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SHORT-TIME WORK, ITS REPLACEMENT RATE AND LABOUR MOBILITY

by Gaetano Basso*, Emanuele Ciani*, Francesco Manaresi* and Lorenzo Pesaresi*

Abstract

This paper studies the impact of the short-time work (STW) replacement rate on workers' job-to-job mobility. Exploiting a discontinuity in the Italian STW benefit schedule and matched employer-employee data, we show that a 10-percentage-point increase in the replacement rate raises the probability of remaining with the current employer by 3 per cent, thereby reducing job-to-job flows. The effect is driven by low-ability workers, who are less likely to move to higher-paying firms, suggesting a deterioration in workforce quality among firms relying on STW. We develop a model to rationalise these results and derive an optimal STW benefit formula. The fiscal externality is positive but smaller than that generated by unemployment benefits.

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1 Introduction¹

By subsidising temporary reductions in working hours, short-time work (STW) schemes preserve job matches and mitigate excessive unemployment during downturns (Kopp and Siegenthaler, 2021; Giupponi and Landais, 2023). The Great Recession and the Covid-19 crisis were no exception: these schemes were either massively expanded or introduced in OECD countries, and a significant share of the workforce was placed under STW arrangements (OECD, 2020; Boeri and Cahuc, 2023). The level of the STW replacement rate influences workers’ ability to smooth consumption when their employers need to operate below full capacity, similar to unemployment insurance benefits in the event of a layoff. However, unlike unemployment insurance, STW schemes are often particularly generous. Many countries even increased the replacement rate during the 2020 recession to compensate for the sharp reduction in hours worked.²

The use of STW is first and foremost a firm-level decision, as employers apply for the scheme and reduce hours worked when facing economic distress. STW provides value to firms by preserving existing matches, but it also generates costs by potentially sustaining unviable jobs and hampering reallocation (Boeri and Cahuc, 2023; Cahuc, 2024). This can occur when STW subsidises low-productivity companies with limited growth potential (Giupponi and Landais, 2023), or due to moral hazard, when employers inflate take-up by misreporting or cross-subsidising across jobs (Cahuc and Nevoux, 2017; Lapeyre, 2024).

At the individual level, high STW benefits dampen fluctuations in workers’ income in their current jobs, thereby facilitating consumption smoothing.³ However, a high replacement rate can also entail costs by diminishing incentives to seek new opportunities beyond distressed firms, thereby hindering job-to-job mobility and slowing down labour reallocation. Due to the lack of suitable identifying variation, the existing literature has largely overlooked the worker-level consequences of STW. In particular,

¹We are grateful to Federico Cingano, Giulia Giupponi, Camille Landais, Attila Lindner, Matteo Paradisi, Jon Piqueras, Lorenzo Pandolfi, Johannes Spinnewijn, Eliana Viviano and participants at several institutions and conferences for insightful comments and discussions. We thank Francesca Carta for sharing with us the data on regional job vacancies used in Carta and Rizzica (2018). The views expressed are those of the authors and do not necessarily reflect those of the Bank of Italy or the Eurosystem. *Corresponding author:* Gaetano Basso (gaetano.basso@bancaditalia.it).

²In April 2020, Germany increased the statutory replacement rate of the *Kurzarbeit* scheme from 60% to 70% from the third month and to 80% from the seventh month. In May 2020, Japan nearly doubled the top ceiling for its STW benefit in large firms. The UK introduced a generous furlough scheme, allowing workers to receive 80% of their wages up to a cap.

³An increase in the benefits might also make it easier for trade unions and workers to accept the STW scheme (Boeri and Bruecker, 2014). This seems to be a relatively minor factor in the context we study — Italy during and after the Great Recession — given the already widespread use of the instrument.

there is no evidence on how the STW replacement rate affects job flows and transitions to new jobs.

In this paper, we provide new evidence on the effect of the STW replacement rate on job-to-job mobility. We focus on the job-to-job mobility choices of workers placed on STW by distressed firms. We develop a theoretical framework where risk-averse workers, differing in ability, set the reservation productivity for alternative employers to warrant a job-to-job transition. The theoretical framework delivers two key predictions. First, higher STW benefits increase the likelihood of remaining with the original firm as workers raise their reservation productivity. Second, holding the current wage constant, whether low- or high-ability workers react more strongly to changes in the replacement rate depends on the shape of the productivity density of outstanding vacancies around the reservation productivity — a local measure of outside options.

Empirically, we provide causal estimates of the effect of the STW replacement rate by exploiting a discontinuity in the schedule of the Italian STW scheme, the Cassa Integrazione Guadagni (CIG). The CIG replacement rate increases by 10 percentage points when the previous month wage crosses a specific threshold that can be treated as quasi-random (as it does not coincide with variation in any other labour market policy). This lends itself naturally to a Regression Discontinuity Design (RDD) that we conduct on a sample of linked employer–employee data from the Italian National Social Security Institute (INPS) spanning more than 10 years (2008–2019). The period of analysis includes the Great Recession and the sovereign-debt crisis, during which the proportion of the Italian workforce on STW rose to as much as 4 per cent of the working-age population. Although the RDD strategy identifies only the effect around the threshold, we argue that this estimate is policy-relevant. This is because (i) a significant portion of the workforce earns wages near this threshold and (ii) the increase in the replacement rate at the threshold is substantial.

Consistent with the first model prediction, our main finding shows that the probability that a worker remains with the same firm one year after being placed on STW increases with the replacement rate, being 2.2 percentage points higher to the right of the threshold, or nearly 3% compared to the baseline likelihood. Consequently, job-to-job mobility is reduced, with the probability of transitioning to a new firm 1.7 percentage points lower above the threshold — a significant reduction of around 20% relative to the baseline likelihood. This effect remains substantial when rescaled to account for flows from workers who are not on STW. We calculate that a 10-percentage-point increase in benefits can lead to a 4% reduction in job-to-job transitions among full-time permanent-contract workers near the mode of the Italian wage distribution. Moreover, the retention effects of higher benefits are more pronounced among younger

workers and in local labour markets with greater slack.

The increased likelihood of workers staying with their original firm is an intended outcome of STW schemes. If a higher replacement rate helps preserve high-quality matches and prevents workers from moving to less favourable firms, reduced mobility could be beneficial. However, limiting the flow of workers towards firms with higher potential may reduce overall efficiency. To provide empirical discipline to the second model prediction on heterogeneous effects across workers, we estimate worker fixed effects in a wage regression using the approach introduced by Abowd et al. (1999, AKM). Our analysis reveals that the reduction in job flows at the threshold is concentrated among workers with a lower AKM fixed effect, leading to a deterioration in the quality of the workforce in the original firm. Additional analyses also suggest that the workers are less inclined to move towards high-AKM-FE firms. This result is rationalised by our theoretical model under *increasing* productivity density of outstanding vacancies, implying that lower-ability workers have more outside options around their reservation productivity.

Finally, we emphasise that maintaining matches through STW entails a fiscal cost, arising both from the risk that firms delay resuming normal activity and from workers remaining on benefits for longer. To quantify the latter, we use our theoretical model to derive the optimal STW replacement rate, which solves a heterogeneous-agent counterpart of the Baily–Chetty formula applied to STW (Baily, 1978; Chetty, 2006). The optimal STW replacement rate balances the marginal social gain from redistributing consumption shares to poorer workers against the fiscal externality, which is inflated by a moral hazard cost. The latter captures the marginal increase in the share of workers on STW induced by a more generous replacement rate through reduced job-to-job transitions. Based on our empirical findings, we estimate a local elasticity of remaining on STW with respect to its replacement rate of 15%, which is lower than standard estimates for unemployment insurance benefits. This result points to limited moral hazard induced by STW through worker behavior. Our findings complement the evidence of similarly mild moral hazard driven by firm behaviour documented by Giupponi et al. (2022). We provide a unified framework to interpret these two margins jointly. Within this framework, the moral hazard effects induced by worker search behavior and firm take-up choices compound each other, bringing the overall fiscal externality generated by STW closer to that typically estimated for UI systems.

