Living Arrangements in Europe: Whether and Why Paternal Retirement Matters*

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Abstract

This paper uses retrospective micro data from eleven European countries to investigate the role of paternal retirement in explaining children's decisions to leave the parental home. To assess causality, I use a bivariate discrete-time hazard model with shared frailty and exploit over time and cross-country variation in early retirement legislation. Overall, the results indicate a positive and significant influence of paternal retirement on the probability of first nest-leaving of children residing in Southern European countries, for both sons and daughters. By contrast, there is no evidence of significant effects on children living in Northern and Central European countries. I then discuss and test empirically the potential mechanisms by which paternal retirement may affect children's co-residence. I find that the increase in children's nest-leaving around paternal retirement does not appear to be justified by changes in parental resources or in the supply of informal child care provided by grandparents. Rather, one must probably look for channels involving negative externalities in preferences between retired fathers and their children.

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1 Introduction

Over the last few years, a substantial body of research has attempted to identify some of the potential determinants that may induce youths to continue living with their parents. While this investigation is particularly relevant for Italy and some other Southern European countries, such as Spain and Greece, where young people tend to remain with their parents until their late 20s and early 30s, leaving home only when they get married, the way children respond to these factors has attracted increasing attention in the public policy debate of most European countries. For example, policymakers may be interested in reducing the adverse impact of delayed cohabitation on an array of children's outcomes, including individual motivations and ambitions, reservation wages, labor market entry and geographical mobility (Billari and Tabellini 2010). A further cause of concern regards the phenomenon of falling fertility rates associated with prolonged co-residence. Combined with the effects of population aging, this phenomenon raises the elderly dependency ratio, thereby contributing to placing extra pressure on the long-term financial sustainability of pension systems.

This issue has also been actively debated among economists. There is consensus in recent literature that in Italy parental retirement induces a significant decline in the number of grown children living with their parents; however, researchers remain puzzled about the possible mechanisms underlying this relationship. There are two major competing explanations for this pattern. On the one hand, Manacorda and Moretti (2006) argue that retired parents are no longer able to make a financial transfer to their children and thus are unable to bribe them to stay at home because of the drop in their post-retirement income. On the other hand, Battistin et al. (2009) emphasize that liquidity considerations are unlikely to play a role because most Italian employees receive a generous lump-sum payment upon retirement. Therefore, they suggest that parents may use part of their severance payment to help their children leave the nest, which may account for most of the decline in consumption around the time of retirement. While these two studies differ in many respects, they have two important common traits. First, they use Italy as a case study. The Italian case is of particular interest because Italy is among the European countries with the highest age for home-leaving and because it is one of the very few European countries in which workers are entitled to receive a large severance payment at the time of retirement. A second similarity is that both studies employ an instrumental variable (IV) approach that obtains identification from Italian pension reforms that substantially changed the eligibility conditions for retirement during the 1990s.

Overall, the lack of a cross-country analysis severely limits the ability to clarify whether the housing emancipation of young adults upon parents' retirement can be attributed to liquidity problems faced by parents, as suggested by Manacorda and Moretti (2006), or to the receipt of a sizeable retirement allowance, as noted by Battistin et al. (2009). Thus, there is the need for empirical work to test which of the channels dominates in practice.

This paper contributes to the extant literature by taking advantage of a European dataset to test and discuss the relative weight of these two competing hypotheses and shed some light into the mechanism. To address problems of endogeneity caused by omitted variables or reverse causation, I estimate a bivariate discrete-time hazard model with shared frailty (Abbring and Van den Berg 2003) for the impact of paternal retirement on the timing of children's nest-leaving. Furthermore, to provide random variation in the timing of paternal retirement, I strengthen my identification strategy by employing changes in eligibility rules for early retirement benefits that were implemented across European countries and during the period 1961 to 2007 as an exclusion restriction. To the best of my knowledge, this is the first paper that makes use of this exogenous source of variation to children's living arrangements to assess whether and to what extent paternal retirement caused their children to leave the nest. Compared to the linear IV strategy, the hazard specification provides a more appropriate statistical framework for modeling time-to-event/survival outcomes and accounting for right-censoring, thereby allowing me to overcome certain limitations faced by previous IV studies. The bivariate hazard model finally offers greater flexibility in handling nonlinear baseline hazards and nonlinear effects of covariates and provides a novel approach to identifying treatment effects by modeling unobserved heterogeneity explicitly through bivariate specification.

To conduct this analysis, I use data from the second wave (2006) of the Survey of Health, Ageing and Retirement in Europe (SHARE). This European dataset has three important features: first, it collects data on current economic, health, and family conditions of over 30,000 individuals aged fifty and above in several European countries; second, it provides retrospective information on the retirement age of the respondents and the nest-leaving ages of their children; and lastly, because it is designed to be cross-nationally comparable, this dataset enables me to properly conduct a multi-country analysis. Furthermore, I employ data about the European early retirement legislation by relying on Angelini et al. (2009), Mazzonna et al. (2012) and the country-specific studies discussed in Gruber and Wise (2004). It should be stressed, however, that across the countries considered in the present investigation there are very different cultural histories, labor market institutions and social characteristics. Such differences may play a lasting role in explaining the substantial heterogeneity in the ages of children when they leave home across Europe (Aassve et al. 2002; Billari et al. 2001) and may not be entirely captured by including country fixed effects in the model estimated on the pooled sample from multiple countries. To mitigate this concern, I conduct the main analysis by European region. These regions correspond to the geographical aggregation into Northern European countries (Sweden, Denmark and the Netherlands), Central European countries (Austria, Germany, Switzerland, France and Belgium) and Southern European countries (Italy, Spain and Greece). According to the previous literature (see, for example, Albertini et al. 2007, 2012), this aggregation is particularly relevant because it reflects profound differences in welfare states and family regimes across the above-mentioned country groups. One implication of this division is that the conditional impact of early retirement eligibility rules on paternal retirement and children's nest-leaving outcomes is allowed to vary between Northern, Central and Southern European countries.

Based on these data, my main results demonstrate the following: a) Paternal retirement has a positive and significant effect on the timing of children's nest-leaving in Southern European countries. In this European region, the magnitude of the effect varies between 1.4% and 5.5%, and there are no significant differences between sons and daughters; b) The mechanism through which this pattern may occur remains an open issue because it cannot be attributed to families' liquidity problems or a severance payment at the time of paternal retirement. One must probably look for channels involving negative externalities in preferences between parents and children; c) In Northern and Central Europe, there is no evidence that children's nest-leaving outcomes are significantly affected by paternal retirement. These findings are robust to a number of specification checks. On the policy side, the results of this paper suggest that in Southern Europe there are potentially unintended and undesirable consequences of pension reforms on moving-out decisions of young people.

The remainder of the paper is organized as follows. The next section discusses the relevant literature on children's nest-leaving. Section 3 presents a description of the data and provides background information on eligibility ages for retirement in Europe. Section 4 describes the empirical specification and identification strategy. The main results of the paper are presented in Section 5, and Section 6 illustrates the robustness checks. I discuss the results in Section 7, and concluding remarks are provided in Section 8.

2 Related Literature

A vast economic literature has investigated the channels that may affect young individuals' living arrangements. Most papers have focused on parental and children's economic resources, youth labor market conditions, the prevailing characteristics in housing markets and cultural factors. Among these channels, the father's resources around the time of retirement plays a relevant role. As discussed herein, although there is consensus that parental retirement encourages the nest-leaving of Italian young adults, less is known about the mechanisms underlying their departure from the parental home. In the literature to date, there are two competing explanations for the change in the pattern of children's leaving home upon paternal retirement. The first explanation, proposed by Manacorda and Moretti (2006), concentrates on the role played by parental preferences for co-residence. Using the Italian pension reforms of the 1990s as a source of exogenous variation in household income, the authors find that the prolonged co-residence of youths can be attributed to parents' desire for cohabitation because they may be willing to give up some of their additional income due to postponed retirement to bribe their children to stay at home longer. This view would imply that once parents retire, they are no longer able to keep their children at home as a result of the decline in their post-retirement income. The second explanation, that of Battistin et al. (2009), suggests a different mechanism. According to these authors, because most Italian employees receive a sizeable severance payment upon retirement, parents may use this money to buy a house for their sons and daughters, who can then leave the parental home.[1]

These two studies, however, limit their analyses to the Italian case and do not test the implications of their findings on other European countries. Therefore, the multi-country analysis and the source of exogenous variation provided by the early retirement legislation in Europe allows this study to address questions that other researchers have not. By exploiting the intergenerational nature of the dataset, I analyze the decline in children's co-residence at the time of their fathers' retirement. In particular, I provide the first empirical test for these two competing explanations and shed some light on the specific mechanism through which this may happen. As noted by Battistin et al. (2013), there could be an additional mechanism that explains the increase in children's nest-leaving around parental retirement: for instance, if pension reforms force grandparents to stay in the labor market longer and thus reduce the time devoted to child care activities with their grandchildren. The authors find heterogeneous effects depending on the gender, with grandmothers having a significantly stronger impact on their children's fertility and nest-leaving outcomes.

[1] Guiso and Japelli (2002) analyze the importance of this channel, finding that economic transfers from parents contribute to earlier nest-leaving of their children.

This paper is also related to other contributions from the economic literature on moving-out decisions. Most notably, Becker et al. (2010) show that high rates of co-residence among young Italians can be the result of higher job insecurity compared to that of their parents, whereas Card and Lemieux (2000) find that poor labor market conditions and lower wages decrease the probability of leaving the parental nest. Another potential determinant of moving-out decisions are housing market features. Analyzing living arrangements in Italy and the Netherlands, Alessie et al. (2006) highlight that the presence of high transaction costs in housing discourages home-leaving. Finally, my paper relates to recent literature in economics that attempts to quantify the impact of culture on economic outcomes, including children's living arrangements. The starting point of this strand of literature is the observation, by Reher (1998), that Western Europe can be divided into two groups: the Southern European countries, which are characterized by the existence of "strong family ties"; and their Northern European counterparts, which are characterized by "weak family ties". According to this scholar, the late departure from the parental home is one of the indicators of "strong" family ties. Giuliano (2007) studies the impact of the sexual revolution of the 1960s on the propensity of adult children to remain in their parents' home and argues that high rates of cohabitation in Southern European countries can be explained by liberalized parental attitudes towards their children's participation in pre-marital sex. She concludes that cultural traits play a major role in determining living arrangements. In a similar vein, Alesina and Giuliano (2011) provide evidence that in societies with strong family ties home production and the proportion of young adults living at home are higher, whereas labor force participation and geographical mobility are lower compared to those of societies with weak family ties.

3 Data and Institutional Context

In my empirical analysis, I draw data from the Survey of Health, Ageing and Retirement in Europe (SHARE). This survey collects key information on demographics, current socio-economic status, health, expectations and social and family networks for nationally representative samples of European individuals aged fifty and above who speak the official language of their respective countries, and who do not live abroad or in an institution, plus their spouses or partners irrespective of age. In this paper, I use data from the second wave collected in 2006/2007. This wave is particularly suitable for my investigation, as it provides retrospective information on the retirement years of the respondents and the year in which their children left their parental houses. The main advantage of this data source concerns the representativeness of the sample of elderly individuals in Europe, because this survey is constructed to ensure the comparability of the analysis across the different countries. In this study, I present evidence from eleven countries for which I was able to collect information on the legislated early and normal ages at which individuals become eligible for a public old-age pension. These countries cover the various regions of continental Europe, ranging from Scandinavia (Sweden and Denmark), through Central Europe (Austria, Belgium, France, Germany, Switzerland and the Netherlands) and the Mediterranean countries (Italy, Spain and Greece).

In my sample selection, I constrain the sample of parents to fathers because of the problems associated with labor market interruptions that typically characterize the careers of women of childbearing age. Battistin et al. (2009) also focus on fathers. Moreover, I restrict my attention to fathers who were either working[2] or retired at the time of the survey, who have at least one biological child, and who were born between 1920 and 1957. Overall, these cohorts of fathers were affected by changes in the eligibility for old-age and early retirement benefits resulting from reforms that gradually came into effect across Europe over the period 1961 to 2007 to respond to the demographic transition. To construct the sample of children, I include all children, both first-born and later-born children,[3] and the cohorts of interest were born between 1940 and 1988. The choice of this interval allows me to consider virtually all the cohorts of children who were at least 18 at the time of the interview. I then link the socio-demographic characteristics of each child to the data of the corresponding father to create an intergenerational dataset. After these restrictions, I obtain a working sample of parents that contains 4,935 fathers and a sample that consists of 10,720 children (5,525 sons and 5,195 daughters). The distribution of the sample of fathers as well as the sample of children across the countries is presented in Table 1.

