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 high 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 upgrade low severance payment contracts into high severance payment ones. We find that workers covered by fixed-term contracts accumulate more financial wealth. An increase of one standard deviation in the probability of losing the job 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

Economic theory predicts that households that are more exposed to the risk of losing their job postpone their consumption and accumulate more assets to build a buffer that permit absorbing income losses associated to unemployment spells (see Caballero, 1990, or Carroll, 2001). The extent of precautionary savings has important consequences for the sensitivity of consumption to increases in income (Hall, 2006) and for understanding the determinants of the distribution of household wealth. From a policy perspective, in a situation of growing unemployment it is important to assess if households exposed to lay-offs have accumulated wealth to sustain consumption during an unemployment spell and how large those buffers are. Our study exploits the large variation in the costs of dismissing workers covered by different contracts in the Spanish labor market to quantify the size of wealth accumulated among households exposed to the risk of job loss.

A large literature has used different methods to establish if households facing (or perceiving) higher chances of losing their job have lower consumption levels and/or accumulate higher levels of wealth. The results are not uncontroversial; Carroll, Dynan and Krane (2003) find that households with higher exposure to the risk of losing their job (and sufficiently high permanent income) have more wealth, consistent with the precautionary saving model. Fuchs-Schündeln and Schündeln (2005) use the reunification of Germany and the transition from a (possibly) risk-free environment to a capitalist economy to examine if affected households accumulated more wealth, using civil servants as a control group. They find evidence consistent with the hypothesis of precautionary saving. Engen and Gruber (2001) document that unemployment subsidies crowd out private wealth accumulation, a finding that is consistent with the idea that workers accumulate precautionary savings. On the other hand, Guiso, Jappelli and Terlizzese (1992) or the survey of of Browning and Lusardi (1996) find little evidence for precautionary savings.

The discrepancy of the results may be due to several problems. First, it is hard to measure to what extent an individual is exposed to the risk of losing his or her job.[1] Second, even when one can find a group that does experience a higher probability of transiting into unemployment, it is not always the case that the higher probability is uncorrelated with the unobserved propensity to save (see Fuchs-Schündeln and Schündeln, 2005 and Lusardi, 1997 for evidence on self-selection). Second, workers who are relatively more

[1] A very interesting approach measures subjective expectations of job loss (Manski and Straub, 2000, Arrondel, 2009). Our study focuses on objective measures of job loss.

exposed to the risk of losing the job are also more likely to have consumed their wealth balances in recent unemployment spells, biasing downward the link between household wealth and exposure to job loss. Finally, credit constraints often generate predictions that are empirically hard to distinguish from a precautionary saving motive: individuals who are more exposed to the risk of losing the job are also riskier borrowers from the bank's perspective.[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, differences in dismissal costs cause that identifiable groups of the population face very different probabilities of transiting into non-employment. During the eighties, Italy, Spain, Germany, Sweden, Portugal and France (among other countries) introduced low firing cost contracts as a way to fight against unemployment. Typically, countries that introduced fixed-term contracts already featured rigid labor markets with very high dismissal costs. Fixed-term contracts allowed firms hire workers paying a small firing cost in the event they needed to downsize (see Dolado, Garcia-Serrano and Jimeno, 2002, for an overview). The introduction of fixed-term contracts has generated labor markets where identifiable groups of individuals face very different probabilities of transiting into unemployment for reasons unrelated to their own choice, but to firm's labor demand. Among all countries that introduced fixed-term contracts, Spain is the country with the highest share of fixed-term contracts (30%), thus providing an ideal setting to analyze the saving decisions of households differently exposed to the risk of losing the job.[3]

Secondly, due to regional regulations in the Spanish labor markets, the incidence of fixed-term contracts varies across regions and demographic groups. In 1997, six out of the 17 Spanish regions implemented subsidies to firms that promoted workers covered by a fixed-term contract into regular (high

[2]For example Carroll, Dynan and Krane (2003) note that a fraction of the buffer stock accumulated by households more exposed to the risk of losing the job could be due to higher loan-to-value requirements by banks. They quantify an upper bound of such confounding effect at most half of the estimated buffer stock. Fuchs-Schündeln and Schündeln (2005) mention the problem, but do not provide an explicit test quantifying the impact.

[3]A large literature has examined the impact of dismissal costs on unemployment rates and employment fluctuations but, to our knowledge, there is little evidence on how household wealth responds to the (lack of) turbulence generated by those labor market regulations.

dismissal costs) open-ended contracts. Between 1998 and 2004, most regions started giving subsidies to contract conversion. Furthermore, different regions targeted different demographic groups and gave very different subsidies. As a result, there is substantial variation in the firm's incentives to offer their workers an open-ended (high dismissal cost) contract. Such arguably exogenous variation permits us to estimate the reaction of wealth to differences in the risk of losing the job among workers who are comparable in other dimensions. Namely, we can compare workers who got a high dismissal cost contract because a subsidy to contract conversion was available to those who did not get a high dismissal cost contract because of the absence of the subsidy in their region-age-gender cell.

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 containing detailed information on households' assets, consumption and on the labor market situation of each household member.[4] In addition, the EFF contains information about a number of outcomes that allows us to test the validity of our approach. In particular, we can use information on credit rejections to examine if wealth differences are due to credit-supply or household's demand factors.[5]

We present the empirical results in three steps. First, we document that the amount of the subsidy to contract conversion in the age-gender-region cell the worker belonged to in the first two years of tenure with the current firm is a significant predictor of the type of contract held by the household head. Second, we present intention-to-treat estimates documenting that workers whose firm could obtain a higher subsidy by converting their fixed-term contract into a permanent one accumulated lower levels of financial wealth. Finally, we construct two-stage least squares estimates indicating that households whose head obtained a high dismissal cost contract as a consequence of the regional subsidies have financial wealth-earnings ratios between 20% and 30% lower than households whose head had a low firing cost contract. Nevertheless, we do not find that high dismissal cost contracts lead to higher wealth

[4]For example, we do not need to construct saving rates (that are typically noisy) or measures of wealth based on interest income, but can examine household wealth levels directly.

[5]In ongoing work, we are examining if consumption and consumption growth responses to the risk of losing the job are consistent with the link we estimate between wealth and having a high dismissal cost contract.

when we include the net value of owner-occupied housing. Finally, we document that subsidies to contract conversion have little predictive ability in experiencing a credit rejection. We argue that those findings are consistent with a precautionary saving motive.

Section 2 summarizes the legislation of dismissal costs in Spain and its implications for wealth accumulation. Section 3 presents the data. Section 4 presents the identification strategy and Section 5 presents the main results.

2 Dismissal costs in Spain

Spain had one of the most rigid labor markets among European countries before 1984. The main form of contract was what from now on we term an "open-ended" contract. Such contracts 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. 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. In 1984, in a context of high unemployment rates, a menu of contracts that were exempted from the general rule of high severance payments were introduced. The legal figure used was the authorization of extending contracts that before 1984 were used to regulate seasonal jobs to other types of labor relationships ("fixed-term" contracts). Such contracts initially had dismissal costs of 12 days per year worker, and the worker had no right to sue the employer claiming that the lay-off was unfair.

Fixed-term contracts were heavily used by Spanish employers and, by 1994, 30% of workers reported to the Spanish Labor Force Survey (EPA, in its Spanish initials) being covered by a fixed-term ("low-firing cost" contract). While subject to small fluctuations, the share has remained stable since, despite of several reforms that tried to address the problem of the duality in the labor market. The use of such contracts has been widespread across industries, and even the Public Administration has made substantial use of such contracts to fill positions (Dolado, Garcia-Serrano and Jimeno, 2002). Not surprisingly, fixed-term contracts are a strong predictor of the probability of transiting into unemployment. According to the Spanish Labor Force Survey (EPA, in its Spanish initials) an individual whose job position is regulated by a fixed-term contract was 6% more likely to be observed in unemployment in the next quarter than a worker covered by an open-ended contract (see Güell and Petrongolo, 2007 or García-Ferreira and Villanueva, 2007 for further reference).

