

The Long-Term Consequences of a Golden Nest: Socioeconomic Status in Childhood and the Age at Leaving Home

Viola Angelini, Marco Bertoni, and Guglielmo Weber

ABSTRACT The age at leaving the parental home has significant implications for social and economic outcomes across the life course, highlighting the importance of examining nest-leaving patterns. We study the role of childhood standard of living on the age at nest leaving. Using data from the Survey of Health, Ageing and Retirement in Europe (SHARE), we show empirically that individuals who grow up in families with a higher socioeconomic status—that is, in a golden nest—leave the parental home later than others. Given that better-off individuals tend to obtain more education, and that young adults generally leave the parental home after completing their education, we also find that a higher level of education delays nest leaving. Nonetheless, the positive relationship between socioeconomic status and nest-leaving age still holds for given education levels, across European countries characterized by different cultural traits, for both males and females, and among urban and rural residents. We use a three-period life cycle model to show that this behavior is consistent with standard assumptions about preferences and resources if earnings increase with age. Moreover, habit-forming preferences that assume that utility depends on the gap between current and past consumption reinforce the delaying effect of a golden nest on nest leaving.

KEYWORDS Nest leaving • Socioeconomic status • Habit-forming preferences • Intergenerational mobility • SHARE

Introduction

Why do some children leave the parental home as soon as possible, whereas others delay their nest leaving to a later age? Is the nest-leaving decision driven by purely cultural factors, or do economic considerations play a role?

Answering these questions is important because the nest-leaving age matters for economic and social outcomes throughout the life course. For example, Billari and Tabellini (2010) found that Italians who leave the parental home earlier in life have higher incomes in their mid-30s. When coupled with late family formation and late entrance into the labor market, late nest leaving may be associated with lower or

postponed fertility and a shorter working life, with negative consequences for the sustainability of pension systems as well as for individual pension wealth accumulation.

Whereas the sociodemographic literature has stressed the role of cultural factors in explaining nest-leaving patterns (e.g., Aassve et al. 2002; Billari et al. 2001; Kiernan 1986), the economics literature has investigated the importance of financial resources (for a review of the relevant literature, see Cobb-Clark 2008). Manacorda and Moretti (2006) found that Italian parents prefer to coreside with their children and, therefore, may reward them financially for staying in the parental home: a \$500 increase in parental income is associated with a 3.5- to 3.9-percentage-point increase in the probability that adult children will live with their parents. As the authors noted, however, this result is unlikely to hold for countries in Central and Northern Europe, which have different cohabitation preferences (Giuliano 2007). Another strand of the economics literature has examined the role of the younger generation's limited access to the housing market (Alessie et al. 2006; Guiso and Jappelli 2002). Capital and housing market imperfections may indeed delay nest leaving, but this effect might be stronger for children who grew up in poorer families, who cannot rely on their parents for help finding suitable accommodations. In fact, Cobb-Clark and Gørgens (2014) showed that wealthier parents are more likely to offer housing support and financial gifts to their young adult children.

In this study, we focus on the role of childhood standard of living on the nest-leaving decision. We use data from the Survey of Health, Ageing and Retirement in Europe (SHARE) on a representative sample of individuals aged 50 or older living in one of 28 European countries or Israel. The data contain retrospective information on socioeconomic conditions at age 10, age at nest leaving, and other major life events. Our interest lies in the relationship between the age at nest leaving and socioeconomic conditions at age 10. We provide strong evidence of a positive relationship between socioeconomic conditions at age 10 and the age at leaving the parental home. The size of the effect differs somewhat across genders and cultures, but the sign and significance of the key estimated coefficient remain the same across different groups. Looking at mechanisms, we find that education only partially mediates the effect of a golden nest on leaving the parental home.

We also show that a standard, stripped-down version of the life cycle model with three periods can easily explain why grown, working children of affluent parents leave home later. We assume that grown children who live at the parental home while working pass their earnings on to the parents and have the same standard of living (consumption) as in their childhood. We also assume that the age profile of earnings is upward sloping, so that income during the first period of the model is lower than the consumption level in the parental home. If the children move out, they can use their current and future earnings to smooth consumption over the whole life cycle. In this context, we show that grown children will move out sooner if the standard of living at the parental home is lower. However, if the parental home is sufficiently attractive in terms of consumption, they will delay home leaving. We also add habit-forming preferences—that is, a specification of the utility function that assumes that current utility depends on the gap between current and past consumption—to our nest-leaving model (see, e.g., Angelini 2009; Diaz et al. 2003) and show that this specification reinforces the dependence of the nest-leaving decision on the childhood consumption level.

Our empirical contribution lies in showing that children of high socioeconomic status (SES) leave home later in many countries irrespective of their culture and that this effect is only partly due to education. Our theoretical contribution consists in

providing evidence that late nest leaving can be explained as a combination of a taste for consumption smoothing (maintaining standards of living over time) and access to economic resources (parents' current income on the one hand and children's lifetime income processes on the other). This induces children who grow up in high-SES families to leave the parental home later, in line with what we observe in the data.

Data

We use data from the Survey of Health, Ageing and Retirement in Europe, a biennial survey that collects information on Europeans aged 50 or older. As of this writing, the survey has collected eight waves of data. We use data from the third and the seventh waves of SHARE release 7.1.0 containing retrospective life course information on all respondents. These data, known as "ShareLife data," cover 28 European countries plus Israel. They contain information on several indicators of the standard of living at age 10, including the number of rooms per capita, the number of books at home, the household's number of amenities (a fixed bath, a cold/hot running water supply, an inside toilet, and central heating), and the occupation of the household's main breadwinner. In addition, information on whether the area of residence at age 10 was urban or rural is also available. Furthermore, the data also report the year each individual left the parental home. This information allows us to investigate age differences at nest leaving.

The initial sample includes all individuals who completed a ShareLife interview (91,774 observations) in Wave 3 or, if they were not yet part of the survey at that time, Wave 7. We keep only those born in 1936–1956 to avoid issues of selective mortality and to obtain a stable gender balance across cohorts (58,922 observations).

We omit 4,528 individuals who were not born in the current country of residence to avoid issues of endogenous mobility, reducing the sample size to 54,394. We also omit records with missing or implausible information on nest-leaving age (453 with missing information and 330 with implausibly low nest-leaving ages of younger than 14) or with missing or implausible information on age at first cohabitation (three observations with missing information and 394 with implausibly low ages of younger than 14). The sample size after all these selections is 53,214.

