

UNIVERSITA' COMMERCIALE "LUIGI BOCCONI"

PhD SCHOOL

PhD program in Public Policy and Administration

Cycle: 36°

Disciplinary Field (code): SPS/07

**Essays on Family Transitions and Income
Dynamics: Cross-National Evidence from Survey
Data**

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Year 2026

Introduction: Essays on Family Transitions and Income Dynamics: Cross-National Evidence from Survey Data

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The field of family demography stands at a critical juncture, as family structures and dynamics have become more complex and less prone to classification. The idea of a single, dominant nuclear family is no longer the standard approach in the discipline, with researchers framing the current observed trends within the concept of the Second Demographic Transition (Lesthaeghe, 1995; Van de Kaa, 1987). This shift in society as a whole is marked by diminished and delayed fertility, a preference for cohabitation, rising union dissolution and re-partnering, and greater variety in household composition, with a higher incidence of single parents and same-sex couples (Castrén et al., 2021). This sort of transformations is the result of a mix of socio-economic and cultural factors, including greater individualization, shifting gender norms and women’s empowerment, and adjustments in welfare policies. (McDonald, 2000; Sobotka & Toulemon, 2008). The reshaping of family structures can have ripple effects as it impacts not only demographic trends but also broader aspects of social inequality, intergenerational solidarity and policy agendas.

Traditional demographic methods relying on aggregate figures or cross-sectional data are proving increasingly insufficient to capture the reality of these changes. Therefore, for contemporary family demography, the challenge is not only in documenting new patterns, but also in explaining how they unfold throughout individual life courses and varying societal settings. As a result, there is a critical need for longitudinal, high-quality micro-level survey data—often managed by so-called ‘life observatories’ on the national and international level—that allow tracking individual trajectories through key family events, and capturing their chronological sequence and duration (Evandrou et al., 2020). The inclusion of time as a dimension adds the analytical depth that is in demand among researchers across disciplines and that the cross-sectional design inherently lack, providing insights into causality and long-term effects of social processes (Primerano et al., 2026) The life course perspective,

as highlighted by Elder (1985), offers a distinct theoretical angle, noting how individual lives are intertwined with history and social relations, and processes such as family formation, well-being and dissolution are rooted in a dynamic interplay between individual agency and structural constraints. Beyond merely documenting change, rich longitudinal datasets track individuals over time in a nuanced and detailed manner and come with specific advantages. They can track individuals before and after significant life changes such as employment, marriage, or childbirth, strengthening causal inferences beyond correlation. In addition, these datasets can reveal heterogeneity in social group experiences and how individuals from distinct social groups, categorized by education or social class, may go through similar life stages, but diverge in outcomes. The identification of these groups is what can enable effective policy formulation. Finally, high-quality longitudinal data enables disentangling, with precision, the period, cohort, and age effects. This way, scholars are better equipped to determine whether changes are rooted in historical events, generational experiences, or universal life-course processes.

A key part of the changing demographic landscape, and a main focus of this dissertation, is the economic foundation of family life. Economic factors like resources, income differences, job trends, and financial support are no longer seen as irrelevant; they are closely linked to people's ability to form relationships, have children, and maintain family stability. The family unit cannot be separated from its economic aspects. As such, families have been viewed as central units of production, consumption, and resource allocation since Becker. Nowadays, shifting job markets, along with rising living costs and different levels of support from welfare programs, can directly affect decisions about partnering, starting a family, or becoming independent from parents. Modern family demographics highlight the connection between the economy and family structure, showing that financial resources and security are vital to family decisions. This perspective also extends to a deeper interrogation of inter-generational relationships, not just in terms of transfers of wealth or care, but also of their broader implications for generational inequality and social cohesion.

The contributions in this dissertation directly address crucial developments in family demography, offering new insights into some of the field's pressing questions and its method-

ological challenges. One major change in the field is the evolving economic dynamics within couples. Dual-earner families have become the standard, and female breadwinner models are also increasingly common. This challenges traditional assumptions about gender specialization, economic dependency, and bargaining power within couples. Understanding how such changes in partners' relative earnings shape union stability remains a key question in family demography. The first paper of the dissertation investigates how income asymmetries between partners relate to the risk of dissolution of first partnerships in the United Kingdom. By building on the comprehensive UK household panel and on competing theories on the independence hypothesis, income effects and gender-institutional framework, the analysis tests whether and under which conditions women's economic advantage relates to dissolution risk. Results from survival models indicate that periods of pronounced female income dominance are associated with moderately higher dissolution. Yet, these associations tend to become smaller and imprecisely estimates as we include a time-varying subjective distress among controls. Relatedly, distress emerges as a strong predictor of separation but it is unlikely to operate as mediator between female income dominance and dissolution, based on the results from decomposition analyses. There is no overall support for a single dominant mechanism whereby women's higher earnings structurally destabilizes unions. The dissolution risk may instead emerge at the intersection of extreme income circumstances, broad economic insecurity, and psychological vulnerability. The findings warrant the need to move beyond simplistic views of economic roles and consider as well household dynamics, financial security and psychological strain when assessing relationship stability. Longitudinal data remain key to explore these complex interactions over time.

Another relevant area of focus in modern family demographics and social stratification is the changing economic situation of young adults and its impact on their move to an independent living and family formation. As the transition to adulthood increasingly delays and becomes uncertain in many high-income countries, it is essential to understand the roots of youth economic vulnerability for both demographic insight and effective policy design, especially given the rising costs of housing and education and the instability of entry-level jobs. This line of research can highly benefit from the comparative analysis of the economic well-being of young adult household heads (aged 18–30) across 15 high-income countries over

two decades. By focusing on those who have already left the parental home, the study offers a clearer view of the challenges facing independent young adults, which are crucial years for forming unions and fertility intentions. Leveraging on the harmonized microdata available through the Luxembourg Income Study (LIS), the study points to a widespread increase in poverty risk among young adults, particularly in Continental Europe, along with a widening intergenerational gap. Importantly, decomposition techniques differentiate between compositional changes, such as shifts in education, employment, or living arrangements, and structural forces like declining labor market and higher education returns. This helps clarify whether youth poverty mainly results from changes in the characteristics of young people or from more structural variations in available economic opportunities. This distinction is critical for designing targeted interventions that may address the sources of economic insecurity among the younger generation. This, in turn, affects their ability to form and sustain families, impacting future demographic trends.

Fertility intentions and behaviors are still central to family demographics, with recent research increasingly emphasizing the interplay between individual aspirations, structural barriers, and informal support networks. Among these, private financial transfers received by family members represent a significant yet understudied factor, particularly in settings with limited welfare provision or access to public services. The field is shifting towards a clearer understanding of how these transfers operate, focusing on their enabling effects and how they originally reflect general social inequalities. My approach complements this current trend in the literature by examining the connection between ascendants' monetary transfers and short-term fertility intentions in five European countries, using harmonized data from two rounds of the Generations and Gender survey. Our findings do not point to a systematic positive relationship between parental transfers and fertility intentions. In both rounds, positive associations often appear in simple specifications but weaken after demographic, socioeconomic, and parental background controls are included. The Panel analyses on the first round do not provide any evidence of stable within-individual effects, suggesting that initial existing association could be driven by compositional differences between recipients and not recipients, rather than by an independent impact of transfers per se. In the second round, instead, the results reveal some evidence of cross-national heterogeneity. Austria

and Germany display more robust positive associations in the pooled analyses, especially for what concerns large and recent parental transfers. Persistent positive coefficients are confirmed by propensity score matching technique for these two economies, implying that financial support by close relatives could function as enabling resources in specific national contexts, rather than broadly in the whole Continent, and in more recent times. Moreover, our results further suggest that the relevance of material support may depend on timing, size, parity and institutional context. This work can shed additional light on the subtle and contingent ways in which a specific dimension of intergenerational ties affect reproductive choices in modern European societies. This is especially relevant given the ongoing economic uncertainty of the past two decades, which has further underscored the importance of family networks as a source of stability and support when institutional protection weakens.

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Response to Reviewers' Comments on Paper 1

Elena Marseglia

1. Abstract

Reviewer 2

The abstract could be more informative regarding the data source (the UKHLS), including years covered and the sample characteristics. Elena may briefly summarize the empirical strategy (e.g., piecewise exponential model, fixed effects, etc.), so that readers can quickly understand the methodological approach. The main findings and their implications could be stated more clearly, linking them back to the central research question.

Response:

In the revised version of the abstract, we now identify more explicitly the data source in the UK Household Longitudinal Study, specifying the years covered (1999-2019) and clarifying that the analysis focuses on the first partnerships observed in the panel. A new concise description of the empirical strategy is included and mentions the combination of nonparametric estimates and semiparametric Cox hazard models, including its interval-based specification. There is a sharper presentation of the main findings: female breadwinner arrangements are associated with moderately higher dissolution risks in several specifications, the association is concentrated in the most unconventional configurations, and the coefficient attenuates once a measure for subjective distress is included. We used decomposition analyses to test whether the association between family arrangement and dissolution risk is mediated by psychological distress, but we found no support. This links the findings to one research question that concerns the existence of a mechanism less anchored to the the economic explanations of independence and income effects. The new abstract should now summarize more clearly the data, methods, findings and contribution.

2. Introduction

Reviewer 1

In the introduction, you emphasize that the chapter focuses on whether women's economic self-sufficiency destabilizes partnerships. While this is an interesting question, the current formulation contains certain presumptions that may limit the scope of your analysis. For instance, the phrasing somewhat assumes that women achieve economic self-sufficiency only after entering stable partnerships. However, there might be alternative pathways: (i) women may establish professional careers before entering stable partnerships and choose partners compatible with their working lives, or (ii) stable partnerships themselves may facilitate women's career advancement. I suggest reformulating the research question more neutrally and acknowledging these alternative pathways from the outset of the chapter. Also, the current framing may inadvertently suggest that women's economic stability or self sufficiency necessarily leads to higher divorce rates or bears significant responsibility for increasing divorce rates. This perspective has been challenged in numerous contexts, and a more nuanced body of evidence has emerged regarding the relationship between women's economic status and union formation/dissolution behaviors. The introduction would benefit from engaging with this more recent and nuanced scholarship. Furthermore, the introduction would benefit from a more substantial discussion of why this research question matters, what gaps or limitations exist in previous studies, and what the actual contribution of this study is. The current motivation for the study feels somewhat disconnected from the research question itself.

Reviewer 2

The way the main research question is phrased ("Does women's economic self-sufficiency destabilize partnerships?") suggests a one-way causal link from women's economic status to union dissolution. But in reality, the relationship could go in both directions, or even be shaped by selection processes. For instance, women who are already career-oriented or economically independent might choose partners whose lifestyles are compatible with their work, which could actually stabilize the relationship rather than destabilize it. Conversely, being in a supportive, stable partnership might help women advance in their careers and achieve self-sufficiency in the first place. In short, the question assumes a direction of influence that is theoretically debatable and empirically difficult to isolate. A brief reflection on these alternative mechanisms, reverse causality and partner selection, maybe in the Discussion section, would make the framing more balanced and credible. One thing that's missing in the chapter is a discussion of selection into marriage and cohabitation, and how that, in turn, shapes selection into union dissolution.

Response :

We have substantially revised the introductory section to avoid stressing on a one-directional relationship between women's self-sufficiency and union dissolution. The original phrasing of the research question could have been interpreted as presuming only destabilizing effects stemming from higher female earnings, so we now have reformulated the central question under a more neutral tone on how, whether, and under which conditions emerging episodes of female income dominance are related to union stability. The section now better acknowledges alternative pathways such as adaptation to evolving gender roles or selection mechanisms; it also emphasizes that new forms of family arrangements could in principle have stabilizing or destabilizing effects. The introduction also serves to set the study within alternative frameworks proposed in the literature (on top of the independence and income effects, a gender-institution perspective is now added) that are better discussed within the Theoretical Background. It also presents an overview of the findings, highlighting its contributions (use of UK data in a debate largely dominated by US evidence, distinction between average and time varying income configurations to better capture the episodes of specific income configurations, incorporation of the subjective distress indicator and use of decomposition techniques to formally test mediation). [pages 2-5].

3. Background**Reviewer 1**

The background relies heavily on the Beckerian family model. Given that this dissertation aims to contribute to demography and sociology, it would benefit from incorporating complementary theoretical frameworks and engaging more substantively with work by family sociologists and demographers. For instance, Oppenheimer (1994) challenges the notion that women's economic position is the primary driver of observed changes in family formation behavior.

Reviewer 2

There is limited discussion of selection into marriage versus cohabitation and how this selection, in return, affects union dissolution patterns. The socioeconomic profiles of married and divorced couples have changed significantly over time and across contexts. A more thorough treatment of these selection processes would allow you to develop more nuanced and even potentially competing hypotheses.

Over time, the kinds of people who marry or divorce have changed quite a bit, not just demographically but also socioeconomically. Ignoring this makes it seem as if marriage and cohabitation are homogeneous institutions, when they're really not. A short reflection on these selection processes could help you develop different (alternative/additional) hypotheses, and it would give the study's motivation and contribution a stronger grounding.

Response :

In the background section we have incorporated complementary perspectives and added more recent studies, including most of the suggested articles. The section engages more with Oppenheimer's critique to the specialization-based explanation, emphasizing the role of economic uncertainty as alternative driver of union change. There is now a third category of studies that situate income asymmetry within a normative context that includes gender- and identity-based mechanisms, or highlight the importance of other aspects such as relationship quality. We include a brief discussion of how selection processes may also influence duration patterns and account for the role of subjective distress as a potential mechanism. [pages 5-8].

We also acknowledge the issue of selection into union type and increasing socioeconomic stratification of marriage. While, the empirical design focuses on first partnerships spells and always include the union type as a covariate, we more clearly discuss that we do not model the transitions between cohabitation and marriage over time. [pages 7, 30].

4. Data, method, analysis

Reviewer 1

Please provide a rationale for using piecewise exponential models. Is this choice driven by variation in divorce risk across marriage duration? Also, please be more explicit about how episodes and intervals are defined. Several suggestions other suggestions for this section:

- Descriptive statistics: Include sociodemographic characteristics of the three family types (male breadwinner, dual-earner, and female breadwinner) to give readers a clearer picture of the populations being compared.
- Data censoring: Consider censoring data after a specified duration (e.g., 10 or 15 years). Please review standard practices in the literature to determine the most appropriate approach.
- Table presentation: Revise regression tables to ensure that variables and models are clearly labeled and easy to interpret.
- Relationship quality and housework division: These variables may be correlated with family type itself, raising important questions about causal ordering. Specifically, relationship quality and housework division could be outcomes of family type rather than independent predictors, suggesting they may play a mediating role. For example, in female-breadwinner households, male partners may have more time available for housework. At the same time, the female-breadwinner arrangement might challenge traditional masculinity norms and create relationship strain, thereby reducing relationship quality. It is crucial to clarify where these variables are placed in your causal framework and when they are measured temporally. Note that if these variables are indeed mediators, they should not be controlled for in the same models as the main

predictor.

Response:

We mainly use piecewise Cox models (and exponential parametric models as robustness check in the Appendix) to test whether observed periods of income dominance affect the dissolution risk at a given time, rather than sticking only with an average measure of the female relative income ratio as predictor. These models allow for the inclusion of richer and more precise controls such as the household income position or the subjective level of distress which are also likely to fluctuate over time. We restructure the union spells into yearly intervals that match the survey observation period (1991-2019) and perform left-truncation. Unions enter the risk at time in which both partnership is ongoing and relative income is observed [page 18].

We include the descriptive statistics of time-invariant characteristics by family type (male breadwinner, mostly egalitarian, female breadwinner) in Table 1 [page 12]. The groups are broadly similar in the key sociodemographic characteristics, with the most notable difference in a lower prevalence of marriage and presence of children among female breadwinner unions. These patterns may mechanically relate to union stability and we control for union status and presence of children in the multivariate modelling. Before running the time-varying models, we also note that female breadwinner episodes co-occur more often with the unemployment status of the respondent or below-median household income, suggesting that these arrangements can often emerge in context of higher vulnerability [pages 18-19].

In the baseline nonparametric estimates, we right-censor at the 75th percentile of the observed union duration to limit distortions from a small number of especially long-lasting unions [page 12]. We further check in the Appendix that censoring at 10 years and 15 years (roughly one-third and one-half of the previous duration) yields consistent shapes of survival curves as well as unchanged conclusions from log-rank tests. This supports the interpretation that the instability gap is not driven by long unions only. In the piecewise empirical specifications, we test whether the association between family arrangement and dissolution is duration-dependent. We do not find proportional-hazards violation in our variable of interest nor significant interaction effects for the (log) duration [page 22].

We lastly address the concern about potential mediators and causal ordering. In the interval-based models, the last time-varying covariate added is the standardized indicator for subjective distress that is a more consistently reported variable compared to unpaid work or relationship satisfaction

(albeit more generic). Once distress is introduced, the female breadwinner coefficient attenuates and becomes statistically insignificant, while distress is a strong and consistent predictor of dissolution. We conduct additional analyses as mediation tests using KHB decompositions. We implement both a contemporaneous decomposition and a lagged specification, and find very limited evidence of mediation. In light of the evidence, distress appears to capture an important dimension of vulnerability (potentially overlapping with the extreme female breadwinning configurations) rather than a mechanism generated by female income dominance per se [pages 25-26].

5. Discussion and conclusions

Reviewer 1

The discussion section is currently quite brief and would benefit from expansion. Please address: (i) why you observe these particular effects, what mechanisms might explain the patterns you find, (ii) how these family types differ socioeconomically and what implications this selection has for your findings, and (iii) the key limitations and strengths of this chapter.

Reviewer 2

The discussion section feels quite brief and doesn't fully unpack the results. It would really benefit from a deeper reflection on why you find the observed effects and what mechanisms might be driving them. For instance, how do the three family types differ socioeconomically, and how might that selection influence the patterns you observe in union stability? Right now, the results are reported but not sufficiently interpreted. It would also help to discuss the limitations (e.g., potential endogeneity, measurement of income, or omitted mediators) and the strengths of your analysis (e.g., longitudinal data, detailed family typology).

Response:

The Discussion section has been thoroughly expanded, it now provides a fuller account of the findings and attempts to clarify the mechanisms that may underlie the observed associations. We discuss that dissolution risks are usually concentrated among the more extreme female breadwinner arrangements, and that episodes of female income dominance may actually reflect shocks on the labor market or broader socioeconomic vulnerabilities. While this income configuration tends to show destabilizing effects in most time-varying models, the pattern largely changes when subjective distress is also accounted for. This latter is a strong predictor of dissolution, but decomposition analyses have provided little support for the existence of a mediation pathway from female income dominance to partnership breakdown. There is therefore the possibility that both distress and extreme income asymmetries stem from shared underlying strains (that would warrant further investigation) rather than a direct causal chain. There is also limited evidence of a pure income effect

or interactions with partnership status. We have further expanded the limitations of the article, including issues of selection and reverse causality, measurement of labor income rather than total household resources, limited availability of relational variables, and the individual and more generic nature of the distress indicator. Among the strengths, we report the advantage of longitudinal data for survival analysis, the distinction between time-invariant and interval-based income configurations, and the integration of the psychological dimension instead of the mere economic measures. These revisions help providing a more comprehensive and grounded interpretation of the findings, as suggested. [pages 28-31].

Do female breadwinner family structures increase partnerships' instability? Evidence from UK survey data

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1 Abstract

This article examines whether and under which conditions income asymmetries within couple, and in particular those arrangements in which women outearn their male partners, are associated to partnerships' duration in the United Kingdom. Drawing on longitudinal data on first partnerships from the UK Household Longitudinal Study between 1991 and 2019, the analysis primarily combines nonparametric estimates and Cox hazard models. Preliminary results with time-invariant measure of relative income indicate that female breadwinner models display higher dissolution rates, with instability concentrated in the most unconventional arrangements. Interval-based models further show that periods of female income dominance are associated with about 20% higher hazard relative to male breadwinner models, while egalitarian arrangements do not differ from the reference group. Yet, once a measure of subjective distress is included, the female breadwinner coefficient attenuates and becomes insignificant. Distress itself emerges as a strong predictor of dissolution, however KBH decomposition analyses provide limited support for a mediation pathway from female income dominance to dissolution via distress. Rather than supporting a single explanatory mechanism, the results point to a multi-faceted role of extreme earnings asymmetries, socioeconomic vulnerabilities and psychological well-being in shaping unions' dissolution risk.

2 Introduction

Over the past decades, women's increased participation in higher education and the labor market has contributed to reshaping the economic structures to be found within partnerships. One of the visible effects of this societal transformation is that there is an increasing percentage of unions where the female partner contributes equally or relatively more than her male partner to the household labor income. Still, this latter constitutes a rather minority configuration in most contexts and hence questions of how, whether, and under which conditions emerging episodes of female income dominance are related to union stability appear to be rather unexplored, both theoretically and empirically. The existence of a link between income dominance and union stability is far from settled or straightforward, with competing perspectives that point to different expectations regarding its potential consequences.

On the one hand, one can stress the economic independence and, with it, the availability of exit options. Women who contribute a larger share of household income face fewer material constraints to remain in unsatisfactory relationships, and this could potentially lead to higher dissolution rates among this group. A related argument may also suggest that women who are economically independent are also likely to prioritize their self-realization and autonomy, thus showing greater selectivity in the choice of partner and more willingness to abandon unions which do not meet their expectations. On the other hand, extending beyond the mere exit capacity of women, there might exist further implications in terms of dynamics of bargaining within couples. Divergences from a more traditional male breadwinner model, where the male partner is the sole earner, often require both partners to adapt to or renegotiate established gender roles. For instance,

the increasing deterioration of a male income dominance model may challenge identity norms tied to provision as well as rise expectations related to collaboration for men, while for women it may entail overburden of roles or tensions in balancing between paid and unpaid labor and childcare. Such adjustments can generate new types of frictions within the couple and undermine relationship stability, particularly in contexts where gendered expectations maintain salience. Alternatively, a stronger position of women on the labor markets can exert stabilizing effects through the availability of more material resources for the household or through selection mechanisms. For instance, women could delay union entry to the achievement of specific educational or occupational goals or become more selective in their partner choice, with a greater number of unions potentially being formed by partners with similar gender attitudes or balanced expectations on work and family roles.

The variety of offered explanations suggest that the relationship between female income dominance and union dissolution is unlikely to be uniform. It may differ by partnership type, degree of income asymmetry, the broader household's socioeconomic position, or the prevailing norms on gender values. Moreover, even if the point of departure is an economic asymmetry in earnings, any observed association with union duration may operate through multiple channels including psychological wellbeing or relational adjustments, or reflect selection of unions formed under specific socioeconomic or attitudinal conditions. Although the consequences of earnings asymmetry within couples have long been debated, much of the empirical literature is restricted to the United States. Few studies have examined these dynamics in other institutional contexts using longitudinal data that allow a dynamic observation of couples' income configurations and dissolution outcomes. To this purpose, the United Kingdom provides a relevant setting as it shares similarities with the United States in terms of female labor force participation, variety of partnership

forms, as well as the availability of rich panel data that allow for the matching of income measurement and partnerships' duration.

This chapter attempts to address the ambiguity surrounding this nexus using 27 waves from UK Household Longitudinal Study (UKHLS), with the analysis proceeding in stages. First, non-parametric and baseline Cox models document survival differences across three diverse family arrangements based on the prevalent earnings ratio across years (male breadwinner, mostly egalitarian, and female breadwinner unions) and further disaggregate female relative income into more detailed brackets to highlight that instability is quite concentrated among the most unconventional income configurations. Second, piecewise Cox models and exponential models test whether observed periods of income dominance affect the dissolution risk. These models allow for the inclusion of a richer set of controls such as the household income position and the subjective level of distress that may fluctuate over time, similarly to our main variable of interest. Our findings indicate that female breadwinner models carry higher risk of breakdown (20-25% higher hazard relative to male breadwinner periods) across several specifications with progressive use of controls. Additionally, we do not find evidence of varying effects by the household's income position or the relationship status in interaction models. A destabilizing effect of female breadwinner models disappears only when an indicator capturing subjective distress is introduced. While distress is a consistent predictor of dissolution, decomposition analyses reveal that this variable explains only a small fraction of the association between female income dominance and union instability, while an earlier income dominance does not significantly predict subsequent distress. While we cannot claim that psychological strain operates as a mediating mechanism, it can highlight a broader vulnerability phenomenon that may overlap with specific income asymmetries, suggesting a rather complex relationship between an extreme earnings

asymmetry and partnership instability that go beyond the economic-oriented approaches.

3 Theoretical Background

In recent decades family structures have undergone profound transformations along with dynamics related to education, employment and attitudes towards the formation and maintenance of intimate relationships. A central driver of these changes has been the expanding economic role of women. From being primarily family carers and providers of unpaid work, in most Western countries women have entered higher education and labor market at unprecedented rates, favored also by exogenous factors like technological advances, the diffusion of household appliances and birth control instruments (Goldin, 2006). In many European countries, women now surpass men in educational attainment, and female labor force participation now stands above 50% percent in most advanced economies (Esteban Ortiz-Ospina & Roser, 2018; Klesment & Van Bavel, 2017). These developments have contributed to a declining prevalence of a traditional model based on complete gender specialization and to the normalization of dual-earner arrangements (Daly, 2011).

The concurrent phenomena of more frequent unions' dissolution and greater female participation in the labor market have spurred more than a two-decade-long investigation into the sociological consequences of women's economic self-sufficiency. The earliest economic theories have emphasized specialization gains within marriage, conceived as an institution that maximizes utility when built upon a gendered division of labor (Becker, 1981; Cherlin, 1979). From this perspective, the increasing economic power of women undermines the spouses' mutual interdependence and weakens the economic

gains stemming from marriage. The relative decrease of the exit costs for women along with the democratization of divorce through diminished legal barriers and social stigma (Ruggles, 1997) has been accordingly associated with heightened dissolution risks - an argument that has been popularized as independence effect. At the same time, the income effect, articulated notably by Ross and Sawhill (1975), offers the counter-prediction of a stabilizing potential of women's earnings, especially when a secondary income improves household economic security and buffers against major financial distress. Subsequent critiques have similarly challenged the rather deterministic reading of the independence effect. For instance, Oppenheimer (1997) argued that economic uncertainty, increasingly affecting men's employment prospects, may be more central to transitions within the family than women's rising employment. Women's earnings in this framework can compensate for labor market volatility or sustain partnerships at times of economic precariousness (Brines & Joyner, 1999; Dechter, 1992). Moreover, later contributions have cautioned against attributing union dissolution solely to women's economic autonomy without accounting for gender ideology, relationship quality or broader contextual factors (Sayer & Bianchi, 2000). Empirical evidence across countries has indeed produced mixed findings, suggesting that the implications of women's earnings depend on institutional settings or socioeconomic conditions (Cooke, 2006; Cooke et al., 2013; Jalovaara, 2003; Liu & Vikat, 2004; Vignoli et al., 2018). Another consideration concerns selection into union type. Marriage and cohabitation cannot be considered homogeneous institutions, and entry into one or the other, as well as the presence or absence of a transition between the two, can be socially patterned. In Western societies, marriages have become more selective of advantaged individuals from a socioeconomic view, whereas cohabitations tend to be more diverse, and include couples facing greater economic uncertainty (Billari & Esping-Andersen, 2015; Lundberg et al., 2016). Therefore, differences in instability

across union types may also mirror some compositional differences of couples entering partnerships, and possible associations between female income dominance and dissolution risk have to be understood within partnership forms which are themselves shaped by socioeconomic selection or evolving patterns related to marriage as an institution.

Recent work has, moreover, advanced a gender institution perspective, with a new focus on gender identity and norm-based mechanisms. Instead of treating earnings only as economic resources, relative income is increasingly conceived also as a signal of gender performance. It has been shown that couples may react strategically to deviations from breadwinning norms by adapting their working or domestic behavior to maintain conformity with prevailing gender expectations (Bertrand et al., 2015; Sayer et al., 2011). Dissatisfaction with partnership and divorce outcome linked to income asymmetry may thus reflect role conflict and identity concerns instead of a narrative of failed economic interdependence. Episodes where women out-earn their male partners can potentially entail new forms of conflict and renegotiation within the relationship with consequences of female breadwinning likely to depend on contextual norms, cohort attitudes or meaning attributed to new structures of income arrangement (Killewald, 2016). Under this framework, female income dominance could have either destabilizing effects if it challenges expectations around masculinity and provision or even stabilizing effects among couples with egalitarian attitudes, particularly among recent cohorts where dual-earners are more common. Relative income asymmetry may prompt adjustment mechanisms that depend on the normative context and partners' expectations, with now a growing literature supporting the argument that earnings are one of many factors influencing dissolution of unions. A recent machine learning analysis of German panel data found that partners' life satisfaction and other intra-household factors are among the most important deter-

minants of union dissolution, alongside employment and union status (Arpino et al., 2022).

In sum, while relative earnings are materially defined, their consequences may operate also through psychological channels. If departures from traditional breadwinner norms generate role renegotiation or identity strain, these processes may manifest in elevated subjective distress, which is also associated with relationship instability but rarely formally tested (Røsland et al., 2013). Whether distress works as a mediator of female breadwinner models or represents an independent risk factor remains a relevant open question. In light of the above discussion emphasizing the ambiguity of female income dominance, it is thus worth considering not only whether but also under what conditions relative income asymmetries correlate with union duration, and whether subjective distress may act as pathway to dissolution. Much of the empirical literature on income asymmetries union dissolution has focused on the United States; we propose the UK as a valuable comparative setting, which combines relative high female labor force participation and detailed longitudinal data tracing both relative earnings and partnership transitions. The study provides an account of union dissolution patterns by relative income ratio configurations, tests whether likelihood of dissolution is tied to periods of female income dominance and operationalizes the mediation pathway of subjective distress as a general predictor for dissolution.

4 Data

To our knowledge, there is no study that has previously tested the assumption of an existing association between the classification of family type based on partners' relative earnings and union dissolution using the UK household longitudinal study. The UKHLS provides detailed information on the social and economic circumstances of individuals and households in the United Kingdom. It combines 18 waves of the British Household

Panel Survey (1991–2008) and 9 subsequent waves of Understanding Society (2009–2019), following nationally representative samples over time. As such, it is one of the most consolidated study, both spatially and temporarily, in the whole continent as it follows representative samples of the population over time and covers all areas of the United Kingdom (Taylor, 2010). In addition to annual survey data up to UKHLS Wave 9 containing individual and household information, we exploit the Understanding Society Marital and Cohabitation Histories dataset which provides retrospective start and end dates for marriage and cohabitation spells in a format suitable for event-history analysis (University of Essex, Institute for Social and Economic Research, 2019, 2023).

One of the key strengths of longitudinal surveys is the ability to monitor changes in the life course of individuals and therefore to link changes in partnership trajectories to economic conditions. Specifically, the UKHLS part of the data set is rich in terms of information on sources of income. Variables for labor and total net income are available at the individual and household level, and inconsistencies in reporting are systematically corrected for. While the earlier waves of the BHPS dataset include less detailed information on income and only provide labor income consistently, the combined dataset enables harmonized variables to be constructed over almost three decades.

Our analysis focuses on respondents’ first co-residential union, whether initiated as cohabitation or marriage. For individuals who cohabited and subsequently married the same partner, the cohabitation and marriage phases are combined into a single union spell. The duration of the union is therefore measured from the date of first co-residential partnership to the date of separation, divorce, or breakup, or to the last observation if the union did not dissolve. Therefore, respondents with multiple union spells are retained only

with respect to their first observed partnership and unions which ended in widowhood or decease of the respondent were excluded since the focus is on dissolution through separation or divorce. We also kept only one respondent per couple to avoid double counting.

The dependent variable is the waiting time (in months) from union formation to dissolution (i.e. separation, divorce, or the termination of partnership). Unions still ongoing at the last interview are treated as right-censored. To avoid including extremely short-lived partnerships that may reflect very transitional living arrangements, unions lasting fewer than six months were excluded from the baseline specification (about 6% of the eligible first unions). Some unions were formed before the start of panel observation in 1991. Because the partnership histories provide retrospective start and end dates, the full duration of these unions is observed. The baseline specification retains all these unions in order to maximize statistical power and exploit complete duration information. Since the covariates are measured instead only during the period of survey participation, in order to evaluate whether the retrospective measurement on relationship spells affects the results later models restrict the sample to unions belonging to the observation window. Our main explanatory variable is defined as the female relative income ratio, computed as the woman's labor income divided by total household labor income across waves. The variable is classified according to the average relative income ratio within each union into three broad categories: male breadwinner (female share 0–0.4), (mostly) egalitarian (0.4–0.6), and female breadwinner (> 0.6). A rich set of covariates is added in the piecewise models. Along more standard indicators on household and individual characteristics (e.g. education, job status, area of residence, number of household members), we include a measure of self-reported distress based on the 12-item General Health Questionnaire

(GHQ-12). Such instrument captures psychological strain linked to feelings of worry, stress and depression and is reported on a 1-36 Likert scale, with higher values corresponding to greater distress. Its repeated administration is distinctive of UK data in both its BHPS and UKHLS components, meaning that psychological wellbeing is tracked longitudinally along with more socioeconomic dynamics, although similarly not available in every wave.

The data preparation process has ensured the accomplishment of a single time-event structure containing at least one observation for the main explanatory variable, with the final analytical sample consisting of 7,826 unique respondent-spells observations in the baseline model.

5 Main Findings

5.1 Descriptive statistics and preliminary evidence

Table 1 reports descriptive statistics by the family arrangements. Unions are classified according to the woman’s average share of household labor income recorded across the survey waves, irrespectively of whether the male or female partner is the respondent retained in the final sample. Male breadwinner unions constitute the majority of the sample (54%), followed by equal shares of (mostly) egalitarian and female breadwinner unions (23%). Dissolution rates differ considerably across the groups, with male and egalitarian unions displaying rather equivalent dissolution rates (around 16%) while female breadwinner unions are characterized by higher instability (26% incidence of separation or breakup). The mean duration in months of the unions follow accordingly. In terms of status of the union, the female breadwinner arrangements are more frequently observed in cohabiting partnerships (30%), as compared to male breadwinner unions

(21%), and are also less likely to involve children, indicating that part of the instability could be potentially driven by the commitment attached to the partnership. Moreover, female breadwinner unions seem not to be characterized by younger cohort or higher educational attainment of the respondent. The similar average birth year and the modest differences in education across groups imply the higher dissolution rate observed in female breadwinner unions is unlikely to be driven by simple compositional differences in age or education. Lastly, female breadwinner arrangements tend to include a higher share of male respondents and a slightly larger proportion of individuals not born in the UK. These descriptive differences motivate the need for multivariate models to better assess the existence of instability gaps by family arrangements.

Table 1: Descriptive Statistics by Family Type

	Male breadwinner	Egalitarian	Female breadwinner	Total
Number of unions	4,202	1,839	1,785	7,826
Dissolution rate (%)	16.0	15.4	26.2	18.2
Mean duration (months)	220.4	197.2	182.2	206.3
Cohabitation (%)	20.5	26.4	30.4	24.2
Marriage (%)	79.5	73.6	69.6	75.8
Ever had children (%)	75.0	67.0	64.7	70.8
Mean birth year respondent	1964.5	1966.1	1965.9	1965.0
Male respondent (%)	34.8	44.6	52.7	41.2
Born in UK (%)	79.3	83.4	78.6	80.1
Mean age left education	18.3	18.8	18.5	18.4

Before moving to the models, we further examine non-parametric Kaplan–Meier survival estimates to assess whether union stability differs across family types. This method allows us to compare the timing of dissolution without the need to impose parametric assumptions, thereby delivering a first descriptive benchmark. Figure 1 presents Kaplan–Meier survival estimates by family type as we limit the distortions from the small

number of very long-lasting partnerships, censoring at the 75th percentile of the observed union duration (335 months or ≈ 28 years). We can note how the survival curve of female breadwinner types diverge early in the duration of partnerships and remain clearly separated throughout the time window. Relatedly, the female breadwinner category is marked by substantially lower survival probability relative to both the other two categories which follow rather overlapping trajectories. This difference emerges already within the first year of union formation. After 12 months, approximately 5.5% of female breadwinner unions have dissolved, compared to 2.7% of male breadwinner and 2% of egalitarian unions. Moreover, the gap tends to widen over time: the cumulative dissolution reaches the rate of 14% among female breadwinner unions after 5 years, compared to more modest figures around 6% among the two counterparts. In ten years time, about one fifth of the total female breadwinner unions come to an end, compared to just 8.5% and 7.4% of the other two family arrangements. Indeed, a log-rank test strongly rejects the null hypothesis of equal survivor functions across family types ($\chi^2 = 195.09$, $p < 0.001$). It results in female breadwinner unions experiencing more failures than expected under the assumption of equality of survival, whereas the male breadwinner and the egalitarian unions show fewer events than expected. This pattern can be considered robust to alternative truncation configurations: as we censor unions after 10 or 15 years, the ordering and shape of the survival functions remain unchanged, together with the rejection of the null hypothesis in the log-rank tests as reported in the Appendix. This provides some confidence that an instability gap for female breadwinning seems to emerge relatively early, and it does not appear to be driven by a limited subset of longer unions.

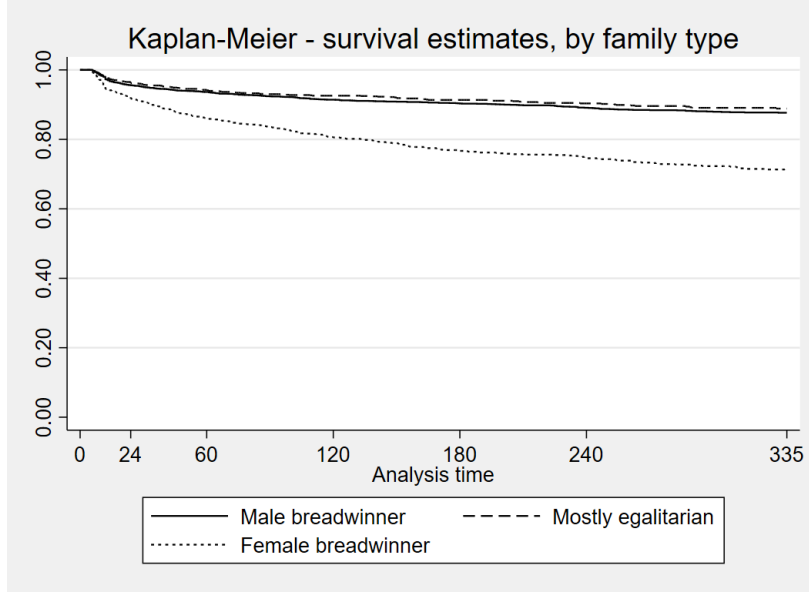


Figure 1: K-M estimates survival functions by different family types

While the previous plot has showed that female breadwinner unions tend to experience higher dissolution risks overall, it remains unclear whether the pattern is uniform across all forms of female income dominance. The division in three categories may in fact conceal some meaningful heterogeneity within female-led family types, and in particular between more moderate and more extreme relative income configurations. We therefore disaggregate the female relative income into 5 different brackets and present new Kaplan-Meier estimates across the redefined sample groups. The survival curves clearly illustrate how the earlier divergence was primarily driven by the most unconventional unions, where the female partner contributes at least 80% of household income. Those unions that are instead characterized by a more moderate female income dominance closely resemble the traditional male breadwinner during the first 10 years of partnership duration. After that threshold, male breadwinner family types with the lowest female ratio start to display a relatively higher dissolution risk. There is thus some indication that the risk does not increase linearly

with the female income ratio, but may concentrate in the most asymmetrical configurations.

