UNIVERSITÀ DEGLI STUDI DI PADOVA

Dipartimento di Scienze Economiche ed Aziendali "Marco Fanno"

LIVING ARRANGEMENTS IN EUROPE: WHETHER AND WHY PATERNAL RETIREMENT MATTERS

LUCA STELLA
University of Padova

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

Luca Stella

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 the time of paternal retirement does not appear to be justified by changes in parental resources. Rather, one must probably look for channels involving the supply of informal child care provided by grandparents or the negative externalities in preferences between retired fathers and their children.

JEL Codes: J13, J26, J01

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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 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 ways in which children respond to these factors have 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 involves the phenomenon of falling fertility rates associated with prolonged co-residence (see, for example, Giuliano 2007, 2010). 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 adult 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 obtain identification from the exogenous variability in the 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 cash 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 a 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 reverse causation and endogeneity of paternal retirement, 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 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 the 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. This information is exploited to construct duration variables that indicate the time elapsed before each event occurs; 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 regarding 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 noted, 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 using 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 the provision of informal child care provided by grandparents or the negative externalities in preferences between retired fathers and their 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 parental resources around the time of retirement play 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. 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 allow 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 pathway through which this may happen.

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 because 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 lies in 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. Mana-

corda and Moretti (2006) and Battistin et al. (2009) also focus on fathers. Moreover, I restrict my attention to fathers who were either working[1] 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,[2] 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]

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,

[1] 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.

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

I assume that individuals can locate past events along the time line with adequate precision. Although 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.[3] 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 one year in Northern and Central Europe to approximately two 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 women.[4]

[Figure 1 - around here]

Table 3 reports the share of adult children that left home after paternal retirement, with Southern Europe showing by far the highest mean level, especially for sons.

[3] The quality of the retrospective information is a feature that has also been investigated 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 in life, Gould et al. (2011) find that retrospective information collected more than 50 years ago is of reasonably high quality.

[4] 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.

[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).[5] 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 emphasizing 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] 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.[6] This novel strategy for identifying treatment effects in the presence of an endogenous treatment when both the treatment and outcome are survival variables of a duration process 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 an 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).[7] Formally, the model can be written in the following manner:

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

where the unit of observation i represents the child-father pair residing in a given country, the outcome $h_{1,it}$ is the hazard that child i leaves the parental home at age t, $h_{2,it}$ 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.[8] Formally, the baseline hazard can be written as follows:

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

[7] These two destination states are assumed to be absorbing. Although this assumption appears 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 the previous literature, I assume that nest-leaving decisions are irreversible.

[8] As pointed out by Van den Berg et al. (2004), a piecewise constant function is the most flexible specification used for duration dependence functions.

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

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 at the time of the interview, 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.[9] 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 $u_{1,i}$ and $u_{2,i}$, I follow the latent class approach adopted by Melberg et al. (2010), who estimate a bivariate hazard model for the impact of cannabis on the risk of consuming hard drugs using a finite mixture framework.[10] Therefore, unobserved heterogeneity is assumed to divide the sample into two latent classes.[11] The intuitive explanation for the presence of these two classes

[9] An additional issue that would arise when controlling for paternal occupation is related to determining how to address 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.

[10] A finite mixture model describes the unobservable heterogeneity in terms of a finite number of latent classes that exist in the population (McLachlan and Peel 2000). Finite mixture models have recently been used by many authors, including, for instance, Bago d'Uva (2006), Bago d'Uva and Jones (2009), Balia (2013), and Angelini et al. (2013).

[11] Following Melberg et al. (2010), I perform the analysis using two latent classes. The reason is that the unobservable heterogeneity is considered at the child-father level. For example, there might be a number of unobservable factors, such as ability, transmitted from fathers to their children that may not be well captured by observable characteristics, and, consequently, they enter the error term. As a robustness check, I used three latent classes obtaining similar results.

is that individuals are clustered into two sub-groups that differ in terms of their unobservable propensity for nest-leaving and retirement. For instance, as I demonstrate in Section 5, one group is composed of young people who appear more likely to leave the nest later, whereas the other is more prone to leave the parental home earlier. Consistent with Melberg et al. (2010), I then allow all the coefficients to differ across the two latent groups; other studies (see, for instance, Pudney 2003; van Ours 2003), 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. 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 after 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.[12] 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 instrumental variable (IV) strategy, recognize this instrument as valid because pension reforms produce variation in paternal retirement that is credibly exogenous and unlikely to be

[12] See, for example, the evidence in Battistin et al. (2009), Attanasio and Brugiavini (2003), Battistin et al. (2014) and Liebman et al. (2009).

related to unobservable characteristics of the fathers that might explain the different nest-leaving outcomes of their offspring. 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.[13]

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. Thus, 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.[14] 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,[15] 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,[16] 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 a value of zero for all but the last of $i's$ periods (i.e., year $1, ..., T_{is} - 1$) and takes a value of one 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).[17] It is worth noting that one of the main advantages of the duration analysis over the linear IV setting adopted by previous studies

[13] 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 materially unchanged.

