CrossMark

Living arrangements in Europe: whether and why paternal retirement matters

Luca Stella[1,2]

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. Focusing on Southern Europe, 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, channels involving the supply of informal child care provided by grandparents or the quality of the home should be the focus of study.

Keywords Living arrangements · Retirement · Pension reforms · SHARE

JEL Classification J13 · J26 · J01

1 Introduction

Over the last several 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. Economists have primarily focused on parental economic resources

✉ Luca Stella
Stella@wiwi.uni-wuppertal.de

[1] University of Wuppertal and IZA, Wuppertal, Germany

[2] Schumpeter School of Business and Economics, Rainer-Gruenter-Str. 21, 42119 Wuppertal, Germany

(Edmonds et al. 2005; Manacorda and Moretti 2006; Battistin et al. 2009), labor market conditions for youth (Card and Lemieux 2000; Becker et al. 2010), prevailing characteristics in housing markets (Börsch-Supan 1986; Ermisch 1999; Alessie et al. 2004) and cultural factors (Giuliano 2007; Alesina and Giuliano 2011). 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 in 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 (Giuliano 2007). 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.

In the literature to date, many studies have shown that parental retirement significantly affects children's living arrangements decisions, thus suggesting that household composition responds to changes in household income. For example, Edmonds et al. (2005), exploiting a special feature of the South African pension system as a source of exogenous variation in household income, show that in African households, when a woman becomes pension eligible, there is a rise in the numbers of young women of childbearing age and children younger than five but a fall in the number of prime working-age women, who leave the parental home in search of work. Using a similar identification strategy from the exogenous variability in the Italian pension reforms during the 1990s, Manacorda and Moretti (2006) and Battistin et al. (2009) also show that there are sizeable changes in household composition when a parent becomes pension eligible. However, while both studies find that parental retirement in Italy induces a significant decline in the number of adult children living with their parents, they offer two competing explanations for this pattern. On the one hand, Manacorda and Moretti (2006) argue that once parents retire, they 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 Italy is one of the few European countries where most Italian employees receive a generous lump-sum payment upon retirement. Therefore, they suggest that parents may use part of their severance payment to buy a house for their sons and daughters, who can then leave the parental home. Nonetheless, not only changes in family income but also social factors may alter children's decisions to leave the parental home. For example, Angelini and Laferrère (2013) find that the nest-leaving age is influenced not only by the parental income but also by the housing quality. The authors point out that the housing quality effects are indeed important: children's nest-leaving is delayed with a more comfortable, less crowded home or which offers more privacy.

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), to the receipt of a sizeable retirement allowance, as noted by Battistin et al. (2009), or to social factors, including, for instance, housing characteristics at the time of children's nest-leaving. 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 using a European dataset to investigate whether and to what extent paternal retirement affects children's nest-leaving, and then to test the relative weight of these alternative hypotheses and shed some light on 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 1961–2007 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. Finally, the bivariate hazard model 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. 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 and Peracchi (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; Albertini and Kohli 2013), 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.

Overall, my results demonstrate that paternal retirement has a positive and significant effect on the timing of children's nest-leaving only in Southern European countries. In this European region, the magnitude of the effect varies between 1.2 and 5 %, and there are no significant differences between sons and daughters. Focusing on Southern Europe, I find that 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 focus on channels involving the provision of informal child care provided by grandparents or the home quality at the time of children's nest-leaving. 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 that raise the retirement age on decisions by young people to move out of their parents' homes. Therefore, it is important that such pension reforms are accompanied by policy programs (e.g., interventions in the housing market) that encourage home-leaving of young adults.

The remainder of this paper is organized as follows. The next section presents a description of the data and provides background information on the eligibility ages for retirement in Europe. Section 3 describes the empirical specification and identification strategy. The main results of the paper are presented in Sects. 4, and 5 illustrates the robustness checks. I discuss the results in Sect. 6, and concluding remarks are provided in Sect. 7.

2 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. 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. Manacorda and Moretti (2006) and Battistin et al. (2009) also focus on fathers. Moreover, I restrict my attention to fathers who were either working 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 from 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, 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 4935 fathers and a sample that consists of 10,720 children (5525 sons and 5195 daughters).[1] The distribution of the sample of fathers as well as the sample of children across the countries is presented in Table 1.

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. 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).

To determine the retirement age of the fathers and age at which children leave the nest, I exploit recall information from the following two questions in the questionnaire asked to the parents: "In what year did you retire?" and "In what year did the child move from the parental household?". The availability of such information relating events that occurred at some point in time before the year of the survey is essential because it allows for the creation of a retrospective panel dataset. For this reason, to conduct the analysis, I assume that individuals can locate past events along the time-line with adequate precision. While these retrospective data are self-reported and may be susceptible to recall error that may bias coefficient estimates, it is reassuring that Korbmacher (2014) finds that overall, the majority of respondents report their retirement year correctly. Some limitations of my data are worth mentioning. First, with the exception of the year of nest-leaving, I lack any

---

[1] Fathers who have more than one child in the sample are overrepresented because each child is treated as a unit in the analysis. In SHARE, questions on the children's nest-leaving age are asked for a maximum of four children.