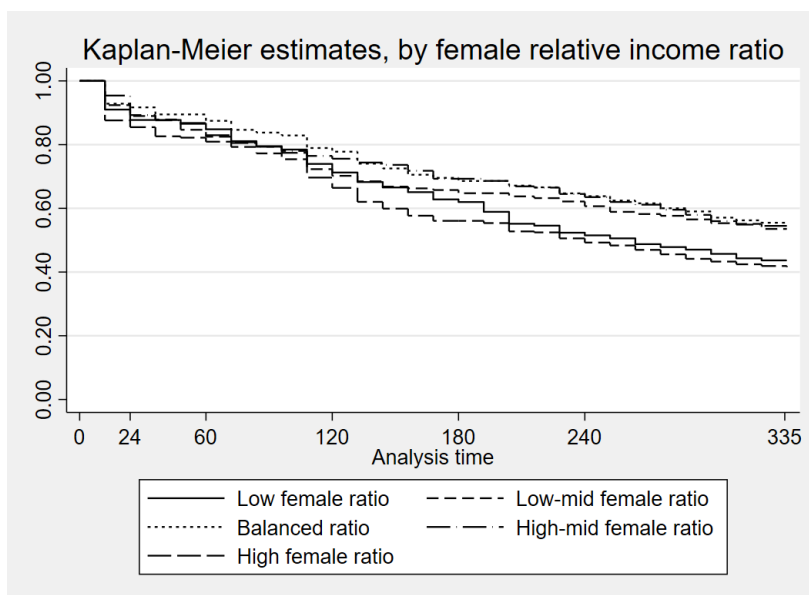


Figure 2: K-M estimates survival functions based on female relative income brackets

5.2 Baseline semi-parametric models

Albeit flexible, the nonparametric models present the issue of reaching quite fast a point of saturation as the number of stratifying factors, which one intends to control for, rises. At that stage, comparing survival functions would become pointless due to the exiguous number of observations left. Even when still feasible, it would be even cumbersome to interpret the graphs. That is why, in the following, we turn to semi-parametric models that enable estimating separate coefficients for the covariates of interest. To assess whether the differences in union stability documented in the Kaplan–Meier estimates persist, we estimate Cox proportional hazard models of union dissolution. The Cox model allows us to compare relative hazards across family types while leaving the baseline hazard unspecified. We begin with a baseline specification

including only family type and then introduce key demographic controls, including the union type, cohort, migration background, parenthood status, and age of leaving education.

Table 2 includes such Cox proportional hazard estimates of union dissolution. In the baseline model, female breadwinner unions face about double the hazard of dissolution relative to male breadwinner unions ($HR = 2.02, p < 0.01$), whereas (mostly) egalitarian family types seem not to differ from the reference category. When introducing the demographic controls instead, the hazard ratio associated with female breadwinner unions declines but maintains a sound significance ($HR = 1.49, p < 0.01$). This stands to indicate that even when accounting for compositional differences between couples, a relevant residual association remains. In this second specification, even egalitarian unions become modestly more stable than male breadwinner unions ($HR = 0.80, p < 0.05$). As expected, both marriage and parenthood are protective statuses, while younger birth cohorts also face significantly higher hazards of union dissolution. Respondents who are female and not born in the UK report lower dissolution risks, while the age of leaving education does not seem to account for differences in stability. When we allow the association between family type and dissolution to vary by the union type, the instability of female breadwinner arrangements become more evident within marriages. Even if marriage is found to be generally protective against dissolution, the relative disadvantage of the female breadwinner unions seems to widen among married couples. The elevated instability of female breadwinner unions could not be explained by the inclusion of key invariant demographic characteristics and is not confined to less institutionalized partnerships in our first empirical specifications.

Table 2: Cox Proportional Hazard Models of Union Dissolution

	(1) Baseline	(2) Controls	(3) + Interaction
Family Type <i>ref. Male Breadwinner</i>			
(mostly) Egalitarian	1.05 (0.09)	0.80** (0.07)	0.81** (0.08)
Female breadwinner	2.02*** (0.14)	1.49*** (0.11)	1.37*** (0.12)
Union Type <i>ref. cohabitation</i>			
Marriage		0.12*** (0.01)	0.11*** (0.01)
Interaction Terms			
Egalitarian $\times$ Marriage			0.92 (0.20)
Female breadwinner $\times$ Marriage			1.47** (0.26)
Female respondent		0.83***	0.84**
Birth year		1.07***	1.07***
Not born in UK		0.85*	0.85*
No children		1.22***	1.21**
Age leaving education		0.99	0.99
Observations	7,826	7,279	7,279
Dissolution events	1,020	909	909

Robust standard errors in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

5.3 Piece-wise semi-parametric models

The previous models relied on the classification based on the average income configuration across the union. However, partners' relative earnings may fluctuate over time. A couple classified as female breadwinner on average may have experienced only a few intervals of female income dominance, while another may have maintained this arrangement for longer. Relying solely on the average relative income ratio therefore risks conflating more long-term arrangements with ones that experience a greater number of fluctuations. To assess whether the observed instability gap aligns with structural arrangements or specific periods of female income dominance and to include more controls, we next estimate models incorporating time-varying measures of partners' relative income. To this end, we restructure the union spells into yearly intervals that match the survey observation period (1991-2019) and perform left-truncation, whereby unions enter the risk set at time in which both the partnership is ongoing and income information is observed. This ensures that the exposure time and the covariates' measurement is fully aligned. Whenever the labor income is available, we classify the couple according to the relative income definition for the observed survey years. This approach enables to better investigate whether the risk of dissolution increases in those precise intervals in which the woman earns more than her partner, rather than among couples who are predominantly female breadwinner on average. These models estimate the hazard of dissolution at a given time, by allowing relative income and a few other covariates such as the qualification and job status of the respondent, the number of components in the household, the area of living, an income below or above the median, and a subjective measure of distress to vary over time within partnerships. The inspection of the interval data highlights some interesting patterns. Family arrangements tend to vary not only between couples, but also over time within couples, indicating that unions may experience

transitions in earnings ratios rather than keeping a fixed breadwinner structure. Moreover, female breadwinner episodes co-occur more often with the unemployment status of the respondent or an household income below the median, suggesting that these arrangements can often emerge in context of higher vulnerability. The subjective distress tend also to fluctuate across intervals but it is not strictly more typical of female breadwinner periods. These features underscore the importance of modeling the income configurations and related covariates at the interval level to better isolate plausible determinants of dissolution.

Table 3: Interval-Level Descriptives

Family type	Below median income (%)	Unemployed (%)	Mean life satisfaction (z)
Male breadwinner	36.79	2.18	-0.02
Mostly egalitarian	16.27	0.16	-0.05
Female breadwinner	57.12	4.48	0.03

Notes: Percentages refer to the share of intervals within each breadwinner configuration. Life satisfaction is standardized (z-score).

Table 4 reports the Cox hazard models that estimate the association between the time varying version of female relative income and union dissolution. The 5 progressive specifications enable to distinguish between the raw associations, some socio-demographic adjustments mirroring the previous empirical exercise, household characteristics, contextual controls related to the area of living and economic situation, and subjective wellbeing. Model 1 estimates the raw association, without additional controls. During observed periods in which couples are classified as mostly egalitarian, the hazard of dissolution is 26% higher than the reference (periods where couples are of the male breadwinner type). Female breadwinner periods also carry similar hazards ($HR = 1.22, p < 0.05$). As noted earlier for the average configuration, the departure from the male breadwinner model - this time not only confined to female income dominance, but also toward a more equal earning division - is associated with a more elevated instability. In model

2, we further include controls for union type, gender, birth cohort, place of birth, and education. As these characteristics are introduced, the egalitarian coefficient loses its significance, while female breadwinner effect remains similar in terms of magnitude and significance ($HR = 1.25, p < 0.05$). This indicates that the initial destabilizing effect of the egalitarian couples may have stemmed from compositional characteristics. Female income dominance instead seems to retain an independent association even when demographic and educational characteristics are accounted for. Marriage remains the strongest protective factors across specifications, as it reduces the hazard of dissolution by nearly 75-80% relative to cohabitation. Its effect is considerably larger than the ones observed for the kind of within-couple income arrangements. Model 3 adds the number of children in the household as well as the household size. The female breadwinner association slightly strengthens, while the one for the egalitarian family types remains neutral. If anything, the number of children is weakly associated with higher dissolution risk, but the effect is unstable to subsequent specifications. Model 4 further incorporates a dummy variable for below-median household income in the 12 months interval, a categorical variable for job status and controls for the area of residence. These additions do not alter the female breadwinner association, and egalitarian temporal arrangements do not statistically differ from the breadwinner ones. Neither the household income position nor the employment status seem to underlie the greater instability observed in periods of female income dominance. Lastly, in Model 5 we inspect the role of subjective distress by adding a standardized time-varying form. For every standard deviation increase in subjective distress, the hazard of union dissolution increases by a significant 18.7%. In this case only, the female breadwinner declines and becomes statistically non-significant ($HR = 1.19, p = 0.105$), while egalitarian arrangements remain neutral. When Model 4 is estimated on the same restricted sample of Model 5, the female breadwinner coefficient

remains marginally significant, therefore the attenuation observed can be associated with the inclusion of subjective distress rather than with the sample restriction itself. This attenuation could be consistent with the idea that instability in unions where the female partner is the primary earner may be linked to elevated levels of distress reported by at least one partner, rather than inherently motivated by economic asymmetries. When adjusting for perceived distress, the relative income configuration no longer acts as a significant predictor of likelihood of dissolution. Moreover, when we further disaggregate the relative income ratio into the 5 finer brackets, we obtain that when in the main models an higher dissolution risk was reported (Models 1-4), this primarily concerned the most unconventional configuration, in which the female partner contributes the largest share of household income. More moderate forms of female income dominance were not found to differ from the male breadwinner reference, consistently with earlier nonparametric results (Table A2, Appendix).

In order to ensure that these main findings do not depend on the semi-parametric approach of the Cox specification, we re-estimate Models 1, 2, 4 using exponential proportional hazard models. Contrarily to the Cox regressions that leave the baseline hazard unspecified, the exponential model assumes a constant hazard over time and imposes a stronger parametric structure on the duration. Reassuringly, the results displayed in Appendix (Table A3) leave the interpretation largely unchanged. The hazard ratios for the female breadwinner arrangements remain positive and similar in magnitude, with only a marginal statistical significance recorded in the equivalent of Model 5. Other than that, subjective distress continues to show a robust association with the dissolution risk, whereas marriage and earlier birth cohorts are protective predictors. This close similarity between Cox and exponential specifications confirm that the core findings do not result

from the flexible hazard design in the Cox model and strengthen confidence of their validity under an alternative distributional assumption. We also test the proportional hazards assumption for the main Model 4. Schoenfeld residual tests point to the absence of violations for the relative income variable; and models that include interactions between female breadwinner model and log duration yield insignificant coefficients for the interaction terms. When a modest association between female income dominance and the dissolution risk emerges, this is not to be considered duration-dependent.

As an additional check, we assess whether an association between female breadwinner arrangement and union dissolution emerges by the income position of the household or by the partnership status. In both interaction models we found no confirmation for the existence of such moderating effects. In table 5, the interaction between the FBW status and marriage is not statistically significant ($p = 0.178$), so the effect was not found to vary between married or cohabiting unions. Likewise, the interaction between the FBW status and the below-median income is imprecisely estimated (Table 6, $p=0.768$). Nonetheless, both models continue to report subjective distress as a meaningful and strong predictor of dissolution, with the same hazard ratios of the main full model ($HR \approx 1.19, p < 0.01$).

Table 4: Piecewise Cox Models of Union Dissolution

	Model 1	Model 2	Model 3	Model 4	Model 5
Family type (<i>ref. Male breadwinner</i>)					
Mostly egalitarian	1.26** (0.12)	1.02 (0.10)	1.04 (0.10)	1.04 (0.10)	1.04 (0.11)
Female breadwinner	1.22** (0.12)	1.25** (0.12)	1.27** (0.13)	1.24** (0.13)	1.19 (0.13)
Marriage (<i>ref. Cohab</i>)					
Marriage	–	0.23*** (0.02)	0.22*** (0.02)	0.22*** (0.02)	0.23*** (0.02)
Female respondent	–	1.09 (0.09)	1.08 (0.09)	1.09 (0.09)	1.05 (0.09)
Birth year	–	1.08*** (0.01)	1.08*** (0.01)	1.07*** (0.01)	1.08*** (0.01)
Not born in UK	–	1.10 (0.15)	1.11 (0.15)	1.10 (0.15)	1.17 (0.17)
Highest qualification (<i>ref. no qualifications</i>)					
Degree	–	1.06 (0.19)	1.07 (0.19)	1.09 (0.20)	1.10 (0.21)
Upper secondary	–	1.01 (0.17)	1.02 (0.17)	1.04 (0.18)	1.03 (0.19)
Lower secondary	–	1.10 (0.19)	1.11 (0.19)	1.09 (0.19)	1.07 (0.19)
Children in the hh	–	–	1.10* (0.07)	1.10 (0.07)	1.05 (0.07)
HH size	–	–	0.95 (0.04)	0.96 (0.04)	0.98 (0.05)
Below median income	–	–	–	1.09 (0.09)	1.11 (0.10)
Job status	–	–	–	✓	✓
Region	–	–	–	✓	✓
Subjective distress score GHQ (std.)	–	–	–	–	1.19*** (0.04)
Subjects	5,131	4,966	4,966	4,962	4,671
Dissolution events	643	618	618	616	570

The table reports throughout hazard ratios with Standard Errors clustered at the individual level in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table 5: Interaction between female breadwinner status and union status (Extension Model 5 – Table 4)

	HR	SE	p
Female breadwinner (ref. male)	1.07	0.14	0.623
Marriage (ref. cohab.)	0.23	0.03	<0.01
FBW $\times$ Marriage	1.32	0.27	0.178
Subjective distress (std.)	1.19	0.04	<0.01

Subjects = 4,671; Dissolution events = 570.

Standard errors are clustered at the individual level.

Table 6: Interaction between female breadwinner status and below-median income
(Extension Model 5 – Table 4)

	HR	SE	p
Female breadwinner (ref. male)	1.21	0.20	0.234
Below median income	1.21	0.14	0.110
FBW $\times$ Below median income	0.94	0.20	0.768
Subjective distress (std.)	1.19	0.04	<0.01

Subjects = 4,671; Dissolution events = 570.
Standard errors are clustered at the individual level.

As noted earlier, the Cox specification from the Model 5 in Table 4 allowed us to observe that the inclusion of the distress indicator attenuates the significance of the temporary FBW family arrangement, however a simple coefficients comparison in nonlinear survival models does not provide a formal decomposition of the effects. Moreover, as the main variable of interest and the supposed mediator of subjective distress are contingent on the same time interval, it remains possible that distress only reflects the anticipatory strain preceding the dissolution rather than the mechanism whereby female income dominance affects union stability. To assess more rigorously whether subjective distress accounts for part of the association between FBW family type and union dissolution, we implement a couple of Karlson-Holm-Breen (KHB) decompositions. The approach is particularly suitable in nonlinear models because it corrects for the rescaling problem that arises when comparing nested logit models, and allows to separate the association in its direct and indirect components.

The results of the first decomposition - where FBW family type and subjective distress are still measured in the same interval - provide only limited support for a mediation pathway (Table 7). While distress is clearly associated with the dissolution risk, it only explains a limited fraction (approximately 7%) of the association between female income dominance and partnership instability. The decomposition also clarifies that the attenuation is greater for the unemployment status of the respondent, as its coefficient declines more considerably in the full model where the subjective distress is included. In contrast, the coefficients for education, relative household income, number of children, household size, and region remain largely unchanged as the dissolution risk is accounted for. To test more accurately whether a pathway such that of female income dominance $\rightarrow$ subjective distress $\rightarrow$ dissolution exists, we further perform a temporally ordered test using our ex-

planatory of interest (FBW arrangement) at interval t-2, the proposed mediator (subjective distress) at t-1, and the outcome (dissolution event) at interval t. In the first OLS test of lagged $x \rightarrow M$, earlier female-breadwinner is not found to predict later levels of subjective distress with a small and insignificant coefficient. Other socio-economic covariates such as the female gender, the unemployment status of the subject or a below-median household income are instead associated with higher future distress. In the second logit model (lagged $M \rightarrow Y$), subjective distress remains a statistically significant predictor of later dissolution ($OR \approx 1.19, p < 0.05$). This second KHB decomposition further reveals the absence of an indirect effect via lagged distress of an earlier period of female relative dominance as well as the loss of a direct effect of the earlier FBW arrangements on the dissolution risk, once distress is accounted for (Table 8). This lagged decomposition is yet to be interpreted with caution, and not as conclusive proof, since both the relative female income ratio and the level of distress are recorded sparsely throughout the available time intervals with the overall number of observations dropping considerably from 21,536 to 12,699.

Table 7: KHB decomposition of the association between female breadwinner family type and union dissolution

	Coef.	SE	p-value	OR	95% CI (OR)
<i>Female breadwinner (vs. male breadwinner)</i>					
Total (reduced)	0.242	0.127	0.057	1.27	[0.99, 1.63]
Direct (full)	0.225	0.127	0.078	1.25	[0.98, 1.61]
Indirect (diff)	0.018	0.066	0.788	1.02	[0.89, 1.16]
<i>Unemployed (vs. employed)</i>					
Total (reduced)	0.224	0.280	0.423	1.25	[0.72, 2.17]
Direct (full)	0.121	0.281	0.667	1.13	[0.65, 1.96]
Indirect (diff)	0.103	0.069	0.137	1.11	[0.97, 1.27]
<i>Mediator (full model)</i>					
Subjective distress	0.202	0.045	<0.001	1.22	[1.12, 1.33]

Notes: Discrete-time logit models with 12-month time bands.

Controls include partnership status, sex, birth cohort, education, income below the median, number of children, household size, job status, and area of residence.

Standard errors clustered by respondents.

Table 8: Temporal mediation test: $X_{t-2} \rightarrow M_{t-1} \rightarrow Y_t$

	Coef.	SE	p	OR	95% CI (OR)
<i>$X_{t-2} \rightarrow Distress_{t-1}$ (OLS)</i>					
Female breadwinner (vs male)	0.050	0.036	0.164	–	–
<i>$Distress_{t-1} \rightarrow Dissolution_t$ (Logit)</i>					
Subjective distress	0.172	0.076	0.023	1.19	[1.02,1.40]
<i>KHB mediation decomposition</i>					
Total (reduced)	-0.218	0.201	0.277	0.80	[0.54,1.19]
Direct (full)	-0.223	0.201	0.267	0.80	[0.54,1.19]
Indirect (diff)	0.004	0.066	0.947	1.00	[0.88,1.14]

Notes: Exposure measured at $t - 2$, mediator at $t - 1$, and dissolution at t .

Sample restricted to intervals with observed lagged exposure and mediator.

Standard errors clustered by respondent.

6 Discussion and Conclusions

This chapter has investigated whether income inequality within couples, and especially those arrangements where women out-earn their male partners, is positively associated with union instability in the United Kingdom. The survival analysis was motivated by a long-standing debate on whether women’s economic autonomy increases the feasibility of exit from unsatisfactory unions (independence effect), whether a second source of income is effective in stabilizing partnerships due to greater economic security (income effect), and whether departures from the most conventional male breadwinner model may generate relational strain or adaptation costs (gender-institution perspective). While these mechanisms are acceptedly complex to isolate empirically, the UK long-established panel enables us to get a closer inspection of how family arrangements and first partnerships unfold over time, and whether psychological processes could be linked to partnerships’ dissolution in this setting.

Preliminary descriptive statistics and Kaplan-Meier estimates based on the average family arrangement have highlighted that unions classified as female breadwinner face higher instability patterns, and that this observed divergence is largely led by the most asymmetric unions where women contribute the largest share of the household labor income. Cox models using the time-invariant family typology aligned closely, with female breadwinner unions remaining at higher risk of dissolution after accounting for a limited set of demographic and partnership characteristics. When switching to interval-based specifications with fuller alignment between measurement of income and observed spells, the results got more nuanced. In these new Cox models, periods of female income dominance were associated with a modestly increased dissolution risk relative to male breadwinner periods (approximately 20–25% in Models 1–4), whereas periods

of egalitarian arrangements generally did not differ from the reference. Consistent with the nonparametric evidence, the disaggregation of relative income in granular brackets confirmed that the dissolution risk was concentrated among the most extreme episodes of female dominance.

One central result concerned an indicator of subjective distress which was found to be a robust predictor of dissolution in the full model. The association between the female breadwinner model and dissolution became imprecisely estimated once distress was included, in accord with the idea that psychological well-being captures a dimension of vulnerability only partially proxied by other socioeconomic controls. At the same time, KHB decomposition analyses have provided limited evidence for the existence of a mediation pathway from female income dominance to dissolution via distress. In the lagged specification, we were unable to establish a causal chain where female income dominance increases distress which then leads to dissolution. These results could suggest that while subjective distress is clearly linked to dissolution, this is not uniquely produced by female income dominance, with the attenuation of this latter open to different interpretations. For instance, precisely those unions which experience extreme patterns in favor of female earnings may be exposed to employment- or health-related shocks that simultaneously affect distress and dissolution. Moreover, distress can also reflect the relational or role adjustment behaviors missing from this analysis rather than acting as mediator of income asymmetry *per se*.

Overall, this study makes three contributions. First, It adds evidence from survival analysis by family types (as defined on partners' relative income ratio) on the UK context that has received less attention relative to the US-dominated debate. Second,

by distinguishing between average and time varying income arrangements, it allows to better clarify whether dissolution risk is more elevated during periods of female income dominance. Third, it integrates a measure of subjective distress and evaluates its role using decomposition techniques. While the mediation evidence is limited, distress is found to be a relevant correlate of dissolution and accounting for it fundamentally changes the association between family arrangements based on relative earnings and partnerships' instability. Yet, the analysis suffers from some limitations that are complex to deal with the current dataset. Despite careful interval construction and left-truncation of partnerships, earnings configurations may also adjust in response to internal relationship dynamics or anticipatory separation, so reverse causality cannot be fully ruled out. Constructs that may be central to the gender and bargaining perspective about gender attitudes, relationship quality or division of unpaid work are available only in rotation modules for a limited sets of waves from 2009 onward, limiting the feasibility of their integration without a substantial loss of observation periods. Even if the distress measure is more consistently reported than relationship satisfaction, it may correlate with other short-term shocks and does not necessarily capture a dyadic process. Lastly, the sourced indicator of labor income does not reflect the division of assets or the total household resources, and we controlled for partnership status but did not model selection into marriage over cohorts.

Despite these constraints, these findings provide a clearer picture of how relative earnings relate to union stability in the UK. While the female breadwinner model predicts higher hazards of dissolution in several specifications, this association is somewhat modest in interval-based models and insignificant when distress is included. This latter emerges as a strong predictor of dissolution but unlikely to mediate any assumed association between female breadwinning and partnerships' dissolution. Rather than pinpointing a single

dominant mechanism specific to the female breadwinner models, the results suggest that the outcome of dissolution lies at the intersection of extreme earnings asymmetries, socioeconomic circumstances and psychological processes, with distress capturing a dimension of vulnerability that could partially overlap with, but not solely confined to, episodes of female income dominance.

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Appendix

A Additional Kaplan–Meier Estimates and log-rank tests

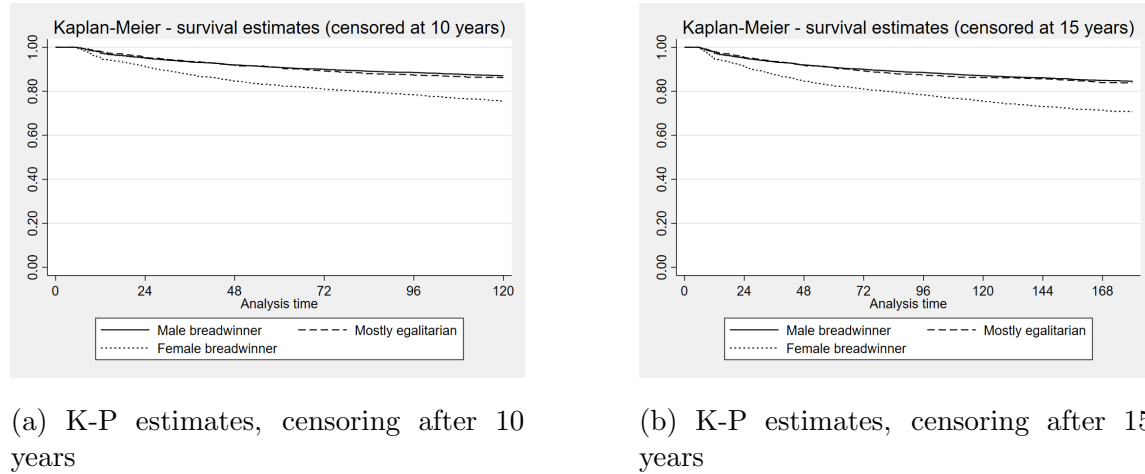


Figure A1: Survival functions with alternative censoring configurations

Table A1: Log-rank tests under alternative censoring thresholds

Censoring	Family Type	Observed	Expected	Obs. - Exp.
10 years (120 months)	Male breadwinner	464	567.23	-103.23
	(mostly) Egalitarian	209	242.26	-33.26
	Female breadwinner	347	210.51	+136.49
	Log-rank test	$\chi^2(2) = 112.23, p < 0.001$		
15t years (180 months)	Male breadwinner	532	648.64	-116.64
	(mostly) Egalitarian	234	275.14	-41.14
	Female breadwinner	395	237.22	+157.78
	Log-rank test	$\chi^2(2) = 132.49, p < 0.001$		

Notes: Observed and expected numbers of dissolution events from log-rank tests comparing survivor functions across family types.

B Additional Cox Models

Table A2: Cox Models of Union Dissolution by Relative Income Brackets

	Model 4	Model 5
Relative income brackets (<i>ref. Low female ratio</i>)		
Low-mid female ratio	0.85 (0.10)	0.85 (0.11)
Balanced ratio	0.94 (0.11)	0.95 (0.12)
High-mid female ratio	0.98 (0.16)	0.96 (0.16)
High female ratio	1.28** (0.17)	1.21 (0.17)
Marriage (<i>ref. Cohab</i>)		
Marriage	0.22*** (0.02)	0.23*** (0.02)
Female respondent	1.09 (0.09)	1.06 (0.09)
Birth year	1.08*** (0.00)	1.08*** (0.01)
Not born in UK	1.10 (0.15)	1.17 (0.17)
Highest qualification (<i>ref. no qualifications</i>)		
Degree	1.11 (0.21)	1.11 (0.21)
Upper secondary	1.04 (0.18)	1.04 (0.19)
Lower secondary	1.10 (0.19)	1.08 (0.19)
Children in the hh	1.08 (0.06)	1.04 (0.06)
HH size	0.96 (0.04)	0.99 (0.05)
Below median income	1.04 (0.09)	1.07 (0.10)
Job status	✓	✓
Region	✓	✓
Subjective distress (std.)	–	1.19*** (0.04)
Subjects	4,962	4,671
Dissolution events	616	570

Hazard ratios with SE clustered at the individual level in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

C Alternative parametric models

Piece-wise exponential models are run as robustness check on the data sample split on 12-months intervals, where $exp(x_{ij})$ is the relative risk for an individual with covariate values x at interval j , compared to the baseline, in any given time interval.

$$\lambda_{ij} = \lambda_j exp(x'_{ij}\beta)$$

Among the time-varying covariates, we include both characteristics of the household (number of household members, number of children in the household, area of residence, relative income position) and individual features such as job status of the respondent, standardized, level of subjective distress, highest qualification achieved.

Table A3: Exponential Proportional Hazards Models of Union Dissolution (Robustness Check)

	Model 1	Model 4	Model 5
Family type (<i>ref. Male breadwinner</i>)			
Mostly egalitarian	1.23** (0.12)	1.09 (0.11)	1.09 (0.12)
Female breadwinner	1.20* (0.12)	1.29** (0.13)	1.24* (0.14)
Marriage (<i>ref. Cohab</i>)			
Marriage	—	0.17*** (0.02)	0.17*** (0.02)
Female respondent	—	1.08 (0.10)	1.05 (0.10)
Birth year	—	1.04*** (0.00)	1.04*** (0.00)
Not born in UK	—	0.99 (0.14)	1.07 (0.16)
Highest qualification (<i>ref. no qualifications</i>)			
Degree	—	0.85 (0.17)	0.84 (0.18)
Upper secondary	—	0.96 (0.18)	0.96 (0.19)
Lower secondary	—	1.09 (0.20)	1.07 (0.21)
Children in the hh	—	1.18*** (0.07)	1.13* (0.07)
HH size	—	0.92* (0.05)	0.94 (0.05)
Below median income	—	1.09 (0.10)	1.13 (0.11)
Job status	—	✓	✓
Region	—	✓	✓
Subjective distress (std.)	—	—	1.15*** (0.04)
Subjects	5,131	4,962	4,671
Dissolution events	643	616	570

Hazard ratios with SE clustered at the individual level in parentheses.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Response to Reviewers' Comments on Paper 2

Elena Marseglia

1. Introduction

Reviewer 1

The introduction should include a clear statement of this chapter's contribution. You are working with a genuinely impressive dataset that enables cross-country comparative analysis on an important policy issue, yet this strength is not adequately highlighted. Please make a more compelling case for why this dataset is particularly valuable for addressing your research question, and be explicit about which specific limitations or gaps in the existing literature your analysis addresses.

Response:

In the revised introduction we now provide a more explicit discussion of the contributions that the paper makes to the literature. Unlike most existing research that treats young adults as a broad demographic group, we focus specifically on those independent young adults who have exited the family of origin. It allows to isolate a subgroup more likely to be exposed to market and institutional risks and to provide a more precise assessment of early-adulthood vulnerability. The analysis overcomes the limitations of short-term or outdated studies by focusing on a long cross-national analysis spanning 20 years (1999-2020). The attempt of the chapter is to go beyond documenting poverty trends and attempt to study micro-macro links by combining regression analysis, predictive margins, quantile regressions, and Oaxaca-Blinder decompositions. It constitutes an integrated framework that has never been presented in cross-national analyses of young adults. [pages 2-3].

2. Data and methods

Reviewer 1

The use of a uniform under-30 cut-off to define young household heads may not capture equivalent stages of independence across countries, as transition to adulthood is very context-specific. I recommend you to provide a clearer rationale for the age 30 threshold. At present, the text notes that the focus is on "young adults under age 30 who have begun independent living," but does not fully explain why 30 is the appropriate cut-off. You may try different conceptualisations of "young adults" as robustness test. If that's not possible, I recommend you to discuss what are the

implications, particularly regarding generalizability, of your findings.

Reviewer 2

- A possible concern is that the use of a uniform under-30 cut-off to define young household leads may not capture equivalent stages of independence across countries. In contexts such as Italy, Spain, or Ireland, very few individuals establish independent households before age 30, so the under-30 group may represent a highly selected minority. By contrast, in Nordic countries, where residential independence typically occurs much earlier, the same age threshold captures a much larger and more representative share of young adults. This asymmetry could affect cross-national comparability, and it would be helpful to acknowledge and discuss this limitation more explicitly than you do (check page 2 and page 7 of paper 2). I would encourage you to clarify the rationale behind setting the age threshold at 30. At present, the text highlights that the focus is on “young adults under age 30 who have begun independent living,” but it does not fully explain why 30 is the appropriate cut-off.

- It is not entirely clear how you handle the fact that some covariates (e.g., subnational region, housing tenure) are missing in the first wave for certain countries (par. 4, page 6, second paper). Do you simply exclude these variables when they are unavailable, or do you adopt another strategy? As it stands, the models may not be fully comparable across waves, and this potential limitation should be explained more explicitly.

- It would be useful to specify that both the square root scale for equalization and the 50% of median poverty threshold are standard conventions in the LIS/OECD literature, so that the choices do not appear arbitrary .

- [...]By contrast, young adults who remain in the parental home or who are partners but not designated as the household lead are excluded. This selection mechanism is important for interpreting your findings, and I would encourage you to make it crystal clear in the text. Related to this, one concern is that the composition of the group “young household leads” may not be stable over time. The descriptive evidence shows that fewer young adults form independent households overall, and among those who do, an increasing share are single-person households rather than couples. This means that in earlier waves the under-30 household leads included more couples (with or without children), whereas in later waves they are increasingly singles living alone. This shift induces a variable sample selection across periods, which may partly account for the observed rise in poverty risks. I suggest making this issue explicit in the discussion, as it affects the interpretation of temporal changes. This phenomenon does not seem to be fully addressed in the Oaxaca–Blinder decomposition section. Although the decomposition separates compositional from structural effects to me there seems to be little explicit discussion of how the shifting balance between singles and couples among young household leads may itself drive the results. If this compositional change is

a key factor behind the rising poverty risks, it would be useful to highlight and interpret it more directly in that section.

- Even though I think your methods are sound, for future work, you might consider alternative strategies that could complement/extend the current approach. A pooled cross-country model with country fixed effects would make you exploit the full dataset in a single framework while controlling for unobserved, time-invariant country characteristics. This could also help circumvent the limitation that subnational region fixed effects are not always available across time and space. Also, multilevel models (with individuals nested in country-wave units) could capture variance components at both the individual and contextual level, which would make you quantify how much of the poverty penalty is attributable to country-level vs. individual-level factors. Although I think these approaches would not remove the data constraints entirely, they may provide additional leverage for addressing unobserved heterogeneity.

Response :

We have now included a short discussion of the threshold selection. The cut-off point of those aged under 30 follows the conventional definition used in much of the literature on the transition to adulthood from the early 2000s onwards, with the twenties generally defined as the critical period for school-to-work transitions, residential independence, and labor market instability (Buchmann & Kriesi, 2011). This cut-off point enables us to establish continuity with previous studies while focusing that on cohorts exposed to labor market instability. While this threshold helps maintaining comparability with earlier research and isolating the age group presenting a distinct social risk, in a small number of countries (Ireland, Italy and Spain) the share of household leads under 30 falls below 10 percent. For those countries only, we report in the Appendix a robustness analysis of the baseline model and of the last decomposition using the alternative threshold of 35 years, which indeed captures more leads and show that the results remain broadly unchanged. [pages 6-7, Appendix A2-A4].

We reran the baseline models using the same empirical specifications for all countries and describe it more clearly [page 14]. For the computation of margins within occupational status, instead we opt for more complete models that incorporate geographical controls, housing status and occupations where available [Appendix, A5-A7].

We specified that the equivalized scale is a standard practice when using LIS data [page 7].

We comment in the descriptive section that in most countries we indeed observe a lower average of household members. In this revised version of the paper the focus has shifted from family structure to favor instead the variation of the effects by occupational categories and housing burden (avoiding therefore stylized family types categorization). All the models throughout the paper

control for the number of household members and number of children. In the Oaxaca-Blinder decompositions, among the different covariates, we now devote some space to the results on the household characteristics which go often in the expected direction (e.g. contribute to poverty with shrinking number of household members, decrease of poverty with declining presence of children). The effects on the coefficients are instead more heterogeneous. We also suggest that the effects are relatively mild, and that the intercept in the unexplained components remain predominant in our decompositions [pages 31-32; 34-35].

We agree that pooled cross-country or multilevel models can offer additional insights into unobserved heterogeneity. However, because our main objective is to identify and interpret cross-national and temporal differences in age-related poverty penalties, we rely on country-specific linear probability models. This approach allows clearer comparisons across institutional contexts and avoids averaging effects that might mask meaningful cross-country variations, although relying on larger number of country-time observations would be beneficial.

3. Findings

Reviewer 2

In Table 7, the reported quantiles differ across countries (e.g., 0.9 and 0.3 for Australia, 0.9 and 0.4 for Austria, 0.9 and 0.1 for Canada, etc.). From the text, I infer that these are the points where the maximum and minimum statistically significant penalties were observed, but this is not made fully explicit. As it stands, the variation in quantiles may confuse readers, since one might expect results to be reported at the same deciles across all countries (e.g., 10th, 30th, 50th, 90th). For clarity, it would be very helpful to explain explicitly why different quantiles are shown in different contexts, and why the choice to report only the extreme statistically significant values is preferable to reporting a uniform set of quantiles. Maybe a small methodological note would help. . .

Response :

To improve clarity, we have replaced the Table with graphical plots (Figure 6) that display, for each country and period, the maximum and minimum statistically significant income penalties across the distribution. This presentation allows for clearer cross-country comparison while highlighting where age-related disadvantage is most pronounced. We do not report a fixed set of deciles as it would not add greater interpretative value, and some estimates could be insignificant. We describe within the text where in the distribution the penalties occur [pages 27-28].

4. Discussion

Reviewer 1

Discussion only summarizes results and lists limitations, and does not adequately make use of the potential of your cross-national data which give you the opportunity to develop a more elaborate discussion. Please develop a more substantive interpretation of why you observe the cross-national differences in your results. How might variation in welfare regimes, youth labour market policies, housing support systems, educational costs, or family support programs explain the patterns you find? Are your results consistent with what we would expect based on different welfare state typologies? Do certain policies appear more protective against the risks you document.

