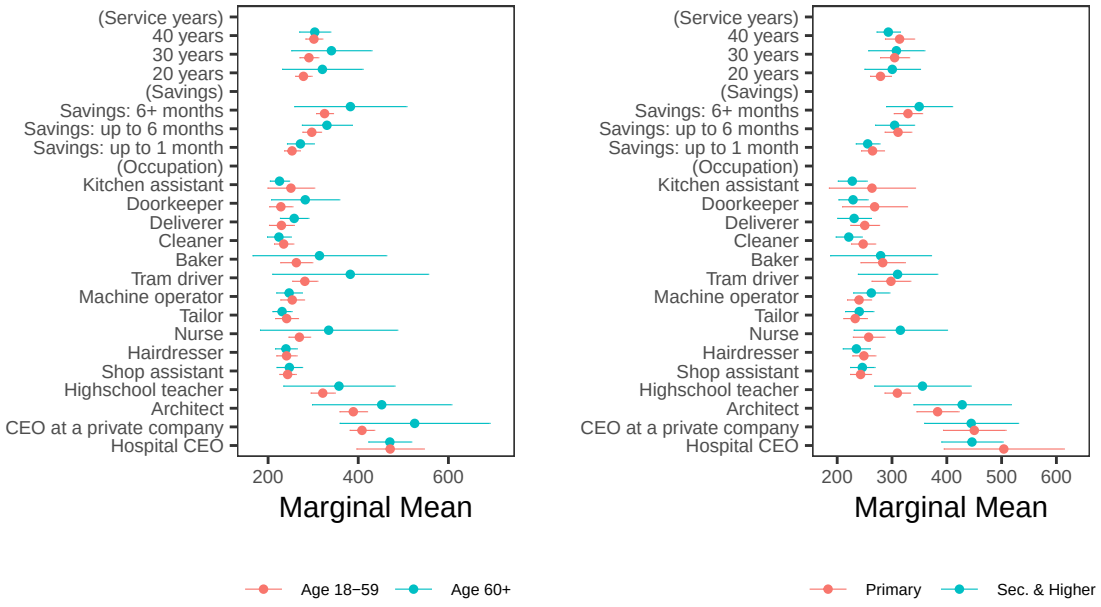
Note: Figure shows the differences in marginal means for three pensioner attributes, by age (on the left) and education level (on the right) of the respondents. The dependent variable is the amount of pension (’000 HUF) the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied. (The average exchange rate in January 2024: 1 EUR = 382 HUF.)

Figure A.14: Differences in conditional marginal means of fair pensions (’000 HUF), by ideology



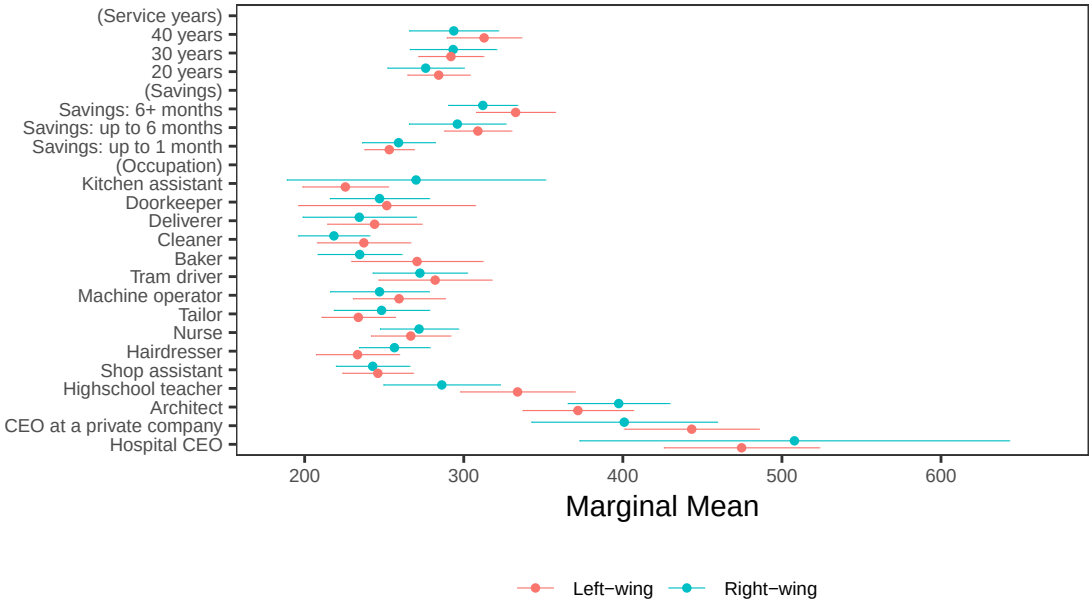
Note: Figure shows the differences in marginal means for three pensioner attributes, by the ideological stance of the respondents. The dependent variable is the amount of pension (’000 HUF) the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied. (The average exchange rate in January 2024: 1 EUR = 382 HUF.)

Figure A.15: The assigned fair pension (’000 HUF) by age and education of respondents



Note: Figure shows the marginal means of pensioner attributes. The dependent variable is the amount of pension (’000 HUF) the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied. (The average exchange rate in January 2024: 1 EUR = 382 HUF.)

Figure A.16: The assigned fair pension ('000 HUF) by ideology of respondents



Note: Figure shows the marginal means of pensioner attributes. The dependent variable is the amount of pension ('000 HUF) the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied. (The average exchange rate in January 2024: 1 EUR = 382 HUF.)

A.3 Appendix - Chapter 3

A.3.1 Descriptive results

Figure A.17: Questions of the conjoint experiment

2. What do you think, what would be a fair pension compared to the last salary of the 1st and 2nd pensioner? Please enter a % from 1 to 100%, where i.e. 25% corresponds to a quarter of the salary, 50% to half, 75% to three-quarters and 100% to the full amount of the salary.

---------- (%) and ---------- (%)

3. Let's say that to sustain the pension system, the pension of one of the pensioners must be cut. One way is to cut the 13th month pension, at least for some pensioners. If you had to choose, which pensioner's 13th month pension would you cut?

- 1st pensioner
- 2nd pensioner

Table A.10: Descriptive statistics about the respondents

Share of...	**'Replacement rate-respondents'**	**'Pension cut-respondents'**
Female	60%	60%
Aged 60+	38%	37%
Pensioner	31%	30%
Secondary and higher educated	57%	57%
Right-wing	42%	43%
Number of observation	5668	4148

Note: Descriptive statistics about the respondents in the conjoint experiment. 'Replacement rate-respondents' are those who responded the questions about a fair replacement rate for the hypothetical pensioners. 'Pension cut-respondents' are those who responded and targeted a pension cut in the conjoint experiment.