Our paper contributes to three strands of literature. First, it complements existing analyses on the effect of STW schemes. The literature, recently reviewed by Cahuc (2024), has mostly focused on decisions to use STW at the sectoral or firm level (Hijzen and Venn, 2011; Boeri and Bruecker, 2014; Kopp and Siegenthaler, 2021; Diàz

et al., 2023; Giupponi and Landaïs, 2023). While most of the literature finds that STW schemes are generally effective in reducing congestions in the labour market (Montenegro and Hijzen, 2023; Mohimont et al., 2024) and preserving jobs (Kopp and Siegenthaler, 2021; Giupponi and Landaïs, 2023; Arranz et al., 2018), the evidence on their impact on overall economic reallocation is more limited and mixed (Giupponi and Landaïs, 2023; Diàz et al., 2023; Lalinsky et al., 2024). The work closest to ours is by Giupponi and Landaïs (2023), which shows that the availability of CIG helps Italian firms preserve employment after a temporary crisis.⁴ However, for low-productivity firms, CIG mostly subsidises jobs that would disappear regardless of its take-up. Unlike these papers, we are the first to study the incentives for individuals to stay with their current employer or search for a new job. We do so by exploiting local changes in STW replacement rates at the worker level around the policy-induced discontinuity. We argue that the generosity of the replacement rate is a key, yet understudied, policy lever that can alter such incentives. Indeed, an optimal STW scheme would subsidise labour hoarding to preserve valuable matches while, at the same time, not distorting the flow of workers toward more productive firms. We contribute to the existing literature by providing clean identification of the impact of the STW replacement rate on match preservation and job-to-job transitions at the worker level.

Second, our paper is related to the literature on the generosity of work insurance schemes. This body of research has focused almost exclusively on unemployment insurance benefits, particularly on determining the optimal level of benefits that balances the insurance value they provide against jobseekers’ moral hazard. Most international evidence indicates that greater generosity — in terms of benefit level or duration — leads to longer unemployment spells, although the magnitude of this effect varies across studies and subpopulations.⁵ More closely related to our work, Grindaker and Simmons (2024) recently studied how unemployment benefits influence workers’ decisions to search for new jobs while employed, finding that higher benefits reduce workers’ propensity to move to a new job. Interestingly, the effect of the STW replacement rate on the job-to-job transition probability that we estimate is very similar to the effect of unemployment benefits they report. Unlike Grindaker and Simmons (2024) and the broader unemployment insurance literature, we focus on STW — a setting in which the replacement rate is of first-order relevance for job-to-job transition decisions.

Third, our paper contributes to the growing literature investigating job-to-job mobility motives, especially when workers face high layoff risk (Jarosch, 2021; Cederlöf

⁴This finding is confirmed by Pavlopoulos and Chkalova (2022), who study the impact of STW in the Netherlands during the Great Recession.

⁵For a review, see Schmieder and von Wachter (2016).

et al., 2024; Simmons, 2025). We provide empirical evidence on workers’ mobility in a context where the risk of unemployment is particularly high, but where a specific form of public insurance aimed at mitigating this risk is available.

The paper is organised as follows. Section 2 briefly discusses how STW works in the Italian context. Section 3 provides a theoretical framework and introduces the empirical approach. Section 4 introduces the data and presents evidence that supports the identification strategy. Section 5 reports the main results. Section 6 draws on insights from the unemployment benefits literature to quantify the fiscal externality associated with the STW benefit and to discuss the optimality of its current level. Section 7 concludes.

2 Institutional details

The Italian STW scheme, *Cassa Integrazione Guadagni* (CIG), has been in place since 1941 to preserve viable job relationships during downturns by subsidising a reduction in hours worked (up to zero hours worked). There are three sub-programmes. The ordinary scheme (CIGO) is traditionally reserved for manufacturing and lasts up to 12 months. If a firm’s crisis lasts longer, the firm can request an “extraordinary” intervention (CIGS) that covers additional 24 months. This latter scheme also partially covers other sectors, including services. Since the Great Recession, the Italian government introduced a supplementary scheme (“in deroga”, CIGD), which worked similarly to the other two and allowed access to firms otherwise not covered. In the remainder of the paper, we consider the three schemes together, as there is no difference in workers’ benefits and replacement rates. This also implies that we cover a wider set of firms not limited to manufacturing.⁶ Until 2017, if a firm exhausted the maximum duration of CIGS and could not reintegrate employees on short-time work, those dismissed could continue to receive the same benefits for an additional 12 to 48 months through a related scheme called *indennità di mobilità*.⁷ Since this scheme follows the same benefit schedule, we can consider it as an extension of CIGS for the purpose of our analysis.

The employer is not free to arbitrarily choose which workers to put in CIG. Before applying for authorisation from the National Social Security Institute, firms must consult with the trade unions and establish objective criteria for selection. These

⁶During the period we analyse, there were some differences in firms’ requisites and contributions and in the administration in charge of granting authorisation (by regional authorities instead of the National Social Security Institute). However, there were no differences in terms of workers’ eligibility and benefits.

⁷The duration is extended for workers over 50 and/or those in southern Italy. The *indennità di mobilità* was abolished in 2017.

criteria need to be consistent with the economic reasons for diminished activity and must be fully transparent and measurable. After consulting with the trade unions, the employer applies for authorisation, which involves only a formal review of the firm’s eligibility. Finally, the National Social Security Institute either pays the subsidy directly to the workers (“CIG diretta”) or reimburses the firm (“CIG a congruaglio”).

The benefit that employees receive while on CIG depends on a *baseline* monthly wage. This baseline wage — which is the running variable z for our RDD — is calculated by removing any additional components, such as overtime pay, from the monthly wage and then increasing it by $1/12$ to accommodate the annual extra payment typical in Italian contracts, known as the 13th-month salary. The benefit covers 80% of the baseline wage, but it is subject to two distinct caps. In 2016, if z was below €2,102.24, the maximum payment was €971.71 for workers who were on a full-time CIG (referred to as “0-hour CIG”). For z values above this threshold, the cap increased to €1,167.91, resulting in an approximate €200 increase in the benefit near the threshold.⁸ The caps are adjusted for inflation every year using a national index that is applied to all social benefits. It is worth mentioning that the threshold is set close to the modal point of the monthly baseline wage distribution and our results are, therefore, representative for a large share of the Italian workforce (Appendix Figure 1).

A worker loses the benefit if he or she finds a new permanent job. While on CIG, a worker may accept a fixed-term position, subject to contractual limitations. However, any earnings from this employment are completely deducted from the benefit. There is, therefore, little incentive to take up a fixed-term job, and it is rare for people to pick up such secondary jobs while on STW. Moreover, employees on CIG are protected against dismissal for economic reasons, except in cases of collective dismissal or other lawful grounds for terminating the individual contract, such as a breach by the employee.

3 Theoretical framework and empirical approach

3.1 Theoretical framework

We develop a theoretical model to illustrate how the STW benefit replacement rate affects workers’ job mobility decisions once the employer reduces their working hours

⁸Workers are frequently placed on short-time work for only a portion of their working hours, so the calculation of the benefit is performed on an hourly basis. This involves dividing both the baseline wage and the caps by the contractual hours. This method ensures that the relationship between the replacement rate (benefit over foregone earnings) and the baseline wage remains consistent, regardless of whether the CIG is full-time (0-hours) or partial. As a result, the relevant thresholds remain unchanged.

and puts them on STW. The model will then be used to derive an optimal STW replacement rate formula à la Baily-Chetty (Section 6).