We further omit 10,865 individuals with missing information on covariates about childhood; the binding constraint is parental occupation. Because the SES distribution has a long but very thin right tail, we eliminate outliers on this variable using Tukey's (1977) proposed criteria.¹ This restriction results in omitting 165 observations in the right tail of the distribution and no observations in the left tail. The final sample contains 42,184 observations.²

Following Mazzonna (2014), we construct a general index for childhood SES using information on individuals' standard of living at age 10. This method is standard among researchers using SHARE data, given that the survey minimizes recall

¹ We define Q1 and Q3 as the 25th and 75th percentiles of the distribution of SES. We omit observations with an SES value above $Q3 + (Q3 - Q1)$ and below $Q1 - (Q3 - Q1)$. The results are unchanged if we instead include them.

² In Table A1 in the online appendix, we show the geographical distribution of the estimation sample and whether a country is present in Wave 3 only, Wave 7 only, or both. For respondents who participated in both Waves 3 and 7, the retrospective life history questions are not asked again in Wave 7.

Table 1 Descriptive statistics

Variable	Number of Observations	Mean	SD
SES at Age 10	42,184	0	1
Left Home by Age 49	42,184	.98	.15
Age at Nest Leaving	41,252	23.13	4.82
Cohabited With a Partner by Age 49	42,184	.96	.19
Age at First Cohabitation	40,527	23.88	4.54
Female	42,184	.55	.50
Year of Birth	42,184	1947	5.76
Lived in Rural Area at Age 10	42,184	.47	.50
Never Had Siblings	42,184	.17	.38
Lived Without at Least One Parent at Age 10	42,184	.10	.31
Years of Education	42,184	11.00	4.07
Present Only in Wave 3	42,184	.18	.38
Present Only in Wave 7	42,184	.59	.49

Note: The omitted wave group is *present in both waves*.

bias by not asking respondents to provide information on parental education or income when they were young, two standard measures of family background used in the literature on intergenerational transmission (see Björklund and Jäntti 2012).³ We construct our index using polychoric principal component analysis to extract the first component from the following four proxies of SES: (1) occupation of the main breadwinner (low-tier occupation or higher, according to the one-digit ISCO-88 coding); (2) the number of rooms per capita (excluding toilets and kitchens); (3) the number of books at home (at least one bookcase); and (4) the number of household amenities (a fixed bath, a cold/hot running water supply, an inside toilet, and central heating). Given that some of these variables are related to institutional factors specific to individual countries, such as running water or central heating for individual dwellings, we de-mean each variable within countries before computing the index.

Consistent with Mazzonna (2014), we obtain only one principal component with an eigenvalue above 1, which explains more than 40% of the total variance. The signs of the scoring coefficients are also consistent: we estimate a negative loading for low parental occupation and positive loadings for the other variables (details are available from the authors). As shown in Table 1, which displays descriptive statistics for the key variables used in this study, we standardize the resulting index to have a zero mean and unit standard deviation in the full sample. The distribution of the index for the full sample, displayed in Figure 1, is right-skewed and has a very granular, nearly continuous support, thereby providing a rich description of the variability in SES. Table A2 in the online appendix reports variation in the index in each country.

Our dependent variable is the age at nest leaving, obtained from responses to the following retrospective question: “In which year did you start to live on your own or

³ Although information on parental education was collected in two standard waves (Waves 5 and 6), it is missing for our sample respondents: (1) those who participated in Wave 3 but not in Waves 5 and 6, and (2) respondents living in countries that joined SHARE in Wave 7.

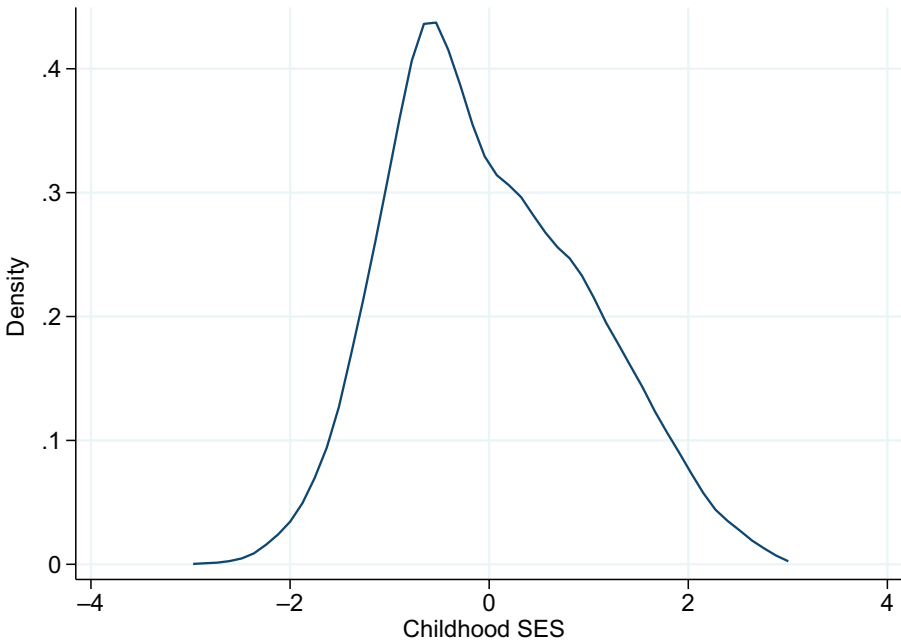


Fig. 1 Kernel density estimate of the distribution of the SES indicator in the full sample. An Epanechnikov kernel is used, with a bandwidth of 0.1070.

establish your own household?” Respondents could provide the year or indicate that they never established their own household. The youngest interviewees in SHARE are aged 50. Thus, to avoid mechanical age effects, we cap nest-leaving age at 49 and recode the answers of older subjects who reported establishing their own household at later ages as “not left by 49.” Although nest-leaving age could suffer from recall bias, measurement error in the dependent variable should not lead to biased estimates. As shown in Table 1, 2% of our final sample had not left home by age 49. In addition, the average nest-leaving age conditional on nest leaving is 23.13 years (22.08 for females and 24.40 for males; not shown). Figure 2 displays the average nest-leaving age by country, which ranges from 20 to 25 years and exhibits a stark North–South gap.