Table A.11: Fair pension replacement rate among primary, secondary, and higher educated respondents (%)

Fair replacement rate (%)	**Education**		
	Primary	Secondary	Higher
Mean	69	71	69
SD	29	29	28
Median	75	80	75
N	2426	2308	934

Note: The table shows the mean and median fair replacement rate by the educational level of the respondents.

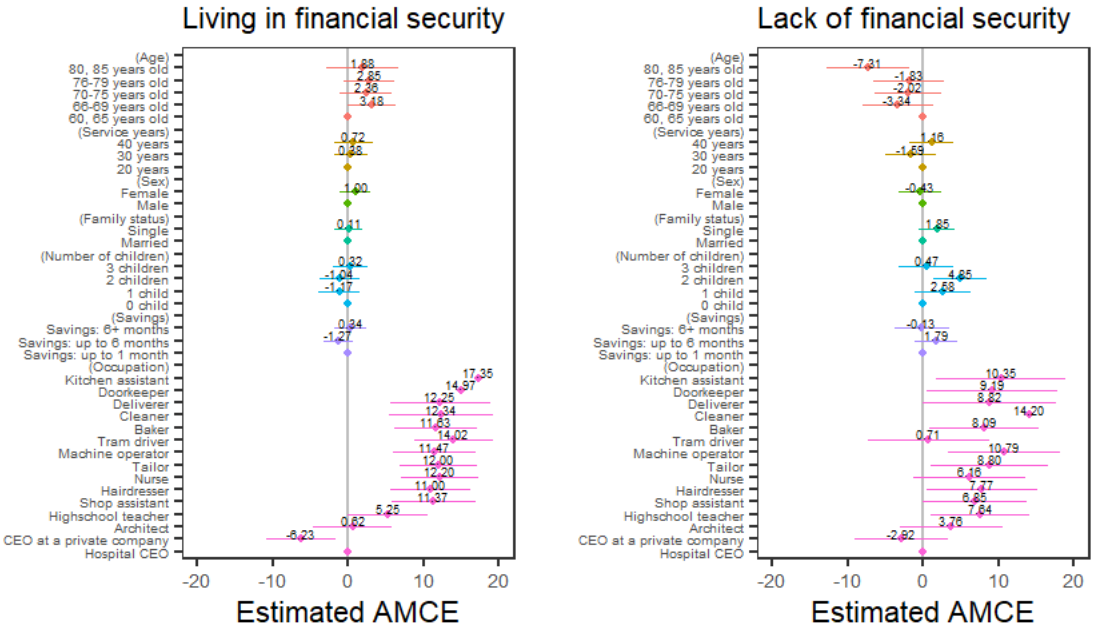
Table A.12: Fair pension replacement rate among left-wing and right-wing respondents (%)

Fair replacement rate (%)	**Ideology**	
	Left-wing	Right-wing
Mean	70	71
SD	29	28
Median	75	80
N	3010	2165

Note: The table shows the mean and median fair replacement rate by the ideological stance of the respondents.

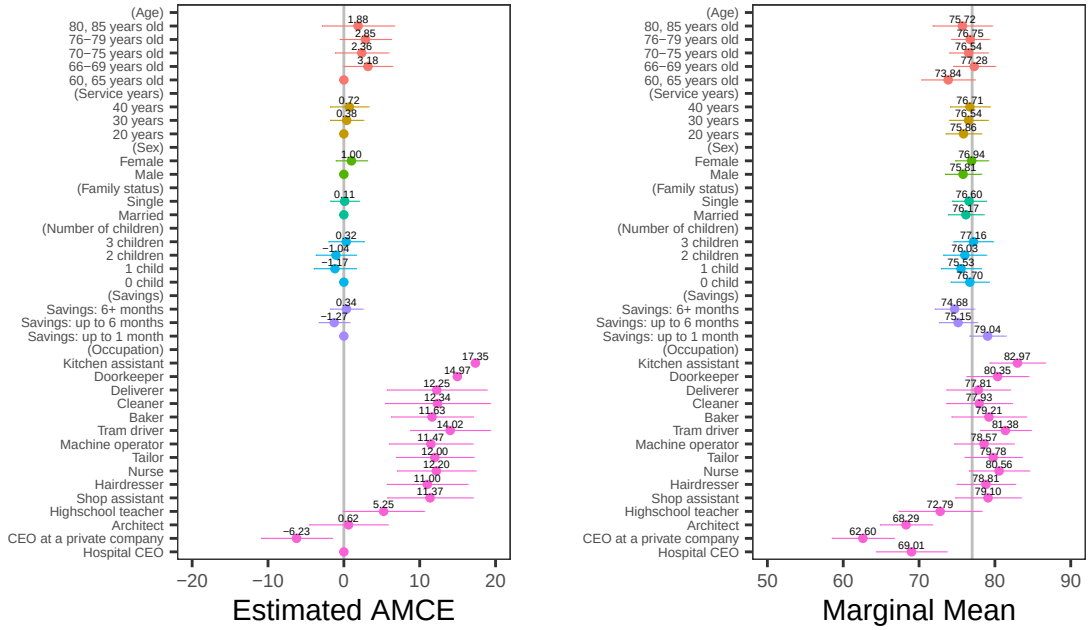
A.3.2 Results of the conjoint experiment

Figure A.18: The effects of pensioner attributes on the assigned fair pension replacement rate (%), by subjective financial security



