Preferences for the commodification of pensions in Chile: the rol of intergenerational social mobility

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Abstract

My abstract

Keywords: Pension commodification · Intergenerational mobility · Causal inference · Meritocracy · Chile

This document was last modified at 2025-11-17 19:08:04 and it was last rendered at 2025-11-17 19:08:04.

1 Introduction

What is the legitimate extent of market inequality in the eyes of the public? Since the early 1980s, many countries have experienced a widespread retreat from universal welfare programs and a shift toward the privatization and commodification of public goods, welfare policies, and social services (Gingrich, 2011; Streeck, 2016). In Latin America, as elsewhere, neoliberal reforms reshaped welfare-state institutions by extending market logic into domains of social reproduction that were traditionally governed by the state (Ferre, 2023). This transformation reduced the role of public provision and increased the presence of private actors in core social services (Ferre, 2023). Echoing Polanyi's (1975) insight that markets constitute a distinct moral order, the institutional diffusion of market rules has fostered a corresponding moral economy: a constellation of norms and values concerning fair allocation, embedded in institutions and shaping individual subjectivities (Mau, 2015; Svallfors, 2006). Within this framework, a growing body of research examines the extent to which, and the mechanisms by which, citizens consider it fair that the allocation of services like health care, pensions, and education be governed by market-based criteria—a phenomenon known as *market justice preferences* (Busemeyer, 2014; Castillo, Laffert, et al., 2025; Castillo et al., 2024; Immergut & Schneider, 2020; Koos & Sachweh, 2019; Lindh, 2015; Lindh & McCall, 2023). Understanding these preferences is crucial, as they contribute to legitimizing economic inequality by framing it as the fair result of individual responsibility and limited state intervention (Mau, 2015).

Existing literature shows that market justice preferences are shaped by both the economic and institutional context and individuals' positions within social stratification. Grounded in the notion that economic institutions influence people's normative attitudes (Immergut, 1998), studies find that countries with stronger public provision or more expansive welfare states exhibit lower levels of market justice preferences (Busemeyer, 2014; Immergut & Schneider, 2020), while more privatized contexts show stronger support for market-based criteria (Castillo et al., 2024; Lindh, 2015). In such contexts, market justice preferences tend to rise as individuals "ascend" the social structure, with those in more privileged positions in terms of class, education, and income holding stronger preferences for market-based solutions compared to those in more disadvantaged or at-risk positions (Castillo et al., 2024; Immergut & Schneider, 2020; Lee & Stacey, 2023; Lindh, 2015; Otero & Mendoza, 2024; Svallfors, 2007; Von Dem Knesebeck et al., 2016).

Market justice preferences are shaped not only by objective socioeconomic conditions but also by popular beliefs about inequality. Among these, meritocracy is a key normative principle underpinning market-based distributive preferences (Mau, 2015). Studies show that individuals with stronger meritocratic beliefs tend to perceive less inequality and legitimize it by attributing economic differences to personal achievement (Batruch et al., 2023; Mijs, 2019; Wilson, 2003). In highly unequal societies where access to services is largely governed by market logic, such beliefs play a critical role in normalizing inequality. Recent evidence from Chile shows that students who believe effort and talent are rewarded in their country express stronger preferences for market-based access to healthcare, pensions, and education (Castillo et al., 2024).

Although it is clear that one's social position influences market justice preferences, the question of how upward or downward mobility within the social structure affects these preferences remains unanswered. This question is far from trivial, especially in Latin America, where many have experienced various forms of mobility amid high economic inequality and deep welfare privatization (Ferre, 2023; López-Roldán & Fachelli, 2021; Torche, 2014). Social origins and destinations affect attitudes toward inequality in distinct ways (Day & Fiske, 2017; Gugushvili, 2016b, 2017; Jaime-Castillo & Marqués-Perales, 2019; Mijs et al., 2022; Wen & Witteveen, 2021), while movement between these positions exposes individuals to different experiences and mechanisms that shape their views on what is fair (Gugushvili, 2014; Mau, 2015). Building on this research, examining the effects of social mobility on market justice preferences can help to illuminate how inequalities in access to social services are justified among individuals who have experienced, or not, changes in their social standing, and what are the normative mechanisms that guide this justification (Mau, 2015).

Beyond their isolated effects, social mobility and meritocratic beliefs interact in complex ways to shape market justice preferences. Among others, a key mechanism proposed in the literature to explain how mobility influences distributive justice preferences is the psychological process of self-serving bias in causal attribution (Gugushvili, 2016a; Schmidt, 2011). This bias suggests that individuals attribute failures—such as downward mobility—to external factors, while crediting successes—such as upward mobility—to their own merit and effort (Miller & Ross, 1975). Those who experience upward mobility tend to view their social position as earned, making them more likely to believe that individuals are responsible for their own success or failure. Research shows that upward mobility is associated with weaker preferences for redistribution (Alesina et al., 2018; Gugushvili, 2016a; Jaime-Castillo & Marqués-Perales, 2019; Schmidt, 2011) and stronger legitimacy of income inequality (Shariff et al., 2016). In contrast, individuals who experience downward mobility tend to blame structural factors like inequality and are more supportive of redistribution while rejecting merit-based explanations (Gugushvili, 2014). Taken together, I argue that meritocratic beliefs may reinforce this self-serving attribution mechanism by legitimizing one's social status as the outcome of personal merit, closely tied to attribution bias.

Against this background, this article pursues two main objectives: first, to analyze the extent to which intergenerational social mobility influences market justice preferences regarding healthcare, pensions, and education; and second, to examine how meritocratic beliefs may moderate this relationship. Building on a theoretical framework that emphasizes how neoliberal transformations—particularly through the privatization and commodification of key areas of social reproduction—have profoundly reshaped processes of subject formation (Mau, 2015), the central argument is that upward mobility increases support for market justice preferences, while downward mobility decreases it. Moreover, meritocratic beliefs are expected to moderate this relationship by reflecting a self-serving attribution bias, whereby individuals justify their social position in terms of personal merit.

This study focuses on Chile, a particularly intriguing case for examining market justice preferences. Despite significant economic growth and poverty reduction, Chile has some of the highest levels of inequality in Latin America and among OECD countries (Chancel et al., 2022; Flores et al., 2020). This inequality coexists with short-range upward mobility among lower-class segments moving into middle strata, though strong barriers remain to reaching higher positions (Espinoza et al., 2013; Torche, 2014). What makes Chile especially salient is that much of this inequality is rooted in deep neoliberal reforms that institutionalized the privatization and commodification of key social sectors (Madariaga, 2020). Introduced during the dictatorship (1973–1989) and expanded in democracy, these reforms enabled the unprecedented emergence of markets in health, pensions, and education, with provision segmented by individuals' ability to pay and supported by public subsidies (Boccardo, 2020). In parallel—and despite waves of protest against inequality and commodification from 2006 to 2019 (Somma et al., 2021)—Chilean subjectivities have been increasingly shaped by neoliberal discourses and market logics, influencing their attitudes toward inequality and welfare distribution (Araujo & Martuccelli, 2012; Canales Cerón et al., 2021).

In this context, the questions that guide this research are as follows:

(1) To what extent does intergenerational social mobility influence market justice preferences regarding healthcare, pensions, and education in Chile?

(2) How do meritocratic beliefs condition or moderate this relationship in the Chilean context?

To address these questions, this study draws on large-scale, representative survey data collected in 2018 from the urban Chilean population aged 18 to 75 (n = 2,726). The next section outlines the theoretical framework linking market justice preferences, social mobility, and meritocratic beliefs, and proposes a set of hypotheses. This is followed by a description of the data, variables, and analytical strategy. The final sections present the empirical findings, offer an interpretation of the results, and conclude with a discussion of their implications.

2 Theoretical and empirical background

2.1 Market justice preferences

Beyond state's capacity to reallocate resources from the advantaged to the vulnerable, market institutions also play a central role in distributing socially valuable goods and rewards (Koos & Sachweh, 2019; Lindh & McCall, 2023). As Polanyi (1975) observed, economic integration in capitalist societies is primarily organized through market exchange, governed by a self-regulating price system embedded in institutional frameworks. These institutions are not mere aggregates of individual behavior, but social realities endowed with rules, mechanisms, and normative meanings that shape everyday thinking (Immergut, 1998; Koos & Sachweh, 2019). In this sense, the economic order is mirrored in a moral economy: collectively shared norms and beliefs about justice in distribution, embedded and reinforced through institutions (Mau, 2015; Svallfors, 2007). While most research from this perspective has focused on welfare institutions, recent scholarship has brought the market back into focus as a site of distributive justice beliefs and institutional responses to inequality (Lindh & McCall, 2023). In many countries, privatization and commodification have expanded market logic into core areas of social reproduction, such as healthcare, education, pensions, and social security, deepening inequality in access to these services (Ferre, 2023; Gingrich, 2011). Yet public support for market-based welfare provision has grown, even in traditional welfare states, particularly among higher-income groups who view private alternatives as more efficient or higher in quality (Busemeyer & Iversen, 2020). This shift calls for broader inquiry into how institutions like markets structure access to resources and legitimize inequalities (Mau, 2015; Satz, 2019).

The legitimacy of market-based inequalities is closely tied to beliefs about distributive justice grounded in market principles. Beyond stratification structures and socioeconomic conditions, the study of inequality also focuses on individuals' beliefs about the origins of inequality, the normative frameworks that sustain those beliefs, the mechanisms that shape them, and their implications for attitudes and behavior (Kluegel & Smith, 1981). In this regard, empirical research on distributive justice focuses on individuals' conceptions of how goods and rewards *should* be distributed in society (Jasso et al., 2016). This line of inquiry allows for examining the extent to which economic inequality is perceived as just, or in a certain sense, legitimate (Castillo, 2011). Most of the research on this field explores the legitimacy of wage inequality, specially salary gaps between jobs at opposite ends of the occupational hierarchy (Castillo, 2011; Jasso, 1978; Jasso & Wegener, 1999). Recently, others examines how individuals justify market-generated inequalities in access to core social services such as healthcare, education, pensions or social security. Here, legitimacy stems from the belief that access to these goods should follow market-based criteria (Castillo et al., 2024; Lindh, 2015). In such views, these services are treated as legitimate commodities; goods that can be traded, priced, and evaluated through market logic (Busemeyer & Iversen, 2020).

Market justice preferences refer to normative beliefs that legitimize the idea that access to core social services—such as healthcare, education, or pensions—should be determined by market-based criteria. Following Janmaat's (2013, p. 359) distinction, these preferences fall under the category of "beliefs," understood as normative ideas about what inequality should look like, as opposed to "perceptions," which refer to subjective evaluations of existing inequality. Market justice preferences reflect the view that access to these services should depend on individuals' ability to pay, thus justifying inequalities generated by market mechanisms (Kluegel et al., 1999; Lindh, 2015). The concept draws on Lane's (1986) classic contrast between market justice and political justice: the former is grounded in the idea of earned deserts—where rewards reflect effort, productivity, and skill—while the latter prioritizes need and equality, typically expressed in welfare state policies. Lane argued that markets and states differ in purpose (efficiency vs. need), logic (individual vs. collective), and fairness criteria (merit vs. equality). Market justice assumes that markets are neutral, self-regulating systems in which fair procedures yield outcomes proportional to merit. Inequality, from this perspective, is not only expected but legitimate—as long as it arises from fair competition. In this way, market justice offers a moral lens through which individuals can view the commodification and stratified access to social services as fair and justified (Kluegel et al., 1999; Lindh, 2015).

The empirical study of market justice preferences has employed various strategies to capture how individuals assess inequalities arising from market allocation. A common approach evaluates whether people find it fair that access to essential social services—such as healthcare, education, or pensions—depends on income. This builds on foundational work by Kluegel and Smith (1999), who examined normative justifications for capitalist systems. Subsequent studies have extended this logic to a wider set of welfare goods. For example, Immergut and Schneider (2020) and von dem Knesebeck et al. (2016) explore whether respondents believe it is just that higher-income individuals receive better healthcare. Similarly, Lee and Stacey (2023) and Castillo, Iturra, et al. (2025) apply this framework to education. Complementing these efforts, more recent research has introduced composite indicators to capture broader orientations toward market-based allocation. Lindh (2015), for instance, constructs an index averaging support for income-based access to healthcare and education in comparative perspective, while Castillo et al. (2024) propose a single-item measure covering health, education, and pensions in Chile. These instruments seek to gauge to what extent individuals view market-generated inequalities as legitimate. Taken together, they capture two core dimensions of market distribution: the role of economic resources as a key determinant of outcomes, and the framing of services as tradable commodities that can be bought and sold according to ability to pay (Lindh, 2015).

