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Abstract

Economic theory predicts that individuals exposed to the risk of losing their job postpone their consumption and accumulate more assets to build a buffer stock of saving. We provide a new test of the hypothesis using substantial variation in severance payments across contracts in the Spanish labor market. Using the 2002 and 2005 waves of a new survey of wealth and consumption we estimate the link between the probability that several household members lose their job and the wealth and consumption of that household. We instrument the type of contract using regional variation in the amount, timing and target groups of subsidies given to firms to hire workers using high severance payment ones. We find that workers covered by fixed-term contracts accumulate more financial wealth. An increase in the probability of losing the job of 8 percentage points increases average financial wealth by 4 months of income. We provide simulations from a simple buffer stock and a permanent income models that suggest that our results are more likely to be generated by the former.

Keywords: precautionary savings, household wealth and consumption, labor firing costs.

JEL classification: D12, D31, D91, J41.

1 Introduction

Do workers exposed to the risk of losing the job accumulate wealth balances in anticipation of future income losses? Quantifying such balances and their distribution over the population is crucial to understand the evolution of consumption when unemployment increases. From a theoretical perspective, whether or not households perceive unemployment risk and react to it by accumulating wealth determines the welfare consequences of income fluctuations (Hall, 2006, Caballero, 1990, or Carroll, 2001). Our study uses the wide variation in the costs of dismissing workers covered by different contracts in the Spanish labor market to quantify the size of wealth accumulated by households differently exposed to the risk of losing the job.

Several studies have documented that, in a variety of nations, households experience substantial consumption drops following job losses, an indication of an imperfect ability to insure against the risk of losing the job - see Gruber (1997), Browning and Crossley (2009) or Bentolila and Ichino (2008). In the absence of insurance, a natural response of households is to accumulate wealth to self-insure against the risk of losing the job. Nevertheless, quantifying the consumption or wealth balances held faces at least three empirical problems. First, workers who are more averse to risk are also more likely to self-select into occupations with lower risks of losing the job (Lusardi, 2007 and Fuchs-Schündeln and Schündeln, 2005). Secondly, if the propensity to lose the job is persistent over time, workers more exposed to that risk are also more likely to have used their savings to sustain consumption during previous unemployment spells. Finally, households with unstable income paths are less attractive customers for banks, casting doubts on whether higher wealth balances reflect a reaction to borrowing constraints.[1] Alternative solutions to those problems have led to different estimation strategies; Carroll, Dynan and Krane (2003) use a sample of US workers who have been at least three years in their jobs and find that an increase in the probability of suffering an unemployment spell of 1 percent leads to an increase in total wealth of about 3 months of earnings, a magnitude that can be reconciled with a

[1] Regarding the second empirical problem, Carroll, Dynan and Krane (2003) present simulations documenting the likely relevance of the extent of wealth depletion after past unemployment shocks. The same authors note that a fraction of the buffer stock accumulated by households more exposed to the risk of losing the job could be due to lower loan-to-value requirements by banks. They quantify an upper bound of such confounding effect of at most half of the estimated buffer stock. Fuchs-Schündeln and Schündeln (2005) mention the problem, but their paper does not report an explicit test quantifying the likely impact of borrowing constraints.

precautionary motive -i.e., households save because they are averse to fluctuations in their marginal propensity to consume. Engen and Gruber (2001) document a modest crowd-out of private wealth by unemployment insurance in the US, well below those predicted by a model with a precautionary saving motive. Fuchs-Schündeln and Schündeln (2005) and Fuchs-Schündeln (2008) find sizable wealth and savings responses to the differential impact of the German reunification shock across groups of the population and show that those can be explained by a precautionary saving model. On the other hand, Guiso, Jappelli and Terlizzese (1992) and the survey of Browning and Lusardi (1996) find less evidence supporting the hypothesis that household saving responds to income risk.[2]

Our study analyzes the wealth response to possible job losses using a well-known but largely unexploited source of variation in the exposure to the risk of losing the job: differences in dismissal costs associated to the type of contract a worker has. Differences in severance payments across contract types are indeed prevalent in OECD countries. France, Germany, Italy, Portugal, Spain and Sweden (among other countries) introduced low firing cost contracts during the eighties as a way of introducing flexibility at the margin in the labor market. As a wide literature has documented, these contracts have generated labor markets where easily identifiable workers face very different probabilities of transiting into unemployment (see, Blanchard and Landier, 2002 or Dolado et al., 2002). Importantly, that variation in exposure to unemployment risk allows us to examine behavioral responses that individuals are well aware of and that occur within occupation and industry groups.[3] Among all countries that introduced fixed-term contracts, Spain is the country with the highest share of fixed-term contracts (OECD, 2004), providing an ideal setting to analyze the saving decisions of households exposed to different dismissal costs.

A second advantage of our approach is that, due to the fact that some

[2] Jappelli, Padula and Pistaferri (2008) reject the buffer stock model, a particular model of precautionary saving.

[3] The use of measures that may not be strong predictors of future transitions into unemployment can explain some discrepancies of the results in the literature. For example, a possible explanation of the low wealth responses to state Unemployment Insurance reported in Engen and Gruber's study is that it is not always the case that individuals who can potentially benefit from higher unemployment benefits actually experience a higher probability of transiting into unemployment. An interesting literature has tried to elicit directly individual's perceptions of losing the job by reporting subjective expectations (Manski and Straub, 2000). In this study, we focus on objective measures.

labor market policies are decentralized in Spain, the incentive to use fixed-term contracts to hire a worker varies across regions, demographic groups and year-of-entry at the firm. In 1997, six out of the 17 Spanish regions started implementing subsidies to firms that hired workers using open-ended contracts (García-Pérez and Rebollo-Sanz, 2009a). By 2004, 15 out of 17 regions had implemented such programs during at least one year. Furthermore, different regions targeted the extension of the use of open-ended contracts among different demographic groups (see Table A.2). Such arguably exogenous variation permits us to compare the wealth holdings of workers who got a high dismissal cost contract because a subsidy was available in their region-age-gender-year of entry cell to the wealth holdings of those workers who did not get a high dismissal cost contract because a subsidy was absent. Such identification strategy mitigates biases due to self-selection and to previous unemployment experiences because it relies on the comparison of workers who were hired in the same year but whose firms faced different incentives to use open-ended contracts. Furthermore, as subsidies to hire using high firing cost contracts differed by gender, we are able to estimate the causal household wealth response to exposure to job loss for different household members.

Finally, we use an unusually rich wealth and consumption survey: the 2002 and 2005 waves of the Spanish Survey of Household Finances (in Spanish, *Encuesta Financiera de las Familias*, EFF), conducted by the Banco de España. The EFF is one of the few surveys in the world containing detailed information on households' assets, consumption and on the labor market situation of each household member. In addition, the EFF contains information about a number of outcomes that allows us to disentangle between alternative saving motives. In particular, we can use information on credit rejections to examine if wealth differences are due to liquidity constraints or household's demand factors.

We present Two-Stage Least Squares (TSLS) estimates of the impact of contract form on household wealth suggesting that households whose head obtained a high dismissal cost contract as a consequence of the regional subsidies have average financial wealth-earnings ratios between 29% and 40% lower than households whose head had a low firing cost contract. Within the sample of households headed by a male, the response of household financial wealth to the risk that the spouse loses her job is remarkably similar to that of the head. Nevertheless, we do not find that low dismissal cost contracts lead to higher wealth holdings when we include the net value of owner-occupied housing. We also document that subsidies to hire workers using open-ended contracts have little predictive ability in experiencing a credit rejection, ruling out explanations of our findings based on credit constraints. Finally, as higher exposure to job loss is not a mean preserving increase in risk, we assess what model is consistent with our findings by comparing our results to simulations from a simple buffer stock saving model and to a simple permanent income model. We conclude that a precautionary saving motive is more consistent with the data.

Section 2 summarizes the legislation of dismissal costs in Spain and its likely implications for wealth accumulation. Section 3 presents the data. Section 4 describes the identification strategy and Section 5 presents the main results. Section 6 discusses what wealth accumulation models can be consistent with our results and Section 7 concludes.

2 Dismissal costs in Spain

Before 1984 "open-ended" contracts were the most common form of regulating an employer-employee relationship in Spain. Those contracts featured a high cost if the firm wanted to dismiss the worker: between 20 and 45 days of wages per year worked. The former applied if the worker appealed to Court and the judges declared the dismissal as "fair". Otherwise, the corresponding severance payment amounted to 45 days per year worked, with a limit of 24 months' wages.[4] In 1984, in a context of high unemployment rates, the Worker's Act (*Estatuto de los Trabajadores*) introduced a menu of contracts that was exempted from the general rule of high severance payments ("fixed-term" contracts). The legal figure used was the extension to other types of labor relationships of contracts that up to that moment were only used to regulate seasonal jobs. "Fixed-term" contracts differed from "open-ended" ones in two main aspects. The first was the cost of dismissing the worker: initially of 12 days per year worker, zero if the firm waited until expiration. The second difference was that, in case of a dismissal, a worker covered by a fixed-term contract had no right to sue the employer claiming that the lay-off was unfair. There were no other differences between contracts in terms of contributions to the old-age or disability social security systems, access to

[4] Izquierdo and Lacuesta (2006) and Galdón-Sánchez and Güell (2000) estimate that between 72% and 75% of cases that arrived to court were declared "unfair" by Spanish judges.

unemployment benefits or the ability to access medical services.

Fixed-term contracts have been heavily used by Spanish employers; by 1994, according to the Spanish Labor Force Survey (EPA, in its Spanish initials) around 30% of workers reported being covered by a "fixed-term" contract. While subject to small fluctuations, the share has remained stable since then. Dolado et al. (2002) document that, while the use of such contracts has been widespread across all industries, they were specially prevalent among female, young and unqualified workers. Not surprisingly, whether the contract regulating the labor relationship is fixed-term or not is a strong predictor of the probability of transiting into unemployment. In Appendix Table A.1, we present computations from the EFF that suggest that holding constant education, an individual whose job position is regulated by an open ended contract was much less likely to experience an unemployment spell in 2004 than a worker covered by a fixed-term contract.[5] According to our EPA estimates, the difference in the yearly probability of entering an unemployment spell is 8% between temporary and permanent workers (10% for workers covered by a fixed-term contract and 2% for those with open-ended contracts).

We use the fact that variation in contract form causes identifiable groups of the population exposed to very different degrees of unemployment risk to estimate the response of household wealth to the contract held by different members of the household. Now, whether or not a worker's contract is open-ended or fixed-term is the result of firm's personnel selection practices, and an extensive literature has shown that workers who obtain an open-ended contract have higher expected productivity than other workers.[6] In this way, our key variable to measure the exposure to the risk of losing the job (the kind of job contract) is endogenous, since it is correlated positively with unobserved factors that make workers accumulate more wealth due to their higher expected permanent income.

Our identification strategy relies then on a set of programs implemented independently by 15 of the 17 Spanish regions to increase the stock of workers covered by high-dismissal cost contracts. In 1997, 5 of the 17 Spanish regions

[5] Güell and Petrongolo (2007), or García-Ferreira and Villanueva (2007) present similar evidence.

[6] Blanchard and Landier (2002) present a model of contract conversions in which firms only convert a fixed-term contract into a permanent one if the future expected productivity is "high enough". Portugal and Varejão (2009) also find evidence in favor of the screening hypothesis.

introduced regional subsidies to incentive firms to use open-ended contracts to hire workers. There were two main forms of subsidies (Table A.2 shows the maximum amount of these two subsidy forms). The first was granted to firms that converted an existing fixed-term contract into an open-ended one during the period in which the subsidy was available. The second form of subsidy was available to firms who hired an unemployed worker using an open-ended contract. Subsidies were granted by the regional administration as either a lump sum in the year when the conversion took place (which was the case in most regions) or a reduction to the pay-roll tax during two to three years (which was the case in three specific regions: Valencia, Castilla la Mancha and Extremadura). Progressively, between 1997 and 2004, all regions but Catalonia and Navarra had implemented for at least a year some form of subsidy to hiring using an open-ended contract. The statutory subsidy amount varied across demographic and gender groups, often being higher if the worker holding a first-term contract was female or either below 30 or above 45, and the amount granted ranged between 1,200 and 14,000 euro. García-Pérez and Rebollo-Sanz (2009a) estimate that the lump sum received by the firm was about 20% of the yearly labor cost of the mean worker.