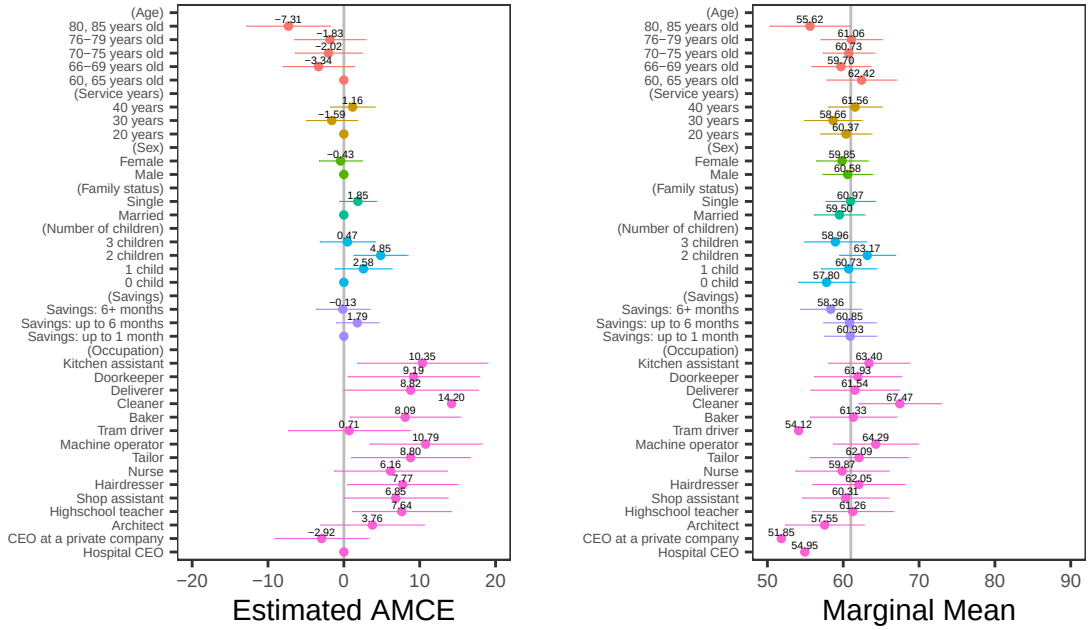
Note: Figure shows the average marginal component effects and marginal means of pensioner attributes. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Point estimates are reported above the dots. The dots on the zero line denote the reference category for each attribute, when estimating AMCEs. The vertical line on the right panel indicated the overall mean of fair replacement rates. Weights are applied.

Figure A.19: Drivers of replacement rate among those who live in financial security



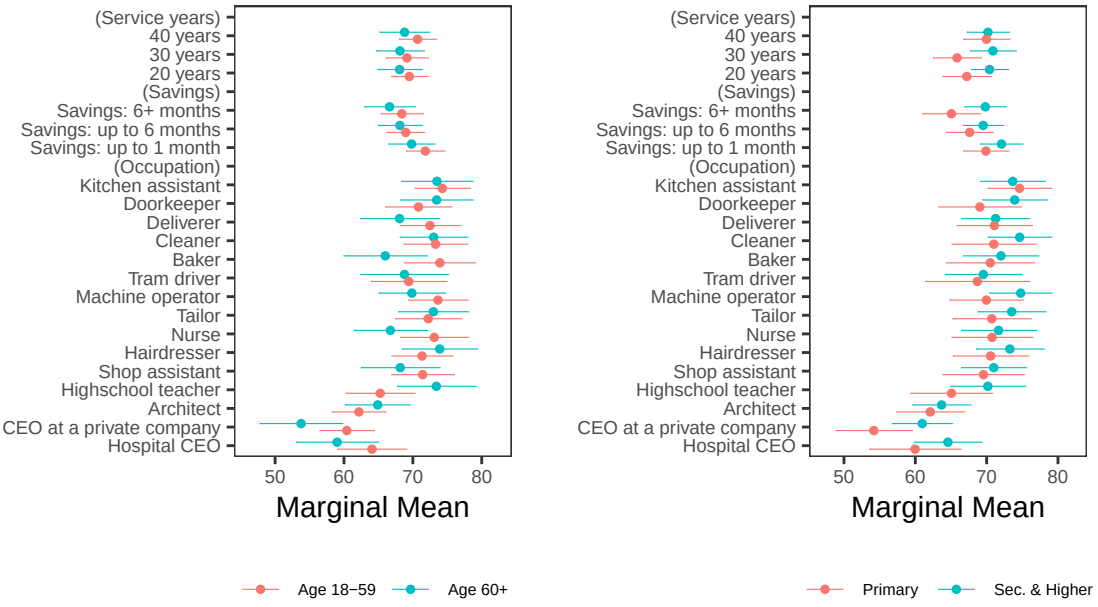
Note: Figure shows the average marginal component effects of pensioner attributes. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.20: Drivers of replacement rate among those who lack financial security



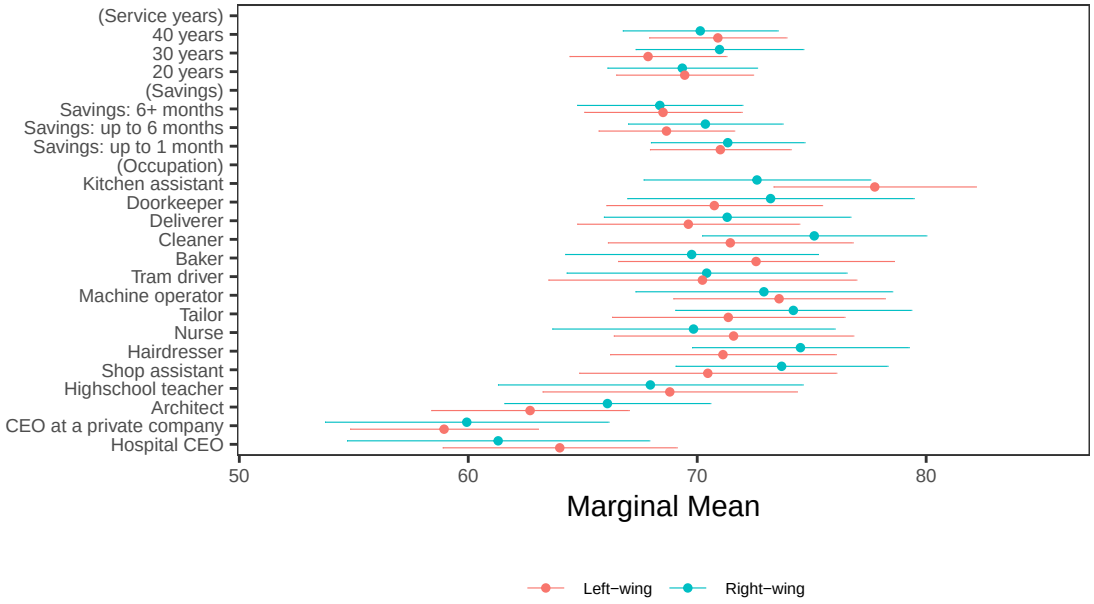
Note: Figure shows the... Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.21: The assigned fair pension replacement rates (%) by age and education of the respondents