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

[15] 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.

[16] 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.

[17] One issue that arises in this particular setting is the possibility that paternal retirement occurs after children leave the nest. Although the majority of my 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 for the hazard of paternal retirement.

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:

$$logL = \sum_{i=1}^{n} \left\{ \sum_{k=1}^{2} \pi_k \left\{ \sum_{j=1}^{2} \left\{ \sum_{t=1}^{T_{i,j}-d_{i,j}} log\left[1 - h(\theta_k)_{kj,it}\right] + d_{i,j} log\left[h(\theta_k)_{kj,it}\right] \right\} \right\} \right\} \qquad (3)$$

where the prior probabilities π_k (each $\pi_k \geq 0$ and $\sum_{k=1}^{2} \pi_k = 1$) represent the proportion of the sample composing each latent class k. The variable $d_{i,j}$ is a dummy with a value of 1 if individuals are non-censored and a value of 0 if observations are right-censored, and θ_k is a vector of parameters that includes β_1, δ, β_2 and γ and that varies also at the latent class level. 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 expectation-maximization (EM) algorithm.[18]

[18] This is a commonly-used iterative procedure for computing the maximum likelihood estimates in problems where the data are incomplete or have missing values. See Jacho-Chávez and Trivedi (2009) and Balia (2013) for a recent discussion of this computational approach within the finite mixture framework. In brief, each iteration of the EM algorithm consists of two steps. First, posterior probabilites are recovered by using the Bayes' rule (E-step). Second, binary choice models are estimated using the individual probabilities of class membership as weights (M-step). The EM algorithm iterates these two steps until the likelihood converges. The convergence properties are discussed in detail by McLachlan and Krishnan (1996). To circumvent problems due to local optima, I repeated the analysis with different initial values of the π_k and the θ_k. The algorithm yielded the same solution, thereby suggesting that the maximum really is a global maximum. Hans Melberg graciously provided me with his Stata program for the EM algorithm.

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.[19] 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.[20] Overall, this figure shows 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;[21] 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 of 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 empirical evidence indicating that Southern European individuals tend to retire earlier (Gruber and Wise 2004).

[Figure 3 - around here]

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

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

[20] Notice that the reason why the smoothed hazard estimate is not depicted for $t < 5$ is associated with the choice of the bandwidth.

[21] 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.

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 appear 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.[22]

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]

[22]Results remain unchanged when excluding either the dummy for paternal bad health or household size or both. For this reason, although these variables are included in the model, they are not reported in the following tables.

5.2 Model with Shared Frailty

The primary concern regarding 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 individual probability of class membership for each European region as well as for the full sample. The prior probabilities, π_1 and π_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 individual probabilities of belonging to Group 1 and Group 2 as weights, respectively. Thus, 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 posterior 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 27% greater than the mean of Group 2

[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.

(25% versus 19%). Such large differences in the fraction of retired fathers can contribute to explaining why young people in Group 1 (labeled "low-propensity" nest-leaving types or "late" nest-leavers) are much more affected by paternal retirement than their counterparts in Group 2 (labeled "high-propensity" nest-leaving types or "early" nest-leavers). Interestingly, these two groups also differ 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 four times larger when 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. While treatment effects of paternal retirement are positive and significant for Group 1, they are close to zero for Group 2. As in 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 (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 prior 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).

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]

[24] In each macro-region, descriptive statistics for sons and daughters belonging to 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.

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. To address these concerns, I estimate the following linear version of model (1) using two stage least squares (2SLS):

$$Pr(S_{it} = 1) = \alpha + \beta Retired_{it} + \gamma X_i + \epsilon_{it} \tag{4}$$

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 S_{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 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

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

to children's living arrangements. In this setup, the first stage regression is given by:

$$Retired_{it} = \delta_0 + \delta_1 Eligibility_{it} + \pi X_i + \nu_{it} \tag{5}$$

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. 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 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

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

[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.

retirement year. To check the robustness of my results, I exploit information on the year in which the father first received pension benefits.[28] Thus, I re-estimate my model using an alternative treatment indicator variable set equal to 1 if father i collects pension income at time t. 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.[29] 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, i.e., "unaffected" by the lump-sum payment upon retirement.[30] My information on severance arrangements is drawn from Holzmann et al. (2011), from personal communications with national experts and from other country-specific sources.[31] As previously mentioned, the literature would attribute this causal relationship mainly to two competing mechanisms. To provide an empirical test for these two mechanisms, I use model (1) and analyze the differential effects of paternal retirement by separating Southern Europe across the above-mentioned two groups.

