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Abstract

Recent cohorts in various developed countries take a longer time to form their own household and display lower rates of home ownership than older cohorts. Previous literature has linked these developments to higher job instability, especially among youths. We exploit the 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 for firms to promote high firing cost contracts between 1997 and 2009. Using data from the 2002-2014 waves of the Spanish Survey of Household Finances, we document that an increase of 1% in the stock of workers with an open-ended contract increases the probability of forming a new household by a similar magnitude (especially through renting new accommodation). The results are consistent with the predictions of precautionary saving models, whereby individuals exposed to the risk of job loss postpone their consumption of housing services.

Keywords: job insecurity, household formation, housing investments.

JEL classification: J1, J2, D91.

Resumen

En varios países desarrollados, se ha observado tanto una caída de la proporción de individuos jóvenes que viven en su propio hogar como, entre estos, una reducción de la tasa de propiedad de la vivienda habitual. Una posible explicación de ambos fenómenos es el aumento de la inestabilidad laboral observado entre los más jóvenes, especialmente a raíz de la crisis que comenzó en 2008. Este estudio analiza cómo reacciona la demanda de vivienda ante cambios en la probabilidad de perder el empleo. Para ello, se utiliza el hecho de que, entre 1997 y 2009, los incentivos de las empresas a utilizar contratos indefinidos variaron sustancialmente entre Comunidades Autónomas y según el grupo de edad de los trabajadores. Utilizando datos de la Encuesta Financiera de las Familias, entre 2002 y 2014, se ofrece evidencia de que incrementos en el número de trabajadores con contrato indefinido llevan consigo aumentos similares en la formación de hogares, en especial mediante el alquiler de una nueva vivienda. Estos resultados sugieren la presencia de un motivo de ahorro de precaución, por el cual los individuos que se enfrentan a una mayor probabilidad de perder el empleo posponen tanto el alquiler como la compra de vivienda.

Palabras clave: inseguridad laboral, formación de hogares, inversión en vivienda habitual.

Códigos JEL: J1, J2, D91.

Introduction

Employment protection legislation and mandated severance payments vary widely across and within countries. Previous literature studying the consequences of that variation has stressed that while high mandated severance payments generate allocation costs by reducing job hiring and firing, they also increase welfare by mitigating the exposure to job loss of risk-averse employees.[1] While the literature assessing the magnitude of precautionary savings is large, much less is known about how employment protection alters important lifetime decisions, such as forming a new household, and the associated consumption decisions, such as becoming a home-owner or a renter.

Housing decisions are important for various reasons. Firstly, the macroeconomic literature has emphasized that increases in the risk of job loss diminish the expenditure on durable goods, which helps to explain the evolution of consumption along the business cycle.[2] Indeed, the empirical literature has documented a negative relationship between home ownership and the exposure to the risk of job loss.[3] In addition, housing is a prominent component of household portfolios, and its size -and the way it is financed- determines the effectiveness of monetary and fiscal policy. Furthermore, fluctuations in the value of housing affect consumption.[4] Hence, understanding the determinants of housing decisions is key to analyzing how macroeconomic shocks propagate. Finally, as we discuss below, among the population of youths the consumption of housing services is tied to household formation, a proxy of lifetime welfare that has attracted recent attention given international trends in several advanced economies.[5]

This paper focuses on Spain, a country with a large variation in the degree employment protection varies across workers -like it is the case in many European countries. As a result of policies aimed at introducing flexibility at the margin of the labor market during the 80s, around 30% of employees are covered by fixed-term contracts, typically with a pre-

[1] See Cozzi and Fella (2016) and Alvarez and Veracierto (2001). In Barceló and Villanueva (2016) we quantify the impact of the reduction of unemployment risk on wealth and non-durable consumption in Spain.

[2] Dunn (2003) and Shore and Sinai (2010) show how the decision to own or rent is affected by exposure to job loss. An earlier study is Eberly (1994) on the response of expenses in automobiles to idiosyncratic income risk.

[3] Gathergood (2011), Diaz-Serrano (2005), Henley et al (1994), Henley (1998) and Duca and Rosenthal (1994) document the negative relationship between home ownership and labor market uncertainty. However, note that these empirical studies have focused on the analysis of households, instead of individuals. The way we deal with the endogeneity of exposure to job loss risk is very different from those papers.

[4] Badarinza et al (2016) and Mathä et al (2017) discuss the prominence of housing. Campbell and Cocco, (2007) discuss how consumption varies with housing wealth and Cloyne et al (2018) discuss how macroeconomic shocks propagate depending on household housing investments.

[5] In the US, Dettling and Hsu (2018), Bleemer et al. (2017), Paciorek (2016), Kaplan (2012), Cooper and Luengo-Prado (2017), or Bell et al. (2007) have documented the increase in co-residence between adult children and their parents. Similar findings have been documented for the United Kingdom by Blundell et al (2017). Martins and Villanueva (2009) studies the link between household formation and access to credit and Adamopoulou and Kaya (2018) the influence of peers on household formation.

specified duration after which they can be terminated with small costs. On the other hand, employees under open-ended contracts are protected against unfair dismissal by court appeal and severance payments. Our study exploits those large differences in employment protection -and thus exposure to the risk of job loss- to make three contributions.

Firstly, we provide a direct test of the impact of exogenous changes in employment protection on the housing choices of employees. We take explicitly into account that the decision of a firm to grant an open-ended contract depends on the worker's expected productivity as well as on local labor demand conditions, variables that in turn also correlate with the propensity to consume housing services.[6] To mitigate the influence of those unobservables, we thus exploit only exogenous variation in the firms' incentives to convert fixed-term contracts into open-ended ones. Namely, as a result of decentralized labor laws, several Spanish regions introduced different subsidies to firms that converted fixed-term contracts into open-ended ones between 1997 and 2009. Those incentives varied by gender, age of the worker and year, a variation that allows us to isolate the causal impact of changes in employment protection on housing choices.

Secondly, we highlight the relevance of examining the household formation, an increasingly important margin in a number of countries.[7] Failing to take into account household formation may lead to sample selection biases in the literature that examines the negative relationship between home-ownership and unemployment focusing on samples of already established households.[8] If household formation increases with job security, the trade-offs between home ownership and renting detected in previous studies may be the artifact of sample selection: more households are created when the risk of job loss decreases, and the characteristics of those selected households lead them to own, rather than to rent. The characteristics that influences on household formation are higher unobserved ability to accumulate income and wealth resources. We present evidence that this is indeed the case.

Thirdly, we use rich household data from the Survey of Household Finances (EFF) a sample that allows us to explore the mechanisms linking the risk of job loss and the housing decisions of Spanish employees. The survey contains retrospective information about both the date of house purchases and of entering into rental housing, and the year of

[6]See Topel (1986) for the relevance of local labor markets.

[7]Kaplan (2012), Bleemer et al (2017), 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.

[8]The literature has studied the relationship between unemployment risk and patterns related to household formation. Becker et al (2010) find a positive relationship between the prevalence of co-residence between children and parents and the incidence of fixed-term contracts. Adsera (2005) discusses whether labor market institutions in Southern European countries account for the drop in fertility (a lifetime event related to household formation). Finally, Deaton and Paxson (2000) and Chiuri and Jappelli (2003) emphasize the need of using a sample of individuals instead of households to analyze housing and saving decisions, in order to avoid sample composition effects.

job hire, what allows us to examine the dynamic response of housing decisions to changes in employment protection.[9] In addition, the survey contains longitudinal information on household wealth that allows us to trace the evolution of wealth on subsamples where a household member has left the residence of the parents.

Our results suggest a strong relationship between exogenous increases in job security and the consumption of housing services -either rented or owned. Namely, an increase in the subsidy for contract conversion of 1,000 euros increases the stock of open-ended contracts three years after the beginning of a job spell by .8 percentage points (pp) and diminishes the stock of individuals living with parents by a similar amount (1.01 pp). Unlike suggested by previous literature, higher employment protection and lower risks of job loss result in a higher incidence of *both* owning and renting. Namely, an extra of 1,000 euros of a subsidy to contract conversion increases the chances of renting between .44 and .57 percentage points while owning increases less than .44 pp. However, we find that home ownership increases while renting decreases with job security when we do not control for the endogeneity of type of labor contract. Finally, when we examine the dynamics of increases in employment protection and housing choices, we find a delayed response of the incidence of household formation -especially, renting- after an exogenous increase in job security. This delay is not explained by credit frictions that discourage owning, but possibly due to couple formation.[10] Therefore, higher job instability helps in explaining the decrease in household formation, while the impact on home ownership is less clear.

Overall, our findings imply that employment protection increases the consumption of all forms of housing services, a prediction that is consistent with standard models of precautionary saving. However, our results do not confirm another prediction of that model, namely that home ownership is specially responsive to increases in job security - funds accumulated for precautionary purposes can be channeled towards a downpayment. The analysis of a subsample of youths who form their own household reveals that youths do not hold sizable wealth holdings in response to job loss risk that can be used for a downpayment.[11] Therefore, one can view our results as confirming that in the absence of substantial wealth accumulation, home ownership is not the dominant response to increases in job security. In any case, a fraction of the population reacts to lower employment protection by postponing housing consumption, an indication of welfare losses.

[9] 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.

[10] An alternative explanation not considered in this paper is peer effects as a mechanism for the delay in household formation, see Adamopoulou and Kaya (2018).

[11] Crossley and Low (2011) and Dunn (2003) simulate models where funds accumulated as precautionary savings can serve as sources for life-cycle saving or a downpayment when the risk of job loss disappears. In that case, the welfare losses due to cutting consumption when risks are higher are lower than in situations when excess savings must all be consumed in good times.

The paper is organized as follows. Section 1 provides some background on employment protection in the Spanish labor market. Section 2 discusses the link between fixed-term contracts and housing choices. Section 3 lays out the empirical strategy in the analysis of household formation and housing demand, Section 4 describes the data used in the estimation, Section 5 presents the empirical results and Section 6 concludes.

1 Employment protection in Spain

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. Fixed-term contracts in Spain featured very low indemnities for termination, that were virtually zero if the firm waited until expiration of the term specified in the contract and constitute the main exit from unemployment -see Bover and Gómez, 2004 or García Pérez et al (forthcoming).

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, three of the 17 regional authorities decided to subsidize firms who signed open-ended contracts, possibly in response to the growing incidence of fixed-term contracts among vulnerable workers - see García Pérez and Rebollo Sanz (2009). Subsidies to contract conversion were typically lump-sum amounts given to firms that proved that they converted a fixed-term contract into an open-ended one. Table A.1, constructed using data from García Pérez and Rebollo Sanz (2009), shows the subsidies by region and demographic groups.[12] In Barceló and Villanueva (2016) we conduct an analysis of the effectiveness of regional subsidies in the conversion of fixed-term contracts into open-ended ones. The average subsidy was around 4,000 euro, that amounts to 20% of the average cost of labor. Further taking into account the increase in labor costs associated to converting a contract into open-ended we estimate that the mean subsidy amounts to 16% of the yearly labor cost that the worker entails to the firm.[13] Moreover, in regions that implemented subsidies, around the 23% of contract conversions were subsidized, what implies that between 5% and 7% of all conversions were subsidized in Spain.

[12] 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 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).

[13] We approximate the expected increase in dismissal cost as

$$\sum_{t=1}^{15} P(stay_{t-1}) * P(dismiss_t = 1|Stay_{t-1}) * \frac{1}{(1+r)^t} * \{.125 * t\}$$

Type of job contract and perceptions of risk

A key point in this study is whether or not workers covered under a fixed-term contract actually *perceive* a higher unemployment risk than those under an open-ended contract.[14] Two pieces of evidence suggest that employees under fixed-term contracts perceive higher insecurity in their jobs. Firstly, a sample drawn from the European Community Household Panel documents that satisfaction with job security increases monotonically with tenure up to contract conversion to remain constant afterwards (Figure A.1). That evidence suggests that workers seem to feel more certain about their jobs when these are covered by a permanent contract, instead of a fixed-term one. As a second piece of evidence, Figure 1 shows the individuals' subjective probability of job loss over the next twelve months in 2011, broken down by the kind of job contract using data from the Spanish Survey of Household Finances. More than 50% of workers holding an open-ended contract assign a probability lower than 10% of losing their job during the next 12 months. Among employees covered by a fixed-term contract, 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%.

2 Fixed-term contracts and housing choices

A natural way to assess the consequences of employment protection legislation on the consumption of housing services is to consider uncertainty about job loss as an uninsurable risk –see Hendren (2017). The literature studying the consequences of the risk of job loss has typically studied the choice between owning and renting in a utility maximizing framework where youths consume housing services and non-durables. Owning a house allows individuals to consume housing services and to finance future expenses. The drawback of owning housing services is that, in the event of job loss, selling the stock of housing and paying-off the mortgage that financed the investment is a costly process. The alternative of renting permits consuming housing services while allowing adjustment in the case of income fluctuations (see Chetty and Szeidl, 2007). Adding the possibility of living with parents to this framework is less common.

$P(stay_{t-1})$ is the probability that the worker stays at the firm at least until year $t-1$. $P(dismiss_t = 1|Stay_{t-1})$ is the probability of dismissal conditional on staying at least until year $t-1$ and is set at 2%. Finally, r is the discount rate and .125 is the yearly increase in firing costs, obtained by dividing 45 (the maximum lay-off cost in wage days per year worked) by 365. We express everything in terms of the current wage, implicitly assuming that labor costs increase at the same pace as productivity.

Setting the expected job duration at the sample mean of 15 years, and a small discount rate, r, of 0.5%, the expected dismissal costs are 8pp, and the expected present value of the subsidy would be 20-8=12pp. With an interest rate of 4pp, the net present value would be 16.2% of the current wage (=20-3.8).