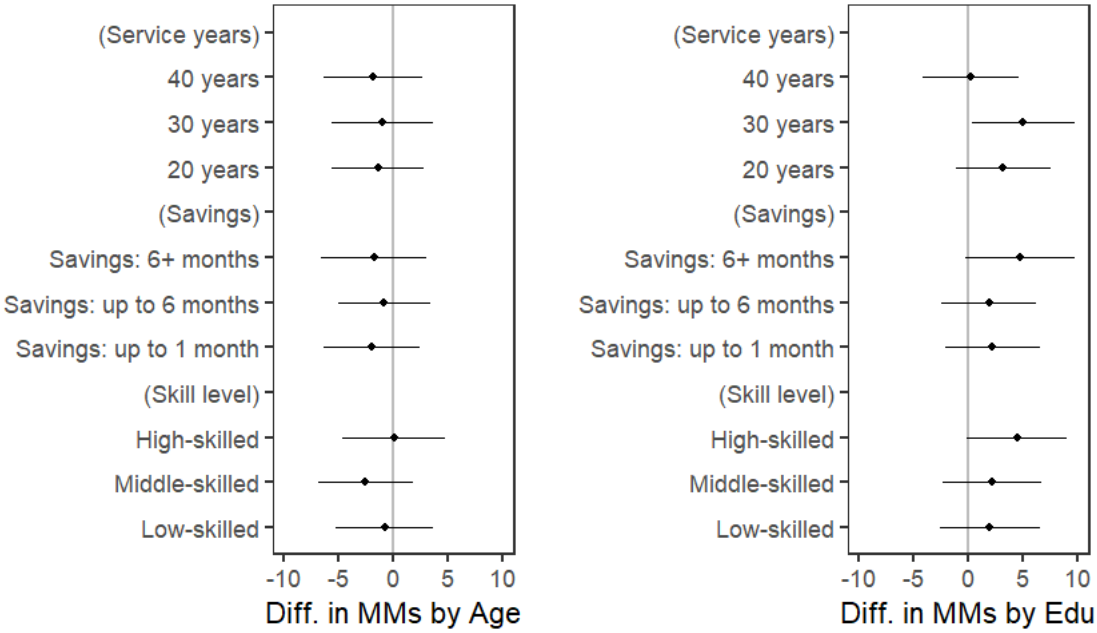
Note: Figure shows the marginal means of pensioner attributes. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.22: The assigned fair pension replacement rates (%) by ideology of the respondents



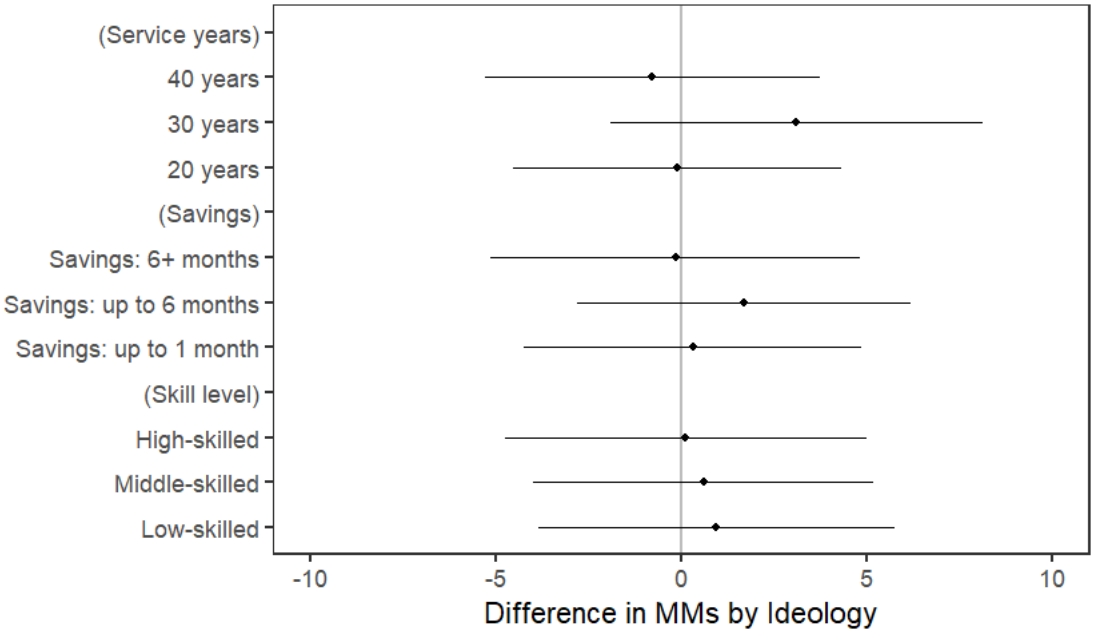
Note: Figure shows the marginal means of pensioner attributes. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.23: Differences in conditional marginal means of replacement rates (%), by the age and by the education level of the respondents



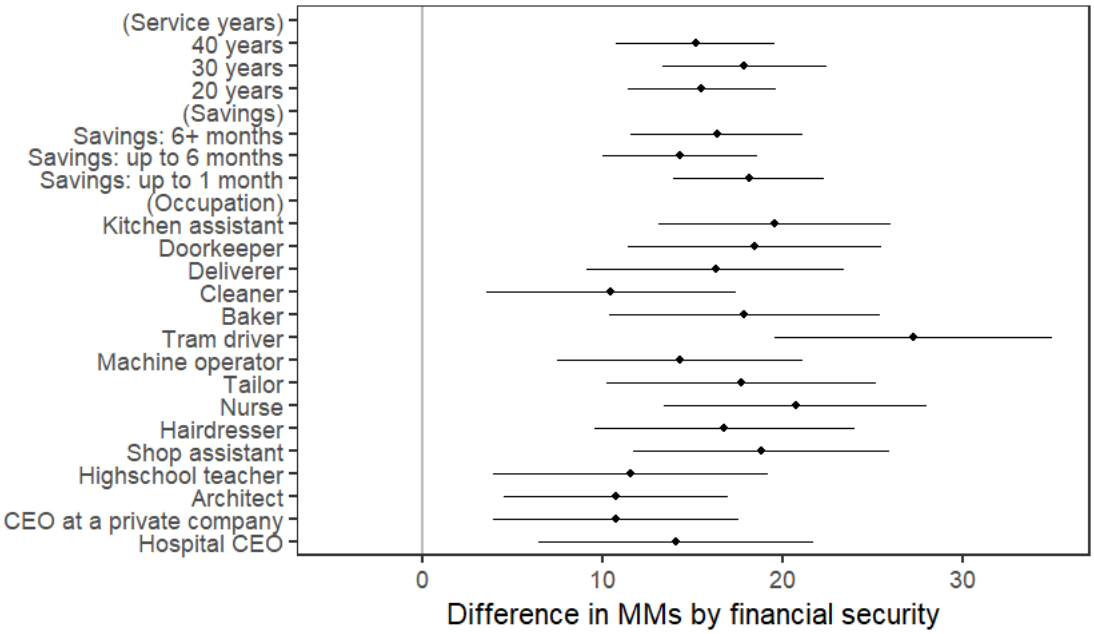
Note: Figure shows the differences in marginal means for three pensioner attributes, by the education level of the respondents. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.24: Differences in conditional marginal means of replacement rates (%), by ideology



