Employment Protection and Consumption: Evidence from Italy

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Abstract

Leveraging an Italian labour market reform known as the *Jobs Act* (JA), I study the effect of employment risk on household consumption and labour supply. The JA reduced protection against unlawful individual termination only for workers hired after March 6, 2015, by firms with at least 15 employees. Using this time-based discontinuity in employment protection as a source of exogenous variation in employment risk, I find that workers subject to the reform consume 8% less than workers hired before March 6, 2015. The effect is stronger among individuals younger than 40 and for those living in Northern Italy, the wealthiest region of the country. There is, on the other hand, no sizable effect on labour supply. Finally, I show that a variant of the Bewley–Huggett–Aiyagari model, augmented with ex-ante employment risk heterogeneity, qualitatively matches the empirical result, and that risk accounts for a sizable part of the effect.

Keywords: Jobs Act, Precautionary Saving, Employment Risk, Employment Protection
JEL Classification: D15, E21, J28, G51

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1 Introduction

People deeply value the security of their jobs (Baghai et al. 2023). Rothwell and Crabtree (2019) surveyed a representative sample of U.S. employees, asking them to rate the importance of job security for their quality of life on a scale from 1 to 5. Over 90% assigned a score of 4 or 5, indicating strong preferences for stable employment. More recently, BCG (2023) surveyed 11,000 employees across both advanced and developing countries and found that job stability is prioritized even above pay levels.

Following the 2008 Financial Crisis, many advanced economies - especially in Southern Europe - implemented labour market reforms aimed at increasing employment and enhancing firm productivity by reducing firing costs for firms (OECD 2016). From the workers' perspective, however, these reforms often meant a reduction in employment protection and, consequently, an increase in employment risk.

In this paper, I examine how changes in employment risk affect household consumption and labour supply. To this purpose, I leverage the *Jobs Act* (JA), an Italian labour market reform that applies only to workers hired after March 6, 2015, by firms with at least 15 employees. Workers hired before that date remain under the previous regime, called the *Chart of Labour* (COL). Under the COL, employees who are unlawfully terminated are reinstated to their previous job with the same wage. In contrast, JA workers are not reinstated in case of unjust termination; instead, they receive a lump-sum monetary payment proportional to their tenure. The reform therefore significantly reduced job security for the latter because it almost zeroed the probability of reinstatement in case of unlawful dismissal.

I use data from the 2014 and 2018 waves of the *Participation, Labour, Unemployment Survey* (PLUS), conducted by the Italian Research Agency INAPP (*Istituto Nazionale per le Analisi delle Politiche Pubbliche*). PLUS is a representative survey of the Italian population aged 18 to 75 that provides detailed information on demographic characteristics, food consumption, weekly hours worked, employees' firm size, and whether individuals were hired before or after March 6, 2015. This last piece of information allows me to identify private sector workers in large firms whose contracts are regulated either by the *Jobs Act* or by the *Chart of Labour*.

My baseline finding is that JA workers reduce their food consumption to the household income rate (the average propensity of food consumption) by 1.5 percentage points. To estimate the impact of the *Jobs Act* on broader measures of consumption, I propose two strategies. First, I adopt an approach similar to Broda and Parker (2014), in which I scale the effect on the average propensity to food consumption by multiplying it by the ratio of total spending to food spending, as reported in the National Accounts and the *Household Budget Survey* (HBS) - the Italian counterpart of the CEX in the United States - containing in depth information on how household spending. This approach yields an effect on the overall saving rate of around 10 percentage points. Alternatively, following the approach of Blundell et al. (2008) and Patterson (2023) - BPP from now on - I impute consumption from HBS into PLUS, finding a causal effect on this imputed measure of consumption of 8%. The magnitude of the effect is in line with Clark et al. (2022) and is qualitatively consistent with the effect of aggregate uncertainty on consumption as recently estimated by Coibion et al. (2024).

To strengthen the credibility of the causal claim, I conduct the following placebo tests. First, I show that the incomes of JA and COL workers are not statistically different. Second, I run the same analysis as before for workers employed in small firms, where employment protection remained unchanged, and find no significant effects. Third, I examine public sector employees, who are arguably less exposed to employment risk (Fuchs-Schündeln and Schündeln 2005), and for whom employment protection against termination does not apply. Similarly, I find no significant results. Additionally, I check that hiring rates before and after the reform exhibit a smooth trend, suggesting no manipulation of hiring dates by either workers or firms and no anticipatory effects of the time-based discontinuity in employment protection. Finally, I exclude employees working for firms with fewer than 20 employees. The rationale behind this test is that the reduction in employment protection was coupled by a hiring subsidy, of which large firms were the first beneficiaries (Sestito and Viviano 2018 and Boeri and Garibaldi 2019). Marginal firms, such as those with 14 employees, might have decided to hire an additional worker - and thus reach the threshold of 15 employees - to exploit the hiring subsidy, whereas firms already well above the threshold did not face such incentives. Results remain significant economically and statistically, indicating that potential firms' dimensional sorting does not compromise the identification strategy.

The results described above mask noticeable heterogeneity across several dimensions. The reduction in consumption is more pronounced among younger workers, consistent with the predictions of Gourinchas and Parker (2002). Their work demonstrates that consumers transition from buffer-stock saving (Deaton 1991; Carroll 1997; Carroll and Samwick 1998) to a certainty-equivalent model (Hall 1978) around the age of 40, as the precautionary saving motive weakens with proximity to retirement. In addition, the effect is significant only for people living in the North, the richest and most developed region of Italy. Finally, the reduction in consumption is stronger for higher-educated and higher-income individuals.

The reason why results are only significant in the North is that people in the South and in the Center are more likely to be hand-to-mouth and therefore off the Euler equation (Christelis et al. 2020). For this reason, prudence does not play a major role in their consumption decisions, which are insensitive to changes in employment risk. On the other hand, results are stronger for higher-income individuals because of the negative association between average income and income risk: better-paying jobs also tend to be less risky (Savoia 2023). Therefore, the reduction in employment protection is, at the margin, more relevant for high-income workers, who have much more to lose from the *Jobs Act*.

On theoretical grounds, higher employment risk impacts consumption via two channels. On the one hand, higher employment risk implies a more volatile income and a stronger precautionary saving motive. This is the channel whose importance I aim to measure in this paper. On the other hand, a higher job separation rate reduces the present discounted value of income (PDV), which in turn reduces consumption. In other words, the *Jobs Act* does not induce a mean-preserving spread in individuals' income distributions. Unfortunately, given that the

PDV is unobservable in my data, it is impossible to disentangle these two mechanisms empirically.

I therefore rely on a variant of the Bewley–Huggett–Aiyagari model (Bewley 1986, Huggett 1993, Aiyagari 1994) - from now on BHA - in which there are two types of workers that are ex-ante heterogeneous in employment risk: safe and risky workers. The latter mimic JA workers and face a higher separation rate (i.e., the probability of losing their job while employed) than the former, who instead mimic COL workers. I calibrate the model to match Italy's job market flows and household wealth between 2016 and 2019 using the Method of Simulated Moments (MSM). After solving the model, I use the policy functions to simulate a panel of consumption, income, and wealth. I then examine the difference in consumption between risky and safe employees, both unconditionally and conditional on permanent income. The unconditional model-predicted average consumption difference between JA and COL workers is 5.4%, reasonably close to the empirical results outlined above. I then run a regression on the simulated data of log consumption against the *Jobs Act* dummy, assets, and lifetime income, finding that the effect of employment risk alone is around 2.2%, thus explaining approximately 40% of the total effect of the JA on consumption.

Finally, I consider the following two extensions of the baseline model. First, I assume recursive preferences à la Epstein and Zin (1989) instead of constant relative risk aversion (CRRA) and find that results are very close to the baseline model. Second, I allow households to save in two types of accounts: liquid and illiquid, with the latter having transaction costs and an infrequent possibility of adjustment as in Auclert et al. (2024). In line with Graves (2025), I find that JAs workers increase their savings more in the liquid account than in the illiquid account. The intuition of this last result is the same as in the seminal paper of Kaplan and Violante (2014). Consumers prefer to save for precautionary reasons, mostly in low transaction costs, ready-to-liquidate assets, as they can easily access them in case of need.

Related Literature and Contribution. This paper is connected to the literature on income risk and precautionary saving[1]. Empirically assessing the relevance of precautionary saving has proven challenging for two reasons. First, consumption growth volatility, the term governing prudence in the Euler equation (see Fagereng et al. 2017), is unobservable. Therefore, income risk is often used as a proxy, assuming that income is the main source of risk individuals face[2]. In turn, Guiso et al. (2002) argue that income risk depends on the distribution of future earnings, the distribution of unemployment benefits, and the probability of job loss (i.e., employment risk). While previous studies often bundle employment and income risk together

[1] Blanchard and Mankiw (1988), Kimball (1990), Caballero (1991), Deaton (1991), Carroll (1997) among the seminal contributions.

[2] See, for instance, Jappelli and Pistaferri (2000). More recently, Guiso and Jappelli (2024a), using the novel *Italian Survey of Consumer Expectations* (ISCE, Guiso and Jappelli 2024b), relate subjective consumption growth risk to earnings, health, energy, and house price risk, as well as aggregate sources of risk such as GDP, interest rate, and inflation risk. Idiosyncratic factors explain about 75% of the variability in consumption risk. In earlier work, Dynan (1993) replaces volatility with realized squared consumption growth, instrumenting it with lagged demographic characteristics. However, these are predictors of the consumption path, and therefore violate the exclusion restriction.

(see Fuchs-Schündeln and Schündeln 2005), this paper specifically differentiates between employment risk (i.e., being under the JA or the COL) and the remaining components of income risk, which mostly depend on idiosyncratic productivity.

The reason it is important to specifically examine employment risk, as opposed to income risk more generally, is the need to separate the effect of exogenous shocks experienced by workers (such as involuntary unemployment) from the endogenous responses to them (such as decisions to stay in or exit the labor market). In this regard, Low et al. (2010), using a life-cycle model with endogenous employment participation and search frictions in the labor market, show that failing to account for job mobility (i.e., the endogenous component of income volatility) leads to an overestimation of income risk.

The second challenge is, however, that employment risk is endogenous, because workers self-select into riskier and safer jobs according to unobservable characteristics, of which preference heterogeneity is the most prominent example (see Jappelli and Pistaferri 2020).

Four approaches have been followed in the Literature to estimate the strength of precautionary saving. The first consists in empirically testing the Euler Equation using survey data to estimate the elasticity of intertemporal substitution (*EIS*) and/or the coefficient of relative prudence[3]. Among the papers following this approach are: Guiso and Jappelli (2024a), Bertola et al. (2005) in Italy, and Christelis et al. (2020) in the Netherlands, Sciacchetano (2024) using data on Germany, France, Italy, Spain, the Netherlands and Belgium and Crump et al. (2022) in the United States. They all find a CRRA coefficient of around 1 and 2, corresponding to a prudence parameter between 2 and 3. Using mortgage rates notches as a quasi-experimental variation in interest rates, Best et al. (2019) estimate an EIS of 0.1 in the United Kingdom.

The second approach is structural estimation. Gourinchas and Parker (2002) simulate a life cycle model assuming CRRA preferences and idiosyncratic income risk, and estimate the prudence coefficient (jointly with the discount rate) by minimizing the distance between the average of the simulated and actual life cycle consumption profiles. Their main result is a risk aversion of 1, which corresponds to a Kimball coefficient of 2. Cagetti (2003), matching median consumption in lieu of the mean, finds a risk aversion coefficient around 4 (and a Kimball coefficient of 5).

The third approach directly leverages survey questions to elicit motives for saving. Jappelli and Pistaferri (2025) show that households revise their target wealth in response to permanent income shocks approximately one-to-one. They furthermore show that convergence to target wealth is faster if households are below the target than when they are above. Finally, they simulate a buffer-stock model with different calibration choices, finding that their results can be replicated using reasonable parameters for preferences and the income process. They elicit target wealth from a direct question in the *Survey of Household Income and Wealth* of the Bank of Italy, asking how much people think their savings need to be to meet unexpected events.

A similar strategy is also followed by Guiso et al. (1992), who use self-reported measures of income uncertainty to show that subjective income risk is negatively associated with con-

[3]Assuming CRRA utility, as in the papers cited here, it can be shown that $EIS = 1/\gamma$ and $p = 1 + \gamma$, where γ is the utility function curvature and p is Kimball's coefficient of relative prudence.

sumption. Graves (2025) finds that higher subjective unemployment risk induces an increase in liquid wealth, but a reduction in illiquid wealth. The intuition is that illiquid assets are not well-suited for smoothing income fluctuations, as they are riskier and characterized by transaction costs (Kaplan and Violante 2014, Auclert et al. 2024). He also shows, with the help of a two-assets HANK model à la Kaplan et al. (2018), that this "flight to liquidity" amplifies business cycle fluctuations.

My paper differs in that it leverages a natural experiment rather than subjective expectations of job loss, which might be endogenous to unobservable, potentially time-varying characteristics such as preferences, health status, and ability. For instance, a worker might expect to be fired soon because her productivity has persistently declined in the last few months.

The last approach consists of estimating the causal effect on savings of exogenous shocks to income risk. Kantor and Fishback (1996) exploit the introduction of workers' compensation in the United States in the 1910s, finding that the introduction of accident insurance decreased households' savings by 25%[4]. Leveraging on the historical event of the German reunification in 1990, Fuchs-Schündeln and Schündeln (2005) find that civil servants in Eastern Germany (i.e., former DDR government employees) tend to accumulate 25% less wealth than private workers, controlling for permanent income and demographics. They also demonstrate that not accounting for workers' self-selection underestimates the strength of precautionary saving by half.

Clark et al. (2022) look at the impact of the 2012 Italian labour market reform on consumption and savings, finding a drop in the former and a rise in the latter, both around 9%. My paper differs along several dimensions. First, I rely on a different source of exogenous variation. The 2012 reform lowered employment protection for all large firm employees. So, they compare large firm workers (exposed to the reform) with small firm workers (their controls) before and after the reform (D-i-D). On the contrary, my analysis only involves large firm workers: I compare large firm workers who have been hired starting from March 7, 2015, with large firm workers who have been hired before March 7. This allows me to address potential workers' selection into small or large firms driven by unobservable, potentially time-varying, risk-aversion heterogeneity. In addition, as Figure 2 shows, the 2015 reform led to a much larger reduction in employees' protection than the 2012 one[5]. Finally, my analysis, with the help of the model, digs deeper into the mechanisms of the effect of employment protection, distinguishing between the sole effect of risk and the permanent income effect.

To the best of my knowledge, this is the first paper to leverage a time-based discontinuity in employment protection to empirically assess the impact of employment risk on consumption. Furthermore, my paper connects the approaches outlined above from an operational point of view: I rely on a structural model to replicate the causal effect of employment risk on consumption, estimating the parameters that regulate job market flows internally to match the

[4]There is another strand of the literature that, instead of assuming incomplete markets, tries to microfound them in terms of a principal-agent problem where there is "hidden trade" from the agent side and therefore moral hazard arises. Therefore, the introduction of insurance might be seen as a shock that "alleviates" moral hazard problems, increases insurance opportunities, and reduces the gap with the complete market benchmark.

[5]The reason why this is the case is offered in Appendix E.

dynamics of the labor market in Italy in the period from 2016 to 2019.

Finally, I contribute to the existing empirical literature leveraging the *Jobs Act* as a natural experiment. De Paola et al. (2021, 2024) show that the reform caused a reduction in women's fertility, both directly (the former paper) and indirectly via workplace peer effects (the latter paper). Bertoni et al. (2023) find that workers hired with the JA perceive their jobs as more precarious and less satisfactory. This prompts workers to engage in on-the-job searching more actively, resulting in higher job mobility rates. Mistrulli et al. (2023) find that being covered by the JA reduced the loan-to-value ratio for mortgages. My incremental contribution to this literature is that the efforts aimed at increasing flexibility in the labour market might have second-order effects on savings, on top of fertility and access to credit.

Structure of the Paper. The paper continues as follows. Section 2 describes the institutional context and shows the importance of the reform. Section 3 describes the PLUS dataset and the sample selection. Section 4 delves into the empirical challenges of estimating the causal effect of employment risk on consumption and motivates the used empirical strategy. Section 5 shows the baseline empirical results and offers a qualitative comparison with the existing reduced form evidence. Section 6 shows a battery of placebos and robustness checks. It also presents some further results in which I control for risk aversion and some split-sample regressions to investigate relevant heterogeneity dimensions. Section 7 shows the results of the model. Section 8 concludes.

2 Institutional Background and Job Mobility in Italy

This Section explains the reform of the *Jobs Act* and its main consequences for workers. Appendix E contains a detailed description of how employment protection is regulated in Italy.