Our study considers one of the attempts to reduce the use of fixed-term contracts to obtain exogenous variation in the fraction of the workforce that is exposed to the risk of losing the job. In 1997, several of the 17 Spanish regions introduced regional subsidies to incentive firms to use open-ended contracts to hire workers (see García-Pérez and Rebollo-Sanz, 2009, for a detailed explanation). A set of subsidies were given to firms that converted fixed-term contracts into open-ended ones. A second set of subsidies were given to firms that hired unemployed workers using open-ended contracts. In this study, we focus only on subsidies to the conversion of fixed-term contracts into open-ended ones. Those subsidies typically took the form of either a lump sum given to the firm by the public administration in the year when the conversion took place or to a reduction to the pay-roll tax. Table A.2 documents how amount that subsidized contract conversion varied across gender and age groups. It is also worth noting that some major regions did not implement those subsidies at all between 1997 and 2004 (Catalonia). Other regions, like Andalucia offered such subsidies to contract conversion to particular age groups. Finally, other major regions like Madrid did not offered them until 2002. The amount subsidized was substantial; García-Pérez and Rebollo-Sanz (2009) estimate that the lump sum received by the firm could amount around 20% of the yearly labor cost of the worker.

In sum, the incentive to the firm to convert a fixed-term contract into an open-ended one varied across regions of residence, with the year in which the contract started and with the age and gender of the worker. Thus, we exploit those differential incentives to obtain arguably exogenous variation in the exposure of workers to the risk of losing the job. If regional subsidies to contract conversion affected the stock of workers covered by an open-ended contract across age, gender, time at the firm and region cells (a hypothesis we explicitly test below), we will be able to compare the wealth accumulation patterns of similar workers who end up covered by different dismissal costs for exogenous reasons.

2.1 Modelling issues

We build on analytical results by Blundell and Stoker (1999). 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 unemployment benefit b 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, and differ a rigorous discussion of the case of hard-to-adjust consumption to a future draft. 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, $pb + (1-p)y$, and variance, $Var_1(Y) = (1-p)p[y-b]^2$. Following Blundell and Stoker (1999), we define the present value of expected wealth in period 1 as the sum of the initial wealth in period 1 and the expected stream of income in period 2, as follows:

$$W = W_1 + pb + (1-p)y$$

and 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

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

We are implicitly assuming that the individual can borrow against the expected value of future income (including expected payments due to layoffs). While perhaps not a realistic assumption, it permits obtaining closed-form solutions. Blundell and Stoker (1999) linearize around the perfect-certainty solution of consumption (that is linear in first-period wealth) and obtain the following consumption levels in the presence of risk:

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

Equation (1) implies that when we compare two individuals A and B, with the same level of expected income, but where the first has a zero probability of losing the job but the second is exposed to a non-zero chance of unemployment, the second one must have a lower level of consumption.

A second implication is that the consumption growth of both individuals is different; the individual who is exposed to the risk of losing the job postpones consumption to the future and hence will exhibit higher consumption growth. Blundell and Stoker (1999) and others derive the following expression for consumption growth

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

In Equation (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. Now, taking expectations in Equation (2a) over the distribution of Y one obtains the following expression:

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

That is, workers who, as of period 1, realize that they are exposed to a higher risk of losing their job are more likely to postpone consumption and thus experience higher consumption growth than workers in safer jobs.

Overall, the discussion thus far suggests three testable hypotheses:

- First, do workers who are more exposed to the risk of losing the job consume less?
- Second, do workers who are more exposed to the risk of losing the job exhibit higher consumption growth?
- Third, do workers who are more exposed to the risk of losing the job hold more (liquid) wealth?

3 Dataset and summary statistics

The main dataset we use is 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).* 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-non-response; in the final version of the paper we will combine the estimates done separately on each of the five data sets to take into account imputation uncertainty and facilitate a correct use of the data –for details on the EFF imputations see Bover (2004) and Barceló (2006).

The dependent variable:

We use various measures of wealth. The first is "liquid" wealth, i.e., a subset of wealth that we assume to be easily cashed in the event of an emergency. It contains amounts held in checking and saving accounts, mutual funds, stock (either listed or not), all types of bonds and other financial assets. Nevertheless, there is a discussion regarding whether or not households are able to use housing equity to finance a period of unemployment. For example, Carroll, Dynan and Krane (2003) argue that US households have access to home equity loans that permit "cashing" housing wealth. Shore and Sinai (forthcoming) argue that if both durable goods (housing) and non-durable goods enter the utility function, prudent agents react to increased risk as "risk-loving" gamblers and *increase* their consumption of housing.[6] Both arguments would lead to include the value of housing in the wealth measure (for different reasons). On the other hand, Engen and Gruber (2001) argue that housing wealth cannot act as a buffer against unemployment risk (and find evidence in that direction). Thus, we also experiment with two alternative measures that include, sequentially and in addition to "liquid" wealth, (a) the value of real estate properties net of associated debts and (b) the net value of main house. Throughout the paper, we assume that vehicles, pension funds (the Spanish version of IRAs, which were not cashable

[6]The reason for such behavior is that in the event of a minor income drop, households who do *not* find it optimal to adjust their housing after a shock must reduce non-durable consumption substantially. An increase of the probability of experiencing a large income drop (that would make a housing adjustment the best option) relative to the probability of a minor income loss diminishes the likelihood of the bad scenario of overconsumption of housing services and underconsumption of nondurable goods. Thus, a household that faces the risk of large losses acts as risk lover toward housing consumption.

in case of unemployment), life insurance and business wealth cannot serve a precautionary saving motive and exclude them from the analysis.

Sample selection in the Wealth Survey:

The main sample is composed of households headed by an employee head between 23 and 65 years of age.[7] We excluded those cases that had total labor earnings below 1,000 euros of 2005. The reason for dropping the self-employed is that the instrument we use (regional subsidy to contract conversion) was only available for employees. The sample contains 3,853 household-years. As we take logarithms of wealth in most of the analysis, we lose another 114 cases that have zero financial wealth.[8]

Our measure of the probability of losing the job: Fixed-term contracts.

We establish the risk of losing the job according to the contract type of in the first job reported by the head of the household. Dolado, García-Serrano and Jimeno (2002) document that the public sector was one of the most active employers in using fixed-term contract, and thus we chose not to exclude individuals who work for the public sector. Additionally, we also treat civil servants as an additional worker covered by an open-ended contract; Fuchs-Schündeln and Schündeln (2005), have documented the self-selection of workers with strong preferences for saving into civil servant status, and we do not want to select our sample on the basis of attitudes toward saving. Finally, we also included as workers covered by "fixed-term contracts" those who declare not being covered by a contract currently.

3.1 Summary statistics

We start by documenting that the type of contract held by a worker does correlate with the probability of transiting into unemployment. We then present summary statistics comparing the income, wealth and consumption of households headed by workers with open-ended and with fixed-term contracts.

[7]The definition of head of the household is not left to the household, but was determined based on the relative incomes of household members.

[8]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.

3.1.1 Documenting differences in the exposure to the risk of losing the job

Table A.1 in the Appendix shows suggestive evidence documenting that workers covered by fixed-term contracts are indeed more likely to transit into unemployment. There we present results of (gender specific) *logit* regressions of the probability of transiting into unemployment on several covariates and, most importantly, a measure of whether or not the individual has a fixed-term contract. The omitted group in the regressions are self-employed workers. Clearly, employees with an open-ended contract face a much lower probability of transiting into non-employment than either employees with a fixed-term contract and similar to that of self-employed workers.

We show various measures of exposure to the risk of losing the job in Table 1. Each cell in the Panel A of Table 1 represents the predicted probability of transiting from employment to unemployment in a quarter for groups of the population defined by the type of contract. The probabilities are estimated using the estimates in Table A.1. Panel B provides an alternative measure of job insecurity. The EFF asks in each wave the number of months that each household member was working during the year prior to the interview (2001 in the case of the 2002 wave and 2004 for the 2005 wave). Using the fact that the EFF has a longitudinal component, we estimated a *logit* model of the probability of spending at least one month in unemployment in 2004 for each employee in 2002 that was also successfully interviewed in the 2005 wave. The explanatory variables are basically the same as in Table A.1. While the statistics in Panel A of Table 1 measure high-frequency moves from employment to unemployment, the statistics in Panel B measure long-run exposure to the risk of losing the job.

