

Aparicio Fenoll, Ainhoa; Oppedisano, Veruska

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## Should I Stay or Should I Go? Sibling Effects in Household Formation

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Ainhoa Aparicio-Fenoll  
Veruska Oppedisano

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# Should I Stay or Should I Go? Sibling Effects in Household Formation

**Ainhoa Aparicio-Fenoll**

*Collegio Carlo Alberto,  
IZA and CHILD*

**Veruska Oppedisano**

*London Metropolitan University,  
IZA and IFS*

Discussion Paper No. 8713  
December 2014

IZA

P.O. Box 7240  
53072 Bonn  
Germany

Phone: +49-228-3894-0  
Fax: +49-228-3894-180  
E-mail: [iza@iza.org](mailto:iza@iza.org)

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## ABSTRACT

### Should I Stay or Should I Go? Sibling Effects in Household Formation<sup>\*</sup>

In Southern Europe youngsters leave the parental home significantly later than in Northern Europe and United States. Policies have been implemented in Southern Europe to incentivize young adults to leave parental home earlier. Do peer effects among siblings amplify the effects of these policies? Estimating peer effects is challenging because of problems of reflection, endogenous group formation, and correlated unobservables. We overcome these issues in the context of a Spanish rental subsidy, exploiting the subsidy eligibility age threshold to analyse peer effects among siblings. Instrumental variable estimates show that peer effects among siblings are negative, and that the effect is explained by the presence of old or ill parents. Findings indicate that policy makers should target the household rather the individual, and combine policies for young adults together with policies for elderly.

JEL Classification: J1, H2, I3

Keywords: peer effects, conditional cash transfer, youth emancipation, living arrangements

Corresponding author:

Ainhoa Aparicio-Fenoll  
Collegio Carlo Alberto  
Via Real Collegio 30  
Moncalieri (TO) 10024  
Italy  
E-mail: [ainhoa.aparicio@carloalberto.org](mailto:ainhoa.aparicio@carloalberto.org)

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# 1 Introduction

There are great disparities in the fraction of young people living with their parents in Europe. In 2010, almost 60 percent of young people in the 18-34 age bracket lived in their parental homes in Italy, Spain, Portugal, and Greece, whilst that statistic is below 40 percent in France, the UK, and the Netherlands, and as low as 20 percent in Norway, Sweden, and Finland.<sup>1</sup> Moreover, the fraction of young people living with their parents has increased sharply in Southern Europe during the last three decades and hence, this disparities have been widening.

Policy makers in Southern European countries are concerned about young adults late household formation because it may critically affect family formation decisions, overall fertility rates, youth labour supply, and the sustainability of pay-as-you-go pension systems. As a consequence, several Southern European governments have implemented measures or advocated the need for incentives to promote household formation. In 2007, the Portuguese government introduced "Porta 65", which offers to young adults in the 18-30 age group a monetary subsidy to cope with rental expenditures. In 2008, the Spanish Government implemented "Renta Basica de Emancipation", which offers to young people in the 22-29 age bracket a monthly monetary subsidy of € 210, conditional on renting accommodation. In 2013 the French government introduced the "Aide Mobili-Jeune", which offers a monthly subsidy of €10-100 to less than 30 years old young adults to help them reducing the rental cost burden and promote young adults emancipation. In 2007, the Italian Minister of Economy defined young adults who still reside with their parents as "big babies" (bamboccioni), advocating financial incentives to induce them leaving their parental home earlier.<sup>2</sup>

Siblings' interactions are particularly interesting for policy makers. In our context, an individual's decision to leave parental home in response to a financial incentive could affect the emancipation decision of her siblings, even when those are not directly affected by the reform. If spillover effects among siblings are positive (negative) the change in individual incentives to leave parental home would be amplified (reduced) and hence the aggregate impact of policies that promote youngsters emancipation would be larger (smaller) than the one implied by individual responses to the policies.

There are several reasons why sibling effects may play a role in amplifying the effect of the policy. First, imitation among siblings may reflect an intrinsic desire to behave like others. Imitation is stronger for smaller age differences and from older to younger siblings (Barr and Hayne, 2003). Similarly, young adults may derive utility from acting in accordance with others if the age at first marriage is a social norm (Di Stefano, 2008). Second, the fact that one sibling responds to the other's emancipation may also reflect interactions in information transmission, so that the choices of any single person modify the information available to all her siblings (Dufflo and Saez, 2003). However, sibling effects may also operate in the opposite direction, and reduce

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<sup>1</sup>Source: Eurostat, EU-SILC.

<sup>2</sup>Observer: flowers and taxes, Financial times 2007

the effects of the policy (Angrist and Lang, 2004). If a sibling has left parental home, the quantity of public goods available for the remaining sibling increases, reducing the incentive to follow the first sibling in her emancipation decision. Parents may also increase household resources to the remaining sibling in response to the higher risk of remaining alone: Manacorda and Moretti (2006) show that if children have a preference for living on their own, some parents are willing to trade off their own consumption to bribe their children into staying at home. Besides, single children of disrupted families are observed leaving the nest at a slower rate (Mencarini et al., 2010). In case of lone parents, emancipation of one sibling may deter the emancipation of the remaining one.

