Employment Protection and Consumption: Evidence from Italy

Nello Esposito*
University of Naples Federico II

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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 on or after March 7, 2015. 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 between 9-10% less and work one additional hour per week compared with workers hired before March 7, 2015. The effect is stronger among young individuals, those living in Northern Italy and with lower income. Finally, I show that a variant of the Bewley–Huggett–Aiyagari model, augmented with employment risk heterogeneity and calibrated to Italian labour market flows and reasonable risk aversion, closely matches the empirical result on consumption, and that risk accounts for most of the effect.

Keywords: Jobs Act, Precautionary Saving, Employment Risk, Labour Market Reform
JEL Classification: D15, E21, J28, G51

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1 Introduction

People deeply value the stability of their jobs (Baghai et al. 2023). Rothwell and Crabtree (2019) surveyed a representative sample of U.S. workers, 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 labor market reforms aimed at increasing employment and enhancing firm productivity by reducing firing costs 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 labor supply. To this purpose, I leverage on the *Jobs Act* (JA), an Italian labor market reform that applies only to workers hired after March 7, 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 one-time compensation payment proportional to their tenure. The reform therefore significantly reduced job security for the latter because 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 and provides detailed information on demographic characteristics, food consumption, weekly hours worked, employees' firm size, and whether individuals have been hired before or after March 7, 2015. This last piece of information allows me to identify private sector workers, working for large firms, whose contract is regulated by the *Jobs Act* or by the *Chart of Labour*.

My baseline finding is that JA workers spend almost 6% less on food than COL workers and work an additional hour per week. While the former result is consistent with precautionary saving and buffer-stock models (Jappelli and Pistaferri 2017, Carroll 1997, Deaton 1991), the increase in working hours, on the other hand, suggests that workers react to higher employment risk also by raising their labor supply on the intensive margin. To estimate the impact of the *Jobs Act* on broader measures of spending, I propose two strategies. I first adopt an approach similar to Broda and Parker (2014) in which I scale the food share of consumption (the ratio of at-home food spending over family income) by multiplying it by the ratio of total spending to food spending, as reported in the National Accounts and the *Household Budgetary Survey* (HBS) - the Italian counterpart of the CEX in the United States - produced by the Italian Statistical Agency (ISTAT).

Alternatively, following the approach by Blundell et al. (2008) and Patterson (2023) - BPP from now on - I impute consumption from HBS into PLUS, finding a reduction in imputed spending of around 9-10%. The magnitude of the effect is in line with Clark et al. (2022) and is qualitatively consistent with the effect aggregate uncertainty 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 incomes of JA and COL workers are not statistically different, controlling for the same variables used in the main analysis. Second, I run the same analysis 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 similarly 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 neither by workers nor by firms and no anticipatory effects. Finally, I exclude employees working for firms with less than 20 employees. The rationale behind this test is that marginal firms, such as those with 14 employees, might have an incentive to hire an additional worker to benefit from the JA, whereas firms well above the threshold would 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, consistently 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 significant only among workers who report the presence of a local trade union representation at their workplace.

On a theoretical ground, higher employment risk impacts consumption via two channels.

On the one hand, higher employment risk means a more volatile income and a stronger precautionary saving motive. This is the channel I want to measure the importance of in this paper. On the other hand, a higher job separation rate reduces the present discounted value of income (PDV), reducing consumption as well. 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's impossible to disentangle empirically these two mechanisms.

I therefore rely on a variation of the Bewley–Huggett–Aiyagari model (Bewley 1986, Huggett 1993, Aiyagari 1994) - from now on BHA -, in which there are two types of workers with ex-ante heterogeneity in employment risk: safe and risky workers. The latter mimic JA workers and face a higher employment risk than the former, that mimic COL workers instead. I calibrate the model to match Italy's job market flows between 2016 and 2019 and use standard preference parameters: CRRA utility function with risk aversion of 2 and discount factor of 0.96 annually. After solving the model, I use the policy functions to simulate a panel of consumption, income and wealth for 20,000 individuals and 200 periods. The model-predicted average consumption difference between JA and COL workers is 8.2%, reasonably close to the empirical results outlined above. I then run a regression on simulated data of (log) consumption against the *Jobs Act* dummy, assets and lifetime income, finding an effect of sole employment risk of around 6%, therefore explaining more than two thirds of the total effect of the JA on consumption.

Finally, I perform a sensitivity analysis in the fashion of Kaplan and Violante (2010, 2014) and Jappelli and Pistaferri (2025) varying the relative risk aversion parameter used in the model. I find that the model is able to match the empirical analysis using values in line with ones used in the HANK literature (see Kaplan and Violante 2022 and Auclert et al. 2024), while it gets farther away from the empirical counterpart once higher and implausible values for risk aversion are used.

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 rather challenging for two reasons. The first reason is that consumption growth volatility, the term governing prudence in the Euler Equation (see Fagereng et al. 2017) is unobservable. Therefore, income risk is used as a proxy, assuming income is the main source of risk individuals face[2]. In turn, Guiso et al. (2002) argue that income risk depends

[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

on the distribution of future earnings, the distribution of unemployment benefits and the probability of job loss (i.e., employment risk). While previous natural experiments bundle employment and income risk together (see Fuchs-Schündeln and Schündeln 2005), this paper specifically differentiate between employment risk (i.e. being under the JA and the COL), and the remaining components of income risk, that mostly depends on idiosyncratic productivity risk.

The reason why is important to specifically look at employment risk, as opposed to income risk more in general, is the need to separate the effect of exogenous shocks happening to workers (like unemployment) from the endogenous response to them (like deciding to stay in or out the labour market). To this regard, Low et al. (2010), with the help of a life cycle model with endogenous employment participation and search frictions, show that failing to account for job mobility (so 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 risk aversion is the most prominent example.

As summarized by Jappelli and Pistaferri (2017), three approaches have been followed by 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 around 1 and 2, corresponding to a prudence parameter between 2 and 3. Using mortgages rates notches as a quasi-experimental variation in interest rates, Best et al. (2019) estimate a 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 (constant relative risk aversion) preferences and idiosyncratic income risk, and estimate the prudence coefficient (jointly with the discount rate) by mini-

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 like GDP, interest rate and inflation risk. Idiosyncratic factors explain about 75% of the variability in consumption risk. In a 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.

[3] Assuming CRRA (constant relative risk aversion) 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.

mizing the distance between the average of the simulated and actual life cycle consumption profiles. Their main result is a risk aversion of 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).

More recently, Jappelli and Pistaferri (2025), using the covariance restrictions dictated by the buffer-stock model, show that households revise their target wealth in response to permanent income shocks approximately by 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 the preferences and for the income process[4].

Graves (2025) finds that higher subjective unemployment risk induces an increase in overall savings, but a reduction in illiquid savings. 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 on a natural experiment and not on subjective expectations of job loss, that might be endogenous to unobservable characteristics like risk aversion, health status and ability. It might be, indeed, that a worker expects to be fired any time soon because her productivity has persistently declined in the last months.

Then, natural experiments have been used. Kantor and Fishback (1996) exploit the introduction of workers' compensation in the United States the 1910s, finding that the introduction of accident insurance decreased households' savings by 25%[5]. 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 self-selection underestimates the strength of precautionary saving by a 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. 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

[4]They elicit target wealth from a direct question of 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), that use self-reported measures of income uncertainty.

[5]There is another strand of the literature that, instead of assuming incomplete markets, tries to microfoud 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.

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, characteristics. In addition, as Figure 2 shows, the 2015 reform entailed a much larger reduction in employees' protection than the 2012 one[6]. 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 on a time-based discontinuity in employment protection to empirically assess the impact of employment risk on consumption. Furthermore, my paper connects the three approaches outlined above from an operational point of view: I indeed rely on a structural model to replicate the causal effect of employment on consumption, finding the range of values of the risk aversion parameter that make the model-implied effect as close as possible to the empirical effect.

Finally, I contribute to the existing empirical literature leveraging on 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 rate. 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 also 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 show a battery of placebos and robustness checks. It also presents some further results in which I control for risk aversion. Section 7 explores the heterogeneity of the baseline results. Section 8 shows the results of the model. Section 9 concludes and draws a path for current research I am currently pursuing.

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

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 more 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, or 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[7]. 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 not insured against job displacement: i.e., the long-lasting income losses due to unemployment. This latter point has proven particularly important and has been highlighted by Jarosch (2023) and Caratelli (2022), 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 finds that Italy is among the countries with the highest and most persistent negative effects of job loss.

The Importance of Job Stability in Italy. The "job for life" (Ureta 1992) is considered a highly valuable asset in Italy, a country historically characterized by relatively low job mobility (Del Bono and Vuri 2011). Bertoni et al. (2023) document that individuals hired

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

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 Employment Protection Legislation (EPL) index across OECD countries. As a consequence of the reform, Italy's EPL score declined in 2015 and 2016. It is important to emphasize that neither the COL nor the JA regulate mass layoffs such as those resulting from firm downsizing or closures. Instead, both contracts 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. In addition, Figure 2 illustrates that, even after the 2015 reform, Italy remains one of the OECD countries with the highest levels of employment protection.

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 80% of them voted in favor of repealing the reform[8].

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 because it explicitly asks respondents whether they were hired before or after March 7, 2015, allowing me to precisely identify JA workers[9]. The 2018 wave of the survey additionally includes

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

[9] 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,

questions on risk propensity 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 select the sample as follows. First, I retain only individuals who are currently employed. I then exclude workers whose employer is a public organization (direct civil servants, employed in state-own 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 8,500 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 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

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.

