The Response of Household Wealth to the Risk of Losing the Job: Evidence from Differences in Firing Costs[1]

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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. While the fraction of workers covered by a high severance payment contract that transit into unemployment is below 2% per quarter, the corresponding estimate among workers covered by low firing cost contracts exceeds 10%. 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 3.6 months of income.

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

JEL codes: D12, D31, D91, J41.

1 Introduction

Do workers who are more exposed to the risk of losing the job accumulate higher wealth balances in anticipation of future income losses? From a policy perspective, understanding the evolution of consumption requires assessing if households that are exposed to lay-offs have accumulated wealth to sustain consumption during an unemployment spell and how large those buffers are. From a theoretical perspective, whether or not households perceive unemployment risk and actually react to it by accumulating wealth is crucial to understand the consumption and 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.

The literature has used different methods to establish if households facing higher chances of losing their job have lower consumption levels and/or accumulate higher levels of wealth, highlighting three empirical problems. First, workers who are more averse to risk are more likely to self-select into occupations with lower risks of losing the job (Lusardi, 2007 and Fuchs-Schündeln and Schündeln, 2005). Second, individuals who are more likely to lose the job are also more likely to be observed holding small wealth holdings if savings have been used 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 the 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 precautionary motive. Engen and Gruber (2001) document that Unemployment Insurance in the US crowds out private wealth accumulation, but the estimated wealth

[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.

responses are 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 are consistent with hypothesis of the precautionary saving model. Nevertheless, 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 has three advantages that permit examining the relationship between the probability of losing the job and household decisions like consumption and wealth. First, we focus on differences in the exposure to the risk of losing the job caused by differences in dismissal costs. During the eighties, Italy, Germany, Spain, Sweden, Portugal and France (among other countries) introduced low firing cost contracts. Those countries typically had rigid labor markets with high dismissal costs, so fixed-term contracts allowed firms to hire workers paying a small firing cost in the event they needed to downsize. A wide literature has documented that in the following years, workers covered by fixed-term contracts faced high probabilities of transiting into unemployment (Blanchard and Landier, 2002 or Dolado, Garcia-Serrano and Jimeno, 2002).[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 differently exposed to different dismissal costs.

Secondly, due to the fact that labor market policy is to some extent 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-Perez and Rebollo-Sanz, 2009a). By 2004, 15 out of 17 regions had implemented such programs

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

[3]For example, a possible explanation of the low wealth responses 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. In contrast, variation in contract type is a strong predictor of future transitions 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). We focus on objective measures here.

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 into more secure jobs and to previous unemployment experiences (because we are comparing workers hired in the same year but whose firms faced different incentives to use open-ended contracts). An additional advantage of our approach is that, as subsidies to hire using high firing cost contracts differed by gender, we are able to estimate the wealth responses of exposure to job loss of 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 then reduced-form TSLS estimates suggesting that households whose head obtained a high dismissal cost contract as a consequence of the regional subsidies have 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 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. Overall, using simulations from a simple buffer stock saving model, we argue that a precautionary saving motive is more consistent with our results than a model with perfect certainty.

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

2 Dismissal costs in Spain

The main form of labor contract in Spain before 1984 was "open-ended" contracts, that featured high dismissal costs: between 20 and 45 days 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 only in two 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 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. Otherwise, there was no difference between those contracts regarding contributions to the old-age or disability social security systems, access to 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, Garcia-Serrano and Jimeno (2002) document that the use of such contracts has been widespread across all industries, including the Public Administration but those contracts are relatively more common among females, 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

[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.

Table A.1, we present computations from the EPA that suggest that holding constant age, industry, gender and education, an individual whose job position is regulated by a fixed-term contract was 7.7% more likely to be observed in unemployment in the next quarter than a worker covered by an open-ended contract.[5]

Our study uses the fact that identifiable groups of the population are exposed to very different degrees of unemployment risk depending on the contract form to analyze whether households where the head or spouse have a labor contract with a higher dismissal cost hold on average lower wealth balances. 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. Therefore, to estimate properly the causal impact of the risk of losing the job on precautionary wealth, we need to make use of instrumental variables instead of looking at Ordinary Least Squares (OLS) estimates.

Hence, our identification strategy relies 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 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 and 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

[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.

two to three years (which was the case in Valencia, Castilla la Mancha or Extremadura). Progressively, between 1997 and 2004, all regions but Catalonia and Navarra had implemented for at least a year some form subsidy to contracting with 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 (2009) estimate that the lump sum received by the firm was about 20% of the yearly labor cost of the mean worker.[7]

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. The new contract was also available for firms that converted a worker's contract from fixed-term into an open-ended one. In the empirical section, we discuss our strategy to disentangle the impact of regional variation in incentives to hire from the national reform.

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. Assume that an individual lives for two periods, does not discount the future, and that there is a zero interest rate. 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

[7] Up to this moment, we only have fragmentary evidence regarding whether firms actually used those subsidized contracts. Information from 2004 suggests that the incidence of such subsidies was the highest in Baleares (accounting for 63% of open-ended contracts created in 2002), Murcia (52 percent of open-ended contracts) or Galicia (14% of contracts). The incidence was very low in Madrid (1 percent of contracts), Basque Country (2 percent) or Aragon (2 percent of all permanent contracts).

unemployment benefits and severance payments $b + F$ 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:

$$\max_{c_1, c_2} \quad \log c_1 + E_1 \log(c_2)$$

Where the expectation is taken over the binary random variable Y, with mean, $p(b + F) + (1 - p)y$, and variance, $Var_1(Y) = (1 - p)p[y - b - F]^2$. Following Blundell and Stoker (1999), we define the present value of expected lifetime resources in period 1 (W) as the sum of the initial wealth accumulated in period 1 (W_1) and the expected stream of income in period 2, as follows:

$$W = W_1 + p(b + F) + (1 - p)y$$

We are implicitly assuming that the individual can borrow against the expected value of future income (including expected payments associated to lay-offs). While perhaps not a realistic assumption, it allows us to obtain a closed-form solution. Linearizing around the perfect-certainty solution of consumption (that is linear in first-period wealth) we obtain the following consumption levels in the presence of risk:

$$c_1 = \frac{W}{2 + \frac{Var_1(Y)}{W^2}} \tag{1}$$

A worker covered by a high firing cost contract exhibits a low value of p and, upon job loss, a high value of F. If we ignore the possibility that a fixed-term contract can be upgraded into an open-ended one, a worker covered by a fixed-term contract has a higher value of p and lower value F than a worker covered by an open ended contract. According to equation (1), there are two channels that lead the first worker to accumulate a lower level of wealth. The first is that, as long as $b + F < y$ the expected future income of workers covered by high firing cost contracts is higher: they expect a lower probability of exiting employment and upon job loss workers with a high firing cost contract have a higher compensation (i.e., their level of W is higher). The second channel that leads the worker covered by a high firing cost contract to save less only operates if the utility function exhibits prudence: 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: $\frac{1}{2+\frac{Var_1(Y)}{W^2}}$ is higher the lower $Var_1(Y)$.[8]

A second implication of the precautionary saving 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 second-period shock ζ_2 as the difference between the realization of second-period income and the expected value of the income stream

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

Blundell and Stoker (1999) also show that under our assumptions, consumption growth can be written as follows:

$$\log(c_2) - \log(c_1) = \frac{Var_1(Y)}{W_1^2} + \frac{1}{c_1}\zeta_2 \tag{2a}$$

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. Taking expectations in (2a) over the distribution of Y one obtains the following expression:

$$E_1[\log(c_2) - \log(c_1)] = \frac{Var_1(Y)}{W_1^2} \tag{2}$$

Individuals who are covered by a low dismissal cost contract are exposed to a higher variability in earnings (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

[8] Some researchers introduce adjustments for expected income, like controlling for p to control for mean effects and examining the response of wealth to $Var_1(Y)$ (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.

zero. Hence, a positive link between consumption growth and the dismissal cost specified in the worker's contract gives indications that a precautionary saving motive is indeed present in the data.