[Table 1 - around here]

Descriptive statistics on the primary variables of interest are reported in Table 2. As expected, the vast majority of the fathers (72%) are retired in the interview year of wave 2, and approximately 30% of the fathers report their general health as being less than good. The individuals in my sample of children's generation are, on average, 38 years old, 52% are men and they have much better educational outcomes than their fathers (approximately 40% of adult children have completed their undergraduate or graduate studies versus 23% of the first generation).

[Table 2 - around here]

[2] I use the term "working" to denote both the employed in the private or public sector and the self-employed at time of the interview of wave 2.

[3] In SHARE, questions on the children's nest-leaving age are asked for a maximum of four children.

To determine the retirement age of the fathers and age at which children leave the nest, I exploit recall information from the following two questions in the questionnaire asked to the parents: "In what year did you retire?" and "In what year did the child move from the parental household?". The availability of such information relating events that occurred at some point in time before the year of the survey is essential because it allows for the creation of a retrospective panel dataset. For this reason to conduct the analysis, I assume that individuals can locate past events along the time line with adequate precision. While these retrospective data are self-reported and may be susceptible to recall error that may bias coefficient estimates, the validation studies by Havari and Mazzonna (2011) and Garrouste and Paccagnella (2010) find that the fraction of memory errors is likely to be low, thereby confirming the overall accuracy of the retrospective information in the SHARE data.[4]

Some limitations of my data are worth mentioning. First, with the exception of the year of nest-leaving, I lack any source of time-varying information on children, such as the year of marriage, the year young people left education or their employment history. Second, I lack information regarding the reason for children's nest-leaving and there is no information on the characteristics of the house at the time of children's moving-out.

As discussed in the introduction, I conduct the main analysis by grouping countries into Southern (Italy, Spain and Greece), Northern (Sweden, Denmark and the Netherlands) and Central (Austria, Germany, Switzerland, France and Belgium) Europe. Figure 1 illustrates the mean age at which children leave the nest by gender and country group. As expected, young adults living in Southern Europe moved out much later than their counterparts in the other regions. To be more specific, compared to youths in Northern European countries, Italians, Spanish and Greek children left approximately five years later (26.9 years in Southern Europe versus 22.1 years in Northern Europe). Young people in the Central European countries fall somewhere between these extremes. The figure also shows the presence of a gender gap in nest-leaving age: daughters leave the parental home earlier than sons, ranging from approximately 1 year in Northern and Central Europe to approximately 2 years in Southern Europe. This gap can partly be explained by the fact that age at marriage, which is positively correlated with the postponement of home-leaving, is lower for

[4]The quality of the retrospective information is a feature that has been investigated also in other surveys. For instance, Smith (2009) confirms the validity and reliability of recalled health questions in the Health and Retirement Survey (HRS) and the Panel Study of Income Dynamics (PSID). Furthermore, in their study of the long-term impact of early life environment on outcomes of individuals later on in life, Gould et al. (2011) find that retrospective information collected more than 50 years ago is of reasonably high quality.

women.

[Figure 1 - around here]

In Figure A1 in Appendix A, I show that the proportion of married daughters is higher than that of married sons across all European regions. Interestingly, in Southern Europe, the fraction of married individuals is markedly higher than that in the other regions.[5] Table 3 reports the share of adult children that left home after paternal retirement, with Southern Europe showing the highest mean level, especially for sons.[6]

[Table 3 - around here]

With regard to the institutional context, I use data on early eligibility ages across the above-mentioned European countries, building on the work by Angelini et al. (2009), Mazzonna et al. (2012) and Gruber and Wise (2004).[7] Figure 2 shows the distribution of the actual paternal retirement age for each country. The vertical red and blue lines denote, respectively, the eligibility ages for old-age and early retirement benefits, whereas the red and blue areas indicate changes in eligibility ages for the cohorts in my sample. As expected, there are sizeable jumps in retirement rates that occur at early and standard retirement ages. The overall picture reveals that across eleven countries with very different social security systems and labor market institutions, there are noticeable differences in many respects. For example, the normal age of eligibility for pension benefits is currently set at 65 in almost all countries, but ranges from a low of 60 in a couple of countries (Italy and France) to a high of 67 in some Nordic countries (Denmark and Sweden). A further feature worth stressing is that there is even larger multi-country variability in early eligibility ages. Especially striking is that the early retirement age ranges from 52 in Italy before 1998 to 61 in Sweden after 1997.

[Figure 2 - around here]

[5] The dummy variable is coded as 1 for married adult children living together with the spouse during the interview year of wave 2 and 0 otherwise.

[6] Table 3 also demonstrates that gender differences within each macro-region are statistically significant.

[7] Information on the retirement legislation in Greece is obtained from Duval (2003).

4 Empirical Specification

4.1 Bivariate Discrete-Time Hazard Model with Shared Frailty

In this section, I describe my approach to investigating the extent to which paternal retirement affects the probability of the first nest-leaving of children. To do this, I use a bivariate discrete-time hazard model with shared frailty.[8] This novel strategy to identify treatment effects in the presence of an endogenous treatment when both the treatment and outcome are survival variables was pioneered by Abbring and Van den Berg (2003). This class of models is specified in terms of the hazard, defined as the conditional probability of the event occurring at a point in time provided that it has not already occurred. In this study, I am interested in jointly estimating a bivariate hazard model for the first episode of a child leaving the nest (first equation) and the first time that the father retires (second equation), allowing for correlations between the unobserved heterogeneity terms that affect these two transitions (shared frailty).[9] Formally, the model can be written in the following way:

$$\begin{cases} \theta_{1,it} & = \lambda_1(t)\,\phi_1\left(X_i\beta_1 + \delta Retired_{it} + u_{1,i}\right) \\ \\ \theta_{2,it} & = \lambda_2(t)\,\phi_2\left(X_i\beta_2 + \gamma Eligible_{it} + u_{2,i}\right) \end{cases} \tag{1}$$

where the unit of observation i represents the child-father pair residing in a given country, the outcome $\theta_{1,it}$ is the hazard that child i leaves the parental home at age t, $\theta_{2,it}$ refers to the hazard that father i retires at age t, and u reflects the individual-level, time-invariant, unobserved heterogeneity. The terms $\lambda_1(t)$ and $\lambda_2(t)$ represent the baseline hazard functions for the first and second equations, respectively. These functions capture the time dependence of the transitions into the two states, and they are modeled using a flexible piecewise constant function.[10] Formally, the baseline hazard can be written as follows:

$$\lambda_j(t) = \sum_{s}^{20} \lambda_{js} I_s(t) \tag{2}$$

[8] The term frailty was first suggested by Vaupel et al. (1979) in the context of mortality studies.

[9] These two destination states are assumed to be *absorbing*. Although this assumption seems to be natural for paternal retirement, it could be somewhat less intuitive for nest-leaving because the child could go back to the parents' home after the first move-out. Because information on whether the child returned home is not available in the SHARE data, consistent with previous studies I assume that nest-leaving is an absorbing state.

[10] Alternatively, consistent with Melberg et al. (2010), I employ a cubic function of time, obtaining similar results.

where j ($j = 1, 2$) refers to the equation, s indexes the 1-year intervals, and $I_s(t)$ are dummy variables that take value 1 if the recorded duration is in the s interval. I use an open interval from $s = 19$ onwards because after 19 years the survival and censoring times occur with insufficient frequency to use finer intervals. Because I include a constant in the model, λ_{11} and λ_{21} are normalized to 0.

As for the hazard functions ϕ_1 and ϕ_2, my preferred specification uses a logistic regression. The variable X_i is a matrix of time-invariant, individual controls that may affect the hazard. Specifically, I include household size, a dummy for poor paternal health that takes value 1 if self-reported health is less than good, and an indicator for the father having a college-level education or above (ISCED$\geq$5, tertiary education) or a high school education (ISCED=3 or 4, secondary and post-secondary education). I do not include paternal occupation because of the large fraction of missing observations (approximately 30% of the cross-sectional sample); however, education is strongly correlated with occupation.[11] Both equations also entail a full set of country dummies that capture country-level, time-invariant confounding factors affecting co-residence and paternal retirement. Such factors might include, for example, cross-national differences in preferences and attitudes regarding co-residence and retirement due to discrepancies in cultural and institutional backgrounds. In the variable X_i, I then add birth cohort fixed effects for fathers (in 1-year intervals) to control for possible cohort trends in retirement, i.e., younger cohorts of fathers are likely to retire later, and include controls for the birth order of the child. $Retired_{it}$ is my variable of interest and is equal to 1 if father i is retired at time t. Thus, the treatment effect δ indicates whether the child becomes more likely to leave the nest upon the father's retirement.

With regard to the unobserved heterogeneity terms $u_{1,it}$ and $u_{2,it}$, I follow the latent class approach adopted by Melberg et al. (2010) regarding the impact of cannabis on the risk of consuming hard drugs.[12] Therefore, unobserved heterogeneity is modeled assuming a discrete distribution that has two unrestricted mass points.[13] The intuitive explanation for the presence of these two mass points is that individuals are clustered into two sub-groups that differ in terms of their unobservable propensity for nest-leaving. For instance, one group is composed of individuals who appear more likely to leave the nest later (labeled $k = 1$, Group 1, "low propensity" nest-leaving types or "late" nest-leavers), while the other is more prone to leave the

[11] An additional issue that would arise when controlling for paternal occupation is related to how to deal with fathers who retired many years before their children's nest-leaving. Moreover, because occupation is an individual variable that usually varies over the life cycle, it is not straightforward to identify the occupational spell that really mattered for children's nest-leaving decisions.

[12] In an empirical study, Angelini et al. (2013) follow the latent class approach proposed by Melberg (2010) to evaluate the effect of illiquid assets holding on the probability of becoming a home-owner.

[13] A discrete distribution with two mass points is a flexible parametric distribution since it does not impose any assumptions about the underlying heterogeneity other than that it can be suitably approximated by two latent classes. As noted by Melberg et al. (2010), using more than two latent classes leads to some convergence problems with the algorithm. For this reason, throughout the paper I perform the analysis using two classes.

parental home earlier (labeled $k = 2$, Group 2, "high propensity" nest-leaving types or "early" nest-leavers). Consistent with Melberg et al. (2010), I then allow all the coefficients to differ across the two latent groups; other studies (Pudney 2003; van Ours 2003; Abbring et al. 2005; Salisbury 2012), in which the unobserved heterogeneity is assumed to affect only the constant term, limit this flexibility.