The Reform. The backbone of Italy's labour market regulation is the *Chart of Labour* (COL), which applies to all employees of large firms (i.e., with more than 15 employees) hired before March 6, 2015. Under the COL, workers must be reinstated in their jobs in all cases of unjust dismissal, as determined by a labour judge. This entitles them to return to their position, receive compensation for lost wages during the period of absence, and collect a fine paid by the employer. As a result, this regulation makes dismissals both difficult and costly for employers. Indeed, labour judges in Italy tend to rule in favor of workers in the majority of cases (Del Punta 2010).

The major novelty of the *Jobs Act* (JA), which applies to employees hired by large firms starting from March 7, 2015, is that reinstatement in the case of unjust dismissal is only granted in instances of discriminatory firing - that is, when a worker is dismissed due to gender, race, sexual orientation, political or religious beliefs. In all other cases of unjust dismissal, as determined by a judge, the employer is not required to reinstate the worker but must just pay a one-time compensation that is proportional to the worker's seniority within the firm.

This change has two main consequences[6]. First, it significantly reduces employment protection in the current job, as reinstatement is much less likely for JA employees compared to those covered by the COL. Second, JA workers are more likely to bear job displacement costs: i.e., the long-lasting income losses due to unemployment. This latter point has proven particularly important and has been highlighted by Pellegrini (2025), Jarosch (2023), and Caratelli (2024), who show that workers take into account the fact that, when they move to another job, at the beginning, the risk of being fired is higher. Moreover, Bertheau et al. (2023), using administrative employee-employer matched data from several advanced economies, compares earnings up to five years after displacement and find that Italy is among the countries with the highest and most persistent negative effects of job loss.

How did the JA change Employment Protection in Italy? Bertoni et al. (2023) document that individuals hired under the JA are more concerned about losing their jobs compared to those hired under the COL. Moreover, using Italian administrative data, they show that job mobility increased following the introduction of the JA, as workers, especially in low-paying sectors, seek better employment opportunities. In other words, workers tend to change employers more quickly than before, particularly in the early months of employment.

Figure 1 shows that average job tenure (residualized from a quadratic time trend) declined after the implementation of the JA, indicating that in the five years preceding the pandemic, workers' attachment to long-term employment weakened.

Figure 2, on the other hand, presents the OECD Employment Protection Legislation (EPL) index in Italy, France, Germany, and Spain. The EPL index measures how strongly protected workers are in a country. Before the reform, Italy's EPL was the highest among the four largest economies of the European Union. However, as a consequence of the *Jobs Act*, Italy's EPL score declined sharply and fell below that of France and Germany in 2016. In addition, the 2015/2016 drop is much more pronounced than the 2011/2012 one, due to the Fornero reform (the one used by Clark et al. 2022).

However, it is important to emphasize that neither the COL nor the JA regulates mass layoffs such as those resulting from firm downsizing or closures. Instead, both laws govern cases in which employers dismiss workers for individual-specific reasons (e.g., disciplinary action or declining individual productivity). Therefore, this analysis does not address the impact of mass layoffs on employment or spending. There were, in the considered period, also reforms on the regulation of mass layoffs that applied to all large firm workers, regardless of the day on which they were hired. Therefore, they are not correlated with the *Jobs Act*.

Ten years after its approval, the *Jobs Act* continues to be a subject of political debate in Italy. In June 2025, a national referendum was held to restore the previous regulation for workers hired after March 7, 2015 - effectively aiming to abolish the *Jobs Act*. Figure B1 shows the text of the referendum. However, the law remained in place, as the referendum was declared invalid due to low voter turnout: less than 30% of eligible voters participated, even though more than

[6] Table B1 in Appendix B compares the two regimes.

80% of them voted in favor of repealing the reform[7].

3 Data and Measures of Spending

Dataset Description and Sample Selection. To conduct the analysis, I use INAPP's *Participation, Labour and Unemployment Survey* (PLUS). PLUS is a representative cross-section of the Italian population aged between 18 and 75 and contains detailed demographic and labour market information. It is particularly well-suited for my purposes and my context because it explicitly asks respondents whether they were hired before or after March 7, 2015, allowing me to precisely identify JA workers[8]. The 2018 wave of the survey additionally includes questions on risk tolerance (the inverse of risk aversion) and patience, which I use to control for unobserved preference heterogeneity (Jappelli and Pistaferri 2020) in a robustness check. The phrasing of these questions is reported in Appendix A.

For the main analysis, I retain only individuals who are currently employed. I then exclude workers whose employer is a public organization (direct civil servants, employed in state-owned business and utilities or public independent agencies) and those on fixed-term contracts, as the *Jobs Act* applies exclusively to open-ended contracts. This yields a pooled sample of approximately 8,000 observations. I then use public employees (as well as small firm workers) for the placebo analysis. Table 1 reports descriptive statistics for the full sample and separately for employees in small and large firms. Individuals under 25 represent around 10% of the sample, while those aged between 30 and 49 comprise roughly 60%. Men and married individuals are overrepresented in large firms, whereas there is no significant difference in family size. Approximately one-fifth of workers in small firms reside in the South, compared to a smaller proportion in large firms. Lower-educated individuals are more concentrated in small firms, while highly educated individuals tend to be employed in larger firms. Both income and weekly working hours are higher in larger firms. Finally, the share of income spent on food is slightly lower in larger firms, likely reflecting their higher average incomes.

Table 2 compares workers in large firms hired under either the COL or JA regime. The most notable difference between the two groups is age: about 30% of JA workers are under 30, compared to just 7% in the COL group. This reflects the fact that younger individuals are more likely, by construction, to be hired under the JA regime. Nevertheless, a considerable share of JA workers are over 40. There is no significant gender difference between the two groups, with males comprising roughly two-thirds of each sample. Geographic distribution is also broadly similar. Average net monthly individual income is nearly identical across the two groups. This is particularly important, as it indicates there are no compensating wage differentials to offset the higher employment risk faced by JA workers. Finally, JA workers report working more

[7]Under the Italian Constitution, a referendum is only valid if at least 50% of eligible voters cast a ballot, known as the *quorum*.

[8]In principle, a worker's contract could be regulated by the COL even after March 7, 2015, if the employer agrees. This would imply that the estimated causal effect shall be seen as an *intention-to-treat* effect. However, this possibility is, in actuality, very rare. Therefore, being hired after March 7 and being subject to the JA are almost collinear among large firm workers.

hours on average than COL employees.

Scaling Food Consumption. To the best of my knowledge, PLUS is the only available Italian dataset that simultaneously provides information on (i) the legal regime governing unlawful termination (COL or JA), (ii) firm size (above or below 15 employees), and (iii) consumption. However, the consumption variable refers only to monthly at-home food expenditures. Following the approach of Broda and Parker (2014), I rescale the effect on food consumption using the average share of food in total consumption, as derived from either national accounts or ISTAT's Household Budget Survey (HBS). This scaling arises from a manipulation of the household budget constraint, whereby income can be expressed as the sum of consumption and savings (Heathcote et al. 2023, Fagereng et al. 2021).

$$Y = C + S$$

Moreover, I can further split consumption into food consumption and other types of consumption (durable and nondurable), therefore having:

$$Y = C_F + C_{NF} + S$$

Where C_F denotes food consumption, while C_{NF} denotes "other" consumption. Dividing both sides by Y and multiplying and dividing C_{NF} by C_F I get:

$$1 = \frac{C_F}{Y}\left[1 + \frac{C_{NF}}{C_F}\right] + \frac{S}{Y}$$

$$\frac{S}{Y} = 1 - \frac{C_F}{Y}\left[1 + \frac{C_{NF}}{C_F}\right]$$

Computing the difference of the saving rates for COL and JA workers using the above equation, I get:

$$\Delta\frac{S}{Y} \equiv \left(\frac{S}{Y}\right)^{JA} - \left(\frac{S}{Y}\right)^{COL} = -\Delta\left(\frac{C_F}{Y}\right)\frac{C}{C_F} = -\beta\frac{C}{C_F} \tag{1}$$

Equation (1) states that the difference in the saving rate equals the negative of the difference in the food share, multiplied by the ratio of total consumption to food consumption. The first term on the right-hand side is the estimated coefficient from a regression of the average propensity to consume food on the *Jobs Act* dummy. This coefficient captures how much less, on average, individuals hired under the JA spend on food compared to those hired under the COL regime.

Figure 3 displays the ratio between total consumption and food consumption, sourced from either national accounts (NIPA) or ISTAT's Household Budget Survey (HBS) - the Italian counterpart to the U.S. Consumer Expenditure Survey. Two noteworthy patterns emerge. First, the two data series are nearly identical. Second, both remain constant over time. This constancy has important microeconomic implications for modeling the complementarity between food and non-food consumption. Specifically, a constant ratio may be interpreted as evidence of a unitary elasticity of substitution between the two consumption bundles, and thus a Cobb-Douglas utility function. Assuming a unitary elasticity of substitution is standard in housing models (see Piazzesi and Schneider 2016[9]) and in incomplete markets models with housing (see Kaplan and Violante 2014, Richard 2024). Under Cobb-Douglas preferences, one can work with the logarithm of consumption. In particular:

$$\frac{C_{NF}}{C_F} = k \iff \log C_{NF} = \log C_F + \log k$$

Where k is the constant ratio between non-food and food consumption, directly retrievable from either national account data or survey data (NIPA and HBS). As a result, the estimated coefficient for the log of food consumption is equivalent to the estimated coefficient for total log consumption.

Alternatively, still exploiting the properties of Cobb-Douglas preferences, one can divide and multiply by income and get:

$$\frac{C_{NF}}{C_F} = \frac{\frac{C_{NF}}{Y}}{\frac{C_F}{Y}} \iff \log\left(\frac{C_{NF}}{Y}\right) = k + \log\left(\frac{C_N}{Y}\right)$$

The equation above tells that estimating the effect on the log of the food consumption share is equivalent to estimating the effect on the log of food consumption and, in turn, the log of total consumption.

The above argument goes under the assumption of homothetic preferences over food and non-food consumption. Conversely, if preferences are non-homothetic - reflecting, for instance, the fact that food is a necessary good and its consumption declines less than proportionally relative to other items - then the estimates based on food consumption represent a lower bound for the true effect of the JA on total consumption. In that case, the scaled coefficient à la Broda and Parker (2014), and the BPP-imputed measure of consumption will more accurately measure consumption and therefore offer a better estimate of the causal effect of interest.

4 Empirical Methodology

I exploit the discontinuity in employment protection introduced by the hiring date to estimate the impact of employment risk on consumption and hours worked. Specifically, I estimate the following equation by OLS for employees in large firms (i.e., firms with at least 15 employees) hired under open-ended contracts in 2018:

$$Y_{is} = \alpha + \beta\, JobsAct_{is} + \varphi \mathbf{X}_{is} + \gamma_s + \varepsilon_{irs} \tag{2}$$

In equation (2), the dependent variable Y_{is} is either the ratio of food consumption over

[9] Moreover, Piazzesi et al. (2007) show that the ratio between spending for housing and overall spending is roughly constant over time. This is true both in NIPA and in CEX. This evidence point to a unitary elasticity of substitution and thus Cobb Douglas aggregation between housing and other types of spending.

family income (i.e., the average propensity to consume food) or its log. The vector $\mathbf{X}_{is}$ includes individual-level controls, including age (in bins), gender, marital status, family size, number of children, and education. γ_s represents sector fixed effects that control for the fact that hiring patterns may vary across industries, and for the unequal sectoral and dimensional distribution of firms across sectors. Finally, ε_{is} is an idiosyncratic error term. I estimate separately two equations like (2): one including only people who live in the North of Italy - the wealthiest region of the country -, and another only including people living in the Centre and in the South - including Sicily and Sardinia. This is the reason why I do not include regional fixed effects.

The coefficient of interest is β, which captures the average difference in the outcome variable between workers hired before and after March 7, 2015. I expect the coefficient to be negative: higher employment risk should increase precautionary savings and thereby push consumption downward.

Causal identification relies on the assumption that neither workers nor firms can manipulate hiring dates precisely around the cutoff, once controlling for demographics potentially affecting the hiring day. In Section 6, I provide evidence in support of this assumption.

Alternative Specification: Difference-in-Differences. On top of equation (2), I also consider a D-i-D strategy in which I exploit variation in both hiring date and firm size (i.e., above vs. below the 15-employee threshold) as in De Paola et al. (2021), Bertoni et al. (2023), and Clark et al. (2022). Table D1 in Appendix D presents the D-in-D estimates. Different concerns are associated with the D-in-D. First, marginal firms (e.g., those with 14 employees) might have strategically adjusted their size to benefit from the *Jobs Act*. The law stipulates that once a firm reaches 15 employees, not only are newly hired workers covered under the JA, but so are those hired before March 7, 2015. Boeri and Garibaldi (2019) finds a large increase in firms with more than 15 employees in the period after the reform, therefore raising a concern of firms' selection in the dimensional threshold. Moreover, the reduction in firing cost for firms was coupled with government hiring subsidies that presumably altered firms' hiring decisions (Boeri and Garibaldi 2019 and Sestito and Viviano 2018). Given that the hiring subsidy was disproportionately used by large firms, this raises a concern about whether large and small firm workers are actually comparable in terms of their wages and the employment policies they are subject to. Second, there are concerns related to self-selection in large or small firms based on consumers' attitude toward risk[10]. The direction of the bias is unclear, so no prediction could be made on the magnitude of the true effect. On the one hand, risk-averse individuals may prefer larger firms, creating a positive correlation between risk aversion and the D-i-D interaction term. This would lead to an overestimation of the magnitude of the effect[11]. On

[10]In a seminal paper, Guiso et al. (2005) show that firms pass-through a only a fraction of their idiosyncratic (permanent and transitory) shocks to wages. In this sense, they represent an important "vehicle of insurance provision" (cit. from the abstract). It's therefore sensible to imagine that risk-averse workers choose to work for the firm that will give them the best wage and employment insurance.

[11]The effect of risk aversion of current consumption is negative, through the Euler Equation, and the coefficient attached to the D-i-D interaction term is assumed to be negative. This means that the estimated causal effect is more negative, i.e., bigger in absolute value. Intuitively, more risk-averse people would consume less anyway and would select into bigger firms. Therefore, it's impossible to tell apart the effect of the JA on consumption.

the other hand, small or family-owned firms may provide higher implicit employment insurance, especially during downturns. Ellul et al. (2018) show that family firms are less likely to lay off workers during bad times due to personal attachments between employers and employees. More than 90% of businesses in Italy employ less than 10 employees and are family-owned. If such firms attract more risk-averse workers, this would introduce a negative correlation between risk aversion and the D-i-D interaction term, thus underestimating the precautionary saving response. Unfortunately, I can control for risk aversion only in 2018, since the question is asked only in the 2018 wave[12]. Third, and lastly, the data at my disposal do not allow me to test for pre-trends, given that I just have one pre-treatment period.

5 Empirical Results

Food Consumption. Table 3 reports estimates for the average propensity to food consumption. Results are reported separately for people living in the North of Italy - the wealthiest region of the country - and the Centre and the South (including Sicily and Sardinia). JA workers' average propensity to food consumption is approximately 1.5 percentage points lower than COL workers in the North of Italy. On the other hand, the effect is not statistically significant in the South. The intuition for this result is the following: people in the South are more likely to be liquidity-constrained. Therefore, the main reason they save is to escape from that situation and to be able to smooth consumption in the future. In other words, prudence doesn't play a major role in their saving decisions, and, at the margin, a change in employment risk does not induce them to consume less.

The results presented so far refer to the food consumption share of household income. Following Broda and Parker (2014), I rescale the estimated coefficient using equation (1), which links food consumption to the saving rate. In particular, I multiply the estimated coefficient by the ratio between total and food consumption. Figure 3 plots the ratio, computed using NIPA data (blue line) or HBS (red line), from 2014 to 2019. Both ratios are constant over time and are equal to 7. Therefore, the saving rate is estimated to increase by approximately 10.5 percentage points (1.5×7). Table 4 reports estimates using the log of the average propensity to food consumption as the dependent variable. Qualitatively, the results are similar: food consumption for JA workers is about 6% lower than for COL workers. This coefficient remains stable across all specifications.