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 straightforward 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 in food consumption and other types of consumption, therefore having:

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

Where C_F denotes food consumption, while C_O denotes "other" consumption. Dividing both sides by Y and multiplying and dividing C_O 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[10]) 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. As a result, the estimated coefficient for the log of food consumption is equivalent to the estimated coefficient for total log 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. I 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:

$$Y_{irst} = \alpha + \beta \; JobsAct_{irst} + \varphi \mathbf{X}_{irst} + \lambda_t + \theta_r + \gamma_s + \varepsilon_{irst} \tag{2}$$

In equation (2), the dependent variable Y_{irst} is either food consumption (measured as a share of income and in log-level) or the individual's average weekly labour supply, both sourced from the PLUS survey. The vector $\mathbf{X}_{irst}$ includes individual-level controls such as age (in bins), gender, marital status, family size and number of children. λ_t denotes time fixed effects, which absorb aggregate time shocks, γ_s and θ_r represent sector and region fixed effects, respectively, controlling for the fact that hiring patterns may vary across industries and regions, and for the unequal sectoral and dimensional distribution of firms across the North and the South of Italy. Finally, ε_{irst} is an idiosyncratic error term.

[10]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.

The coefficient of interest is β, which captures the average difference in the outcome variable between workers hired before and after March 7, 2015. 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 informal evidence supporting this assumption.

Why not a difference-in-differences (D-i-D)? Rather than estimating equation (2), one might consider a D-i-D strategy by exploiting variation in both hiring date and firm size (i.e., above vs. below the 15-employee threshold) as in De Paola et al. (2021) and Clark et al. (2022). However, I restrict the sample to large firms for several reasons, in line with the approach of Mistrulli et al. (2023), who collect the considerations on the *Jobs Act* made by Boeri and Garibaldi (2019).

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. This provision, coupled with government hiring subsidies introduced in the meantime (Boeri and Garibaldi 2019), likely influenced firm behavior, including wage setting. In fact, it might be that some workers would not have been hired - or would have been hired at different wage levels - absent the JA. Second, there are concerns related to self-selection in large or small firms on the basis of consumers' attitude toward risk[11]. 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[12]. On 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. Third, and lastly, the data at my disposal

[11] 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.

[12] The effect of risk aversion of current consumption is negative, trough 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. Risk-averse people would consume less anyways.

do not allow me to test for pre-trends, given that I just have one pre treatment period.

For these reasons, I restrict my analysis to workers in large firms to mitigate self-selection concerns from both the employer and employee sides[13].

5 Empirical Results

Food Consumption and Labour Supply. Table 3 reports estimates for the ratio of food consumption to family income. Column 1 presents a naive specification including only time fixed effects and no additional controls. The food share for JA employees is 2.30 percentage points lower than for COL employees, and this estimate is statistically significant at 1%. Given a 2014 benchmark food share of approximately 30%, this corresponds to a 7% reduction in the ratio of food consumption to family income.

Subsequent columns progressively add controls to address potential confounders. Column 2 includes demographic variables, with age brackets being particularly important given that younger individuals are more likely to be hired under the JA. The coefficient decreases in magnitude to 1.30 percentage points (corresponding to a 4% reduction in the food share) but remains statistically significant. Columns 3 and 4 add job sector and regional fixed effects respectively and do not change the estimated coefficient, suggesting that the only observable characteristic that potentially selects workers into the *Jobs Act* come from demographics, and in particular age.

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 6. Therefore, the saving rate is estimated to increase by approximately 10 percentage points (1.38×7).

Table 4 reports estimates using the log of food consumption (multiplied by 100) as the dependent variable. Qualitatively, the results are similar: food consumption for JA workers is about 6% lower than COL workers. This coefficient remains stable across all specifications that control for age.

The second outcome of interest is labour supply. Table 5 presents OLS estimates of weekly labour hours. The first column shows the baseline specification without demographic controls; the second includes demographics. Across specifications, JA employees work roughly one hour

[13] Table D1 in Appendix D nonetheless presents OLS estimates from a D-i-D specification for completeness.

more than COL employees in a typical week, corresponding to a normal five-day workweek.

The control vector doesn't include education, even though it's available in PLUS. The reason for this is that the age bins are quite large (especially the 50-64 one) and absorb almost all of the variation of education levels. Therefore, the age bins and the education levels are practically collinear. This, however, doesn't represent an econometric problem because the day on which the contract is signed is arguably independent on the education level. Nevertheless, if I perform the analysis only on the over 25 (and thus dropping individuals aged between 18 and 24, and for whom there are very few COL workers), I eliminate much of the correlation between the age bins and the education levels. Therefore estimates, controlling for the years of education results, are still significant both economically and statistically. I perform this robustness check in Appendix B in Table B3.

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 in general, 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 on a more comprehensive measure of consumption: total consumption. To get such a measure, I follow the apporach 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:

$$\log(f_{it}) = \alpha + \mu \mathbf{X}_{it} + \delta \log(c_{it}) + \varepsilon_{it} \tag{3}$$

In the above equation, f_{it} is food spending, $\mathbf{X}_{it}$ is a vector of demographic controls available in both surveys[14]. δ is the elasticity of food consumption against total nondurable spending. 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 6 show the estimated coefficient of equation (2), for the (imputed) log of consumption and the average propensity to consume respectively. 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 9%. This estimate is higher in magnitude

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

than the baseline on log of food consumption. This might 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 had, 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.

It's also worth remarking these results are not so distant from the previous ones: 9-10% reduction in BPP consumption against a 6% reduction in food consumption and a 10 percentage points reduction in the average propensity to consume. This suggest two things. The first is that the Cobb-Douglas approximation is, overall, not a bad one, at least on average. Second, that income differences between JA and COL workers are not big enough to influence the results, as I will more formally show in the next section.

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*[15]. 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' 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 qualitatively 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

[15] 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.

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. To 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 in 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.

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.

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 €3 (the unit is €100), equal to 0.33% 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 less 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 7, Table 8, and Table 9 show that the estimates are statistically indistinguishable from zero. This result is confirmed by the last three columns of Table 6, where only small firm workers are considered.

Another 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 10 and Table 11 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. Similarly, Table 12 reports the same regression for hours worked, and no effect is found.

Bigger Firms. Following the robustness checks in 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 labor 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 13) 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 estimated effect for 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 labor 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 nether 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.

Risk Aversion. 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:

$$Y_{irs} = \alpha + \beta Jobs\ Act_{irs} + \varphi \mathbf{X}_{irs} + \theta_r + \gamma_s + \rho_{irs} + r_{irs} + \varepsilon_{irs} \tag{4}$$

In equation (4), ρ_{irs} and r_{irs} are dummies representing different levels of patience and risk aversion, respectively. The trade-off for this richer specification is a reduction in sample size and time variation: these questions are only asked in 2018, yielding approximately 4,000 observations, or roughly half of the full food consumption sample. Additionally, since the analysis is cross-sectional, common shocks occurring in 2018 cannot be accounted for. OLS estimates of equation (4) are shown in Table 14. Columns 1 and 2 report the estimated coefficients for the food expenditure-to-income ratio, with and without controls, respectively. The β coefficient is very close to the baseline. This is true also for the log of food consumption (columns 3 and 4) and for labor supply (columns 5 and 6). Finally, column 3 of Table 6 controls for risk aversion and patience heterogeneity. The estimated coefficient remains practically unchanged, as well as its statistical significance.

In sum, adding controls for preference heterogeneity does not meaningfully alter the results. Because the sample is restricted to employees of large firms, selection concerns are minimal. These results also help rule out the possibility that more risk-averse individuals selected into treatment by timing their hiring, anticipating the *Jobs Act*.

7 Heterogeneity Analysis

Average results mask significant heterogeneity along several dimensions. Table 15 presents the results of regressions performed on sub-samples defined according to the column titles.

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 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.

Next, I examine heterogeneity by geographical area, dividing the sample into the North (the wealthiest and most industrialized part of the country) and the Centre-South (relatively underdeveloped). Interestingly, the effect is significant only for workers in the North. I interpret this result as follows: Northern Italy has a substantially higher GDP per capita—about 50% greater—compared to the South. Individuals in the North tend to have higher cash-on-hand levels and are therefore better equipped to smooth consumption. Furthermore, the stronger employment protection provided by the COL contract likely mitigates precautionary saving motives more effectively than the JA. In other words, people living in the North are the ones who have to lose from lower employment protection. Conversely, individuals in the South are more likely to be hand-to-mouth, i.e., constrained in their borrowing and consumption decisions. For such individuals, employment protection plays a more limited role in shaping saving behavior, since their precautionary savings are mainly motivated by the fear of hitting the liquidity constraint (Deaton 1991) in case of a bad income realization next period[16]. This explains why the effect is not statistically significant in the South[17].

The PLUS dataset also asks respondents whether their workplace has union representation. The effect is significant only for unionized firms. This result speaks to the literature investigating the importance of union representation at the firm level as an insurance mechanism. Kim et al. (2018) show 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's capacity to protect workers, maybe because it altered the relationship between employees and employers in a more favorable way to the latter.

[16] These people are not optimally choosing their consumption path through the Euler Equation, but are on a corner solution

[17] Table B6, Table B7 and Table B8 show the OLS estimates of an equation where the *Jobs Act* dummy is interacted with age bins, geographical and income dummies. They confirm that the effect is mostly concentrated among younger individuals, people living in the North of Italy and with incomes below the median.

8 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. From and 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 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 rules out layoffs, not covered by the *Jobs Act* regulation. The economy is populated by two types of consumers/workers: *Chart of Labour* workers and *Jobs Act* workers. Their preferences are defined by a CRRA utility function with the same relative risk aversion γ. They value future at the discount factor β. 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 is 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.

