The Risk of Job Loss, Household Formation and Housing Demand: Evidence from Differences in Severance Payments[1]

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Abstract

Sharp falls in uncertainty about job loss ought to increase the expenditure of durable goods and the likelihood of undertaking costly-to-reverse decisions, such as becoming a homeowner. We exploit large differences in firing costs across contract types in the Spanish labor market to identify the causal link between sharp changes in the risk of job loss and the timing of different forms of household formation among youths. Our identification strategy uses variation in regional incentives to promote high-firing cost contracts between 1997 and 2009. Hiring subsidies predict job stability. An increase of 1% in the stock of permanent workers increases the probability of forming a new household by 1.2%. The probability of forming a household increases in 0.5% through home ownership and in 0.7% through renting the new accommodation. We also encounter that the probability of forming a new household peaks two years after an exogenous increase in job security, mainly through renting. The delay in renting is consistent with the decision of living with a partner among individuals that have not accumulated enough savings for a downpayment on a house. The analysis of individuals that demand for new accommodation due to regional subsidies for job-contract conversions indicates that these individuals have not been able to save for a downpayment yet. We discard the delay in rental is due to the presence of borrowing constraints among renters.

JEL Codes: J1, J2, D91

Keywords: Job insecurity, household formation, housing investments.

Introduction

The housing decisions in several countries around the world are changing. That pattern is specially noticeable among young households. In Spain, between 2002 and 2012, the percentage of home owners below 35 years of age has fallen by between 10 and 15 percentage points. In Germany, the age of becoming a homeowner has increased by two years over the last two decades (Arrondel et al, 2016). Going beyond the choice of owning vs renting, the living arrangements of young adults is changing as well. In the United States, a number of studies have documented a slowdown in the rate of household formation.[1]

Those changes are relevant for several reasons. Housing is the most prominent component of portfolios for most households (Badarinza et al, 2016, Matthä et al, 2017). Fluctuations in the value of housing is a key driver of household consumption (see Campbell and Cocco, 2007) and the fraction of homeowners shapes the effectiveness of monetary and fiscal policy, as (indebted) homeowners are sensitive to monetary and fiscal policy surprises (Cloyne et al. 2017).

Candidate explanations for changes in home ownership and, more generally, in living arrangements are changes in job stability, credit conditions, cohort-specific experiences or demographic changes.[2] This paper analyzes in detail the role of job insecurity, or the perceived probability of job loss, on the decision to form a household, owning or renting it.[3] Bloom (2009) shows that uncertainty leads agents to postpone costly-to-reverse decisions. Becoming a homeowner or forming a household may be viewed as examples of such decisions (Becker et al, 2010).[4] Measuring the individual perception of job loss is difficult, as it is hard to identify individuals who are exposed to the risk of job loss before that risk realizes. To that end, we exploit an institutional feature in European labor markets and more concretely, of the Spanish one. Namely, depending on their legal status of their job contract, individuals face very different probabilities of job loss. Laying-off an employee with an open-ended contract is a process that typically entails going to court and incurring in costly severance payments -in Spain, between 33 and 45 wage days per year worked. On the contrary, dismissing workers on fixed-term

[1] See Dettling and Hsu (2018), Bleemer et al. (2014), Paciorek (2013), Kaplan (2012), Cooper and Luengo-Prado (2016), or Bell et al. (2006).

[2] Focusing on youths, Giuliano (2007) discusses the influence of cultural factors in the living arrangements, while Martins and Villanueva (2009) and Dettling and Hsu (2018) discuss the role of the cost of credit. Rosenzweig and Wolpin (1994) find that low-income children are more likely to live with parents (holding parental income constant). Manacorda and Moretti (2006) document that parents with higher income levels are more likely to live with their children, consistent with the predictions of an exchange model of the extended family.

[3] Some authors have documented that jobs are becoming more unstable [see Valletta (1999) for the US and García-Pérez and Rebollo-Sanz (2009) for Spain].

[4] Shore and Sinai (2010) show that, in the presence of adjustment costs, increases in income risk may lead owner-occupier households to *increase* their consumption of housing services. Their study focuses on the intensive margin (the amount of housing units consumed, measured by their price). Here, we focus on the intensive margin.

contracts involves much lower firing costs -between 0 and 8, and possibly zero at the termination of the contract. That difference causes that easily observable individuals face very different probabilities of job loss.

The decision of a firm to grant an open-ended (rather than a fixed-term) contract depends on the worker's expected productivity as well as on local labor demand conditions that also correlate with the propensity to consume housing services (see Topel, 1986). We thus exploit exogenous variation in the firms' incentives to convert fixed-term contracts into open-ended ones. As a result of decentralized labor laws, several Spanish regions introduced different subsidies to firms that converted fixed-term (insecure) contracts into open-ended (secure) ones. Those incentives were introduced between 1997 and 2009 and varied by gender, age of the worker and year.[5] The staggered introduction and subsequent evolution of the incentives to provide workers with job security provides us with a source of exogenous variation across regions, age and gender groups that permit identifying causal impacts of exposure to job insecurity on housing choices.

Thus, our paper contributes to the empirical literature on household formation and housing tenure in mainly three aspects.[6] First, we try to study the causal impact of job security on both household formation and housing tenure choice by exploiting exogenous variation in the risk of job loss and estimating using instrumental variables. With respect to the literature that measures the risk of job loss using regional data, such as the unemployment rate or the incidence of fixed-term contracts, our study analyzes the particular risk of job loss the individuals face. With respect to those studies that measure the risk of job loss using subjective perceptions of job security, our proxy of the risk is the kind of job contracts individuals hold (whether a fixed-term contract or a permanent one). As said before, these two contracts entail very different firing costs and, thus, very different probabilities of entering an unemployment spell. Unlike subjective perceptions of job security, the promotion of a kind of job contract and the firing costs are exogenously affected very often by labor market reforms, which provide us enough exogenous variation to study the causal link between housing demand (household formation) and the risk of job loss.

The second contribution of our paper to the literature of housing demand is that we use a sample of individuals instead of a sample of households in order to analyze how the risk of job loss affect the choice of housing tenure.[7] The sample of household heads is

[5] García Pérez and Rebollo Sanz (2009) coded those subsidies. They study the impact of the amount of the subsidy on worker flows using administrative data and document similar findings to ours.

[6] Kaplan (2012), Bleemer et al (2014), Becker et al (2010), Rosenzweig and Wolpin (1994) and Martínez-Granado and Ruiz-Castillo (2002) have studied the link between job security and household formation and related decisions, such as fertility and marriage.

[7] See Gathergood (2011), Diaz-Serrano (2005), Henley et al (1994), Henley (1998) and Duca and Rosenthal (1994) for some related papers that have addressed on the issue of housing demand and labor market uncertainty. These empirical studies have focussed on the analysis of households, instead of individuals.

not representative of the population of individuals at younger ages, since a non-negligible fraction of individuals aged under 35 at least in Spain are still living with their parents, so the estimates of the impact of job security will suffer from sample composition biases [see Deaton and Paxson (2000) and Chiuri and Jappelli (2003)]. As we use a sample of individuals, we do not study the determinants of the decision of owning vs renting, because we also consider the decision of living with parents or other relatives (coresidence). In this way, home ownership and rental may change in the same direction through the household formation margin.

And finally, there is a growing literature on inferring risks from income histories [see Guvenen and Smith (2014)] to link different sources of income risk to consumer decisions. Our third contribution is to this empirical literature, as we can infer the timing of the risk resolution (when the job conversion takes place). This enables us to study lagged responses of the risk of job loss on the household formation and housing tenure decisions by using retrospective information about the year of hiring in the current job, the exact year in which individuals bought their owner-occupied house or the year in which they rented their accommodation. Therefore, we can study whether both household formation and housing tenure respond simultaneously to an exogenous increase in job security or whether there is a delay in undertaking these decisions.

We use rich survey data from the 2002-2014 waves of the Spanish Survey of Household Finances (EFF), with retrospective information on housing choices as well as detailed information on labor market status to implement several identification strategies. The first strategy examines if the availability of subsidies to contract conversion during the first two years of the employment relationship (when most contract conversions happen) increases the fraction of open-ended contracts, the fraction of adults who live with relatives, and the fraction of households formed, distinguishing between those who are owner-occupiers and renters. The second empirical strategy uses retrospective information on the date of housing purchases and rentals to build a duration model of housing purchases and rentals during job tenure.[8] Namely, leveraging on the fact that most contract conversions happen during the first two years of the contract, we examine whether there are changes in the probability of a house purchase or rental immediately after an exogenous increase of job security or whether there is a delay. Finally, using administrative data from Social Security records, we corroborate that most of job contract conversions from fixed-term contracts to open-ended ones occur in the first two years of workers' job tenure.

Our results suggest a strong relationship between exogenous increases in job security and the probability of forming a new household. An increase in the subsidy for contract

[8] Bover (2010) has also used retrospective information of the EFF data to estimate hazard models for studying housing purchases, housing price expectations and the dynamics of the household housing wealth.

conversion of 1,000 euros increases the stock of open-ended contracts by .8 percentage points and diminishes the stock of individuals living with parents by a similar amount. The increase in household formation results in a similar increase in owning and renting (about .5 percentage points each). When we examine the dynamics of the events, we find that, after an increase in job security, the response in the probability of renting a new accommodation is higher among youths (aged 25-45) than in the full sample of employees (aged 25-64). Secondly, the increase in renting after an increase in job security is rather sticky, as it happens between one and two years after the year when most contract conversions happen. On the contrary, the increase in the probability of owning a house and renting following an increase in job security is quite similar in the full sample of adults. Complementary evidence suggests that the youths affected by the subsidy program, while exposed to risk of job loss, did not accumulate enough savings to be used toward a downpayment when that risk disappears. In addition, we find little evidence that the delay in the response of renting to increases in job security is due to credit constraints. Instead, we encounter evidence that the decision of living with a partner and not having accumulated enough savings for a downpayment on a house might be one potential explanation behind the delay of forming a household through rental.

Our results are not consistent with the notion that young households accumulate precautionary savings that can be used for life-cycle saving when risk disappears, like Crossley and Low (2011) find. In our study, affected youths do not accumulate precautionary savings. Regarding whether job insecurity can explain the fall in home ownership that we detect in Spain, the results are nuanced. For a substantial fraction of our sample, when job security decreases the margin that falls the most is not owning, but renting. Finally, the sticky response of household formation to job security offers an explanation for the slow recoveries of housing markets after the Great Recession.

The paper is organized as follows. Section 1 provides some background on the Spanish labor market and on living arrangements. Section 2 lays out the empirical strategy in the analysis of household formation and housing demand in the medium run, Section 3 describes the data used in the estimation, Section 4 presents the empirical results and Section 5 concludes.

1 Fixed-term contracts and housing choices

1.1 Living arrangements in Spain

Figure 1 displays that the fraction of Spanish males who are renters or homeowners as a fraction of all males (either renting, owning or living with parents, friends or other relatives) in three moments in time: 2002, 2005 (an expansionary period) and 2014 (the end of a severe recession). We note two facts. The 2002 cross-sectional profile of home ownership is rather steep, and home owners reach 80% of the population by age 45. On the other hand, the 2002 fraction of renters hovered around 20% in all ages. Note that both graphs imply that a substantial fraction of the population of youths live with parents (around 20% at age 35 and 70% at age 25). Comparing across waves, we see a substantial fall in home ownership (15 percentage point fall at age 27) as well as an increase in the fraction of renters (25 percentage point increase at age 27). Those big changes are often attributed to the Great Recession, and may be due to several reasons, such as unrealistically high housing prices, difficulty in access to credit and job instability, among others. This paper tries to address the causal link of both housing demand and household formation with job security.

1.2 Fixed term contracts: legal framework

Fixed-term contracts were introduced in various European countries as a way of introducing flexibility -at the margin in labor markets with severe firing costs (see Dolado et al., 2002, Güell and Petrongolo 2007). Contracts with low firing costs could be used for new employment relationships while not changing the firing costs of other existing contract types. Spain was the European country with the strongest prevalence of such fixed-term contracts, providing a laboratory to examine the consequences of high exposure to the risk of losing the job. Fixed-term contracts featured very low indemnities for termination, that were virtually zero if the firm waited until expiration of the term specified in the contract. Bover and Gómez (2004) document that the main exit from unemployment is through a fixed-term contract.

In 1997 a national-wide reform reduced the cost of firing permanent workers from 45 wage-days per year worked to 33 wage-days (see Kugler et al., 2005). At the same time, some of the 17 regional authorities decided to subsidize firms who signed permanent contracts, possibly in response to the growing incidence of fixed-term contracts among vulnerable workers - see García Pérez and Rebollo Sanz (2009), who also examine the impact of those subsidies on labor market flows. Subsidies to contract conversion were typically lump-sum amounts given to firms that proved that a new permanent contract was signed (either by an existing worker whose job was regulated by a fixed-term contract or by a new worker who was unemployed). In some cases, the subsidies took the form of a reduction in the payroll tax . Table A.1, taken from García Pérez and Rebollo Sanz (2009) shows the subsidies by region and demographic groups.[9]

[9] Table A.1 also documents that the size of subsidies varied over time (see the case of Canarias, where subsidies were removed after 1999), and also among demographic groups; Andalucia had special subsidies

Barceló and Villanueva (2016) make an analysis of the effectiveness of regional subsidies in the conversion of fixed-term contracts into open-ended ones. Assuming an interest rate of 4%, they estimate that the mean subsidy amounts to 16% of the yearly labor cost that the worker entails to the firm. Moreover, in regions that implemented subsidies around the 23% of contract conversions were subsidized, what implies that between 5% and 7% of all hires were subsidized in Spain.