We consider a stationary economy populated by a measure one of risk-averse workers, who differ in ability $x \in [0, 1]$, and by a continuum of one-job firms, which differ in productivity $y \in [0, 1]$. Let $G(y)$ denote the productivity distribution of outstanding job vacancies.

Match output p is linearly increasing and additively separable in worker ability x , firm productivity y , and a zero-mean match-specific shock ϵ . To generate sorting between workers and firms, we further assume that workers of ability x are qualified for all jobs of productivity $y \leq x$:⁹

$$p(x, y, \epsilon) = \begin{cases} x + y + \epsilon & x \geq y, \\ 0 & \text{else.} \end{cases}$$

We assume that workers appropriate a constant share of match output, which yields a wage function $w = w(x, y, \epsilon)$ that is linearly increasing in all entries and has zero cross-derivatives (i.e., $w_{xy} = w_{x\epsilon} = w_{y\epsilon} = 0$). We further posit that match-specific shocks are drawn from an exogenous distribution after mobility decisions are taken and remain fixed for the entire duration of the match thereafter.

Some jobs are in STW, meaning that the worker works a fraction $h \in (0, 1)$ of full-time hours (normalized to 1). Let $\theta \in (0, 1)$ denote the (exogenous) probability that a job currently on STW remains on STW in the next period,¹⁰ β the discount factor, and δ the exogenous match destruction probability, upon which the worker enters unemployment with value U . The government subsidizes non-worked hours at replacement rate $\alpha \in (0, 1)$. Hence, workers on STW earn a wage $(h + \tilde{\alpha})w$, where $\tilde{\alpha} \equiv (1 - h)\alpha$. We refer to $\tilde{\alpha}$ as the *gross* STW replacement rate.

Employed workers randomly search the labour market for alternative job opportunities. Let λ denote the probability that an employed worker meets a vacancy in a given period. Suppose that firms post vacancies for full-time jobs only.

⁹The functional form of the production technology we adopt is commonly used in the statistical discrimination literature (Jarosch and Pilossoph, 2019; Lalive et al., 2025).

¹⁰We interpret the parameter θ as capturing the firm-level decision of whether to continue STW. Intuitively, it reflects the persistence of the negative shock that motivates the use of STW. Consistent with our empirical analysis, we treat θ as given and focus on worker-level decisions.

The value of an employed worker on STW reads:

$$W^{STW}(x, y, \epsilon; \tilde{\alpha}) = u((h + \tilde{\alpha})w(x, y, \epsilon)) + \beta(1 - \delta) \left[\lambda(G(x) - G(y^*)) \int_{y^*}^x \mathbb{E}_{\tilde{\epsilon}}[W(x, \tilde{y}, \tilde{\epsilon})] \right. \\ \left. dG(\tilde{y} | \tilde{y} \in (y^*, x)) + (1 - \lambda(G(x) - G(y^*))) \mathbb{E}_{\theta}[W^{STW}(x, y, \epsilon; \tilde{\alpha})] \right] + \beta\delta U.$$

where the instantaneous utility function is increasing and concave ($u' > 0, u'' < 0$), and $\mathbb{E}_{\theta}[W^{STW}(x, y, \epsilon; \tilde{\alpha})] = \theta W^{STW}(x, y, \epsilon; \tilde{\alpha}) + (1 - \theta)W(x, y, \epsilon)$ is the continuation value from staying at the current, STW-using firm. In words, an employed worker meets a vacancy with probability λ . Conditional on meeting, she is qualified for the vacancy with probability $G(x)$. Finally, she decides to move to the new job if its value exceeds the one of the job she currently holds. This mobility decision is summarized by the reservation productivity y^* , which is the only choice variable of the worker.

The reservation productivity $y^*(x, y, \epsilon; \tilde{\alpha})$ makes a worker with ability x and match-specific shock ϵ currently on STW at a firm y with gross replacement rate $\tilde{\alpha}$ indifferent between staying at her current job and moving to a new full-time job. In Appendix B we derive the exact indifference condition. To obtain transparent comparative statics, we assume throughout that workers set their reservation productivity statically from the indifference condition between moving to a full-time job in another firm of that productivity and remaining in the current STW-using firm:

$$\mathbb{E}_{\tilde{\epsilon}}[u(w(x, y^*(x, y, \epsilon; \tilde{\alpha}), \tilde{\epsilon}))] = \mathbb{E}_{\theta}[u^{STW}(w(x, y, \epsilon; \tilde{\alpha}))],$$

where $\mathbb{E}_{\theta}[u^{STW}(w(x, y, \epsilon; \tilde{\alpha}))] \equiv \theta u((h + \tilde{\alpha})w(x, y, \epsilon)) + (1 - \theta)u(w(x, y, \epsilon))$.

Denote the probability of moving as:

$$m(x, y, \epsilon; \tilde{\alpha}) \equiv \lambda[G(x) - G(y^*(x, y, \epsilon; \tilde{\alpha}))].$$

Notice that our production technology entails that, holding the current wage fixed, high-ability workers (both full-time and on STW) are expected to move more frequently. We empirically validate this distinctive prediction of our production technology in the data (see Section 5.2 and Appendix Table 7).

By explicitly modeling workers' reservation productivity policy, we are in the position to characterize the effect of the replacement rate on the frequency of job-to-job transitions. All proofs are relegated to Appendix B.

Proposition 1 (Average effect of STW replacement rate on job-to-job transitions.)

The probability of moving is decreasing in the replacement rate:

$$\frac{\partial m(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} < 0.$$

Our theoretical model predicts that workers in STW should reduce their job-to-job transitions if the replacement rate is higher. This prediction is common across virtually all the alternative modelings of worker search behaviour.

We are not only interested in the average effect of the replacement rate on the job-to-job transition probability but also in the heterogeneous effects by worker ability. Specifically, holding constant both the full-time wage and the employer firm, we would like to figure out whether changing the STW replacement rate affects worker search behaviour differently depending on ability.

Proposition 2 (Heterogeneous effects of STW replacement rate on job-to-job transitions across worker ability.)

Holding the wage and firm productivity constant, the response of the job-to-job transition probability to changes in the STW replacement rate is locally increasing in worker ability x if and only if the density of the productivity distribution is increasing around the associated reservation productivity threshold:

$$\left. \frac{\partial^2 m(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w > 0 \iff G''(y^*(x, y, \epsilon)) > 0.$$

Our theoretical model predicts that the cross-sectional impact of an increase in the STW replacement rate depends on the shape of the productivity distribution of outstanding vacancies — a local measure of worker outside options. Notice that the reservation productivity for a full-time job is likely lower than the productivity of the STW-using firm where a worker is currently employed. Therefore, it seems plausible that the productivity of most outstanding vacancies exceeds the reservation productivity thresholds, which makes the density of the productivity distribution locally increasing. It follows that high-ability workers are expected to react *less* than low-ability workers to changes in the replacement rate.

The intuition behind Proposition 2 is that a marginal increase in reservation productivity translates into a larger decline in job-to-job transitions, the larger the pool of outside options that the worker stops considering. For a given employer and wage, a higher-ability worker must have drawn a worse match-specific shock and is therefore expected to move more frequently regardless of benefit generosity, as her reservation productivity is lower. Hence, if the productivity density is locally increasing, lower-ability workers stop considering more outside options for a given marginal increase in

reservation productivity.