In Panel B, employed heads of household that are employees covered by a fixed-term contract are 7.7 percentage points more likely to move from employment to unemployment than similar workers with open-ended contracts. From a longer-run perspective, workers covered by a fixed-term contract were 13.2 percentage points more likely to experience a spell of unemployment of at least a month two years later than workers with an open-ended contract. The differences are present for all levels of skill. Table 1 suggests that the differences in the exposure to the risk of losing the job are substantially different according to the type of contract.

3.1.2 Differences in income, wealth and consumption

The summary statistics of the EFF sample are presented in Table 2. There, we split the sample according to our measure of "exposure to unemployment risk". The first group are households whose the 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 suggest that the group of households headed by an employee with an open-ended contract are older and wealthier than the group of households where a member has a fixed-term contract. Both groups differ among many dimensions. Heads covered by an open-ended contract are older than heads covered by a fixed-term contract (44 vs 40 years of age). They are also more likely to be in the firm (14 years of tenure vs 3.5) and have higher household earnings. Given those comparisons, it is not surprising that households headed by an individual with an open-ended contract are more likely to own a house (87 percent vs 69 percent), have higher wealth-earnings ratios.

The summary statistics 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. Thus, we need to use an instrument that changes the contract of a worker holding constant those characteristics that lead to higher wealth accumulation ability.

3.1.3 The instrument: eligibility to subsidies to contract conversion

We compute the subsidy an individual is eligible to by using the reported time at the job, the current (exact) age, gender, and region of residence.[9] We have used the subsidy available during the first and second years of the time at the firm, as that is the time when most contracts were converted according to the 2003-2004 waves of the Spanish Labor Force Survey (see below).

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

4 Identification strategy

There are several problems in estimating the link between the risk of losing the job and the amount of wealth accumulated by the worker for a precautionary motive. First, several authors have convincingly argued that workers less averse to risk are more prone to end up in a job or industry where transitions into unemployment are more prevalent (Lusardi, 1997, Fuchs-Schündeln and Schündeln, 2005). Even controlling for risk aversion, another problematic issue is that workers who are more exposed to the risk of losing the job will most likely have experienced recent unemployment spells and used the accumulated buffer of wealth to sustain consumption during the spell. Thus, in a cross-section of data workers more exposed to the risk of losing the job will be more likely to be observed with small wealth holdings, even if a precautionary motive is indeed present. Carroll, Dynan and Krane (2003) present simulations that document a sharp and relatively persistent drop in household wealth following an unemployment spell.

Our identification strategy exploits the fact that firms differential incentives to use a high or low severance payment contract to hire a worker depending on the year when the contract was made, its region of residence and the demographic group the worker belonged to. In an economy with limited geographical mobility, such variation is in principle not related to the preferences of workers to avoid risk. Secondly, we test whether wealth holdings differ among workers who have been employed by the firm during the same number of periods but whose employers had access to different incentives to hire using a high dismissal cost contract. By holding tenure at the firm constant, we control for sources of differences in wealth due to recent uses of precautionary wealth holdings.

We describe the exact identification strategy below.

4.1 The first stage: Do subsidies to contract conversion 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 use two main strategies. Firstly, we compare two workers who started to work in the same year in two different regions and examine if the worker who was hired in a region that implemented a higher subsidy to conversion of fixed-term into open-ended ones was more likely to be observed covered by an open ended contract in 2002 or 2005. In principle, firms could use the subsidy in any year after hiring the worker. We use as our instrument the average of the subsidy during the first two years of tenure at the firm. The decision to use the first two years was guided by the evidence from the 2003-2004 waves of the quarterly Spanish Labor Force Survey (EPA). According to that source, 18% of employed heads of households covered by high firing cost contract were first contracted using a fixed-term contract. Within that fraction, 90% had their contract converted into an open-ended one during the first two years of tenure at the current firm. We also experiment using an average of the subsidy amount during the first three years at the firm, with little impact on the results.[10]

The first stage regression we run is the following linear probability model:

$$\begin{aligned} Open\_ended = {} & \alpha_0 + \alpha_1 Subsidy_{R,g,t_0} + \alpha_2 Subsidy_{R,g,t_0} \cdot 1(Age\_head \geq 35) + \\ & \alpha_3 Subsidy_{R,g,t_0} \cdot Fem\_head + \sum_{g=1}^{g=4} \alpha_{4,g} Age\_head_g + \alpha_5 Fem\_head + \\ & \alpha_6 Hired\_post97_{t_0} + f(Tenure - 3) + X'\alpha_7 + \varepsilon \end{aligned} \tag{C.1}$$

For simplicity of exposition, we remove the subindices i and t denoting households and time, respectively, from all equation variables. The dependent variable indicates whether the worker is observed 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.[11] The key variable is $Subsidy_{R,g,t_0}$, denoting the average maximum statutory amount a firm coul get by converting a fixed-term contract

[10] We also 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.

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

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). 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 younger workers. Thus, we interact the subsidy with a dummy for workers over 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). 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 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.[12] It also includes a dummy for whether or not the contract was signed after 1997. The dummy controls for the fact that subsidies started in almost all 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 Jimeno, Kugler 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 contract conversion. 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. α_1 is an intention-to-treat effect that measures the impact of the statutory amount of the subsidy to contract conversion on the probability that a male head of household worker between 36 and 45 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

[12] 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 analyst 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.

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 contract conversion increased the fraction of workers covered by open-ended contracts, α_1 would be positive.

The parameter α_1 could capture the impact of variables unrelated to the subsidies to contract conversion if employer in different regions exhibited different propensities to use high dismissal cost contracts (for example, due to differences in the industry specialization) or if there were different regional trends driving the decision to hire a worker with a type of contract. We mitigate the second problem by including the unemployment rate in the region in the gender and age-band of the worker at the time of the hiring.

To avoid the concern that α_1 really 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 using an alternative identification strategy that also includes region dummies:

$$\begin{aligned} Open\_ended = {} & \alpha_0 + \alpha_1 Subsidy_{R,g,t_0} + \alpha_2 Subsidy_{R,g,t_0} 1(Age\_head \geq 35) + \\ & \alpha_3 Subsidy_{R,g,t_0} \cdot Fem\_head + \sum_{g=1}^{g=4} \alpha_{4,g} Age\_head_g + \alpha_5 Fem\_head + \\ & \alpha_6 Hired\_post97_{t_0} + f(Tenure - 3) + X' \alpha_7 + \varepsilon \end{aligned} \tag{C.1}$$

In this second case, identification of the parameter α_1 is achieved by comparing the relative chances of having currently covered by an open-ended contract of workers who were hired in the same year by a firm within the same region, but who belong to different demographic groups that were entitled to different levels of the subsidy.

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)].

4.2 Intention-to-treat effects: Do subsidies to contract conversion reduce the amount of household wealth?

Second, we examine intention-to-treat responses of(the logarithm of) household wealth to the presence of regional subsidies when the worker was hired.

This is done to study whether our exogenous variable of subsidies that captures the risk of losing the job by workers initially employed with a fixed-term contract also moves their wealth. In particular, the experiment we think of is the following. Imagine two groups of comparable workers who started working at different firms in the same year. The first group started working in a region that subsidized the conversion of fixed-term contracts into open-ended ones for that worker's demographic group, while the second group was hired in a region that did not introduce such subsidies. Does the group of workers in the "high subsidy" region hold a lower amount of wealth on average? If precautionary motives are indeed operative in the data, the group of workers whose employer had access to a higher subsidy should hold less wealth on average, because a higher fraction of them will have been covered by high firing cost contracts and thus are relatively more protected from the risk of transiting into unemployment.