To the extent that the Manacorda and Moretti mechanism is at play, I expect paternal retirement to bribe Italian and Greek adult children to stay at home longer as a consequence of the positive shock to the

[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] As noted by Bolin et al. (2008), Southern European countries 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.

[30] 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 used to elicit this information was 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.

[31] 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.

family's liquidity associated with the retirement severance payment. However, the results reported in Panel A of 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 magnitudes of 6.1% and 1.5%, respectively. This result indicates that cash 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 emphasized 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 of Panel A 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 at the 10% level. This result is what I expected given the substantial reduction in sample size.

One may still be concerned that Spain is not a comparable group or that in Italy and Greece self-employed workers are not entitled to retirement severance payment. To address these concerns, I propose an additional test: for Italy and Greece, I use the employed as the group which is characterized by the presence of the retirement severance pay and self-employed as the group which is unaffected by the retirement severance pay.[32] The results reported in Panel B of Table 9 indicate that there are positive causal effects of paternal retirement on the timing of children's nest-leaving for both the employed (columns 1 to 3) and the self-employed (columns 4 to 6), which I interpret as corroborating evidence that the drop in paternal post-retirement income or the boost in family's income due to retirement severance payment does not provide a satisfactory explanation for the mechanism behind the decline in children's cohabitation at paternal retirement.

[Table 9 - 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. (2014) 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. In particular, the authors find that an

[32] 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. In addition, 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 provide an appropriate comparable group for analyzing the differential effects of paternal retirement.

additional grandparent at home increases the likelihood of children's nest-leaving by approximately 3%;[33] however, the authors do not consider grandmaternal and grandpaternal effects separately. Although this scenario can be applied to other Southern European countries, including Spain and Greece,[34] there is general consensus that grandmothers are the main providers of informal child care arrangements for their grandchildren (see, for instance, Richter et al. 1994). As discussed previously in the paper, female partners are excluded from the present analysis. Nevertheless, empirical literature has increasingly provided evidence that coupled individuals tend to plan their retirement decisions jointly (see, for example, Hurd 1990; Gustman and Steinmeier 2000; Stancanelli 2012). To account for the joint retirement hypothesis, I have demonstrated in Table A5 (Appendix A) that, when focusing on fathers whose spouses have never worked, there is a positive and quantitatively similar causal effect of paternal retirement on the likelihood of children's nest-leaving but only for "late" nest-leaving types. Therefore, this result reveals the potential effect of grandparents' supply of informal child care alongside other unexplained factors.

Anecdotal evidence invites the hypothesis that there may be a number of preference-related reasons that concern negative externalities between retired fathers and their offspring: children's departures from the parental home potentially stem from conflicting relationships with their fathers, which likely result 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 reasons for children's nest-leaving. However, to partially address this limitation in the data, I can use a measure for overcrowding at the time of children's nest-leaving as a proxy for preferences' negative externalities. 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.[35] To allow for the presence of household overcrowding, I estimate model (1) for Southern Europe, in which I include the interaction between paternal retirement and the dummy variable for overcrowding. If the coefficient on the interaction term is positive and statistically significant, then it does appear that preference-related reasons are likely to play a role in explaining children's decisions to leave their parental homes. Table 10 shows the parameter estimates. I find that the estimated

[33] Both the magnitude and the precision of this estimate is strikingly similar to those reported in Table 5 for Southern Europe.

[34] In Southern European countries, leaving the nest only at the time of marriage and childbearing is a widespread trend.

[35] To be more precise, SHARE provides information on the number of rooms available in the household's accommodation (including bedrooms but excluding kitchen, bathrooms, and hallways) at the interview year of wave 2. SHARE also contains information on the number of years of residence in the current accommodation, which enables me to retain only child-father pairs where the current accommodation was the same to that at the time of children's nest-leaving (approximately 84% of the cross-sectional sample). However, SHARE does not provide 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 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. Overall, I note that Greece is the country with the lowest median of the number of rooms per person at the time of children's nest-leaving (0.75), whereas Italy and Spain present the same median (0.8).

effect of paternal retirement remains substantially unchanged with respect to the benchmark specification (see Table 5) and that the coefficient on the interaction term is positive and significant (at the 1% level for Group 1 and the 10% for Group 2). This result suggests that more children leave the nest upon paternal retirement with overcrowding.