BPP Imputation. The analysis above assumes Cobb-Douglas aggregation of food and non-food consumption. Even if this assumption appears to be reasonable on average, as Figure 3 demonstrates, it might not be true for some subgroups of the population. More generally, it's sensible to think that food is a necessary good, and therefore that preferences are not homothetic (and thus not Cobb-Douglas). Therefore, it's still valuable to perform the analysis

[12] The real preference heterogeneity to control is prudence. However, if CRRA preferences are assumed, as stated in the literature review, the coefficient of relative prudence is equal to the coefficient of relative risk aversion plus 1.

on a more comprehensive measure of consumption: total consumption. To get such a measure, I follow the approach proposed by Blundell et al. (2008) and modified by Patterson (2023). More specifically, I estimate in HBS the following demand for food spending, available both in HBS and PLUS:

$$\log(f_i) = \alpha + \mu \mathbf{X}_i + \delta \log(c_i) + \varepsilon_i \tag{3}$$

In the above equation, f_i is food spending, $\mathbf{X}_i$ is a vector of demographic controls available in both surveys[13]. δ is the elasticity of food consumption against total spending - durable and nondurable. The error term captures both measurement error in food consumption and unobserved heterogeneity in the demand for food (Blundell et al. 2008). I invert this demand function to back out an imputed measure of spending in PLUS using the coefficients estimated by OLS. Panel A and Panel B of Table 5 show the estimated coefficient of equation (2), for the (imputed) log of consumption and the average propensity to consume, respectively. Both dependent variables are multiplied by 100. The first three columns report estimation results for large firm workers without sector fixed effects (column 1), with job sector fixed effects (column 2), and with risk aversion dummies (column 3). The last three columns report the same specifications for small firm employees. The effect on imputed spending of the JA is around 8%. This estimate is higher in magnitude than the baseline on the log of food consumption. This result is consistent with the fact that food consumption is a necessary good, and so people cut it proportionately less than other non-essential consumption items.

Appendix C shows the details of the imputation. Table C1 shows the summary statistics of HBS, while Table C2 shows the results of the imputation. Figure C1 compares the two distributions. They both have the same mean (around 7.7), but the imputed one has fatter tails. Moreover, Figure C2, left panel, shows that there is a strong correlation between the imputed and the actual measure: the coefficient of a regression of one measure onto the other without the constant is very close to 1 with an R^2 of 0.98. The right panel of Figure C2, on the other hand, shows the scatter plot and confirms that the imputed distribution has fatter tails than the actual one, as there are many bin points far away from the 45-degree line.

How do the empirical results compare to the existing literature? Because this reduced-form approach does not structurally estimate preference parameters, the coefficients cannot be directly interpreted as the *EIS*[14]. Nonetheless, the magnitude and sign of the coefficients can be (qualitatively) compared to related reduced-form studies.

The estimated effects on food consumption and on the BPP-imputed consumption measure are qualitatively in line with Clark et al. (2022), who find a roughly 9% decline in nondurable consumption for large firm employees after the 2011 labour market reform. Unlike their study, I also detect a statistically significant decline in food consumption.

Kantor and Fishback (1996) find a 25% reduction in savings after the introduction of workers'

[13] The vector includes: age bins, gender, place of residence, Italian citizenship, education levels, marital status, and family size.

[14] In the model section, I show that a variant of the standard Bewley-Huggett-Aiyagari framework with plausible preference parameters (*EIS* of 0.5 in the baseline calibration) can replicate the empirical findings.

compensation in the 1910s, while Fuchs-Schündeln and Schündeln (2005) report that public sector employees in the former East Germany accumulated 25% less wealth than their private-sector counterparts, controlling for permanent income and demographics.

It is also informative to compare the estimates in this paper with the effects of aggregate risk on consumption. In their survey RCT in which they randomly expose survey respondents to information about European Union GDP growth in 12 months, Coibion et al. (2024) find that an increase in subjective uncertainty about GDP growth of one percentage point reduces monthly household nondurable consumption by 3.43% one month after the treatment and 3.10% four months later. They also find no significant change in the composition of nondurable spending, supporting the Cobb-Douglas assumption used in the first part of this paper. As they argue, GDP growth uncertainty might affect consumption through different channels, like future interest rates, taxes, and, most importantly, given my context, income uncertainty stemming from job uncertainty. In this regard, Patterson (2023) documents wide sectoral differences of employment sensitivity to business cycle fluctuations, fueled by the positive correlation between MPC (marginal propensity to consume) and individual earnings elasticity to GDP (what she calls the matching multiplier). In other words, while the average effect of aggregate uncertainty also contains the response of those whose income is completely insensitive to short-run fluctuations (think about public employees, for instance), individual employment risk goes deeper into this heterogeneity. In addition, I include the richest specifications sector FEs, and thus purging the analysis from sector-lever heterogeneity in job security. Moreover, Savoia (2023) documents a positive correlation between income risk and MPC conditional on the same level of assets. Therefore, Patterson (2023)'s channel is likely to be in place in the context of the JA as well.

The Effect on Labour Supply. Table B2 shows results for the intensive margin of labour supply, intended as how many hours individuals work on average per week. While the first two or three coagulums suggest an increase in weekly hours worked, the effect disappears once one controls for tenure. Indeed, either by controlling for tenure or by keeping only the individuals with maximum years of tenure, the effect is no longer distinguishable from 0 economically and statistically.

6 Placebos and Robustness Checks

In this section, I perform a battery of placebo tests and robustness checks to strengthen the causal claim made above. I also look at how the estimated causal effect varies across some heterogeneity dimensions.

Income Differences. I first examine whether the *Jobs Act* is correlated with income. It is important to confirm that incomes did not change as a result of the reform. Figure 4 plots the estimated coefficients from a regression of income (in both levels and logs) on the *Jobs Act* dummy, controlling for the same covariates used in the main analysis. The 95% confidence intervals are also displayed. In both cases, the coefficients are neither statistically nor economically different from zero: the difference in levels is just €5 (the unit is €100), equal to 0.38% of monthly income. Of course, as discussed in the model section, this does not imply that the present discounted value (PDV) of income is the same for JA and COL workers. Since JA workers face a higher probability of dismissal, they receive unemployment insurance more frequently, leading to lower cash flows to be discounted at the same interest rate. Nonetheless, it is important to establish *prima facie* that there are no systematic differences in earnings (or productivity) between JA and COL workers.

Small Firms and Public Employees. As explained in Section 2, the *Jobs Act* changed the way individual unlawful dismissals are regulated for employees of large firms. However, it did not affect the regulation for employees of small firms, employing fewer than 15 workers. Therefore, the latter ones represent a natural placebo group. I estimate the same specification as equation (2) for small firm employees. Since nothing changed in terms of individual employment protection for employees of firms with fewer than 15 workers, I should observe no difference in consumption between workers hired after or before March 7, 2015. Table 6 shows that the estimates are statistically and economically indistinguishable from zero. This result is confirmed by the last three columns of Table 5, where only small firm workers are considered. An additional placebo test is to verify that there is no effect for public sector employees, since the *Jobs Act* does not apply to them. I therefore estimate equation (2) for public employees only. Table 7 displays the results for the food share and the log of food consumption across the same specifications as before. The coefficient is both economically and statistically indistinguishable from zero.

Bigger Firms. Following De Paola et al. (2021), I restrict the sample to employees in firms with at least 20 workers and re-estimate the main equations for both food consumption and labour supply. While firms slightly below the 15-worker threshold may have some incentive to hire two or three additional workers to benefit from lower firing costs, certainly, firms well above the dimensional threshold do not have such an incentive. The estimated coefficients (in Table 8) are essentially unchanged and remain statistically significant. This serves as further evidence that firm size self-selection is not driving the results. Figure 7 shows the effect including only workers of firms with 20+, 30+ and 40+ employees.

Hiring Rates. A final concern, connected to the one raised above, is whether firms altered their hiring behavior around the time of the reform because they were aware of it. Although a reform aimed at reducing labour market rigidity and duality had long been anticipated, it is unlikely that firms or workers could have foreseen the specific discontinuous change in employment protection implemented on March 7, 2015. Additionally, the law was published in the *Gazzetta Ufficiale* on March 6, 2015, and took effect the very next day, leaving neither workers nor firms time to adjust. Figure 5, produced using data sourced from Italy's *Labour Force Survey*, shows the share of workers in the sample hired in the quarters before and after the first quarter of 2015 ($t = 0$). The hiring rate appears unchanged around this period, and the distribution is fairly uniform for both large and small firms. Additional evidence is provided in Figure 6, which plots the monthly growth rate of employment from March 2012 to March 2018. The average growth rate before the reform was 0.02%, compared to 0.06% after the reform. This change is not statistically different from 0, since the time series is very volatile. This last finding is in line with Fana et al. (2015), who argue that the JA did not change employment growth significantly.

Risk Aversion and Tenure. As anticipated in Section 2, in 2018, the survey includes questions aimed at eliciting individual discount factors and relative risk aversion, allowing me to control for preferences heterogeneity (Jappelli and Pistaferri 2020). The wording of these questions is provided in Appendix A. The first question allows for the identification of the discount factor, assuming it equals one over one plus the interest rate. The second question elicits the certainty equivalent, allowing one to infer the coefficient of relative risk aversion under CRRA utility. I estimate the following specification:

$$Y_{ir\tau s} = \alpha + \beta \, JobsAct_{ir\tau s} + \varphi \mathbf{X}_{ir\tau s} + \theta_s + \gamma_\tau + \sigma_r + \varepsilon_{ir\tau s} \tag{4}$$

In equation (4), γ_τ and σ_r are dummies representing different levels of tenure and risk aversion, respectively. The tenure FEs are 5-year bins. Table 9 shows the estimated coefficients of (4). As columns 1 and 4 suggest, controlling for preference heterogeneity makes little difference in the estimated coefficient, both for the level of the food ratio and its log. This is reassuring in the fact that workers do not self-select in or out of the *Jobs Act* based on their preferences.

Controlling for tenure, on the other hand, increases the magnitude of the estimated effect by about one percentage point: to 2.61 from 1.51. Finally, including in the sample only individuals with a maximum of 5 years of tenure, despite reducing the sample size significantly, does not influence the results.

Heterogeneity by Age. The first dimension of heterogeneity I explore is age. Gourinchas and Parker (2002) argue that households transition from buffer-stock saving behavior (Carroll 1997) to certainty-equivalent consumption plans (Hall 1978) around the age of 40–45. As a result, income risk should matter less—or not at all—for older individuals. The empirical findings (in Table 10, columns 1 and 2) are consistent with this prediction: the reduction in consumption is statistically significant only for individuals under 40, while it is indistinguishable from zero for those aged 40 and above.

Heterogeneity by education and income. The next dimension I look at is education. Columns 3 and 4 of Table 10 suggest that the effect is higher for people who hold a college degree. This is explained by the fact that higher education typically leads to higher and more stable incomes (Savoia 2023). Therefore, on the margin, a reduction in employment risk is going to matter more for the top part of the income distribution. To dig deeper into this potential explanation, columns 5 and 6 split the sample according to whether income is below or above the median. The effect is stronger for above-median incomes, providing evidence in favor of the fact that better-educated individuals have more to lose from the *Jobs Act*.

Union Representation. The PLUS dataset also asks respondents whether their workplace has union representation. The effect is significant only for unionized firms (Table 10, columns 7 and 8). This result speaks to the literature investigating the importance of union representation at the firm level as an insurance mechanism. Kim et al. (2018) shows that the presence of a union representation at the workplace fosters risk-sharing and protects workers against adverse shocks. It might be that the JA reduced, within the workplace, unions' capacity to protect workers, maybe because it altered the relationship between employees and employers in a more favorable way to the latter.

7 The Model

Why a model? As mentioned in the introduction, higher employment risk can affect consumption via two channels. The first channel is a stronger precautionary saving motive (Blanchard and Mankiw 1988, Kimball 1990). More volatile income - induced by the JA - leads consumers to save more to ensure against negative employment shocks. Furthermore, a higher separation rate lowers the PDV of income, leading to a reduction in consumption predicted by the standard PIH (Friedman 1957). Given that both channels drive consumption down, and given that I cannot observe the PDV in PLUS, it's impossible to disentangle the two channels in the above empirical analysis. This point is even more pressing given that I do not observe wealth in PLUS. From an econometric point of view, in the above empirical part, I essentially estimate this reduced form model (with no controls for expositional convenience):

$$\log\left(C_{it}\right) = \alpha + \beta JA_{it} + \varepsilon_{it} \tag{5}$$

However, the true model is likely to more closely resemble the following equation:

$$\log\left(C_{it}\right) = \alpha + \beta JA_{it} + \gamma PDV_i + \delta A_{it} + \xi_{it} \tag{6}$$

Where PDV_i is the present discounted value of income and A_{it} is assets. A straightforward application of OLS algebra (Angrist and Pischke 2009) yields:

$$\hat{\beta} = \beta + \frac{Cov\left[\gamma PDV_i + \delta A_{it}, JA_{it}\right]}{Var\left(JA_{it}\right)}$$

$$= \underbrace{\beta}_{\text{True effect of Employment Risk}} + \underbrace{\frac{Cov\left[\gamma PDV_i + \delta A_{it}, JA_{it}\right]}{Var\left(JA_{it}\right)}}_{\text{PIH Channel}}$$

The above formula shows that the OLS coefficient estimated above, even though causally identified, bundles together the two components. To quantify the contribution to the total effect, I therefore rely on the model below.

Model Set Up. Time is discrete, and the horizon is infinite. Each period is one year. There is no aggregate risk and no mortality risk. The former assumption is particularly crucial in my setting because it rules out layoffs, not covered by the *Jobs Act* regulation[15]. The economy is populated by two types of consumers/workers: a share $1-\omega$ of *Chart of Labour* workers and a share ω of *Jobs Act* workers. Their preferences are defined by a CRRA utility function with the same relative risk aversion coefficient γ[16]. They value the future at the discount factor β. At each point in time, they draw their idiosyncratic productivity from the same stochastic process that will be defined below. They face, however, two different employment risks. Employment transitions for COLs are regulated by the matrix Π_{COL}, with job finding probability f and separation rate s_{COL}. The matrix for JAs is, instead, Π_{JA} with the same f but with $s_{JA} > s_{COL}$. It's worth remarking that the "outer" employment transition matrix governing employment transitions and the matrix governing productivity draws are independent. This means that employees face the same unemployment risk regardless of their productivity draw. This rules out that the employer's decisions to fire workers are contingent on their productivity. The rationale for this assumption is that I want COL and JA employees to be completely comparable in terms of their individual characteristics.

$$
\begin{gathered}
\max_{\{a_{j,t+1},c_{j,t}\}} \quad \mathbb{E}_0 \sum_{t=0}^{\infty} \beta^t \left[\frac{c_{j,t}^{1-\gamma}-1}{1-\gamma}\right] \\
a_{j,t+1} = (1+r)a_{j,t} + y\big(e_t(e_{t-1},\Pi_j),p_t\big) - c_{j,t} \\
a_{j,t+1} \geq 0 \\
y\big(e_t(e_{t-1},\Pi_j),p_t\big) = \begin{cases} w \times p_t & \text{if} \quad e_t = 1 \\ b(p_t) \times p_t & \text{if} \quad e_t = 0 \end{cases} \\
b(p_t) = \begin{cases} b_0 \times p_t & \text{if} \quad p_t \leq p_{50} \\ 0.5 \times p_{50} & \text{if} \quad p_t > p_{50} \end{cases} \\
\log(p_t) = \rho \log(p_{t-1}) + \varepsilon_t, \qquad \varepsilon_t \sim N(0,\sigma_\varepsilon^2) \\
\Pi_j = \begin{bmatrix} 1-f & f \\ s_j & 1-s_j \end{bmatrix}, \quad j \in \{COL, JA\}
\end{gathered} \tag{7}
$$

When employed ($e_t = 1$), workers earn their productivity draw times w (normalized to 1); when unemployed ($e_t = 0$), their productivity draw times $b(p_t)$, i.e. the replacement rate of the unemployment insurance. Unemployment insurance is equal to 0.75 of productivity ($b_0 = 0.75$) if productivity is below the cross-sectional median (p_{50}). Otherwise, it's capped at one-half of

[15]There were reforms on the regulation of mass layoffs procedures, but they apply regardless of the hiring date. So, they do not interact with the JA.

[16]In Appendix F, show that results are unchanged with Epstein-Zin preferences.

the median productivity. This design is meant to replicate as closely as possible the Italian regulation of unemployment benefits. The log of productivity $\log(p_t)$ is modeled as an AR(1) with persistence ρ and variance of the stochastic component σ^2_ϵ. Consumers can save in a liquid asset that gives a fixed return of r. They choose consumption each period to maximize the discounted sum of expected utility subject to the dynamic budget constraint and the borrowing constraint, respectively, the second and the third line of equation (7).

Calibration. The calibration of the model is in two steps. First, I chose a set of parameters externally, choosing values commonly used in the literature. I then choose a second set of parameters internally with indirect inference methods. I set the risk aversion γ at 2, in line with recent Euler-Equation estimations for Italy (Guiso and Jappelli 2024a, Bertola et al. 2005) and other European Countries (Sciacchetano 2024). Consistent with Jappelli et al. (2024) and the standard of the literature (Auclert et al. 2024), I model the productivity process with an AR(1) with 7 states. The persistence parameter ρ is set to 0.95, while the standard deviation of the productivity is $\sigma_\varepsilon = 0.5$. The wage w is normalized to 1. The borrowing constraint is set to 0. The share of JAs ω is 0.17, to match the amount of large firm workers hired after 2015 as of 2018.