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, i.e. the replacement rate of the unemployment insurance. The log of productivity $\log(p_t)$ is modeled as an AR(1) with persistence ρ and volatility of the stochastic component σ^2_ϵ. Consumers can save in a liquid asset that gives the fixed return r. They choose consumption each period to maximize lifetime utility subject to the budget constraint and the borrowing constraint.

Table 16 shows the calibration for the model. For the annual calibration, I follow Jappelli et al. (2024) and model productivity 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, and unemployment income is 0.75, i.e., the replacement rate of the Italian Unemployment Insurance. The interest rate is set at 2%. The rates s_{COL} and f are set to 0.048 and 0.76 respectively, to match the quarterly EU transition rate and the job finding probability estimated by D'Amuri et al. (2022). On the other hand, s_{JA} is calibrated to meet the 2016-2019 average unemployment rate of 11.2%[18]. It's set to 0.431, corresponding to an average tenure of 2.32 years. The model can be written as in equation (7)[19].

I solve the model by endogenous grid method (see Carroll 2006), separately for employed and unemployed states. 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. JAs workers' consumption, conditioning on the same wealth, productivity and employment status, is always below the one of COL workers. In other words, JA workers choose to consume less than COLs, even when they have the same

[18] I start from $u_t - u_{t-1} = (1 - u_t)\tilde{s} - u_{t-1}f$, where $\tilde{s} = \omega_{COL}s_{COL} + \omega_{JA}s_{JA}$. Knowing f, ω_{COL}, ω_{JA} and s_{JA}, and setting $u_t = u_{t-1} = 11.2\%$, I solve for s_{JA}.

[19] I drop the subscript i for notational convenience. The recursive formulation and computational details are in Appendix F.

resources as the latter ones. Following the above discussion, this result might be either due to higher precautionary saving or a lower incomes.[20]

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

I simulate the model with $N = 20{,}000$ and $T = 200$ (I simulate 250 years but I discard the first 50 as burn-in period). The share of COL is set to $\omega_{COL} = 0.875$, while the share of JA is $\omega_{JA} = 0.125$. Therefore, there are 17,500 COL and 2,500 JA workers respectively. I start the simulation with an initial wealth equal to 0.

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.[21] It can be seen that the COLs first-order stochastically dominates 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. I then estimate (5) and (6) with simulated data and report the results in Table 17. Column 1 simply shows the mean difference in consumption between JA and COL workers. This difference of 8.04% is close to the estimated result of around 9% with imputed consumption. Column 2 controls for the PDV of income for lagged wealth and for tenure. The last one is computed in the model as the number of consecutive years in employment. I find that a coefficient attached to *Jobs Act* slightly below 6%. This means that the sole effect of employment risk accounts for more than two thirds of the total effect of the *Jobs Act*. Column 3 allows for a more flexible interaction between employment risk,

[20] Figure B3 plot the mean income for JA and COL workers over the simulated period od the model.

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

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 that 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.

The importance of precautionary saving motive due to the JA can be also appreciated in the model by regressing wealth against the JA dummy, a specification that resembles the one by Fuchs-Schündeln and Schündeln (2005). I do this in columns 4 and 5 of Table 17. In column 4, the coefficient is negative, suggesting that on average JAs hold fewer assets than COLs. However, when one controls for the PDV, and therefore providing a mean-preserving spread of the income distribution, the coefficient turns positive. Controlling for the PDV and for tenure, JAs hold 0.72 more in wealth than COLs. This implies a 16% increase relative to the COL mean of 4.268.

Sensitivity Analysis. Figure 10 plots the model-implied consumption difference as a function of γ. The baseline calibration is the point of the graph corresponding to $\gamma = 2$ (the vertical dotted line). The graph conveys two main messages. First, the effect is generally decreasing in the relative risk aversion. The intuition for this result is the following: the higher is γ, the stronger is the precautionary saving motive (and thus prudence), independently from the degree of employment security. Therefore, on the margin, the effect of the JA on savings is lower. The second result is that the model-implied effect matches well the empirical estimates for values of γ between 1 and 3 , while is distant for higher values. This means that the values of risk aversion that discipline the model to replicate a causally identified effect are very close to those commonly used in estimating macro models, especially HANK models (see Auclert et al. 2024 and Kaplan and Violante 2022). To a certain extent, the approach outlined here is specular to the one by the seminal paper by Gourinchas and Parker (2002), who estimate the preference parameters by matching observed mean life cycle profiles using CEX data. In this paper, instead of using indirect inference to estimate some parameters of the model, I directly assign the parameters and see how close I get to the empirical estimates. For the same reason, this paper also contributes to the large literature estimating the EIS, outlined in the above literature review. In other words, my paper might be seen as a bridge between the structural and the reduced form literature on income risk and consumption.

9 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 1.30 percentage points, corresponding to a drop in spending between 6 and 9%, depending on the measure of consumption used as dependent variable (either the log of food spending or the BPP-imputed log of consumption).

These effects are stronger for younger workers, those who live in the North of Italy and for those employed in firms with union representation.

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 preferences (CRRA utility with risk aversion of 2 and a discount factor of 0.96), it generates a difference in consumption between risky and safe employees that closely matches the estimated causal effect. Running a regression on simulated data, controlling for PDV and asset levels, I find that employment risk heterogeneity accounts for more than two-thirds of the observed consumption difference, while the remaining third is attributable to the PIH channel.

This work can be improved along several dimensions. 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) nondurable consumption data, rather than imputed spending, - and possibly wealth - would represent a substantial data improvement.

These findings also pave the way for a broader research agenda I am currently pursuing. First, recent evidence by Graves (2025) suggests that employment risk amplifies business cycle dynamics, as it represents a key component of counter cyclical income risk. Given that JA workers are expected to constitute the majority of the labour force in the coming years, an important open question is whether structural labour market reforms affect not only long-run growth and productivity but also the magnitude of short-run macroeconomic fluctuations.

It is also valuable to extend the model to incorporate an illiquid asset, as in Kaplan and Violante (2014), to study the channels of precautionary saving. Graves (2025) shows that individuals primarily engage in precautionary saving using liquid assets, which can be easily accessed in the case of income shocks due to the absence of transaction costs. This is especially relevant in the short run[22]. It would be particularly interesting to test whether this holds in the Italian context, where financial literacy is relatively lower than in the U.S., potentially increasing reliance on liquid assets regardless of employment protection levels.

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.

[22] Graves (2025) shows that the extent of consumption smoothing during unemployment spells depends critically on liquid wealth: individuals with more liquid assets prior to unemployment (i.e., not wealthy hand-to-mouth) experience smaller consumption declines.

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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: OLS results for food consumption - share × 100

	(1)	(2)	(3)	(4)
Jobs Act	-2.32***	-1.27**	-1.29**	-1.33***
	(0.51)	(0.52)	(0.51)	(0.51)
Male		-1.29***	-1.40***	-1.57***
		(0.27)	(0.27)	(0.28)
Married		-1.01***	-0.98***	-0.89***
		(0.34)	(0.34)	(0.34)
Family Size		2.27***	2.21***	2.17***
		(0.16)	(0.16)	(0.16)
N. of Children		0.24	0.22	0.20
		(0.21)	(0.21)	(0.21)
Between 25 and 29		-0.33	-0.51	-0.38
		(0.77)	(0.77)	(0.77)
Between 30 and 39		-0.25	-0.46	-0.30
		(0.76)	(0.76)	(0.76)
Between 40 and 49		0.81	0.75	0.86
		(0.77)	(0.77)	(0.77)
Between 50 and 64		2.73***	2.65***	2.86***
		(0.79)	(0.78)	(0.79)
Observations	8338	8328	8328	8328
Regional Dummies			✓	✓
Job Sector				✓

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 include only large firm workers. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 4: OLS Results for food consumption - log level $\times$ 100

	(1)	(2)	(3)	(4)
Jobs Act	-10.56***	-5.70***	-5.75***	-5.86***
	(2.13)	(2.11)	(2.10)	(2.10)
Male		-4.08***	-4.29***	-5.01***
		(1.03)	(1.03)	(1.08)
Married		0.44	0.35	0.42
		(1.30)	(1.30)	(1.29)
Family Size		11.48***	11.27***	11.05***
		(0.65)	(0.65)	(0.65)
N. of Children		-1.49*	-1.46*	-1.40*
		(0.78)	(0.79)	(0.79)
Between 25 and 29		-2.11	-2.42	-2.41
		(3.07)	(3.07)	(3.08)
Between 30 and 39		0.18	-0.15	0.02
		(2.99)	(3.00)	(3.01)
Between 40 and 49		3.83	3.79	3.84
		(3.05)	(3.05)	(3.06)
Between 50 and 64		12.03***	11.90***	12.20***
		(3.07)	(3.08)	(3.09)
Household Income (Log x 100)	0.87***	0.80***	0.80***	0.81***
	(0.01)	(0.01)	(0.01)	(0.01)
Observations	7665	7656	7656	7656
Regional Dummies			✓	✓
Job Sector				✓

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

Table 5: OLS results for weekly hours worked

	(1)	(2)	(3)	(4)
Jobs Act	2.12***	1.25***	1.27***	1.20***
	(0.33)	(0.34)	(0.34)	(0.33)
Male		4.18***	4.23***	3.57***
		(0.18)	(0.18)	(0.18)
Married		-0.37*	-0.39*	-0.37*
		(0.21)	(0.21)	(0.21)
Family Size		-0.23**	-0.21**	-0.25***
		(0.09)	(0.09)	(0.09)
N. of Children		-0.45***	-0.44***	-0.42***
		(0.12)	(0.12)	(0.12)
Between 25 and 29		0.52	0.57	0.59
		(0.48)	(0.48)	(0.48)
Between 30 and 39		0.45	0.52	0.48
		(0.47)	(0.47)	(0.47)
Between 40 and 49		-0.12	-0.08	-0.23
		(0.49)	(0.49)	(0.48)
Between 50 and 64		-0.58	-0.53	-0.60
		(0.49)	(0.49)	(0.48)
Individual Net Income (€100)	0.74***	0.65***	0.64***	0.65***
	(0.02)	(0.02)	(0.02)	(0.02)
Observations	8375	8362	8362	8362
Regional Dummies			✓	✓
Job Sector				✓