The prediction that the cross-sectional impact of an increase in the STW replacement rate on the job-to-job transition probability depends on the shape of the productivity distribution is a distinctive feature of random search models with two-sided heterogeneity. In Appendix B.1.2 we discuss alternative models of search behaviour and characterize the conditions under which high-ability workers react less to a change in the replacement rate.¹¹

3.2 Empirical implementation

To identify how the generosity of benefits influences workers' decisions, it is necessary to exploit variation in the benefit amount, b , that is independent of other aspects of the employment relationship, such as worker productivity and the likelihood of continued hours reduction in subsequent periods.

We do so by exploiting the €2,102 threshold of the benefit schedule. As previously explained, a worker on a 0-hour CIG earns up to €971 if their baseline wage falls below this threshold and €1,168 if it is above. Consequently, while the replacement rate decreases as the wage increases, it rises by approximately 10 percentage points at the threshold (see Figure 1). Within a Regression Discontinuity Design (RDD) framework, we exploit this change by comparing the outcomes of workers on STW whose wages are just above and just below the threshold.

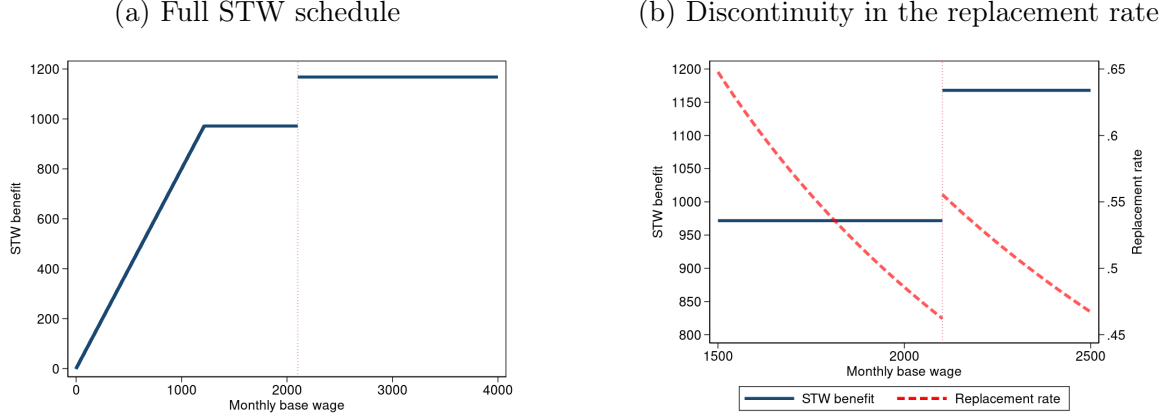
In the analysis, we focus only on workers who are in STW, and for each worker-firm pair we select the first episode of CIG over the period of analysis. We do so to avoid possible sorting across the threshold once the worker discovers the benefit rules (which are otherwise little known to the workforce). For each worker h working in firm i , we estimate the following equation:

$$y_{h(i)} = \alpha_0 + \beta 1(z_{h(i)} > 0) + \sum_{k=1}^K [(\gamma_k z_{h(i)}^k + \delta_k z_{h(i)}^k * 1(z_{h(i)} > 0))] + \varepsilon_{h(i)} \quad (1)$$

where y is the outcome of interest (probability of retention at the original firm or of moving to a new firm) estimated 12 months after the STW event. In some specifications, we also estimate the outcomes at different time intervals t ($t = 3, \dots, 12, \dots, 24$) since the STW event (which is $t = 0$). In the rest of the analysis we rescale the running variable z by taking the difference between the baseline wage at the onset of the STW

¹¹Under risk neutrality and ex ante identical jobs, the corresponding condition is that the search effort cost function is more convex than quadratic with endogenous search effort and random search, and that the elasticity of the meeting function with respect to vacancies is less than 0.5 with directed search.

Figure 1: STW schedule, replacement rate and discontinuity



Note: (0-hour) CIG full benefit schedule (panel (a)) and focus around the €2,102 threshold with the replacement rate for hours not worked (panel (b)); 2016 values.

event (i.e., the one used to calculate the replacement rate) and the threshold that determines the discontinuity in the benefits (€2,102). To preserve comparability over the years, we transform the running variable into real terms using 2016 as the basis.¹² In our baseline estimates, we use a local linear polynomial with a triangular kernel in z ($K = 1$) and the optimal bandwidth of Calonico et al. (2014) (see also Calonico et al., 2017).

Our goal is to identify β : the effect of the generosity of the replacement rate on the probabilities that the worker is still working with the initial firm i at time t , or that he or she works for another firm instead (i.e., that there is a job separation). As an additional dependent variable, we also consider the probability that the worker is no longer in salaried employment.

To this aim, it is important to highlight that the benefit schedule threshold for CIG does not align with other firm-level benefits or policy rules. At the worker level, however, this threshold was also used to cap unemployment benefits until 2012. We believe that this overlap is unlikely to significantly bias our results for three main reasons. First, unemployment benefits are only accessible to workers who are fired, not those who resign. Since employers cannot dismiss workers for economic reasons while claiming CIG, and because CIG entails coordination with trade unions and reduces the likelihood of judicial intervention initiated by dismissed workers, CIG remains — to date — a more common instrument to deal with economic downturns.¹³ To limit

¹²As a deflator, we use the index that the Social Security Institute uses to adjust all the main benefit schedules, including the CIG thresholds.

¹³Workers seeking unemployment benefits would need to be fired for contract violations. To this end, the most common strategy used by workers to force their employers to fire them is to no-show,

the concerns about the interactions between STW and unemployment insurance, we focus on workers already on STW for whom unemployment insurance did not apply. Second, unemployment benefits were of much shorter duration than CIG, despite having the same replacement rate around the threshold, and were thus unattractive also for workers and not only for the firm.¹⁴ Workers thus had little incentive to transition to unemployment benefits; in other words, their main insurance against income shocks was the CIG. Third, the 2012 reform of unemployment insurance — implemented in 2013 — altered the benefit schedule, eliminating the alignment with the CIG threshold. In Section 5.2 we show that our results are similar for CIG spells initiated before or after 2013.

4 Data and validation of the empirical strategy

4.1 Data

We carry out the analysis on a sample of linked employer–employee data from the Italian National Social Security Institute. The dataset covers all non-farm private employees born on 24 calendar dates (the 1st and the 9th of each month) and links them to the firms they are employed in. For each year and each contract between a worker and a firm, the dataset contains information about annual wages, in which months the contract was active, and key job characteristics (e.g. part- or full-time, permanent or fixed-term). It further contains information on when each worker was in STW and for how many hours.

We select workers with at least one CIG episode started between 2008 and 2017. In this period, which began with the Great Recession, STW utilisation was particularly strong (see Figure 2 in the Appendix). Our sample is mainly composed of CIGO events (about 84%). CIGS accounts for most of the remaining events.

First, we turn this dataset into a monthly panel. In the case of multiple job relationships between a firm and a worker in the same month — which is possible in the case of multiple subsequent contracts — we select the highest-paying one. We apply the same selection criterion in the case of multiple job relationships with different firms in the same month.¹⁵

but this entails losing weeks of pay before dismissal and then waiting out the statutory one-week gap before receiving benefits. We see this strategy as highly unlikely.

¹⁴CIG benefits can extend up to 36 months, plus additional years after dismissal if workers qualify for the *indennità di mobilità* extension (see Section 2). In contrast, unemployment benefits lasted a maximum of 12 months until 2012.