At the same time, in 1997 a national policy was also implemented at the national level introducing a new form of contracts with lower firing costs. That new contract was available for unemployed workers hired when they were below 30 years of age or above 45 years of age or for firms that converted a worker's contract from fixed-term into an open-ended one. In the empirical section, we disentangle the impact of regional variation in incentives to hire from the national reform by introducing a separate intercept for all contracts signed after that year, and experiment interacting that dummy with age trends.

Summarizing, as a result of the policy of contract subsidization, the incentive to the firm to convert a fixed-term contract into an open-ended one varied across regions of residence, the year in which the contract started and with the age and gender of the worker. We exploit those differential incentives to obtain exogenous variation in the exposure of workers to the risk of losing the job and to estimate the associated wealth response.

2.1 Modelling issues

We build on analytical results by Blundell and Stoker (1999) to understand the link between household saving decisions and exposure to different dismissal costs. For simplicity of exposition, we assume that an individual lives for two periods, does not discount the future, and that there is a zero interest rate, but the results do not depend on such assumptions. The individual has an inelastic labor supply and is subject only to a single source of income risk: job loss. Namely, second-period income Y can either be the sum of unemployment benefits b and severance payments f if the individual loses his or her job or the current level of earnings y if the individual keeps his or her job. The first event happens with probability p. We further assume that consumption is adjustable. The utility function of the individual is the following:

$$\max_{c_1,c_2} \quad \frac{(c_1)^{1-\rho}}{1-\rho} + E\frac{(c_2)^{1-\rho}}{1-\rho}$$

Where the expectation is taken over the binary random variable Y, with mean, $E(Y) = p(b+F)+(1-p)y$, and variance, $Var(Y) = (1-p)p[y-b-f]^2$. The present value of expected lifetime resources in period 1 (W) is the sum of the initial wealth (W_0) and the expected stream of income in period 2, as follows:

$$W = W_0 + p(b+f) + (1-p)y$$

Blundell and Stoker (1999) show that, assuming that individuals can borrow freely against their future expected income (including against the value of their compensation package if laid-off), and taking a second-order approximation of the Euler condition evaluated at the solution under perfect-certainty, a closed form for consumption levels can be obtained:

$$c_1 = \frac{1}{1+\left[1+2\rho\left(1+\rho\right)\sigma^2\right]^{\frac{1}{\rho}}}W \tag{1}$$

Where σ^2 equals $\frac{Var(Y)}{W^2}$. Note that if the worker perceives no income uncertainty, σ^2 equals zero and lifetime wealth is split equally between the two periods. A worker covered by a high firing cost contract exhibits a lower value of p and, upon job loss, a higher value of f than a worker covered by a fixed-term contract. According to equation (1), such difference generates two channels that lead the first worker to accumulate a lower level of wealth. The first one holds as long as $b+f<y$ and is due to the fact that W diminishes with the probability of losing the job: discounted lifetime income is lower among workers who, everything else equal, are more likely to experience

an unemployment spell in the future. Therefore, workers holding fixed-term contracts will accumulate more wealth to smooth their consumption path due to their lower expected future income. The channel operates even if workers are certain about job loss, or σ^2 equals zero. A second channel leading a worker covered by a high firing cost contract to save less is that workers who are relatively more exposed to higher uninsurable income risk accumulate a higher buffer of wealth to minimize the fluctuations in the marginal utility of consumption. That reason shows up in the coefficient linking consumption in the first period to lifetime resources: $\frac{1}{1+[1+2\rho(1+\rho)\sigma^2]^{\frac{1}{\rho}}}$ is higher the lower σ^2, all else equal.[7]

Finally, a second implication of the model is that individuals who are exposed to a higher risk of losing the job postpone consumption to the future and hence will exhibit higher consumption growth than workers covered by high firing cost contracts. To see this, define the binary variable "second-period shock" ζ_2 as the difference between the realization of second-period income and the expected value of the income stream

$$\zeta_2 = Y - [p(b+f) + (1-p)y]$$

Under our assumptions, if $\rho = 1$, consumption growth can be written as follows:

$$\log(c_2) - \log(c_1) = \kappa_0 + \kappa_1 \frac{Var(Y)}{W^2} + \frac{1}{c_1}\zeta_2 \tag{2a}$$

In (2a), consumption growth of an individual exposed to the risk of losing the job is a stochastic variable. It may take positive or negative values depending on whether or not the individual experiences the unemployment shock. The parameters κ_0 and κ_1 are functions of the structural parameters of the model. Taking expectations in (2a) over the distribution of Y -that is,

[7] Some researchers introduce adjustments for expected income, like controlling for p to control for mean effects and examining the response of wealth to σ^2 (see Lusardi, 1997). Other researchers have used of married working couples and used the fact that if both members work in the same industry, the variance of overall household income is higher to obtain mean preserving shocks (Shore and Sinai, 2009). While both are very interesting approaches, in our application we preferred neither to focus on married couples or to rely on non-linearities in p to identify the main effects, and follow a different strategy to gauge what underlying savings motives explain our results.

aggregating consumption growth among households who did experience and did not experience a job loss episode- one obtains the following expression:

$$E_1[\log(c_2) - \log(c_1)] = \kappa_0 + \kappa_1 \frac{Var(Y)}{W^2} \tag{2}$$

Individuals who are covered by a low dismissal cost contract are exposed to a higher variability in earnings (i.e., they have a higher value of p), so they tend to postpone consumption. Interestingly, under perfect certainty (that is, if job loss is a perfectly anticipated event), expected consumption growth should be zero. Hence, a positive link between consumption growth and the dismissal cost specified in the worker's contract gives an additional indication of whether a precautionary saving motive is indeed present in the data.

Summarizing, we test two hypotheses. The first is that workers covered by contracts that make them costly to dismiss accumulate less wealth than workers covered by low dismissal cost contracts. The second is that workers in "protected" jobs exhibit lower consumption growth than workers in low dismissal cost contracts.

It is worth noting that both predictions could be generated by a model where households do not react to risk, but lenders prefer lend only to workers with stable jobs. Workers with unstable jobs may then accumulate more assets and postpone consumption to build collateral. We test for that possibility in Section 5.3 by examining the response of credit rejections to changes in contractual firing costs.

3 Data set

The main data set we use contains the 2002 and 2005 waves of the Spanish Survey of Household Finances, conducted by the Banco de España (in Spanish, *Encuesta Financiera de las Familias*, EFF). The EFF surveys around 5,000 households in each wave, obtaining detailed information about wealth holdings, debt, payment habits and consumption at the household level and individual information about demographics, income and labor income status.[8] Unless otherwise noted, all the calculations reported in this study

[8] Based on the wealth tax, there is over-sampling of wealthy households. Around 40% of the sample corresponds to households liable to the wealth tax. To assess the consequences of the oversampling for our estimates, we experimented weighting the regressions

make use of the five multiply imputed data sets provided by the Banco de España as a way of dealing with item nonresponse – for details on the EFF imputations see Bover (2004) and Barceló (2006).

The dependent variable: We use various measures of wealth. The first is gross "liquid" financial wealth, i.e., the subset of wealth that we consider to be easily cashed during an unemployment spell. It contains amounts held in checking and saving accounts, mutual funds, stock (either listed or not), all types of bonds and other financial assets. Throughout the paper, we exclude from the analysis vehicles (hard to cash, as there is a limited second-hand market for them), pension funds (the Spanish version of IRAs, not cashable in the event of unemployment until 2006), life insurance and business wealth. The second measure of wealth is a bit broader and adds to the former the value of real estate properties net of associated debts, but does not include owner occupied housing -see Bover (2005) for evidence that other real estate properties may serve a precautionary motive. There is a discussion regarding whether or not households are able to use owner-occupied housing equity to finance an unemployment spell; Carroll, Dynan and Krane (2003) argue that housing wealth is "liquid" because of the possibility of extracting equity, others disagree - Engen and Gruber (2001) or Shore and Sinai (2009), who treat housing wealth as consumption. In the Spanish case, the possibility of extracting housing equity from owner-occupied housing was rather limited. Furthermore, Spain has experienced a sharp increase in housing prices since the late nineties, with yearly increases well above 10%. Much of the variation in housing wealth would pick up those capital gains, which we found hard to interpret. We experiment thus with our broadest measure that includes the net value of owner-occupied housing, but we are more confident about the first two measures: gross financial wealth and gross financial wealth plus the net value of real estate other than housing equity.

The main sample is composed of households headed by an employee head between 23 and 65 years of age. We excluded those cases that had total labor earnings below 1,000 euros of 2005 or who were hired after 2005 (we did not collect the regional subsidies after 2004). The reason for dropping the self-employed is that the instrument we use (regional subsidy to hire workers using an open-ended contract) was only available for employees. The sample contains 3,660 household-years on average across the five multiply imputed data sets. As we take logarithms of wealth in most of the analysis, we lose

by population weights. The results were qualitatively similar but noisier.

another 114 cases that have zero financial wealth.[9] Finally, we establish the risk that the household head or a working spouse (if present) lose the job according to the contract type of in the first job reported (the EFF asks about the characteristics of the first three most important jobs). For the estimation of the impact of the risk of losing the job on consumption growth, we use the subsample of the EFF for which we have two observations. The sample is composed by 625 households in all data sets imputed multiply.

The instrument: statutory amount to hire workers using open-ended contracts. We compute the subsidy an individual is eligible to by using the reported time at the job, the age when the worker entered the firm, the gender and region of residence of the worker.[10] Now, in the EFF we do not know whether the worker was unemployed before entering the firm or got the subsidy through a contract conversion. Hence, we impute to each individual the maximum amount of both subsidies to contract conversion and to hire unemployed workers. Second, the timing when the subsidies impact contract conversion is not unambiguous. A subsidy to hiring unemployed workers could only affect the chances of being observed today with an open-ended contract in the first year of the labor relationship. Nevertheless, the subsidy to contract conversion could affect the chances during several years, because firms could use the subsidy to contract conversion in any year after hiring the worker. We have used the average subsidy available in the region during the first and second years of the time at the firm. Our decision was guided by the evidence from the 2003-2004 waves of the quarterly Spanish Labor Force Survey, that suggests that 18% of employed heads of households covered by high dismissal cost contract were first contracted using a fixed-term contract. Within that group, 90% of workers had their contract converted into an open-ended one during the first two years of tenure at the current firm.[11] We convert the monetary magnitude of the subsidy into constant euro of 2005 using regional deflators of household gross disposable income.[12]

[9] We suspect that the impact of those restrictions will be small, given the small number of cases involved.

[10] Due to confidenciality reasons, region of residence is not available in the public version of the EFF.

[11] We also experiment using an average of the subsidy amount during the first three years at the firm, with little impact on the results. Additionally, we experimented including separate variables for the maximum subsidy available in each of the first two, three and fourth year at the firm. Nevertheless, possibly due to the limited sample size, we could not identify separately the impact of each subsidy.

[12] The database used is BDMORES, elaborated by the Spanish Ministry of Finance.

3.1 Summary statistics

Table 1 splits our EFF sample according to our measure of "exposure to unemployment risk". The first group are households whose head is an employee with an open-ended (or high dismissal cost) contract. The second group is composed by households whose head is an employee with a fixed-term contract. All the monetary magnitudes in the paper are expressed in thousands of 2005 euro. The summary statistics in Table 1 illustrate that the group of households headed by an employee with an open-ended contract differs from the group of households headed by an employee whose position is covered by a fixed-term contract. Heads covered by an open-ended contract are older than heads covered by a fixed-term contract (44 vs 40 years of age). They work in their current firm for a lengthier period of time: 14 vs almost 4 years of tenure, and their households also receive higher earnings. Given those comparisons, it is not surprising that households headed by an individual with an open-ended contract are more likely to own a house (87 percent vs 70 percent) and thus have higher wealth-earnings ratios.

The summary statistics presented stress the idea that simple differences in contract status alone cannot be used to test for a precautionary saving motive. Households headed by individuals with an open-ended and fixed-term contracts differ in many of the observable (and, most likely, unobservable) characteristics that one would expect to result in higher wealth accumulation.