I internally calibrate four parameters with the Method of Simulated Moments (MSM). I choose the parameters of interest to minimize the distance between key model-implied moments and their empirical counterpart. This amounts to solving the following minimization problem. Let θ and $\hat{\theta}$ be respectively the vectors of the true parameters and the estimated ones. Let $\mathbf{W}$ be a weighting matrix and ψ the vector of the externally set parameters. Let $g(\theta;\psi)$ be the vector of theoretical moments and m the vector of data moments. The minimization problem can be written as:

$$
\begin{gathered}
\hat{\theta} = \arg\min_{\theta} \sum_{k=1}^{4} \left[\frac{g_k(\theta;\psi) - m_k}{0.5 g_k(\theta;\psi) + 0.5 m_k} \right] W_k \\
\theta = \Bigg[\beta, s_{COL}, s_{JA}, f \Bigg]
\end{gathered}
\tag{8}
$$

I estimate with MSM the following four parameters: the discount factor β, the separation rates for COLs and JAs s_{COL}, s_{JA}, and the job finding rate f. I also impose that $s_{JA} > s_{COL}$. I target four moments: the average wealth-to-income ratio (including housing) of Italian households in 2016 of 4.44, the average tenure of COL workers of 18.3 years, the average unemployment rate, and its standard deviation between 2016 and 2019, respectively at 10.75% and 0.87%. The source of the first moment is the 2016 *Survey of Household Income and Wealth* (SHIW), the source of the second moment is the 2018 waves of the LFS, while to obtain the remaining two moments, I look at aggregate statistics on unemployment sourced from the Federal Reserve of St. Louis website.

Model Results. I solve the model by the endogenous grid points method (see Carroll 2006), separately for JA and COL workers. Figure 8 shows the consumption policy function for JA (red line) and COL (blue line) in the states of employment (left panel) and unemployment (right panel), both in the median productivity state. JA workers' consumption - conditioning on the same wealth, productivity, and employment status - is always below that of COL workers. In other words, JA workers choose to consume less than COLs, even when they have the same resources as the latter ones. Following the above discussion, this result might be either due to higher precautionary saving or a lower permanent income.[17]

Table 11 shows the calibration for the model. Table 12, on the other hand, shows the performance of the MSM, comparing targeted empirical moments and the theoretical counterparts. In the table, the estimated coefficients are reported. Overall, the model manages to match the targeted parameters decently closely. The discount factor β is 0.96, while the separation rate for COLs is 0.05, close to the separation rate found (transitions from Employment to Unemployment) by D'Amuri et al. (2022), who find it to be around 4.8% (annualized) in the period before the financial crisis. The separation rate for JAs is around 10%, while the finding rate is 50%.

Figure 9, on the other hand, shows the distribution of consumption (left panel) and the present discounted value of income (right panel) for JA and COL workers, in red and blue, respectively. The PDV of income is calculated by summing all incomes an individual earns over the simulated period, discounted by $1 + r$ to the power of the period.[18]. It can be seen that the COLs almost first-order stochastically dominate the JAs both for consumption and for the PDV. This finally confirms that the concerns raised at the beginning of this section were well posed. To understand which of the two channels prevails, I estimate (5) and (6) with simulated data and report the results in Table 13. Column 1 simply shows the mean difference in the log of consumption between JA and COL workers. This difference of 5.4% not far from the empirical estimate of imputed consumption. Column 2 controls for the PDV of income and wealth. I find that a coefficient attached to *Jobs Act* is slightly below 3%. This means that the sole effect of employment risk accounts for roughly 40% of the total effect of the *Jobs Act*. Column 3 allows for a more flexible specification with the interaction between employment risk, PDV, and assets, finding that the effect of the JA decreases in magnitude as the PDV and assets increase. This is very intuitive: as a consumer is on higher levels of cash on hand, she is more able to smooth consumption, and therefore the precautionary saving motive weakens (i.e., she is on the "linear" side of the consumption policy function). The coefficients attached to tenure and its interaction with the *Jobs Act* are positive. The intuition of this result is that the higher the tenure, the longer the period during which unemployment does not actually occur. Therefore, individuals who buffer up savings for precautionary reasons, when the risk does not materialize, end up with extra savings they can use to consume more. Column 3 of Table 13 also shows that the positive effect on consumption of an increase in the PDV is higher

[17] Figure B3 plots the mean income for JA and COL workers over the simulated period of the model.

[18] $PDV_i = \sum_{t=1}^{200} y_{it} \times (1 + r)^{-(t-1)}$

for JA workers as opposed to COLs. Indeed, the coefficient associated with the interaction term between the PDV and the JA is equal to 0.129, suggesting that the same unit increase in PDV spurs consumption of JA workers by 0.13 percentage points more compared to the COL. This result echoes the one by Savoia (2023), who shows that high-income risk individuals have higher MPCs, even conditioning on wealth.

The importance of precautionary saving motive due to the JA can also be appreciated in the model by regressing wealth (scaled down by income) against the JA dummy, a specification that resembles the one by Fuchs-Schündeln and Schündeln (2005). I do this exercise in columns 4 and 5 of Table 13. When I do not control for income - i.e., I simply look at the unconditional mean difference in assets between JAs and COLs - the coefficient is negative. However, when I control for the PDV, and so I condition on the same amount of permanent resource, the difference changes sign and becomes slightly positive. This change in the sign of the coefficient is the effect of the precautionary saving motive induced by the JA.

Model Extensions. Appendix F proposes two variants of the baseline model. The first variant is in the assumed preferences: recursive preferences à la Epstein and Zin (1989) instead of CRRA. With recursive preferences, risk aversion, and the elasticity of intertemporal substitutions are governed by two distinct parameters. The first parameter governs how individuals are willing to smooth consumption over time, while the second parameter governs how individuals want to smooth consumption over states. A comparison between Table F2 and Table 13 shows that there are no significant differences between the two model specifications.

The second extension I propose is to introduce a second, illiquid account as in Auclert et al. (2024), Kaplan and Violante (2022), and Graves (2025). In this class of more realistic models, consumers can save in a liquid asset and an illiquid asset. The last one pays a higher return, but can be accessed infrequently, and is subject to transaction costs. Consumers maximize the same utility function - assumed to be CRRA -, but face the following borrowing constraints[19]:

$$
\begin{gathered}
c_t + b_{t+1} + \big[\psi(a_{t+1}, a_t) + d_t\big]\, \mathbb{I}_{adjust=1} = (1 + r_b)b_t + y\big(e_t(e_{t-1}, \Pi_j), p_t\big) \\
a_{t+1} = (1 + r_a)a_t + d_t\, \mathbb{I}_{adjust=1} \\
\psi(a_{t+1}, a_t) = k\, |a_{t+1} - a_t|,\ \ a_{t+1} \geq 0,\ \ b_{t+1} \geq 0
\end{gathered}
\tag{9}
$$

When consumers have the possibility of adjusting the illiquid account ($\mathbb{I}_{adjust=1}$), they deposit an amount d_t and pay a transaction cost $\psi(a_{t+1}, a_t)$, that is equal to the percentage k of the absolute difference between the positions on $t+1$ and t. If there is no possibility of adjustment ($\mathbb{I}_{adjust=0}$), the problem is the same as in the one-asset case. Neither the liquid nor the illiquid account can be shorted. Appendix F contains the recursive formulation of the model and the details on the calibration.

The insight from the papers mentioned above is that people save in the liquid account to

[19] I omit the j index in the budget constraint.

have cash available in the case of a temporary negative shock occurring. The illiquid account, on the other hand, is only used to save for intertemporal reasons, and it's only withdrawn if big shocks occur. In particular, Graves (2025) shows that when there is an increase in the separation rate, households increase their holdings in the liquid account model to buffer up more wealth to deal with employment risk. In my setting, I should expect JAs workers to hold a higher share of their wealth in the illiquid account, as the risk of becoming unemployed is larger for them. This is exactly what I found in the model. While the difference in consumption is very close to the empirical result (6.78%), JAs workers have a ratio of liquid wealth to income around 24% higher than the COLs, as shown by Table F4.

8 Conclusion and Direction for Further Research

In this paper, leveraging the Italian *Jobs Act* (JA), I investigate the relationship between employment risk, consumption, and labour supply. I find that workers affected by the reform reduce their food spending-to-income ratio by 1.5 percentage points, corresponding to a 8% drop in spending and a 10 percentage point increase in the saving rate. These effects are stronger for younger, unmarried, and less educated workers in Northern Italy.

To disentangle the pure effect of employment risk from the permanent income hypothesis (PIH) channel - arising from a lower present discounted value (PDV) of income - I develop a Bewley-Huggett-Aiyagari model with ex-ante heterogeneity in employment risk. Calibrating the model to match Italy's labour market flows and assuming standard CRRA preferences, it generates a difference in consumption between risky and safe employees that is both qualitatively and quantitatively in line with the estimated causal effect. Running a regression on simulated data, controlling for PDV and asset levels, I find that employment risk heterogeneity accounts for around 40% of the observed consumption difference. An alternative version with recursive preferences as in Epstein and Zin (1989) gives similar results. Finally, a model in which consumers are allowed to save in two accounts - one liquid and one illiquid - suggests that JAs workers accumulate more wealth in liquid assets than COLs.

This work can be improved along several dimensions. To the best of my knowledge, currently, no dataset provides joint information on firm size, employment protection, and detailed spending categories beyond food. While the results presented here are informative and consistent with theoretical predictions - further validated by model simulations - access to actual (e.g., self-reported) consumption data, rather than imputed spending, and possibly wealth, would represent a substantial data improvement.

Finally, a comprehensive welfare evaluation of the JA should combine the demand-side effects examined here (i.e., the impact on the precautionary saving motive) with the supply-side impacts that the reform aimed to achieve.

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Tables

Table 1: Summary Statistics by firm size

	Large Firms		Small Firms		All	
	Mean	Sd	Mean	Sd	Mean	Sd
Between 18 and 24	0.03	0.16	0.04	0.20	0.03	0.17
Between 25 and 29	0.06	0.24	0.09	0.28	0.07	0.26
Between 30 and 39	0.26	0.44	0.29	0.45	0.27	0.44
Between 40 and 49	0.37	0.48	0.34	0.48	0.36	0.48
Between 50 and 64	0.28	0.45	0.24	0.43	0.27	0.44
Male	0.65	0.48	0.51	0.50	0.61	0.49
Married	0.62	0.49	0.58	0.49	0.60	0.49
Family Size	3.02	1.18	3.05	1.19	3.03	1.18
N. of Children	1.10	1.02	1.05	1.01	1.08	1.02
South and Islands	0.17	0.38	0.21	0.41	0.19	0.39
Centre	0.21	0.40	0.23	0.42	0.21	0.41
North	0.62	0.49	0.56	0.50	0.60	0.49
Low Education	0.36	0.48	0.43	0.50	0.38	0.49
Middle Education	0.49	0.50	0.48	0.50	0.49	0.50
High Education	0.15	0.35	0.09	0.28	0.13	0.33
Union Representation	0.69	0.46	0.05	0.21	0.47	0.50
Jobs Act	0.06	0.25	0.09	0.28	0.07	0.26
Food Share (%)	29.37	12.25	30.29	12.76	29.66	12.42
Food Consumption (€100)	6.72	4.01	6.16	3.86	6.55	3.97
Individual Net Income (€100)	13.97	5.23	10.94	4.47	13.01	5.19
Family Net Income (€100)	23.21	10.63	20.31	9.74	22.29	10.44
Weekly average hours worked	37.47	8.26	34.63	10.76	36.55	9.24
Employed in a large Firm	.	.	.	.	0.68	0.47

Notes: This Table displays the summary statistics for the 2014 and 2018 PLUS waves by firm size.

Table 2: Summary Statistics by contract type - only large firm employees

	JA		COL	
	Mean	Sd	Mean	Sd
Between 18 and 24	0.12	0.33	0.02	0.14
Between 25 and 29	0.20	0.40	0.05	0.23
Between 30 and 39	0.31	0.46	0.26	0.44
Between 40 and 49	0.24	0.43	0.37	0.48
Between 50 and 64	0.13	0.34	0.29	0.45
Male	0.63	0.48	0.66	0.48
Married	0.41	0.49	0.63	0.48
Family Size	2.99	1.22	3.02	1.18
N. of Children	0.71	0.98	1.12	1.02
South and Islands	0.20	0.40	0.17	0.38
Centre	0.20	0.40	0.21	0.41
North	0.60	0.49	0.62	0.48
Low Education	0.23	0.42	0.37	0.48
Middle Education	0.52	0.50	0.49	0.50
High Education	0.26	0.44	0.14	0.35
Union Representation	0.46	0.50	0.70	0.46
Food Share (%)	27.14	12.81	29.55	12.19
Food Consumption (€100)	6.49	4.12	6.74	4.00
Individual Net Income (€100)	13.43	5.33	14.01	5.22
Family Net Income (€100)	24.42	10.69	23.13	10.62
Weekly average hours worked	38.24	9.05	37.42	8.20

Notes: This Table displays the summary statistics for the 2014 and 2018 PLUS waves by contract type (JA and COL). Only large firm employees are included. Sample weights are used.

Table 3: Food Share ($\times$ 100)

	North			Centre and South		
	(1)	(2)	(3)	(4)	(5)	(6)
Jobs Act	-1.63**	-1.72**	-1.57**	0.74	0.59	0.72
	(0.68)	(0.69)	(0.68)	(1.03)	(1.05)	(1.04)
Male	-1.59***	-1.53***	-1.14**	-1.66**	-1.45**	-1.33*
	(0.51)	(0.54)	(0.53)	(0.68)	(0.70)	(0.69)
Married	-1.26**	-1.18*	-0.44	-0.70	-0.64	-0.01
	(0.62)	(0.63)	(0.62)	(0.87)	(0.88)	(0.87)
Family Size	1.94***	1.89***	2.82***	2.36***	2.37***	2.97***
	(0.31)	(0.32)	(0.33)	(0.34)	(0.35)	(0.36)
N. of Children	0.91**	0.94**	0.34	-0.22	-0.28	-0.55
	(0.39)	(0.39)	(0.40)	(0.48)	(0.48)	(0.48)
Between 25 and 29	-2.51***	-2.57***	-3.41***	-2.93**	-3.46***	-4.56***
	(0.97)	(0.97)	(0.95)	(1.29)	(1.32)	(1.33)
Between 30 and 39	-1.02	-1.07	-1.64**	-1.62*	-1.88**	-2.63***
	(0.69)	(0.70)	(0.69)	(0.90)	(0.92)	(0.92)
Between 40 and 49	-0.81	-0.90	-1.17*	-1.19	-1.37	-2.11**
	(0.65)	(0.66)	(0.65)	(0.84)	(0.85)	(0.84)
Years of Education	-0.58***	-0.55***	-0.33***	-0.74***	-0.68***	-0.43***
	(0.08)	(0.08)	(0.09)	(0.10)	(0.10)	(0.10)
Net Family Income (€100)			-0.23***			-0.23***
			(0.03)			(0.03)
Observations	2227	2227	2227	1461	1461	1461
Controls	✓	✓	✓	✓	✓	✓
Job Sector FE		✓	✓		✓	✓

Notes: This Table shows OLS estimates of (2). The dependent variable is food consumption as a share of family income. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample includes only large firm workers in 2018. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4: Log of Food Share (× 100)

	North			Centre and South		
	(1)	(2)	(3)	(4)	(5)	(6)
Jobs Act	-6.25**	-6.58**	-5.92**	-0.34	-0.89	-0.70
	(2.87)	(2.89)	(2.88)	(3.75)	(3.81)	(3.76)
Male	-6.57***	-6.09***	-4.70**	-4.57*	-3.64	-3.28
	(1.96)	(2.10)	(2.07)	(2.52)	(2.64)	(2.60)
Married	-5.46**	-5.04**	-1.95	-2.35	-1.99	0.62
	(2.43)	(2.43)	(2.43)	(3.21)	(3.23)	(3.16)
Family Size	7.09***	6.89***	10.38***	8.99***	8.96***	11.46***
	(1.25)	(1.25)	(1.31)	(1.32)	(1.34)	(1.39)
N. of Children	3.96***	4.07***	1.80	-0.70	-0.93	-2.10
	(1.45)	(1.44)	(1.45)	(1.73)	(1.73)	(1.72)
Between 25 and 29	-11.55***	-11.83***	-14.82***	-10.34**	-12.03**	-15.98***
	(4.17)	(4.18)	(4.11)	(4.86)	(4.94)	(4.93)
Between 30 and 39	-3.00	-3.27	-5.28**	-4.60	-5.42	-8.20**
	(2.63)	(2.65)	(2.64)	(3.27)	(3.33)	(3.32)
Between 40 and 49	-4.10*	-4.43*	-5.35**	-4.49	-5.12*	-7.82**
	(2.48)	(2.49)	(2.47)	(3.06)	(3.10)	(3.05)
Years of Education	-2.43***	-2.32***	-1.52***	-2.66***	-2.50***	-1.58***
	(0.30)	(0.32)	(0.33)	(0.36)	(0.37)	(0.39)
Net Family Income (Log x 100)			-0.21***			-0.22***
			(0.03)			(0.03)
Observations	2227	2227	2227	1461	1461	1461
Controls	✓	✓	✓	✓	✓	✓
Job Sector FE		✓	✓		✓	✓

