# Fleeing the Crowded Nest: Siblings and Leaving the Parental Home*

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**Abstract**

We investigate whether the decision of young adults on when to leave the parental home is influenced by the number of siblings they have, in the context of European countries over the last seventy years. Exploiting random variation in sibship size induced by twin births, we identify the causal effect of having an extra sibling on the timing of home-leaving. We find that one additional sibling speeds up the transition to independent living by one year. We provide evidence that the main mechanism behind our results is a decrease in the value of inter-generational co-residence, as having an extra sibling entails higher privacy costs associated with living in a more crowded nest. We furthermore document response heterogeneity by groups of countries, with Western European countries displaying the largest responses. Our findings suggest that demographic trends may have attenuated the gap in co-residence rates across different European regions.

*Keywords:* parental home, inter-generational co-residence, youth emancipation, siblings.

*JEL classification numbers:* D10, J11, J12, J13

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# 1 Introduction

The proportion of young adults (individuals aged 18-34) living in the parental home has been rising across the world over the starting years of the twenty-first century (Esteve and Reher (2021)). Ranging from the US and Canada to Latin America and Mediterranean Europe, individuals from recent cohorts left their parental homes much later than those from previous generations did. As the timing of youth emancipation[1] has been found to be a relevant determinant of many life-course outcomes (ranging from fertility choices to geographic mobility and wages), a vast literature emerged trying to understand the causes of the observed postponement of home-leaving: notable explanations include the spread of unstable employment (Becker et al. (2010)), the postponement of union formation (Mazurik, Knudson, and Tanaka (2020)) and shifts in cultural norms (Giuliano (2007)).

In this paper, we study an additional channel that might have driven the increase in inter-generational co-residence: past fertility trends and their effect on family structure. Due to the plunge in fertility rates that occurred in the second half of the last century, youths born after the seventies have, on average, fewer siblings with respect to individuals from previous cohorts. We ask whether the number of sibling one grows up with (*sibship size*) has a causal effect on home-leaving patterns. Growing up in larger families (which were more prevalent before the seventies) could either speed up or slow down the home-leaving process, depending on two sets of channels. First, sibship size may affect the value of living with parents: indeed, having an extra sibling could increase privacy costs associated with co-residence, but, at the same time, if family ties are strong, an extra sibling could increase the benefit from living in the parental household, leading to a delayed exit. Second, resource dilution resulting from an extra sibling (Becker (1973)) could affect life-course trajectories: on one hand, higher parental investments may lead to better labor market opportunities and faster exit; on the other, it could result in a longer transition to adulthood as a consequence of higher educational achievement and a delayed entry in the labor market.

Macro-level evidence for Europe[2] suggests a possible link between past fertility trends

[1]Throughout the paper, we use the terms *home-leaving* and *youth emancipation* interchangeably, always with the same meaning: we define the age at which an individual *emancipates* as the age in which she/he leaves the parental home in order to start living independently.

[2]We focus on Europe for multiple reasons. First, as noted by Esteve and Reher (2021), it is among the areas of the world where the increase in rates of inter-generational co-residence has been more sustained in the first two decades of this century. Second, despite this aggregate trend, there is sizeable variation across regions, with Southern and Eastern countries being characterized by completely different shares of young adults co-residing with their parents. Third, the availability of two rich, household-level datasets combining information on family structure and

and current changes in the prevalence of inter-generational co-residence: shares of young adults living with parents are on the rise in those European regions where fertility fell the most in the latest decades of the last century. To understand if these macro-level trends are actually connected, we try to establish whether there is a causal link between sibship size and home-leaving decisions at the micro-level.

We combine data from two different cross-national studies, the *Survey of Health, Ageing and Retirement* and the *Generations and Gender Survey*, obtaining detailed information on family structure and on the timing of youth emancipation of children for a large sample of European households over many decades. We exploit random variation in the number of siblings induced by twin births at the second parity to identify the causal effect of having two siblings (as opposed to one) on home-leaving patterns. We combine an instrumental variables approach with survival analysis to estimate the causal effect of an additional sibling on age profiles of home-leaving and on the average emancipation age.

We find that having an additional sibling speeds up home leaving of first-born children, decreasing their expected emancipation age by approximately one year. Exploiting information on educational achievement and employment, we provide evidence that rules out mechanisms based on resource dilution. Therefore, our results are consistent with an extra sibling leading to a fall in the value of co-residing with parents: the increase in costs associated with co-residence (e.g., privacy costs rise when living in a *crowded nest*) offsets the higher benefits from sharing the nest with an additional sibling (e.g., spending time with her). We corroborate this interpretation by performing heterogeneity with respect to the distance between the first and second parity, exploiting age differences as a proxy of *siblings ties*[3]. Consistently with the idea that the *crowded nest* effect resulting from a rise in sibship size is increasingly offset by the direct benefit of living with siblings as ties get stronger, we find that the speeding-up effect of an extra sibling is weaker when age differences are smaller. Our results are heterogeneous by country group, with the effect being strongest for Western European countries and weakest for Southern European countries. We argue that the observed heterogeneity is a consequence of the institutional settings. In Mediterranean countries, family ties and structural economic conditions do not leave room for the marginal contribution of increasing cohabiting costs to be relevant. Similarly, in Nordic

home-leaving processes in European countries is a key resource that makes our identification strategy applicable to the European but not other contexts.

[3]Several papers in demography and sociology find that age differences are a relevant predictor of ties among siblings, in terms of contact frequency, emotional closeness, and practical support. In particular, smaller age differences are systematically associated with stronger ties among siblings (see for example Voorpostel et al. (2007) and Tanskanen and Rotkirch (2019)).

countries, early emancipation seems to be an independent phenomenon from sibship size. Finally, our results suggest that demographic trends, if anything, attenuated the gap in co-residence rates across European regions: in a counterfactual scenario with no differential changes in fertility, regional differences would be even wider, suggesting that strong, heterogeneous structural changes are driving the observed divergence in co-residence rates across Europe.

**Related literature.** Delayed home-leaving has been found to be a determinant of birth postponements, eventually leading to low fertility (Giuliano (2007)). Besides its effects on birth rates, the evidence on the implications of a late transition to adulthood is mixed: on one hand, it is associated with reduced degrees of geographic mobility (Becker et al. (2010)) and lower incomes (Billari and Tabellini (2011)); on the other, studies show that in some countries inter-generational co-residence is positively associated with happiness levels of both parents and children (Manacorda and Moretti (2006)) and that a longer stay in the parental household allows adult children to insure against unemployment shocks (Kaplan (2012)).

As these results showed the relevance of youth emancipation choices, a vast literature emerged aimed at investigating the determinants of home-leaving decisions, in order to understand the causes of such widespread increases in co-residence rates. Existing papers have shown that the timing of youth emancipation is affected by many factors, including parental resources (Manacorda and Moretti (2006), Angelini and Laferrère (2013), Stella (2017)), youths' labor market opportunities (Aassve, Billari, and Ongaro (2001)), housing prices (Modena and Rondinelli (2012)) and credit constraints (Fogli (2004)). Recent research has shown rising co-residence rates are the consequence of structural changes such as labor market uncertainty and the spread of unstable employment (Becker et al. (2010)), the postponement of union formation (Mazurik, Knudson, and Tanaka (2020)), rising house prices relative to income (Cooper and Luengo-Prado (2018)) and shifts in cultural norms (Giuliano (2007)).