[14]Following Manski and Straub (2000), it can be argued that employees whose employment relation is regulated by a permanent contract may perceive high risk of unemployment, in the sense that if they lose their job, there are few chances of finding a comparable one.

Consider the case of a youth who lives for two periods and her future income in period 2 is uncertain: she can be working either at the same job or unemployed. That is, we abstract from income fluctuations within the job. Let us assume first that housing services can be adjusted freely and that housing is not an investment. In that simple model, a higher risk of job loss leads consumers to diminish both their consumption of housing services and non-durables in the first period because the marginal utility of future wealth increases with the probability of job loss. That is the typical prediction of models with precautionary savings, and a higher degree of employment protection -by making layoffs less likely- expands the consumption of housing services in the first period.

A simple way of adding coresidence is to allow youths to stay with parents - possibly consuming a suboptimal amount of housing services from their point of view. In that case, the suboptimal amount of housing services consumed for free at parents' house acts as a floor. In such model, youths exposed to a higher risk of job loss would be more likely to stay with parents in the first period for two reasons: a precautionary motive to avoid large fluctuations in the marginal utility of wealth as in the case above and, in addition, coresiding permits consuming more housing services in the second period -due to the possibility to accumulate higher savings during the first period. So a higher degree of employment protection increases the fraction of renters (the only housing tenure considered in this simple model) and diminishes the fraction of coresidents in that model.

What would happen if housing services were costly to adjust? Fernandes et al (2008) present a two period model where youths choose between coresiding with parents or consuming housing services, and the housing decision of the first period is a commitment for the second period. While the relationship between employment risk and living arrangements is ambiguous, for most relevant cases, they find that increases in the risk of job loss leads to a higher prevalence of co-residence between parents and children.[15] Shore and Sinai (2010) model the existence of fixed costs of changing the consumption of housing services if the fall in income upon job loss is large enough -for example, liquidating a mortgage on the present house to move to a smaller dwelling. They provide simulations indicating that higher levels of risk lead consumers to prefer renting than owning in the first period because the monetary loss of adjusting in the case of realization of bad shocks is too costly.[16]

Dunn (2003) studies the link between the probability of job loss in a model where consumers choose between owning and renting. In her model housing is a store of value that is costly to adjust, while renting permits consuming frictionless housing services. In

[15] The ambiguity in the relationship between the child's job loss risk and living arrangements in Fernandes et al (2008) arises because altruistic parents may provide transfers to emancipated children. We abstract from those transfers.

[16] Those authors show that conditional on owning, risk may still increase the consumption of housing services. We do not examine the intensive margin of consuming housing services in this study.

her simulations, an increase in the probability of job loss leads households to postpone the decision to purchase a new dwelling. The reason is that home ownership in that model follows an S-s rule, where housing wealth depreciates until a threshold is reached and consumers purchase a new house. As housing investments are costly to reverse and must be financed with savings for a downpayment, in periods with a high risk of job loss liquid resources are specially valuable to support non-durable consumption. As a result, the level of housing wealth that triggers the purchase of a new house becomes higher -i.e., purchases are postponed. While she does not discuss the consequences of employment protection, her simulations imply that a decrease in the probability of job loss increases home ownership over renting -as funds set aside for precautionary purposes can be used for a downpayment. Dunn (2003) also documents youths react more to a decrease in the risk of job loss than old households -as the latter may have accumulated savings for other purposes already.[17]

In summary, most of the models reviewed here predict that, regardless of the existence of adjustment costs, decreases in the risk of job loss lead individuals to decrease co-residence and increase the consumption of housing services, either owned or rented. Some of the papers stress that when the risk of job loss falls, households resort preferably to purchase new accommodations -rather than renting them. Secondly, accumulated precautionary savings seem to play an important role in the decision to buy. Thus, following this review we test three hypotheses:

- Does lower employment protection decrease the consumption of housing services –either owned or rented?
- Does higher employment protection increase more the consumption of owner-occupied housing than the one of rented housing?
- Does lower employment protection increase the wealth holdings of marginal consumers (i.e., those living with parents)?

3 The identification strategy

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 impact of job loss risk on housing arrangements. For example, workers who are observed with a fixed-term contract are more likely to have been unemployed in the past, having depleted any accumulated wealth during prior unemployment spells. Hence, they have less resources for a downpayment and are more likely to rent or to live with parents. In that

[17] Crossley and Low (2011) make a similar point in the context of saving for life-cycle purposes.

case, the different housing choices across workers with different contract types mainly pick up different labor market histories.[18] We comment on those biases later on.

Our study exploits variation in contract type that is unrelated 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 fixed-term 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 employers faced different incentives to hire workers with a high severance payment contract. 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 fixed-term contract into an open-ended one.[19]

The effects of the risk of job loss on housing tenure and household formation are analyzed using two different estimation approaches. The first uses a stock sample of individuals and allows us to examine how risk of job loss 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 three years after being hired (the reduced-form or intention-to-treat).[20] The second approach estimates a duration model using a sample of multiple transitions to a new accommodation, distinguishing between owning and renting. That sample allows us to study the timing of decisions, in particular how impacts vary by the number of years elapsed since the job contract was signed.

The reason why we first show estimates from a stock sample of employees instead of focusing only on the sample of transitions to a new accommodation, which allow us to study the timing of decisions, is that the stock sample is useful to assess the validity and quality of our instrumental variable. For the sample of transition data, we cannot provide first-stage estimates of the impact of subsidies on job contract conversions along job spells, since we do not know the year of conversion.[21]

[18] Alternatively, employees in tight local labor markets may be more likely to be hired on an open-ended contract -due to firm's competition for scarce labor- and those better income prospects may lead workers to become home owners. The contrary would happen in less active labor markets. 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.

[19] 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.

[20] Throughout the paper, the first-stage and intention-to-treat estimates provided are computed for a reference person that has a job tenure of three years in the survey year.

[21] With transition data, we can only provide intention-to-treat estimates of the impact of subsidies for job contract conversions on the decision of moving to a new accommodation, emancipated from parents, along the individuals' job spell.

3.1 Estimates from a stock sample of individuals

To investigate if job security affects the housing tenure regime, we show causal evidence mainly from the 2002-2011 waves of the triennial Spanish Survey of Household Finances (in Spanish, *Encuesta Financiera de las Familias*, EFF) covering the period spanned between 2002 and 2012. This sample period covers an expansion of the labor and housing markets (2002-2008) as well as a severe crisis (2009-2011). In some analysis we also include the 2014 wave. As described below, the survey contains retrospective information about the year when the current labor spell started.

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 has to upgrade a contract into an open-ended one. As the source of variation depends on the year of hire, region, gender and age when the worker was hired, we include fixed effects of all those characteristics.

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

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 dependent variables of interest (Y_{it}) are four binary variables in this empirical strategy. In the first stage, Y_{it} is the indicator of whether the employee holds an open-ended contract. In the second stage (the reduced-form model) and in the instrumental variable estimates, Y_{it} consists of three binary variables that take the values 0 and 1 indicating whether the individual lives with their parents, whether he is a homeowner and whether he is a renter.[22]

The key variable identifying the risk of job loss is $Subsidy_{r,a,t_0}$, which measures the economic incentive a firm in region r and year t_0 faces to upgrade a fixed-term contract into an open-ended one for an individual with age a. We do not observe if the firm for which the young adult works actually got the subsidy, so we use the amount of the subsidy the firm was eligible for, presented in Table A.1. For workers covered by an open-ended

[22] We estimate three separate models for each one of these three outcomes (decision of living with parents vs not, decision of being a homeowner vs not and decision of renting vs not).

contract, we do not know when the contract was converted. However, previous studies have documented that (a) most job spells start with a fixed-term contract and (b) most conversions happen during the first two years of the job spell.[23] Hence, we assign the mean subsidy during the first two years of the match between firm and the employee. We test the latter assumption using administrative data from Social Security records -see the Appendix. We use a sample of EFF respondents that contains individuals hired between 1995 (two years prior to the introduction of regional subsidies to contract conversion for the first time in Spain) and 2009 (the last year for which the subsidies were collected).[24]

The parameter of interest in this specification, θ_1, measures the impact of 1000 euro subsidy to convert a fixed-term contract into an open-ended one on the probability of occurrence of the dependent variable. In the first stage, θ_1 is the marginal impact of 1,000 euro subsidy on the incidence of open-ended contracts. In the intention-to-treat specification, the corresponding parameter for θ_1 is the impact of 1000 euro subsidy on the probability of living with parents, owning or renting. The choice of the year-of-hire dummies (excluding 1999) and time dummies (excluding 2002) implies that the constant reflects choices three years after the job spell started -which is the mean tenure of our sample, as described below.[25] Hence, the thought experiment we make is whether the availability of a subsidy during the first two years of a job spell results in a higher fraction of workers with an open-ended contract and, subsequently, in a higher fraction of employees renting or owning their own accommodation.

We make three notes on the specification. The first is that our focus is the fraction of employees choosing their living arrangements, which can change at any point in time (renters can choose to become owners or viceversa). Thus, employees who have bought or rented their own accommodation before their current job spell started could have also changed their living arrangements during the period we analyze, so they must be included in the sample. Secondly, a concern is that employees may change locations in response to labor and housing market conditions. In principle, identification is achieved through changes in the incentive to hire that happen during the first two years of the current job spell, so determinants of mobility prior to that job spell should not affect the estimates unless they are correlated with the subsidy. However, we re-estimate the model on a panel of youths who originally live with parents. The location of such youths is based on their parents' decisions, rather than their own, so biases due to sorting are less likely in this sample. A final issue is the existence of other trends across regions or cohorts that affect

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

[24] The inclusion of two years prior to the introduction of the subsidies allows us to better estimated trends in regional housing and labor market conditions.

[25] This paper uses different time spans 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.

housing and labor markets. As subsidies to contract conversions vary in cells defined by age-at-hire, gender, region and year-of-hire, we introduce in the first stage and in the intention-to-treat specifications interactions between year-of-hire $\times$ region dummies (that absorb region-specific business cycles), age-at-hire $\times$ region dummies and age-at-hire $\times$ year of hire dummies (to absorb changes in cohort-level characteristics).[26]

3.2 Estimates from a sample of transitions

We use retrospective information on the year when the individual started their current job and the exact year when individuals moved to a new house living as a emancipated person (by owning or renting) to construct a duration model. 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 ($D_i = 2$). We assume that both risks are uncorrelated. 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 (and given that the individual has not moved previously during the job spell). This transition intensity follows a multinomial logit specification:

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

The hazard rate, $\theta\left[t \mid Z_i\right]$ 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 given that she has not moved previously during the job spell), and it is equal to the sum of both transition intensities, as follows:

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

The individual characteristics (Z_i) include regional subsidies ($Subsidy_{r,a,t_0}$), indicators of region, age at hire, year of hire and other time-invariant characteristics, such as the indicator of the individual's gender and education level. As calendar year dummies are

[26] One may also consider other changes in the welfare state that interact with the decision to form a household and the choice of rental vs owning. Anti-poverty programs are small in Spain, and target a very small fraction of the population. Tax incentives for housing are mainly established at national level.

perfectly collinear with dummies of year of hire and dummies of the yearly duration of the job spell, we control for the business cycle by including the national unemployment rate at year $(t_0 + t)$.

In this duration model, censored observations correspond with two kinds of individuals: firstly, individuals living with their parents and, secondly, individuals who 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 and has not moved previously during the job spell.[27]

4 Sample data used 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). 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. All the summary statistics shown in the paper are weighted to be representative of the Spanish population. The estimates of the empirical models are not weighted, instead.[28]

Before describing the samples, we give some background for housing choices in Spain. Figure 2 displays that the fraction of Spanish males who are renters or homeowners as a fraction of all males (either renting, owning or living with another person in a house they

[27] 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 exit rates conditional on not moving with the other alternative 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 among individuals that stay in their initial accommodation at least at t years of their job spell. For estimating competing-risk models, exits to 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 logit models.

[28] 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)].

neither own nor rent). We do this 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).

4.1 Stock sample of individuals

Using the EFF survey we construct a stock sample of household members between 25 and 64 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. To obtain a sample of individuals whose attachment to the labor market is strong, we exclude respondents that declare to be inactive (other than education), self-employed, or who do not contribute to Social Security. We also drop from 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). Finally, we exclude from the sample workers who 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.

Employment protection is measured by the kind of job contract the young workers hold, whether a open-ended contract or a fixed-term one. The variables of regional subsidies to contract conversion are expressed in thousand euros of 2005 using regional deflators of the gross household disposable income.[29] 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 an open-ended one during the first two years of the individuals' job tenure, since almost 75% of the contract conversions occurred in the first two years -we test this assumption below.

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

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

(parents-in-law) or other relatives (and in the latter case, the other relative is either the reference person or the spouse).[30]

Table 1 presents descriptive statistics of the main characteristics of the estimation sample. A third 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, 64% of the individuals hold an open-ended contract. Employees holding a fixed-term contract are 36 years of age, 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 - workers under an open-ended contract have mean job tenures of three 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 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%).

Concerning accommodation characteristics, the current value of the owner-occupied house represents the labor earnings of around 9 years a half on median for permanent workers and the median of 11 years for temporary workers. For indebted homeowners, the median loan-to-value (LTV) ratio at the moment of the home purchase (the initial value of the loans taken out for the purchase relative to the acquisition value of the house in %) was of around 96% for permanent workers and 100% for temporary workers. The share of earnings that households spend on accommodation (either rents or loan repayments for the owner-occupied house) is higher for temporary workers than the one for permanent workers. Renters devote a lower percentage of resources (31.2% vs 34.9% in the full sample).