As an alternative outcome variable, we use age at first cohabitation with a partner. Ninety-six percent of our sample had cohabited with a partner by age 49, and the conditional average age at first cohabitation is approximately 24 years. In addition, roughly one fourth of nest leavers did so as singles, and the remaining three fourths left the parental home to start cohabiting with their first partner (not shown). Other variables used in our analysis are gender (55% of our sample is female), year of birth (the average respondent was born in 1947), lived in a rural area at age 10 (47%), never had siblings (17%), and lived without at least one parent at age 10 (10%). We also control for self-reported math and language ability at age 10, expressed in terms of being much better, better, about the same, worse, or much worse than one’s classmates; for both subjects, roughly 50% reported being about the same as their peers, but less than 15% reported being worse or much worse than their peers (not shown). In some of our analyses, we also explore the mediating effect of educational achievement. Our

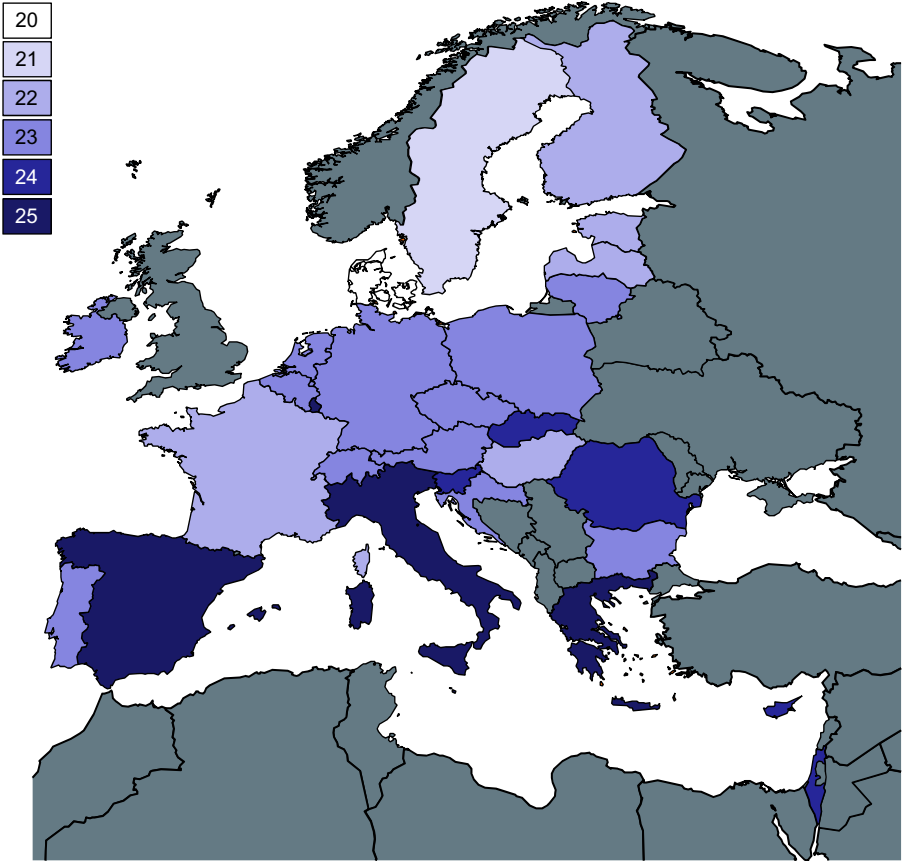


Fig. 2 Distribution of the average age at nest leaving across countries

sample completed an average of 11 years of education.⁴ Table 1 also reports the distribution of the sample by wave: 18% of the sample was present only in Wave 3, and 59% of the sample participated only in Wave 7. The remaining 23% of respondents participated in both waves (but life history information was collected only in Wave 3).

The Effect of SES on Nest-Leaving Age

Main Estimates

The thought experiment we would like to conduct to identify the effect of SES on nest leaving would be to randomly allocate at birth two identical twins to two families with different levels of SES and compare their nest-leaving decisions. We mimic

⁴ Roughly 11% of our sample left the parental home before completing education. Approximately 75% of these individuals have a tertiary degree, and 60% are women. Only 6% of individuals started cohabiting with a partner before finishing education. The composition by gender and educational achievement is very similar. These data are not shown.

Table 2 OLS regressions of nest-leaving age

	(1)	(2)	(3)
SES at Age 10	0.269*** (0.024)	0.274*** (0.024)	0.276*** (0.024)
Mean Dependent Variable	23.13	23.13	23.13
Number of Observations	41,252	41,252	41,252
Individual-Level Controls	Yes	Yes	Yes
Cohort Fixed Effects	Yes	Yes	No
Country Fixed Effects	Yes	Yes	No
Country-Specific Cohort Trends	No	Yes	No
Country × Cohort × Gender Fixed Effects	No	No	Yes

Notes: Column 1 includes the individual-level observable controls and country and cohort dummy variables. Columns 2 and 3 include additional variables: country-specific linear cohort trends in column 2 and the interaction among year of birth, country, and gender dummy variables in column 3. Each column reports the coefficient on SES at age 10 from a different OLS regression. SES at age 10 is standardized to have zero mean and unit standard deviation in the full sample. All models include the following individual-level controls: dummy variables for gender, self-rated math and language ability, residence in a rural versus urban area at age 10, never had siblings, lived without at least one parent at age 10, and wave dummy variables. The fixed effects and trends indicated at the bottom of each column are also included. The estimation sample excludes 932 individuals who had not established their own household by age 49. Heteroskedasticity-robust standard errors are shown in parentheses.

*** $p < .001$

this experiment with our observational data by estimating ordinary least-squares (OLS) regressions of the age at nest leaving on childhood SES, controlling for interview wave (coded as dummy variables), living in a rural area at age 10, self-reported ability in math and language at age 10 (coded as dummy variables), never had siblings, and lived without at least one parent at age 10. We also include gender, year of birth, and country dummy variables. Controlling for rural residence and gender is important because age at nest leaving varies across these characteristics: individuals from rural areas leave the nest about one half year later than those from urban areas, and females leave the parental nest roughly two years sooner than males. Including measures of math and language ability helps assuage concerns about self-sorting into education due to higher ability and hence the postponement of nest leaving, and controlling for family structure (the presence of siblings and living without at least one parent in childhood) is important because it determines the availability of care. Cohort (year of birth) dummy variables are crucial to control for macro-level shocks that are common to all countries, whereas country dummy variables account for all country-specific cultural and institutional factors. In some specifications, we also include either country-specific linear cohort trends or interactions of cohort, gender, and country dummy variables. These trends and interaction terms control for country-specific changes in SES and nest-leaving age that took place over time, as well as for changes in macroeconomic and institutional factors (e.g., compulsory schooling, divorce, or retirement laws) that vary by cohort, country, and gender. Finally, we also include interview wave dummy variables to account for time effects.