Comparative empirical research has identified several individual-level factors that influence support for market justice. Individuals in more advantaged socioeconomic positions—those with higher income, education, and occupational status—are consistently more likely to endorse market-based distributive principles (Koos & Sachweh, 2019; Lindh, 2015; Svallfors, 2007). For example, Lindh (2015) finds that individuals from the service class are more likely to support market-based access to healthcare and education than skilled and unskilled workers across 17 relatively affluent countries. In a comparative analysis, Svallfors (2007) observes that this expected class pattern appears clearly only in Sweden, where support for private education and healthcare varies systematically by class. Busemeyer (2014) similarly shows that support for private education is stronger among high-income groups, while Immergut and Schneider (2020) and von dem Knesebeck et al. (2016) report comparable findings for healthcare, suggesting that wealthier individuals perceive private provision as a means of maintaining relative advantage. In Chile, Otero and Mendoza (2024) show that individuals with higher income and university education express stronger support for market allocation in healthcare, education, and pensions. Beliefs about inequality and political orientation also matter. Castillo et al. (2024) and Castillo, Laffert, et al. (2025) show that individuals who have strong meritocratic beliefs are more likely to support market-based distribution in Chile, while Lee and Stacy (2023) in Australia suggest that these preferences are also greater as people lean toward economic conservatism. In this sense, market justice preferences are shaped by the interplay between structural position and normative reasoning.

However, individual characteristics alone do not fully explain variation in support for market justice. A growing body of cross-national research highlights the importance of institutional arrangements in shaping these preferences. For instance, Immergut and Schneider (2020) find that in countries with higher public spending on healthcare, individuals are less likely to view income-based access as fair. Similarly, Busemeyer (2014) shows that increased public investment in education is associated with lower support for privatized provision. Conversely, Lindh (2015) finds that in countries with more market-oriented welfare systems, support for market-based distribution tends to be higher, suggesting that individual attitudes often align with institutional outputs. These findings are consistent with neo-institutionalist and policy feedback theories (Campbell, 2020), which argue that institutions do more than redistribute resources—they also shape the normative categories through which individuals assess who is deserving of support (Immergut, 1998). In this view, dominant values and preferences are both shaped by and embedded in institutional configurations (Busemeyer, 2014), reinforcing the notion that institutions are not neutral structures, but active producers of the moral frameworks that legitimize or challenge inequality.

2.2 Social mobility

The study of social mobility, its drivers and consequences has long been central to sociology. Classic theorists explored not only movement across social hierarchies but also its broader implications for class conflict, norm stability, and institutional change (Breen & Ermisch, 2024). Sorokin (1927) introduced the concept formally, defining mobility as the shift of individuals, values, or objects between positions within a stratification system, and distinguishing between horizontal and vertical forms. Later work differentiated intergenerational from intragenerational mobility, as well as absolute mobility—driven by structural change—from relative mobility, which captures the extent to which origins constrain destinations (Eyles et al., 2022). In Latin America, research shows high absolute but low relative mobility (Bucca, 2016): although educational expansion and economic modernization have enabled some upward movement, status reproduction remains strong, especially among elites (López-Roldán & Fachelli, 2021; Torche, 2014). This reflects deep structural inequality, segmented education systems, stratified labor markets, and legacies of dependent development that restrict access to mobility channels and reinforce the intergenerational transmission of advantage.

Beyond mapping mobility patterns, growing research has examined its subjective and attitudinal effects. Mobility effects—defined as outcomes resulting from movement between origin and destination classes (Breen & Ermisch, 2024)—have long attracted theoretical interest. Sorokin's (1959) dissociative hypothesis posits that mobility, whether upward or downward, may produce psychological strain due to conflicting norms between class contexts, leading to lower life satisfaction. Similarly, Lenski (1954) argued that status inconsistency—mismatches among education, income, and occupation—can undermine well-being. These ideas underpin extensive empirical work linking intergenerational mobility to outcomes such as life satisfaction, mental health, and stress (Gugushvili, 2024; Hadjar & Samuel, 2015; Präg & Gugushvili, 2021).

Research on intergenerational social mobility has increasingly examined its effects on attitudes toward economic inequality. A key area of inquiry focuses on how upward and downward mobility influence support for redistribution, though findings remain mixed. Alesina et al. (2018) show that individuals with pessimistic expectations about their mobility—particularly those anticipating downward movement—are more likely to support generous redistributive policies. Similarly, Ares (2020) finds that upwardly mobile individuals tend to be less supportive of state-led redistribution compared to those who have experienced downward mobility. Comparative studies by Schmidt (2011) and Gugushvili (2017) likewise report that subjective upward mobility is associated with weaker preferences for redistribution, while downward mobility strengthens redistributive support. However, recent causal evidence from Breen and Ermisch (2024) suggests the opposite: upward mobility may increase redistributive preferences, while downward mobility may reduce them.

Beyond redistribution, other studies have explored the impact of mobility on broader beliefs about inequality. Gugushvili (2016b) finds that upwardly mobile individuals are more likely to adopt individualistic attributions of poverty and to legitimize income inequality, particularly in post-socialist societies. Similarly, Bucca (2016) shows that subjective upward mobility reinforces individualistic explanations of wealth in seven Latin American countries. At the macro level, Shariff et al. (2016) demonstrate that higher levels of national economic mobility correlate with greater tolerance of inequality. In contrast, Day and Fiske (2017) find that low perceived mobility undermines belief in meritocracy and a just world, thereby weakening system justification. Taken together, this body of research suggests that mobility shapes attitudes toward inequality and justice through multiple, and sometimes contradictory, mechanisms.

The literature on the effects of social mobility has proposed various mechanisms to explain how and why changes in social position may influence individual outcomes (Helgason & Rehm, 2025). One of the most prominent is the self-interest mechanism, which posits that individuals who experience upward or downward mobility undergo a shift in their material interests, thereby altering their perceptions and preferences (Ares, 2020; Helgason & Rehm, 2023; Langsæther et al., 2022). Closely related to this logic is the Prospect of Upward Mobility (POUM) hypothesis, which suggests that individuals may oppose redistribution not because of their current position, but because they anticipate improving their status in the future (Benabou & Ok, 2001). A second line of explanation draws on the framework of theories of socialization. Within this tradition, hypotheses such as acculturation, socialization, and status maximization propose that individuals adjust their attitudes based on the norms and values of either their class of origin or their destination, or a combination of both (Jaime-Castillo & Marqués-Perales, 2019). However, most of these mechanisms focus primarily on the indirect effects of origin and destination positions, rather than on the direct effect of experience of movement itself. In response to this gap, recent research has highlighted the role of self-serving bias in causal attribution processes, suggesting that individuals tend to explain their mobility trajectories in ways that justify their current position, which in turn shapes their beliefs and preferences (Gugushvili, 2016a; Molina et al., 2019; Schmidt, 2011).

The self-serving bias mechanism builds on a value-oriented perspective, emphasizing that individuals' experiences of social mobility shape their causal attributions, which in turn influence their beliefs about justice and distributive preferences (Gugushvili, 2014). Causal attribution refers to the process through which individuals generate explanations for their own behavior and outcomes, as well as those of others (Gugushvili, 2016b). In this view, people's interpretations of economic inequality depend on whether they believe such disparities reflect unequal individual contributions. Individuals who adopt an internal attribution framework tend to see success or failure as rooted in personal characteristics such as effort, talent, or merit. In contrast, those who regard inequality as unjust are more likely to adopt an external attribution model, viewing outcomes as the result of structural barriers beyond individual control (Kluegel & Smith, 1981). This mechanism, often described as intrapersonal causal attribution, reflects how people explain their own socioeconomic positions—typically attributing their successes to internal qualities while blaming failures on external circumstances (Miller & Ross, 1975). Over time, individuals may revise their beliefs and attitudes: while early views are shaped by their social origin, these are later adjusted in light of personal experiences of mobility and the perceived role of ascribed versus achieved factors in determining socioeconomic outcomes (Gugushvili, 2016b).

Empirical research has provided support for the self-serving bias mechanism in shaping redistributive preferences and attitudes toward the legitimacy of inequality. For instance, Schmidt (2011) finds that individuals who experience upward mobility are more likely to interpret their success as the result of personal effort or merit, and consequently perceive redistribution as less necessary. Conversely, individuals who experience downward mobility tend to attribute their decline to external circumstances—such as structural inequality or unemployment—and show stronger support for redistribution. In a comparative analysis across different welfare domains, Gugushvili (2017) demonstrates that upward mobility is associated with lower support for government spending on housing and pensions, while individuals who experience downward mobility express lower support for healthcare and education spending, but favor increased investment in housing and pensions, reflecting the material nature of these domains. Moreover, Gugushvili (2016a) finds that upward mobility is linked to greater justification of income inequality, suggesting that improved social standing reinforces an attributional view in which success is seen as the result of individual characteristics—thus legitimizing inequality as a fair outcome.

Consequently, considering this theoretical and empirical background, the first hypothesis of this research is that:

$H1$: Experiencing upward (downward) social mobility is positively (negatively) associated with greater support for market justice in healthcare, pensions, and education.

2.3 Meritocracy

Meritocracy constitutes a central ideological framework for legitimizing different types of social inequality, for instance through market justice beliefs. Rooted in the belief that rewards and positions should be allocated based on individual effort and talent, meritocracy operates as a normative ideal and a descriptive belief about how society functions. As initially conceptualized by Michael Young (1958), the term was meant to critique a system in which merit-based stratification becomes a new form of inequality. However, over time, meritocracy has been widely supported in many societies as a fair and desirable principle of distribution, particularly within liberal democracies and market-oriented economies (Mijs, 2019; Sandel, 2020). From a sociological perspective, the belief in meritocracy is more than a cognitive assessment; it reflects a moral lens through which individuals interpret inequality. People who believe that success results from hard work and talent are more likely to view social and economic disparities as legitimate (Batruch et al., 2023; Castillo et al., 2019). Conversely, if they see outcomes as driven by luck, social origin, or systemic barriers, inequality is more likely to be perceived as unjust. This distinction becomes crucial in societies with persistent structural inequality, where public narratives often emphasize personal responsibility and merit while overlooking entrenched disadvantages.

I adopt a multidimensional perspective on meritocracy, distinguishing between two key dimensions: effort-based and talent-based perceptions. This distinction is essential, as it captures different pathways through which individuals justify inequality (Young, 1958). Effort-based meritocracy emphasizes hard work and perseverance as the basis for social rewards, aligning closely with cultural narratives of personal responsibility. A talent-based meritocracy, by contrast, emphasizes intelligence and innate abilities, which are often perceived as less malleable and more unequally distributed. Both dimensions have been shown to correlate with acceptance of inequality, but they may carry distinct implications for how inequality is justified in specific domains (Castillo et al., 2023). The relevance of this distinction is supported by recent studies, which show that individuals respond differently to these dimensions. For instance, perceptions that effort is rewarded in society are more strongly associated with positive evaluations of fairness and acceptance of unequal outcomes (Batruch et al., 2023; Wiederkehr et al., 2015; Wilson, 2003). This may be because effort is seen as a controllable and morally virtuous trait, whereas talent is often perceived as a natural advantage. Consequently, effort-based meritocracy is likely more potent in legitimizing inequality, particularly in neoliberal contexts.

These dimensions of meritocracy reflect how respondents perceive society's distributive logic, regardless of whether they endorse meritocratic principles. This distinction aligns with recent findings indicating that individuals distinguish between how merit is perceived in society and how it should ideally operate, which in turn shapes their preferences for redistribution and justice (Tejero-Peregrina et al., 2025). Meritocratic beliefs serve as symbolic justifications for unequal outcomes, particularly when access is stratified by income or social opportunity. Prior studies in Chile have shown that individuals who perceive higher levels of meritocracy tend to express stronger support for unequal distributions that reflect market outcomes in healthcare, education and pensions (Castillo, Laffert, et al., 2025; Castillo et al., 2024).

In addition to influencing individual attitudes toward inequality, meritocratic beliefs can contribute to social division and the stigmatization of disadvantaged groups. Recent research has demonstrated that exposure to meritocratic narratives can reinforce the belief that poverty results from individual failings rather than systemic conditions, reducing support for redistributive measures and increasing the stigmatization of the poor (Hoyt et al., 2023). This reinforces negative stereotypes and reduces empathy toward individuals from lower socioeconomic backgrounds. Moreover, Busemeyer et al. (2021) argues that meritocratic narratives can serve as feedback mechanisms that shape public opinion and well-being by framing individuals' understanding of welfare outcomes as deserved or undeserved within existing institutional structures. This psychological mechanism highlights the normative power of meritocracy in stabilizing unequal systems by shaping political attitudes and personal perceptions of success and failure.

Importantly, recent research has explored how meritocratic beliefs interact with experiences of intergenerational mobility to shape distributive attitudes. The belief that one's success is earned can lead upwardly mobile individuals to internalize meritocratic narratives and justify existing inequalities, reinforcing support for market justice (Gugushvili, 2016a; Molina et al., 2019). Conversely, downwardly mobile individuals who maintain strong meritocratic beliefs may interpret their status as a personal failure, reducing their support for redistribution (Day & Fiske, 2017). At the macro level, Shariff et al. (2016) show that higher perceived mobility increases tolerance for inequality, suggesting that meritocracy and mobility are mutually reinforcing.