[Table 10 - 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 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).

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 specify 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 considered 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 appear to be driven by changes in parental economic resources. Rather, one needs to look for channels involving the supply of informal child care provided by grandparents or the negative externalities in preferences between retired 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. However, the results of this paper suggest that in Southern Europe policy makers should also be aware that there may potentially be 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

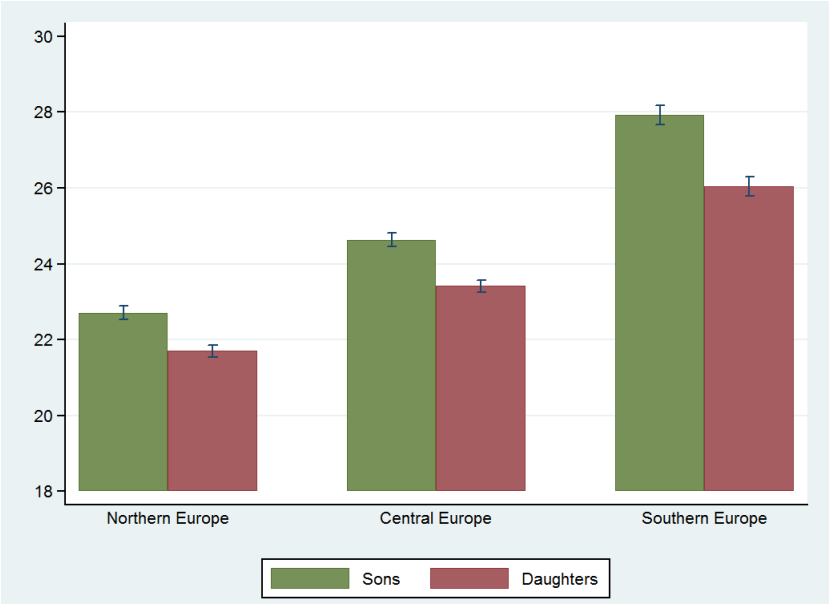
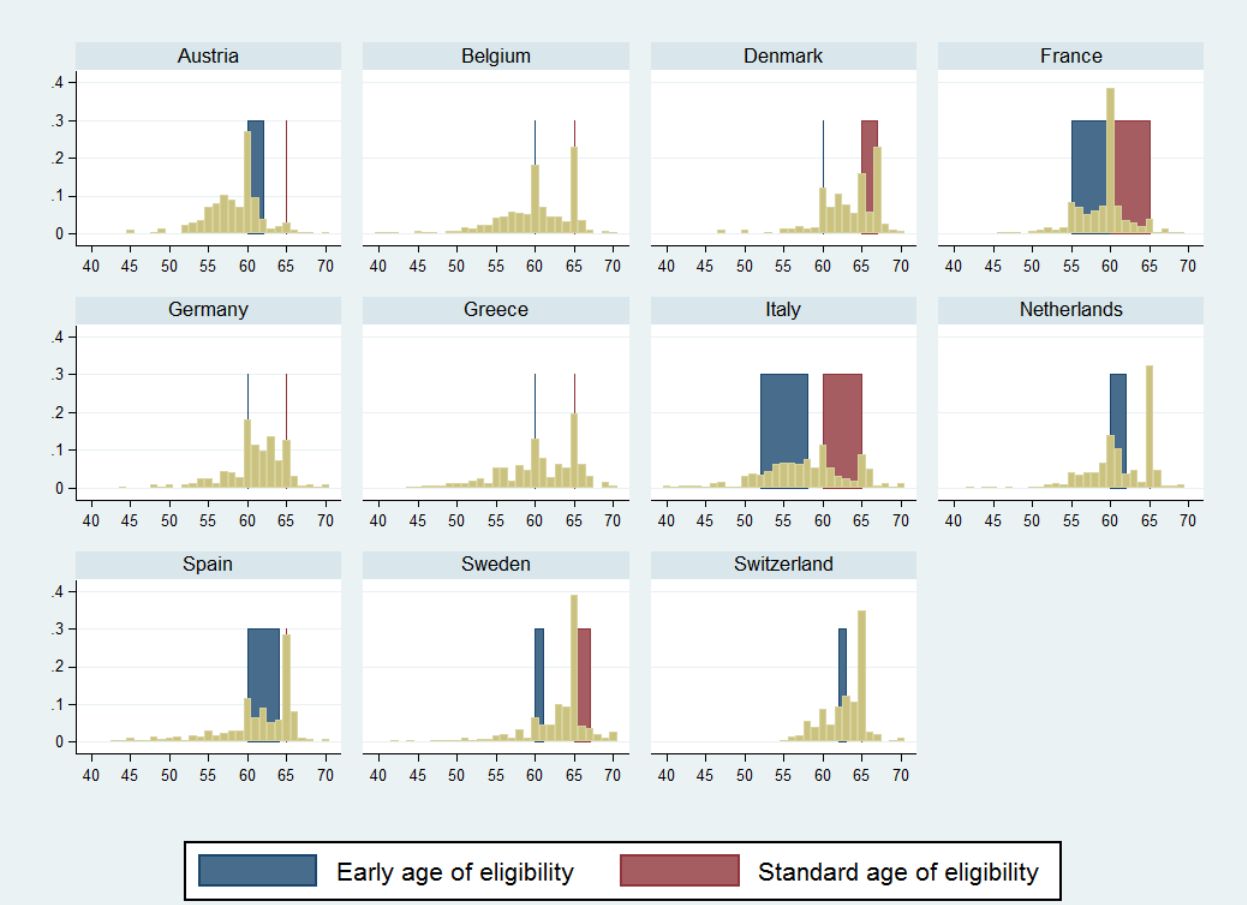
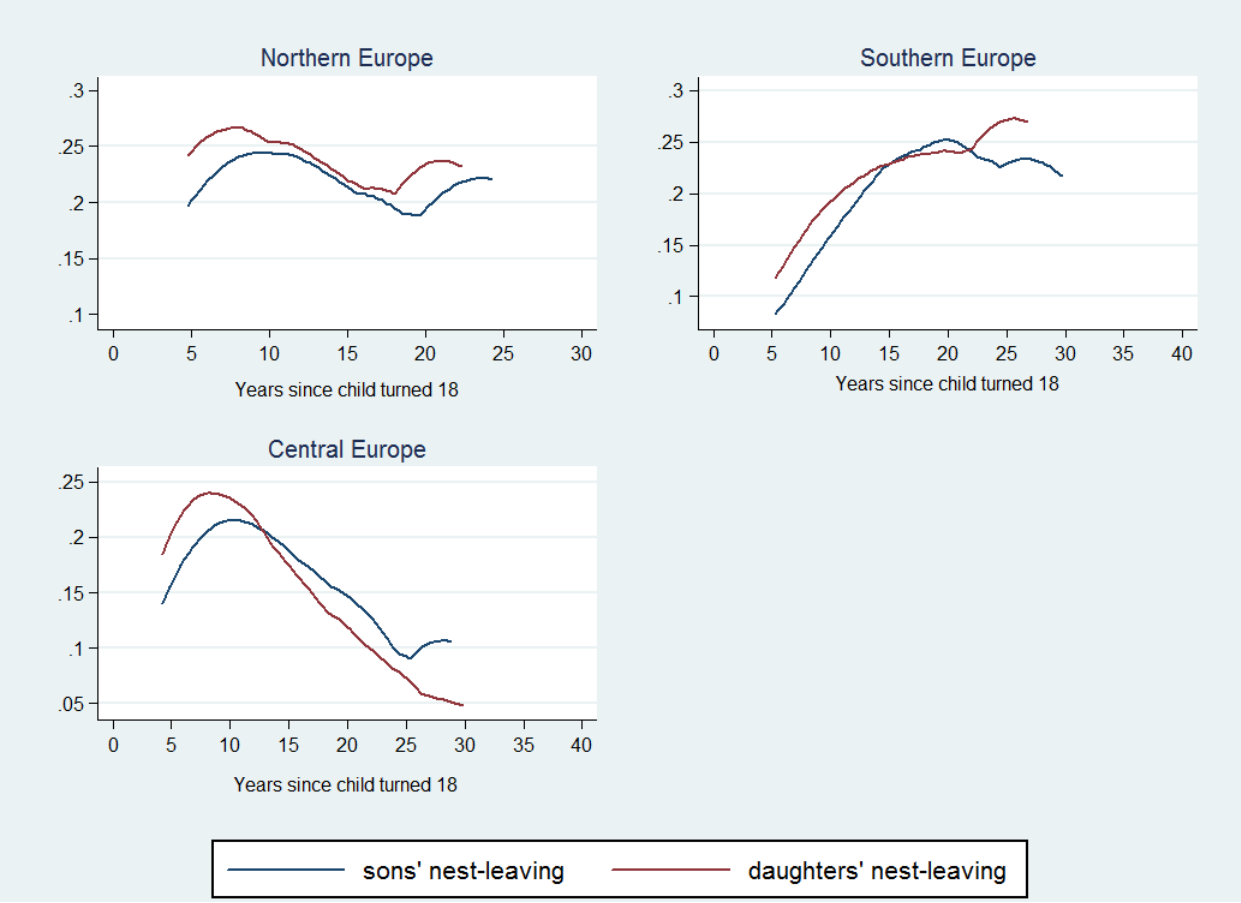


Figure 2: Histograms of father's retirement age, by country



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). Recall that children who were less than 18 (i.e., $t < 0$) are left-truncated. Notice that the reason why the smoothed hazard estimate is not depicted for $t < 5$ is associated with the choice of the bandwidth.