The exact model we run is the following:

$$\begin{aligned}\log(\frac{W}{Y}) \quad = \quad & \delta_0 + \delta_1 Subsidy_{R,g,t_0} + \delta_2 Subsidy_{R,g,t_0} 1(Age\_head \geq 35) + \\ & \delta_3 Subsidy_{R,g,t_0} \cdot Fem\_head + \sum_{g=1}^{g=4} \delta_{4,g} Age\_head_g + \delta_5 Fem\_head + \\ & \delta_6 Hired\_post97_{t_0} + g_1(Tenure - 3) + X'\delta_7 + u \qquad \text{(C.2)}\end{aligned}$$

Dependent variable: The ratio of household wealth over household earnings. 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: gross financial wealth (checking and saving accounts, mutual funds, stocks and bonds). The second measure adds real estate other than owner-occupied housing, substracting associated debts. Bover (2005) provides some evidence consistent with the notion that Spanish households use real estate wealth as a buffer against certain forms of risk. The third measure adds the net value of owner-occupied housing to the former measure. Business wealth and non-cashable pension funds are excluded from all the definitions. Finally, 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.[13]

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: 128 out of 3,912 households (3.2 percent of the number of original households). We leave a full assessment of working with other transformations of the wealth variable, like the hyperbolic sine function to a future draft.[14] 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 with high dismissal costs 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.

4.3 Assessing the magnitude: how much more wealth do workers covered by low firing cost contracts hold?

As shown later in the empirical results on Section 5, once we have checked that our exogenous variable of regional subsidies is a good instrument, i.e. it is positively correlated with the stock of open-ended contracts and moves downwards the household wealth of those workers entitled to regional subsidies, we estimate the causal impact of the risk of losing the job on the household wealth by the method of instrumental variables. The OLS estimates of equation (W1) would be biased upwards for the various reasons

[13] 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, possibly because the coefficient of household earnings was very close to 1 in such specification.

We plan to construct a measure of permanent household earnings in a future version of this draft.

[14] We have done a limited number of experiments using the hyperbolic sine transformation of the wealth variable (that preserves zeroes and negative values), obtaining qualitatively similar results. Still, a complete assessment of how to handle the skewness of the wealth variable is left to a future draft of the paper.

mentioned earlier. 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:

$$\log(\frac{W}{Y}) = \gamma_0 + \gamma_1 Open\_ended + g_2(Tenure - 3) + X_i'\gamma_2 + v \tag{W1}$$

$$\begin{aligned} Open\_ended = {} & \alpha_0 + \alpha_1 Subsidy_{R,g,t_0} + \alpha_2 Subsidy_{R,g,t_0} 1(Age\_head \geq 35) + \\ & \alpha_3 Subsidy_{R,g,t_0} \cdot Fem\_head + \sum_{g=1}^{g=4} \alpha_{4,g} Age\_head_g + \alpha_5 Fem\_head + \\ & \alpha_6 Hired\_post97_{t_0} + f(Tenure - 3) + X'\alpha_7 + \varepsilon \end{aligned} \tag{C.1}$$

The parameter of interest is γ_1, which measures the response of (the logarithm of) household wealth over household earnings to holding an high dismissal cost contract. The causal estimation of this coefficient only exploits variation in open-ended contracts that is due to the fact that there were different incentives to use those contracts across regions and demographic groups. The error term of equation (W1) 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 average unconditional wealth-income ratio held by households with a fixed-term contract: 0.117.[15]

5 Results

5.1 The quality of the instrument

Table 3 presents OLS regressions of the type of contract held on our key identifying variable: the statutory subsidy amount that the firm could get in

[15] We also experimented evaluating the results taking antilogs in W1, and estimating the amount of precautionary wealth as:

$$\Pr\text{ecaut\_wealth} = \exp(\gamma_0)[1 - \exp(\gamma_1)]$$

This is an approximation that ignores the variance of the residual of the log of wealth. The results were rather similar to the ones we report below.

the first two years of the contract in the region where the household lives. The standard errors are presented in parentheses and take into account that there can be correlation between the error term due to the imputation of a variable ($Subsidy_{R,g,t_0}$) that varies across regions, age and gender. Thus, the standard errors correct for autocorrelation at that level. In Table 3, row 1, column 1, the estimate is 0.0181 (standard error: 0.0045). 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 $Subsidy_{R,g,t_0}$ and a dummy for age below 35 years is -0.0054 (standard error: 0.0078). While not significantly different from zero, the magnitude of the estimate implies that subsidies to conversion had a smaller impact on the probability of observing relatively younger workers covered by an open-ended contract in 2002 or 2005: an increase in the subsidy of 1,000 euro increases the stock of young workers covered by an open-ended contract in 2002-2005 by 1.27%. A smaller impact among younger workers is consistent with the findings of García-Pérez and Rebollo (2009). 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 0.54% (=0.0181-0.0127). 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.[16] 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 sixteen regional dummies, with Madrid as the excluded group. The estimate of the variable $Subsidy_{R,g,t_0}$ is 0.0133 (standard error: 0.0049). That is, workers belonging to demographic groups that were entitled to a subsidy to contract conversion 1,000 euros higher than a benchmark group in the same region are 1.33% 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 7.37, resulting in a weaker instrument than in the previous specification.

[16]We use the definition of household head provided for the EFF by Banco de España (2005). The household 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 case the household head is the partner.

Given the pattern of our results, in specifications (3) and (4) we 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 $Subsidy_{R,g,t_0}$ predicts the share of head male employees covered by an open-ended contract by between 1.67 percentage points (standard error: 0.0051) and 2.08 percentage points (standard error: 0.0048) in a specification that includes and excludes region dummies, respectively.

Overall, we conclude that the instrument "subsidy to contract conversion" works best for the sample of mature male heads (above 35 years of age). The result can be rationalized by economic theory. The decision of an employer to convert a fixed-term contract into an open-ended one may be less responsive to drops in labor costs brought by a subsidy in the case of a young worker than in the case of a mature one. 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 can make the employer more certain about the expected future productivity of the match.

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 ($Subsidy_{R,g,t_0}$) 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 4.5 percentage points (standard error: 0.22). 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 $Subsidy_{R,g,t_0}$ and an indicator of the household head aged below 35 is 0.036 (standard error: 0.028), positive but not very precise. Adding this estimate to that of the variable $Subsidy_{R,g,t_0}$ yields an estimate of 0.009 (=0.045-0.036), 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 2 of Table 4 introduces indicators of the region of residence.[17] The estimate of the variable $Subsidy_{R,g,t_0}$ is -0.0327 (standard error: 0.0239), still negative and consistent with a precautionary saving motive, but not significantly different from zero.

The third column, we present results from our preferred sample, that composed by male heads. The estimate in Table 4, row 1, column 3 of the instrumental variable $Subsidy_{R,g,t_0}$ is -0.065 (standard error: 0.023), 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 wealth to income ratio of 6.5 percentage points. The result is somewhat smaller when we add regional indicators: -0.053 (standard error 0.023), shown in Table 4, column 4, row 1.

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, possibly because the incentive to convert fixed-term into open-ended contracts is a less powerful predictor of the type of contract held for those groups.

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. OLS estimates compare wealth holdings between workers who hold a high dismissal cost contract and those with a low dismissal cost contract. TSLS estimates do the same comparison but focusing on the unobserved group of workers who changed their contract because of the presence of the public subsidy to contract conversion.

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

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 over income ratio is -0.026 (standard error: 0.0945). Multiplying that estimate by 0.117 (the median wealth-income ratio), the estimate suggests that workers covered by fixed-term contracts have wealth holdings about 0.3 percent of their annual earnings higher than comparable workers covered by open-ended contracts (Table 5, Panel B, row 3, column 1). That is a rather small estimate that may be affected by biases due to different previous employment histories. For example, workers whose position is currently covered by a fixed-term contract are relatively more likely to have experienced a recent unemployment spell that affected their wealth holdings than the typical worker covered by an open-ended contract.

The TSLS estimate of the causal impact of holding a fixed-term contract is -2.297 (standard error: 1.342), 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 26.9% of their gross earnings in financial wealth. Our explanation of the stark difference between the OLS estimate of -0.026 and the TSLS estimate of -2.297 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 contract conversions, 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 workers currently covered by "open-ended contracts" are likely to have had employment histories similar to those of workers who are currently covered by fixed-term contracts.