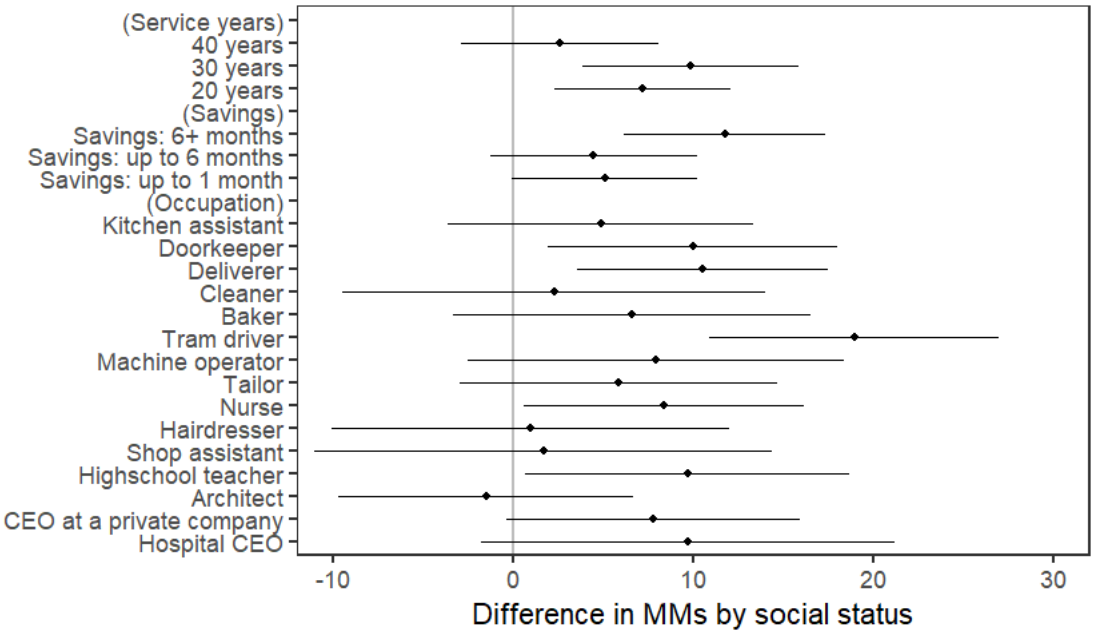
Note: Figure shows the differences in marginal means for three pensioner attributes, by the ideological stance of the respondents. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.25: Differences in conditional marginal means of replacement rates (%), by the subjective financial security of the respondents



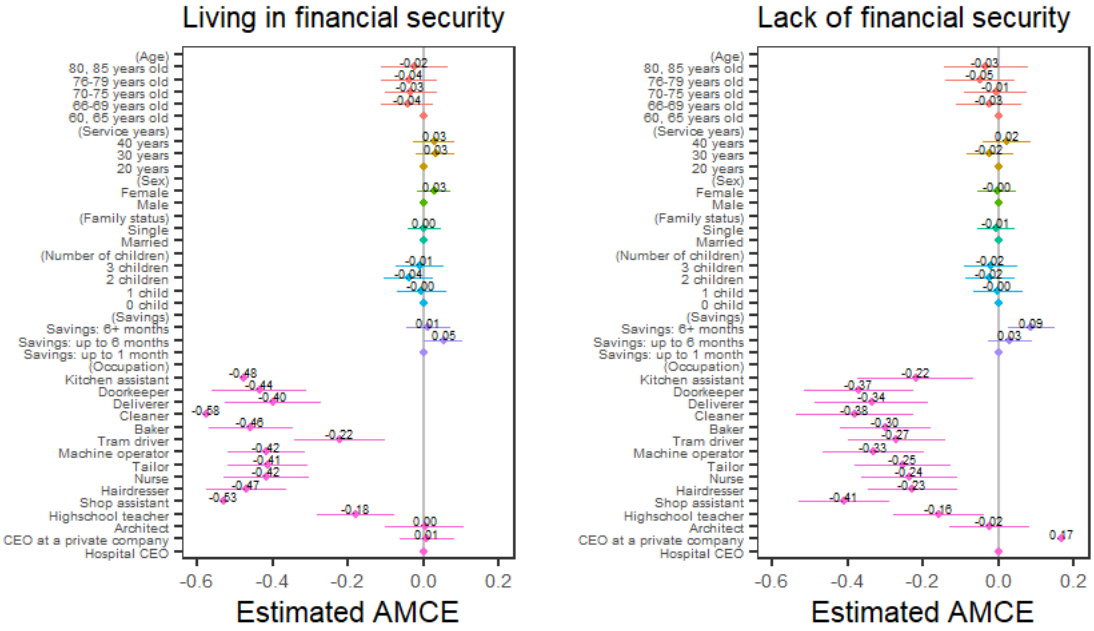
Note: Figure shows the differences in marginal means for three pensioner attributes, by the subjective financial security of the respondents. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.26: Differences in conditional marginal means of replacement rates (%), by the subjective social status of the respondents