Table 1 Sample of fathers and children, by country

Sample	Fathers	Sons	Daughters	Total
Austria	242	278	255	533
Belgium	664	704	686	1390
Denmark	407	478	421	899
France	543	588	606	1194
Germany	568	585	546	1131
Greece	300	339	298	637
Italy	629	655	673	1328
Netherlands	518	593	590	1183
Spain	361	442	385	827
Sweden	455	573	464	1037
Switzerland	248	290	271	561
Total	4935	5525	5195	10,720

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	SD
Sons			
Age	5525	38.15	8.22
Nest-leaving age	5525	24.92	4.83
High school	5525	0.46	0.50
College or more	5525	0.37	0.48
Married	5525	0.72	0.45
Never left home	5525	0.01	0.10
Daughters			
Age	5195	37.77	8.42
Nest-leaving age	5195	23.61	4.30
High school	5195	0.46	0.50
College or more	5195	0.40	0.49
Married	5195	0.77	0.42
Never left home	5195	0.01	0.10
Fathers			
Age	4935	66.89	8.60
Retired	4935	0.72	0.45
Working	4935	0.28	0.45
Retirement age	3553	60.34	4.73
High school	4935	0.34	0.47
College or more	4935	0.23	0.42
Household size	4935	2.23	0.57

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

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 5 years later (26.9 years in Southern Europe vs 22.1 years in Northern Europe). Young people in the Central European countries fall somewhere between these extremes. The figure also shows the presence of a gender gap in nest-leaving age: daughters leave the parental home earlier than sons, ranging from approximately 1 year in Northern and Central Europe to approximately 2 years in Southern Europe. This gap can partly be explained by the fact that age at marriage, which is positively correlated with the postponement of home-leaving, is lower for women.[2]

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 and Peracchi (2012) and Gruber and Wise (2004).[3] 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.

3 Empirical specification

3.1 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. 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

---

[2] See Chiuri and Boca (2010) for a discussion on the gender differences in nest-leaving decisions across Europe.

[3] Information on the retirement legislation in Greece is obtained from Duval (2003).
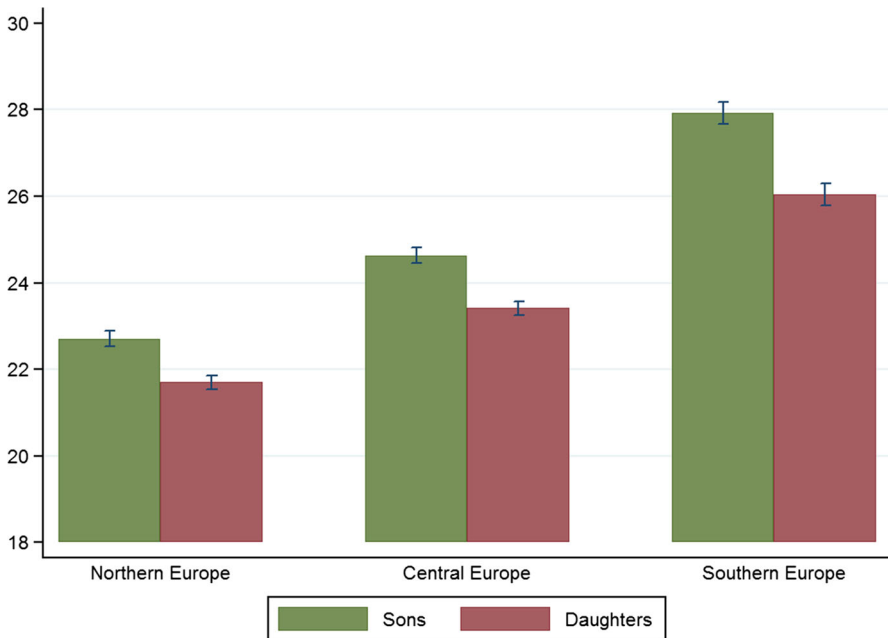
**Fig. 1** Children's nest-leaving mean age, by European region

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).[4] Formally, the model can be written in the following manner:

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

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

---

[4] 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.
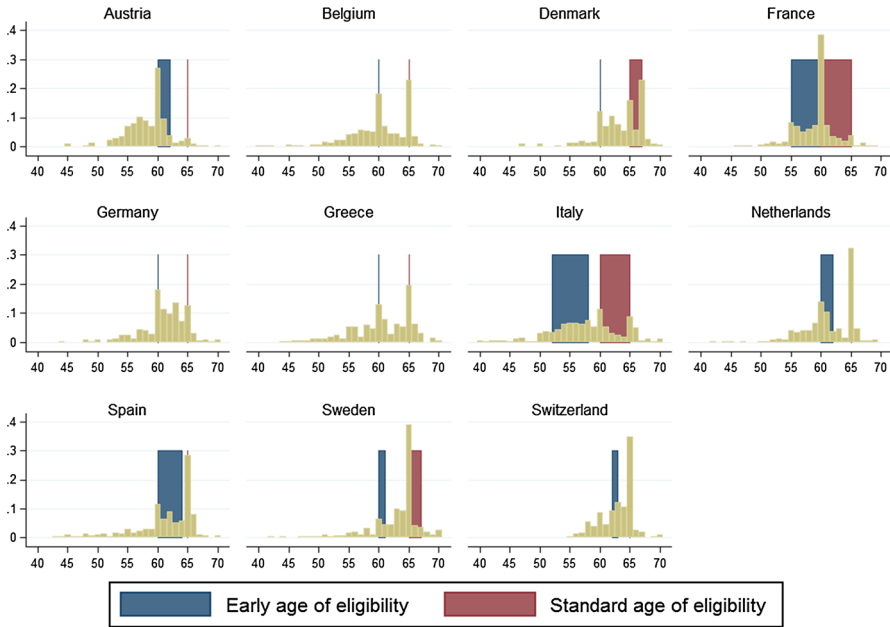
Fig. 2 Histogram of father's retirement age, by country. *Notes*: *Source*: Angelini et al. (2009), Mazzonna and Peracchi (2012), Gruber and Wise (2004) and Duval (2003). The *vertical blue* and *red lines*, respectively, mark the eligibility ages for early and normal retirement age, whereas the *blue* and *red areas* represent changes in the eligibility ages for the cohorts in my sample (Color figure online)

are modeled using a flexible piecewise constant function.[5] Formally, the baseline hazard can be written as follows:

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

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_{it} is a matrix of individual controls that may affect the hazard. Specifically, I include child's age and age squared, child's gender, a dummy that takes value 1 if the child is married at the time of the interview, household size, and an indicator for the father having a college-level education or above (ISCED ≥ 5, tertiary education) or a high school education (ISCED $= 3$ or 4, secondary and post-secondary education). I do not include paternal occupation

---

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

because of the large fraction of missing observations (approximately 30 % of the cross-sectional sample); however, education is strongly correlated with occupation. 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_{it}, 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.[6] Therefore, unobserved heterogeneity is assumed to divide the sample into two latent classes.[7] The intuitive explanation for the presence of these two classes 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 Sect. 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

---

[6] 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 and Jones (2009), Balia (2014) and Angelini et al. (2013).

[7] 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.

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. 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 related to unobservable characteristics of the fathers that might explain the different nest-leaving outcomes of their offspring.[8] 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.[9]

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 $(i = 1, \ldots, n)$ is associated with multiple time periods t_i $(t_i = 1, \ldots, T_{is})$, where T_{is} is the total number of years subject i was at risk for the event.[10] 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,[11] 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

---

[8] People may have expected the early retirement age to be raised at some point in the future, but the exact timing was most likely not anticipated.

[9] 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.

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

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

the father's age when his child is 18 as the onset of risk,[12] 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 $1, \ldots, T_{is} - 1$) 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). It is worth noting that one of the main advantages of the duration analysis over the linear IV setting adopted by previous studies is the allowance for censoring, which leads to the elimination of any constraints on the age at which children left their parents' home. For example, Manacorda and Moretti (2006) focus only on youths aged 18–30, whereas Billari and Tabellini (2010) 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:

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

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.[13]

---

[12] 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.

[13] 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) for a recent discussion of this computational approach within the finite mixture framework. Hans Melberg graciously provided me with his Stata program for the EM algorithm.

4 Main results

Before presenting the results, I provide a visual analysis of the evolution of the estimated hazard function for nest-leaving. In particular, Fig. 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. 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; (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.[14]

4.1 Instrumental variable analysis

Although the bivariate hazard model described in Sect. 3 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 begin my analysis with the following linear version of model (1) estimated using two stage least squares (2SLS):

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

where the outcome variable H_{it} is a dummy taking the value 1 if a child i residing in a given country left the parental home at age t. The treatment dummy $Retired_{it}$ and the variable X_{it} are defined in the same way as in Sect. 3. Therefore, X_{it} contains child's age and age squared, child's gender, a dummy that takes value 1 if the child is married at the time of the interview, household size, paternal education, country fixed effects, cohort fixed effects for fathers, and controls for the birth order of the child. Following Manacorda and Moretti (2006), I focus on youth aged 18–30 years. 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.

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 1961–2007 as an instrument for paternal retirement. As discussed in Sect. 3, this instrument is recognized to be relevant and arguably exogenous to children's living arrangements. In this setup, the first stage regression is given by:

---

[14] A visual analysis of the dynamics of the hazard for paternal retirement shows that as expected, in all European regions the hazard of paternal retirement increases with time. Due to space considerations, this figure is not reported here, but is available from the author upon request. It is also available in an earlier working paper (Stella 2014).
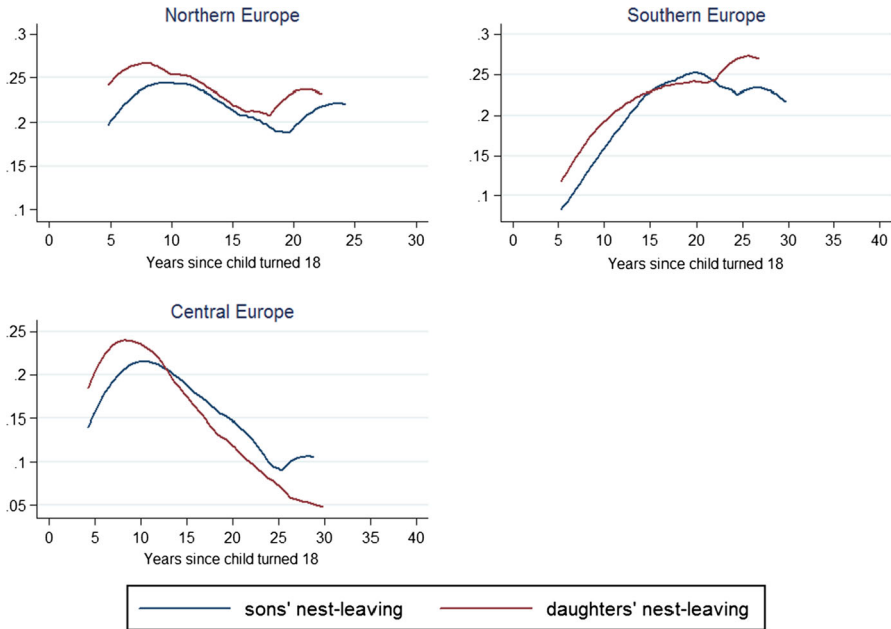
Fig. 3 Empirical hazard rate of children's nest-leaving, by European region. *Notes* This figure plots the estimated hazard function of nest-leaving of children by European region. This hazard function is estimated using a nonparametric kernel-smoothing methodology (STS package in STATA). Recall that children who were <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