1.3 Are fixed-term contracts a good proxy for the risk of job loss?

Workers hired under fixed-term contracts were much more likely to experience a transition to non-employment than comparable workers hired under an open-ended contract. Own computations from the Spanish Labor Force Survey suggest that the risk of job loss does vary substantially between both types of workers; while workers covered by a fixed-term contract faced a yearly probability of transiting into unemployment of about 10%, the corresponding probability for a worker covered by a permanent contract was about 2%.[10]

Whether or not workers covered under a fixed-term contract actually *perceive* a higher chance of transiting into unemployment than other workers is less clear (see Manski and Straub, 2000). Workers whose employment relation is regulated by a permanent contract may perceive that if they lose their job, there are few chances of finding a comparable one, because most exits from unemployment typically happen through fixed-term contracts. Alternatively, some workers covered by a fixed-term contract may still perceive small chances of moving into unemployment because they work in a local labor market with a strong demand for their particular skills. We settle the issue by examining whether changes in the type of job contract increases the worker's perception of job security by examining how satisfaction with job security varies around the upgrade of a fixed-term contract into an open-ended one. We use a sample drawn from the European Community Household Panel. Figure A.1 documents that satisfaction with job security increases monotonically with tenure up to contract conversion. However, the increase in job security is specially high during the year when their fixed-term contract is upgraded into an open-ended one. Interestingly, the relationship between tenure on the job and satisfaction fluctuates around zero after contract conversion. That pattern suggests that once the worker obtains a high firing cost contract, additional years of tenure do not add much more in terms of the perception of job security. That evidence suggests that workers seem

for firms who changed the contract of workers below 30 years of age into a permanent one. Some regions had higher subsidies for females (Comunidad Valenciana, Cantabria and Galicia, for example).

[10] Regressions of the probability of transiting into unemployment on an indicator of fixed-term contract, age, occupation and industry dummies deliver similar results.

to feel more certain about their jobs when these are covered by a permanent contract, instead of a fixed-term one.

Figure 2 shows the individuals' subjective probability of losing their job over the next twelve months in 2011 broken down by the kind of job contract using data from the Spanish Survey of Household Finances. For each job contract, a histogram shows the proportion of workers that expects to lose their job with zero probability, a probability of 10%, 20%, and so on, up to a probability of 100% (they expect to lose the job with total certainty). More than 50% of workers holding an open-ended contract expect to lose their job with a probability of less than 10%. Among the temporary workers, the median worker expects to lose her job with a probability of 50%. Thus, the difference in the expected probability of losing the job between (the median) temporary and permanent workers is 40%. Therefore, the kind of job contract seems to be a good proxy of the risk of job loss.

2 The identification strategy in the empirical analysis: regional variation in subsidies to contract conversion

This paper investigates if job insecurity leads individuals to change their demand for housing services. There are several reasons why different individuals respond differently to the probability of job loss. Consider youths living with their parents. Young adults living with parents may react to the presence of the risk of job loss by delaying household formation while accumulating funds until the source of risk disappears. Alternatively, they may choose to rent their accommodation, as that form of housing consumption involves lower adjustment costs in the case of an income drop. On the contrary, home ownership involves large adjustment costs in case of an income downturn; selling a house is a costlier process than leaving a rented accommodation (see Fogli, 2004, Fernandes et al, 2008). Later in the life-cycle, mature-age individuals who rent their accommodation may react to risk of job loss by delaying house purchase while accumulating funds toward a downpayment -again, the idea is that selling a house is costly (see Chetty, 2012 or Shore and Sinai, 2010). Finally, we model the decision of being neither an owner nor a renter (for example, because an individual lives with parents), as well as buying and renting.

For various reasons, the simple comparison of housing choices made by workers with different labor contracts (fixed-term vs open-ended) is a misleading indicator of the choices outlined above. For example, workers who are observed with a fixed-term contract are more likely to have been unemployed in the past, and thus have depleted any accumulated wealth during prior unemployment spells. Hence, they have less resources for a downpayment and be more likely to rent or to live with parents (if young enough). In that case, the different housing choices across workers with different contract types mainly pick up different labor market histories. Alternatively, workers may be located in different local labor markets. Employees in tighter local labor markets may be more likely to be hired on an open-ended contract -due to firm's competition for scarce labor- and better income prospects may lead workers to become home owners. In that case, different housing choices across workers with different labor contracts would merely reflect differences in local labor markets, rather than exposure to risk. We comment on those biases later on.

Our study exploits variation in contract type that is weakly correlated with previous labor market histories or the local labor market. In particular, we exploit variation in firing costs due to the existence of regional subsidies for the conversion of temporary contracts into permanent ones as documented in Section 1. In that case, we compare workers hired in the same year, or same region, but whose firms faced different incentives to hire workers with a high severance payment contract. That variation is likely to be unrelated to local labor markets or previous histories (in Barceló and Villanueva, 2016, we use a similar strategy to analyze the households' precautionary wealth against the risk of job loss). Basically, we assume that the evolution over time of those subsidies is uncorrelated with decisions of household formation for channels other than the conversion of a temporary contract into a permanent one.[11]

The long-run effects of job security on housing tenure and household formation are analyzed using two different estimation approaches. The first uses a stock sample of individuals to study these decisions. That sample allows us to examine how our key identifying variable of exogenous risk of job loss (regional subsidies for contract conversions) affected the evolution of the stock of workers with a high severance payment contract (the first-stage), and the fraction of coresidents, home owners and renters of their main residence (the reduced-form or intention-to-treat). The second approach estimates a duration model using a sample of multiple transitions to a new accommodation, distinguishing between owning and renting. In this way, we study the number of years elapsed since the job contract was signed until the individual forms a new household with one of the two housing tenure regimes considered. This specification allows us to examine the timing of decisions and the dynamic impacts of job security on housing choices.

[11] As mentioned above, the introduction of those subsidies coincided with a major, national-wide reform that diminished firing costs for workers who were employed under a permanent contract. As done in Barceló and Villanueva (2016), the reduction in firing costs is converted into a subsidy-Euro equivalent and added to the regional subsidies in order to take into account this permanent policy change — once again, the results are very similar with and without the addition.

2.1 Estimates from a stock sample of individuals

To investigate if job security affects the housing tenure regime that young workers choose to live when they form a household independent of their parents, we show causal evidence from the 2002-2011 waves of the triennial Spanish Survey of Household Finances (in Spanish, *Encuesta Financiera de las Familias*, EFF). We follow a similar estimation strategy to that carried out in Barceló and Villanueva (2016) to estimate the household wealth response to the risk of job loss. To study the impact of job security on the housing tenure regime, we use the sample as a series of cross-sections, where the dependent variable is the living arrangement of the youth (staying with parents, owning or renting his or her accommodation), and the independent variable, $Subsidy_{r,a,t_0}$, is the incentive that the employer of the youth has to upgrade a contract into a high firing cost one. As the source of variation depends on the year of hire, region, gender and age when the worker was hired, we saturate the model with fixed effects of all those characteristics.

$$Y_{i,t} = \theta_0 + \theta_1 Subsidy_{r,a,t_0} + \theta_2 X_{i,t} + \mu_r + \mu_a + \mu_{t_0} + \mu_t + \varepsilon_{r,a,t_0,i,t} \tag{1}$$

That is, we regress the dependent variable of interest ($Y_{i,t}$) on region indicators (μ_r, omitted region: Madrid), age-at-hire indicators (μ_a, the omitted group is 31-40 years of age) and year-of-hire indicators (μ_{t_0}, the omitted year is 1999). The model also includes calendar year dummies (μ_t, the omitted year is 2002) and some explanatory variables included in $X_{i,t}$, such as indicators of individual's gender and education level and a third-order polynomial on the logarithm of the total labor earnings received in the previous year. The subindex i refers to individuals in the sample and $\varepsilon_{r,a,t_0,i,t}$ denotes the error term in the equation, distributed with a zero mean. The specification permits examining how the incentive to convert a fixed-term contract into an open-ended one in the first two years of the individuals' job tenure affects their household formation decision, i.e. our parameter of interest is θ_1.

The key variable identifying the risk of job loss is $Subsidy_{r,a,t_0}$, which measures the economic incentive a firm in a given region r and in a given year t_0 faces to upgrade a fixed-term contract into a permanent one for an individual with age a. We do not observe if the firm for which the young adult works actually got the subsidy, we only use the amount of the subsidy the firm was eligible for, presented in Table A.1. For workers covered by an open-ended contract, we do not know when the contract was converted. However, previous studies have documented that most conversions happen during the first two years of the labor contract.[12] Hence, we assign the mean subsidy during the first two years of the match between firm and the employee. As shown in Section 4, estimates from

[12] See Güell and Petrongolo (2007) and Izquierdo and Jimeno (2015), among others.

a stock sample of employees using administrative data from Social Security records also give evidence for the fact that most of conversions happen in the first years of job tenures.

Our parameter of interest in this specification, θ_1, measures the causal impact of a decrease in the probability of losing the job on the dependent variable of interest, which is contract type (open-ended or fixed-term) in the first stage and three indicators of housing services in the intention-to-treat specification, where we analyze separately three outcomes: neither owning nor renting (for example, living with parents), owning and renting. The main reason why we make use of the stock sample of individuals is to assess the validity of our instrument, regional subsidies for contract conversions, to obtain causal estimates of the impact of job security on the decisions of household formation and housing demand.

Finally, this paper uses a longer time span to analyze the response of household formation to changes in job security, not only in the adjacent years to the conversion, but also in the medium or long run. We consider periods up to seven years of job tenure to study household formation or housing tenure in order to take into account that this decision may be delayed several years after contract conversion.

2.2 Estimates from a sample of multiple transitions to a new house as an emancipated person

In order to investigate further the long-run effects of job security on the decision of housing tenure and on household formation, we study the relationship between the exact year in which individuals move to a new house living as a emancipated person (by owning or renting) and the number of years elapsed since the individuals started to work in the current firm. We use retrospective information of the year in which individuals started to work at the current job, the year in which they acquired the owner-occupied house and the year in which they started to live in the rented accommodation. We estimate a discrete choice duration model with two alternative exits to a new house as an emancipated person, we consider a move to an owner-occupied house ($D_i = 1$) and a move to a rented house ($D_i = 2$). The specification of the duration model is similar to those implemented by Bover and Gómez (2004) and Barceló (2006a). The transition intensity to emancipation with alternative k (home ownership or rental) is defined as the probability of moving to a new house with that alternative at year t after having started to work at current job given that the individual has been working for at least t years. This transition intensity follows a multinomial logit specification:

$$\theta_k [t \mid Z_i] = \Pr (T_i = t, D_i = k \mid T_i \geq t, Z_i) = \frac{\exp (Z_i' \beta_k)}{1 + \sum_{j=1}^{2} \exp \left(Z_i' \beta_j \right)}, \quad k = 1, 2 \qquad (2)$$

The hazard rate, $\theta [t \mid Z_i]$ is defined as the probability of leaving home at t years after having started to work at the current job given that the individual i has a job tenure of at least t years, and it is equal to the sum of both transition intensities, as follows:

$$\theta [t \mid Z_i] = \Pr (T_i = t \mid T_i \geq t, Z_i) = \sum_{j=1}^{2} \theta_j [t \mid Z_i] \qquad (3)$$

The individual characteristics (Z_i) considered in the duration model are the following: regional subsidies $(Subsidy_{r,a,t_0})$, indicators of region (ξ_r), age at hire (ξ_a), year at hire (ξ_{t_0}) and other time-invariant characteristics, such as the indicator of the individual's gender and education level. As calendar year dummies are perfectly collinear with dummies of year at hire and dummies of the yearly duration of the job spell, we control for the business cycle by using the unemployment rate at year $(t_0 + t)$ in Spain as another covariate of the model.

In this duration model, censored observations correspond with two kinds of individuals: first, individuals living with their parents and, second, individuals living in the same house as the one in which they live when they started to work in the current job (i.e. individuals that had previously rented or bought their dwelling before starting to work in the current job). Our parameter of interest in β_k from Equation (2) is the one associated with the regional subsidies, $Subsidy_{r,a,t_0}$, which measures the causal impact of an exogenous increase in regional subsidies for job-contract conversions (an exogenous increase in job security) on the probability of moving to a new accommodation through each alternative housing tenure regime k at t years of the job spell, given that the individual has a job tenure of at least t years or more.

As Jenkins (1995) emphasizes, when the transition intensities follow a multinomial logit specification, we can estimate a competing-risk model for each exit separately, and then the conditional exit rates follow a logit binary specification with the same parameters, β_k. The conditional exit rate using alternative k gives the probability of moving to a new house with that housing tenure regime k at year t of the job spell given that the individual has been working at least t years and does not move to a new house with the other housing tenure regime during their current job spell. For estimating competing-risk models, exits using the alternative housing tenure regime are also treated as censored observations when we concatenate the survival subsamples on each duration for estimating the parameters of the conditional exit rate of interest using a logit model.

3 Sample data used in the empirical analysis

We use the 2002-2011 waves of the Spanish Survey of Household Finances (in Spanish, *Encuesta Financiera de las Familias*, EFF). The EFF is a triennial survey conducted by the Banco de España, which interviews around 6,000 households and obtains detailed information about their wealth holdings, debt and consumption, as well as individual information about personal characteristics, earnings, labor status and other labor market characteristics. This survey allows us to examine the specific route of household formation, as it contains information about whether youths live with their parents, own their own accommodation or rent it.

In order to allow different studies of the household savings and wealth portfolio, the EFF has an oversampling of the rich population. All the summary statistics shown in the paper are weighted to be representative of the Spanish population. The estimates of the empirical models are unweighted as they analyze the individuals' behavior, how they make their decision of household formation and housing demand when they face uncertainty about their risk of job loss. Finally, as a typical way of dealing with item non-response in wealth surveys, the EFF provides five different values imputed stochastically for each missing observation in order to take into account the uncertainty about the imputed data [for more details about the EFF imputation, see Barceló (2006b)]. All calculations reported in this paper make use of the five data sets imputed multiply by combining estimates using Rubin's rules [see Rubin (1987)].