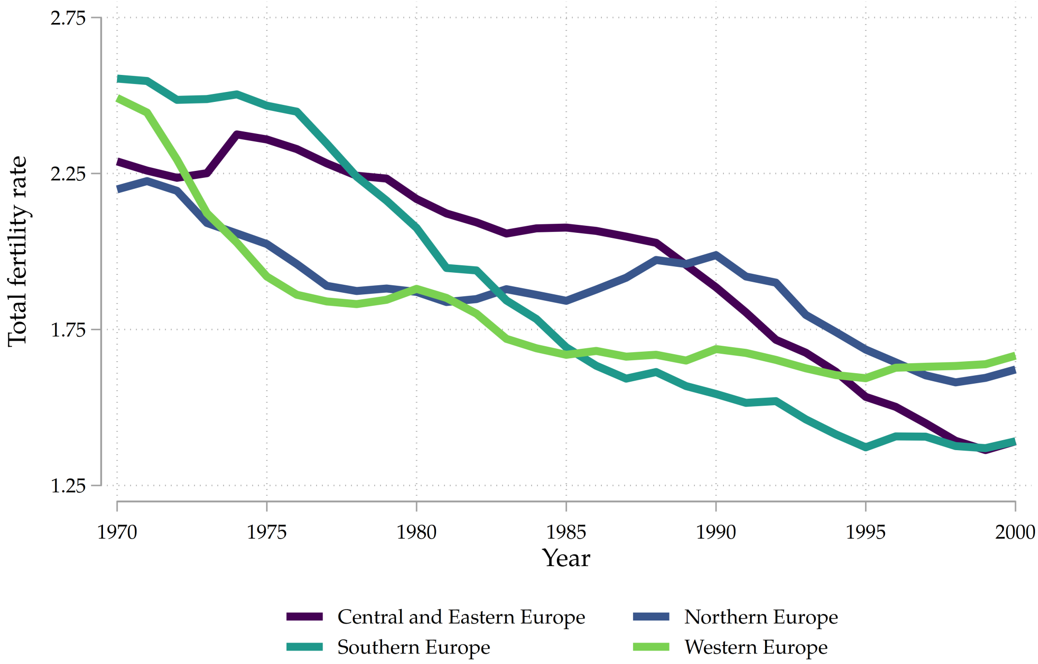
Some studies have instead focused on the effect that different characteristics of the family of origin have on the decision to move out for independent living: examples include the presence of step-parents or step-siblings (Mitchell, Wister, and Burch (1989), Goldscheider and Goldscheider (1998)) and the quality of parent-child and marital relations (Ward and Spitze (2007)). In this paper, we study the role played for home-leaving decisions by a relevant characteristic of the family of origin: the number of siblings an individual grows up with. We are not the first to analyze the relationship between sibship size and emancipation decisions:

however, existing studies (Gierveld, Liefbroer, and Beekink (1991), Ward and Spitze (2007), Chiuri and Del Boca (2010)) only include the number of siblings as a control variable in their empirical models, obtaining mixed results. Other papers study the impact of siblings' choices on nest-leaving decisions: examples include Aparicio-Fenoll and Oppedisano (2016) and Her, Vergauwen, and Mortelmans (2022). However, there are no studies whose main focus is to establish the effect of sibship size on the timing of youth emancipation.

We contribute to the literature in a twofold way. On one hand, we evaluate the causal effect of an important characteristic of the family of origin (sibship size) on home-leaving decisions of daugthers and sons; therefore, we contribute to the literature that studies the determinants of emancipation choices. Second, we find micro-level evidence of a potential macro-level relationship between two demographic variables: past fertility rates and current rates of inter-generational co-residence. We are the first to point out that changes in family structure due to falling fertility rates could have a long-term effect on home-leaving decisions of youths, if sibship sizes influence emancipation ages. Moreover, we contribute to the literature that studies cross-national differences in inter-generational co-residence, as our results suggest that the divergence in co-residence rates across European countries would be even stronger if it wasn't for past demographic trends.

**Structure of the paper.** In Section 2, we provide suggestive evidence that links rates of inter-generational co-residence with past fertility trends across European regions. In Section 3 we describe our data sources and our sample restrictions. In Section 4 we define the estimand, we present our empirical framework and we illustrate our identification strategy. In Section 5 we present our main results and in Section 6 we discuss possible mechanisms. Section 7 concludes.

Figure 1: Fertility rates in Europe, 1970 - 2000.

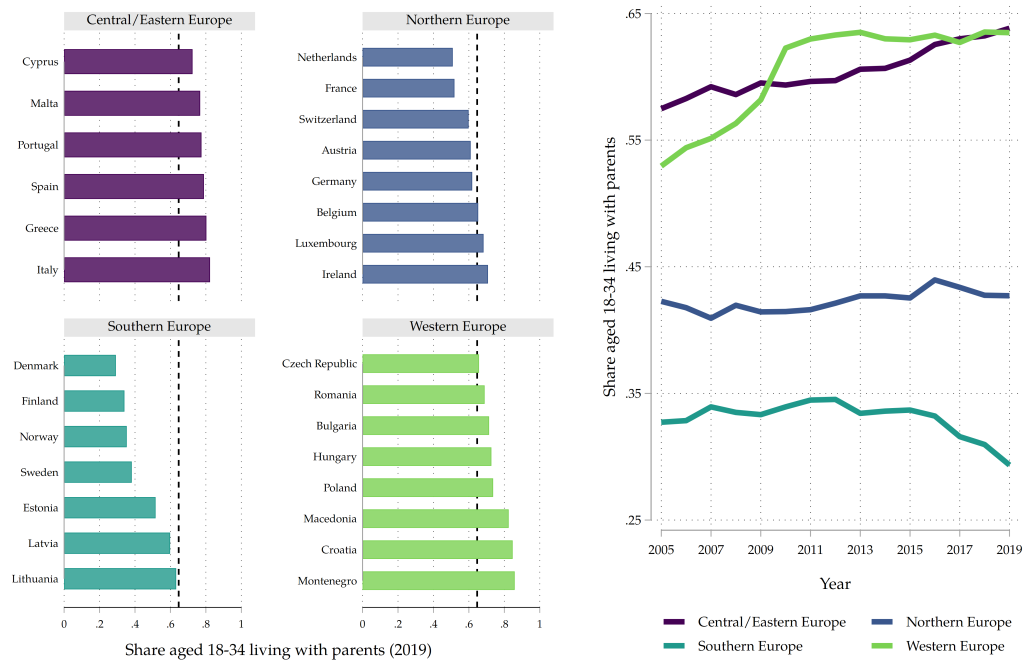



## 2 Fertility and inter-generational co-residence in Europe

As Figure 1 shows, over the last portion of the twentieth century, fertility has been falling across all European regions. Until the mid-seventies, the decrease in fertility was stronger for Western and Northern countries, with the other two regions experiencing little to no plunge in TFR. As fertility kept falling in South and East Europe throughout the 1985-2000 period, TFRs in these country groups ultimately fell below those prevailing in Northern and Western countries, with the two pairs of regions stabilizing on different levels from the early 2000s.

The prevalence of inter-generational co-residence and its evolution over time is also widely heterogeneous across European regions, as Figure 2 illustrates. First, as shown in the left panel of the Figure, the share of young adults living with their parents changes a lot across European countries: in 2019, in countries such as Denmark, Finland, Norway and Sweden, only a minority (less than 40%) of individuals aged 18-34 were still living in their parental home; in Greece, Italy, Spain or Croatia, the vast majority of similarly aged youths (around 80%) are still co-residing with parents. Grouping countries in four regions, one can easily see that co-residence is widely prevalent in Southern and Central/Eastern Europe, while it is much less common in Western and (especially) Northern countries. Second, regional differences in terms of co-residence levels are associated with different trends in the 2005 - 2019 period: as the right panel

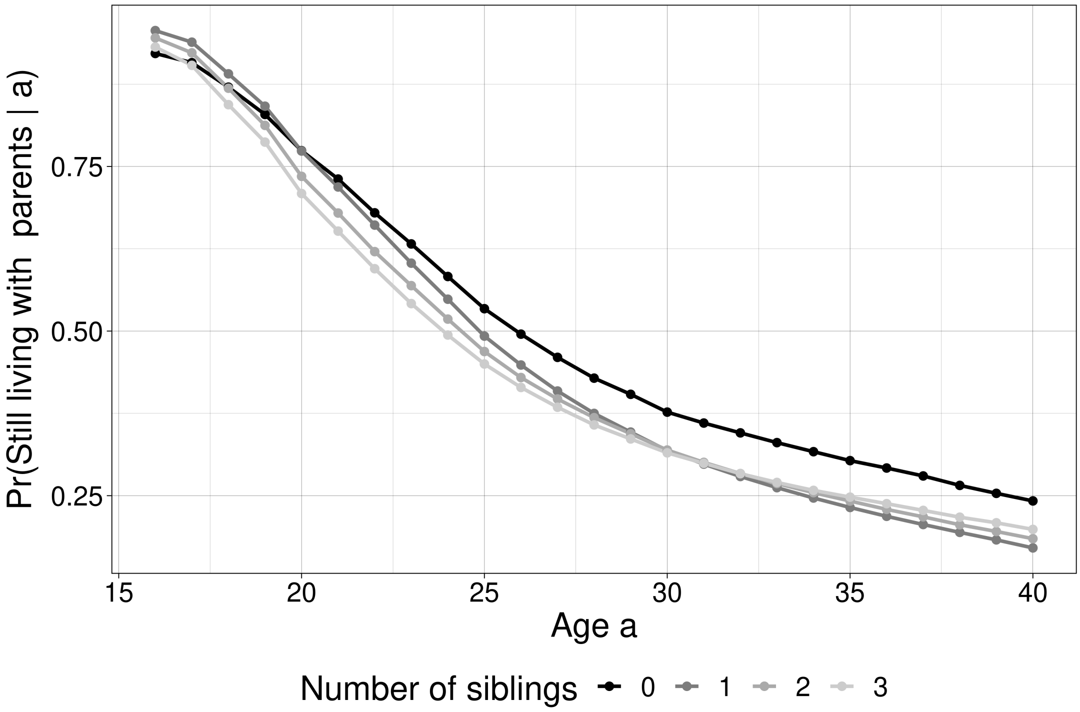
Figure 2: Rates of inter-generational co-residence across Europe, 2015 - 2019.




of the Figure shows, the regions characterized by a slow transition to adulthood (Southern and Central/Eastern Europe) are the ones in which home-leaving patterns are becoming even slower over time; on the other hand, in Nordic countries, where most young adults leave the parental home early, co-residence shares are falling even more. In other words, the gap in home-leaving patterns across European regions is widening.