Taken together, this literature supports the idea that meritocratic beliefs can moderate the relationship between mobility and market justice preferences. Individuals who experience mobility—especially upward—may draw on meritocratic narratives to legitimize both their own status and broader inequalities, thereby strengthening their support for market-based distribution. Accordingly, the second hypothesis of this study is:

$H2$: The positive (negative) relationship between upward (downward) social mobility and support for market justice in healthcare, pensions, and education is moderated by meritocratic beliefs; specifically, this association is stronger (weaker) among individuals with higher perceptions of meritocracy.

2.4 The Chilean context

Chile offers a compelling case for examining individual preferences for market justice in the access to core social services. Since the early 1990s, the country experienced a long period of sustained economic growth, with per capita GDP increasing by nearly 4% annually until the mid-2010s (Llorca-Jaña & Miller, 2021). This growth brought notable improvements in living standards, including a strong decline in income poverty—from 38.6% in 1990 to 6.5% in 2022 (MIDEPLAN, 2017). However, this success story has been tempered by growing concerns over the unequal distribution of its benefits. In recent years, economic growth has slowed—exacerbated by the global impact of the COVID-19 pandemic—and structural inequalities have become more visible (Barozet et al., 2021). Despite its economic achievements, Chile remains one of the most unequal countries in Latin America and the OCDE. The poorest 50% of the population captures just 10% of total income and holds negative net wealth, while the wealthiest 1% receives nearly 27% of all income and controls almost half (49.6%) of the country's wealth (Chancel et al., 2022), a figure that has seen little change since the 1990s (Flores et al., 2020). These stark disparities extend beyond income, reflecting deep inequalities in access to essential social services (PNUD, 2017).

A significant share of Chile's inequality stems from neoliberal reforms that privatized and commodified key domains of social reproduction (Arrizabalo, 1995; Ferre, 2023). Introduced during the military dictatorship (1973–1989) and expanded under democratic governments, these reforms embedded market logic into public services through concessions, subsidies, demand-side vouchers, and pro-private regulatory frameworks (Boccardo, 2020; Madariaga, 2020). This "crowded-out" welfare model benefits higher-income groups, leaving lower-income individuals to rely on limited public options. Scholars argue that this neoliberal shift repurposed the state to create niches of capitalist accumulation -from water to education-, giving rise to a model of "public-service capitalism" heavily reliant on state funding (Boccardo, 2020). In healthcare, around 79% of the population is covered by the public insurer FONASA, while 15% are enrolled in private ISAPREs, which offer faster and higher-quality services to wealthier groups (MIDEPLAN, 2024). The pension system is based on individual capitalization, with mandatory contributions managed by private administrators investing in the financial market, currently involving 11 million contributors through private fund administrators (AFPs), yet 27% of the labor force remains outside the system due to informality (Superintendencia de Pensiones, 2024). Education is similarly stratified: only 30.6% of students attend fully public schools, while 54% are in voucher-funded private institutions and 9.3% in fully private schools that largely serve affluent families (Ministerio de Educación, 2023). This arrangement has created a structurally segmented system, where access and quality of services depend on ability to pay, reinforced by state subsidies to private providers (PNUD, 2017).

Over recent decades, Chile has undergone a significant transformation in its class structure, shaped by sustained economic growth, the expansion of the service sector, the massification of education, and major changes in the role of the state and social policy (Ruiz & Boccardo, 2014). These changes contributed to the growth of the middle sectors, a phenomena labelled as "mesocratrization" of society (Espinoza et al., 2013). However, this expansion did not reduce structural inequalities nor facilitate upward mobility into elite positions (Espinoza et al., 2013; López-Roldán & Fachelli, 2021). While Chile shows high absolute mobility, most flows occur between adjacent classes—particularly from working classes into lower service-class positions—while access to upper classes remains limited, indicating strong elite closure and reproduction (Espinoza et al., 2013; Espinoza & Núñez, 2014; López-Roldán & Fachelli, 2021; Pérez-Ahumada, 2019; Torche, 2005). Though often described as "unequal but fluid", recent evidence points to growing rigidity in relative mobility, especially after the economic slowdown, increased household debt, labor market precarity, and the COVID-19 crisis (Barozet et al., 2021). However, the majority of the population identifies themselves as members of the middle class (Castillo et al., 2013). These dynamics have fueled a legitimacy crisis: meritocracy is increasingly questioned, and many perceive that advancement depends more on informal networks than on effort or earned credentials (Barozet et al., 2021). The combination of constrained mobility, persistent inequality, and concentrated opportunity has eroded expectations of upward mobility and intensified social discontent (PNUD, 2017).

Despite recurrent social unrest, Chile presents a paradoxical coexistence between strong conflict over inequality and widespread public legitimation of it. The October 2019 "social outburst", a period of mass protests and severe political repression throughout the country, crystallized discontent over the privatization and commodification of social services, alongside a crisis of political legitimacy (Somma et al., 2021). A survey during the protests identified pensions, healthcare, and education as the top demands (Núcleo de Sociología Contingente, 2020), reflecting dissatisfaction with Chile's stratified welfare regime. Yet, this unrest coexists with strong meritocratic beliefs and high tolerance for income inequality. Castillo (2011) shows that wage gaps between occupations are consistently justified, and the larger the perceived gap, the greater its legitimation, an effect amplified by Chile's structural inequality. Moreover, perceptions of unfairness due to non-meritocratic factors often reinforce belief in effort and talent as legitimate bases for success (Castillo et al., 2023), which are associated with lower perceived injustice (Castillo et al., 2019). Mac-Clure et al. (Mac-Clure et al., 2024) find that low and lower-middle status individuals are less likely than upper-status groups to view educational income differences as unfair. Qualitative research further reveals the internalization of the enterprising self (Mau, 2015), understood as a way in which the subjectivization process arises in line with positioning the self in the market, generating its own market value and subjecting it to competition. Canales et al. (2021) describes how families navigate school choice as investment decisions, while Panes (2020) finds that many workers frame pension contributions as individual investments aimed at maximizing future returns. These representations suggest that market-based moral orientations are deeply embedded in Chilean subjectivities, offering fertile ground to study how support for market justice varies across welfare domains.

Although still limited, recent research on market justice preferences in Chile indicates that support for these beliefs has increased in recent years across domains such as healthcare, pensions, and education, and that both objective and subjective dimensions of inequality shape them. Castillo et al. (2025) show that agreement with the notion that it is fair for higher-income individuals to access better services was relatively low in 2016, but rose significantly by 2023—especially in the pension domain. Otero and Mendoza (2024) provide evidence that social class, measured using the EGP scheme, influences these preferences: individuals in lower or subordinate class positions are less supportive of market-based access to welfare services compared to those in higher or privileged positions. Additionally, greater diversity in social class networks is associated with lower support for market justice. This is particularly relevant in the Chilean context, where one of the main determinants of class-diverse networks is intergenerational social mobility, which exposes individuals to relationships with people from different class backgrounds (Otero et al., 2022). Regarding subjective factors, Castillo et al. (Castillo, Laffert, et al., 2025; Castillo et al., 2024) find that individuals who endorse meritocratic beliefs—such as the idea that effort and talent are rewarded in Chile—are more likely to support market justice across welfare domains.

3 Method

Social mobility effects—understood as impacts on an outcome arising from movements between an origin state and a destination state (e.g., social class)—have long been a focus of sociological research (Eyles et al., 2022; Langsæther et al., 2022). Yet, as Breen and Ermisch (2024, p. 467) emphasize, most of the mobility hypotheses are, at their core, individual-level counterfactual comparisons between the observed outcome under a mobility trajectory and the outcome the same person would have shown if they had remained in their origin class (or moved to an alternative destination). Standard specifications (e.g., SAM, DRM, and mobility-contrast models) struggle to retrieve these inherently counterfactual quantities due to identification constraints and their reliance on between-group contrasts constructed from additive terms and interactions, thereby yielding primarily associational evidence (Breen & Ermisch, 2024; Song & Zhou, 2025). Following the causal framework of Breen and Ermisch (2024), I conceive of the destination class as a treatment, conditioning on the origin class and estimating the heterogeneous effects of the destination with observational data under explicit identification assumptions. To align design and target, I follow the MIDA template (Blair et al., 2023): set out the causal model and assumptions (M), define the inquiry and estimand (I), describe the data and variables (D), and detail the answer strategy (A) used to identify the causal effect of intergenerational occupational mobility on preferences for the commodification of pensions.

3.1 A Causal Model for Mobility Effects

Breen and Ermisch (2024), building on a critical reassessment of the inferential limits of standard mobility models, propose a causal framework that reconceptualizes the "mobility effect" as the treatment effect of reaching a destination class, with impacts heterogeneous by class of origin. The core claim is that typical mobility hypotheses are fundamentally within-person counterfactual comparisons rather than between-person contrasts. Defining mobility as $M = D - O$; a change from origin (O) to destination (D), rather than an additive or interactive combination, places the focus squarely on the Neyman–Rubin problem: for any individual, we observe the result in the social position they currently occupy, but not in the situation of other alternative destinations or immobility.

Exploiting the temporal ordering of origin, destination and outcome, Breen and Ermisch (2024) frame mobility as a treatment process that motivates the causal question: "how would the outcome among people from origin *j* who entered destination *k* have been different if those people had, counterfactually, entered destination *k'* instead?" (p.472). The corresponding estimand is the conditional causal effect of destination given origin. It is of particular interest when *k'* = *j*, i.e., immobility. This estimand compares movers' observed outcomes with the hypothetical outcomes those same individuals would have exhibited had they remained in their origin class. This formulation provides a coherent potential-outcomes basis for studying how social mobility shape individual preferences and attitudes.

Following Breen and Ermisch (2024), the identification of causal mobility effects from observational data requires the following assumptions:

Positivity: For all (j, c) in the support of (O, C) and for each relevant destination (k), the probability of receiving $D = k$ is strictly between 0 and 1; substantively, each type (O, C) has a nonzero probability of entering each comparison destination.

$$0 < P(D = k \mid O = j, C = c) < 1 \tag{1}$$

Stable Unit Treatment Value Assumption (SUTVA): each unit's potential outcomes $Y_i(D)$ does not depend on the mechanism used to assign treatments (destinations) and by the treatments assigned to other units (also called no interference), assuming a single, well-defined version of each treatment (consistency).

$$Y_i(D_i) = Y_i(D_i, D_{-i}) \quad \forall, D_{-i} \tag{2}$$

Conditional Independence: Also known as conditional unconfoundedness or exchangeability, this assumption emphasizes that, conditional on origin O and pre-treatment covariates C (e.g., parental education, ethnicity, cohort, early-life factors), assignment to D is as good as random. This underlies IPW and regression adjustment, which seek exchangeability between mobile (treated) and immobile individuals (controls).

$$Y(D) \perp\!\!\!\perp D \mid (O, C) \tag{3}$$

Establishing the causal relationship between objective social mobility and subjective outcomes, such as preferences, poses several challenges for causal inference. Below, I use directed acyclic graphs (DAGs) to illustrate some of these challenges and evaluate possible strategies for identifying the effects of social mobility. Figure 1 depicts the causal model guiding identification. In this model, a respondent's preference Y is directly affected by her class destination D, and indirectly influenced by a set of pre-treatment attributes C rooted in childhood—family resources, household composition, parental education, and other ascriptive characteristics. Class origin O is itself shaped by these background factors C and, in turn, affects destination D. I allow for unobserved determinants U that influence background factors C, but I assume that any such unobserved variation operates only through C. Thus, there are no direct paths $U \rightarrow D$ or $U \rightarrow Y$ beyond those mediated by C.

Under this structure, conditioning on (O, C) blocks all relevant backdoor paths from D to Y—notably $D \leftarrow O \leftarrow C \rightarrow Y$. This renders (O, C) a minimal sufficient adjustment set for identifying the causal effect of mobility on preferences, in line with the framework proposed by Breen & Ermisch (2024). Identification therefore relies on the assumption that background factors C adequately summarize pre-treatment characteristics that jointly shape both mobility and attitudes.