Notes: This Table shows OLS estimates of (2). The dependent variable is weekly average labour supply, defined as how many hours the respondents declares to work on a typical week. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample include only large firm workers. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 6: Results with BPP-Imputed Measures of Consumption

	large Firms			Small Firms		
	(1)	(2)	(3)	(4)	(5)	(6)
Panel A: Log of Consumption						
Jobs Act	-10.38***	-9.40**	-9.65**	-2.70	-1.15	-0.60
	(3.71)	(3.72)	(3.75)	(5.34)	(5.37)	(5.42)
Observations	7665	7665	3753	3440	3440	1460
R^2	0.006	0.016	0.017	0.006	0.038	0.050
Job Sector Dummies		✓	✓		✓	✓
Risk Aversion			✓			✓
Panel B: Consumption to Income Ratio						
Jobs Act	-9.41***	-9.54***	-9.32***	-6.02	-5.48	-4.61
	(2.97)	(2.97)	(2.99)	(4.23)	(4.28)	(4.36)
Observations	7665	7665	3753	3440	3440	1460
R^2	0.001	0.004	0.012	0.001	0.013	0.027
Job Sector Dummies		✓	✓		✓	✓
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 7: Food Consumption in small firms - share × 100

	(1)	(2)	(3)	(4)
Jobs Act	-0.85	-0.85	-0.95	-0.88
	(0.73)	(0.72)	(0.71)	(0.72)
Male		-1.70***	-1.75***	-1.99***
		(0.43)	(0.43)	(0.45)
Married		-1.01*	-1.00*	-1.09**
		(0.53)	(0.53)	(0.53)
Family Size		2.24***	2.17***	2.13***
		(0.25)	(0.25)	(0.25)
N. of Children		0.09	0.13	0.22
		(0.33)	(0.33)	(0.33)
Between 25 and 29		-0.47	-0.72	-0.66
		(1.04)	(1.03)	(1.04)
Between 30 and 39		-0.47	-0.68	-0.57
		(1.03)	(1.03)	(1.04)
Between 40 and 49		1.00	0.86	0.92
		(1.08)	(1.08)	(1.08)
Between 50 and 64		1.17	0.99	1.06
		(1.12)	(1.11)	(1.12)
Observations	3778	3774	3774	3774
Regional Dummies			✓	✓
Job Sector				✓

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

Table 8: Food consumption in small firms - log level × 100

	(1)	(2)	(3)	(4)
Jobs Act	-2.57	-2.07	-2.33	-1.89
	(2.89)	(2.77)	(2.78)	(2.81)
Male		-6.71***	-6.81***	-7.79***
		(1.57)	(1.57)	(1.67)
Married		0.43	0.25	-0.24
		(2.00)	(2.00)	(1.99)
Family Size		10.16***	9.90***	9.81***
		(0.95)	(0.96)	(0.96)
N. of Children		-1.71	-1.53	-0.99
		(1.22)	(1.22)	(1.22)
Between 25 and 29		-0.30	-0.79	-0.64
		(3.93)	(3.93)	(3.94)
Between 30 and 39		0.49	0.12	0.22
		(3.94)	(3.93)	(3.95)
Between 40 and 49		5.44	5.17	5.14
		(4.11)	(4.10)	(4.12)
Between 50 and 64		5.87	5.57	5.60
		(4.27)	(4.26)	(4.28)
Household Income (Log x 100)	0.96***	0.88***	0.89***	0.88***
	(0.02)	(0.02)	(0.02)	(0.02)
Observations	3440	3437	3437	3437
Regional Dummies			✓	✓
Job Sector				✓

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

Table 9: Weekly hours worked in small firms

	(1)	(2)	(3)
Jobs Act	0.50	0.46	0.46
	(0.59)	(0.59)	(0.59)
Male	4.47***	4.42***	3.99***
	(0.35)	(0.36)	(0.38)
Married	-2.03***	-2.02***	-1.88***
	(0.42)	(0.42)	(0.42)
Family Size	-0.50***	-0.52***	-0.50***
	(0.17)	(0.17)	(0.17)
N. of Children	-0.16	-0.16	-0.25
	(0.26)	(0.26)	(0.26)
Between 25 and 29	0.00	-0.03	-0.02
	(0.73)	(0.73)	(0.73)
Between 30 and 39	-0.10	-0.13	-0.09
	(0.72)	(0.72)	(0.72)
Between 40 and 49	-1.30*	-1.31*	-1.35*
	(0.77)	(0.77)	(0.76)
Between 50 and 64	-1.43*	-1.46*	-1.41*
	(0.79)	(0.79)	(0.79)
Individual Net Income (€ 100)	1.06***	1.07***	1.09***
	(0.05)	(0.05)	(0.05)
Observations	3913	3913	3913
Regional Dummies		✓	✓
Job Sector			✓

Notes: This Table shows OLS estimates of (2). The dependent variable is weekly average labour supply, defined as how many hours the respondents declares to work on a typical week. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample include only small firm workers. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 10: Food consumption in public sector - share × 100

	(1)	(2)	(3)
Jobs Act	-1.39*	-0.52	-0.60
	(0.77)	(0.79)	(0.78)
Male		-1.36***	-1.54***
		(0.27)	(0.27)
Married		0.25	0.24
		(0.35)	(0.35)
Family Size		1.87***	1.86***
		(0.17)	(0.17)
N. of Children		0.03	-0.02
		(0.19)	(0.19)
Between 25 and 29		-0.17	-0.17
		(1.01)	(1.01)
Between 30 and 39		-0.90	-0.88
		(0.94)	(0.94)
Between 40 and 49		0.61	0.65
		(0.93)	(0.93)
Between 50 and 64		1.16	1.03
		(0.91)	(0.91)
Observations	8519	8506	8506
Regional Dummies			✓

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

Table 11: Food consumption in public sector - log level × 100

	(1)	(2)	(3)
Jobs Act	-3.64	0.58	0.39
	(2.97)	(2.99)	(2.98)
Male		-6.59***	-6.97***
		(1.00)	(1.01)
Married		6.81***	6.58***
		(1.31)	(1.32)
Family Size		8.87***	8.77***
		(0.65)	(0.65)
N. of Children		-1.54**	-1.58**
		(0.73)	(0.73)
Between 25 and 29		-3.76	-3.78
		(4.01)	(4.01)
Between 30 and 39		-5.57	-5.50
		(3.73)	(3.74)
Between 40 and 49		-0.47	-0.36
		(3.66)	(3.67)
Between 50 and 64		3.64	3.30
		(3.61)	(3.62)
Household Income (Log x 100)	0.90***	0.81***	0.82***
	(0.01)	(0.01)	(0.01)
Observations	7881	7870	7870
Regional Dummies			✓

Notes: This Table shows OLS estimates of (2). The dependent variable is weekly average labour supply, defined as how many hours the respondents declares to work on a typical week. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample include only public sector employees. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 12: Weekly hours worked in public sector

	(1)	(2)
Jobs Act	-0.87	-0.84
	(0.71)	(0.71)
Male	3.44***	3.56***
	(0.19)	(0.19)
Married	0.14	0.13
	(0.24)	(0.24)
Family Size	-0.17	-0.16
	(0.11)	(0.11)
N. of Children	-0.23*	-0.18
	(0.13)	(0.13)
Between 25 and 29	-0.06	-0.13
	(0.65)	(0.64)
Between 30 and 39	-1.12*	-1.22**
	(0.61)	(0.60)
Between 40 and 49	-1.46**	-1.55***
	(0.60)	(0.59)
Between 50 and 64	-2.13***	-2.13***
	(0.59)	(0.58)
Individual Net Income (€100)	0.43***	0.42***
	(0.02)	(0.02)
Observations	8934	8934
Regional Dummies		✓

Notes: This Table shows OLS estimates of (2). The dependent variable is weekly average labour supply, defined as how many hours the respondents declares to work on a typical week. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample include only public sector employees. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 13: Including only firms with at least 20 employees

	Food Share		Log of Food		Labour Supply	
	(1)	(2)	(3)	(4)	(5)	(6)
Jobs Act	-2.37***	-1.20**	-10.87***	-5.36**	2.21***	1.32***
	(0.54)	(0.55)	(2.27)	(2.09)	(0.34)	(0.35)
Male		-1.76***		-5.68***		3.51***
		(0.30)		(1.16)		(0.20)
Married		-0.95**		0.56		-0.31
		(0.37)		(1.40)		(0.22)
Family Size		2.15***		11.04***		-0.23**
		(0.17)		(0.66)		(0.10)
N. of Children		0.28		-1.23		-0.38***
		(0.23)		(0.84)		(0.13)
Between 25 and 29		-0.61		-3.82		0.67
		(0.85)		(3.13)		(0.53)
Between 30 and 39		-0.76		-1.75		0.60
		(0.83)		(3.09)		(0.51)
Between 40 and 49		0.55		2.94		-0.22
		(0.85)		(3.17)		(0.53)
Between 50 and 64		2.67***		11.51***		-0.70
		(0.87)		(3.19)		(0.53)
Household Income (Log x 100)			0.89***	0.80***		
			(0.01)	(0.01)		
Individual Net Income (€100)					0.71***	0.65***
					(0.02)	(0.02)
Observations	7300	7291	6705	6697	7253	7243
Regional Dummies		✓		✓		✓
Job Sector Dummies		✓		✓		✓