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.

3 Datasets

The main dataset 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. 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. In this preliminary version of the paper, all the calculations reported make use of only the first 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. Finally, there is a discussion regarding whether or not households are able to use owner-occupied housing equity to finance a period of unemployment; while Carroll, Dynan and Krane (2003) argue that housing wealth is "liquid" wealth, due to the possibility of extracting housing equity, Engen and Gruber (2001) argue that housing wealth cannot act as a buffer against unemployment risk (and find evidence in that direction). Recent work on precautionary wealth even considers housing as *consumption*, rather than wealth (Shore and Sinai, 2009). In our case, the considerations that lead researchers to consider owner-occupied housing as a source of precautionary wealth are unlikely to apply: the possibility of extracting housing equity from owner-occupied housing was rather uncommon in Spain during our sample period. 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 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 3,776 household-years. As we take logarithms of wealth in most of the analysis, we lose 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 up to three jobs).

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

[9]Some of the sample restrictions, like the exclusion of unemployed heads or households with zero wealth, merit further investigation. Nevertheless, 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.

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. According to that source, 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. Hence, if subsidies to contract conversion had an impact, they would be most likely to affect on the first two years of the life of the contract.[11] We convert the monetary magnitude of the subsidy into constant euro of 2005 using regional deflators of household gross disposable income.[12]

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. The summary statistics in Table 2 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

[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.

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 69 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.

3.2 A secondary source: The Spanish Labor Force Survey

We also make limited use of a secondary source of information: the *Spanish Labor Force Survey* (EPA). As described above, the EPA is a rotating panel that interviews around 20,000 households, per quarter, keeping track of each member for up to six quarters. Due to its large sample size, demographic and labor market information in the EPA permits a reliable computation of the relevance of flows from employment to unemployment at the beginning of the contract for a worker qualified for the subsidy. To maintain some comparability across samples, we span the waves of the EPA between 2000 and 2004 and we focus on a sample of family heads aged 25 or over who are employees and were hired using any form of contract between 1994 and 2004.[13]

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

[13]The EPA reports the age of the worker in five-year brackets. This poses some problems when inferring the age of the worker at the time when the contract started. We assigned the subsidy using the mean age of the interval, what is admittedly a noisy imputation.

of the working spouse, if one exists. This allows us to examine wealth responses to different exposures to risk depending on which household member is exposed to the risk.

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.

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

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 in 2002 or 2005 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.[14] 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).[15] Finally, X_i is a vector of covariates that includes three dummies with the educational level attained by the worker (primary education or less (omitted category), first

[14] 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 EPA and found that controlling for tenure fixed effects or for a third-order polynomial yielded very similar results.

[15] García-Pérez and Rebollo-Sanz (2009) 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.

stage of secondary education, upper secondary school and college), three industry dummies (agriculture, industry, construction and service sector, the last is the omitted category), 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 of the household head is employed.[16] 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 head of household 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

[16] 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.

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 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)].

4.1.1 Impact on transitions from employment to unemployment

We also examine flows from employment to unemployment to check if workers who were exposed to obtain a subsidized high firing cost contract had more stable jobs than workers covered by fixed-term contracts. The reason is that 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. 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 EPA to have a large sample. The coefficient of interest is α_1, that 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.

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

Second, 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:

$$\log(\frac{W}{Y}) = \delta_0 + \delta_1 Subsidy^h_{R,g,t_0} + \delta_2 Subsidy^h_{R,g,t_0} 1(Age^h \geq 35) + \delta_3 Subsidy^h_{R,g,t_0} \cdot Female^h + \sum_{g=1}^{g=4} \delta_{4,g} Age^h_g + \delta_5 Female^h + \delta_6 Hired\_post97^h_{t_0} + g_1(Tenure^h - 3) + X'\delta_7 + u \quad \text{(C.2)}$$

The dependent variable is the logarithm of the ratio of household wealth to earnings of head and spouse, if one is present. 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. To be able to compare our results to those in previous literature that has measured precautionary wealth holdings as the extra fraction of yearly household earnings held as wealth by households whose head is exposed to the risk of losing the job, we normalize wealth holdings by gross household earnings.[17]

Given the strong skewness of the wealth distribution, we decided to work with logarithm of wealth selecting out of the sample a relatively small number of households that have zero "liquid" wealth: 117 out of 3,858 households (3.2 percent of the number of original households). Finally, 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 at the beginning of her contract.

[17] 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.1. The results hardly changed. Rather than constructing measures of permanent income (limited 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.

4.3 Assessing the magnitude: 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 toward zero 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. Namely, we estimate the following system of equations:

$$\begin{aligned} \log(\frac{W}{Y}) &= \gamma_0 + \gamma_1 Open\_ended_{head} + \gamma_2 Open\_ended_{spouse} \\ &+ g_2(Tenure_{head} - 3) + g_2(Tenure_{spouse} - 3) + X'\gamma_2 + v \end{aligned} \tag{W.1}$$

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 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: 0.117.[18]

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

[18]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 log of wealth. The results were similar to those reported here.

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 a variable ($Subsidy^{head}_{R,g,t_0}$) -see Moulton (1986). Thus, the standard errors are corrected for arbitrary autocorrelation across panel households and at the age-region-year of tenure-gender level. In Table 2, row 1, column 1, the estimate of the variable $Subsidy^{head}_{R,g,t_0}$ is 0.0181 (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 16.18. The estimate of the interaction of $Subsidy^{head}_{R,g,t_0}$ and a dummy for age below 35 years is -0.013 (standard error: 0.007). The magnitude of the estimate implies that subsidies to conversion had a smaller impact on the probability of observing relatively younger workers as heads covered by an open-ended contract in 2002 or 2005: 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 0.54%. While not significantly different from zero, the estimate is also lower among female heads: 1,000 extra euros increase the stock of female heads with a high dismissal cost contract by 1.6% (=0.0181-0.002). In principle, females are the most benefitted from the subsidy, but the group of female heads of household is arguably a very selected one according to our definition of household head in the EFF.[19] Overall, the instrument we use seems to work best for male heads of household above 35 years of age.

Specification 2 in Table 3 adds household earnings, 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 0.011 (standard error: 0.005), shown in row 1, Column 3 of Table 3. 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 4.84, resulting in a weaker instrument than in the previous specification.

[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.

Specifications (4) and (5) turn to the group for whom the instrument is strongest: the sample of male heads. 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: 0.006) and 1.9 percentage points (standard error: 0.005) in a specification that excludes and includes region dummies, respectively. Column (6) investigates the impact of the incentive to hire using high firing cost contracts on the contract form of both head and spouse. 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 when she started the contract, using a sample of male heads. The estimate is 0.02, very similar to the estimate in Column (4) of Table 3.

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 point. The impact is larger than among male heads. Reassuringly, the subsidy that the head (spouse) was eligible for at the time of starting the current contract explains little of the variation in current contract form of the spouse (head), suggesting that our instrument is not picking up other 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

[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 employment 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.