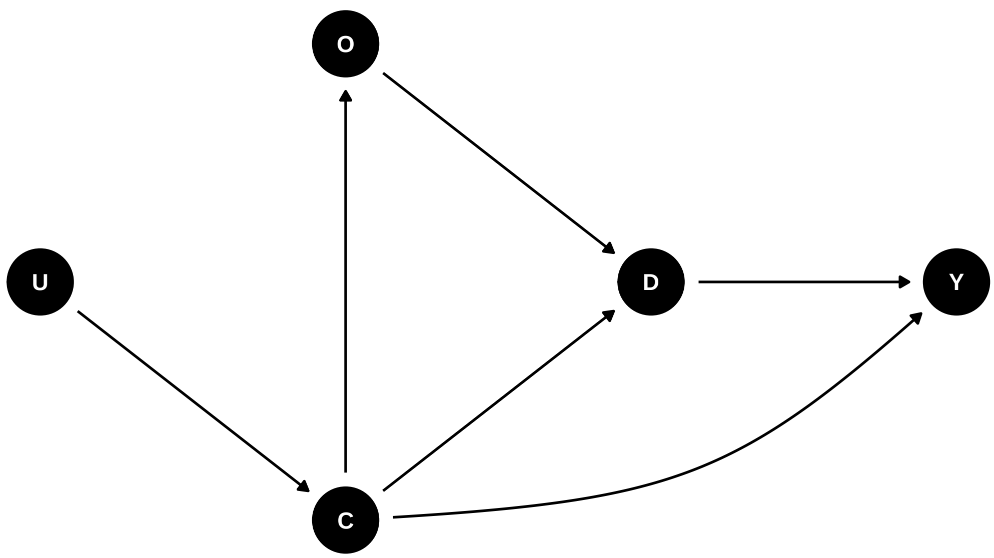

Figure 1: Causal graph of the effect of intergenerational social mobility on preferences for commodification. Y = preferences for pension commodification, D = an individual's class of destination, O = an individual's class of origin, C = different attributes determined in childhood or earlier that affects an individual's class of origin and destination. Finally, U = unobserved factors influencing origin attributes.

Why are the assumptions plausible here? Conditional independence is targeted by controlling for a plausible sufficient adjustment set (O, C) and by using inverse probability weights obtained via entropy balancing (Hainmueller, 2012), which enforce covariate balance within each origin stratum and mitigate selection on observables. Positivity is assessed by inspecting the distribution of the entropy-balancing weights and trimming observations with extreme weights, thus restricting inference to regions of common support. SUTVA holds by treating the destination class as a single exposure measured before (Y) and assuming no interference. Together, the DAG and assumptions define conditions for identifying the causal effect of intergenerational social mobility on preferences for commodification.

3.2 Inquiry and estimand

The causal inquiry guiding this study asks: How does intergenerational social mobility affect individuals' preferences for the commodification of pensions in Chile? Substantively, I am interested in the effect of mobility relative to immobility within a given class of origin.

Following the potential outcomes framework proposed by Breen and Ermisch (2024, p. 473), the estimand of interest is an average treatment effect on the treated (ATT) defined within each origin class. I focus on the specific case in which the counterfactual destination corresponds to class immobility $(k' = j)$. In other words, I ask how the outcome among people from origin *j* who entered destination k would have bee different if those people had, counterfactually, remained in origin *j* instead (inmobility).

Formally, for individuals with origin $(O = j)$ who attain destination $(D = k)$, the estimand is:

$$
ATT_{j,k,j} = E\left[Y(D = k) \mid O = j, D = k\right] - E\left[Y(D = j) \mid O = j, D = k\right]
$$
{#eq-att}

which represents the mean causal effect of moving from origin class (j) to destination class (k), compared to the counterfactual outcome those same individuals would have exhibited had they instead remained in (j). Thus, this ATT captures the counterfactual contrast within origin that lies at the heart of mobility effects: how preferences for market-based pensions would differ if, for the same upwardly or downwardly mobile individuals, their observed destination class were replaced by immobility in their origin class.

3.3 Data and variables

3.3.1 Data

This study draws on data from the Chilean Longitudinal Social Survey (ELSOC) of the Center for Social Conflict and Cohesion Studies (COES). This survey is a nationally representative panel study of the urban adult population in Chile, conducted annually between 2016 and 2023. Designed to examine individuals' attitudes, emotions, and behaviors regarding social conflict and cohesion, ELSOC employs a probabilistic, stratified, clustered, and multistage sampling design covering both major urban centers and smaller cities. The sampling frame was proportionally stratified into six categories of urban population size (e.g., large and small cities), followed by a random selection of households within 1,067 city blocks. The target population includes men and women aged 18 to 75 who are habitual residents of private dwellings.

Because respondent occupation[1] is not measured in every wave, I restrict the analysis to the 2016, 2018, and 2023 waves. After listwise deletion and restricting to key variables (respondent occupation and the outcome measure), the final analytic sample comprises $N = 3{,}435$ observations nested within $N = 1{,}787$ individuals (2016: 914; 2018: 1,377; 2023: 1,144). Consistent with the study design, estimation proceeds within trajectory-specific subsamples (e.g., low→low; low→middle), retaining movers

[1] Parental occupation was collected only in 2023 as open-text labels. I coded these texts to the ISCO-08 two-digit scheme using the National Institute of Statistics of Chile's automated coding API and retained cases with high-confidence matches (>95%). Treating parental occupation as a time-invariant origin attribute, I carried these codes back to 2016 and 2018.

and matched non-movers from the same origin class. Then, I apply weights to construct the corresponding counterfactual control sample for each trajectory. Following Breen and Ermisch (2024), I use the three waves in the estimations that follow, allowing for correlation between the observations for each individual in calculating the standard errors (i.e. clustering on the personal identifier).

3.3.2 Outcome variable

The outcome variable measures preferences regarding the commodification of pensions, operationalized with a single item addressing how strongly individuals justify conditioning access to old-age pension benefits based on individual income. Respondents were asked: "Is it fair in Chile that people with higher incomes have better pensions than people with lower incomes?" and answered on a five-point Likert scale ranging from 1 ("strongly disagree") to 5 ("strongly agree"). For estimation and interpretability, I construct a binary indicator in which values 4–5 indicate agreement with market-based access to old-age pensions (coded 1), while values 1–3 indicate non-agreement (coded 0). Substantively, the dependent variable is meant to capture *endorsement* of pension commodification: only responses in the upper tail of the scale represent an active, explicit justification of market-based differentiation, whereas neutrality corresponds to an absence of such endorsement and is therefore treated as closer to non-support than to support.[2] This binary coding thus yields a clear contrast between respondents who *support* market-based differentiation in pensions and those who do not. I use this item for two main reasons: first, to enable comparisons with existing work on market-based justice in social policy (Castillo, Laffert, et al., 2025; Lindh, 2015; Otero & Mendoza, 2024); and second, because it taps two core dimensions of market-based welfare distribution—(i) the centrality of economic resources as a criterion for allocating outcomes and (ii) the framing of pensions as tradable commodities that can be bought and sold according to ability to pay (Lindh, 2015).

3.3.3 Treatment

3.3.3.1 Intergenerational occupational mobility

I treat intergenerational occupational mobility as an exposure indicating whether respondents occupy a different occupational status than their fathers, closely following Breen and Ermisch's (2024) causal framework. Occupational assignment proceeds in two steps. First, I derive occupational status for both origin (father) and destination (respondent) from two-digit ISCO-08 codes using the International Socio-Economic Index of Occupational Status (ISEI). Second, I group these ISEI scores into terciles (low, middle, high), yielding a three-category schema for both origin and destination strata (see Table 1). Substantively, I use ISEI because occupational status is a core indicator of stratification and a reliable mobility measure: it locates jobs on a hierarchical continuum defined by incumbents' typical education level and earnings (Hauser, 2010; Salgado et al., 2025), and thus approximates long-run socio-economic position more closely than volatile, single-year income, which is also prone to recall and reporting error for both parental and own resources (Barone et al., 2022). This occupational focus is consistent with a long tradition that interprets the occupational structure as the backbone of the stratification system and a key determinant of life chances (Wright, 2015). Moreover, ISEI was explicitly designed to harmonize occupational stratification across class schemes and SES measures (Ganzeboom & Treiman, 1996), supporting cross-study comparability. In short, using ISEI aligns

[2] A potential concern is that the neutral category may reflect genuine ambivalence rather than disagreement. To address this, I conduct robustness checks using alternative codings: (i) excluding neutral respondents from the analysis; (ii) combining the neutral and agreement categories (3–5 vs. 1–2); and (iii) estimating models with the original 5-point scale using linear specifications. The main mobility effects reported below are substantively unchanged across these specifications (see Supplementary Material).

Table 1: Occupational mobility by occupational groups.

Father↓	Offspring→	Low	Middle	High	Total
Low		44.7% (501)	33.6% (377)	21.7% (244)	100.0% (1,122)
Middle		35.1% (388)	34.9% (385)	30.0% (331)	100.0% (1,104)
High		24.6% (297)	28.8% (348)	46.7% (564)	100.0% (1,209)
Total		34.5% (1,186)	32.3% (1,110)	33.2% (1,139)	100.0% (3,435)

the measurement of origin and destination on a common vertical status continuum that is theoretically meaningful for social mobility research and empirically feasible given the available information of ELSOC's data.

3.3.3.2 Pre-treatment control variables for selection into treatment

Within each origin stratum ($O = j$), I estimate average treatment effects on the treated (ATT) by comparing movers to otherwise similar non-movers from the same origin. To reduce bias from non-random selection into mobility based on observed characteristics, I preprocess the data using entropy balancing, reweighting only the control group (the immobile) so that the distribution of pre-treatment origin covariates C among non-movers matches that of movers on selected moments. Entropy balancing implements a maximum-entropy reweighting scheme that chooses unit weights for non-movers to satisfy a set of balance constraints (e.g., equality of means and, where relevant, higher moments), while keeping the new weights as close as possible to the original weights (Hainmueller, 2012). Substantively, this procedure addresses selection on observables: it reweights immobile respondents so that, within each origin class, they resemble mobile respondents in their observed background characteristics, making any remaining differences in the outcome more plausibly attributable to mobility rather than to pre-existing measured advantages.

The selection of the adjustment set C follows two principles. First, it mirrors the logic in Breen and Ermisch (2024), who condition on attributes determined in childhood or earlier (e.g., parental resources, household structure, ascriptive traits). Second, it is constrained to origin-side characteristics available in ELSOC. Guided by theory and Chilean evidence on the determinants of relative mobility (mainly Brunori et al., 2025; Espinoza & Núñez, 2014; Salgado et al., 2025; Torche, 2005), I include six covariates covering family resources, household composition, and ascriptive status: (a) parental education (highest of father/mother), in 10 ordinal categories from no schooling to postgraduate studies; (b) co-residence with both parents at age 15 (0 = no, 1 = yes); (c) nationality (0 = non-Chilean, 1 = Chilean); (d) age in years; (e) sex (0 = male, 1 = female); and (f) indigenous ethnicity (0 = no, 1 = yes). This set of covariates is fixed prior to the destination class and constitutes the maximum set on the origin side in ELSOC, thus allowing for the plausible capture of both family and contextual influences that may affect career trajectories, making the conditional independence hypothesis more credible within each $O = j$. Descriptive statistics for all variables are presented in Supplementary Material Table A1.

After achieving balance, I construct stabilized IPW-ATT weights by setting treated (mobile) cases to weight 1 and assigning the entropy-balancing weights to controls, rescaled within each origin stratum so that the mean weight remains close to one. This aligns the counterfactual distribution of non-movers with the covariate profile of movers in each origin class (see Supplementary Material for balance diagnostics). The resulting weights are then used in the outcome stage to estimate the causal effect of intergenerational mobility on preferences for the commodification of pensions.

3.3.4 Effect heterogeneity

To test heterogeneity in mobility effects, I examine two moderators: meritocratic perceptions and time. Meritocracy is captured with two items (Young, 1958), one referring to effort ("In Chile, people are rewarded for their efforts") and one to talent ("In Chile, people are rewarded for their intelligence and skills"), each answered on a five-point Likert scale (1 = "strongly disagree" to 5 = "strongly agree"). I average the two items into a single index and dichotomize it into low meritocracy (≤ 3) and high meritocracy (≥ 4). Time is measured by survey wave (2016, 2018, 2023) and entered as a categorical (dummy) variable. Both variables are used as moderators by interacting them with the mobility treatment to assess conditional average treatment effects—i.e., whether the effect of mobility varies across levels of meritocratic perceptions or across survey waves.

3.3.5 Controls

All models include the same pre-treatment covariates C used in the IPW construction, not to re-balance groups (entropy balancing already does so), but to (i) block backdoor paths from C to the outcome Y as implied by the DAG, and (ii) achieve double robustness: estimates remain consistent if either the weighting model or the outcome model is correctly specified. Concretely, I adjust for (a) father's educational level, (b) co-residence with both parents at age 15, (c) nationality, (d) age, (e) sex, and (f) ethnicity, along with wave fixed effects to absorb secular trends.

3.4 Analytical strategy

Because the dependent variable—the preference for pension commodification—is binary ($Y_{it} \in 0, 1$), I estimate weighted linear probability models (WLS) using stabilized inverse probability weights (IPW) obtained via entropy balancing. Although OLS is often questioned with binary outcomes, it consistently estimates the conditional mean $E(Y|X)$ under standard exogeneity $E[\varepsilon_i|X_i] = 0$, and heteroskedasticity can be handled with robust or clustered standard errors (Wooldridge, 2009, ch. 7.5). WLS with IPW further improves efficiency and implements the ATT estimand by recreating the counterfactual distribution of non-movers (Gelman et al., 2020, pp. 270–272). Linear models are preferred here because coefficients are directly interpretable as average differences in predicted probabilities, whereas non-linear models complicate both weighting and interpretation (Gelman et al., 2020, ch. 13).