3.1 Stock sample of individuals in the analysis of their decision of household formation and housing tenure

Using the EFF survey we construct a stock sample of household members between 25 and 45 years, who are employees with a job tenure of seven years or less and who earned at least 2,500 euros in 2005 constant terms in the year prior to the one of the survey interview. In order to obtain an homogeneous sample of individuals whose attachment to the labor market is strong, we exclude from the sample individuals that declare an economic inactivity (other than education), individuals who are self-employed or who do not contribute to Social Security. We also drop out of the sample individuals whose job tenure is longer than one year at the same time that they earned on average less than the minimum wage in the previous year (taking into account their job working time -whether part-time or full-time). We also remove from the sample those individuals that live in a household where neither the reference person nor the spouse are their relatives. Finally, we exclude from the sample workers that started to work in 2010 or later, as we do not know the amount of subsidies they were eligible for in case of job-contract conversions.

Job security is measured by the kind of job contract the young workers hold, whether a permanent contract or a fixed-term contract. As the kind of job contract is an endogenous variable, we obtain causal estimates by instrumenting the stock of permanent workers with the mean regional subsidies that firms were eligible for the conversion of fixed-term contracts into open-ended ones (permanent contracts) in period 1997-2009. The variables of regional subsidies are expressed in thousand euros of 2005 using regional deflators of the gross household disposable income.[13] Labor income earned in the previous period is converted in thousand euros of 2005 by using the Consumer Price Index, provided by the Spanish National Statistics Institute (INE). Finally, our measure of regional subsidies refers to the mean subsidy the firms can benefit from the conversion of a fixed-term contract into a permanent one in the first two years of the individuals' job tenure, since almost 75% of the contract conversions occurred in the first two years [see Izquierdo and Jimeno (2015)].

In the empirical analysis, we consider that an individual is a homeowner (a renter) when he or she is either the reference person in the survey or the spouse and the household owns (rents) their main residence (and they do not live with their parents or parents-in-law). Finally, an individual is a coresident, when he or she lives with their parents (parents-in-law) or other relatives (and in the latter case, the other relative is either the reference person or the spouse).

Table 1 presents the descriptive statistics of the main characteristics of the estimation sample. The 33% of the individuals aged between 25 and 64 who are employees live with their parents, 51% are homeowners and the remaining 16% live emancipated in a rented house. In the sample, the 64% of the individuals hold an open-ended contract and 36% a temporary contract. Employees holding a fixed-term contract are less than half-a-year younger than workers under an open-ended contract, and they have been hired on average two years prior to the date of the interview, while workers under an open-ended contract have mean job tenures of four years and a half. Eligibility, as mentioned before, mainly depends on age at hire, year of hire, region and gender. The 56% of employees under a fixed-term contract and under an open-ended one were eligible to benefit from regional subsidies during the first two years of their job tenure. The population of workers under fixed-term contracts overrepresents females (47% vs. 42% in the subsample of employees with an open-ended contract) and employees with lower wages (almost 12 thousand euro

[13] These regional deflators come from the database BDMORES, elaborated by the Spanish Ministry of Finance.

vs 16 thousand in the subsample of employees with an open-ended contract). Moreover, fixed-term workers have achieved lower education levels. Finally, 36% of employees with a fixed-term contract live with parents, while the share among workers with an open-ended contract is 31%. Similarly, employees under fixed-term contracts are more likely to live in a rented house than employees under open-ended contracts (21% rent versus 14%).

In order to investigate if the pattern of household formation according to the risk of job loss differs highly from the choice of housing demand of older workers, we also estimate the same empirical models with a sample of youths, individuals aged between 25 and 45. The summary statistics of this sample are documented in Table 2. In this sample of young individuals, the percentage of workers holding a permanent contract is very similar to that of the full sample, 64%. Younger individuals have also higher levels of attained education, and their rate of home ownership is lower (47% vs 51% in Table 1). Among the sample of individuals aged between 25 and 64 that live with their parents or other relatives, only 12.7% correspond to individuals aged over 45. The patterns of individual characteristics across subsamples of workers broken down by their kind of job contract are very similar to those in the sample of older individuals in Table 1.

3.2 Analysis of the timing of household formation and housing tenure choice using transition data

When estimating the duration model, the sample is formed by those transitions to a new accommodation since the individual has started to work at current job, i.e. the number of years elapsed from his/her current job spell until the individual moves to a new accommodation with one of the two housing tenure regimes considered (home ownership or rental). The individuals that take part into the sample are those household members aged between 25 and 64 in the year of the interview, who have a job tenure of ten years at most, and who live with their parents (or other relatives), have acquired an owner-occupied house or have rented an accommodation before or after having started to work in the current job. Thus, we do not drop out of the sample those observations of individuals that formed a household independent from their parents before starting to work in the current job. These observations are treated as censored in the transition data analysis, since we do not observe any changes of residence for these individuals after having started to work at their current job. This sample of transition data allows us to analyze the exact year after starting to work at current job in which the household formation is produced and the time path in which the housing tenure regime is chosen.

Table 3 shows the summary statistics of the transition data sample of employees aged between 25 and 64, who have worked in the current job for ten years or less. The 20%

of individuals moved to a new accommodation as an emancipated person during their current job spell, 13.6% of the movements being to an owner-occupied house (67% of the changes of main residence occur among individuals that become homeowners). In the total sample, 68% of transitions correspond to individuals holding a permanent contract, and more than 80% of the transitions happen in the first six years of the individuals' job tenure. Individuals who move mainly to an owner-occupied house are more likely to hold a permanent contract (85%) than individuals moving to a rented house are (65.3%). Most of the exits to a new accommodation as an emancipated person occur in the first five years of job tenure (almost 90% of purchases of the owner-occupied houses and 85% of the exits to a rented house).

4 Empirical Results of the Medium-Run Analysis of Household Formation and Housing Demand

This section presents the empirical results obtained using data from the 2002-2011 waves of the Spanish Survey of Household Finances. Subsection 4.1 describes the causal evidence of the impact of the risk of job loss on the household formation decision and the housing tenure choice drawn from the stock sample of employees at different age intervals (25-64 and 25-45). We use a sample of younger employees, aged between 25 and 45, in order to investigate if the pattern of household formation and housing tenure choice differs greatly from the decision of housing demand of the population of older employees, concerning the risk of job loss. Subsection 4.2 documents the empirical results obtained from the estimation of a duration model using retrospective information of event years. Finally, the standard errors of the estimated parameters shown in parentheses in all Tables take into account that there can be group correlation in the error term within each region [see Moulton (1986)].

4.1 Causal evidence from a stock sample of individuals

First-stage estimates and intention-to-treat estimates Table 4 shows reduced-form estimates of linear probability models of the stock of permanent workers and the decision of the housing tenure regime in a sample of employees at different age intervals (25-45 and 25-64). We estimate linear probability models instead of nonlinear discrete choice models, such as probit or logit models, because the usual tests that measure the quality of the instruments used are based on linear regression models in a setting of instrumental variable estimation. However, we obtain the same results when we estimate probit and logit models of these outcome variables.

We consider three different housing tenure regimes: home ownership, rental and coresidence (living in the parental home). We show reduced-form estimates of the impact of regional subsidies on the decision of housing tenure regime before providing Instrumental Variable (IV) estimates, because the intention-to-treat estimates are much more precise.[14] Panel A shows the first-stage estimates of the effect of regional subsidies for the conversion of fixed-term contracts into open-ended ones on the probability of observing individuals with permanent contracts. Panel B shows the reduced-form estimates of the effect of regional subsidies on the different decisions of housing tenure regime.

Panel A of Table 4 shows the Ordinary Least Square (OLS) estimates of the indicator of whether the household member holds a permanent contract on the mean subsidy in the first two years of the worker's job tenure that the firm can benefit from its region for converting a fixed-term contract into an open-ended one. The estimates are 0.006 in the sample of young individuals (column (1)) and 0.008 in the full sample of workers (column (2)), and they are statistically significant at the 5% and 1% level of significance, respectively. Note that monetary variables, such as regional subsidies, enter the model in thousand euros of 2005 in constant terms. Thus, the estimates imply that an increase of 1,000€ in the subsidy rises the stock of permanent workers by 0.6%-0.8% depending on the age of the worker (what suggests an increase of 1%-1.3% in the prevalence of permanent workers in the sample according to summary statistics in Tables 1 and 2). The F-statistic of the significance of the instruments in the full sample is over 10. However, the estimates are less precise and the corresponding F-statistic is much lower in the sample of young individuals, perhaps mainly due to the small sample size. The first-stage estimates in both samples of individuals at different ages are very similar.

In order to assess the validity and weakness of our instrument in the sample of employees aged between 25 and 45, we reproduce our first-stage in another data coming from the 2004-2015 waves of the Continuous Sample of Working Histories (in Spanish, *Muestra Continua de Vidas Laborales*, MCVL). These data consist of a random sample of 4% of the administrative Social Security records, which are representative of total population. The MCVL data contains information on pension earners, recipients of unemployment benefits, and information about the jobs of employees and self-employed workers. The MCVL also collects longitudinal information, past labor histories of all individuals included in the sample. Table A.2. shows the estimates of the first-stage in a sample of employees aged between 25 and 45, constructed in a similar way to that drawn from the EFF, but using administrative data from the MCVL. The first four columns of Table A.2

[14] In our sample of household members aged between 25-45 whose job tenure is not longer than 7, the minimum sample size is 3,974.

show estimates of the impact of subsidies on the stock of permanent workers aged between 25-45 who were hired since 2001 and observed in period 2002-2012. The estimates are obtained in different subsamples of individuals according to the length of their job tenure (lower than one year, two, three and four years of job tenure), in order to obtain an indication of the year in which the job contract conversions usually happens. We can see that the coefficient estimates associated with subsidies are very stable across subsamples of different job tenures, the estimates are around 0.0060 and significant at the 1% level of significance. Almost of all the F-test statistics associated with the significance of the instrument have a value over 10. These estimates are identical to those obtained using data from the EFF for employees aged 25-45 (column (1) in Table 4). Column (5) of Table A.2. shows the estimates equivalent to the first-stage estimates obtained in the EFF data, where the sample is formed by individuals with a job tenure of seven years or less in period 2002-2012. The estimated impact of subsidies is 0.0059, very similar to the estimate of 0.0062 encountered in the EFF. The stability of the coefficient estimates associated with subsidies across different lengths of job tenure indicates that contract conversions usually happen in the first years of job tenure.

Panel B of Table 4 shows the reduced-form estimates of the probability of the decision of coresidence (living with parents) and the housing tenure regime in which emancipated individuals are observed to live. First of all, the decision of living with parents seems to be negatively affected by regional subsidies. An increase of 1,000€ in regional subsidies decreases significantly the probability of coresidence from 1% to 1.3% depending on the age of workers, the impact seems to be higher among younger workers. These estimates suggest a decrease from 3% ($\frac{0.01}{0.332} \cdot 100$) to 3.5% ($\frac{0.013}{0.371} \cdot 100$) in the rate of coresidence due to an exogenous increase of job security given by the rise of 1,000€ in the regional subsidies for contract conversion. Individuals are equally likely to form a new household by the purchase of their owner-occupied house (the rate (%) of home ownership increases in 0.58% for young workers and in 0.44% for older workers) or by renting a house (the percentage of young people renting their dwelling increases in 0.7% and in 0.6% for older workers), when subsidies increase by 1,000€. These figures represent a rate at which the probability of becoming a homeowner increases of about 0.8%-1.3% ($\frac{0.004}{0.507} \cdot 100$ and $\frac{0.006}{0.466} \cdot 100$), being the increase higher for young workers, according to the summary statistics shown in the first column of Tables 1 and 2. The rate at which the probability of renting increases is 4.3% ($\frac{0.007}{0.162} \cdot 100$) for young workers aged between 25 and 45 and 3.7% ($\frac{0.006}{0.161} \cdot 100$) among older workers aged up to 64. These figures are more precisely estimated for the decision of rental than for home ownership. Thus, we encounter significant effects of exogenous increases of job security on the decision of forming a new household and on the choice of housing tenure among individuals with long job tenures, up to seven years of tenure.

The conversion of job contracts is highly expected to happen in the first two years of job tenure. Therefore, these intention-to-treat estimates may capture the fact that young individuals postpone the decision of forming a new household several years after that the risk of job loss disappears.

Instrumental variable estimates Furthermore, Table 5 shows estimates of the impact of holding a permanent job on the probability of each housing tenure regime for employees at different ages, between 25 and 45 in columns (1) and (3) and between 25 and 64 years of age in columns (2) and (4). The first two columns show suggestive evidence of this impact from Ordinary Least Square estimates (OLS). Workers holding a permanent contract has a probability of being observed living in a rented accommodation 0.7% and 0.5% lower than that of temporary workers in the two samples of young workers and total workers, respectively. On the contrary, permanent workers are 0.4% more likely to be observed living in an owner-occupied house than temporary workers.

When we compare these estimates with those obtained by using instrumental variable estimates (two-step least square estimates, TSLS), the empirical results are very different. Mainly in column (4), the estimates of the causal impact of a permanent contract on the decision of housing tenure regime indicate that an exogenous increase in job security that leads to a conversion of job contract into an open-ended one makes permanent workers more likely to form a household by owning their own accommodation and by renting it, instead of living in their parental home or with other relatives. An increase in the probability of holding an open-ended contract of 1% increases the probability of forming a household by home ownership or by renting in 0.54% and 0.70%, respectively. On the contrary, the probability of living with parents decreases in 1.25%. That is, workers holding permanent contracts are more prone to live in an emancipated household not only by owning, but also by renting, unlike the OLS estimates indicate. This result gives evidence that the OLS estimates are biased, since temporary workers have less chances of forming a household by home ownership. This is due to the fact that temporary workers have been able to accumulate less assets for buying a house than permanent workers have, as fixed-term workers have had to use their buffer stock to sustain consumption in past unemployment spells more often than permanent workers have due to their higher risk of losing the job.