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 jobs. The probit estimates for males is smaller: -0.00045 (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.

Panel A in 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 convert a fixed term contract into an open ended ones diminishes household financial wealth by 5.8 percentage points (standard error: 0.23). The estimate is consistent with the notion of precautionary wealth holdings: groups of the population that experiment 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 0.066 (standard error: 0.03), positive but not very precise. Adding this estimate to that of the variable $Subsidy^{head}_{R,g,t_0}$ yields an estimate of 0.008 (=0.058-0.066), suggesting that the incentive to convert fixed-term contracts for workers currently below the age of 35 reduces wealth by less than 0.9%, 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. A possible reason for this small estimate is that the instrument is not very powerful predictor of the stock of workers covered by a high dismissal cost contract below the age of 35.

Column 3 of Table 4 introduces indicators of the region of residence.[22]

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

The estimate of the variable $Subsidy^{head}_{R,g,t_0}$ is -0.038 (standard error: 0.025), 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 3 of the instrumental variable $Subsidy^{head}_{R,g,t_0}$ is -0.071 (standard error: 0.024), negative and significantly different from zero at the 1 percent confidence level. Thus, when we use a sample of male heads in their mature age, an increase of 1,000 euro of the incentive to convert a fixed-term contract into a permanent one results in a drop of the logarithm of the household financial wealth to income ratio of 7.1 percentage points. The result is smaller but similar when we add regional indicators: -0.049 (standard error 0.027), 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-earnigns ratio by 5 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: -0.084 (standard error: 0.023).

Overall, our interpretation of the results in Table 4 is that households headed by a male employee over the age of 35 react to variables measuring an exogenous increase of the probability of being 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. We also find that households react to exogenous increases in the job stability of spouses by reducing wealth-earnings ratios.

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 by 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 3 and

4. Here we only display our parameters of interest.

The OLS estimate of the impact of "open-ended contract" on the log of household wealth to income ratio is -0.004 (standard error: 0.091). Multiplying that estimate by 0.117 (the median wealth-income ratio), the estimate suggests that workers covered by fixed-term contracts have no additional gross financial wealth holdings than comparable workers covered by open-ended contracts (Table 5, Panel B, row 3, 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 -2.504 (standard error: 1.493), 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 about 29.3% of their gross earnings in financial wealth. Our explanation of the stark difference between the OLS estimate of -0.004 and the TSLS estimate of -2.50 is due to the population affected by our instrument. By exploiting only the variation in current type of contract that is related to the amount of subsidies to open-ended contracts, the wealth responses in Columns (2)-(5) of Table 5 only use the fraction of workers currently covered by open-ended contracts who started their current employment spell as fixed-term contract employees. Such heads of households currently covered by "open-ended contracts" are likely to have had labor histories similar to those of workers who are currently covered by fixed-term contracts. Therefore, the different wealth responses of workers according to their kind of contract in specifications (2)-(5) are mainly due to their different degree of protection against unemployment risk determined exogenously by the presence of these regional subsidies.

The estimate of the impact of the type of contract on average wealth-earnings ratios is comparable to that in Carroll, Dynan and Krane (2003). They 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. Nevertheless, one must take into account that the differences in the chances of transiting into unemployment we exploit are much higher than the differences in employment flows between different workers in the US in the study of Carroll et al. (2003).

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 3.71 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 43.4 percent, or 5.2 months' income (a substantial amount). The estimate is not very sensitive to the inclusion of region dummies (presented in column 5, Table 5, Panel B row 3).

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 -2.831 (standard error: 1.587), around 4 months of yearly earnings. The response of household wealth to the risk that the spouse loses the job is similar in magnitude.

The evidence from Table 5 suggests that households react to the risk of job loss of the main and secondary earner by accumulating financial wealth. The evidence is strongest among households headed by male employees, and the average size of the excess of wealth kept with respect to workers covered by high dismissal cost contracts is between 29% and 44% of gross household earnings.

5.3 Robustness checks

Alternative measures of wealth Table 6 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 unappropriate instrument should have little impact on accumulated household wealth. Otherwise, we could be suspicious that the instrument is picking particular regional trends in wealth and employment quality. The estimate shown in Table 6, Column 1, Row 1 is close to zero: 0.771 (standard error: 1.210). Hence, we infer that the results in Table 5 are unlikely to be driven by spurious regional trends.

We then turn to analyzing alternative wealth measures. First, we use an expanded wealth measure that includes gross financial wealth and net housing wealth that excludes home equity. The rationale for that measure is that other real estate is less costly to liquidate in case of an unemployment spell than own housing wealth. The estimate of the coefficient "Head covered by high dismissal costs contract" in Table 6, Row 1, Column 2 is -3.651

(standard error: 1.751). The estimate is slightly smaller than that using financial wealth, and is consistent with a "buffer stock saving" of about 42% of gross annual earnings (after multiplying the coefficient by the median wealth of households covered by a fixed-term contract, 0.12).

In the third column of Table 6 we use the broadest wealth concept: net wealth (excluding non-cashable pension funds and life insurance products). The point estimate is 1.324 (standard error: 1.335), 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. One can argue that workers covered by a fixed-term contract are less likely to obtain credit from banks and must then accumulate wealth through lower consumption to purchase a house, an explanation that does not rely necessarily on a precautionary saving motive. A second explanation is related to demand for credit in the presence of employment risk. 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. That second explanation is consistent with a precautionary wealth motive. We disentangle between both hypothesis below.[23]

Do the estimates reflect the prevalence of credit constraints? We test if the results in Tables 5 and 6 are driven by credit constraints binding for individuals whose job position is covered by a fixed-term contract. For this purpose, we examine how loan rejections vary with the statutory subsidy to open-ended contracts. In the EFF, we identify three forms of credit constraints, following Jappelli (1990). The first is whether the household did not ask for a loan during the last two years because the loan is thought to be rejected. The second is whether the interviewed household asked for a loan, but was rejected. The third form of credit constraint is whether the loan was not rejected, but it was given a lower amount than the one asked. The three latter outcomes denote a credit constrained household. The estimation sample is formed by male household heads and the set of regressors is the

[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.

same shown in Tables 3 and 4.

We estimate a multinomial logit model with five different outcomes. The first is not having asked for a loan for the first two years, the second is having asked for a loan and the application accepted. The third outcome is not asking for a loan because of the fear of having it rejected. The fourth is having the loan rejected and the fifth is having received a lower amount than that asked. If the estimates in Table 5 and 6 are picking up the responses of credit contrained households, we would expect that the variable $Subsidy^{head}_{R,g,t_0}$ causes a drop in the relative chances of being credit constrained.

The first row of Table 7 shows the summary statistics: 28.2 percent of the households have requested (and obtained) a loan in the last two years during the sample period, 1.1 percent of the households did not ask because they feel they would be rejected, 1 percent were actually rejected, and 1.5 percent got less than what they asked. According to this measure, 3.6 percent of all households were credit constrained, and 11.3 percent of potential loan applicants [=3.6/(28.2+3.6)*100].