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 "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 exceed by 2.62 those 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 30.7 percent, or 3.7 month's income (a substantial amount). The estimate is not very sensitive to the inclusion to region dummies (presented in column 5, Table 5, Panel B row 3)

The evidence from Table 5 is consistent with the notion that households headed by workers who are exposed to a large chance of losing their job react by accumulating financial wealth. The evidence is stronger 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 about 30.7% of gross household earnings.

5.3 Robustness checks

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 employment 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: -174 (standard error: 1,129). 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. Bover (2005) infers from the age profile of the marginal propensity to consume out of an increase in the value of housing wealth that Spanish households use real estate as a buffer against risk (not necessarily unemployment risk). Thus, we start by experimenting with 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.[18] The estimate of the coefficient "Head covered by high dismissal costs contract" in Table 6, Row 1, Column 2 is -2.24 (standard error: 1.179). The estimate is slightly smaller than that using financial wealth, and is consistent with a "buffer stock saving" of about 26% of gross annual earnings (after multiplying the coefficient by the median wealth of households covered by a fixed-term contract, 0.180).

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 1.88 (standard error: 1.20), positive and not significantly different from zero.[19]

We find two possible explanations for that result. First, 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 home. Such explanation of the findings in Table 5 and 6 does not rely necessarily on a precautionary saving motive. A second explanation is related to demand of credit in the presence of employment risk. (Locally) prudent households will 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.

5.4 Precautionary saving or credit constraints?

A liquidity-constrained individual is willing to borrow at the market interest rate, but is rationed in the sense that either a bank would reject the application for a loan or give an amount lower than that asked (Jappelli, 1990 and many others). 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 by examining how loan rejections vary with the subsidy to contract conversion. We identify three forms of credit constraints using the EFF. The first is whether an individual did not ask for a loan during the last

[18]See Carroll et al. (2003) for a different reasoning. We discuss the issue below.

[19]Another two robustness checks done are to study whether our instrument given by the regional subsidies also moves other undesired outcomes, such as the household head's earnings and the fact that the spouse works or not. For this purpose, we regress both outcomes on the variables related to subsidies in our specifications and we obtain coefficient estimates near zero not being statistically different from zero.

two years because he or she thinks it would be rejected. The second form of credit constraints is whether the interviewed member of the household asked for a loan, but was rejected. The third form of credit constraint is whether the member of the household interviewed was not rejected, but 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 same shown in Tables 3 and 4.

We examine if heads of households who were more likely to get an open-ended contract due to the existence of a higher subsidy to contract conversion were less likely to be affected by liquidity constraints.We use 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 $Subsidy_{R,g,t_0}$ causes a drop in the relative chances of being credit constrainted.

Table 7 shows predicted probabilities of each of the events. 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]

We also include for further reference the results from a model that includes 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: 17.8 percent among all applicants holding fixed-term contracts while only 7.3 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 length of the job contract, is an endogenous variable. For this reason, we estimate Model 2 replacing the indicator of having a permanent contract by other exogenous proxy of the risk of job loss: the regional subsidies to the conversion of fixed-term contracts into permanent ones.

The pattern of coefficients shown 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 12.8 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 risk of losing the job measured by the regional subsidies moves the liquidity constraints. 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 consistent with a precautionary saving motive more than with liquidity contraints (their impossibility of investing in real estate due to the rejection of their loans applied for).

5.5 The response to the risk of losing the job at various points of the wealth distribution

Is the response homogenous over the wealth distribution

We interpret from Table 5 that workers exposed to a higher risk of losing the job keep on average higher (liquid) wealth balances. Now, such average response may reflect a situation in which all households exposed to the risk of losing the job keep uniformly higher balances or, alternatively, a situation in which most households keep small responses but a small fraction keep substantial amounts. In the second situation, precautionary motives would be present for only a minority of households/workers, leading to substantial consumption and welfare losses upon the event of unemployment. 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 and preliminary (the results are still a bit imprecise), the results suggest a non-uniform impact over the wealth distribution, the response of liquid wealth is stronger at the bottom tail as it would be predicted by a buffer-stock model based on a precautionary saving motive. Finally, the results are very imprecise in the 75th quantile.

6 Further evidence from consumption (growth) responses

This section provides evidence of the household consumption response to the risk of losing the job by testing the main hypotheses formulated in Section 2. In this version of the paper, we only give suggestive evidence of the consumption response looking at Ordinary Least Squares estimates that are affected by endogeneity biases as explained later. In the following version of the paper, we will implement the identification strategy explained in Section 4 to obtained causal estimates of the household consumption response.

The theoretical model suggests two hypothesis about the household consumption response to the risk. Firstly, workers more exposed to the risk of losing the job postpone their consumption to build a buffer stock against future unexpected income losses. Secondly, workers more exposed to the risk of unemployment will exhibit higher consumption growth once the uncertainty about the future is solved. Moreover, the model also predicts that the consumption growth of all workers exposed to the risk (independently of whether they will become unemployed or not) will be higher on average [see Equation (2)].

To constrast both hypotheses, we make use of the consumption information collected in the EFF and exploit the panel component of the EFF to estimate the average household consumption growth. We use various measures of consumption. The first is a comprehensive PSID-like question about expenditure on food in a typical week. The second is a comprehensive question based on expenditure on non-durable goods. Finally, we also experiment with a broader definition of consumption that includes non-durable goods and the service flow of selected durables (jewellery, works of art, cars and other means of transport, furniture and housing equipment). The rates of depreciation in Fraumeni (1997), mostly based on the Hulten and Wykoff (1981) rates, are used to derive consumption measures from the household's stock of equipment and vehicles (see Bover, 2005, for a similar strategy).

The key regressor in our estimates that control for the risk of losing the job is the probability that an individual transits from employment to non-employment. This probability is estimated using the 1998-2001 waves of the EPA. The dependent variable takes the value of 1 if the individual is employed in quarter q but not in quarter $q+1$. The independent variables are common across both data sets: occupation, industry, age dummies and whether or not employment in quarter q was covered by a fixed-term contract. We run separate *logit* models for males and females (see Table A.1).

In a second step, we use those predicted probabilities to impute in the EFF the probability that the head of the household (and spouse, if one exists) loses his or her job over the following quarter. We then run regressions of the outcomes of interest on the predicted probability that the head and spouse (if one exists) lose their job.

6.1 Tests based on household consumption

The first outcome of interest is the logarithm of consumption. For the level of consumption, our main specification is:

$$\log C_{it} = \beta_0 + \beta_1 P_{it}(U_h = 1) + \beta_2 P_{it}(U_s = 1) + X_{it}'\gamma + \varepsilon_{it}^c \qquad \text{(C1)}$$

$P_{it}(U_h = 1)$ measures the probability that the head of the household transits into unemployment. $P_{it}(U_s = 1)$ measures the corresponding probability for an employed spouse (if one is present).

X_{it} contains various sets of regressors. First, it includes variables that are associated with transitions into unemployment but that we do not use for the identification of β_1 and β_2. These include dummies with the head and spouse's schooling, industry and occupation dummies.[20] We also include a dummy for spouse not employed, to properly interpret the magniture of $P_{it}(U_s = 1)$. In the estimates, the reference person is a married head of household whose spouse also works. Finally, we include a dummy for the kind of self-employment [an independent professional or self-employed worker (omitted category), an owner of a family business, and a partner in a non-family partnership]. Second, we include variables that pick up life-cycle accumulation of assets due to aging, income and demographic shifters: four dummies in 10 year age bands, three separate intercepts for single, divorced and widow head and female-head, and 5 dummies capturing different household sizes. X_{it} also contains total household income accrued last year. Finally, Equation

[20] See Lusardi (1997), for a detailed analysis of why occupation-specific variance in income does not properly identify the income risk an individual is exposed to.

(C1) is identified by assuming that the variable "type of contract" held by the household head and spouse enters the consumption equation only through its impact on the probability of losing the job.

According to the life-cycle model including the risk of losing a job, β_1 and β_2 should be negative, as explained in Section 2. We experiment with two measures of consumption: total non-durable consumption and a broader measure that includes durables.