Response :

In the revised discussion section, we expanded the interpretation of the cross-national differences observed in the results. We more clearly relate the divergent patterns across European countries and other liberal welfare regimes to differences in the labor market structures, employment, occupational stratification and role of education. We elaborate on how, in several Continental European contexts, youth disadvantage appears more broadly distributed across employed groups and with weaker returns within lower-skilled occupations—whereas in liberal regimes employment and higher education continue to function as comparatively stronger protective factors despite existing structural pressures. We also discuss the contrasting case of Poland in light of its macroeconomic growth and structural transformation. This revision should better leverage the comparative potential of the cross-national results, while still acknowledging the limitations of our decompositions [Pages 38-41].

Understanding Changing Economic Conditions of Young Adults Over Time: A Comparative Analysis Using the Luxembourg Income Study

Elena Marseglia, Bocconi University

1 Abstract

This study investigates how the economic position of young adults has evolved across 15 high-income countries over two decades. It focuses on individuals aged 18 to 30 who have moved out of their parental homes and become heads of different households. We analyze changes in relative poverty and income inequality over the 2000-2020 time span of this population subgroup by combining descriptive statistics, pooled linear probability models, quantile regressions, and Oaxaca-Blinder decompositions on the harmonized data available through the Luxembourg Income Study (LIS). Our findings reveal a persistent poverty risk among young adult household heads across most advanced economies. In parts of Continental Europe, the poverty risk for young heads has increased and income penalties have become more concentrated at the lower end of the distribution, widening intergenerational gaps. In Southern Europe, high overall poverty levels across groups coexist with elevated penalties for young heads. By contrast, in many liberal regimes poverty rates among the young have remained relatively stable over time, albeit at comparatively high baseline levels. The observed patterns in the decomposition analyses are not driven by unfavorable demographic composition, but mostly by changes in returns to observed characteristics. In European societies, improvements in compositional effects are largely countered by deteriorating returns, in contrast to liberal societies elsewhere, where education and labor attachment continue to provide relatively robust protection.

2 Introduction

Over the last two decades, the transition to adulthood has become more fragmented and uncertain, especially in high-income countries. The traditional life-cycle model - which used to show a clear path from education to stable jobs, independent living, and family formation - is now less representative of the real experiences of many young people. Delays in reaching life milestones have turned early adulthood into a time marked more by instability than by upward mobility, particularly after the global financial crisis. In this situation, understanding how young adults are faring economically has become a pressing concern. Recent studies show a change in the geography and demographics of poverty, with younger groups facing new social risks that include but are not limited to insecure jobs, restricted welfare support, and rising housing costs.

Despite increasing academic and policy interest, there is a scarce comparative and up-to-date account of the economic conditions of young adults who have exited the family of origin. The study intends to isolate this subgroup which is greatly exposed to market and institutional risk for a more precise assessment of economic vulnerability during early adulthood. By leveraging on the harmonized microdata from the Luxembourg Income Study (Luxembourg Income Study (LIS) Database, 2026), the analysis provides a new cross-national perspective of the economic situation over two decades (1999-2020) using consistent measurement of income dynamics across advanced economies. We further move beyond documenting trends in poverty prevalence by examining distributional dynamics and decomposing poverty changes into compositional and structural components, which allow to better assess not only whether and in which contexts young adults are more exposed to relative poverty, but also whether returns to education, employment, or household structure have shifted over time. We begin by mapping relative poverty trends across countries and age groups, before estimating pooled linear probability models to assess the age gradient in poverty, controlling for key individual and household characteristics. We then compute marginal effects for occupational categories (restricting the sample to employed leads) and for housing cost burden (restricting the sample to renters), apply quantile regressions and use Oaxaca-Blinder decompositions to disentangle compositional and structural drivers of poverty among the general population

and the subgroup of interest.

Our findings suggest that there is a persistent, and in some cases, increasingly large disadvantage of young adult household heads, with some cross-national differences. In some Continental and Nordic European countries, the poverty gap between young and older cohorts has widened, and income penalties have concentrated further at the lower end of the income distribution. In Southern Europe, high levels of poverty across the whole population coincide with great income penalties for the young, implying that intergenerational disadvantage is situated in vulnerable labor markets. In some liberal economies, the poverty risk of the young has remained more stable, although often at comparatively high baseline level. The observed patterns in the decomposition analyses are not driven by unfavorable demographic composition, but mostly by changes in returns to observed characteristics. In European societies, improvements in compositional effects are largely countered by deteriorating returns, in contrast to liberal societies, where education and labor attachment continue to provide relatively robust protection. The predominance of the unexplained component in several cases, however, suggests that there are deeper structural transformations in how economic risk is allocated. The next section discusses the relevant literature and the context of this study in the broader debate surrounding young adults.

3 Literature and Research Contribution

The period between 2000 and 2020 under analysis represents a critical shift for young adults across advanced economies. Unlike the straightforward progress shown in traditional life-cycle models (Modell et al., 1976), with consecutive and reachable life milestones, the younger generation has experienced a less standard but more protracted transition to adulthood—one marked by shorter and more unpredictable economic cycles, along with frequent challenges in finding stable jobs and affordable housing (Aassve, Cottini, & Vitali, 2013; Standing, 2011). One clear sign of this shift is that many young adults are leaving their parents' homes later, a phenomenon that has been well documented across high-income countries (Goldscheider, 2014; Iacovou, 2010). A protracted living within the family of origin is often dictated not by personal preferences but by economic necessity. Unstable jobs and increasing housing costs have mainly been indicated as the drivers for this delay in moving out on their own.

The trend worsened significantly during the global financial crisis, which occurred halfway through this period and hit the young job seekers the hardest (Bell & Blanchflower, 2011; Verick, 2009). Youth unemployment rates increased and stayed high even after the recovery started. Younger workers often encountered more challenges in finding stable jobs, had weaker labor protection than older counterparts, and faced a heightened risk of falling into poverty (Chen et al., 2018; O’Higgins, 2012). Moreover, those early labor market setbacks are likely to generate a cumulative disadvantage over the life course, with cohorts exposed to recessions at labor market entry experiencing persistent earning penalties (Buchmann & Kriesi, 2011) . As a result, economic insecurity has increasingly influenced important life choices, with delays in starting families and partnerships becoming examples of an adaptive response to ongoing financial instability (Billari & Liefbroer, 2010; Furlong, 2016; Sobotka et al., 2011).

These developments show a wider change toward what scholars call new social risks. These are types of economic vulnerability that come not from traditional sources like old age, but from life-course transitions and changing labor markets. Brady et al. (2017) highlight that young adulthood itself has become a risk factor for poverty in many advanced economies. Moving towards self-sufficiency has been acknowledged as a increasingly complex and long process for the youth, and one characterized by multiple episodes of unemployment and family support (Côté & Bynner, 2008). Moreover, there is little evidence of a between-countries convergence with regard to the timing or the sequencing of events, with enduring clear differences with regards to age at union formation, entering a first job, or entry into parenthood (Pailhé et al., 2014). While exposure to poverty risk associated with age may vary across individuals and contexts, what matters even more is the penalty attached to it: in systems with weak institutional protection, being young, less experienced, or temporarily employed can sharply influence one’s likelihood of ending up in poverty. Furthermore, while higher education has seen a substantial expansion across many high-income nations, its protective effects against poverty have weakened (Barone & Ortiz, 2011). On this matter, Parolin et al. (2024) provide relevant insight using longitudinal U.S. data on the diminished capacity of the conventional benchmarks like education or stable employment to reduce the individual poverty risk. The severity of risks may vary considerably across countries. In

some high-income nations, strong safety nets —such as unemployment benefits, affordable housing, and active labor market policies— help buffer young adults from income shocks or the labor market exclusion (Esping-Andersen, 1990). In others, particularly those with more familialistic welfare regimes, the lack of robust public support shifts the burden of protection to private transfers and extended family networks, often prolonging economic dependence (Aassve, Arpino, & Billari, 2013; Aassve et al., 2006; Ferrera, 1996).

This study contributes to a growing literature on youth economic vulnerability by offering a cross-national analysis of poverty trends among young adults who have exited the parental home and assumed the role of household leads. In line with calls for stronger macro–micro integration in the study of the transition to adulthood (Buchmann & Kriesi, 2011), this paper examines how institutional contexts shape the economic vulnerability of young adults. Unlike most existing studies, which treat young adults as the broad demographic category, we focus on a specific subgroup — those under age 30 who have begun an independent living. The subgroup is of particular interest as it is no longer sheltered by the family of origin but often lacks the economic security and the institutional protection of the older adults. The youngest cohort is the most highly educated in recent history, yet, as discussed, it has recently faced a growing number of barriers into stable employment or affordable housing. In spite of their relevance for the understanding of intergenerational mobility and inequality, economically independent young adults remain understudied in comparative poverty research — especially over time and across institutional contexts. Drawing on the harmonized microdata from the Luxembourg Income Study (LIS), this paper investigates whether the incidence of the relatively poor among young household leads (ages 18–30) has increased between 1999 and 2020, whether this trend varies across countries, and what factors explain such variation.

Specifically, we ask:

- I. Has the economic position of young adult household leads deteriorated relative to older groups?
- II. To what extent do poverty and income penalties differ across countries?

III. Are changes in poverty rates driven primarily by compositional shifts (education, employment, household characteristics) or by changing returns to these characteristics?

To address these questions, we combine descriptive analyses with pooled linear probability models, predictive margins for relevant characteristics such as the occupational category and the housing burden, quantile regressions and Oaxaca-Blinder decompositions (Blinder, 1973; Oaxaca, 1973). This multifaceted approach offers a timely reassessment of youth poverty in the post financial crisis era and it traces the risks faced by young adults across a diverse pool of institutional contexts and family systems.

4 Data and empirical strategy

To analyze changes in poverty among young adults across countries and over time, we retrieved data from the Luxembourg Income Study (LIS), selecting two survey waves per country: an initial wave between 1999 and 2001 and a final wave between 2018 and 2020. The LIS database offers a unique advantage for this kind of analysis: it provides harmonized, high-quality microdata across a wide range of high-income countries (not limited to Europe), enabling direct cross-national and longitudinal comparisons of household income and demographic characteristics with consistent variable definitions. It is widely recognized as a premier resource due to its rigorous standardization of national household surveys by applying consistent definitions of disposable household income; the availability of several individual- and household level variables further enables a closer inspection of subgroup disparities or the decomposition of poverty trends (Atkinson et al., 1995; Gornick & Jäntti, 2013).

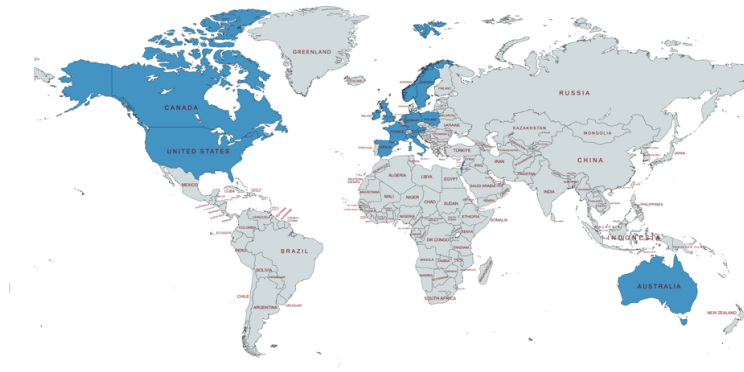


Figure 1: Countries included in the LIS sample used in the analysis.

For the scope of the analysis, in each wave we merge household and individual datasets using unique household identifiers. To preserve sample size—especially in smaller countries with limited observations in the initial wave— we retain all adult individuals aged 18–64 in the national samples. The selected individual-level variables reflect the characteristics of the household lead, defined as the adult with the highest labor income. In households where no member reported labor income, we assign the household head as the lead and to avoid redundancy, we ensure only one lead per household. We have comparable variables on income, age, sex, employment status, education, household size, and number of children and most of the times on occupations as classified on 3 simplified categories and housing status. We first exclude all children and non-working-age individuals from the analysis and we categorize age of the lead into four groups (under 30, 31–40, 41–50, and 51–64). The under 30 threshold for our age group of interest reflects the conventional definition adopted in the early 2000s literature on transition to adulthood, with the the twenties remaining the period in which transitions like school-to-work or residential independence are most likely to occur (Buchmann & Kriesi, 2011). While this threshold helps maintaining comparability with earlier research and isolating the age group most exposed to early-career instability, in a small number of countries (Ireland, Italy and Spain) the share of household leads under 30 falls below 10 percent. For those countries only, where independent living protracts usually longer, we report in the Appendix robustness analysis using the alternative threshold of 35 years as to capture many more leads. We use wherever possible the simplified ISCED education into three levels (low, medium, and high), while the equivalized income is calculated

using the square root scale, which is a standard practice in the applications of LIS data (Brady et al., 2017). Poverty is defined relative to the national median (at 50% threshold) and we apply LIS-provided household weights in all estimations.

We begin by examining descriptive trends in relative poverty rates by age groups across countries and waves, offering an initial overview of the evolving economic risks faced by young adults. We then estimate baseline linear probability models (LPMs) for each country, pooling the two waves and clustering standard errors at the household level to improve inference. In addition to the first probability models, we estimate extended specifications to examine heterogeneity within employment among the employed leads and housing conditions among renters. To further explore the distributional effects, we run quantile regressions of (log) equivalized income at various deciles, identifying where in the income distribution the penalties for age were highest or lowest. Finally, we apply two Oaxaca-Blinder decompositions to disentangle observed changes in poverty into parts driven by population composition (explained) and parts driven by changing returns to these characteristics (unexplained), providing some insight into both demographic shifts and structural transformations.

5 Descriptive evidence

Tables 1,2 and 3 present an initial descriptive portrait of households led by young adults (age up to 30) across our sample of 15 countries from the Luxembourg Income Study (LIS). We focus on five core indicators — the share of households headed by young adults, their tertiary education attainment, the employment rate of the leads, poverty incidence, and household size — observed at two time points spanning roughly two decades (1999-2001 to 2018-2020). These simple statistics provide a valuable first step in understanding the evolving economic conditions of younger generations, and they help set the later model-based analyses in a general framework.

Across most countries, the data yield some consistent signals of tangible demographic, educational, and economic shifts among the under-30 group in the 20-years span. Overall, the share of households led by someone under 30 has declined in most countries, attaining

low levels particularly in Southern Europe (e.g., Italy, Spain) and Ireland, a pattern that hints to delayed transitions to independent living. This stands in contrast to Nordic countries like Sweden and Norway, where young adults continue to represent a relatively stable and substantial portion of household leads (around 17%). Across all the observed countries then the proportion of young household heads with tertiary education has increased markedly. In Canada, Ireland, Israel and the Netherlands, more than half of young leads sampled in the last period hold a tertiary degree. These gains reflect broader trends of educational expansion, though the translation into stable employment and income may remain uneven. The young leads are generally highly attached to the labor market in most countries; yet, their employment rates have slightly declined in several settings despite the described increase of education attainment, including Sweden, Norway, and the Netherlands, pointing to more prolonged education spells or higher precariousness when entering the job market.

Some countries (Australia, Canada, Poland, Israel, Italy, the UK) have an experience of falling poverty rates among young household leads; others, like Germany or Sweden, have recorded sharp increases of poverty levels among young leads in the latest wave. This raises questions about low wages in entry-level jobs and labor market exclusion, especially when the proportion of tertiary education holders is limited such as in these two national samples. Together with a general shrinking independent young population, the household size among young leads has also generally decreased, particularly in Western and Nordic countries, with the average young-led household often consisting of fewer than two members. This could reflect living alone for longer, with delayed family formation and first child births. Overall, these descriptive patterns point to a generation that is generally more educated and slower in forming independent households compared to the past. In certain European contexts, there is some preliminary indication of a heightened exposure to economic risks. In the upcoming section we build on this descriptive foundation through regression-based and decomposition approaches aimed at disentangling structural from compositional changes.

In what follows, we present a more comprehensive picture of average poverty rates (defined as below 50% of the median equivalized disposable income) for the four age groups of the household lead — under 30, aged 31–40, 41–50, and 51–64 — across the earliest and

Table 1: Profile of Households Led by Young Adults (Aged 18–29) I

Indicator	AU 2001	AU 2018	AT 2000	AT 2019	CA 2000	CA 2019	FR 2000	FR 2019	DE 2000	DE 2019
% HHs w/ lead <30	14.3%	12.4%	10.3%	11.0%	13.7%	11.5%	12.9%	13.4%	11.2%	9.1%
% tertiary edu (lead <30)	–	44.4%	15.5%	27.8%	53.1%	67.7%	37.3%	45.9%	14.6%	22.3%
% employed (lead <30)	82.8%	88.7%	89.9%	81.9%	85.4%	87.7%	79.4%	78.4%	84.4%	78.9%
% in poverty (lead <30)	14.2%	11.8%	12.2%	19.1%	22.6%	18.3%	20.2%	22.0%	20.4%	37.5%
Mean HH size (lead <30)	2.40	2.50	2.26	2.03	2.36	2.31	2.10	1.97	2.03	1.94

Table 2: Profile of Households Led by Young Adults (Aged 18–29) II

Indicator	IE 2000	IE 2019	IL 2001	IL 2019	IT 2000	IT 2020	NL 1999	NL 2019	NO 2000	NO 2019
% HHs w/ lead <30	14.6%	4.9%	14.4%	12.3%	4.8%	3.0%	14.5%	14.2%	17.3%	16.6%
% tertiary (lead <30)	27.3%	56.1%	45.2%	54.4%	11.9%	23.4%	43.4%	55.9%	26.8%	35.9%
% employed (lead <30)	82.2%	83.3%	82.6%	90.9%	85.7%	81.4%	86.1%	74.9%	74.9%	65.8%
% in poverty (lead <30)	14.8%	15.2%	26.6%	25.7%	19.3%	17.6%	15.2%	22.0%	32.2%	35.3%
Mean HH size (lead <30)	2.43	2.53	2.83	2.81	2.54	2.26	1.98	1.66	1.87	1.71

latest years selected for each country. The figures are adjusted for households weights and take into account the structure of the data, allowing an accurate view on how economic hardship is distributed across different age groups in 15 countries.

The three plots capture a rather heterogeneous evolution in the distribution of economic disadvantage, rather than a uniform generational shift. Already in the early 2000s young adult household leads (under age 30) were in many cases more likely to be exposed to poverty than older age groups. In only a few instances, like Australia, Spain or Poland the poverty rates among young adults were in the early 2000s somewhat comparable to those observed among individuals aged 31–40 or 41–50. By the last period, the trajectories had diverged greatly across the examined economies. In a number of observed cases, youth poverty has declined or remained generally stable; Australia (14.2% to 11.8%), Canada (22.6% to 18.3%), Poland (10.3% to 7.1%), Italy (19.3% to 17.6%), and the United Kingdom (18.7% to 13.2%) all recorded a reduction in poverty levels among under-30 leads. At the same time, the United States and Israel exhibited relative stability (around 25%), while the increases recorded in Ireland or Spain were modest. By contrast, a subset of European countries has experienced more noticeable peaks in youth poverty over time. Austria stands out, with the under-30 poverty rate rising from 12.2% to 19.1%. The Netherlands registered a sharp increase (15.2% to 25.6%), as did Norway (22.0% to 32.2%) and Sweden (19.6% to 27.6%). Similarly, Germany exhibits a marked deterioration, with youth poverty rising from 20.4% to 37.5%.

Table 3: Profile of Households Led by Young Adults (Aged 18–29) III

Indicator	PL 1999	PL 2019	ES 2000	ES 2019	SE 2000	SE 2019	UK 2000	UK 2019	US 2000	US 2019
% HHs w/ lead <30	10.6%	9.2%	6.0%	3.1%	17.3%	17.8%	14.4%	13.0%	15.2%	13.8%
% tertiary (lead <30)	10.7%	33.6%	36.4%	51.2%	15.9%	29.0%	28.0%	46.0%	31.9%	46.8%
% employed (lead <30)	92.1%	90.2%	87.7%	78.0%	71.9%	67.6%	82.4%	83.3%	89.6%	87.6%
% in poverty (lead <30)	10.3%	7.1%	15.9%	16.2%	19.6%	27.6%	18.7%	13.2%	26.0%	25.5%
Mean HH size (lead <30)	3.45	2.91	2.20	2.28	1.71	1.65	2.35	2.41	2.72	2.49

France has comparatively a more moderate increase, from 20.2% to 22.0%.

The evolution of poverty among older age groups has not followed a uniform pattern either. Several countries where youth poverty increased at relatively high rates (the aforementioned Austria, Germany, Norway, Sweden or Netherlands) have also showed increases, usually by a smaller margin, among middle-aged or older groups. Comparatively, instead, countries like Australia, Canada, Poland and the UK have experienced a decline across most age brackets, highlighting what seem general economic improvements in the population rather than age-specific reversals. When computing averages across all countries, youth poverty has increased by roughly 2.4 percentage points between the two periods, compared to 1.4 points among the 41-50 group and 0.5 points among those aged 51-64. This gradient points to a relative worsening of the position of young adults, yet the magnitude remains quite moderate and driven by a subset of countries characterized by more pronounced increases. Overall, our descriptive evidence does not point to a generalized upward shift in youth poverty across high-income countries. Rather, it reveals a more fragmented landscape in which parts of Northern and Continental Europe have seen a marked deterioration in the economic position of young adults, while other countries in the Anglo-Saxon world have experienced stabilization or even improvement, even if commonly starting at high rates.

These contrasting trajectories raise important questions about the institutional and labor-market mechanisms shaping intergenerational vulnerability. To gain a greater understanding of the structural drivers behind this trend, the next section turns to regression analysis. By pooling data from the initial and most recent waves for each country, and interacting age group with the latest pre-COVID period, we aim to assess to what extent the relative poverty risks of young adults have changed over time when we additionally control for key demographic and socioeconomic characteristics.

Figure 2: Descriptives of mean poverty rates by age group and country - I

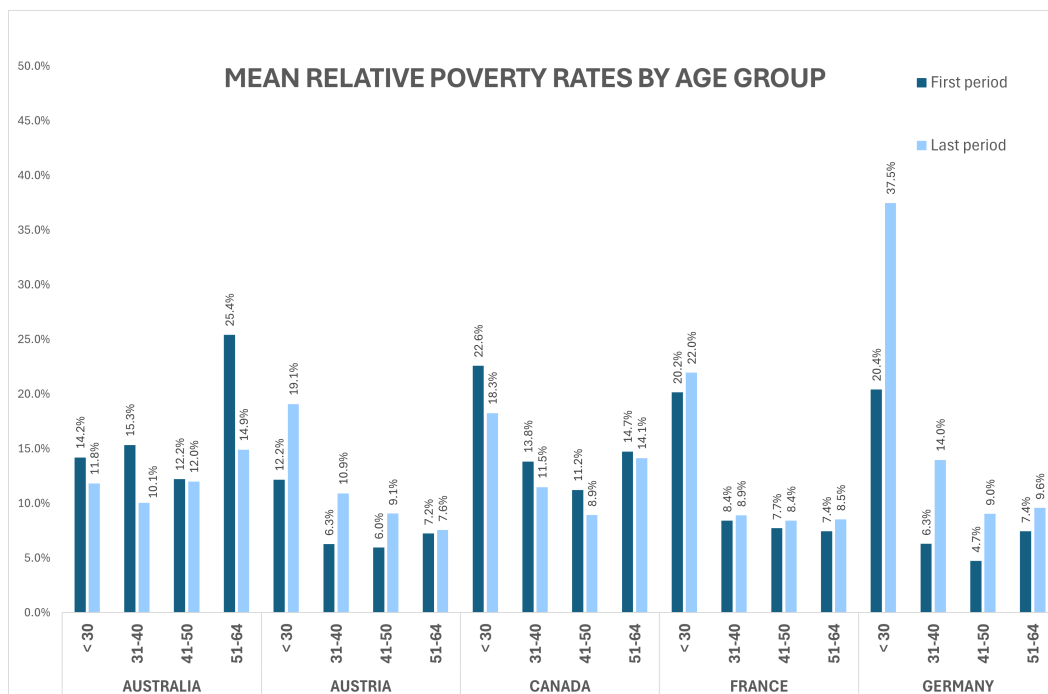


Figure 3: Descriptives of mean poverty rates by age group and country – II

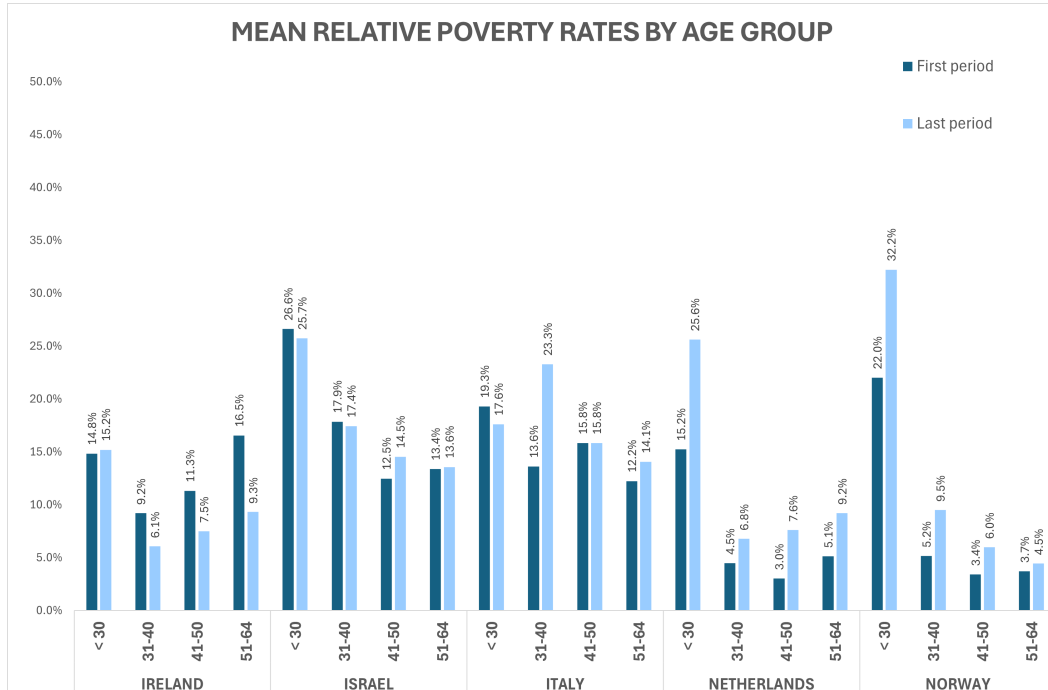
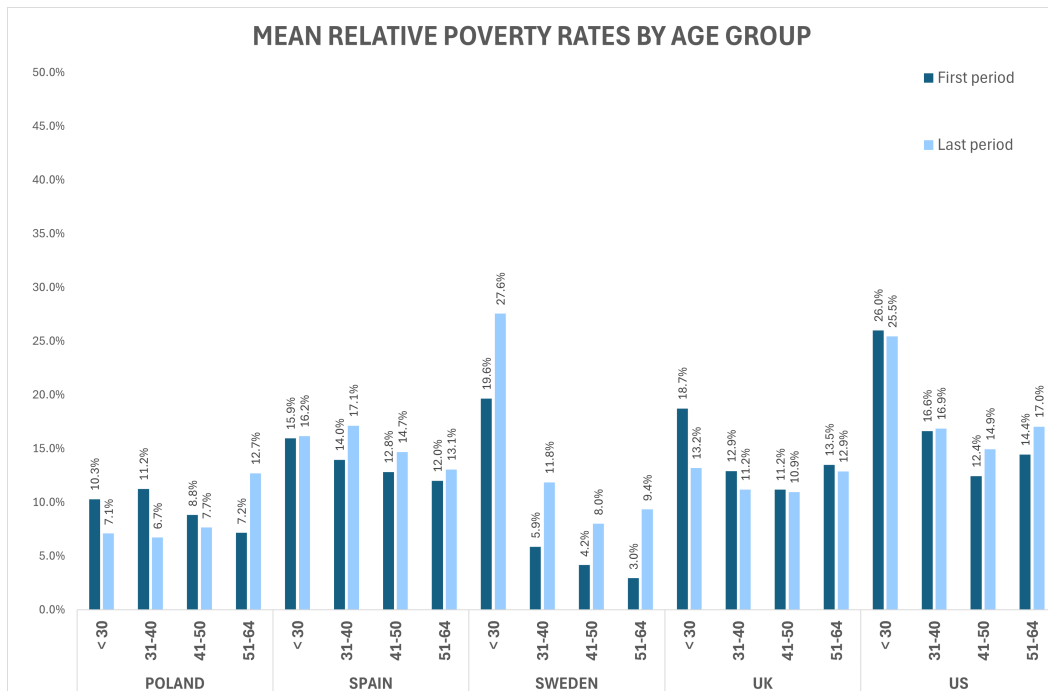


Figure 4: Descriptives of mean poverty rates by age group and country – III



6 Findings

6.1 National regression analyses

To better understand the drivers of poverty among young adults across countries and over time, we estimate a series of pooled linear probability models for each of the 15 countries in the study. The models combine the data from two survey waves—an early wave (around 2000) and a more recent one (2019 for most countries, 2018 and 2020 in a few instances)—and assess the likelihood that a household lead falls below 50% of the national median equivalized income. The analysis controls for the characteristics of the lead and the household, including sex, age, the employment status and education level of the lead, the household size, and number of children. Moreover, each model incorporates interaction terms between the age group of the lead and survey year, allowing for a comparison of how the poverty risk associated with age has evolved over time. The inclusion of an interaction between age group and survey year in fact captures whether age-based poverty differentials have widened, narrowed, or remained stable over time. The same regression specification is estimated separately for each country to ensure full comparability of results across national contexts, using household sampling weights and robust standard errors clustered at the household level to account for within-household correlation. A slightly different output is presented for Australia, where information on educational attainment is not available in the first wave (2001), and therefore education was excluded in that specification.

The pooled regression estimates provide a comparative perspective on age-related economic vulnerability and its dynamics over time. A consistent pattern emerges in the baseline period, with nearly all 15 countries showing significantly higher probability of falling below the relative poverty threshold for the young household leads. The age penalty at baseline appears statistically robust, with countries like France, Germany, Israel or Italy displaying coefficients for older groups exceeding -0.15, meaning that older leads in these countries were at minimum 15 percentage points less likely to be poor than comparable young leads. Only in a handful of cases (such as that of Ireland, and partially of Spain and Poland for the the 31–40 group), there were statistically insignificant reported age differences for the earlier

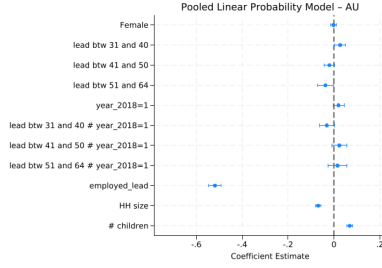
wave. A youth disadvantage seemed to be already present and structural at the turn of the millennium, with younger household leads facing a higher risk of poverty, holding employment status, education or household characteristics constant.

Turning to the dynamic component, the coefficient for the latest year of observation - capturing the changes for the reference group of under-30 leads - reveals a more heterogeneous trend than what initially anticipated. Youth poverty increased significantly in a subset of countries, ranging from about 3% in the United States, to 6-7% in the Netherlands and Nordic countries, and to a considerable 16% in Germany. In contrast, we observe a significant and mild decline for the United Kingdom and Poland and no robust changes in Australia, Canada, Ireland, Italy, and Spain. The evidence does not speak in favor of a generalized deterioration in youth poverty across advanced economies and points instead to different national trajectories, with the interaction terms between age groups and the latest wave providing a relevant test of whether age disparities have widened or not over time. Indeed, Germany and Norway are the only two countries where interaction effects are consistently negative and statistically significant, indicating a widening gap. A weaker version of this pattern appears in the Netherlands and Australia, where negative and significant interactions are confined mainly to the 31–40 age group. On an opposite note, the United Kingdom and Poland display positive and statistically significant interaction terms for the older age groups, indicating narrowing age-based disparities over time. In the UK, in particular, the relative advantage of older cohorts diminishes across all age groups, suggesting a more marked flattening of the generational gradient in poverty risk. In the remaining countries, interaction terms are generally small and statistically insignificant, pointing to somewhat parallel trends across age groups, the absence of deterioration of the relative position of older adults or reconfigurations of age-based inequalities. Given that in a number of countries (Italy, Spain, and Ireland) the number of household leads under age 30 is comparatively limited with a proportion of less than 10%, we performed robustness checks adopting an alternative definition of youth by extending the cutoff to age 35. The results, reported in the Appendix (Tables A1-A3) confirm that the direction and overall magnitude of the age penalties remain stable across specifications, with only a minor variation in statistical significance for one interaction factor in Italy.

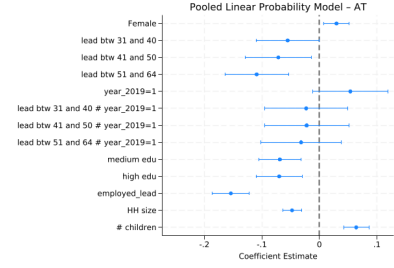
Beyond age dynamics, the outputs underscore the importance of the lead’s employment status, which emerges as the strongest protective factor against poverty across all countries. The corresponding coefficient ranges from around -0.15 in Austria and Poland to approximately -0.40 in countries such as the UK, the US, and Canada. These latter economies also share the highest constant terms, signaling particularly severe poverty risks for jobless young household leads and steeper income penalties associated with non-employment. Education also exerts significant protective effects in most cases despite different magnitudes. In countries such as the United States, Italy, Israel or Poland, tertiary education substantially reduces the poverty risk, whereas in others - such as France, Netherlands or Norway there is a more compressed educational gradient. In many cases, the magnitude of the youth penalty relative to older cohorts in the baseline year is comparable to — and sometimes even exceeds — the protective effect associated with completing tertiary education, reinforcing the view of young age as a structural vulnerability factor. The household arrangement also plays a critical role in shaping the poverty outcome. A larger household size is generally associated with slightly reduced poverty risks, likely due to the presence of multiple earners. In contrast, the number of children in the household is a strong and consistent predictor of higher poverty risk across countries. *Ceteris paribus*, each additional child increases the likelihood of falling below the poverty threshold, with effects typically ranging between 5 and 10 percentage points, highlighting dependent children as a contributing factor.

The baseline regression results point to an enduring and widespread age-related poverty gradient across advanced economies. Yet, they do not reveal a uniform deterioration in poverty levels over time, as three distinct trajectories emerge. A pattern of generational divergence is evident in a limited subset of countries (Germany, Norway, and the Netherlands), while convergence characterizes the United Kingdom and Poland. In the majority of cases, relative stability prevails, and the distribution of young age penalties remain largely unchanged. Being a young household lead continues to represent a disadvantage in most contexts—often comparable in magnitude to other structural determinants such as educational attainment. The consistency in both direction and significance of the baseline age coefficients emphasizes the economic vulnerability faced by the youngest household leads.

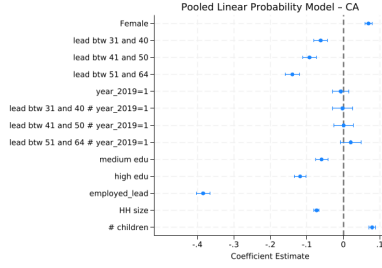
Figure 5: Pooled linear probability models - Part I



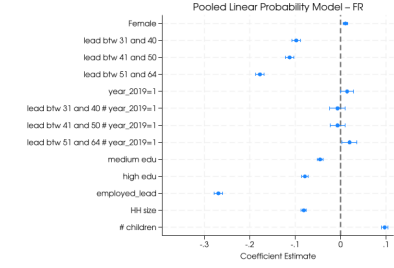
(a) Australia (2001 and 2018)



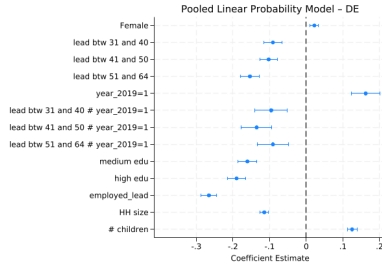
(b) Austria (2000 and 2019)



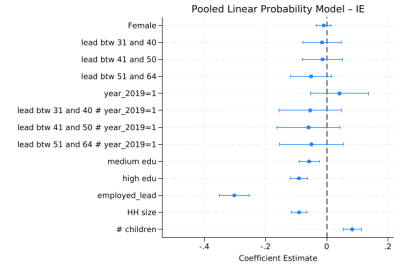
(c) Canada (2000 and 2019)



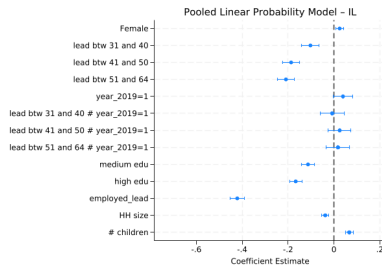
(d) France (2000 and 2019)



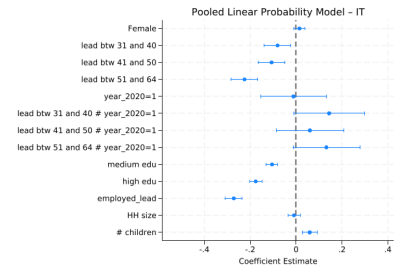
(e) Germany (2000 and 2019)



(f) Ireland (2000 and 2019)

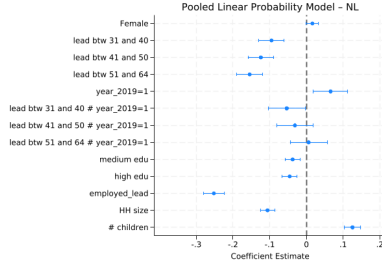


(g) Israel (2001 and 2019)

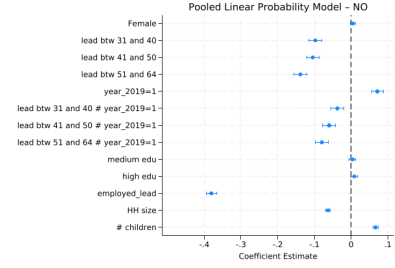


(h) Italy (2000 and 2020)

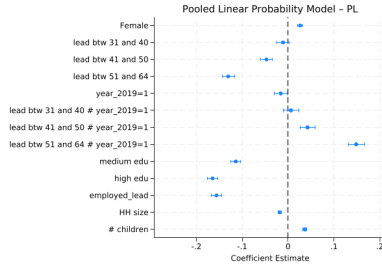
Figure 6: Pooled linear probability models - Part II



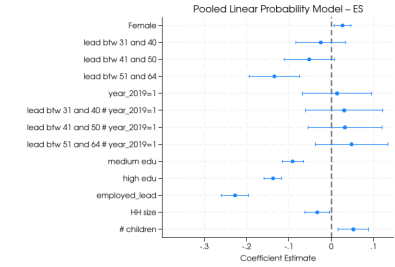
(a) Netherlands (1999 and 2019)



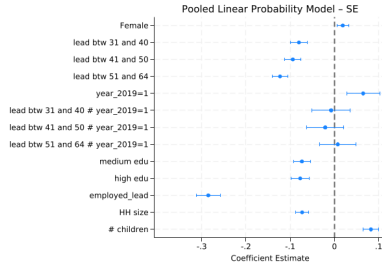
(b) Norway (2000 and 2019)



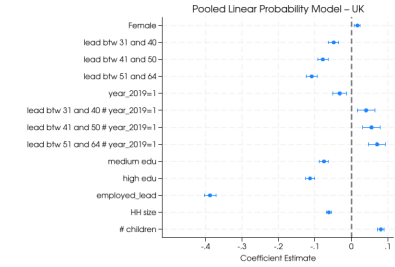
(c) Poland (1999 and 2019)



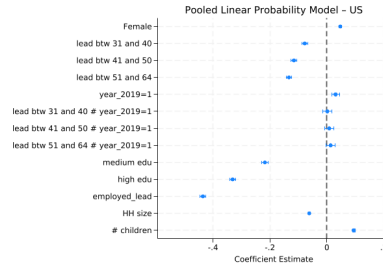
(d) Spain (2000 and 2019)



(e) Sweden (2000 and 2019)



(f) UK (2000 and 2019)



(g) US (2000 and 2019)

6.2 Occupational Stratification

The previous regression analyses have established that young household leads tend to face sustained poverty penalties over the examined period. They primarily distinguished between employed and non-employed leads but they did not explicitly account for heterogeneity within employment. To explore whether age-related disadvantage operates through stratification in the labor market, we further restrict the sample to employed household leads and run more complete models that introduce occupational categories, where possible, in the specifications. Focusing on the employed leads allows us to evaluate whether the explored poverty dynamics were attributable to differential access to employment or whether they also reflect differences in the type and quality of jobs. We therefore distinguish between three professional categories - namely managers/professionals, skilled occupations, and elementary occupations - and we interact the occupation with survey year. In these extended models we also control for education, household size, number of children, housing status, and regions where available. Although education and occupation are both structured into three ordinal variables, they represent different dimensions of economic stratification. Education captures formal qualifications, while occupation reflects how the qualifications may translate into labor market positions. Controlling for both allows to examine whether young workers face higher poverty risks because they are less educated, because they are concentrated in lower occupational segments, or because returns to these positions differ across the age groups.