Panel A includes the results from a model that uses the actual form of contract as the regressor. The results of that specification suggest that households headed by a male employee covered by a fixed-term contract are more likely to be credit constrained: 18.9 percent among all applicants holding fixed-term contracts while only 8.0 percent among households headed by an open-ended contract. The stronger presence of liquidity constraints among fixed-term workers may not only be due to their higher risk of losing the job, but also due to their past labor history (a higher propensity of having experienced past unemployment spells and unobserved characteristics that make them less able to accumulate wealth and earnings). Thus, our measure of the risk of losing the job, the form of the job contract, is an endogenous variable. For that reason, Panel B estimates Model 2 replacing the indicator of having a permanent contract by other exogenous determinant of the risk of job loss: the regional subsidies to the conversion of fixed-term contracts into open-ended ones.

The pattern of coefficients in Model 2 shows 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 12.9 percent (Table 7, Model 2, Column 6, row 1). Among households who are eligible for a 1,000 euro subsidy, the fraction of credit constrained households is slightly *higher.* 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. Fixed-term workers seem to be more affected by credit constraints due to other factors like their past labor history and unobserved characteristics rather than their risk of losing the job. 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 likely rejection of the loans applied for.

The response to the risk of losing the job at various points of the wealth distribution A given average wealth response to the risk of losing the job may be due to the fact that *all* households exposed to the risk of losing the job keep uniformly higher wealth balances. Alternatively, it may be the case that exposure to job loss generally leads to a negligible wealth response for most households, with a small fraction of households accumulating very high wealth balances. The distinction matters. First, if only a minority of households accumulate wealth, there could be substantial consumption losses following an increase in the unemployment rate, as most households would have accumulated little balances. From a distributional point of view, if wealth responses are concentrated at the top of the wealth distribution, public policy targeted to low wealth households would be optimal.

We distinguish between both situations by estimating Instrumental Variable-Quantile Regression Models of the response of wealth to the risk of losing the job (see Chernozhukov and Hansen, 2004 and 2008). The results are shown in Table 8. While tentative (the results are imprecise at the top), the results suggest a non-uniform impact over the wealth distribution, the response of liquid wealth is most precise and largest in relative terms at the 25th centile than at the median. Finally, the results are very imprecise in the 75th quantile, where we obtain sizable estimates but cannot reject a zero response.

5.4 Additional evidence from consumption growth

A second prediction from Section 2 is that average consumption growth should be higher among workers who are relatively more exposed to the risk of losing the job than among workers who are relatively more protected, because uncertainty about the risk of losing the job induces individuals to postpone consumption. A key element of that prediction is that the expectation of consumption growth should be taken among all possible states (including unemployment). Hence, to test 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. Next, we use the panel component of the EFF to track these individuals into the 2005 wave, *regardless* of whether they are employed or not. The exact model we run is

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

The dependent variable denotes consumption growth of household i between 2005 and 2002. The information in the EFF permits us to construct three consumption measures. The first is a measure of regular food expenditure based on a recall question about regular amounts spent on food (without disaggregating between food at home and at restaurants). Our second measure is again a recall question on total non-durable consumption, including food but also other items. Finally, 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), we 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]

[24] A concern with the test is that we do not use exogenous variation in $Opterm^{hd}_{2002}$ or $Opterm^{sp}_{2002}$to conduct the analysis. The reason is that the instrument used thus far is not a powerful predictor of contract form in this (small) sample. Still, we think that the biases that preclude a reliable estimation of the link between household wealth and the exposure to job loss due to different dismissal costs are less likely to affect a test

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 -0.19 (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 0.192-0.19=0.002). Conversely, households headed by an employee covered by a low dismissal cost contract would experience consumption growth of 1.92 percent (standard error: 0.088).

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. Firstly, holding else equal, workers whose contract is covered by high firing cost contract have higher expected income than those covered by a fixed-term contract. Even if job loss is a perfectly anticipated event, lower future income relative to current one should lead households to accumulate more wealth. Secondly, prudent households exposed to the uncertain event of job loss 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 a model with precautionary savings. We keep the model deliberately simple and abstract from a number of issues (like demographics or pension arrangements) to stress how a model readily comparable to others in the literature can generate wealth

based on consumption *changes*. While wealth is a stock variable that reflects past shocks, consumption growth is less likely to be affected by such shocks. Secondly, unobserved heterogeneity in risk aversion combined with self-selection into more secure jobs would bias the test against finding any impact of contractual dismissal costs on the slope of consumption growth. The reason is that risk averse workers self-select into positions where they are more likely to be promoted to a high dismissal cost contract.

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

$$\max_{c_t} U = \sum_{t=0}^{t=\infty} \frac{(c_t)^{1-\rho}}{1-\rho}$$

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

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

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

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 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 $\sigma_N = .1$. 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 1 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 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 $E(P)$: one where the probability of transiting into unemployment is 0.02 (intended to mimick the unemployment chances of workers covered by open-ended contracts) and another where the probability of transiting into unemployment

[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 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 in the panel component of the EFF.

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 panel of consumption data in Spain.

Solving the model: The model is solved by specifying, for 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]

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

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 $\{\frac{c_t}{Y_t^P}, \frac{A_t}{Y_t^P}\}$.

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.

[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}.

6.2 A model without uncertainty

Finally, we compare our estimates to those by a model *without* uncertainty. Carroll, Dynan and Krane (2003) show that the consumption policy in a model similar to that in the previous section when $r = \beta$ is $C_t = \frac{r}{1+r}A_t + \frac{r}{r-g}Y_t$. Sticking to our assumption that r equals 2% and there is no growth in permanent income, $\frac{r}{r-g}$ equals 1. The expected income drop would be $E\Delta Y_t = P(jobloss) \cdot (time\_unemployed + wage\_drop)$. 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 0.10·(0.5+0.1)·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 unfavourable to our case, we choose the 45 days compensation package and assume 3 years of tenure. The consumption drop would be 0.02·(0.5+0.1-45·3/365)·1=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 4 and 6 times larger.

Overall, our simple computations lead us to conclude that the wealth responses we estimate are much more likely to be generated by a precautionary saving motive than by lower future expected income.

7 Conclusions

We use the large dispersion in dismissal costs in the Spanish labor market and a new dataset 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 head exogenously obtains a high dismissal cost contract accumulate less financial wealth than comparable workers covered by low dismissal cost contracts. 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 most precise 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.

8 References

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Table 1: Summary statistics, combined EFF2002 and EFF2005

	Total sample	Head, open-ended	Head, fixed-term
Head with open-ended contract	0.805	--	--
	(0.396)		
Head with fixed-term contract	0.195	--	--
	(0.396)		
Age of household head	43.412	44.308	39.704
S.D.	(9.742)	(9.562)	(9.606)
Married	0.799	0.815	0.733
	(0.401)	(0.389)	(0.443)
Household size	3.218	3.244	3.107
	(1.239)	(1.210)	(1.346)
Prob. job loss (quarter),head			
Mean:	0.030	0.016	0.086
S.D.	(0.034)	(0.011)	(0.041)
# Years at current job	12.202	14.225	3.824
	(10.481)	(10.405)	(5.407)
Household earnings	27.205	22.784	13.103
	(19.031)	(14.488)	(7.572)
Whether head eligible for subsidy	0.278	0.261	0.579
	(0.448)	(0.490)	(0.494)
Amount head is eligible for	1.063	0.900	1.741
	(1.962)	(1.869)	(2.181)
Spouse covered by open-term	0.642	0.683	.408
(sample of working spouses)	(.479)	(.465)	(.49)
Spouse eligible for subsidy	0.379	0.311	0.543
(sample of working spouses)	(0.485)	(0.463)	(0.499)
Amount spouse is eligible for	1.357	1.128	1.913
	(2.169)	(2.067)	(2.308)
Non-durable expenditure	12.555	13.119	10.218
S.D.	(7.353)	(7.669)	(5.254)
Owns real estate	0.832	0.866	0.693
	(0.374)	(0.341)	(0.462)
Net worth			
Median	119.452	132.223	65.522
Mean	166.896	185.143	91.333
Net worth to earnings ratio			
Median	4.909	5.147	3.476
Mean	7.276	7.118	7.929
Financial wealth			
25th centile	0.977	1.167	0.522
Median	3.381	4.347	1.646
Mean	15.739	18.053	6.156
Financial wealth to earnings ratio			
Median	0.156	0.169	0.117
Mean	0.550	0.565	0.488