Table 2 reports the OLS estimates of the effect of SES at age 10 in three specifications, with heteroskedasticity-robust standard errors shown in parentheses. Column 1 includes the individual-level observable controls and country and cohort dummy variables. Columns 2 and 3 include additional variables: country-specific linear

cohort trends in column 2 and the interaction among year of birth, country, and gender dummy variables in column 3. These regressions are estimated on the sample that excludes the 934 individuals who had not left the parental home by age 49. However, we also estimated an OLS regression model for the probability of nest leaving by age 49 using the same specification adopted in column 2 of Table 2. Reassuringly, we found no effect of SES at age 10 on this probability (coefficient=0.0006; standard error=0.0007).

The results are remarkably stable across columns. Focusing on the point estimate in column 2, we see that a high SES in childhood (SES=1) makes an individual postpone nest leaving by roughly half a year compared with a low SES in childhood (SES=-1). Although SES at age 10 may imperfectly capture SES at potential nest-leaving ages, this source of measurement error would cause attenuation bias and imply that our point estimates can be interpreted as a lower bound for the true effect.

Heterogeneous Effects

To assess whether our main result in the full population applies in general or is specific to population subgroups, we conduct several heterogeneity analyses using our baseline OLS specification reported in column 2 of Table 2. The results of these analyses are shown in Table 3.⁵

Columns 1a and 1b show that although females are significantly more sensitive to childhood SES, the positive effect of childhood SES on nest-leaving age holds for both genders. Similarly, columns 2a and 2b show that the childhood SES effect holds across urban and rural residence at age 10, but those who lived in an urban area were more sensitive to childhood SES.

Columns 3a, 3b, and 3c report heterogeneous effects by the main cultural features of the country of residence in childhood. Following Inglehart and Welzel (2005), we classify each country in our sample into one of three cultures as follows:

- Catholic: Austria, Spain, Italy, France, Greece, Belgium, Israel, Ireland, Luxembourg, Portugal, Cyprus, and Malta;
- Ex-communist: the former East Germany, the Czech Republic, Poland, Hungary, Slovenia, Estonia, Croatia, Lithuania, Bulgaria, Latvia, Romania, and Slovakia; and
- Protestant: the former West Germany, Sweden, the Netherlands, Denmark, Switzerland, and Finland.

We find that although the positive and significant effect of childhood SES is significantly smaller in magnitude in Protestant countries, it is not peculiar to Catholic countries but is generalizable to the entire set of European countries considered in this

⁵ Table 3 also reports the mean of the outcome by group, as well as the ratio between the estimated effects and the mean outcome. Overall, the relative size of the effects across groups is the same in absolute and percentage terms.

Table 3 OLS regressions of nest-leaving age by gender, rural/urban residential area during childhood, and country's culture

	Gender		Childhood Residential Area		Country's Culture		
	Female (1a)	Male (1b)	Rural (2a)	Urban (2b)	Catholic (3a)	Ex-communist (3b)	Protestant (3c)
SES at Age 10	0.373*** (0.030)	0.155*** (0.037)	0.197*** (0.037)	0.312*** (0.031)	0.310*** (0.036)	0.298*** (0.046)	0.185*** (0.040)
Mean Dependent Variable	22.09	24.40	22.94	23.31	24.01	23.10	21.56
Effect / Mean (%)	1.69	0.63	0.86	1.34	1.29	1.29	0.86
H_0 : Female = Male	$p < .01$						
H_0 : Rural = Urban	$p = .02$						
H_0 : Catholic = Ex-communist = Protestant	$p = .04$						
H_0 : Catholic = Ex-communist	$p = .83$						
H_0 : Catholic = Protestant	$p = .02$						
H_0 : Ex-communist = Protestant	$p = .06$						
Number of Observations	22,632	18,620	19,672	21,580	17,629	14,114	9,509
Individual-Level Controls	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Cohort Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country-Specific Cohort Trends	Yes	Yes	Yes	Yes	Yes	Yes	Yes
Country $\times$ Cohort $\times$ Gender Fixed Effects	No	No	No	No	No	No	No

Notes: The subsample used is defined at the top of each column. Each column reports the coefficient on SES at age 10 from a different OLS regression. SES at age 10 is standardized to have zero mean and unit standard deviation in the full sample. All models include the following individual level controls: dummy variables for gender, self-rated math and language ability, residence in a rural versus urban area at age 10, never had siblings, lived without at least one parent at age 10, and wave dummy variables. The fixed effects and trends indicated at the bottom of each column are also included. The estimation sample excludes 932 individuals who had not established their own household by age 49. Heteroskedasticity-robust standard errors are shown in parentheses.

*** $p < .001$

study.⁶ Therefore, our results are not driven by the fact that parents in some cultures might see coresidence with their adult children as a normal good and bribe their children to stay at home with them, as in the case of Italy (Manacorda and Moretti 2006).

In additional analyses (not reported but available from the authors), we also estimated heterogeneous effects by cohort (respondents of the silent generation born in 1936–1945 vs. baby boomers born in 1946–1956), country of birth (Southern Europe, Central/Northern Europe, the ex-communist bloc), birth order (firstborn, middle-born, or last-born sibling),⁷ whether subjects had siblings, whether subjects lived without at least one parent at age 10, and a self-rated assessment of whether either parent ever harmed them physically in childhood. In all cases, we found a positive effect of childhood SES and no evidence of significantly heterogeneous effects. The lack of heterogeneous effects across cohorts speaks in favor of the external validity of our main result. Respondents born in 1936–1945 left the parental home primarily during the period of economic growth that Europe experienced in the 1950s–1960s, and those born in 1946–1956 left home during the oil crisis–induced recessions of the 1970s. Our findings thus suggest that our main result holds irrespective of the prevailing macroeconomic conditions.

The Mediating Role of Education

Individuals with high SES in childhood are likely to acquire more education and may choose to postpone family formation to complete their education. We therefore investigate the mediating role of education in explaining the relationship between childhood SES and nest-leaving age.

In line with the standard approach to mediation analysis (see Baron and Kenny 1986), we assess (1) the impact of childhood SES on years of education; (2) the impact of years of education on nest-leaving age, conditional on childhood SES; and therefore (3) whether the impact of childhood SES on nest-leaving age is explained entirely by its indirect mediating effect on education or whether there is also a direct link between childhood SES and nest-leaving age.