Allowing for correlated unobserved heterogeneity is crucial to the identification of the treatment effect δ, because there may be a potential problem of reverse causality or because there may be individual-level, unobservable factors, such as paternal ability, that determine both paternal retirement and children's decisions to leave home. In particular, if unobservable heterogeneity exists and is ignored, the estimated coefficient may be vulnerable to omitted variable bias. Moreover, the direction of the bias on the timing of nest-leaving would be unclear. For example, higher ability fathers may be more prone to retire later and may provide their children with more opportunities, thus making them more likely to leave home earlier; however, these children may also be more selective and hence more resistant to moving-out. Abbring and Van den Berg (2003) show that an appealing feature of the shared frailty model is that it is identified without the need for any exclusion restrictions or assumptions about the functional form of either the baseline hazard or the joint distribution of the unobserved heterogeneity, as long as the actual timing of the treatment (paternal retirement) is random and is unaffected by the anticipation of the subsequent outcome (children's nest-leaving). However, there may still exist concerns that these two latter conditions are not entirely satisfied in model (1). The main threat to identification is that, even once correlation between frailty terms has been corrected for, the precise timing of the treatment may not occur randomly at year t, i.e., the "no anticipation" assumption is unlikely to hold. As is well known, retirement is a life event that affects various decisions of the family, including consumption, saving, fertility and labor supply.[14] For this reason, children may be able to predict when their fathers will retire, and in response to this expected event, they may modify their lifestyle behaviors and their propensity to become independent. Hence, the anticipation of paternal retirement by adult children would violate one of the key identification assumptions described above, thereby producing biased estimates. To circumvent this problem, I strengthen the identification by providing an exclusion restriction for paternal retirement. The exclusion restriction that I use is based on cross-country early retirement rules and is measured by the indicator $Eligible_{it}$, which equals 1 if father i residing in a given country was eligible for early retirement benefits at age t. These early retirement rules are not only correlated with retirement decisions (Gruber and Wise 2004), but they also provide a potentially valid instrument. Manacorda and Moretti (2006) and Battistin et al. (2009), using an IV strategy, recognize this instrument as valid because

[14] See, for example, the evidence in Battistin et al. (2009), Attanasio and Brugiavini (2003), Battistin et al. (2013) and Liebman et al. (2009).

pension reforms produce variation in paternal retirement that is credibly exogenous and unlikely to be related to unobservable characteristics of the fathers that might explain the different nest-leaving outcomes of their offspring. More importantly, consistent with the previous literature, it seems reasonable to argue that the timing of pension reforms came as a surprise to the fathers directly affected as well as their children. As a result, once the correlation between unobserved factors across both equations and the non-randomness of the timing of the treatment have been corrected for, the remaining difference between the probability of nest-leaving before and after paternal retirement can be interpreted as a causal effect of paternal retirement. To account for within family correlation, all standard errors are clustered at the household level.[15]

To estimate model (1) using maximum likelihood, I expand the data from a cross-section to a panel dataset by exploiting the retrospective information on the year in which the father retired and his child left home. This means that each individual i ($i = 1, ..., n$) is associated with multiple time periods t_i ($t_i = 1, .., T_{is}$), where T_{is} is the total number of years subject i was at risk for the event.[16] For simplicity of exposition, it is useful to distinguish between the two equations ($j = 1, 2$) because they refer to two different outcomes. For the first equation, age 18 is assumed to be the initial period in which the exposure to the risk of nest-leaving begins,[17] such that t_i goes until the age at which the first event is observed (the child's departure from the parental home). If this event does not occur by the end of the survey, then the child is a right-censored observation and t_i lasts until her age at the time of the interview. A similar reasoning applies to the second equation, where I now define the father's age when his child is 18 as the onset of risk,[18] thereby allowing t_i to go until either the father's age at which the second event occurs (his retirement) or the father's age at the time of the survey if the father is employed at the end of the observation period (right-censored case). As a result of this reorganization of the data, I obtain an unbalanced panel, as each individual in the two equations is associated with a different number of time units. Furthermore, a new binary dependent variable y_{it} must be created. If individual i is right-censored, then y_{it} is always equal to zero. If individual i is not censored, y_{it} takes value zero for all but the last of $i's$ periods (i.e., year $1, ..., T_{is} - 1$) and takes value 1 in the last period (i.e., year T_{is}). After having experienced the event, the subject no longer contributes to the risk set and is dropped from the sample (right-truncated cases). One issue that arises in this particular setting is the possibility that paternal retirement occurs after children leave the nest. While the majority of my

[15]Alternatively, given that eligibility rules vary by country and paternal age, I cluster the standard errors by these two dimensions and find that the results remain virtually unchanged.

[16]This construction follows Jenkins (2005) and Melberg et al. (2010).

[17]This starting age for children is consistent with prior research (among others, Manacorda and Moretti 2006; Billari and Tabellini 2008; Becker et al. 2010). In my duration analysis, this assumption implies that children under the age of 18 years are left-truncated.

[18]The vast majority of fathers considered in my sample are at least in their 40s when their child is 18. The rationale for this lower bound is that even fathers in their 40s experience a positive, albeit small, risk of transition into retirement.

sample is composed of fathers who retire after the departure of their children, these time observations would no longer contribute to explaining the hazard of children's nest-leaving, which is the relevant focus of this study. For this reason, these time units are excluded from the second equation. It is worth stressing that one of the main advantages of the duration analysis over a linear IV setting adopted by previous studies is the allowance for censoring, which leads to the elimination of any constraints on the age at which children left their parents' home. For example, Manacorda and Moretti (2006) focus only on youths aged 18 to 30, whereas Billari and Tabellini (2008) and Becker et al. (2010) limit their analysis to adult children aged up to 35 years old.

Consistent with Melberg et al. (2010), the overall log-likelihood function for the bivariate model (1) depends on both the hazard function and the survival function and is given by:

$$\mathcal{L} = \sum_{i=1}^{n} \left\{ \sum_{k=1}^{2} \pi_k \left\{ \sum_{j=1}^{2} \left\{ \sum_{t=1}^{T_{i,j}-d_{i,j}} log\left[1-\theta_{j,it}\right] + d_{i,j} log\left[\theta_{j,it}\right] \right\} \right\} \right\} \tag{3}$$

where the probabilities π_k represent the proportions of the sample composing each latent class, and $d_{i,j}$ is a dummy variable with a value of 1 if individuals are non-censored and a value of 0 if observations are right-censored. It is worth noting that the likelihood of the non-censored individuals differs from that of the censored ones. For the former group, the likelihood is composed of two elements: the survival function from $t = 1$ to $t = T - 1$ and the hazard function in the last period $t = T$ the subject was exposed to the risk. For the latter group, because the censored individuals are never exposed to the event, the likelihood is given solely by the survival function from $t = 1$ to $t = T$.

To maximize (3) under the presence of unobserved heterogeneity, I follow Melberg et al. (2010) and employ the EM algorithm.[19] This method begins with a vector of parameters, α_0, which includes β_1, β_2, δ, γ, $u = (u_1, u_2)$, and the probability weights, $p = (p_1, p_2)$, associated with each of the two latent classes into which my observations may fall. Using these parameters, I create a set of weights for each observation as follows:

$$\pi^0_{k,i} = \frac{p^0_k L^0_{k,i}}{\sum_{k=1}^{2} p^0_k L^0_{k,i}} \tag{4}$$

[19]This is a commonly-used iterative procedure for computing the maximum likelihood estimates when the data are incomplete or have missing values. See, for example, Heckman and Singer (1984) and Ng et al. (1995).

where $\pi_{k,i}$ represents the probability that individual i is assigned to unobserved heterogeneity group k. Thus, individuals are sorted into the most likely latent class to which they belong, based on their observed outcomes. When probabilities of class membership are estimated, I then construct an expected log-likelihood function, which I maximize over α to obtain α_1. Using α_1, I create a new set of weights, π^1, and repeat the algorithm until convergence.

5 Main Results

Before presenting estimates of the model described in the previous section, I provide a visual analysis of the evolution of the estimated hazard functions for nest-leaving and paternal retirement, which are estimated non-parametrically using a kernel-smoothing methodology.[20] In particular, Figure 3 illustrates the pattern of nest-leaving for each European region, with the variable time measured in terms of the number of years since the child turned 18.[21] Overall, this figure notes a number of cross-region differences. These differences include the following: a) in the beginning, in Northern Europe, the hazard of nest-leaving for sons and daughters is considerably higher compared to that in the other country regions; b) in all country groups, daughters initially have significantly higher rates of nest-leaving compared to those of sons;[22] c) in Southern Europe, there is a proportion of adult children who are at high risk of leaving home even when they are in their 40s, thereby providing further evidence about the prolonged cohabitation of Mediterranean youths in their parents' homes.

Finally, Figure A2 in Appendix A displays the dynamics of the hazard for paternal retirement. As expected, in all European regions, the hazard of paternal retirement increases with time. It is also evident that fathers living in Southern Europe are initially at higher risk of transition into retirement. This result is consistent with the evidence indicating that Southern European individuals tend to retire earlier.

[Figure 3 - around here]

[20]This is carried out using the STS package in STATA. A detailed discussion of this package can be found in Cleves et al. (2010).

[21]Notice that the reason why the smoothed hazard estimate is not depicted for $t < 5$ has to do with the choice of the bandwidth.

[22]For each country group, the log-rank and Wilcoxon tests clearly reject the null hypothesis that the survivor functions of sons and daughters are the same.

5.1 Model without Shared Frailty

I begin by estimating a discrete-time duration model for the hazards of children leaving the nest and paternal retirement without correcting for correlated unobserved heterogeneity. Thus, each equation in model (1) is estimated using a separate logistic hazard equation. Table 4 contains the results, with average marginal effects of covariates on the hazard associated with retirement listed next to their average marginal effects on the hazard of children's nest-leaving. In each specification, I include country fixed effects, cohort fixed effects for fathers and a set of controls such as household size, an indicator for paternal poor health and educational achievement. Specifically, in columns 1, 3 and 5, I estimate the equation explaining the probability of leaving the nest for the first time by dividing the sample into Southern, Northern and Central European countries. When examining Southern Europe (see column 1), I find that the estimated effect of paternal retirement is positive and strongly statistically significant (at the 1% level). Paternal retirement implies an increase in the probability of children's nest-leaving of 2.3%. However, when focusing on the Northern and Central European countries (see columns 3 and 5), the coefficient on paternal retirement becomes insignificant, and the magnitude is reduced to 0.017 and 0.003, respectively. As expected, in each macro-region, the eligibility status for early retirement benefits matters for the hazard of paternal retirement (see columns 2, 4 and 6). While eligible fathers are more likely to retire, the differences in the magnitude of the coefficient on paternal eligibility are remarkable, ranging from 3.2% in Northern Europe to 8.9% in Southern Europe. In columns 7 and 8, I separately estimate the two equations in model (1) using the pooled sample. Interestingly, the point estimate of the coefficient of interest remains positive and significant, with a magnitude of 0.021. It seems clear that this significant impact on the full sample is driven by the highly significant effects of paternal retirement obtained from the regression on the sample of Southern European countries (see column 1). Moreover, I find that coefficients on household size are quite small in magnitude and change signs across the various subsamples for both risks, indicating that household size is not the most important factor for children's nest-leaving or paternal retirement. A similar observation applies to the coefficients on fathers' poor health, which seems to play a very limited role in explaining these two risks. Overall, it is difficult to extrapolate any systematic or interesting patterns from these coefficients.

In sum, although these correlations may suffer from problems of confounding, they provide a first indication that paternal retirement is associated with a higher probability of first nest-leaving by children (first equation) only in the Mediterranean countries, and that early retirement rules strongly predict the hazard of paternal retirement (second equation). In the next subsection, I attempt to establish whether this positive correlation has a causal interpretation.

[Table 4 - around here]

5.2 Model with Shared Frailty

The primary concern about the point estimates presented in Table 4 is that they may not adequately account for the correlation between unobserved characteristics that affect children's nest-leaving and unobserved factors that determine paternal retirement, thereby generating omitted variable bias.

To address this concern, I allow for the possibility of correlated unobserved heterogeneity terms across both equations by using the latent class approach suggested by Melberg et al. (2010), in which individuals are divided into two sub-groups of the population. Table 5 presents the estimation results of logistic regressions on the hazard of nest-leaving. As mentioned in the previous subsection, average marginal effects are calculated for each European region (columns 1 to 9) and for the pooled sample (columns 10 to 12). To account for unobservable differences between Southern, Northern and Central Europe, I allow the frailty to vary across these regions. Thus, I separately estimate the probability weights attached to the unobserved heterogeneity Group 1 and Group 2 for each European region as well as for the full sample. The estimated probabilities, $\hat{\pi}_1$ and $\hat{\pi}_2$, are also listed in Table 5.

[Table 5 - around here]

In particular, in columns 1 to 3, I focus on Southern European countries. To facilitate comparisons, in column 1, I report the average marginal effects corresponding to the model in which unobserved heterogeneity is ignored (see, also, column 1 of Table 4). In columns 2 and 3, I present the same predicted effects when unobserved heterogeneity is allowed for by using the probabilities of belonging to Group 1 and Group 2 as weights, respectively. This means that a different logistic hazard regression is estimated for each of the two groups. The results suggest that paternal retirement is a statistically significant predictor of children's nest-leaving. For those belonging to Group 1, the treatment effect of paternal retirement is positive and strongly statistically significant (at the 1% level). With respect to the magnitude, paternal retirement increases the probability of children's first nest-leaving by 5.5%. The treatment effect remains highly significant, albeit quantitatively less important (1.4%), for those who belong to Group 2.

To learn more about the characteristics of the two groups, Table 6 displays summary statistics on selected covariates.[23] Specifically, individuals in the sample with a predicted probability of falling into Group 1 below the median are assigned to that group, whereas the remaining individuals are placed in Group 2. As evidenced in Panel A (Southern Europe), these two groups differ substantially with respect to the proportion of retired fathers. For Group 1, this proportion is approximately 12% greater than the mean of the entire sample (25% versus 22%) and approximately 27% greater than the mean of Group 2 (25% versus 19%). Such significant differences in the fraction of retired fathers can contribute to explaining why young people in Group 1 ("low propensity" nest-leaving types) are much more affected by paternal retirement than their counterparts in Group 2 ("high propensity" nest-leaving types). Interestingly, these two groups also differ significantly in a number of other observable characteristics, such as educational outcomes and children's age at time of leaving home. For instance, adult children in Group 1 are more likely to leave the parental home later and have better outcomes in terms of their own and their fathers' education.