Comparing Figure 1 and Figure 2, it can be noticed that across Europe, the regions that experienced an increase in the share of young adults living with their parents are those where fertility fell more sharply in the last decades of the twentieth century. This evidence suggests a possible link between demographic trends and changes in home-leaving patterns across European regions: if having an additional sibling accelerates the transition to adulthood, it is possible that the stronger decrease in family sizes experienced by Southern and Central/Eastern European countries is a driving factor behind the widening gap in home-leaving patterns across European regions. Using microdata from SHARE and GGS, in Figure 3, we can look at the association between the number of siblings an individual has and the timing at which she decides to leave the parental home to start living independently. The Figure shows that a negative association exists between sibship size and the timing of home-leaving: only children and individuals with only one siblings leave the parental household later than those coming from larger families.

Figure 3: Sibship size and home-leaving patterns.




As family size is clearly endogenous, this association could very well be spurious, masking a different causal relationship between the two variables. In the next Sections, we will present an empirical analysis aimed at identifying the causal effect of sibship size on the timing of exit from the parental home. Estimating this effect across different country groups will allow us to establish if demographic trends actually contributed to the divergence of home-leaving patterns across European regions.

## 3 Data

We use information from two different data sources: the first seven waves of the *Survey of Health, Ageing and Retirement in Europe* (SHARE), and the first two waves of the *Generations and Gender Survey* (GGS). SHARE is a large-scale, cross-national longitudinal study representative of the population aged 50 or older (and their partners) in 28 European countries. The survey contains a wide array of information on the life course of around 140,000 respondents. Similarly, GGS is a longitudinal survey containing information on around 200,000 adults coming from 19 European countries. Starting from 2004, both surveys collect information on individuals working status, partnership, and fertility, among others.

These sources are particularly suitable for our study as they contain detailed information

Table 1: Summary statistics

| | Mean | SD | Min | Median | Max | *N* |
|---|---|---|---|---|---|---|
| Age | 35.3183 | 11.6652 | 16 | 35 | 70 | 91051 |
| Birth year | 1973.9515 | 11.5885 | 1934 | 1975 | 2002 | 91051 |
| Distance between first and second parity | 3.7796 | 2.5852 | 1 | 3 | 16 | 91051 |
| Emancipated | 0.7504 | 0.4328 | 0 | 1 | 1 | 91051 |
| Emancipation age | 23.1554 | 4.4407 | 16 | 22 | 40 | 91051 |
| Female | 0.4914 | 0.4999 | 0 | 0 | 1 | 91051 |
| N. of siblings | 1.5232 | 0.8181 | 1 | 1 | 5 | 91051 |
| Parent birth year | 1949.0947 | 11.4881 | 1905 | 1950 | 1987 | 91051 |
| Twins at second parity | 0.0116 | 0.1071 | 0 | 0 | 1 | 91051 |

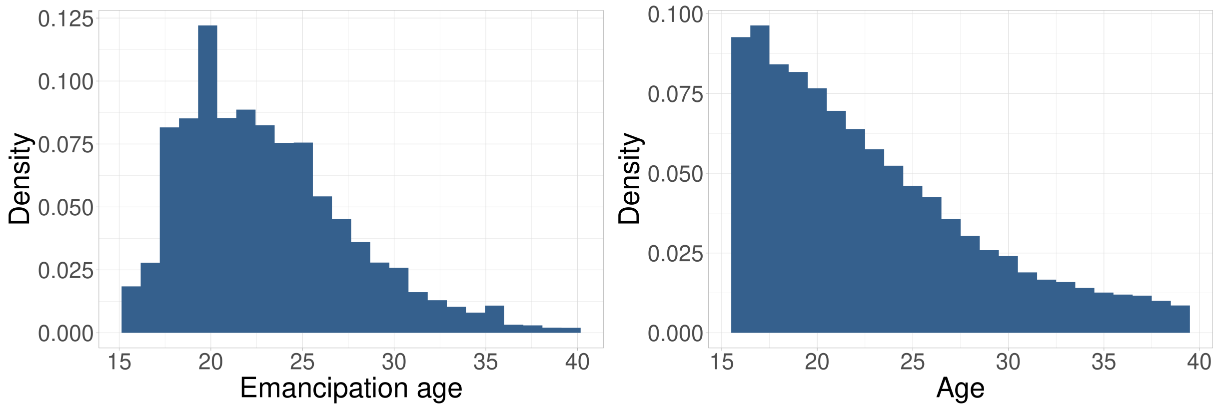
*Note:* Sample of first-born children used in main results.

on all the respondents' children: in particular, in both surveys respondents are asked about the year in which each of their children moved out from the parental house[4]. Interviewers are explicitly asked to instruct respondents to refer only to the last move-out episode, in case an adult child came back home after a failed emancipation attempt (boomeranging). Moreover, temporary departures from the parental home (for example, for a study period or military service) do not count as leaving-home episodes. Family structure is the second key ingredient for our analysis. This is strictly related to our identification strategy relying on twin births. We define twins as individuals born in the same year and born from the same mother. Despite being potentially subject to measurement error (people can have multiple, non-twin births in the same year), we show that our definition leads to twin birth probabilities that are in line with population-level estimates.

**Sample restrictions and summary statistics.** Throughout the analysis, we impose some restrictions on our sample. First, we select couples without foster, adopted, or step-children, and couples whose parent age at childbirth is above 13 and below 70. Moreover, as we focus on an outcome (youth emancipation) which is usually determined in early adulthood, we exclude agents aged under 16. Moreover, we exclude agents whose reported emancipation age is below 14 or above 40, as it is likely that their decision to move out was determined by peculiar circumstances. Further restrictions are motivated by our identification strategy. As detailed in Section 4,

[4]We exploit the panel nature of the surveys to get the most updated information on each child, eliciting the emancipation details from the last wave in which each family appears. For instance, if a family is observed both in 2007 and in 2010, we could note that children still living with their parents in 2007 moved out before 2010. In those cases, the repeated observations enable us to obtain additional information on the timing of emancipation.

Figure 4: Distribution of emancipation age or age at interview (if not emancipated)




(a) Emancipation age

(b) Age at interview, non-emancipated

it is common in studies leveraging the same source of variation to restrict the analysis to non-twin children. The analysis is carried out by exploiting twin births at the second. Therefore, the population of interest is composed of non-twin firstborns in families with at least two children.

In Table 1 we show our summary statistics. The sample is balanced in terms of sex, and the average age is around 36. The probability of experiencing a twin birth at the second parity is around 1.1. Around 75 percent of individuals left the parental home, and the average age at which they did it is around 23. In Figure 4 we display the distribution of home-leaving ages for individuals that had already left home when observed. It appears to be right-skewed, with few individuals emancipating when older than 30.

## 4 Empirical Strategy

As our outcome of interest is the timing of a life-course event, we make use of survival analysis as the main modeling tool. Moreover, survival analysis provides a suitable framework to deal with censored observations, which constitute a sizeable portion of our sample. Censoring may happen for two reasons: either (i) an adult child has not moved out yet when the parent is interviewed (*right-censoring*), or (ii) the adult child has already moved out when the interview takes place, but the responding parent does not precisely remember (or refuses to share) the year in which her child moved out[5].

[5]When information on emancipation status is missing, we are able to derive it by exploiting a key variable available in the data: geographical proximity between the responding parent and each child. The *Proximity* variable can take several values (*In the same household*, *In the same building*, *Less than 1km away*, ... , *More than 500km away*). If it takes a value different from *In the same household*, we code the individual as emancipated.