4 Methodology

4.1 First stage: Did subsidies increase the pool of workers covered by high severance payment contracts?

We start by examining whether the amount of the subsidy to hire a worker using a high severance payment contract during the first two years of the contract relationship is a good instrument for the prevalence of such contracts. We examine the response of households to both the job risk of the head and of the working spouse, if one exists.

The first-stage regressions are thus linear probability models for whether the head or the working spouse are covered by an open-ended contract, estimated using pooled OLS regressions.

$$Open\_ended^h = \alpha_0 + \alpha_1 Subsidy^h_{R,g,t_0} + \alpha_2 Subsidy^h_{R,g,t_0} \cdot 1(Age^h \geq 35) + \\ \alpha_3 Subsidy^h_{R,g,t_0} \cdot Female^h + \sum_{g=1}^{g=4} \alpha_{4,g} Age^h_g + \alpha_5 Female^h + \\ \alpha_6 Hired\_post97^h_{t_0} + f(Tenure^h - 3) + X'\alpha_7 + \varepsilon;\ h = \text{head, spouse} \tag{C.1}$$

For simplicity of exposition, we remove the subindices i and t denoting households and time, respectively, from all equation variables. The dependent variable indicates whether the worker is observed with a high severance payment contract. The function $f()$ is a third-order polynomial of $Tenure$, the time spent working at the current firm. $Tenure$ is a key covariate, that allows us to compare workers who entered at the firm in the same year.[13] The key variable is $Subsidy_{R,g,t_0}$, denoting the average maximum statutory amount a firm could get by converting a fixed-term contract into an open-ended one during the first two years of tenure of the worker. Note that subsidies vary across regions (indexed by R), age group (indexed by g) and the time when the contract started (indexed by t_0). We interact the subsidy with a dummy for workers under 35 and with the gender of the household head (and we also include separate dummies for 10-year band age groups of the head and a separate dummy for female).[14] We differentiate the impacts by age of the household head because heads under 35 years of age are very selected groups in Spain (50% of working young adults below 35 live with their parents). X is a vector of covariates that includes three dummies with the educational level attained by the worker (primary education or less (omitted category), first stage of secondary education, upper secondary school and college), the logarithm of household income, indicators of the household size up to six members or more and an indicator of whether the spouse or partner

[13] Ideally, we could also control for time at the firm in a non-parametric fashion by including tenure fixed-effects, but our sample size is a bit small to allow for this. As a validation check, we replicated regression (C.1) using the Spanish Labor Force Survey (EPA) and found that controlling for tenure fixed effects or for a third-order polynomial yielded very similar results.

[14] García-Pérez and Rebollo-Sanz (2009a) have documented that the impact of the subsidies on contract conversion varied with the age and gender of the worker. In particular, they find limited effects among young male workers.

of the household head is employed.[15] It also includes a dummy indicating if the contract was signed after 1997 ($Hired\_post97_{t_0}$) to controls for the fact that subsidies started in some regions in 1997 and that in 1997 there was a national-level reform that introduced a new set of open-ended contracts with lower firing costs and established a set of payroll deductions to the conversion of fixed-term contract into open-ended ones (see Kugler, Jimeno and Hernanz, 2002). By including the post-1997 dummy, we make sure that α_1 captures mainly regional variation in the availability of subsidies to open-ended contracts. The error term of the equation is denoted by ε.

Coefficients of interest: The coefficients of interest are α_1 and the interaction terms α_2 and α_3. The parameter α_1 is an intention-to-treat effect that measures the impact of the statutory amount of the subsidy to open-ended contract on the probability that a male worker between 36 and 65 years of age is currently covered by an open-ended (high dismissal cost) contract. The parameter α_1 is identified by comparing the chances of being observed with a high dismissal cost contract of two workers hired at the same time, but whose employer had access to different subsidies due to (a) being hired in a different region or (b) belonging to a different age group at the time of the hire or (c) belonging to a different gender group at the time of the hire. If the subsidies to open-ended contracts increased the fraction of workers covered by open-ended contracts, α_1 would be positive.

Regional dummies: To avoid the concern that α_1 actually captures the impact of long-run regional characteristics, rather than the firm's incentive to hire the worker using an open-ended contract, we experiment including region dummies. In this second case, identification of the parameter α_1 is achieved by comparing the relative chances of having currently covered by an open-ended contract of workers who were hired in the same year by a firm within the same region, but who belong to different demographic groups that were entitled to different levels of the subsidy. Finally, to avoid any remaining trends in contract form due to different industrial specialization across regions, we also include the unemployment rate in the region in the gender and age-band of the worker at the time of the hiring.

Arguably, the dependent variable is binary, and linear methods may

[15] Household earnings is a rather dubious regressor, because one would expect that workers who are able to obtain a high firing cost contract are selected by the firm on the basis of characteristics that are unobserved by the econometrician and that may also lead to a higher wage. Nevertheless, excluding income from the first-stage regression has little impact on our estimate of α_1.

present problems of extrapolation outside the 0-1 range. Still, we present results from OLS specifications because the literature has provided a variety of tests of quality of instruments in a linear setting (see Staiger and Stock, 1997).

Impact on transitions from employment to unemployment: Firms could upgrade fixed-term contracts into open-ended ones just to benefit from the subsidy and dismiss those workers afterwards. Were that the case, a job with a high nominal firing costs would not necessarily be more stable than a job regulated by a fixed-term contract (see García-Pérez and Rebollo-Sanz, 2009b). To test for the presence of "churning" effects, we estimate gender-specific regressions similar to (C.1) but using as a dependent variable an indicator of whether the worker experienced a transition from employment to unemployment. We use the Spanish Labor Force Survey to have a large sample. The coefficient of interest is α_1, which measures if workers who, at the beginning of the current job spell, belonged to a period, demographic group and regions where more generous subsidies were available have a lower probability of transiting from employment into unemployment. As a transition from employment to unemployment is a low probability event, we experiment using both OLS and Probit models.

4.1.1 Intention-to-treat effects: Did regional subsidies reduce the amount of household wealth?

We examine intention-to-treat responses of (the logarithm of) the household wealth to earnings ratio to the statutory amount of regional subsidies when the worker was hired. The exact model we estimate is the following:

$$\begin{aligned}
\log(\frac{W}{Y}) \quad = \quad & \delta_0 + \delta_1 Subsidy^{head}_{R,g,t_0} + \delta_2 Subsidy^{\text{spouse}}_{R,g,t_0} + \delta_3 Subsidy^{head}_{R,g,t_0} 1(Age^{head} \geq 35) \\
& + \delta_4 Subsidy^{head}_{R,g,t_0} \cdot Female^{head} + \sum_{g=1}^{g=4} \delta_{5,g} Age^{head}_g + \delta_6 Female^{head} \\
& + \delta_7 Hired_post97^{head}_{t_0} + g_1(Tenure^{head} - 3) + g_1(Tenure^{\text{spouse}} - 3) \\
& + X'\delta_8 + u \qquad \text{(C.2)}
\end{aligned}$$

Following previous studies, we use as the dependent variable is the logarithm of the ratio of household wealth to earnings of head and spouse, if one is present (3.2 percent of the number of original households were lost due to taking logs).[16] As we discussed above, there are reasons to examine the response of various measures of household wealth to the risk of losing the job. We present the results sequentially starting with the strictest measure of wealth that can be cashed. According to the model briefly discussed in Section 2, the coefficients associated with the risk of losing a job, δ_1, should be negative: workers who (for exogenous reasons) obtained a contract that protects them from transiting into unemployment end up holding lower amounts of precautionary wealth.

In equation (C.2) the alternative measures of household wealth are regressed on the variables based on subsidies and on all covariates introduced in the first-stage equation (C.1). The error term of the wealth equation is denoted by u. As in the first stage equation, we examine the robustness of the estimates of δ_1 when we include region dummies and the statutory amount that the spouse was eligible for.

4.2 Instrumental variables: how much more wealth do workers covered by low firing cost contracts hold?

We estimate the causal impact of the risk of losing the job on the household wealth by the method of instrumental variables. The OLS estimates of equation (W1) would be biased upwards for the various reasons mentioned in the Introduction. Thus, we quantify the average impact of holding a high dismissal cost contract on the amount of wealth held using Two Stage Least Squares estimates.

$$\begin{aligned}\log(\frac{W}{Y}) &= \gamma_0 + \gamma_1 Open\_ended^{head} + \gamma_2 Open\_ended^{spouse} \qquad \text{(W.1)}\\ &+g_2(Tenure^{head} - 3) + g_2(Tenure^{spouse} - 3) + X'\gamma_2 + v\end{aligned}$$

where $Open\_ended^{head}$ and $Open\_ended^{spouse}$ are instrumented separately using linear probability models like (C.1). The parameters of interest are γ_1 and γ_2, which measure the response of (the logarithm of) household

[16] We tested if normalizing by current income was restrictive by examining the sensitivity of the estimate of δ_1 in a specification where the dependent variable was the logarithm of household wealth, but that was otherwise similar to (C.2). The results hardly changed. Rather than constructing measures of permanent income (noisy in the absence of long panel data), we used last year's earnings and control for determinants of permanent income, like three dummies with the educational attainment of head and spouse, if the latter is present.

wealth over household earnings to holding a high dismissal cost contract. The causal estimation of this coefficient only exploits variation in open-ended contracts that is due to the fact that firms faced different incentives to use those contracts in different years, regions and demographic groups. The error term is denoted by v. Finally, we quantify how many months of household earnings are kept as precautionary wealth by households relatively more exposed to a job loss by multiplying γ_1 by the unconditional median wealth-income ratio held by households with a fixed-term contract: 0.125.[17]

Response at different quantiles: A priori, there is little reason to expect that the response of household wealth to the risk of losing the job is similar across centiles. The precautionary buffer stock model suggests that the consumption-cash on hand relationship is concave, with larger share of precautionary wealth for households with small cash-on-hand holdings. Secondly, whether the response of wealth is due to a large reaction by a small set of households or it is widespread among the population matters for the path of consumption during an unemployment spell, because most households would have accumulated little in the first scenario. We address the issue estimating Instrumental Variable-Quantile Regression Models of the response of wealth to the risk of losing the job (see Appendix for details).

5 Results

5.1 The quality of the instrument

Table 2 presents OLS regressions of the type of contract held on our key identifying variable: the statutory subsidy amount that the firm could get in the first two years of the contract in the region where the household lives. The standard errors are presented in parentheses and take into account that there can be group correlation in the error term due to the imputation at the region, age, gender and year of initial contract of the variable $Subsidy^{head}_{R,g,t_0}$ -see Moulton (1986).[18] In Table 2, row 1, column 1, the estimate of the

[17] We also experimented evaluating the results taking antilogs in (W.1), and estimating the amount of precautionary wealth as: Precaut_wealth= $\exp(\gamma_0)[1-\exp(\gamma_1)]$

This is a first-order approximation that ignores the variance of the residual of the logarithm of wealth. The results were similar to those reported here.

[18] Unless otherwise noted, all specifications reported in the paper make use of the five imputed datasets of the EFF. Namely, the point estimates shown in Tables 2, 4, 5, 7, 8 and 9 are the average across the five datasets of the EFF. As for the standard errors, the

variable $Subsidy^{head}_{R,g,t_0}$ is 0.018 (standard error: 0.005). The estimate implies that an increase in the subsidy to the conversion of fixed-term contracts into open-ended ones in the first two years of the life of the contract increases the chances of observing the worker being covered by open-ended contracts by 1.8%. The estimate is significant at the 1 percent confidence level, and the F-statistic is 13.91.

The estimate of the interaction of $Subsidy^{head}_{R,g,t_0}$ and a dummy for age of the head below 35 years is -0.013 which, combined with the 0.018 estimate implies a limited impact for that group: an increase in the subsidy of 1,000 euros increases the stock of young workers covered by an open-ended contract in 2002-2005 by 0.5%. The point estimate is also lower among female heads: 1,000 extra euros increase the stock of female heads with a high dismissal cost contract by 1.7% (=0.018-0.001). While females are the most benefitted group from the subsidy, one must take into account that the group of female heads of household is arguably a very selected one according to our definition of household head in the EFF.[19]

Specification 2 in Table 2 adds household earnings as a covariate, without noticeable impact on the estimates, and specification 3 adds sixteen regional dummies, with Madrid as the excluded group. The estimate of the variable $Subsidy^{head}_{R,g,t_0}$ is now 0.011 (standard error: 0.005), shown in row 1, Column 3 of Table 2. That is, workers belonging to demographic groups that were entitled to a subsidy 1,000 euros higher than a benchmark group in the same region are 1.1% more likely to be observed in 2002 and 2005 with an open-ended contract. The F-statistic of the instrument in this new specification is 4.97, resulting in a weaker instrument than in the previous specification.