[Table 6 - around here]

When restricting the analysis to Northern Europe (columns 4 to 6 of Table 5) and Central Europe (columns 7 to 9 of Table 5), I find that the dummy variable for paternal retirement is no longer statistically significant in any of the two unobserved groups. This lack of significance can likely be explained by looking at the differences in the fraction of adult children who left the nest after paternal retirement. Table 3 reveals that such differences across European regions are enormous, ranging from 42% in Southern Europe to 15% in Central Europe and to 6% in Northern Europe. In other words, when fathers retire, only a very limited share of adult offspring in Northern and Central European countries is still living with their parents, thus raising concerns about the lack of power in my identification strategy for these two macro-regions.

Descriptive statistics (see Panel B for Northern Europe and Panel C for Central Europe in Table 6) confirm that young people in Group 1 can still be viewed as "low propensity" nest-leaving types, with a much larger fraction of retired fathers. To be more precise, in Northern and Central Europe, these fractions are approximately 60% higher compared to the mean of the full sample, and they are four times larger when they are compared to the mean of the respective Group 2. Moreover, in Northern and Central Europe, young people belonging to Group 1 tend to leave the nest later relative to their counterparts in Group 2.

In columns 10 to 12 of Table 5, I report the estimated coefficients obtained from the pooled sample.

[23] To conserve space, household size and paternal health status are not reported. However, they are not found to display any significant differences between Group 1 and Group 2.

While treatment effects of paternal retirement are positive and significant for Group 1, they are close to zero for Group 2. Similar to the analysis ignoring unobserved heterogeneity (see column 7 in Table 4), it seems evident that the significant effect for Group 1 on the pooled sample is driven by the strongly significant effect obtained for the same group in Southern Europe. As expected, when examining the descriptive statistics (see Panel D in Table 6), individuals in Group 1 are characterized by a markedly larger share of retired fathers compared to those belonging to Group 2 (17% higher with respect to the mean of the full sample and 40% higher relative to the mean of Group 2) and are more likely to leave the nest later. It is also worth noting that the estimated probability of belonging to Group 1 varies substantially with the associated macro-region and is much higher in Southern Europe (33%) as opposed to Northern (6%) and Central (21%) Europe. This result confirms that young people sharing some latent characteristics that make them belong to the latent class of "late" nest-leavers (Group 1) are concentrated in Southern European countries. Overall, the evidence presented above suggests that, although quantitatively small, there are positive causal effects of paternal retirement on the timing of children's nest-leaving only for Southern European countries. The non-significant effects obtained for Northern and Central Europe are presumably because most youths have already left their parental homes at the time of their fathers' retirement. In the discussion section, I explain why these findings may differ so largely by European region.

Moreover, Table A1 in Appendix A presents the estimates for the hazard of paternal retirement. In accordance with the model in which unobserved heterogeneity is not allowed for (see Table 4), the coefficients on eligibility status reveal the significant influence of eligibility rules on actual retirement. These findings are consistent with the available empirical evidence on the relevance of early retirement incentives (Gruber and Wise 2004). Interestingly, in the Southern European countries, the strength of the estimated effects is larger compared to that of the other country groups. This may be because Italian, Spanish and Greek workers have more financial incentives to retire early due to their particularly generous early retirement benefits with respect to those of other European regions.

Finally, in an attempt to disentangle the treatment effects of paternal retirement on sons from the effects on daughters in Southern Europe, I separately consider the samples of male and female children. The results for sons and daughters are presented in Table A2 in Appendix A. When restricting the analysis to sons (see columns 2 and 3), the coefficient on paternal retirement varies between 5.5% for individuals in Group 1 and 1.3% for those belonging to Group 2. A similar pattern is observed in the regressions for daughters (see columns 5 and 6), with the difference being that the magnitude for daughters in Group 1 is slightly smaller compared to sons in Group 1 (4.9% vs. 5.5%) and the treatment effect for daughters in Group 2 is no longer significant, which may be partly due to the smaller sample size. However, these differences between sons and daughters are not significantly different from zero. In Tables A3 and A4 in Appendix A, I show that paternal retirement has no significant positive effects on sons and daughters in Northern and Central Europe.[24]

6 Sensitivity Analysis

Before proceeding to discuss and test empirically the potential mechanisms, I perform a variety of robustness checks to determine if the results change when I modify the estimation strategy or use a different specification of the model (see Tables 7 and 8).

[Table 7 - around here]

[Table 8 - around here]

6.1 Instrumental Variable Analysis

Although the bivariate hazard model described in section 4 provides the most appropriate description of the relationship between paternal retirement and the timing of children's nest-leaving, there may still be concerns regarding the sensitivity of my results to their stability or to the parametric assumptions made in the estimation. For example, as noted by Melberg et al. (2010), in latent class models, the convergence of the likelihood can be vulnerable to problems due to local optima. To address this concern, I estimate the following linear version of model (1) using two stage least squares (2SLS):

$$Pr(L_{it} = 1) = \alpha + \beta Retired_{it} + \gamma X_i + \epsilon_{it} \tag{5}$$

where the treatment dummy $Retired_{it}$ and the variable X_i are defined in the same way as in Section 4. Here, the outcome variable L_{it} is a dummy taking the value 1 if a child i residing in a given country left the parental home at age t. Following Manacorda and Moretti (2006), I focus on youth aged 18 to 30 years.[25] Finally, ϵ_{it} represents an idiosyncratic error term, which is presumably correlated with the outcome

[24] In each macro-region, descriptive statistics for Group 1 and Group 2 confirm the conclusions obtained for the full sample (see Table 6). These tables are available from the author upon request.

[25] As a robustness check, I considered children aged 18 to 35, obtaining similar results.

variable because it embodies unobserved factors of fathers, including ability, which might affect children's home-leaving decisions. Consistent with previous analysis, I would expect to find a positive and significant effect of paternal retirement only in Southern Europe.

I identify the causal effect of paternal retirement on children's nest-leaving using cross-country changes in eligibility rules for early retirement benefits for the period 1961 to 2007 as an instrument for paternal retirement. As discussed in Section 4, this instrument is recognized to be relevant and arguably exogenous to children's living arrangements. In this setup, the first stage regression is given by:

$$Retired_{it} = \delta_0 + \delta_1 Eligibility_{it} + \pi X_i + \nu_{it} \tag{6}$$

where the dummy $Eligibility_{it}$ represents the instrument introduced in Section 4. It is important to acknowledge that this instrumental variable strategy is relevant only for the subset of compliers, i.e., fathers who retire as a consequence of early retirement schemes.

Panel A of Table 7 reports the 2SLS results. The treatment dummy on paternal retirement is positive and significant at the 5% level only for Southern Europe (see column 1). This dummy variable, however, becomes non-significant and negative for Northern and Central European countries (see columns 2 and 3). Panel B contains the first-stage results. As expected, these estimates indicate that eligibility for early retirement benefits is an important determinant for paternal retirement. Altogether, the IV analysis lends some additional evidence that only for Southern Europe there is a positive causal relation between paternal retirement and children's nest-leaving, a finding that calls for further explanation.

6.2 Additional Sensitivity Checks

As a further check, I investigate the robustness of my estimates to the use of an alternative definition of the treatment dummy for paternal retirement. A common concern is that as children age, they are more likely to leave the parental home regardless of their fathers' retirement status. In order to allow for this possibility, I define a time frame of three years, and construct a binary variable that is set to 1 if the father retired prior to the child's first move-out within the time frame[26] and 0 otherwise. This approach is similar in spirit to that of van Ours (2003), who refers to this time frame as the "incubation period" to identify

[26]The results are similar when considering time frames of 2 or 4 years. These tables are available from the author upon request.

a gateway effect of cannabis on cocaine. The results are presented in Panel A of Table 8.[27] Reassuringly, these parameter estimates resemble those obtained in the benchmark specification (see Table 5), with the only difference being that in Southern Europe the magnitude of the estimated effects of paternal retirement becomes slightly smaller.

An additional concern is that the father may start receiving pension benefits only some years after his retirement year. To check the robustness of my results, I use information on the year in which the father first received pension benefits.[28] Thus, I employ an alternative treatment indicator variable that takes value 1 if the father i collects pension income at time t and re-estimate my model. As the coefficients reported in Panel B show, the evidence remains substantially unchanged relative to the benchmark specification, although for individuals in Group 1 the magnitude of the coefficient of interest is slightly reduced.

7 Discussion

In the literature on moving-out decisions, what remains largely unexplained is the mechanism regulating the positive causal relationship between paternal retirement and children's nest-leaving. In this section, I start to fill this gap by focusing the analysis on Italy, Greece and Spain, countries for which I found a positive causal effect of paternal retirement. A unique feature of these Southern European countries is that they can be divided into two groups. One group is composed of Italy and Greece, where there is a large bonus payment at the time of retirement that amounts to approximately three times the gross annual salary. The second group includes only Spain, where such severance payment does not exist.[29] My information on severance arrangements is drawn from Holzmann et al. (2011), from personal communications with national experts and from other country-specific sources.[30] As previously mentioned, the literature would attribute this causal relationship mainly to two competing mechanisms. To provide an empirical test for these two

[27]To save space, this table reports the estimated coefficients only for the hazard of children's nest-leaving, which is the outcome of main interest in this paper. Results for the hazard of paternal retirement remain substantially unchanged and are available from the author upon request.

[28]The following question was asked: "In which year did you first receive this pension?". Approximately 20% of the cross-sectional sample reports a retirement year that differs from the year in which pension benefits were first received.

[29]García-Gómez et al. (2013) document that Spanish employed who leave employment and transit into unemployment may receive a severance payment from the employer. To overcome this issue, I excluded from the sample Spanish individuals who declare themselves as retired because they were made redundant. The exact question to elicit this information is stated as follows: "Please look at card 21. For which reasons did you retire?". However, as shown in Table A5 in Appendix A, the main results still hold if these individuals are included.

[30]For Italy, information on retirement severance payment is obtained from Miniaci et al. (2003). For Greece and Spain, I acknowledge that institutional details have been integrated by personal communications with Olympia Bover, Pilar García-Gómez, Athanasios Tagkalakis and Platon Tinios.

mechanisms, I adopt a differences-in-differences strategy, where Italy and Greece constitute the treatment group and Spain is the control group, "unaffected" by the lump-sum payment upon retirement. The key identification assumption for Spain to be a valid control group is that children's nest-leaving behavior of Spain and Italy and Greece would have followed similar trends over time, in the absence of retirement severance pay. It is plausible to justify this assumption, given that, conditional on country fixed effects, the Southern European countries included in my sample were not only undergoing similar economic conditions and were very similar in terms of welfare state regime, family structure and culture, but they also had similar demographic patterns of intra-generational co-residence and patterns of support for the elderly (Bolin et al. 2008).

To the extent that the Manacorda and Moretti mechanism is at play, I expect paternal retirement to bribe Italians and Greek adult children to stay at home longer as a consequence of the positive shock to the family's liquidity associated with the retirement severance payment. However, the results reported in Table 9 (columns 1 to 3) are in the opposite direction. For individuals belonging to Groups 1 and 2, the dummy variable for paternal retirement remains positive and highly statistically significant (at the 1% level), with a magnitude of 6.1% and 1.5%, respectively. This result indicates that liquidity problems faced by fathers at the time of retirement do not provide an entirely satisfactory explanation. On the other hand, if retirement severance payment mattered, as stressed by Battistin et al. (2009), I would expect to find no evidence of significant effects of paternal retirement for Spain. Nevertheless, the coefficient estimates presented in columns 4 to 6 largely contradict the prediction of this second hypothesis: for individuals in Group 1, the estimated coefficient on paternal retirement retains its significance, whereas for those in Group 2, the magnitude of the coefficient of interest remains substantially unchanged with respect to the estimate in column 3, but is significant only at the 10% level. This result is what I expected given the substantial reduction in sample size.

[Table 9 - around here]

The main conclusion that I draw from this empirical test is that the decline in children's cohabitation at paternal retirement cannot be entirely ascribed to liquidity problems or a boost in family's income due to severance payment.