The TSLS estimates for the sample of younger workers in column (3) of Table 5 suggest higher impacts of job security on the decision of household formation and housing tenure choice, but the estimates are much less precise, mainly in rental. Table 5 also shows the empirical confidence intervals for the TSLS estimates in finite samples, proposed by Chernozhukov and Hansen (2008), which are robust to the presence of weak instruments, instead of looking at only TSLS confidence intervals defined properly for large samples. These robust confidence intervals indicate that the TSLS estimates in the full sample of workers are significant at the same confidence level, except for the estimated effect of holding a permanent contract on the probability of home ownership, as the value of zero is inside the confidence interval, although very near to the lower bound of the interval. Concerning the TSLS estimates in the sample of young workers aged 25-45, the estimates are very imprecise, but they are statistically significantly different from zero at the 10% level of significance, since the value of zero is outside the interval in the three housing tenure regimes considered. Thus, exogenous increases in job security have a significant impact on the decision of forming a household and on the housing tenure regime chosen.

When we compare our results with those obtained in the empirical literature, our estimated impacts are much higher. Becker et al. (2010) document that an increase of 10% in the probability of job security has a simultaneous effect on the probability of forming a new household of 3.4% in that year the risk of job loss is reduced. In our case, taking into account that the difference in the subjective probability of job loss is 40% between median permanent and temporary workers (see Figure 2), the estimate of an increase of 0.5% in the probability of forming a household with one housing tenure implies that an increase of 10% in job security rises the probability of leaving parental home in 15.6% ($\frac{0.005}{0.008 \cdot 0.4} \cdot 10\%$). This discrepancy of results may be due to the fact that our estimates measure the accumulated effect of job security on household formation during three years, not only the instantaneous effect, since the reference person in our estimates is an individual interviewed in 2002 that was hired in 1999. In order to reconcile both findings, we study the timing and the dynamic of the events using a duration model to investigate whether there is a delay in undertaking the decision of household formation and housing tenure choice. García-Ferreira and Villanueva (2007) do not encounter a simultaneous effect of a sharp increase in job security on the decision of household formation in Spain using data from the Labor Force Survey.

Robustness checks Finally, in all models estimated in the paper, we have made some robustness checks in order to assess how the estimates are robust to the inclusion of second-order fixed effects among the dimensions in which regional subsidies vary, i.e. indicators of region (μ_r), bands of age at which the individual was hired (μ_a) and year of hire (μ_{t_0}) in Equation (1). In this way, we include sequentially in the estimates the following second-order fixed effects: age at hire and year at hire ($\mu_a \times \mu_{t_0}$), region and age at hire ($\mu_r \times \mu_a$) and region and year at hire ($\mu_r \times \mu_{t_0}$). The estimates of our outcome variable, the regional subsidies that firms are eligible for the conversion of fixed-term contracts into permanent ones, are very stable across different models of fixed-effects, and most of the fixed effects are not statistically significant.

We have also carried out a falsification exercise of our measure of the exogenous variation of job security, the regional subsidies for the upgrades of fixed-term contracts into permanent ones, by analyzing placebo subsidies using randomization inference as done by Bertrand, Duflo and Mullainathan (2002). We generate random laws by reassigning the regional subsidies of our sample completely randomly or randomly by region, age at hire, gender and year of hire (we also assign the subsidies randomly by pairs of these covariates), and we reestimate our empirical models with these placebo subsidies as our key identifying variable. We generate 200 independent random draws of subsidies, in order to construct properly empirical confidence intervals that tests the hypothesis of a zero impact of subsidies on the outcome variables. In all regressions, the estimated coefficients associated with the placebo subsidies are near zero, and our estimated effects of regional subsidies in the first-stage and intention-to-treat equations shown in Table 4 almost always lie outside the empirical confidence intervals at the 1% or 5% levels of statistical significance computed by the randomization inference approach, which test the null hypothesis of a zero value of the coefficient associated with the placebo subsidies. This means that our instrumental variable, the regional subsidies, are not capturing other unobserved effects not taken into account in the analysis and potentially correlated with regions, age at hire, gender and year of hire.

4.2 Causal evidence from a sample of transitions

We further study the exact year of the job spell in which individuals decide to move to a new accommodation as an emancipated person by estimating transition data models by considering two alternative exits, a move to an owner-occupied house or a move to a rented house. Table 6 shows the estimates of how the probability that an individual moves to a new house, as an emancipated person, varies with the number of years elapsed since the individual started to work at the current job. The specification (i) considers a time-invariant effect of regional subsidies on the probabilities of moving and specification (ii) allows for a different effect of subsidies that varies along the time elapsed since the job contract was signed. Columns (1) and (4) consider a move to a new house (irrespective of the housing tenure regime), and columns (2) and (3) (and columns (5) and (6) in specification (ii)) distinguish the move by the housing tenure regime, whether to an owner-occupied house or to a rented house, respectively. We can see that regional subsidies are significant to explain transitions to a new house at the 1% level of significance, also considering both types of housing tenure, unlike the previous estimates using the stock sample of individuals.

When we allow for time-varying effects of subsidies in columns from (4) to (6) of Table 6, note that the interaction of the subsidy with the first two years of job tenure is negative in columns (4) and (6). The interaction of the subsidy with the indicator of job tenure above four years is also negative in column (4) and (6). That pattern suggests that the demand for new accommodation (possibly household formation) peaks three years after the beginning of the job, about one year later than the period where most conversions are expected to happen. That is, there is a delayed effect of increased job security on household formation. Interestingly, the pattern is most noticeable among renters. Home ownership seems to present a flat pattern according to job tenure.

Figure 3 shows the marginal effects of an increase in 1,000€ regional subsidies to contract conversion on the probability of each housing tenure regime (rental vs. home ownership) implied by the estimates of specification (ii) in Table 6, when we allow the subsidy effect to be time-varying across the job spell durations. Marginal effects are expressed in terms of the hazard rates, the probability of moving to each tenure regime in each year of the job spell. The vertical line indicates the second year of the job spell, when the contract conversion is expected to have occurred.[15] The full black line shows the marginal effects on the probability of renting a new accommodation. We can see that the decision of renting a house seems to be delayed between one and two years after the contract conversion is expected to happen. That is, the pattern of the increase in the likelihood of renting seems to be an inverse U-shaped across job tenure. However, the pattern of the marginal effects on the probability of home ownership (dashed and dotted line) is flat across the duration of the job spell, without a clear shape. The delay in the response of renting is somewhat surprising as renting -the housing tenure regime that entails smaller adjustment costs- is the most delayed in time after an increase in job security.

Figure 4 shows the marginal effects implied by the estimates of the duration data model in the sample of employees between 25 and 45. Here we can also appreciate an inverse U-shaped pattern of exits to a rented house, but the differences in the pattern with respect to exits to home ownership are much more apparent in this sample. Firstly, within the sample of 25-45, the probability of renting is higher than in the full sample. A 1,000 euro subsidy to contract conversion during the first two years of the contract increases the probability of renting by more than .002 in the 25-45 sample, while it does not reach .002 in the full sample. This result suggests that renting is much more responsive to increases

[15] Almost the 75% of the contract conversions from fixed-term contracts into open-ended ones are estimated to happen in the first two years of the workers' job tenure in 2013 in Spain using firms' information from the third wave of the Wage Dynamic Network Survey conducted by the European Central Bank (ECB) and the National Central Banks of the European Union Member States [see Izquierdo and Jimeno (2015)].

in job security in the youth sample than in the overall one. Secondly, in both samples, the response of renting to increases in job security is delayed in both the 25-45 sample and in the 25-64 sample. Thirdly, the impact of 1,000 euro subsidy to contract conversion on home ownership is rather similar across both samples.

One interpretation of those results is that renting is a much more flexible way of forming a new household when security on the job increases (i.e., when the alternative is to live with parents), while in the full sample, containing also mature workers, renting is equally attractive as owning the main residence. The fact that we encounter a delay in undertaking the decision of household formation, mainly by rental, may help to reconcile our estimated impacts of job security on the decision of household formation in the stock sample of individuals with those encountered in the literature, like Becker et al. (2010).

Robustness checks Similar robustness checks to those carried out for the estimates of the impact of job security in the stock sample of workers have been implemented in the duration model. The estimates are stable across different specifications of the duration model, which include second-order fixed effects of those factors on which subsidies vary. Moreover, the coefficient estimate associated with regional subsidies are generally outside the empirical confidence intervals that test a zero impact of subsidies at the 1% and 5% significance level, constructed using 200 realizations of placebo subsidies drawn independently by randomization inference.

4.3 Possible explanations of the mechanisms behind the empirical results

Heterogeneity in responses among youths This subsection examines the characteristics of youths who form a new household, to get some insights on possible reasons for the patterns in household formation detected above. In Table 7, we use the panel component of the EFF2002-2014 to examine the characteristics of youths who left the household between one wave and the next. Namely, we regress an indicator variable of whether a youth left the household on the instrument (the mean subsidy available for the first two years of job tenure) on various subsamples. A broad view of Table 7 indicates that the youths whose household formation decisions were most affected by the instruments were between 31 and 45 years of age (column 3 of Table 7), had low schooling levels (Column 4), come from households with little wealth (Column 6) and tend to have labor earnings below the median (Column 8). Many of those characteristics indicate that those youths are unlikely to have accumulated much wealth towards a down-payment and offer an explanation for the patterns detected in Figure 4: upon forming a new household, they are most likely to respond by renting a new accommodation.

Table 8 provides further information about the wealth levels of affected youths. While we cannot observe the wealth of the individual who leaves the household, the panel component allows us to observe the evolution of the household wealth once the individual has left the household. Any systematic change in that wealth level likely captures the wealth levels of the youth who formed his or her own household. Using an array of wealth measures, we find that, while subsidies to contract conversion successfully predict household formation (the youth leaving the parental household), they have little or no predictive power in the evolution of household wealth. Our interpretation is that youths who see an increase in job security through these subsidies hold modest wealth levels, and are thus unlikely to move to an owner-occupied dwelling.

Presence of credit constraints In Section 4.2, we encounter that the most delayed way of forming a household is rental, surprisingly the least hard-to-reverse housing tenure regime in case of materializing the risk of job loss. One possible explanation of this delay in rental is that individuals that make this decision several years after their job contract conversion are discouraged homeowners, i.e. individuals that wished to form a household by buying their own accommodation, but they could not do it due to credit constraints. In order to test this hypothesis, we consider a definition of liquidity constraints similar to that used by Jappelli (1990) and Barceló and Villanueva (2016). An individual is credit constrained if he or she satisfies one of these three conditions: in the last two years (1) the individual did not ask for a loan due to the fear of being rejected; (2) she or he asked for a loan, but it was accepted with an initial capital lower than the one requested; and (3) the loan was fully rejected. Another two alternative situations considered on which the individuals are not credit constrained are the following: (1) when the individuals did not ask for a loan in the last two years because they did not need it and (2) when they asked for a loan that was fully accepted. To address the issue of liquidity constraints, we estimate a multinomial logit model in which the omitted category is that the individual did not ask for any loans in the last two years because they were not needed.

Panel A of Table 9 shows estimates of the multinomial logit model of credit constraints using the same set of covariates as in Table 4 and using the sample of individuals aged between 25 and 45. The estimates do not give evidence for the alleviation of borrowing constraints due to a sharp fall in the risk of job loss coming from an increase of the regional subsidies firms can benefit from the conversion of fixed-term contracts into permanent ones, since the estimated impact of 1,000€ subsidy on the probability of being credit constrained is near zero, 0.0005, when this probability is 6% and the estimates are not statistically significant. Panel B of Table 9 estimates a joint model of the presence of borrowing constraints and the choice of housing tenure regime. For this purpose, we construct a multinomial variable that is the result of interacting the three indicators of being credit constrained or not (did not ask for a loan, asked for a loan fully accepted or be credit constrained) with the indicators of the housing tenure choice (coresidence, home ownership and rental). Thus, this multinomial dependent variable can take nine values, and the omitted category is to live with parents and not to have asked for a loan for two years. Again, in our sample period of 2002-2011, we do not find evidence that renters are more likely to be credit constrained than homeowners are, since the estimated value of the coefficient associated with the category of renting an accommodation and being credit constrained is 0.080 and is not statistically significant. Moreover, the estimated impact of a subsidy of 1,000€ on the probability of being credit constrained in each housing tenure regime chosen is very low, near zero.

Therefore, we do not encounter evidence for the fact that the presence of borrowing constraints associated with the risk of job loss was a potential explanation of the delay of individuals in undertaking the decision of living emancipated from parents in a rented dwelling.

Decision of living with a partner in a household emancipated from parents Another potential explanation of the delay in undertaking the decision of household formation by renting is that, once the risk of job loss disappears, single individuals prefer to continue living with their parents in order to save for a down payment on a house, and then they emancipate from their parents by buying their owner-occupied house. This decision of living with parents is hard to be made if individuals wish to live with their partner; thus, they can decide to form a new household by renting a house if they have not accumulated enough savings for a down payment. The delay in rental may be associated with these individuals that wish to live with a partner. Unfortunately, we do not have information of the exact year in which individuals got married or started to cohabit with a partner. Therefore, we cannot estimate a transition joint model of the decision of housing tenure (household formation) and marriage or cohabitation. Instead, using the stock sample of individuals, we estimate a model of the joint decision of living with a partner and being emancipated from parents (i.e. by owning or renting the own accommodation).

Panel A of Table 10 shows the estimates in the sample of young individuals aged between 25 and 45 and panel B does for the full sample of individuals aged between 25 and 64. The model is estimated using linear probability models for each joint decision of living with a partner or not and housing tenure, and we use the same set of covariates as those used in Tables 4 and 5. Table 10 shows two specifications of the model for the decision of living with a partner. The first specification is shown in column (1), where we only consider the decision of living with a partner, irrespective of the choice of housing tenure (the omitted category is not living with a partner in an emancipated household from parents). The second specification of the model is shown in columns from (2) to (5) where the omitted category is coresidence or not living with a partner. Columns (2) to (5) shows linear probability model estimates of each alternative joint decision individuals can choose.