Specifications (4) and (5) turn to the group for whom the instrument is strongest: the sample of households headed by a male. Within such group, an increase of 1,000 euros in the variable $Subsidy^{head}_{R,g,t_0}$ predicts the share of head male employees covered by an open-ended contract increases by between 1.4 percentage points (standard error: 0.005) and 1.9 percentage points (standard error: 0.005) in specifications that include and exclude region dummies,

Huber-White correction for group-correlation is computed at the household, year-at-the job, age band, region and gender level for each replicate. The five standard errors obtained are subsequently combined using the adjustment suggested by Li et al. (1991).

[19] We use the definition of household head provided for the EFF by Banco de España (2005). The household head is defined as the reference person designated by the household for replying to the survey except for the case that the reference person is a woman and her partner lives in the household, in such a case the household head is the partner.

respectively. Columns (6) and (7) investigate the impact of the incentive to hire using high firing cost contracts on the contract form of both head and spouse, respectively. Column (6) examines if the estimate of $Subsidy^{head}_{R,g,t_0}$ is affected when we include the subsidy that the spouse qualified for, using a sample of male heads. The estimate is 0.02, very similar to the estimate in Column (4) of Table 2.

Response of other family members: Finally, column (7) shows the estimates of the model (C.1) using as the dependent variable the indicator of whether the spouse has a high firing cost contract.[20] The coefficient $Subsidy^{spouse}_{R,g,t_0}$ is 0.032, suggesting that an extra 1,000-euro subsidy in the first two years of the contract increases the chances of observing the spouse covered by a high firing cost contract by 3.2 percentage points. The impact is larger than among male heads. Reassuringly, the subsidy that the head (spouse) was eligible for explains little of the variation in current contract form of the other member of the couple, spouse (head), suggesting that our instrument is not picking up spurious regional trends unrelated to the program.

Overall, we conclude that the instrument "subsidy to open-ended contracts" works best for the sample of mature male heads and their female spouses. The subsidies to young male heads of households seem to have had a lower impact on current contract form.[21]

The response of employment stability: Table 3 gives further evidence from the Spanish Labor Force Survey regarding the impact of exposure to higher statutory subsidies at the beginning of the contract on subsequent employment stability. The OLS estimate of $Subsidy^{head}_{R,g,t_0}$ is shown in Column 1 is -0.0010 (standard error: 0.0005) significant at the 5 percent level. The point estimate confirms that male employees who could benefit from higher subsidies at the beginning of their contract indeed ended up in more stable

[20]The sample we use in this case contains both married and single households headed by a male. The reason for doing so is that the sample size is not large enough. We also include in the sample married females who do not work (introducing a dummy indicating whether or not the spouse works).

[21]A way of rationalizing the differential effects of subsidies to contract conversion on young adults and mature workers is the following: an employer, considering whether or not to convert a fixed-term contract into a high dismissal cost one, may decide to postpone the decision for a young worker until more information about the productivity of the match is revealed. Nevertheless, in the case of a mature worker, previous labor history and references reveal much information about the employee's expected future productivity of the match, so contract upgrades are more sensitive to labor costs.

jobs. The probit estimates for males are smaller: they imply a reduction in the probability of job loss of -0.0005 (standard error 0.00016), shown in the first row, Column 2 of Table 3. The evidence for female spouses also points at more stable jobs, albeit the estimate are not significant at the 5% in all specifications.

5.2 The response of wealth to the risk of losing the job.

Table 4 documents intention-to-treat estimates of the response of "liquid" household wealth to the incentive to convert low dismissal cost contracts into high dismissal costs one ($Subsidy^{head}_{R,g,t_0}$) as shown in equation (C.2). The estimate displayed in the first row and first column of Table 4 shows that a higher incentive to hire using open-ended contracts diminishes household financial wealth by 5.4 percentage points (standard error: 2.4 percentage points). The estimate is consistent with the notion of precautionary wealth holdings: groups of the population that experience an exogenous increase in the degree of protection of their job accumulate less financial wealth. The estimate of the interaction between the variable $Subsidy^{head}_{R,g,t_0}$ and an indicator of the household head aged below 35 is 0.060 (standard error: 0.032), positive but not very precise. Adding this estimate to that of the variable $Subsidy^{head}_{R,g,t_0}$ yields an estimate of 0.006 (=-0.054+0.060), suggesting that the incentive to hire using open-ended contracts for workers currently below the age of 35 does not reduce precautionary wealth (it even increases wealth in 0.6%), a small number statistically not different from zero. The estimate suggests very limited wealth responses among the group of workers below 35 years of age. Such lack of response is again reassuring, as subsidies to open-ended contracts were not powerful predictors of the stock of young workers covered by a high dismissal cost contract, those who are head of household aged below 35. Column 3 of Table 4 introduces indicators of the region of residence.[22] The estimate of the variable $Subsidy^{head}_{R,g,t_0}$ is -0.036 (standard error: 0.027), still negative and consistent with a precautionary saving motive, but not significantly different from zero.

The fourth column presents results from our preferred sample, that composed by male heads. The estimate in Table 4, row 1, column 4 of the

[22] Region indicators allow to control for unobserved characteristics that correlate with wealth, like tastes of inhabitants in a particular region.

instrumental variable $Subsidy^{head}_{R,g,t_0}$ is -0.066 (standard error: 0.025), negative and significantly different from zero at the 1 percent confidence level. The result is smaller but similar when we add regional indicators: -0.045 (standard error 0.029), shown in Table 4, column 5, row 1.

Finally, the sixth column of Table 4 examines the separate responses of household wealth to the subsidies that the firm of the head and spouse could benefit from. An increase of 1,000 euro in the statutory amount of the subsidy that the head was eligible for during the first two years of the contract diminishes current wealth-earnings ratio by 4.8 pp. Perhaps surprisingly, the point estimate of the response of household wealth-earnings ratio to the incentive for the spouse to have a high firing cost contract is even significantly larger: -0.075 (standard error: 0.024).

Overall, the results in Table 4 suggests that households headed by a male employee over the age of 35 react to variables measuring an exogenous increase of the probability that *either the head or the spouse* are protected from lay-offs by accumulating less wealth in "liquid" financial wealth. We find less evidence of responses among households headed by females or by younger workers.

5.2.1 Two Stage Least Squares Estimates

Table 5 presents OLS and Two Stage Least Squares estimates of the magnitude of the average response of financial wealth to holding a low dismissal cost contract. Table 5 Panel A presents first-stage estimates of how much it is more likely to observe a worker with an open-ended contract due to regional subsidies for the conversion for each of the groups considered, and Table 5 Panel B examines how much households reduce their (log) wealth-income ratios when the head holds a high dismissal cost contract. The estimates in both panels are done using the same controls as those shown in Tables 2 and 4, but we only display the parameters of interest: the causal impact of having a high dismissal cost contract on the logarithm of financial wealth over earnings ratio.

The OLS estimate of the impact of "open-ended contract" on the log of household wealth to income ratio is -0.039 (standard error: 0.095). Multiplying that estimate by 0.125 (the median wealth-income ratio), the estimate suggests that workers covered by fixed-term contracts have half a percentage point of the total household earnings as financial wealth holdings more than comparable workers covered by open-ended contracts (Table 5, Panel C, row

1, column 1). Above we have already discussed the possible downward biases of the OLS estimates of precautionary wealth holdings.

The TSLS estimate of the causal impact of the head holding an open ended contract on wealth-earnings ratio is -2.341 (standard error: 1.477), and is shown in Table 5, Panel B, row 1 column 2. Evaluated at the median wealth-earnings ratio, the estimate suggests that households headed by a male with a fixed-term contract hold ratios of financial wealth over earnings that are 29.3% higher than households where the head has an open-ended contract. The stark difference between the OLS and the TSLS estimates may be due to fact that the latter identifies wealth responses from a set of workers who started their job spell as fixed-term contract employees. Such individuals are likely to have had labor histories similar to those of workers who are currently covered by fixed-term contracts, so biases due to wealth used prior to their current job are mitigated considerably.

Our estimate of the variable "Head covered by high firing cost contract" becomes larger when we examine households headed by male workers. On average, the average log-wealth-earnings ratio held by households headed by a male worker with a fixed-term contract exceeds by 3.44 that held by workers covered by open-ended contracts (Table 5, Panel B, row 1, column 4). For that particular group, the average buffer of liquid wealth exceeds that of open-ended contract by 43 percent, or 5.2 months' income (a substantial amount).

Finally, in Column (6), we present the household wealth to earnings response to the exposure of the risk of losing the job by the head and spouse (if present). The point estimate of the impact of exposure to head's job loss on the log of the wealth-earnings ratio is -2.67 (standard error: 1.48). The response of household wealth to the risk that the spouse loses the job is strinkingly similar in magnitude. In both cases, the wealth accumulated in response to the risk that each member loses the job is similar: around 4 months of yearly earnings.

The response at various points of the wealth distribution The results of the impact of whether both members of the couple, the head and spouse, have a high dismissal cost contract on the log wealth to earnings ratio at various centiles are shown in Table 6. In the case of the household head, the response is most precisely estimated at bottom half of the wealth distribution than at the 75th centile, where the estimate is large and imprecise. Hence, we concentrate on the responses at the bottom half of the wealth distribution. Being covered by a high firing cost contract diminishes the 25th centile of wealth-earnings ratio by 1.7, but the median wealth-earnings ratio by 1. Both magnitudes represent around 10% of household earnings as extra wealth held when the head of household is exposed to unemployment risk, an estimate that is below that estimated for the mean. The response is proportionally larger at the bottom of the wealth distribution. That pattern is consistent with a precautionary saving motive, as we document below.

Panel B of Table 6 shows the response of household wealth to the risk that the spouse loses the job at various points of the wealth distribution. Again, they are very much in line with those found by the head, and we do not comment them in detail.

Overall, our estimates of the impact of exposure to job loss on average wealth-earnings ratios are lower but comparable to those in Carroll, Dynan and Krane (2003), who estimate that households in the US react to a percentage point increase in the risk of losing the job by accumulating between 2.5 and 3 months of income. The average response we estimate is also around 33% of household earnings, but the point estimates for most of the wealth distribution are in the ballpark of 10% and one must take into account that the differences in the chances of transiting into unemployment we exploit are much larger than the differences in unemployment flows between different workers in the US in the study of Carroll et al. (2003). Our estimates are definitely larger than those of Engen and Gruber (2001), who estimate that reducing UI to a half would increase financial wealth holdings by $241. The magnitude of the buffer we estimate is between 1,770 euros (=0.10·17,700, assuming the median wealth response of 10% off average household earnings among workers covered by fixed-term contracts) and 5,841 euros (assuming the mean wealth response of 33%) -see footnote 3.

5.3 Robustness checks

Alternative measures of wealth Table 7 conducts a series of robustness checks to the specification (4) in Table 5. We start by falsifying our empirical strategy by using as an instrument the subsidy amount during the fourth year of the contract. Very few conversions happen at the fourth year at the firm, according to the Spanish Labor Force Survey (EPA), so the variation in current labor status generated by an inappropriate instrument should have little impact on accumulated household wealth. Otherwise, we could be suspicious that the instrument is picking up particular regional trends in wealth and employment quality. The estimate shown in Table 7, Column 1, Row 1 is close to zero: 0.854 (standard error: 1.251). Hence, we infer that the results in Table 5 are unlikely to be driven by spurious regional trends.

The results in Table 5 still hold when we use an expanded wealth measure that includes gross financial wealth and net housing wealth that excludes home equity. The point estimate of the coefficient "Head covered by high dismissal costs contract" in Table 7, Row 1, Column 2 is -3.56 (standard error: 1.81). The estimate is slightly smaller than that using financial wealth, and is consistent with a "buffer stock saving" of about 45% of gross annual earnings. Nevertheless, they do not hold when we use net wealth as a dependent variable (excluding non-cashable pension funds and life insurance products). The point estimate is shown in Column 3 of Table 7 and is 1.77 (standard error: 1.44), positive and not significantly different from zero, contrary to the hypothesis that workers react to the risk of losing the job by saving more total wealth. The result is due to the fact that subsidies seem to increase the probability that a worker is observed as a homeowner (not shown).