One may still be concerned that Spain is not a comparable control group or that Italy and Greece do not represent an appropriate treatment group because self-employed workers are not entitled to retirement severance payment. In order to address these concerns, I propose an additional test: for Italy and Greece, I use the employed as the treatment and self-employed[31] as the control group. Descriptive statistics presented in Table A6 in Appendix A demonstrate that employed and self-employed do not differ significantly in a large number of observable characteristics, thus providing empirical evidence in support for the claim that self-employed workers are a valid counterfactual. The results reported in Table 10 indicate that there are positive causal effects of paternal retirement on the timing of children's nest-leaving for the treatment (columns 1 to 3) and control group (columns 4 to 6), which I interpret as corroborating evidence that the drop in paternal post-retirement income or retirement severance payment do not provide a satisfactory explanation for the mechanism behind children's nest-leaving upon paternal retirement.

[Table 10 - around here]

For this reason, it seems worthy to investigate other potential channels. In their study on the intergenerational effects of Italian pension reforms on fertility, Battistin et al. (2013) argue that the rise in retirement age has reduced the amount of informal child care provided by grandparents, which in turn has determined an increase in the children's age at first child and of home leaving. While this scenario can be applied to other Southern European countries, including Spain and Greece, there is a general consensus that grandmothers are the main provider of informal child care arrangements to their grandchildren (see, for instance, Richter et al. 1994). Battistin et al.'s findings move in the same direction by showing that it is the grandmothers' provision of informal child care that plays a primary role in their children's fertility decisions. Therefore, this conjecture does not seem particularly relevant for this study because, as discussed earlier in the paper, female partners are excluded from the analysis. Nevertheless, it is still possible that individuals in a couple plan their retirement closely together (Stancanelli 2012). To address this concern, in Table A7 in Appendix A, I demonstrate that, even when focusing only on fathers whose spouses have never worked, there still exists a positive and quantitatively similar causal effect of paternal retirement on the hazard of children's nest-leaving. It may be argued, however, that some evidence in favor of this interpretation cannot be totally ruled out given that for individuals in Group 2 the coefficient on paternal retirement is close to zero, thus revealing the potential presence of an effect originating from the grandparents' supply of informal child care on the top of other unexplained factors for the "early" nest-leaving types.

To sum up, these findings indicate that the decline in children's co-residence immediately after paternal retirement does not appear to be justified by changes in parental resources or in the supply of informal

[31] As in Angelini et al. (2013), the term "self-employed" refers to those individuals who have been self-employed at any stage during their career. To recover this information, I use SHARE data provided by the job episodes panel. See Brugiavini et al. (2013) for a full description of this panel dataset.

child care provided by grandparents. The evidence, therefore, points to the hypothesis that there may exist a number of preference-related reasons that concern the negative externalities between retired fathers and their offspring. Anecdotal evidence suggests that children's departure from parental home can be explained by the conflicting relationship with their fathers, likely resulting from the paternal presence in the house upon retirement. Unfortunately, it is difficult to verify this hypothesis with my data because, as already mentioned, the SHARE questionnaire does not provide information regarding the reason for children's nest-leaving. To partially address this data limitation, I can, however, use a measure for overcrowding at the time of children's nest-leaving as a proxy for the negative externalities in preferences. More specifically, I create an indicator variable that is equal to 1 if the number of rooms per person is below the median for the given country.[32] I then examine whether the treatment effects of paternal retirement on children's nest-leaving vary with the presence or absence of overcrowding in the house. If the estimated coefficients are statistically different with a larger magnitude in the presence of overcrowding, then it does appear that preference-related reasons are likely to play a role in explaining children's decisions to leave the parental home. The parameter estimates are contained in Table 11. Overall,

[Table 11 - around here]

Although it is not a contribution of this paper, it remains to be explored why the coefficient on paternal retirement is not statistically significant in Northern and Central Europe. As argued in Section 5, a plausible explanation is that there is not enough power in my identification strategy for these two macro-regions because only a very limited share of adult offspring left their parental home after paternal retirement. However, this finding raises the issue of why young people living in Northern and Central Europe leave home much earlier relative to their counterparts in Southern Europe. Such disparities in the age of home-leaving can be reconciled with the strand of the literature that analyzes the presence of a European North-South gradient in family ties (see, for instance, Reher 1998; Alesina and Giuliano 2011), labor market conditions (Card and Lemieux 2000) and cross-regional differences in housing markets (Alessie et al. 2006).

[32]To be more precise, SHARE provides information on the number of rooms available in the household's accomodation at the interview year of wave 2. SHARE also contains information on the number of years of residence in the current accomodation, which enables me to retain only child-father pairs where the current accomodation was the same to that at the time of children's nest-leaving (approximately 84% of the cross-sectional sample). However, SHARE does not give information on the number of persons in the household at the time of children's nest-leaving. To overcome this lack of information, I created a proxy variable by summing up the household size at the interview year of wave 2 and the number of children that have already left home at the year of the interview.

8 Conclusion

In this paper, I examine the relationship between paternal retirement and the timing of housing emancipation of young adults in Europe, with the aim of testing empirically which of the mechanisms proposed in the literature dominates in practice. Taking advantage of the retrospective dimension of my micro data, I use a bivariate discrete-time hazard model with shared frailty and exploit cross-country variation in early retirement legislation. Overall, my regression results suggest that there is a significant influence of paternal retirement on the probability of first nest-leaving of children living in Southern European countries. However, there is no evidence of significant effects on children residing in Northern and Central European countries. I interpret this evidence as indicating that paternal retirement is a relevant explanatory variable of co-residence decisions only in Southern Europe, once differences in institutions, culture and other unobservables are controlled for.

To shed some light into the mechanism, I provide an empirical test for the two main competing channels by which paternal retirement may be thought to affect children's co-residence. Comparing my cross-country evidence for Southern Europe with important country-specific evidence obtained for Italy from two other studies (Manacorda and Moretti 2006; Battistin et al. 2009), it seems plausible to conclude that the increase in children's nest-leaving around paternal retirement does not seem to be driven by changes in parental economic resources. In addition, I discuss the plausibility of the hypothesis proposed by Battistin et al. (2013); however, I do not find conclusive evidence that the supply of informal child care provided by grandparents is a major determinant for children's moving-out. Rather, one needs to look for channels involving negative externalities in preferences between fathers and their children.

Empirical evidence that paternal retirement can affect children's nest-leaving has relevant policy implications. It is well-known that because the population is rapidly aging in Europe, it is becoming increasingly important to maintain the long-term financial sustainability of pension systems. To achieve this goal, in the recent past European governments have primarily adopted a number of pension reforms that have raised the retirement age or removed financial incentives to early retirement. However, the results of this paper suggest that in Southern Europe policy makers should also be aware that there may be potential unintended and undesirable consequences of pension reforms on moving-out decisions of young people.

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Figures and Tables

Figure 1: Children's nest-leaving mean age, by European region

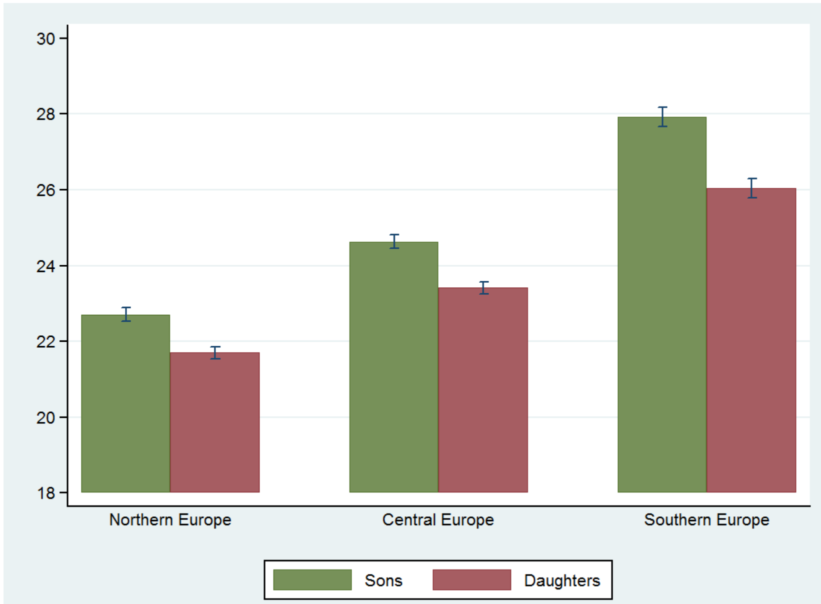
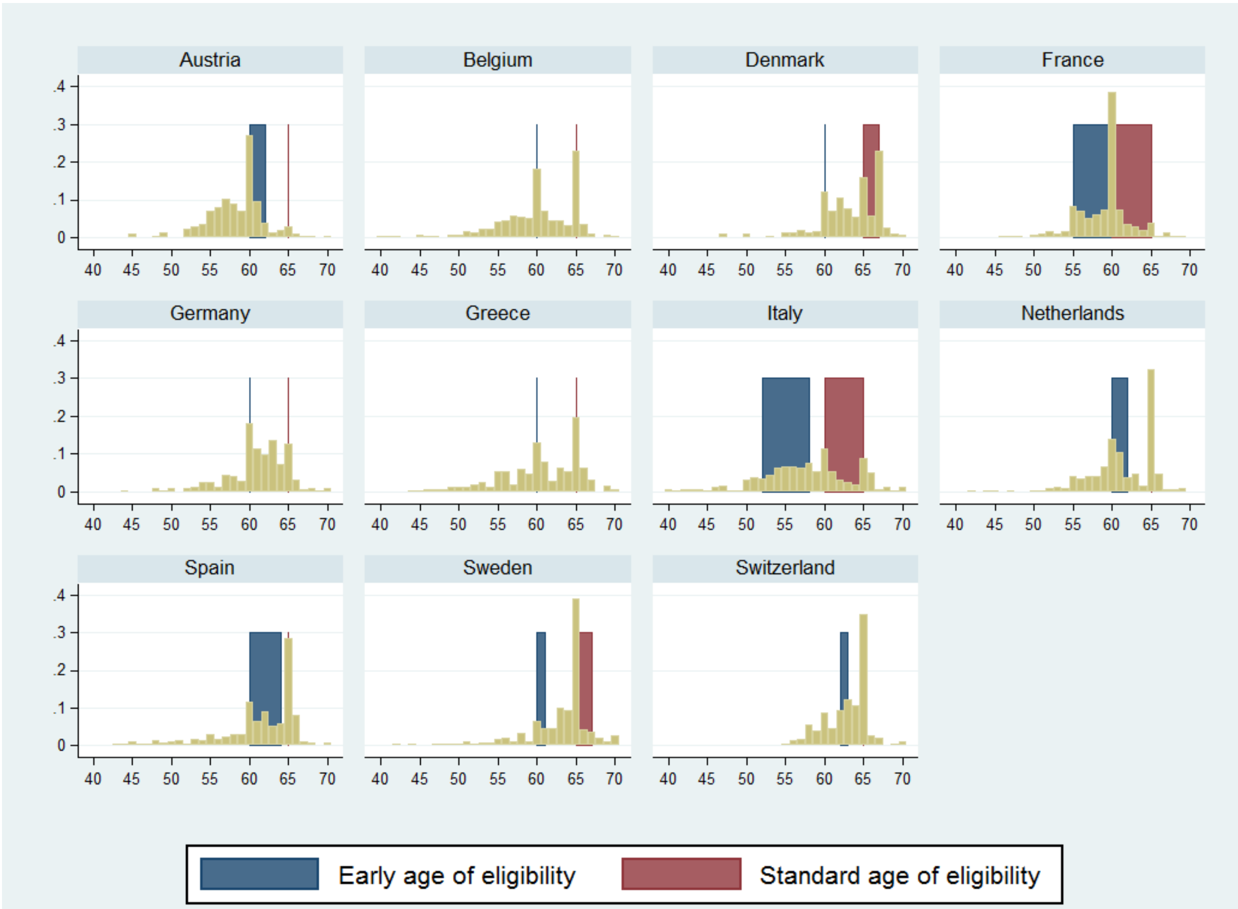
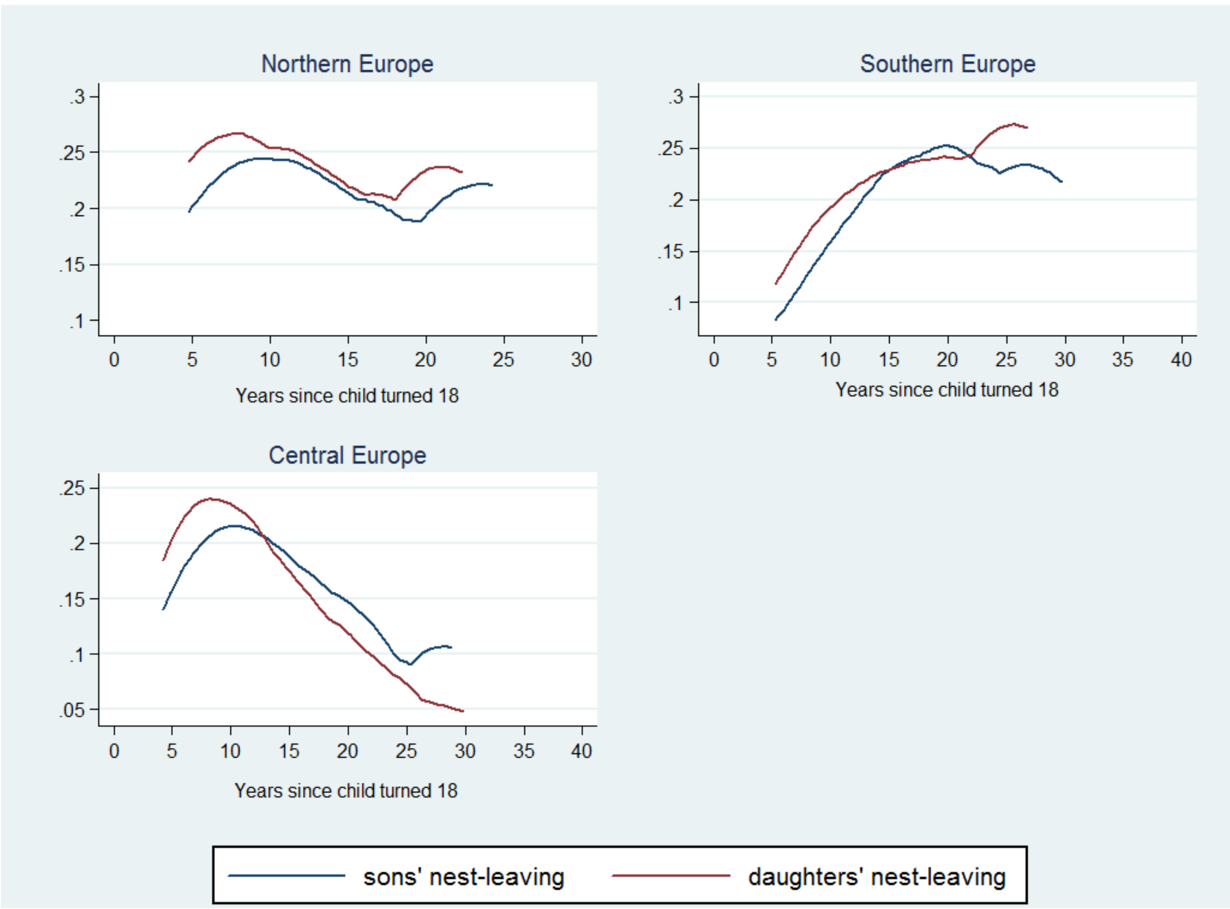