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 younger 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 younger than 18. The paternal sample consists of all individuals who are either working or retired.

Table 3: Summary Statistics, fraction of children who left home after paternal retirement

Sample	Sons			Daughters			Overall		
	Obs.	Mean	Std. Dev.	Obs.	Mean	Std. Dev.	Obs.	Mean	Std. Dev.
Southern Europe	1,436	0.45	0.49	1,356	0.38	0.48	2,792	0.42	0.49
Northern Europe	1,644	0.07	0.26	1,475	0.05	0.22	3,119	0.06	0.24
Central Europe	2,445	0.16	0.37	2,364	0.13	0.33	4,809	0.15	0.35
Overall	5,525	0.21	0.41	5,195	0.17	0.38	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 younger 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. 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)
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
π_1		0.334			0.065			0.210			0.319	
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	13,197	28,698	28,698	28,698	66,425	66,425	66,425

Notes: Logit estimations; average marginal effects reported. The marginal effects are unweighted (col. 1, 4, 7, 10) and weighted, using as weights the individual probability to belong to Group 1 (col. 2, 5, 8, 11) or Group 2 (col. 3, 6, 9, 12). 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 6: Model with shared frailty - Differences between clusters, by European region and full sample

Variable	Group 1		Group 2	
	Mean	Std. Dev.	Mean	Std. Dev.
Panel A: Southern Europe ($\pi_1 = 0.33$)				
Father is retired	0.247	0.431	0.195	0.397
Male (child)	0.570	0.495	0.578	0.494
High school (father)	0.150	0.357	0.136	0.342
College or more (father)	0.084	0.277	0.073	0.259
High school (child)	0.403	0.490	0.423	0.494
College or more (child)	0.301	0.459	0.235	0.424
Nest-leaving age	30.078	5.268	29.325	5.262
Panel B: Northern Europe ($\pi_1 = 0.07$)				
Father is retired	0.072	0.259	0.018	0.132
Male (child)	0.610	0.488	0.563	0.496
High school (father)	0.277	0.448	0.350	0.477
College or more (father)	0.213	0.409	0.282	0.450
High school (child)	0.463	0.499	0.459	0.498
College or more (child)	0.350	0.477	0.388	0.487
Nest-leaving age	26.308	5.196	23.704	4.104
Panel C: Central Europe ($\pi_1 = 0.21$)				
Father is retired	0.159	0.366	0.040	0.197
Male (child)	0.580	0.494	0.539	0.498
High school (father)	0.445	0.497	0.429	0.495
College or more (father)	0.272	0.445	0.253	0.435
High school (child)	0.511	0.500	0.456	0.498
College or more (child)	0.430	0.495	0.488	0.500
Nest-leaving age	29.024	7.055	25.326	4.286
Panel D: Full sample ($\pi_1 = 0.32$)				
Father is retired	0.172	0.377	0.123	0.328
Male (child)	0.574	0.495	0.561	0.496
High school (father)	0.334	0.472	0.277	0.448
College or more (father)	0.217	0.412	0.164	0.370
High school (child)	0.469	0.499	0.438	0.496
College or more (child)	0.392	0.488	0.351	0.477
Nest-leaving age	28.560	6.299	26.807	5.172

Notes: Descriptive statistics are computed using the longitudinal sample. Individuals with a posterior 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)
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)
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 window 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)
π_1		0.334			0.065			0.210			0.319	
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	13,197	28,698	28,698	28,698	66,425	66,425	66,425
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)
π_1		0.334			0.065			0.210			0.319	
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	13,197	28,698	28,698	28,698	66,425	66,425	66,425

Notes: Logit estimations; average marginal effects reported. The marginal effects are unweighted (col. 1, 4, 7, 10) and weighted, using as weights the individual probability to belong to Group 1 (col. 2, 5, 8, 11) or Group 2 (col. 3, 6, 9, 12). 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. 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

Panel A: Italy and Greece vs. Spain

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)
Country F.E.	YES	YES	YES	N/A	N/A	N/A
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
π_1		0.334			0.334	
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

Panel B: 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)
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
π_1		0.334			0.334	
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. The marginal effects are unweighted (col. 1, 4), and weighted, using as weights the individual probability to belong to Group 1 (col. 2, 5) or Group 2 (col. 3, 6). Notice that in Spain I exclude individuals who declare themselves as retired because they were made redundant since they may receive severance pay. Standard errors in parentheses are clustered at the household level.