¹⁵If a worker has two job contracts at the beginning of a CIG spell, we select the contract with CIG at $t = 0$, irrespective of the pay. This is done to avoid excluding from the analysis workers who

Second, we use this dataset to track the job history of workers who have a CIG episode. For each worker–firm pair, we focus on the first CIG episode during 2008–2017, as it might happen that firms put workers in STW repeatedly. Starting from this first episode, we follow the worker’s employment history every three months for the following two years (hence up to 2019).

We focus on permanent employees only, as the duration of temporary contracts is usually short and firms have no incentive to put them on STW as their contract would terminate anyway.¹⁶ We also focus on full-time employees only. This is done to avoid measurement error, given that in their case we would need to calculate a full-time equivalent wage using poor information on part-time intensity recorded in our data.

Finally, firms below 15 employees should not in principle access the measure, although there are exceptions (in particular for the CIGD), and the “legal” measure of firm size does not correspond to the average annual size we observe in our data. In fact, approximately one fourth of firms in our sample are below 15 employees. To avoid very small firms affecting the results, we truncate the sample at the 5th percentile of firm size, which corresponds to nearly 3 employees. Results are similar if we use all firms. Our selected sample is composed of 145,594 workers. A large fraction of them is located next to the threshold, with the mode of the running variable being just slightly below it.

4.2 The impact on replacement rate and CIG intensity

We begin by analysing the empirical relationship between the benefit amount and the running variable to document the presence of a significant jump at the threshold. The goal of this exercise is to validate the empirical design, despite we can run it only on a subset of workers. We focus here, in fact, on the subsample of the minority of workers to whom the Social Security Institute directly paid the benefits. For them, the administrative archives provide detailed records of individual benefits and exact STW hours.¹⁷ This subsample has, however, one limitation: the baseline wage is recorded only annually in December, or at contract termination if it ends mid-year. This recording frequency may introduce measurement error when calculating the threshold. In

leave the firm as soon as they are put on STW.

¹⁶According to Basso et al. (2021), 41.5 per cent of temporary contracts in the private non-financial service sector last less than 6 months. In the South of Italy, only 7.6 per cent of temporary contracts last more than 12 months (Bank of Italy, 2019).

¹⁷Typically, employers advance the cash, but in case of a firm’s financial distress, the Social Security Institute makes direct payments. In the rest of the analysis, we use the larger sample of workers for whom the firms advance the benefits on behalf of the Social Security Institute, and the running variable is recorded monthly with no measurement error.

the rest of the analysis, we are not able to test for the benefit jump at the threshold as we lack detailed information on the benefit amount and hours in STW, but we observe the running variable much more precisely reducing the concerns about measurement error. Therefore, our main analysis, can be considered a ‘reduced-form’ RDD or, alternatively, as a sharp design assuming that the theoretical jump is effective for all workers (as it should be since these data are the same as the Social Security Administration uses to pay STW benefits based on the provisions of the law and the observed running variable).¹⁸

Indeed, in the direct payment subsample we observe outliers with a replacement rate above 80%, which is theoretically not possible. Even discarding these observations (which are 6.6% of the sample), substantial measurement error remains, as can be observed by plotting the observed replacement rate against the theoretical benefit schedule (Appendix Figure A.3). To address the measurement error in the observed replacement rate, in what follows we report the results based on two samples: the first dropping only outliers with a replacement rate above 80%, and the second excluding also those that have a more than 10 per cent difference between the observed and the theoretical monthly benefit.¹⁹

Figure 2 shows the benefit amount vis-à-vis the running variable z_{it} . To account for the fact that some people are on the scheme only for some hours, all benefits are rescaled as full-time equivalent by multiplying the hourly amount by the monthly contractual hours. There is clear evidence of a sizeable increase in the benefit level and replacement rate at the threshold regardless of the correction applied (see Appendix Table A.2 for the full set of results and Appendix Figure A.4 for robustness checks in terms of polynomial degree and size of the bandwidth). In the first sample, excluding outliers with a more-than-80% replacement rate, the jump is about 55 euro for the full-time-equivalent benefit, about 2.6 percentage points in terms of the replacement rate — which is less than what is expected from the theoretical schedule. Once we impose a tighter restriction on outliers, the discontinuities become about 182 euro and

¹⁸Ideally, we would have run a fuzzy-RDD but the two samples do not contain the same information. An alternative solution could have been that of using a two-sample-2SLS. However, the measurement error in the running variable observed in the direct-payment subsample would have not been necessarily the same as that we might have in the main sample; most likely the former is much larger than the latter. Such issue makes the two-sample approach not ideal. However, we adopt a similar two-step approach – under the assumption of no measurement error in the main-sample running variable – when we quantify the model results to calculate the change in job mobility as a function in the change in benefits (Section 6). That is, we divide our job mobility estimates by the theoretical jump at the threshold provided by the law, 0.093.

¹⁹Without dropping the outliers with a replacement rate above 80%, the estimate of the jump in the monthly benefits ($\hat{\beta}=51.355$, p-value=0.000) is quite similar to Figure 2, panel (a); the estimated jump in the replacement is smaller compared to panel (c) but still significant ($\hat{\beta}=0.018$, p-value=0.000).

8 percentage points, respectively, very close to the values predicted by the theoretical schedule (196 euro and 9.3 percentage points, respectively). Evaluated at the average benefit to the left of the threshold, the increase in benefits is 18.0%.

Although we are unable to provide the same estimate for the main sample used in our analysis, we can extrapolate these estimates accounting for observable characteristics of workers. To this end, we replicate the local linear regression in Figure 2, panel d (excluding outliers), using the same bandwidth and implicit weights but fully interacting both the running variable and the threshold dummy with covariates from a vector X_{it} including age, log of the average wage in the firm, and indicators for gender, foreign citizenship, and the manufacturing sector. This allows for heterogeneity in the estimated jump in benefit at the threshold. We can then extrapolate it to the main sample we use in the rest of the analysis. The value is 183.5 euro, indicating that there is no substantial difference in terms of observable characteristics between the restricted sample we use for the first stage and the main sample of analysis.

The employer may consider the difference in benefit generosity and try to equalise the earnings loss across employees by assigning higher CIG intensity (in terms of share of hours in STW over contractual hours) to those above the threshold who benefit from a higher replacement rate. While there is no direct gain for employers in doing so, this strategy may be used to retain employees and respond to trade union pressures to minimise wage disparities. However, estimates reported in Table A.2 in the Appendix show no evidence of a significant change in the CIG intensity around the threshold in either samples. We take this as a sign of the lack of worker–firm collusion that would invalidate our analysis.

These estimates also underscore that the variation in benefit generosity underlying our RDD strategy is sizeable and relevant for the workers. Those near the threshold on average are on CIG for 52% of their contractual schedule. Coupled with the increase in the replacement rate at the threshold, this results in an earnings loss that is slightly more than 4 percentage points lower for workers on the right-hand side of the threshold.²⁰

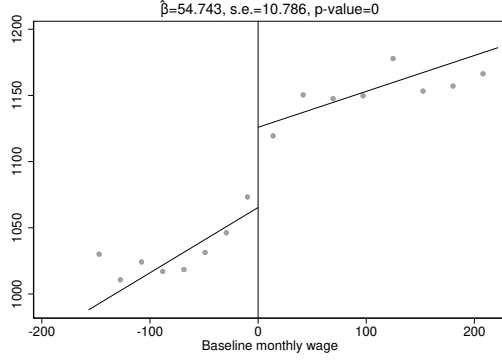
4.3 Evidence of no manipulation nor sorting

The RDD strategy hinges on workers having similar characteristics around the threshold. This might be violated if the running variable is manipulated or if there are systematic differences that lead some groups of workers to be over-represented on

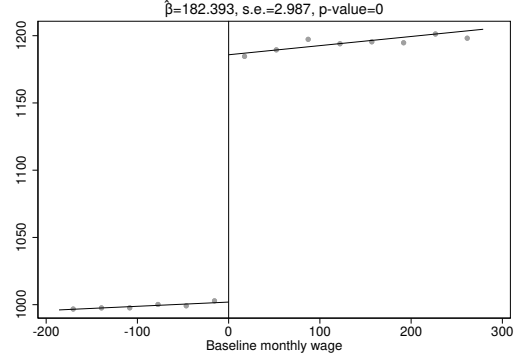
²⁰This is calculated as the jump in the CIG replacement rate times 45% (i.e., the loss they would have because of the average hours reduction).