Figure 2: Histograms of father's retirement age, by country



Notes: Source: Angelini et al. (2009), Mazzonna and Peracchi (2012), Gruber and Wise (2004) and Duval (2003). The vertical blue and red lines, respectively, mark the eligibility ages for early and normal retirement age, whereas the blue and red areas represent changes in the eligibility ages for the cohorts in my sample.

Figure 3: Empirical hazard rate of children's nest-leaving and fathers' retirement, by European region



Notes: This figure plots the estimated hazard function of nest-leaving of children and that of paternal retirement by European region. These hazard functions are estimated using a nonparametric kernel-smoothing methodology (STS package in STATA). Notice that the reason why the smoothed hazard estimate is not depicted for $t < 5$ has to do with the choice of the bandwidth. Recall that children who were less than 18 are left-truncated.

Table 1: Sample of Fathers and Children, by Country

Sample	Fathers	Sons	Daughters	Total
Austria	242	278	255	533
Belgium	664	704	686	1,390
Denmark	407	478	421	899
France	543	588	606	1,194
Germany	568	585	546	1,131
Greece	300	339	298	637
Italy	629	655	673	1,328
Netherlands	518	593	590	1,183
Spain	361	442	385	827
Sweden	455	573	464	1,037
Switzerland	248	290	271	561
Total	4,935	5,525	5,195	10,720

Notes: This table reports the observations from the cross-sectional sample before reshaping it as a longitudinal dataset. All of the samples contain fathers for whom information on education is not missing and exclude children who were less than 18.

Table 2: Summary Statistics, Sample of Fathers and Children

Variable	Observations	Mean	Std. Dev.
Sons			
Age	5,525	38.15	8.22
Nest-leaving age	5,525	24.92	4.83
High school	5,525	0.46	0.50
College or more	5,525	0.37	0.48
Married	5,525	0.72	0.45
Never left home	5,525	0.01	0.10
Daughters			
Age	5,195	37.77	8.42
Nest-leaving age	5,195	23.61	4.30
High school	5,195	0.46	0.50
College or more	5,195	0.40	0.49
Married	5,195	0.77	0.42
Never left home	5,195	0.01	0.10
Fathers			
Age	4,935	66.89	8.60
Retired	4,935	0.72	0.45
Working	4,935	0.28	0.45
Retirement age (retired)	3,553	60.34	4.73
High school	4,935	0.34	0.47
College or more	4,935	0.23	0.42
Bad health	4,935	0.29	0.45
Household size	4,935	2.23	0.57

Notes: This table reports the observations from the cross-sectional sample before reshaping it as a longitudinal dataset. All of the samples contain individuals for whom information on children's nest-leaving age and paternal education is not missing and exclude children who were less than 18. The paternal sample consists of all individuals who are either working or retired.

Table 3: Summary Statistics, Children who left home after paternal retirement

Sample	Sons			Daughters				Overall		
	Obs.	Mean	Std. Dev.	Obs.	Mean	Std. Dev.	p-value difference	Obs.	Mean	Std. Dev.
Southern Europe	1,436	0.45	0.49	1,356	0.38	0.48	0.00	2,792	0.42	0.49
Northern Europe	1,644	0.07	0.26	1,475	0.05	0.22	0.00	3,119	0.06	0.24
Central Europe	2,445	0.16	0.37	2,364	0.13	0.33	0.00	4,809	0.15	0.35
Overall	5,525	0.21	0.41	5,195	0.17	0.38	0.00	10.720	0.19	0.39

Notes: This table reports the observations from the cross-sectional sample before reshaping it as a longitudinal dataset. All of the samples contain individuals for whom information on children's nest-leaving age and paternal education is not missing and exclude children who were less than 18.

Table 4: Model without shared frailty - Determinants of the Hazard of Nest-Leaving and Retirement

Sample	**Southern Europe**		**Northern Europe**		**Central Europe**		**Full sample**	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Outcome	Nest-leaving	Ret.	Nest-leaving	Ret.	Nest-leaving	Ret.	Nest-leaving	Ret.
Father is retired	0.023***		0.017		0.003		0.021***	
	(0.005)		(0.030)		(0.009)		(0.005)	
Father is eligible		0.089***		0.032***		0.043***		0.055***
		(0.005)		(0.003)		(0.004)		(0.002)
Household size	-0.006**	0.002	0.013***	-0.001	-0.012***	-0.006**	-0.008***	-0.001
	(0.003)	(0.002)	(0.003)	(0.002)	(0.004)	(0.003)	(0.003)	(0.001)
Bad health (father)	0.005	0.004	-0.029***	0.004*	-0.005	0.002	-0.003	0.003*
	(0.004)	(0.004)	(0.010)	(0.003)	(0.006)	(0.003)	(0.003)	(0.002)
Country F.E.	YES	YES	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES	YES	YES	YES	YES
Log-likelihood	-7,883	-3,185	-6,950	-710	-12,236	-2,298	-27,684	-6,485
Observations	24,530	18,806	13,197	12,597	28,698	23,682	66,425	55,085

Notes: Logit estimations; average marginal effects reported. The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. All specifications include time dummies representing duration dependence. Standard errors in parentheses are clustered at the household level.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table 5: Model with shared frailty - Determinants of the Hazard of Nest-Leaving, by European region and overall sample

Sample	**Southern Europe**			**Northern Europe**			**Central Europe**			**Full sample**		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Father is retired	0.023***	0.055***	0.014***	0.017	0.023	-0.097	0.003	0.009	-0.026	0.021***	0.026***	0.002
	(0.005)	(0.007)	(0.005)	(0.030)	(0.025)	(0.067)	(0.009)	(0.009)	(0.021)	(0.005)	(0.007)	(0.005)
Household size	-0.006**	-0.011***	-0.004	0.013***	0.032***	-0.027	-0.012***	-0.011***	-0.001	-0.008***	0.002	-0.010***
	(0.003)	(0.004)	(0.003)	(0.003)	(0.005)	(0.019)	(0.004)	(0.004)	(0.004)	(0.002)	(0.003)	(0.002)
Bad health (father)	0.005	0.015***	0.008*	-0.029***	-0.013	-0.113**	-0.006	-0.007	-0.004	-0.003	0.004	-0.003
	(0.004)	(0.006)	(0.004)	(0.009)	(0.010)	(0.046)	(0.006)	(0.006)	(0.006)	(0.004)	(0.005)	(0.004)
Mass points :												
$\hat{\pi}_1$	0.334			0.065			0.210			0.319		
	(0.325)			(0.196)			(0.290)			(0.336)		
$\hat{\pi}_2$	0.666			0.935			0.790			0.681		
	(0.325)			(0.196)			(0.290)			(0.336)		
Wald test p-value for diff. btw. $\delta^{(2)}$ *and* $\delta^{(3)}$	0.000											
Country F.E.	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Log-likelihood	-7,883	-2,726	-4,905	-6,950	-5,391	-1,444	-12,236	-9,851	-2,022	-27,684	-8,522	-17,856
Observations	24,530	24,530	24,530	13,197	13,197	10,623	28,698	28,698	22,114	66,425	66,425	57,267

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1, 4, 7, 10), and weighted, using as weights $\hat{\pi}_1$ (col. 2, 5, 8, 11) or $\hat{\pi}_2$ (col. 3, 6, 9, 12). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. All specifications include time dummies representing duration dependence. Notice that observations for which the probability of belonging to Group 1 and Group 2 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.

* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table 6: Model with shared frailty - Differences between clusters, by European region and full sample

Variable	**Group 1**		**Group 2**			**Full sample - No Het.**	
	Mean	Std. Dev.	Mean	Std. Dev.	p-value difference	Mean	Std. Dev.
Panel A: Southern Europe ($\hat{\pi}_1 = 0.33, \hat{\pi}_2 = 0.67$)							
Father is retired	0.247	0.431	0.195	0.397	0.000	0.221	0.415
Male (child)	0.570	0.495	0.578	0.494	0.180	0.574	0.495
Married (child)	0.834	0.372	0.831	0.374	0.518	0.833	0.373
High school (father)	0.150	0.357	0.136	0.342	0.001	0.143	0.349
College or more (father)	0.084	0.277	0.073	0.259	0.000	0.078	0.268
High school (child)	0.403	0.490	0.423	0.494	0.001	0.413	0.492
College or more (child)	0.301	0.459	0.235	0.424	0.000	0.268	0.442
Nest-leaving age	30.078	5.268	29.325	5.262	0.000	29.701	5.278
Panel B: Northern Europe ($\hat{\pi}_1 = 0.07, \hat{\pi}_2 = 0.93$)							
Father is retired	0.072	0.259	0.018	0.132	0.000	0.045	0.207
Male (child)	0.610	0.488	0.563	0.496	0.000	0.587	0.492
Married (child)	0.708	0.455	0.678	0.467	0.000	0.693	0.461
High school (father)	0.277	0.448	0.350	0.477	0.000	0.314	0.463
College or more (father)	0.213	0.409	0.282	0.450	0.000	0.247	0.431
High school (child)	0.463	0.499	0.459	0.498	0.656	0.461	0.498
College or more (child)	0.350	0.477	0.388	0.487	0.000	0.369	0.482
Nest-leaving age	26.308	5.196	23.704	4.104	0.000	25.006	4.858
Panel C: Central Europe ($\hat{\pi}_1 = 0.21, \hat{\pi}_2 = 0.79$)							
Father is retired	0.159	0.366	0.040	0.197	0.000	0.100	0.299
Male (child)	0.580	0.494	0.539	0.498	0.000	0.560	0.496
Married (child)	0.706	0.456	0.715	0.451	0.084	0.711	0.453
High school (father)	0.445	0.497	0.429	0.495	0.005	0.437	0.496
College or more (father)	0.272	0.445	0.253	0.435	0.000	0.263	0.440
High school (child)	0.511	0.500	0.456	0.498	0.000	0.483	0.499
College or more (child)	0.430	0.495	0.488	0.500	0.000	0.459	0.498
Nest-leaving age	29.024	7.055	25.326	4.286	0.000	27.175	6.122
Panel D: Full sample ($\hat{\pi}_1 = 0.32, \hat{\pi}_2 = 0.68$)							
Father is retired	0.172	0.377	0.123	0.328	0.000	0.147	0.354
Male (child)	0.574	0.495	0.561	0.496	0.000	0.567	0.495
Married (child)	0.724	0.447	0.779	0.415	0.000	0.751	0.432
High school (father)	0.334	0.472	0.277	0.448	0.000	0.306	0.460
College or more (father)	0.217	0.412	0.164	0.370	0.000	0.190	0.393
High school (child)	0.469	0.499	0.438	0.496	0.000	0.453	0.498
College or more (child)	0.392	0.488	0.351	0.477	0.000	0.371	0.483
Nest-leaving age	28.560	6.299	26.807	5.172	0.000	27.684	5.829

Notes: Descriptive statistics are computed using the longitudinal sample. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. Observations with an estimated probability below the median are assigned to Group 1, whereas the remaining individuals are assigned to Group 2.