A possible source of biases regarding the test in Equation (C1), $\beta_1 < 0$ and $\beta_2 < 0$, is that workers covered by an open-ended contract are more likely to have had continued labor market spells and lifetime income, which we cannot fully control for. The omission of lifetime income creates a negative link between $P_{it}(U_h = 1)$ and ε_{it}^c and between $P_{it}(U_s = 1)$ and ε_{it}^c biasing the OLS estimates of β_1 and β_2 toward a more negative number. In other words, the estimates of the consumption equation (C1) may be biased in favor of the null hypothesis, which is the reason we turn to alternative tests.

Our second test examines if households headed by a worker who has a higher probability of transiting into unemployment in 2002 had higher consumption *growth* between 2002 and 2005. Using the household panel sample, we estimate an equation for the household consumption growth with the following functional form:

$$\begin{aligned}\log C_{i,2005} - \log C_{i,2002} = \alpha_0 + \alpha_1 P_i(U_h = 1|year = 2002) + \\ +\alpha_2 P_i(U_s = 1|year = 2002) + X_i^{\Delta c'}\alpha_3 + \varepsilon_i^{\Delta c} \qquad \text{(DC1)}\end{aligned}$$

Equation (DC1) does not come from transforming consumption equation (C1) into first differences. The variable $P_i(U_h = 1|year = 2002)$ is the probability that the head of household i employed in 2002 loses her or his job next quarter. The same applies to $P_i(U_s = 1|year = 2002)$ when the household head's spouse was employed in 2002. The vector of explanatory variables, $X_i^{\Delta c}$, contains household and personal characteristics in levels and in first-differences, such as an indicator of whether the spouse did not work in 2002; the family head's gender, age band, marital status, economic sector and nature of the business if self-employed; and the education level of the couple. The covariates in first-differences control for a three-year change in the household size and the number of children by age, and the three-year household income growth. Finally, the error term of the equation is denoted by $\varepsilon_i^{\Delta c}$, which may also include measurement errors in the consumption growth.

According to the Euler equation governing the consumption growth in (DC1), households exposed to risk postpone consumption to the future. Thus, individuals who hold low firing cost contracts should experience higher consumption growth over a two year horizon than workers whose job is regulated by a high firing cost contract. Three comments are in order.

First, rather than modelling the variance of the income process, we only include the probability of losing the job, so our test is a very reduced form of the second-order approximation to the Euler equation. Second, we include a set of covariates that do not belong to an Euler equation, like the growth of total household income. The reason for doing so is to avoid biases associated to reversion to the mean: workers covered by fixed-term contracts have lower incomes and may mechanically experience higher income and consumption growth than higher-income workers. Third, note that we do not condition on labor market attachment in 2005. The prediction of higher average consumption growth holds after averaging across all states of the world, including unemployment.

6.2 Empirical results of the consumption responses

6.2.1 Consumption levels

Table A.3 shows the relationship between the probability of losing the job on two measures of consumption. The first is a measure of (recall) non-durable consumption. The second is a broader measure that adds to non-durable consumption an estimate of the flow value of services from car and furniture holdings. The rationale is to allow for adjustments to the risk of losing a member of the couple's job by delaying the purchase of durable goods. We report both the impact of the probability of losing the job on mean consumption (using OLS) and median consumption (using median regressions).

The coefficient of "the probability that the head of the household loses the job over the next *quarter*" is -.004 (standard error: .004), shown in the first column, first row in Table A.3. The negative sign implies that a higher exposure to the risk of losing the job correlates negatively with non-durable consumption. In our sample, the change from the 50th centile to the 90th centile in the probability of transiting into unemployment in the following quarter is about 4 percentage points. Thus, the estimate in row 1 of Table A.3 implies that households would cut non-durable expenses by 1.44 percent as a response to a 4 percent increase in the probability of losing the job. The estimate seems small.

The coefficient measuring the impact on non-durable consumption of the probability that an employed spouse in a married household loses his or her job over the next *quarter* is -0.002 (standard error: 0.003). It is shown in the first column, second row in Table A.3. The specification contains controls for a dummy that takes the value of 1 if the secondary earner does not work. The estimate is positive, contrary to the precautionary savings hypothesis. Neither estimate of the impact of the risk of job loss is very precise.

In column (2), row 1 of Table A.3, we turn to the impact of the probability that the head transits into unemployment on *total* consumption. The coefficient is now -0.010 (standard error: 0.003), significantly different from zero at the 1 percent confidence level. The magnitude suggests that households react to the risk that the household head transits into non-employment by either cutting or delaying durable expenses, like cars or housing equipment. We quantify the magnitude of the estimate as in the case with non-durables: an increase in the quarterly probability that the head loses the job of 4 percentage points per quarter (basically, from the 50th to the 90th centile of the distribution of the probability of entering an unemployment spell in the next quarter) leads to a drop in durable consumption of 4 percent. The magnitudes of estimates of the impact of unemployment risk on median consumption are similar to mean impacts, and we do not comment them in detail.

Overall, the evidence in Table A.3 is consistent with the notion that households respond to the risk that the head loses his or her job by cutting mainly durable expenses. The response for the risk that the spouse loses her job (when a spouse is present and works) is somewhat smaller and also confined to durable goods. As we mention above, the potential biases in the previous specifications go in favor of finding evidence supporting precautionary savings, which is the reason we now turn to examine consumption growth and balance sheet responses.

6.2.2 Consumption growth

Table A.4 presents estimates of the impact of exposure to the risk of losing the job on various measures of consumption growth. The results in column 1 suggest that a 1 percent increase in the chance of losing the job of the head over the next quarter led households to increase food consumption growth by 3.3 percentage points between 2002 and 2005. Taking the 4 percent difference between open-ended and fixed-term contracts, one obtains a 13.2% relative increase in consumption growth, but the estimate is very imprecise.

Now, the estimates are much more reliable when we examine total non-durable consumption and total consumption. The estimate in row 1 and column 2 of Table A.4 implies that a shift of 4 percentage points in the exposure to lose the job leads to an increase in non-durable consumption of 13.2 percentage points. The relative increase in the growth of our broadest measure of consumption (including the flow of services from cars and housing equipment) following a 4 percent increase in the probability that the head loses the job is smaller, around 9%. Again, the evidence in Table A.4 is consistent with the idea that households exposed to the risk of losing the job delay mostly non-durable and durable consumption. The evidence for changes in food consumption is much less clear-cut. We find little evidence for responses of household consumption growth to the spouse's risk of losing the job.

7 Conclusions

This draft has used the large dispersion in dismissal costs in the Spanish labor market 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 converted low dismissal cost contracts into high dismissal cost ones. We obtain that workers who exogenously obtain a high dismissal cost contract accumulate less financial wealth than comparable workers covered by low dismissal cost contracts. The size of the buffer stock estimated amounts to around 20-30% of gross yearly earnings. Nevertheless, 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. Our results are consistent with the notion that households headed by a worker exposed to the risk of losing the job accumulate higher balances of liquid wealth than comparable households covered by "safer" contracts.

A number of issues is still pending. First, we need to develop a theoretical framework to properly interpret the magnitudes we estimate. We plan to examine those issues in the next draft.

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Table 1.A.: 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 1.B.: 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:		
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 2: Summary statistics, combined EFF2002 and EFF2005

	Total sample	Open-ended	Fixed-term
Head with open-ended contract	.805	--	--
	(.396)		
Head with fixed-term contract	.194	--	--
	(.370)		
Age of household head	43.412	44.308	39.903
S.D.	(9.473)	(9.563)	(9.537)
Married	.807	.813	.718
	(.394)	(.389)	(.477)
Household size	3.217	3.244	3.107
	(1.238)	(1.21)	(1.347)
Prob. job loss (quarter),head			
Mean:	.0299	.0065	.0859
S.D.	(.0342)	(.0036)	(.0407)
# Years at current job	--	14.16	3.529
		(10.42)	(5.187)
Household earnings	27.212	29.561	17.485
	(19.112)	(19.754)	(12.0277)
Head eligible for subsidy	.271	.142	.431
Amount eligible	1.003	.888	1.501
Non-durable expenditure	12.565	13.298	10.228
S.D.	(7.344)	(7.833)	(5.253)
Owns real estate	.832	.865	.693
Net worth			
Median	120.044	133.138	66.034
Mean	167.622	186.217	90.623
Net worth to earnings ratio			
Median	4.909	5.157	3.476
Mean	7.322	7.129	8.119
Financial wealth			
25th centile	966	1.598	.480
Median	3.354	4.120	1.598
Mean	16.270	19.339	6.008
Financial wealth to earnings ratio			
Median	.155	.17	.117
Mean	.563	.581	.465

Sample: 3,583 household-years in two EFF waves (2002 and 2005).
All statistics weighted. S.D. are standard deviations. Monetary magnitudes in 1000s of 2005 euro. Subsidy amounts deflacted using regional industrial price index.