The margins analysis reports the change in probability of falling in relative poverty between waves within each age group-occupation cell (for the under-30 and 51-64 for brevity), indicating once again uneven patterns of the age-related poverty dynamics across the labor market. Those countries where the earlier models pointed to a greater divergence, such as Germany and the Netherlands, tend to record higher increases in poverty risk concentrated among the young adults in lower-skilled occupations. In Germany, most notably, young leads in elementary occupations have experienced a substantial rise over time in the likelihood of falling under the relative poverty threshold, compared to the older counterparts or equivalent young professionals. In the case of Germany, the observed deterioration in the economic position seems to disproportionately affect both the young at the lower end and at the mid

part of the occupational structure. A similar pattern can be observed also in Austria, the Netherlands, and Sweden where the increase in economic vulnerability is particularly more pronounced among young workers employed in elementary occupations than for other age groups. These findings suggest that even within employment, the protective effect seen for having a job may depend heavily on its quality and income-generation capacity. By contrast, in countries characterized by generational convergence in the earlier regressions—such as the United Kingdom and Poland—the margins point to more stable or even improving poverty outcomes among young employed leads. In the UK in particular, poverty risks among young workers in elementary occupations decline over time, reinforcing the earlier evidence of a flattening age gradient. This indicates that the narrowing of generational disparities is not limited to employment rates but may possibly extend to within-employment dynamics.

The full regression tables reported in the Appendix further highlight some mechanisms. Across most countries, occupational status is a significant determinant of poverty risk even after controlling for education. Skilled and especially elementary occupations are associated with higher poverty probabilities relative to managerial positions in several countries, including the United States, the United Kingdom, Poland, and Sweden. In Germany and France, elementary occupations are strongly associated with higher poverty risk, underscoring the structural vulnerability of low-wage segments. The interaction terms between occupation and year show that, in some contexts, poverty risks increased disproportionately within lower occupational groups over time—particularly in Germany and Austria—while in others (such as the UK), they remained stable or even declined. The inclusion of education in the model does not eliminate occupational effects, indicating that poverty risk cannot be reduced to educational attainment alone. In countries such as the United States and Poland, tertiary education remains a strong protective factor even among employed leads, but occupational stratification exerts an independent influence as well. Returns to education are mediated by allocation on the labor market, with young workers likely to face other structural barriers that prevent education from fully translating into economic security.

These findings refine the conclusions drawn from the baseline regressions. The earlier models demonstrated that being a young household lead constitutes a persistent disadvan-

tage and, in a limited set of cases, an increasingly severe one. The occupational analysis shows that this disadvantage may be also closely linked to the stratification of the labor market. Generational divergence appears most acute where young workers are disproportionately concentrated in lower-skilled occupations and where the returns to those positions have weakened over time. Conversely, in contexts where poverty risks among lower occupational categories have stabilized or declined, generational convergence becomes more likely. Employment alone does not guarantee protection against economic vulnerability; rather, the type of employment and related income security seem to be crucial. The generational divide observed in some countries is best understood as an interaction between age and occupational stratification, pointing to deeper structural differences in how advanced economies allocate economic risk across cohorts.

Table 4: Change in relative poverty probability over the 20y period by age group and occupation (employed household leads) - Part I

	AU	AT	CA	FR	DE
Under 30					
Managers / Professionals	-1.3 (1.6)	-8.8 (11.1)	–	10.4*** (2.5)	8.1 (5.3)
Skilled Occupations	4.4*** (1.6)	7.2* (4.0)	–	6.9*** (2.3)	12.4*** (2.8)
Elementary Occupations	5.9*** (4.1)	29.4** (11.5)	–	8.3* (4.9)	33.8*** (10.9)
51–64					
Managers / Professionals	-1.1 (1.9)	0.1 (3.2)	–	-0.0 (0.8)	1.1 (1.0)
Skilled Occupations	1.3*** (2.1)	-0.8 (2.3)	–	3.9*** (1.0)	2.4*** (0.8)
Elementary Occupations	12.7*** (5.3)	-2.5 (7.5)	–	-10.4*** (1.5)	5.5 (4.7)

Notes: Entries report average marginal effects of year (2001 vs 2018 for Australia; 2019 vs. 2000 for other countries) from weighted linear probability models estimated separately by country among employed household leads. Effects expressed in percentage points. Robust standard errors clustered at the household level in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 5: Change in relative poverty probability over the 20y period by age group and occupation (employed household leads) - Part II

	IE	IL	IT	NL	NO
Under 30					
Managers / Professionals	-3.9 (3.2)	5.7 (3.6)	4.2 (8.2)	-26.8* (13.7)	–
Skilled Occupations	9.6 (8.2)	4.6 (3.1)	-4.7 (3.8)	6.8** (3.4)	–
Elementary Occupations	-25.4 (17.7)	16.4 (14.8)	-14.3* (7.3)	21.4* (12.1)	–
51–64					
Managers / Professionals	0.4 (1.6)	1.1 (1.1)	-2.8 (2.4)	4.1*** (1.5)	–
Skilled Occupations	-0.4 (2.5)	2.0 (1.5)	6.0*** (2.0)	1.3 (1.1)	–
Elementary Occupations	-17.6* (10.3)	11.6* (7.1)	-0.1 (2.9)	5.5* (3.1)	–

Notes: Entries report average marginal effects of year (2020 vs. 2000 for Italy; 2019 vs. 2000 for other countries) from weighted linear probability models estimated separately by country among employed household leads. Effects expressed in percentage points. Robust standard errors clustered at the household level in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 6: Change in relative poverty probability over the 20y period by age group and occupation (employed household leads) - Part III

	PL	ES	SE	UK	US
Under 30					
Managers / Professionals	1.77** (0.76)	-14.49* (8.25)	5.15 (3.48)	-0.18 (1.02)	2.66*** (0.99)
Skilled Occupations	-2.96*** (0.78)	-0.11 (6.47)	7.42** (2.93)	-0.61 (1.23)	3.53*** (1.05)
Elementary Occupations	4.54 (3.50)	24.01 (15.23)	25.97* (13.62)	-11.50** (5.36)	-3.00 (2.96)
51–64					
Managers / Professionals	1.67*** (0.48)	3.20 (2.03)	2.11* (1.08)	1.43* (0.75)	0.05 (0.47)
Skilled Occupations	1.05 (0.76)	3.33 (2.48)	4.04** (1.69)	0.35 (0.79)	3.60*** (0.64)
Elementary Occupations	9.17*** (2.26)	0.70 (4.90)	-1.97* (1.12)	2.69 (2.97)	4.52** (2.16)

Notes: Entries report average marginal effects of year (2019 vs. 2000) from weighted linear probability models estimated separately by country among employed household leads. Effects expressed in percentage points. Robust standard errors clustered at the household level in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

6.3 Housing burden as mediator

The previous analysis focused on occupations has conveyed that patterns of generational divergence have close links with labor market stratification. Yet, the employment condition cannot fully account for evolution of the economic vulnerability among the young leads. We acknowledge in fact that in many advanced economies, rising housing costs have disproportionately affected the younger cohorts, who are more likely to rent and devote a larger share of their disposable income to housing. To examine in particular whether housing costs may contribute to generational poverty dynamics in our samples, we introduce in this section a measure of housing burden, which we define as housing expenditures (rent and utilities) exceeding 30 percent of household disposable income. To this end, we estimate new weighted LPM models on relative poverty that incorporate an interaction term between the age group, the housing burden and survey year among employed renter household leads. This allows us to assess, wherever possible, whether the evolution of relative poverty risk differed in time by age and housing costs pressure. Unfortunately, this supplementary analysis could be run only for those 10 countries which provide housing cost information through the *hc4* variable in LIS.

Across the ten countries for which housing cost information is available, descriptive evidence from Tables 7-9 indicate that rental has become more prevalent in several contexts. The share of leads on rent has increased more evidently in Ireland (+15.6 pp), Spain (+11.8 pp), and the UK (+11 pp). More moderate rises have been instead observed in Australia, Austria and Israel. Rental shares have remained relatively stable in Italy while they declined among German, Polish and Swedish leads. The general trend suggests that a growing fraction of employed household leads is potentially exposed to pressures on the rental market.

The margins shown in Tables 10-12 indicate whether differences in changes of relative poverty over time vary by housing burden and age groups for the group of renter leads. While the results are often statistically insignificant across countries, some trends are observable. For leads under 30 years old, the risk of increased relative poverty probabilities over time appears to be greater for those with high housing burdens than for those with lower housing burdens. This is most evident in Australia, Germany, and Sweden, where the

increased poverty probability for high-burden young renters is considerably greater than for low-burden young renters. In other countries, the difference between high- and low-burden young renters is smaller or not statistically different, although the point estimates often suggest that high-burden young renters are more vulnerable to poverty. For people aged 51-64 years, the changes in relative poverty probabilities are generally smaller and less systematically related to housing burden. In some instances, poverty probabilities remain relatively unchanged over time for older renters with low housing burdens. In some countries (e.g. Sweden), however, older renters with high housing burdens experience a significant increase in poverty probabilities, although in most cases the estimates are imprecise. Although the extent and significance of the impact differ among the economies, these findings indicate that housing cost pressure could be a contributing factor to the observed trend of relative poverty, especially among the younger working group. In countries where the risk of poverty has risen among those under 30, these increments seem to be more significant for those allocating a larger proportion of their disposable income to rental and housing costs. Moreover, this evidence, despite suffering incomplete coverage and limited power in some contexts, is in line with the notion that housing affordability acts as additional factor to the economic vulnerability affecting the younger age groups.

Table 7: Share of Renter Household Leads (%), first and last period

	AU	AT	CA	FR	DE
First period	30.9	41.6	31.6	46.1	57.7
Second period	37.5	47.3	33	50.3	56.9
Change (pp)	6.6	5.7	1.4	4.2	-0.8

Notes: Shares of renter household leads, with changes expressed in percentage points.

Table 8: Share of Renter Household Leads (%), first and last period

	IE	IL	IT	NL	NO
First period	19.9	26.4	23.2	38.1	–
Second period	35.5	35.3	23.9	38.4	–
Change (pp)	15.6	8.9	0.7	0.3	

Notes: Shares of renter household leads, with changes expressed in percentage points.

Table 9: Share of Renter Household Leads (%), first and last period

	PL	ES	SE	UK	US
First period	20.4	10.6	45.4	28.9	33.7
Second period	12.1	22.4	41.7	39.9	36.8
Change (pp)	-8.3	11.8	-3.7	11	3.1

Notes: Shares of renter household leads, with changes expressed in percentage points.

Table 10: Change in relative poverty probability over the 20-year period by age group and housing burden (employed renter household leads)

	AU	AT	CA	FR	DE
Under 30					
Low Housing Burden	1.18* (0.67)	3.99 (26.05)	–	–	4.18* (2.34)
High Housing Burden	11.72** (5.7)	8.13 (12.38)	–	–	17.12*** (4.67)
51–64					
Low Housing Burden	≈ 0 (5.70)	≈ 0 (1.96)	–	–	1.18* (0.72)
High Housing Burden	4.20 (13.51)	≈ 0 (17.26)	–	–	5.84 (4.59)

Average marginal effects of survey year (second period vs. first period) from weighted linear probability models with year $\times$ housing burden $\times$ age interactions, estimated among employed renter household leads. Canada and France present missing information on the hc4 variable. Effects in percentage points with standard errors clustered at the household level in parentheses.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 11: Change in relative poverty probability over the 20-year period by age group and housing burden (employed renter household leads)

	IE	IL	IT	NL	NO
Under 30					
Low Housing Burden	-11.39 (12.05)	6.8* (4.2)	≈ 0 (6.5)	–	–
High Housing Burden	10.89 (13.25)	1.3 (6.4)	-14.3 (15.2)	–	–
51–64					
Low Housing Burden	≈ 0 (15.38)	1.3 (5.2)	5.6 (5.0)	–	–
High Housing Burden	19.67* (11.27)	10.7 (7.8)	-17.9 (12.2)	–	–

Notes: Average marginal effects of survey year (second period vs. first period) from weighted linear probability models with year $\times$ housing burden $\times$ age interactions, estimated among employed renter household leads. The Netherlands present missing information on the hc4 variable; Norway has missing information on housing for the first period. Effects in percentage points with standard errors clustered at the household level in parentheses.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table 12: Change in relative poverty probability over the 20-year period by age group and housing burden (employed renter household leads)

	PL	ES	SE	UK	US
Under 30					
Low Housing Burden	-1.87** (1.14)	-1.32 (5.23)	1.57 (2.27)	-7.09*** (1.52)	–
High Housing Burden	≈ 0 (3.07)	-16.58 (13.41)	23.0*** (6.7)	0.18 (4.05)	–
51–64					
Low Housing Burden	1.57 (2.05)	0.2 (3.1)	-0.65 (0.66)	-1.67 (1.34)	–
High Housing Burden	15.97 (6.28)	0.16 (3.08)	22.1*** (7.03)	-13.1 (5.88)	–

Notes: Average marginal effects of survey year (second period vs. first period) from weighted linear probability models with year $\times$ housing burden $\times$ age interactions, estimated among employed renter household leads. The US present missing information on the hc4 variable. Effects in percentage points with standard errors clustered at the household level in parentheses

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

6.4 Quantile regression results

To complement the linear probability models presented earlier, Figure 6 reports the maximum and minimum statistically significant income penalties for young adult household leads (under 30) relative to their older counterparts aged 51–64. We derive these estimates from quantile regressions of log equivalized household income at all deciles of the distribution, for both the earliest and most recent selected years in each country and using the same covariates of the baseline LPM models. Unlike linear regressions, this approach captures how income disparities by age can manifest across the income ladder and how their location may have shifted over time.

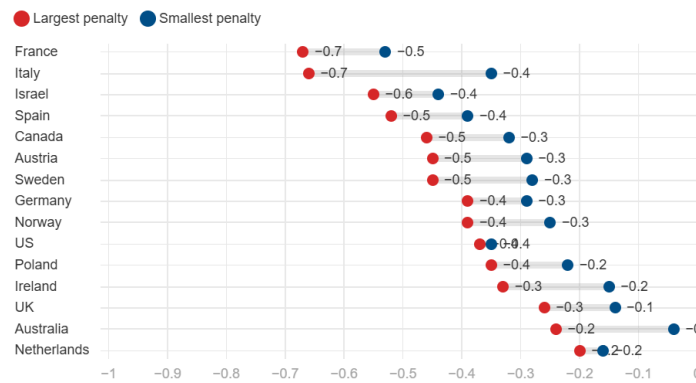
In the first period, the penalties for young household leads are meaningful and statistically significant in all countries. In most advanced economies, such as France, Israel, Canada, Austria, Norway, the United States, Ireland, the United Kingdom, Australia, and the Netherlands, the maximum penalty occurs at the 90th percentile. This suggests that the youngest leads were especially underrepresented at the top of the income distribution. In France and Israel, in particular, the penalties in the top tail were quite large (0.67 and 0.55, respectively), suggesting that there were large barriers to entry for the highest income positions. By contrast, in Italy, Spain, Sweden, Germany, and Poland, the largest penalties occur at the bottom of the distribution (around the 10th percentile). In these countries, the age difference is most pronounced in low-income households, with young individuals especially vulnerable at the bottom of the income distribution. In contrast, the minimum penalties in the first period are generally found in the middle of the distribution. Even in those countries with a relatively smaller gap, such as Australia, the UK, Ireland, and the Netherlands, young adults still face some kind of income disadvantage compared to older leads. In sum, earning disparities based on age in the early 2000s were widespread, though the patterns of their concentration varied from country to country, with some concentrated at the top, while others more acute at the bottom.

In the most recent period, penalties remain a persistent feature in almost all countries, but their shape changes in significant ways. In France, Germany, and Israel, the largest

penalties are now clearly concentrated at the lower end of the distribution (10th percentile), with France having the highest maximum penalty in the entire sample (0.92). This indicates that there is a trend toward lower-end concentration, which implies that the relative disadvantage of young adults may have become more extreme for low-income families. This trend is particularly worrisome, as the gaps in this part of the distribution are more directly related to material deprivation and economic insecurity. Conversely, a number of nations, such as the US, Norway, Austria, the Netherlands, Australia, and the UK, continue to display their largest penalties at the top of the distribution. In these settings, younger individuals seem to be facing stronger barriers to accessing high-paying jobs, even if the gaps in the lower end are reduced. Italy and Spain also present large penalties, although there is some shift towards the top of the distribution compared to the first period. It is important to note that the minimum penalties in the second period are smaller in a number of countries. The UK and Poland have the most limited minimum gaps, and Ireland is the only country where no statistically significant penalties are found. These examples indicate a form of convergence between younger and older household leads, at least in some points of the income distribution, and are aligned to the earlier evidence on changes in the likelihood of relative poverty.

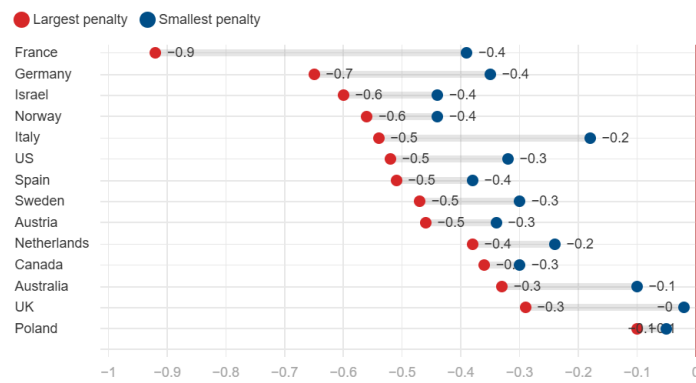
Comparing the two periods reveals some remarkable developments. First, in many countries, the magnitude of the largest penalties has grown over time, especially in France, Germany, Norway, and the US. Second, while maximum penalties in the earlier period were often concentrated at the upper end of the income distribution, the more recent period shows a clearer shift toward bottom-concentrated penalties in several non-Anglo-Saxon countries, particularly France and Germany. This suggests that intergenerational disparities are increasingly rooted in lower-tail disadvantage rather than solely in barriers to accessing high-income positions. Increased disadvantages at the bottom may carry important social implications in terms of heightened risks of economic precariousness and delayed economic independence for younger generations.

QR (first period): Under-30s vs 51-64 leads



(a) Age earnings penalties relative to 51-64, first period

QR (last period): Under-30s vs 51-64 leads



(b) Age earnings penalties relative to 51-64, last period

Figure 7: Max and min earnings penalties across the distribution in the two periods

6.5 Aggregate decompositions of relative poverty

In the last part of our analysis, to better understand the mechanisms behind poverty trends across high-income countries, we use Oaxaca–Blinder decompositions to disentangle observed changes into components attributable to shifts in population characteristics (“explained” or endowments) and components due to changes in the returns to these characteristics (“unexplained” or structural factors). This approach allows us to distinguish whether variations in relative poverty stem primarily from improvements in employment, education, or household composition—or from broader transformations in how these traits are rewarded. Concretely, we estimate linear probability models of relative poverty (our dependent variable, i.e. 50% of median disposable income) for the two periods in each country and then apply the Oaxaca–Blinder decomposition to the change in predicted poverty risk between the first and last survey waves. The baseline specification includes demographic and socio-economic characteristics: gender of the household lead, age group of the lead, his or her employment status and education level, the household size, and number of children. We first perform the decomposition for the general population of household leads to capture aggregate poverty dynamics. In a second step, we also focus specifically on young household leads (under 30), extending the specification to account for occupational categories and housing tenure status (renting), in order to better isolate those channels that might be particularly relevant for younger cohorts. All models are weighted and standard errors are clustered at the household level. The decomposition expresses the overall change in poverty as the sum of three components: an endowment effect (changes in characteristics), a coefficient effect (changes in returns), and an interaction term capturing the joint contribution of both. As the overall difference in Tables 13-15 is computed as the first period minus the second period, a negative overall difference indicates that relative poverty has increased over time.

Across the 15 countries, poverty dynamics in the general population reveal a clear primacy of structural over compositional drivers, even though the magnitudes differ considerably. Germany shows the largest increase (roughly 6 pp), driven almost entirely by the coefficients component despite the protective effects of both employment and education. The same pattern emerges for Austria (3 pp) and France (near 1 pp), where the structural decline

overshadows the small compositional changes. In the case of France, employment presents a positive coefficient effect, although the structural part is negative, which confirms that the increase is a systemic change rather than a deterioration along a single observable margin. Other European economies like the Netherlands, Norway, Sweden and Italy show moderate increases in poverty rates, ranging between 3 and 6 p.p., mainly through structural forces. In all these cases, significant coefficient effects show up across several dimensions (such as higher education or employment) along some stable or positive endowments, yet the overall unexplained component points to a general baseline shift in poverty risk. The United States follows this same pattern, although the observed poverty change is very modest (just above 1 p.p.). Israel, Spain, and Poland, on the other side, present non significant results on aggregate poverty rates, as improvements in the endowments are almost fully offset by structural deterioration. Their cases illustrate that improvements in education and, in some instances, in employment can coexist with broader weakening factors to economic security. Economies such as Canada or Ireland report declines in poverty rates mostly through improvements in higher education or employment, with only minimal evidence of structural deterioration, whereas Australia and the UK show the same positive contributions through the endowments, which are however partially neutralized by weakening returns (e.g. both education and employment in the UK).

The results from the decomposition analysis conducted on the general population point to a broader structural transformation, largely captured by the constant term in the unexplained component, and indicate that households with similar observable characteristics faced higher poverty risk in the later period despite an often protective role led by higher education or - less frequently - by employment. Liberal economies appear somewhat better positioned, as compositional improvements usually offset part of the underlying structural weakening. However, the recurrent relevance of the intercept across countries highlights the difficulty of attributing poverty dynamics solely to observable margins such as employment or education, and instead suggests systemic changes in how labor market position, age or household characteristics translate into economic security. In particular, the age structure plays a relevant role in selected European countries. The compositional effect of older age groups (especially 51–64) is often positive, reflecting population aging would if anything mit-

igate the aggregate economic vulnerability. At the same time, coefficient effects for older age are in a few instances positive and significant in the unexplained component (UK, Germany, Norway) and would point to a rise in the returns accrued to this specific age group in those contexts. The household composition contributes mildly to poverty dynamics depending on the considered dimension. The explained component is very frequently negative and significant across countries as shrinking households have likely contributed to poverty increases due to a single source of income. At the same time, the economic returns to household size move in more heterogeneous directions, and in a few instances (US, Norway, Sweden) there is an observed robust weakening gradient linked with household size. Less frequently in comparison, a lower number of children has mitigated the poverty increase, even if this may go in hand with worsening returns associated with parenthood.

Table 13: Blinder–Oaxaca Decomposition of the Change in Relative Poverty - Part I

	AU	AT	CA	FR	DE
Overall Difference	0.030 (0.010)***	-0.030 (0.010)***	0.0216 (0.0049)***	-0.0091 (0.0025)***	-0.0579 (0.0058)***
<i>Explained (Endowments)</i>	0.050 (0.004)***	0.005 (0.004)	0.0213 (0.0031)***	0.0079 (0.0016)***	0.0170 (0.0039)***
Female (lead)	0.002 (0.001)**	0.000 (0.001)	-0.0028 (0.0005)***	-0.0011 (0.0004)**	-0.0017 (0.0007)*
Age 31–40	0.001 (0.001)	0.002 (0.001)**	-0.0014 (0.0005)**	-0.0029 (0.0004)***	-0.0170 (0.0024)***
Age 41–50	0.001 (0.001)	0.001 (0.001)	-0.0033 (0.0007)***	-0.0025 (0.0005)***	-0.0030 (0.0019)
Age 51–64	0.003 (0.001)*	0.010 (0.002)***	-0.0002 (0.0007)	0.0083 (0.0007)***	0.0331 (0.0033)***
Medium Education	—	-0.004 (0.002)	-0.0047 (0.0016)**	-0.0001 (0.0002)	-0.0159 (0.0025)***
High Education	—	0.006 (0.002)**	0.0246 (0.0033)***	0.0116 (0.0010)***	0.0146 (0.0025)***
Employed	0.040 (0.004)***	-0.004 (0.002)	0.0069 (0.0018)***	0.0006 (0.0008)	0.0139 (0.0025)***
Household Members	0.001 (0.001)	-0.020 (0.005)***	0.0009 (0.0013)	-0.0136 (0.0012)***	-0.0156 (0.0030)***
Children	0.001 (0.001)	0.015 (0.004)***	0.0012 (0.0010)	0.0076 (0.0010)***	0.0087 (0.0025)***
<i>Unexplained (Coefficients)</i>	-0.030 (0.010)***	-0.027 (0.009)***	0.0049 (0.0046)	-0.0175 (0.0023)***	-0.0650 (0.0051)***
Female (lead)	-0.070 (0.020)***	0.024 (0.016)	0.0413 (0.0137)***	-0.0015 (0.0067)	0.0071 (0.0160)
Age 31–40	0.010 (0.004)**	0.006 (0.009)	-0.0045 (0.0039)	0.0013 (0.0021)	0.0122 (0.0043)**
Age 41–50	0.003 (0.003)	0.008 (0.009)*	-0.0047 (0.0037)	0.0014 (0.0023)	0.0241 (0.0053)***
Age 51–64	-0.010 (0.007)	0.021 (0.010)**	-0.0062 (0.0038)*	-0.0058 (0.0030)*	0.0438 (0.0099)***
Medium Education	—	0.029 (0.014)*	-0.0030 (0.0040)	-0.0031 (0.0032)	0.0292 (0.0136)*
High Education	—	0.026 (0.010)**	0.0077 (0.0135)	-0.0001 (0.0031)	0.0291 (0.0096)**
Employed	-0.030 (0.020)	0.041 (0.027)	-0.0020 (0.0173)	0.0209 (0.0079)**	0.1326 (0.0196)***
Household Members	-0.050 (0.040)	0.075 (0.045)	-0.0921 (0.0212)***	-0.0184 (0.0161)	0.0465 (0.0291)
Children	0.010 (0.020)	-0.033 (0.019)	0.0507 (0.0091)***	0.0060 (0.0069)	-0.0070 (0.0104)
<i>Interaction</i>	0.010 (0.004)**	-0.007 (0.004)*	-0.0046 (0.0026)*	0.0005 (0.0014)	-0.0100 (0.0027)***
<i>N - first period</i>	7,698	3,182	37,039	82,542	15,461
<i>N - second period</i>	16,433	6,547	32,723	51,230	23,096

Table 14: Blinder–Oaxaca Decomposition of the Change in Relative Poverty - Part II

	IE	IL	IT	NL	NO
Overall Difference	0.0474 (0.0124)***	-0.0123 (0.0089)	-0.0282 (0.0122)**	-0.0449 (0.0068)***	-0.0316 (0.0030)***
<i>Explained (Endowments)</i>	0.0268 (0.0093)**	0.0441 (0.0058)***	0.0709 (0.0088)***	0.0038 (0.0041)	-0.0086 (0.0021)***
Female (lead)	0.0015 (0.0011)	-0.0013 (0.0006)*	-0.0024 (0.0021)	-0.0010 (0.0005)	-0.0005 (0.0001)***
Age 31–40	-0.0004 (0.0014)	0.0011 (0.0012)	0.0062 (0.0066)	-0.0077 (0.0018)***	-0.0045 (0.0007)***
Age 41–50	0.0030 (0.0022)	-0.0047 (0.0018)**	0.0003 (0.0008)	-0.0031 (0.0014)**	-0.0002 (0.0008)
Age 51–64	0.0071 (0.0036)*	-0.0110 (0.0021)***	0.0113 (0.0080)	0.0140 (0.0023)***	0.0103 (0.0011)***
Medium Education	-0.0000 (0.0001)	0.0006 (0.0014)	0.0072 (0.0025)***	-0.0043 (0.0019)**	-0.0002 (0.0003)
High Education	0.0171 (0.0067)**	0.0264 (0.0037)***	0.0168 (0.0030)***	0.0085 (0.0026)***	-0.0005 (0.0002)
Employed	0.0058 (0.0039)	0.0372 (0.0042)***	0.0281 (0.0050)***	0.0046 (0.0027)	-0.0128 (0.0016)***
Household Members	-0.0288 (0.0080)***	0.0016 (0.0016)	-0.0108 (0.0047)**	-0.0123 (0.0031)***	-0.0069 (0.0012)***
Children	0.0216 (0.0069)**	-0.0058 (0.0031)	0.0141 (0.0044)***	0.0052 (0.0027)	0.0066 (0.0012)***
<i>Unexplained (Coefficients)</i>	-0.0005 (0.0110)	-0.0433 (0.0079)***	-0.0885 (0.0109)***	-0.0426 (0.0068)***	-0.0230 (0.0028)***
Female (lead)	0.0146 (0.0353)	-0.0571 (0.0223)**	-0.0223 (0.0317)	-0.0180 (0.0223)	-0.0166 (0.0089)
Age 31–40	0.0161 (0.0142)	0.0069 (0.0082)	-0.0276 (0.0139)*	0.0060 (0.0058)	0.0093 (0.0021)***
Age 41–50	0.0181 (0.0170)	0.0004 (0.0068)	-0.0205 (0.0229)	0.0015 (0.0063)	0.0161 (0.0023)***
Age 51–64	0.0076 (0.0189)	0.0000 (0.0055)	-0.0476 (0.0367)	0.0028 (0.0092)	0.0233 (0.0030)***
Medium Education	-0.0239 (0.0084)**	0.0060 (0.0095)	-0.0144 (0.0120)	-0.0010 (0.0076)	0.0057 (0.0037)
High Education	-0.0394 (0.0161)**	0.0419 (0.0152)**	0.0063 (0.0053)	0.0083 (0.0105)	0.0057 (0.0039)
Employed	-0.0877 (0.0407)*	0.0640 (0.0294)*	0.1906 (0.0347)***	0.1480 (0.0253)***	0.0498 (0.0126)***
Household Members	-0.0277 (0.0827)	-0.0328 (0.0611)	0.1247 (0.0726)*	-0.0847 (0.0535)	0.0847 (0.0185)***
Children	0.0089 (0.0396)	-0.0179 (0.0349)	-0.0535 (0.0343)	0.0387 (0.0208)*	-0.0508 (0.0071)***
<i>Interaction</i>	0.0210 (0.0120)*	-0.0132 (0.0044)**	-0.0106 (0.0089)	-0.0061 (0.0037)	-0.0000 (0.0011)
<i>N - first period</i>	3,593	8,051	9,927	6,670	19,005
<i>N - second period</i>	4,360	9,904	5,865	12,954	274,474

Table 15: Blinder–Oaxaca Decomposition of the Change in Relative Poverty - Part III

	PL	ES	SE	UK	US
Overall Difference	0.0025 (0.0026)	-0.0148 (0.0100)	-0.0563 (0.0067)***	0.0149 (0.0040)***	-0.0134 (0.0027)***
<i>Explained (Endowments)</i>	0.0532 (0.0021)***	0.0400 (0.0066)***	0.0062 (0.0039)	0.0274 (0.0027)***	0.0266 (0.0016)***
Female (lead)	-0.0007 (0.0002)***	-0.0023 (0.0017)	-0.0011 (0.0005)**	-0.0005 (0.0004)	-0.0021 (0.0002)***
Age 31–40	-0.0006 (0.0002)*	-0.0001 (0.0023)	-0.0000 (0.0006)	-0.0003 (0.0003)	-0.0018 (0.0003)***
Age 41–50	0.0009 (0.0004)*	0.0015 (0.0016)	-0.0014 (0.0008)	0.0001 (0.0001)	-0.0052 (0.0005)***
Age 51–64	-0.0002 (0.0003)	0.0052 (0.0022)**	0.0012 (0.0009)	0.0019 (0.0006)***	0.0113 (0.0007)***
Medium Education	-0.0103 (0.0010)***	0.0057 (0.0016)***	-0.0100 (0.0020)***	-0.0202 (0.0031)***	-0.0189 (0.0012)***
High Education	0.0394 (0.0025)***	0.0239 (0.0031)***	0.0162 (0.0029)***	0.0278 (0.0027)***	0.0456 (0.0018)***
Employed	0.0231 (0.0012)***	-0.0064 (0.0027)**	0.0006 (0.0026)	0.0178 (0.0016)***	-0.0049 (0.0009)***
Household Members	-0.0020 (0.0004)***	-0.0254 (0.0101)**	0.0010 (0.0027)	0.0046 (0.0011)***	-0.0047 (0.0007)***
Children	0.0035 (0.0007)***	0.0380 (0.0099)***	-0.0002 (0.0025)	-0.0039 (0.0012)**	0.0074 (0.0009)***
<i>Unexplained (Coefficients)</i>	-0.0400 (0.0024)***	-0.0456 (0.0101)***	-0.0591 (0.0059)***	-0.0201 (0.0036)***	-0.0350 (0.0023)***
Female (lead)	-0.0391 (0.0078)***	0.0353 (0.0301)	-0.0245 (0.0184)	0.0265 (0.0109)**	0.0059 (0.0069)
Age 31–40	-0.0110 (0.0027)***	-0.0030 (0.0096)	-0.0044 (0.0056)	-0.0099 (0.0031)***	-0.0024 (0.0021)
Age 41–50	-0.0186 (0.0025)***	-0.0023 (0.0154)	-0.0015 (0.0051)	-0.0143 (0.0033)***	-0.0036 (0.0020)
Age 51–64	-0.0328 (0.0030)***	-0.0087 (0.0185)	-0.0122 (0.0070)	-0.0272 (0.0042)***	-0.0041 (0.0029)
Medium Education	0.0318 (0.0085)***	-0.0099 (0.0059)	0.0465 (0.0102)***	-0.0052 (0.0037)	0.0014 (0.0048)
High Education	0.0139 (0.0041)***	-0.0018 (0.0088)	0.0320 (0.0081)***	-0.0171 (0.0062)**	0.0237 (0.0063)***
Employed	0.1163 (0.0103)***	0.0365 (0.0274)	0.1302 (0.0238)***	-0.0347 (0.0147)**	0.0196 (0.0096)**
Household Members	-0.0186 (0.0117)	0.0050 (0.0804)	0.1522 (0.0381)***	0.0072 (0.0207)	0.0446 (0.0110)***
Children	0.0390 (0.0058)***	-0.0230 (0.0344)	-0.0679 (0.0149)***	0.0003 (0.0086)	-0.0098 (0.0046)**
<i>Interaction</i>	-0.0107 (0.0023)***	-0.0092 (0.0090)	-0.0034 (0.0024)	0.0077 (0.0025)***	-0.0050 (0.0009)***
<i>N - first period</i>	45,249	5,768	16,896	28,173	104,810
<i>N - second period</i>	42,331	16,870	6,611	20,014	71,629

6.6 Decompositions of relative poverty among the youngest age group

To assess the changes in relative poverty among young household leads, we apply a Oaxaca-Blinder decomposition to linear probability models estimated separately by survey year and exclusively on the youngest age group (under 30). Compared to the decomposition applied to the general population, we further explore here the occupational and housing dimensions. Across the fifteen countries examined, youth poverty trends show again some remarkable heterogeneity. Compared to the general population case, the negative significant estimates on the overall difference are characterized by greater magnitude and are usually driven by the unexplained component. In several economies, located especially in Europe, growing shares of the highly educated would have contributed to reducing poverty rates, but weakening returns to the characteristics actually worked in the opposite direction. Youth poverty in Germany was the highest (14.5 pp) observed over the two decades period and almost entirely shaped by structural deterioration, notwithstanding general improvements in tertiary education. France, Sweden, and Norway also featured growing and similar shares of relative youth poverty (around 10%) led by structural factors, often partially related to the housing market. In these cases, the expansion of higher education (or skilled occupations in France) did not translate into stronger protection against poverty and was accompanied by a broader weakening of the institutional framework. Israel further illustrates a case where a significant increase in poverty among the young (7.3 p.p.) stems mostly from declining returns to advanced education. While in the Netherlands the poverty rate seems to have remained relatively more stable, the significant negative interaction term stands to show that compositional improvements in education and skilled occupations translated less effectively into poverty protection due to more unstable institutional conditions.