Figure 2: Discontinuity in the replacement rate and the benefit at the threshold

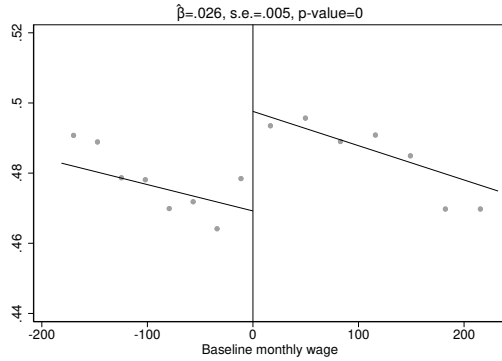
(a) Monthly benefits. Direct payment subsample; weaker outlier detection rule



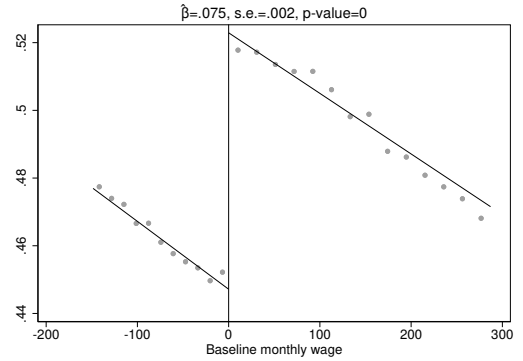
(b) Monthly benefits. Direct payment subsample; stronger outlier detection rule



(c) Replacement rate. Direct payment subsample; weaker outlier detection rule



(d) Replacement rate. Direct payment subsample; stronger outlier detection rule



Note: sample of workers with direct payment from the Social Security Institute. The Figure plots, in panels a and c, the RD estimates at the threshold determined by the baseline wage at the time of the worker's first STW event for the sample of CIG direct payment workers, dropping only observations with a replacement rate above 80% (weaker outlier detection rule), which is theoretically not possible. Panels b and d report the same RD estimates at the threshold, but further removing all observation with a monthly benefit more than 10% different from the theoretical value. The monthly benefit is rescaled by CIG intensity (the amount of time actually spent in CIG) so that the values refer to a 0-hour STW (all time spent in STW). The discontinuity is estimated with a local linear regression with different bandwidth on the left and on the right of the threshold selected through the optimal mean squared error robust estimator of Calonico et al. (2014, 2017). See Appendix Table A.2 for the full set of results.

either side.

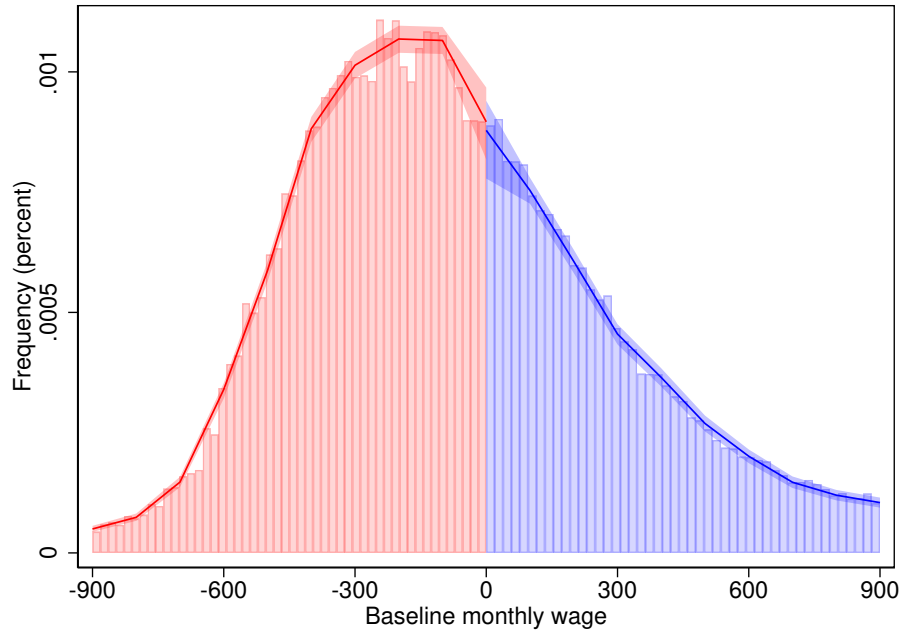
First, workers to the left of the threshold may be able to influence their employers to provide wage increases that take their earnings above the threshold and thus secure a higher CIG benefit. If such strategic behaviour were widespread, one would expect a higher concentration of observations immediately above the threshold. In contrast,

the absence of strategic wage manipulation would not manifest itself in a significant change in the frequency distribution above the threshold.

To evaluate the presence of discontinuities in the distribution of observations around the threshold, we apply a McCrary (2008) density test. This method assesses the continuity of the earnings density function at the threshold by fitting separate distributions to the data on either side. We adopt the approach proposed by Cattaneo et al. (2018, 2020), using a local polynomial estimator with asymmetric bandwidths to accurately reflect the density near the threshold.

Our analysis fails to reject the null hypothesis of no discontinuity in the density of observations, yielding p-values of 0.498 and 0.275 under local linear and quadratic estimation methods, respectively. This finding is further corroborated by visual evidence from a density plot (see Figure 3), which also does not indicate any manipulation of wages around the threshold.²¹

Figure 3: Distribution of workers by baseline wage (McCrary test)



Note: The Figure plots the distribution of workers' monthly wages at $t = 0$ in our main sample (binned). The discontinuity is estimated using the local linear regression and the robust mean squared error optimal bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold.

One related concern is that the likelihood of workers being placed in CIG may

²¹As noted above, the wage distribution local to threshold represents the modal part of the overall wage distribution (Appendix Figure 1).

change at the threshold. The absence of a discontinuity in the distribution of workers in our main sample provides indirect evidence against this possibility: if firms on one side of the threshold were systematically more likely to place workers in CIG, we should have observed excess mass on that side of the threshold among workers experiencing a CIG event. We also investigated this issue directly by using the full sample of worker-firm observations and estimating an RDD for the probability that a worker experiences a first CIG event with a given firm. Following a standard discrete-time survival approach, for each worker-firm spell we retain observations for all months in which the worker is at risk, up to the month preceding the first CIG event with that firm. We then define indicators for whether the worker enters CIG one, three, six, nine, or twelve months ahead.

Unfortunately, as discussed in Section 4.2, in the full sample of workers the running variable can only be constructed using the baseline wage measured annually in December and is therefore subject to measurement error. This is unlikely to prevent the test from being informative. In particular, using this alternative running variable, we previously found evidence of a discontinuity in the amount of CIG benefits despite the measurement error. We would therefore expect a sufficiently large discontinuity in the probability of entering CIG to remain detectable, although measurement error may attenuate or blur the estimated effect. The results, reported in Appendix Table A.5, provide no evidence of such a discontinuity: the estimated effects are small and statistically insignificant at all horizons.