We find two possible explanations for that finding. First, workers covered by a fixed-term contract are less likely to obtain credit from banks and accumulate wealth through lower consumption to purchase a house. Alternatively, prudent households refrain to borrow to invest in owner-occupied housing, because the net value of the asset is hard to cash in the event of an involuntary job loss. We disentangle between both hypotheses below.[23]

Do the estimates reflect the prevalence of credit constraints? Next, we examine how loan rejections vary with the statutory subsidy to open-ended contracts. Following Jappelli (1990), we identify three forms of credit constraints. The first is whether the household did not ask for a loan during the last two years because the loan was thought to be rejected. The second is whether the household asked for a loan, but was rejected. The third form of credit constraint is whether the loan was not rejected, but the household was given a lower amount than the one asked.

[23]We have also examined whether regional subsidies to contract conversion may have affected wealth through other channels, like household earnings and female labor force participation. We estimated regressions otherwise similar to (C.2), but when the outcome variable was whether or not the spouse participates in the labor market and overall household earnings (both specifications excluded household earnings from the set of regressors). We obtained small coefficient estimates not being statistically different from zero.

We estimate a multinomial logit model with five different outcomes: the three mentioned above and two new ones: not having asked for a loan for the last two years and having asked for a loan with the application accepted. If the estimates in Table 5 and 6 are picking up the responses of credit constrained households, we would expect that the variables $Subsidy^{head}_{R,g,t_0}$ and $Subsidy^{spouse}_{R,g,t_0}$ are associated to a drop in the relative chances of being credit constrained. The set of regressors is the same shown in the fourth column of Table 2.

The first row of Table 8 shows the summary statistics: 28.1 percent of the households have requested (and obtained) a loan in the last two years during the sample period, 1% of the households did not ask because they feel they would be rejected, 0.8% were actually rejected, and 1.4% got less than what they asked. According to this measure, 3.2 percent of all households, who represent 10.2 percent of the potential loan applicants [=3.2/(28.1+3.2)·100] were credit constrained.

Model 1 uses the actual form of contract as the key regressor. The cells shown are the predicted probability of the outcome in each column for a childless household with two earners with three years of tenure at their job, the main earner is between 36 and 45 years of age, have both completed basic schooling and earn average earnings. The results of Model 1 in Table 8 suggest that households headed by a male employee covered by a fixed-term contract are more likely to be credit constrained: 8.4 percent among all applicants holding fixed-term contracts while only 3.8 percent among households headed by an open-ended contract. The stronger presence of liquidity constraints among fixed-term workers may be due to factors other than the higher risk of losing the job, like a higher propensity of having experienced past unemployment spells that make applicants with a fixed-term contract less able to accumulate collateral. To avoid biases due to such unobserved factors, Model 2 replaces the indicator of having a permanent contract by an exogenous determinant of the risk of job loss: the regional subsidies to the conversion of fixed-term contracts into open-ended ones.

The fitted probabilities shown in Model 2 implies that subsidies did not move the fraction of households that are credit constrained. Among households that are not eligible for subsidies, the fraction of liquidity constrained households is 5.7 percent (Table 8, Model 2, Column 6, row 1). Among households whose head or spouse are eligible for a 1,000 euro subsidy, the fraction of credit constrained households is 5.2%, slightly *smaller* (but the difference is small and not significantly different from zero). Hence, there is little evidence that the reduction in the risk of losing the job provoked by the regional subsidies changes the liquidity constraints faced by households. Therefore, the accumulation of more liquid wealth among fixed-term workers is unlikely to be generated by the impossibility of investing in real estate due to their more likely rejection of the loans applied for.

5.4 Additional evidence from consumption growth

We test if consumption growth is higher among workers more exposed to the risk of losing the job - equation (2) in Section 2. We select a sample of 625 employed individuals in the 2002 wave of the EFF who are covered either by open-ended or by fixed-term contracts. We use the panel component of the EFF to track these individuals into the 2005 wave, *regardless* of whether they are employed or not (a key element of the prediction of higher consumption is that the expectation of consumption growth should be taken among all possible states, including unemployment). The exact model we run is

$$\log(c_{2005,i})-\log(c_{2002,i}) = \gamma_0+\gamma_1 Opterm^{head}_{2002,i}+\gamma_2 Opterm^{spouse}_{2002,i}+\delta X_i+u_{2005,i}-u_{2002,i} \tag{3}$$

The dependent variable denotes consumption growth over a three-year period of household i between 2005 and 2002. The information in the EFF permits us to construct three consumption measures: regular food expenditure -based on a recall question about regular amounts spent on food without disaggregating between food at home and at restaurants, and a recall question on total non-durable consumption. The third measure adds to the former an imputation of consumption on durable goods. The EFF asks households separate questions about the value of furniture and home appliances and about the value of their stock of cars. Using the depreciation rates in Fraumeni (1997), one can obtain an imputed value of the flow of services associated to the stock of those values.

The key covariates denote measures of the variance of the income process by including an indicator of the type of contract held by the head of the household and the spouse (if present) in 2002. Households whose head (or spouse) were covered by an open-ended contract in 2002 were exposed to ex-ante higher job security than workers covered by a fixed-term contract. If precautionary saving motive is present in the data, γ_1 and γ_2 ought to be negative.

As additional covariates that are likely to impact the marginal utility of consumption, we include detailed demographic indicators, like four dummies with the age of the head in 10-year brackets, changes in household size and in household composition. Finally, to control for the fact that more educated individuals may have different rates of patience, we also include three dummies with the educational attainment of the household head.[24]

The results are shown in Table 9. Across all consumption measures, we find significantly higher average consumption growth among households headed by an employee covered by a low dismissal cost contract than among those whose head is covered by a high dismissal cost contract. The estimate of γ_1 shown in row 1, column 4 of Table 9 is -0.18 (standard error: 0.073). This suggests that, over a three-year period, households headed by an employee covered by a high dismissal cost contract would experience basically no increase in non-durable consumption (adding up the constant in column 4 of Table 9 to the estimate of γ_1 results in an estimate of 0.190-0.183=0.007). Conversely, households headed by an employee covered by a low dismissal cost contract would experience consumption growth of 19 percent between 2002 and 2005.

6 What model of wealth accumulation is consistent with our findings?

As stressed in Section 3, differences in severance payments in case of worker's lay-offs affect household wealth-earnings ratios through two main channels: different falls in their expected future income and the response of prudent households that accumulate more wealth to buffer against the risk of changes in their marginal utility to consume. Understanding which of those two channels prevails is important, because uncertainty about future income losses leads to a welfare loss above and beyond lower expected income. This section addresses the issue comparing our estimates to the predictions of models with and without precautionary savings. We keep the model deliberately simple and abstract from a number of issues (like demographics or pension

[24] A concern with the test is that we do not use exogenous variation in $Opterm_{2002}^{head}$ or $Opterm_{2002}^{spouse}$ to conduct the analysis. The reason is that the instrument used so far is not a powerful predictor of contract form in this (small) sample. Hence, these results should be viewed as suggestive.

arrangements) to stress whether a model readily comparable to others in the literature can generate wealth responses similar to those in the data.

6.1 A model with uncertainty

We simulate a following (simple) buffer stock model as in Carroll (2001). Assume that individuals live forever and solve the following problem

$$\max_{c_t} U = \sum_{t=0}^{t=\infty} \delta^t E_0 \left[\frac{(c_t)^{1-\rho}}{1-\rho} \right]$$

$$A_{t+1} = (1+r)[A_t - c_t] + Y_{t+1}$$

$$Y_t = GPS_t Y_t^P \qquad Y_t^P = N_t Y_{t-1}^P$$

A_t is the level of beginning-of-period wealth (that we assume to be liquid), and r is the riskless interest rate. P is a binary random variable representing the chances of transiting into unemployment. If P equals 1, the temporary income level equals $S_{\min}$. G denotes permanent income growth. S_t is an independent and identically distributed (iid) lognormally shock with mean $\frac{1-E(P)*S_{min}}{1-E(P)}$ and the standard deviation of the associated normal distribution is σ_S.[25] N_t are iid, log normally distributed shocks with mean zero and standard deviation of the associated normal distribution $\sigma_N = 0.1$. We simulate the model using the usual strategy of normalizing the state variable A_t by the level of permanent income Y_t^P.

Parameter choice: A period in the model is assumed to be one year. The coefficient of relative risk aversion, ρ, is set to be 2, following standard estimates of the elasticity of intertemporal substitution. We set r as 0.02, the discount factor, δ, as 0.95 and G to be 1.[26] We also set $V_{\min}$ to be 0.6, implicitly assuming a 60% replacement rate of the level of permanent income. We conduct separate simulations for different values of P: one where the annual probability of transiting into unemployment is 0.02 (intended to

[25] By using that mean, we ensure that we keep the mean of the process constant when we change the probability of losing the job, P.

[26] Several sources point out that labor income growth in Spain has been at best modest since 1995 (see, for example, Carrasco, Jimeno and Ortega, 2008). Using the EFF, Bover (2008) documents zero earnings growth on average in the panel component of the EFF.

mimick the unemployment chances of workers covered by open-ended contracts) and another where the probability of transiting into unemployment is 0.10 (unemployment chances of workers covered by fixed-term contracts). We obtain those estimates from regressions in the Spanish Labor Force Survey predicting chances of transiting into unemployment by household heads above 35 years of age. We follow Casado-Garcia (2009) who estimates a value of $\sigma_S = 0.17$ and $\sigma_N = 0.1$. using a Spanish rotating panel of consumption data in Spain. See Appendix 2 for details.

Results: Table 10 presents the simulated distribution of household wealth over permanent earnings in both scenarios: when the probability of job loss equals 0.02 and when it equals 0.10. Note that in such simple exercise, we keep the mean of the process constant and also ignore the role of severance payments in helping sustain consumption during an unemployment spell. The first column in Table 10 presents the distribution of wealth in the steady state when the probability of experiencing unemployment risk is 2%, and the second column presents the same distribution when the probability is 10%. To match the empirical estimates, those distributions are computed among consumers with three periods of "tenure" (three periods not in unemployment). On average, the wealth-earnings ratio of workers covered by fixed-term contracts exceeds that of workers covered by permanent contracts by 0.24. The simulation results also suggest that wealth responses are proportionally larger at the bottom of the wealth distribution, possibly due to the concavity of the consumption policy function. We find both findings surprisingly in line with our estimates.

6.2 A model without uncertainty

Finally, we compare our estimates to those generated by a model *without* uncertainty. Carroll, Dynan and Krane (2003) show that the consumption policy of an infinite-lived consumer when the market interest rate r equals the discount rate β is $C_t = \frac{r}{1+r}A_t + \frac{r}{r-g}Y_t$. A_t is cash-on and Y_t is earnings as in the previous section. Sticking to our assumptions that r equals 2% and there is no growth in permanent income, $\frac{r}{r-g}$ equals 1. The expected income drop would be $E\Delta Y_t = \frac{r}{(r-g)} \cdot P(jobloss) \cdot (time\_unemployed + wage\_drop)$. Assume also that the event we are considering is a job loss that lasts two quarters, followed by a subsequent permanent wage loss of 10% in their following job (an unrealistically high wage drop, as workers covered by fixed-term contracts accumulate little specific capital, given the tenures reported in Table 1). In a certainty equivalence world, consumption of workers covered by a fixed-term contract would fall–and wealth would rise–by 1·0.10·(0.5+0.1)=0.06. The corresponding loss for a worker covered by a high dismissal cost contract is lower for various reasons. First, that worker has a lower exposure to risk. Second, in the event of job loss, the worker would receive a severance payment that may vary between 20 days per year worked and 45 days. To make the results more unfavorable to our case, we choose the 45 days compensation package and assume 3 years of tenure. The consumption drop would be 1·0.02·(0.5+0.1-45·3/365)=0.0046. That is, the excess wealth-earnings ratio held by a worker with a fixed-term contract would be 0.055 (=0.06-0.0046), while our estimate is between 0.29 and 0.40, between 5 and 7 times larger.