In Panel A of Table 10, column (1) shows that regional subsidies are significant to explain the decision of living with a partner in the sample of young individuals aged 25-45. The eligibility of a subsidy of 1000€ increases the probability of living with a partner in a household emancipated from their parents in 0.6%, and this estimate is significant at the 5% level of significance. Once we interact this decision with housing tenure (columns (2) to (5)), an exogenous decrease in the risk of job loss due to the eligibility of a subsidy of 1000€ leads to increase the probability of individuals that decide to live with a partner by owning in 0.21% and by renting in 0.45%. The estimates are only statistically significant at the 5% level in the joint decision of renting. On the contrary, alone individuals that wish to form a household are more likely to do it by owning their own accommodation than by renting, the estimated impact of regional subsidies on this probability is 0.37% and significant at the 1% level. The estimated impact on the probability of renting by an individual living alone is lower and not significant. The 93% of young individuals aged 25-45 that do not live with a partner correspond to single individuals, and 85% of them live with their parents and only 9% live in a rented accommodation.

Panel B of Table 10 shows the estimates for the full sample of workers. Column (1) indicates that regional subsidies are not significant to explain the decision of getting married or living with a partner in a sample of older workers, since the estimated impact on this probability, 0.0041, is not statistically significant. However, exogenous increases in job security due to the existence of a subsidy of 1000€ are significant to explain the decisions of stay in a household emancipated without living with a partner by owning and less by renting, with a significant impact on these probabilities of 0.35% and 0.27%, respectively. Once again, exogenous increases in job security make older workers more prone to live with a partner in a rented house at the 5% level of significance. The demand for new accommodation of individuals aged 25-64 living alone might also come mainly from separated, divorced and widowed individuals. In this full sample of workers, single individuals represent only 64% of the individuals living alone, 83% of them live with their parents and 8% live in an owner-occupied house. On the contrary, the 36% of the individuals that live alone in their own household are separated, divorced or widowed.

Therefore, these estimates give evidence that one reason why individuals decide to form a new household, emancipated from their parents, mainly among young individuals is their wish to live with a partner. This decision is estimated to be made more likely once the uncertainty about the risk of losing the job disappears, due to the existence of regional subsidies for the conversion of fixed-term contracts into permanent ones. This mechanism might be behind the delay in undertaking the decision of forming a household by renting.

5 Conclusions

This study exploits two institutional features of the Spanish labor market to address the question: Does the growing incidence of job insecurity account for their recent patterns of housing arrangements? The advantage of working with Spanish labor market data is the widespread use of low-firing cost contracts, that allows us to identify who is exposed to the risk of job loss. We can also exploit legal changes that influence the labor demand of firms for workers with different contract types to obtain arguably exogenous variation in job insecurity. Thus, we are able to estimate the link between obtaining a more secure job and the decision to establish a new household controlling for other confounding factors.

Our strategy to identify the causal link between job insecurity and household formation uses temporal and regional variation in subsidies that firms were eligible for the conversion of low-firing cost contracts into high firing costs ones between 1997 and 2009.

In the empirical analysis, we use the 2002-2014 waves of the Spanish Survey of Household Finances (in Spanish, *Encuesta Financiera de las Familias*, EFF). Our results suggest that the rate at which employees aged between 25 and 64 live with parents or other relatives decreases by 3% after an exogenous increase in 1,000€ in the regional subsidies for job contract conversions. Moreover, and relative to baseline, increased job security increases the rate at which youths people form a new household by owning their main residence or by renting their dwelling by 1% and 3%, respectively. Finally, we provide causal evidence indicating that there is a sluggish response of the decision to rent to increases in job security. Namely, the demand for new accommodation happens one or two years after the time in which most conversions have occurred, mainly for those individuals who move to a rented new accommodation as an emancipated person. However, the response of home ownership to increases in job security is slightly smaller among youths than among the full population.

Among young individuals aged 25-45, those who respond to the presence of regional subsidies for the conversion of job contracts into permanent ones by forming a new household or demanding for accommodation emancipated from their parents are individuals aged 31-45, having only attained a low level of education and coming from a household with a level of wealth below the median. We have also analyzed the change of wealth of parents' household when young individuals leave the parental home, and we conclude that these individuals may not have accumulated enough savings for a down payment to buy a house, so they are more likely to leave the nest by renting than owning. Finally, we have investigated whether the delay in forming a household by renting may be due to the presence of borrowing constraints among renters, and the estimates do not give evidence for this hypothesis. Instead, it seems that a significant factor for forming a new household is the wish to get married or to cohabit with a partner, by renting or by owning if they have accumulated enough savings for a down payment. The analysis of compliers indicates that these individuals have not accumulated enough wealth for buying a house. Thus, the choice of living with a partner might be behind the delay in undertaking the decision of forming a new household by renting.

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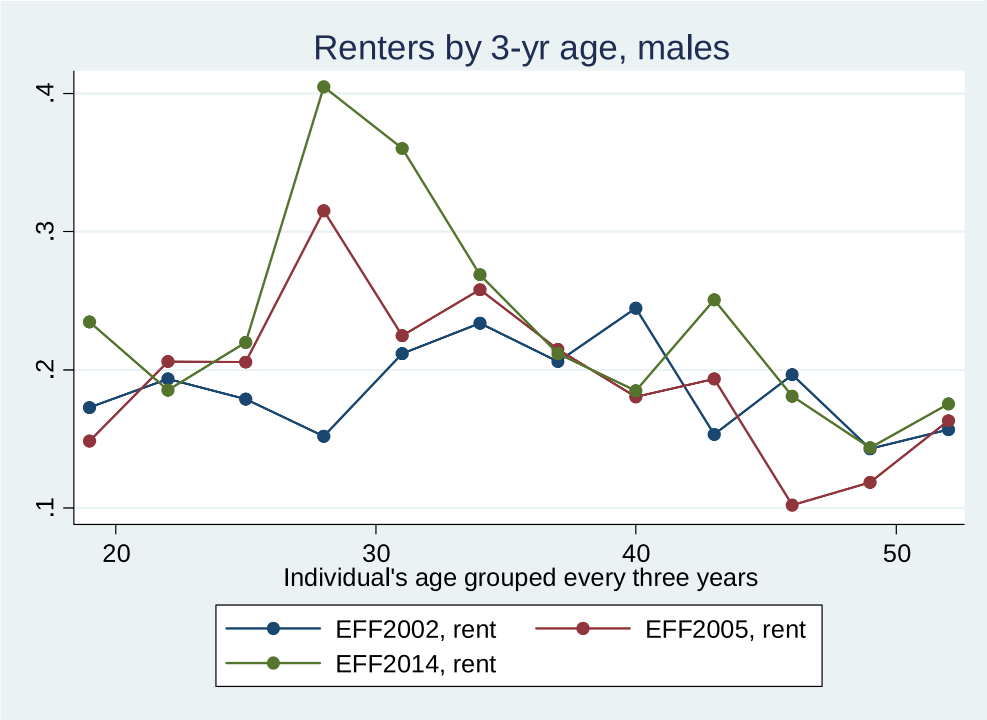
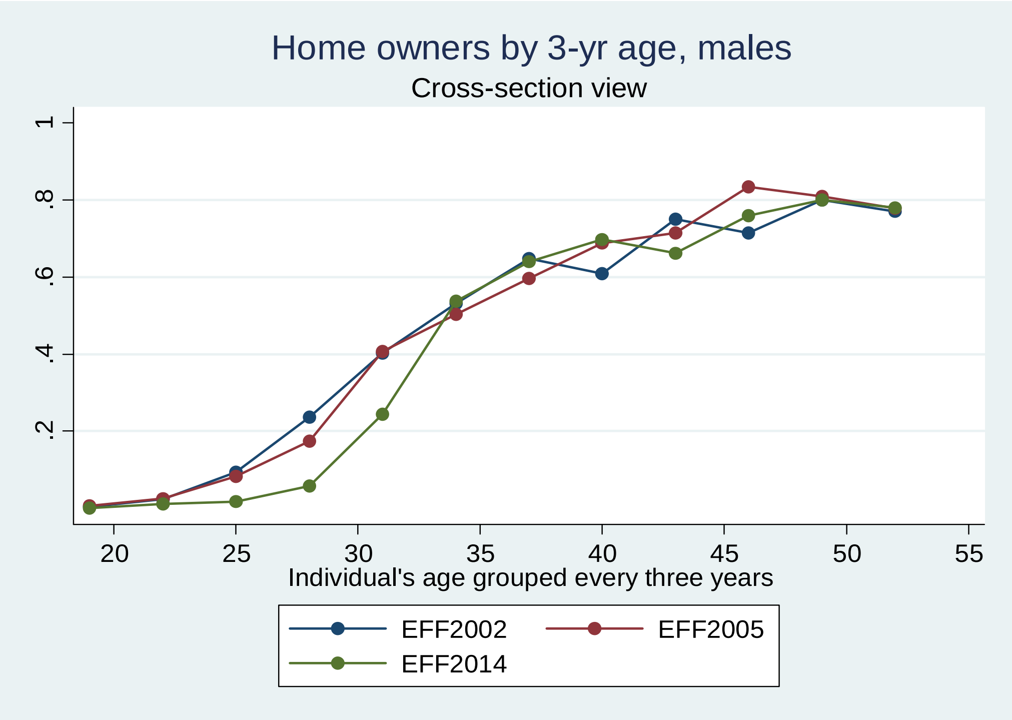
Figure 1: Evolution of living arrangements of individuals 20-55 in Spain between 2002 and 2014





The panel on the left contains the fraction of males who rent their house of residence as a fraction of all males at that age (including owners, renters and those living with another person in a house they neither own or rent). The panel on the right contains the fraction of males who own their house of residence.
Source: Spanish Survey of Household Finances, EFF2002-EFF2014.

Figure 2: Subjective probability of job loss over the next 12 months in 2011, by type of contract

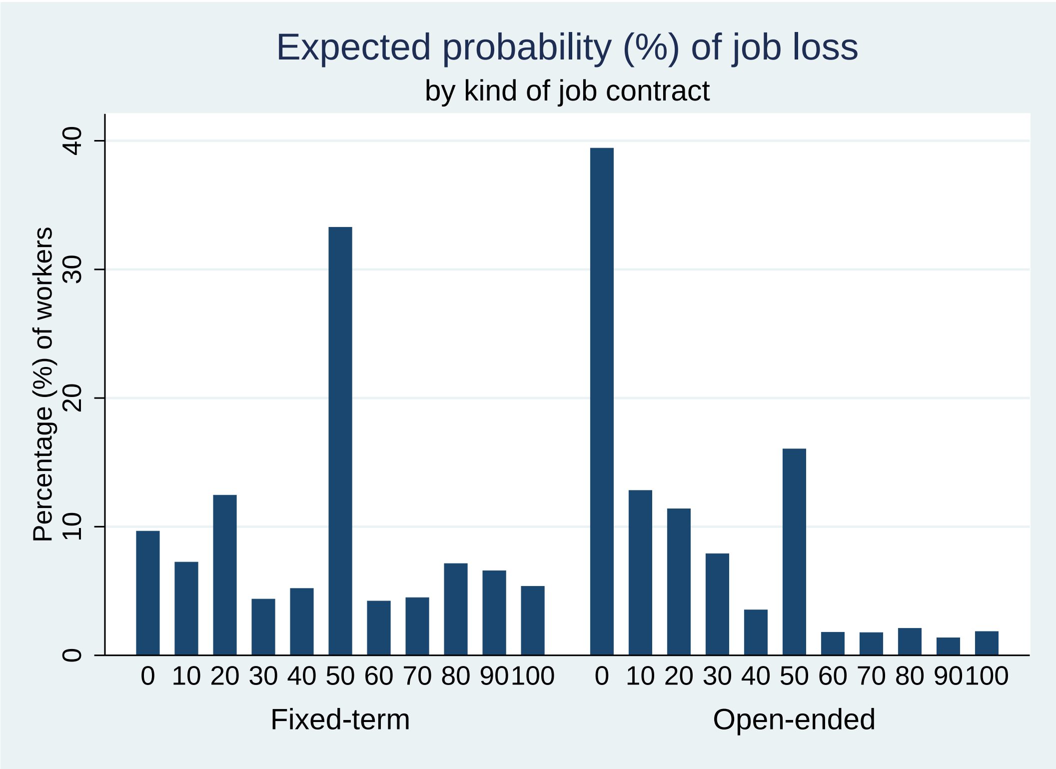
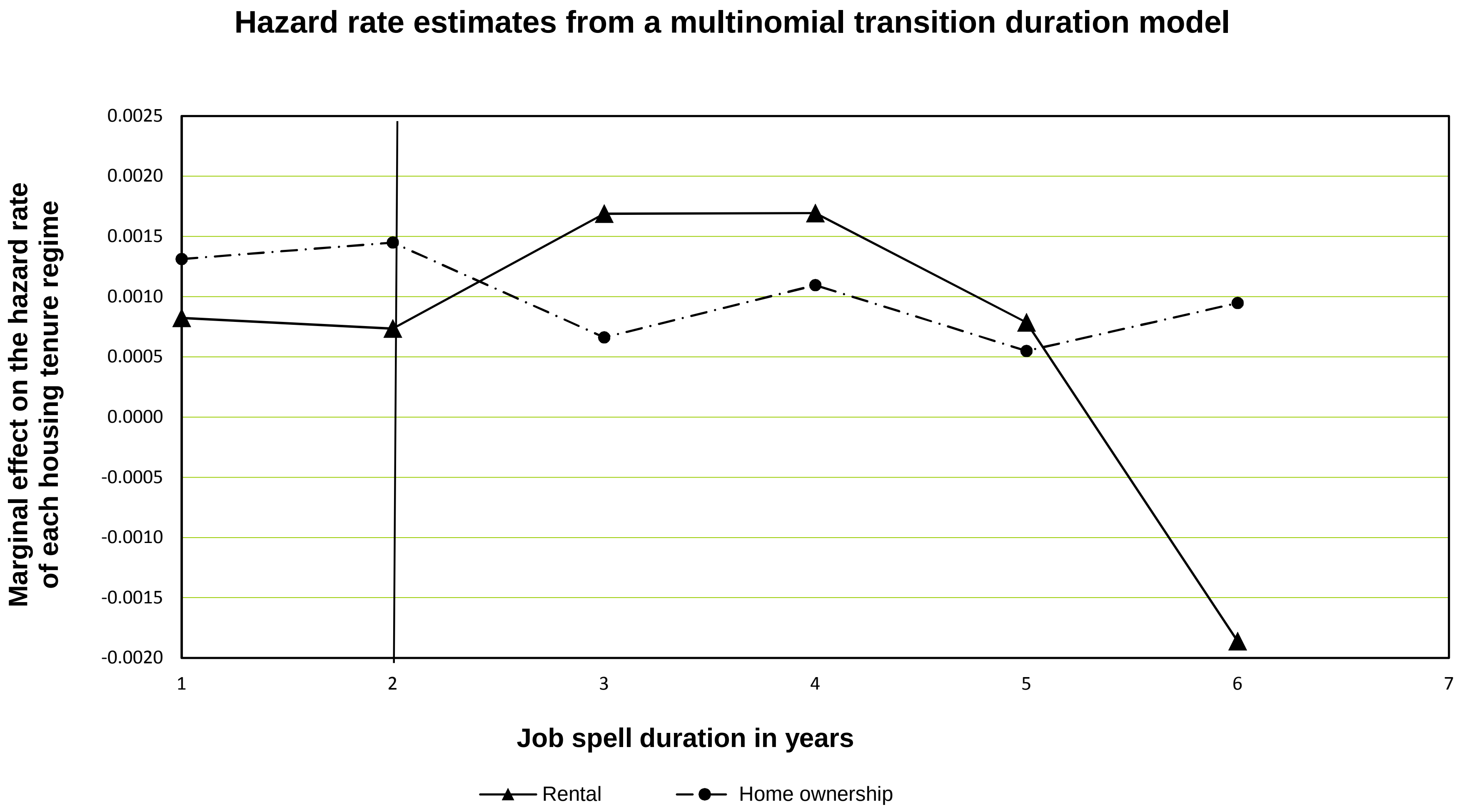