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, Table 1 also presents a sample of individuals aged between 25 and 45, a slightly younger sample (33 years on average vs 36 in the full one). The full sample of individuals oversamples the youth population (the 78% of observations come from individuals aged between 25

[30] We drop out of sample those observations of individuals that are not the reference person (nor the spouse) and that live in the household with other household members that are not related in a family, such as individuals that live in a shared house, rent a room, etc. The reason is that we do not have information of the degree in which these individuals contribute to the household finances, whether they can be considered as emancipated persons or coresidents.

and 45), since we restrict our sample to individuals with short job tenures in order to link housing decisions with their risk of job loss. We are interested in studying not only household formation, but also the consumption of housing services of recently hired employees of the whole population that come from other housing transitions not related only to emancipation from parents' house.[31]

4.2 Analysis using transition data

The duration sample is formed by 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 sample includes all 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 either live with their parents, have acquired an owner-occupied house or have rented an accommodation before or after having started to work in the current job.[32] 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.

Table 2 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. 21% of individuals moved to a new accommodation as an emancipated person during their current job spell, 14% 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 do (65%). 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).

Descriptive statistics of transitions from young employees aged between 25 and 45 are very similar (see Table A.2), since the 77% of the transitions in the full sample come from the sample of young employees.

[31] In order to analyze the household formation decision of young individuals, we face very small sample sizes to obtain precise estimates if we narrow the age bands of young individuals under the age of 45, such as 40 or 35. Moreover, given the small sample, we are not able to make any robustness about gender.

[32] In the paper, when we refer to individuals living with their parents, we are also considering individuals living with their parents-in-law and other relatives (the latter case only when these relatives are the reference person of the survey or the spouse).

5 Results

This section presents the empirical results obtained using data from the 2002-2011 waves of the Spanish Survey of Household Finances (the 2014 wave is used in a later sample). Subsection 5.1 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 also use the subsample of employees aged between 25 and 45 in order to investigate if the patterns of household formation and housing tenure choice differ from those in the full population. Subsection 5.2 documents the empirical results obtained from the estimation of a duration model using retrospective information of event years. In all Tables, the standard errors of the estimated parameters shown in parentheses take into account that there can be group correlation in the error term within each region [see Moulton (1986)].[33]

5.1 Causal evidence from a stock sample of individuals

Before turning to the regression-based estimates, we illustrate the identification strategy in the first stage using a particular episode when Andalucia and Extremadura stopped granting subsidies between 2003 and 2004 (see Table A.1). These regions resumed granting subsidies in 2005. The rationale behind our estimation is that employees are typically hired using a fixed-term contract, that is either converted into an open-ended one at some point (typically in the first two years) or it is simply terminated at expiration. Hence, if subsidies to contract conversion were effective in hiring decisions, one would expect that the fraction of employees with an open-ended contract three years after the start of the job spell would fall in Andalucia and Extremadura among workers hired in 2002 (who could only benefit in the first year), 2003 (in neither of the first two years) or in 2004 (could benefit in the second year only). This would not happen in the rest of the regions that continued granting subsidies.

The upper graphs in Figure 3 plot the differential fraction of employees with an open-ended contract three years after being hired in the treatment regions (Andalucia and Extremadura) with respect to the control regions that continued providing subsidies (Madrid, Comunidad Valenciana, Galicia, Basque Country and La Rioja), separately for women and for eligible men (i.e. those not aged between 31 and 40). To benefit from large samples, Figure 3 uses a 4% sample of Social Security records, described in Appendix A. Among workers starting their job spells between 1999 and 2001, the differential share of employees with an open-ended contract three years after being hired was about 10%

[33] There are seventeen regions in Spain, which may be a too low number of groups for clustered standard errors to be asymptotically valid. We explore the issue by computing empirical standard errors simulating placebo subsidies at various levels, adapting Bertrand et al (2002).

lower in Andalucia and Extremadura (treatment regions) than the one in the control group regions, both evolving in a similar fashion. However, among employees hired in 2002, the probability of having an open-ended contract three years after fell by 7% in Andalucia and Extremadura, while it remained basically constant in the control regions. That gap remained in 2003 and 2004. To quantify the fall in terms of the subsidy amount, consider the case of females. Conversion of a fixed-term contract into an open-ended one was subsidized by 6012 euro (see Table A.1). A 7% fall in the fraction of employees with an open-ended contract divided by 6012 euro implies that 1,000 euro subsidy increases the fraction of workers under an open-ended contract by 1.2 percent ($0.012 = \frac{.07}{6.012}$).

The lower graphs in Figure 3 show the differential change of the fraction of employees aged 25-45 living with their parents or other relatives three years after having been hired in Andalucia and Extremadura with respect to the control group regions. The estimates are not significant due to small sample sizes we have using data from the Spanish Labor Force Survey (LFS).[34] The rate of coresidence of eligible male workers increases after 2002 (around 7%) in comparison with their counterpart demographic group in control regions. However, for women, we observe a long-run decrease in the incidence of coresidence of women in Andalucia and Extremadura relative to women in the control group regions for the whole period analyzed.

First-stage estimates and intention-to-treat estimates Table 3 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 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 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 (ITT) estimates rely on less stringent assumptions than IV. For example, assume that the increase in the potential demand for housing altered the relative price of rents vs real estate. In that case, ITT estimates would still measure the overall response of housing decisions to increases in job security and the general equilibrium response of prices. On the contrary, IV estimates require that subsidies affect the consumption of housing services only through the increase in the stock of employees with

[34] The LFS is the data source that provide us with the largest sample size to analyze the rate of coresidence of young individuals in this particular event study. This study cannot be accomplished using data from the EFF.

open-ended contracts.[35] 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 3 shows the Ordinary Least Square (OLS) estimates of the impact on the probability of holding an open-ended contract of the mean subsidy in the first two years of the worker's job tenure that the firm can benefit from. An increase of 1,000 euro in the subsidy to convert fixed-term contracts into open-ended ones increases the fraction of employees with an open-ended contract by .008 (standard error .0024) in the full sample (Table 3, Panel A, row 1, column 1). In the sample of respondents aged 25-45 the corresponding increase is .006 (Table 3, Panel A, row 1, column 2). Both estimates are statistically significant at the 1% and 5% level of significance, respectively. 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 lower in the sample of young individuals, perhaps due to the small sample size. Appendix A replicates the analysis using administrative Social Security data in the youth sample, and report similar -but more precise- estimates to those in Table 3 column 2 (see Tables A.3 and A.4).

How large is the size of the first stage? The estimates imply that an increase of 1,000€ in the subsidy raises the stock of permanent workers by 0.6%-0.8% depending on the age of the worker (the weighted sample average is 64% of open-ended contracts). To put these numbers in perspective, the mean subsidy is about 4,000 euro per contract converted. Our simulations imply that such average subsidy amounts to between 12% and 16% of the labor cost of hiring a worker with an open-ended contract, once one takes into account the associated increase in severance payments. So 1,000 euro would represent a reduction in labor costs of 4% ($\frac{1,000}{4,000} \cdot 0.16 = 0.04$). The first-stage estimates imply an elasticity of the percentage of open-ended contracts to labor costs of .31 ($.31 = \frac{0.008}{0.64 \cdot 0.04}$). An elasticity of .3 is in line with estimates of the demand for labor in Spain and other European economies.[36]

The response of living arrangements

Panel B of Table 3 shows the reduced-form estimates of the probability of the decision of living with parents and the housing tenure regime. An increase of 1,000€ in regional subsidies decreases significantly the probability of coresidence by 1% in the total sample of individuals aged 25-64 -see Table 3, Panel B, row 1a column 1. We note that the 1% fall in the fraction of coresidents in the full sample mirrors the increase in the share

[35] Furthermore, ITT estimates are much more precise than IV estimates. This is particularly relevant in our sample of young household members aged between 25-45 whose job tenure is not longer than 7, where the minimum sample size is 3,974 in one of the five data sets imputed multiply.

[36] Kugler et al (2005) report elasticities of the demand for labor of .2 for unemployed Spanish workers below 30 years of age. Saez et al (2017) report elasticities in Sweden of about .2 as well.

of workers with open-ended contracts documented in Table 3, Panel A, row 1, column 1. Thus, at face value, it seems that over the three years of tenure we consider, there is a close correspondence between contract conversions and the fall in the proportion of coresidents. It can be argued that most -if not all- individuals end up forming their own household, so the relevant margin is how long is that decision anticipated because of employment protection. We discuss the issue in the context of the duration model below. For now, we note that models of precautionary savings predict that increased employment protection increases the demand for housing services, as Table 3 suggests.

When we consider the particular form of housing services consumed, an increase of 1,000 euro in the incentive to convert a fixed-term contract into an open-ended one increases the fraction of home-owners by .44 percent (standard error: .22), while renting increases by .57 percent (standard error: .16 percent) -see Table 3, Panel B, rows 2a and 3a, respectively. Taking into account the weighted mean of ownership in the full 25-64 age sample, the estimate implies that an increase of 1,000 euro in the subsidy to convert a fixed-term contract into an open-ended one increases the probability of being a home-owner by 0.8% ($\frac{0.004}{0.507} \cdot 100$). However, the rate at which the probability of renting increases is four times as large, 3.7% ($\frac{0.006}{0.161} \cdot 100$). These figures are more precisely estimated for the decision of rental than for home ownership. The estimates are similar, but larger in magnitude in the sample of employees in ages 25-45.

Robustness checks 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) (i.e., including sequentially $\mu_{t_0} \times \mu_a$, $\mu_r \times \mu_a$ and $\mu_r \times \mu_{t_0}$) -see Bleemer et al (2017) for a discussion on the roles of trends in local labor markets. Tables 4 shows estimates of the impact of regional subsidies on the stock of permanent workers and on housing demand in the full sample. One can view our tests as an indication of whether the estimates are affected by cohort-specific trends (the interaction of year and age at hire), cohort-region specific trends ($\mu_r \times \mu_a$) or region-specific business cycles ($\mu_r \times \mu_{t_0}$).

The estimates of the first stage are stable across specifications, indicating that a 1,000€ subsidy to contract conversion increases the stock of workers with an open-ended contract by between .8 percent and 1.1 percent -the latter estimate comes from a specification that controls for region $\times$ age at hire, region $\times$ year of hire, and age-at-hire $\times$ year of hire dummies.[37] The impact of 1,000 extra euro in subsidies to contract conversion increases

[37] In addition, each column presents the p-value of a test that examines if the residuals of our baseline specification (presented in column 1) correlate with the interactions of fixed effects. For most specifications, the residuals from our main specification are uncorrelated with the fixed-effect pairs. In those cases, in which these second-order fixed effects are significant, the estimates do not vary greatly. The exception is the estimated impact on home-ownership, which diminishes substantially when we add all three pairs of fixed-effects.

the probability of forming a new household between 1 percent and .5 percent (the latter case in a specification that includes the full set of second-order interactions). Regarding the exact form of household formation, and in keeping with the previous results, subsidies to contract conversion mainly increase the probability of renting a new accommodation -compare columns 1-4 in Table 4, panel B to those in columns 5-8.

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 200 independent 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). Then we reestimate our empirical models with these placebo subsidies to construct properly empirical confidence intervals that tests the hypothesis of a zero impact of subsidies on the outcome variables. The empirical confidence intervals computed by the randomization inference approach at the 1% level of significance shown in Table A.5. show that the first-stage and intention-to-treat estimates in Table 3 almost always lie outside the empirical confidence intervals at the 1% level or at the 5% level (the latter not shown here). We interpret that our instrumental variable, the regional subsidies, are not capturing other unobserved trends not taken into account in the analysis and potentially correlated with regions, age at hire, gender and year of hire.[38]

Instrumental variable estimates Table 5 shows estimates of the impact of holding an open-ended contract on the probability of each housing tenure regime for employees between 25 and 64 years of age in columns (1) and (3). The first two columns show suggestive evidence of this impact from Ordinary Least Square estimates (OLS). Workers holding an open-ended contract have a probability of being observed living in a rented accommodation 0.5% and 0.7% lower than that among temporary workers in the full sample and in the young sample, respectively. On the contrary, workers under an open-ended contract are 0.4% more likely to be observed living in an owner-occupied house than temporary workers. Those results are broadly in line with previous literature estimating that decreases in the probability of job loss increases home-ownership at the expense of rentals, while household formation can be safely ignored.

However, the results change when one uses instrumental variable estimates (two-step least square estimates, TSLS). The estimates in Column 3 of Table 5 show that an exogenous increase in job security due to the conversion of a fixed-term contract into an

[38] This happens except for the confidence intervals computed using placebo subsidies assigned randomly across regions and year of hire at the same time, where the estimates are near the bounds of the intervals. However, Table 4 shows that the estimates of the impact of subsidies when we consider second-order fixed effects of region and year of hire are very similar to our baseline estimates in Table 3 (the estimates are inside the confidence intervals, what gives evidence for similar estimates at the 5% significance level).

open-ended one makes employees more likely to form a household by owning their own accommodation and by renting it -at the expense of the fraction who lives with parents 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 1.25%, and our estimates cannot reject a one-for-one response. TSLS estimates suggest that the bulk of the response comes from renting (1% extra open-ended contracts increase the fraction of renters by .7%). The impact on home ownership is imprecisely estimated. Given the low F statistic of the first stage, we do not present TSLS estimates for the sample 25-45.[39]

An interpretation of the discrepancy between OLS and TSLS estimates is that, relative to employees covered by secure contracts, workers under fixed-term ones may have been more able to experience unemployment spells during which buffer stocks have been depleted. Thus, their lower propensity to own a house is the result of an inability to accumulate funds towards a downpayment. Once we compare workers who have had an open-ended contract because of higher incentives for employers, differences in labor market histories and ability to accumulate wealth may disappear.