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

Table 14: Risk Aversion and Discount Rate

	Food Share		Log of Food		Labour Supply	
	(1)	(2)	(3)	(4)	(5)	(6)
Jobs Act	-2.39***	-1.26**	-11.04***	-4.85**	1.96***	0.86**
	(0.51)	(0.55)	(2.12)	(2.26)	(0.33)	(0.37)
Male		-1.21***		-3.36**		3.85***
		(0.41)		(1.62)		(0.28)
Married		-1.20**		-0.92		-0.07
		(0.47)		(1.91)		(0.30)
Family Size		2.17***		11.51***		-0.26*
		(0.22)		(0.94)		(0.14)
N. of Children		0.53*		0.04		-0.44**
		(0.28)		(1.10)		(0.18)
Between 25 and 29		-2.17		-10.93*		0.82
		(1.45)		(6.09)		(0.97)
Between 30 and 39		-0.98		-1.94		-0.21
		(1.44)		(6.05)		(0.97)
Between 40 and 49		-0.21		-0.40		-0.87
		(1.48)		(6.16)		(1.00)
Between 50 and 64		1.25		7.20		-1.34
		(1.49)		(6.23)		(1.00)
Household Income (Log x 100)			0.87***	0.76***		
			(0.02)	(0.02)		
Individual Net Income (€100)					0.68***	0.60***
					(0.03)	(0.03)
Observations	4118	4118	3753	3753	3686	3686
Regional Dummies		✓		✓		✓
Job Sector Dummies		✓		✓		✓
Risk Aversion	✓	✓	✓	✓	✓	✓
Discount Rate	✓	✓	✓	✓	✓	✓

Notes: This Table shows the OLS estimates for equation (3). The dependent variables are indicated on the columns. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 15: Heterogeneity Results

	By Age		By Geographical Area		Union Presence	
	≤ 40	> 40	South and Centre	North	No Union	Union
Panel A: Food Consumption (share)						
Jobs Act	-1.51**	-1.44	-0.41	-1.98***	-0.73	-1.83***
	(0.66)	(0.92)	(0.90)	(0.61)	(0.76)	(0.71)
Observations	3513	4815	3431	4896	2706	5622
Panel B: Food Consumption (log level)						
Jobs Act	-7.57***	-4.79	-2.41	-8.37***	-3.04	-8.30***
	(2.77)	(3.57)	(3.42)	(2.67)	(3.10)	(2.93)
Observations	3229	4427	3199	4456	2463	5193
Panel C: Weekly Hours Worked						
Jobs Act	1.44***	0.55	1.38**	1.33***	1.19**	1.18**
	(0.48)	(0.65)	(0.62)	(0.43)	(0.55)	(0.49)
Observations	3700	4854	3571	4983	2803	5751

Notes: This Table shows OLS estimates of (2). The dependent variable is either food consumption (Panel A) or weekly labour supply (Panel B). Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample include only large firm workers. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table 16: Model Baseline Calibration

Parameter	Description	Value	Source/Target
Preference Parameters			
γ	Relative Risk Aversion	2	Standard
β	Discount Factor	0.96	Standard
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_{JA}	Annual JA Job Separation Rate	0.393	Unemployment Rate in 2016-2019
s_{COL}	Annual COL Job Separation Rate	0.048	D'Amuri et al. (2022)
f	Annual Job Finding Rate	0.759	D'Amuri et al. (2022)
Asset Grid			
a_{min}	Borrowing Constraint	0	
a_{max}	Maximum Value of Asset	50	
n_a	Number of Asset Grid-points	100	
Prices			
w	Wage Rate	1	Normalization
b	UI Replacement Rate	0.75	UI Replacement Rate
r	Annual Interest Rate	0.02	Jappelli et al. (2024)
Simulation Parameters			
N	Number of Individuals	20,000	
T	Number of Periods	200	
ω_{JA}	Share of JA workers	0.125	PLUS 2018
ω_{COL}	Share of COL workers	0.875	PLUS 2018

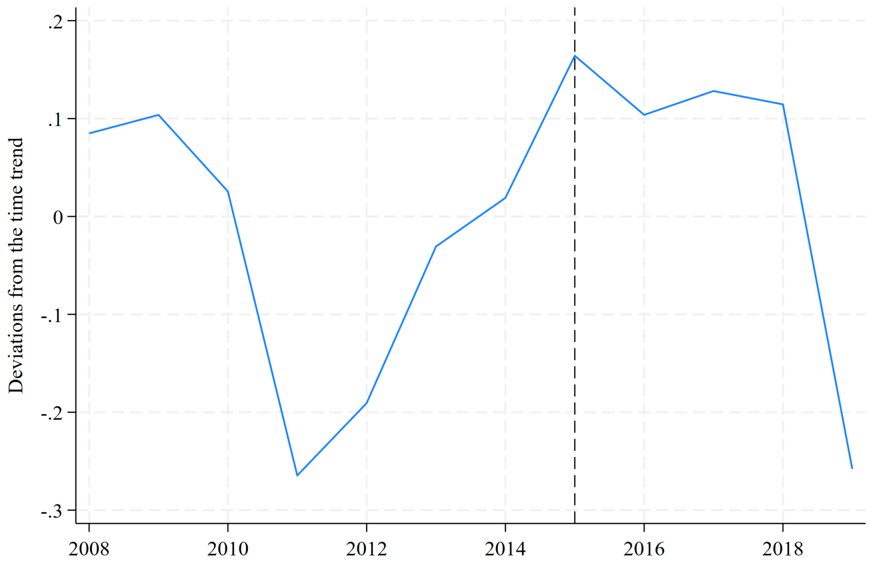
Table 17: Regression with Simulated Data (Baseline Calibration)

Dependent Variable:	Log of Consumption × 100			Wealth	
	(1)	(2)	(3)	(4)	(5)
Jobs Act	-8.043***	-5.799***	-7.636***	-0.203***	0.722***
	(0.069)	(0.055)	(0.189)	(0.008)	(0.008)
PDV of Income		0.398***	0.392***		0.268***
		(0.002)	(0.002)		(0.000)
L.Wealth		4.957***	4.967***		
		(0.004)	(0.004)		
Tenure		0.005***	0.004***		0.009***
		(0.001)	(0.001)		(0.000)
Jobs Act x PDV of Income			0.050***		
			(0.006)		
Jobs Act x L.Wealth			-0.085***		
			(0.011)		
Jobs Act x Tenure			0.552***		
			(0.034)		
Observations	4000000	3980000	3980000	4000000	4000000
Baseline Mean	.	.	.	4.268	4.268

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

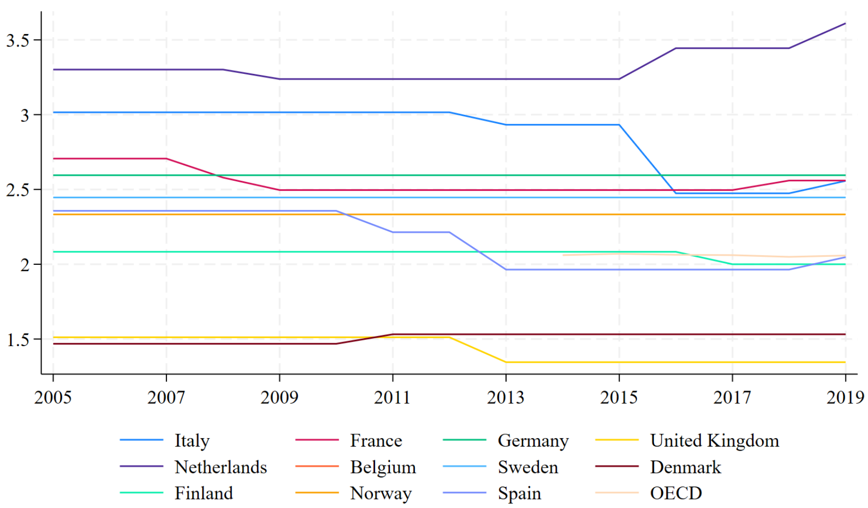
Figures

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



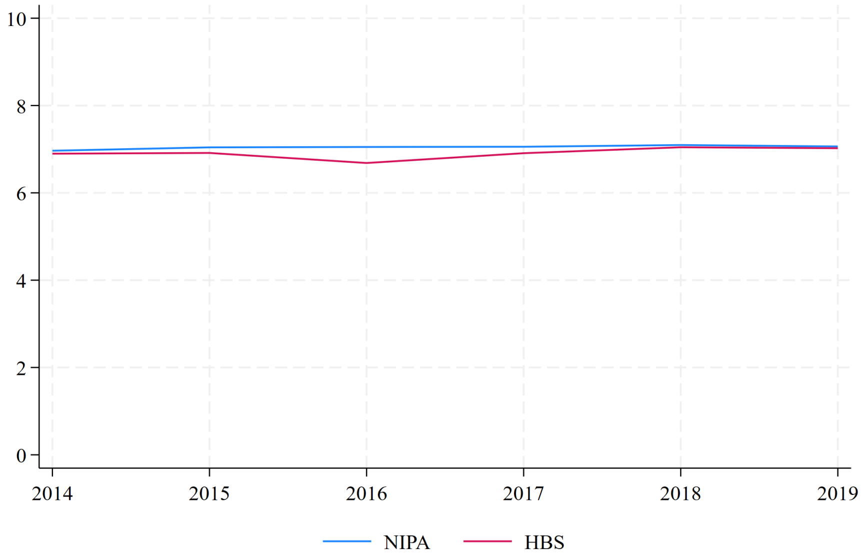
Notes: The Figure shows average tenure of Italians employed in large firms between 2008 and 2019. Data are de-trended to account for the business cycle (i.e. the 2008-2011 double-deep recessions). Average tenure increases until 2016 and then goes down sharply.