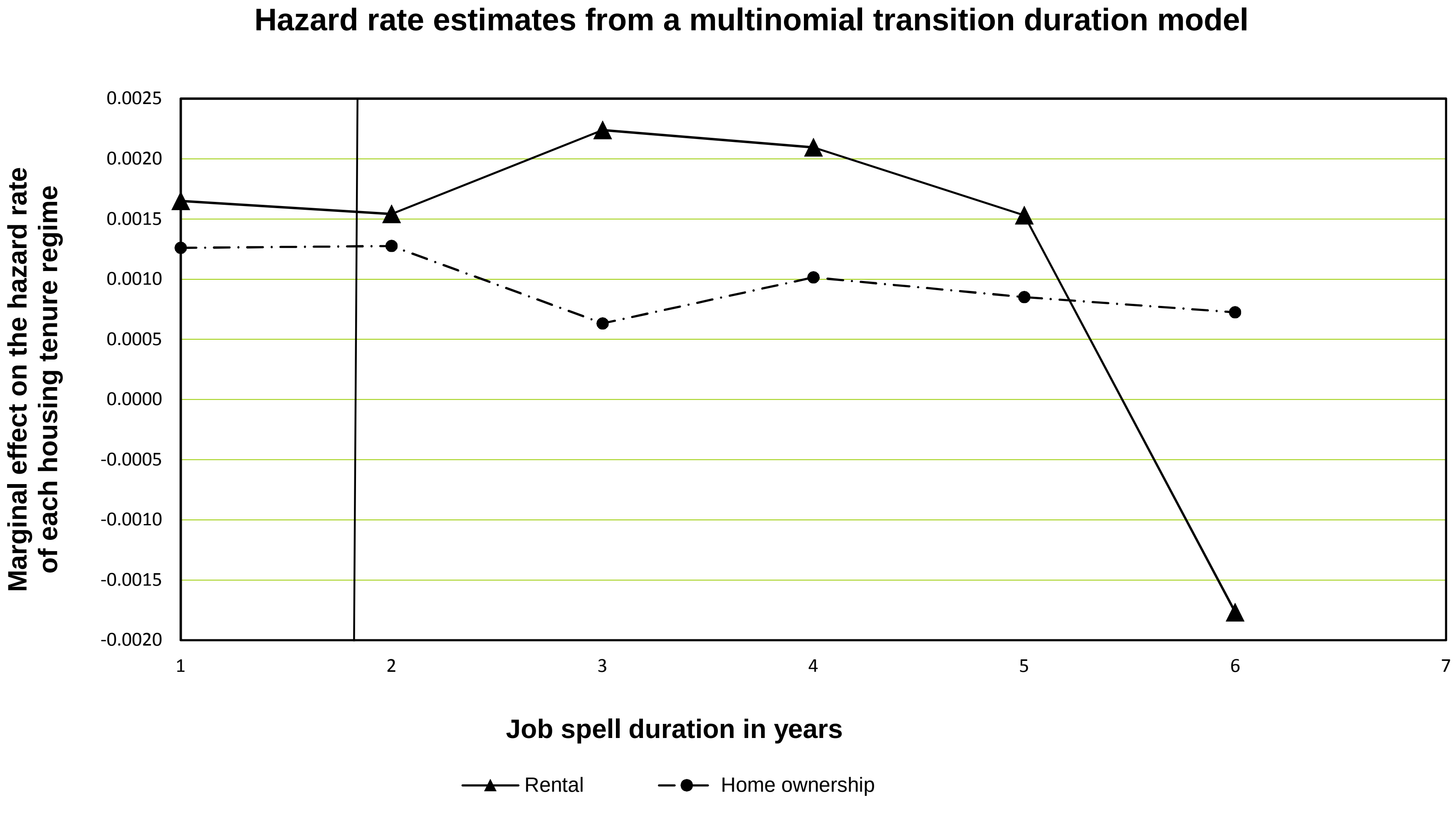
Figure 3: Marginal effects of 1000€ regional subsidies to contract conversion on the probability of household formation (rental vs home ownership) across job spell durations in a sample of transitions to a new accommodation of individuals aged between 25 and 64.



Each point is the marginal impact of a 1000 euro increase in the mean subsidy to contract conversion during the first two years of the contract on the demand of housing each year of job tenure.

The full line is the impact on the probability of renting a new accommodation and the dashed line if the impact on owning. The omitted outcome is living with parents.

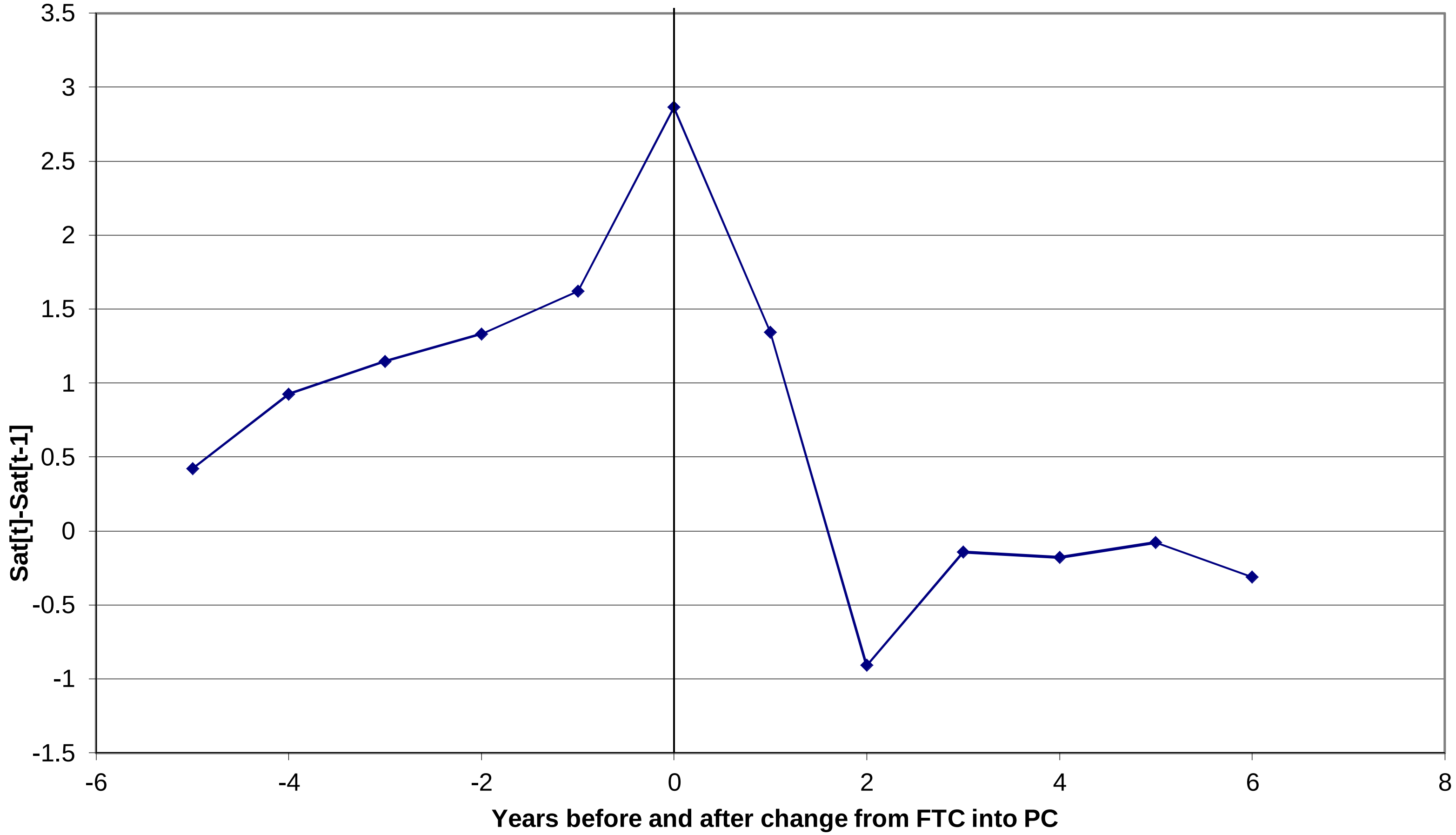
Figure 4: Marginal effects of 1000€ regional subsidies to contract conversion on the probability of household formation (rental vs home ownership) across job spell durations in a sample of transitions to a new accommodation of individuals aged between 25 and 45.



Each point is the marginal impact of a 1000 euro increase in the mean subsidy to contract conversion during the first two years of the contract on the demand of housing each year of job tenure.

The full line is the impact on the probability of renting a new accommodation and the dashed line if the impact on owning. The omitted outcome is living with parents.

Figure A.1 One-year change in satisfaction with job security



The graph shows the yearly change in satisfaction (Sat) with security on the job (measured from 0 to 10) on a window between 5 years before (Sat[t-1]) and 6 years after an employee obtains an upgrade from a fixed-term contract (FTC) into a permanent one (PC) (Sat[t]). The sample is an unbalanced panel drawn from the 2004-2011 waves from the European Community Household Panel.

Table 1: Summary statistics of the stock sample of employed individuals aged between 25 and 64 with a job tenure of seven years at most.

	Total sample	By type of job contract: Open-ended	By type of job contract: Fixed-term
Holding an open-ended contract	0.636	1.000	0.000
Holding a fixed-term contract	0.364	0.000	1.000
Age	36.351	36.445	36.187
	(8.69)	(8.69)	(8.69)
Age at hire	33.387	32.972	34.113
	(8.77)	(8.77)	(8.72)
No. of years at current job	2.964	3.473	2.074
	(1.84)	(1.71)	(1.70)
Individual eligible for subsidy	0.562	0.562	0.562
Amount individual was eligible for	1.896	1.993	1.726
(include zeroes)	(2.42)	(2.50)	(2.25)
Individual labor earnings:			
Mean	14.699	16.316	11.873
	(9.52)	(10.46)	(6.71)
Median	12.811	14.098	10.872
Attained education level:			
Primary education or less	0.152	0.113	0.220
First stage of secondary educ.	0.456	0.447	0.472
Second stage of secondary educ.	0.151	0.169	0.119
Tertiary education	0.240	0.269	0.189
Male individual	0.562	0.583	0.525
Housing tenure regime:			
Living with parents	0.332	0.314	0.363
Home ownership	0.507	0.550	0.432
Rental	0.161	0.136	0.205
Minimum sample size	5,087	3,408	1,678

Source: Pooled sample of the 2002-2011 waves of the Spanish Survey of Household Finances (EFF).

Sample: A minimum sample size of 5,087 household-year observations in each one of the five datasets imputed multiply in the four waves of the EFF data.
All summary statistics are weighted. Standard deviations are in parentheses. Monetary values are expressed in thousands of 2005 euro. Subsidy amounts are in real terms using deflators of the regional gross disposable income.

Table 2: Summary statistics of the stock sample of employed individuals aged between 25 and 45 with a job tenure of seven years at most.

		By type of job contract	
	Total sample	Open-ended	Fixed-term
Holding an open-ended contract	0.635	1.000	0.000
Holding a fixed-term contract	0.365	0.000	1.000
Age	33.301	33.381	33.160
	(5.56)	(5.51)	(5.64)
Age at hire	30.366	29.937	31.114
	(5.75)	(5.69)	(5.76)
No. of years at current job	2.935	3.444	2.046
	(1.83)	(1.72)	(1.67)
Individual eligible for subsidy	0.565	0.563	0.570
Amount individual was eligible for	1.936	2.030	1.772
(include zeroes)	(2.45)	(2.53)	(2.30)
Individual labor earnings:			
Mean	14.839	16.522	11.903
	(9.37)	(10.20)	(6.77)
Median	13.024	14.234	10.854
Attained education level:			
Primary education or less	0.113	0.079	0.172
First stage of secondary educ.	0.450	0.428	0.488
Second stage of secondary educ.	0.164	0.185	0.127
Tertiary education	0.272	0.306	0.213
Male individual	0.560	0.580	0.524
Housing tenure regime:			
Living with parents	0.371	0.353	0.404
Home ownership	0.466	0.511	0.388
Rental	0.162	0.135	0.209
Minimum sample size	3,974	2,639	1,335

Source: Pooled sample of the 2002-2011 waves of the Spanish Survey of Household Finances (EFF).