Second, we test for differences in workers' and firms' observable characteristics across the threshold. Focusing on the individual socio-demographic characteristics available in the administrative archives, Tables A.3 and A.4 show that such differences are small and not statistically significant. The only exception is a small drop in the fraction of female workers at the threshold, which, however, is not statistically significant at conventional levels with the robust variance estimator (and only borderline so, with p-value 0.097, without the robust variance estimator).²² Third, we test that proxies of workers' and firms' productivity are continuous at the threshold, measured through the methodology first advanced by Abowd et al. (1999). The results, presented in more detail in Section 5.2.2 and reported in Table A.6, show no evidence of significant discontinuities at the threshold. Finally, in Table A.8 in the Appendix we also look at differences in local labour market characteristics, such as urbanisation, unemployment rate, vacancy rate and degree of tightness, and we find no evidence of

²²We replicate the main results restricting the sample to males only. The point estimates are largely unchanged, although they are less precise as we now rely on a 20% smaller sample. Results available upon request.

a discontinuity.

5 Results

5.1 The impact on job separations

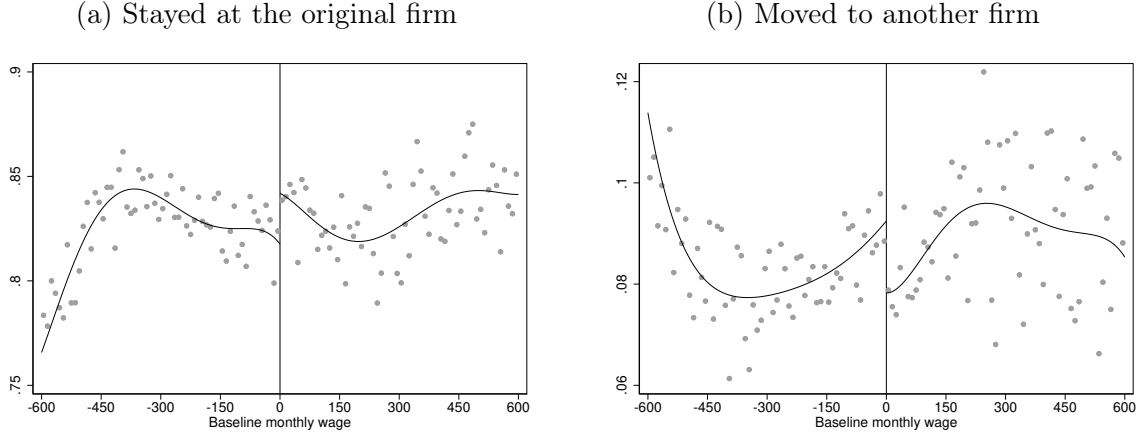
The first prediction of our theoretical framework is that workers are more likely to stay (and less likely to leave) if benefits are more generous. We therefore start by estimating the discontinuity at the threshold in the likelihood that the worker is still with the original firm 12 months after the CIG episode.

Following the steps suggested by Calonico et al. (2015), we first visually inspect the overall relationship between the outcomes and the running variable before proceeding to full RDD estimates. The left panel of Figure 4 plots the probability that the worker is still with the same firm 12 months after the beginning of her first CIG event against the baseline wage. It displays both the binned scatter and the regression function, separately on the left and on the right of the threshold, estimated through a global polynomial of order 5. There appears to be a significant increase in this probability above the threshold. The right panel of the Figure plots the same relationship, but for the probability that the worker moved to another firm. This probability declines significantly on the right-hand side of the threshold. These results are in line with the theoretical framework: an increase in benefit generosity reduces the likelihood that workers leave the firm after having been put on STW.

To estimate the size of the discontinuity, we use a local linear regression with a triangular kernel and separate bandwidths on the left and on the right of the threshold. We select the bandwidth according to the optimal mean squared error criterion of Calonico et al. (2014), and we report conventional RDD estimates, bias-corrected estimates, and bias-corrected estimates with a robust variance estimator. Results using polynomials of different orders and alternative bandwidth selection criteria and length are extremely similar (see Section 5.4 and Appendix Figures 5 and 6).

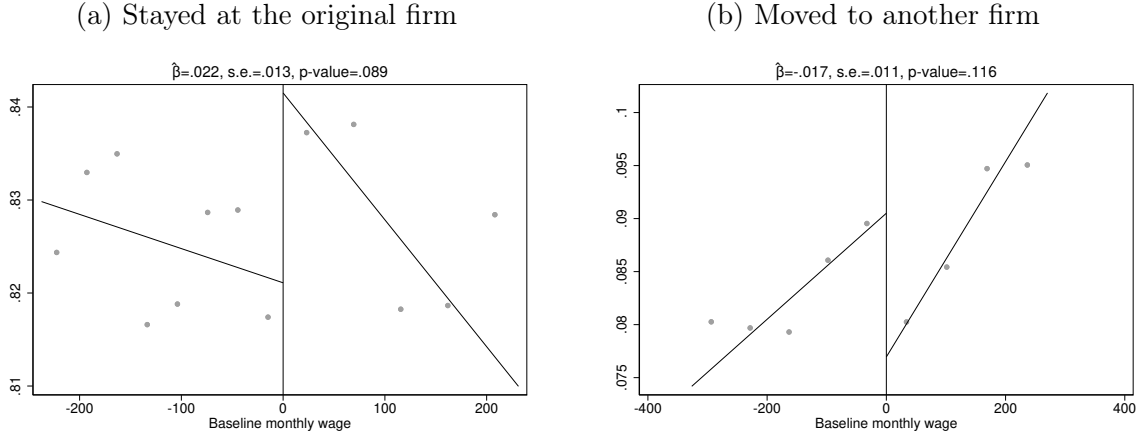
Figure 5 plots the RDD estimates. The regression lines are consistent with our theoretical predictions regarding the effect of the replacement rate. When the running variable increases, the replacement rate decreases, and so does the probability that workers stay with the same firm (panel a). Importantly, this relation is discontinuous at the threshold, where we know that the replacement rate increases sharply. Workers above the threshold have a 2.2 percentage point higher probability of staying at the original firm after 12 months of being put on STW, which is an increase of 2.7% compared to the baseline probability calculated to the left of the threshold. Given

Figure 4: Probability to stay and leave after 12 months



Note: Main sample. The Figure plots the binned plot and the regression functions of the main outcomes of interest as a function of the baseline wage at 12 months from the first worker's STW event. Bins are chosen at equally sized intervals of 10 euro. The regression function reported is estimated as a global polynomial regression of order 5 over the $[-500, +500]$ support and separately on either side of the threshold (as suggested by Calonico et al., 2015).

Figure 5: Discontinuity in the probability of staying and leaving after 12 months



Note: Main sample. The Figure plots the RD estimates of the main outcomes of interest at the baseline wage threshold after 12 months from the worker's first STW event. The discontinuity is estimated with a local linear regression with different bandwidth on the left and on the right of the threshold selected through the optimal mean squared error estimator of Calonico et al. (2014, 2017). For the binned plots, bins are chosen using the IMSE-optimal evenly-spaced method using polynomial regression proposed by Calonico et al. (2017). See Table 1 for the full set of results.

that after one year only 37% of workers who are still with the firm are also on STW, this implies that these jobs are predominantly still viable.