**Estimands.** Let $a$ denote age and $n$ denote the number of siblings an individual has. The survival function $S(a|n)$ gives the probability that an individual with $n$ siblings is still living in the parental home at age $a$, i.e.,

$$S(a \mid n) = \Pr(T > a \mid n),$$

with the random variable $T$ denoting the home-leaving age. We are interested on the effect of $n$ on home-leaving profiles: our object of interest is the function

$$\Delta S(a \mid n) = \frac{\partial S(a \mid n)}{\partial n} \tag{1}$$

i.e., the derivative of the survival function with respect to the number of siblings $n$. The second estimand of interest is the implied change in the average leaving-the-parental-home age. Let $P(a,n)$ be the probability of leaving home at age $a$ for an individual with $n$ siblings. We have that

$$P(a,n) = S(a,n) - S(a+1,n)$$

Define now as $E(n)$ the average emancipation age for an individual with $n$ siblings. Let $\mathcal{A} = \{\underline{a}, \underline{a}+1, ..., \overline{a}-1, \overline{a}\}$ denote the set of ages at which it is possible to experience the event. We assume that $S(\underline{a}) = 1$ and $S(\overline{a}) = 0$ (everybody leaves home after age $\underline{a}-1$ and before age $\overline{a}+1$).

$$E(n) = \sum_{a \in \mathcal{A}} P(a,n) \cdot a.$$

The second estimand is therefore defined as

$$\frac{\partial E(n)}{\partial n} = \sum_{a \in \mathcal{A}} \frac{\partial P(a,n)}{\partial n} \cdot a = \sum_{a \in \mathcal{A}} \left( \frac{\partial S(a,n)}{\partial n} - \frac{\partial S(a+1,n)}{\partial n} \right) \cdot a \tag{2}$$

**Estimation.** Let $Age_{i,t}$ denote the age of individual $i$ at time $t$ and define $S_{i,t} = \mathbb{1}(T_i > Age_{i,t})$, a dummy variable equal to one if individual $i$ is still living with her parents at age $Age_{i,t}$ and equal to zero otherwise. Moreover, let $NSiblings_i$ represent the number of siblings that individual $i$ has. A natural empirical counterpart of (4) is the linear probability model

$$S_{i,t} = \sum_{j=0}^{k} \beta_{a,j} Age_{i,t}^{j} + \sum_{j=0}^{k} \beta_{n,j} (NSiblings_i \times Age_{i,t}^{j}) + \gamma \mathbf{X}_{i,t} + \epsilon_{i,t}, \tag{3}$$

Table 2: First stage, pooled and by country groups

| Dependent Variable: | Number of Siblings | | | | |
|---|---|---|---|---|---|
| | Full sample | Central and Eastern | Northern | Southern | Western |
| | (1) | (2) | (3) | (4) | (5) |
| Twins | 0.8415*** | 0.8439*** | 0.8838*** | 0.7404*** | 0.8423*** |
| | (0.0233) | (0.0394) | (0.0530) | (0.0462) | (0.0426) |
| N. individuals | 91,057 | 30,406 | 19,030 | 14,634 | 26,987 |

*Note:* The Table shows first-stage estimates of twin births at second parity. The number of individuals refers to children in our sample. Additional controls include country fixed effect, age of the mother and birth year.

where $\mathbf{X}_{i,t}$ is a vector of controls including country fixed effect, age of the mother and year of birth. We allow baseline survival probabilities to depend on a $k$th order polynomial of age, and we adopt a flexible specification in which we allow the effect of an additional sibling to have a different effect at different stages of the early life cycle. Suppose that we restrict the sample to individuals such that $NSiblings_i \in \{\bar{n}, \bar{n}+1\}$ and that we estimate (3). Then, a natural estimator of (1) is given by

$$\Delta \widehat{S(a \mid n)} = \sum_{j=0}^{k} \hat{\beta}_{n,j} a^j \tag{4}$$

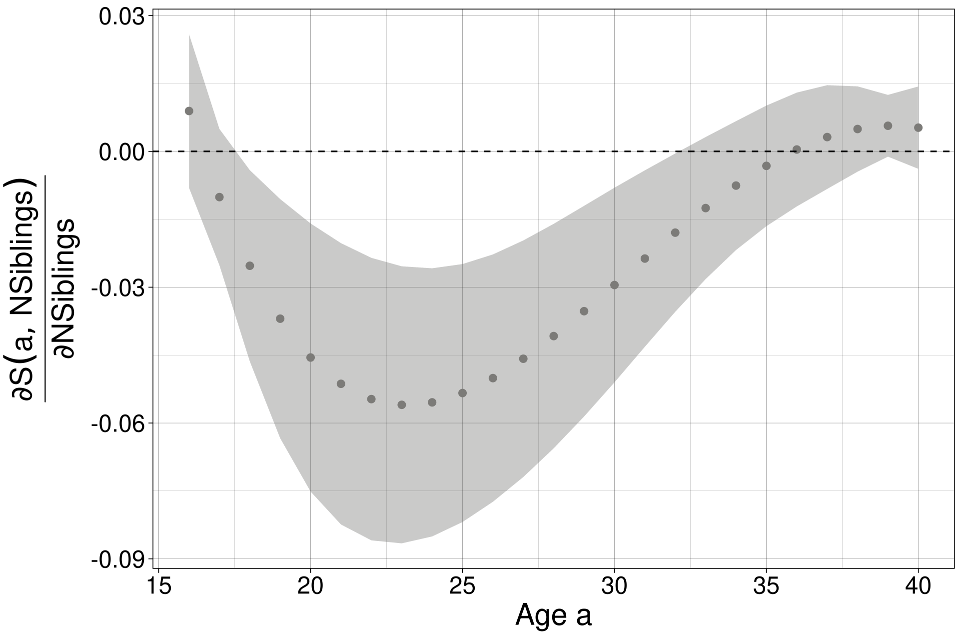
Clearly, a major concern related to the estimation of such a model is the endogeneity of family size. Families with a different number of children are likely to differ in unobserved characteristics we cannot adjust for, which could lead to biased estimates for the causal effect of an extra sibling.

**Identification strategy.** We estimate $\beta_{n,j}$ using a Two Stage Least Squares (2SLS) identification strategy, instrumenting $NSiblings$ with the variable $Twins_i$, which is equal to one if the parents of individual $i$ experience a twin birth at their second parity and equal to zero otherwise. In Table 2 we show estimates from the first-stage equation. The effect is heterogeneous across countries, with southern countries displaying a lower first stage.[6].

Two features of the instrument $Twins$ are particulary attractive for our setting. First, the average causal effect is computed by exploiting fertility variation which induces families to increase the number of children almost uniquely at the parity of twin birth. To be more precise, families with twin birth at second parity, for example, are estimated almost equally likely to

[6] The estimates are slightly larger compared to Black, Devereux, and Salvanes (2005) and Angrist, Lavy, and Schlosser (2010), but comparable to other studies as Angrist and Evans, 1998.

Figure 5: Marginal effects of an extra sibling on survival probabilities




Note: The figure presents estimates of marginal effects as shown in equation **??**. The sample is composed of non-twin first born with at least one sibling. Standard errors are clustered at the individual level. Confidence intervals shown are at the 95% confidence level.

decide to have a child at third parity [7]. Secondly, and as a consequence of the first feature, the local average treatment effect (LATE) for the compliers coincides with the average effect on the non-treated (ATNT), where the treatment is having an additional child[8]. We conjecture that the ATNT, defined for a very large group of individuals in this setting (around 99% of families do not have twins), is arguably close to the average treatment effect (ATE). [9].