Estimating causal effects in social interactions has proven challenging. As Manski (1993) pointed out, estimation of these effects needs to deal with problems of simultaneous causality, correlated unobservables, and endogenous group membership. Researchers have used different strategies to tackle these issues. Some authors attempt to control for as many observable characteristics as possible, or use instrumental variables.<sup>3</sup> Others identify peer effects by exploiting exogenous group assignment.<sup>4</sup> A third approach consists in studying peer effects in naturally occurring groups, and exploiting random variation in exposure to the treatment for a random subset of individuals. This last strategy is called partial population approach and was advocated by Moffitt (2001).<sup>5</sup>

We follow the partial population approach and examine the causal effect of a sibling's emancipation decision on the individual's own probability of leaving parental home. We use data from the Spanish Survey on Income and Living Conditions which follows individuals over time, even when forming a new household. Our identification strategy makes use of the panel structure of the data and the exogenous increase in household formation rates induced by the rental subsidy. The panel data nature of our sample allows us to difference out any individual or household time-invariant unobservable characteristic; while the exogenous eligibility criteria for the rental subsidy allow us to deal with other omitted variables as well as reverse causality concerns. Our findings suggest that there are negative siblings' effects on household formation: a youngster who leaves the parental home delays her siblings' decision to leave. However, the negative sibling effect disappears if siblings are close in age. When looking at the mechanisms, our results suggest that the negative sibling effect can be explained by the remaining individual staying longer in the parental home in the presence of an old or ill parent. Moreover, sibling effects turn positive when the influence goes from older to younger siblings and siblings are close-in-age. Overall, siblings' interactions

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<sup>3</sup>See, for instance Bayer et al., (2008), Burke and Sass (2013), Carrell et al., (2008), Gaviria and Raphael (2001), Hensvik and Nilsson (2010), Maurin and Moschion (2009) and Nicoletti and Rabe (2014).

<sup>4</sup>See, for instance, Carrell et al., (2011), Hoxby (2000), Katz et al., (2001) and Sacerdote, (2001).

<sup>5</sup>See Angelucci et al. (2010), Baird et al. (2012), Dahl et al. (2014), Hesselius et al., (2009) and Lalive and Cattaneo (2009).

reduce the impact of policies that foster household formation unless age differences between siblings are small.

Numerous studies have produced empirical evidence documenting the existence of relevant siblings' interactions in many areas.<sup>6</sup> However, little evidence has been produced on the effects of peer behavior in living arrangements. Using Italian data, Di Stefano (2008) estimates a structural model in which young adults simultaneously choose labor supply, residential arrangement and marital status conditional on the social norm on the age at first marriage, endogenously determined as an equilibrium outcome. Her results indicate that young adults, and especially women, tend to conform to each other. Adamopoulou and Kaya (2013), using peers' characteristics as an instrument for the fraction of emancipated peers, finds evidence of positive peer effects among North-American high school friends. To the best of our knowledge, we are the first to explore the role of sibling effects on the decision to leave parental home.

The rest of the paper proceeds as follows. Section 2 describes the institutional setting and data. Section 3 presents the empirical methodology. Section 4 provides a discussion of the empirical results and Section 5 concludes.

## 2 Institutions and data

### 2.1 Institutional background

Announced in September 2007 and enacted since January 2008, the Basic Rent for Emancipation is a monetary subsidy introduced by the Spanish Ministry of Housing with the aim of fostering youngsters' household formation. The government expected to achieve this goal by helping young individuals to cope with rental expenses. The policy also aimed at promoting youngsters' economic independence and geographical mobility.

The subsidy pays € 210 monthly for a maximum period of four years. Eligibles may also benefit from an additional € 120 to pay the bank guarantee associated with the rental contract, and a one-time € 600 loan to pay the rent deposit in case they sign a new rental contract. To appreciate the magnitude of the subsidy, it can be useful to compare it with the average Spanish youngsters' monthly earnings. Average gross monthly earnings of young people in the 20-24 age brackets amount to € 1,100 in 2008.<sup>7</sup> The subsidy is therefore equivalent to almost 20 percent of the average gross salary of a young person. Moreover, young people who receive the subsidy devote on average 25 percent of their income to pay the rent, while they would devote 42 percent to pay the same amount in the absence of the subsidy. Finally, the subsidy is

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<sup>6</sup>Solon et al., (1991) Bjorklund et al., (2002) Mazumder (2008), and Schnitzlein (2014) investigate sibling effects in income. Björklund and Salvanes (2011) and Nicoletti and Rabe (2014) in school achievement. Kuziemko (2006) in fertility, and Altonji et al., (2013) in teenage substance use.

<sup>7</sup>Source: Spanish Wage Structure Survey, 2008.

likely to make household formation affordable for many youngsters, as the maximum affordable rent for the average young household is € 560, while the average rent is € 626. By July 2011, the subsidy was given to 35 percent of households headed by an individual aged 22 to 29.<sup>8</sup> The total cost of the program from January 2008 to December 2011 was € 400 million (approximately, \$523 million).

To be eligible for the subsidy, youngsters need to be in the 22-29 age bracket and have a rental contract. This includes individuals that had a rental contract before becoming eligible.<sup>9</sup> Those who do not have a rental contract may request the subsidy conditional on providing the contract signed in three months time. Eligibles need to certify that they are employed, autonomous workers, grant holders, or receivers of a periodic social benefit (including unemployment benefit). The latter are also required to have worked for at least six months or provide evidence that the social benefit will last for at least six months. For all the eligibles, the net source of income must not exceed € 1,500 per month. EU citizens and non-EU citizens with a permanent resident permit are eligible. If several individuals are sharing accommodation, each young adult entitled to the subsidy receives a share of the subsidy proportional to the number of people who sign the rental contract. Individuals who rent out from close family members are not eligible.