Sample: Average sample size of 3,853 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 at start of relationship 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	0.018	0.017	0.011	0.019	0.014	0.020	-0.001
two first years of job tenure -head	(0.005)***	(0.005)***	(0.005)**	(0.005)***	(0.006)**	(0.006)***	(0.005)
2. Subsidy amount	-0.013	-0.012	-0.011	-0.022	-0.022	-0.022	-0.006
*(Age< 35) - head	(0.007)**	(0.007)*	(0.007)*	(0.007)***	(0.007)***	(0.007)***	(0.007)
3. Subsidy * Female head	-0.002	-0.002	-0.001	--	--	--	--
-head	(0.007)	(0.007)	(0.007)				
4. Mean subsidy amount in	--	--	--	--	--	-0.005	0.032
two first years of job tenure -spouse						(0.005)	(0.006)***
Head is a female	-0.011	0.012	0.000	--	--	--	--
	(0.024)	(0.023)	(0.024)				
Head aged under 25	0.003	0.015	0.007	0.139	0.125	0.137	-0.025
	(0.067)	(0.069)	(0.069)	(0.073)*	(0.074)*	(0.073)*	(0.046)
Head aged 26-35	0.014	0.014	0.014	0.033	0.032	0.030	0.031
	(0.023)	(0.023)	(0.023)	(0.024)	(0.024)	(0.024)	(0.019)
Head aged 46-55	-0.037	-0.035	-0.035	-0.028	-0.028	-0.026	0.005
	(0.016)**	(0.016)**	(0.016)**	(0.017)	(0.018)	(0.017)	(0.014)
Head aged 56-65	-0.020	-0.020	-0.017	-0.012	-0.010	-0.011	0.007
	(0.020)	(0.019)	(0.020)	(0.021)	(0.022)	(0.022)	(0.015)
Household size 1	0.020	0.010	0.009	-0.021	-0.023	-0.020	-0.013
	(0.025)	(0.025)	(0.025)	(0.040)	(0.040)	(0.040)	(0.009)
Household size 3	0.004	0.008	0.010	0.033	0.033	0.035	-0.009
	(0.016)	(0.016)	(0.016)	(0.017)**	(0.017)*	(0.017)**	(0.014)
Household size 4	-0.005	0.000	0.004	0.017	0.020	0.018	-0.024
	(0.017)	(0.016)	(0.016)	(0.017)	(0.017)	(0.017)	(0.014)*
Household size 5	-0.009	-0.007	-0.001	0.013	0.017	0.012	-0.015
	(0.022)	(0.022)	(0.022)	(0.023)	(0.023)	(0.023)	(0.018)
Household size 6+	-0.026	-0.017	-0.006	0.003	0.013	0.005	-0.009
	(0.031)	(0.031)	(0.031)	(0.032)	(0.032)	(0.032)	(0.021)

Table 2: (continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Head started contract after 1997	0.072	0.064	0.080	0.054	0.070	0.053	-0.015
	(0.030)**	(0.029)**	(0.030)***	(0.032)*	(0.032)**	(0.032)*	(0.025)
Unemployment rate in region	-0.001	-0.001	0.000	-0.001	-0.001	-0.001	0.000
(year entered current firm)	(0.000)***	(0.000)**	(0.001)	(0.001)**	(0.001)	(0.001)**	(0.000)
Head entered labor market	-0.063	-0.053	-0.054	-0.041	-0.041	-0.041	0.039
after 1984	(0.016)***	(0.016)***	(0.016)***	(0.018)**	(0.018)**	(0.018)**	(0.014)***
Logarithm of earnings	--	0.096	0.090	0.082	0.076	0.083	0.035
(head and spouse, if present)		(0.011)***	(0.011)***	(0.012)***	(0.012)***	(0.012)***	(0.009)***
Spouse works	-0.002	-0.034	-0.036	-0.030	-0.033	-0.008	0.098
	(0.012)	(0.013)***	(0.013)***	(0.013)**	(0.013)**	(0.019)	(0.018)***
Secondary school, head	0.083	0.069	0.070	0.077	0.078	0.076	-0.015
	(0.022)***	(0.021)***	(0.021)***	(0.023)***	(0.023)***	(0.023)***	(0.014)
Upper secondary school, head	0.130	0.100	0.096	0.100	0.096	0.100	-0.018
	(0.023)***	(0.023)***	(0.023)***	(0.025)***	(0.024)***	(0.025)***	(0.016)
College, head	0.134	0.067	0.070	0.068	0.069	0.067	-0.037
	(0.022)***	(0.022)***	(0.022)***	(0.024)***	(0.024)***	(0.024)***	(0.017)**
Works for the public sector, head	-0.005	0.009	0.012	0.003	0.007	0.003	0.012
	(0.010)	(0.010)	(0.011)	(0.011)	(0.011)	(0.011)	(0.011)
Secondary school, spouse	0.040	0.032	0.035	0.031	0.034	0.031	0.018
	(0.022)*	(0.022)	(0.021)	(0.022)	(0.022)	(0.022)	(0.013)
Upper secondary school, spouse	0.028	0.009	0.012	0.014	0.016	0.015	0.034
	(0.024)	(0.024)	(0.023)	(0.024)	(0.024)	(0.024)	(0.018)*
College, spouse	0.043	0.005	0.009	0.017	0.020	0.019	0.034
	(0.024)*	(0.024)	(0.024)	(0.025)	(0.025)	(0.025)	(0.017)**
Works for the public sector, spouse	0.015	0.013	0.016	0.018	0.021	0.026	-0.029
	(0.013)	(0.013)	(0.013)	(0.013)	(0.013)*	(0.014)*	(0.016)*
Single	-0.021	-0.005	-0.003	0.009	0.012	0.012	0.011
	(0.033)	(0.032)	(0.032)	(0.038)	(0.038)	(0.038)	(0.013)
Widow/er	0.049	0.064	0.066	0.099	0.101	0.098	0.026
	(0.039)	(0.038)*	(0.038)*	(0.038)***	(0.038)***	(0.038)**	(0.018)
Divorced/separated	-0.005	0.005	0.010	0.078	0.080	0.078	0.014
	(0.034)	(0.034)	(0.034)	(0.045)*	(0.045)*	(0.045)*	(0.017)
Year 2003	0.024	0.019	0.012	0.017	0.010	0.016	-0.014
	(0.015)	(0.015)	(0.016)	(0.016)	(0.017)	(0.016)	(0.012)

Table 2: (continued)