## 5 Results

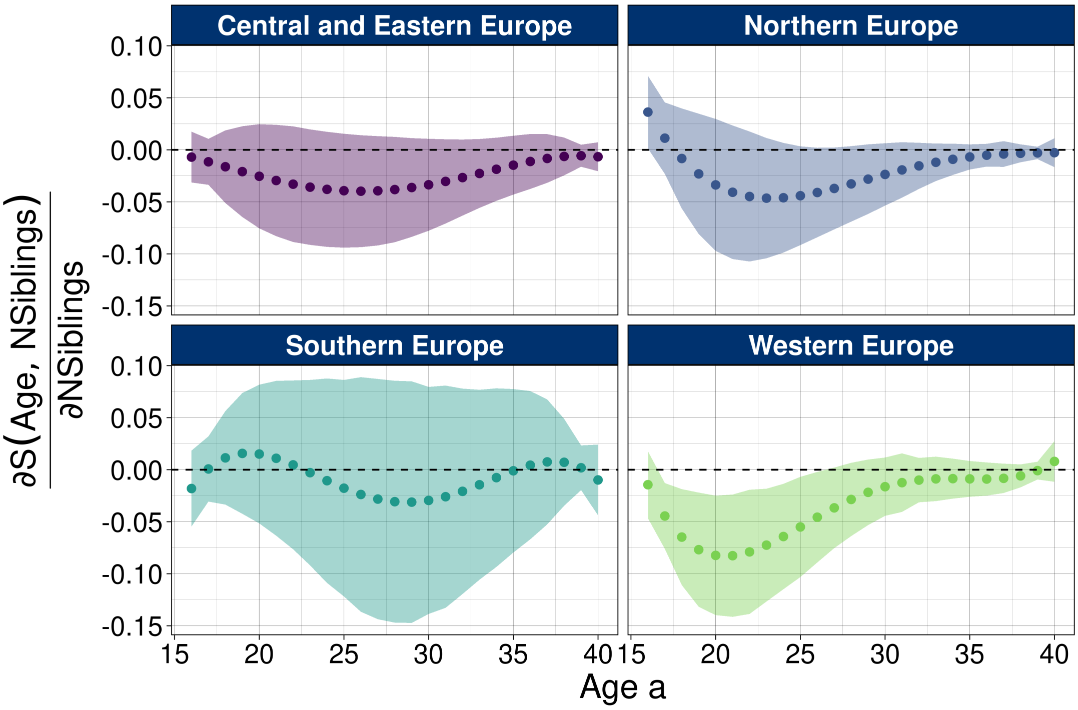
In Figure 5 we plot the marginal effect of sibship size on the probability of still living with parents by age $a$ (equation 1). Having an additional sibling speeds up the emancipation process. In terms of magnitude, the effect is estimated to be the largest around 24 years old, when children are around 5 percentage points (p.p.) less likely to still live with their parents. As stressed in the previous section , estimating differences in survival function allows us to recover the difference in

[7] In **??** we show the first stage equation using samples of couples with at least $n$ children.

[8] The result follows the reasonable assumption of one-sided perfect compliance as a consequence of the absence of never takers (see Angrist Levy (2010) for a discussion). In this setting, never-takers are defined as families who decide to have a child and, despite a twin pregnancy, decide to have only one child.

[9] Instruments relying on endogenous treatment take-up (i.e. gender composition of the first two children) identify the treatment effect for a subset of the population which may be not of particular interest, given the theory we propose connecting sibship size and emancipation age. For example, families may adjust housing size in response to a preference shock implying an additional child, or being more likely to comply if their house is larger, mitigating the crowded nest channel

Figure 6: Marginal effect of an extra sibling, by country groups




Note: The figure presents estimates of marginal effects as shown in equation **??**. The sample is composed of non-twin first born with at least one sibling. Standard errors are clustered at the individual level and shown at the 95% confidence level.

the change in the probability of leaving the parental home at age $a$, and consequently the change in the average age at which individuals with different numbers of siblings emancipate. The estimates shown in Table 3 suggest that having an additional sibling decreases leave-parental-home age by approximately 1 year.

Moving to potential heterogeneous responses by countries, we present the same results conditioning on European regions. Similarly for the evidence presented in Section 1, we categorize countries into groups: Northern, Southern, Western Europeans, and Central/Eastern[10]. Cultural, institutional, and economic factors which are likely correlated with emancipation decisions contribute to this definition[11]. In Figure 6, we plot the marginal effect of having an additional sibling for each of those groups. As we can notice, the average effect described above masks some heterogeneity in responses. More precisely, it seems to be driven by North-West countries mostly. The effect in the Mediterranean countries, even if imprecise, is close to zero

[10]Northern countries include Sweden, Denmark, Estonia, Lithuania, and Norway. Southern countries include Italy, Spain, Greece, and Portugal. Central/ Eastern countries include Czech Republic, Croatia, Hungary, Poland, Slovenia, Bulgaria, and Romania. Finally, Western countries include Belgium, Switzerland, Germany, France, Ireland, Luxembourg, and the Netherlands.

[11]Not surprisingly, there exist large differences in home-leaving decisions pattern between countries also in our sample, as showed in Figure 6, where we plot Kaplan-Meier survival functions by country group.

Table 3: Average effects on emancipation age

| | Average effect | 95% bounds | Baseline |
|---|---|---|---|
| Full sample | -1.02 | [ -1.48 , -0.56 ] | 24.04 |
| Central and Eastern Europe | -0.54 | [ -1.41 , 0.32 ] | 24.49 |
| Northern Europe | -0.22 | [ -1.34 , 0.9 ] | 21.87 |
| Southern Europe | -0.23 | [ -1.76 , 1.31 ] | 27.19 |
| Western Europe | -1.52 | [ -2.49 , -0.55 ] | 23.42 |

*Note:* The Table shows average effects of an extra sibling on emancipation age computed as outlined in Section 4. We exclude ages 16 and 40 from the computation of the average effect because of the instability of the polynomial fit at the boundaries. The bounds for each age are computed by substituting the upper and lower 95% confidence intervals for $\frac{\partial S(a,n)}{\partial n}$ in (2).

Table 4: Effects on emancipation age, breakdown by sex

| | Male | | | Female | | |
|---|---|---|---|---|---|---|
| | Average effect | 95% bounds | Baseline | Average effect | 95% bounds | Baseline |
| Full sample | -1.11 | [ -1.67 , -0.56 ] | 24.98 | -0.79 | [ -1.47 , -0.1 ] | 23.07 |
| Central and Eastern Europe | -0.65 | [ -1.67 , 0.38 ] | 25.62 | -0.25 | [ -1.54 , 1.05 ] | 23.32 |
| Northern Europe | -0.07 | [ -1.65 , 1.51 ] | 22.61 | -0.23 | [ -1.69 , 1.22 ] | 21.12 |
| Southern Europe | -0.79 | [ -2.77 , 1.18 ] | 28.08 | 0.2 | [ -2.28 , 2.68 ] | 26.21 |
| Western Europe | -1.84 | [ -3.02 , -0.66 ] | 24.23 | -1.03 | [ -2.21 , 0.16 ] | 22.58 |

*Note:* The Table shows average effects of an extra sibling on emancipation age computed as outlined in Section 4. We exclude ages 16 and 40 from the computation of the average effect because of the instability of the polynomial fit at the boundaries. The bounds for each age are computed by substituting the upper and lower 95% confidence intervals for $\frac{\partial S(a,n)}{\partial n}$ in (2).

over the entire age profile. In Northern and Central-Eastern the effect is smaller in absolute value. It is worth noticing that the impact picks up around the average emancipation age. In Northern and Western countries it is stronger in the early twenties, while in the Mediterranean and Central-Eastern around 28 and 26 respectively. This evidence suggests that decisions about leaving home are more likely to be influenced by the sibship size around the average age at which individuals decide to live on their own. The difference in magnitudes is easier to grasp by combining this evidence with the estimated impact on the average age at which individuals leave the parental home. In Table 3, we summarize its effect on average, and by country group. In Western Europe, the average age decreases by one and a half years approximately. In Central-Eastern countries, (decreased around 6 months), Northern European countries (decreased by about 2 months), and Southern European countries (decreased by about two months) the average effects are imprecisely estimated.

The effects seem to differ slightly by sex. In Table 4, we show the same results when splitting the sample between males and females. On average, males seem to respond more strongly, namely, they move out faster than females. More precisely, the effect for them is 11 months approximately while 9 for females. Sex differences are more pronounced in Western Europe, where males exit 10 months faster than females in response to an additional sibling.

**Robustness checks.** We test the robustness of the results by relaxing our parametric assumptions regarding the survival function. As shown in Figure 9, assuming a particular functional form for $S(a|n)$ does not modify our results. We estimate, separately for each age $a$ a linear probability model using $S_{i,t}$ as the outcome, estimating a set of $\beta_a$ for each $a$. As we can see, the pattern showed Figure 9 and Figure 5 are fairly similar. The same patterns hold true for each sub-sample analysis carried out so far. Secondly, the results are not sensitive to the exclusion of families with a high number of children. The reader may think our definition of twins may misleadingly include some high-fertility families that decided to have two children over one year. Their exclusion does not modify our results, implying that mis-classification issues are unlikely to matter[12].