In our empirical analysis we define subsidy eligibility exclusively on the basis of age and time survey, the only criteria that are impossible to manipulate.<sup>10</sup> Omitting the employment status when defining eligibility does not constitute a threat to our identification strategy. First, because employment is not a necessary condition for receiving the subsidy: eligibles include grants holders and social benefits recipients. Second, because the requirement holds only when the individual applies for the subsidy. After that, lack of employment does not imply the benefit's withdrawal.

## 2.2 Data

Our main dataset consists of the 2005-2012 waves of the Spanish data from the European Union Statistics on Income and Living Conditions (EU-SILC). The data contain a wide range of information on individual's and household's characteristics. Individuals are surveyed yearly and stay in the sample for four consecutive years. If an individual forms a new household during that time, both the old and the new households are interviewed. If the new household does not respond (in cases, for instance, in which the individual has died or moved abroad), the old household reports whether the individual has formed a new household.

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<sup>8</sup>Source: Spanish Ministry of Housing.

<sup>9</sup>Note that the policy generates no incentives to 21 year-olds to postpone emancipation. Emancipated 21 year-olds will be entitled to the same amount of subsidy as soon as they become eligible. Aparicio and Oppedisano (2014) provide an empirical test that confirms that postponement was not significant.

<sup>10</sup>Aparicio and Oppedisano (2014) show that the majority of individuals who were eligible in terms of age actually fulfilled all other criteria.



The estimation sample includes 14,298 individuals in the 18-26 age group. We choose the 18-26 age range because it is the widest interval for which all treated individuals are entitled to the subsidy for the four years. Aparicio and Oppedisano (2014) show that the effect of the subsidy is very similar across the 21-22, 20-23, 18-26 and 18-33 age groups. The panel is unbalanced: individuals who become 18 during the survey period enter the sample only after they turn 18. Similarly, individuals who become 27 during the survey period exit the sample as soon as they turn 27. Building up an independent household is measured by a dummy equal to one if the individual has moved out of parental home over the period 2006-2012.<sup>11</sup> Sibling's household formation is defined with a dummy equal to one if the sibling has left the parental home in the same time period. Our analysis focuses on how the policy affects flows out of the parental home, and therefore indirectly the stock of individuals living independently from their parents.

Table 1 shows descriptive statistics for our sample, i.e., young adults in the 18-26 age bracket in each of the six waves of the EU-SILC data. Around 4.9 percent of the sample has left parental home over the period 2006-2012, and 5.6 percent of individuals have at least one sibling who has left parental home over the same time period. Around 30 percent of individuals and siblings are eligible for the rental subsidy. Slightly more than half of respondents are male.

The basic idea behind the identification strategy is illustrated in Figure 1, which shows the trends of household formation rates for individuals with siblings in the eligible and not eligible age group. The figure shows that the trends were parallel before the policy was implemented, and diverge after the introduction of the policy, illustrating the source of exogenous variation of our identification strategy. The trend in household formation rates of individuals who have eligible siblings lies below that of those who have non-eligible siblings: individuals with eligible siblings are on average less likely to form new households than those who have non-eligible siblings. This happens because individuals with eligible siblings are younger on average, although the age difference is not significant. The two trends diverge slightly from 2008 to 2009 and start diverging significantly after 2009. This delay can be explained by the nature of our data. Individuals that report having formed a household in 2009 have actually done it at any point in time between their interview in 2008 and the interview in 2009.

### 3 Empirical strategy

Peer effects occur when an individual's action influences the action of another individual in the same social group. However, measuring peer effects has proven difficult (Manski, 1993). First, if peers  $i$  and  $j$  affect each other, then it is difficult to separate

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<sup>11</sup>The time frame does not include 2005 because, for individuals interviewed in 2005, we observe household formation decisions from 2006.

out the actual causal effect that individual  $i$ 's outcome has on individual  $j$ 's outcome. This is commonly called the reflection problem and it is likely to arise whenever individual and peer behaviour are determined simultaneously. Second, unobserved group characteristics and individual traits that are correlated within the group may induce artificial correlation among peer outcomes. Third, self-selection in peer groups represents another challenge in the estimation of peer effects.

In our setting, which looks at living arrangements within the household, the first two challenges are present: siblings are likely to influence each other; and unobservable family and individual characteristics, such as strength of family ties, taste for independence and privacy, are likely unobserved and correlated among siblings. However, self-selection is not an issue in the context of exogenously-formed peer groups as siblings.

Our objective is to estimate the causal effect of sibling's household formation on individual's household formation. Ideally, we would estimate the following equation:

$$y_{i,t} = \alpha_0 + \alpha_1 y_{j,t} + \alpha_2 X_{i,t} + \alpha_3 X_{j,t} + \lambda_{i,j} + \varepsilon_{i,t} \quad (1)$$

where  $y_{i,h,t}$  is a dummy equal to one if individual  $i$  has left parental home in period  $t$  and  $y_{j,t}$  is the corresponding value for sibling  $j$ . The coefficient  $\alpha_1$  captures the effect of having a sibling who has formed a new household on the individual's probability of leaving parental home. The vectors  $X_{i,t}$  and  $X_{j,t}$  contain individual's and sibling's control variables,  $\lambda_{i,j}$  represents the vector of the sibling-pair fixed effects and  $\varepsilon$  is the error term. Unfortunately, the OLS estimated  $\alpha_1$  coefficient would be biased due to reflection and correlated unobservables.

In order to obtain an estimate that is informative about the causal effect of sibling's choices on individual's household formation, we take advantage of the exogenous increase in the propensity to form a new household for individuals in the 22-26 age range induced by the introduction of the Spanish rental subsidy. Our identification strategy relies on exogenous variation in siblings' eligibility for the rental subsidy. We exploit two sources of variation. One source of variation is determined by the year of the interview. Individuals interviewed before 2008 did not benefit from the program, since the rental subsidy only came into force in January 2008 and hence, only some individuals interviewed after that date were fully eligible. The other source of variation arises from age. Due to the eligibility criteria established by the law, only individuals in the 22-26 age group were entitled to the subsidy.