Figure 2: Employment Protection in OECD countries



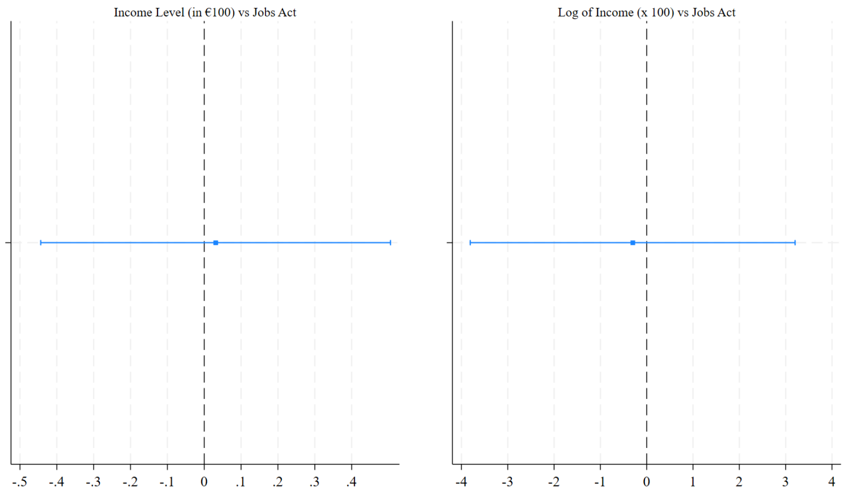
Notes: The Figure shows the employment protection index computed by the OECD. It can be seen a sharp drop in 2015 and 2016 for Italy, represented by the blue line on top. It's interesting to notice that even after the 2012 reform, Italy had a more rigid labour market than many of the OECD countries.

Figure 3: Ratio between Non Food and Food Consumption



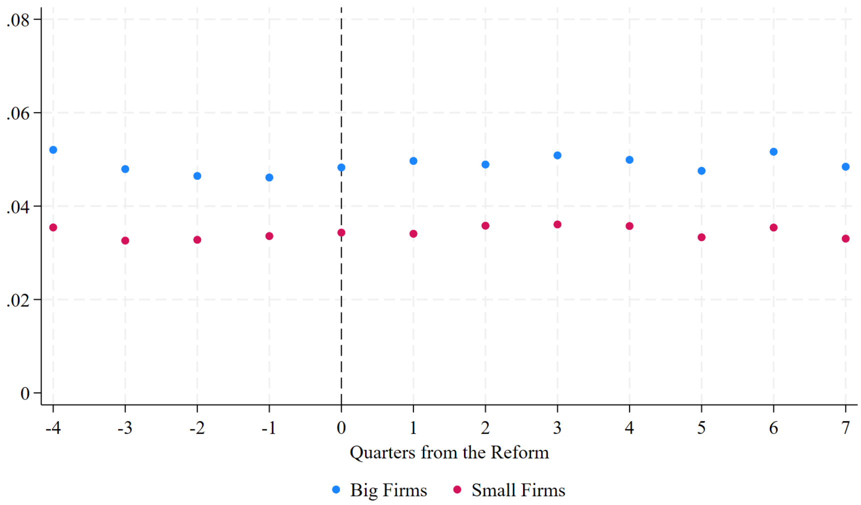
Notes: The Figure shows the ratio between non food consumption and food consumption from national accounts and from ISTAT's HBS from 2014 to 2019. The HBS series is constructed by dropping the top and bottom deciles of the pooled cross section distribution.

Figure 4: Income vs Jobs Act



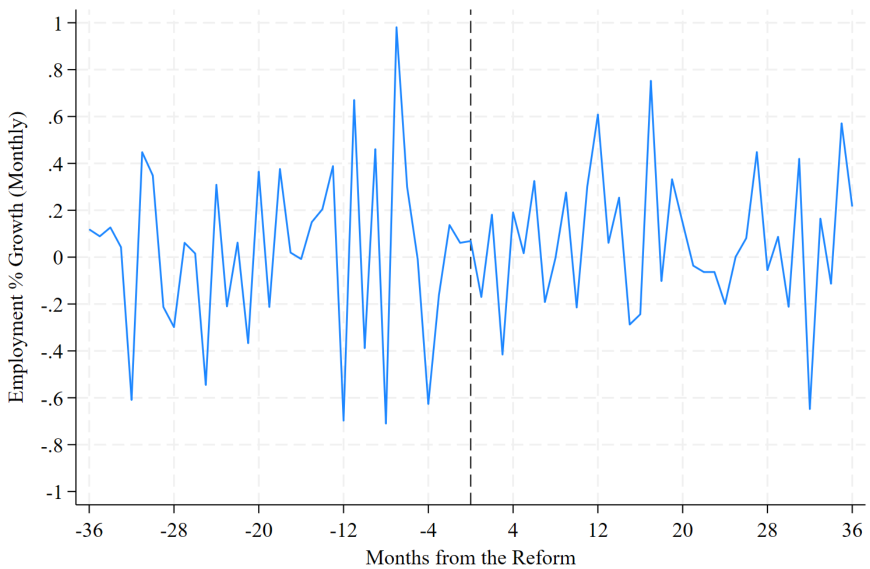
Notes: This Figure shows the OLS estimated coefficient of income (both in levels and in logs) and the Jobs Act dummy, controlling for the same variables used in the main analysis.

Figure 5: Hiring Rates in Small and Large Firms



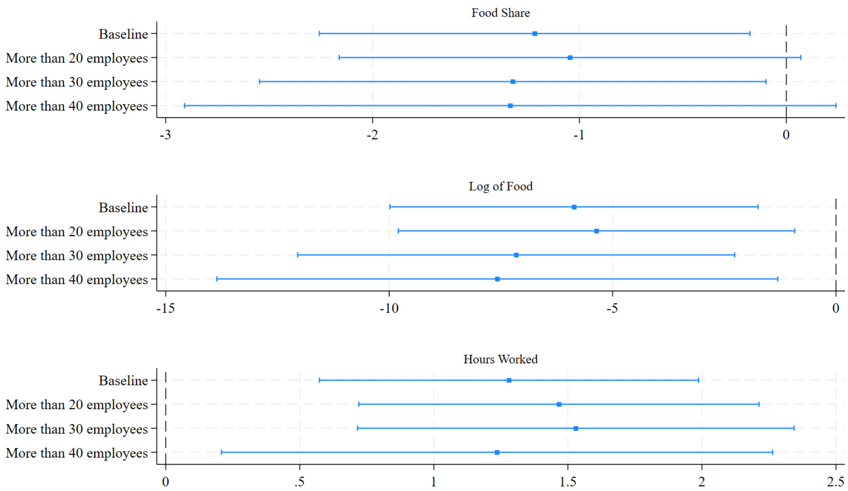
Notes: This Figure shows the sample hiring rates before and after the reform (first quarter 2015, t = 0). The horizontal axis is the quarter from 2014 (-4 is 2014-q1) to 2019 (7 is 2019-q4). The Source is the Italian Labour Force Survey.

Figure 6: Monthly Net Employment Growth



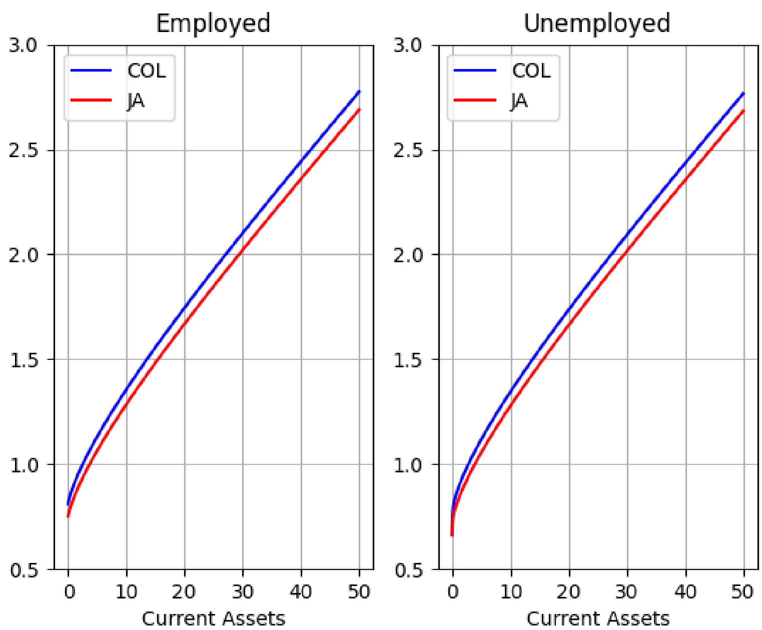
Notes: This Figure shows the monthly growth rate of the level of employment. The Source is the Italian Statistical Office.

Figure 7: Estimated Causal Effects for Bigger Firms



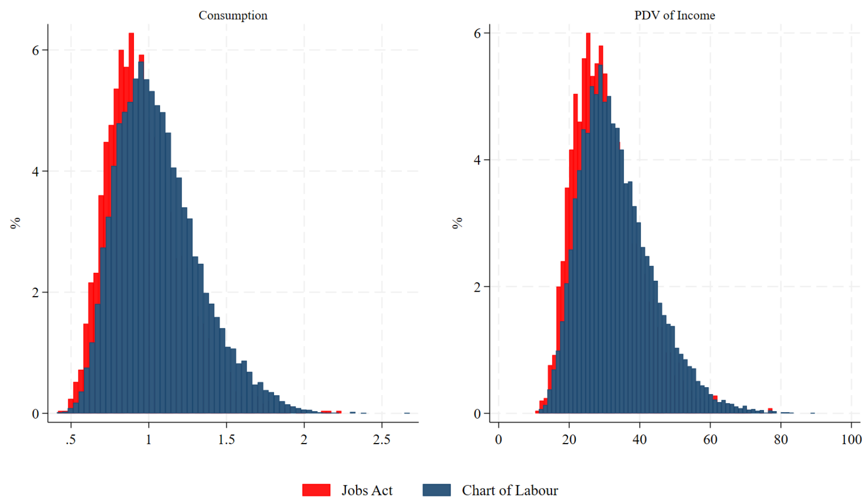
Notes: This Figure shows the effect on food consumption and labour supply for employees of firms with +20, +30 and +40 workers.