## 6 Mechanisms

We conjecture there may be two sets of explanations relating to the delay in emancipation induced by higher sibship size, which we broadly categorize in *resources* and *preferences*. Following the quality-quantity line of reasoning (Becker (1973))[13], families with a large number of children invest less in their human capital accumulation. The effect of higher resources on youth emancipation is not clear ex-ante, as pointed out in the Introduction. Another set of explanations concerns preferences for cohabitation with siblings. Individuals may face higher privacy costs raising from higher sibship size, or forming stronger family ties leading them to accelerate/postpone the leaving-home process. In this section, we aim to test the relative importance of the two channels.

**Resource dilution.** For any SHARE respondent (around 33.000 individuals), we have information on marital status, labor market outcomes, fertility, and education at the moment of the interview. In Table 5 we show 2SLS estimates of the effect of a number of siblings on a number of children, an indicator for being married, employed, and having a university degree and

[12] We carried the analysis excluding families with more than 7, 8, 9, and 10 children

[13] The twin instrument has been broadly used to test the existence of a quality-quantity trade-off, see for example Black, Devereux, and Salvanes (2005), Angrist, Lavy, and Schlosser (2010) and Bagger et al. (2021)

Table 5: Effect of Number of Siblings on Marital, Labor, Educational outcomes

| | | Marriage | Employment | Fertility | University | High school |
|---|---|---|---|---|---|---|
| Full Sample | | 0.031 | -0.029 | -0.026 | -0.017 | 0.015 |
| | | (0.024) | (0.025) | (0.067) | (0.03) | (0.019) |
| | Mean outcome | 0.746 | 0.816 | 1.400 | 0.393 | 0.863 |
| | N. observations | 33001 | 32898 | 32569 | 33062 | 32576 |
| Western Europe | | 0.029 | -0.036 | -0.131 | -0.041 | 0.014 |
| | | ( 0.036 ) | ( 0.037 ) | ( 0.097 ) | ( 0.044 ) | ( 0.027 ) |
| | Mean outcome | 0.73 | 0.831 | 1.347 | 0.439 | 0.891 |
| | N. observations | 13802 | 13744 | 13630 | 13837 | 13583 |
| Central and Eastern Europe | | 0.024 | -0.037 | 0.088 | 0.042 | 0.054* |
| | | ( 0.046 ) | ( 0.056 ) | ( 0.129 ) | ( 0.064 ) | ( 0.029 ) |
| | Mean outcome | 0.807 | 0.807 | 1.52 | 0.31 | 0.883 |
| | N. observations | 7052 | 7033 | 6996 | 7065 | 7008 |
| Northern Europe | | 0.031 | -0.005 | -0.231 | -0.044 | 0.038 |
| | | ( 0.058 ) | ( 0.051 ) | ( 0.159 ) | ( 0.068 ) | ( 0.034 ) |
| | Mean outcome | 0.709 | 0.84 | 1.62 | 0.438 | 0.9 |
| | N. observations | 5736 | 5718 | 5665 | 5747 | 5701 |
| Southern Europe | | 0.032 | -0.024 | 0.336 | -0.012 | -0.076 |
| | | ( 0.067 ) | ( 0.078 ) | ( 0.21 ) | ( 0.081 ) | ( 0.077 ) |
| | Mean outcome | 0.744 | 0.774 | 1.162 | 0.347 | 0.744 |
| | N. observations | 6411 | 6403 | 6278 | 6413 | 6284 |

*Note:* The Table shows 2SlS estimates of the effect of the number of siblings on Marital, Labor and Educational outcomes. The sample is composed of firstborns of families with at least two children, interviewed in SHARE. Heteroskedasticity-robust standard-errors in parentheses. Signif. Codes: ***: 0.01, **: 0.05, *: 0.1

Table 6: Effects on emancipation age, breakdown by parity spacing

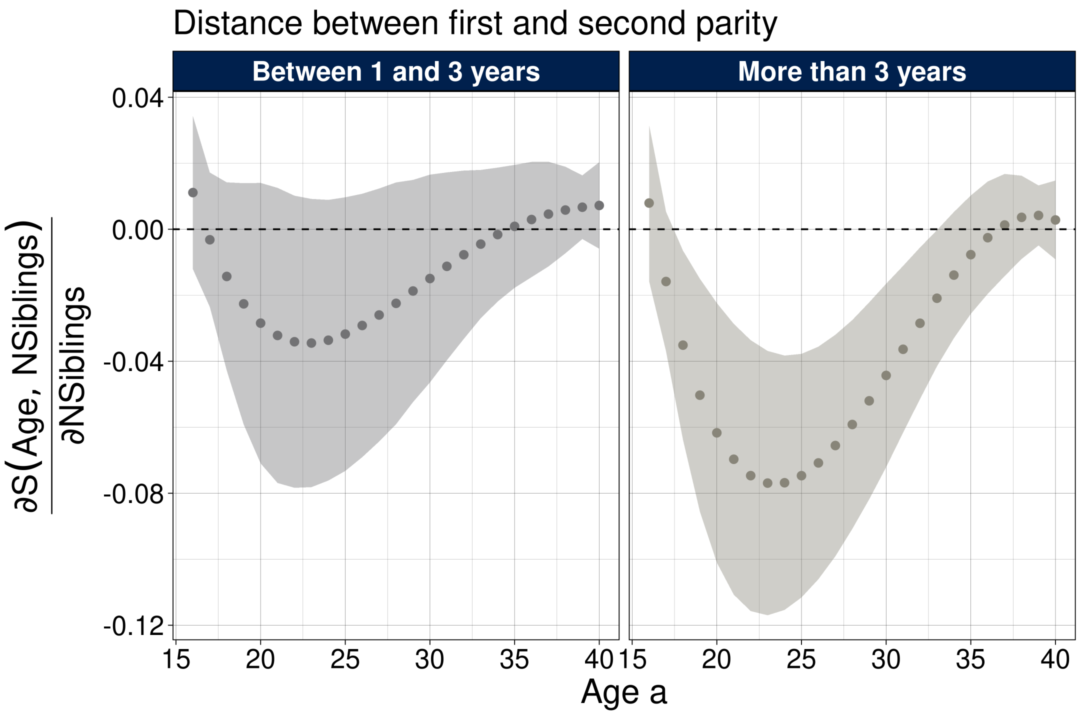
| | Between 1 and 3 years | | | More than 3 years | | |
|---|---|---|---|---|---|---|
| | Average effect | 95% bounds | Baseline | Average effect | 95% bounds | Baseline |
| Full sample | -0.66 | [ -1.31 , -0.02 ] | 23.94 | -1.35 | [ -1.97 , -0.72 ] | 24.18 |
| Central and Eastern Europe | 0.54 | [ -0.56 , 1.63 ] | 24.45 | -1.38 | [ -2.36 , -0.39 ] | 24.54 |
| Northern Europe | -0.04 | [ -1.7 , 1.61 ] | 21.8 | -0.32 | [ -1.62 , 0.97 ] | 21.96 |
| Southern Europe | 0.13 | [ -1.73 , 1.98 ] | 26.77 | -0.92 | [ -2.58 , 0.73 ] | 27.69 |
| Western Europe | -1.14 | [ -2.27 , -0.01 ] | 23.54 | -1.71 | [ -3.16 , -0.26 ] | 23.17 |

*Note:* The Table shows average effects of an extra sibling on emancipation age computed as outlined in Section 4. We exclude ages 16 and 40 from the computation of the average effect because of the instability of the polynomial fit at the boundaries. The bounds for each age are computed by substituting the upper and lower 95% confidence intervals for $\frac{\partial S(a,n)}{\partial n}$ in 2

high-school diploma. As in the previous section, we show results aggregating all countries and by country group. The estimates point out that first-born with an additional sibling induced by twins being born at second parity do not show systematic differences with respect to the ones with one sibling. The only coefficient appearing to be different from zero at the conventional level is the probability of having a high-school diploma for Central/Eastern individuals, and its sign contradicts the prediction of the theory. We want to stress that our goal is not testing the quality-quantity theory per see, an exercise that has been extensively performed in different contexts and with more appropriate data. The purpose is to rule out that labor market, educational and marital decisions, which have been shown to be influenced by the sibship size, are not the only or main mediators of the effect on youth emancipation.