At the same time, in several cases such as Canada, Germany, and the Netherlands, the coefficient effects for household members are negative and significant, indicating that the marginal association between household size and poverty became more unfavorable over time. This points to a structural intensification of the poverty gradient linked to household composition rather than a simple demographic shift. Comparatively, in the Anglo-Saxon contexts, the relative stability observed in aggregate poverty conceals similar redistribution

dynamics. For instance, while Canada shows no significant overall change (1.6 pp), yet this pattern is the outcome of compositional improvement in education offset by structural deterioration. Similarly, the United Kingdom and the United States feature poverty-reducing compositional effects that are counterbalanced by weakening returns, especially among lower-skilled occupations.

As for the household structure, even in the case of young leads, a decline in the household size has contributed to rising poverty through endowments (Norway, US). Additionally, we get negative coefficients standing for unfavorable association between household size and poverty in places such as Canada, Germany, and the Netherlands. Compositional effects for children are generally modest but occasionally present a positive sign, and only Norway presents clear worsening returns to parenthood. Poland stands out as a clear case where youth poverty reduction stems from compositional improvements through the expansion of higher education and no offsetting mechanisms through the labor market. Interpretation should remain particularly cautious instead for Italy, Austria, Ireland, and Spain due to very limited samples. Estimates in those cases are often imprecise and unstable and prevent us from deriving solid conclusions about the role of education or skills and their structural change.

Compared to the general population, youth poverty appears to be driven less by demographic composition and more by transformations in how education, skills, and housing conditions translate into economic security. In many advanced economies (Continental Europe, Nordic countries, and Israel), rising youth poverty reflects structural deterioration in the returns to education, occupation, or household conditions, while liberal contexts seem to have experienced stronger compositional improvements that have partially offset similar weakening returns. At the same time, although the coefficients component is often statistically significant, the individual mechanisms within it are imprecise or only weakly identified because of smaller and sometimes unbalanced samples. This makes it difficult to attribute evolution in youth poverty to specific channels for most economies. In a few countries - Canada, US, Australia - compositional improvements have cushioned declining returns in the characteristics, but this can also coexist with widening inequality between

those who have upgraded their education or skills and those who did not. Poverty reduction appears more credibly linked to educational expansion only in Poland, while we refrain from over-interpreting the results from countries with too few observations in the separate yearly samples.

Table 16: Blinder–Oaxaca Decomposition of the Change in Relative Poverty among Young Household Heads (Under 30) - Part I

	AU	AT	CA	FR	DE
Overall Difference	-0.0143 (0.0135)	-0.0736 (0.0383)*	-0.0161 (0.0175)	-0.0982 (0.0192)***	-0.1452 (0.0285)***
<i>Explained (Endowments)</i>	0.0089 (0.0051)*	0.0053 (0.0177)	0.0393 (0.0125)***	-0.0068 (0.0130)	-0.0099 (0.0143)
Skilled Occupation	0.0021 (0.0015)	0.0202 (0.0201)	—	0.0013 (0.0126)	-0.0006 (0.0053)
Elementary Occupation	0.0048 (0.0027)*	-0.0036 (0.0058)	—	-0.0092 (0.0064)	-0.0083 (0.0053)
Medium Education	—	-0.0201 (0.0186)	0.0080 (0.0097)	-0.0065 (0.0053)	-0.0455 (0.0137)***
High Education	—	0.0121 (0.0172)	0.0287 (0.0137)**	0.0084 (0.0088)	0.0388 (0.0135)***
Rent	-0.0051 (0.0031)	-0.0136 (0.0090)	-0.0025 (0.0026)	0.0102 (0.0036)***	-0.0018 (0.0030)
Household Members	0.0057 (0.0028)**	0.0007 (0.0158)	-0.0011 (0.0013)	-0.0205 (0.0080)**	-0.0032 (0.0070)
Children	0.0013 (0.0023)	0.0136 (0.0211)	0.0015 (0.0035)	0.0128 (0.0061)**	0.0131 (0.0071)*
<i>Unexplained (Coefficients)</i>	-0.0246 (0.0129)*	-0.0451 (0.0539)	-0.0595 (0.0152)***	-0.0963 (0.0191)***	-0.1424 (0.0274)***
Skilled Occupation	-0.0054 (0.0141)	-0.0885 (0.0840)	—	0.0063 (0.0171)	-0.0033 (0.0603)
Elementary Occupation	0.0028 (0.0047)	-0.0137 (0.0134)	—	0.0041 (0.0160)	-0.0132 (0.0100)
Medium Education	—	0.1425 (0.0934)	-0.0189 (0.0168)	-0.0078 (0.0240)	-0.0517 (0.0526)
High Education	—	0.1087 (0.0577)*	-0.0044 (0.0544)	-0.0300 (0.0367)	-0.0077 (0.0310)
Rent	0.0139 (0.0189)	-0.0624 (0.0536)	-0.0354 (0.0181)*	-0.1147 (0.0215)***	-0.1206 (0.0543)**
Household Members	-0.0429 (0.0459)	0.0237 (0.1145)	-0.1297 (0.0395)***	0.0549 (0.0612)	-0.1138 (0.0669)*
Children	0.0102 (0.0096)	-0.0034 (0.0194)	0.0271 (0.0070)***	-0.0026 (0.0107)	0.0163 (0.0110)
<i>Interaction</i>	0.0014 (0.0061)	-0.0338 (0.0393)	0.0040 (0.0092)	0.0048 (0.0128)	0.0071 (0.0131)
<i>N - first period</i>	686	168	5,143	4,744	921
<i>N - second period</i>	1,120	378	2,896	799	1,461

Table 17: Blinder–Oaxaca Decomposition of the Change in Relative Poverty among Young Household Leads (Under 30) - Part II

	IE	IL	IT	NL	NO
Overall Difference	-0.0270 (0.0600)	-0.0729 (0.0304)**	0.0985 (0.0516)*	0.0320 (0.0327)	-0.1045 (0.0120)***
<i>Explained (Endowments)</i>	0.0233 (0.0542)	0.0294 (0.0185)	-0.0477 (0.0479)	0.0570 (0.0246)**	-0.0065 (0.0059)
Skilled Occupation	0.0241 (0.0220)	0.0233 (0.0094)**	-0.0001 (0.0077)	0.0191 (0.0140)	—
Elementary Occupation	0.0002 (0.0030)	0.0093 (0.0080)	-0.0008 (0.0345)	0.0366 (0.0190)*	—
Medium Education	0.0152 (0.0177)	0.0013 (0.0034)	0.0158 (0.0454)	0.0041 (0.0138)	-0.0069 (0.0017)***
High Education	0.0176 (0.0359)	0.0026 (0.0081)	-0.0883 (0.0722)	0.0038 (0.0191)	0.0084 (0.0012)***
Rent	-0.0264 (0.0251)	-0.0037 (0.0085)	0.0048 (0.0130)	-0.0024 (0.0056)	—
Household Members	0.0002 (0.0021)	0.0021 (0.0045)	0.0260 (0.0207)	-0.0158 (0.0083)*	-0.0542 (0.0097)***
Children	-0.0100 (0.0182)	-0.0037 (0.0116)	-0.0015 (0.0278)	0.0114 (0.0071)	0.0489 (0.0060)***
<i>Unexplained (Coefficients)</i>	-0.0505 (0.0632)	-0.1094 (0.0306)***	0.0703 (0.0587)	0.1389 (0.0711)*	-0.0959 (0.0139)***
Skilled Occupation	-0.1122 (0.0627)*	-0.0511 (0.0380)	-0.0427 (0.0542)	-0.1021 (0.0617)*	—
Elementary Occupation	0.0082 (0.0121)	-0.0037 (0.0053)	-0.1798 (0.1557)	-0.0156 (0.0085)*	—
Medium Education	-0.1034 (0.0543)*	-0.0772 (0.0603)	-0.1258 (0.0913)	0.0118 (0.0312)	0.0107 (0.0199)
High Education	-0.0569 (0.1009)	-0.1530 (0.0859)*	-0.1607 (0.1067)	0.0805 (0.0731)	0.0163 (0.0195)
Rent	-0.0174 (0.0824)	-0.0232 (0.0464)	-0.0096 (0.0324)	-0.0049 (0.0321)	—
Household Members	—	-0.0422 (0.1256)	-0.1049 (0.1031)	-0.2670 (0.0999)***	0.1199 (0.0394)***
Children	—	-0.0350 (0.0419)	0.0384 (0.0366)	0.0088 (0.0102)	-0.0245 (0.0072)***
<i>Interaction</i>	0.0003 (0.0628)	0.0071 (0.0216)	0.0759 (0.0566)	-0.1639 (0.0600)***	-0.0021 (0.0081)
<i>N - first period</i>	168	496	151	492	2,010
<i>N - second period</i>	44	467	37	479	43,997

Table 18: Blinder–Oaxaca Decomposition of the Change in Relative Poverty among Young Household Leads (Under 30) - Part III

	PL	ES	SE	UK	US
Overall Difference	0.0258 (0.0070)***	-0.0105 (0.0610)	-0.1029 (0.0254)***	0.0139 (0.0108)	0.0052 (0.0089)
<i>Explained (Endowments)</i>	0.0113 (0.0028)***	0.0086 (0.0327)	0.0562 (0.0390)	-0.0137 (0.0074)*	0.0569 (0.0050)***
Skilled Occupation	0.0011 (0.0010)	-0.0104 (0.0101)	-0.0044 (0.0213)	0.0015 (0.0018)	0.0038 (0.0012)**
Elementary Occupation	0.0021 (0.0011)*	-0.0031 (0.0133)	0.0511 (0.0374)	-0.0029 (0.0031)	-0.0015 (0.0009)
Medium Education	-0.0227 (0.0078)***	0.0018 (0.0112)	-0.0156 (0.0167)	-0.0131 (0.0108)	-0.0171 (0.0039)***
High Education	0.0275 (0.0092)***	0.0167 (0.0227)	0.0174 (0.0120)	0.0121 (0.0079)	0.0537 (0.0073)***
Rent	0.0006 (0.0006)	-0.0003 (0.0131)	0.0064 (0.0044)	-0.0136 (0.0031)***	-0.0010 (0.0012)
Household Members	0.0016 (0.0028)	0.0050 (0.0086)	-0.0134 (0.0079)	0.0038 (0.0019)*	-0.0134 (0.0022)***
Children	0.0022 (0.0024)	-0.0006 (0.0035)	0.0132 (0.0085)	-0.0011 (0.0028)	0.0343 (0.0035)***
<i>Unexplained (Coefficients)</i>	0.0035 (0.0064)	-0.0084 (0.0647)	-0.0965 (0.0268)***	0.0392 (0.0132)***	-0.0459 (0.0082)***
Skilled Occupation	0.0250 (0.0103)**	0.0209 (0.1145)	0.0039 (0.0487)	0.0021 (0.0100)	-0.0001 (0.0108)
Elementary Occupation	-0.0035 (0.0027)	-0.0177 (0.0237)	-0.0094 (0.0083)	0.0189 (0.0093)**	0.0073 (0.0034)**
Medium Education	-0.0160 (0.0279)	0.0111 (0.0324)	-0.0114 (0.0406)	-0.0186 (0.0192)	0.0085 (0.0198)
High Education	-0.0119 (0.0175)	0.1959 (0.0902)**	0.0021 (0.0319)	-0.0497 (0.0312)	0.0267 (0.0229)
Rent	-0.0026 (0.0053)	0.0032 (0.0588)	-0.0511 (0.0292)*	0.0123 (0.0118)	-0.0120 (0.0111)
Household Members	-0.0124 (0.0192)	-0.1703 (0.1087)	0.0577 (0.0701)	0.0160 (0.0324)	0.0023 (0.0213)
Children	0.0166 (0.0066)**	0.0637 (0.0284)**	-0.0129 (0.0126)	-0.0046 (0.0096)	0.0026 (0.0051)
<i>Interaction</i>	0.0111 (0.0039)***	-0.0107 (0.0470)	-0.0626 (0.0396)	-0.0115 (0.0098)	-0.0058 (0.0041)
<i>N - first period</i>	2,704	276	955	1,765	8,517
<i>N - second period</i>	2,262	188	341	1,011	4,827

7 Discussion and conclusions

This study examined how the economic position of young adults evolved across 15 high-income countries between 1999 and 2020, using harmonized microdata from the Luxembourg Income Study. We focused on individuals who had already exited the parental home and assumed the role of household leads, those for whom independent living marks a critical transition in the life course. Our analysis combined descriptive statistics, baseline linear probability models and margins effects for more complex specifications, quantile regressions, and Oaxaca-Blinder decompositions to assess changes in relative poverty and income inequality, as well as some potential underlying drivers. The results reveal a somewhat consistent and concerning pattern. While not universal, as anticipated upfront, a general worsening trend is observed, especially in selected Western and Northern European countries (Norway, Germany, Netherlands, France). Across most of the examined 15 economies, household leads under 30 continue to face higher poverty risks compared to the older cohorts, a pattern that was evident already in the early 2000s, often accompanied by widening maximum penalties in earnings. Hence, the age-related poverty gradient has remained quite strong over time, with effects often comparable to those of tertiary education - though in opposite direction - and very few signs of convergence. Even in those contexts where relative poverty rates among the young have remained relatively stable over time, such as the United States, Canada or Israel that can be classified as liberal welfare regimes, the baseline risk continues to be particularly elevated. The descriptive statistics report on average 20-25% poverty levels among the youngest leads, and their constant term in the comparative linear probability models are among the highest in the sample (around or above 0.7). Therefore, what appears as stability in trends does not signify also favorable conditions, as it rather reflects the persistence of a marked vulnerability among the independent young adults. We acknowledge more favorable transitions for countries like Australia, the UK, and Poland but emphasize a continued disadvantage for young adults in most areas, with implications in terms of growing inequality in economic outcomes across generations. Moreover, although the descriptive statistics for most Northern and Western European countries may point to a rather short-lived exposure to relative poverty, as the prevalence of the condition diminishes markedly for the next group aged 31-40, the regression estimates show that age penalties are not negligible even

with respect to the closest age group. In many instances, these penalties amount roughly to 60% of the penalties observed when comparing the youngest group to the 51-64 age category.

When we extend the analysis to employed household leads, we further highlight that the disadvantage is not only shaped by access to employment, but also through the stratification within the labor market. The European countries, and especially Germany or the Netherlands, display some deteriorating returns in lower-skilled occupations affecting more disproportionately the younger cohorts. In those contexts that earlier showed generational convergence (UK, Poland), poverty risks among young adults have remained stable or even declined across occupational groups. In several country models, occupational status retains an independent effect, indicating that qualifications alone may not offer sufficient protection and that not only employment itself, but also its quality and the position occupied in the labor market play a central role for poverty dynamics. Moreover, among the 10 countries with available housing cost data, renting has sometimes become more prevalent and could have increased the exposure of the young cohorts to housing market pressures. The regression analyses performed among the employed renters provide some evidence that indeed poverty risk tends to be more pronounced for the young cohorts who face higher housing burdens - particularly for countries like Germany, Sweden, or Australia. By contrast, changes in poverty rates are less systematically associated with the housing burden for older leads. Based on the estimated margins, we can observe that young adults face, particularly in Europe, a double poverty exposure in terms of labor market stratification and declining housing affordability.

We also documented through quantile regressions a partial shift in both the size and the patterns of concentration of income penalties for the young household leads. In the most recent period, earnings penalties have generally remained persistent or intensified, while also increasingly shifting toward the lower end of the distribution for Continental European countries. This suggests a growing concentration of the disadvantage at lower deciles among young adults, as compared to liberal countries where the largest penalties continue to be observed predominantly at the top of the distribution. In Ireland, the UK, Poland, and Australia, the last period is remarkably characterized by very limited or absent earning penalties.

In an attempt to better clarify underlying mechanisms behind persistent age penalties across poverty risks and income distributions, we have turned to Oaxaca-Blinder decompositions. Again, the results show some heterogeneity in how the demographic or socioeconomic indicators relate to poverty trends. For the general population, the prevailing pattern across most European countries is one where modest or positive compositional shifts, such as higher education and in a few instances improved employment attachment, are outweighed by general negative returns. It would signify that households with otherwise similar observable characteristics face higher poverty risks in the later period. The youth-specific decomposition partially reinforce this pattern. In Continental and Nordic contexts, small or insignificant endowment effects are offset by negative structural coefficients, with higher poverty outcomes among young household leads. By contrast, in liberal welfare regimes such as the United Kingdom, the United States, Canada, and Australia (but not Israel), structural weakening returns are also present but more effectively neutralized by stronger compositional improvements, often through educational expansion. In these contexts, improvements in the groups' characteristics continue to contribute meaningfully to poverty reduction even if returns have broadly deteriorated and identification of precise channels remains limited.

Overall, we point out what seems an important difference between Continental Europe and liberal economies. In the former, youth poverty is not only high in prevalence terms but more broadly distributed. Being employed or holding a low- or mid-level occupational position does not provide to date strong protection for young adults. This suggests that young age disadvantage may be not confined to those outside employment, but extend to wider segments of working leads. In liberal countries such as the UK, US, Canada, and Australia, instead, employment and higher education continue to operate as strong protective factors, even if their returns have weakened over time. Poverty risks are thus more likely to remain concentrated among those with weaker labor market attachment and low qualification. The divergence across regimes is less tied to the existence of structural pressures or differences in household structures, but more to how effectively education or employment continue to translate into economic security for the young leads. To what pertains Southern European countries such as Italy and Spain, the poverty rates are relatively high across age groups

but young household display elevated earnings and poverty penalties relative to the oldest cohort of leads. If anything, the magnitude of the youth disadvantage could be underestimated because of a positive selection, as a smaller (and probably more affluent) subset exits the parental household, but the results hold as we extend the youth classifier to 35. By contrast, the strong economic growth and structural transformation characterizing Poland over the last two decades speak of a favorable outcome for the young cohort mainly through compositional improvement.

The conclusions stemming from the Oaxaca–Blinder decomposition have to be interpreted as indicative. They help to distinguish between more compositional and structural forces in their influencing role over poverty, however they rarely allow us to pinpoint specific institutional or market mechanisms behind the unexplained component. In many cases, the contribution of the individual covariates is not much robust, with the residual term capturing a combination of factors (that may go from wage-setting institutions to finer changes in the labor market segmentation) that are not easily observable with the data at hand. Missing dimensions on collective bargaining or contract stability may be relevant contributors to the unexplained component. Future research could develop this framework by building richer institutional indicators and finding alternatives to address the absence of coverage for certain variables in the early waves of the dataset. Nonetheless, the study outlines a comprehensive approach to examining the evolution of the economic conditions of young adults across high-income countries. By focusing on household leads under 30 — those who have exited the parental home and assumed economic responsibility — we capture a group for whom poverty and income disparities can have severe implications. The results suggest that in several advanced economies, the transition to independent adulthood increasingly unfolds under constrained labor markets and weakened returns to otherwise protective factors, pointing to the need for more targeted policies to support the youngest during their move to economic independence.

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Appendix

Table A1: Availability of key variables across LIS countries

Country	Years	Educ	Emp	Occ	Own	Region	H costs
Australia	2001–2018	P	✓	✓	✓	✓	✓
Austria	2000–2019	✓	✓	✓	✓	–	✓
Canada	2000–2019	✓	✓	–	✓	–	P
France	2000–2019	✓	✓	R	✓	✓	–
Germany	2000–2019	✓	✓	✓	✓	✓	✓
Ireland	2000–2019	✓	✓	✓	✓	–	✓
Israel	2001–2019	✓	✓	R	✓	✓	✓
Italy	2000–2020	✓	✓	R	✓	✓	✓
Netherlands	1999–2019	✓	✓	R	✓	–	P
Norway	2000–2019	✓	✓	–	P	–	–
Poland	1999–2019	✓	✓	✓	✓	✓	✓
Spain	2000–2019	✓	✓	✓	✓	–	✓
Sweden	2000–2019	✓	✓	R	✓	–	✓
UK	2000–2019	✓	✓	✓	✓	✓	✓
US	2000–2019	✓	✓	✓	✓	✓	–

Notes: ✓ = available in both waves.

R = reconstructed from auxiliary variable (`occ1_c`).

P = missing in one wave.

– = not available or not well harmonized.

Column labels correspond to the following variables: Educ = education level (three categories); Emp = employment status; Occ = occupation (`occa1`, three categories); Own = housing tenure (owner, rent, provided); Region = regional identifier (`region_c`); H costs = housing costs including rent and utilities (`hc4`). Variables such as sex, age, household size, and number of children are available for all countries and waves.

Table A2: Robustness Check: Alternative Youth Cutoff (Italy)

	Baseline (under 30)	Alternative (under 35)
Year 2020	-0.011 (0.073)	0.061 (0.044)
Age Groups (ref. youngest)		
Age 41–50	-0.107*** (0.029)	-0.076*** (0.021)
Age 51–64	-0.225*** (0.029)	-0.194*** (0.021)
Age × Year Interactions		
Age 41–50 × 2020	0.060 (0.075)	-0.012 (0.048)
Age 51–64 × 2020	0.132* (0.074)	0.060 (0.046)
Sex		
Female	0.015 (0.012)	0.015 (0.012)
Education		
Medium education	-0.105*** (0.013)	-0.104*** (0.013)
High education	-0.176*** (0.014)	-0.175*** (0.014)
Employment		
Employed	-0.272*** (0.019)	-0.272*** (0.019)
Household Characteristics		
Household members	-0.009 (0.013)	0.009 (0.013)
Children	0.059*** (0.016)	0.060*** (0.016)

Notes: Linear probability models with robust standard errors clustered at household level in parentheses. The baseline specification defines young household leads as aged 30 or below. In the alternative specification, the cutoff is extended at age 35.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A3: Robustness Check: Alternative Youth Cutoff (Spain)

	Baseline (30)	Alternative (35)
Year 2019	0.014 (0.042)	0.023 (0.027)
Age Groups (ref. youngest)		
Age 41–50	-0.052* (0.030)	-0.039* (0.022)
Age 51–64	-0.135*** (0.030)	-0.122*** (0.022)
Age × Year Interactions		
Age 41–50 × 2019	0.032 (0.045)	0.022 (0.032)
Age 51–64 × 2019	0.048 (0.044)	0.038 (0.031)
Sex		
Female	0.027** (0.010)	0.027** (0.010)
Education		
Medium education	-0.091*** (0.013)	-0.091*** (0.013)
High education	-0.138*** (0.010)	-0.138*** (0.010)
Employment		
Employed	-0.227*** (0.017)	-0.228*** (0.017)
Household Characteristics		
Household members	-0.033 (0.015)	-0.033 (0.015)
Children	0.052*** (0.018)	0.052*** (0.018)

Notes: Linear probability models with robust standard errors clustered at household level in parentheses. The baseline specification defines young household leads as aged 30 or below. In the alternative specification, the cutoff is extended at age 35.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A4: Robustness Check: Alternative Youth Cutoff (Ireland)

	Baseline (30)	Alternative (35)
Year 2019	0.042 (0.048)	0.025 (0.025)
Age Groups (ref. youngest)		
Age 41–50	-0.013 (0.033)	0.001 (0.027)
Age 51–64	-0.051 (0.034)	-0.038 (0.028)
Age × Year Interactions		
Age 41–50 × 2019	-0.059 (0.052)	-0.043 (0.032)
Age 51–64 × 2019	-0.050 (0.053)	-0.033 (0.033)
Sex		
Female	-0.009 (0.012)	-0.009 (0.012)
Education		
Medium education	-0.058*** (0.017)	-0.055*** (0.017)
High education	-0.091*** (0.014)	-0.091*** (0.014)
Employment		
Employed	-0.302*** (0.025)	-0.303*** (0.025)
Household Characteristics		
Household members	-0.090*** (0.013)	-0.090*** (0.013)
Children	0.083*** (0.015)	0.083*** (0.015)

Notes: Linear probability models with robust standard errors clustered at household level in parentheses. The baseline specification defines young household leads as aged 30 or below. In the alternative specification, the cutoff is extended at age 35.

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A5: Linear Probability Model of Relative Poverty among Employed Household Leads

	AU	AT	CA	FR	DE
Sex					
Female	0.01* (0.01)	0.02** (0.01)	0.05*** (0.01)	0.01** (0.00)	0.03*** (0.01)
Age group (ref. under 30)					
31–40	0.03 (0.02)	-0.16 (0.11)	-0.04*** (0.01)	-0.02** (0.01)	-0.09*** (0.03)
41–50	0.00 (0.02)	-0.16 (0.11)	-0.06*** (0.01)	-0.02** (0.01)	-0.10*** (0.03)
51–64	0.01 (0.02)	-0.16 (0.11)	-0.05*** (0.01)	-0.01 (0.01)	-0.09*** (0.03)
Education (ref. low)					
Medium education		-0.05** (0.02)	-0.03*** (0.01)	-0.03*** (0.00)	-0.12*** (0.02)
High education		-0.03 (0.03)	-0.08*** (0.01)	-0.04*** (0.00)	-0.14*** (0.02)
Occupation (ref. managers/professionals)					
Skilled occupations	-0.01 (0.01)	-0.11 (0.11)		0.03*** (0.01)	0.01 (0.03)
Elementary occupations	0.01 (0.03)	-0.21* (0.11)		0.13*** (0.02)	0.09 (0.09)
Year × Occupation					
Skilled × second period	0.06*** (0.02)	0.16 (0.12)		-0.04 (0.03)	0.04 (0.06)
Elementary × second period	0.07* (0.04)	0.38** (0.16)		-0.02 (0.06)	0.26** (0.12)
Household characteristics					
Household size	-0.02*** (0.00)	-0.02* (0.01)	-0.04*** (0.00)	-0.04*** (0.00)	-0.07*** (0.01)
Number of children	0.04*** (0.01)	0.03** (0.01)	0.05*** (0.00)	0.05*** (0.00)	0.08*** (0.01)
Tenure (ref. own)					
Rent	0.01 (0.01)	-0.02 (0.02)	0.08*** (0.01)	0.03*** (0.00)	0.02*** (0.01)
Rent × second period	0.01 (0.01)	0.06*** (0.02)	0.02 (0.01)	0.03*** (0.01)	0.03*** (0.01)
Region control	Yes	No	No	Yes	Yes
Occupation x Age Group x Year	Yes	Yes	No	Yes	Yes
Observations	12,418	4,748	59,510	44,093	17,284
R-squared	0.03	0.04	0.06	0.07	0.14

Table A6: Linear Probability Model of Relative Poverty among Employed Household Leads

	IE	IL	IT	NL	NO
Sex					
Female	0.01 (0.01)	0.04*** (0.01)	0.01 (0.01)	0.02*** (0.01)	0.02*** (0.00)
Age group (ref: under 30)					
31–40	-0.01 (0.03)	-0.06** (0.03)	-0.08 (0.05)	-0.33** (0.14)	-0.04*** (0.01)
41–50	0.00 (0.04)	-0.10*** (0.03)	-0.09 (0.05)	-0.32** (0.13)	-0.04*** (0.01)
51–64	0.00 (0.03)	-0.05* (0.03)	-0.07 (0.06)	-0.30** (0.13)	-0.04*** (0.01)
Second Period	-0.03 (0.04)	0.06 (0.04)	0.04 (0.08)	-0.28** (0.14)	0.01 (0.01)
Education (ref. low)					
Medium education	-0.01 (0.02)	-0.10*** (0.02)	-0.02 (0.02)	-0.03*** (0.01)	-0.02*** (0.00)
High education	-0.03* (0.02)	-0.12*** (0.02)	-0.05*** (0.02)	-0.02 (0.01)	-0.03*** (0.00)
Occupation (ref: managers/professionals)					
Skilled occupations	0.04 (0.05)	0.11*** (0.04)	0.00 (0.06)	-0.25* (0.14)	
Elementary occupations	0.21 (0.18)	0.14* (0.08)	0.08 (0.07)	-0.22 (0.14)	
Year × Occupation					
Skilled × second period	0.13 (0.09)	-0.01 (0.05)	-0.09 (0.09)	0.34** (0.14)	
Elementary × second period	-0.21 (0.18)	0.11 (0.15)	-0.19* (0.11)	0.48*** (0.18)	
Household characteristics					
Household size	-0.06*** (0.01)	-0.04*** (0.01)	-0.01 (0.01)	-0.04*** (0.01)	-0.04*** (0.00)
Number of children	0.07*** (0.01)	0.07*** (0.01)	0.06*** (0.01)	0.06*** (0.01)	0.04*** (0.00)
Tenure (ref. own)					
Rent	0.06 (0.04)	0.03* (0.02)	0.02 (0.02)	0.06*** (0.01)	
Rent × second period	-0.02 (0.04)	0.00 (0.02)	0.03 (0.03)	0.03* (0.02)	
Region control	No	Yes	Yes	No	No
Occupation x Age Group x Year	Yes	Yes	Yes	Yes	No
Observations	3,746	8,699	5,131	9,411	247,493
R-squared	0.08	0.12	0.15	0.10	0.04

Notes: Norway models do not include occupation, tenure, region controls, or occupation/year interaction terms. Standard errors clustered at the household level. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.10$.

Table A7: Linear Probability Model of Relative Poverty among Employed Household Leads

	PL	ES	SE	UK	US
Sex					
Female	0.01* (0.00)	0.04*** (0.01)	0.02*** (0.01)	0.02*** (0.00)	0.06*** (0.00)
Age group (ref: under 30)					
31–40	0.00 (0.01)	-0.12* (0.07)	-0.02 (0.01)	0.00 (0.01)	-0.02** (0.01)
41–50	-0.01* (0.01)	-0.12* (0.07)	-0.03** (0.01)	-0.01* (0.01)	-0.02*** (0.01)
51–64	-0.01 (0.01)	-0.13* (0.07)	-0.02** (0.01)	0.01 (0.01)	0.01 (0.01)
Second period	0.02** (0.01)	-0.15* (0.08)	0.03 (0.04)	0.01 (0.01)	0.03** (0.01)
Education (ref: low)					
Medium education	-0.10*** (0.01)	-0.08*** (0.01)	-0.04*** (0.01)	-0.05*** (0.01)	-0.18*** (0.01)
High education	-0.12*** (0.01)	-0.10*** (0.01)	-0.04*** (0.01)	-0.06*** (0.01)	-0.24*** (0.01)
Occupation (ref. managers/professionals)					
Skilled occupations	0.06*** (0.01)	-0.04 (0.09)	0.04* (0.02)	0.03*** (0.01)	0.11*** (0.01)
Elementary occupations	0.08*** (0.02)	-0.04 (0.10)	0.04** (0.02)	0.18*** (0.04)	0.24*** (0.02)
Year × Occupation					
Skilled × second period	-0.05*** (0.01)	0.14 (0.10)	0.02 (0.05)	-0.00 (0.02)	0.01 (0.01)
Elementary × second period	0.03 (0.04)	0.38** (0.17)	0.21 (0.14)	-0.11** (0.05)	-0.06* (0.03)
Household characteristics					
Household size	-0.01*** (0.00)	-0.03* (0.02)	-0.05*** (0.01)	-0.03*** (0.00)	-0.04*** (0.00)
Number of children	0.02*** (0.00)	0.05*** (0.02)	0.06*** (0.01)	0.05*** (0.00)	0.08*** (0.00)
Tenure (ref. own)					
Rent	0.00 (0.01)	0.05* (0.03)	-0.00 (0.01)	0.09*** (0.01)	0.10*** (0.00)
Rent × second period	0.00 (0.01)	0.04 (0.04)	0.06*** (0.02)	-0.04*** (0.01)	0.00 (0.01)
Region control	Yes	No	No	Yes	Yes
Occupation x Age Group x Year	Yes	Yes	Yes	Yes	Yes
Observations	41,592	10,696	10,347	24,027	92,227
R-squared	0.04	0.07	0.08	0.06	0.16

Table A8: Blinder–Oaxaca Decomposition of the Change in Relative Poverty among Young Household Leads (Under 35)

	IE	IT)	ES
Overall Difference	0.0332 (0.0259)	0.0185 (0.0383)	0.0005 (0.0286)
<i>Explained (Endowments)</i>	0.0007 (0.0178)	-0.0143 (0.0308)	0.0106 (0.0129)
Skilled Occupation	0.0112 (0.0071)	-0.0035 (0.0057)	0.0013 (0.0026)
Elementary Occupation	-0.0002 (0.0017)	0.0137 (0.0152)	-0.0004 (0.0025)
Household Members	-0.0012 (0.0033)	-0.0003 (0.0046)	-0.0025 (0.0035)
Children	0.0021 (0.0068)	0.0138 (0.0176)	0.0041 (0.0046)
Medium Education	0.0271 (0.0128)**	0.0141 (0.0164)	0.0074 (0.0058)
High Education	-0.0240 (0.0164)	-0.0565 (0.0405)	0.0189 (0.0107)
Rent	-0.0144 (0.0093)	0.0013 (0.0047)	-0.0150 (0.0101)
<i>Unexplained (Coefficients)</i>	0.0310 (0.0295)	-0.0047 (0.0345)	-0.0097 (0.0293)
Sex (Lead)	0.1384 (0.0644)**	0.0204 (0.0578)	0.0355 (0.0816)
Skilled Occupation	-0.0260 (0.0208)	-0.0180 (0.0321)	-0.0073 (0.0477)
Elementary Occupation	0.0095 (0.0101)	-0.0308 (0.0550)	0.0037 (0.0110)
Household Members	-0.0741 (0.1085)	0.0005 (0.0662)	-0.2123 (0.0850)**
Children	0.0527 (0.0386)	-0.0368 (0.0428)	0.0973 (0.0295)***
Medium Education	-0.0609 (0.0197)***	-0.0772 (0.0528)	0.0073 (0.0203)
High Education	-0.1408 (0.0504)***	-0.0869 (0.0584)	0.0761 (0.0482)
Rent	0.0410 (0.0335)	-0.0233 (0.0257)	-0.0140 (0.0283)
<i>Interaction</i>	0.0015 (0.0254)	0.0375 (0.0269)	-0.0004 (0.0170)
<i>N - first period</i>	376	571	823
<i>N - second period</i>	211	134	771

Response to Reviewers' Comments on Paper 3

Elena Marseglia

1. Introduction

Reviewer 1

Please be more explicit about the key contribution of this chapter. How does your study differentiate from existing research on family transfers and fertility intentions? What specific gaps does this chapter address that previous studies have not adequately covered? A clearer statement of contribution would help readers understand the value added by your analysis.

Response:

In the refined introduction we frame the issue of persistently low fertility and the gap between fertility intentions and outcomes, emphasizing the role of economic pressures or perceived uncertainty for prospective parents. We point out that despite research on the economic and institutional determinants of fertility, little attention has been paid to financial transfers per se as a potential buffer against these constraints. We stress that transfers are potentially socially stratified, and this motivates our effort to disentangle the role of financial support from family advantages. The contribution is further elaborated in the theoretical background where we bring together economic, intergenerational solidarity and institutional perspectives while we claim the absence of studies which systematically examine the relationship between private financial transfers and short-term fertility intentions using harmonized, cross-national data across multiple survey rounds. [pages 2-3, 7].

2. Background

Reviewer 1

- The literature review would benefit from addressing which parents are able to transfer financial support to their children. Given that these parents represent a selective segment of the population (and their children are therefore also selected), this warrants explicit discussion particularly since some of your significant findings disappear once you adjust for parental background. This suggests that parental background is an important variable that should not be overlooked in the literature review. Even a few sentences acknowledging this selectivity would strengthen your theoretical framework and help motivate your choice of control variables. While you review substantial literature on social support more broadly, your focus is on financial support specifically. It would be helpful to discuss whether different support types (e.g., care support versus financial support) might have differential or even substitutive effects on fertility intentions. For instance, care support might play a more important role than financial support in certain contexts. Does one type of support prevail in influencing fertility intentions, or might they serve as substitutes for one another? Addressing this question would provide a more comprehensive theoretical foundation.

- Please clarify which literature supports Hypothesis 3, particularly regarding how institutional and economic conditions mediate the relationship between financial transfers and fertility intentions across periods or cohorts. The current background section lacks discussion of these contextual factors, yet they appear central to your third hypothesis. A more explicit engagement with this literature would strengthen the theoretical justification for H3.

Reviewer 2

- I found the Social Support Theory and the Cognitive social model pertinent. The theoretical framing could be strengthened if you consider more perspectives. Both highlight how material and emotional support reduce stress and shape fertility intentions. However, they remain general and do not fully capture the specific mechanisms at play with financial transfers you want to investigate. For example, Intergenerational Solidarity Theory (Bengtson Roberts, 1991) could give you a direct lens for understanding monetary transfers as part of broader patterns of exchange and reciprocity between generations. Also, Elder's Life Course Perspective (1994) would allow you to situate these transfers within "linked lives", emphasizing how parental resources shape children's family transitions over time.

- I agree with you that most literature on fertility has focused on public policies, employment, and

childcare provision (see Matysiak 2011, Thévenon 2011, Gauthier 2007) and that studies on private transfers and fertility are few and fragmented (page 5, paper 3). So, the statement that research on private financial transfers and fertility intentions is still limited is correct, but the literature review could be expanded slightly, in my opinion. There is also a small body of work on private intergenerational support and related life transitions (e.g., Albertini, Kohli & Vogel 2007 on transfers in Europe; Aassve, Cottini & Vitali 2013 on family support during recessions; Mulder & Smits 2013; Kuin & Mulder 2021 on family ties and housing pathways). Maybe you want to quote them – if useful – to broaden the extent of your analysis or if you find them useful to better frame your research question.

Response :

Thanks to reviewers' comments, we have clarified and enriched the theoretical framework [Pages 4-7].

Following the argument surrounding the selection into transfers in the introduction, the background section now better situates financial transfers within intergenerational resource flows and motivates our attention to parental background in the empirical strategy. We also specified the distinction between financial transfers and other forms of support. Bridging the different theories, we argue that monetary transfers primarily operate through liquidity provision and risk reduction, which may complement or substitute time-based support depending also on the institutional context, building on the framework by Saraceno and Keck (2010).

We also strengthen the justification for Hypothesis 3 by integrating the suggested studies on intergenerational transfers and life-course transitions that help contextualize financial support within patterns of housing transitions, early adulthood, and economic vulnerability. We therefore argue that financial transfers are likely to influence fertility intentions when received in closer proximity to family transitions and suggest that private transfers may operate more strongly in recent contexts, building on the literature on economic uncertainty and post-Great Recession instability.