Linking our results to previous literature on housing tenure, authors like Diaz-Serrano (2005) and Gathergood (2011) among others has emphasized that risk of job loss leads to lower home ownership rates. However, these empirical studies do not take into account two relevant issues in their analysis. Firstly, these empirical papers do not measure of the causal impact of the risk of job loss on housing demand, since they do not exploit exogenous variations in job security to estimate the causal link between the risk of job loss and housing tenure choice. As shown above, we find very different results between suggestive evidence from OLS and causal impacts estimated by instrumental variables. Secondly,.previous literature on the impact of the risk of job loss on housing tenure does not deal with the endogeneity and the self-selection problems that arise from ignoring the decision of household formation, the decision of housing tenure of individuals that are still leaving with their parents. When we disregard coresidents, OLS estimates using a sample of households already formed (estimates not shown here) provide similar results, such as those found by Diaz-Serrano (2005) and Gathergood (2011), that job security deters households already formed from renting. Once we take into account casuality and the endogeneity of household formation, this result is reverted, that is, exogeneous increases of job security encourage not only home ownership but also renting a new accommodation through the household formation channel.

[39] 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.

Magnitude of the cross-section estimates

When we compare our results with those obtained in previous papers, our estimated impacts are much higher. Bleemer et al (2017) find that an increase of 1% in unemployment in the county increases the prevalence of coresidence among 23-25 year olds by .05%. Becker et al. (2010) document that an increase of 10% in perceived of job security has a contemporaneous effect on the probability of forming a new household of 3.4%. 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 1). Hence, an increase of 10% in the perception of job insecurity rises the probability of leaving parental home in 30% $0.31 = \frac{1.245}{40}10$). This discrepancy of results may arise because of several reasons. Firstly, the measures of risk differ.[40] Secondly, 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. As we discuss below, the response of living arrangements to changes in employment protection happens with a delay, so the time window matters. The next section studies the timing and the dynamics 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.[41]

5.2 Evidence from a sample of transitions

Next, we study the exact year of the job spell in which individuals decide to move to a new accommodation as an emancipated person by estimating competing-risk models considering two alternative exits, a move to an owner-occupied house and a move to a rented house. Table 6 shows the estimates of how the hazard 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 hazard of moving and specification (ii) allows for a different effect of subsidies that varies along the time elapsed since the job contract was signed. The estimates shown are the latent coefficients of the different Logit models, so we postpone interpretation of the magnitude for now and focus on the signs. Finally,

[40] Bleemer et al (2017) use the county-level unemployment rate, which may not coincide with the perception of job loss risk of a particular individual. For example, Becker et al (2010) find that coresidence decisions in Italy respond more to changes in the fraction of workers under a fixed-term contract in the province than to changes in the unemployment rate. Becker et al's aggregate analysis uses qualitative answers like "a lot" to the question "During the last year, have you (or someone in your household) worried about losing a job or not finding a job?". Qualitative measures differ from the measures we present (as the meaning of "a lot" may be different for different respondents).

[41] García-Ferreira and Villanueva (2007) do not find 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.

the coefficient of the variable $Subsidy_{r,a,t_0}$ in row 1, Column 4 of Table 6 reflects the impact of 1,000 euro of subsidy to contract conversion on the hazard of moving to a new accommodation during the third year of the job spell (the omitted duration dummy in the estimates). To obtain the corresponding estimate during the first year, for example, one must add the estimates in row 1 and 2 in Table 6, Column 4.

Columns (1) and (4) in Table 6 consider a move to a new house (irrespective of the housing tenure regime), and concerning housing tenure columns (2) and (3) and columns (5) and (6) show competing-risk model estimates for each conditional exit, a move through home ownership or a move to a rented house, respectively. 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.

Once we allow for time-varying effects of subsidies in column (4) in Table 6, we note that the coefficient estimate of the interaction of the subsidy with the first two years of job tenure is negative. 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 happen. In other words, there is a delayed effect of increased job security on household formation, most noticeable among renters. The tenure profile of becoming a home owner is flat, instead.

The upper graph in Figure 4 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 conditional hazard rates, the probability of moving to each tenure regime in each year of the job spell conditional on individuals that have not moved previously in their job spell and whose job lasts at least that number of years. The vertical line indicates the second year of the job spell, when the contract conversion normally occurs.[42] The full black line shows the marginal effects on the probability of renting a new accommodation. The decision of renting a house lags 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. The delay in the response of renting is somewhat surprising as search for a rental place is not very costly -and certainly less costly than owning.

[42] Table A.3 uses administrative data from Social Security records to document that the impact of subsidies on the stock of open-ended contracts stabilizes after the second year of the tenure relationship.

The lower graph in Figure 4 shows the marginal effects implied by the estimates of the duration data model in the sample of employees between 25 and 45 (see the model estimates in Table A.6 of Appendix). 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 salient in this sample. Within the sample of individuals aged 25-45, 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, while it does not reach .002 in the full sample.

Overall, we draw three conclusions from this analysis. The first is that both the probability of purchasing a new house and of renting a new accommodation increase with employment protection. The impact of 1,000 euro subsidy to contract conversion on home ownership is rather similar in the full sample and in the 25-45 sample. The second is that renting is more responsive to increases in job security in the youth sample than in the overall one. This is contrary to the prediction in Dunn (2003), who simulates a life cycle model augmented with uncertainty and finds that the decision to buy a house is more responsive to increases in job security among youths than among mature individuals. Finally, in both samples, the response of renting to increases in job security lags contract conversions in both the full sample and in the 25-45 sample. The delay in household formation to increases in job security helps to reconcile our estimates with the smaller impacts in studies looking at contemporaneous effects.

Magnitude of the estimates

To assess the magnitude of the estimates, we compute by how much individuals anticipate the consumption of housing services because of the presence of the subsidies. To that end, we use the estimates in Table 6 and reconstruct the probability of either renting or owning of an individual with average characteristics with and without the average subsidy to contract conversion of 4000 euro. In the absence of the subsidy, and assuming that the worker lives with parents at the beginning of the job spell, she would either purchase or rent a new accommodation with probability of 20% at the end of the 7th year of the job spell. The same individual working in a firm that could benefit from the mean subsidy of 4000 euro would purchase or rent a new accommodation with 20% probability by the end of the 5th year of the job spell. So according to that measure, the mean subsidy to contract conversion advances the consumption of new housing services by 2 years.[43]

Wald estimates

Tentatively, we can construct an informal Wald estimator of the causal impact of conversions from fixed-term contracts into open-ended ones on household formation using

[43] To compute these numbers, we take the estimates of the hazard rate in Table 6, column 4. We compute the hazard as $\frac{\exp[\beta_0+\beta_t 1(\text{job tenure}=t)+\beta_1 Subsidy_{r,a,t_0}+\beta_{1,t} Subsidy_{r,a,t_0}\cdot 1(\text{job tenure}=t)]}{1+\exp[\beta_0+\beta_t 1(\text{job tenure}=t)+\beta_1 Subsidy_{r,a,t_0}+\beta_{1,t} Subsidy_{r,a,t_0}\cdot 1(\text{job tenure}=t)]}$. We calculate the estimated survival probabilities and compare the resulting probability of not consuming housing services with the same amount when we assume a 4,000 euro subsidy.

the estimates of Table A.4, although the sample population in Table 6 (Table A.6), i.e. a sample of transition data of (permanent and temporary) employees, is not the same as that in Table A.4 (a sample of job spells until job contract conversions of fixed-term workers occur). The estimates from specification (ii) in Table A.4 imply that the probability of conversion in the first five years of job tenure is 59.6% for a temporary worker whose firm can benefit from a 4,000 euro subsidy. This probability diminishes to 53.7% for a non-eligible worker, i.e. the probability of job contract conversion increases in 5.9% for the reference worker being eligible for the mean subsidy of 4,000€. Similarly, Table A.6 estimates that the probability of moving in the first five years of their job tenure is 5.1% higher for the reference worker being eligible for the mean subsidy (2.3% higher in the case of moving to an owned dwelling and 3.4% higher through rental). These estimated impacts seem to be in line with those estimated using the stock sample of employees.

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, shown in Tables A.7 and Tables A.8. The results are similar to those in Tables 4 and A.5 and we do not comment them in detail.

5.3 Possible explanations of the underlying mechanisms

Wealth levels of youths Table 7 provides further information about the wealth levels of affected youths. While we cannot observe the wealth of the particular individual who leaves the household, the panel component allows us to observe the evolution of the household wealth once the individual has left the sample. Hence, we regress the log change in household wealth (Panel A) and an indicator of whether an individual 25-45 has established a new household between waves (in Panel B) on the same set of regressors as in Table 3. We include wave 2014 in this analysis. If youths affected by the subsidies had accumulated a substantial amount of wealth and had taken it with them as they formed a new household, we would expect a negative estimate of the variable $Subsidy_{r,a,t_0}$ on the growth of wealth between waves. Panel B of Table 7 indeed finds that subsidies predict the event of an employee living with parents leaves the household, and the estimate, shown in row 1, Panel B of Table 7 is very similar in size to that in Table 3, Panel A, row 1. Thus, we do observe that subsidies explain household formation in this sample.[44] However, using an array of wealth measures the incentives to hire using an open-ended contract have little or no predictive power in the evolution of household wealth (median household wealth even increases slightly in cases where youths leave the household). Our

[44]Incidentally, note that the magnitude is similar to that in Table 3. As this is a sample of youths living with parents, it is unlikely to be affected by endogenous mobility of the youth in reaction to the subsidies. That similarity gives us confidence that our estimates are not driven by selection.

interpretation is that youths who experiment an increase in job security through these subsidies hold modest wealth levels to start with, and are thus unlikely to move to an owner-occupied dwelling.

Presence of credit constraints One possible explanation of the delay in rental documented in Section 5.2 is that renters are discouraged homeowners, i.e. individuals that wished to form a household by buying their own accommodation but could not do it due to credit constraints. In order to test that hypothesis, we consider a definition of liquidity constraints similar to that used by Jappelli (1990). 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. We estimate a multinomial logit model with three categories: credit constrained, having asked for a loan fully accepted and the omitted category, the choice of not having asked for any loans for the last two years because they were not needed.

Panel A of Table 8 shows estimates of the multinomial logit model of credit constraints using the same set of covariates as in Table 3 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 the existence of regional subsidies to contract conversion. Furthermore, when we interact the outcome "credit constrained" with the different forms of housing tenure in Panel B of Table 8 we do not find evidence that renters are more likely to be credit constrained than homeowners are.

Decision of living with a partner in a household emancipated from parents An alternative explanation of the delay in rental may be associated with individuals that wish to live with a partner, which may lead to coordination issues in the couple (like, waiting until both obtain an open-ended contract -see Bentolila et al, 2017). While we do not have information of the exact year in which individuals got married or started to cohabit with a partner, we can use the stock sample of individuals to 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).

Panels A of Table 9 shows the estimates in the sample of individuals aged between 25 and 64 and panel B does for the sample of individuals aged between 25 and 45. 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 Table 3. Table 9 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 emancipated.

Column (1) in Panel B of Table 9 indicates that higher incentives to convert fixed-term contracts into open-ended ones increase the fraction of individuals living with a partner among 25-45 year-olds. Namely, an increase of 1,000 euro in the incentive to hire using open-ended contracts increases the probability of individuals who rent and live with a partner by 0.45%. Interestingly, we do not find a similar increase among owners who live with a partner. In other words, the increase in the fraction of individuals who live with their partner is only observed for renters -the group for whom consumption of housing services lags behind increases in job security. Hence, the need to coordinate both members might be behind the delay in undertaking the decision of forming a household by renting.

6 Conclusions

This study exploits regional and time variation between 1997 and 2009 in the incentives to hire workers using high employment protection labor contracts to understand how risk of job loss affects housing decisions. 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 incentives to strengthen employees' employment protection increase the fraction of workers covered by a high severance payment contract by about .8% and diminishes the fraction of individuals living with parents by a similar amount three years after the beginning of a job spell. Those estimates suggest a one-to-one response between obtaining a high employment protection contract and moving to a new accommodation, most of the increase is due to increases in the fraction of renters.

Those results can be only partially explained by the existence of precautionary savings: while we find evidence that higher employment protection anticipates the consumption of housing services, we do not find that youths are specially more likely to purchase a house when job security increases. A possible reason for the lack of a response is that young individuals living with their parents do not accumulate much savings while exposed to job loss risk. We discarded that the higher response of renting is due to the presence of borrowing constraints. Instead, it seems that a significant factor for forming a new household through rental is the wish to get married or to cohabit with a partner, and this is not observed in the case of home ownership. A recent literature has stressed that mandated severance payments increase welfare by reducing the need of precautionary saving (see Cozzi and Fella, 2016). Our results support the notion that lower severance payments result in significant delays in housing consumption and in lifetime decisions like forming a new household.

Appendix A: Evidence from Administrative records

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 administrative 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 of individuals with any record in the Social Security Administration. 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.3. 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.3 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 happen. 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. Most 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 (2) in Table 3). Column (5) of Table A.3. 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 0.0057, 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.