7 Conclusions

We use the large dispersion in dismissal costs in the Spanish labor market and a new data set of household finances to estimate the link between the probability of losing the job and household consumption and wealth. We obtain exogenous variation in the type of contract by exploiting the different timing and target groups of regional subsidies for firms that hired workers using open-ended contracts. Our results suggest that households whose two main earners exogenously obtain a high dismissal cost contract accumulate less financial wealth than other comparable households where members are exposed to unemployment risk. The magnitude of the wealth response is similar for the exposure to the risk of losing the job for heads and spouses and amounts to around 30% of gross yearly earnings. Instrumental Variable Quantile regression estimates suggest that the response of gross financial wealth is proportionally larger at the bottom of the wealth distribution.

We do not find that workers covered by high dismissal cost contracts accumulate more wealth when the net value of owner occupied housing is included in the measure, a fact that we attribute to a preference for saving in assets that are easier to cash. Finally, simple simulations of a buffer stock model suggest that the pattern and magnitude of the responses we estimate are consistent with a model with a precautionary saving motive, but less so with a model with perfect certainty where consumers save in anticipation of expected income losses.

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Appendices

Appendix 1: The implementation of the Instrumental Variable Quantile Regression

In our case, the estimator at the median proposed by Chernozhukov and Hansen (2004, 2008) is based on the following moment conditions:

$$\frac{1}{NT}\sum_i\sum_t\left(1[\log(\frac{W}{Y})_{it} \leq \gamma_0 + \gamma_1 Open\_ended^h_{it} + X'_{it}\gamma_2] - \tau\right)X_{it} \to 0$$

$$\frac{1}{NT}\sum_i\sum_t\left(1[\log(\frac{W}{Y})_{it} \leq \gamma_0 + \gamma_1 Open\_ended^h_{it} + X'_{it}\gamma_2] - \tau\right)Subsidy^h_{R,g,t_0,it} \to 0$$

$$h = \text{head, spouse}$$

The indices i and t denote households and periods, respectively, and N and T indicate the total number of households and periods in the pooled sample, respectively. We implement this estimation method by running a set of standard quantile regressions of the dependent variable $\log(\frac{W}{Y})_{it} - d_1 Open\_ended^h_{it}$ on the covariates X_{it} and $Subsidy^h_{R,g,t_0,it}$ for a grid of values of d_1. We choose $\widehat{\gamma}_1$as the value for which the impact of $Subsidy^h_{R,g,t_0,it}$ is the smallest statistically. The interval considered is [-10,10] and each point in the grid is 0.1 apart from the following one.

We run separate models for head and spouse, holding constant in each case the variable $Subsidy^h_{R,g,t_0,it}$ of the other member of the couple. The confidence intervals are constructed as the set of points in the grid of d_1 for which the coefficient of $Subsidy^h_{R,g,t_0,it}$ is not significantly different from zero for each level of confidence.

Appendix 2: The solution of the precautionary saving model

The model in Section 6.1 is solved by specifying a pre-specified grid of points of $\frac{A_t}{Y_t^P}$, an initial guess of V_{t+1} equal to zero and then iterating the Bellman Equation to find the value function V_t that solves[27]

[27] Exactly, we iterate the Bellman equation while the value function in a given iteration V_t differs from the value of V_{t+1} by more than 1e^{-7}.

$$V_t(\frac{A_t}{Y_t^P}) = \max_{c_t} \quad \frac{1}{1-\rho}(\frac{c_t}{Y_t^P})^{1-\rho} + V_{t+1}[(\frac{A_t}{Y_t^P} - \frac{c_t}{Y_t^P})\frac{(1+r)}{GN_t} + PS_t]$$

After obtaining the optimal sequence of $\{\frac{c_t}{Y_t^P}, \frac{A_t}{Y_t^P}\}$ for the pre-specified grid of points, we simulate 100 periods shock histories of N_t, P and S_t of 1,000 individuals, who start their working life with an initial wealth of two times $\frac{S_{\min}}{Y_t^P}$. We compute their beginning-of period wealth, and consumption in each period by interpolating between the optimal gridpoints $\{\frac{c_t}{Y_t^P}, \frac{A_t}{Y_t^P}\}$.

Table 1: Summary statistics, combined EFF2002 and EFF2005

	Total sample	Head, open-ended	Head, fixed-term
Head with open-ended contract	0.822	--	--
	(0.383)		
Head with fixed-term contract	0.178	--	--
	(0.383)		
Age of household head	43.585	44.355	40.029
S.D.	(9.721)	(9.571)	(9.621)
Married	0.803	0.818	0.736
	(0.398)	(0.386)	(0.441)
Household size	3.220	3.244	3.108
	(1.234)	(1.206)	(1.348)
Prob. job loss (quarter),head			
Mean:	0.029	0.016	0.086
S.D.	(0.033)	(0.011)	(0.041)
# Years at current job	12.607	14.414	4.265
	(10.479)	(10.419)	(5.563)
Household earnings	27.716	29.876	17.745
	(19.247)	(19.816)	(12.120)
Whether head eligible for subsidy	0.289	0.233	0.547
	(0.453)	(0.423)	(0.498)
Amount head is eligible for	0.948	0.790	1.676
	(1.865)	(1.749)	(2.184)
Spouse covered by open-term	0.712	0.763	0.441
(sample of working spouses)	(0.453)	(0.426)	(0.497)
Spouse eligible for subsidy	0.377	0.352	0.509
(sample of working spouses)	(0.485)	(0.478)	(0.500)
Amount spouse is eligible for	1.355	1.250	1.913
	(2.170)	(2.080)	(2.525)
Non-durable expenditure	12.668	13.179	10.308
S.D.	(7.424)	(7.707)	(5.347)
Owns real estate	0.839	0.868	0.701
	(0.368)	(0.338)	(0.458)
Net worth			
Median	122.330	134.204	66.946
Mean	170.244	186.964	93.051
Net worth to earnings ratio			
Median	4.954	5.156	3.569
Mean	7.140	7.004	7.765
Financial wealth			
25th centile	1.097	1.300	0.639
Median	3.841	4.645	1.803
Mean	16.398	18.558	6.424
Financial wealth to earnings ratio			
Median	0.167	0.180	0.125
Mean	0.561	0.581	0.472

Sample: Average sample size of 3,660 household-year observations in five data sets imputed multiply in two EFF waves (2002 and 2005). All statistics weighted.
S.D. are standard deviations. Monetary magnitudes in 1000s of 2005 euro.
Subsidy amounts in real terms using deflators of the regional gross disposable income.

Table 2: First stage: the impact of subsidies to open-ended contracts on the share of open-ended contracts in 2002-2005.

Estimation method: OLS (Linear probability models)							
Sample:	Total sample			Sample of male heads			
Dependent variable: head covered by an open-ended contract							Spouse has open-ended contract
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
1. Mean subsidy amount in two first years of job tenure -head	0.018 (0.005)***	0.017 (0.005)***	0.011 (0.005)**	0.019 (0.005)***	0.014 (0.005)***	0.020 (0.005)***	-0.001 (0.005)
2. Subsidy amount *(Age< 35) - head	-0.013 (0.007)**	-0.012 (0.006)*	-0.011 (0.006)*	-0.022 (0.007)***	-0.022 (0.007)***	-0.022 (0.007)***	-0.007 (0.006)
3. Subsidy * Female head -head	-0.001 (0.007)	-0.001 (0.007)	0.000 (0.007)	--	--	--	--
4. Mean subsidy amount in two first years of job tenure -spouse	--	--	--	--	--	-0.005 (0.005)	0.032 (0.006)***
Head is a female	-0.012 (0.025)	0.012 (0.024)	-0.001 (0.024)	--	--	--	--
Head aged under 25	0.004 (0.066)	0.015 (0.068)	0.007 (0.068)	0.141 (0.071)**	0.127 (0.072)*	0.139 (0.071)**	-0.027 (0.047)
Head aged 26-35	0.015 (0.023)	0.015 (0.022)	0.014 (0.022)	0.034 (0.023)	0.033 (0.023)	0.031 (0.023)	0.031 (0.019)
Head aged 46-55	-0.038 (0.017)**	-0.036 (0.016)**	-0.036 (0.017)**	-0.028 (0.018)	-0.029 (0.018)	-0.027 (0.018)	0.004 (0.015)
Head aged 56-65	-0.023 (0.020)	-0.022 (0.020)	-0.020 (0.020)	-0.015 (0.022)	-0.013 (0.022)	-0.014 (0.022)	0.005 (0.016)
Household size 1	0.019 (0.025)	0.009 (0.025)	0.009 (0.025)	-0.021 (0 .040)	-0.023 (0.040)	-0.020 (0.040)	-0.014 (0.009)
Household size 3	0.004 (0.016)	0.007 (0.015)	0.009 (0.015)	0.033 (0.016)**	0.033 (0.016)**	0.035 (0.016)**	-0.010 (0.014)
Household size 4	-0.005 (0.017)	-0.001 (0.017)	0.003 (0.017)	0.016 (0.017)	0.019 (0.017)	0.017 (0.017)	-0.025 (0.014)*
Household size 5	-0.009 (0.022)	-0.007 (0.022)	-0.001 (0.022)	0.013 (0.022)	0.018 (0.022)	0.013 (0.022)	-0.016 (0.018)
Household size 6+	-0.026 (0.031)	-0.018 (0.031)	-0.006 (0.031)	0.003 (0.031)	0.013 (0.031)	0.004 (0.031)	-0.009 (0.021)

Table 2: (continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Head started contract after 1997	0.070	0.063	0.079	0.052	0.069	0.051	-0.013
	(0.031)**	(0.030)**	(0.030)***	(0.033)	(0.033)**	(0.033)	(0.024)
Unemployment rate in region	-0.001	-0.001	0.000	-0.001	-0.001	-0.001	0.000
(year entered current firm)	(0.0005)***	(0.0005)**	(0.001)	(0.001)***	(0.001)	(0.001)**	(0.0005)
Head entered labor market	-0.063	-0.054	-0.055	-0.041	-0.041	-0.042	0.039
after 1984	(0.017)***	(0.016)***	(0.016)***	(0.018)**	(0.018)**	(0.018)***	(0.014)*^**
Logarithm of earnings	--	0.094	0.087	0.081	0.075	0.082	0.034
(head and spouse, if present)		(0.011)***	(0.012)***	(0.012)***	(0.012)***	(0.012)***	(0.009)***
Spouse works	-0.002	-0.034	-0.036	-0.031	-0.033	-0.008	0.098
	(0.012)	(0.013)***	(0.013)***	(0.013)**	(0.013)***	(0.020)	(0.018)***
Single	-0.021	-0.005	-0.003	0.009	0.012	0.012	0.010
	(0.033)	(0.032)	(0.032)	(0.037)	(0.038)	(0.037)	(0.013)
Widow/er	0.049	0.063	0.066	0.099	0.101	0.097	0.026
	(0.038)	(0.038)*	(0.038)*	(0.037)***	(0.037)	(0.037)***	(0.018)
Divorced/separated	-0.004	0.005	0.009	0.076	0.078	0.076	0.013
	(0.034)	(0.033)	(0.033)	(0.044)*	(0.045)	(0.044)*	(0.017)
Tenure on the job-3, head	0.070	0.065	0.064	0.063	0.062	0.063	-0.001
	(0.005)***	(0.005)***	(0.005)***	(0.006)***	(0.006)***	(0.006)***	(0.004)
Tenure on the job squared, head	-0.004	-0.003	-0.004	-0.003	-0.003	-0.003	0.000
	(0.0003)***	(0.0003)***	(0.0002)***	(0.0003)***	(0.0003)***	(0.0003)***	(0.0002)
Tenure on the job cubed, head	0.000	0.000	0.000	0.000	0.000	0.000	0.000
	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(0.000)***	(3.2e-06)
Tenure on the job-3, spouse	--	--	--	--	--	-0.002	0.092
						(0.003)	(0.004)***
Tenure on the job squared, spouse	--	--	--	--	--	0.000	-0.005
						(0.0003)	(0.0004)***
Tenure on the job cubed, spouse	--	--	--	--	--	0.000	0.000
						(6.0e-06)	(9.7e-06)***
Constant	0.494	0.555	0.584	0.534	0.561	0.525	0.339
	(0.038)***	(0.038)***	(0.039)***	(0.041)***	(0.042)***	(0.043)***	(0.030)
Region dummies	No	No	Yes	No	Yes	No	No
Average sample size	3,660	3,660	3,660	3,144	3,144	3,144	3,144
Average R-squared	0.27	0.29	0.30	0.29	0.28	0.27	0.74

Source: Five replicates of EFF2002 and 2005, sample of households headed by an employee between 23 and 65 years of age. Schooling and year dummies included but not shown. In each replicate, the standard errors are corrected for heteroscedasticity and arbitrary correlation at the level of both household and cell at which subsidies are imputed. Standard errors (in parentheses) of estimates combined across replicates are computed as in Li et al. (1991). Earnings are the deviation from the sample mean.