3. Data and methods

Reviewer 1

- In the GGS longitudinal component, do new people join in every wave? How did you address this issue? Did you focus only on respondents who were present in all waves? Please clarify your sample construction.
- How did you address attrition? How substantial was attrition for each country from the first wave to the last wave? This information is important for assessing the robustness of your findings.
- I am puzzled by your choice of pooled logit regression for longitudinal data (Round 1) instead of

individual fixed effects models, which would be a substantially stronger approach. You may choose to continue with pooled regression for this dissertation, but if you intend to develop this chapter into a journal article, you will likely receive this comment from reviewers. At a minimum, you need to make an explicit case for why pooled logit is a more appropriate approach than individual fixed effects for your research question and data structure.

Reviewer 2

- In the GGS Round II, the variable on financial transfers is indeed very limited: in most countries it is formulated as “Have you ever received a large financial transfer or gift from your parents or relatives?”, with no information on when the transfer took place. Only in a few countries (such as Austria or Germany) is there a follow-up item distinguishing whether the transfer occurred recently (for example, within the last three or five years) or earlier, but this information is not harmonized and is not available for all countries. So, when you write “where possible, we distinguish more recent from older transfers” you are probably referring to those few countries where the question includes some indication of timing, but it is not clear how this was implemented or in which cases. In other words, the sentence to me sounds a bit ambiguous, as it suggests that temporal information (when the bequests were made) that is in fact missing in most Round II datasets. Also, you should make more clear in which countries, and by what procedure, it was possible to distinguish between recent and older transfers. To sum up, it is not entirely clear how the analysis distinguishes “more recent” from “older” transfers in Round II, given that the GGS variable on financial support generally captures whether respondents have ever received a large transfer, with no harmonized reference to timing. If some countries include additional questions on when the transfer occurred, you could specify which ones and how this distinction was operationalized. As written, the phrase could be misleading, since the available data do not consistently allow such differentiation.

- You apply two different analytical strategies for Round I (logistic and fixed-effects models) and Round II (propensity score matching). This makes sense given the different data structures (panel versus cross-sectional) but I wonder to which extent the models, and their findings could be comparable. The results from Round I and Round II are not directly comparable, since the estimation methods rely on different identifying assumptions and produce different types of coefficients (log-odds versus matched differences). Also, the fixed-effects models in Round I capture within-person changes over time, whereas the PSM in Round II adjusts only for observable cross-sectional differences. I understand why you justify this two-facet approach, but it would be more informative for your readers that you acknowledge its implications more explicitly. Readers should be reminded that differences between Round I and Round II results may reflect methodological rather than substantive variation. You might also note that PSM cannot replicate the advantages of panel

estimation, especially the ability to control for unobserved, time-invariant heterogeneity, so the interpretation of results across rounds should remain cautious. If your long-term goal is a publishable and methodologically solid paper, then the FE approach from Round I would be the stronger choice. You control for unobserved time-invariant heterogeneity and you have a more “causal” interpretation. At the same time, I reckon that the receipt of parental transfers in the GGS is not homogeneous across waves. In round I, the variable captures a recent, short-term transfer. In round II, you have a lifetime measure. What if the answer is “yes” to both? Can you assume that a child received transfers both before Round I and during Round I and II? Not really, because the Round II “yes” does not imply the bequest reception in between the two rounds. To wrap it up, I see some downside in both strategies (two-round FE and cross section PSM), and I would not present a paper burdened by a potential double liability. I’d rather work on two separate pieces of research to limit their cons as much as you can.

Response :

In the revised section, we have clarified more thoroughly both GGS design and our strategies to deal with the data structure, and better harmonize the two rounds for the questions of interest without overstating strict comparability between the 2 Rounds [Pages 8-11].

In Round I, the GGS follows a closed panel design, so there is no refreshment with new respondents entering in subsequent waves. We better specify that the sample includes the individuals observed at least once within the eligible age range (no strict balanced panel). We therefore exploit the existing longitudinal variation while minimizing the loss of observations due to attrition.

We now better document the attrition patterns for each country, and report the country-specific retention rates (Tables 2-4). We acknowledge the issue of selection attrition, in particular we find that attrition follows usual life-course patterns (stayers are usually in better economic condition and more stable family arrangements). Yet, we also observe that only in the Netherlands recipients of parental resources are somewhat over-represented while there are minimal differences in the reported family intentions between stayers and attriters. We partially address this issue by running in this first Round random-effects models with Mundlak term to account for the time-invariant unobserved heterogeneity.

To better exploit the longitudinal variation in round I, we have now opted for random-effects models with a Mundlak correction rather than individual fixed-effects logit models. While this latter offers a stricter identification strategy, our data structure and the relatively rare transitions in transfer receipt would reduce excessively the statistical power. This new approach allows us to still

control for time-invariant heterogeneity, but also to retain both the between- and within-individual heterogeneity which is relevant for our research question [Pages 10, 20].

We have better clarified the operationalization of the financial transfers using the scheme in Table 1. In Round I, financial transfers are measured as support received 12 months preceding the interview, with a clear reference window. We have 2 possible indicators for the preceding year: any parental transfer, and the more restrictive large parental transfer. In Round II, respondents report whether they have received any large inter vivos transfer, by whom and the year of the most recent transfer. Therefore, we have again two possible indicators: one defined over the life time, and the other restricted on the same year or year preceding the interview. The large recent transfer is the only one that presents some degree of comparability across rounds and the most important in the analysis. [Pages 9-10].

We agree that in the last part of our analyses we rely on different identification strategies depending on the round under consideration, as the two rounds are not directly comparable given the nature of the data and the transfer measurement. The two approaches are meant to address mainly the common concern of non-random selection into transfer receipt making use of the structure of the data, therefore we interpret cross-round differences more cautiously and specifically mention it as a limitation. In the analysis part, we present the two rounds sequentially and independently, with the aim of extracting the most from each data structure rather forcing comparability. [Pages 10, 42-43].

4. Discussion

Reviewer 1

The conclusion remains quite limited, primarily summarizing results without sufficient interpretation. This is particularly notable given that you analyze data from five different countries with distinct institutional contexts and welfare regimes. I would have expected more elaboration on why you find different results across countries and how these differences can be explained by country-specific characteristics. Additionally, since your research question concerns fertility intentions, it would be worthwhile to provide more information about each country's fertility policies. It would be valuable to discuss how the influence of financial support might interact with the presence or absence of policies designed to encourage higher fertility levels. For instance, might parental financial transfers matter more in contexts where state support for young families is limited? Addressing these questions would significantly strengthen the contribution of this chapter and help readers

understand the policy relevance of your findings.

Reviewer 2

I noticed that the analytical sample size for some countries changes substantially between Models 1–2 and 3–4. Let me take Austria as an example: 4,997 observations and 3,111 respondents (1-2) and (4,359 / 2,709; 3-4). It would be useful to clarify why these differences occur, for instance, whether they result from missing data on additional covariates introduced in later models, sample attrition across waves, or other restrictions applied during model estimation. Similarly, the main explanatory variable “Transfer received” varies in its statistical significance across models: significant at the 1% level in Model 1, non-significant in Model 2, again non significant in Model 3, and significant at 5% in Model 4. I could not understand whether these shifts are primarily due to the sequential inclusion of control variables or to changes in the analytical sample. As it is currently presented, to me it’s not entirely clear how these two sources of variation interact.

Response:

We have deeply changed the conclusion to move beyond a descriptive summary of results and attempt to provide an explicit interpretation of cross-country differences in light of the institutional context or patterns of fertility. We elaborate that the positive associations concentrated in Austria and Germany may relate to their position as regimes of moderate supported familialism, where parental responsibilities in childrearing remain central despite quite developed family policies. Yet, formal tests of institutional moderation, using country-level indicators of family spending, childcare costs for dual earners and female labor force participation, do not reveal stable interaction effects. We interpret the finding as evidence that private transfers may not operate through single policy instruments but more through the broader institutional setting [Pages 42-43].

We acknowledge the issue on sample size in the limitations at the end. The analytical sample may vary because of missing information (especially among socio-economic and parental background variables), rather than because of panel attrition. In Round I, baseline and demographic models rely on nearly identical samples, which means the attenuation of the transfer effects in the main models is primarily linked to the covariate adjustment rather than to sample composition. In the second Round the reduction in sample size is more substantial, and the limited fluctuations seen in statistical significance could be attributable also to estimation precision and composition, but are still unlikely to mirror reversals in the relationship between parental transfers and fertility intentions based on estimates and standard errors. [Page 43].

Do Private Family Transfers Shape Fertility Intentions? Evidence from the Generations and Gender Survey

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1 Abstract

This paper explores the relationship between receiving private financial transfers from parents (or from parents and grandparents jointly) and short-run fertility intentions in Austria, Germany, the Netherlands, Estonia, and the Czech Republic. We employ longitudinal data from the 2000s and the more recent cross-sectional data collected after 2020 which are available through the internationally renowned Generations and Gender Survey. Fertility intentions are measured as respondents' plans to conceive a (another) child within the next three years. Our strategy follows the structure of the data, with pooled and country-specific logistic regression analyses with ample control variables accompanying random-effects models for Round I and propensity score matching for Round II. Although in the earlier data large transfers are positively associated with fertility intentions in Germany—and more weakly in the Netherlands—in single-country and pooled logistic models, these effects do not persist in random-effects specifications, likely reflecting limited statistical power. When observed, the associations are also stronger among parents than among childless respondents. By contrast, evidence from the latest period points to more consistent positive associations with large and recent parental transfers in specific national contexts. In Germany and Austria, large parental transfers are associated with higher probabilities of intending a child, which range between 3% and 5%, and go as high as 8% when restricted to recent large transfers alone. Only for Germany we find the association to be stronger for parents. Our findings moreover suggest that the role of private family is at most context-dependent; the concentration of positive effects in Western Europe, and notably in Austria and Germany in the last years, can be consistent with settings where public support for families coexists with parental involvement, although the direct measures of institutional moderation yield short evidence. In doing so, the article provides an empirical assessment of both the potential and the limits of the existing survey data for studying the link between informal financial support and fertility intentions in Europe.

2 Introduction

Over the last decades, declining fertility rates and delayed family formation have emerged as defining demographic traits of advanced European economies. Throughout Europe, most countries have seen their total fertility rates (TFR) drop below the replacement level of 2.1 children per family with Southern and Central Eastern Europe displaying even "lowest-low" fertility levels (TFRs below 1.5) (Frejka & Sobotka, 2008; Kohler et al., 2006). These demographic patterns present considerable social and economic challenges like rapid aging of the population and greater strain on social welfare systems, raising concerns about the economic sustainability and social cohesion within the continent.

A growing body of work refers to the existing gap between fertility intentions and fertility outcomes as individuals often report to desire more children than they ultimately have. (Goldstein et al., 2003; Testa, 2014). This disparity indicates that factors beyond personal preferences - such as economic pressures and perceived uncertainty - may play a decisive role in shaping reproductive outcomes (Adsera, 2005). In this context, financial resources that reduce economic pressures and enhance a sense of security may empower individuals to overcome these barriers and have their desired number of children. Yet, despite mounting research on the social and economic determinants of fertility, there is a limited focus in the literature on how private support from one generation to the next affects people's childbearing intentions and behavior. In many European societies, families continue to represent relevant providers of informal financial security for younger generations, and especially during pivotal life-course transitions like partnership formation, housing acquisition, and childbearing (Aassve et al., 2013; Albertini et al., 2007). Moreover, the capacity to provide financial support is itself socially stratified. Those parents able to pass on substantial resources to their children tend to come from more advantaged background, which implies a positive selection of adult children into exposure to parental resources. As a result, the association between parental transfers and fertility intentions may reflect underlying family advantages rather than the transfer itself. Whether and to what extent financial transfers from ascendants are associated with short-term fertility planning remains an open empirical question. By explicitly accounting for a rich set of parental background characteristics, this study seeks to address this issue and disentangle the association between financial support and fertility intentions from pre-existing family advantages, allowing for a more direct evaluation of private transfers as a distinct resource.

We attempt to do so by using harmonized data from two rounds of the Generations and Gender Survey (GGS). We focus on five countries—Austria, Germany, the Netherlands, Estonia, and the

Czech Republic—which are observed in both survey rounds conducted roughly a decade apart. The focus on these specific countries is primarily driven by data availability rather than by a predetermined comparative design. Yet, this group of countries serves as a theoretically informative and relevant sample. In fact, all five share to some degree below-replacement fertility and delayed family formation, while at the same time exhibiting variation in the welfare state arrangements and in the ways obligations toward children are distributed between public institutions and private households.

Methodologically, we combine country-specific and pooled logistic regression models with random-effects specifications for the longitudinal data in Round I, and propensity score matching for the cross-sectional data in Round II. This strategy allows us to some extent to address selection issue, test the robustness of association in different contexts, and examine the existence of heterogeneity by parity or timing of transfers in the second round. Across these complementary approaches, the results show a modest, and limited to few settings, association between parental (or ascendants') financial transfers and fertility intentions. In the earlier period, there is little systematic evidence of the relevance of family transfers—likely also reflecting the shortage of large transactions—whereas in more recent years we observe more consistent associations, and especially for large and recent transfers, in Austria and Germany. This study is meant to provide a new comparative insight into the role of informal family resources in low-fertility settings and underscores the importance of accounting for the institutional context and timing when evaluating the informal welfare function of the family in contemporary Europe.

3 Background

To understand how private financial transfers can affect fertility intentions, there is a need to bring together ideas of economic, sociological, and psychological theories. Fertility intentions and behavior, across these perspectives, are not motivated solely by individual preferences, but by the interaction between perceived costs and resources, as well as the role of intergenerational relationships. Theoretical explanations about low fertility in advanced societies have developed along complementary lines (Tanturri, 2014). On one side, economic approaches rooted in rational choice theory have emphasized how fertility is negatively correlated with economic development, mainly due to the rising direct and opportunity costs of childbearing (Becker, 1981; Willis, 1973). Prolonged education, constraints on housing in urban areas or interruptions in employment histories of mothers are all factors that would drive up the financial burden associated with parenthood, according to this view. The perceived affordability of children is also driven by subjective economic uncertainty rather than just by current income levels or structural conditions (Kreyenfeld, 2010). On the other side, cultural explanations linked to the Second Demographic Transition remark how societal changes and the spread of post-materialist values underlie a broader de-standardization of family trajectories and an individualistic life course, alleviating the social obligation attached to reproduction and making it increasingly a matter of choice (Lesthaeghe, 1995; Van de Kaa, 1987).

Bridging somehow these views, perspectives on the life course insist on how delayed transition to adulthood, labor market uncertainty, and expansion of tertiary education combined lead individuals to postpone reproductive decisions until key relational and economic conditions are met, with structurally reduced fertility levels in the long-term (Blossfeld et al., 2005; Kohler et al., 2002). New insights from institutional and gender-equity theories also underscore how expanded opportunities for women in the public sphere have not always been matched by a renegotiation of the division of labor within the family, with the resulting double burden potentially discouraging fertility (McDonald, 2000). In this context, the level and extent of welfare support for greater gender equity and work-family reconciliation have been identified as relevant levers in influencing reproductive choices across Europe. Further on, studies on well-being support the idea that childbearing is more likely when individuals feel more supported as key life decisions reflect perceptions of security and expectations about future life satisfaction (Aassve et al., 2012; Billari, 2009).

Across these perspectives we can derive the implication that fertility outcomes are not solely driven by changing preferences, but also by constraints that may prevent the realization of reproduc-

tive plans. Low fertility should therefore not only be interpreted as showcasing a decline in desire, but mainly as the result of unmet intentions arising from perceived or material obstacles. In this context, financial resources that relax economic pressures and enhance a sense of security may represent one instrument whereby individuals overcome barriers and achieve their intended family size. Family transfers - especially inter vivos support received early in life - can help offset part of the rising cost of child rearing. These transfers may be seen as means of easing credit restrictions or lowering the perceived financial risk of having children.

The Intergenerational Wealth Flows theory by Caldwell (1976) provides a macro-level lens on this process. He argues that fertility patterns reflect the dominant direction of intergenerational transfers: when flows move from younger to older generations, high fertility is economically rational; when they reverse—as in modern welfare states—children become a net economic cost, leading hence to lower fertility. In low-fertility contexts, then, private transfers from parents to children can compensate for this reversal, especially when received before or during the childbearing years. Unlike bequests, which arrive too late to influence early family formation, early-life transfers provide liquidity at more crucial moments for family planning. The way private financial transfers are embedded in intergenerational relationships is further conceptualized by the Intergenerational Solidarity Theory. According to it, transfers are an exchange in a multidimensional spectrum between parents and their adult children, encompassing functional, affective, and normative dimensions (Bengtson & Roberts, 1991). Monetary transfers are not to be considered in isolation, but as belonging to broader expectations over family cohesion, which can have an influence on how the individuals perceive both feasibility and desirability of childbearing. Transfers received in contexts of strong solidarity between generations may be associated in this sense with a deep commitment of family continuation, which reinforces the fertility intentions beyond what would be otherwise predicted by mere individual cost-benefit calculations.

The linked lives perspective taken by (Elder, 1994) goes further emphasizing relevant aspects like timing and agency under constraint, with parental resources and support more likely to matter when they occur in proximity of family transitions and individuals actively plan their behavior. This opens the way to the hypothesis that transfers may have also different implications depending on life stage, relational closeness, and parity. Linked to this idea of momentum, Social Support Theory also insists on how both practical and emotional help from family can lessen the mental stress tied to major life changes like having children (House, 1981). In this view, fertility intentions are influenced by material constraints as well as personal feelings about security and available support. The Cognitive-Social Model also highlights how people create and adjust their fertility plans throughout their lives. This

adjustment happens in response to their sense of stability, relational support, and social expectations (Bachrach & Morgan, 2013).

The relevance of private financial transfers is further conditioned by institutional context. Studies of the welfare state show that there is considerable cross-national variation when it comes to the division of responsibilities between the family, the market, and the state regarding the provision of care and economic support (Esping-Andersen, 1990; Gauthier, 2002). In countries where there is widespread and consistent support from the state, private transfers may serve to reinforce or complement public provision, whereas in familialistic or transforming welfare regimes, private transfers may substitute for limited public assistance (Albertini & Kohli, 2013; Dykstra & Fokkema, 2011; Fenger, 2007). Building on this tradition, the influential framework by Saraceno and Keck (2010) on intergenerational policy regimes emphasize specifically the degree of familialism embedded in welfare states, that is the extent to which families are expected to act as primary providers of economic support and care to the younger and the older. Their typology distinguishes between contexts characterized by stronger de-familialization, where the market or the state relieve families of care, and those marked by supported familialism or familialism by default, where families act as the key providers due to respectively ad hoc policy design or absence of family policies altogether for what concerns parental leaves, childcare coverage and child-related income transfers. Applied to the countries under examination, their perspective indeed pinpoints some cross-national variation. Czech Republic and Estonia, for instance, are among the countries with the highest degree of supported familialism through the instruments of leaves and transfers. Austria and Germany occupy a middle position in all dimensions, displaying a more moderate supported familialism profile despite generous cash transfers. By contrast, the Netherlands is often classified as familialistic by default despite extensive defamilialization in childcare services. This way to frame the institutional diversity provides a valuable reference to examine whether private financial transfers from parents act as compensatory mechanisms.

In spite of these theoretical contributions, studies on private financial transfers and fertility intentions are meager. Past research has primarily examined the effects of public policy provision, employment, and childcare on fertility behavior (Matysiak, 2011; Thévenon & Gauthier, 2011). The role of family support has received comparatively less attention. There is a growing strand of research that has examined how intergenerational resources interact with life-course transitions, related to housing acquisition and residential independence, supporting the idea that family support can ease early-life economic constraints. (Dräger et al., 2024; Lersch & Luijkx, 2015; Mulder & Smits, 2013). Other studies point to a persistent and even growing association between income and fertility

in high-income contexts, as shown in Swedish register data where children increasingly resemble a “normal good” primarily accessible to higher-income individuals (Kolk, 2023). This trend concerned especially Southern and Central Europe in the aftermath of the Great Recession of 2008, as economic instability deprived young adults from the safety nets they typically relied upon (Matysiak et al., 2020).

However, few studies have explicitly linked private financial support to fertility intentions, as opposed to fertility outcomes or other life transitions. One comparative analysis using earlier GGS data examined the influence of informal networks on second-birth intentions in Italy, Bulgaria, and Germany (Di Giulio et al., 2012). The findings were mixed and detected cross-national variation in reliance on kin support. Still, that work did not directly assess financial transfers, nor did it incorporate recent or harmonized measures of timing and parity. To date, no study has systematically examined the relationship between private financial transfers and short-term fertility intentions using harmonized, cross-national data across multiple survey rounds. This paper addressed the data availability gap by leveraging both rounds of the GGS to assess whether parental financial support is associated with fertility planning, and how this association varies by parity, timing of transfers, and national context. The inclusion of two rounds of data can moreover provide some preliminary insight into how private support may have shifted over time, particularly in the period following the 2008 financial crisis and preceding the COVID-19 pandemic.

Drawing on the frameworks outlined above, we formulate the following hypotheses:

- H1.** Receiving a private financial transfer from parents (or ascendants) is positively associated with the intention to have a child in the near term.
- H2.** The association between parental financial transfers and fertility intentions differs by parity, with stronger associations expected among childless individuals.
- H3.** The association is stronger when transfers are large and received closer to the survey year, and is more pronounced in the more recent survey period.
- H4.** The association between parental financial transfers and fertility intentions varies across national contexts, reflecting differences in welfare regimes and type or degree of familialism.

4 Data and analytical strategy

This study is based on data extracted from the Generations and Gender Survey (GGS), a collection of harmonized and longitudinal surveys now covering more than 30 countries (not limited to Europe) and designed with the main aim of studying demographic behavior and family dynamics in a comparative perspective. Since its launch in the early 2000s, the GGS has become a key and widely used data repository in demographic research to investigate themes such as fertility, partnership dynamics, and intergenerational ties (Gauthier, 2025). Owing to its broad coverage and an harmonized conceptual framework, it constitutes a thorough data source for research questions that seek to link life-course decisions to embedded family or societal structures. One key distinctive strength on which we build is the priority attached to the reconstruction of ties between adult respondents and their ascendants, including financial support, frequency of contacts or co-residence and parental attributes. All these features render the GGS particularly valuable for analyzing how private family resources intersect with early-life transitions such as fertility planning, central to this study. What sets the GGS apart from other datasets carrying household-level indicators is then its privileged direct observation of intergenerational exchanges together with a consistent focus on fertility intentions as a recurrent item across the national surveys. In spite of some limitations explained in the remainder, this distinctive design that combines information on individuals and their family relations across different institutional contexts provides a solid foundation for addressing the increasingly relevant question on how private transfers may have contributed to shaping reproductive plans over the last two decades.

The Generations and Gender Programme (GGP), the research infrastructure underlying the survey, has conducted two major rounds of data collection. The first round (GGS-I), fielded between 2004 and 2011, targeted over 200,000 individuals aged 18 to 79, and provides rich longitudinal datasets for most of its 19 participating countries. Within this round, the survey follows a closed-panel design, meaning that respondents are observed repeatedly over time, but there is no refreshment of the sample through the entry of new individuals. The second round (GGS-II) was launched in 2020 and aims to extend the analytical potential of the survey by offering to date predominantly one wave per country, with some updated instruments to better capture emerging family arrangements and demographic changes. It is to be noted that the two rounds are fully independent samples.

For the purposes of this study, the analysis focuses on five countries—namely Austria, Germany, the Netherlands, Estonia, and the Czech Republic—that are covered by both GGS rounds. The eligible sample consists of individuals who were at most 45 years old at the time of the (first) interview,

in order to restrict the focus to those who are most likely to be within a reproductive time window. Short-term fertility intentions, the outcome variable, are collected in a comparable fashion across the two rounds, as respondents falling within the eligible sample are asked in the fertility module whether they intend to have a (another) child within the next three years. The key explanatory variable - receipt of a private financial transfer from parents (and, in the Appendix, from ascendants more broadly) — is not directly comparable across the two rounds of the Generations and Gender Survey. Differences encompass the reference period and the framing of transfers; Round I focuses on recent financial support (12 months), whereas Round II rather captures only large inter vivos transfers over the life course. In an attempt to better harmonize the two, we adopt a stepwise strategy (Table 1) that respects the original survey design while highlighting the point of convergence in the construction of the transfer receipt variables.

Table 1: Conceptual scheme of parental transfers variables across GGS rounds

Round I (GGS-I, 2002–2013)	Round II (GGS-II, 2020–2023)
Any transfer in last 12 months	Ever received large transfer (€5,000)
↓	↓
Identify provider (parents R or partner’s*)	Identify provider (parents R or partner’s*)
↓	↓
Exclude inheritances when possible	Use year of last transfer (same year or preceding year)
↓	↓
<i>Output 1: Any parental transfer (Round I only)</i>	<i>Output 1: Ever large parental transfer (Round II only)</i>
↓	↓
Classify by size (€500, baseline NL**/ €5,000)	—
↓	↓
Output 2: Recent large parental transfer	Output 2: Recent large parental transfer

Note: Only Output 2 is conceptually comparable across GGS rounds. *In Estonia (Round 1), transfers received in the last 12 months are observed together with their amount, but not the identity of the provider (whenever it exists, the transfer is a proxy for parental support). **In the Netherlands, grandparents do not figure in the list of potential providers, and the amount is recorded as above or below 500 euro.

In Round I (2002–2013), respondents were asked whether they had received any financial transfer in the 12 months prior the interview. Follow-up questions helped identifying the provider and in most cases foregoing inheritances. This structure enables the construction of two variables: one broad measure of any parental transfer and a more restrictive one on large parental transfers. At baseline, the size threshold is anchored to the Dutch definition of a large transfer (€500), while an alternative threshold equivalent to €5,000—the reference amount in Round II—is computed for Austria and Germany. In general, to ensure both cross-national and overtime comparability all monetary amounts are expressed in real euro. Nominal values are deflated using country-specific Harmonised Index of Consumer Prices (HICP) indices, sourced from Eurostat (2026). For each country, the HICP entry corresponds to the survey month and year most relevant for Round II interviews, thus reflect prices in effect when large transfers were observed in the latest round. For Czech Republic and Estonia which

had not yet adopted the euro during the first Round, threshold amounts were first converted from the national currency into euros using the official exchange rate for the survey year and subsequently deflated. This way of proceeding ensures that the meaning attached to large parental transfers indeed reflects a comparable purchasing power across both context and timing.

On the other hand, Round II (2020–2023) asked whether responders had ever received a large financial transfer in the lifetime and the year of occurrence of the most recent transfer. This design hinders the identification of any transfers and those of a small amount, yet the question on timing allows to reconstruct a measure of recent large transfers (i.e. restricting to year of interview or the preceding one). Only this latter construct can be considered as conceptually similar across the two rounds and therefore represents the most relevant explanatory variable in the main analyses.

The analytical strategy is organized sequentially by survey round. A first empirical section makes use of data from GGS Round I by exploiting the longitudinal structure available for most countries. We then turn to GGS Round II, whose more recent but cross-sectional design calls for a slightly different approach. Given the highlighted differences between the two rounds, at no point the relevant datasets are pooled together. Each round is presented and analyzed independently, and pooled models are estimated across the five countries within the given round. Across the next two sections, the analysis follows three common stages. Descriptive statistics are presented to introduce the analytical samples, document the cross-country variation of parental transfers and fertility intentions, among other characteristics, and panel attrition for Round I. Second, we estimate country-specific regression models with a gradual approach, and study the associations between parental transfers and fertility intentions net of an expanding set (4) of individual- and family-related controls. Lastly, we pool all countries within each round to estimate average marginal effects and explore cross-national patterns and a potential institutional gradient in the link between transfers and fertility intentions, using some country-level indicators from the OECD (2026) and the World Bank (2026).

On top of this overarching structure, the analysis is complemented by round-specific choices that reflect the discussed divergences in survey design. It follows that the availability of repeated observations from Round I can be better exploited through random-effect models with Mundlak term to account for unobserved time-invariant heterogeneity that may jointly influence the receipt of parental transfers and fertility intentions; In Round II, instead, as the available data are to date mostly cross-sectional, we employ propensity score matching to mitigate selection into parental transfer receipt by balancing treated and untreated individuals on a rich set of observed characteristics.

Despite the methodological differences, the two approaches are motivated by a similar identification concern on a non-random allocation of family financial support. This mix of techniques within an otherwise common and sequential approach is meant to extract the most out of each dataset, while aiming to preserve a broad conceptual comparability across rounds. Accordingly, Section 5 presents the results from GGS Round I, beginning with descriptive statistics and attrition patterns and proceeding to country-specific and pooled analyses. Section 6 reports the corresponding results for GGS Round II and incorporates propensity score matching. The final section summarizes the main findings and offers concluding remarks.

5 Analysis - Round I

Tables 2-4 report the descriptive statistics and attrition patterns related to Round I. Only the Netherlands present a complete round with three waves at three-years intervals. The Czech Republic, Austria, and Germany are observed for two waves each, and Estonia contributes with only a single wave. There is some evident substantial cross-national variety both in the overall sample composition and in the participation over time. Attrition between wave 1 and 2 is exceptionally elevated in the Czech Republic and Germany (Table 4), where more than 70% of the eligible respondents drop out from one wave to another. By contrast, Austria presents quite the opposite scenario, with an attrition rate of about 22%. The Netherlands also present a comparable level of attrition (almost 26% from the first wave to the second and slightly less between the second to last). Looking at the baseline socio-demographic characteristics, there are some key differences as well. Respondents in Germany and the Netherlands tend to be on average older than in the Czech Republic. At the same time, Germany and the Netherlands present also relatively smaller households. Germany and Estonia are the only two countries with about a third of childless respondents, whereas this figure rises significantly (above 40%) in the Czech Republic, the Netherlands, and Austria. The prevalence of respondents who intend to have a child is not markedly different, and, at baseline, ranges between around 20% in Germany to 26% in the Czech Republic, partially mirroring the patterns highlighted for childlessness. Moreover, childlessness is much more common among attriters than among stayers in all countries, pointing to some degree of selection in attrition along family trajectories. Similarly, in the Netherlands higher attrition affects non-owners, unmarried, and less economically stable respondents. Education-wise, in Estonia, the Netherlands, and Germany there is a greater share of highly educated respondents. Some notable cross-country differences are also observed for the key explanatory variable. The Netherlands shows the highest prevalence of parental support, with roughly one-third of respondents reporting any transfer and around one-fifth receiving large transfers (above 500 €). By contrast, the Czech

sample presents parental transfers as a rare phenomenon, with fewer than 3% of respondents reporting any parental transfer and large transfers being virtually absent. Transfers are relatively more common in Estonia and Austria, but only in the latter do large transfers above 500 € and above the high threshold (equivalent to 5,000 € in Round I) reach non-negligible shares of about 5% and 2%, respectively. Germany, in comparison, displays similar but slightly lower levels. It is worth noting that only in the Netherlands do recipients of parental transfers tend to be somewhat over-represented among stayers, a pattern that may underlie a modest over-representation of respondents with greater family resources in the overall panel. Despite these limitations— particularly high attrition and the low prevalence of parental transfers in the Czech Republic—the data remain informative for the questions at hand. Attrition follows usual life-course patterns, while reported fertility intentions differ only modestly between stayers and attriters.

Table 2: Netherlands (Round I) – Attrition patterns and sample characteristics

Panel A. Attrition - summary		
Eligible respondents at Wave 1 (age ≤ 45)	4141	
Stayed from Wave 1 to Wave 2	3073	
Attrited from Wave 1 to Wave 2	1068	25.79%
Stayed from Wave 2 to Wave 3	2246	
Attrited from Wave 2 to Wave 3	827	19.97%
Panel B. Descriptives by attrition status		
	<i>Wave 1 $\rightarrow$ Wave 2</i>	
	Stayed	Attrited
Age (mean)	34.52	32.49
Household size (mean)	3.01	2.65
Female (%)	62.19	54.31
Education: Low (%)	21.64	32.67
Education: Medium (%)	41.00	38.70
Education: High (%)	37.36	28.63
Married (%)	51.07	35.58
Childless (%)	41.78	53.75
Employed (%)	73.82	69.87
Inactive (%)	22.53	23.63
Homeowner (%)	66.87	46.25
Rented housing (%)	33.13	53.75
Intends a child within 3 years (%)	22.27	22.25
Received any parental transfer (%)	31.76	27.53
Received large parental transfer > 500€ (%)	21.64	17.51
Number of siblings (mean)	2.53	2.40
Father high education (%)	24.18	22.53
Mother high education (%)	10.39	10.88
Parents married (%)	85.01	79.52
	<i>Wave 2 $\rightarrow$ Wave 3</i>	
	Stayed	Attrited
Age (mean)	38.49	37.31
Household size (mean)	3.27	3.12
Female (%)	63.85	57.32
Education: Low (%)	16.56	23.58
Education: Medium (%)	39.36	41.48
Education: High (%)	44.08	34.95
Married (%)	59.13	49.15
Childless (%)	30.77	36.28
Employed (%)	81.30	77.87
Inactive (%)	16.42	17.41
Homeowner (%)	77.70	64.70
Rented housing (%)	22.30	35.31
Intends a child within 3 years (%)	19.78	17.73
Received any parental transfer (%)	33.79	29.99
Received large parental transfer > 500 €(%)	20.44	17.78
Number of siblings (mean)	2.39	2.41
Father high education (%)	24.53	23.19
Mother high education (%)	10.78	9.26
Parents married (%)	86.58	80.73

Notes: Percentages are shown in points (0–100). Means are shown for continuous variables.

Table 3: Estonia (Round I) – Sample characteristics at Wave 1

Panel A. Sample size	
Eligible respondents at Wave 1 (age ≤ 45)	3,602
Panel B. Descriptives	
Age (mean)	33.06
Household size (mean)	3.31
Female (%)	62.08
Education: Low (%)	12.27
Education: Medium (%)	57.86
Education: High (%)	29.87
Married (%)	43.06
Childless (%)	30.18
Employed (%)	76.32
Inactive (%)	18.10
Homeowner (%)	84.20
Rented housing (%)	9.69
Intends a child within 3 years (%)	25.22
Received any parental transfer (%)	9.58
Received large parental transfer > 500 (%)	0.56
Number of siblings (mean)	1.53
Father high education (%)	16.51
Mother high education (%)	15.75
Parents married (%)	66.64

Notes: Percentages are shown in points (0–100). Means are reported for continuous variables. Estonia is observed in a single wave and no attrition analysis is viable.

Table 4: Countries with 2 waves (Round I) –Attrition patterns and sample characteristics

	Czech Republic		Austria		Germany	
	Stayed	Attrited	Stayed	Attrited	Stayed	Attrited
Panel A. Attrition – summary						
Eligible respondents at Wave 1 (age ≤ 45)	5,289		5,000		4,883	
Stayed from Wave 1 to Wave 2	1,533		3,912		1,432	
Attrited from Wave 1 to Wave 2		71.02%		21.76%		70.67%
Panel B. Descriptives by attrition status						
Age (mean)	31.46	30.43	32.89	32.14	35.27	32.37
Household size (mean)	3.28	3.10	3.25	3.02	3.12	2.73
Female (%)	50.82	50.56	59.30	62.59	61.03	53.78
Education: Low (%)	17.98	21.48	11.73	16.64	9.05	14.49
Education: Medium (%)	68.87	65.22	70.53	66.27	61.11	64.57
Education: High (%)	13.15	13.31	17.74	17.10	29.84	20.93
Married (%)	44.44	39.66	46.31	34.54	58.16	45.03
Childless (%)	43.12	51.54	42.46	47.15	31.15	47.29
Employed (%)	63.41	63.29	73.93	71.32	64.87	61.17
Inactive (%)	28.25	28.34	21.96	22.43	26.89	28.46
Homeowner (%)	52.51	47.84	56.69	47.19	47.49	30.71
Rented housing (%)	26.56	33.98	36.45	46.83	51.19	67.17
Good financial situation (%)	35.37	38.30	74.57	70.47	62.54	58.99
Intends a child within 3 years (%)	26.18	26.57	28.79	30.91	20.45	19.30
Received any parental transfer (%)	2.15	3.17	7.39	9.47	4.47	4.64
Received large parental transfer > 500 (%)	0.13	0.13	5.16	5.88	3.14	2.61
Received large parental transfer > 5,000 (%)			2.61	2.48	1.33	0.93
Received any transfer from ascendants (%)	2.48	3.49	8.79	10.94	5.10	5.13
Received large transfer from ascendants (%)	0.20	0.13	5.93	6.62	3.63	2.96
Number of siblings (mean)	1.35	1.30	2.23	2.22	1.94	1.75
Father high education (%)	12.68	12.31	9.25	9.39	28.40	23.95
Mother high education (%)	8.03	7.17	6.47	7.06	10.67	10.62
Parents married (%)	58.38	57.02	77.56	71.80	77.97	79.99

Notes: Percentages are shown in points (0–100). Means are reported for continuous variables.

Table 5 reports country-specific logistic regressions estimating the association between parental transfer receipt and short-term fertility intentions in Round I. Model 1 includes parental transfers and wave fixed effects, while Model 2 adds some basic demographic controls (sex, age, marital status, and childlessness). All models incorporate country-specific weights and regional stratification, reflecting the GGS complex survey design. In the baseline specification (Model 1), receipt of a parental transfer is positively and statistically significantly associated with fertility intentions in four out of five countries. The estimated odds ratios range from 1.39 in the Netherlands to 1.62 in Germany, suggesting substantially higher levels of desired fertility among the transfer recipients. Estonia represents a partial exception: in the single available wave, its point estimate is positive ($OR \approx 1.21$), but not statistically significant at conventional levels.

Once the demographic controls are included (Model 2), the association between parental support and fertility intentions is considerably diminished and is no longer statistically significant in all the countries. This would indicate the fact that the initial association estimated by Model 1 could be explained by the differences in the demographic composition of the recipients and the non-recipients of support (age, marital status, presence of children). Across the economies, demographic predictors display highly consistent patterns. Age is negatively associated with fertility intentions in all samples, while being childless strongly increases the likelihood of intending a child within the next three years. Another variable that is positively related to fertility intentions in most countries is marital status, with large effects in Germany and in the Netherlands. Gender differences are generally small and statistically insignificant; in Estonia, fertility intentions are reported only for women, and the coefficient for female is therefore omitted due to collinearity. In sum, the information that can be gained from Table 5 is that the correlation between parental transfers and fertility intentions is not in the end robust once controlling for the key demographic variables.