For our strategy to be meaningful, we first need the subsidy to be effective in promoting household formation in the sample of siblings. We check this using the following specification:

$$y_{j,t} = \beta_0 E_j + \beta_1 T_t + \beta_3 E_j T_t + \beta_4 E_i + \beta_5 E_i T_t + \beta_6 X_{j,t} + \beta_7 X_{i,t} + \lambda_{i,j} + \varepsilon_{jt} \quad (2)$$

where  $E_j$  is a dummy variable for sibling  $j$  being in the 22-26 age group,  $T_t$  is a year dummy equal to one for individuals interviewed in 2008 and after, and zero otherwise.

The coefficient  $\beta_3$  captures the effect of the sibling being eligible for the subsidy on the probability that the sibling forms a household. The specification includes the following individual and sibling controls: survey year dummies, individual’s and sibling’s age, individual’s and sibling’s male dummy, dummy for individual’s region of residence, dummy for the month of interview, a dummy capturing the number of times the individual has been interviewed in the panel, and sibling-pair fixed effects. We cluster residuals at the household level to account for this common unobservable household component, which may capture taste for independence or attachment to the family correlated among siblings affected by the same parents.

In our baseline specification, we assume household formation is a function of siblings’ subsidy eligibility and individual controls. The equation reads as follows:

$$y_{i,t} = \gamma_0 E_i + \gamma_1 T_t + \gamma_3 E_i T_t + \gamma_4 E_j + \gamma_5 E_j T_t + \gamma_6 X_{i,t} + \gamma_7 X_{j,t} + \lambda_{i,j} + \varepsilon_{i,t} \quad (3)$$

where the coefficient  $\gamma_5$  captures the causal impact of a sibling’s eligibility on the individual’s probability of leaving parental home, conditional on the individual’s eligibility for the subsidy. Note that in our main specification we estimate the effect of sibling’s eligibility on the individual’s likelihood of leaving the nest in the same year. To assess whether the effect persists over time, we check whether a sibling’s forming a new household affects the individual’s decision of leaving parental home one year later.

The validity of the estimation proposed in Equation 3 relies on the use of panel data and the exogeneity of the rental subsidy. The panel data nature of our sample allows us to difference out any sibling-pair fixed over time unobservable component.<sup>12</sup> However, in the context of a standard fixed-effect estimation, it remains difficult to rule out the possibility of reverse causality. The exogenous change induced by the rental subsidy addresses concerns arising both from potential omitted variables and reverse causality.

Equation 3 is informative about how policies promoting individuals’ household formation affect their siblings. However, it is also a reduced form approach to estimate sibling effects on household formation. For the latter purpose, we could have opted for a two-stage least-squares (TSLS) estimate in which subsidy eligibility serves as an instrument for sibling’s household formation. To consistently estimate the size of the sibling effect via TSLS, one also needs to assume that the only channel through which individuals are affected by siblings’ eligibility is by siblings’ household formation. This could be problematic if household formation means something different before and after the subsidy implementation, with individuals forming a household under the subsidy sending a different signal to their siblings. TSLS also requires the monotonicity assumption that the subsidy would not induce any young individuals to

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<sup>12</sup>We also estimated all the outcomes using household and individual fixed effects and results are invariant.

stay longer at parental home which may have happened if the subsidy increased competition for accommodation. Moreover, the assumptions required for the estimation of average treatment effects by TSLS are incompatible with the discrete nature of the outcome, the endogenous variable and the instrument (Chesher and Rosen, 2013). Finally, the alternative option of non-parametric instrumental variable approach as in Chesher (2009) delivers too wide intervals in our case. We discuss the results from the TSLS estimation in Section 4, keeping in mind these caveats in interpreting the coefficients.

## 4 Results

### 4.1 Individual and sibling’s household formation: OLS regression

We first report the results of the naive estimation of Equation 1 by OLS in Table 2. In our analysis, we explore whether sibling effects change with siblings’ age difference, which is inversely related to the willingness to imitate each other. To this, we estimate Equation 1 separately for the following three samples: all siblings, siblings who are five and two years apart. In all columns, we control for the full set of individual and sibling’s characteristics and for sibling-pair fixed effects. The OLS estimates show positive and significant correlations between sibling’s household formation in all specifications, with the effect increasing the lower is the age difference between siblings.

The positive correlation between individual and sibling’s household formation could be easily justified by the common background shared by siblings or be an outcome of the reflection problem. To learn about causal effects, we next interpret the specifications that use subsidy eligibility as an explanatory variable.