Following Baron and Kenny (1986), the total effect of childhood SES on nest-leaving age can be written as the sum of a direct and an indirect effect. The indirect effect, which is mediated by education, is obtained as the product of the effect of childhood SES on education and the effect of years of education on nest-leaving age. As a result, the direct effect would be zero if the effect of education on nest-leaving age were large enough to make the indirect effect of SES on nest-leaving age equal to the total effect, for a given effect of SES on years of education. Hence, the effect of years of education on nest-leaving age that would lead the direct effect of SES on nest-leaving age to be equal to zero is given by the ratio between the total effect of SES on nest-leaving age and the effect of SES on years of education.

⁶ Given that some countries might display regional variation in the prevailing religion, we also adopted a slightly different definition of culture that codes as Catholic the German *länder* of Bavaria, Rheinland-Palatinate, Nordrhein-Westfalen, and Baden-Württemberg; the Southern region of the Netherlands; and the French- and Italian-speaking cantons of Switzerland. This definition produced comparable results, although the differences between Protestant and the other countries were no longer significant.

⁷ This information is not available for 6,227 observations, mostly belonging to the Wave 7 only sample, which were dropped for this analysis.

Table 4 OLS regressions of the mediating role of education on nest-leaving age

	Dependent Variable		
	Nest-Leaving Age, Baseline Model (1)	Years of Education (2)	Nest-Leaving Age (3)
SES at Age 10	0.274*** (0.024)	1.012*** (0.019)	0.192*** (0.025)
Years of Education			0.068*** (0.007)
Mean Dependent Variable	23.13	11.00	23.13
H_0 : Effect of SES Is Equal in Columns 1 and 3			$p < .01$
Proportional Selection Coefficient (δ) for Zero Effect of Years of Education			$\delta = 1.04$
Proportional Selection Coefficient (δ) for Zero Direct Effect of SES			$\delta = -13.26$
Number of Observations	41,252	41,252	41,252
Individual-Level Controls	Yes	Yes	Yes
Cohort Fixed Effects	Yes	Yes	Yes
Country Fixed Effects	Yes	Yes	Yes
Country-Specific Cohort Trends	Yes	Yes	Yes
Country $\times$ Cohort $\times$ Gender Fixed Effects	No	No	No

Notes: Each of the coefficients on SES at age 10 in each column and the coefficient on years of education in column 3 is from a different OLS regression. SES at age 10 is standardized to have zero mean and unit standard deviation in the full sample. All models include the following individual-level controls: dummy variables for gender, self-rated math and language ability, residence in a rural versus urban area at age 10, never had siblings, lived without at least one parent at age 10, and wave dummy variables. The fixed effects and trends indicated at the bottom of each column are also included. The estimation sample excludes 932 individuals who had not established their own household by age 49. Heteroskedasticity-robust standard errors are shown in parentheses.

*** $p < .001$

The results of our mediation analysis are reported in Table 4. Column 1 reports the total effect of childhood SES obtained from our baseline specification in column 2 of Table 2. Column 2 of Table 4 uses the same empirical specification and shows that high-SES individuals obtain more years of education. Column 3 displays the effect of years of education on nest-leaving age (conditional on childhood SES) and the additional controls included in our baseline specification. This effect is positive and significant: one additional year of education leads individuals to postpone nest leaving by 0.07 years (roughly one month).

The indirect effect of childhood SES on nest-leaving age is obtained as the product of the effect of SES on years of education shown in column 2 and the effect of years on education on nest-leaving age shown in column 3. This effect is equal to 0.069 years (1.012×0.068) for a 1 standard deviation change in SES and is statistically significant ($p < .01$; not reported in the table and obtained after joint estimation of the three equations).

Table 4 also shows that the direct effect of childhood SES, reported in column 3, holds even after we take into account the mediating effect of education. However, this direct effect is significantly smaller than the total effect reported in the baseline model

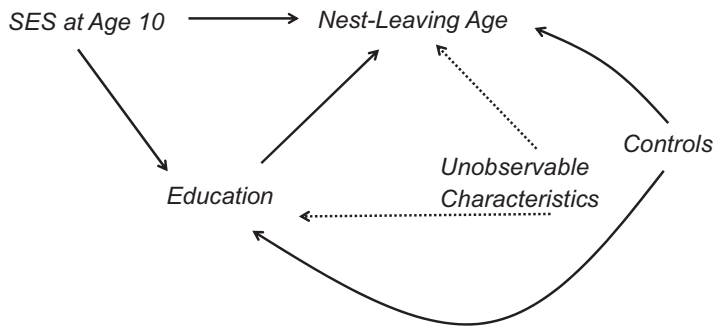


Fig. 3 Mediating role of education

in column 1, suggesting that education partially mediates the effect of childhood SES on nest-leaving age.

One empirical difficulty in carrying out this exercise relates to the potential endogeneity of years of education, as there may be omitted unobservable traits that jointly determine both educational achievement and nest-leaving age, as shown by the directed acyclic graph shown in Figure 3.

Thus far, we have assumed selection on observable characteristics: we have assumed that the controls included in our model (childhood SES, gender, self-reported math and English ability, residence in a rural area, family structure, country and cohort dummy variables, and country-specific linear trends) are such that the remaining variation in years of education is as good as random. However, we can assess the sensitivity of our estimates to selection on unobservable characteristics using Oster’s (2019) proportional selection test, devised following the logic of Altonji et al. (2005). This test aims to assess how strong the impact of unobserved variables would have to be with respect to the impact of observed variables to drive our estimated treatment effects down to zero. We let δ indicate this ratio of proportional selection.⁸ For example, $\delta=1$ would imply that the impact of unobservable characteristics should be as large as the impact of the included controls to nullify our estimated effect. $\delta=1$ is also the threshold Altonji et al. (2005) suggested using to rule out that the bias due to omitted unobservable characteristics is strong enough to explain away the estimated effect.

The last column of Table 4 shows the value of $\bar{C}$ for (1) the estimated effect of education on nest-leaving age and (2) the direct effect of SES on nest-leaving age. The first test corroborates our strategy to estimate the effect of education on the

⁸ Let us assume the following simple model (see Oster 2019): $Y = \beta X + \gamma w_1 + W_2 + \epsilon$, where X is the treatment, w_1 is an observable control, and W_2 is an index of unobserved controls. The proportional selection relationship is then $\delta \frac{Cov(w_1, X)}{Var(\gamma w_1)} = \frac{Cov(W_2, X)}{Var(W_2)}$, where δ is the coefficient of proportionality. The test works by computing the value of δ for which β would be 0 under the assumption that $w_1 \perp W_2$. To conduct this test, we need to state a maximum attainable value of the R squared (R_{max}), which indicates the maximum share of variance of the outcome that could be explained by any set of observable and unobservable covariates (w_1 and W_2). Following Oster’s (2019) suggestions, we set the value of R_{max} to 1.3 times the value of the R squared of the model that includes all the observable controls (only w_1).

dependent variable. The second test is more directly relevant for our case, given that we are investigating the *mediating* effect of education.