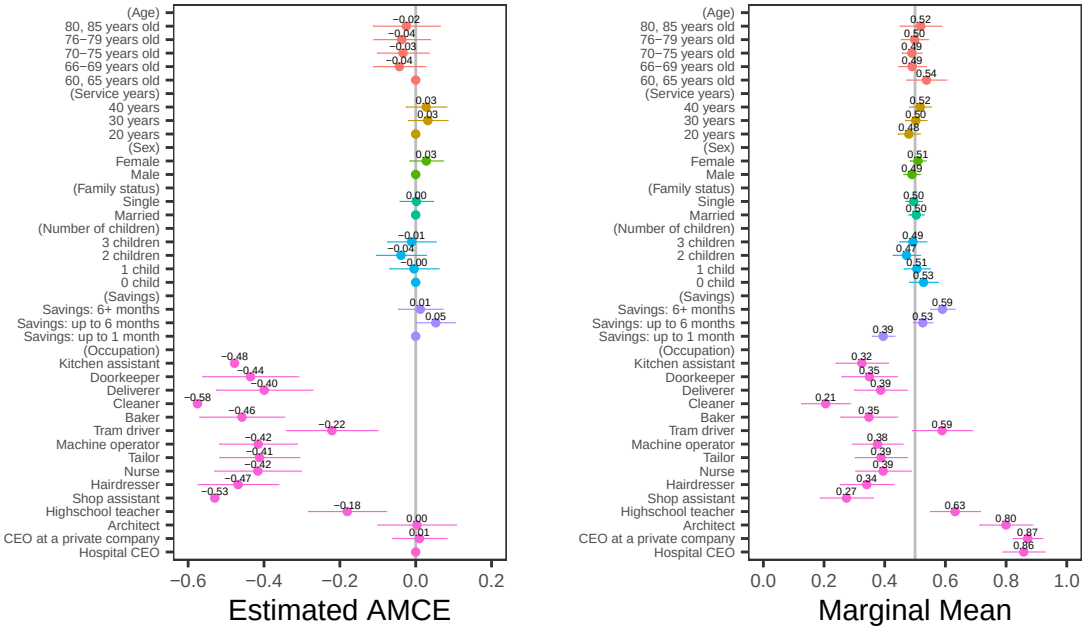
Note: Figure shows the differences in marginal means for three pensioner attributes, by the subjective social status of the respondents. The dependent variable is the pension replacement rate the respondent considers fair. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.27: The effects of pensioner attributes on a forced pension cut-back, by financial security of the respondents



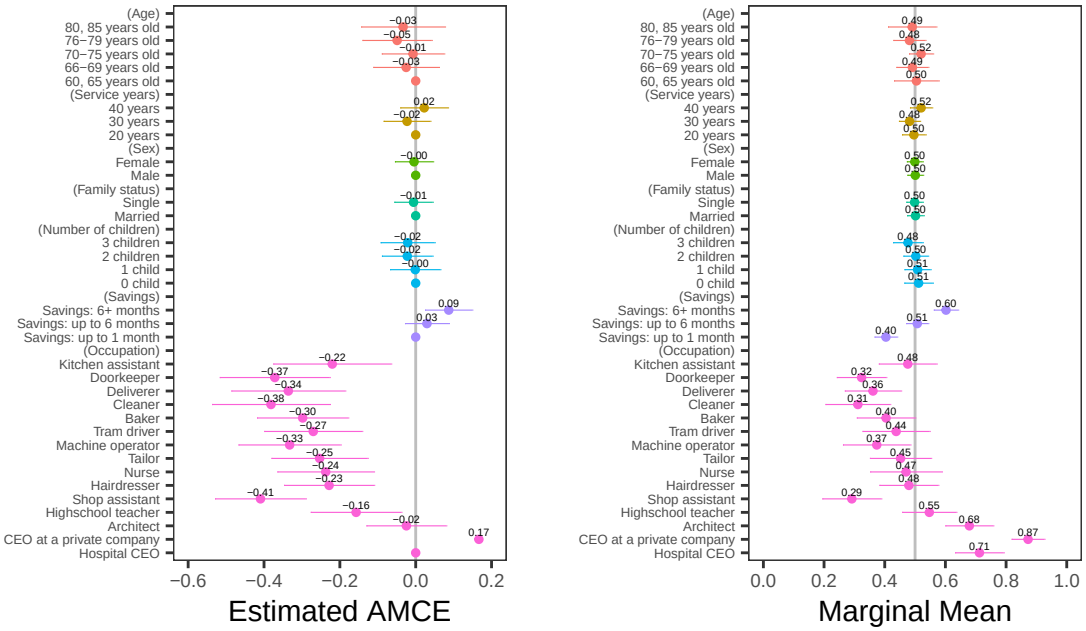
Note: Figure shows the average marginal component effects and marginal means of pensioner attributes. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Point estimates are reported above the dots. The dots on the zero line denote the reference category for each attribute. Weights are applied.

Figure A.28: Drivers of pension cut among those who live in financial security



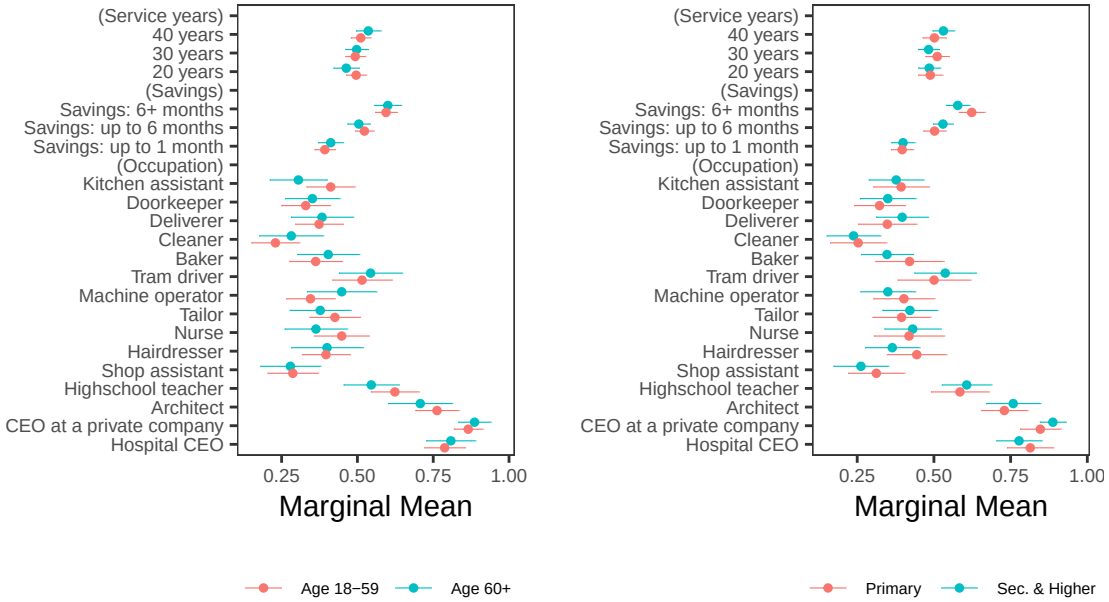
Note: Figure shows the... Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.29: Drivers of pension cut among those who lack financial security