Notes: This Table shows OLS estimates of (2). The dependent variable is the log of the food consumption as a share of family income. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The controls are the same as in the previous table. The sample includes only large firm workers in 2018. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 5: Imputed Consumption

	Large Firms			Small Firms		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Log of Imputed Consumption						
Jobs Act	-8.58**	-8.68**	-7.91**	-1.65	0.44	-1.90
	(3.58)	(3.59)	(3.62)	(5.18)	(5.30)	(5.13)
Panel B: Consumption to Income Ratio						
Jobs Act	-6.64***	-7.09***	-6.64***	-3.85	-2.97	-2.80
	(2.45)	(2.46)	(2.46)	(3.52)	(3.64)	(3.63)
Observations	3753	3753	3753	1460	1460	1460
Sector FE		✓	✓		✓	✓
Risk Aversion			✓			✓

Notes: This Table shows the OLS estimates for equation (2). The dependent variable is the BPP-imputed log of consumption (panel A) and the imputed consumption to income ratio (panel B). Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 6: Small Firm Workers

	Food Share			Log Food Share		
	(1)	(2)	(3)	(4)	(5)	(6)
Jobs Act	-0.24	-0.14	-0.18	0.30	0.83	0.56
	(0.77)	(0.80)	(0.79)	(2.91)	(3.01)	(2.98)
Family Net Income (€100)			-0.22***			
			(0.04)			
Household Income (Log x 100)						-0.18***
						(0.03)
Observations	1428	1428	1428	1428	1428	1428
Controls	✓	✓	✓	✓	✓	✓
Job Sector FE		✓	✓		✓	✓

Notes: This Table shows OLS estimates of (2). The dependent variable is either the food consumption as a share of family income (first three columns) or its log (last three columns). Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The controls are the same as in the previous table. The sample includes only small firm workers in 2018. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 7: Public Sector Workers

	Food Share			Log Food Share		
	(1)	(2)	(3)	(4)	(5)	(6)
Jobs Act	0.15	0.10	0.31	1.45	1.16	1.54
	(0.87)	(0.87)	(0.86)	(3.25)	(3.23)	(3.18)
Family Net Income (€100)			-0.23***			
			(0.02)			
Household Income (Log x 100)						-0.21***
						(0.02)
Observations	3632	3632	3632	3632	3632	3632
Controls	✓	✓	✓	✓	✓	✓
Job Sector FE		✓	✓		✓	✓

Notes: This Table shows OLS estimates of (2). The dependent variable is either the food consumption as a share of family income (first three columns) or its log (last three columns). Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The controls are the same as in the previous table. The sample includes only workers of the public sector in 2018. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 8: Bigger Firms ($\geq$ 20 employees)

	Food Share			Log Food Share		
	(1)	(2)	(3)	(4)	(5)	(6)
Jobs Act	-1.58**	-1.70**	-1.66**	-6.30**	-6.64**	-6.35**
	(0.74)	(0.75)	(0.74)	(3.14)	(3.16)	(3.15)
Family Net Income (€100)			-0.22***			
			(0.03)			
Household Income (Log x 100)						-0.20***
						(0.03)
Observations	2041	2041	2041	2041	2041	2041
Controls	✓	✓	✓	✓	✓	✓
Job Sector FE		✓	✓		✓	✓

Notes: This Table shows OLS estimates of (2). The dependent variable is either the food consumption as a share of family income (first three columns) or its log (last three columns). Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The controls are the same as in the previous table. The sample workers of firms with at least 20 employees in the North of Italy. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 9: Controlling for Risk Aversion and Tenure

	Food Share			Log Food Share		
	(1)	(2)	(3)	(4)	(5)	(6)
Jobs Act	-1.51**	-2.61**	-2.83**	-5.79**	-10.12**	-10.72**
	(0.69)	(1.24)	(1.26)	(2.88)	(5.02)	(5.08)
Family Net Income (€100)	-0.23***	-0.23***	-0.11**			
	(0.03)	(0.03)	(0.05)			
Household Income (Log x 100)				-0.21***	-0.21***	-0.12**
				(0.03)	(0.03)	(0.06)
Observations	2227	2227	435	2227	2227	435
Controls	✓	✓	✓	✓	✓	✓
Preference FE	✓	✓	✓	✓	✓	✓
Tenure FE		✓	$\leq$ 5 years		✓	$\leq$ 5 years

Notes: This Table shows OLS estimates of (2). The dependent variable is either the food consumption as a share of family income (first three columns) or its log (last three columns). Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The controls are the same as in the previous table. The sample workers of firms with at least 20 employees in the North of Italy. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 10: Heterogeneity Analysis

	By Age		By Education		By Income		Union Representation	
	$<$ 40 y.o.	$\geq$ 40 y.o.	No College	College	Below Median	Above Median	No	Yes
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Panel A: Food Share								
Jobs Act	-1.94***	-1.35	-1.57**	-2.22***	-2.28***	-3.14***	-1.60*	-3.41***
	(0.72)	(0.99)	(0.76)	(0.82)	(0.75)	(0.84)	(0.83)	(0.78)
Panel B: Log of Food Share								
Jobs Act	-10.22***	-5.24	-6.11**	-11.51***	-9.49***	-15.11***	-7.38**	-15.27***
	(2.93)	(3.72)	(2.84)	(3.48)	(2.83)	(3.55)	(3.24)	(3.17)
Observations	1304	2385	2562	1126	1872	1818	1244	2445

Notes: This Table shows OLS estimates of (2). The dependent variable is either the food consumption as a share of family income (Panel A) or its log (Panel B). Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample is split according to the definitions of the columns. The whole sample includes people from the North and the South of Italy. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 11: Model Baseline Calibration

Parameter	Description	Value	Source/Target
Preference Parameters			
γ	Relative Risk Aversion	2	Standard
β	Discount Factor	0.96	Average Wealth-to-Income ratio
Productivity Process			
ρ	Persistence of AR(1)	0.95	Jappelli et al. (2024)
σ_ε	Std of the shock	0.50	Jappelli et al. (2024)
n_e	Number of States	7	
Job Market Flows			
s_{COL}	COL Job Separation Rate	0.05	Average tenure of COLs
s_{JA}	JA Job Separation Rate	0.10	Average Unemployment Rate in 2016-2019
f	Job Finding Rate	0.50	Volatility of Unemployment Rate in 2016-2019
Asset Grid			
a_{min}	Borrowing Constraint	0	Standard
a_{max}	Maximum Value of Asset	50	Standard
n_a	Number of Asset Grid-points	100	Standard
Prices			
w	Wage Rate	1	Normalization
b_0	UI Replacement Rate	0.75	UI Replacement Rate
r	Annual Interest Rate	0.02	Jappelli et al. (2024)

Table 12: Targeted Moments

Parameter	Estimate	Moment	Source	Model	Data
β	0.96	Wealth to Income Ratio	SHIW 2016	4.48	4.44
s_{COL}	0.05	COLs Years of Tenure	LFS 2018	18.11	18.31
s_{JA}	0.10	Unemployment Rate	FRED-FED 2016-2019	10.31	10.75
f	0.50	Std Unemployment Rate	FRED-FED 2016-2019	0.85	0.87

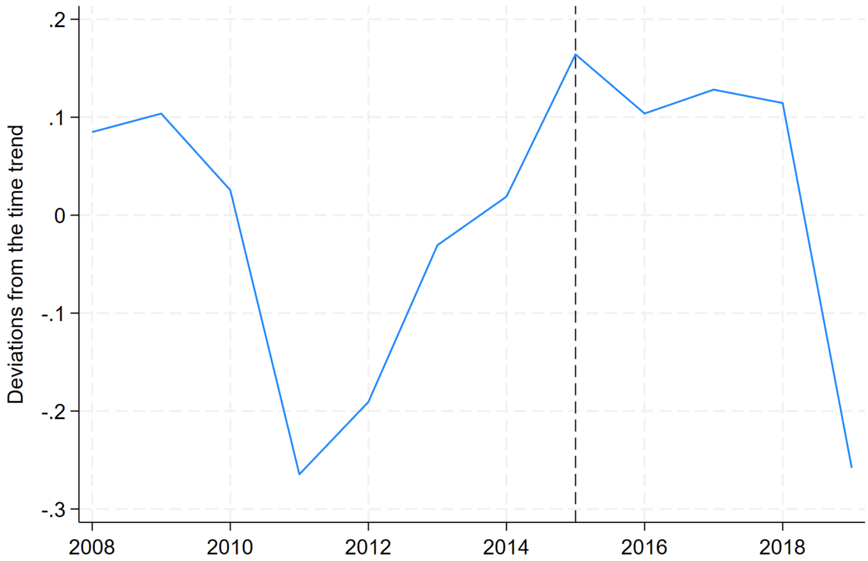
Table 13: Regression with Simulated Data

Dependent Variable:	Log of Consumption × 100			Wealth / Income	
	(1)	(2)	(3)	(4)	(5)
Jobs Act	-5.382***	-2.197***	-5.642***	-0.334***	0.097***
	(0.115)	(0.107)	(0.389)	(0.016)	(0.015)
PDV of Income		1.713***	1.694***		0.195***
		(0.005)	(0.005)		(0.001)
Initial Wealth		0.197***	0.218***		0.204***
		(0.010)	(0.011)		(0.001)
Jobs Act x PDV of Income			0.129***		
			(0.013)		
Jobs Act x Init. Wealth			-0.157***		
			(0.030)		

Notes: This Table shows the OLS estimated coefficient of equation (5), run with simulated data. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

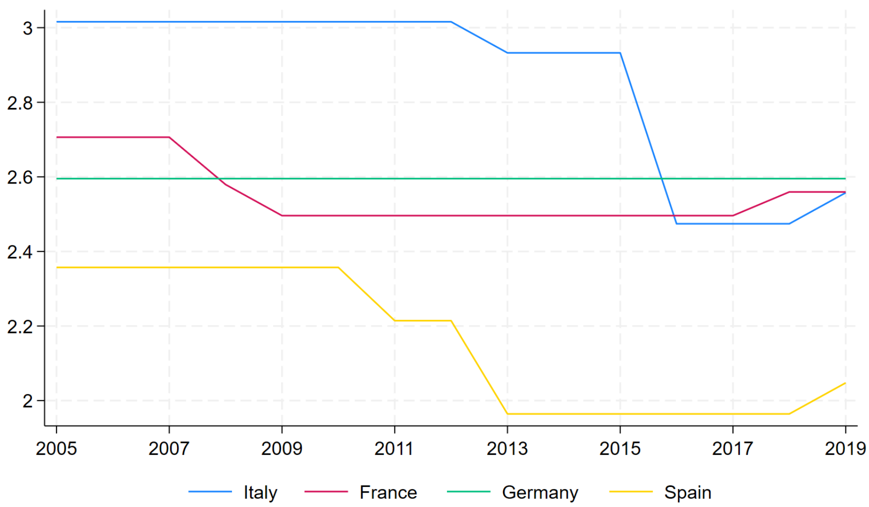
Figures

Figure 1: Workers' tenure in large firms over time.



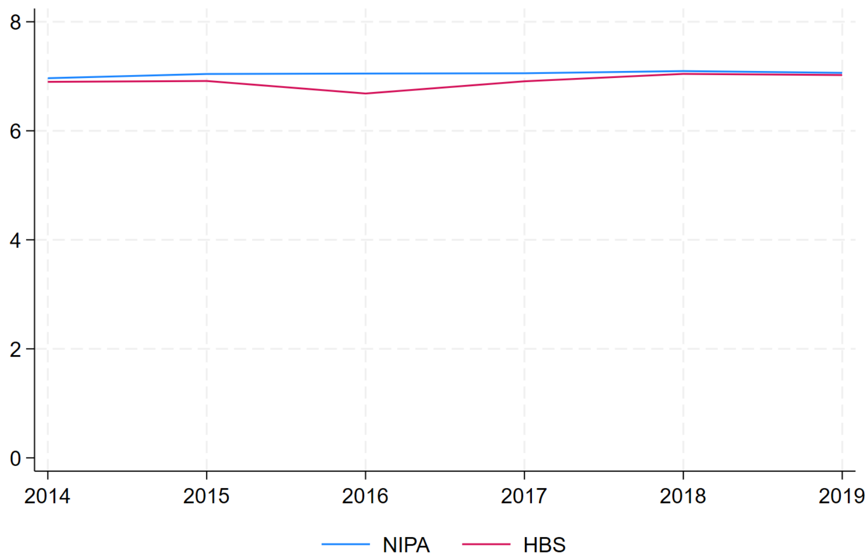
Notes: The Figure shows the average tenure of Italians employed in large firms between 2008 and 2019. Data are de-trended to account for the business cycle (i.e., the 2008-2011 double-deep recessions). Average tenure increases until 2016 and then goes down sharply.

Figure 2: Employment Protection in OECD countries



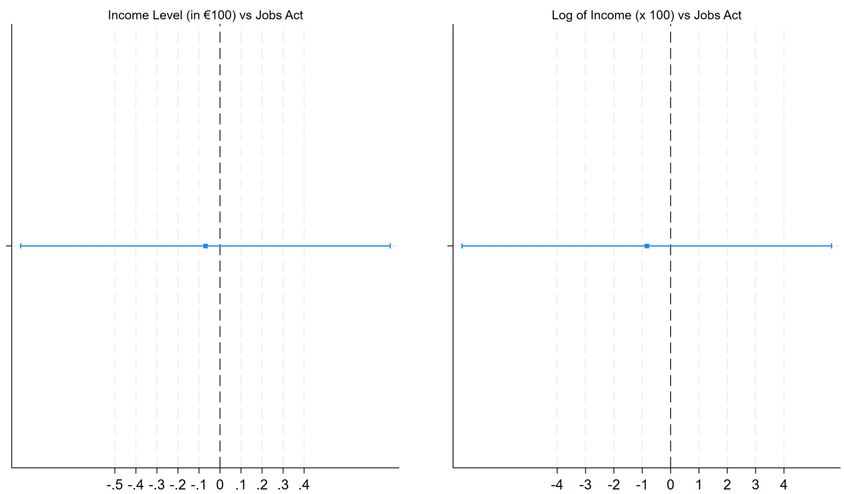
Notes: The Figure shows the employment protection index computed by the OECD. A sharp drop can be seen in 2015 and 2016 for Italy, represented by the blue line on top. It's interesting to notice that even after the 2012 reform, Italy had a more rigid labour market than many of the OECD countries.

Figure 3: Ratio between Non-Food and Food Consumption



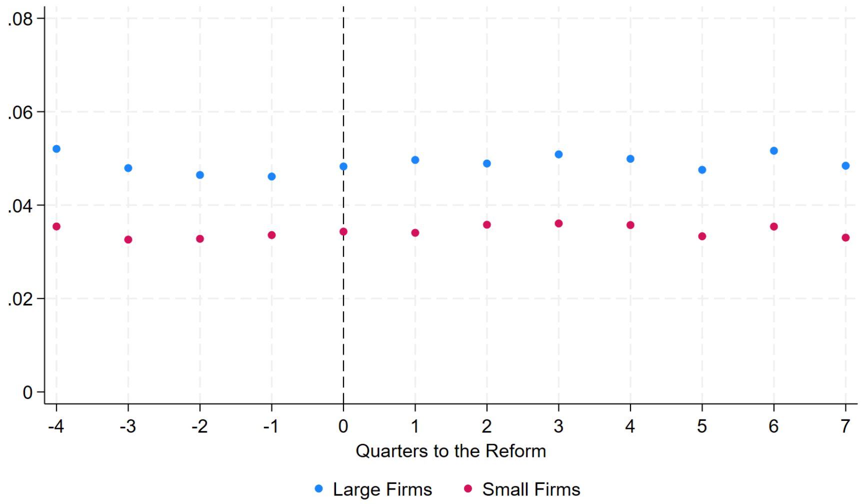
Notes: The Figure shows the ratio between non-food consumption and food consumption from national accounts and from ISTAT's HBS from 2014 to 2019. The HBS series is constructed by dropping the top and bottom deciles of the pooled cross-section distribution.