**Crowded nest and sibling ties.** We indirectly test whether the value of cohabiting with parents decreases due to an increase in the number of siblings. Under the assumption that the value of cohabitation with a sibling is decreasing in their age distance, we perform the same analysis carried so far focusing on the heterogenous effect by this proxy of *sibling ties*. If individuals trade-off costs (for example due to a *crowded nest*) and benefits from cohabiting with siblings when deciding to leave the parental home, we would expect the results to be more pronounced with siblings being born more years apart. Figure 7 corroborates this theory by showing faster exit induced by an additional sibling for individuals with, relatively to them, younger siblings. In particular, if individuals are exposed to random variation in the number of siblings after they were three years old, they decide to leave their parental home 14 months earlier on average, contrary to the 7 months in the opposite case. The effect doubles on average, and it is consistent across country groups. The shape of the survival function shown in Figure 7 is similar, but the

Figure 7: Marginal Effects Full Sample, by distance from second-born sibling




Note: The figure presents estimates of marginal effects as shown in (**??**). The sample is composed of non-twin first born with at least one sibling. The left panel shows estimates for first born children who are 1 to 3 years older than their second born sibling, while the right panel shows the same results for first born children who are at least 4 years older than their nearest sibling. Standard errors are clustered at the individual level. Confidence intervals shown are at the 95% confidence level.

probability change estimates are consistently larger for every age if the age distance between siblings is larger. The results are similar if we change the age distance threshold defining different *siblings ties*.

## 7 Conclusion

In this paper, we ask whether past fertility trends and their impact on family structure may have an effect on the home-leaving choices of young individuals. Exploiting random variation in sibship size induced by twin births, we estimated the causal effect of having an additional sibling on the timing of home-leaving in Europe over the last six decades of the twentieth century, combining two datasets that contain rich information on family structure and youth emancipation. Our results indicate that having more siblings speeds up the transition to independent living. We provide evidence that rules out a resource dilution mechanism à la Becker (1973), and we interpret our findings as evidence that when the number of siblings increases, the value of co-residence with parents goes down. We indirectly test this mechanism by performing a heterogeneity analysis with respect to the age distance between the first and second parity within each household, which has been shown to be a proxy for *siblings ties*: effects are stronger when siblings are born more years apart, consistently with individuals taking into account costs and benefits of co-residing with parents when deciding over their living arrangements. Therefore, our findings indicate that having an extra sibling increase the costs of co-residence, possibly through a *crowded nest* effect: as the house gets more crowded, privacy costs rise. We furthermore document response heterogeneity by groups of countries, with Western countries displaying the largest responses. Our findings suggest that demographic trends may have attenuated the gap in co-residence rates in European countries.

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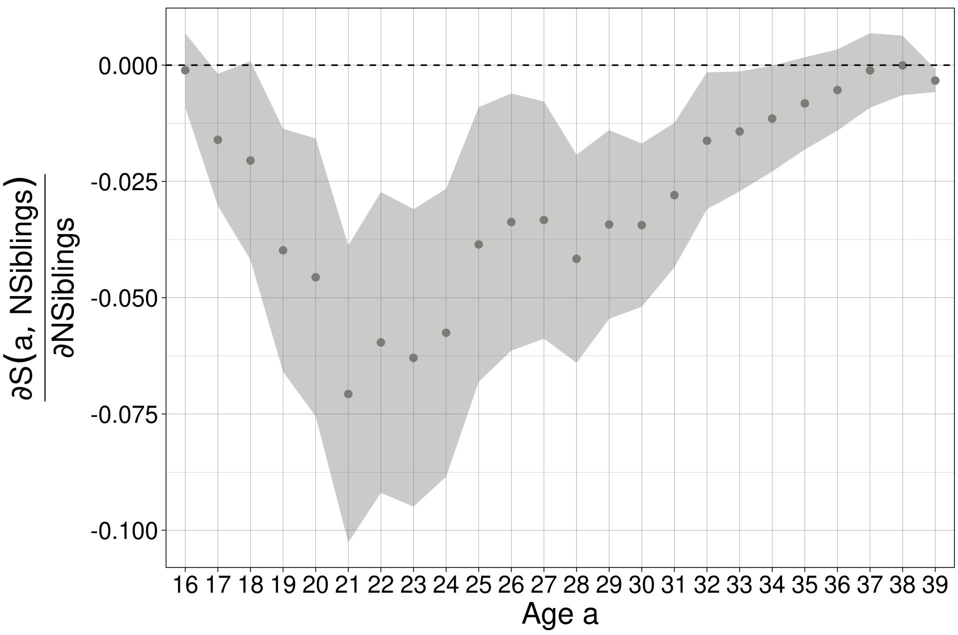
# A Appendix

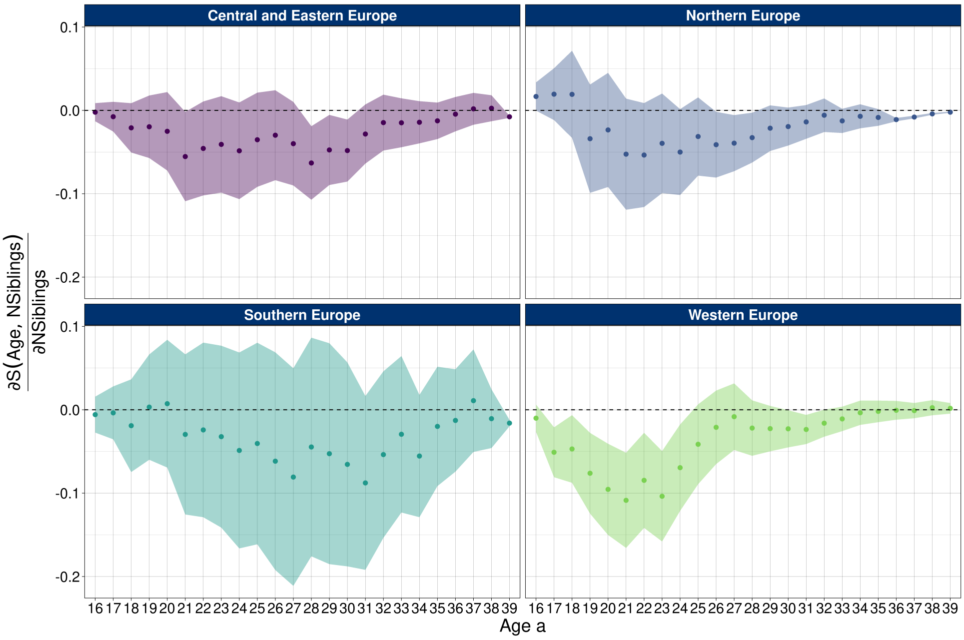
## A.1 Additional Figures

Figure 8: Marginal effect of an extra sibling, non-parametric



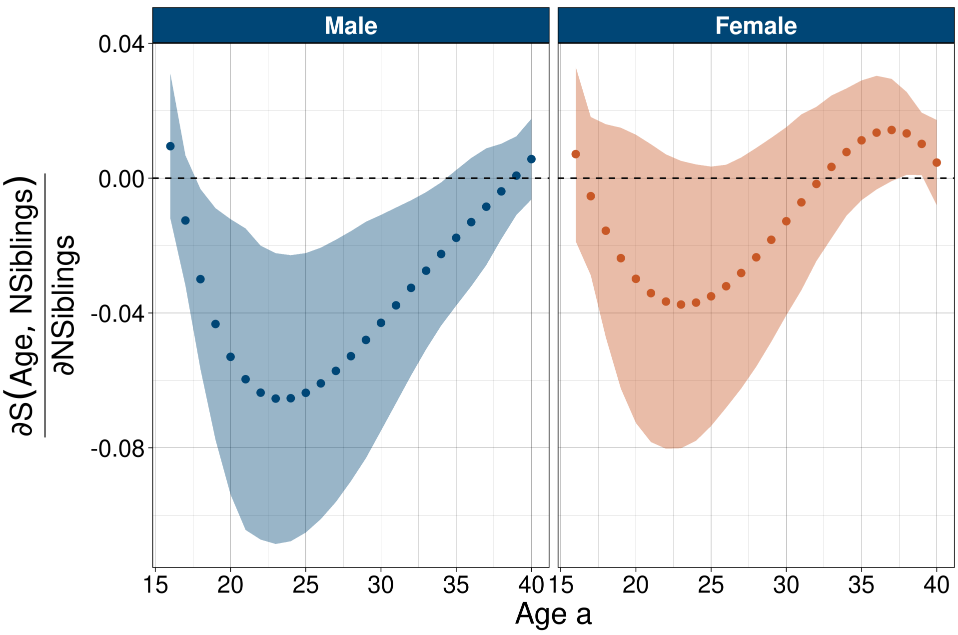

Note: The figure presents estimates of a linear probability model of living with parents by age $a$ using twin birth at second parity as an instrument for the number of siblings. The sample consists of individuals who moved out at most at age 40. The model is estimated seperately for every $a$.