Figure 8: Consumption Policy Functions by employment status



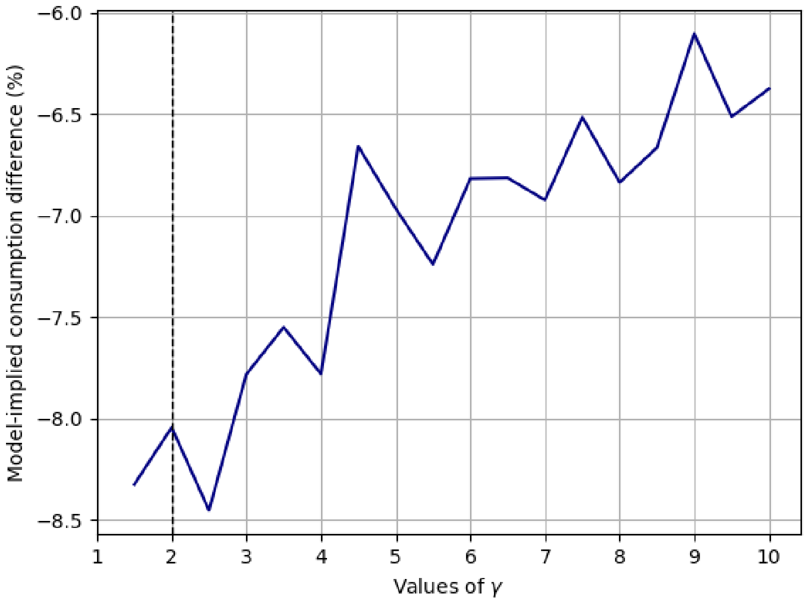
Notes: This Figure plots consumption as a function of current period assets for JA and COL workers (red and light) for the median productivity state for employed and unemployed people (left and right panel respectively). Baseline calibration: $\beta = 0.96$, $\gamma = 2$ and $r = 0.02$

Figure 9: Distribution of Consumption and PDV of Income



Notes: This Figure plots the consumption difference between JA and COL workers as a function of γ.

Figure 10: Sensitivity Analysis



Notes: This Figure plots the model-implied difference of the log of consumption between JA and COL workers as a function of risk aversion. The vertical dotted line indicates the baseline calibration with $\gamma = 2$. The remaining parameters of the model are the same as in the baseline calibration.

Appendix A: Patience and Risk Aversion

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

Imagine of 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	Fine (Function of Tenure)

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

Table B2: Income Against Jobs Act - Full Regression

	Individual Income Level (€100)		Log of Individual Income (× 100)	
	(1)	(2)	(3)	(4)
Jobs Act	0.03	0.12	-0.30	-0.92
	(0.24)	(0.27)	(1.79)	(2.00)
Male	3.46***	3.80***	25.87***	27.29***
	(0.12)	(0.19)	(0.90)	(1.38)
Married	0.76***	0.76***	5.55***	4.69***
	(0.14)	(0.23)	(1.05)	(1.66)
Family Size	-0.34***	-0.35***	-2.87***	-2.78***
	(0.06)	(0.10)	(0.47)	(0.71)
N. of Children	0.50***	0.64***	3.46***	4.42***
	(0.08)	(0.13)	(0.64)	(0.97)
Between 25 and 29	0.56**	1.11**	6.91***	8.12*
	(0.27)	(0.47)	(2.29)	(4.16)
Between 30 and 39	0.67**	1.47***	7.86***	9.43**
	(0.26)	(0.48)	(2.27)	(4.22)
Between 40 and 49	1.34***	2.20***	11.50***	13.24***
	(0.28)	(0.51)	(2.38)	(4.38)
Between 50 and 64	1.63***	1.96***	12.65***	10.42**
	(0.29)	(0.52)	(2.39)	(4.42)
Regional Dummies	✓	✓	✓	✓
Job Dummies	✓	✓	✓	✓
Observations	8362	3684	8362	3684
R^2	0.230	0.219	0.230	0.222
Sample	2014 and 2018	Only 2018	2014 and 2018	Only 2018

Notes: This Table shows the OLS estimation results of a regression of individual income (both the log and the level) against the *Jobs Act* dummy along the demographic variables used in the analysis. The regressions include time, geographical and sector FEs

Table B3: Dropping under 25

	Food Share	Log of Food x 100	Weekly Hours Worked	BPP-Log
	(1)	(2)	(3)	(4)
Jobs Act	-1.07**	-5.38**	1.12***	-9.68***
	(0.52)	(2.16)	(0.35)	(3.16)
Male	-1.49***	-4.90***	3.53***	
	(0.29)	(1.10)	(0.19)	
Married	-1.02***	-0.52	-0.37*	
	(0.35)	(1.31)	(0.21)	
Family Size	2.26***	10.98***	-0.23**	
	(0.17)	(0.68)	(0.10)	
N. of Children	0.06	-1.60**	-0.50***	
	(0.21)	(0.80)	(0.12)	
Between 30 and 39	0.18	2.74	-0.10	
	(0.44)	(1.77)	(0.28)	
Between 40 and 49	0.81*	5.14***	-0.90***	
	(0.46)	(1.84)	(0.29)	
Between 50 and 64	2.42***	12.37***	-1.35***	
	(0.49)	(1.88)	(0.29)	
Years of Education	-0.50***	-1.38***	-0.10***	
	(0.04)	(0.17)	(0.03)	
Household Income (Log x 100)		0.84***		
		(0.01)		
Individual Net Income (€100)			0.66***	
			(0.02)	
Observations	8014	7367	8005	7324

Notes: This Table shows the OLS estimates of (2), but dropping under 25 observations. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B4: Excluding People from the Centre-South and under 25

	Food Share	Log of Food x 100	Weekly Hours Worked	BPP-Log
	(1)	(2)	(3)	(4)
Jobs Act	-1.64***	-7.43***	1.15***	-11.17***
	(0.62)	(2.73)	(0.42)	(4.20)
Male	-1.69***	-5.82***	3.66***	
	(0.37)	(1.44)	(0.24)	
Married	-0.92**	0.38	-0.69***	
	(0.43)	(1.70)	(0.27)	
Family Size	2.13***	11.10***	-0.01	
	(0.22)	(0.94)	(0.12)	
N. of Children	0.56**	-0.44	-0.70***	
	(0.28)	(1.08)	(0.16)	
Between 30 and 39	-0.19	1.92	-0.25	
	(0.57)	(2.36)	(0.34)	
Between 40 and 49	0.17	3.37	-1.31***	
	(0.59)	(2.43)	(0.36)	
Between 50 and 64	1.86***	10.77***	-1.99***	
	(0.62)	(2.48)	(0.36)	
Years of Education	-0.51***	-1.47***	-0.12***	
	(0.06)	(0.23)	(0.04)	
Household Income (Log x 100)		0.83***		
		(0.02)		
Individual Net Income (€100)			0.64***	
			(0.03)	
Observations	4694	4270	4617	3844

Notes: This Table shows the OLS estimates of (2), but dropping under 25 and people living in the South. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B5: Excluding People from the North and under 25

	Food Share	Log of Food x 100	Weekly Hours Worked	BPP-Log
	(1)	(2)	(3)	(4)
Jobs Act	-0.30	-2.66	1.16*	-7.96*
	(0.93)	(3.54)	(0.59)	(4.74)
Male	-1.12**	-3.20*	3.22***	
	(0.46)	(1.70)	(0.31)	
Married	-1.19**	-1.90	0.10	
	(0.59)	(2.10)	(0.34)	
Family Size	2.29***	10.58***	-0.45***	
	(0.25)	(0.98)	(0.15)	
N. of Children	-0.53	-3.08**	-0.37*	
	(0.34)	(1.20)	(0.19)	
Between 30 and 39	0.72	4.04	0.15	
	(0.70)	(2.67)	(0.45)	
Between 40 and 49	1.69**	7.63***	-0.32	
	(0.75)	(2.82)	(0.49)	
Between 50 and 64	3.43***	15.48***	-0.44	
	(0.78)	(2.87)	(0.49)	
Years of Education	-0.47***	-1.23***	-0.08	
	(0.07)	(0.26)	(0.05)	
Household Income (Log x 100)		0.85***		
		(0.02)		
Individual Net Income (€100)			0.67***	
			(0.03)	
Observations	3319	3096	3388	3480

Notes: This Table shows the OLS estimates of (2), but dropping under 25 and people living in the North. Standard errors in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B6: Interaction results - food share × 100

	(1)	(2)	(3)
Jobs Act	-2.18***	-2.07***	-0.60
	(0.78)	(0.59)	(0.59)
Jobs Act x Age 30-49	2.12**		
	(1.04)		
Jobs Act x Over 50	-1.18		
	(1.41)		
Jobs Act x Centre		1.84	
		(1.16)	
Jobs Act x South		1.96	
		(1.38)	
Jobs Act x High Income			-2.41**
			(0.98)
Observations	8328	8328	8328
Controls and FEs	✓	✓	✓