	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Year 2005	-0.006	-0.011	-0.014	-0.004	-0.007	-0.003	-0.027
	(0.017)	(0.017)	(0.018)	(0.018)	(0.019)	(0.018)	(0.014)**
Year 2006	0.019	0.013	0.012	0.004	0.002	0.004	0.002
	(0.015)	(0.015)	(0.015)	(0.015)	(0.016)	(0.015)	(0.011)
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)***	(0.005)***	(0.005)***	(0.005)***	(0.004)
Tenure on the job squared, head	-0.004	-0.004	-0.004	-0.004	-0.003	-0.003	0.000
	(.000249)***	(.000246)***	(.000245)***	(.00027)***	(.000267)***	(0.0003)***	(.00042)
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)***	(0.000)
Tenure on the job-3, spouse	--	--	--	--	--	-0.002	0.092
						(0.003)	(0.004)***
Tenure on the job squared, spouse	--	--	--	--	--	0.000	-0.005
						(0.000)	(0.000)***
Tenure on the job cubed, spouse	--	--	--	--	--	0.000	0.000
						(0.000)	(0.000)***
Constant	0.491	0.555	0.583	0.533	0.559	0.524	0.341
	(0.038)***	(0.038)***	(0.039)***	(0.041)***	(0.042)***	(0.043)***	(0.030)***
Region fixed-effects	No	No	Yes	No	Yes	No	No
Sample size	3,662	3,662	3,662	3,144	3,144	3,144	3,144
R-squared	0.27	0.29	0.29	0.27	0.28	0.27	0.74

Source: Spanish Survey of Household Finances, sample of households headed by an employee between 23 and 65 years of age.
Standard errors are corrected for heteroscedasticity and arbitrary correlation among observations belonging to the cell at which subsidies are imputed: years at the job, region, age group and gender. Household earnings are the deviation from the weighted 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 first year of job tenure	-.00101	-.00045	-.00101	-.00035
	(.0005)**	(.00016)**	(.00041)**	(.00025)
2. Subsidy amount *(Age< 35)	-0.00013	.00006	--	--
	(.00015)	(.00013)		
Contract started after 1997	0.0066	0.0044	0.0082	-0.0004
	(.00254)	(.0021)	(.0046)	(.0028)
Age between 40-44	0.00315	0.0028	0.0015	0.00126
	(.0014)	(.00051)	(.0013)	(.0008)
Aged between 45-49	0.0025	0.0024	-.0064	-0.0034
	(.0010)	(.0010)	(.0036)	0.00093
Age between 50-54	0.00327	0.00304	-0.0029	-0.0012
	(.0015)	(.0007)	(.0016)	(.00178)
Age between 55-59	0.00762	0.0058	-0.0079	-0.004
	(.0060)	(.0022)	(.00446)	(.0016)
Primary or less	0.0102	0.0057	0.0095	0.0042
	(.0041)	(.00044)	(.0036)	(.0008)
Secondary school	0.00204	0.0016	0.00196	.00097
	(.00066)	0.0008	(.0017)	(.00189)
College	-0.00417	-0.00522	-.0107	-0.0091
	(.0027)	0.0012	(.0048)	(.00067)
Widow	-0.00727	-0.00427	-0.0079	
	(.0045)	(.0015)	(.0071)	
Divorced	-0.0004	-3.58E-06	0.00188	
	(.0046)	(.0032)	(.0018)	
Single	-0.0098	-0.0083	0.0013	
	(.0039)	(.0004)	(.0027)	
Year 2000	0.00976	0.0084	0.0159	0.0139
	(.00457)	(.001)	(.0067)	(.0007)
Year 2001	0.0087	0.008	.0182	0.0158
	(.0026)	(.0014)	(.0079)	(.0008)
Year 2002	0.0083	0.0077	0.0131	0.012
	(.0034)	(.0009)	(.0067)	(.0014)
Year 2003	0.0081	0.008	0.011	0.011
	(.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:	137008		87720	

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	-0.058	-0.059	-0.038	-0.071	-0.049	-0.051
two first years of tenure -head	(0.023)**	(0.023)**	(0.025)	(0.024)***	(0.027)*	(0.024)**
2. Subsidy amount	0.066	0.068	0.079	0.080	0.097	0.088
*(Age< 35) -head	(0.030)**	(0.030)**	(0.031)**	(0.033)**	(0.034)***	(0.033)***
3. Subsidy * Female head	0.070	0.070	0.068	--	--	--
	(0.037)*	(0.038)*	(0.038)*			
4. Mean subsidy amount	--	--	--	--	--	-0.084
in two first years of tenure- spouse						(0.023)***
Head is a female	-0.352	-0.329	-0.384	--	--	--
	(0.140)**	(0.142)**	(0.143)***			
Head aged under 25	-0.352	-0.341	-0.437	-0.394	-0.575	-0.410
	(0.231)	(0.234)	(0.235)*	(0.284)	(0.280)**	(0.284)
Head aged 26-35	-0.214	-0.214	-0.255	-0.215	-0.288	-0.239
	(0.108)**	(0.108)**	(0.109)**	(0.116)*	(0.117)**	(0.115)**
Head aged 46-55	0.298	0.301	0.326	0.215	0.244	0.216
	(0.090)***	(0.090)***	(0.090)***	(0.100)**	(0.100)**	(0.100)**
Head aged 56-65	0.738	0.738	0.788	0.651	0.703	0.647
	(0.117)***	(0.117)***	(0.117)***	(0.131)***	(0.131)***	(0.132)***
Household size 1	-0.757	-0.766	-0.770	-0.873	-0.852	-0.873
	(0.132)***	(0.133)***	(0.130)***	(0.206)***	(0.205)***	(0.206)***
Household size 3	0.227	0.231	0.218	0.157	0.148	0.160
	(0.088)**	(0.088)***	(0.087)**	-0.096	-0.095	(0.095)*
Household size 4	0.117	0.121	0.105	0.058	0.036	0.046
	(0.091)	(0.091)	(0.090)	(0.096)	(0.096)	(0.096)
Household size 5	0.287	0.290	0.283	0.224	0.210	0.207
	(0.128)**	(0.128)**	(0.127)**	(0.132)*	(0.131)	(0.133)
Household size 6+	-0.060	-0.051	-0.017	-0.107	-0.076	-0.121
	(0.184)	(0.184)	(0.182)	(0.193)	(0.192)	(0.193)
Contract started after 1997	-0.032	-0.040	-0.057	0.096	0.090	0.072
	(0.148)	(0.147)	(0.148)	(0.160)	(0.161)	(0.161)