Sample: A minimum sample size of 3,974 household-year observations in each one of the five datasets imputed multiply in the four waves of the EFF data.
All summary statistics are weighted. Standard deviations are in parentheses. Monetary values are expressed in thousands of 2005 euro. Subsidy amounts are in real terms using deflators of the regional gross disposable income.

Table 3: Individual characteristics in the sample of job spells of individuals aged 25 to 64 with job tenures below 10 years.

	Total Sample	Subsample conditional on: Moving to a new house	Individual moves to: An owned house	Individual moves to: A rented house
Move to an owned house	0.136	0.671	1.000	0.000
Move to a rented house	0.066	0.329	0.000	1.000
Holding an open-ended contract	0.682	0.783	0.846	0.653
Holding a fixed-term contract	0.318	0.217	0.154	0.347
Age	36.754	35.561	35.857	34.957
	(8.73)	(6.75)	(6.80)	(6.61)
Age at hire	32.860	30.166	29.946	30.615
	(8.75)	(6.77)	(6.80)	(6.69)
No. of years at current job	3.893	5.395	5.910	4.342
	(2.55)	(2.38)	(2.12)	(2.52)
Individual eligible for subsidy	0.567	0.616	0.610	0.627
Amount individual was eligible for	1.998	2.491	2.437	2.603
(include zeroes)	(2.54)	(2.83)	(2.71)	(3.07)
Individual labor earnings:				
Mean	15.219	17.717	18.779	15.551
	(9.84)	(11.07)	(11.73)	(9.21)
Attained education level:				
Primary education or less	0.151	0.091	0.073	0.128
First stage of secondary educ.	0.447	0.414	0.380	0.483
Second stage of secondary educ.	0.157	0.194	0.212	0.158
Tertiary education	0.244	0.301	0.335	0.231
Male individual	0.556	0.563	0.546	0.598
No. of years elapsed from the job spell until a move occurs or a job spell ends:				
One year	12.880	27.720	27.850	27.450
Two years	16.690	22.540	22.300	23.030
Three years	16.510	13.960	13.530	14.860
Four years	13.760	13.850	14.750	12.030
Five years	11.580	9.790	10.990	7.360
Six years	9.510	5.510	5.290	5.970
Seven years	7.000	4.360	3.680	5.750
Eight years or more	14.590	2.260	3.690	5.390
Minimum sample size	5,997	1,072	703	368

Source: The sample is formed by all individuals aged between 25 and 64 years, who are employees with a job tenure not longer than 10 and who have been interviewed in the 2002-2011 waves of the Spanish Survey of Household Finances (EFF). Job spells of movers to a new house before starting to work at the current job are considered as censored observations, as well as job spells of individuals living with their parents.

All summary statistics are weighted. Standard deviations are in parentheses. Monetary values are expressed in thousands of 2005 euro. Subsidy amounts are in real terms using deflators of the regional gross disposable income.

Table 4: The effect of subsidies for job contract conversions on the stock of open-ended contracts and on the decision of housing tenure during period 2002-2012.

Estimation method: Ordinary Least Squares estimates		
PANEL A: FIRST-STAGE ESTIMATES		
Dependent variable: Indicator of whether the individual has an open-ended contract		
Sample:	Individuals aged between:	
	25 and 45 years	25 and 64 years
	(1)	(2)
1. Subsidy to contract conversion	0.0062	0.0082
(standard error)	(.0026)**	(.0024)***
2. Constant	0.553	0.554
(standard error)	(.049)***	(.060)***
F-test of instruments	5.96	11.21
Minimum sample size	3,974	5,087
PANEL B: INTENTION-TO-TREAT ESTIMATES		
Sample:	Individuals aged between:	
	25 and 45 years	25 and 64 years
Dependent variable (in italics below):	(1)	(2)
1. The individual lives with parents:		
a. Subsidy to contract conversion	-0.0131	-0.0102
(standard error)	(.0022)***	(.0018)***
b. Constant	0.304	0.228
(standard error)	(.052)***	(.041)***
2. The individual is a homeowner:		
a. Subsidy to contract conversion	0.0058	0.0044
(standard error)	(.0024)**	(.0022)*
b. Constant	0.445	0.534
(standard error)	(.039)***	(.034)***
3. The individual is a renter:		
a. Subsidy to contract conversion	0.0073	0.0057
(standard error)	(.0030)**	(.0016)***
b. Constant	0.251	0.238
(standard error)	(.059)***	(.043)***
Minimum sample size	3,974	5,087

Source: The sample is formed by all household members aged between 25 and 64 years, who are employees with a job tenure not longer than 7 and who have been interviewed in the 2002-2011 waves of the Spanish Survey of Household Finances (EFF). The interviews ended up in 2012.

Other covariates included in the model not shown here are the following: indicators of age at hire, indicators of year at hire, year dummies, indicators of the household member's gender and education level, a third-order polynomial based on the logarithm of the household member's labor income and region dummies. Standard errors are corrected for heteroscedasticity, also take into account arbitrary correlation among regions and combined across 5 implicates.

Table 5: The average effect of job contract on the decision of housing tenure during period 2002-2012 in a sample of employees.

Estimation method: Ordinary Least Squares (OLS) and Two-Step Least Squares (TSLS) estimates				
	OLS estimates		TSLS estimates	
Sample:	Individuals aged between:		Individuals aged between:	
	25-45 years	25-64 years	25-45 years	25-64 years
Dependent variable (in italics below):	(1)	(2)	(3)	(4)
1. The individual lives with parents:				
a. Open-ended contract	0.029	0.018	-2.108	-1.245
(standard error)	(0.026)	(.021)	(1.057)**	(.506)***
Robust confidence interval	--	--	[-16.031,-.911]	[-4.433, -.471]
at the confidence level of:	--	--	90%	95%
b. Constant	0.219	0.163	1.471	0.918
(standard error)	(.054)***	(.041)***	(.699)**	(.373)***
2. The individual is a homeowner:				
a. Open-ended contract	0.037	0.035	0.927	0.543
(standard error)	(.021)*	(.021)	(.465)**	(.327)*
Robust confidence interval	--	--	[.058,7.121]	[-.066, 1.69]
at the confidence level of:	--	--	90%	90%
b. Constant	0.453	0.537	-0.068	0.233
(standard error)	(.036)***	(.030)***	(.289)	(.224)
3. The individual is a renter:				
a. Open-ended contract	-0.066	-0.053	1.181	0.701
(standard error)	(.018)***	(.017)***	(.837)	(.337)**
Robust confidence interval	--	--	[.357,9.406]	[.089, 2.747]
at the confidence level of:	--	--	90%	95%
b. Constant	0.328	0.300	-0.402	-0.151
(standard error)	(.062)***	(.046)***	(.540)	(.243)
F test of instruments in first-stage	--	--	5.96	11.21
Minimum sample size	3,974	5,087	3,974	5,087

Source: The sample is formed by all household members aged between 25 and 64 years, who are employees with a job tenure not longer than 7 and who have been interviewed in the 2002-2011 waves of the Spanish Survey of Household Finances (EFF). The interviews ended up in 2012.

The set of covariates is identical to that in Table 4. Other covariates not shown here are the following: indicators of age at hire, indicators of year at hire, year dummies, indicators of the household member's gender and education level, a third-order polynomial based on the logarithm of the household member's labor income and region dummies. Standard errors are corrected for heteroscedasticity, also take into account arbitrary correlation among regions and combined across 5 implicates.

A confidence interval that is robust to the presence of weak instruments and that tests the hypothesis of a zero impact of regional subsidies is computed using the approach of Chernozhukov and Hansen (2008) at the 95% confidence level.

Table 6: Estimates of the probability that individuals move to a new house at each year of their job tenure.

Estimation method: logit model estimates (competing-risk models)

Sample: Individuals aged between 25 and 64 years being employees with job tenures shorter than 10.

	Time-invariant effects of subsidies (i)			Time-varying effects of subsidies (ii)		
Dependent variable:	Indicator of whether the individual moves to:					
	A new house (1)	An owned house (2)	A rented house (3)	A new house (4)	An owned house (5)	A rented house (6)
Subsidy to contract conversion	0.064 (.011)***	0.065 (.016)***	0.063 (.019)***	0.089 (0.018)***	0.499 (0.024)**	0.132 (0.023)***
Subsidy * 1 year of job tenure	--	--	--	-0.037 (0.016)**	0.466 (0.026)	-0.095 (0.046)**
Subsidy * 2 years of job tenure	--	--	--	-0.023 (0.020)	0.281 (0.029)	-0.093 (0.048)**
Subsidy * 4 years of job tenure	--	--	--	-0.002 (0.015)	0.347 (0.020)	-0.016 (0.047)
Subsidy * 5 years of job tenure	--	--	--	-0.053 (0.024)**	-0.034 (0.037)	-0.080 (0.054)
Subsidy * 6 years of job tenure	--	--	--	-0.053 (0.041)	-0.041 (0.036)	-0.213 (0.086)***
Subsidy * 7 or more years of job tenure	--	--	--	-0.025 (0.045)	-0.753 (0.051)	-0.117 (0.042)***
One year of job tenure	0.428 (0.118)***	0.499 (0.121)***	0.237 (0.161)	0.557 (0.138)***	0.451 (0.172)***	0.595 (0.217)***
Two years of job tenure	0.337 (0.105)***	0.466 (0.120)***	0.070 (0.208)	0.423 (0.152)***	0.370 (0.178)**	0.424 (0.319)
Four years of job tenure	0.201 (0.099)**	0.281 (0.128)**	0.070 (0.135)	0.209 (0.147)	0.223 (0.157)	0.135 (0.289)
Five years of job tenure	0.192 (0.111)*	0.347 (0.110)***	-0.098 (0.223)	0.380 (0.146)***	0.414 (0.200)**	0.223 (0.354)
Six years of job tenure	-0.040 (0.142)	-0.034 (0.178)	0.029 (0.388)	0.146 (0.193)	-0.143 (0.206)	0.776 (0.437)*
Seven years of job tenure	-0.177 (0.229)	-0.041 (0.313)	-0.389 (0.235)*	-0.084 (0.232)	-0.137 (0.359)	0.077 (0.220)
Eight years of job tenure	-0.516 (0.287)*	-0.753 (0.373)**	-0.031 (0.489)	-0.422 (0.282)	-0.849 (0.421)**	0.439 (0.366)
Nine or ten years of job tenure	-0.778 (0.372)**	-1.093 (0.427)***	-0.116 (0.616)	-0.685 (0.319)**	-1.189 (0.446)***	0.354 (0.479)
Constant	-3.589 (.331)***	-4.325 (.368)***	-4.235 (.384)***	-3.682 (.339)***	-4.275 (.377)***	-4.524 (.378)***
Minimum number of spells	5,997	5,997	5,997	5,997	5,997	5,997

Source: The sample is formed by all household members aged between 25 and 64 years, who are employees with a job tenure not longer than 10 and who have been interviewed in the 2002-2011 waves of the Spanish Survey of Household Finances (EFF). The interviews ended up in 2012.

Other covariates included in the model not shown here are the following: indicators of age at hire, indicators of year at hire, year dummies, indicators of the household member's gender and education level and region dummies.

Standard errors are corrected for heteroscedasticity, also take into account arbitrary correlation among regions and combined across 5 implicates.

Table 7: Impact of regional subsidies for the conversion of job contracts on the decision of leaving the parental house, by different individual characteristics.

Sample: Individuals aged between 25 and 45 that live with their parents and that are employees with job tenures shorter than 7 years

Estimation method: Probit Estimates

Dependent variable: Indicator of leaving parents' house three years after having been interviewed in the survey.

		By individual's age		By individual's level of education		By initial household total net wealth		By initial household member labour income	
Samples:	Total sample	Aged 25-30	Aged 31-45	<=1st stage secondary educ.	>=second stage secondary educ.	<= median	>median	<= median	>median
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)	(9)
1. Subsidy to contract conversion	0.029	0.013	0.068	0.072	0.007	0.038	0.022	0.037	0.019
(standard error)	(.014)**	(.020)	(.028)***	(.041)*	(.023)	(.029)	(.025)	(.022)*	(.017)
2. Constant	-0.409	0.038	-0.598	-1.064	-0.256	-0.336	-0.482	-0.327	-0.195
(standard error)	(.442)	(.394)	(.760)	(.687)	(.341)	(.549)	(.526)	(.429)	(.388)*
Marginal impact of 1000€ subsidy	0.010	0.005	0.021	0.016	0.003	0.015	0.007	0.014	0.007
Minimum sample size:	1203	761	441	472	726	415	780	545	642
Weighted sample mean:	0.480	0.518	0.409	0.425	0.537	0.450	0.507	0.453	0.513

Notes: The panel sample is formed by all household members that live with their parents, aged 25-45 years, that work as employees with a job tenure of seven years at most and whose parental households have been interviewed in the 2002-2014 waves of the Spanish Survey of Household Finances (EFF). The covariate of interest is the subsidy for job contract conversions that each individual was eligible for in the first two years of job tenure. The set of regressors is the same as thosed used in Table 4.

The marginal effects are computed for an individual aged between 31 and 40 years when was hired in 1999 and observed in the sample of 2002. The remaining covariates were evaluated at their sample means. Standard errors (in parentheses) clustered at the region level. Estimates combined across five implicates.