The results show that the estimated value of δ for the estimated effect of education on nest-leaving age in column 3 is larger than 1, putting us in a safe position according to the threshold Altonji et al. (2005) proposed. By contrast, δ is very large and negative for the effect of childhood SES on nest-leaving age. Hence, the effect of SES on nest-leaving age could be driven down to zero by the presence of unobservable characteristics only if these characteristics were much more relevant drivers of selection bias than the included controls and their correlations with SES were of the opposite sign with respect to one of the included controls. This possibility is implausible because our set of controls is rather comprehensive.

Robustness Tests

Before presenting our economic model, we report the results of several sensitivity tests that support the robustness of our previously presented results. First, our baseline OLS estimates in Table 2 are obtained from the sample of 41,252 subjects who left the parental home by age 49. To safeguard against this potential censoring problem, we repeat our main analysis using a logistic regression to estimate a discrete duration model for the probability of nest leaving at any given age, conditional on not having left already. Unlike the OLS regression, this model uses information on all 42,184 individuals. However, because each individual contributes an observation for each year from age 14 until nest-leaving age or until 49 if the individual has not left yet (right-censored observations), the panel estimation sample has 451,563 observations. In these equations, we use the same set of time-invariant controls as in the OLS regressions, as well as standard error estimators that are robust to heteroskedasticity.

Table A3 in the online appendix displays estimates of the effect of SES on the odds ratio of nest leaving for three specifications that differ with respect to the functional form chosen for the baseline hazard and for the set of trends and fixed effects included, as reported at the bottom of each column. All columns include the individual-level observable covariates included in our main specification in Table 2. In all cases, the key parameter is significantly smaller than unity, indicating that children from a higher SES leave the parental nest later.

Figure A1 in the online appendix shows the estimated hazard rate (panel a) and survival function (panel b) by SES resulting from the model shown in column 1 of Table A3. These estimates are averages obtained after we assign $SES = -1$ (low SES) and $SES = 1$ (high SES) to all units in the sample. The results show that the hazard rate (the nest-leaving probability for those still living in the parental home) is always higher for low-SES individuals, and the difference between groups peaks at around age 25.

Second, we replicate the OLS analysis shown in Table 2 but introduce all four measures of SES at age 10 that we use to build our index instead of using the index itself. The results shown in Table A1 in the online appendix indicate that each measure attracts a significant coefficient, suggesting that no single facet of SES (e.g., the number of rooms per capita in the parental home, the home's amenities) is driving our results.

Finally, we address a remaining concern about our estimates on the mediating role of education: some individuals may have answered the question on nest-leaving

age by reporting the year when they left home to attend a university. If so, then our results might be picking up an effect of education on nest-leaving age that is smaller than it should be, given that individuals with more education would report leaving home earlier than they should. Although this pattern is not very common in our data (occurring for 11% of the respondents), we conduct a robustness check to dispel this concern. We replicate the exercise shown in Table 4, this time using age at first cohabitation as the dependent variable. Because it refers to a specific living arrangement, this measure is less amenable to a subjective interpretation of *nest leaving*. Reassuringly, the results (shown in Table A5 in the online appendix) are wholly consistent with our earlier findings. If anything, we obtain larger coefficients of proportional selection, lending greater support to our empirical strategy.

Model and Simulation Results

To provide an economic explanation of our robust finding that a golden nest delays nest leaving, we consider a life cycle model in which the young base their nest-leaving decision on the utility they enjoy during their lifetime. For simplicity, we consider a three-period life cycle model. Period 0 is childhood and is predetermined: the childhood consumption level ($\bar{C}$) is outside individuals' control. In Period 1, the young people can choose whether to stay in the parental home or to leave. Individuals who stay share any income with their parents and consume the same amount as in their childhood ($\bar{C}$). Individuals who move out smooth their resources across all three periods. In Period 2, we assume that no child remains with their parents.

In this context, individuals always leave home in Period 1 if their income in that period is higher than their childhood consumption level ($\bar{C}$). This result is not surprising, because nest leaving allows them to enjoy higher lifetime resources and the freedom to optimize consumption and saving decisions over their entire adulthood (three periods).

If individuals' income in Period 1 is instead lower than their childhood consumption, lifetime resources would be higher if they stayed with their parents in Period 1. However, nest leaving may still be attractive if their Period 1 income is just below their childhood consumption because nest leaving confers the possibility to smooth consumption. A numerical example helps explain why.

Assume the interest rate equals the time preference parameter, so the optimal plan is characterized by the same consumption level across periods. Assume that the utility function is well behaved (we use the constant absolute risk-aversion (CARA) function in our numerical simulations).

Suppose that income is increasing with age, such that it is 0.5 in Period 1, 1 in Period 2, and 1.5 in Period 3. If $\bar{C}$ is 0.6, then nest leaving is optimal because individuals can consume 1 in all three periods (they would consume 0.6 in Period 1 and 1.25 in Periods 2 and 3 if they stayed in the parental home in Period 1). However, if $\bar{C}$ is 0.9 (a golden nest), the nest-leaving consumption profile of (1, 1, 1) will be less attractive than the home-staying profile of (0.9, 1.25, 1.25) for sensible preference parameters.

We also add habits to our nest-leaving model. The standard way in which habits are modeled is that utility is affected by a stock of habits that depends on past consumption. Rational consumers will take this dependence in consideration when deciding today's consumption. We show that habits reinforce the dependence of the

nest-leaving decision on the childhood consumption level. In our model, we follow Angelini (2009) and define period utility over the current period and the previous period consumption:

$$U_t = U(C_t, C_{t-1}).$$

A commonly adopted specification is

$$U_t = U(C_t - \gamma C_{t-1}),$$

where $0 \leq \gamma < 1$ denotes the force of habits. U_t could be of the CARA type, as in Angelini (2009):

$$U_t = -\frac{1}{\theta} e^{-\theta(C_t - \gamma C_{t-1})}, \quad (1)$$

where θ is the absolute risk-aversion parameter.