Table 3: First stage: the impact of subsidies on contract conversion.

Dependent variable is binary (1 if head has an open-ended contract)
Estimation method: OLS (Linear probability model)

	Full sample		Sample of male heads	
	(1)	(2)	(3)	(4)
1. Mean subsidy to conversion in two first years of tenure	.0181 (.0045)**	.0133 (.0049)**	.0208 (.0048)***	.0167 (.0051)***
2. Subsidy to conversion *(Age< 35)	-.0054 (.0078)	-.013 (.0059)	-.024 (.0064)	-.0247 (.006)**
3. Subsidy * Female head	-.0127 (.0064)	-.0045 (.008)	--	--
Head is a female	-.003 (.023)		--	--
Head aged 18-25	-.0215 (.0616)		.092 (.071)	.093 (.071)
Head aged 26-35	-.0012 (.0216)		.0146 (.023)	.017 (.0226)
Head aged 46-55	-.030 (.014)		-.021 (.016)	-.025 (.0163)
Head aged 56-65	-.0041 (.0188)		.010 (.020)	.0063 (.020)
Household size 1	.0215 (.0241)		-.012 (.037)	-.014 (.037)
Household size 3	.0048 (.015)		.026 (.016)	.0252 (.016)
Household size 4	.0066 (.0159)		.0212 (.017)	.0236 (.017)
Household size 5	-.0014 (.0205)		.0146 (.022)	.018 (.022)
Household size 6+	-.0088 (.030)		.007 (.030)	.0163 (.0305)
Contract started after 1997	.0787 (.0308)	.0923 (.028)	.068 (.031)	.0798 (.030)
Unemployment rate in region (year entered current firm)	-.0011 (.0005)			
Head entered labor market after 1984	-.0526 (.0158)		-.036 (.017)	-.035 (.0176)
Logarithm of earnings (head and spouse, if present)	.0964 (.010)		.077 (.011)	.0717 (.011)
Spouse works	-.0342 (.0126)			
Secondary school, head	.0484 (.021)		.048 (.022)	.0498 (.0219)
Upper secondary school, head	.079 (.023)		.0772 (.024)	.0735 (.0239)
College, head	.049 (.022)		.0438 (.023)	.044 (.023)
Works for the public sector	-.012 (.011)		-.029 (.0118)	-.025 (.0123)
Industry, head	-.0088 (.0135)		-.0228 (.0134)	-.030 (.013)

Table 3: First stage: the impact of regional subsidies on contract conversion (Contd.).

Dependent variable is binary (1 if head has an open-ended contract)
Estimation method: OLS

Instrument:	Amount head was eligible during the first two years of contract (average over the two years)			
	(1)	(2)	(3)	(4)
Agriculture, head	-.112		-.144	-.134
	(.031)		(.0287)	(.028)
Construction, head	-.121		-.140	-.138
	(.020)		(.0218)	(.0218)
Year 2003	.0179		.0146	.0065
	(.0162)		(.0155)	(.0163)
Year 2005	-.021		-.0159	-.018
	(.0177)		(.0175)	(.018)
Year 2006	.0113		-.0003	-.0038
	(.0147)		(.0149)	(.015)
Tenure on the job-4, head	.064		.055	.055
	(.0049)		(.0044)	(.0043)
Tenure on the job squared, head	-.0033		-.003	-.0031
	(.00026)		(.00025)	(.0002)
Tenure on the job cubed, head	5.00E-05		4.80E-05	3.00E-05
	(4.05e-06)		(4.31e-6)	(2.4e-06)
Constant	.593		0.588	.614
	(.0374)		(.038)	(.040)
Region controls?	No	Yes	No	Yes
R.squared	.332			

Notes: 3853 cases, sample of households headed by an employee between 20 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 4: The impact of subsidies to contract conversion on household financial wealth

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

	(1)	(2)	(3)	(4)
1. Mean subsidy to conversion in	-.045	-.0327	-.065	-.053
two first years of tenure	(.022)**	(.0239)	(.0229)***	(.0252)**
2. Subsidy to conversion	.0360	.0423	.0544	.0634
*(Age< 35)	(.028)	(.0287)	(.030)	(.030)**
3. Subsidy * Female head	.0426	.0439	--	--
	(.035)	(.035)		
Head is a female	-.378	-.399	--	--
	(.137)	(.137)		
Head aged 18-25	-.559	-.543	-.567	-.610
	(.223)	(.222)	(.275)	(.269)
Head aged 26-35	-.197	-.204	-0.191	-.22
	(.105)	(.106)	(.112)	(.113)
Head aged 46-55	.346	.335	.255	.241
	(.089)	(.088)	(.098)	(.098)
Head aged 56-65	.816	.809	.720	.706
	(.115)	(.115)	(.129)	(.129)
Household size 1	-.760	-.755	-.858	-.837
	(.132)	(.13)	(.205)	(.205)
Household size 3	.214	.203	.144	.135
	(.088)	(.087)	(.095)	(.095)
Household size 4	.124	.111	.069	.0486
	(.0915)	(.090)	(.097)	(.096)
Household size 5	.290	.283	.231	.213
	(.127)	(.127)	(.132)	(.131)
Household size 6+	-.020	.012	-.067	-.039
	(.180)	(.127)	(.187)	(.187)
Contract started after 1997	-.0057	-.029	.125	.116
	(.148)	(.151)	(.165)	(.162)
Head entered labor market	.1248	.114	.002	-.0003
after 1984	(.090)	(.090)	(.101)	(.10)
Logarithm of earnings	.078	.077	.147	.147
(head and spouse, if present)	(.069)	(.071)	(.078)	(.08)
Spouse works	-.228	-.228	-.245	-.247
	(.076)	(.075)	(.077)	(.0767)
Secondary school, head	.020	-.0025	-.016	-.037
	(.099)	(.100)	(.110)	(.11)
Upper secondary school, head	.209	.173	.211	.175
	(.117)	(.117)	(.127)	(.127)
College, head	.797	.774	.820	.803
	(.122)	(.121)	(.138)	(.137)
Works for the public sector	-.140	-.123	-.145	-.124
	(.083)	(.082)	(.093)	(.091)
Head works in manufacturing sector.	.111	.032	.128	.0586
	(.082)	(.083)	(.086)	(.0876)

Table 4: The impact of subsidies to contract conversion on household financial wealth

Dependent variable is binary (1 if head has an open-ended contract)
Estimation method: OLS
Instrument: Amount head was eligible during the first two years of contract (average over the two years)

	(1)	(2)	(3)	(4)
Head works in agriculture	.085	.067	.150	.144
	(.128)	(.133)	(.134)	(.137)
Head works in construction sector	-.185	-.207	-.213	-.226
	(.098)	(.096)	(.101)	(.099)
Year 2003	.0249	.076	.0038	.037
	(.084)	(.086)	(.090)	(.093)
Year 2005	.345	.305	.104	.314
	(.097)	(.100)	(.360)	(.108)
Year 2006	.0776	.133	.040	.083
	(.0801)	(.080)	(.086)	(.085)
Tenure on the job-3, head	.0212	.022	.0316	.035
	(.018)	(.0186)	(.021)	(.021)
Tenure on the job squared, head	-.00013	-.00038	-.0011	-.0014
	(.001)	(.0011)	(.0012)	(.0013)
Tenure on the job cubed, head	-6.36e-06	-2.19e-06	1.50E-05	2e-5
	(2.2e-5)	(2e-5)	(2e-5)	(2e-5)
Constant	-2.631	-2.751	-2.51	-2.63
	(.187)	(.196)	(.203)	(.215)
Region controls?	No	Yes	No	Yes
Sample size	3739	3739	3207	3207
R-squared	.158	.177	.164	.182