The symbols *, ** and *** denote the estimate is significant at the 10% of significance level, at the 5% and at the 1%, respectively.

Table 8: The impact of regional subsidies for the conversion of job contracts on the variation of household wealth and on the decision of leaving the parental home three years after using a panel sample.

Sample: Individuals living with their parents, aged between 25 and 45 years and working as a employee with a job tenure shorter than 7 years.
Estimation method: Quantile Regression Estimates for the median

PANEL A: Median estimates of the change of various measures of household wealth in logarithm

Dependent variable: Variation of the household wealth in logarithm between two consecutive triennial survey waves (three years)

	Gross financial liquid wealth	Gross financial liquid wealth +			
		+ pension schemes	+ pension schemes + life insurance	+ pension schemes + life insurance+ net value of real assets other main house	Total net wealth
	(1)	(2)	(3)	(4)	(5)
1. Subsidy to contract conversion (standard error)	0.020 (.037)	0.012 (.030)	0.011 (.034)	0.033 (.024)	0.012 (.009)
2. Constant (standard error)	0.630 (.611)	0.139 (.578)	0.168 (.571)	-0.392 (.364)	0.215 (.142)
Marginal impact of 1000€ subsidy on wealth (in 1000€)	0.122	0.130	0.121	1.120	2.368
Minimum sample size:	1138	1148	1148	1001	1157

PANEL B: Ordinary Least Square Estimates of the decision of leaving the parental home three years after being interviewed for each sample of Panel A

Dependent variable: Indicator of whether the individual has left her or his parents' house three years after.

	Leaving parental home (1)	Leaving parental home (2)	Leaving parental home (3)	Leaving parental home (4)	Leaving parental home (5)
1. Subsidy to contract conversion (standard error)	0.011 (.0049)**	0.012 (.0049)**	0.012 (.0049)**	0.012 (.0063)*	0.012 (.0051)**
2. Constant (standard error)	0.365 (.154)**	0.361 (.158)**	0.361 (.158)**	0.322 (.206)	0.319 (.179)*
Minimum sample size:	1138	1148	1148	1001	1157

Notes: The panel sample is formed by all household members that live with their parents, aged between 25 and 45 years, that work as employees with a job tenure of seven years at most and whose parental households have been interviewed in the 2002-2014 waves of the Spanish Survey of Household Finances (EFF). The covariate of interest is the subsidy for job contract conversions that each individual was eligible for in the first two years of job tenure. The set of regressors is the same as in Table 4.

The marginal effects are computed for a parental household holding the median wealth and for individuals aged 31-40 years when was hired in 1999, observed in 2002. The remaining covariates were evaluated at their sample means. Standard errors (in parentheses) clustered at the region level. Estimates combined across five implicates.

The symbols *, ** and *** denote the estimate is significant at the 10% of significance level, at the 5% and at the 1%, respectively.

Table 9: The impact of regional subsidies for job contract conversion on the access to credit markets of employees aged between 25 and 45 years.

Sample: Individuals aged between 25 and 45 working as employees with job tenures shorter than 7 years.

PANEL A: Multinomial logit estimates of the probability of whether the household has borrowing constraints.

Dependent variable:	Indicator of being credit constrained:	
	Asked for a loan fully accepted	Credit constrained
1. Subsidy to contract conversion (standard error)	-0.015 (.023)	0.036 (.038)
2. Constant (standard error)	-0.870 (.274)***	-4.309 (1.308)***
Marginal impact of 1000€ subsidy	-0.0030	0.0005
Minimum sample size:	3974	
Weighted sample mean:	0.295	0.059

PANEL B: Multinomial logit estimates of the probability of borrowing constraints and the decision of housing tenure.

Dependent variable:	Indicator of being credit constrained or not and the housing tenure chosen:							
	Coresidence		Home ownership			Rental		
	Asked for a loan fully accepted	Credit constrained	Not asked for a loan	Asked for a loan fully accepted	Credit constrained	Not asked for a loan	Asked for a loan fully accepted	Credit constrained
1. Subsidy to contract conversion (standard error)	0.015 (.033)	0.044 (.045)	0.081 (.017)***	0.042 (.023)	0.140 (.087)	0.101 (.025)***	0.051 (.043)	0.080 (.079)
2. Constant (standard error)	-0.817 (.365)**	-3.701 (1.211)***	0.586 (.325)*	-0.155 (.387)	-21.382 (10.439)	0.064 (.487)	-1.982 (1.125)*	-15.471 (13.089)
Marginal impact of 1000€ subsidy	-0.0039	0.0000	0.0064	-0.0004	0.0000	0.0129	0.0001	0.0000
Minimum sample size:				3974				
Weighted sample mean:	0.097	0.017	0.284	0.161	0.022	0.105	0.038	0.020

Notes: The indicator of credit constrained means that the household where the individual lives has at least a loan rejection in the two years prior to the survey interview, has not asked for any loans due to the fear of rejection or has been granted a loan with smaller capital than that asked for. The omitted category in the estimates of Panel A is not having asked for a loan during the last two years, and in Panel B it is not having asked for a loan and living with their parents.

The sample is formed by all household members aged between 25 and 45 years, that work as employees with a job tenure of seven years at most and that have been interviewed in the 2002-2011 waves of the Spanish Survey of Household Finances (EFF). The covariate of interest is the subsidy for job contract conversions that each individual was eligible for in the first two years of job tenure. The set of regressors is the same as thosed used in Table 4. The marginal effects are computed for an individual aged between 31 and 40 years when was hired in 1999, observed in the sample of 2002 and their remaining covariates were evaluated at their sample means. Standard errors (in parentheses) clustered at the region level. Estimates combined across five implicates.

The symbols *, ** and *** denote the estimate is significant at the 10% of significance level, at the 5% and at the 1%, respectively.

Table 10: Impact of subsidies for job contract conversions on the decisions of living with a partner and housing tenure during period 2002-2012.

Estimation method: Ordinary Least Squares estimates

INTENTION-TO-TREAT ESTIMATES

PANEL A: SAMPLE OF INDIVIDUALS AGED 25-45 WITH A JOB TENURE OF SEVEN YEARS OR LESS

	Model (i)	Model (ii)			
Dependent variable:	Living with a partner	Owning and living with a partner	Owning and not living with partner	Renting and living with a partner	Renting and not living with partner
	(1)	(2)	(3)	(4)	(5)
1. Subsidy to contract conversion	0.0059	0.0021	0.0037	0.0045	0.0028
(standard error)	(.0023)**	(.0018)	(.0012)***	(.0019)**	(.0017)
2. Constant	0.636	0.420	0.024	0.177	0.074
(standard error)	(.032)***	(.042)***	(.014)*	(.040)***	(.026)***
Minimum sample size	3,974	3,974	3,974	3,974	3,974

PANEL B: SAMPLE OF INDIVIDUALS AGED 25-64 WITH A JOB TENURE OF SEVEN YEARS OR LESS

	Model (i)	Model (ii)			
Dependent variable:	Living with a partner	Owning and living with a partner	Owning and not living with partner	Renting and living with a partner	Renting and not living with partner
	(1)	(2)	(3)	(4)	(5)
1. Subsidy to contract conversion	0.0041	0.0010	0.0035	0.0030	0.0027
(standard error)	(.0025)	(.0022)	(.0012)***	(.0012)**	(.0014)*
2. Constant	0.671	0.466	0.068	0.171	0.066
(standard error)	(.022)***	(.035)***	(.025)**	(.028)***	(.021)***
Minimum sample size	5,087	5,087	5,087	5,087	5,087

Source: The sample is formed by all household members aged between 25 and 64 years, who are employees with a job tenure not longer than 7 and who have been interviewed in the 2002-2011 waves of the Spanish Survey of Household Finances (EFF). The interviews ended up in 2012.

The set of regressors is identical to that in Table 4. Other covariates not shown here are the following: indicators of age at hire, indicators of year at hire, year dummies, indicators of the household member's gender and education level, a third-order polynomial based on the logarithm of the household member's labor income and region dummies. Standard errors are corrected for heteroscedasticity, also take into account arbitrary correlation among regions and combined across 5 implicates.

Table A.1: Subsidies for conversion of fixed-term contracts into open-ended ones, by region and year

Region / Year	1997	1998	1999	2000
1. Andalucia	Between 1997 and 2000: 4200 if age<30 , 3000 if female >30, 2400 if male >30			
2. Aragon	None	4200 if female or age>45	5160 if female or age>45	5160 if female or age>45
		3000 if male 41<=age<=44	4500 if 41<=age<=44	4500 if 41<=age<=44
			3600 if male age<30	3600 if male age<46
3. Asturias	4500 for all	4500 for all	None	4,200 if female or age>=46
				3600 otherwise
4. Baleares	None	None	None	1652.78 if female
5. Canarias	None	3,600 if female or age<=25	3,600 if female or age<=25	None
			3,000 otherwise	
6. Cantabria	None	3900 if female or age<=30	None	5408 if age>=46
		3300 if male 30<age<=40		3606 if age<=30
		3,600 if male age>=41		2163 otherwise
7. Castilla-Leon	None	5112 if female or age <30	5115 if age <30	4507.59 if age <30
		3300 rest of males	3900 if female age>=30	3305.57 if female age>=30
			1800 if male age >=41	1803 if male age>=41
8. Castilla-La Mancha	None	3600 if females	None	3600 if female
		3000 if age<30		3000 if age>45 or age<30
10. Valencia	None	30% of payroll tax	30% of payroll tax	30% of payroll tax
11. Extremadura	10655.94 if age<45	10100 if age<=30	14027.62 if age>46	5217.076 if female age>46
	13402.57 if age>45	11180 if age>30 and age<=40	11178.83 if age<45	4296.416 if male age>46
		14027 if age>40		
12. Galicia	None	4200 euro if female or age<30	1800 for males	1800 for males
		3000 if age>45	2400 for females	2400 for females
13. Madrid	None	6000 euro if female	7800 if female	9000 if female
		6000 euro if age<30 or age>40	7800 if age<25 or age>40	6600 if age<25 or age>40
14. Murcia	None	None	None	2100 if age<30
				1800 for the rest
16. Basque country	None	7512 for all	7512 for all	7512 for all
17. Rioja	None	4500 for all	4491 for all	6011 for all

Table A.1: Subsidies for conversion of fixed-term contracts into open-ended ones, by region and year (continued)

Region / Year	2001	2002	2003	2004
1. Andalucia	4200 if age<30	6012 for females or age<30	None	None
	3000 if females >30	3607 if male age>40		
	2400 if males >30			
2. Aragon	5160 if female or age>45	4500 if female	4500 if female	3750 if male, 7250 if female
	4500 if age>=41	4125 if age<30 or age>=41	4125 if age<30 or age>=41	
	3600 if male age<30			
3. Asturias	4,200 if female or age>=46	4200 if female of age>46	4200 if female of age>46	None
	3600 otherwise	3600 otherwise	3600 otherwise	
4. Baleares	1652.78 for females	1800 for females	4808 for females	4808 for females
5. Canarias	None	None	None	None
6. Cantabria	4808 for females	same as previous year	same as previous year	same as previous year
	3005 if male age <=30			
	4207 if age >45, 1803 otherwise			
7. Castilla-Leon	4507.59 if age <30	same as previous year	same as previous year	same as previous year
	3305.57 if female age>31			
	1803 if male age>41			
8. Castilla-La Mancha	3600 if female	same as previous year	same as previous year	None
	3000 if age>45 or age<30			
10. Valencia	4808.1 for all	1800 for females	2000 for females	4000 if female
			1500 for the rest	2000 if age<30, 1500 otherwise.
11. Extremadura	5410.086 if female >45	6010 for all	None	None
	4455.365 if male > 45			
	2386.802 otherwise			
12. Galicia	1800 for males	1800 for males	1800 for males	1800 for males
	2400 for females	2400 for females	2400 if females	2400 for females
13. Madrid	12000 if above 45 (males)	12000 for all	9000 if age<=45	3000 euro, all
	12000 if above 40 (females)		12000 if above 45	
	10800 for the rest			
14. Murcia	4800 for all	4800 for all	4800 if age above 30	2400 for all
			5400 if female or age 30 or below	
16. Basque country	7512 for all	7512 for all	7512 for all	6000 for males, 7500 for females
17. Rioja	6011 for all	6011 for all	6011 for all	6011 for all

Table A.2: The effect of subsidies for job contract conversions on the stock of open-ended contracts using a sample of Security Social records for period 2002-2012.

Estimation method: Ordinary Least Squares (OLS) estimates					
Dependent variable: Indicator of whether the employee has an open-ended contract					
Sample:	Individuals aged 25-45 whose job tenure is:				
	1 year	<=2 years	<=3 years	<=4 years	<=7 years
	Hired since 2001				First-stage in EFF
	(1)	(2)	(3)	(4)	(5)
1. Subsidy to contract conversion	0.0055	0.0061	0.0056	0.0054	0.0059
(standard error)	(.0016)***	(.0017)***	(.0018)***	(.0020)**	(.0022)**
2. Constant	0.377	0.476	0.431	0.438	0.730
(standard error)	(.0209)***	(.025)***	(.031)***	(.035)***	(.023)***
F- test of instruments	11.80	12.59	9.84	6.97	6.84
Number of individuals	321,855	396,767	418,163	427,114	499,209
Average no. of observations per individual	1.82	2.52	3.11	3.58	4.28

Source: The sample is formed by all employees aged between 25 and 45 years and holding a job tenure not longer than 7 years. The data come from the 2004-2015 waves of the Continuous Sample of Working Histories (MCVL).

Other covariates included in the model not shown here are the following: indicators of age at hire, indicators of year at hire, year dummies, indicators of the household member's gender and occupation groups, a third-order polynomial based on the logarithm of the labor income earned last year and region dummies.
Standard errors are corrected for heteroscedasticity and also take into account arbitrary correlation among regions.
Notes: * denotes the estimates are significant at the 10% level, ** at the 5% level and *** at the 1% level.