Figure 4: Income vs Jobs Act



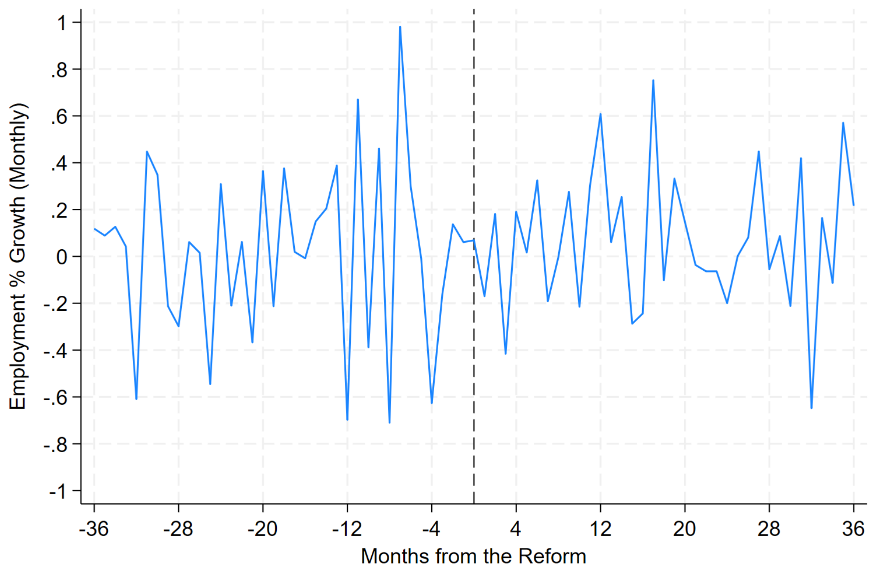
Notes: This Figure shows the OLS estimated coefficient of income (both in levels and in logs) and the Jobs Act dummy, controlling for the same variables used in the main analysis.

Figure 5: Hiring Rates in Small and Large Firms



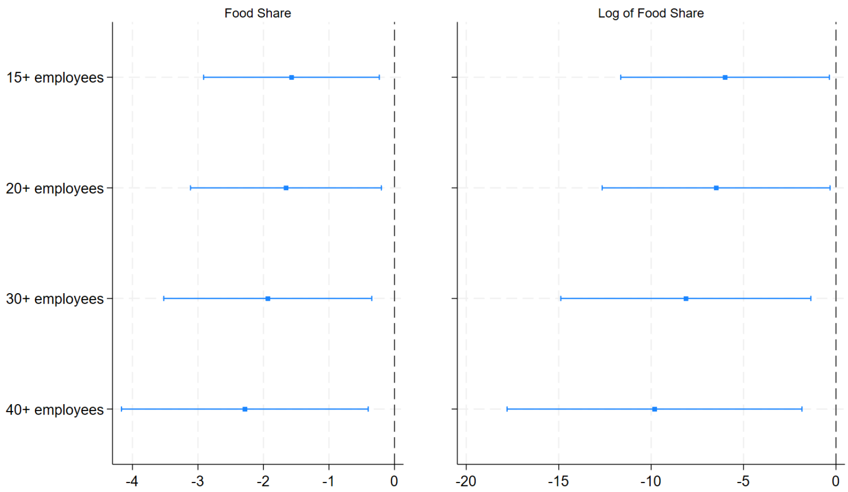
Notes: This Figure shows the sample hiring rates before and after the reform (first quarter 2015, t = 0). The horizontal axis is the quarter from 2014 (-4 is 2014-q1) to 2019 (7 is 2019-q4). The Source is the Italian Labour Force Survey.

Figure 6: Monthly Net Employment Growth



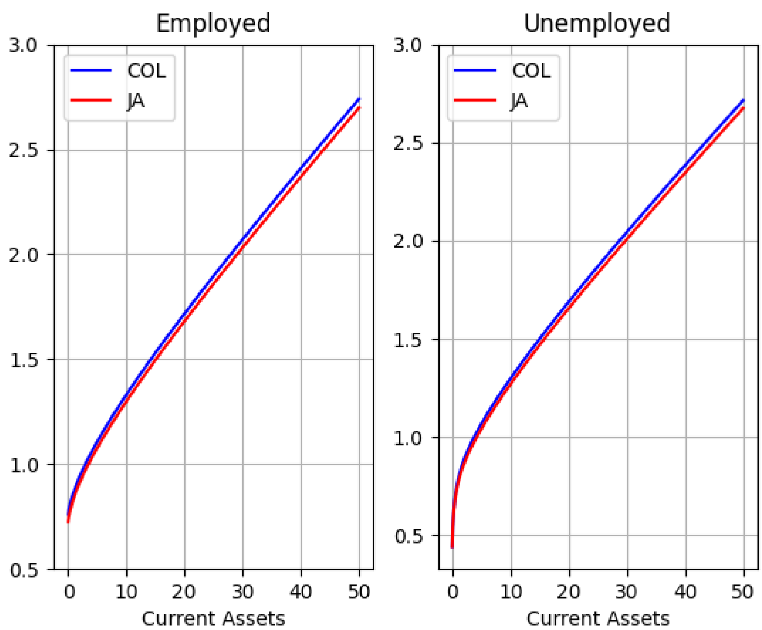
Notes: This Figure shows the monthly growth rate of the level of employment. The Source is the Italian Statistical Office.

Figure 7: Estimated Causal Effects for Bigger Firms



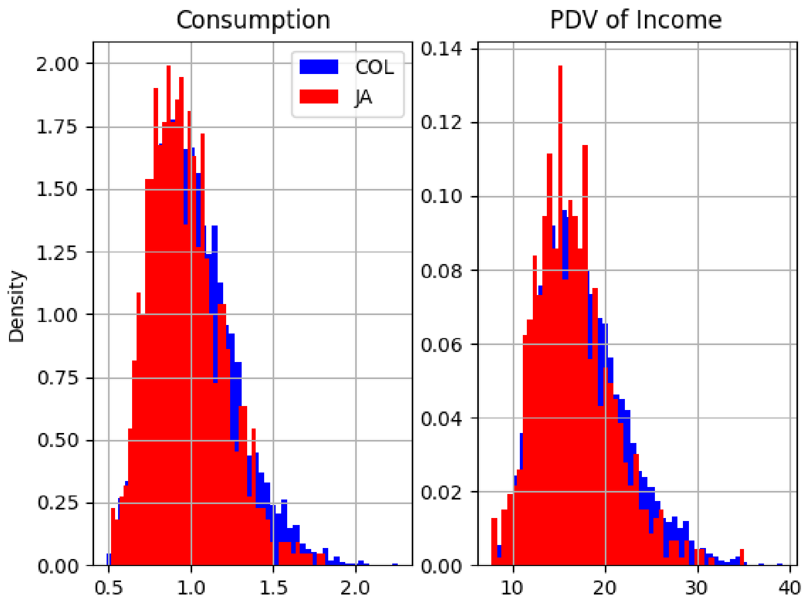
Notes: This Figure shows the effect on food consumption for employees of firms with +20, +30, and +40 workers.

Figure 8: Consumption Policy Functions by employment status



Notes: This Figure plots consumption as a function of current period assets for JA and COL workers (red and light) for the median productivity state for employed and unemployed people (left and right panel, respectively). Baseline calibration: $\beta = 0.96$, $\gamma = 2$ and $r = 0.02$

Figure 9: Distribution of Consumption and PDV of Income



Notes: This Figure plots the distribution of consumption and of the PDV of income for JA and COL workers as a function of γ.

Appendix A: Patience and Risk Aversion

The questions on the discount rate and on the certainty equivalent.

Imagine receiving as a gift an amount of money equal to your annual net income. Would you give up this gift in exchange for the same amount next year, increased by...

1. *1 % of your annual income*
2. *5 % of your annual income*
3. *10 % of your annual income*
4. *50 % of your annual income*
5. *100 % of your annual income*
6. *300 % of your annual income*

.

Imagine you are being offered a lottery ticket for which you have 50 % chance of winning your net annual income. Would you exchange the ticket for...

1. *5 % of your annual income*
2. *10 % of your annual income*
3. *25 % of your annual income*
4. *50 % of your annual income*
5. *75 % of your annual income*

Appendix B: Additional Tables and Figures

Table B1: What happens in case of unjust dismissal in JA vs COL

	Chart of Labour	**Jobs Act**
Who applies to:	Large Firms Workers	Large Firm Workers
Hiring Day:	Until March 6, 2015	March 7, onwards
Discriminatory Firing:	Reinstatement (i.e. Back in the Job)	Reinstatement
Disciplinary Firing:	Reinstatement	No Reinstatement
Other Terminations:	Probable Reinstatement	No Reinstatement
Monetary Insurance:	Arrears + Fine	Monetary Compensation (Function of Tenure)

Notes: This Table summarizes the main difference between *Jobs Act* and *Chart of Labour* workers.

Table B2: The Effect on Labour Supply

	(1)	(2)	(3)	(4)	(5)
Panel A: North of Italy					
Jobs Act	1.31***	1.10**	1.08**	0.09	-0.36
	(0.47)	(0.47)	(0.46)	(0.86)	(0.88)
Individual Net Income (€100)	0.62***	0.60***	0.60***	0.60***	0.64***
	(0.04)	(0.04)	(0.04)	(0.04)	(0.08)
Observations	2105	2105	2105	2105	414
Panel B: South and Centre					
Jobs Act	0.62	0.60	0.57	-0.78	-1.24
	(0.65)	(0.64)	(0.64)	(0.99)	(1.02)
Individual Net Income (€100)	0.69***	0.68***	0.67***	0.68***	0.83***
	(0.05)	(0.05)	(0.05)	(0.05)	(0.12)
Observations	2105	2105	2105	2105	414
Controls	✓	✓	✓	✓	✓
Job Sector FE		✓	✓	✓	✓
Risk Aversion			✓	✓	✓
Tenure				✓	≤ 5

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B3: Income Differences between JAs and COLs

	Income Level (€100)		Log of Income (× 100)	
	(1)	(2)	(3)	(4)
Jobs Act	-0.07	0.04	-0.84	-0.66
	(0.45)	(0.45)	(3.32)	(3.38)
Male	3.99***	3.53***	28.45***	25.76***
	(0.18)	(0.39)	(1.31)	(2.98)
Married	0.74***	1.27**	4.26***	6.23
	(0.21)	(0.55)	(1.57)	(4.11)
Family Size	-0.36***	-0.20	-2.80***	-1.74
	(0.09)	(0.17)	(0.68)	(1.29)
N. of Children	0.71***	0.86***	4.82***	5.85**
	(0.13)	(0.32)	(0.90)	(2.37)
Between 25 and 29	-1.51***	-0.71	-5.33**	0.14
	(0.36)	(0.76)	(2.65)	(5.53)
Between 30 and 39	-1.04***	-0.43	-3.29	0.69
	(0.27)	(0.72)	(2.00)	(5.33)
Between 40 and 49	-0.03	0.16	1.40	2.45
	(0.24)	(0.80)	(1.69)	(5.98)
Years of Education	0.63***	0.68***	4.10***	4.80***
	(0.03)	(0.07)	(0.20)	(0.49)
Observations	3505	703	3505	703
Job Sector	✓	✓	✓	✓
Tenure FE	All	≤ 5 years	All	≤ 5 years

Standard errors in parentheses

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B4: Further Heterogeneity

	By Age		By Civil Status		By City Size		By Education	
	(1)	(2)	(3)	(4)	(5)	(6)	(7)	(8)
Jobs Act	-2.94***	-12.75***	-2.89***	-11.81***	-3.14***	-12.26***	-2.36**	-8.70**
	(0.86)	(3.72)	(0.86)	(3.83)	(0.81)	(3.31)	(0.92)	(3.59)
Jobs Act x Above 45	0.72	4.42						
	(1.44)	(5.88)						
Above 45	2.26***	8.14***						
	(0.60)	(2.38)						
Jobs Act x Married			-0.25	-1.02				
			(1.34)	(5.61)				
Married			2.92***	11.12***				
			(0.57)	(2.24)				
Jobs Act x Big City					-1.41	-7.11		
					(1.39)	(6.04)		
Big City					-0.42	-1.03		
					(0.62)	(2.41)		
Jobs Act x College							-0.42	-2.90
							(1.32)	(5.61)
College							-4.24***	-18.07***
							(0.63)	(2.55)
Observations	2228	2228	2228	2228	2228	2228	2228	2228

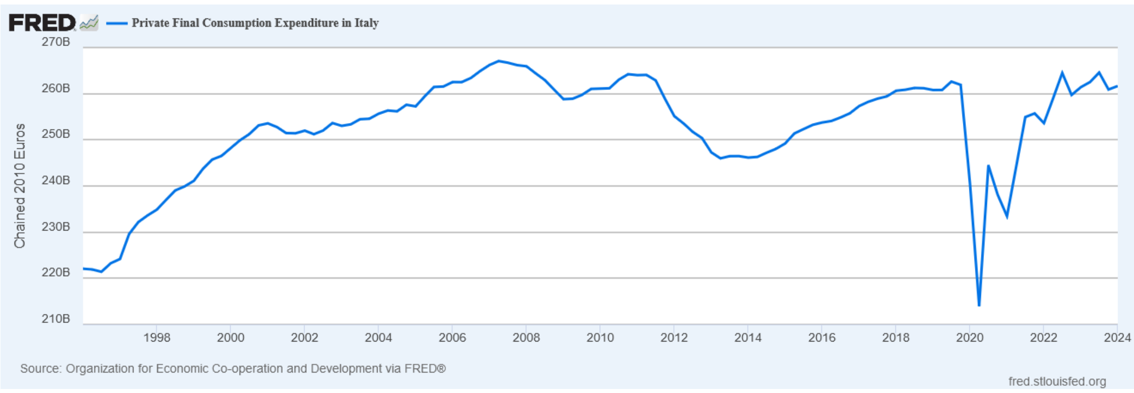
Notes: This Table show the estimates of (2) where the *Jobs Act* dummy is interacted with the heterogeneity dimension of interest. The dependent variable in odd columns is the food consumption as a share of family income, and its log in even columns. The sample only includes large firm workers in the North. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The controls are the same as in the previous table. The sample workers of firms with at least 20 employees in the North of Italy. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Figure B1: Proposed Referendum on the Jobs Act

«Do you want the repeal of Legislative Decree no. 4 March 2015? 23, containing "Provisions regarding permanent employment contracts with increasing protection, in implementation of law 10 December 2014, n. 183" in its entirety?".

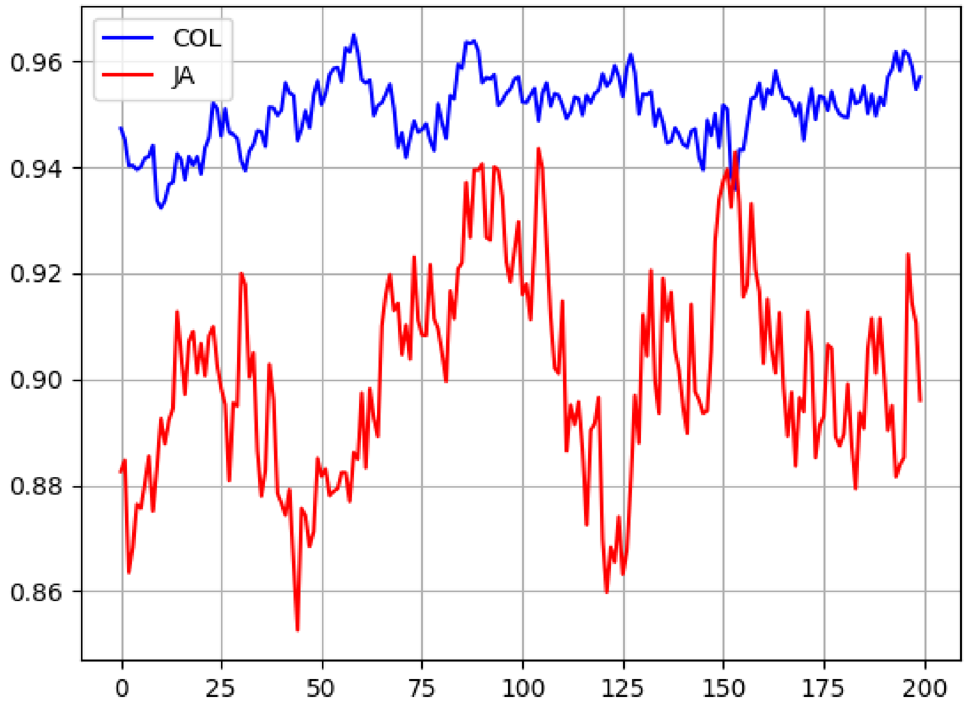
Notes: This picture shows the text (translated in English) of the referendum promoted by Maurizio Landini, today's head of CGIL. The union collected signatures for a petition to propose a referendum to repeal the *Contratto a Tutele Crescenti*. At the moment this paper is written, the petition is under the supervision of the Italian Court of Cassation, which, according to the law on popular referenda, is responsible for counting the signatures, making sure they are genuine, and evaluating the admissibility of the quest from a Constitutional point of view.