Table 3. The impact of subsidies to open-ended contracts on transitions to unemployment

Sample:	Male heads		Female spouses	
Dependent variable has value 1 if individual is observed transiting from employment to unemployment				
Estimation method:	OLS	Probit	OLS	Probit
	(1)	(2)	(3)	(4)
1. Mean subsidy amount in	-0.0010	-0.0005	-0.0010	-0.0004
first year of job tenure	(0.0005)**	(.00016)**	(.00041)**	(.00025)
2. Subsidy amount	-0.0001	0.0001	--	--
*(Age< 35)	(.00015)	(.00013)		
Contract started after 1997	0.0066	0.0044	0.0082	-0.0004
	(.00254)	(.0021)	(.0046)	(.0028)
Age between 40-44	0.0032	0.0028	0.0015	0.0013
	(.0014)	(.00051)	(.0013)	(.0008)
Aged between 45-49	0.0025	0.0024	-0.0064	-0.0034
	(.0010)	(.0010)	(.0036)	0.00093
Age between 50-54	0.0033	0.0030	-0.0029	-0.0012
	(.0015)	(.0007)	(.0016)	(.00178)
Age between 55-59	0.0076	0.0058	-0.0079	-0.0040
	(.0060)	(.0022)	(.00446)	(.0016)
Primary or less	0.0102	0.0057	0.0095	0.0042
	(.0041)	(.00044)	(.0036)	(.0008)
Secondary school	0.0020	0.0016	0.0020	0.0010
	(.00066)	0.0008	(.0017)	(.00189)
College	-0.0042	-0.0052	-0.0107	-0.0091
	(.0027)	0.0012	(.0048)	(.00067)
Widow	-0.0073	-0.0043	-0.0079	
	(.0045)	(.0015)	(.0071)	
Divorced	-0.0004	0.0000	0.0019	
	(.0046)	(.0032)	(.0018)	
Single	-0.0098	-0.0083	0.0013	
	(.0039)	(.0004)	(.0027)	
Year 2000	0.0098	0.0084	0.0159	0.0139
	(.00457)	(.001)	(.0067)	(.0007)
Year 2001	0.0087	0.0080	0.0182	0.0158
	(.0026)	(.0014)	(.0079)	(.0008)
Year 2002	0.0083	0.0077	0.0131	0.0120
	(.0034)	(.0009)	(.0067)	(.0014)
Year 2003	0.0081	0.0080	0.0110	0.0110
	(.0036)	(.0007)	(.0046)	(.0013)
Constant	0.0412	--	0.0608	--
	(.0056)		(.0075)	
Region dummies	Yes	Yes	Yes	Yes
Time at the job dummies	Yes	Yes	Yes	Yes
Sample size:	137,008		87,720	

Sample: Spanish Labor Force Survey (EPA). The first two columns use a sample of heads of households employees and older than 25 years of age. Columns (3) and (4) use a sample of married spouses, employed and older than 25 years of age. In all specifications, the dependent variable takes value 1 if the individual is unemployed in the following quarter, and zero otherwise. The estimates shown in Columns (2) and (4) are marginal impacts on the probability of job loss holding the rest of the variables at their sample means. Standard errors are corrected for arbitrary autocorrelation at the time at the job level.

Table 4: OLS estimates of the impact of subsidies to open-ended contracts on household financial wealth

Dependent variable: Logarithm of wealth held in "liquid" financial assets over household earnings

	Total sample			Sample of male heads		
	(1)	(2)	(3)	(4)	(5)	(6)
1. Mean subsidy amount in	-0.054	-0.056	-0.036	-0.066	-0.045	-0.048
two first years of tenure -head	(0.024)**	(0.024)**	(0.027)	(0.025)***	(0.029)	(0.026)*
2. Subsidy amount	0.060	0.062	0.073	0.075	0.091	0.082
*(Age< 35) -head	(0.032)*	(0.032)**	(0.032)**	(0.034)**	(0.034)***	(0.034)**
3. Subsidy * Female head	0.066	0.066	0.064	--	--	--
	(0.037)*	(0.037)*	(0.037)*			
4. Mean subsidy amount	--	--	--	--	--	-0.075
in two first years of tenure- spouse						(0.024)***
Head is a female	-0.289	-0.289	-0.345	--	--	--
	(0.143)**	(0.143)**	(0.144)**			
Head aged under 25	-0.277	-0.277	-0.367	-0.320	-0.490	-0.334
	(0.233)	(0.233)	(0.234)	(0.280)	(0.277)*	(0.280)
Head aged 26-35	-0.171	-0.171	-0.211	-0.184	-0.256	-0.207
	(0.113)	(0.113)	(0.113)*	(0.116)	(0.117)**	(0.116)*
Head aged 46-55	0.326	0.326	0.350	0.238	0.267	0.242
	(0.094)***	(0.094)***	(0.095)***	(0.105)**	(0.106)***	(0.105)**
Head aged 56-65	0.806	0.806	0.854	0.719	0.768	0.716
	(0.128)***	(0.128)***	(0.127)***	(0.142)***	(0.141)***	(0.144)***
Household size 1	-0.732	-0.743	-0.746	-0.862	-0.842	-0.862
	(0.135)***	(0.135)***	(0.133)***	(0.212)***	(0.213)***	(0.212)***
Household size 3	0.225	0.229	0.214	0.158	0.147	0.159
	(0.092)***	(0.092)***	(0.091)**	(0.103)	(0.103)	(0.102)
Household size 4	0.134	0.139	0.121	0.077	0.053	0.068
	(0.094)	(0.094)***	(0.093)	(0.100)	(0.099)	(0.100)
Household size 5	0.300	0.303	0.294	0.240	0.224	0.225
	(0.129)**	(0.129)**	(0.129)**	(0.134)*	(0.133)*	(0.135)*
Household size 6+	-0.022	-0.013	0.016	-0.060	-0.036	-0.069
	(0.186)	(0.187)	(0.184)	(0.198)	(0.195)	(0.198)
Contract started after 1997	-0.043	-0.043	-0.064	0.073	0.065	0.053
	(0.148)	(0.148)	(0.146)	(0.161)	(0.159)	(0.162)

Table 4: OLS estimates of the impact of subsidies to open-ended contracts on household financial wealth

	(1)	(2)	(3)	(4)	(5)	(6)
Unemployment rate in region	-0.011	-0.011	-0.005	-0.010	-0.003	-0.009
(year entered current firm)	(0.004)***	(0.004)***	(0.004)	(0.004)***	(0.004)	(0.004)***
Head entered labor market	0.135	0.135	0.124	0.014	0.017	0.027
after 1984	(0.093)	(0.093)	(0.092)	(0.103)	(0.102)	(0.103)
Logarithm of earnings	--	0.106	0.110	0.156	0.161	0.153
(head and spouse, if present)		(0.070)	(0.072)	(0.078)	(0.080)**	(0.079)**
Spouse works	-0.184	-0.220	-0.212	-0.226	-0.234	-0.024
	(0.077)**	(0.079)***	(0.078)***	(0.080)***	(0.079)***	(0.112)
Single	1.507	1.526	1.512	1.648	1.638	1.634
	(0.170)***	(0.170)***	(0.169)***	(0.206)***	(0.205)***	(0.207)***
Widow/er	0.862	0.878	0.835	0.671	0.538	0.668
	(0.241)***	(0.241)***	(0.242)***	(9.365)*	(0.373)	(0.366)*
Divorced/separated	0.515	0.525	0.539	0.525	0.538	0.514
	(0.191)***	(0.191)***	(0.190)***	(0.294)*	(0.292)*	(0.295)*
Tenure on the job-3, head	0.050	0.050	0.042	0.064	0.056	0.064
	(0.022)**	(0.022)**	(0.022)*	(0.024)***	(0.024)**	(0.025)***
Tenure on the job squared, head	-0.002	-0.002	-0.002	-0.003	-0.003	-0.003
	(0.001)	(0.001)	(0.001)	(0.001)**	(0.001)*	(0.001)**
Tenure on the job cubed, head	0.000	0.000	0.000	0.000	0.000	0.000
	(0.000)	(0.000)	(0.000)	(0.000)*	(0.000)	(0.000)*
Tenure on the job-3, spouse	--	--	--	--	--	-0.022
						(0.019)
Tenure on the job squared, spouse	--	--	--	--	--	0.001
						(0.002)
Tenure on the job cubed, spouse	--	--	--	--	--	0.000
						(0.000)
Constant	-2.782	-2.712	-2.883	-2.598	-2.773	-2.686
	(0.194)***	(0.200)***	(0.211)***	(0.219)***	(0.234)***	(0.227)***
Region dummies	No	No	Yes	No	Yes	No
Average sample size	3,660	3,660	3,660	3,144	3,144	3,144
Average R-squared	0.16	0.16	0.18	0.17	0.19	0.17

Five replicates of EFF2002 and 2005, sample of households headed by an employee between 23 and 65 years of age. Schooling and year dummies included in the regressions but not shown. See notes to Table 2.

Table 5: The average effect of being covered by high firing costs contract on the log of financial wealth over earnings ratio

Estimation method:	OLS	Two Stage Least Squares				
Sample:	All households	All households		Headed by a male		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A Dependent variable takes value 1 if the household head has an open-ended contract (first stage).						
1. Subsidy amount head was eligible for	--	0.017	0.011	0.019	0.014	0.020
		(0.005)***	(0.005)**	(0.005)***	(0.005)***	(0.005)***
2. Subsidy amount * (Age <=35)	--	-0.012	-0.011	-0.022	-0.022	-0.022
		(0.006)*	(0.006)*	(0.007)***	(0.007)***	(0.007)***
3. Subsidy amount * (Head is female)	--	-0.001	0.000	--	--	--
		(0.007)	(0.007)			
4. Subsidy amount spouse was eligible for	--	--	--	--	--	-0.005
						(0.005)
5. Constant	--	0.555	0.584	0.534	0.561	0.525
		(0.038)***	(0.039)***	(0.041)***	(0.042)***	(0.043)***
Panel B Dependent variable is the logarithm of financial wealth over earnings of head and spouse						
1. Head covered by high firing cost contract	-0.039	-2.341	-2.826	-3.439	-3.918	-2.667
	(0.095)	(1.477)	(2.488)	(1.427)**	(1.934)**	(1.481)*
2. Spouse covered by high firing cost contract	--	--	--	--	--	-2.756
						(1.028)***
3. Constant	-2.717	-1.434	-1.231	-0.760	-0.566	-0.354
	(0.207)***	(0.851)*	(1.497)	(0.811)	(1.144)	(0.923)
Panel C: Fraction of extra gross earnings held as financial wealth (at the median)						
1. Head has an open-ended contract	**0.005**	**0.293**	**0.353**	**0.430**	**0.490**	**0.333**
2. Spouse has an open-ended contract	--	--	--	--	--	**0.345**
Region dummies	No	No	Yes	No	Yes	No
Average sample size:	3,660	3,660	3,660	3,144	3,144	3,144

Estimates in Panel B are Two Stage Least Squares estimates of the contract type held by head and spouse on financial wealth over earnings ratio. The same set of regressors in Tables 2 and 4 is used in all specifications, but not shown to save space. Standard errors (in parentheses) of estimates combined across replicates are computed as in Li et al. (1991), and corrected in each replicate for arbitrary autocorrelation at the household, age-region-gender-year of entry at the firm level.