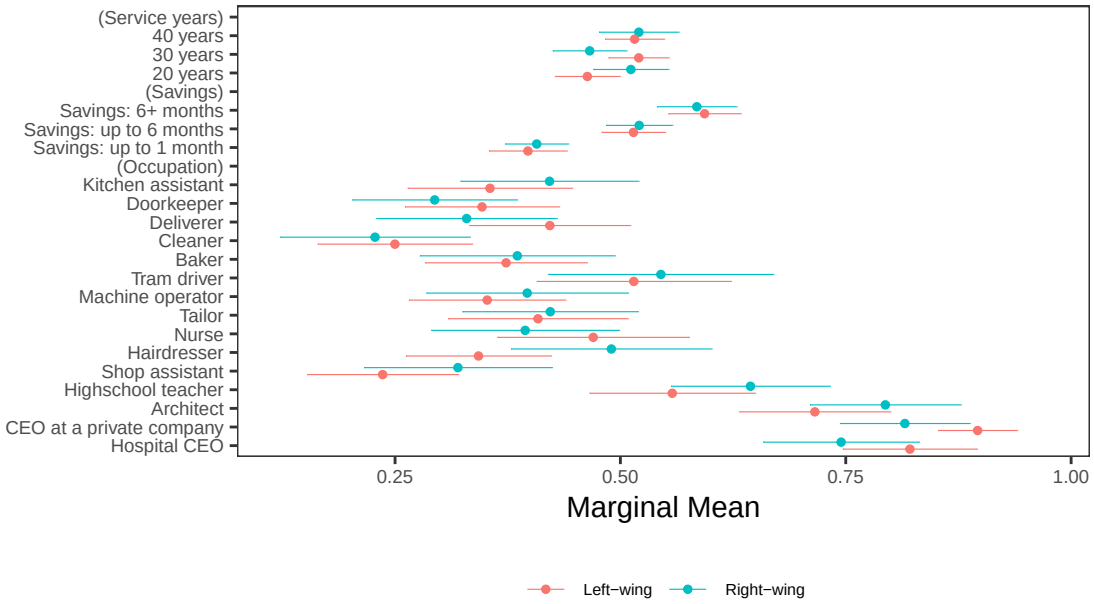
Note: Figure shows the... Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.30: Probability of a pension cut-back by age and education of the respondents



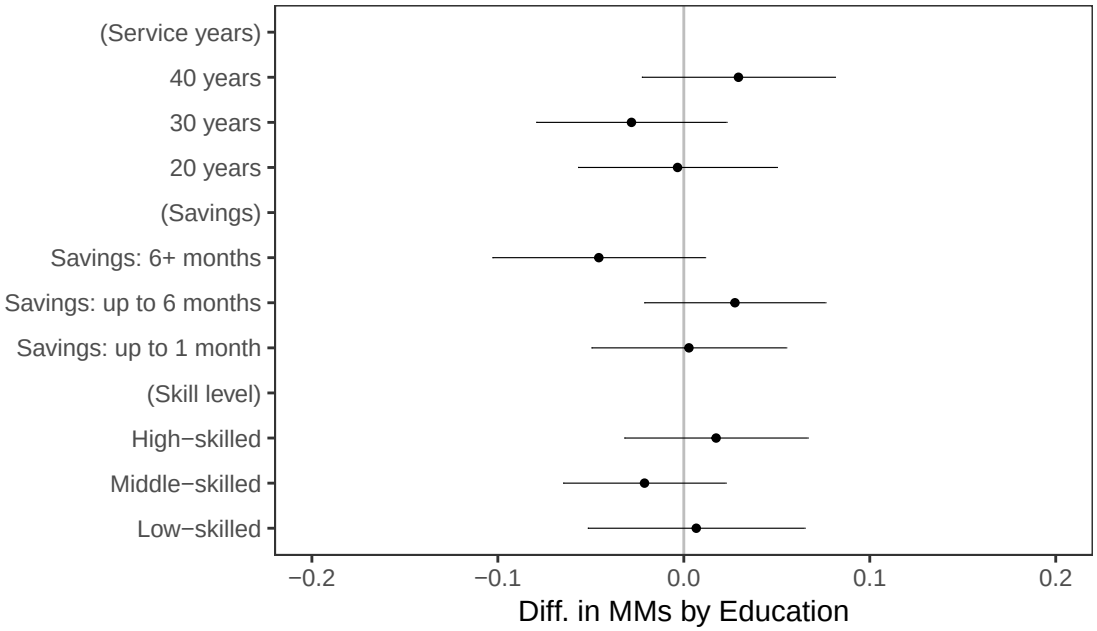
Note: Figure shows differences in the marginal means of pensioner attributes. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.31: Probability of a pension cut-back by ideology of the respondents



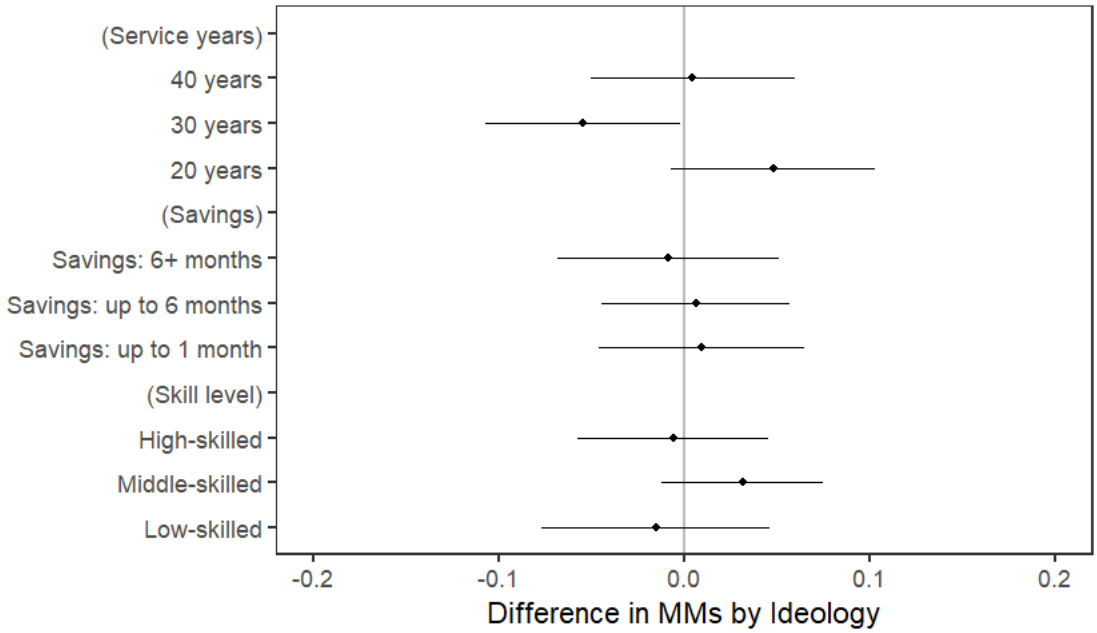
Note: Figure shows the marginal means of pensioner attributes. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.32: Differences in conditional marginal means of pension cut-back probability, by education level of the respondents