Figure B2: Aggregate Private Final Expenditure in Italy



Notes: This picture shows the time series of aggregate private consumption at the quarterly level. Aggregate consumption in Q4 2024 was still below the peak reached in Q2 2007, a quarter prior to the Great Financial Crisis. *Source:* St. Louis Federal Reserve.

Figure B3: Simulated Income for JA and COL workers



Notes: This picture shows the simulated incomes of JA and COL workers over time in the model.

Appendix C: Blundell Pistaferri Preston Imputation

In their seminal work, Blundell et al. (2008) impute nondurable consumption in the PSID from CEX. Their methodology consists of estimating a demand equation of food consumption - available in both surveys - using non-food consumption, demographic characteristics, and prices as RHS variables. Under some conditions, this demand function can be inverted, and non-food consumption might be pinned down by applying the estimated coefficients to the variables in PSID. A recent work by Patterson (2023) uses a similar approach.

In this Appendix, I impute nondurable consumption into PLUS by estimating a demand function using HBS. In practice, HBS is the counterpart of CEX in BPP, and PLUS is the counterpart of PSID in BPP. Table D1 shows the summary statistics of HBS, while Table D2 the results of the imputation (OLS estimation results of (3). Finally, Figure C1 compares the distribution of the actual log of spending (right panel) with the imputed (left panel).

Table C1: Summary Statistics of HBS

	Mean	Sd
Age 18-34	0.12	0.32
Age 35-64	0.88	0.32
Male	0.68	0.47
Married or Civil Union	0.54	0.50
South	0.29	0.45
Centre	0.21	0.41
North	0.50	0.50
Total spending (monthly)	2625.06	1346.57
Food-only spending (monthly)	470.06	274.42
Total/Food Consumption Ratio	6.75	4.08

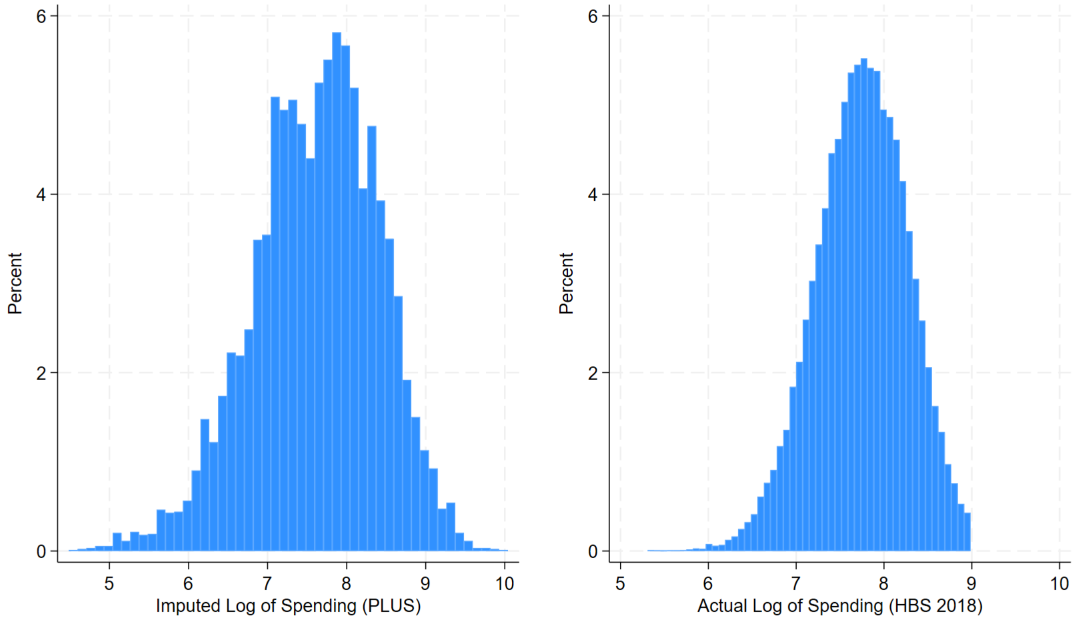
Table C2: Results of the Imputation

	(1)
Log of Nondurable Spending	0.70***
	(0.01)
Family size	0.10***
	(0.00)
Male	-0.00
	(0.00)
Age 35-64	0.06***
	(0.01)
Married or Civil Union	0.06***
	(0.01)
Italian citizenship	-0.07***
	(0.01)
Centre	-0.11***
	(0.01)
North	-0.19***
	(0.00)
Middle education	-0.07***
	(0.01)
High education	-0.18***
	(0.01)
Constant	0.48***
	(0.04)
Observations	52774
R^2	0.472

Standard errors in parentheses

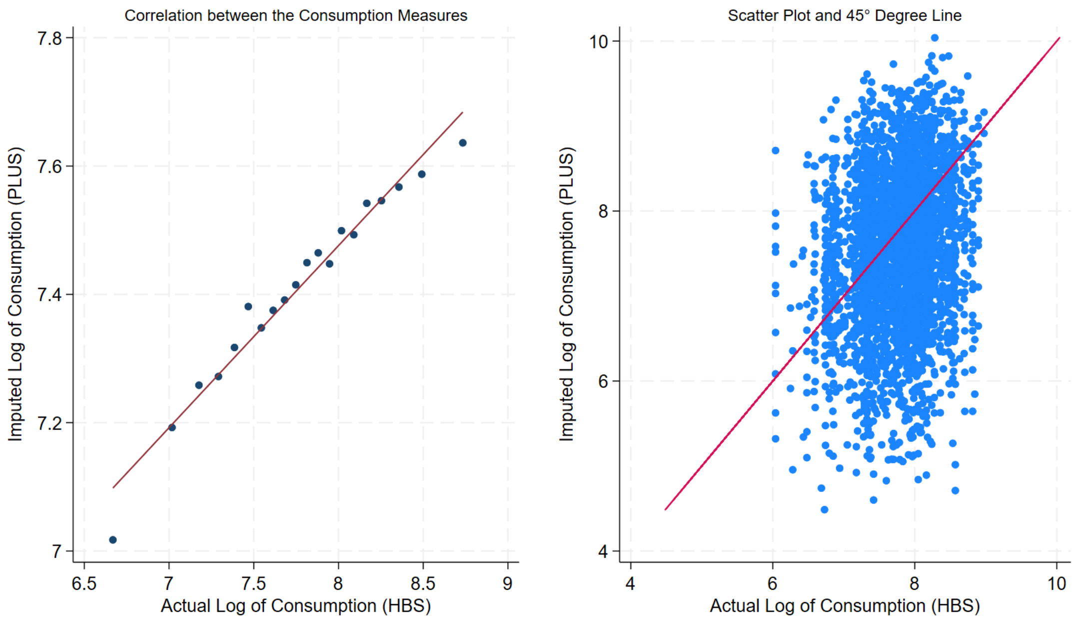
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Figure C1: Distribution of Imputed and Actual Consumption



Notes: This Figure shows the pooled distribution of the imputed measure of consumption, obtained using the BPP procedure.

Figure C2: Percentiles and Scatter Plot of the Distributions



Notes: The left panel shows the correlation between the percentiles of the actual distribution and the percentiles of the imputed distribution. The right graph shows the scatter plots with a 45-degree line.

Appendix D: Differences in Difference

I estimate the following equation via OLS:

$$Y_{it} = \beta Jobs\ Act_{it} \times Large_{it} + \delta Jobs\ Act_{it} + \gamma Large_{it} + \varphi \mathbf{X}_{it} + \lambda_t + \varepsilon_{it}$$

Where Y_{it} is the food share, the log of food spending, and weekly hours worked. The following Table shows the estimation results. The first two columns show result for the food share, the middle ones the log of food spending, and the last one the hours worked. Within each dependent variable, I perform two specifications: without and with controls. To limit selection problems as much as possible, I drop from the sample workers working for firms with between 10 and 14 employees.

Table D1: Difference-in-Differences Estimates

	Food Share		Log of Food		Hours Worked		Imputed
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Jobs Act x Large	-2.04**	-1.03	-10.86***	-7.31**	1.51*	0.74	-17.09***
	(0.92)	(0.89)	(3.63)	(3.45)	(0.80)	(0.78)	(5.74)
Large Firm	-1.37***	-1.00***	-2.35**	-0.78	3.74***	2.26***	-4.22**
	(0.29)	(0.29)	(1.07)	(1.07)	(0.23)	(0.22)	(1.70)
Jobs Act	-0.23	-0.38	0.38	0.82	-0.16	0.28	0.64
	(0.80)	(0.77)	(3.07)	(2.93)	(0.73)	(0.71)	(4.88)
Observations	11376	11363	10446	10435	11877	11859	10446
Controls		✓		✓		✓	.

Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Appendix E: Detailed Institutional Framework

The Italian labour market has historically been characterized by a certain degree of rigidity. In 1970, the COL was passed. The COL contains general rules on safety in the workplace, unionization, and, most importantly for the argument of the paper, on workers' reinstatement in case of unjust dismissal by the employer. Article 18, in its first version, stated that employees of large firms got reinstated in their jobs in all cases of unjust dismissal, ruled by a labour judge. This means that workers who were fired unlawfully by their employers were entitled to get their job back, plus all the forgone wages during the period of absence and a damage repayment by the employer.

This labour regulation made firing very hard and costly for employers. Indeed, judges in Italy have a strong pro-labour attitude (Del Punta 2010), and the overwhelming majority of labour cases are ruled in favor of workers. It also created a dual labour market: on one hand, there was a category of over-privileged workers, for whom being fired was virtually impossible; on the other hand the other workers with a much lower level of employment protection (and therefore a high level of employment risk).

Article 18 was rewritten in 2012. The new version of Article 18 covers all large firm employees hired until March 6, 2015. After the reform, which was named after Labour Minister Elsa Fornero[20], the compulsory reinstatement remained only in some circumstances of unjust dismissal, while it was abolished for others. In case of the latter ones, the employer only has to pay a fine to the worker, but is not obliged to reinstate her anymore. This new version of Article 18 gives large discretionary power to judges, who not only have to rule whether a firing is unlawful or not, but also have to decide on the reason for the unlawfulness. Given judges' pro-labour attitude, they typically motivate the unjust dismissals based on those circumstances for which reinstatement was compulsory, given that workers' most preferred option is to get the job back in most of the cases. Therefore, employment protection remained high in comparison with the other EU countries. This can be seen from Figure 2: the drop in the employment protection index in 2012 is modest compared to the drop in 2015/2016 due to the *Jobs Act*.

Given that the 2012 reform failed at increasing flexibility in the labour market, considered in the mid 2010s an economic policy priority in Italy, the Government led by Prime Minister Matteo Renzi passed the *Jobs Act*, applied to large firm employees starting from March 7th 2015. Article 18 still applies to workers hired before this day, still by large firms. The major novelty is that, in case of unjust dismissals, workers get reinstated in their job only in case of discriminatory firing, i.e., in a circumstance in which a worker is fired because of her gender, race, sexual orientation, political and religious beliefs. In all other cases of unjust dismissal ruled by the judge, the employer is not forced either to reinstate or to hire the worker back, but only to pay her a fine proportional to her tenure within the firm. This intervention had two strong effects. First, it eliminated or significantly reduced judges' discretionary power, given that now, in almost all circumstances, the employer doesn't have to reinstate the worker anymore. Second,

[20]This reform shall not be confused with the Fornero pension reform, another reform Minister Fornero is famous for.

it reduced, or made extremely much more predictable, firing costs for firms. Indeed, firms know perfectly the cost they would incur in case of "negative" ruling, given that the amount of money payable to workers can be computed with certainty using the formula provided by the law. Looked from the other side of the coin, this led to a reduction in employment protection. The *Jobs Act* is only part of one of the packages of reforms undertaken by PM Renzi's Government to reform Italy's labour market. In December 2014, the Italian Parliament passed a law that allowed the Government to reform the labour market without further future approval by the Parliament (this kind of law is called *legge delega*, and it's sometimes justified by the complexity of the reforms, which is therefore best dealt with by experts from the Government). Other reforms dealt with the regulation of mass layoffs, safety and privacy in the workplace, and unemployment insurance. Two concerns arise from this point. The first is the relation between the JA and these other reforms, and the second is whether both firms and workers expected such a reform. Regarding the former, these other pieces of legislation involve all workers regardless of the size of the firm they work for and the day on which they were hired. They are therefore uncorrelated with the JA and might affect consumption only by means of a fixed effect. Regarding the latter, the passing of the *legge delega*, of course, created expectations of a reform of Article 18. However, the exact content of the reform was intentionally kept secret by the Government and was released on March 6, 2015. Therefore, there was no scope for manipulation around March 7, 2015, neither from the workers nor the firm's side.

Appendix F: Model Appendix

Recursive Formulation and Numerical Details

There are two types of consumers: JA and COL. They are homogeneous in their preference parameters and productivity, but different in their employment risk: the former have a higher job separation s_j (with $j \in \{ JA, COL\}$ and $s_{JA} > s_{COL}$). The job finding rate f is the same for both. For both JAs and COLs, the Bellman Equations for the employed and unemployed workers are the following:

$$V_{jt}^{empl}(a,p) = \max_{c,a'} u(c) + \beta \left\{ s_j \mathbb{E}\left[V_{jt+1}^{unem}(a',p')|p\right] + (1-s_j)\mathbb{E}\left[V_{jt+1}^{empl}(a',p')|p\right] \right\}$$

$$a' + c = (1+r)a + w \times p$$

$$V_{jt}^{unem}(a,p) = \max_{c,a'} u(c) + \beta \left\{ (1-f)\mathbb{E}\left[V_{jt+1}^{unem}(a',p')|p\right] + f\mathbb{E}\left[V_{jt+1}^{empl}(a',p')|p\right] \right\}$$

$$a' + c = (1+r)a + b \times p$$

The FOCs of the two Bellman equations are:

$$u'(c_t) \geq \beta \left\{ s\mathbb{E}\left[V_{a',t+1}^{unem}(a',p')|p\right] + (1-s)\mathbb{E}\left[V_{a',t+1}^{empl}(a',p')|p\right] \right\}$$

$$u'(c_t) \geq \beta \left\{ (1-f)\mathbb{E}\left[V_{a',t+1}^{unem}(a',p')|p\right] + f\mathbb{E}\left[V_{a',t+1}^{empl}(a',p')|p\right] \right\}$$

The Envelope Conditions are:

$$(1+r)u'(c) = V_{a,t}^{empl}$$

$$(1+r)u'(c) = V_{a,t}^{unem}$$

I find the consumption policy functions for JA and COL workers - in employment and unemployment - via the endogenous grid point method (see Carroll 2006). After finding the policy functions, using the transition matrix for employment Π_{JA}, I simulate a series of 0 (if unemployed) and 1 (if employed) for N_{JA} households and 200 periods. Then, I compute the mean difference between the log of consumption of JA and COL workers. Finally, I perform standard OLS with the outcome of the simulation, controlling for wealth, tenure, and the present discounted value of income, computed as the discounted sum of all incomes earned by the individual in the model, using the interest rate r for discounting.

Epstein-Zin Preferences

The baseline model assumes isoelastic preferences, in which the *EIS* and the relative risk aversion are each other's reciprocals. In this appendix, I assume recursive preferences in the fashion of Epstein and Zin (1989), in which intertemporal substitution and risk aversion are governed by two separate parameters: the *EIS* θ and the risk aversion γ. The model with recursive preferences is the following:

$$\max_{c_j, a_{jt+1}} V_j = \left[(1-\beta) c_{jt}^{\frac{\theta-1}{\theta}} + \beta \mathbb{E}_t \left[V_{jt+1}^{1-\gamma} \right]^{\frac{\theta-1}{\theta(1-\gamma)}} \right]^{\frac{\theta}{\theta-1}}$$

Subject to the same constraints as in Section 7. The Table below shows the calibration, while the next table shows the estimated coefficients of equations (5) and (6).