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

Table 10: Mechanisms: Negative externalities in preferences

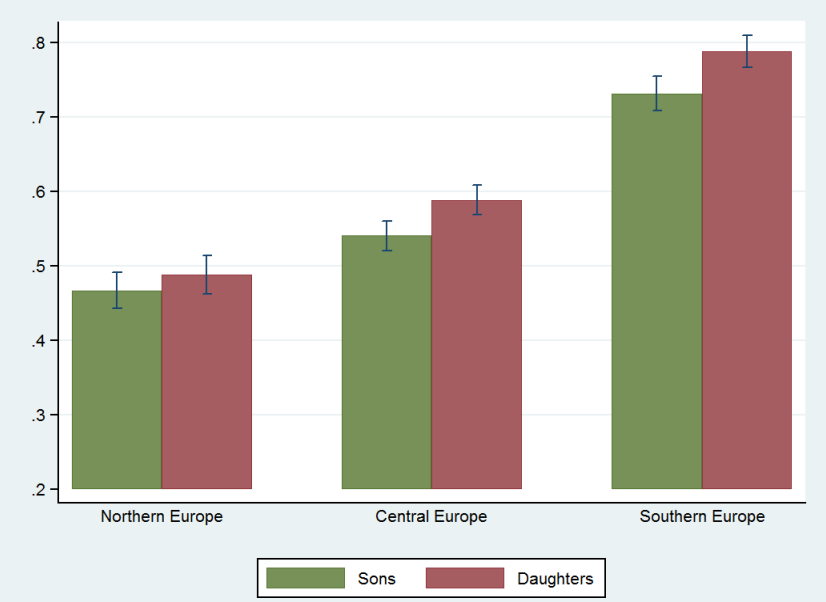
Sample	**Southern Europe**		
	(1)	(2)	(3)
Unobserved Group	No Het.	Group 1	Group 2
Father is retired	0.026***	0.064***	0.015**
	(0.006)	(0.009)	(0.006)
Father is retired*overcrowding	0.019***	0.039***	0.013**
	(0.006)	(0.010)	(0.006)
Overcrowding	0.009*	0.015**	0.007
	(0.005)	(0.007)	(0.005)
Country 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
π_1		0.334	
Log-likelihood	-6,646	-2,338	-4,098
Observations	21,348	21,348	21,348

Notes: Logit estimations; average marginal effects reported. The marginal effects are unweighted (col. 1), and weighted, using as weights the individual probability to belong to Group 1 (col. 2) or Group 2 (col. 3). Standard errors in parentheses are clustered at the household level.