Table 6: The effect of an open-ended contract on the ratio of financial wealth over income

Estimation method: Instrumental variable quantile regression (Chernozhukov and Hansen, 2004, 2008)

	25th centile (1)	50th centile (2)	75th centile (3)
Panel A: The response of financial wealth over earnings to the contract type of the head of household			
1. Head covered by an open-ended contract	-1.7	-1.0	-2.9
90% confidence interval	[-5.4, -1.0]	[-3.7, 0.4]	[-10, 2.3]
2. Constant	-2.221	-1.760	0.806
3. Fraction of extra gross yearly earnings held as wealth when head covered by fixed-term	**0.089**	**0.109**	**-**
Panel B: The response of financial wealth over earnings to the contract type of the working spouse			
1. Working spouse covered by an open-ended contract	-3.7	-2.3	-2.2
90% confidence interval	[-8.3, -0.7]	[-4,2, -0.4]	[-3.2,-0.9]
2. Constant	-2.467	-2.040	-0.759
3. Fraction of extra gross yearly earnings held as wealth when working spouse covered by fixed-term	**0.083**	**0.117**	**0.417**

Estimates shown are Instrumental Variable Quantile Regression estimates of the contract type held by head on the ratio of financial wealth over earnings estimated separately for the head and the spouse (controlling in each case for the subsidy that the other member qualifies for). The same set of regressors used in Tables 2 and 4 is included in all specifications, but not shown to save space. The estimates use only the first of the five replicates.

Table 7: The average effect of being covered by high severance payments on various measures of household wealth

	Falsification exercise	Alternative dependent variables	
	Subsidy available during the 4th year subsidy	Net wealth minus home value and debts associated	Net wealth as dependent variable
	(1)	(2)	(3)
1. Head covered by high dismissal cost	0.854	-3.557	1.770
	(1.251)	(1.813)**	(1.443)
Head aged under 25	-0.252	-0.359	-1.494
	(0.299)	(0.473)	(0.405)***
Head aged 26-35	-0.064	-0.221	-0.333
	(0.103)	(0.133)*	(0.124)***
Head aged 46-55	0.253	0.356	0.297
	(0.111)**	(0.138)***	(0.092)***
Head aged 56-65	0.718	1.105	0.683
	(0.141)***	(0.170)***	(0.107)***
Contract started after 1997	-0.082	0.180	-0.180
	(0.198)	(0.253)	(0.197)
Unemployment rate in region	-0.009	-0.012	-0.001
(year entered current firm)	(0.004)**	(0.006)**	(0.004)
Head entered labor market	0.043	-0.287	-0.149
after 1984	(0.118)	(0.153)*	(0.116)
Logarithm of earnings	0.079	0.561	-0.207
	(0.132)	(0.179)***	(0.145)
2. Fraction of gross earnings as financial wealth (at the median)	**-0.107**	**0.445**	**--**
Average sample size:	3,144	3,136	3,038

Two-stage-least squares estimates, "Subsidy to conversion" and its interaction with age of the head below 35 as instruments. Sample of male heads.The standard errors (in parentheses) of combined estimates are corrected in each replicate for arbitrary autocorrelation at the household, age-region-gender-year of entry at the firm level and then computed as in Li et al. (1991).

Table 8: The average effect of being covered by a high severance payment contract on access to credit markets

Estimation method: multinomial logit (base outcome: asked not for a loan in the last 2 years)

		Kinds of "credit constrained" households			All constrained households	
	Asked for a loan and fully accepted (1)	Did not ask, fears rejection (2)	Asked and was rejected (3)	Given less than asked (4)	Overall (2)+(3)+(4) (5)	Among potential borrowers (6)=(5)/[(1)+(5)]
Sample means:	0.281	0.010	0.008	0.014	0.032	0.102
Model 1: Open-ended contract as a regressor						
1. Fixed-term contract, both	0.272	0.000	0.002	0.022	0.025	0.084
2. Open-ended contract, head	0.281	0.000(***)	0.002	0.009 (*)	0.011	0.038
3. Open-ended contract, spouse	0.334 (*)	0.000 (*)	0.001	0.032	0.032	0.089
Model 2: Subsidy amount as a regressor						
1. Zero subsidies, both	0.295	0.000	0.002	0.016	0.018	0.057
2. 1,000-euro subsidies, head	0.287	0.000	0.002 (*)	0.013	0.016	0.052
3. 1,000-euro subsidies, spouse	0.295	0.000	0.002	0.018	0.019	0.061

Entries are fitted probabilities of a multinomial logit that has "Not asked for a loan" as the base outcome. Each cell contains the probability of the outcome in each column, predicted for a married household that has no children and both members have basic schooling, no children, the head is aged between 36 and 45 and both members of the couple have three years of tenure at their job. (***), (**) and (*) mean that the coefficient of the latent variable row is significant at the 1, 5 or 10% of confidence level. Model 1 uses "Open-ended contract" as a regressor. Model 2 uses our instrument (subsidies). Rest of covariates: age, marital status, logarithm of income, schooling of head and spouse, family size, and third order polynomial in tenure minus 3. Probabilities are predicted using the combined estimates across the five replicates of the EFF.

Table 9: The impact of the risk of losing the job on consumption growth over a three year-period

Dependent variable:	Log (Food t+3) -Log(Food t)		Log(Non durables t+3) -Log(Non durables t)		Log(Total Cons. t+3) -Log(Total Cons. t)	
Estimation method: OLS						
	(1)	(2)	(3)	(4)	(5)	(6)
1. Head covered by open-ended contract	-0.129	-0.128	-0.189	-0.183	-0.125	-0.120
	(0.059)**	(0.059)**	(0.074)***	(0.073)***	(0.064)**	(0.064)*
2. Spouse covered by open-ended	--	-0.010	--	-0.056	--	-0.050
contract		(0.066)		(0.077)		(0.064)
Spouse works	-0.035	-0.029	-0.050	-0.014	-0.042	-0.009
	(0.046)	(0.066)	(0.051)	(0.068)	(0.040)	(0.054)
Head between 23 and 25	0.027	0.024	-0.276	-0.293	-0.271	-0.286
	(0.220)	(0.220)	(0.245)	(0.249)	(0.234)	(0.237)
Head between 26 and 35	0.085	0.084	-0.016	-0.019	-0.001	-0.004
	(0.075)	(0.075)	(0.078)	(0.078)	(0.065)	(0.065)
Head between 46 and 55	-0.036	-0.036	-0.009	-0.011	0.035	0.034
	(0.050)	(0.050)	(0.065)	(0.065)	(0.054)	(0.054)
Head between 56 and 65	0.048	0.048	0.024	0.026	0.020	0.021
	(0.073)	(0.074)	(0.076)	(0.076)	(0.062)	(0.062)
Change in household size	0.147	0.147	0.173	0.173	0.139	0.139
	(0.033)***	(0.033)***	(0.036)***	(0.036)***	(0.030)***	(0.031)***
Change in number of children 0-3	-0.130	-0.131	-0.126	-0.129	-0.116	-0.119
	(0.066)**	(0.066)**	(0.077)*	(0.078)*	(0.064)*	(0.064)*
Change in number of children 4-7	-0.064	-0.064	-0.172	-0.171	-0.135	-0.135
	(0.066)	(0.066)	(0.080)**	(0.080)**	(0.067)**	(0.067)**
Change in number of children 8-11	0.012	0.012	-0.085	-0.085	-0.051	-0.051
	(0.060)	(0.060)	(0.074)	(0.074)	(0.059)	(0.059)
Change in number of children 12-15	0.002	0.002	-0.124	-0.123	-0.080	-0.079
	(0.053)	(0.053)	(0.066)*	(0.066)*	(0.054)	(0.054)

Table 9 (continued)

Dependent variable:	Log (Food t+3) -Log(Food)		Log(Non durables t+3) -Log(Non durables t)		Log(Total Cons. t+3) -Log(Total Cons. t)	
Estimation method: OLS						
	(1)	(2)	(3)	(4)	(5)	(6)
Change in number of children 16-18	-0.067	-0.067	-0.089	-0.087	-0.066	-0.064
	(0.050)	(0.050)	(0.058)	(0.058)	(0.047)	(0.046)
Single	-0.061	-0.060	-0.101	-0.099	-0.163	-0.162
	(0.113)	(0.113)	(0.132)	(0.132)	(0.122)	(0.122)
Widow	0.275	0.275	0.329	0.327	0.280	0.279
	(0.213)	(0.213)	(0.236)	(0.235)	(0.249)	(0.249)
Divorced	0.239	0.239	0.220	0.220	0.129	0.129
	(0.078)***	(0.079)***	(0.161)	(0.162)	(0.131)	(0.132)
Secondary education, head	0.011	0.012	0.044	0.046	0.030	0.031
	(0.060)	(0.060)	(0.069)	(0.069)	(0.057)	0(.057)
Upper secondary, head	0.062	0.063	0.074	0.079	0.029	0.034
	(0.064)	(0.065)	(0.074)	(0.074)	(0.060)	(0.060)
College, head	0.048	0.050	0.017	0.023	-0.010	-0.005
	(0.060)	(0.061)	(0.069)	(0.069)	(0.057)	(0.057)
Constant	0.092	0.090	0.198	0.190	0.167	0.160
	(0.074)	(0.074)	(0.086)**	(0.085)**	(0.074)**	(0.073)**

Notes: Average sample size: 625. Standard errors are in parentheses. The dependent variable is consumption growth over a three year period. The estimates are combined across the five replicates of the EFF and standard errors are computed as in Li et al. (1991).

Table 10: Simulated steady state distribution of wealth by probability of job loss

	Prob. Job loss 0.02 (1)	Prob. Job loss 0.10 (2)	Absolute change (3)	Relative change (4)
Mean W /Y of:				
All households	0.404	0.643	0.239	0.592
1.. 20th-30th W /Y Percentile	0.175	0.353	0.178	1.019
2. 40th-50th W /Y Percentile	0.315	0.534	0.218	0.693
3. 60th-70th W /Y Percentile	0.467	0.725	0.258	0.553
4. 80th-90th W /Y Percentile	0.703	1.026	0.323	0.459

Table A.1 Exposure to the risk of an unemployment spell of at least one month, by contract form

Probability of experiencing an unemployment spell in 2004 by the type of contract in 2002 (EFF 2002-2005)

	Open-ended contract	Fixed-term contract
Head:		
Total	0.055	0.187
Primary school	0.117	0.289
Secondary school	0.050	0.138
Upper secondary school	0.046	0.130
College	0.027	0.079
Spouse:		
Total	0.105	0.511
Primary school	0.170	0.589
Secondary school	0.148	0.550
Upper secondary school	0.112	0.469
College	0.057	0.300

The probabilities are predicted from weighted logit estimates obtained separately for the head and the spouse and using the type of contract and the level of education. The probabilities come from the estimates combined across the five replicates of the EFF.

Table A.2: 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<=age<=44	5160 if female or age>45 4500 if 41<=age<=44 3600 if male age<30	5160 if female or age>45 4500 if 41<=age<=44 3600 if male age<46
3. Asturias	4500 for all	4500 for all	None	4,200 if female or age>=46 3600 otherwise
4. Baleares	None	None	None	1652.78 if female
5. Canarias	None	3,600 if female or age<=25	3,600 if female or age<=25 3,000 otherwise	None
6. Cantabria	None	3900 if female or age<=30 3300 if male 30<age<=40 3,600 if male age>=41	None	5408 if age>=46 3606 if age<=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>=30 1800 if male age >=41	4507.59 if age <30 3305.57 if female age>=31 1803 if male age>=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. Valencia	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<=30 11180 if age>30 and age<=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	None	None
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	1800 for all 2400 if age<30	1800 for all 2100 if age<30	1800 for all 2100 if age<30
16. Basque country	None	7512 for all	7512 for all	7512 for all
17. Rioja	None	4500 for all	4491 for all	6011 for all

Table A.2: 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 of 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 for all, 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. Valencia	4808.1 for all	1800 for females	2000 for females 1500 for the rest	4000 if female 2000 if age<30, 1500 ow.
11. Extremadura	5410.086 if female >45 4455.365 if male > 45 2386.802 otherwise	6010 for all	None	None
12. Galicia	None	None	4200 euro if female or age<30 2400 if age>45 3000 if age>24 & age<45	4500 if female or age>50 3000 if 25<age<50
13. Madrid	10800 for all 12000 if above 45 (males) 12000 if above 40 (females)	12000 for all	9000 for all 12000 if above 45	3000 euro, all
14. Murcia	4800 for all	4800 for all	None	None
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

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