Notes: This Table shows OLS estimates of an equation in which the *Jobs Act* is interacted with age bins, geographical and income distribution. High income is 1 if the individual belongs to the second half of the income distribution. 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 include only large firm workers. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B7: Interaction results - log of food consumption $\times$ 100

	(1)	(2)	(3)
Jobs Act	-10.05***	-9.01***	-6.44***
	(3.36)	(2.58)	(2.44)
Jobs Act x Age 30-49	8.39*		
	(4.34)		
Jobs Act x Over 50	1.46		
	(5.90)		
Jobs Act x Centre		9.42**	
		(4.80)	
Jobs Act x South		6.24	
		(5.30)	
Jobs Act x High Income			1.73
			(4.17)
Observations	7656	7656	7656
Controls and FEs	✓	✓	✓

Notes: This Table shows OLS estimates of an equation in which the *Jobs Act* is interacted with age bins, geographical and income distribution. High income is 1 if the individual belongs to the second half of the income distribution. The dependent variable is the log of food consumption. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample include only large firm workers. Standard errors, clustered at the individual level, in parentheses. * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

Table B8: Interaction results - labour supply

	(1)	(2)	(3)
Jobs Act	2.17***	1.54***	1.16***
	(0.51)	(0.39)	(0.40)
Jobs Act x Age 30-49	-1.54**		
	(0.67)		
Jobs Act x Over 50	-1.79*		
	(0.98)		
Jobs Act x Centre		-0.74	
		(0.78)	
Jobs Act x South		-0.99	
		(0.86)	
Jobs Act x High Income			0.12
			(0.59)
Observations	8362	8362	8362
Controls and FEs	✓	✓	✓

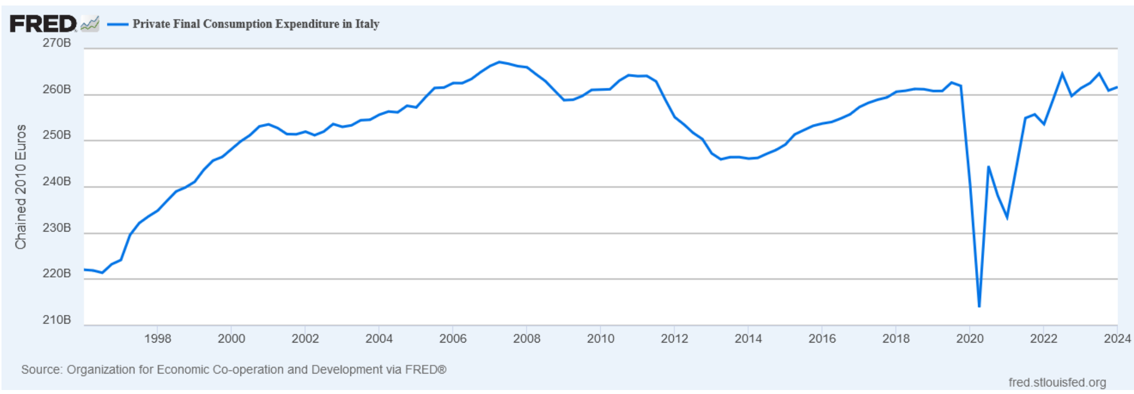
Notes: This Table shows OLS estimates of an equation in which the *Jobs Act* is interacted with age bins, geographical and income distribution. High income is 1 if the individual belongs to the second half of the income distribution. The dependent variable is weekly hours worked. Jobs Act is a dummy equal to 1 if the worker is hired after March 7 2015. The sample include only large firm workers. 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 Maurizo Landini, today's head of CGIL. The union collected signatures to a petitions 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, that, according to the law on popular referenda, is responsible for counting the signatures, making sure they are genuine, and evaluate 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 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 in 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 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 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 actual log of spending (right panel) with the imputed (left panel).

Table C1: Summary Statistics of HBS

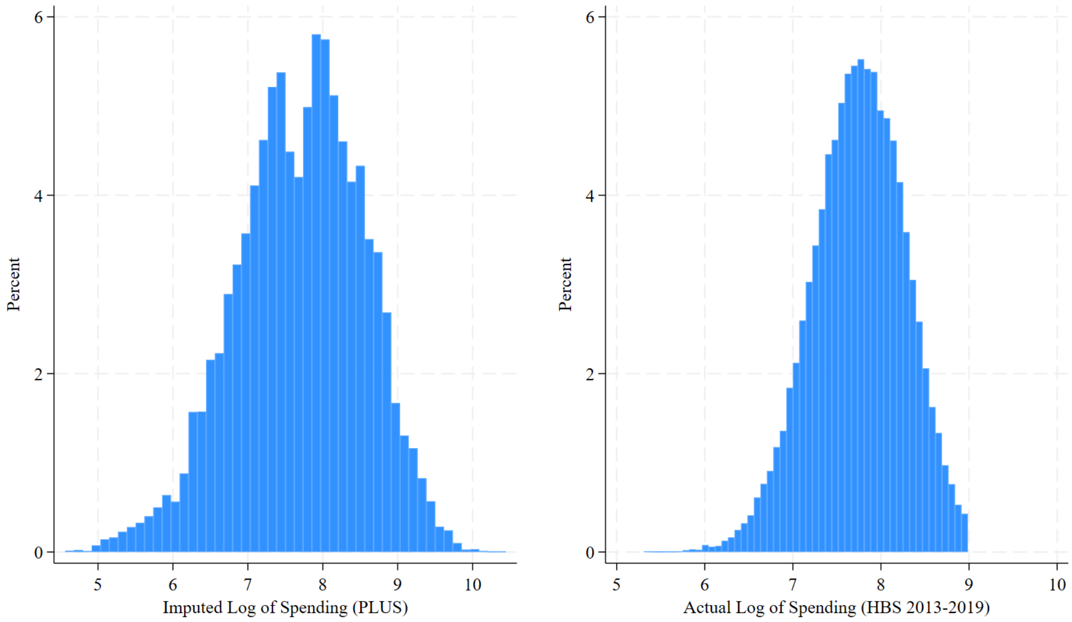
	Mean	Sd
Age 18-34	0.12	0.32
Age 35-64	0.88	0.32
Male	0.68	0.47
Married or Civil Union	0.54	0.50
South	0.29	0.45
Centre	0.21	0.41
North	0.50	0.50
Total spending (monthly)	2625.06	1346.57
Food-only spending (monthly)	470.06	274.42
Total/Food Consumption Ratio	6.75	4.08

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

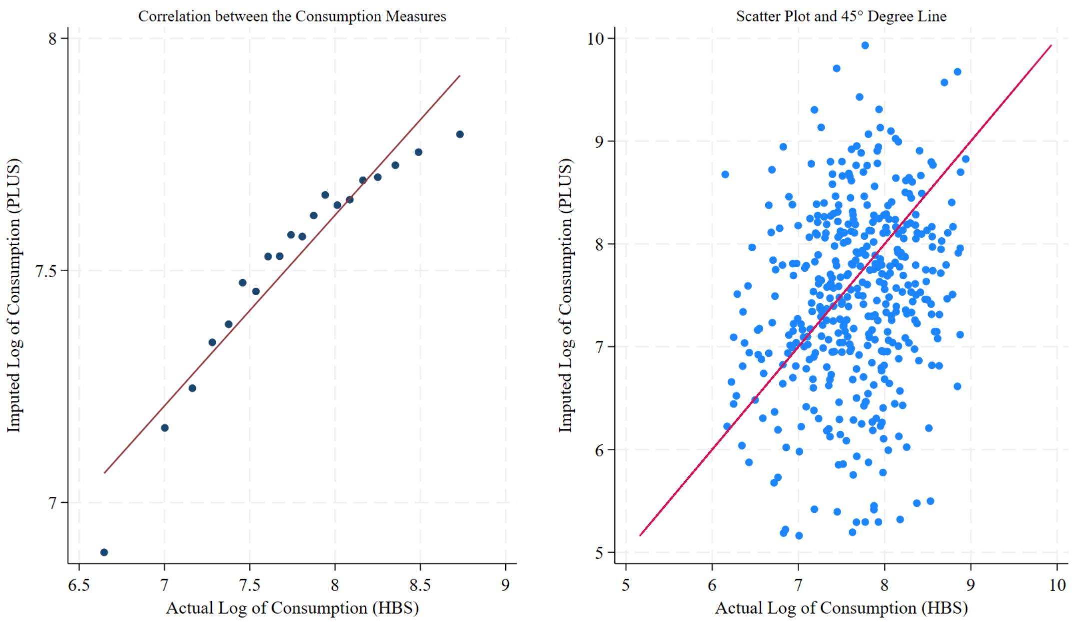
Standard errors in parentheses
* $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$

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 degrees 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 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 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 on 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 job 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 labor 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 by the Monti Government, a technocratic Government appointed after Prime Minister Silvio Berlusconi's resignation in November 2011. After the reform, that was named after Labour Minster Elsa Fornero[23], 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 of the unlawfulness. Given judges' pro labour attitude, they typically motivate the unjust dismissals on the basis of those circumstances for which reinstatement was compulsory, given that workers' most preferred option is to get the job back 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

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

ruled by the judge, the employers is not forced neither to reinstate nor 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, 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 lead 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 labour market reforms 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, that is therefore best dealt with by experts from the Government). Other reforms dealt with the regulation of mass layoffs, safety and privacy on the workplace and unemployment insurance. Two concerns raise 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 to 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 have been hired. They are therefore uncorrelated with the JA and might affect consumption only by means 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 firms side.

Appendix F: Recursive Formulation of the Model and Details on the Algorithm

In this Appendix, I show the recursive formulation of the problem. 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 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.