Table F1: Calibration - EZ Preferences

Parameter	Value	Moment	Source	Model	Data
θ	0.5	-	-	-	-
γ	5	-	-	-	-
β	0.95	Wealth to Income Ratio	SHIW 2016	4.48	4.44
s_{COL}	0.05	COLs Years of Tenure	LFS 2018	18.11	18.31
s_{JA}	0.10	Unemployment Rate	FRED-FED 2016-2019	10.32	10.75
f	0.50	Std Unemployment Rate	FRED-FED 2016-2019	0.85	0.87

Table F2: Results - EZ Preferences

Dependent Variable:	Log of Consumption × 100			Wealth / Income	
	(1)	(2)	(3)	(4)	(5)
Jobs Act	-5.310***	-2.301***	-6.511***	-0.204***	0.123***
	(0.113)	(0.106)	(0.383)	(0.013)	(0.012)
PDV of Income		1.652***	1.630***		0.153***
		(0.004)	(0.005)		(0.001)
Initial Wealth		0.026**	0.041***		0.123***
		(0.011)	(0.012)		(0.001)
Jobs Act x PDV of Income			0.149***		
			(0.013)		
Jobs Act x Init. Wealth			-0.104***		
			(0.032)		
Observations	1000000	1000000	1000000	1000000	1000000

* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Two-account model

Following Auclert et al. (2024) and Graves (2025), I also extend the model by allowing consumers to save in two accounts: liquid and illiquid. The illiquid account yields a higher return r_a than the liquid account r_b, but each consumer has an i.i.d. probability χ with which they can tap-in or tap-out the illiquid account. When they adjust their illiquid account, consumers also pay a transaction cost equal to k percentage points of the absolute value of the difference between the wealth they decide to hold next period and the current wealth. The rest of the model is the same as before. The recursive formulation of the model follows below. The utility function is CRRA with risk aversion γ. In the Bellman equations, a and b represent the holdings of the illiquid and the liquid account, respectively. In this model, there are 5 states: i) whether the consumer can adjust or not, ii) the holdings of the illiquid account, iii) the holdings of the liquid account, iv) employment status, and v) idiosyncratic productivity. I treat the first one as a state just for numerical convenience, but since χ is iid, it can be disregarded as a state. When the individuals can adjust, they face the following Bellman equation, with the respective budget constraints:

$$V_{jt}^{empl,adj}(a,b,p) = \max_{c,a',b'} +\beta\chi \left\{ s_j \mathbb{E}\left[V_{jt+1}^{unem,adj}(a',b',p')|p\right] + (1-s_j)\mathbb{E}\left[V_{jt+1}^{empl,adj}(a',b',p')|p\right]\right\} +$$

$$\beta(1-\chi)\left\{ s_j \mathbb{E}\left[V_{jt+1}^{unem,noadj}(a',b',p')|p\right] + (1-s_j)\mathbb{E}\left[V_{jt+1}^{empl,noadj}(a',b',p')|p\right]\right\}$$

$$a' + b' + k\ |a-a'| = (1+r_a)a + (1+r_b)b + y_t(empl)$$

$$V_{jt}^{unem,adj}(a,b,p) = \max_{c,a',b'} +\beta\chi \left\{ f \mathbb{E}\left[V_{jt+1}^{empl,adj}(a',b',p')|p\right] + (1-f)\mathbb{E}\left[V_{jt+1}^{unempl,adj}(a',b',p')|p\right]\right\} +$$

$$\beta(1-\chi)\left\{ f \mathbb{E}\left[V_{jt+1}^{eml,noadj}(a',b',p')|p\right] + (1-f)\mathbb{E}\left[V_{jt+1}^{unem,noadj}(a',b',p')|p\right]\right\}$$

$$a' + b' + k\,|a-a'| = (1+r_a)a + (1+r_b)b + y_t(empl)$$

When they cannot adjust, they face the following equations:

$$V_{jt}^{empl,noadj}(a,b,p) = \max_{c,a',b'} +\beta\chi \left\{ s_j \mathbb{E}\left[V_{jt+1}^{unem,adj}(a',b',p')|p\right] + (1-s_j)\mathbb{E}\left[V_{jt+1}^{empl,adj}(a',b',p')|p\right]\right\} +$$

$$\beta(1-\chi)\left\{ s_j \mathbb{E}\left[V_{jt+1}^{unem,noadj}(a',b',p')|p\right] + (1-s_j)\mathbb{E}\left[V_{jt+1}^{empl,noadj}(a',b',p')|p\right]\right\}$$

$$c + b' = (1+r_b)b + y_t(empl), \quad a' = (1+r_a)a$$

$$V_{jt}^{unem,noadj}(a,b,p) = \max_{c,a',b'} +\beta\chi \left\{ f \mathbb{E}\left[V_{jt+1}^{empl,adj}(a',b',p')|p\right] + (1-f)\mathbb{E}\left[V_{jt+1}^{unempl,adj}(a',b',p')|p\right]\right\} +$$

$$\beta(1-\chi)\left\{ f \mathbb{E}\left[V_{jt+1}^{eml,noadj}(a',b',p')|p\right] + (1-f)\mathbb{E}\left[V_{jt+1}^{unem,noadj}(a',b',p')|p\right]\right\}$$

$$c + b' = (1+r_b)b + y_t(empl), \quad a' = (1+r_a)a$$

Numerical Solution and Calibration. I obtain the policy functions of the model via VFI. Then I set the transaction costs k to be 0.10, following Graves (2025). Then I choose β and χ to match the illiquid wealth to income and the liquid wealth to income ratio, respectively. I classify assets into the liquid and the illiquid accounts following Kaplan et al. (2018): the liquid account contains cash, government and corporate bonds, while the illiquid account contains stocks, managed funds, and foreign-denominated assets. The discount factor is 0.92. On the other hand, the parameter of adjustment is 0.24, meaning that individuals adjust their illiquid account once every for year on average. The interest rate on the illiquid account is set at 10%, while the interest rate on the liquid account is -2%. Neither the liquid nor the illiquid account can be shorted. The parameters of the transition matrices are calibrated in the same way as before. To reduce computation time, I set the grid size to 50 and the maximum of the grid to 10. Table F3 shows the calibration of the model, while Table F4 shows the differences in the wealth-to-income ratio for the liquid, the illiquid, and the total accounts. JA workers have a higher wealth-to-income ratio in general, but this difference is mostly explained by the liquid account: in other words, people prefer to save precautionarily with liquid wealth, a result that is consistent with Kaplan and Violante (2014). Figure F1 provides the graphical intuition: while the distribution of illiquid wealth for JA and COL almost overlaps, the distribution of liquid wealth for JA almost first-order stochastically dominates the one for COLs. Finally, Figure F2 depicts the distributions of consumption and all wealth, i.e., the sum of the liquid and the illiquid accounts.

Table F3: Calibration of the Two-Account Model

Parameter	Value	Moment	Source	Model	Data
β	0.92	Illiquid Wealth to Income Ratio	SHIW 2016	4.12	4.14
χ	0.24	Illiquid Wealth to Income Ratio	SHIW 2016	0.46	0.38
γ	1	-	Standard	-	-
k	0.10	-	Graves (2025)	-	-
r_b	- 0.02	-	-	-	-
r_a	0.10	-	-	-	-
s_{COL}	0.04	COLs Years of Tenure	LFS 2018	18.57	18.31
s_{JA}	0.13	Unemployment Rate	FRED-FED 2016-2019	10.29	10.75
f	0.50	Std Unemployment Rate	FRED-FED 2016-2019	0.73	0.87

Notes: The other parameters of the calibration, excluding the account grids, are the same as in the baseline model.

Table F4: Difference in the Wealth to Income ratios

%Δ Illiquid Account	%Δ Liquid Account	%Δ Total Wealth	%Δ Consumption
10.53	24.20	11.96	-6.78

Notes: These are the percentage differences of the wealth-to-income ratio.

Figure F1: Distribution of Illiquid and Liquid accounts

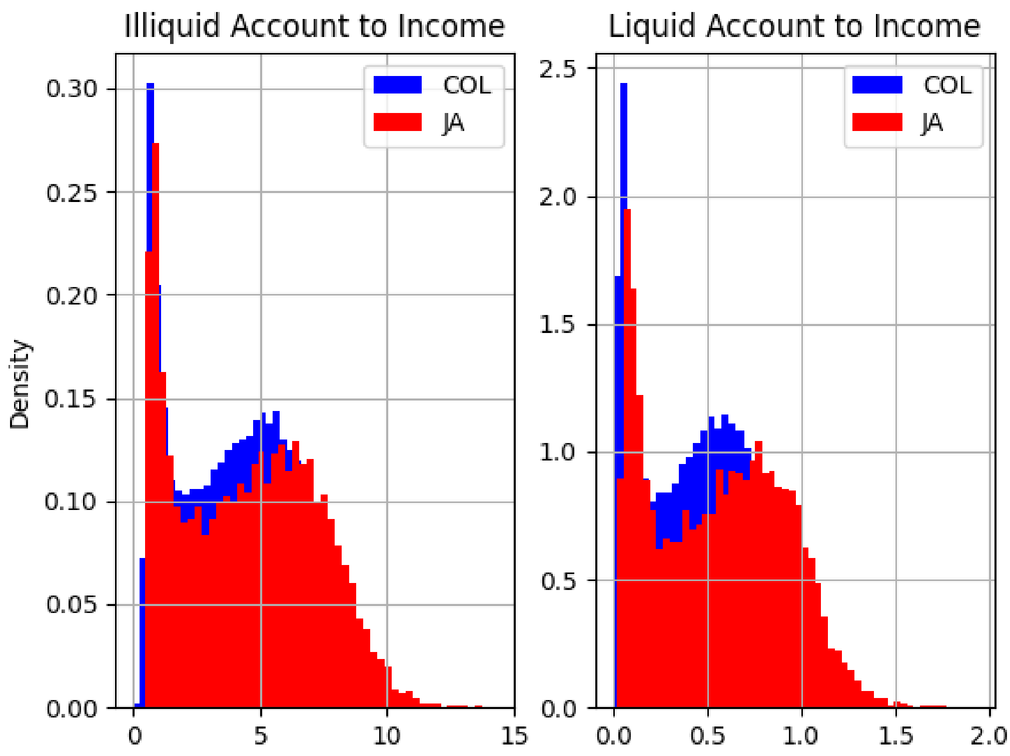
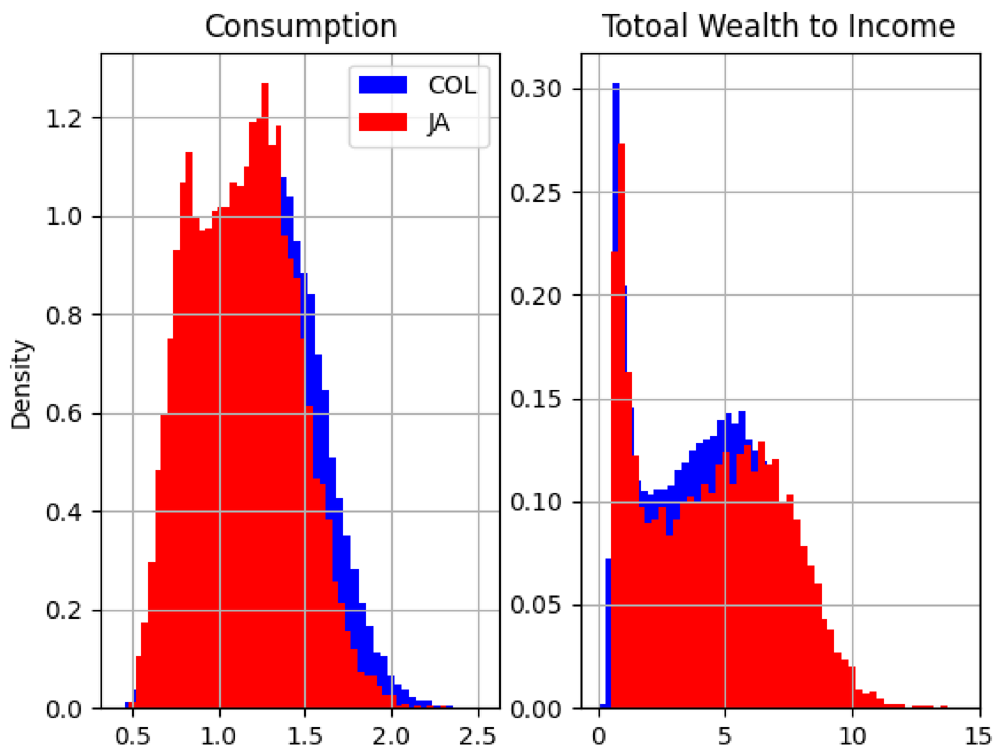


Figure F2: Total Wealth (Liquid + Illiquid)



Appendix G: Data Collection in detail

1. Participation, Labour and Unemployment Survey (PLUS)

The *Participation, Labour and Unemployment Survey* (PLUS) is a representative cross-section of the Italian population aged between 18 and 75 years of age. Access to the data might be requested on the INAPP website: https://www.inapp.gov.it/.

It contains demographics, information on the number of employees of the firms that workers work in, and information on whether they were hired before or after March 7, 2015. I include in the main analysis only workers of the private sector employed in large firms in 2018 with an open-ended contract. Below is a list of variables I use in the analysis:

- **Food Consumption to Family Income Ratio:** The percentage of family income spent on food consumption at home.
- **Family Income:** Average monthly net family income.
- **Food Consumption:** Food Consumption to Family Income Ratio $\times$ Family Income (for the imputation)
- **Tenure:** How many years the respondent has been working for the current employer (in bins):
  1. less than 5 years
  2. 5 - 10
  3. 10 - 14
  4. 15 - 19
  5. 20 - 29
  6. 30 - 39
  7. 40 - 49
  8. more than 50
- **Age** (in bins):
  1. 18 - 24
  2. 25 - 29
  3. 30 - 39
  4. 40 - 49
  5. 50 - 64

2. Household Budget Survey (HBS)

The *Household Budget Survey* (HBS) is a yearly repeated cross-section that reports information on household spending composition. It details durable and nondurable consumption, as well as demographic information on the composition of the household. It's publicly available at the ISTAT website https://www.istat.it/. HBS is available from 2013 to 2024. I use the 2019 wave to impute consumption into PLUS following Blundell et al. (2008) and Patterson (2023). The imputation procedure, and its performance, is explained in Appendix C, which also contains the summary statistics. In the procedure, I use the following variables:

- **Food Consumption:** Average Monthly Spending of the Household on at-home food
- **Total Consumption:** Average Monthly Spending of the Household on food, other non-durables, and durables

3. Labour Force Survey (LFS)

The *Labour Force Survey* (LFS) is a quarterly survey containing detailed information on workers' demographics, employment status, tenure, and whether they work for a big or a small firm (above or below 15 employees). There are two versions of LFS: the

The LFS data can be easily downloaded from the ISTAT website: https://www.istat.it/.

I use all quarterly waves from 2008 to 2019, and I restrict the sample to include working-age individuals hired in the private sector with an open-ended contract. My sample includes almost 1.3 million pooled individual-quarter observations. I use the following variables:

- **Large:** Dummy equal to 1 if the individual works in a firm with at least 15 employees
- **Jobs Act:** Dummy equal to 1 if the individual was starting from March 7, 2015
- **Tenure:** The time difference (in years) between the current year and the year the respondent started her current occupation

Table G1: Summary Statistics

	Mean	Sd	Min	P50	Max
Jobs Act	0.07	0.25	0.00	0	1
Large Firm	0.59	0.49	0.00	1	1
Tenure	14.31	10.63	0.00	12	50

4. Survey of Household Income and Wealth (SHIW)

The *Survey of Household Income and Wealth* (SHIW) is a biannual survey conducted by the Bank of Italy. Data and documentation are available at the following website: https://www.bancaditalia.it/.

I use the dataset to calibrate the model to match the wealth-to-income ratio. The SHIW contains detailed information on demographics, income, wealth, and consumption. There are two versions of the SHIW dataset: the historical dataset and the yearly waves. I use a combination of both, and I use the 2016 cross-section, which surveys around 8k households and 16k individuals. I restrict the sample to employees working in the private sector with an open-ended contract. This yields a sample size of around 2k individuals. The Table below shows the descriptive statistics of my sample. The variables I use are the following:

- **Liquid Wealth**: Bank Accounts + Government Bonds + Corporate Bonds
- **Illiquid Wealth**: Stocks + Managed Funds + Foreign Assets + Real Estate
- **Income**: Labour Income + Government Transfers
- **Liabilities**: Loans + Mortgages + Other Debts
- **Total Wealth to Income**: Illiquid Wealth / Income
- **Marginal Propensity to Consume**: the answer to the following question (used in Jappelli and Pistaferri 2014, 2020 and Auclert et al. 2024):

> *"Imagine you unexpectedly receive a reimbursement equal to the amount your household earns in a month. How much of it would you save, and how much would you spend? Please give the percentage you would save and the percentage you would spend."*

Table G2: Summary Statistics of 2016 SHIW

	Mean	Sd	P50	Min	Max
Liquid Wealth	12495.80	50698.04	3699	0	1945613
Illiquid Wealth	155868.05	233304.64	121000	0	5162034
Liabilities	11525.73	35629.43	0	0	476000
Income	32912.07	19706.30	27952	802	213438
Total Wealth to Income	4.44	4.79	4	0	73
Marginal Propensity to Consume	0.49	0.34	0	0	1