Alternatively, the impact of subsidies for job-contract conversions on the incidence of permanent workers can be analyzed looking at their impact on conversions from fixed-term contracts into open-ended ones in the firm where the employees work. As in Section 5, we focus in the first ten years of job tenure in the duration model of conversions. Table A.4 shows the estimates of the hazard rates using alternative specifications of Logit duration models for a sample of yearly transition durations until the conversion of the fixed-term job contract into open-ended one occurs in the firm where the temporary employee aged between 25 and 45 works. Censored duration observations correspond with employees that continue holding a fixed-term contract along their job spell analyzed and with individuals that leave the firm (in situations such as entering unemployment or a non-economic activity, change of job, etc.). The model covariates are similar to those considered in the study of the transition data of household formation and housing tenure choice in Section 5.

Table A.4 shows that the coefficient estimates associated with the subsidy variable are positive (0.0392 in column 1) and statistically significant at the 1% significance level. The estimates in column (1) imply that an increase of 1,000 euro on the average subsidy raises the likelihood of job contract conversions in 0.5% when we consider that the effect of subsidies is constant across years of job tenure. When we allow the impact of subsidies to be time varying along the job spell (in column (2)), the predicted probability of conversion for a worker being eligible for the mean subsidy of 4,000€ is 5.3% in the first year, 28.7% in the second and 11% in the third. The corresponding predicted probabilities of conversion for an identical worker not eligible for benefitting from any subsidies are 4.5% in the first year, 27.6% in the second year and 8.9% in the third.

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Table 1: Summary statistics of the stock sample of employees

	Age 25-64			Age 25-45		
		By type of job contract			By type of job contract	
	Total sample	Open-ended	Fixed-term	Total sample	Open-ended	Fixed-term
Holding an open-ended contract	0.64			0.64		
Holding a fixed-term contract	0.36			0.36		
Age	36.35	36.45	36.19	33.30	33.38	33.16
	(8.69)	(8.69)	(8.69)	(5.56)	(5.51)	(5.64)
Age at hire	33.39	32.97	34.11	30.37	29.94	31.11
	(8.77)	(8.77)	(8.72)	(5.75)	(5.69)	(5.76)
No. of years at current job	2.96	3.47	2.07	2.93	3.44	2.05
	(1.84)	(1.71)	(1.70)	(1.83)	(1.72)	(1.67)
Individual eligible for subsidy	0.56	0.56	0.56	0.57	0.563	0.570
Amount individual was eligible for	1.90	1.99	1.73	1.94	2.03	1.77
(include zeroes)	(2.42)	(2.50)	(2.25)	(2.45)	(2.53)	(2.30)
Individual labor earnings:						
Mean	14.70	16.32	11.87	14.84	16.52	11.90
	(9.52)	(10.46)	(6.71)	(9.37)	(10.20)	(6.77)
Median	12.81	14.10	10.87	13.02	14.23	10.85
Attained education level:						
Primary education or less	0.15	0.11	0.22	0.11	0.08	0.17
First stage of secondary educ.	0.46	0.45	0.47	0.45	0.43	0.49
Second stage of secondary educ.	0.15	0.17	0.12	0.16	0.19	0.13
Tertiary education	0.24	0.27	0.19	0.27	0.31	0.21
Male	0.56	0.58	0.53	0.56	0.58	0.52
Housing tenure regime:						
Living with parents	0.33	0.31	0.36	0.37	0.35	0.40
Home ownership	0.51	0.55	0.43	0.47	0.51	0.39
Rental	0.16	0.14	0.21	0.16	0.14	0.21
Median value conditional on corresponding tenure:						
Home value to individual labor earnings	10.04	9.65	11.23	10.00	9.51	11.23
Loan-to-value (LTV) at home purchase (%)	97.87	95.58	100.00	97.70	96.24	100.00
Yearly loan payments to labor earnings (%)	34.89	32.31	41.27	34.59	32.31	41.99
Ratio(%) of rent to individual labor earnings	31.24	27.00	33.74	32.14	28.63	35.16
Minimum sample size	5,087	3,408	1,678	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 5,087 household-year observations in each one of the five datasets imputed multiply in the four waves of the EFF data. Sample of employees with at most seven years of tenure on the job. 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: Individual characteristics in the sample of job spells of individuals aged 25-64.

	Total Sample	Subsample conditional on: Moving to a new house	Individual moves to: Owned house	Individual moves to: Rented house
Move to an owned house	0.14	0.67		
Move to a rented house	0.07	0.33		
Holding an open-ended contract	0.68	0.78	0.85	0.65
Holding a fixed-term contract	0.32	0.22	0.15	0.35
Age	36.75	35.56	35.86	34.96
	(8.73)	(6.75)	(6.80)	(6.61)
Age at hire	32.86	30.17	29.95	30.62
	(8.75)	(6.77)	(6.80)	(6.69)
No. of years at current job	3.89	5.39	5.91	4.34
	(2.55)	(2.38)	(2.12)	(2.52)
Individual eligible for subsidy	0.57	0.62	0.61	0.63
Amount individual was eligible for	2.00	2.49	2.44	2.60
(include zeroes)	(2.54)	(2.83)	(2.71)	(3.07)
Individual labor earnings:				
Mean	15.22	17.72	18.78	15.55
	(9.84)	(11.07)	(11.73)	(9.21)
Attained education level:				
Primary education or less	0.15	0.09	0.07	0.13
First stage of secondary educ.	0.45	0.41	0.38	0.48
Second stage of secondary educ.	0.16	0.19	0.21	0.16
Tertiary education	0.24	0.30	0.34	0.23
Male	0.56	0.56	0.55	0.60
No. of years elapsed from the job spell until a move occurs or a job spell ends (% cases):				
One year	12.88	27.72	27.85	27.45
Two years	16.69	22.54	22.30	23.03
Three years	16.51	13.96	13.53	14.86
Four years	13.76	13.85	14.75	12.03
Five years	11.58	9.79	10.99	7.36
Six years	9.51	5.51	5.29	5.97
Seven years	7.00	4.36	3.68	5.75
Eight years or more	12.07	2.26	1.63	3.56
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 3: 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 64 years	25 and 45 years
	(1)	(2)
1. Subsidy to contract conversion	0.0082	0.0062
(standard error)	(.0024)***	(.0026)**
2. Constant	0.554	0.553
(standard error)	(.060)***	(.049)***
F-test of instruments	11.21	5.96
Minimum sample size	5,087	3,974
PANEL B: INTENTION-TO-TREAT ESTIMATES		
Sample:	Individuals aged between:	
	25 and 64 years	25 and 45 years
Dependent variable (in italics below):	(1)	(2)
1. The individual lives with parents:		
a. Subsidy to contract conversion	-0.0102	-0.0131
(standard error)	(.0018)***	(.0022)***
b. Constant	0.228	0.304
(standard error)	(.041)***	(.052)***
2. The individual is a homeowner:		
a. Subsidy to contract conversion	0.0044	0.0058
(standard error)	(.0022)*	(.0024)**
b. Constant	0.534	0.445
(standard error)	(.034)***	(.039)***
3. The individual is a renter:		
a. Subsidy to contract conversion	0.0057	0.0073
(standard error)	(.0016)***	(.0030)**
b. Constant	0.238	0.251
(standard error)	(.043)***	(.059)***
Minimum sample size	5,087	3,974

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 of 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 4: Fixed-effect estimates of the impact of regional subsidies on the stock of open-ended contracts and living arrangements

Sample: Individuals aged between 25 and 64, broken down by different lengths of job tenure
Estimation method: Ordinary Least Square estimates

Panel A: First-stage Estimates (FS) and Intention-to-Treat Estimates (ITT)

	First-stage estimates (FS)				Intention-to-treat estimates (ITT)			
Dependent variable: Indicator of:	Open-ended contract				Living with parents			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Subsidy to contract conversion	0.0082	0.0082	0.0083	0.0108	-0.0102	-0.0098	-0.0097	-0.0054
	(.0024)**	(.0026)***	(.0027)**	(.0037)***	(.0018)***	(.0020)***	(.0022)***	(.0029)*
Minimum R-squared	0.22	0.22	0.23	0.25	0.29	0.30	0.309	0.32
F-test of instruments	11.21	10.08	9.41	8.50	–	–	–	–
P-value of joint significance on residuals	–	0.94	0.81	0.72	–	0.03	0.37	0.81
Minimum sample size	5087	5087	5087	5087	5087	5087	5087	5087
(Age at hire × year of hire) dummies	No	Yes	Yes	Yes	No	Yes	Yes	Yes
(Region × age at hire) dummies	No	No	Yes	Yes	No	No	Yes	Yes
(Region × year of hire) dummies	No	No	No	Yes	No	No	No	Yes

Source: The sample is formed by all household members aged between 25 and 64 years, being an employee and 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 those used in Table 3.

Standard errors (in parentheses) are clustered at the region level. Estimates combined across five implicates to take into account the uncertainty about imputation. The symbols *, ** and *** denote the estimate is significant at the 10% of significance level, at the 5% and at the 1%, respectively.

Table 4: Fixed-effect estimates of the impact of regional subsidies on the stock of open-ended contracts and living arrangements (cont.)

Sample: Individuals aged between 25 and 64, broken down by different lengths of job tenure								
Estimation method: Ordinary Least Square estimates								
Panel B: Intention-to-Treat Estimates (ITT)								
	Intention-to-treat estimates (ITT)							
Dependent variable: Indicator of:	Homeowner				Renter			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Subsidy to contract conversion	0.0044	0.0042	0.0041	0.0013	0.0057	0.0056	0.0056	0.0041
	(.0022)*	(.0026)	(.0027)	(.0034)	(.0016)***	(.0018)***	(.0019)***	(.0022)*
Minimum R-squared	0.23	0.246	0.25	0.27	0.04	0.05	0.06	0.08
P-value of joint significance on residuals	–	0.03	1.00	0.99	–	0.87	0.32	0.51
Minimum sample size	5087	5087	5087	5087	5087	5087	5087	5087
(Age at hire × year of hire) dummies	No	Yes	Yes	Yes	No	Yes	Yes	Yes
(Region × age at hire) dummies	No	No	Yes	Yes	No	No	Yes	Yes
(Region × year of hire) dummies	No	No	No	Yes	No	No	No	Yes

Source: The sample is formed by all household members aged between 25 and 64 years, being employees with a job tenure not longer than 7 and 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 those used in Table 3.

Standard errors (in parentheses) are clustered at the region level. Estimates combined across five implicates to take into account the uncertainty about imputation. The symbols *, ** and *** denote the estimate is significant at the 10% of significance level, at the 5% and at the 1%, respectively.

Table 5: The average effect of job contract on the decision of housing tenure

Estimation method: Ordinary Least Squares (OLS) and Two-Step Least Squares (TSLS) estimates			
	OLS estimates		TSLS estimates
Sample:	25-64 years	25-45 years	25-64 years
Dependent variable (in italics below):	(1)	(2)	(3)
1. The individual lives with parents:			
a. Open-ended contract	0.018	0.029	-1.245
(standard error)	(.021)	(0.026)	(.506)***
Robust confidence interval	–	–	[-4.433, -.471]
at the confidence level of:	–	–	0.95
b. Constant	0.163	0.219	0.918
(standard error)	(.041)***	(.054)***	(.373)***
2. The individual is a homeowner:			
a. Open-ended contract	0.035	0.037	0.543
(standard error)	(.021)	(.021)*	(.327)*
Robust confidence interval	–	–	[-.066, 1.69]
at the confidence level of:	–	–	0.9
b. Constant	0.537	0.453	0.233
(standard error)	(.030)***	(.036)***	(.224)
3. The individual is a renter:			
a. Open-ended contract	-0.053	-0.066	0.701
(standard error)	(.017)***	(.018)***	(.337)**
Robust confidence interval	–	–	[.089, 2.747]
at the confidence level of:	–	–	0.95
b. Constant	0.300	0.328	-0.151
(standard error)	(.046)***	(.062)***	(.243)
F test of instruments in first-stage	–	–	11.21
Minimum sample size	5087	3974	5087

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 3. Other covariates not shown are the following: indicators of age at hire, indicators of year of 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 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

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)
1. Subsidy to contract conversion	0.064	0.065	0.063	0.089	0.050	0.132
	(.011)***	(.016)***	(.019)***	(0.018)***	(0.024)**	(0.023)***
2. Subsidy * 1 year job tenure	–	–	–	-0.037	0.047	-0.095
				(0.016)**	(0.026)	(0.046)**
3. Subsidy * 2 years job tenure	–	–	–	-0.023	0.028	-0.093
				(0.020)	(0.029)	(0.048)**
4. Subsidy * 4 years job tenure	–	–	–	-0.002	0.035	-0.016
				(0.015)	(0.020)	(0.047)
5. Subsidy * 5 years job tenure	–	–	–	-0.053	-0.034	-0.080
				(0.024)**	(0.037)	(0.054)
6. Subsidy * 6 years job tenure	–	–	–	-0.053	-0.041	-0.213
				(0.041)	(0.036)	(0.086)***
7. Subsidy * 7 or more years of job tenure	–	–	–	-0.025	-0.753	-0.117
				(0.045)	(0.051)	(0.042)***
One year of job tenure	0.428	0.499	0.237	0.557	0.451	0.595
	(0.118)***	(0.121)***	(0.161)	(0.138)***	(0.172)***	(0.217)***
Two years of job tenure	0.337	0.466	0.070	0.423	0.370	0.424
	(0.105)***	(0.120)***	(0.208)	(0.152)***	(0.178)**	(0.319)
Four years of job tenure	0.201	0.281	0.070	0.209	0.223	0.135
	(0.099)**	(0.128)**	(0.135)	(0.147)	(0.157)	(0.289)
Five years of job tenure	0.192	0.347	-0.098	0.380	0.414	0.223
	(0.111)*	(0.110)***	(0.223)	(0.146)***	(0.200)**	(0.354)
Six years of job tenure	-0.040	-0.034	0.029	0.146	-0.143	0.776
	(0.142)	(0.178)	(0.388)	(0.193)	(0.206)	(0.437)*
Seven years of job tenure	-0.177	-0.041	-0.389	-0.084	-0.137	0.077
	(0.229)	(0.313)	(0.235)*	(0.232)	(0.359)	(0.220)
Eight years of job tenure	-0.516	-0.753	-0.031	-0.422	-0.849	0.439
	(0.287)*	(0.373)**	(0.489)	(0.282)	(0.421)**	(0.366)
Nine or ten years of job tenure	-0.778	-1.093	-0.116	-0.685	-1.189	0.354
	(0.372)**	(0.427)***	(0.616)	(0.319)**	(0.446)***	(0.479)
Constant	-3.589	-4.325	-4.235	-3.682	-4.275	-4.524
	(.331)***	(.368)***	(.384)***	(.339)***	(.377)***	(.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. The estimates shown are latent index model coefficients. See Figure 4 for marginal effects.