Formally, the model is:

$$Y_{it} = \alpha + \beta T_i + X_i' \gamma + \lambda_t + \varepsilon_{it} \tag{4}$$

where Y_{it} indicates whether individual i in wave t supports more market-based pension access; α is the baseline probability for the reference categories; βT_i captures the intergenerational mobility contrast: $T_i = 1$ when the observed destination is $D_i = k$ (mobile to k) and $T_i = 0$ when $D_i = j$ (immobile), so that β is the ATT within origin $O = j(\hat{\beta} = \widehat{ATT} j, k \mid j)$, i.e., the percentage-point change in the probability of preferring pension commodification from reaching k rather than remaining in j. The term $X_i' \gamma$ includes the pre-treatment covariates (C) used to construct the entropy-balancing weights; λ_t are wave fixed effects; and ε_{it} is idiosyncratic error. The analytic sample is restricted to individuals sharing the same origin $(O = j)$. Therefore, estimation uses pooled WLS with stabilized IPW–ATT from entropy balancing and CR2 standard errors clustered by individual to address heteroskedasticity and within-person dependence.

The specification is doubly robust: entropy-balancing weights are combined with covariate adjustment so consistency holds if either the weighting model or the outcome model is correctly specified, while also improving efficiency (Wooldridge, 2009). These results are virtually identical to baseline models without covariates (see Supplementary Material for complete models), with no substantive changes in the mobility coefficients. Heterogeneous effects by meritocratic perceptions and survey wave are examined through interaction models described in the previous section.

To evaluate the robustness of the findings, I implement two sets of sensitivity analyses. First, I conduct standard robustness checks by re-estimating the models under alternative codings of the outcome: (i) excluding neutral responses from the analysis, (ii) combining neutral responses with agreement (3–5 vs. 1–2), and (iii) using the original 5-point item in a linear specification. Second, I assess sensitivity to unmeasured confounding using the omitted-variable–bias framework of Cinelli and Hazlett (2020), which quantifies how strong an unobserved confounder would need to be to explain away the estimated ATT.

4 Results

4.1 Descriptive statistics

Figure 2 shows the annual frequencies of preferences for pension commodification in 2016, 2018, and 2023. Each year presents stacked percentages frequencies for the level of agreement and disagreement. Overall, a large majority rejects market-based access to old-age pension benefits, but this opposition has eased over time: 83.5% in 2016, 81.2% in 2018, and 71.8% in 2023. The 2016–2018 change is modest, whereas 2023 registers a 9.4 point drop in disagreement relative to 2018, mirrored by rising agreement: 16.5% (2016), 18.8% (2018), 28.2% (2023). Substantively, while most respondents continue to oppose the idea that higher-income individuals should obtain better pensions via the market, a non-trivial and growing portion endorses this statement, with the sharpest expansion concentrated in the latest wave (+9.4 points from 2018 to 2023).

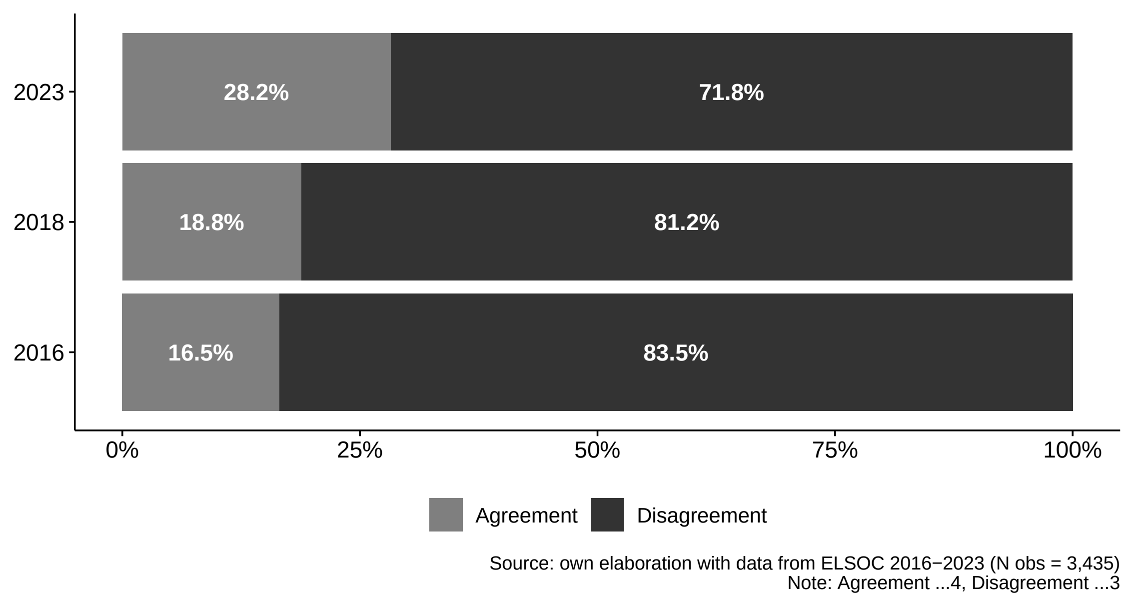

Figure 2: Change in preferences for pension commodification over time (2016, 2018, and 2023).

Regarding the relationship between preferences for the commodification of pensions and intergenerational occupational mobility, Figure 3 shows the percentage of agreement with market-based access to pension benefits according to mobility trajectories and survey waves. Averaging across waves, agreement is highest among the High-High immobile (27.8%), followed by the Middle-High upwardly mobile (24.6%) and the High-Middle downwardly mobile (24.0%). The lowest levels are found among ascending Low-High (17.3%), and immobile Low-Low (17.4%) groups. The specific profiles of each wave accentuate these gradients in most trajectories, especially in 2023: agreement reaches 31.9% for High-High, 29.4% for upward Middle-High, and peaks at 36.8% for downward High-Middle (the highest of all trajectories for that year). Descriptively, immobility at the top is associated with greater support for commodification; among those who move, support is greatest for the upward Middle-High trajectory and the downward High-Middle trajectory, patterns that intensify in 2023. These shifts anticipate the heterogeneity documented below.

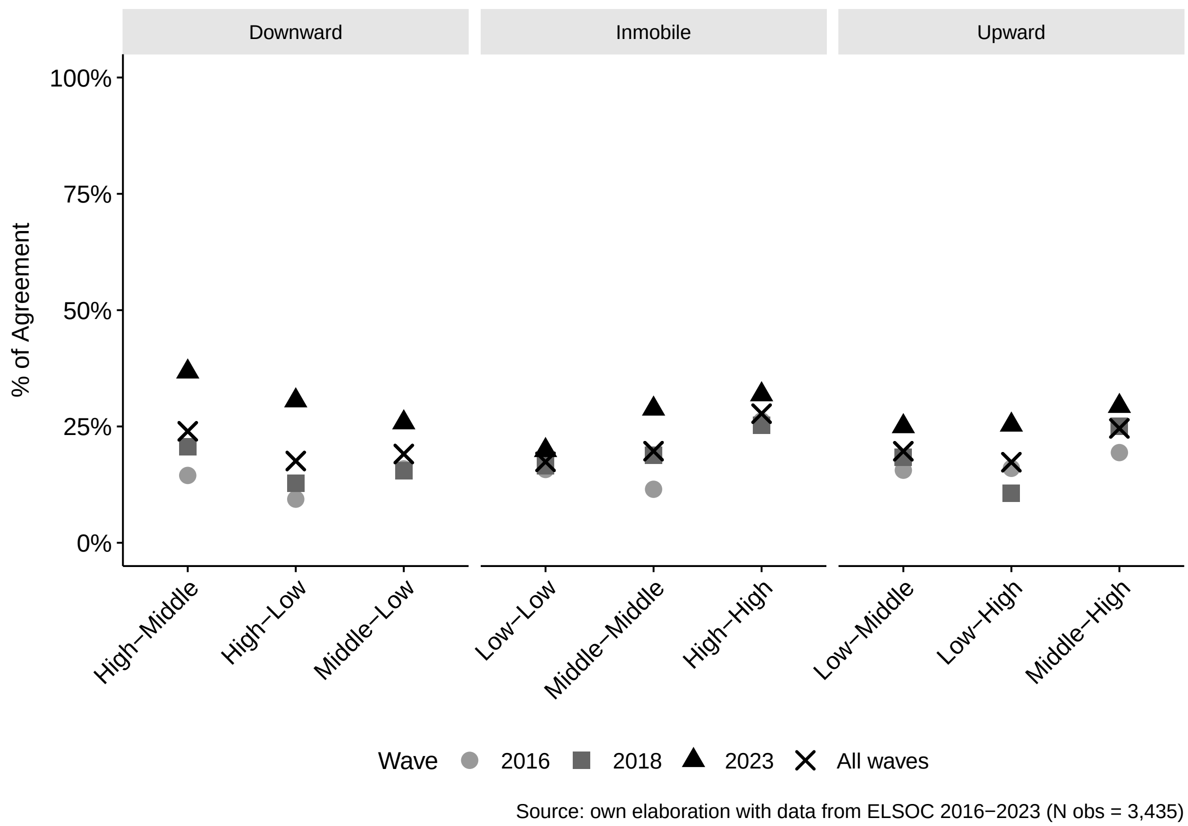

Figure 3: Percentage agreement on preferences for pension commodification according to mobility trajectory and time (2016, 2018, and 2023).

4.2 Mobility effects models

Figure 4 reports the double-robust estimates of the effect of intergenerational occupational mobility on preferences for pension commodification, obtained from weighted models that combine stabilized IPW with covariate adjustment and fixed wave effects. Results reveal a clear directional asymmetry between upward and downward mobility trajectories. Among upwardly mobile respondents, only the Middle→High trajectory exhibits a significant positive effect on support for pension commodification (β = 0.09, 95% CI [0.02, 0.16]). Interpreted as an origin-specific ATT, this estimate compares—for individuals who actually moved from the middle to the high socioeconomic status—their observed endorsement with the counterfactual endorsement they would have expressed had they remained immobile in the middle socioeconomic status. Holding all other predictors constant, the point estimate implies a 9 percentage point higher probability of endorsing market-based access to pension benefits for Middle→High movers relative to their own immobility counterfactual, statistically significant at the 95% confidence level. In contrast, two downward trajectories: High→Middle (β = −0.07, CI [−0.15, −0.00]) and High→Low (β = −0.10, CI [−0.18, −0.02]), show significant negative effects, suggesting that individuals who experience downward mobility from higher-status origins express weaker preferences for market-based pension benefits. The remaining pathways (Low→Middle, Low→High, Middle→Low) display no significant differences relative to their non-mobile counterparts. These patterns indicate that upward movement within the upper segment (Middle→High) reinforces pro-commodification attitudes, whereas downward movement from privileged origins reduces them.

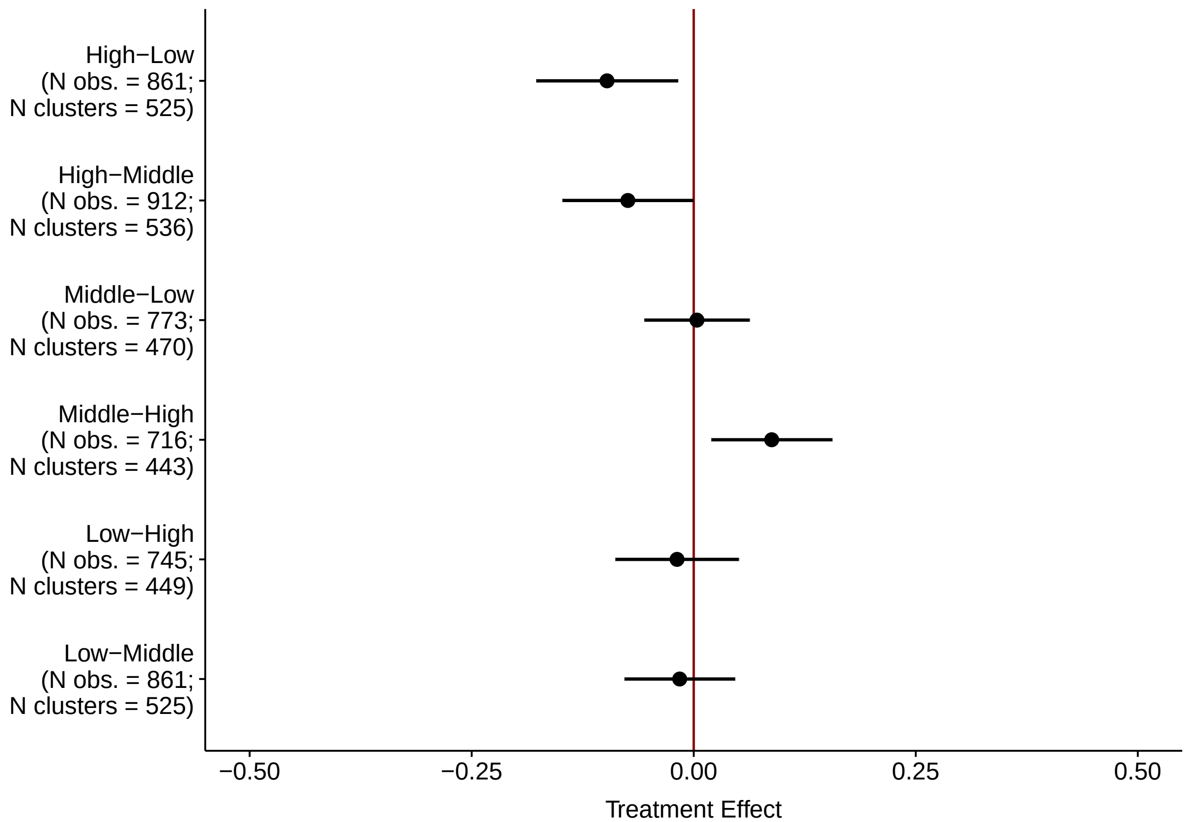

Figure 4: Effects of intergenerational occupational mobility on preferences for pension commodification. Coefficient plot of origin-specific ATT estimates comparing mobility versus immobility. Estimates from pooled WLS with doubly robust adjustment, wave fixed effects, and standard errors clustered by individual; bars show 95% confidence intervals.