### 4.2 The impact of the subsidy on sibling’s household formation

Our identification strategy relies on the effectiveness of the rental subsidy in fostering household formation among siblings. Table 3 presents estimates of the coefficient  $\beta_3$  in Equation 2. The different columns replicate the structure of Table 2. In each column we control for the whole set of individual and sibling’s characteristics, and for sibling-pair fixed effect. The dependent variable is a dummy equal to one if an individual’s sibling has left the parental home, and zero otherwise. The coefficient of interest is the interaction between the dummy for being interviewed after 2008, and the dummy for the sibling’s being in the 18-26 age group, which captures the sibling’s eligibility to the subsidy. The coefficient is positive and statistically significant in all three specifications. The size of the coefficient decreases for shorter age

differences between siblings. If individuals with close-in-age siblings enjoy living with their siblings, they may be less likely to respond to the rental subsidy. The coefficient in the full sample indicates that subsidy eligibility increases the propensity to leave the nest by 6 percent. These estimates are higher than those obtained by Aparicio and Oppedisano (2014), who estimated a lower bound effect of 3 percent. A possible explanation for the higher coefficient could rely on the fact that the time frame used in the two papers is different: while in the previous paper we focused on the 2006-2009 time period, here we look at the 2006-2012 time period. If it takes time for the policy to be known among eligible young adults and for the applications for the subsidy to be processed, the effect of the policy should increase over time, and be on average larger if a wider time frame is considered.<sup>13</sup>

### 4.3 The impact of sibling’s eligibility on household formation

Results in Table 3 show that siblings’ eligibility for the rental subsidy significantly affects the probability that the sibling leaves the nest. In this section, we exploit the exogenous variation in exposure to the subsidy across youngsters, to assess the causal impact of the sibling’s eligibility to the subsidy on the probability of forming a new household.

We estimate Equation 3 and focus on the interaction between the post policy dummy and the dummy equal to one if one of the siblings is in the eligible 22-29 age group. Table 4 shows that the estimate of the impact of sibling’s subsidy eligibility on household formation differs across specifications. The estimated effect is negative for the full sample of siblings and for siblings that are at most five years apart, but becomes positive for siblings that are two years apart. The latter effect is not significant, either because of the small sample size or the lower effectiveness of the subsidy for this subsample. Although the effect is not significant, the coefficient is positive, suggesting that imitation effects may be present among siblings with small age difference. We explore this possibility in the next Subsection. In terms of magnitude, sibling’s subsidy eligibility decreases the probability of leaving the nest by 3 percentage points when the sample includes all siblings. For siblings at most five years apart the effect declines to 2.5 percentage points. The direction of the effect estimated in the reduced form specification is opposite to the positive effect delivered by the naive OLS estimation.

The second row of the Table also shows the effect of the individuals’ subsidy eligibility on their probability of leaving the nest: the effect is positive and significant, consistently with the estimates obtained in Table 3.

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<sup>13</sup>Spanish newspapers documented that in some regions there was a delay in processing the subsidy during the first months of its application due to lack of communication between administrative entities (El Pais, 05/02/2008).

## 4.4 Mechanisms and alternative specifications

Next, we explore the mechanisms behind the estimation results in Table 4. First, we shed more light on the negative sibling effects in the specification with the full sample of siblings. Secondly, we explore whether imitation can be behind the absence of negative sibling effects for close-in-age siblings. To achieve the first objective, we interact the sibling's eligibility dummy with a set of time varying parental characteristics. Results are reported in Table 5. In the first column, we include the interaction of the sibling's eligibility dummy with a dummy indicating whether the youngest parent is younger than fifty years old. This interaction has a positive and significant effect on individual's household formation. The magnitude of the coefficient is such that it offsets 80 percent of the overall negative effect, suggesting that parental age can be behind the negative sibling effect: if the youngest partner is more than 50 years old then the household formation decision of one sibling will induce the other one to remain in the parental home. Differently, the presence of a younger than fifty years old parent, who may take care of the older partner, will not deter the individual's decision to leave parental home after the sibling has left.

In the second column, the sibling's eligibility dummy is interacted with a dummy equal to one if at least one of the parents is healthy. This dummy is constructed from the item response that assesses the individual's general health status. Respondents can define their health status as very good, good, regular, bad or very bad. We define an individual to be healthy if her health status is regular, good, or very good. The coefficient of the interaction between the sibling's eligibility dummy and the dummy for healthy parents is positive and statistically significant. Again, the magnitude of the estimated coefficient is such that it offsets more than 80 per cent of the overall effect. This result indicates that another channel through which negative siblings spillover effects arise is through parental health: if both parents are ill and one sibling leaves the nest, the other sibling will respond by remaining in the parental home. If at least one parent is healthy, the remaining sibling does not show a lower probability of following her sibling.

We explore whether negative sibling effects are attenuated by the presence of other siblings in the household, who can share the burden of taking care of the parents. When considered, the interaction between sibling's eligibility and the number of siblings who still live in the household, although positive, is not significant.

Negative siblings' effects may also arise because the remaining sibling may take advantage of the higher quantity of public goods available to her as a consequence of her sibling having left the parental home. We try to assess this effect using the number of rooms in the house at the time the household is interviewed for the first time as a proxy for public good. We do not find that individuals living in smaller houses are more likely to leave the nest than those living in larger houses, and we therefore tend to count out that public goods as measured by the space in the house explain negative siblings spillover effects. We also look at whether household income, adjusted by household size, affects the coefficients of interest, but we do not find any

significant pattern.

Note that negative siblings' effects can also be explained by an increase in household resources enjoyed by the remaining children if parents are willing to bribe their children into staying at home. Unfortunately, our data do not convey information on transfers from parents to the child, and therefore we cannot assess the importance of this mechanism in siblings' interactions.

Next, we explore whether imitation can play a role in the sample of close-in-age siblings. We depart from the work of Barr and Hayne (2003) and as in Altonji et al. (2013), we assume that younger siblings are more likely to imitate older ones. We define a dummy variable equal to one if the individual is the older sibling. Results are reported in column 3 of Table 5. The coefficient of this interaction is negative, although statistically significant. However, the coefficient of the sibling's eligibility dummy remains positive and becomes statistically significant. This indicates that sibling effects become positive when the effect operates from older to younger siblings and siblings are close-in-age: when individuals leave the nest, their less-than-two-years younger siblings are more likely to follow her in the decision to form a new household.