Figure 9: Marginal effects by country group, non-parametric



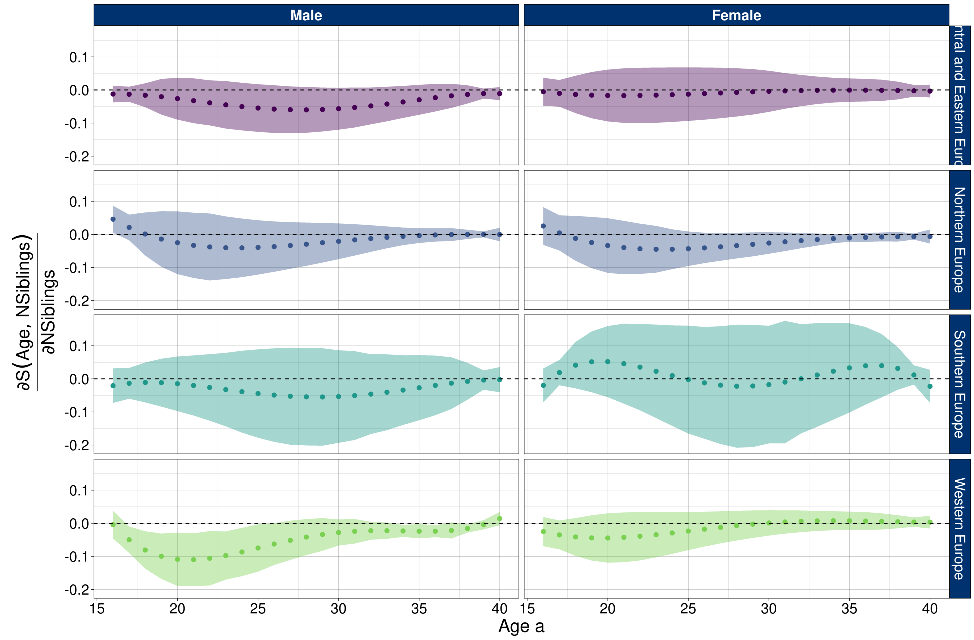

Note: The figure presents estimates of a linear probability model of living with parents by age $a$ using twin birth at second parity as an instrument for the number of siblings. The sample consists of individuals who moved out at most at age 40. The model is estimated seperately for every $a$.

Figure 10: Marginal effects, by sex



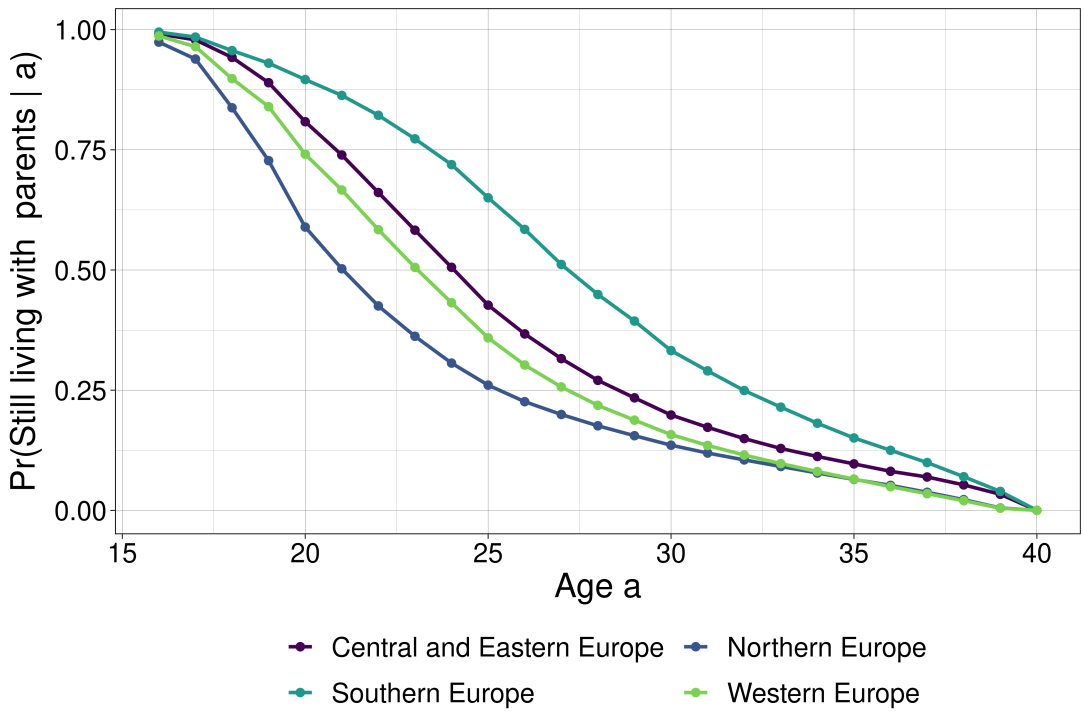

Note: The figure presents estimates of marginal effects as shown in equation **??**. The sample is composed of non-twin first born with at least one sibling. Standard errors are clustered at the individual level and shown at the at the 95% confidence level.

Figure 11: Marginal effects, by country group and sex



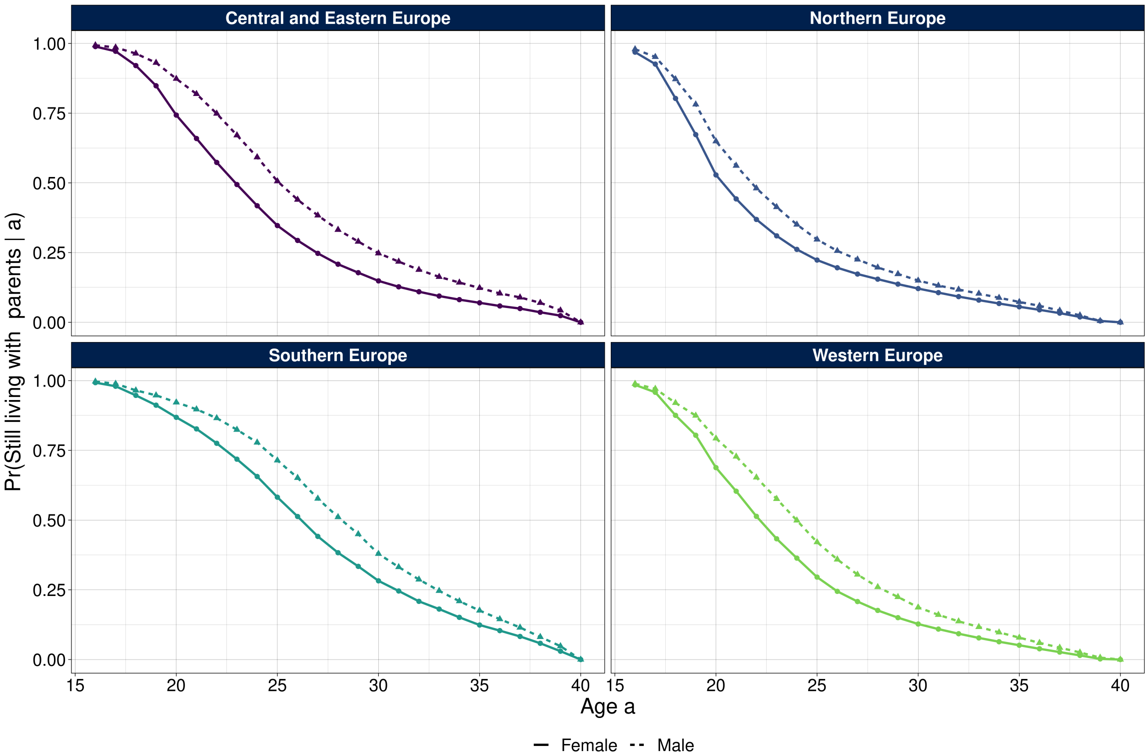

Note: The figure presents estimates of marginal effects as shown in equation **??**. The sample is composed of non-twin first born with at least one sibling. Standard errors are clustered at the individual level, and shown at the at the 95% confidence level.

Figure 12: Kaplan-Maier survival functions, by country group




Note: The figure presents estimates of the proportion of individuals still living with parents by age $a$. The sample consists of individuals who moved out at most at age 40.

Figure 13: Kaplan-Maier Survival Function by Country Group and Sex




Note: The figure presents estimates of the proportion of individuals still living with parents by age $a$. The sample consists of individuals who moved out at most at age 40.