$$Retired_{it} = \delta_0 + \delta_1 Eligibility_{it} + \pi X_{it} + v_{it} \tag{5}$$

where the dummy *Eligibility*$_{it}$ represents the instrument introduced in Sect. 3. 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 3 reports the 2SLS results. The treatment dummy on paternal retirement is positive and significant at the 1 % level only for Southern Europe (see column 1). This dummy variable, however, becomes negative and non-significant for Northern Europe (see column 2), and negative and significant only at the 10 % level for Central European countries (see column 3). In the pooled sample the dummy variable becomes positive although not significant (see column 4). 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 provides 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 analysis and explanation.

Table 3 Effects of paternal retirement on children's nest-leaving, IV analysis

Sample	(1) Southern Europe	(2) Northern Europe	(3) Central Europe	(4) Full sample
Panel A: 2SLS				
Dep. Var.:				
Child leaves home (dummy)				
Father is retired	0.176***	−0.291	−0.087*	0.024
	(0.067)	(0.249)	(0.046)	(0.038)
Household size	−0.015***	−0.007**	−0.031***	−0.020***
	(0.003)	(0.003)	(0.004)	(0.002)
Married (child)	0.082***	0.056**	0.077***	0.073***
	(0.006)	(0.023)	(0.006)	(0.004)
Female (child)	0.128***	0.099***	0.090***	0.103***
	(0.005)	(0.005)	(0.003)	(0.002)
Observations	34,462	37,135	54,976	126,573
R^2	0.277	0.339	0.354	0.376
First stage F statistic	61.86	10.35	86.47	144.88
Panel B: First stage				
Dep. Var.:				
Father is retired				
Father is eligible	0.115***	0.132***	0.244***	0.152***
	(0.015)	(0.034)	(0.026)	(0.013)
Household size	0.006*	−0.003	−0.019***	−0.004*
	(0.004)	(0.004)	(0.002)	(0.002)
Married (child)	−0.021***	0.003	−0.012***	−0.009***
	(0.005)	(0.004)	(0.003)	(0.002)
Female (child)	0.014***	0.002	0.005**	0.007***
	(0.003)	(0.002)	(0.002)	(0.001)
Observations	34,462	37,135	54,976	126,573
R^2	0.344	0.402	0.441	0.401

All regressions include child's age (introduced as a second-order polynomial), country dummies, father's education, cohort fixed effects for fathers, and controls for the birth order of the child

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

4.2 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

Table 4 Model for the hazard of children's nest-leaving and paternal retirement—without shared frailty

Sample	Southern Europe		Northern Europe		Central Europe		Full sample	
Dep. Var.: Hazard of:	(1) Nest-leaving	(2) Retirement	(3) Nest-leaving	(4) Retirement	(5) Nest-leaving	(6) Retirement	(7) Nest-leaving	(8) Retirement
Father is retired	0.020***		0.036		0.007		0.022***	
	(0.005)		(0.025)		(0.009)		(0.005)	
Father is eligible		0.082***		0.027***		0.037***		0.047***
		(0.005)		(0.002)		(0.004)		(0.002)
Household size	−0.007***	0.002	−0.007	0.010***	−0.012***	−0.008**	−0.008***	0.000
	(0.003)	(0.002)	(0.006)	(0.003)	(0.004)	(0.003)	(0.002)	(0.001)
Married (child)	0.032***	−0.007	0.045***	−0.004***	0.030***	−0.011***	0.035***	−0.008***
	(0.006)	(0.004)	(0.006)	(0.001)	(0.005)	(0.002)	(0.004)	(0.002)
Female (child)	0.036***	0.002	0.066***	−0.002	0.043***	0.001	0.044***	0.001
	(0.004)	(0.003)	(0.003)	(0.004)	(0.004)	(0.002)	(0.003)	(0.001)
Log-likelihood	−7807	−3154	−6894	−660	−12,164	−2237	−27,482	−6313
Observations	24,530	18,806	13,197	12,597	28,698	23,682	66,425	55,085

Logit estimations; average marginal effects reported. All regressions include time dummies representing duration dependence, child's age (introduced as a second-order polynomial), country dummies, father's education, cohort fixed effects for fathers, and controls for the birth order of the child. Standard errors in parentheses are clustered at the household level

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

specification, I include child's age and age squared, child's gender, a dummy that takes value 1 if the child is married at the time of the interview, household size, paternal education, country fixed effects, cohort fixed effects for fathers, and include controls for the birth order of the child. 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 %. 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 becomes 0.036 and 0.007, respectively. Overall, 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.[15] Moreover, I find that coefficients on household size are negatively correlated with the probability of children's nest-leaving, and that the probability to leave the parental home is larger for married or female adult children.[16] 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 2.7 % in Northern Europe to 8.2 % 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 0.022. 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).

In sum, although these correlations may suffer from problems of confounding, they provide a first indication that paternal retirement is significantly 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 and significant correlation obtained from the regression on the sample of Southern European countries has a causal interpretation.

---

[15] Specifically, I found that the cross-region differences in the share of adult children that left the home after paternal retirement 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 proportion of adult offspring in Northern and Central European countries is still living with their parents. This table is available in an earlier working paper (Stella 2014), and is also available from the author upon request.