Results in Table 7 reports estimates from the TSLS, which is valid under the additional assumptions of monotonicity and the exclusion restriction. The coefficients of the two specifications with the full sample and that of siblings five years apart are negative and statistically significant at 10 and 10.5 percent levels, indicating that the emancipation of one sibling reduces the individual's probability of emancipating by approximately 50 percent, confirming the direction of the sibling effects. The effect becomes positive and not significant in the third column, where only close-in-age siblings are considered, but the F-test indicates the weakness of the first stage for this subsample. As we mentioned in Section 3, the assumptions required for the estimation of average treatment effects by TSLS are incompatible with the discrete nature of the outcome, the endogenous variable, the instrument and most of our controls. Therefore, we omit a detailed interpretation of the size of the TSLS coefficients, as it is uninformative of the true size of the effect in this setting.

In all our analysis we study contemporaneous sibling effects, showing the effect of a sibling's eligibility on the individual's choice to form a new household in the same year. However, it may be interesting to look at whether these effects persist, or dissipate over time. As individuals are interviewed only four times, we can analyze the effect of sibling's eligibility in one year on next year probability that the individual will leave parental home. In Table 7 we estimate Equation 3 with lagged (rather than contemporaneous) sibling's eligibility dummies. Results show that siblings' eligibility one year before reduces the probability of the individual's forming a new household by almost 3 percent, with the effect being significant at conventional levels only for siblings less than 5 years apart.<sup>14</sup> The effect is positive and significant in the sample of siblings that are at most 2 years apart and the estimated coefficient indicates that the effect is slightly higher than 3 percentage points. Hence, our results confirm the

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<sup>14</sup>In the specification using the full sample, the estimated coefficient is significant at the 11%.

findings from the contemporaneous effects specification and indicate that the effects persist at least one year later.

Finally, we provide further evidence on the validity of our difference-in-differences estimation strategy and perform a placebo test pretending that the policy was implemented in 2007 rather than 2008. The results of this exercise are reported in Table 8. Estimated coefficients are small in magnitude and non-significant, indicating that there are no pre-existing trends in the data that could drive our results.

## 5 Conclusion

The transition to adulthood is a complex process made of several interrelated steps such as leaving school, finding a job, finding a partner, etc. It culminates with the formation of an independent household, possibly with a partner, and usually implies moving out of the parental residence. The increasing late age at which young adults in Southern Europe postpone household formation decisions has led governments, in the last decade, to implement policies that foster the decision to leave the nest, by helping young adults coping with rental expenses. If spillover effects among siblings in the choice of leaving parental home exist, then these incentives may amplify or reduce the aggregate impact of these policies depending on whether sibling effects are positive or negative, which remains an open empirical question.

We empirically analyzed the role of sibling effects on household formation decisions in the context of Spain, a Southern European country characterized by late household formation. To this, we make use of the exogenous variation in household formation for a subset of young individuals eligible for the rental subsidy, and exploit the panel data dimension of the EU-SILC data. Our results suggest that on average siblings' interactions reduce the impact of policies that foster household formation, except for the case in which subsidy recipients are close-in-age siblings, consistently with the hypothesis that the willingness to imitate a sibling is stronger in correspondence of small age gaps. The sibling effects are negative and significant for individuals with ten to five years apart. When exploring the channels through which negative sibling effects are exerted, we find that individuals who further delay the decision to form a new household after a sibling has left do so in presence of old or ill parents. We cannot rule out with available data that the enjoyment of higher public goods for the remaining sibling, or transfers from the parents that try to bribe the remaining children at home are other mechanisms at play.

Overall, in the context of Southern European countries, where family ties are strong, there is more reliance on home production and less participation in market activities as individuals tend to trust more family members (Alesina and Giuliano, 2010). Caring for the elderly is a typical activity demanded to household production in these countries. A policy that aims at fostering the household formation process should account for household composition as well. In particular, our finding indicate that policy makers should target the household rather than the individual, and



combine policies for young adults with policies for elderly.

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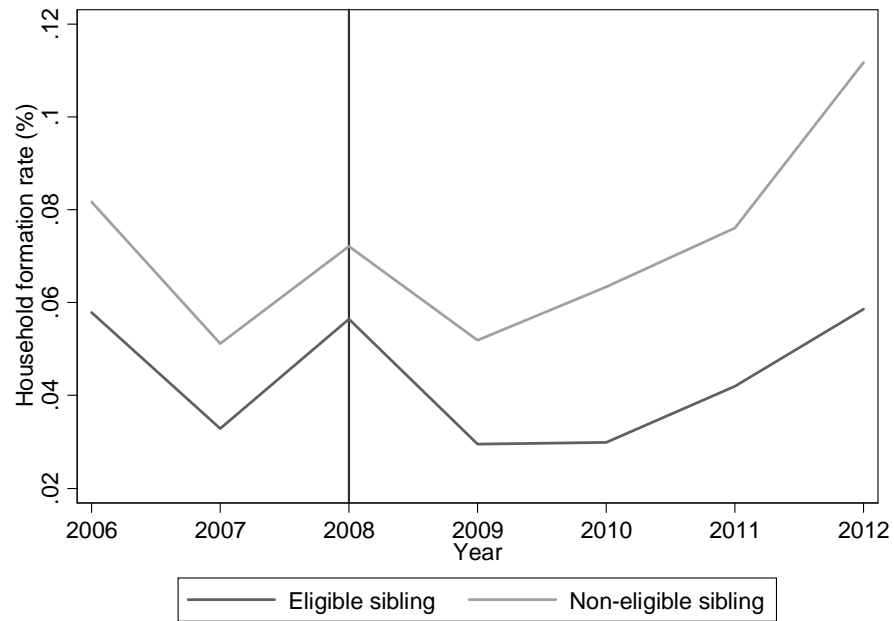
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## Figures