4.3 Effect heterogeneity

To probe potential mechanisms underlying the mobility effects, I estimate origin-specific ATT models that interact the mobility treatment with high meritocratic perceptions (ref. = low), using stabilized IPW weights, covariate adjustment, and wave fixed effects. I also estimate an analogous specification interacting mobility with survey waves (ref. = 2016) to assess temporal heterogeneity in the treatment effect (see Supplementary Material for complete models).

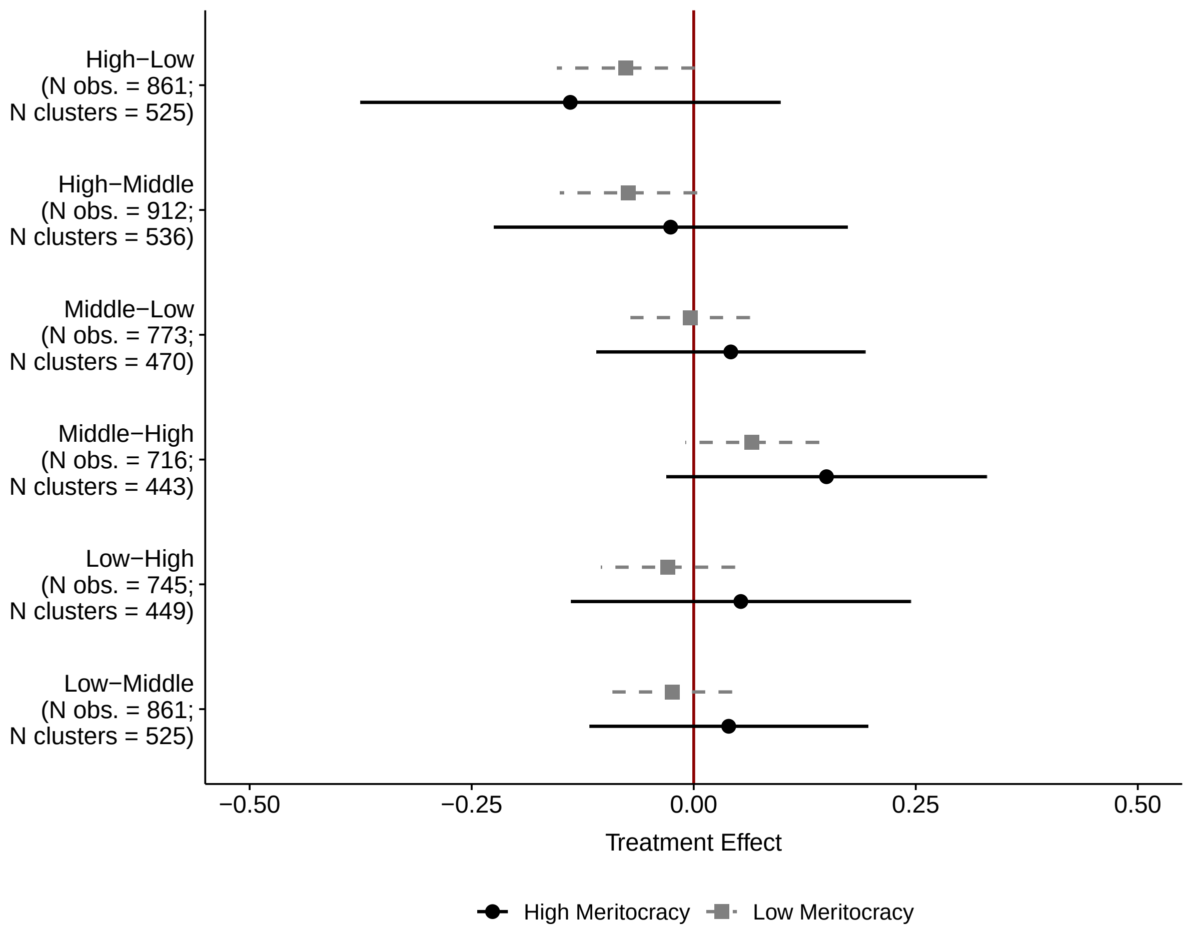

Figure 5: Effects of intergenerational occupational mobility on preferences for pension commodification, by meritocratic perception. Coefficient plot of origin-specific ATT estimates comparing mobility versus immobility, estimated from interaction models (T × Merit). Points show simple slopes of the mobility effect at low merit and high merit; bars are 95% confidence intervals. Estimates from pooled WLS with doubly robust adjustment, wave fixed effects, and standard errors clustered by individual.

Figure 5 shows that neither the interaction terms nor the mobility coefficients reach conventional significance across trajectories, indicating no detectable moderation of the causal mobility effect by meritocratic perceptions at the 95% level. Directionally, in the left panel, the interaction is positive for Middle→High ($\beta = 0.15$, 95% CI [−0.03, 0.33]) and Low→High ($\beta = 0.05$, [−0.14, 0.24]), suggesting, if anything, somewhat larger pro-commodification effects among high-merit respondents when upwardly mobile, whereas it is negative for High→Low ($\beta = -0.14$, [−0.38, 0.10]) and High→Middle ($\beta = -0.03$, [−0.23, 0.17]), showing weaker pro-commodification effects under downward mobility at high merit. Focusing on the treatment effect at low meritocracy (the reference category) at the right panel, the mobility coefficient captures the origin-specific ATT comparing movers to their immobility counterfactual among low-merit respondents. These effects are small and imprecise across trajectories: Low→Middle ($\beta = -0.02$, [−0.09, 0.04]) and Low→High ($\beta = -0.03$, [−0.10, 0.05]) are near zero. Middle→High is positive but not significant ($\beta = 0.07$, [−0.01, 0.14]), Middle→Low is null (β = -0.00, [−0.07, 0.06]), and downward from high shows negative, borderline estimates: High→Middle ($\beta = -0.07$, [−0.15, 0.00]) and High→Low ($\beta = -0.08$, [−0.15, 0.00]). Substantively, I do not find evidence that meritocratic perceptions moderate the causal effect of mobility, although the signs align with the idea that upward mobility at high merit tilts attitudes toward commodification, and downward mobility at high

Table 2: Effects of intergenerational occupational mobility on preferences for pension commodification, by survey wave

	Low-Middle	Low-High	Middle-High	Middle-Low	High-Middle	High-Low
Intercept	0.18	−0.04	0.28	0.20	0.18	0.15
	[−0.36; 0.73]	[−0.36; 0.29]	[−0.01; 0.57]	[−0.11; 0.52]	[−0.15; 0.50]	[−0.27; 0.58]
Mobility treatment	−0.05	−0.03	0.13*	0.06	−0.19*	−0.19*
	[−0.16; 0.06]	[−0.15; 0.10]	[0.05; 0.22]	[−0.03; 0.15]	[−0.33; −0.05]	[−0.32; −0.05]
Wave (Ref.= 2016)						
Wave 2018	0.01	0.01	0.10*	0.09*	−0.04	−0.02
	[−0.10; 0.12]	[−0.10; 0.12]	[0.03; 0.18]	[0.01; 0.17]	[−0.15; 0.07]	[−0.15; 0.10]
Wave 2023	0.03	0.01	0.19*	0.18*	0.05	0.04
	[−0.09; 0.14]	[−0.12; 0.13]	[0.10; 0.28]	[0.08; 0.27]	[−0.09; 0.18]	[−0.11; 0.19]
Mobility treatment x Wave (Ref.= 2016)						
Mobility treatment x Wave 2018	0.03	−0.06	−0.04	−0.09	0.11	0.06
	[−0.11; 0.17]	[−0.21; 0.09]	[−0.17; 0.09]	[−0.20; 0.03]	[−0.04; 0.26]	[−0.09; 0.22]
Mobility treatment x Wave 2023	0.07	0.10	−0.09	−0.07	0.19*	0.18
	[−0.09; 0.23]	[−0.08; 0.28]	[−0.23; 0.06]	[−0.22; 0.08]	[0.01; 0.38]	[−0.01; 0.37]
Controls	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.04	0.05	0.07	0.03	0.06	0.06
Adj. R^2	0.03	0.04	0.06	0.02	0.05	0.05
Num. obs.	878	745	716	773	912	861
RMSE	0.40	0.37	0.39	0.39	0.44	0.42
N Clusters	497	449	443	470	536	525

Note: Cells contain regression coefficients with confidence intervals in parentheses. * Null hypothesis value outside the confidence interval..

merit tilts them away.

Table 2 shows the causal heterogeneity of the mobility effect across the different waves of the survey. Regarding temporal heterogeneity, the only statistically significant interaction term is observed for the High→Middle trajectory in 2023 ($\beta_{\text{int}} =$ 0.19), 95% CI [0.01, 0.38]). However, when this interaction is added to the baseline effect in 2016 (-0.19), the marginal effect of High→Middle mobility in 2023 is approximately zero, and its confidence interval includes the null. Thus, rather than indicating a reversal toward a positive effect, the evidence suggests that the negative impact of downward mobility from high backgrounds is confined to 2016 and attenuates over time. No comparable temporal changes are observed in the remaining trajectories. Overall, the results provide little support for strong time-conditional heterogeneity in mobility effects, beyond the fading of the initially negative High→Middle effect.

4.4 Robustness check and sensitivity analysis

As a robustness check, I re-estimate the models under three alternative codings of the outcome: (i) recoding the neutral category as agreement (3–5 vs. 1–2), (ii) excluding neutral responses from the analysis, and (iii) using the original 5-point item in a linear specification. Across these alternative models (see Supplementary Material for complete tables), the substantive pattern of the main effects is preserved: the Middle→High trajectory remains positively associated with support for pension commodification, and the High→Low trajectory remains negatively associated with it, with both coefficients generally increasing in magnitude. By contrast, the negative High→Middle effect becomes statistically indistinguishable from zero in the robustness specifications, indicating that this particular estimate is more sensitive to how the outcome is coded than the other two trajectories.

To assess the extent to which the estimated mobility effects may be driven by unmeasured confounding, I implement the sensitivity analysis proposed by Cinelli and Hazlett (2020), which quantifies how strong an unobserved confounder would need to be—in terms of its joint explanatory power for both treatment and outcome—to attenuate or explain away the ATT, using gender as a benchmark covariate (the predictor with the largest partial association with the outcome aside from treatment). For the Middle→High trajectory ($\beta \approx 0.09$), partial $R^2 \approx 0.013$, an unobserved confounder would need to explain about 10–11% of the residual variance of both treatment and outcome to reduce the effect to zero and roughly 4% to render it statistically insignificant, substantially more than gender does. For the High→Low effect ($\beta \approx -0.10$), partial $R^2 \approx 0.012$, the required strength is of similar magnitude (around 10–11% to explain away the effect and about 4% to remove significance), again exceeding the explanatory power of the main observed covariates. By contrast, the High→Middle effect ($\beta \approx -0.07$), partial $R^2 \approx 0.007$ is somewhat less robust: a confounder explaining around 8% of residual variance in both treatment and outcome could nullify the estimate, and one explaining about 2% could make it non-significant. Overall, the sensitivity analysis suggests that the positive Middle→High and negative High→Low effects are moderately robust to unmeasured confounding, whereas the High→Middle effect is more vulnerable to relatively modest omitted variables (see Supplementary Material for complete tables).

5 Discussion

6 Conclusion

7 References

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8 Supplementary material

This section presents the supplementary material for this study.

8.1 Descriptive statistics

Table 3: Descriptive statistics for all variables.