Table 6 extends the analysis by sequentially introducing socio-economic characteristics (Model 3) and parental background variables (Model 4). This progression allows us to assess whether the association between parental transfers and fertility intentions persists once differences in education, labor market attachment, housing conditions, financial situation, and family background are taken into account. Mirroring what found in Model 2, the estimated relationship between receiving parental transfers and fertility intentions is statistically insignificant for all countries (Model 3). While point estimates remain above one in many cases – and especially in Austria and Germany – the inclusion of socio-economic controls completely absorbs the residual association present under the reduced specification. This pattern remains robust after the inclusion of parental background variables in

Model 4, indicating that the association between receipt of a transfer and fertility intentions at Round I could be primarily driven by observable individual and familial characteristics, rather than by a potential independent effect of the transfer receipt.

In contrast, several demographic and socio-economic predictors display strong and consistent associations with fertility intentions across countries. Age still strongly and negatively impacts fertility intentions in all models, and childlessness is the most consistent predictor of fertility intentions, especially in Austria, Germany and Estonia. There is a strong positive gradient for educational level: high-educated people are much more likely to report fertility intentions, especially in the Czech Republic and Germany. Labor market status is also important, but relevant thresholds differ across countries. In the Netherlands and the Czech Republic, inactivity is strongly and negatively associated with fertility intentions while in Austria a reversed pattern emerges, indicating that what not being employed stands for may also differ between countries. Housing conditions and subjective financial situation further contribute to explaining fertility intentions, particularly in Austria and Germany, where renting or living in provided housing is positively associated with intended fertility even after conditioning for age. These results align with prior expected relevance of housing arrangements as key correlates of short-term fertility planning.

The inclusion of parental background variables in Model 4 — namely parental education, parental marital status, and the number of siblings — does not substantially alter the main conclusions. While the size of the coefficients remains roughly preserved, above unity for all countries, the attenuation effect in significance already observed in earlier models cannot be attributed to the omission of family background characteristics. Moreover, these individual indicators of parental background do almost never carry an independent explanatory effect. There are a few notable exceptions to this general remark. In the Czech Republic, respondents whose mothers attain a higher education are 40 to 55% less likely to report fertility intentions, all else equal. Conversely, in Germany and Austria, individuals whose parents were married when they were 15 years old exhibit a higher propensity to have fertility plans, with the odds ratios indicating increases of about 33% and 41%, respectively. Across all countries, the likelihood of reporting fertility intentions is systematically higher in the second survey wave, with odds ratios ranging from a minimum of 1.13 in Austria to 1.74 in Germany. The time effect for the Netherlands in the third wave instead does not reach significance. Consistent with earlier models, the most significant predictors for higher fertility intentions remain having a tertiary education and not yet being a parent. Hence, even when considering more complete models, receiving parental transfers of any kind during Round I does not seem to exert independent effects on short-term fertility intentions among European respondents. Instead, the initially observed association

between transfers and fertility intentions appears to reflect selection. From the evidence collected in the first four models, transfers may be correlated with fertility intentions primarily because they are commonly received by individuals who are younger, childless, and more highly educated but not necessarily more advantaged in socio-economic terms.

The overall null result might, however, conceal some heterogeneity by parity status. When, in fact, we allow the effect of parental transfers to differ between childless respondents and parents, a less uniform pattern emerges. Across the examined countries, fertility intentions are consistently higher among the childless, yet the association between parental transfers and fertility intentions is not necessarily stronger for respondents planning a first birth. Transfers, in particular, tend to be more strongly associated with fertility intentions among respondents who already have children. This is particularly evident for Germany, where parental transfers are strongly associated with fertility intentions overall, but the interaction between transfers and childlessness is negative ($OR = 0.21$) and strongly significant. This reflects a marked association between fertility intentions and parental transfers among parents, and a negligible one among the childless. Similar dynamics are also observed in the Netherlands and in Austria, with respectively 70% higher and nearly double the odds of intending a child among parents. In Estonia and the Czech Republic, instead, parental transfers remain unrelated to fertility intentions regardless of parity status. Hence, intergenerational financial support of the early 2000s appears to primarily accompany family expansion decisions rather than transitions to parenthood restrictively to Western European contexts.

Table 5: Country-specific pooled logit models of fertility intentions - Round I (Models 1-2)

Model	Variable	Netherlands	Estonia	Czech Rep.	Austria	Germany
Model 1: Baseline						
	Transfer received	1.39*** (0.11)	1.21 (0.21)	1.51** (0.27)	1.61*** (0.27)	1.62*** (0.30)
	Wave FE	✓		✓	✓	✓
	Observations	7,004	2,002	4,779	7,585	4,080
	Respondents	3,954	2,002	4,093	4,701	3,481
Model 2: + Demographic controls						
	Transfer received	1.04 (0.09)	1.17 (0.21)	1.16 (0.22)	1.31 (0.24)	1.25 (0.26)
	Female	1.17* (0.10)	—	0.91 (0.07)	0.94 (0.10)	0.83* (0.09)
	Age	0.92*** (0.01)	0.92*** (0.01)	0.94*** (0.01)	0.95*** (0.01)	0.93*** (0.01)
	Married	1.62*** (0.19)	0.95 (0.13)	1.53*** (0.19)	0.93 (0.13)	2.72*** (0.51)
	Childless	1.40*** (0.17)	2.75*** (0.40)	1.86*** (0.25)	2.24*** (0.33)	2.81*** (0.48)
	Wave FE	✓		✓	✓	✓
	Observations	7,002	1,998	4,703	7,579	4,047
	Respondents	3,953	1,998	4,029	4,696	3,463

Note: Odds ratios (with standard errors in parentheses) from country-specific logistic regressions with individual-level data pooled across survey waves from Round I of the GGS. Model 2 additionally controls for sex, age, childlessness and marital status. For Estonia, fertility intentions are reported only for women and the coefficient for female is omitted due to collinearity.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 6: Country-specific pooled logit models of fertility intentions - Round I (Models 3-4)

Model	Variable	Netherlands	Estonia	Czech Rep.	Austria	Germany
Model 3: + Socio-economic controls						
	Transfer received	1.05 (0.09)	1.07 (0.20)	1.14 (0.24)	1.25 (0.23)	1.28 (0.29)
	Female	1.27*** (0.10)	—	1.02 (0.08)	0.84* (0.09)	0.92 (0.10)
	Childless	1.13 (0.14)	2.17*** (0.33)	1.62*** (0.23)	2.15*** (0.31)	2.61*** (0.45)
	Age	0.89*** (0.01)	0.91*** (0.01)	0.90*** (0.01)	0.95*** (0.01)	0.89*** (0.01)
	Married	1.61*** (0.19)	0.93 (0.13)	1.54*** (0.20)	0.92 (0.13)	2.71*** (0.52)
	High education (vs low)	1.85*** (0.21)	1.57** (0.36)	4.14*** (0.74)	1.88*** (0.41)	2.82*** (0.62)
	Inactive (vs emp.)	0.26*** (0.04)	0.78* (0.11)	0.40*** (0.05)	1.48*** (0.22)	0.76* (0.11)
	Unemployed (vs emp.)	1.01 (0.22)	0.58* (0.18)	0.85 (0.13)	1.27 (0.23)	0.68* (0.15)
	Rented housing (vs ownership)	0.99 (0.10)	1.18 (0.21)	1.22** (0.11)	1.28** (0.14)	1.63*** (0.23)
	Provided housing (vs ownership)	—	0.94 (0.21)	1.10 (0.12)	1.74*** (0.33)	0.94 (0.39)
	Good financial situation (vs bad)	—	—	0.99 (0.18)	1.29* (0.14)	1.17
	Wave FE	✓		✓	✓	✓
	Observations	6,997	1,998	4,329	7,547	3,677
	Respondents	3,949	1,998	3,780	4,690	3,156
Model 4: + Parental background controls						
	Transfer received	1.07 (0.11)	1.09 (0.21)	1.02 (0.55)	1.36 (0.26)	1.36 (0.36)
	Female	1.31*** (0.11)	—	1.20 (0.17)	0.78** (0.09)	0.96 (0.13)
	Childless	1.28* (0.17)	2.40*** (0.38)	1.22 (0.35)	2.06*** (0.32)	3.19*** (0.66)
	Age	0.88*** (0.01)	0.92*** (0.01)	0.91*** (0.01)	0.95*** (0.01)	0.90*** (0.01)
	Married	1.79*** (0.24)	0.91 (0.14)	0.92 (0.24)	0.88 (0.13)	2.91*** (0.73)
	High Education (vs low)	2.17*** (0.29)	1.90** (0.53)	6.73*** (1.93)	1.62* (0.41)	3.30*** (1.04)
	Inactive (vs emp.)	0.23*** (0.03)	0.85 (0.13)	0.38*** (0.08)	1.43** (0.23)	0.79 (0.13)
	Unemployed (vs emp.)	1.16 (0.30)	0.56 (0.21)	0.94 (0.24)	1.30 (0.25)	0.77 (0.23)
	Rented housing (vs ownership)	0.94 (0.10)	1.16 (0.22)	1.22 (0.20)	1.39*** (0.17)	1.63*** (0.27)
	Provided housing (vs ownership)	—	0.98 (0.23)	0.90 (0.18)	1.82*** (0.37)	0.87 (0.41)
	Good financial situation (vs bad)	—	—	1.13 (0.16)	1.26 (0.19)	1.12 (0.16)
	Wave FE	✓		✓	✓	✓
	Parental Background	✓	✓	✓	✓	✓
	Observations	5,974	1,644	1,609	6,524	2,278
	Respondents/Clusters	3,332	1,644	1,443	4,035	1,945

Note: Odds ratios (with standard errors in parentheses) from country-specific logistic regressions with individual-level data pooled across survey waves from Round I of the GGS. Model 3 additionally controls for education, activity status, housing, and financial condition (making ends meet). Model 4 further includes parental background characteristics (education of mother and father, marital status of parents and number of siblings) - not reported for brevity. For Estonia, fertility intentions are reported only for women and the coefficient for female is omitted due to collinearity.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 7 shifts the focus on whether larger parental transfers (exceeding 500 euro) are associated with short-term fertility intentions in Round I. This threshold reflects the structure of the transfer amount question asked in the Dutch survey. For Estonia, the equivalent cutoff in its unique wave is set at 7,900 EEK, corresponding to the same real value. The Czech Republic is excluded from this analysis, as almost no such large parental transfers were reported for this country during the first Round. In the remaining contexts, large transfers are more strongly associated with fertility intentions than transfers of any amount, although the evidence remains heterogeneous. In Western Europe, such large transfers display a slightly more robust positive association with fertility intentions, as showed by the marginal significance reached in the Netherlands and Germany. This would suggest that the size of the financial support matters more than the transfers receipt per se, and particularly so in those countries where the support was likely to be more relevant for family expansion. If we were to include the interaction factor with the childlessness indicator, large transfers would be again more strongly related to fertility intentions among respondents who already have children, particularly in Germany. Austria and the Netherlands exhibit weaker but consistent findings, whereas in Estonia results are imprecisely estimated due also to the very limited amount of large parental transfers, and no parity-specific patterns can be identified.

Table 7: Country-specific pooled logit models of fertility intentions - Round I (Models 5: large transfers)

Model	Variable	Netherlands	Estonia	Czech Rep.	Austria	Germany
Model 5: Large parental transfers (> 500€)						
	Large transfer received	1.19* (0.12)	1.05 (0.97)	—	1.29 (0.28)	1.63* (0.48)
	Female	1.31*** (0.11)	—	1.19 (0.17)	0.79** (0.09)	0.96 (0.13)
	Childless	1.28* (0.17)	2.4 *** (0.38)	1.22 (0.35)	2.06*** (0.32)	3.17*** (0.66)
	Age	0.88*** (0.01)	0.92*** (0.01)	0.91*** (0.01)	0.95*** (0.01)	0.90*** (0.01)
	Married	1.78*** (0.24)	0.91 (0.14)	0.92 (0.24)	0.87 (0.13)	2.89*** (0.73)
	High education (vs low)	2.14*** (0.29)	1.91** (0.53)	6.71*** (1.93)	1.63* (0.42)	3.29*** (1.04)
	Inactive (vs emp.)	0.23*** (0.03)	0.85 (0.13)	0.38*** (0.08)	1.43** (0.23)	0.79 (0.13)
	Unemployed (vs emp.)	1.15 (0.29)	0.56 (0.21)	0.94 (0.24)	1.30 (0.25)	0.78 (0.24)
	Rented housing (vs ownership)	0.94 (0.11)	1.17 (0.22)	1.22 (0.20)	1.40*** (0.17)	1.62*** (0.27)
	Provided housing (vs ownership)	—	0.86 (0.92)	0.98 (0.23)	1.81*** (0.37)	0.86 (0.41)
	Good financial situation (vs bad)	—	—	1.13 (0.16)	1.25 (0.19)	1.11 (0.16)
	Wave FE	✓		✓	✓	✓
	Parental Background	✓	✓	✓	✓	✓
	Observations	5,974	1,644	1,608	6,524	2,278
	Respondents	3,332	1,644	1,442	4,035	1,945

Note: Odds ratios (with standard errors in parentheses) from country-specific logistic regressions with individual-level data pooled across survey waves from Round I of the GGS. Large transfers are defined as parental transfers exceeding 500€ (using question wording in the Netherlands as reference). The model includes demographic, socio-economic, parental background controls and wave fixed effects. In the Czech Republic, large transfers perfectly predict the outcome and the corresponding coefficient is therefore omitted. For Estonia, fertility intentions are reported only for women and the coefficient for female is omitted.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

As a complementary approach, we re-estimate the main full specification for large transfers using a random-effects logistic model, thereby better exploiting the longitudinal structure of the Round I data. This other strategy treats more explicitly repeated observations on the same individuals, and at the same time it allows to control for unobserved time-invariant characteristics at the individual level. The fixed-effects logit specification would not be appropriate to the case at hand, since it discards all the individuals with no transition in the dependent variable as well as forgoing the estimation of time-invariant covariates. The random-effects framework, instead, retains both within- and between-individual variation and is overall more suitable due to the limited number of waves and the inclusion of several individual- and family-level characteristics in the main model. In order to relax the assumption of strict exogeneity that underlies standard random-effects models, we adopt a random-effects specification following Mundlak (1978). By including the individual-specific mean of the time-varying parental transfer indicator across waves, we explicitly account for differences between individuals in their likelihood of receiving parental support. This is meant to accommodate the potential correlation between transfer receipt and the unobserved individual traits, for instance family orientation from the discussed sociological perspective, which may jointly influence the propensity to receive transfers as well as fertility intentions (Wooldridge, 2010). The Mundlak term is only applied to parental transfers as the primary variable of interest and plausibly correlated with other unobserved individual characteristics.

Overall, the evidence presented from the countries with more than one wave hints to the conclusion that large parental transfers have no independent effect on the fertility intentions, once we properly control for individual heterogeneity. Taking the Netherlands, contrarily to the earlier pooled models, neither the within-person effect of receiving a large transfer in a given wave nor the between-person effect captured by the mean of transfer receipt is statistically significant. For Germany, similarly, once the individual-specific heterogeneity is introduced via random effects, the previously observed association weakens and loses the statistical significance. This attenuation in significance but not in magnitude is likely to reflect the limited within-individual variation in transfer receipt rather than a real reversal of the underlying association. Parental transfers may thus be concurrent with fertility planning, accompanying family-related decisions rather than actively driving them. For Austria, the coefficient for large parental transfers remains almost unchanged in the panel specification and continues to be statistically insignificant. Likewise, the Mundlak term carries no explanatory power. Fertility intentions are instead strongly structured by educational attainment, housing conditions, labor market status, and selected demographic characteristics. Finally, for the Czech Republic, the analysis confirms that the effect of large parental transfers cannot be meaningfully estimated, as the

lack of within-individual variation leads to perfect prediction. The estimated effects for individual-specific means are imprecise and lack a clear pattern across countries: people who are more exposed to large transfers show no systematic difference in fertility plans from people who receive none. Alongside the weakening of transfer effects, this result would imply also that the associations seen in the former country-specific pooled models stem mainly from compositional effects or stage in the personal life. However, in countries such as Germany, where attrition between waves is substantial, the loss of weak statistical significance may also result from reduced power (and some selective panel retention common to all countries), rather than being revelatory of a genuine absence of any association.

Table 8: Random-effects logistic regression predicting fertility intentions – Round I (large transfers)

Variable	Netherlands	Czech Republic	Austria	Germany
<i>Transfers (within + between via Mundlak)</i>				
Large parental transfer (> 500 €)	1.05 (0.15)	–	1.23 (0.30)	3.66 (3.27)
Mean large transfer (Mundlak term)	1.16 (0.33)	–	0.75 (0.26)	0.44 (0.44)
<i>Demographics</i>				
Female	1.07 (0.10)	1.30 (0.31)	0.77*** (0.08)	0.87 (0.14)
Childless	1.38*** (0.17)	1.31 (0.50)	2.15*** (0.26)	3.51*** (0.82)
Age	0.84*** (0.01)	0.86*** (0.03)	0.93*** (0.01)	0.85*** (0.02)
Married	1.49*** (0.17)	0.80 (0.28)	0.95 (0.11)	3.78*** (0.90)
<i>Socio-economic status</i>				
High education (vs low)	2.01*** (0.29)	21.10*** (13.33)	4.04*** (0.80)	5.55*** (2.00)
Unemployed (vs employed)	1.50* (0.32)	0.76 (0.32)	0.75* (0.13)	0.72 (0.21)
Inactive (vs employed)	0.22*** (0.03)	0.18*** (0.07)	0.73*** (0.08)	0.68** (0.13)
Rented housing (vs ownership)	0.81** (0.08)	1.42 (0.39)	1.79*** (0.18)	1.90*** (0.34)
Provided housing (vs ownership)	0.81** (0.08)	1.42 (0.29)	1.68*** (0.18)	0.76 (0.39)
Good financial situation (vs bad)			1.12 (0.12)	1.03 (0.17)
Parental background	✓	✓	✓	✓
Observations	5,974	1,608	6,524	2,278
Respondents	3,332	1,442	4,035	1,945
ρ	0.23	0.63	0.47	0.38

Note: Odds ratios (standard errors in parentheses) from random-effects logistic regressions with correlated random effects (Mundlak specification). Models include wave fixed effects (omitted for brevity) and parental background controls. In the Czech Republic, large transfers perfectly predict the outcome; the corresponding coefficient and the Mundlak term are therefore omitted.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

We now move to an analysis that focuses on those large parental transfers which are intended to approximate the idea underlying a substantial financial support used in Round II, with a threshold set at €5,000. For the two economies where similar relevant transfers are recorded, we construct the country and wave-specific thresholds by adjusting the Round II reference prices of €5,000 using Harmonised Index of Consumer Prices (HICP) values. For Germany, the resulting cut-offs correspond

to €3,900 in Round I and €4,200 in Round II, while for Austria they are set to €3,600 in Round I and to €3,900 in Round II. These adjusted thresholds enable us to identify large transfers in a more consistent way over time as well as to more closely examine whether the sizable intergenerational financial support is differentially associated with the fertility intentions, once we also capture unobserved individual heterogeneity. Table 9 provides short summary estimates for full random-effects logistic regressions with Mundlak correction. In both countries, large transfers do not exhibit a stable and statistically significant association with fertility intentions in the random-effects specification. In Germany, the main coefficient for large transfers is not statistically significant and small, highlighting that among parents receiving a large transfer is not associated with higher or lower fertility intentions in a systematic way when both observed and unobserved individual characteristics are taken into account. At the same time, the interaction between large transfers and childlessness is negative and significant ($OR \approx 0.06$), indicating that, if anything, for childless respondents receiving large transfers may even be associated with lower odds in fertility intentions relative to parents, once all individual differences are taken care of. Furthermore, the Mundlak term itself is not statistically significant, with the association observed in pooled cross-sectional models not driven by differences between individuals in exposure to large parental transfers. For Austria, the association between large transfers and fertility intentions is even weaker. The odds ratios for large transfers is below one and statistically insignificant. Likewise, the interaction with childless is insignificant and presents a different direction from what observed in Germany, with the Mundlak term carrying no explanatory power.

The panel results together point to the absence of a clear and systematic link between large parental transfers and fertility intentions, and indicate that the weak associations observed earlier in selected cases are unlikely to be primarily driven by stable unobserved individual heterogeneity, while remaining consistent with selection related to life-course timing and other observed characteristics.

Table 9: Random-effects logistic regression predicting fertility intentions – Large transfers and childlessness

Variable	Austria	Germany
Large transfer received (upper threshold)	0.87 (0.42)	1.01 (1.10)
Childless	3.29*** (0.39)	3.68*** (0.87)
Large transfer $\times$ Childless	1.61 (0.81)	0.06** (0.07)
Mean large transfer (Mundlak term)	0.85 (0.40)	9.05 (12.35)
Observations	6,524	2,278
Respondents	4,035	1,945
ρ	0.47	0.37

Note: Odds ratios (standard errors in parentheses) from random-effects logistic regressions with correlated random effects (Mundlak specification). Large transfers this time correspond to country- and wave-specific thresholds comparable to €5,000 in Round II. Models include all the earlier demographic, socio-economic, parental background controls and wave fixed effects (omitted for brevity).

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

After having highlighted some relevant cross-country heterogeneity and individual-level persistence in fertility intentions using country-specific pooled and random-effects models, we further present evidence from pooled analyses across all countries in Round I. This serves two purposes: to assess whether the patterns observed at the country level are generalizable when the individual-level data are analyzed jointly and to provide a novel framework to analyze the role of country-level institutional factor. This way, the new following pooled specifications complement the random-effects results by exploring potential cross-national regularities. Empirically, this time we interact parental transfer receipt with country indicators and include country-specific wave fixed effects. We further cluster standard errors at the country level to account for within-country correlation and apply sampling weights. This strategy enables the estimation of average associations across countries while allowing for differences in baseline fertility levels, timing, or transfers regimes.

The pooled model for the receipt of any transfer yields overall a small but significant association with short-term fertility intentions ($OR \approx 1.07$), but this effect masks some existing cross-national heterogeneity (Table 10). We use the Netherlands as the reference country, but final effects for each are reported. Taking Germany, recipients of any parental transfers present about 60% higher odds of intending a child. In contrast, the corresponding association in Austria is just below unity and mildly significant and no distinguishable differences emerge in comparison with Estonia and the Czech Republic. The marginal effects from Figure 1 - Panel (a) closely mirror the odds-ratios point estimates from the interaction modes. In fact, Germany is the only country displaying a sizable and significant positive association (almost 9%) with fertility intentions, while the effects in the remaining countries are very small (-2% for Austria at 10% confidence level) or virtually null. The estimated effects in Table 10 for large transfers only do not hint to radically different conclusions. In this case the association for the Netherlands is more precisely estimated although not much sizable ($OR \approx 1.18$); the effect for Germany is much more robust and the one for Austria loses the weak significance obtained for generic transfers. A similar evidence is provided by Figure 2 - Panel (b), with relatively weak margins of 2.5% predicted this time for the Netherlands and a much stronger effect of about 15% for Germany. Overall, the pooled odds-ratios coefficients across countries amount to a significant 1.19 and are largely driven by the German experience. This pooled analysis makes Germany standing out as a context in which parental transfers seem more strongly aligned with fertility planning, even though earlier panel evidence does not support a causal interpretation. This German effect is better explained also by the more mature age structure and lower baseline childlessness of respondents, in line with the general tighter association observed in Round I among parents. Moreover, robust positive associations with a lower magnitude, as it was expected, are to be found for Germany and

Austria also in the evidence on transfers from ascendants (taking together those from parents or grandparents), presented in the Appendix.

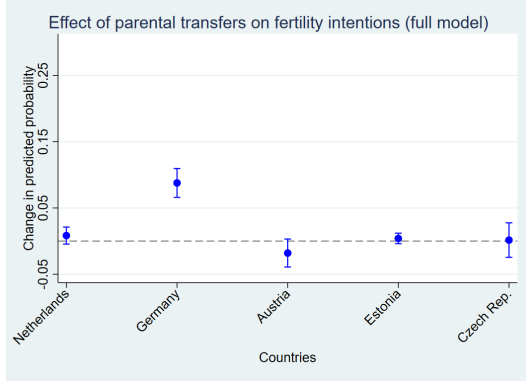
Further on, to assess whether the association between parental transfers and fertility intentions might depend also on the broader institutional context, in the last empirical simpler specifications we interact the transfer receipts with country-year level indicators meant to capture respectively the public family spending (on GDP), early net childcare costs for a dual-earner family (with respect to the average wage) and female labor force participation rate. Table 11 reports that, across the three different models inspecting any transfer receipt, the interaction terms are small and insignificant, signaling that the relationship between the parental support and fertility intentions is not altered by these particular dimensions of the policy environment. Among the institutional factors, a higher female participation rate contributes positively and the generosity of the welfare system contributes negatively to fertility intentions of the European respondents without further mediation on the role of parental transfers. Childcare costs, by contrast, do not display a statistically significant direct association. It is moreover to be noted that only the empowering effect of female participation to the labor force is retained in a full model where most individual characteristics are considered. The nullified effects of transfers might also suggest that earlier correlations tended to reflect more underlying structural conditions and that, when taken individually, generic private transfers do not act necessarily as complements or substitutes of family policies but are contingent to the same structural and normative conditions that used to govern fertility planning in the early 2000s. Yet, the results are quite dissimilar when the focus switches to large parental transfers exceeding €500 (Table 12). Across the three models, large transfers are positively and significantly associated with fertility intentions with about the same size. The institutional factors seem again to shape baseline fertility intentions, but only the female participation rate remains mildly relevant in the complete model. As in the case of any transfers, a larger support is not however necessarily mediated by the quality of the institutional context. Overall, parental transfers of a sizable amount in some contexts can operate as an additional independent resource to sustain family-related decisions.

Table 10: Pooled logistic regression predicting fertility intentions – parental transfers (Round I)

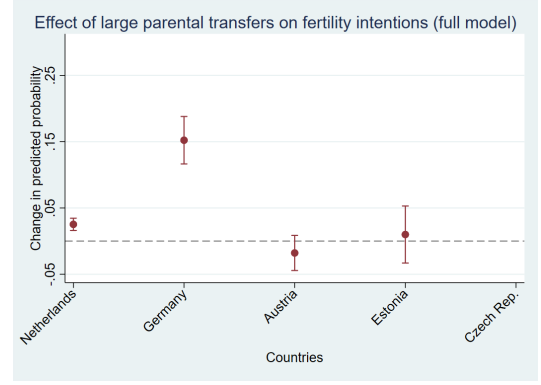
	Any transfer	Large transfer (> 500€)
Model 5: Full controls + country interactions		
<i>Parental transfer (Netherlands, ref.)</i>	1.06 (0.05)	1.18*** (0.03)
Transfer × Germany	1.60*** (0.08)	2.19*** (0.16)
Transfer × Austria	0.90* (0.06)	0.91 (0.07)
Transfer × Estonia	1.02 (0.02)	1.05 (0.13)
Transfer × Czech Republic	1.00 (0.08)	–
Female	1.14 (0.11)	1.15 (0.11)
Age	0.92*** (0.03)	0.92*** (0.03)
Married	1.46** (0.28)	1.54** (0.31)
Childless	1.74*** (0.31)	1.72*** (0.33)
High education (vs low)	2.76*** (0.46)	2.78*** (0.50)
Unemployed (vs employed)	0.85 (0.10)	0.88 (0.10)
Inactive (vs employed)	0.42*** (0.13)	0.39*** (0.12)
Rented housing (vs ownership)	1.23 (0.18)	1.24 (0.18)
Provided housing (vs ownership)	1.06 (0.15)	1.06 (0.17)
Parental background	✓	✓
Country fixed effects	✓	✓
Country-specific wave FE	✓	✓
Pooled average (no interactions)	1.07*** (0.02)	1.19*** (0.06)
Observations	15,319	13,827

Notes: Odds ratios (standard errors clustered at the country level in parentheses) from pooled logistic regressions with interaction terms between transfer coefficients and country fixed effects, using the Netherlands as the reference category. The two models control for individual demographic and socio-economic characteristics, parental background, country fixed effects, and country-specific wave fixed effects (omitted for brevity). Estimates for large (> 500 €) parental transfers in the Czech Republic are not reported due to perfect prediction.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.



(a) Any parental transfers



(b) Large parental transfers (> 500€)

Figure 1: Marginal effects of parental transfers on fertility intentions (Pooled logistic regressions - full models)

Table 11: Pooled logit models with contextual factors – Round I

	Family spending	Childcare costs	Female labor force part.
Parental transfer received	1.09 (0.07)	1.08 (0.07)	1.06 (0.06)
Institutional variable (centered)	0.53*** (0.10)	1.00 (0.00)	1.08*** (0.02)
Transfer × institutional var.	1.34 (0.48)	1.00 (0.01)	1.02 (0.02)
Female	1.03 (0.04)	1.03 (0.04)	1.01 (0.04)
Age	0.94*** (0.00)	0.94*** (0.00)	0.94*** (0.00)
Married	1.46*** (0.08)	1.45*** (0.08)	1.46*** (0.08)
Childless	1.91*** (0.12)	1.91*** (0.11)	1.93*** (0.12)
Year fixed effects	✓	✓	✓
Observations	22,140	22,140	22,140

Notes: Odds ratios from pooled logistic regressions with robust standard errors in parentheses. Institutional indicators vary by country and year and are centered at their mean. All models control for sex, age, marital status, childlessness, and year fixed effects.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 12: Pooled logit models with contextual factors (large transfers) - Round I

	Family spending	Childcare costs	Female labor force part.
Large transfer received (> 500)	1.18** (0.09)	1.18** (0.09)	1.14* (0.09)
Institutional variable (centered)	0.55*** (0.11)	1.01*** (0.00)	1.05* (0.03)
Large transfer × institutional var.	1.14 (0.51)	1.01 (0.01)	1.04 (0.03)
Female	1.00 (0.04)	1.00 (0.04)	1.00 (0.04)
Age	0.94*** (0.00)	0.94*** (0.00)	0.94*** (0.00)
Married	1.52*** (0.09)	1.52*** (0.09)	1.52*** (0.09)
Childless	1.89*** (0.12)	1.90*** (0.12)	1.89*** (0.12)
Observations	20,321	20,321	20,321

Notes: Odds ratios from pooled logistic regressions with robust standard errors in parentheses. Institutional indicators vary by country and year and are centered at their mean. All models control for sex, age, marital status, childlessness, and year fixed effects.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

6 Analysis - Round II

Table 13 reports descriptive statistics for the Round II samples for the five countries, with age restrictions applied (respondents aged 45 or younger). Across all countries, respondents are on average in their early thirties, ranges of mean ages span from 32 in Austria and the Netherlands to slightly over 34 in the Czech Republic. Women are slightly overrepresented in all samples, resembling the gender composition observed in Round I (see Tables 2-4)). Education levels differ greatly across all countries; high education is most prevalent in the Netherlands and Germany, while medium education dominates in Austria. Family formation patterns also display pronounced cross-national heterogeneity. Marriage rates are highest in the Czech Republic and Germany, whereas childlessness exceeds 55 percent in the Netherlands, Germany, and Austria, but remains considerably lower in the Czech Republic and Estonia. Importantly, childlessness status is substantially more prevalent than what documented in Round I, despite broadly comparable age structures and, in some cases, similar or even higher proportions of married respondents. Labor market attachment is strong across all countries, with employment rates exceeding 75% everywhere. Nonetheless, inactivity is more common in Estonia, Czech Republic, and Austria rather than in the Netherlands or Germany. Housing tenure continues to be characterized by sharp cross-national differences. Homeownership rates are particularly high in the Czech Republic and Estonia, while renting is substantially more prevalent in Germany and Austria, in line with housing market patterns already evident in Round I. Fertility intentions within the next three years range from approximately one quarter of respondents in Austria to nearly one third in Estonia and the Czech Republic, in the latter cases higher than in the past since less commonly already realized. Compared to Round I, there is also a more pervasive existence of intergenerational

transfers in all countries, as they capture all lifetime events by the new design. Between 18 and 31 percent of respondents report having received a large transfer from parents, with the highest incidence observed in the Czech Republic (contrarily to the previous round). In greater consistency with Round I, recent large transfers are also relatively uncommon, affecting between 4% and 8.5% of the respondents, but show some meaningful heterogeneity for the multivariate analysis that follows.

Table 13: Round II – Sample characteristics

	Netherlands	Estonia	Czech Rep.	Austria	Germany
Panel A. Sample size - unique wave by country					
Eligible respondents (age ≤ 45)	3876	4772	2745	4263	14832
Panel B. Descriptives					
Age (mean)	32.46	33.60	34.46	32.12	32.57
Female (%)	62.59	55.24	59.85	58.76	55.83
Education: Low (%)	8.05	9.49	15.26	7.71	4.37
Education: Medium (%)	34.46	41.05	37.04	57.52	46.33
Education: High (%)	57.49	49.45	47.70	34.76	49.30
Married (%)	45.15	46.17	57.89	45.37	51.42
Childless (%)	57.07	45.33	38.87	59.82	60.48
Employed (%)	87.83	79.59	78.61	77.85	85.53
Inactive (%)	8.51	15.80	16.75	16.98	11.28
Homeowner (%)	55.02	70.14	68.86	39.71	38.23
Rented housing (%)	29.67	19.59	22.93	49.66	55.97
Good financial situation (%)	70.92	71.60	66.74	63.37	80.80
Intends a child within 3 years (%)	25.25	32.82	30.56	23.89	27.22
Received any large transfer from parents (%)	21.26	18.73	31.33	22.94	23.48
Received any large transfer from ascendants (%)	25.15	25.52	38.47	27.26	30.79
Received recent large transfer from parents (%)	6.66	4.17	8.42	6.29	7.75
Received recent large transfer from ascendants (%)	7.66	5.39	10.05	7.11	9.28
Number of siblings (mean)	1.81	1.79	1.53	1.74	1.70
Father high education (%)	39.17	38.49	25.56	26.28	45.70
Mother high education (%)	31.47	48.73	22.51	18.86	27.76
Parents married (%)	74.40	53.79	62.63	68.16	68.26

Notes: Percentages are shown in points (0–100). Means are reported for continuous variables. Round II countries are observed in a single wave each; no attrition analysis is applicable. “Large transfer” and “recent” follow the paper’s harmonized definitions.

Tables 14 and 15 present the findings from Round II, based on country-level logistic regressions estimating the likelihood of intending to have a child within the following three years. Mirroring what done for Round I, the four models are estimated separately by country, progressively adjusting for demographic, socio-economic, and parental background characteristics. In Model 5, we replace the indicator for having ever received a parental transfer with a measure capturing recent large parental transfers, while maintaining the full set of other controls. Focusing on recent transfers allows us to assess whether the timing of financial support matters for short-term fertility planning and aligns better with the wording of the questionnaires presented during the first Round waves in order to draw some light comparisons.

In the baseline model, the association between having ever received a large parental transfer and the short-term fertility intentions expressed by the respondents varies across countries. A positive and statistically significant association emerges in the Netherlands, Austria, and Germany, with odds

ratios ranging between 1.22 and 1.37, thus indicating higher fertility intentions among transfer recipients. By contrast, in Estonia and the Czech Republic the estimated odds ratios are equal or close to unity and statistically insignificant, suggesting there is no systematic relationship between parental transfers and fertility intentions. Once demographic characteristics are introduced, the baseline patterns this time largely persist. In the Netherlands and Germany, the estimated association between parental transfers and fertility intentions remains positive and statistically significant, with effect sizes comparable to or slightly larger than in Model 1. The stability of the estimated coefficients in Austria also hints that the baseline relationship is not driven by potential compositional differences in the age structure, marital status, or parity between recipients and non-recipients of parental transfers. In contrast, Estonia and the Czech Republic still show no meaningful relationship between transfer receipt and fertility intentions. Individual characteristics such as the young age and childlessness status are relevant predictors anywhere of fertility intentions, whereas being a male, all the rest equal, contributes positively to fertility intentions in the Netherlands, Estonia, and the Czech Republic. The inclusion of education, employment status, subjective financial situation, and housing conditions leads to a general attenuation of the former estimated transfer effects (Table 15). In the Netherlands and Germany, the association between parental transfers and fertility intentions becomes statistically insignificant, suggesting that earlier correlations may be partially mediated by respondents' own socioeconomic position and housing circumstances. In Austria, the coefficient weakens only slightly ($OR \approx 1.31$) but remains positive and statistically significant, pointing to a more robust association. Across the inspected countries, education, childlessness status, young age, housing arrangement and at times the financial situation act as key determinants of fertility intentions. Adding parental education, parental marital status, and number of siblings reveals some different patterns. In the Netherlands, the association between parental transfers and fertility intentions is again fully absorbed, with housing tenure and labor market attachment remaining among the strongest predictors, while parental background variables themselves do not exert independent effects. In Austria, the association between transfers and fertility intentions is stable and statistically significant, even after accounting for family background. In Germany, instead, the inclusion of parental background reverses the attenuation observed in Model 3. The transfer coefficient regains statistical significance with a magnitude similar to the baseline, suggesting that parental background may act as a confounder that masks the transfer–fertility relationship when omitted. Finally, the transfer coefficient of the Czech Republic becomes marginally significant at the 10% level, hinting at a weak and conditional association, while in Estonia the estimate remains null.

When switching to recent transfers alone in Model 5, a differentiated picture emerges once again. In the Netherlands, although higher than 1, the estimate for recent transfer remains statistically

insignificant, and indicates that the timing of the parental support does not modify the overall pattern once all the remaining controls are included. The Dutch context stands for a case where parental financial help does not seem to exert a direct effect on fertility intentions. When examining Estonia, the estimate has a higher magnitude than in the previous models, but it is only marginally significant ($OR \approx 1.42, p < 0.10$). The timing of the financial receipt might be more relevant than the lifetime exposure, but the effect remains not precisely estimated and weaker than in other countries except for the Dutch case. In the Czech Republic, while lifetime transfers play at most a secondary role, recent parental support is associated with nearly 90% higher odds of intending a child, underscoring the potential importance of short-term financial assistance in this context. In Austria and Germany, the association is notably stronger, above 50% higher odds of intending to have a child for the recipients, when the focus is restricted to recent transfers, pointing to the relevance of financial assistance to facilitate childbearing decisions.