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

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

Table 7: The impact of regional subsidies for the conversion of job contracts on the variation of household wealth.

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	Leaving parental home	Leaving parental home	Leaving parental home	Leaving parental home
	(1)	(2)	(3)	(4)	(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 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 in Table 3. The marginal effects computed for a parental household holding the median wealth and for individuals aged 31-40 years when were hired in 1999, and interviewed in 2002. The remaining covariates 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 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 the one 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 those used in Table 3. 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 9: 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 25-64

	Model (i)	Model (ii)			
Dependent variable:	Living with a partner (1)	Owning and living with a partner (2)	Owning and not living with partner (3)	Renting and living with a partner (4)	Renting and not living with partner (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

PANEL B: SAMPLE OF INDIVIDUALS AGED 25-45

	Model (i)	Model (ii)			
Dependent variable:	Living with a partner (1)	Owning and living with a partner (2)	Owning and not living with partner (3)	Renting and living with a partner (4)	Renting and not living with partner (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

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 3. Other covariates not shown here are: indicators of age at hire, indicators of year of 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 $>$30			
2. Aragon	None	4200 if female or age$>$45 3000 if male 41$\leq$age$\leq$44	5160 if female or age$>$45 4500 if 41$\leq$age$\leq$44 3600 if male age$<$30	5160 if female or age$>$45 4500 if 41$\leq$age$\leq$44 3600 if male age$<$46
3. Asturias	4500 for all	4500 for all	None	4,200 if female or age$\geq$46 3600 otherwise
4. Baleares	None	None	None	1652.78 if female
5. Canarias	None	3,600 if female or age$\leq$25	3,600 if female or age$\leq$25 3,000 otherwise	None
6. Cantabria	None	3900 if female or age$\leq$30 3300 if male 30$<$age$\leq$40 3,600 if male age$\geq$41	None	5408 if age$\geq$46 3606 if age$\leq$30 2163 otherwise
7. Castilla-Leon	None	5112 if female or age $<$30 3300 rest of males	5115 if age $<$30 3900 if female age$\geq$30 1800 if male age $\geq$41	4507.59 if age $<$30 3305.57 if female age$\geq$30 1803 if male age$\geq$41
8. Castilla-La Mancha	None	3600 if females 3000 if age$<$30	None	3600 if female 3000 if age$>$45 or age$<$30
10. Comunidad Valenciana	None	30% of payroll tax	30% of payroll tax	30% of payroll tax
11. Extremadura	10655.94 if age$<$45 13402.57 if age$>$45	10100 if age$\leq$30 11180 if age$>$30 and age$\leq$40 14027 if age$>$40	14027.62 if age$>$46 11178.83 if age$<$45	5217.076 if female age$>$46 4296.416 if male age$>$46
12. Galicia	None	4200 euro if female or age$<$30 3000 if age$>$45	1800 for males 2400 for females	1800 for males 2400 for females
13. Madrid	None	6000 euro if female 6000 euro if age$<$30 or age$>$40	7800 if female 7800 if age$<$25 or age$>$40	9000 if female 6600 if age$<$25 or age$>$40
14. Murcia	None	None	None	2100 if age$<$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

Notes: The number associated with each Spanish region corresponds with the one assigned by the Spanish Statistical Office (in Spanish, Instituto Nacional de Estadística, INE).

Catalonia (region 9) and Navarre (region 15) have not provided subsidies for the conversion of fixed-term contracts into open-ended ones in period 1997-2004.

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 3000 if females >30 2400 if males >30	6012 for females or age<30 3607 if male age>40	None	None
2. Aragon	5160 if female or age>45 4500 if age≥41 3600 if male age<30	4500 if female 4125 if age<30 or age≥41	4500 if female 4125 if age<30 or age≥41	3750 if male, 7250 if female
3. Asturias	4,200 if female or age≥46 3600 otherwise	4200 if female of age>46 3600 otherwise	4200 if female of age>46 3600 otherwise	None
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 3005 if male age ≤30 4207 if age >45, 1803 otherwise	same as previous year	same as previous year	same as previous year
7. Castilla-Leon	4507.59 if age <30 3305.57 if female age>31 1803 if male age>41	same as previous year	same as previous year	same as previous year
8. Castilla-La Mancha	3600 if female 3000 if age>45 or age<30	same as previous year	same as previous year	None
10. Comunidad Valenciana	4808.1 for all	1800 for females	2000 for females 1500 for the rest	4000 if female 2000 if age<30, 1500 otherwise
11. Extremadura	5410.086 if female >45 4455.365 if male > 45 2386.802 otherwise	6010 for all	None	None
12. Galicia	1800 for males 2400 for females	1800 for males 2400 for females	1800 for males 2400 if females	1800 for males 2400 for females
13. Madrid	12000 if above 45 (males) 12000 if above 40 (females) 10800 for the rest	12000 for all	9000 if age≤45 12000 if above 45	3000 euro, all
14. Murcia	4800 for all	4800 for all	4800 if age above 30 5400 if female or age 30 or below	2400 for all
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

Notes: The number associated with each Spanish region corresponds with the one assigned by the Spanish Statistical Office (in Spanish, Instituto Nacional de Estadística, INE). Catalonia (region 9) and Navarre (region 15) have not provided subsidies for the conversion of fixed-term contracts into open-ended ones in period 1997-2004.

Table A.2: Individual characteristics in the sample of job spells of individuals aged 25-45.

	Total Sample	Subsample conditional on: Moving to a new house	Individual moves to: Owned house	Rented house
Move to an owned house	0.15	0.67		
Move to a rented house	0.07	0.33		
Holding an open-ended contract	0.68	0.78	0.85	0.64
Holding a fixed-term contract	0.32	0.22	0.15	0.36
Age	33.61	34.05	34.25	33.66
	(5.55)	(4.86)	(4.82)	(4.93)
Age at hire	29.80	28.70	28.36	29.37
	(5.73)	(4.95)	(4.80)	(5.18)
No. of years at current job	3.82	5.36	5.89	4.30
	(2.53)	(2.36)	(2.07)	(2.53)
Individual eligible for subsidy	0.57	0.61	0.61	0.63
Amount individual was eligible for	2.04	2.56	2.49	2.70
(include zeroes)	(2.57)	(2.88)	(2.74)	(3.13)
Individual labor earnings:				
Mean	15.35	17.71	18.69	15.76
	(9.52)	(10.68)	(11.15)	(9.40)
Attained education level:				
Primary education or less	0.11	0.08	0.06	0.11
First stage of secondary educ.	0.44	0.41	0.37	0.50
Second stage of secondary educ.	0.17	0.20	0.22	0.16
Tertiary education	0.27	0.31	0.35	0.23
Male	0.56	0.56	0.54	0.60
No. of years elapsed from the job spell until a move occurs or a job spell ends (% cases):				
One year	13.62	27.95	27.80	28.24
Two years	17.50	22.16	21.72	23.02
Three years	16.92	14.09	14.13	13.99
Four years	13.87	13.95	15.12	11.64
Five years	11.76	9.76	10.88	7.53
Six years	9.26	5.26	4.96	5.86
Seven years	6.48	4.51	3.65	6.23
Eight years or more	10.60	2.33	1.75	3.50
Minimum sample size	4,637	917	598	317

Source: The sample is formed by all individuals aged between 25 and 45 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 A.3: 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	0.0052	0.0064	0.0060	0.0059	0.0057
(standard error)	(.0017)***	(.0018)***	(.0018)***	(.0021)***	(.0019)***
2. Constant	0.517	0.602	0.554	0.559	0.762
(standard error)	(.023)***	(.028)***	(.033)***	(.036)***	(.023)***
F- test of instruments	9.17	12.39	10.70	8.20	9.55
Number of individuals	343,542	409,198	427,620	436,556	519,517
Average no. of observations per individual	1.87	2.57	3.17	3.64	5.00

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 of 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% significance level.

Table A.4: Estimates of the probability of conversion of a fixed-term contract into an open-ended one at each year of the individual's job tenure in their current job and firm.

Estimation method: logit duration model estimates
Sample: Temporary job contracts of employees aged 25-45 with job tenures shorter than 10
Dependent variable: Indicator of conversion of a fixed-term contract into an open-ended one

	Time-invariant effects of subsidies (1)	Time-varying effects of subsidies (2)
1. Subsidy to contract conversion	0.0392	0.0747
	(.0045)***	(.0066)***
2. Subsidy * 1 year of job tenure	–	-0.0306
		(.0084)***
3. Subsidy * 2 years of job tenure	–	-0.0574
		(.0120)***
4. Subsidy * 4 years of job tenure	–	-0.0063
		(.0072)
5. Subsidy * 5 years of job tenure	–	-0.0241
		(.0147)*
6. Subsidy * 6 years of job tenure	–	-0.0456
		(.0216)**
7. Subsidy * 7 or more years of job tenure	–	-0.0563
		(.0205)***
One year of job tenure	-1.2210	-1.1650
	(.0430)***	(0.0506)***
Two years of job tenure	0.8970	0.9964
	(.0664)***	(0.0600)***
Four years of job tenure	0.0242	0.0404
	(.0381)	(0.0378)
Five years of job tenure	-0.4199	-0.3740
	(.0494)***	(0.0600)***
Six years of job tenure	-0.5060	-0.4293
	(.0676)***	(0.0790)***
Seven years of job tenure	-0.5374	-0.4462
	(.0613)***	(0.0760)***
Eight years of job tenure	-0.6552	-0.5636
	(.1181)***	(0.1161)***
Nine or ten years of job tenure	-0.7456	-0.6519
	(.1170)***	(0.1298)***
Constant	-1.8253	-1.8965
	(.0280)***	(.0326)***
Number of individuals	391,923	391,923
Number of spells	782,317	782,317

Source: The data come from the 2004-2015 waves of the Continuous Sample of Working Histories (MCVL). The sample is formed by the duration of the job spell until the employee holding a fixed-term contract is promoted with an open-ended contract in his firm during the first ten years of the job tenure. Estimates are coefficients from the latent index model, see text for predicted marginal probabilities.
Other covariates included in the model not shown here are the following: indicators of age at hire, indicators of year of hire, year dummies, indicators of the household member's gender and occupation groups, region dummies and the yearly national unemployment rate. Standard errors are corrected for heteroscedasticity and also take into account arbitrary correlation among regions.

Table A.5: Randomization Inference of the impact of placebo regional subsidies on the stock of permanent contracts and on forms of housing tenure regime.

Estimation method: Randomization Inference (RI)				
Panel A: Sample of young individuals aged 25-64				
	First-stage (FS)	Intention-to-treat estimates (ITT)		
Dependent variable: Indicator of:	Open-ended	Coresidence	Home ownership	Rental
OLS Estimate:				
Subsidy to contract conversion	0.0082	-0.0102	0.0044	0.0057
	(.0024)***	(.0018)***	(.0022)*	(.0016)***
RI confidence interval at the 1% significance level:				
Completely random	[-.0020,.0023]	[-.0036,.0023]	[-.0036,.0029]	[-.0020,.0015]
Region	[.0009,.0057]	[-.0040,.0015]	[-.0039,.0014]	[-.0001,.0040]
Age at hire	[-.0020,.0024]	[-.0038,.0024]	[-.0035,.0029]	[-.0020,.0016]
Year of hire	[-.0026,.0029]	[-.0035,.0030]	[-.0036,.0035]	[-.0027,.0021]
Gender	[-.0020,.0024]	[-.0038,.0023]	[-.0037,.0030]	[-.0020,.0015]
Region x age at hire	[.0009,.0058]	[-.0050,.0005]	[-.0034,.0019]	[.0006,.0046]
Region x Year of hire	[.0058,.0078]	[-.0117,-.0092]	[.0024,.0049]	[.0054,.0073]
Region x Gender	[.0014,.0065]	[-.0037,.0020]	[-.0038,.0014]	[-.0000,.0040]
Age at hire x Year of hire	[-.0032,.0025]	[-.0034,.0024]	[-.0027,.0033]	[-.0029,.0022]
Age at hire x Gender	[-.0018,.0024]	[-.0038,.0024]	[-.0034,.0029]	[-.0022,.0014]
Year of hire x Gender	[-.0031,.0027]	[-.0034,.0025]	[-.0034,.0033]	[-.0026,.0024]
Minimum sample size:	5087	5087	5087	5087
Panel B: Full sample of workers aged 25-45				
OLS Estimate:				
Subsidy to contract conversion	0.0062	-0.0131	0.0058	0.0073
	(.0026)**	(.0022)***	(.0024)**	(.0030)**
RI confidence interval at the 1% significance level:				
Completely random	[-.0037,.0024]	[-.0027,.0031]	[-.0029,.0028]	[-.0025,.0021]
Region	[-.0008,.0058]	[-.0052,.0008]	[-.0035,.0022]	[.0000,.0045]
Age at hire	[-.0037,.0023]	[-.0026,.0028]	[-.0030,.0028]	[-.0028,.0021]
Year of hire	[-.0041,.0030]	[-.0042,.0032]	[-.0035,.0026]	[-.0039,.0024]
Gender	[-.0036,.0023]	[-.0026,.0029]	[-.0028,.0027]	[-.0026,.0021]
Region × age at hire	[-.0004,.0060]	[-.0055,-.0001]	[-.0031,.0027]	[.0003,.0050]
Region × Year of hire	[.0030,.0055]	[-.0152,-.0128]	[.0038,.0064]	[.0075,.0096]
Region × Gender	[.0004,.0063]	[-.0046,.0011]	[-.0038,.0018]	[-.0002,.0044]
Age at hire × Year of hire	[-.0037,.0028]	[-.0048,.0026]	[-.0030,.0030]	[-.0035,.0027]
Age at hire × Gender	[-.0034,.0024]	[-.0026,.0029]	[-.0028,.0027]	[-.0028,.0020]
Year of hire × Gender	[-.0041,.0030]	[-.0042,.0031]	[-.0038,.0031]	[-.0035,.0022]
Minimum sample size:	3974	3974	3974	3974

Source: The 2002-2011 waves of the Spanish Survey of Household Finances (EFF). The OLS estimates come from Table 3. Estimates combined across five implicates to take into account the uncertainty about imputation. 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 the one used in Table 3.