Table 7: Effects of paternal retirement, IV analysis

	(1)	(2)	(3)	(4)
Sample	South	North	Central	Overall
Panel A: 2SLS				
Dep. Var.: Child leaves home				
Father is retired	0.159**	-0.253	-0.046	0.042
	(0.075)	(0.235)	(0.066)	(0.066)
Household size	-0.007**	-0.022***	-0.033***	-0.022***
	(0.003)	(0.004)	(0.011)	(0.007)
Bad health (father)	0.014**	0.000	0.014	0.010
	(0.007)	(0.012)	(0.012)	(0.008)
Observations	34,462	37,135	54,976	126,573
R^2	0.223	0.201	0.221	0.258
First stage F statistic	82.06	9.12	98.99	159.68
Panel B: First stage				
Dep. Var.: Father is retired				
Father is eligible	0.442***	0.132*	0.246***	0.454***
	(0.020)	(0.044)	(0.025)	(0.009)
Household size	0.005	-0.003	-0.017***	-0.001
	(0.004)	(0.007)	(0.005)	(0.006)
Bad health (father)	0.046***	0.033*	0.028***	0.027***
	(0.005)	(0.010)	(0.008)	(0.007)
Observations	34,462	37,135	54,976	126,573
R^2	0.175	0.188	0.214	0.202
For all panels:				
Country F.E.	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES

Notes: Standard errors in parentheses are clustered at the household level.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table 8: Sensitivity of Estimates

Sample	Southern Europe			Northern Europe			Central Europe			Full sample		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Panel A: Gateway Effect - Narrow windows around paternal retirement (+/- 3 years)												
Father is retired	0.020***	0.038***	0.011*	0.017	0.021	-0.055	0.004	0.009	-0.011	0.021***	0.028***	0.001
	(0.006)	(0.009)	(0.007)	(0.023)	(0.020)	(0.114)	(0.011)	(0.011)	(0.031)	(0.006)	(0.009)	(0.007)
Mass points :												
$\hat{\pi}_1$	0.334			0.065			0.210			0.319		
	(0.325)			(0.196)			(0.290)			(0.336)		
Log-likelihood	-4,599	-1,866	-2,711	-4,242	-210	-3,872	-8,058	-2,165	-5,467	-17,088	-7,489	-8,709
Observations	24,530	24,530	24,530	13,197	13,197	10,623	28,698	28,698	22,114	66,425	66,425	57,267
Panel B: Alternative definition of the treatment dummy - Year in which the father receives pension benefits												
Father receives	0.019***	0.037***	0.016**	0.010	0.025	0.004	0.008	0.010	0.010	0.020***	0.026***	0.005
pension benefits	(0.005)	(0.005)	(0.007)	(0.039)	(0.051)	(0.033)	(0.009)	(0.026)	(0.009)	(0.005)	(0.006)	(0.005)
Mass points :												
$\hat{\pi}_1$	0.334			0.065			0.210			0.319		
	(0.325)			(0.196)			(0.290)			(0.336)		
Log-likelihood	-4,599	-1,866	-2,711	-4,242	-210	-3,872	-8,058	-2,165	-5,467	-17,088	-7,489	-8,709
Observations	24,530	24,530	24,530	13,197	13,197	10,623	28,698	28,698	22,114	66,425	66,425	57,267

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. All specifications include controls for paternal education, country dummies, birth order of the child, birth cohort dummies for fathers (in 1-year interval) and time dummies representing duration dependence. To save space, $\hat{\pi}_2$ (1-$\hat{\pi}_1$) is not reported. Notice that observations for which the probability of belonging to Group 1 and Group 2 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table 9: Mechanisms: Manacorda and Moretti (2006) vs. Battistin et al. (2009) hypotheses

Sample	**Italy and Greece**			**Spain**		
	(1)	(2)	(3)	(4)	(5)	(6)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Father is retired	0.024***	0.061***	0.015**	0.031***	0.049***	0.020*
	(0.006)	(0.009)	(0.006)	(0.011)	(0.016)	(0.012)
Household size	-0.005	-0.007	-0.005*	-0.011**	-0.019***	-0.005
	(0.003)	(0.006)	(0.003)	(0.005)	(0.006)	(0.006)
Bad health (father)	0.005	0.020***	0.005	0.010	0.007	0.012
	(0.005)	(0.008)	(0.005)	(0.009)	(0.009)	(0.010)
Mass points :						
$\hat{\pi}_1$	0.334			0.334		
	(0.325)			(0.325)		
$\hat{\pi}_2$	0.666			0.666		
	(0.325)			(0.325)		
Country F.E.	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES	YES	YES
Log-likelihood	-5,508	-1,942	-3,388	-2,337	-767	-1,501
Observations	16,960	16,960	16,960	6,820	6,820	6,820

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1, 4), and weighted, using as weights $\hat{\pi}_1$ (col. 2, 5) or $\hat{\pi}_2$ (col. 3, 6). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. Notice that observations for which the probability of belonging to Group 1 and Group 2 is equal to zero are not included in the sample. Notice that in Spain I exclude individuals who declare themselves as retired because they were made redundant since they may get severance pay. Standard errors in parentheses are clustered at the household level.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

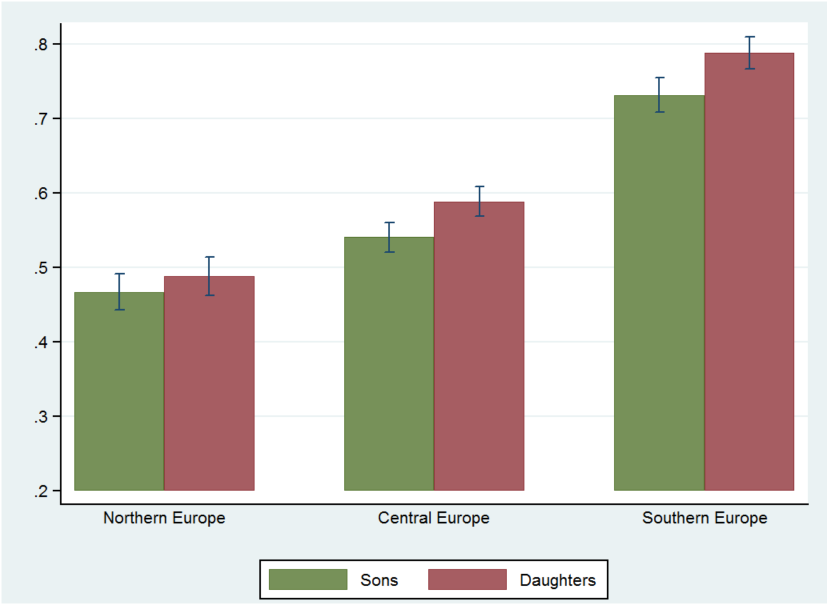
Table 10: Mechanisms: Employed vs. Self-employed in Italy and Greece

Sample	**Employed**			**Self-employed**		
	(1)	(2)	(3)	(4)	(5)	(6)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Father is retired	0.018***	0.057***	0.011*	0.037***	0.053***	0.028*
	(0.007)	(0.010)	(0.007)	(0.012)	(0.018)	(0.015)
Household size	-0.009**	-0.005	-0.007	0.003	-0.005	0.001
	(0.004)	(0.006)	(0.005)	(0.007)	(0.011)	(0.009)
Bad health (father)	0.007	0.030***	0.003	-0.005	-0.013	0.010
	(0.006)	(0.008)	(0.006)	(0.012)	(0.020)	(0.013)
Mass points :						
$\hat{\pi}_1$	0.334			0.334		
	(0.325)			(0.325)		
$\hat{\pi}_2$	0.666			0.666		
	(0.325)			(0.325)		
Country F.E.	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES	YES	YES
Log-likelihood	-5,508	-1,942	-3,388	-2,337	-767	-1,501
Observations	12,901	12,901	12,901	4,059	4,059	4,059

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1, 4), and weighted, using as weights $\hat{\pi}_1$ (col. 2, 5) or $\hat{\pi}_2$ (col. 3, 6). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. Notice that observations for which the probability of belonging to Group 1 and Group 2 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

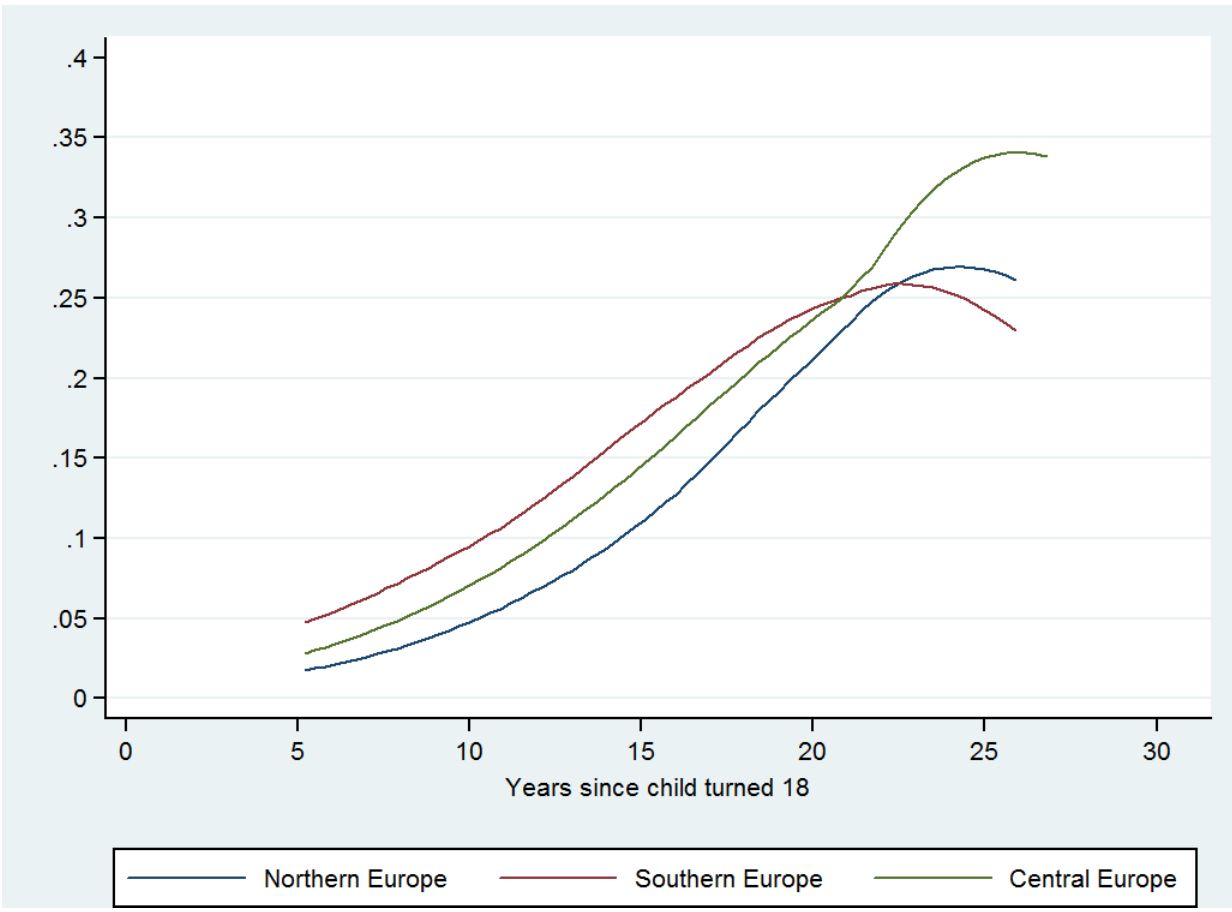
Appendix A: Supplemental Figures and Tables

Figure A1: Fraction of adult children who are married, by European region



Notes: Marital status refers to the interview year of wave 2. This variable is coded as 1 for married adult children living together with the spouse. Unfortunately, information on the year in which the child got married is not collected in SHARE data.