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	-0.010
(year entered current firm)	(0.004)***	(0.004)***	(0.004)	(0.004)***	(0.005)	(0.004)**
Head entered labor market	0.140	0.149	0.139	0.029	0.033	0.045
after 1984	(0.090)	(0.090)*	(0.089)	(0.099)	(0.099)	(0.100)
Logarithm of earnings	--	0.094	0.099	0.160	0.167	0.156
(head and spouse, if present)		(0.070)	(0.072)	(0.080)**	(0.082)**	(0.081)*
Spouse works	-0.192	-0.224	-0.216	-0.238	-0.234	-0.001
	(0.075)**	(0.078)***	(0.078)***	(0.079)***	(0.079)***	(0.111)
Secondary school, head	0.055	0.041	0.028	0.003	-0.009	-0.007
	(0.099)	(0.100)	(0.101)	(0.110)	(0.111)	(0.109)
Upper secondary school, head	0.264	0.234	0.201	0.232	0.198	0.234
	(0.116)**	(0.118)**	(0.117)*	(0.129)*	(0.128)	(0.128)*
College, head	0.859	0.794	0.788	0.807	0.806	0.798
	(0.112)***	(0.123)***	(0.123)***	(0.139)***	(0.139)***	(0.139)***
Works for the public sector	-0.158	-0.145	-0.115	-0.153	-0.124	-0.146
	(0.078)**	(0.079)*	(0.078)	(0.086)*	(0.085)	(0.087)*
Secondary school, spouse	0.192	0.185	0.154	0.190	0.165	0.192
	(0.113)*	(0.114)	(0.113)	(0.119)	(0.118)	(0.118)
Upper secondary school, spouse	0.368	0.349	0.316	0.326	0.295	0.341
	(0.134)***	(0.136)**	(0.136)**	(0.141)**	(0.141)**	(0.140)**
College, spouse	0.582	0.545	0.510	0.496	0.460	0.511
	(0.141)***	(0.143)***	(0.143)***	(0.150)***	(0.150)***	(0.149)***
Works for the public sector, spouse	-0.116	-0.118	-0.146	-0.120	-0.142	-0.105
	(0.100)	(0.099)	(0.099)	(0.100)	(0.099)	(0.101)
Single	1.448	1.464	1.451	1.565	1.558	1.546
	(0.168)***	(0.168)***	(0.166)***	(0.205)***	(0.203)***	(0.205)***
Widow/er	0.901	0.916	0.865	0.692	0.545	0.692
	(0.228)***	(0.227)***	(0.226)***	(0.306)**	(0.315)*	(0.307)**
Divorced/separated	0.474	0.484	0.501	0.453	0.469	0.440
	(0.188)**	(0.188)**	(0.187)***	(0.289)	(0.285)	(0.290)
Year 2003	0.006	0.000	0.060	-0.029	0.013	-0.029
	(0.083)	(0.083)	(0.085)	(0.089)	(0.092)	(0.089)
Year 2005	0.345	0.339	0.292	0.350	0.295	0.362
	(0.097)***	(0.097)***	(0.100)***	(0.103)***	(0.107)***	(0.103)***

Table 4: OLS estimates of the impact of subsidies to contract conversion on household financial wealth

	(1)	(2)	(3)	(4)	(5)	(6)
Year 2006	0.053	0.047	0.107	0.004	0.050	0.013
	(0.081)	(0.081)	(0.082)	(0.087)	(0.087)	(0.087)
Tenure on the job-3, head	0.053	0.048	0.040	0.061	0.054	0.062
	(0.022)**	(0.022)**	(0.022)*	(0.025)**	(0.025)**	(0.025)**
Tenure on the job squared, head	-0.002	-0.002	-0.001	-0.003	-0.002	-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	--	--	--	--	--	-0.030
						(0.018)*
Tenure on the job squared, spouse	--	--	--	--	--	0.002
						(0.002)
Tenure on the job cubed, spouse	--	--	--	--	--	0.000
						(0.000)
Constant	-2.734	-2.672	-2.860	-2.562	-2.762	-2.683
	(0.183)***	(0.188)***	(0.199)***	(0.204)***	(0.218)***	(0.217)***
Region fixed effects	No	No	Yes	No	Yes	No
Sample size	3,662	3,662	3,662	3,144	3,144	3,144
R-squared	0.16	0.16	0.18	0.17	0.19	0.17

Notes: Sample of households headed by an employee between 18 and 65 years of age. We pool the 2002 and 2005 waves. Standard errors are corrected for heteroscedasticity and arbitrary correlation among observations belonging to the cell at which subsidies are imputed: years at the job, region, age group and gender. Household earnings are the deviation from the weighted sample mean.

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)***	(0.006)**	(0.006)***
2. Subsidy amount * (Age <=35)	--	-0.012	-0.011	-0.022	-0.022	-0.022
		(0.007)*	(0.007)*	(0.007)***	(0.007)***	(0.007)***
3. Subsidy amount * (Head is female)	--	-0.002	-0.001	--	--	--
		(0.007)	(0.007)			
4. Subsidy amount spouse was eligible for	--	--	--	--	--	-0.005
						(0.005)
5. Constant	--	0.555	0.583	0.533	0.559	0.524
		(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	-0.004	-2.504	-3.157	-3.712	-4.174	-2.831
	(0.091)	(1.493)*	(2.478)	(1.483)**	(1.939)**	(1.587)*
2. Spouse covered by high firing cost contract	--	--	--	--	--	-2.992
						(1.008)***
3. Constant	-2.700	-1.307	-1.016	-0.584	-0.418	-0.184
	(0.197)***	(0.856)	(1.483)	(0.834)	(1.130)	(0.966)
Panel C: Fraction of gross earnings held as financial wealth (at the median)						
3. Head has an open-ended contract	**0.000**	**0.293**	**0.369**	**0.434**	**0.488**	**0.331**
4. Spouse has an open-ended contract	--	--	--	--	--	**0.350**
Region dummies?	No	No	Yes	No	Yes	No
Sample size:	3,662	3,662	3,662	3,144	3,144	3,144

The same set of regressors used in Tables 3 and 4 is used in all specifications, but not shown to save space. Standard errors (in parentheses) are corrected for arbitrary autocorrelation at the age-region-gender-year of entry at the firm level.

Table 6: 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	0.771	-3.651	1.324
high dismissal cost	(1.210)	(1.751)**	(1.335)
Head aged under 25	-0.305	-0.387	-1.418
	(0.290)	(0.453)	(0.407)***
Head aged 26-35	-0.086	-0.225	-0.334
	(0.098)	(0.137)*	(0.116)***
Head aged 46-55	0.226	0.345	0.279
	(0.105)**	(0.133)***	(0.087)***
Head aged 56-65	0.647	1.086	0.661
	(0.133)***	(0.163)***	(0.102)***
Contract started after 1997	-0.062	0.184	-0.143
	(0.192)	(0.246)	(0.183)
Unemployment rate in region	-0.01	-0.012	-0.002
(year entered current firm)	(0.004)**	(0.005)**	(0.004)
Head entered labor market	0.053	-0.285	-0.189
after 1984	(0.111)	(0.145)*	(0.105)*
Logarithm of earnings	0.09	0.58	-0.16
	(0.130)	(0.174)***	(0.135)
2. Fraction of gross earnings as financial wealth (at the median)	**-0.090**	**0.427**	**--**
Sample size:	3,144	3,135	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.

Table 7: 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 hh	
	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 (5)/[(1)+(5)]
Sample means:	0.282	0.011	0.010	0.015	0.036	0.113
Model 1: Open-ended contract (head) as a regressor						
1. Fixed-term contract	0.269	0.047	0.003	0.012	0.063	0.189
2. Open-ended contract	0.276	0.016 (***)	0.002	0.006 (**)	0.024	0.080
Model 2: Subsidy amount (head) as a regressor						
1. Zero subsidies	0.271	0.037	0.002	0.009	0.040	0.129
2. 1,000-euro subsidies	0.261	0.036	0.003 (***)	0.008 (*)	0.047	0.153

Entries are fitted probabilities of a multinomial logit that has "Not asked for a loan" as the base outcome. (***), (**) and (*) mean that the latent variable coefficient is significant at the 1, 5 and 10 percent, respectively. Model 1 uses "Open-ended contract" as a regressor, model 2 uses our instrument (subsidies). Rest of covariates: age dummies, marital status, logarithm of income, schooling of head and spouse, family size, third order polynomial in tenure minus 3.