[16] Including controls for the presence of kids at the time of the interview yields substantially unchanged results.

Table 5 Model for the hazard of children's nest-leaving in Southern Europe—with shared frailty

Sample	Southern Europe			Southern Europe—sons			Southern Europe—daughters		
Unobserved group	(1) No Het.	(2) Group 1	(3) Group 2	(4) No Het.	(5) Group 1	(6) Group 2	(7) No Het.	(8) Group 1	(9) Group 2
Dep. Var.:									
Hazard of nest-leaving									
Father is retired	0.020***	0.049***	0.012**	0.023***	0.052***	0.011*	0.016*	0.045***	0.011
	(0.005)	(0.008)	(0.005)	(0.006)	(0.009)	(0.007)	(0.008)	(0.012)	(0.009)
Household size	−0.007***	−0.012***	−0.005*	−0.011***	−0.016***	−0.007**	−0.003	−0.006	−0.003
	(0.003)	(0.004)	(0.003)	(0.003)	(0.005)	(0.003)	(0.005)	(0.007)	(0.005)
Married (child)	0.032***	0.022**	0.025***	0.029***	0.028***	0.018**	0.040***	0.024	0.036***
	(0.006)	(0.009)	(0.006)	(0.007)	(0.010)	(0.007)	(0.010)	(0.015)	(0.010)
Female (child)	0.036***	0.045***	0.034***						
	(0.004)	(0.006)	(0.004)						
π_1		0.334			0.334			0.334	
Log-likelihood	−7807	−2700	−4862	−4095	−1428	−2520	−3657	−1253	−2293
Observations	24,530	24,530	24,530	14,076	14,076	14,076	10,454	10,454	10,454

Logit estimations; average marginal effects reported. The marginal effects are unweighted (col. 1, 4, 7) and weighted, using as weights the individual probability to belong to Group 1 (col. 2, 5, 8) or Group 2 (col. 3, 6, 9). All regressions include time dummies representing duration dependence, child's age (introduced as a second-order polynomial), country dummies, father's education, cohort fixed effects for fathers, and controls for the birth order of the child. Standard errors in parentheses are clustered at the household level

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

4.3 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 subgroups of the population. Table 5 presents the estimation results of logistic regressions on the hazard of children's nest-leaving for Italy, Greece and Spain, countries for which I found a positive and significant association between paternal retirement and the probability of first nest-leaving.

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 4.9 %. The treatment effect remains significant, albeit quantitatively less important (1.2 %), for those who belong to Group 2.[17]

To learn more about the characteristics of the two groups, Table 6 displays summary statistics on selected covariates for Southern Europe. 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 Table 6, 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 (25 vs 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

---

[17] In order to save space and increase readability, Table 5 reports the estimated coefficients only for the hazard of children's nest-leaving, which is the outcome of main interest in this paper. Overall, the results for the hazard of paternal retirement suggest that in accordance with the model in which unobserved heterogeneity is not allowed (see Table 4), eligibility rules have a significant influence on actual retirement. These results are available from the author upon request, and are also available in an earlier working paper (Stella 2014).

Table 6 Differences between unobserved groups in Southern Europe

Unobserved group	Group 1		Group 2	
	Mean	SD	Mean	SD
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
Married	0.834	0.372	0.831	0.374

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

home later and have better outcomes in terms of their own and their fathers' education.

It is worth noting that when restricting the analysis to Northern Europe and Central Europe, I find that as expected, the dummy variable for paternal retirement is generally no longer statistically significant; therefore, for brevity, I do not report results from these regressions (the results are available from the author upon request).[18] Overall, the evidence presented above suggests that, although quantitatively small, there are positive causal effects of paternal retirement on the timing of first nest-leaving of children living in Southern European countries.

Finally, in an attempt to disentangle the treatment effects of paternal retirement on sons from the effects on daughters in Southern Europe, I consider the samples of male and female children separately. The results for sons and daughters are presented in columns 4–9 of Table 5. When restricting the analysis to sons (see columns 4–6), the coefficient on paternal retirement is 2.3 % in the model in which unobserved heterogeneity is ignored (see column 4), and varies between 5.2 % for individuals in Group 1 (see column 5) and 1.1 % for those belonging to Group 2 (see column 6). A similar pattern is observed in the regressions for daughters (see columns 7–9), with the difference being that 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.

5 Sensitivity analysis

Before proceeding to discuss and test empirically the potential mechanisms, I perform a robustness analysis for Southern Europe to determine if the results change when I use a different specification of the model.

I begin by investigating 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

---

[18] Furthermore, the full set of results for Northern and Central Europe, as well as for the full sample are available in an earlier working paper (Stella 2014).

regardless of their fathers' retirement status. To allow for this possibility, I define a time frame of 3 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 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 7. Reassuringly, these parameter estimates resemble those obtained in the benchmark specification (see columns 1–3 of Table 5), with the only difference being that the magnitude of the estimated effects of paternal retirement becomes slightly smaller.

I then address the concern that the father may start receiving pension benefits only some years after his retirement year. To check the robustness of my results, I exploit information on the year in which the father first received pension benefits.[19] 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 of Table 7 show, the evidence remains substantially unchanged relative to the benchmark specification.

6 Discussion for Southern Europe

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.[20] 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.[21] My information on severance arrangements is drawn from Holzmann et al. (2011), from personal communications with national experts and from other country-specific sources.[22] As previously mentioned, the literature

---

[19] 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.