When we further allow the association between parental transfers and fertility intentions to vary by parity status, we get some heterogeneity by the type of transfers. In the fully specified models interacting any large parental transfers with childlessness, we observe that — across all countries except the Netherlands — the association between transfers and fertility intentions is significantly stronger among childless respondents. This suggests that in most contexts receiving large parental transfers during lifetime may be particularly relevant for individuals who have not yet entered parenthood. By contrast, the interaction with childlessness for recent parental transfers does not yield the same evidence: the interaction terms are statistically insignificant, indicating that the proximity of parental financial support does not differentially influence fertility intentions among childless respondents. In Germany, recent parental transfers are even associated with approximately 50% higher odds of intending to have an additional child among parents.

Table 14: Country-specific pooled logit models of fertility intentions – Round II (Models 1–2)

Model	Variable	Netherlands	Estonia	Czech Republic	Austria	Germany
Model 1: Baseline						
	Transfer from parents	1.25** (0.13)	1.00 (0.09)	0.96 (0.12)	1.37*** (0.14)	1.22 *** (0.07)
	Observations	3,250	4,233	2,106	3,562	12,031
Model 2: + Demographic controls						
	Transfer from parents	1.33** (0.15)	1.13 (0.11)	1.23 (0.17)	1.36*** (0.16)	1.18*** (0.07)
	Female	0.76*** (0.08)	0.84* (0.07)	0.62*** (0.08)	1.00 (0.11)	0.97 (0.06)
	Age	0.94** (0.01)	0.94*** (0.01)	0.95*** (0.01)	0.94*** (0.01)	0.96*** (0.01)
	Married	1.47*** (0.17)	0.95 (0.09)	1.08 (0.18)	1.47*** (0.20)	1.96*** (0.16)
	Childless	1.71*** (0.24)	2.23** (0.24)	2.98*** (0.57)	2.20*** (0.32)	2.94*** (0.25)
	Observations	2,413	3,188	1,880	2,619	9,245

Note: Odds ratios (with standard errors in parentheses) from country-specific logistic regressions. Survey-weighted estimates using individual weights and stratified by region with single-unit strata centered. Model 2 additionally controls for sex, age, childlessness and marital status.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 15: Country-specific pooled logit models of fertility intentions – Round II (Models 3–4)

Model	Variable	Netherlands	Estonia	Czech Republic	Austria	Germany
Model 3: +Socio-economic controls						
	Transfer from parents	1.13 (0.13)	1.03 (0.10)	1.21 (0.19)	1.31** (0.16)	1.09 (0.07)
	Female	0.82 * (0.09)	0.78 *** (0.07)	0.68 (0.10)	0.94 (0.11)	0.97 (0.06)
	Age	0.89*** (0.01)	0.92*** (0.01)	0.94*** (0.01)	0.94*** (0.01)	0.94*** (0.01)
	Married	1.31** (0.16)	0.83 ** (0.08)	1.00 (0.18)	1.40** (0.20)	2.13*** (0.19)
	Childless	1.87*** (0.28)	2.26 *** (0.28)	2.88*** (0.60)	1.88 *** (0.30)	2.62*** (0.24)
	High education (vs low)	1.39 (0.34)	1.93*** (0.36)	3.47*** (1.00)	1.68* (0.47)	3.12*** (0.52)
	Unemployed (vs emp.)	0.47* (0.20)	0.75 (0.19)	0.47 (0.19)	0.92 (0.31)	0.95 (0.19)
	Inactive (vs emp.)	0.40*** (0.11)	0.73** (0.13)	0.94 (0.24)	1.15 (0.19)	0.97 (0.11)
	Rented housing (vs ownership)	0.56*** (0.10)	0.68* (0.10)	1.52** (0.30)	1.36** (0.18)	1.60*** (0.10)
	Provided housing (vs ownership)	0.17*** (0.05)	1.14 (0.18)	0.95 (0.26)	1.18 (0.24)	1.50** (0.23)
	Good financial situation (vs bad)	1.12 (0.15)	1.38*** (0.08)	0.84 (0.14)	0.96 (0.13)	0.84 ** (0.07)
	Observations	2,314	2,993	1,758	2,366	8,839
Model 4: + Parental Background controls						
	Transfer from parents	1.08 (0.14)	1.06 (0.12)	1.34 * (0.21)	1.33 ** (0.18)	1.24 *** (0.09)
	Female	0.84 (0.10)	0.75** (0.08)	0.61 *** (0.85)	0.93 (0.11)	1.02 (0.07)
	Age	0.90*** (0.01)	0.92*** (0.01)	0.93*** (0.01)	0.93*** (0.01)	0.94*** (0.01)
	Married	1.23 (0.16)	0.82* (0.09)	1.02 (0.19)	1.44** (0.22)	1.93*** (0.19)
	Childless	2.05*** (0.34)	2.29 *** (0.32)	3.11 *** (0.68)	2.07 *** (0.37)	2.77 *** (0.29)
	High education (vs low)	1.47* (0.40)	1.41 (0.33)	4.90 *** (1.64)	2.33 ** (0.78)	3.81*** (0.86)
	Unemployed (vs emp.)	0.42* (0.20)	0.74 (0.22)	0.53 (0.24)	0.93 (0.29)	1.06 (0.25)
	Inactive (vs emp.)	0.40 *** (0.13)	1.03 (0.17)	0.91 (0.24)	1.03 (0.19)	0.97 (0.12)
	Rented housing (vs ownership)	0.61*** (0.11)	0.69** (0.18)	0.61** (0.13)	1.34** (0.19)	1.65*** (0.12)
	Provided housing (vs ownership)	0.20*** (0.06)	1.25 (0.22)	1.01 (0.29)	1.09 (0.24)	1.39** (0.23)
	Good financial situation (vs bad)	1.16 (0.17)	1.47*** (0.18)	0.85 (0.15)	1.06 (0.15)	0.93 (0.09)
	Parental background	✓	✓	✓	✓	✓
	Observations	1,972	2,329	1,649	2,104	6,843

Note: Odds ratios (with standard errors in parentheses) from country-specific logistic regressions. Survey-weighted estimates using individual weights and stratified by region with single-unit strata centered. . Model 3 additionally controls for education, activity status, housing, and financial condition (making odds meet). Model 4 further includes parental background characteristics (education of mother and father, marital status of parents and number of siblings) - not reported for brevity.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

Table 16: Country-specific pooled logit models of fertility intentions – Round II (Model 5)

Model	Variable	Netherlands	Estonia	Czech Republic	Austria	Germany
Model 5:	Transfer timing					
	Recent transfer from parents	1.27 (0.24)	1.42* (0.30)	1.87*** (0.42)	1.53** (0.29)	1.58*** (0.16)
	Female	0.84 (0.10)	0.75*** (0.08)	0.61*** (0.10)	0.92 (0.11)	1.02 (0.07)
	Age	0.90*** (0.01)	0.92*** (0.01)	0.94*** (0.01)	0.94*** (0.01)	0.95*** (0.01)
	Married	1.24 (0.17)	0.81* (0.09)	1.01 (0.19)	1.43** (0.22)	1.93*** (0.19)
	Childless	2.05*** (0.34)	2.28*** (0.32)	3.04*** (0.67)	2.03*** (0.36)	2.76*** (0.29)
	High education (vs low)	1.45** (0.40)	1.39 (0.32)	4.97*** (1.65)	2.39** (0.81)	3.75*** (0.85)
	Unemployed (vs emp.)	0.42* (0.20)	0.74 (0.22)	0.54 (0.24)	0.94 (0.29)	1.06 (0.25)
	Inactive (vs emp.)	0.40*** (0.13)	1.03 (0.17)	0.90 (0.23)	1.03 (0.19)	0.98 (0.12)
	Rented housing (vs ownership)	0.61*** (0.11)	0.70** (0.12)	1.57** (0.32)	1.29* (0.18)	1.63*** (0.12)
	Provided housing (vs ownership)	0.20*** (0.06)	1.26 (0.23)	1.00 (0.28)	1.04 (0.23)	1.38* (0.23)
	Good financial situation (vs bad)	1.15 (0.17)	1.47*** (0.18)	0.85 (0.15)	1.08 (0.16)	0.93 (0.09)
	Parental background	✓	✓	✓	✓	✓
	Observations	1,972	2,329	1,649	2,104	6,843

Note: Odds ratios (with standard errors in parentheses) from country-specific logistic regressions. Survey-weighted estimates using individual weights and stratified by region with single-unit strata centered. Recent transfers are defined as parental transfers received either in the interview year or in the preceding year.

* $p < 0.1$, ** $p < 0.05$, *** $p < 0.01$.

To complement the country-specific analyses and to address limitations related to sample size and statistical power in some national samples, we additionally estimate pooled models combining all five countries in Round II. Pooling the data allows us to exploit substantially greater variation while retaining a within-country identification strategy through the inclusion of country fixed effects, and to formally test whether the association between parental transfers and fertility intentions varies across national contexts. This approach is particularly useful given that some country samples—most notably Estonia—contain relatively few transfer recipients, which may limit the precision of country-specific estimates.

The pooled results closely mirror the patterns already identified in the single-country analyses. When the effect of parental transfers is constrained to be homogeneous, the pooled model yields a positive average association between transfer receipt and fertility intentions ($OR \approx 1.17$). However, once interactions with countries are introduced, substantial heterogeneity emerges. The association remains positive and statistically significant in the Netherlands, Germany, and Austria, with effect sizes comparable to those obtained from separate national regressions. It is instead markedly weaker and statistically indistinguishable from zero in Estonia and the Czech Republic (Table 17). Importantly, adding individual-level controls for age, gender, marital status, and childlessness leaves the qualitative pattern unaffected, and the pooled models confirm the findings from the earlier country-specific models. In Model 3 from Table 18, once a rich set of demographic and socioeconomic controls

is introduced and the effect of parental transfers is allowed to vary by country, the country-specific estimates become smaller but retain significance for Western European countries. By contrast, no meaningful association emerges for Estonia and the Czech Republic, where the estimated odds ratios become closer to one. In a specification without interactions, the average odds ratio equal to 1.13, confirming a modest but statistically significant association between parental transfers and fertility intentions when ignoring cross-country heterogeneity. Model 4 augments the specification by introducing the usual parental background characteristics. The inclusion of those variables does not nullify the association between parental transfers and fertility intentions, and if anything, the estimated effects present a larger magnitude for the Netherlands, Germany and Austria. There is no statistically significant relationship found for Estonia or the Czech Republic, indicating that the effect of generic intergenerational transfers have most probably little impact on fertility intentions in these contexts. The pooled estimate without interactions increases to 1.21, reinforcing the conclusion that parental transfers are correlated with fertility intentions on average, but that this relationship is mainly driven by Western Europe.

Table 17: Pooled logistic regression predicting fertility intentions – country-specific effects and individual controls (Round II, Models 1–2)

	Odds Ratio	Std. Error
Model 1: Baseline		
Transfer (Netherlands, ref.)	1.17**	(0.04)
Transfer × Germany	1.22***	(0.01)
Transfer × Austria	1.37***	(0.07)
Transfer × Estonia	1.00	(0.05)
Transfer × Czech Rep.	0.96	(0.12)
Pooled average (no interactions)	1.17***	(0.04)
Observations	25,172	
Model 2: Individual controls		
Transfer (Netherlands, ref.)	1.36***	(0.15)
Transfer × Germany	1.23***	(0.04)
Transfer × Austria	1.33***	(0.09)
Transfer × Estonia	1.08	(0.09)
Transfer × Czech Rep.	1.12	(0.16)
Female	0.89**	(0.04)
Age	0.95***	(0.01)
Married	1.48**	(0.24)
Childless	2.45***	(0.20)
Pooled average (no interactions)	1.22***	(0.04)
Observations	19,345	

Notes: Odds ratios (with standard errors in parentheses). Survey-weighted logistic regressions accounting for clustering at the country–region level. All models include country fixed effects and interactions between parental transfer receipt and country dummies (the Netherlands act as the reference country). Model 2 additionally controls for sex, age, childlessness and marital status.

*** $p < 0.01$, ** $p < 0.05$.

Table 18: Pooled logistic regression predicting fertility intentions – extended controls (Round II, Models 3–4)

	Odds Ratio	Std. Error
Model 3: Socioeconomic controls		
Transfer (Netherlands, ref.)	1.28*	(0.16)
Transfer × Germany	1.13***	(0.03)
Transfer × Austria	1.23***	(0.04)
Transfer × Estonia	1.00	(0.10)
Transfer × Czech Rep.	1.07	(0.16)
Female	0.88**	(0.04)
Age	0.94***	(0.01)
Married	1.46*	(0.29)
Childless	2.21***	(0.14)
Medium education	1.36**	(0.14)
High education	2.27***	(0.37)
Unemployed (vs emp.)	0.77*	(0.11)
Inactive (vs emp.)	0.91	(0.08)
Rented housing (vs ownership)	1.27	(0.18)
Provided housing (vs ownership)	0.96	(0.21)
Good financial situation (vs bad)	1.00	(0.08)
Pooled average (no interactions)	1.13***	(0.03)
Observations	18,270	
Model 4: Family background controls		
Transfer (Netherlands, ref.)	1.24**	(0.11)
Transfer × Germany	1.26**	(0.02)
Transfer × Austria	1.29***	(0.07)
Transfer × Estonia	1.05	(0.07)
Transfer × Czech Rep.	1.14	(0.17)
Female	0.89	(0.05)
Age	0.94***	(0.01)
Married	1.41	(0.24)
Childless	2.36***	(0.16)
High education (vs low)	2.56**	(0.53)
Unemployed (vs empl.)	0.77	(0.13)
Inactive (vs empl.)	0.90	(0.08)
Rented housing (vs ownership)	1.30	(0.18)
Provided housing (vs ownership)	0.96	(0.17)
Good financial situation (vs bad)	1.08	(0.07)
Parental background	✓	
Pooled average (no interactions)	1.21***	(0.03)
Observations	14,897	

Notes: Odds ratios (with standard errors in parentheses). Survey-weighted logistic regressions accounting for clustering at the country–region level. All models include country fixed effects and interactions between parental transfer receipt and country dummies (the Netherlands act as the reference country). Model 3 additionally controls for education, activity status, housing, and financial condition (making ends meet). Model 4 further includes parental background characteristics (education of mother and father, marital status of parents and number of siblings) - not reported for brevity.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

The results for recent receipt of large transfers in Round II reveal a stronger and more homogeneous association with short-term fertility intentions compared to earlier models. In the Netherlands, respondents who recently received a parental transfer show about 37% higher odds of intending to have a child within the next three years (significant at 5% level) but the relationship appears even more pronounced in other countries. In Germany and Austria the odds of fertility intentions exceed 50%, and positive and significant effects are this time observed in Estonia and the Czech Republic (respectively, $OR \approx 1.41$ and $OR \approx 1.70$). Moreover, the pooled specification without any country interactions yields an average odds ratio of 1.54, reflecting the overall robust relationship between recent parental financial support and fertility intentions. Read together with findings for ascendants' transfers in the Appendix, the intergenerational financial support therefore seems to matter most

when it is temporally close to fertility planning. This is evident also from the estimated margins for the full models (Figure 2). For large parental transfers received at any point in the past, the effects are heterogeneous across countries. Meaningful increases of fertility intentions (about 4–5 p.p.) are detectable in Germany, Austria, and the Netherlands, while the effects in Estonia and the Czech Republic are smaller and not significant. By contrast, recent parental transfers are consistently and strongly associated with higher intended fertility across all countries, with percentages ranging from 6% in the Netherlands to 11% in the Czech Republic (although with less precision). The

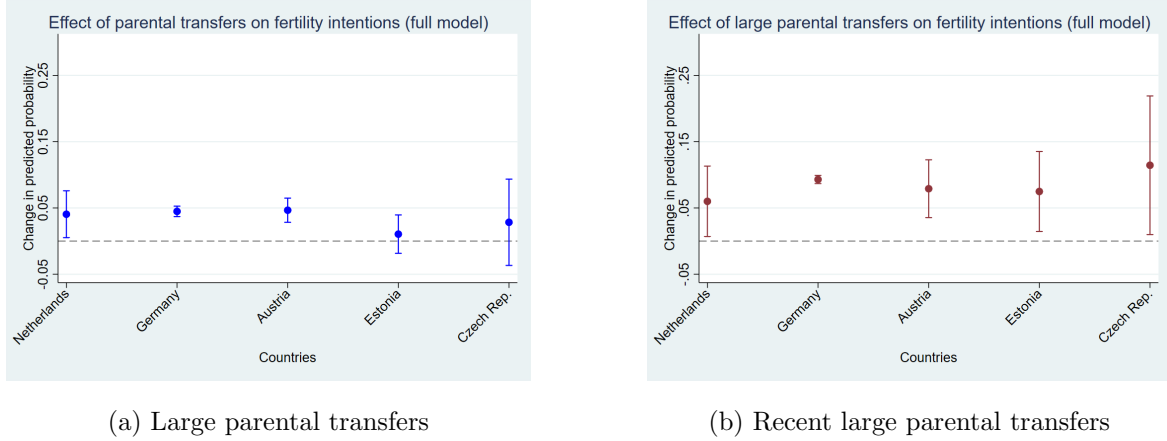


Figure 2: Marginal effects of parental transfers on fertility intentions (Pooled logistic regressions - full models)

Table 19: Pooled logistic regression predicting fertility intentions – Recent parental transfers (Round II, Model 5)

	Odds Ratio	Std. Error
Model 5: Recent parental transfers + full controls		
Recent transfer (Netherlands, ref.)	1.37**	(0.17)
Recent transfer × Germany	1.59 ***	(0.02)
Recent transfer × Austria	1.50***	(0.17)
Recent transfer × Estonia	1.41**	(0.18)
Recent transfer × Czech Rep.	1.70**	(0.39)
Female	0.89	(0.05)
Age	0.94***	(0.01)
Married	1.40	(0.24)
Childless	2.35***	(0.16)
High education (vs low)	2.54**	(0.53)
Unemployed (vs employed)	0.77	(0.13)
Inactive (vs employed)	0.90	(0.08)
Rented housing (vs ownership)	1.28	(0.18)
Provided housing (vs ownership)	0.95	(0.17)
Good financial situation (vs bad)	1.08	(0.07)
Parental background	✓	
Pooled average (no interactions)	1.54***	(0.07)
Observations	14,897	

Notes: Odds ratios (with standard errors in parentheses). Survey-weighted logistic regressions accounting for clustering at the country–region level. All models include country fixed effects and interactions between parental transfer receipt and country dummies (the Netherlands act as the reference country. Recent transfers are defined as parental transfers received either in the interview year or in the preceding year.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

In order to better inspect whether the cross-national differences observed between the parental

financial transfers and fertility intentions reflect the broader institutional environment, we further include the contextual-level factors already discussed for Round I. They are drawn from the relevant databases (OECD, 2026; World Bank, 2026) at the baseline year of 2021, and they are centered at the cross-country mean to facilitate the interpretation. This approach allows to exploit the between-country variation in the policy environment on family matters while maintaining a tighter specification (comparable to Models 2 from previous analyses) focused on the interaction between the financial support from the family of origin and the policy context.

The pooled analyses for Round II show some consistent evidence of a positive association between the receipt of parental transfers and fertility intentions. For all the three inspected models, in fact, individuals who received a large parental transfer display about 20% higher odds of intending to have a child (Table 20). This association becomes even stronger, when focusing on recent transfers only, with odds ratios increasing to 1.58-1.60 (Table 21). Thus, the timing of receipt seems to play again a relevant role in this second Round of data collection and to facilitate decision-making when it comes to fertility planning. These same estimates are to be confirmed in more complete models, when we further account for all the other individual characteristics. By contrast, the included institutional factors show weaker and inconsistent effects in Round II. In fact, childcare costs only seem to display some robust association with fertility intentions, and similarly to the previous Round, in the unexpected direction. Moreover, there is some limited evidence for the existence of insitutional moderation. A mild and negative interaction between parental transfers and family spending may be revelatory of an attenuation of the role of all large transfers in contexts of a more generous public support. This same effect, however, disappears completely when more complete specifications are examined or when focusing on the recent transfers alone. All together, the short-term parental support operates quite independently of the institutional environment, at least as captured by the selected indicators.

When compared to Round I, large, and particularly recent, parental transfers display a greater association with fertility intentions. This shift in magnitude of the coefficient is for sure partially a reflection of the definition used for such large transfers (above 5000 €) in the latest wave, but it may also stand for an increasing reliance on family-based resources in fertility planning. At the same time, the evidence of weaker institutional main effects and interaction terms further raises suspects on the fact that transfers may have become over time an immediate resource available to sustain reproductive decisions with the institutional context playing a comparatively secondary role.

Table 20: Pooled logit models with contextual factors – Round II

	Family spending	Childcare costs	Female labor force part.
Large parental transfer	1.22*** (0.04)	1.20*** (0.04)	1.22*** (0.04)
Institutional variable (centered)	1.35 (0.28)	1.02 *** (0.01)	1.03 (0.03)
Transfer × institutional var.	0.84 * (0.08)	1.00 (0.01)	1.02 (0.02)
Female	0.90 ** (0.04)	0.89 ** (0.04)	0.89 ** (0.04)
Age	0.95*** (0.00)	0.94*** (0.00)	0.95*** (0.00)
Married	1.41** (0.22)	1.44** (0.24)	1.44*** (0.24)
Childless	2.29*** (0.20)	2.34*** (0.11)	2.34*** (0.21)
Year fixed effects	✓	✓	✓
Observations	19,345	19,345	19,345

Notes: Odds ratios from pooled logistic regressions and year fixed effects. Institutional indicators vary by country and are centered at their mean (year=2021). All models control for sex, age, marital status, childlessness, and year fixed effects.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 21: Pooled logit models with contextual factors (recent transfers) – Round II

	Family spending	Childcare costs	Female labor force part.
Large parental transfer	1.60*** (0.06)	1.58*** (0.06)	1.60*** (0.06)
Institutional variable (centered)	1.31 (0.28)	1.03 *** (0.01)	1.03 (0.03)
Transfer × institutional var.	1.00 (0.12)	1.00 (0.07)	0.98 (0.02)
Female	0.89 ** (0.04)	0.90 ** (0.04)	0.90 ** (0.04)
Age	0.95*** (0.00)	0.95*** (0.00)	0.95*** (0.00)
Married	1.40** (0.23)	1.43** (0.24)	1.41*** (0.23)
Childless	2.28*** (0.21)	2.32*** (0.21)	2.26*** (0.21)
Year fixed effects	✓	✓	✓
Observations	19,345	19,345	19,345

Notes: Odds ratios from pooled logistic regressions and year fixed effects. Institutional indicators vary by country and are centered at their mean (year=2021). Recent transfers are defined as parental transfers received either in the interview year or in the preceding year. All models control for sex, age, marital status, childlessness, and year fixed effects.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Propensity score matching results

To assess whether the link between intergenerational support and fertility intentions is affected by selection bias, we complement the previous regression analyses with a propensity score matching (PSM) approach. We estimate the propensity score using a logistic regression model in which the treatment variable is defined first as having received any large parental transfer and second for having received recent large parental transfers (consistently with the approach used for Model 5). In the two cases, the propensity score models included gender, age, educational attainment, childlessness, parental education, parental marital status, and the number of siblings. These factors are selected to capture key demographic, socioeconomic, and family-background characteristics that are plausibly related both to the likelihood of receiving parental transfers and to fertility intentions, while avoiding the ones that may lie on the causal pathway between transfers and fertility planning (such as the housing condition). Matching is then performed using one-to-one nearest-neighbor matching with replacement, imposing a caliper of 0.05 and restricting the analysis to the region of common support, to ensure that treated and untreated respondents are comparable in terms of observed characteristics. Fertility intentions within the next three years are used as the outcome variable, with treatment effects correspond to the average treatment effect on the treated (ATT). This matching strategy is applied separately for each country and for each definition of parental transfers, allowing for a comparison of results both across different contexts and according to timing. While the number of treated observations—particularly for recent transfers—is relatively small in some countries and may limit statistical power, the matching procedure ensures good covariate balance, and the cross-country variation in results suggests that the findings are not driven solely by sample size constraints.

The PSM results depict cross-country differences in the association between parental financial transfers and short-term fertility intentions, mostly corroborating the patterns that emerged earlier in the regression-based analyses. For the Netherlands and Estonia, there is no robust evidence that parental transfers - whether defined in the lifetime or as recent support - are associated with fertility intentions, once the observable characteristics are balanced. The estimated ATT effects are small in magnitude and statistically insignificant, with no meaningful difference in the propensity to intend a child between recipients of parental transfers and comparable non-recipients.

A similar conclusion can be reached when focusing on the Czech Republic. While standard regression models in the complete specifications suggested a positive association between recent parental transfers and fertility intentions, this relationship disappears entirely after matching. Both lifetime

and recent large transfers provide treatment effects close to zero and statistically insignificant, despite good covariate balance and substantial overlap between treated and control groups. This divergence between regression and matching results suggests that initial associations in the Czech case are likely driven by selection into transfer receipt along the socio-economic and family-background dimensions rather than reflecting an independent association.

The matching estimates for Austria and Germany are the only ones to provide evidence of a positive association between parental financial transfers and short-term fertility intentions. In Austria, receiving generic large parental transfers is associated with an increase of approximately 5.5 percentage points in the probability of having short-time fertility intentions, and effect that is statistically significant after matching. Once we restrict the attention to recent transfers only, the estimated effect grows larger (around 8%) although it is less precisely estimated (significant at the 10% level), due to a smaller number of treated observations. Germany displays a similar but more pronounced pattern. Matched estimates indicate a modest and significant association for the receipt of generic large transfers (around 3%), while the effect of recent parental transfers is substantially larger—close to 9 p.p.—and highly statistically significant. The consistency of findings across methods and across transfers typology in the two countries suggests that the timing of parental support plays a consequential role in shaping short-term fertility intentions. The matching analysis reinforces to some extent the cross-national diversity detailed so far for Round II: parental financial transfers are not systematically associated with fertility intentions in the Netherlands, Estonia, or the Czech Republic. Austria and Germany stand out instead as national settings where parental transfers may accompany reproductive decisions more closely than elsewhere in recent years.

Table 22: Propensity score matching estimates of transfer receipt on fertility intentions

Country	ATT	SE	Treated (N)	Controls (N)
Netherlands	0.042	0.028	631	2,041
Estonia	0.019	0.028	675	2,525
Czech Republic	0.025	0.029	700	1,265
Austria	0.055**	0.025	800	2,200
Germany	0.031**	0.014	2,342	6,507

Note: ATT = Average Treatment Effect on the Treated. Estimates from 1:1 nearest-neighbor matching with a 0.05 caliper. ** $p < 0.05$.

Table 23: Propensity score matching estimates of recent transfer receipt on fertility intentions

Country	ATT	SE	Treated (N)	Controls (N)
Netherlands	-0.005	0.047	200	2,473
Estonia	0.042	0.055	155	3,045
Czech Republic	0.002	0.048	195	1,775
Austria	0.082*	0.043	218	2,782
Germany	0.089***	0.022	781	8,068

Note: ATT = Average Treatment Effect on the Treated. Estimates from 1:1 nearest-neighbor matching with a 0.05 caliper. *** $p < 0.01$, * $p < 0.10$.

7 Discussion

This study has examined whether receiving financial help from parents (and ascendants) is associated with short-term fertility intentions in five European countries, using the available harmonized panel and cross-sectional data from the two rounds of the Generations and Gender Survey (GGS). Our findings broadly indicate there is no positive systematic relationship between transfer receipt and fertility intentions, especially when considering evidence from Round I. Both the effects for generic transfers and for larger transfers (defined on a low threshold of 500 euro) largely attenuate in multivariate models (H1), despite an initial positive association observed in the simplest bivariate specification. Some heterogeneity nevertheless emerges when considering parity, running counter to the expectation that associations would be stronger among childless respondents (H2). In the Netherlands, Austria, and especially Germany, both generic and large parental transfers are strongly associated with fertility intentions among respondents who already have children, while associations among the childless are weak or absent. When countries are pooled, Germany stands out as the only case with consistent positive associations (about 9% higher probability of intending a child for generic and 15% for large transfers), whereas the estimates for the other countries were unstable (Netherlands and Austria) and generally imprecisely estimated. Panel analyses based on Round I using random-effects models with Mundlak corrections do not provide any evidence in favor of a stable association between large transfers and fertility intentions, irrespective of the large threshold adopted, when controlling for the unobserved individual heterogeneity. This suggests that the effects observed in pooled specifications may be more closely related to compositional differences between recipients and non-recipients, as well as the timing of transfer receipt, rather than to an independent effect of transfers per se. At the same time, features of the German panel (relatively mature and non-childless respondents, high attrition rate) may have limited our ability to precisely assess within-individual dynamics.

In contrast to Round I, Round II reveals more consistent patterns of association between parental financial transfers and short-term fertility intentions, providing partial support to H1 and H3. Austria and Germany, and to a less extent the Czech Republic, display robust positive associations in full

models. When allowing the association to vary by parity, there is some mixed support for H2. In fact, when considering all lifetime large transfers, the interactions with childlessness point to a stronger association among childless respondents in all countries except for the Netherlands. By contrast, when the focus shifts to recent large transfers — more comparable to the Round I setting — the interaction terms are generally insignificant, with Germany an exception, where recent transfers are linked to higher intentions to have an additional child, reproducing a pattern observed in the earlier round. Recent parental transfers are also associated with markedly higher odds of intending a child in Austria, Germany, and the Czech Republic, and more weakly in Estonia. The pooled models across countries align closely with the country-specific findings. For generic large transfers, the association with fertility intentions is modest and largely driven by Western European countries (with estimated margins around 4–5%). Recent large parental transfers have a more homogeneous association across contexts and considerably larger effect sizes, providing support for H3. When we further address observable selection through propensity score matching, only Austria and Germany have persistent positive and significant treatment effects, particularly for recent large parental transfers. This suggests that, in the Netherlands and the Czech Republic, earlier positive estimates are unlikely to reflect independent transfer effects. The matching analysis for the recent years would therefore corroborate the idea that financial support is mostly relevant when sizable and timely in specific national contexts rather than as a regularity across the whole Europe. In line with the limited evidence in Round I, the institutional moderation, as captured by three selected variables, continues to be weak in Round II (H4). There are no stable patterns in the interactions between parental transfers and the policy indicators, and the mild attenuation of transfer effects in contexts of high family spending disappears in complete models. At the same time, we note an interesting geographic concentration of the effects in Austria and Germany, which would be consistent with a setting of moderate supported familialism (Saraceno & Keck, 2010). Where institutional support for families is relatively strong but parental responsibilities in child rearing remain central, large timely parental transfers may act as a complementary resource that assists reproductive choices.

We acknowledge several limitations when discussing these findings. First, the relative absence of evidence for Central-Eastern European countries has to be analyzed through a specific lens: in both Czech Republic and Estonia fewer observations enter in the complete models, and the incidence of large recent parental transfers is comparatively low, thus limiting statistical power. This particularly applies to Estonia, where the more restrictive definition of transfers is rare in both rounds, and to the Czech Republic during Round I. We also note that respondents in these two settings tend to report higher fertility intentions than in Western European counterparts, despite lower rates of

childlessness in the analytical samples. This suggests that fertility planning may be not as much influenced by uncommon intergenerational financial support in these societies, but by expectations on parenthood linked to different socially normative or structural factors. On the contrary, in Western Europe, notably in Austria and Germany, sizable and timely parental transfers carry more salience and appear to function as additional enabling resources for prospective parents, especially when they plan for more children. This picture is drawn mainly on evidence from Round II and cannot be interpreted causally as the analysis has relied on cross-sectional observations and propensity score matching, which cannot eliminate unobserved confounding. Moreover, no real comparisons across rounds are possible given the analyzed differences in sample compositions (no repeated observations across rounds) and measurement of transfers. Substantial efforts were made to better harmonize the explanatory variable - focusing on large recent transfers where possible -, yet remaining limitations include the incomplete information on transfer amount (Netherlands, Round I) and the low incidence of transfers that are really comparable to €5,000 threshold in Round I (available only for Austria and Germany). Furthermore, these constraints imply that the results are to be interpreted as indicative of contextual associations rather than direct causal effects. Despite the different strategies used to deal with selection and timing, we cannot completely disentangle transfers that are provided just ahead of planned family transitions or other anticipatory mechanisms. Likewise, the analytical sample adopted reflect the progressive inclusion of socio-economic and parental background variables that are not consistently observed among all respondents. In Round I, baseline and demographic models rely on nearly identical samples, which means the attenuation of the transfer effects in the main models is primarily linked to the covariate adjustment rather than to sample composition. In the second Round the reduction in sample size is more substantial, and the limited fluctuations seen in statistical significance could be attributable also to estimation precision and composition, but unlikely to mirror reversals in the relationship between parental transfers and fertility intentions.

These considerations also serve to stress the relevance of a continued investment in collecting richer data on transfer histories and fertility intentions in Europe. Data infrastructures that enable a precise reconstruction of the existence, magnitude, and timing of private transfers—potentially through linkages to register-based sources—would greatly advance our understanding of how familial resources intersect with fertility planning and, more generally, help in the assessment of dynamic and stratifying effects of intergenerational material support across the life course. By providing a comparative analysis of the relationship between private financial transfers and fertility intentions in selected European countries, and by explicitly accounting for aspects such as timing, parity, and selection, this article illustrates both the potential and the limits of survey data like the GGS for studying informal family resources. In doing so, it provides an empirical groundwork for future efforts

to extend this approach to a broader set of countries—particularly Southern European societies, where familialism is strong, as well as more defamilialized contexts—in order to obtain a more comprehensive picture of how and where informal support may be linked to a crucial dimensions of low-fertility settings like the divergence between desired and achieved family sizes. The second Round of the GGS appears especially promising to advance this research agenda and to better assess the real contribution of material family support in fertility planning, owing to a better harmonization across countries and the prospective availability of the additional waves.

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Appendix

Alternative pooled specification with transfers from ascendants - Round I

Table A1: Pooled logistic regression predicting fertility intentions – ascendants' transfers (Round I)

	Any transfer	Large transfer (> 500€)
Model 5: Full controls + country interactions		
<i>Ascendants' transfer (Austria, ref.)</i>	1.08*** (0.00)	1.15*** (0.00)
Transfer × Germany	1.04*** (0.00)	1.52*** (0.00)
Transfer × Czech Republic	0.64*** (0.00)	–
Female	0.79*** (0.00)	0.79*** (0.00)
Age	0.95*** (0.00)	0.95*** (0.00)
Married	0.88*** (0.00)	0.88*** (0.00)
Childless	2.08*** (0.00)	2.09*** (0.00)
High education (vs low)	1.80*** (0.01)	1.79*** (0.01)
Unemployed (vs employed)	1.31*** (0.00)	1.31*** (0.00)
Inactive (vs employed)	1.39*** (0.01)	1.39*** (0.01)
Rented housing (vs ownership)	1.37*** (0.00)	1.37*** (0.00)
Provided housing (vs ownership)	1.80*** (0.00)	1.79*** (0.00)
Parental background	✓	✓
Country fixed effects	✓	✓
Country-specific wave FE	✓	✓
Pooled average (no interactions)	1.08***	1.15***
Observations	10,453	10,452

Notes: Odds ratios (standard errors clustered at the country level in parentheses) from pooled logistic regressions with interaction terms between ascendants' (parents or grandparents) transfer receipt and country fixed effects, using Austria as the reference category. Models control for individual demographic and socio-economic characteristics, parental background, country fixed effects, and country-specific wave fixed effects (omitted for brevity). Estimates for large ascendants' transfers (> 500 €) in the Czech Republic are not reported due to perfect prediction.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Alternative pooled specification with transfers from ascendants - Round II

Table A2: Pooled logistic regression predicting fertility intentions – Recent ascendants' transfers (Round II, Model 5)

	Odds Ratio	Std. Error
Model 5: Recent ascendants' transfers + full controls		
Recent transfer (Netherlands, ref.)	1.33*	(0.21)
Recent transfer × Germany	1.50***	(0.31)
Recent transfer × Austria	1.37**	(0.16)
Recent transfer × Estonia	1.52***	(0.15)
Recent transfer × Czech Rep.	1.47*	(0.30)
Female	0.85**	(0.06)
Age	0.94***	(0.01)
Married	1.28	(0.22)
Childless	2.35***	(0.18)
Medium education (vs low)	1.36	(0.27)
High education (vs low)	2.41***	(0.54)
Unemployed (vs employed)	0.73	(0.14)
Inactive (vs employed)	0.91	(0.08)
Good financial situation (vs bad)	1.10	(0.08)
Rented housing (vs ownership)	1.24	(0.18)
Provided housing (vs ownership)	0.99	(0.16)
Parental background	✓	
Pooled average (no interactions)	1.46***	(0.08)
Observations	18,875	

Notes: Odds ratios from survey-weighted logistic regressions with standard errors in parentheses. Models account for the complex survey design and clustering at the country–region level. All specifications include country fixed effects and interactions between recent ascendants' transfer receipt and country (the Netherlands is the reference category). Recent transfers refer to financial support received from parents or grandparents in the interview year or the preceding year.

*** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Descriptives of institutional moderators

Table A3: Country-level indicators (Round I)

	Netherlands	Estonia	Czech Republic	Austria	Germany
Family spending (% GDP, wave 1)	1.6	1.8	1.7	2.0	2.7
Family spending (% GDP, wave 2)	1.8		2.2	1.9	2.7
Family spending (% GDP, wave 3)	1.6				
Childcare cost (% of average net income, wave 1)	13	3	31	7	18
Childcare cost (% of average net income, wave 2)	13		40	12	3
Childcare cost (% of average net income, wave 3)	10				
Female labor force participation (% , wave 1)	55.9	53.3	50.5	49.4	53.5
Female labor force participation (% , wave 2)	56.1		49.3	52.6	55.1
Female labor force participation (% , wave 3)	58.7				

Notes: Indicators are reported by survey wave. Definitions of indicators are consistent with those used in Table 11.

Table A4: Country-level indicators (Round II)

	Netherlands	Estonia	Czech Republic	Austria	Germany
Family spending (% GDP, 2021)	1.89	3.14	2.23	3.44	3.46
Childcare cost (% of average net income, 2021)	15	5	20	2	1
Female labor force participation (% , 2021)	61.9	57.8	51.6	55.4	55.0

Notes: Definitions of indicators are consistent with those used in Table 20.