Label	Stats / Values	Freqs (% of Valid)	Valid
Preference for pension commodification	1. Disagree	2702 (78.7%)	3435
	2. Agree	733 (21.3%)	(100.0%)
Father stratum	1. Low	1122 (32.7%)	3435
	2. Middle	1104 (32.1%)	(100.0%)
	3. High	1209 (35.2%)	
Offspring stratum	1. Low	1186 (34.5%)	3435
	2. Middle	1110 (32.3%)	(100.0%)
	3. High	1139 (33.2%)	
Parental education	Mean (sd) : 4 (2.2)	1 : 222 (6.5%)	3435
	min < med < max:	2 : 905 (26.3%)	(100.0%)
	1 < 4 < 10	3 : 576 (16.8%)	
	IQR (CV) : 3 (0.6)	4 : 321 (9.3%)	
		5 : 826 (24.0%)	
		6 : 30 (0.9%)	
		7 : 214 (6.2%)	
		8 : 75 (2.2%)	
		9 : 245 (7.1%)	
		10 : 21 (0.6%)	
Co-residence with both parents at age 15	1. No co-residence	996 (29.0%)	3435
	2. Co-residence	2439 (71.0%)	(100.0%)
Nacionality	1. Non-Chilean	62 (1.8%)	3435
	2. Chilean	3373 (98.2%)	(100.0%)

Label	Stats / Values	Freqs (% of Valid)	Valid
Sex	1. Male	1507 (43.9%)	3435
	2. Female	1928 (56.1%)	(100.0%)
Age (in years)	Mean (sd) : 42.6 (12.5)	58 distinct values	3435
	min < med < max:		(100.0%)
	18 < 43 < 75		
	IQR (CV) : 21 (0.3)		
Indigenous ethnicity	1. Non-indigenous	3035 (88.4%)	3435
	2. Indigenous	400 (11.6%)	(100.0%)
Meritocracy perception	1. Low	2741 (79.8%)	3435
	2. High	694 (20.2%)	(100.0%)
Wave	1. 2016	914 (26.6%)	3435
	2. 2018	1377 (40.1%)	(100.0%)
	3. 2023	1144 (33.3%)	

8.2 Balance evaluation

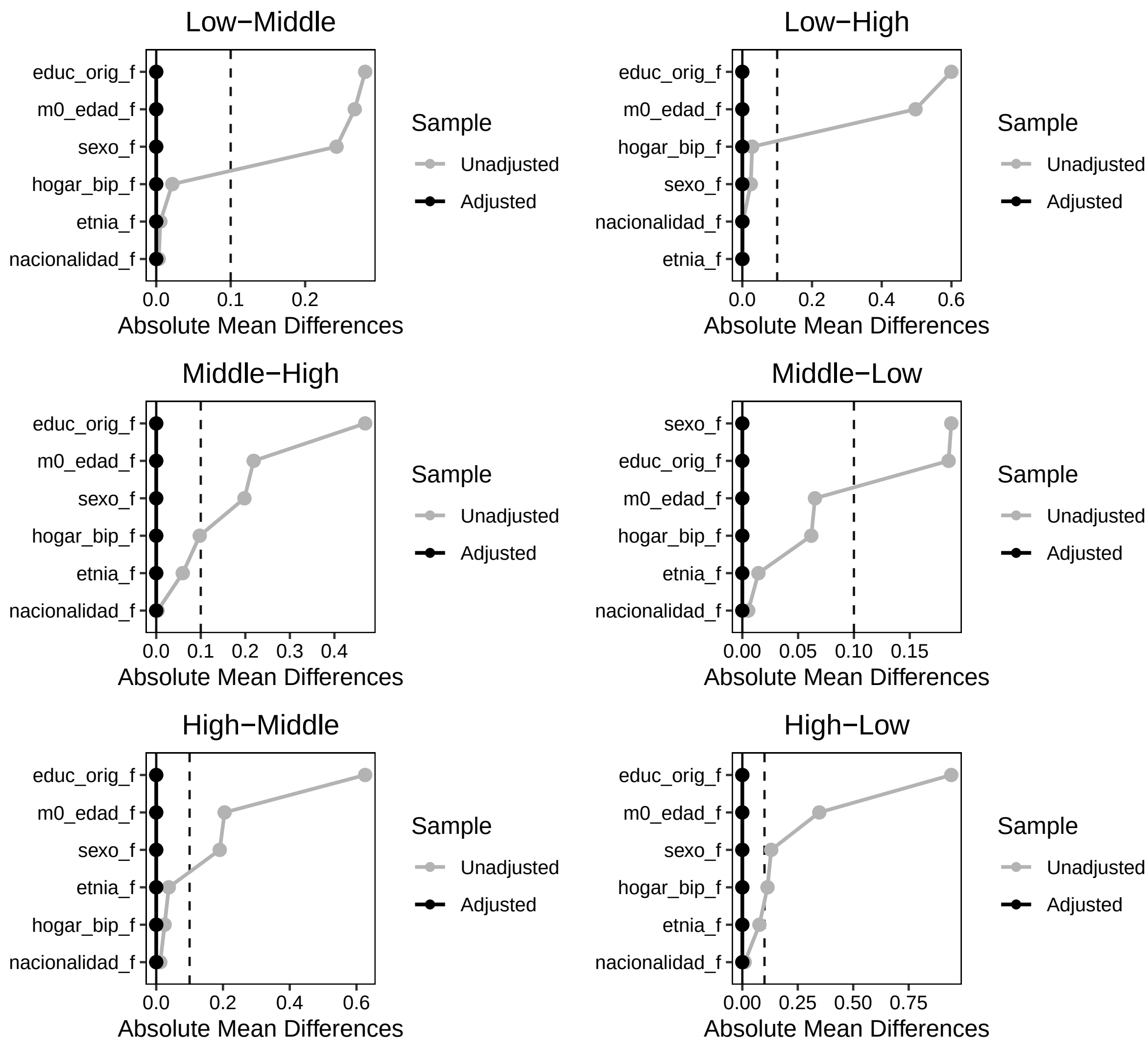

Figure 6: Balance SMD — Mobility treatment (ATT)

8.3 Mobility effects models

Table 4: Effects of intergenerational occupational mobility on preferences for pension commodification, with covariates and wave fixed effects.

	Low-Middle	Low-High	Middle-High	Middle-Low	High-Middle	High-Low
Intercept	0.18	−0.02	0.30*	0.23	0.13	0.14
	[−0.36; 0.71]	[−0.35; 0.31]	[0.01; 0.58]	[−0.08; 0.54]	[−0.19; 0.46]	[−0.28; 0.56]
Mobility treatment	−0.02	−0.02	0.09*	0.00	−0.07*	−0.10*
	[−0.08; 0.05]	[−0.09; 0.05]	[0.02; 0.16]	[−0.06; 0.06]	[−0.15; −0.00]	[−0.18; −0.02]
Age (in years)	0.00	0.00	−0.00	−0.00	0.00	0.00
	[−0.00; 0.00]	[−0.00; 0.00]	[−0.00; 0.00]	[−0.00; 0.00]	[−0.00; 0.01]	[−0.00; 0.01]
Female (Ref. = Male)	−0.14*	−0.14*	−0.13*	−0.07*	−0.11*	−0.12*
	[−0.20; −0.08]	[−0.22; −0.06]	[−0.20; −0.06]	[−0.14; −0.01]	[−0.19; −0.03]	[−0.21; −0.03]
Chilean nacionality (Ref. = Non-Chilean)	−0.03	0.20	0.08	0.04	−0.00	0.06
	[−0.63; 0.57]	[−0.05; 0.46]	[−0.19; 0.36]	[−0.26; 0.34]	[−0.27; 0.27]	[−0.30; 0.42]
Indigenous ethnicity (Ref. = Non-indigenous)	0.03	−0.05	0.05	−0.01	0.07	0.01
	[−0.06; 0.12]	[−0.15; 0.05]	[−0.08; 0.18]	[−0.10; 0.08]	[−0.08; 0.22]	[−0.13; 0.14]
Co-residence with both parents (Ref. = No co-residence)	−0.00	0.04	−0.06	−0.01	0.10*	0.10*
	[−0.08; 0.07]	[−0.04; 0.12]	[−0.14; 0.01]	[−0.09; 0.06]	[0.01; 0.19]	[0.01; 0.19]
Parental education	0.02	0.01	−0.02*	−0.01	−0.00	−0.01
	[−0.00; 0.05]	[−0.01; 0.04]	[−0.03; −0.01]	[−0.02; 0.01]	[−0.02; 0.02]	[−0.03; 0.02]
Wave (Ref.= 2016)						
Wave 2018	0.02	−0.01	0.08*	0.04	−0.00	−0.00
	[−0.05; 0.09]	[−0.09; 0.07]	[0.02; 0.14]	[−0.02; 0.10]	[−0.08; 0.08]	[−0.09; 0.09]
Wave 2023	0.06	0.04	0.15*	0.14*	0.12*	0.10
	[−0.02; 0.13]	[−0.06; 0.13]	[0.08; 0.22]	[0.07; 0.21]	[0.02; 0.21]	[−0.00; 0.21]
R^2	0.04	0.04	0.07	0.03	0.05	0.06
Adj. R^2	0.03	0.03	0.06	0.02	0.04	0.05
Num. obs.	878	745	716	773	912	861
RMSE	0.40	0.37	0.39	0.39	0.44	0.42
N Clusters	497	449	443	470	536	525

Note: Cells contain regression coefficients with confidence intervals in parentheses. * Null hypothesis value outside the confidence interval..

Table 5: Effects of intergenerational occupational mobility on preferences for pension commodification.

	Low-Middle	Low-High	Middle-High	Middle-Low	High-Middle	High-Low
Intercept	0.21*	0.18*	0.17*	0.19*	0.32*	0.28*
	[0.16; 0.26]	[0.13; 0.23]	[0.12; 0.21]	[0.15; 0.24]	[0.26; 0.38]	[0.21; 0.35]
Mobility treatment	−0.01	−0.02	0.08*	−0.00	−0.07	−0.09*
	[−0.08; 0.05]	[−0.08; 0.05]	[0.01; 0.15]	[−0.06; 0.06]	[−0.15; 0.00]	[−0.17; −0.02]
R^2	0.00	0.00	0.01	0.00	0.01	0.01
Adj. R^2	−0.00	−0.00	0.01	−0.00	0.01	0.01
Num. obs.	878	745	716	773	912	861
RMSE	0.40	0.38	0.40	0.39	0.45	0.43
N Clusters	497	449	443	470	536	525

Note: Cells contain regression coefficients with confidence intervals in parentheses. * Null hypothesis value outside the confidence interval..

Table 6: Interactions effects between intergenerational occupational mobility and perceived meritocracy on preferences for pension commodification, with covariates and wave fixed effects.

	Low-Middle	Low-High	Middle-High	Middle-Low	High-Middle	High-Low
Intercept	0.15	−0.04	0.30*	0.21	0.08	0.05
	[−0.36; 0.67]	[−0.40; 0.31]	[0.04; 0.56]	[−0.09; 0.51]	[−0.25; 0.40]	[−0.30; 0.41]
Mobility treatment	−0.02	−0.03	0.07	−0.00	−0.07	−0.08
	[−0.09; 0.04]	[−0.10; 0.05]	[−0.01; 0.14]	[−0.07; 0.06]	[−0.15; 0.00]	[−0.15; 0.00]
High meritocracy perception (Ref.= Low)	0.05	0.06	−0.01	0.03	0.17*	0.26*
	[−0.06; 0.16]	[−0.07; 0.19]	[−0.09; 0.08]	[−0.07; 0.13]	[0.00; 0.33]	[0.04; 0.48]
Mobility treatment x High meritocracy perception (Ref.= Low)	0.04	0.05	0.15	0.04	−0.03	−0.14
	[−0.12; 0.20]	[−0.14; 0.24]	[−0.03; 0.33]	[−0.11; 0.19]	[−0.23; 0.17]	[−0.38; 0.10]
Controls	Yes	Yes	Yes	Yes	Yes	Yes
R^2	0.04	0.05	0.08	0.04	0.07	0.10
Adj. R^2	0.03	0.04	0.07	0.02	0.06	0.08
Num. obs.	878	745	716	773	912	861
RMSE	0.40	0.37	0.39	0.39	0.44	0.41
N Clusters	497	449	443	470	536	525

Note: Cells contain regression coefficients with confidence intervals in parentheses. * Null hypothesis value outside the confidence interval..

Table 7: Effects of intergenerational occupational mobility on preferences for pension commodification coded as 3-5 vs. 1-2, with covariates and wave fixed effects.