[20] 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.

[21] García-Gómez et al. (2014) 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, the main results still hold if these individuals are included.

[22] 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.

Table 7 Robustness checks, model for the hazard of children's nest-leaving in Southern Europe

Sample	Southern Europe		
Unobserved group	(1) No Het.	(2) Group 1	(3) Group 2
Panel A			
Dep. Var.:			
Hazard of nest-leaving			
Father is retired	0.018***	0.034***	0.010
	(0.006)	(0.009)	(0.006)
Household size	−0.007**	−0.014***	−0.003
	(0.003)	(0.004)	(0.003)
Married (child)	0.032***	0.023**	0.023***
	(0.006)	(0.009)	(0.006)
Female (child)	0.037***	0.046***	0.034***
	(0.004)	(0.006)	(0.005)
π_1		0.334	
Log-likelihood	−7584	−2634	−4718
Observations	24,530	24,530	24,530
Panel B			
Dep. Var.:			
Hazard of nest-leaving			
Father receives pension benefits	0.017***	0.025***	0.014*
	(0.005)	(0.005)	(0.007)
Household size	−0.007***	−0.009***	−0.007
	(0.003)	(0.003)	(0.005)
Married (child)	0.034***	0.031***	0.031***
	(0.005)	(0.006)	(0.009)
Female (child)	0.035***	0.035***	0.040***
	(0.004)	(0.004)	(0.006)
π_1		0.334	
Log-likelihood	−8052	−4431	−3332
Observations	24,530	24,530	24,530

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). All regressions include time dummies representing duration dependence, child's age (introduced as a second-order polynomial), country dummies, father's education, cohort fixed effects for fathers, and controls for the birth order of the child. Standard errors in parentheses are clustered at the household level

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

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.

Table 8 Changes in parental income, model for the hazard of children's nest-leaving in Southern Europe

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
Panel A: Italy and Greece versus Spain						
Dep. Var.:						
Hazard of nest-leaving						
Father is retired	0.021***	0.054***	0.012**	0.030***	0.045***	0.022*
	(0.006)	(0.009)	(0.006)	(0.011)	(0.015)	(0.013)
Household size	−0.008**	−0.010*	−0.007**	−0.014**	−0.020***	−0.010
	(0.004)	(0.006)	(0.003)	(0.006)	(0.007)	(0.007)
Married (child)	0.024***	0.013	0.022***	0.042***	0.068***	0.019
	(0.007)	(0.011)	(0.007)	(0.012)	(0.015)	(0.013)
Female (child)	0.040***	0.047***	0.037***	0.033***	0.038***	0.034***
	(0.005)	(0.007)	(0.006)	(0.008)	(0.014)	(0.009)
π_1		0.334			0.334	
Log-likelihood	−5453	−1922	−3356	−1699	−620	−1044
Observations	16,960	16,960	16,960	6820	6820	6820

Sample	Employed			Self-employed		
Dep. Var.:						
Panel B: Employed versus self-employed in Italy and Greece						
Hazard of nest-leaving						
Father is retired	0.015**	0.049***	0.008	0.051***	0.072***	0.033**
	(0.007)	(0.010)	(0.007)	(0.013)	(0.018)	(0.014)
Household size	−0.011***	−0.008	−0.009**	0.002	−0.015	0.004
	(0.004)	(0.006)	(0.004)	(0.008)	(0.012)	(0.007)
Married (child)	0.018**	0.013	0.011	0.044**	0.020	0.069***
	(0.008)	(0.012)	(0.008)	(0.017)	(0.023)	(0.017)
Female (child)	0.032***	0.038***	0.029***	0.080***	0.088***	0.075***
	(0.006)	(0.008)	(0.006)	(0.011)	(0.017)	(0.012)
π_1		0.334			0.334	
Log-likelihood	−4106	−1419	−2549	−1289	−475	−766
Observations	12,901	12,901	12,901	4059	4059	4059

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 regressions include time dummies representing duration dependence, child's age (introduced as a second-order polynomial), country dummies, father's education, cohort fixed effects for fathers, and controls for the birth order of the child. Standard errors in parentheses are clustered at the household level

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

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 family's liquidity associated with the

retirement severance payment. However, the results reported in Panel A of Table 8 (columns 1–3) 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, with magnitudes of 5.4 and 1.2 %, 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 4–6 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 becomes slightly larger with respect to the estimate in column 3, although is less precisely estimated. 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.[23] The results reported in Panel B of Table 8 indicate that overall, there are positive causal effects of paternal retirement on the timing of children's nest-leaving for both the employed (columns 1–3) and the self-employed (columns 4–6), 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.

For this reason, in parallel to financial considerations, 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 additional grandparent at home increases the likelihood of children's nest-leaving by approximately 3 %; 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,[24] there is general consensus that grandmothers are the main providers of informal child care

---

[23] 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. In addition, I 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. For brevity, these descriptive statistics are not reported here, but are available from the author upon request, and are also available in an earlier working paper (Stella 2014).