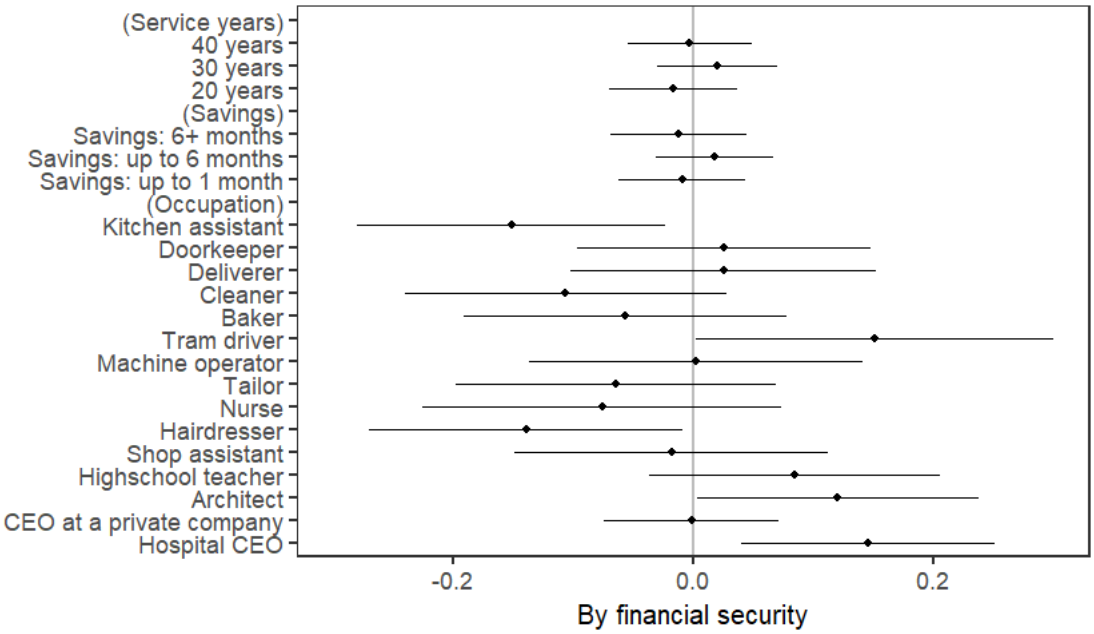
Note: Figure shows the differences in marginal means for three pensioner attributes, by age (on the left) and education level (on the right) of the respondents. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.33: Differences in conditional marginal means of pension cut-back probability, by ideology



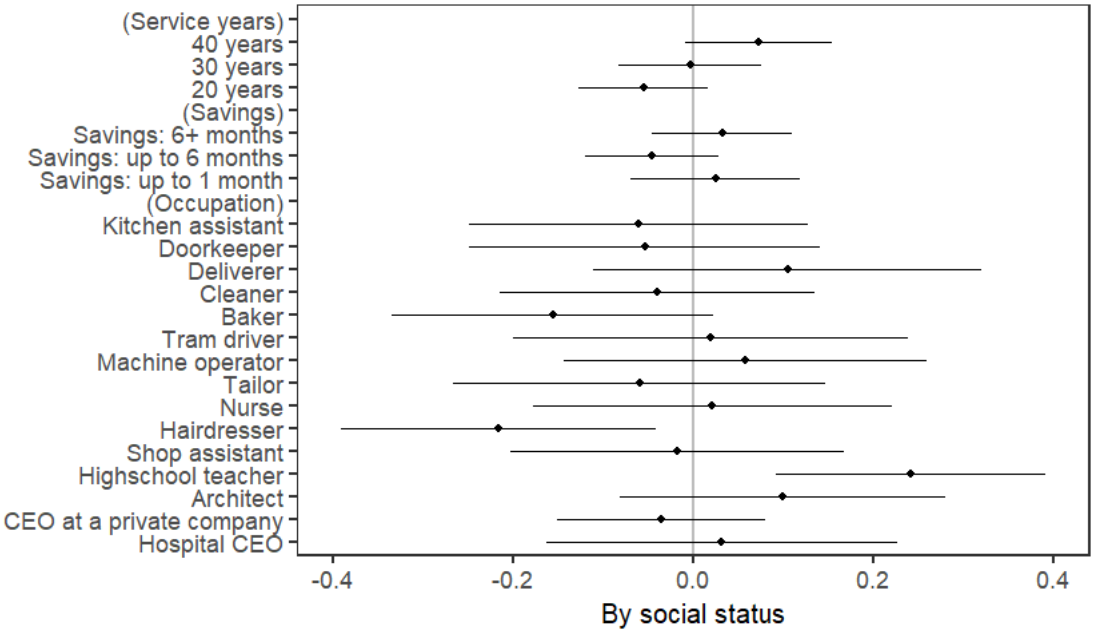
Note: Figure shows the differences in marginal means for three pensioner attributes, by the ideological stance of the respondents. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.34: Differences in conditional marginal means of pension cut-back probability, by the subjective financial security of the respondents



Note: Figure shows the differences in marginal means for three pensioner attributes, by the subjective financial security of the respondents. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.

Figure A.35: Differences in conditional marginal means of pension cut-back probability, by the subjective social status of the respondents



Note: Figure shows the differences in marginal means for three pensioner attributes, by the subjective social status of the respondents. The dependent variable is a forced choice variable, it equals one if the respondent chose to cut the 13th month extra pension of a pensioner and zero otherwise. Dots with horizontal lines indicate point estimates with cluster-robust 95% confidence intervals from a linear probability model estimated via least squares regression. Weights are applied.