Figure A2: Empirical hazard rate of fathers' retirement, by European region



Notes: This figure plots the estimated hazard function of nest-leaving of children and that of paternal retirement by European region. These hazard functions are estimated using a nonparametric kernel-smoothing methodology (STS package in STATA). Notice that the reason why the smoothed hazard estimate is not depicted for $t < 5$ has to do with the choice of the bandwidth. Recall that children who were less than 18 are left-truncated.

Table A1: Model with shared frailty - Determinants of the Hazard of Retirement, by European region and overall sample

Sample	**Southern Europe**			**Northern Europe**			**Central Europe**			**Full sample**		
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)	(10)	(11)	(12)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Father is eligible	0.089***	0.100***	0.087***	0.032***	0.039***	0.023***	0.043***	0.047***	0.037***	0.055***	0.065***	0.061***
	(0.005)	(0.000)	(0.005)	(0.003)	(0.004)	(0.004)	(0.004)	(0.004)	(0.004)	(0.002)	(0.003)	(0.003)
Household size	0.002	-0.009	0.011***	-0.001	-0.000	0.004	-0.006**	-0.009**	0.003***	-0.001	0.001	-0.004***
	(0.002)	(0.000)	(0.002)	(0.002)	(0.003)	(0.003)	(0.003)	(0.004)	(0.001)	(0.001)	(0.002)	(0.001)
Bad health (father)	0.004	-0.025	0.019***	0.004*	0.004	0.002	0.002	0.001	0.008***	0.003*	0.005**	0.005***
	(0.004)	(0.000)	(0.004)	(0.003)	(0.003)	(0.003)	(0.003)	(0.003)	(0.002)	(0.002)	(0.002)	(0.002)
Mass points :												
$\hat{\pi}_1$	0.334			0.065			0.210			0.319		
	(0.325)			(0.196)			(0.290)			(0.336)		
$\hat{\pi}_2$	0.666			0.935			0.790			0.681		
	(0.325)			(0.196)			(0.290)			(0.336)		
Country F.E.	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES	YES
Log-likelihood	-3,185	-892	-2,080	-710	-618	-319	-2,298	-2073	-110	-6,485	-1,777	-4,186
Observations	18,806	18,806	18,806	12,597	12,597	12,597	23,682	23,682	18,419	55,085	55,085	49,822

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1, 4, 7, 10), and weighted, using as weights $\hat{\pi}_1$ (col. 2, 5, 8, 11) or $\hat{\pi}_2$ (col. 3, 6, 9, 12). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. All specifications include time dummies representing duration dependence. Notice that observations for which the probability of belonging to Group 1 and Group 2 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.

* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table A2: Model with shared frailty - Hazard of Nest-Leaving in Southern Europe, Sons and Daughters

Sample	**Sons**			**Daughters**		
	(1)	(2)	(3)	(4)	(5)	(6)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Father is retired	0.024***	0.055***	0.013**	0.017**	0.049***	0.011
	(0.006)	(0.009)	(0.007)	(0.008)	(0.011)	(0.008)
Household size	-0.009***	-0.014***	-0.006*	-0.001	-0.004	-0.002
	(0.003)	(0.005)	(0.003)	(0.005)	(0.007)	(0.005)
Bad health (father)	0.007	0.020***	0.011**	0.004	0.009	0.005
	(0.005)	(0.007)	(0.006)	(0.007)	(0.010)	(0.007)
Mass points :						
$\hat{\pi}_1$	0.334			0.334		
	(0.325)			(0.325)		
$\hat{\pi}_2$	0.666			0.666		
	(0.325)			(0.325)		
Wald test p-value for diff. btw. $\delta^{(2)}$ *and* $\delta^{(3)}$	0.000					
Wald test p-value for diff. btw. $\delta^{(5)}$ *and* $\delta^{(6)}$	0.00					
Country F.E.	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES	YES	YES
Log-likelihood	-4,115	-1,431	-2,529	-3,672	-1,255	-2,304
Observations	14,076	14,076	14,076	10,454	10,454	10,454

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1, 4), and weighted, using as weights $\hat{\pi}_1$ (col. 2, 5) or $\hat{\pi}_2$ (col. 3, 6). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. All specifications include time dummies representing duration dependence. Notice that observations for which the probability of belonging to Group 1 and Group 2 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table A3: Determinants of the Hazard of Nest-Leaving in Northern Europe, Sons and Daughters

Sample	**Sons**			**Daughters**		
	(1)	(2)	(3)	(4)	(5)	(6)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Father is retired	0.028	0.029	-0.099	-0.005	0.012	-0.100
	(0.038)	(0.027)	(0.078)	(0.020)	(0.036)	(0.088)
Household size	0.012	0.033***	-0.037	0.008	0.028*	-0.032
	(0.013)	(0.004)	(0.027)	(0.010)	(0.016)	(0.025)
Bad health (father)	-0.014	-0.002	-0.071**	-0.050***	-0.029***	-0.162***
	(0.010)	(0.010)	(0.036)	(0.009)	(0.009)	(0.044)
Mass points :						
$\hat{\pi}_1$	0.065			0.065		
	(0.196)			(0.196)		
$\hat{\pi}_2$	0.935			0.935		
	(0.196)			(0.196)		
Country F.E.	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES	YES	YES
Log-likelihood	-3,837	-2,984	-785	-3,050	-2,360	-634
Observations	7,740	7,740	6,105	5,453	5,453	4,508

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1, 4), and weighted, using as weights $\hat{\pi}_1$ (col. 2, 5) or $\hat{\pi}_2$ (col. 3, 6). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. Notice that observations for which the probability of belonging to Group 1 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.

* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table A4: Determinants of the Hazard of Nest-Leaving in Central Europe, Sons and Daughters

Sample	**Sons**			**Daughters**		
	(1)	(2)	(3)	(4)	(5)	(6)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Father is retired	0.012	0.014	0.010	-0.012	-0.005	-0.058
	(0.011)	(0.011)	(0.031)	(0.014)	(0.014)	(0.036)
Household size	-0.017***	-0.016**	-0.007	-0.011**	-0.010*	0.000
	(0.006)	(0.007)	(0.007)	(0.006)	(0.005)	(0.006)
Bad health (father)	-0.004	-0.002	-0.010	-0.009	-0.013	0.004
	(0.007)	(0.007)	(0.009)	(0.009)	(0.009)	(0.010)
Mass points :						
$\hat{\pi}_1$	0.210			0.210		
	(0.290)			(0.290)		
$\hat{\pi}_2$	0.790			0.790		
	(0.290)			(0.290)		
Country F.E.	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES
Birth order F.E, (child)	YES	YES	YES	YES	YES	YES
Log-likelihood	-6,380	-5,011	-1,038	-5,760	-4,489	-950
Observations	16,069	16,069	12,062	12,629	12,629	10,052

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1, 4), and weighted, using as weights $\hat{\pi}_1$ (col. 2, 5) or $\hat{\pi}_2$ (col. 3, 6). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. Notice that observations for which the probability of belonging to Group 2 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.

* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table A5: Mechanisms: Manacorda and Moretti (2006) vs. Battistin et al. (2009) hypotheses

Sample	**Italy and Greece**			**Spain**		
	(1)	(2)	(3)	(4)	(5)	(6)
Unobserved Group	No Het.	Group 1	Group 2	No Het.	Group 1	Group 2
Father is retired	0.024***	0.061***	0.015**	0.025***	0.047***	0.015
	(0.006)	(0.009)	(0.006)	(0.010)	(0.015)	(0.011)
Household size	-0.005	-0.007	-0.005*	-0.008*	-0.017***	-0.004
	(0.003)	(0.006)	(0.003)	(0.005)	(0.006)	(0.006)
Bad health (father)	0.005	0.020***	0.005	0.009	0.008	0.014
	(0.005)	(0.008)	(0.005)	(0.008)	(0.009)	(0.009)
Mass points :						
$\hat{\pi}_1$	0.334			0.334		
	(0.325)			(0.325)		
$\hat{\pi}_2$	0.666			0.666		
	(0.325)			(0.325)		
Country F.E.	YES	YES	YES	YES	YES	YES
Education F.E. (father)	YES	YES	YES	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES	YES	YES	YES
Log-likelihood	-5,508	-1,942	-3,388	-2,337	-767	-1,501
Observations	16,960	16,960	16,960	7,570	7,570	7,570

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1, 4, 7), and weighted, using as weights $\hat{\pi}_1$ (col. 2, 5, 8) or $\hat{\pi}_2$ (col. 3, 6, 9). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. Notice that observations for which the probability of belonging to Group 1 and Group 2 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.

* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table A6: Summary Statistics, Employed vs. Self-employed in Italy and Greece

Variable	Employed			Self-employed			
	Obs.	Mean	Std. Dev.	Obs.	Mean	Std. Dev.	p-value difference
Age (father)	689	69.869	7.199	240	70.222	6.723	0.534
Household size	689	2.334	0.653	240	2.320	0.718	0.799
Retired	689	0.932	0.252	240	0.872	0.335	0.006
Retirement age	642	58.555	4.719	209	61.701	4.287	0.000
Bad health	689	0.412	0.493	240	0.325	0.470	0.026
High school (father)	689	0.192	0.394	240	0.123	0.329	0.024
College or more (father)	689	0.075	0.264	240	0.044	0.206	0.123
High school (child)	689	0.492	0.500	240	0.463	0.500	0.469
College or more (child)	689	0.266	0.442	240	0.227	0.420	0.255
Nest-leaving age (child)	689	27.145	5.121	240	26.931	5.139	0.601
Married (child)	689	0.774	0.419	240	0.818	0.387	0.180

Notes: This table reports the observations from the cross-sectional sample before reshaping it as a longitudinal dataset. All the samples contain individuals for whom information on children's nest-leaving age and paternal education is not missing and exclude children who were less than 18. The paternal sample consists of all individuals who are either working or retired.

Table A7: Mechanisms: Fathers whose wives never worked

Sample	**Southern Europe**		
	(1)	(2)	(3)
Unobserved Group	No Het.	Group 1	Group 2
Father is retired	0.016**	0.054***	0.005
	(0.008)	(0.011)	(0.008)
Household size	-0.016***	-0.026***	-0.015**
	(0.006)	(0.008)	(0.006)
Bad health (father)	0.004	0.013	0.010
	(0.007)	(0.010)	(0.008)
Mass points :			
$\hat{\pi}_1$	0.334		
	(0.325)		
$\hat{\pi}_2$	0.666		
	(0.325)		
Country F.E.	YES	YES	YES
Education F.E. (father)	YES	YES	YES
Cohort F.E. (father)	YES	YES	YES
Birth order F.E. (child)	YES	YES	YES
Log-likelihood	-7,883	-2,726	-4,905
Observations	9,435	9,435	9,435

Notes: Logit estimations; average marginal effects reported. $\hat{\pi}_1$ and $\hat{\pi}_2$ are the estimated probability to belong to unobserved heterogeneity Group 1 and Group 2, respectively. The marginal effects are unweighted (col. 1), and weighted, using as weights $\hat{\pi}_1$ (col. 2) or $\hat{\pi}_2$ (col. 3). The sample sizes take into account the longitudinal structure of the data. Education is an indicator for father's college or more (ISCED$\geq$ 5, tertiary education) and high school education (ISCED=3 or 4, secondary and post-secondary education). Bad health is an indicator that takes value 1 if father's self-reported health is less than good. Notice that observations for which the probability of belonging to Group 1 and Group 2 is equal to zero are not included in the sample. Standard errors in parentheses are clustered at the household level.

* Significant at 10%; ** significant at 5%; *** significant at 1%.