Table 8: The effect of an open-ended contract on the ratio of financial wealth over income

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

	25th centile	50th centile	75th centile
1. Covered by an open-ended contract	-1.7	-1.0	-2.9
95% confidence interval	[-5.8, 2.4]	[-4.1, 0.4]	[-10, 2.7]
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 gross yearly income held as wealth	**0.089**	**0.109**	**-**

Table 9: The impact of the risk of losing the job on 3-year consumption growth

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)
1. Head covered by open-ended	-0.122	-0.12	-0.198	-0.19	-0.133	-0.128
contract	(.064)*	(.0642)*	(.073)***	(.073)**	(.065)**	(.061)**
2. Spouse covered by open-ended	--	-0.0065	--	-0.0528	--	-0.0359
contract		(.0639)		(.0737)		(.0607)
Spouse works		-0.0307	0.062	-0.0518	0.077	-0.022
		(.0606)	(0.052)	(.0735)	(0.043)	(.0576)
Head between 23 and 25	0.026	0.0309	-0.303	-0.309	-0.293	-0.296
	(.216)	(.2178)	(.252)	(.250)	(.234)	(.207)
Head between 26 and 35	0.085	0.0877	-0.0251	-0.0242	-0.002	-0.0004
	(.075)	(.0757)	(.0868)	(.087)	(.066)	(.072)
Head between 46 and 55	-0.0202	-0.0218	-0.0086	-0.012	0.034	0.0307
	(.0527)	(.0528)	(.0652)	(.0607)	(.0544)	(.050)
Head between 56 and 65	0.054	0.0447	0.0229	0.0104	0.028	0.0172
	(.066)	(.0679)	(.0768)	(.0781)	(.0625)	(.065)
Change in household size	0.155	0.155	0.1717	0.1755	0.143	0.1429
	(.0313)***	(.0313)***	(.0361)***	(.0360)***	(.0295)***	(.0298)***
Change in number of children 0-3	-0.149	-0.147	-0.1279	-0.133	-0.126	-0.1259
	(.0678)**	(.068)**	(.078)*	(.0783)*	(.0609)**	(.0647)**
Change in number of children 4-7	-0.078	-0.0769	-0.1828	-.175	-0.141	-0.139
	(.0705)	(.0706)	(.0817)**	(.0811)**	(.0672)**	(.0671)*
Change in number of children 8-11	-0.0074	-0.004	-0.1074	-0.0892	-0.067	-0.0623
	(.0628)	(.063)	(.073)	(.0725)	(.0581)	(.0599)
Change in number of children 12-15	0.0025	0.004	-0.1365	-.126	-0.089	-0.0867
	(.0535)	(.0536)	(.062)**	(.0617)**	(.0509)*	(.051)

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	-0.0752	-0.0751	-0.0957	-.0886	-0.072	-0.0713
	(.0454)	(.0455)	(.052)	(.0523)*	(.0442)	(.0432)
Single	-0.0375	-0.0562	-0.0481	-0.057	-0.107	-0.13
	(.1049)	(.1077)	(.1212)	(.123)	(.110)	(.1023)
Widow	0.2881	0.267	0.37	0.326	0.314	0.287
	(.2277)	(.229)	(.262)	(.264)	(.250)	(.218)
Divorced	0.2497	0.2331	0.2468	0.223	0.156	0.135
	(.193)	(.194)	(.222)	(.224)	(.128)	(.1848)
Secondary education, head	0.0113	0.0122	0.054	0.053	0.041	0.0429
	(.0627)	(.0628)	(.0722)	(.072)	(.0536)	(.0597)
Upper secondary, head	0.0338	0.0506	0.0758	0.0784	0.032	0.0329
	(.0618)	(.0696)	(.080)	(.0800)	(.0549)	(.066)
College, head	-0.0052	0.0431	0.0099	0.0306	-0.016	-0.0007
	(.0421)	(.063)	(.0718)	(.0726)	(.0542)	(.060)
Constant	0.0717	0.0846	0.179	0.192	0.15	0.1633
	(.0737)	(.0765)	(.085)**	(.0879)**	(.072)**	(.072)

Notes: Sample size: 625. Standard errors are in parentheses.

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 (4)
Mean W /Y of:				
All households	0.4042	0.6433	0.2391	0.5915
1.. 20th-30th W /Y Percentile	0.1750	0.3533	0.1783	1.0188
2. 40th-50th W /Y Percentile	0.3154	0.5339	0.2185	0.6927
3. 60th-70th W /Y Percentile	0.4666	0.7249	0.2582	0.5534
4. 80th-90th W /Y Percentile	0.7032	1.0257	0.3226	0.4587

Table A.1.: The distribution of the probability of losing the job, by education

Panel A: Probability of transiting into unemployment in the next quarter (Source: EPA 1998-2001)

	Open-ended contract	Fixed-term contract
Total	0.011	0.088
Primary school	0.018	0.111
Secondary school	0.012	0.082
Upper secondary school	0.009	0.074
College	0.006	0.062

Predicted quarterly probabilities of transiting from employment into unemployment in a sample of working individuals. Independent variables are employee, employee covered by open-ended contract, 3 education, dummies, 5 age dummies, 3-quarter dummies and 2 year dummies (1999 and 2001).

Table A.1 (cted) The distribution of the probability of losing the job, by education

Panel B: 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 in Panel B 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.

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	All periods: 4200 if age<30 , 3000 if female >30, 2400 if male >30			
2. Aragon	None	4200 if female or age>45	5160 if female or age>45	5160 if female or age>45
		3000 if age>=41	4500 if age>=41	4500 if age>=41
			3600 if male age<30	3600 if male age<30
3. Asturias	4500 for all	4500 for all	None	4,200 if female or age>=45
				3600 otherwise
4. Baleares	None	None	None	1652.78 if female
5. Canarias	None	3,600 if female or age<25	3,600 if female or age<25	None
			3,000 otherwise	
6. Cantabria	None	3900 if female or age< 30	None	5408 if age>46
		3300 if male 30<age<44		3606 if age<30
		3,600 if male above 40		2163 otherwise
7. Castilla-Leon	None	5112 if female or age <30	5112 if age <30	4507.59 if age <30
		3300 rest of males	3900 if female age>=30	3305.57 if female age>31
			1800 if male age >45	1803 if male age>41
8. Castilla-La Mancha	None	3600 if females	None	3600 if female
		3000 if age<30		3000 if age>45 or age<30
9. Catalonia	None			
10. Valencia	None	30% of payroll tax	30% of payroll tax	30% of payroll tax
11. Extremadura	10655.94 if age<45	10100 if age<=30	14027.62 if age>46	5217.076 if female age>46
	13402.57 if age>45	11180 if age>30 and age<=40	11178.83 if age<45	4296.416 if male age>46 (…)
		14027 if age>40		
12. Galicia	None	4200 euro if female or age<30	None	None
		3000 if age>45		
13. Madrid	None	6000 euro if female	7800 if female	9000 if female
		6000 euro if age<30 or age>40	7800 if age<25 or age>40	6600 if age<25 or age>40
14. Murcia	None	1800 for all	1800 for all	1800 for all
		2400 if age<30	2100 if age<30	2100 if age<30
15. Navarra	None	None	None	None
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	5161 if female or age>45 4501 if age>=41 3601 if male age<30	4500 if female 4125 if age<30 or age>45	4500 if female 4125 if age<30 or age>40	1200 for all, 1800 if age>45
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 4207 if age >45 1803 otherwise	same as previous year	same as previous year	3005 if age<30 4207 if age>45 4808 if female
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