**Figure 1: Trends in household formation rates for eligibles and non-eligibles**



Data source: EU-SILC 2005-2010

## Tables

**Table 1: Descriptive Statistics**

<b>Variable</b>	<b>Mean</b>	<b>Std. Dev.</b>	<b>Min.</b>	<b>Max.</b>
Parental home leaving	0.049	0.217	0	1
Sibling's parental home leaving	0.056	0.23	0	1
Sibling's subsidy eligibility	0.302	0.459	0	1
Subsidy eligibility	0.301	0.459	0	1
Post-policy period	0.556	0.497	0	1
Age eligibility	0.554	0.497	0	1
Sibling's age eligibility	0.556	0.497	0	1
Male	0.517	0.5	0	1

The number of observations is 14,298.

**Table 2: OLS regression of sibling household formation on individual’s household formation**

	Full sample	5 years apart	2 years apart
Sibling’s household formation	0.136 (0.03)***	0.158 (0.032)***	0.205 (0.047)***
Subsidy eligibility	0.07 (0.016)***	0.063 (0.017)***	0.039 (0.02)*
Post-policy period	-.916 (0.109)***	-.854 (0.116)***	-0.089 (0.109)
Age eligibility	-.051 (0.015)***	-.044 (0.015)***	-.017 (0.018)
Sibling’s age eligibility	-.0002 (0.009)	0.003 (0.009)	-.012 (0.012)
Obs.	14288	12525	5632
$R^2$	0.096	0.101	0.116

Notes: EU-SILC data. This table shows the fixed effects OLS regression of individual’s household formation on sibling’s household formation. Controls include dummy for being interviewed after 2008, individual and sibling’s age eligibility dummy and the interaction between the individual’s age eligibility dummy and the dummy for being interviewed after 2008. Other controls include: individual’s and sibling’s gender, individual’s and sibling’s age, survey year dummies, month of interview dummies, regional dummies, and a dummy for the number of times the individual was interviewed. Standard errors clustered at the household level in brackets. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

**Table 3: The impact of sibling’s eligibility on sibling’s household formation**

	Full sample	5 years apart	2 years apart
Sibling’s eligibility	0.064 (0.016)***	0.06 (0.017)***	0.037 (0.022)*
Subsidy eligibility	-.029 (0.013)**	-.025 (0.013)*	0.017 (0.019)
Post-policy period	-1.070 (0.12)***	-1.029 (0.132)***	-.069 (0.093)
Age eligibility	0.021 (0.012)*	0.022 (0.012)*	-.013 (0.018)
Sibling’s age eligibility	-.041 (0.016)***	-.035 (0.016)**	-.006 (0.02)
Obs.	14288	12525	5632
$R^2$	0.085	0.084	0.081

Notes: EU-SILC data. This table shows the fixed effects regression of sibling’s eligibility (the interaction between the dummy for being interviewed after 2008, and the dummy for the sibling’s age eligibility) on sibling’s household formation. Controls include dummy for being interviewed after 2008, individual and sibling’s age eligibility dummy and the interaction between the individual’s age eligibility dummy and the dummy for being interviewed after 2008. Other controls include: individual’s and sibling’s gender, individual’s and sibling’s age, survey year dummies, month of interview dummies, regional dummies, and a dummy for the number of times the individual was interviewed. Standard errors clustered at the household level in brackets. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .



**Table 4: The impact of sibling’s eligibility on household formation**

	Full sample	5 years apart	2 years apart
Sibling’s household formation	-0.031 (0.013)**	-0.025 (0.013)*	0.021 (0.019)
Subsidy eligibility	0.06 (0.016)***	0.056 (0.017)***	0.038 (0.021)*
Post-policy period	-1.043 (0.116)***	-1.000 (0.128)***	-0.094 (0.09)
Age eligibility	-0.044 (0.015)***	-0.038 (0.016)**	-0.017 (0.019)
Sibling’s age eligibility	0.017 (0.012)	0.019 (0.012)	-0.021 (0.018)
Obs.	14288	12525	5632
$R^2$	0.079	0.079	0.077

Notes: EU-SILC data. This table shows the fixed effects regression of sibling’s eligibility (the interaction between the dummy for being interviewed after 2008, and the dummy for the sibling’s age eligibility) on individual’s eligibility. Controls include dummy for being interviewed after 2008, individual and sibling’s age eligibility dummy and the interaction between the individual’s age eligibility dummy and the dummy for being interviewed after 2008. Other controls include: individual’s and sibling’s gender, individual’s and sibling’s age, survey year dummies, month of interview dummies, regional dummies, and a dummy for the number of times the individual was interviewed. Standard errors clustered at the household level in brackets. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

**Table 5: Mechanisms**

	Full sample		2 years apart
	Parental age	Parental health	Older than sibling
Sibling's eligibility	-0.055 (0.022)**	-0.109 (0.033)**	0.037 (0.021)*
Sibling's eligibility*Parent < 50	0.044 (0.025)*		
Parent < 50	-0.012 (0.022)		
Sibling's eligibility*Healthy parent		0.09 (0.032)**	
Healthy parent		0.003 (0.021)	
Sibling's eligibility*Older than sibling			-0.013 (0.034)
Subsidy eligibility	0.06 (0.016)**	0.051 (0.015)**	0.021 (0.023)
Post-policy period	-1.058 (0.115)**	-0.987 (0.121)**	
Age eligibility	-0.042 (0.015)**	-0.040 (0.015)**	0.0003 (0.02)
Sibling's age eligibility	0.03 (0.02)	0.048 (0.022)**	-0.054 (0.02)**
Obs.	14288	14020	5632
$R^2$	0.08	0.068	0.079