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

Estimation method: logit model estimates

Sample: Individuals aged between 25 and 45 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)
1. Subsidy to contract conversion	0.086	0.083	0.093	0.110	0.060	0.180
	(.010)***	(.019)***	(.025)***	(.022)***	(0.028)**	(0.032)***
2. Subsidy * 1 year job tenure	–	–	–	-0.033	0.030	-0.109
				(0.020)*	(0.031)	(0.052)**
3. Subsidy * 2 years job tenure	–	–	–	-0.023	0.032	-0.095
				(0.030)	(0.044)	(0.051)*
4. Subsidy * 4 years job tenure	–	–	–	-0.008	0.021	-0.039
				(0.018)	(0.029)	(0.069)
5. Subsidy * 5 years job tenure	–	–	–	-0.042	0.001	-0.096
				(0.029)	(0.032)	(0.068)
6. Subsidy * 6 years job tenure	–	–	–	-0.073	0.018	-0.240
				(0.054)	(0.050)	(0.096)***
7. Subsidy * 7 or more years of job tenure	–	–	–	-0.021	0.047	-0.151
				(0.047)	(0.053)	(0.042)***
One year of job tenure	0.367	0.383	0.250	0.488	0.278	0.686
	(0.122)***	(0.139)***	(0.192)	(0.154)***	(0.201)	(0.275)***
Two years of job tenure	0.278	0.381	0.043	0.363	0.267	0.439
	(0.102)***	(0.131)***	(0.238)	(0.183)**	(0.247)	(0.385)
Four years of job tenure	0.161	0.243	0.021	0.193	0.169	0.190
	(0.101)	(0.129)*	(0.215)	(0.160)	(0.173)	(0.471)
Five years of job tenure	0.188	0.284	0.040	0.342	0.276	0.452
	(0.111)*	(0.113)***	(0.266)	(0.161)**	(0.191)	(0.462)
Six years of job tenure	-0.019	-0.077	0.179	0.235	-0.144	1.081
	(0.166)	(0.189)	(0.428)	(0.221)	(0.237)	(0.503)**
Seven years of job tenure	-0.071	0.000	-0.124	0.014	-0.163	0.515
	(0.224)	(0.328)	(0.298)	(0.216)	(0.368)	(0.240)**
Eight years of job tenure	-0.461	-0.643	-0.010	-0.375	-0.807	0.636
	(0.292)	(0.383)*	(0.528)	(0.274)	(0.425)*	(0.380)*
Nine or ten years of job tenure	-0.661	-1.018	0.128	-0.576	-1.182	0.776
	(0.344)*	(0.439)**	(0.729)	(0.279)**	(0.456)***	(0.564)
Constant	-3.706	-4.551	-4.221	-3.801	-4.468	-4.612
	(.377)***	(.407)***	(.500)***	(.395)***	(.418)***	(.487)***
Minimum number of spells	4,637	4,637	4,637	4,637	4,637	4,637

Source: The sample is formed by all household members aged between 25 and 45 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. The estimates shown are latent index model coefficients. See Figure 4 for marginal effects.

Other covariates in the model not shown here (as those used in Table 6) are: indicators of age at hire, indicators of year of hire, year dummies, indicators of the household member's gender and education level, region dummies and yearly national unemployment rate.

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

Table A.7: Fixed-effect estimates of the time-invariant impact of regional subsidies for contract conversions on the probability of moving to a new house at each year of job tenure.

Sample: employees aged between 25 and 64 with job tenures shorter than 10.				
Estimation method: logit model estimates of transitions to a new accommodation				
Panel A: Coefficient estimates of the probability of moving to a new accommodation, as an emancipated person				
Dependent variable:	Indicator of whether the individual moves to:			
	A new house			
	(1)	(2)	(3)	(4)
1. Subsidy to contract conversion	0.064	0.067	0.067	0.068
	(.011)***	(.012)***	(.012)***	(.017)***
2. Constant	-3.589	-3.590	-3.497	-3.373
	(.331)***	(.339)***	(.343)***	(.362)***
(Age at hire × year of hire) dummies	No	Yes	Yes	Yes
(Region × age at hire) dummies	No	No	Yes	Yes
(Region × year of hire) dummies	No	No	No	Yes
P-value of significance of second-order effects	–	0.00	0.17	0.00
Minimum pseudo R-squared	0.03	0.04	0.04	0.05
Minimum sample size of transitions	5,997	5,997	5,997	5,997

Data source: The 2002-2011 waves of the Survey of Household Finances.

The covariate of interest is the subsidy for contract conversions that each individual was eligible for in the first two years of job tenure. The set of regressors is the same as those used in Table 6.

Standard errors (in parentheses) are clustered at the region level. Estimates are combined across five implicates to take into account the uncertainty about imputation. P-values of the test of joint significance of the additional second-order effects included in each specification is computed using the approach of Meng and Rubin (1992) with data sets imputed multiply.

The symbols *, ** and *** denote that the estimates are significant at the 10%, 5% and 1% significance level, respectively.

Table A.7: Fixed-effect estimates of the time-invariant impact of regional subsidies for contract conversions on the probability of moving to a new house at each year of job tenure (Contd.).

Panel B: Coefficient estimates of the probability of moving to a new accommodation distinguishing by housing tenure regime (competing risk models).

Dependent variable:	Indicator of whether the individual moves to:							
	An owned house				A rented house			
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
1. Subsidy to contract conversion	0.065	0.067	0.068	0.070	0.063	0.069	0.068	0.069
	(.016)***	(.017)***	(.018)***	(.023)***	(.019)**	(.019)***	(.019)***	(.025)***
2. Constant	-4.325	-4.306	-4.236	-4.170	-4.235	-4.280	-4.168	-3.910
	(.368)***	(.360)***	(.360)***	(.398)***	(.384)***	(.410)***	(.412)***	(.383)***
(Age at hire × year of hire) dummies	No	Yes	Yes	Yes	No	Yes	Yes	Yes
(Region × age at hire) dummies	No	No	Yes	Yes	No	No	Yes	Yes
(Region × year of hire) dummies	No	No	No	Yes	No	No	No	Yes
P-value of significance of second-order effects	–	0.04	0.21	0.19	–	0.21	0.27	0.06
Minimum pseudo R-squared	0.05	0.05	0.06	0.07	0.03	0.03	0.04	0.06
Minimum sample size of transitions	5,997	5,997	5,997	5,997	5,997	5,997	5,997	5,997

Data source: The 2002-2011 waves of the Survey of Household Finances.

The covariate of interest is the subsidy for contract conversions that each individual was eligible for in the first two years of job tenure. The set of regressors is the same as those used in Table 6.

Standard errors (in parentheses) are clustered at the region level. Estimates are combined across five implicates to take into account the uncertainty about imputation. P-values of the test of joint significance of the additional second-order effects included in each specification is computed using the approach of Meng and Rubin (1992) with data sets imputed multiply.

The symbols *, ** and *** denote that the estimates are significant at the 10%, 5% and 1% significance level, respectively.

Table A.8: Randomization Inference of the time-invariant impact of placebo regional subsidies on the probability of moving to a new house at each year of job tenure.

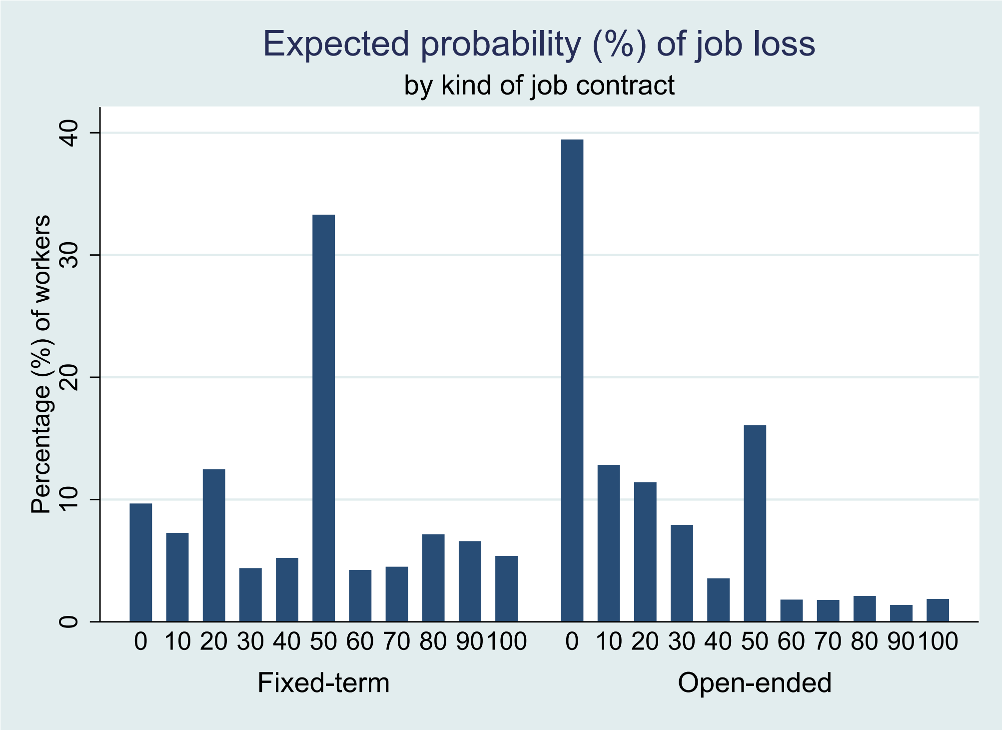
Sample: Job spells from employees aged between 25 and 64 with job tenures shorter than 10.
Estimation method: logit model of transitions to a new accommodation

Panel A: Coefficient estimates of the probability of moving to a new accommodation			
Dependent variable:	Indicator of whether the individual moves to:		
	A new house	An owned house	A rented house
	(1)	(2)	(3)
Logit duration model estimate:			
Subsidy to contract conversion	0.0640	0.0652	0.0627
	(.011)***	(.016)***	(.019)***
RI confidence interval:			
At the 1% significance level:			
Completely random	[.0003,.0233]	[.0003,.0321]	[-.0178,.0245]
Region	[.0111,.0342]	[.0093,.0410]	[-.0067,.0360]
Age at hire	[.0003,.0230]	[.0004,.0319]	[-.0189,.0253]
Year at hire	[.0023,.0263]	[.0051,.0353]	[-.0166,.0217]
Gender	[.0007,.0229]	[.0001,.0317]	[-.0180,.0237]
Region × age at hire	[.0137,.0362]	[.0104,.0428]	[-.0044,.0395]
Region × Year of hire	[.0585,.0689]	[.0600,.0701]	[.0561,.0704]
Region × Gender	[.0114,.0348]	[.0096,.0420]	[-.0041,.0353]
Age at hire × Year of hire	[.0024,.0272]	[.0060,.0362]	[-.0135,.0248]
Age at hire × Gender	[-.0001,.0229]	[.0012,.0324]	[-.0202,.0232]
Year of hire × Gender	[.0029,.0255]	[.0059,.0343]	[-.0165,.0241]
Minimum sample size of transitions:	5,997	5,997	5,997

Data source: The 2002-2011 waves of the Survey of Household Finances. 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 those used in Table 6. Standard errors (in parentheses) are clustered at the region level. Estimates are combined across five implicates to take into account the uncertainty about imputation. The logit estimates come from Table 6.

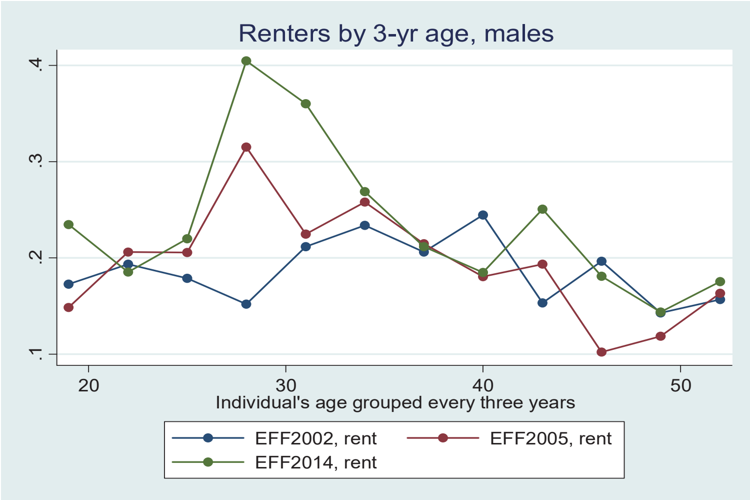
The symbols *, ** and *** denote that the estimates are significant at the 10%, 5% and 1% significance level, respectively.