	Low-Middle	Low-High	Middle-High	Middle-Low	High-Middle	High-Low
Intercept	0.16	0.01	0.59^{*}	0.57^{*}	0.48^{*}	0.29
	$[-0.38; 0.70]$	$[-0.37; 0.40]$	$[0.17; 1.01]$	$[0.23; 0.92]$	$[0.04; 0.92]$	$[-0.32; 0.90]$
Mobility treatment	0.01	0.05	0.09^{*}	-0.02	-0.08	-0.13^{*}
	$[-0.07; 0.08]$	$[-0.02; 0.13]$	$[0.01; 0.16]$	$[-0.09; 0.05]$	$[-0.15; 0.00]$	$[-0.21; -0.04]$
Age (in years)	0.00	-0.00	-0.00^{*}	-0.00	0.00	0.00
	$[-0.00; 0.00]$	$[-0.01; 0.00]$	$[-0.01; -0.00]$	$[-0.00; 0.00]$	$[-0.00; 0.00]$	$[-0.00; 0.01]$
Female (Ref. = Male)	-0.15^{*}	-0.16^{*}	-0.17^{*}	-0.09^{*}	-0.11^{*}	-0.10^{*}
	$[-0.22; -0.08]$	$[-0.24; -0.07]$	$[-0.25; -0.09]$	$[-0.16; -0.02]$	$[-0.19; -0.03]$	$[-0.20; -0.00]$
Chilean nacionality (Ref. = Non-Chilean)	0.03	0.28	-0.10	-0.24	-0.21	-0.02
	$[-0.57; 0.63]$	$[-0.01; 0.57]$	$[-0.54; 0.33]$	$[-0.56; 0.08]$	$[-0.71; 0.29]$	$[-0.79; 0.76]$
Indigenous ethnicity (Ref. = Non-indigenous)	-0.03	-0.10	0.02	-0.02	0.02	-0.07
	$[-0.12; 0.07]$	$[-0.22; 0.01]$	$[-0.11; 0.15]$	$[-0.12; 0.08]$	$[-0.13; 0.16]$	$[-0.21; 0.07]$
Co-residence with both parents (Ref. = No co-residence)	0.01	0.06	0.01	0.00	0.15^{*}	0.12^{*}
	$[-0.08; 0.09]$	$[-0.03; 0.14]$	$[-0.07; 0.10]$	$[-0.08; 0.08]$	$[0.06; 0.24]$	$[0.02; 0.22]$
Parental education	0.02	0.02	-0.02	0.00	0.00	0.00
	$[-0.01; 0.05]$	$[-0.01; 0.05]$	$[-0.03; 0.00]$	$[-0.02; 0.02]$	$[-0.01; 0.02]$	$[-0.02; 0.02]$
Wave (Ref.= 2016)						
Wave 2018	0.05	0.00	0.07	0.03	0.02	-0.00
	$[-0.02; 0.12]$	$[-0.08; 0.08]$	$[-0.01; 0.15]$	$[-0.04; 0.11]$	$[-0.06; 0.10]$	$[-0.09; 0.09]$
Wave 2023	0.13^{*}	0.11^{*}	0.23^{*}	0.20^{*}	0.15^{*}	0.15^{*}
	$[0.04; 0.22]$	$[0.00; 0.21]$	$[0.14; 0.32]$	$[0.12; 0.28]$	$[0.05; 0.25]$	$[0.05; 0.26]$
R^2	0.05	0.07	0.09	0.05	0.06	0.07
Adj. R^2	0.04	0.05	0.08	0.04	0.05	0.06
Num. obs.	878	745	716	773	912	861
RMSE	0.44	0.43	0.45	0.44	0.48	0.45
N Clusters	497	449	443	470	536	525

Note: Cells contain regression coefficients with confidence intervals in parentheses. * Null hypothesis value outside the confidence interval..

8.4 Robustness check and sensitivity analysis

Table 8: Effects of intergenerational occupational mobility on preferences for pension commodification excluding intermediate category, with covariates and wave fixed effects.

	Low-Middle	Low-High	Middle-High	Middle-Low	High-Middle	High-Low
Intercept	1.84^*	1.76^*	2.39^*	2.40^*	2.14^*	2.35^*
	$[1.08; 2.60]$	$[1.02; 2.49]$	$[1.80; 2.97]$	$[1.88; 2.92]$	$[1.45; 2.82]$	$[1.53; 3.18]$
Mobility treatment	-0.03	-0.03	0.24^*	-0.01	-0.12	-0.22^*
	$[-0.16; 0.09]$	$[-0.16; 0.11]$	$[0.10; 0.38]$	$[-0.13; 0.11]$	$[-0.26; 0.01]$	$[-0.37; -0.07]$
Age (in years)	0.00	-0.00	-0.00	-0.00	0.00	-0.00
	$[-0.01; 0.01]$	$[-0.01; 0.01]$	$[-0.01; 0.00]$	$[-0.01; 0.00]$	$[-0.00; 0.01]$	$[-0.01; 0.01]$
Female (Ref. = Male)	-0.24^*	-0.23^*	-0.27^*	-0.16^*	-0.16^*	-0.22^*
	$[-0.35; -0.12]$	$[-0.39; -0.07]$	$[-0.42; -0.12]$	$[-0.29; -0.03]$	$[-0.30; -0.02]$	$[-0.39; -0.06]$
Chilean nacionality (Ref. = Non-Chilean)	0.07	0.28	-0.16	-0.34	-0.20	-0.07
	$[-0.73; 0.87]$	$[-0.33; 0.89]$	$[-0.69; 0.37]$	$[-0.76; 0.07]$	$[-0.82; 0.43]$	$[-0.75; 0.61]$
Indigenous ethnicity (Ref. = Non-indigenous)	0.03	-0.08	-0.02	0.04	0.05	-0.21
	$[-0.14; 0.20]$	$[-0.28; 0.12]$	$[-0.28; 0.25]$	$[-0.15; 0.23]$	$[-0.20; 0.31]$	$[-0.49; 0.07]$
Co-residence with both parents (Ref. = No co-residence)	0.00	0.05	-0.09	-0.01	0.15	0.15
	$[-0.14; 0.15]$	$[-0.10; 0.21]$	$[-0.23; 0.06]$	$[-0.16; 0.13]$	$[-0.01; 0.32]$	$[-0.03; 0.33]$
Parental education	0.01	0.02	-0.04^*	-0.01	0.00	-0.01
	$[-0.04; 0.07]$	$[-0.04; 0.07]$	$[-0.08; -0.01]$	$[-0.05; 0.02]$	$[-0.03; 0.04]$	$[-0.05; 0.03]$
Wave (Ref.= 2016)						
Wave 2018	0.02	-0.13	0.10	-0.01	-0.08	-0.18
	$[-0.11; 0.16]$	$[-0.28; 0.03]$	$[-0.05; 0.26]$	$[-0.14; 0.13]$	$[-0.24; 0.07]$	$[-0.37; 0.01]$
Wave 2023	0.21^*	0.03	0.39^*	0.37^*	0.30^*	0.16
	$[0.06; 0.36]$	$[-0.15; 0.21]$	$[0.23; 0.54]$	$[0.23; 0.51]$	$[0.13; 0.46]$	$[-0.04; 0.36]$
R^2	0.04	0.04	0.10	0.06	0.06	0.08
Adj. R^2	0.03	0.02	0.09	0.05	0.05	0.07
Num. obs.	812	677	629	693	808	770
RMSE	0.77	0.73	0.75	0.75	0.82	0.79
N Clusters	480	431	415	445	499	494

Note: Cells contain regression coefficients with confidence intervals in parentheses. * Null hypothesis value outside the confidence interval..

Table 9: Effects of intergenerational occupational mobility on preferences for pension commodification using the original 5-point likert scale, with covariates and wave fixed effects.

	Low-Middle	Low-High	Middle-High	Middle-Low	High-Middle	High-Low
Intercept	1.99^*	1.76^*	2.91^*	2.86^*	2.53^*	2.59^*
	$[0.81; 3.17]$	$[0.82; 2.70]$	$[2.19; 3.64]$	$[2.24; 3.47]$	$[1.59; 3.47]$	$[1.34; 3.83]$
Mobility treatment	-0.03	0.03	0.30^*	-0.02	-0.19	-0.33^*
	$[-0.21; 0.15]$	$[-0.16; 0.22]$	$[0.11; 0.48]$	$[-0.18; 0.14]$	$[-0.38; 0.00]$	$[-0.54; -0.12]$
Age (in years)	0.00	-0.00	-0.00	-0.00	0.00	0.00
	$[-0.01; 0.01]$	$[-0.01; 0.01]$	$[-0.01; 0.00]$	$[-0.01; 0.00]$	$[-0.01; 0.01]$	$[-0.01; 0.01]$
Female (Ref. = Male)	-0.37^*	-0.37^*	-0.40^*	-0.23^*	-0.26^*	-0.32^*
	$[-0.54; -0.20]$	$[-0.58; -0.15]$	$[-0.60; -0.21]$	$[-0.41; -0.06]$	$[-0.46; -0.06]$	$[-0.55; -0.08]$
Chilean nacionality (Ref. = Non-Chilean)	0.10	0.56	-0.25	-0.49^*	-0.32	-0.05
	$[-1.17; 1.38]$	$[-0.08; 1.20]$	$[-0.89; 0.39]$	$[-0.93; -0.05]$	$[-1.23; 0.59]$	$[-1.24; 1.15]$
Indigenous ethnicity (Ref. = Non-indigenous)	-0.00	-0.18	-0.00	0.02	0.08	-0.26
	$[-0.24; 0.24]$	$[-0.46; 0.10]$	$[-0.35; 0.35]$	$[-0.24; 0.28]$	$[-0.29; 0.44]$	$[-0.66; 0.13]$
Co-residence with both parents (Ref. = No co-residence)	0.01	0.11	-0.06	-0.01	0.28^*	0.26^*
	$[-0.20; 0.22]$	$[-0.11; 0.33]$	$[-0.27; 0.15]$	$[-0.21; 0.19]$	$[0.05; 0.50]$	$[0.01; 0.50]$
Parental education	0.04	0.03	-0.06^*	-0.01	0.01	-0.01
	$[-0.04; 0.11]$	$[-0.04; 0.11]$	$[-0.10; -0.01]$	$[-0.05; 0.04]$	$[-0.04; 0.05]$	$[-0.07; 0.05]$
Wave (Ref.= 2016)						
Wave 2018	0.08	-0.12	0.16	0.03	-0.05	-0.17
	$[-0.11; 0.26]$	$[-0.32; 0.09]$	$[-0.04; 0.36]$	$[-0.15; 0.21]$	$[-0.26; 0.16]$	$[-0.42; 0.09]$
Wave 2023	0.33^*	0.14	0.55^*	0.53^*	0.42^*	0.30^*
	$[0.11; 0.54]$	$[-0.11; 0.39]$	$[0.35; 0.76]$	$[0.35; 0.72]$	$[0.18; 0.65]$	$[0.02; 0.57]$
R^2	0.04	0.05	0.10	0.06	0.06	0.08
Adj. R^2	0.03	0.04	0.09	0.05	0.05	0.07
Num. obs.	878	745	716	773	912	861
RMSE	1.10	1.04	1.06	1.07	1.17	1.12
N Clusters	497	449	443	470	536	525

Note: Cells contain regression coefficients with confidence intervals in parentheses. * Null hypothesis value outside the confidence interval..

Table 10: E-values for ATT estimates for Middle-High trajectory (Cinelli & Hazlett, 2020 approximation)

Outcome: *y*						
Treatment:	Est.	S.E.	t-value	$R^2_{Y\sim D\|\mathbf{X}}$	$RV_{q=1}$	$RV_{q=1,\alpha=0.05}$
t	0.088	0.029	2.995	1.3%	10.7%	3.8%
df = 706				*Bound (1x sexo)*: $R^2_{Y\sim Z\|\mathbf{X},D}$ = 2.7%,	$R^2_{D\sim Z\|\mathbf{X}}$ = 3.6%	

Table 11: E-values for ATT estimates for High-Middle trajectory (Cinelli & Hazlett, 2020 approximation)

Outcome: y						
Treatment:	Est.	S.E.	t-value	$R^2_{Y\sim D\|\mathbf{X}}$	$RV_{q=1}$	$RV_{q=1,\alpha=0.05}$
t	-0.074	0.03	-2.447	0.7%	7.8%	1.6%
df = 902			*Bound (1x sexo)*: $R^2_{Y\sim Z\|\mathbf{X},D}$ = 1.5%, $R^2_{D\sim Z\|\mathbf{X}}$ = 3.9%			

Table 12: E-values for ATT estimates for High-Low trajectory (Cinelli & Hazlett, 2020 approximation)

Outcome: y						
Treatment:	Est.	S.E.	t-value	$R^2_{Y\sim D\|\mathbf{X}}$	$RV_{q=1}$	$RV_{q=1,\alpha=0.05}$
t	-0.098	0.03	-3.245	1.2%	10.5%	4.3%
df = 851			*Bound (1x sexo)*: $R^2_{Y\sim Z\|\mathbf{X},D}$ = 1.7%, $R^2_{D\sim Z\|\mathbf{X}}$ = 0.8%			