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

Table 9 Fathers whose wives never worked, model for the hazard of children's nest-leaving in Southern Europe

Sample	Southern Europe		
Unobserved group	(1) No Het.	(2) Group 1	(3) Group 2
Dep. Var.:			
Hazard of nest-leaving			
Father is retired	0.017**	0.054***	0.007
	(0.008)	(0.012)	(0.009)
Household size	−0.017***	−0.023***	−0.015***
	(0.005)	(0.006)	(0.005)
Married (child)	0.017	−0.012	0.020*
	(0.011)	(0.019)	(0.012)
Female (child)	0.041***	0.044***	0.045***
	(0.007)	(0.009)	(0.008)
π_1		0.334	
Log-likelihood	−2900	−1013	−1785
Observations	9435	9435	9435

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). All regressions include time dummies representing duration dependence, child's age (introduced as a second-order polynomial), country dummies, father's education, cohort fixed effects for fathers, and controls for the birth order of the child. Standard errors in parentheses are clustered at the household level

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

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, in Table 9 I show 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 the conflicting relationship 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. In addition, Angelini and Laferrère (2013) emphasize that not only parental income but also the quality of the house or the desire for privacy are relevant determinants of nest-leaving outcomes. Unfortunately, it is difficult to verify this hypothesis with my data because, as already mentioned, the SHARE questionnaire does not provide information

Table 10 Housing quality effects, model for the hazard of children's nest-leaving in Southern Europe

Sample	Southern Europe		
	(1)	(2)	(3)
Unobserved group	No Het.	Group 1	Group 2
Dep. Var.:			
Hazard of nest-leaving			
Father is retired	0.017**	0.030**	0.001
	(0.007)	(0.012)	(0.007)
Father is retired*overcrowding	0.016*	0.011*	0.015*
	(0.008)	(0.006)	(0.008)
Overcrowding	0.004	0.009	0.005
	(0.004)	(0.013)	(0.004)
Household size	−0.008***	0.000	−0.009***
	(0.002)	(0.004)	(0.002)
Married (child)	0.023***	0.016**	0.019***
	(0.004)	(0.006)	(0.004)
Female (child)	0.043***	0.043***	0.039***
	(0.003)	(0.004)	(0.003)
π_1		0.334	
Log-likelihood	−6705	−2252	−4411
Observations	21,348	21,348	21,348

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). All regressions include time dummies representing duration dependence, child's age (introduced as a second-order polynomial), country dummies, father's education, cohort fixed effects for fathers, and controls for the birth order of the child. Standard errors in parentheses are clustered at the household level

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

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 the housing quality. 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.[25] 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.

---

[25] 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).

If the coefficient on the interaction term is positive and statistically significant, then it does appear that the housing quality is likely to play a role in explaining children's decisions to leave their parental homes. Table 10 shows the parameter estimates. Overall, I find that the coefficient on the interaction between paternal retirement and the dummy variable for overcrowding is positive and significant (at the 10 % level). This result suggests that more children leave the nest upon paternal retirement with overcrowding, which can be interpreted as a housing quality or privacy effect.

7 Conclusion

In this paper, I examine the relationship between paternal retirement and the timing of housing emancipation of young adults in Europe. Taking advantage of the retrospective dimension of my micro data, I specify a bivariate discrete-time hazard model with shared frailty and exploit over time and cross-country variation in early retirement legislation. Overall, my regression results suggest that there is a positive and significant influence of paternal retirement on the probability of first nest-leaving of children living in Southern European countries. However, there is no clear evidence of positive and 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 unobservable factors are controlled for.

Focusing on Southern Europe, I provide an empirical test for the two main financial channels by which paternal retirement may be considered to affect children's co-residence. Comparing my cross-country evidence 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 home quality at the time of children's nest-leaving.

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. Therefore, pension reforms should be accompanied by policy programs (e.g., interventions in the housing market) that encourage moving-out of young adults.

References

Aassve, A., Billari, F. C., Mazzucco, S., & Ongaro, F. (2002). Leaving home: A comparative analysis of ECHP data. *Journal of European Social Policy*, *12*(4), 259–275.

Abbring, J. H., & Van den Berg, G. J. (2003). The nonparametric identification of treatment effects in duration models. *Econometrica*, *71*(5), 1491–1517.

Albertini, M., Kohli, M., & Vogel, C. (2007). Intergenerational transfers of time and money in European families: Common patterns—Different regimes? *Journal of European Social Policy*, *17*(4), 319–334.

Albertini, M., & Kohli, M. (2013). The generational contract in the family: An analysis of transfer regimes in Europe. *European Sociological Review*, *29*(4), 828–840.

Alesina, A., & Giuliano, P. (2010). The power of the family. *Journal of Economic Growth*, *15*(2), 93–125.

Alessie, R., Brugiavini, A., & Weber, G. (2004). Saving and cohabitation: The economic consequences of living with one's parents in Italy and the Netherlands. In R. H. Clarida, J. A. Frankel, F. Giavazzi, & K. D. West (Eds.), *NBER International Seminar on Macroeconomics* (pp. 413–441). Cambridge, MA: MIT Press.

Angelini, V., Brugiavini, A., & Weber, G. (2009). Ageing and unused capacity in Europe: Is there an early retirement trap? *Economic Policy*, *24*(59), 463–508.

Angelini, V., & Laferrère, A. (2013). Parental altruism and nest leaving in Europe: Evidence from a retrospective survey. *Review of Economics of the Household*, *11*(3), 393–420.

Angelini, V., Bucciol, A., Wakefield, M., & Weber, G. (2013). *Can temptation explain housing choices in later life?* Netspar Discussion Paper.

Bago d'Uva, T., & Jones, A. (2009). Health care utilisation in Europe: New evidence from ECHP. *Journal of Health Economics*, *28*, 265–279.

Balia, S. (2014). Survival expectations, subjective health and smoking: Evidence from SHARE. *Empirical Economics*, *47*(2), 753–780.