We define $NL=1$ if the consumer leaves the parental nest in Period 1, and $NL=0$ if the consumer stays with the parents in Period 1 and leaves the parental nest one year later. Let the time preference discount factor be β and the gross interest rate be R , and assume that the consumer can freely lend and borrow at that interest rate after leaving the parental home. The optimization problem then is

$$\max_{C_1, C_2, C_3; NL} \sum_{t=1}^3 \beta^{t-1} U_t,$$

which is subject to the following:

$$C_0 = \bar{C}$$

$$C_1 = \bar{C} \text{ if } NL = 0$$

$$C_1 = Y_1 - A_1 \text{ if } NL = 1$$

$$C_2 = Y_2 - A_2 \text{ if } NL = 0$$

$$C_2 = Y_2 - A_2 + RA_1 \text{ if } NL = 1$$

$$C_3 = Y_3 + RA_2 \text{ if } NL = 0.$$

The consumer leaves the parental home in Period 1 if their lifetime utility is higher when $NL=1$ but otherwise leaves in Period 2. The issue is whether $NL=1$ is more likely if $\bar{C}$ is high or low for a given income at Periods 1, 2, and 3. We can work out the analytical solution if we adopt the CARA specification as in Eq. (1) and use it to calculate and compare the utility of the nest-leaving and home-staying strategies.

It is worth stressing that the way the constraints are written, the cohabiting child passes along the entire Period 1 income to the parents in exchange for enjoying a given consumption level, $\bar{C}$. This model predicts that for a given common income profile, golden nest individuals are less likely to move out in Period 1 than the less fortunate.

In Figure 4, we show the utility gain associated with nest leaving (rather than staying) when income takes values of 0.5 in Period 1, 1 in Period 2, and 1.5 in Period 3. The interest rate and the time preference parameter are 2% (hence, $R = 1.02$, $\beta = \frac{1}{1.02}$).

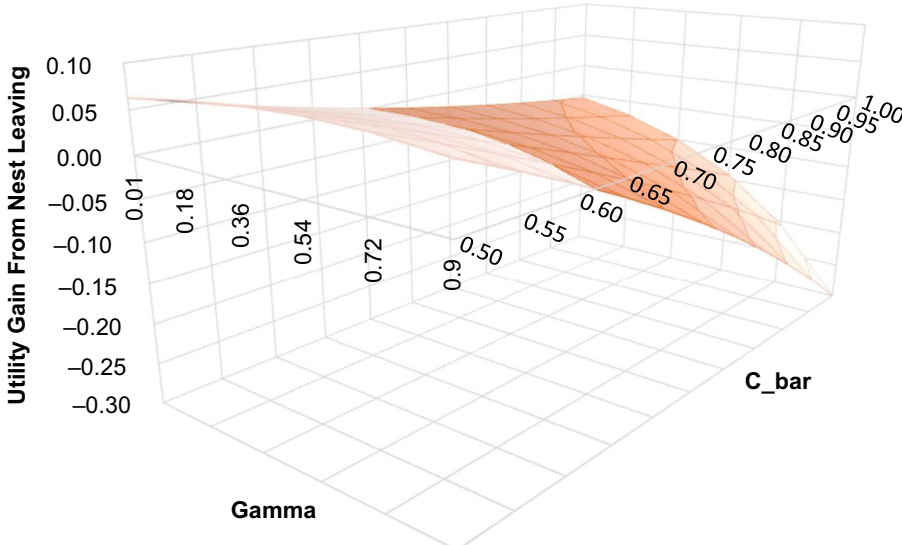


Fig. 4 Utility gain from nest leaving as a function of $\bar{C}$ and γ

The absolute risk aversion parameter, θ , is set equal to 2. We consider how the utility gain changes as a function of $\bar{C}$ (which is allowed to vary between 0.5 and 1) and the force of habits. The habits parameter, γ , is allowed to vary between 0 (in the standard model) and 0.9 (indicating very strong habit dependence).

We see that the utility gain from nest leaving is a decreasing function of both $\bar{C}$ and γ : the higher the consumption at the parental home and the stronger the habit, the less appealing nest leaving becomes. The utility gain from nest leaving is always positive for the lowest values of $\bar{C}$, implying that individuals prefer leaving the parental nest in Period 1 if their standard of living in the parental home is sufficiently low. However, there is a $\bar{C}$ high enough that nest leaving is delayed: in the model without habits, this threshold is $\bar{C} = 0.75$. When habits are particularly strong ($\gamma = 0.9$), individuals prefer staying in the parental home even if $\bar{C}$ is as low as 0.65. We obtain similar results for a less steeply ascending income path (if income takes values 0.8 in Period 1, 1 in Period 2, and 1.2 in Period 3). In that case, nest leaving is more frequently the preferred option, but the utility gain shown in Figure 4 is still a decreasing function of $\bar{C}$ and γ . We find that changes of the risk-aversion parameter and the interest are of little consequence for the nest-leaving decision.

Our model ignores bequests: parents can strategically use (the promise of) bequests to induce their children to postpone the nest leaving. This feature of the model should strengthen its main prediction that children who grow up in a golden nest leave the parental home at later ages. It also ignores the possibility that the utility function depends on living arrangements. However, if preferences change upon nest leaving the same way across individuals, our model prediction should be unaffected.⁹

⁹ For instance, Fernandes et al. (2008) developed a similar but more complex model that considers the role of income insecurity as well as parental utility and altruism.

Conclusions

We used data covering 28 European countries plus Israel to show empirically that individuals who grew up in a golden nest leave the parental home later. This result is remarkable for two reasons. First, it contradicts the commonly held view (in the sociodemographic literature) that the nest-leaving decision is mostly determined by cultural factors. Second, it contradicts the typical hypothesis in the economics literature that capital and housing market imperfections explain why young people find it hard to leave the parental home. If culture were the driving force, the SES gradient should not be similar across countries that markedly differ in their cultural background. If limited access to credit or cheap housing were the key issue, we would expect poorer children to delay nest leaving, not richer ones.

We then solved a three-period intertemporal optimization model to demonstrate the conditions under which this behavior is consistent with standard assumptions on preferences and resources. Our key result is that a standard life cycle model without borrowing constraints predicts this type of behavior if earnings increase with age and the standard of living in the parental home is higher than the income earned early in the life cycle. We also find that habit-forming preferences reinforce the delaying effect of a golden nest on nest leaving because these preferences make drops in the standard of living particularly unattractive to the consumer.

An implication of our results is that to the extent that early nest leaving has positive longer-run consequences on children's economic welfare (see, e.g., Billari and Tabellini 2010), the earlier nest leaving by children of low-SES families may contribute to increases in intergenerational mobility in income and wealth.