Notes: EU-SILC data. This table shows the results of regressing sibling's eligibility on individual's household formation controlling for time varying household characteristics and their interactions with sibling's eligibility. In the first column we add a dummy for at least one parent younger than 50 years old. In the second column, we include an indicator for at least one parent healthy. In the third column, we add the control for the sibling's leaving home being older than the individual. Controls include dummy for being interviewed after 2008, individual and sibling's age eligibility dummy and the interaction between the individual's age eligibility dummy and the dummy for being interviewed after 2008. Other controls include: individual's and sibling's gender, individual's and sibling's age, survey year dummies, month of interview dummies, regional dummies, and a dummy for the number of times the individual was interviewed. Standard errors clustered at the household level in brackets. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

**Table 6: The impact of sibling’s household formation on household formation (TSLS)**

	Full sample	5 years apart	2 years apart
Sibling’s household formation	-.493 (0.262)*	-.449 (0.277)	0.564 (0.588)
Subsidy eligibility	0.044 (0.025)*	0.042 (0.024)*	0.028 (0.033)
Post-policy period	-1.170 (0.222)***	-1.081 (0.228)***	.157 (0.259)
Age eligibility	-.032 (0.021)	-.027 (0.021)	-.010 (0.026)
Sibling’s age eligibility	-.001 (0.012)	0.004 (0.012)	-.018 (0.016)
Obs.	11269	10104	4532
$F$ test	$F(1, 1680)=15.00$	$F(1, 1511)=12.08$	$F(1, 725)=3.08$
Prob $>F$	0.000	0.000	0.080

Notes: EU-SILC data. This table shows the fixed effects regression of individual’s household formation on sibling’s household formation. Controls include dummy for being interviewed after 2008, individual’s age eligibility dummy and the interaction between the individual’s age eligibility dummy and the dummy for being interviewed after 2008, the dummy for the sibling’s age eligibility. Sibling’s household formation is instrumented with the interaction between the sibling’s age eligibility dummy and the dummy for being interviewed after 2008. Other controls include: individual’s and sibling’s gender, individual’s and sibling’s age, survey year dummies, month of interview dummies, regional dummies, and a dummy for the number of times the individual was interviewed. Standard errors clustered at the household level in brackets. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

**Table 7: The impact of sibling’s eligibility in t-1 on household formation**

	Full sample	5 years apart	2 years apart
Sibling’s eligibility in t-1	-0.027 (0.017)	-0.028 (0.017)*	0.033 (0.02)*
Subsidy eligibility	0.063 (0.022)***	0.054 (0.023)**	0.064 (0.037)*
Post-policy period	-0.055 (0.019)***	-0.054 (0.02)***	-0.038 (0.023)
Post-policy period in t-1	-0.05 (0.061)	-0.044 (0.06)	-0.11 (0.082)
Age eligibility	-0.063 (0.022)***	-0.054 (0.022)**	-0.045 (0.035)
Sibling’s age eligibility in t-1	-0.002 (0.015)	0.0005 (0.015)	-0.042 (0.021)**
Obs.	6984	6374	2951
$R^2$	0.083	0.082	0.096

Notes: EU-SILC data. This table shows the fixed effects regression of individual’s household formation on sibling’s eligibility the year preceding the survey (the interaction between the dummy for being interviewed after 2009, and the dummy for the sibling’s age eligibility in t-1). Controls include dummy for being interviewed after 2008, individual’s age eligibility dummy and the interaction between the individual’s age eligibility dummy and the dummy for being interviewed after 2008, the lagged dummy for the sibling’s being interviewed in 2008, the lagged dummy for the sibling’s age eligibility dummy, and the interaction between the two. Other controls include: individual’s and sibling’s gender, individual’s and sibling’s age, survey year dummies, month of interview dummies, regional dummies, and a dummy for the number of times the individual was interviewed. Standard errors clustered at the household level in brackets. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .

**Table 8: Placebo test**

	Full sample	5 years apart	2 years apart
Sibling's eligibility	0.019 (0.018)	0.008 (0.019)	-0.040 (0.029)
Subsidy eligibility	0.057 (0.017)***	0.054 (0.018)***	0.062 (0.026)**
Post-policy period	-0.032 (0.018)*	-0.029 (0.019)	-0.075 (0.088)
Age eligibility	-0.057 (0.019)***	-0.043 (0.019)**	-0.002 (0.023)
Sibling's age eligibility	-0.0009 (0.009)	0.003 (0.01)	-0.009 (0.013)
Obs.	14288	12525	5632
$R^2$	0.079	0.078	0.078

Notes: EU-SILC data. This table shows the fixed effects regression of sibling's placebo eligibility (the interaction between the dummy for being interviewed after 2007, and the dummy for the sibling's age eligibility) on the probability of household formation. Controls include dummy for being interviewed after 2007, individual and sibling's age eligibility dummy and the interaction between the individual's age eligibility dummy and the dummy for being interviewed after 2007. Other controls include: individual's and sibling's gender, individual's and sibling's age, survey year dummies, month of interview dummies, regional dummies, and a dummy for the number of times the individual was interviewed. Standard errors clustered at the household level in brackets. \*\*\*  $p < 0.01$ , \*\*  $p < 0.05$ , \*  $p < 0.1$ .