Battistin, E., Brugiavini, A., Rettore, E., & Weber, G. (2009). The retirement consumption puzzle: Evidence from a regression discontinuity approach. *The American Economic Review*, *99*(5), 2209–2226.

Battistin, E., De Nadai, M., & Padula, M. (2014). *Roadblocks on the road to Grandma's house: Fertility consequences of delayed retirement*. IZA Discussion Paper, n. 8071.

Becker, S. O., Bentolila, S., Fernandes, A., & Ichino, A. (2010). Youth emancipation and perceived job insecurity of parents and children. *Journal of Population Economics*, *23*(3), 1047–1071.

Billari, F. C., Philipov, D., & Baizán, P. (2001). Leaving home in Europe: The experience of cohorts born around 1960. *International Journal of Population Geography*, *7*(5), 339–356.

Billari, F., & Tabellini, G. (2010). Italians are late: Does it matter? In J. B. Shoven (Ed.), *Demography and the economy, NBER Books* (pp. 371–412). Chicago: University of Chicago Press.

Bolin, K., Lindgren, B., & Lundborg, P. (2008). Informal and formal care among single-living elderly in Europe. *Health Economics*, *17*(3), 393–409.

Börsch-Supan, A. (1986). Household formation, housing prices, and public policy impacts. *Journal of Public Economics*, *30*(2), 145–164.

Card, D., & Lemieux, T. (2000). Adapting to circumstances: The evolution of work, school, and living arrangements among North American youth. In R. Freeman & D. Blanchflower (Eds.), *Youth employment and joblessness in advanced countries*. Chicago: University of Chicago Press.

Chiuri, M. C., & Del Boca, D. (2010). Home-leaving decisions of daughters and sons. *Review of Economics of the Household*, 8(3), 393–408.

Duval, R. (2003). *The retirement effects of old-age pension and early retirement schemes in OECD countries*. Paris: OECD.

Edmonds, E. V., Mammen, K., & Miller, D. L. (2005). Rearranging the family? Income support and elderly living arrangements in a low-income country. *The Journal of Human Resources*, 40(1), 186–207.

Ermisch, J. (1999). Prices, parents, and young people's household formation. *Journal of Urban Economics*, 45(1), 47–71.

García-Gómez, P., Jiménez-Martín, S., & Vall Castelló, J. (2014). Financial incentives, health and retirement in Spain. NBER Chapters. In *Social security programs and retirement around the world: Disability insurance programs and retirement*. National Bureau of Economic Research Inc.

Giuliano, P. (2007). Living arrangements in Western Europe: Does cultural origin matter? *Journal of the European Economic Association*, 5(5), 927–952.

Gruber, J., & Wise, D. (2004). *Social security programs and retirement around the world: Micro-estimation*. Chicago: University of Chicago Press.

Gustman, A., & Steinmeier, T. (2000). Retirement in dual-career families: A structural model. *Journal of Labor Economics*, 18, 503–545.

Holzmann, R., Pouget, Y., Vodopivec, M., & Weber, M. (2011). *Severance pay programs around the world: History, rationale, status, and reforms*. IZA Discussion Paper No. 5731.

Hurd, M. (1990). The joint retirement decision of husbands and wives. In D. Wise (Ed.), *Issues in the economics of aging* (pp. 231–258). NBER.

Jacho-Chávez, D., & Trivedi, P. (2009). Computational considerations in empirical microeconometrics: Selected examples. In T. C. Mills & K. Patterson (Eds.), *Palgrave handbook of econometrics, vol. 2. Applied econometrics. Part V*. Basingstoke: Palgrave Macmillan.

Jenkins, S. P. (2005). *Survival analysis*. Unpublished Manuscript. Institute for Social and Economic Research, University of Essex, Colchester, UK.

Korbmacher, J. M. (2014). *Recall error in the year of retirement*. SHARE Working Paper Series 21-2014.

Manacorda, M., & Moretti, E. (2006). Why do most Italian youths live with their parents? Intergenerational transfers and household structure. *Journal of the European Economic Association*, 4(4), 800–829.

Mazzonna, F., & Peracchi, F. (2012). Aging, cognitive abilities and retirement. *European Economic Review*, 56(4), 691–710.

McLachlan, G. J., & Peel, D. (2000). *Finite mixture models*. New York: Wiley.

Melberg, H. O., Jones, A. M., & Bretteville-Jensen, A. L. (2010). Is cannabis a gateway to hard drugs? *Empirical Economics*, 38(3), 583–603.

Miniaci, R., Monfardini, C., & Weber, G. (2003). *Is there a retirement consumption puzzle in Italy?* IFS Working Paper.

Pudney, S. (2003). The road to ruin? Sequences of initiation to drugs and crime in Britain. *Economic Journal*, 113(486), 182–198.

Richter, K., Chai, P., Apichat, C., & Kusol, S. (1994). The impact of child care on fertility in urban Thailand. *Demography*, 31(4), 651–662.

Stancanelli, E. (2012). *Spouses' retirement and hours outcomes: Evidence from twofold regression discontinuity with differences-in-differences*. IZA Discussion Paper, n. 6791.

Stella, L. (2014). *Living arrangements in Europe: Whether and why paternal retirement matters*. IED Discussion Paper Series, n. 251. Boston University.

Van den Berg, G., Van der Klaauw, B., & Van Ours, J. C. (2004). Punitive sanctions and the transition rate from welfare to work. *Journal of Labor Economics*, 22(1), 211–241.

Van Ours, J. C. (2003). Is cannabis a stepping-stone for cocaine? *Journal of Health Economics*, 22(4), 539–554.