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

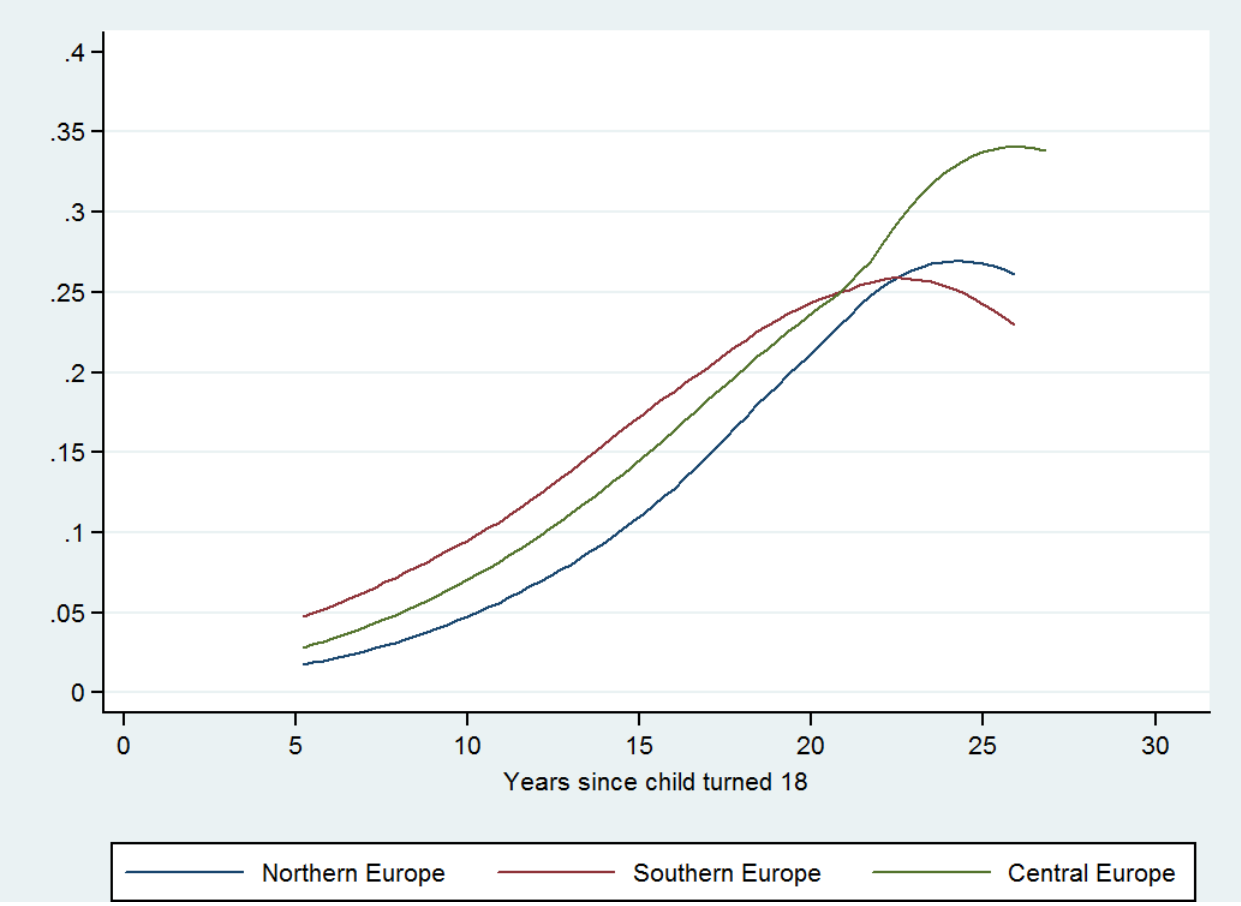
Appendix A: Supplemental Figures and Tables

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). Recall that children who were less than 18 (i.e., $t < 0$) are left-truncated. Notice that the reason why the smoothed hazard estimate is not depicted for $t < 5$ is associated with the choice of the bandwidth.

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)
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
π_1		0.334			0.065			0.210			0.319	
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	23,682	55,085	55,085	55,085

Notes: Logit estimations; average marginal effects reported. The marginal effects are unweighted (col. 1, 4, 7, 10), and weighted, using as weights the individual probability to belong to Group 1 (col. 2, 5, 8, 11) or Group 2 (col. 3, 6, 9, 12). 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 A2: Model with shared frailty - Hazard of Nest-Leaving for Sons and Daughters, by European region

Panel A: Southern Europe

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)
π_1		0.334			0.334	
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

Panel B: Northern Europe

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)
π_1		0.065			0.065	
Log-likelihood	-3,837	-2,984	-785	-3,050	-2,360	-634
Observations	7,740	7,740	7,740	5,453	5,453	5,453

Panel C: Central Europe

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)
π_1		0.210			0.210	
Log-likelihood	-6,380	-5,011	-1,038	-5,760	-4,489	-950
Observations	16,069	16,069	16,069	12,629	12,629	12,629
For all panels:						
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

Notes: Logit estimations; average marginal effects reported. The marginal effects are unweighted (col. 1, 4), and weighted, using as weights the individual probability to belong to Group 1 (col. 2, 5) or Group 2 (col. 3, 6). 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 A3: 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)
Country F.E.	YES	YES	YES	N/A	N/A	N/A
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
π_1		0.334			0.334	
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. The marginal effects are unweighted (col. 1, 4), and weighted, using as weights the individual probability to belong to Group 1 (col. 2, 5) or Group 2 (col. 3, 6). Standard errors in parentheses are clustered at the household level.
* Significant at 10%; ** significant at 5%; *** significant at 1%.

Table A4: 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 younger than 18. The paternal sample consists of all individuals who are either working or retired.

Table A5: 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)
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
π_1		0.334	
Log-likelihood	-7,883	-2,726	-4,905
Observations	9,435	9,435	9,435

Notes: Logit estimations; average marginal effects reported. The marginal effects are unweighted (col. 1), and weighted, using as weights the individual probability to belong to Group 1 (col. 2) or Group 2 (col. 3). Standard errors in parentheses are clustered at the household level.

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