To the extent that late nest leaving is explained in terms of preferences and economic resources (budget sets), a laissez-faire attitude is justified in purely economic terms. If policymakers believe that late nest leaving should be discouraged (for instance, because of its association with low fertility, which leads to a demographic imbalance), they can use fiscal measures to increase better-off young adults' incomes or decrease their parents' income. Examples could be a reduced or even negative earnings taxation for young individuals living independently, accompanied by the progressive taxation of total household income (defined as the sum of earnings of cohabiting parents and children). However, these policies would favor at least some of those who are better off in a lifetime sense and may increase wealth inequality. ■

Acknowledgments We are grateful to Francesco Maura for excellent research assistance and to the audiences at a seminar at the University of Padua, the 2021 annual meeting of the Population Association of America, and the 2021 meeting of the Society of the Economics of the Household for useful comments. This study uses data from SHARE Waves 3 and 7 (DOIs: 10.6103/SHARE.w3.700, 10.6103/SHARE.w7.700); see Börsch-Supan et al. (2013) for methodological details. The SHARE data collection was funded by the European Commission through FP5 (QLK6-CT-2001-00360), FP6 (SHARE-I3: RII-CT-2006-062193, COMPARE: CIT5-CT-2005-028857, SHARELIFE: CIT4-CT-2006-028812), FP7 (SHARE-PREP: GA N°211909, SHARE-LEAP: GA N°227822, SHARE M4: GA N°261982), and Horizon 2020 (SHARE-DEV3: GA N°676536, SERISS: GA N°654221), and by DG Employment, Social Affairs & Inclusion. Additional funding from the German Ministry of Education and Research, the Max Planck Society for the Advancement of Science, the U.S. National Institute on Aging (U01 AG09740-13S2, P01 AG005842, P01 AG08291, P30 AG12815, R21 AG025169, Y1 AG-4553-01, IAG BSR06-11, OGHA 04-064, HHSN271201300071C), and various national funding sources is gratefully acknowledged (see www.share-project.org).

References

- Aassve, A., Billari, F. C., Mazzucco, S., & Ongaro, F. (2002). Leaving home: A comparative analysis of ECHP data. *Journal of European Social Policy*, 12, 259–275.
- Alessie, R., Brugiavini, A., & Weber, G. (2006). Household saving and cohabitation. In R. H. Clarida, J. A. Frankel, F. Giavazzi, & K. D. West (Eds.), *NBER international seminar in macroeconomics 2004* (pp. 413–441). Cambridge, MA: MIT Press.
- Altonji, J. G., Elder, T. E., & Taber, C. R. (2005). Selection on observed and unobserved variables: Assessing the effectiveness of Catholic schools. *Journal of Political Economy*, 113, 151–184.
- Angelini, V. (2009). Consumption and habit formation when time horizon is finite. *Economics Letters*, 103, 113–116.
- Baron, R. M., & Kenny, D. A. (1986). The moderator–mediator variable distinction in social psychological research: Conceptual, strategic, and statistical considerations. *Journal of Personality and Social Psychology*, 51, 1173–1182.
- Billari, F. C., Philipov, D., & Baizán, P. (2001). Leaving home in Europe: The experience of cohorts born around 1960. *International Journal of Population Geography*, 7, 339–356.
- Billari, F. C., & Tabellini, G. (2010). Italians are late: Does it matter? In J. B. Shoven (Ed.), *Demography & the economy* (pp. 371–412). Chicago, IL: University of Chicago Press.
- Björklund, A., & Jäntti, M. (2012). Intergenerational income mobility and the role of family background. In B. Nolan, W. Salverda, & T. M. Smeeding (Eds.), *Oxford handbook of economic inequality* (pp. 491–521). Oxford, UK: Oxford University Press. <https://doi.org/10.1093/oxfordhb/9780199606061.013.0020>
- Börsch-Supan, A., Brandt, M., Hunkler, C., Kneip, T., Korbmayer, J., Malter, F., . . . Zuber, S. (2013). Data resource profile: The Survey of Health, Ageing and Retirement in Europe (SHARE). *International Journal of Epidemiology*, 42, 992–1001.
- Cobb-Clark, D. A. (2008). Leaving home: What economics has to say about the living arrangements of young Australians. *Australian Economic Review*, 41, 160–176.
- Cobb-Clark, D. A., & Görgens, T. (2014). Parents' economic support of young-adult children: Do socioeconomic circumstances matter? *Journal of Population Economics*, 27, 447–471.
- Diaz, A., Pijoan-Mas, J., & Rios-Rull, J.-V. (2003). Precautionary savings and wealth distribution under habit formation preferences. *Journal of Monetary Economics*, 50, 1257–1291.
- Fernandes, A., Becker, S. O., Bentolila, S., & Ichino, A. (2008). Income insecurity and youth emancipation: A theoretical approach. *B.E. Journal of Economic Analysis & Policy*, 8. <https://doi.org/10.2202/1935-1682.1783>
- Giuliano, P. (2007). Living arrangements in Western Europe: Does cultural origin matter? *Journal of the European Economic Association*, 5, 927–952.
- Guiso, L., & Jappelli, T. (2002). Private transfers, borrowing constraints and the timing of homeownership. *Journal of Money, Credit and Banking*, 34, 315–339.
- Inglehart, R., & Welzel, C. (2005). *Modernization, cultural change, and democracy: The human development sequence*. Cambridge, UK: Cambridge University Press.
- Kiernan, K. E. (1986). Leaving home: Living arrangements of young people in six West-European countries. *European Journal of Population*, 2, 177–184.
- Manacorda, M., & Moretti, E. (2006). Why do most Italian youths live with their parents? Intergenerational transfers and household structure. *Journal of the European Economic Association*, 4, 800–829.
- Mazzonna, F. (2014). The long-lasting effects of family background: A European cross-country comparison. *Economics of Education Review*, 40, 25–42.
- Oster, E. (2019). Unobservable selection and coefficient stability: Theory and evidence. *Journal of Business & Economic Statistics*, 37, 187–204.
- Tukey, J. W. (1977). *Exploratory data analysis*. Reading, MA: Addison-Wesley.

Angelini • Department of Economics, Econometrics and Finance, University of Groningen, Groningen, the Netherlands; Netspar, Tilburg, the Netherlands; <https://orcid.org/0000-0002-5177-3447>

Bertoni • Department of Economics and Management, University of Padova, Padova, Italy; IZA, Bonn, Germany; <https://orcid.org/0000-0003-0783-3972>

Weber • Department of Economics and Management, University of Padova, Padova, Italy; Institute for Fiscal Studies, London, UK; <https://orcid.org/0000-0002-2137-9676>