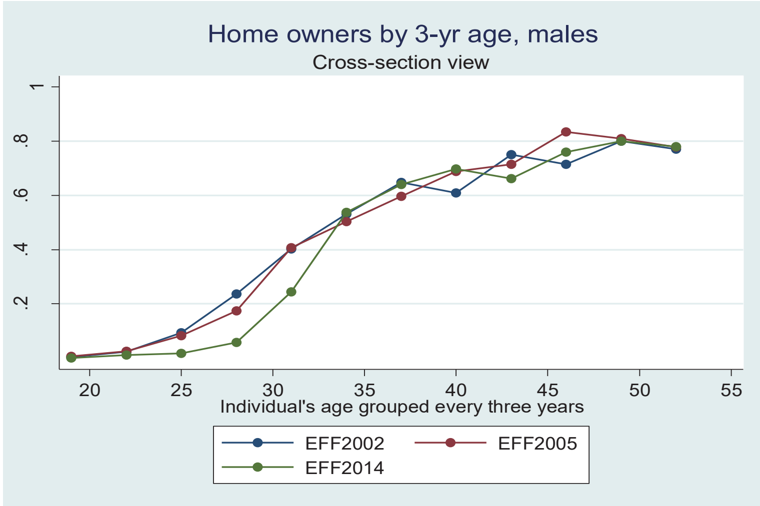
Figure 1: Subjective probability of job loss over the next 12 months, by type of contract



Source. Survey of Household Finances (EFF), 2011 wave. The sample contains all employees. The question asks the respondent to put on a scale from 0 to 100 the chances of losing the job, where 0 is not possible at all and 100 most certain.

Figure 2: Evolution of living arrangements of individuals 20-55 in Spain between 2002 and 2014

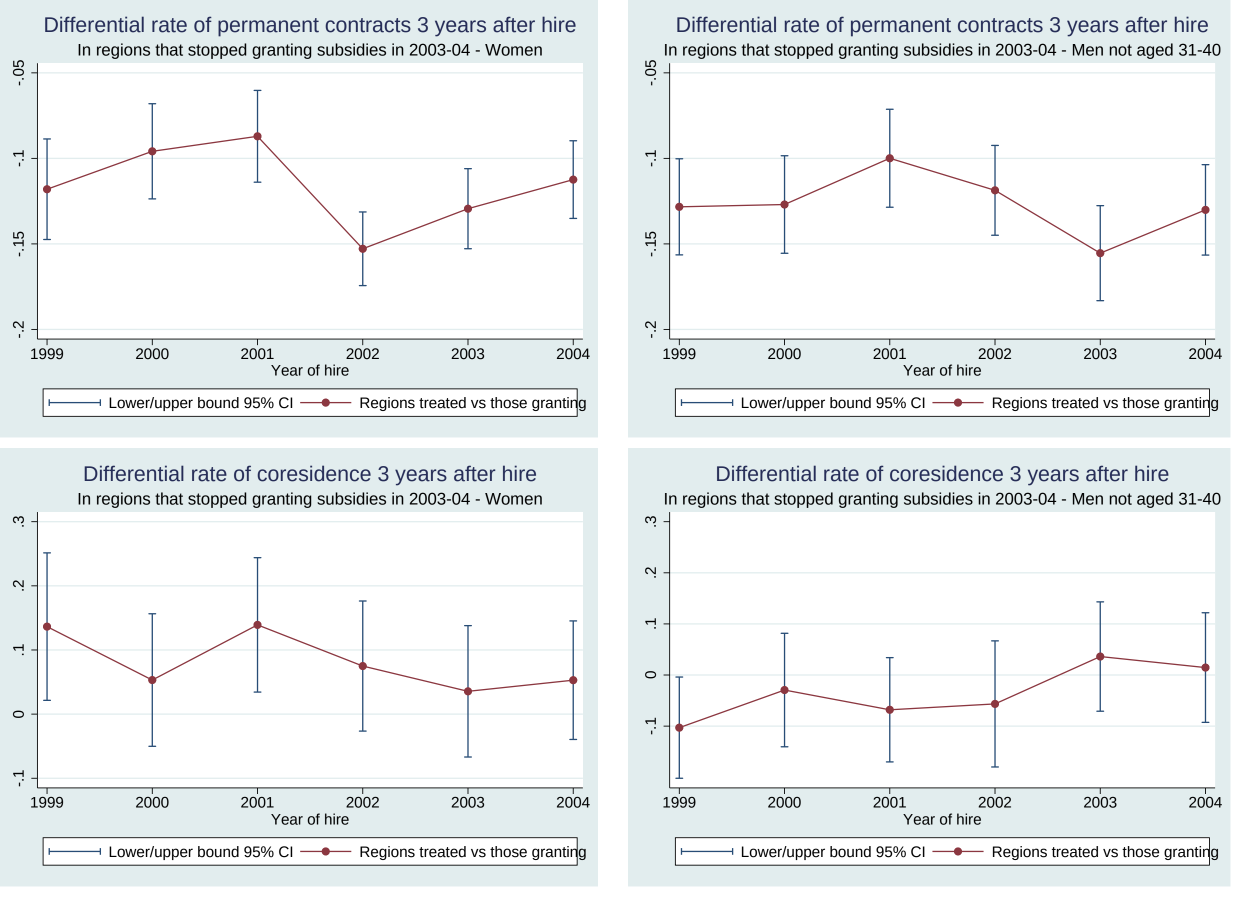




The panel on the left contains the fraction of males who rent their house of residence as a fraction of all males at that age (including owners, renters and those living with another person in a house they neither own nor rent). The panel on the right contains the fraction of males who own their house of residence.

Source: Spanish Survey of Household Finances, EFF2002-EFF2014.

Figure 3: An illustration of the evolution of the share of employees with an open-ended contract and the percentage of coresidents in the third year of the job spell in regions that stopped granting subsidies with respect to those granting continuously.



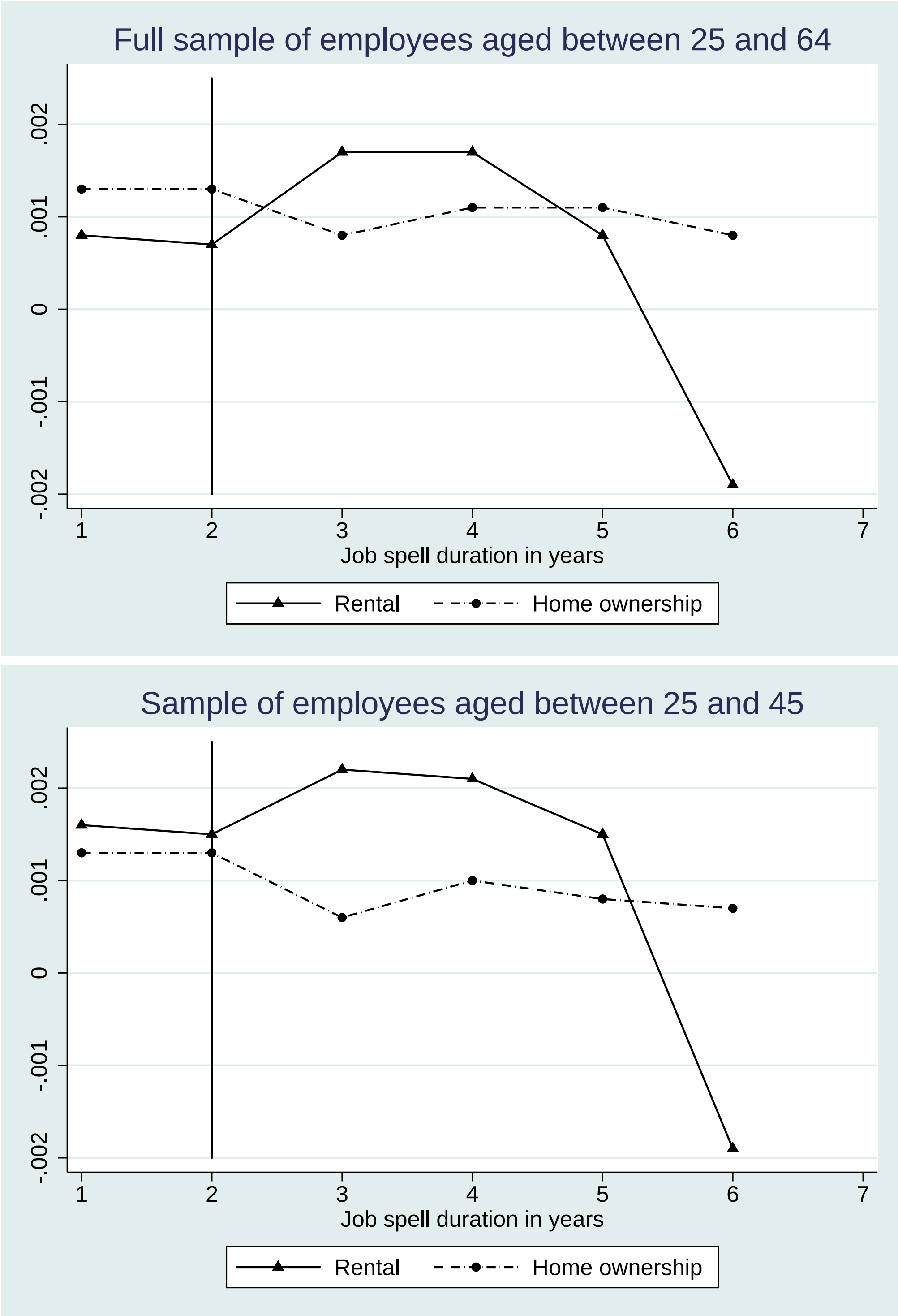
Source: The 2004-2015 waves of the Continuous Sample of Working Histories (MCVL) and the 2000-2008 waves of the Spanish Labor Force Survey (LFS) for employees aged 25-45.

The graphs above show the fraction of workers hired in the year in the horizontal axis who have an open-ended contract three years after the hire, by region (data from MCVL). The graphs below display the share of workers living with their parents or other relatives (coresidence) three years after being hired (year of hire in horizontal axis), data from LFS.

Andalucia and Extremadura are the regions that not granted subsidies the whole period analyzed, 1999-2004, since they stopped granting subsidies for the conversion of job contracts in 2003-2004, resuming in 2005.

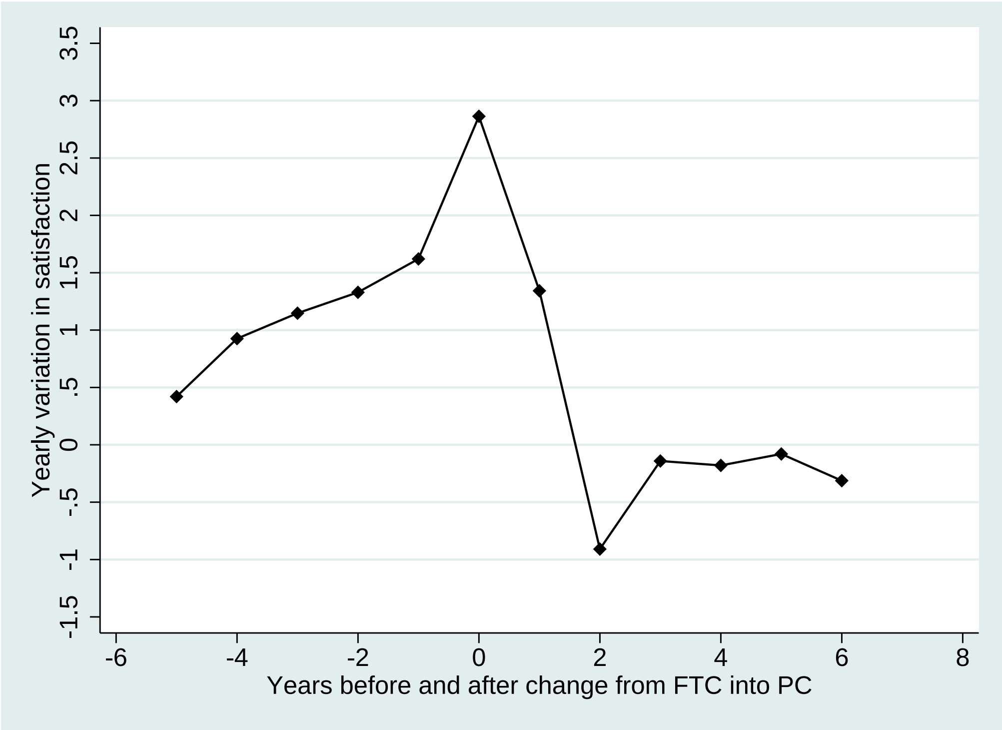
The group of regions that granted subsidies continuously for the same targeted group of individuals (all women and males only aged below 30 and over 40) throughout the period 1999-2004 is formed by Madrid, Comunidad Valenciana, Galicia, Basque Country and La Rioja.

Figure 4: Hazard rate estimates from a multinomial transition duration model.



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 is 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 with security on the job (measured from 0 to 10) on a window between 5 years before and 6 years after an employee obtains an upgrade from a fixed-term contract (FTC) into a permanent one (PC). The sample is an unbalanced panel drawn from the 1994-2001 waves from the European Community Household Panel.

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