Notes: 3853 cases, 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)
Panel A Dependent variable takes value 1 if the household head has an open-ended contract (first stage).					
1. Subsidy to conversion amount head was eligible for	--	.0176	.0127	.0202	.0177
		(.0040)***	(.0044)**	(.0042)***	(.0047)***
2. Subsidy to conversion * (Age <=35)	--	-.0133	-.0129	-0.025	-.0225
		(.0052)*	(.0052)**	(.0058)**	(.0091)***
3. Subsidy to converison * (Head is female)	--	-.0048	-.004	--	--
		(.0059)	(.0058)		
Panel B Dependent variable is the logarithm of financial wealth over earnings of head and spouse					
1. Head covered by high firing cost contract	-.026	-2.297	-2.61	-2.62	-2.62
	(.095)	(1.342)*	(1.975)	(1.164)**	(1.37)*
2. Constant	-2.592	-1.249	-1.133	-.948	-1.065
	(.2074)	(.813)*	(1.252)	(.702)	(.865)
3. Fraction of gross earnings held as financial wealth (at the median)	**0.003**	**0.269**	**0.305**	**0.307**	**0.307**
Region dummies?	No	No	Yes	No	Yes
Sample sizes:	3739	3739	3739	3207	3207

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 high dismissal cost	-.174 (1.279)	-2.24 (1.179)*	1.88 (1.20)
Head aged 18-25	-.363 (.269)	-0.33 (.364)	-1.346 (.4272)
Head aged 26-35	-.060 (.100)	-.0379 (.123)	-.187 (.120)
Head aged 46-55	.159 (.099)	.244 (.120)	.211 (.091)
Head aged 56-65	.548 (.131)	.783 (.157)	.456 (.115)
Contract started after 1997	.023 (.202)		-.184 (.185)
Unemployment rate in region (year entered current firm)	-.0107 (.0042)		-.0007 (.0037)
Head entered labor market after 1984	-.0259 (.1059)		-.0207 (.099)
Logarithm of earnings	.158 (.126)		-.185 (.143)
Logarithm of earnings, squared		--	.214 (.055)
Logarithm of earnings, cubed		--	-.0715 (.0205)
3. Fraction of gross earnings as financial wealth (at the median)	**0.020**	**0.262**	**--**

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 as a regressor						
1. Fixed-term contract	0.273	0.042	0.003	0.014	0.059	0.178
2. Open-ended contract	0.279	0.014 (***)	0.002	0.006 (**)	0.022	0.073
Model 2: Subsidy to contract conversion as a regressor						
1. Zero subsidies	0.273	0.029	0.002	0.009	0.040	0.128
2. 1,000-euro subsidies	0.262	0.030 (*)	0.003 (***)	0.008 (**)	0.041	0.135

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	-2.1	-1.2	-3.1
95% confidence interval	[-5.8, -0.9]	[-4.9, .4]	[-10, 2.5]
2. Constant	-2.244	-1.684	0.905
3. Fraction of gross yearly income held as wealth	**0.093**	**0.130**	**--**

Table A.1: Determinants of the transition from employment to unemployment (EPA)

Dependent variable takes value 1 if there is a transition from employment to unemployment
Estimation method: Logit

	(1)	(2)
Sample:	Males	Females
Employee with open-ended contract	-0.937	-0.880
	(0.017)	(0.018)
Open-ended contract after 1997	0.285	0.190
	(0.023)	(0.024)
Employee	0.922	0.836
	(0.022)	(0.032)
Public sector	0.148	0.086
	(0.027)	(0.021)
Public sector * Open-ended contract	-0.358	-0.286
	(0.041)	(0.033)
Constant	-2.408	-2.002
	(0.038)	(0.037)
Sample size:	326,648	176,633

Notes: Other 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.2: Subsidies for conversion of temporary contracts into permanent ones, by region and year

Region / Year	1997	1998	1999	2000	2001
1. Andalucia			All years, 1,800 euro if age < 30		
2. Aragon			All years, 1,200 euro for females		
3. Asturias	2,100 euro	2,100 euro, all workers	2,100 euro, all workers	2,100 euro, all workers	None
		2,400 if "learning contract"	2,400 euro if "learning contract"	2,400 if "learning contract"	
		600 extra if female in male job	600 extra if female in male job	plus 600 if female in male job	
4. Baleares			None		
5. Canarias	None	3,600 if age<25 or if female	None	None	None
6. Cantabria	None	1,800	None	None	None
		2,400 if age<30 or female			
		3,600 if above 40			
7. Castilla-Leon	None	1,800 euro	1,800 euro	1,803 if age<30	1,803 if age<30
		2,400 if apprenticeship contract	2,400 if apprenticeship contract		2,040 if female
8. Castilla-La Mancha			None		
9. Catalonia			None		
10. Valencia	None	None	30% of payroll tax	30% of payroll tax	1400, practice contr.
					1,800 if "practice c."
					and female
11. Extremadura	4908	3545	3618	2100 if training	2101 if "practice c."
12. Galicia	None	3000 euro if age<30	None	None	None
		4200 if female in male job	None	None	None
13. Madrid			None		
14. Murcia	1800	2100 if age<=30	2100 if age<=30	2100 if age<=30	2100 if age<=30
	2400 if age<30	1500 if age>30	1800 if age>30	1800 if age>30	1800 if age>30
15. Navarra	None	1800	None	Payroll subsidy depending on age	
16. Basque country	None	3000 for age<40	3000 for age<40	Both years: Former+ 6009 euro if age<30	
	None	150 extra if female	150 extra if female	Former+ 4507 euro if age<30 & female	
17. Rioja	None	Depends on # conversions	Depends on # conversions	Depends on # conversions	

1. "Apprenticeship contract" (contrato de aprendizaje): contract typically offered to low-skilled young workers
2. "Learning contract" (contrato de formación): contract typically used for workers between 16 and 18 years of age.
3. "Practice contract" (contrato en prácticas) Contract typically used for qualified young workers without labor market experience

Table A.3: Consumption responses to the risk of losing the job

Dependent variable:	Non-durable consumption (log)	Total consumption (log)	Non-durable consumption (log)	Total consumption (log)
Estimation method:	OLS		QR	
	(1)	(2)	(3)	(4)
1. (Prob job loss, head - .012)	**-.004**	**-.01**	**-.003**	**-.008**
*100	**(.004)**	**(.003)*****	**(.005)**	**(.004)****
2. (Prob job loss, spouse - .032)	**.002**	**-.002**	**.006**	**0.000**
*100	**(.003)**	**(.003)**	**(.004)**	**(.004)**
Spouse does not work	0.052	0.043	0.062	0.042
Constant	2.132	2.476	2.175	2.507
	(0.036)	(0.032)	(0.044)	(0.043)

Notes: Sample size: 5294. Standard errors corrected for heteroscedasticity.
Additional covariates not shown: household income, family head's age, gender, education, economic sector, number of years contributed to Social Security, indicators of whether the family head works self-employed, nature of the business if self-employed and marital status; and the following covariates referred to the family head's spouse: education, economic, number of years contributed to Social Security and the indicator of whether she or he worked continuously last year.

Table A.4.: 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	OLS	OLS
	(1)	(2)	(4)
1. (Prob job loss, head -.012)	**.033**	**.033**	**.023**
***100**	**(.019)**	**(.013)*****	**(.011)****
2. (Prob job loss, spouse -.032)	**.006**	**.008**	**.01**
***100**	**(.012)**	**(.007)**	**(.006)**
Spouse does not work	-0.047	0.062	0.077
	(0.062)	(0.052)	(0.043)
Constant	0.078	0.023	0.028
	(0.088)	(0.069)	(0.059)

Notes: Sample size: 976. Standard errors are in parentheses.
Other covariates included but not shown: family head's age bands of 20-25, 26-35, 46-55 and 56-65, and three-year changes: in logarithm of household wealth, household size and number of children by age groups.