The increase in job retention translates almost one-to-one into a lower probability

Table 1: Discontinuity in the main outcomes at the threshold at $t = 12$

	Stayed at the original firm	Moved to a different firm	Not in salaried employment anymore
	(1)	(2)	(3)
Panel a. Conventional estimator			
$\beta_{threshold}$	0.020*	-0.014	-0.003
s.e.	0.012	0.010	0.006
p-value	0.083	0.160	0.577
Panel b. Bias-corrected estimator			
$\beta_{threshold}$	0.022*	-0.017*	-0.003
s.e.	0.012	0.010	0.006
p-value	0.061	0.080	0.594
Panel c. Robust estimator			
$\beta_{threshold}$	0.022*	-0.017	-0.003
s.e.	0.013	0.011	0.006
p-value	0.089	0.116	0.635
Actual obs. left	35020	48149	27524
Actual obs. right	24930	27996	29803
Left bandwidth	237.4	326.4	189.7
Right bandwidth	231.3	270.3	293.7
Avg. outcome left	0.821	0.090	0.086
Average outcome	0.816	0.093	0.092
Observations	145595	145595	145595

Note: Main sample. Standard errors are clustered at the firm level. The discontinuity is estimated using a local linear regression and the robust optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

of leaving the firm. This drop is estimated at 1.7 percentage points, which represents a sizeable 18.7% decrease relative to the baseline mean. However, the estimate is not statistically significant at conventional levels when using the robust estimator for the variance–covariance matrix.

The decline in the likelihood of moving is also sizeable when considering its impact on total job-to-job transitions, i.e., including workers who are not on CIG. To assess this, we examine all new hires using the full set of Social Security data, including also workers who have never been on CIG. We focus on hires of workers with previous employment experience (since 2008) and, consistent with our main estimates, we limit the analysis to those with a permanent full-time contract and a baseline wage within a

€200 range around the CIG threshold. On average, during the entire analysis period, 22.4% of the total hires in this selected group involved workers who had been in CIG at least once in the previous 12 months (this fraction was higher during the Great Recession in 2010, reaching 36.8%, but decreased to slightly more than 10% in 2018–19). Based on the results in Section 5.1, a 10 percentage point increase in the replacement rate would reduce the hires of CIG workers by 18.7%. This corresponds to reducing total hires of permanent workers with wages between €1,900 and €2,300 by 4.2% (i.e., $22.4\% \times 18.7\%$). Focusing specifically on the Great Recession period (2009–13), this would result in a 5.5% reduction in hires.

The last column of Table 1 reports the results for not being in salaried employment after 12 months (which represents the residual category of the first two outcomes). The effects are very close to zero and not statistically significant.²³

The theoretical model outlined in Section 3.1 suggests not only that higher benefits should reduce job-to-job transitions, but also that workers become more selective, opting to move only if higher wages are offered. Table A.9 in the Appendix shows that, among those who do leave, workers just above the threshold do receive slightly higher wages, even if the effect is not statistically significant. Therefore, drawing definitive conclusions on wage gains is difficult because the estimates are quite imprecise, which may stem from the limited sample size.

5.2 Heterogeneity by workers’ quality

As posited in our theoretical framework, the likelihood that workers will move depends on their outside options. Consistently with the empirical wage decomposition of Abowd et al. (1999), our model features an additively separable wage function in workers’ and firms’ fixed effects and a match-specific component. Hence, holding the wage constant, it must be that workers with a higher personal component (worker fixed effect) — which we may be willing to interpret as a proxy for quality — have received a negative match-specific draw. As our analysis focuses on a sharp discontinuity, in fact, we compare workers who receive virtually the same salary: around the threshold, high-AKM-effect workers are receiving the same wage as those with lower individual fixed effects. Therefore, high-AKM-type workers are more likely to move regardless of the replacement rate, because their current situation is unfavourable and they likely have better outside options, closer to their qualifications. Conversely, mid- or low-AKM-type workers are likely receiving a salary that is already close to their

²³The results in the three columns do not exactly sum to one because the estimates are non-linear and the effective bandwidth differ across outcomes.

outside option. Therefore, a change in the replacement rate is more likely to cross their indifference point and change their decision to leave the firm or not. Within the context of our theoretical model, this reasoning may be formalized by an increasing productivity density of outstanding vacancies around the reservation productivity thresholds of the different worker ability types.

To verify this in the data, we estimate the workers' AKM effects using the entire linked employer–employee sample dataset from the Italian National Social Security Institute, ranging from 1998 to 2019.²⁴ We exclude worker-year observations in which we observe a spell of STW to avoid the estimation of worker and firm effects to be affected by it. As is common in the literature, we also restrict the estimation to observations in the connected sample, i.e., the sample for which the mobility of workers creates a link across all firms.²⁵ We then split the workers into two groups according to whether their fixed effects is below or above the median in the final sample. Before running estimates by subsample, we test that the share of high/low effects is continuous at the threshold. Appendix Table A.6 shows no evidence of significant discontinuities in these shares at the threshold.²⁶ In Appendix Table 7 we also show odds ratio from a logit regression of the likelihood of moving to another firm at $t = 12$ on a dummy for worker fixed effects above the median, the log of the current wage and individual characteristics. Results support the assumption, imposed in section 3.1, that — holding the current wage fixed — high-ability workers are expected to move more frequently.

Table 2 shows that, in line with our conjecture, the effect of a higher replacement rate is stronger for workers with the lowest personal component.²⁷ The positive discontinuity in the likelihood of staying with the same firm is higher among workers with low fixed effects and shrinks closer to zero for workers with effects above the median. Symmetrically, the negative effect on the likelihood of moving to another firm is stronger among low-AKM fixed-effect types. There is no difference in the (small) effect on the likelihood of moving to non-employment. Similar results are obtained by splitting the sample by quintiles of worker fixed effect (see Appendix Table A.11).

Overall, these results suggest that the average effect of the replacement rate is mostly driven by a reduction in the job-to-job flows of workers with lower AKM effects.

²⁴In calculating the effects, we also control for year and region fixed effects, as well as a polynomial in age, a blue/white-collar indicator, whether the worker has a part-time or temporary job, whether they were a new hire, and the interaction of gender, the age polynomial, and contract type indicators.

²⁵We lose about 5% of observations of the main sample.

²⁶In the same table, we also confirm that the origin firms' AKM effects are not discontinuous around the threshold.

²⁷Conversely, Appendix Table A.12 shows no clear heterogeneity in the likelihood of moving to another firm in terms of the firm's wage component.

This is consistent with Proposition 2 from our theoretical model when the density of the productivity distribution of outstanding vacancies is increasing.

To gain a full picture of job-to-job mobility and labour reallocation in our setting, we can also look at destination firms, as high-AKM workers would have predominantly moved towards higher-paying firms. Indeed, job-to-job transitions tend to be towards firms with higher fixed effects. [TO BE CHECKED] In a 100-euro interval around the wage thresholds in our sample, the large majority (68%) of workers who effectively moved 12 months after the STW episode joined a firm in the top fixed-effects tercile, 24% one in the middle tercile, and only 8% an employer in the bottom tercile. We also look at differential effects by the quality of the firm of destination (as proxied by terciles of firms' AKM fixed effects). The results in Appendix Table A.13 are quite imprecise, but overall indicate that the discontinuity in the outcome is driven by workers to the right of the threshold being less likely to move to higher-quality firms.

Table 2: Discontinuity in the main outcomes at the threshold at $t = 12$, by worker fixed effects

	Worker fixed effect:	
	Below median (1)	Above median (2)
Panel a. Outcome: Stay at the original firm		
$\beta_{threshold}$	0.033**	0.013
s.e.	0.016	0.016
p-value	0.032	0.432
Panel b. Outcome: At a different firm		
$\beta_{threshold}$	-0.028**	-0.010
s.e.	0.011	0.016
p-value	0.016	0.524
Panel c. Outcome: Not in salaried employment		
$\beta_{threshold}$	-0.007	-0.004
s.e.	0.010	0.007
p-value	0.494	0.516
Observations	69307	69307

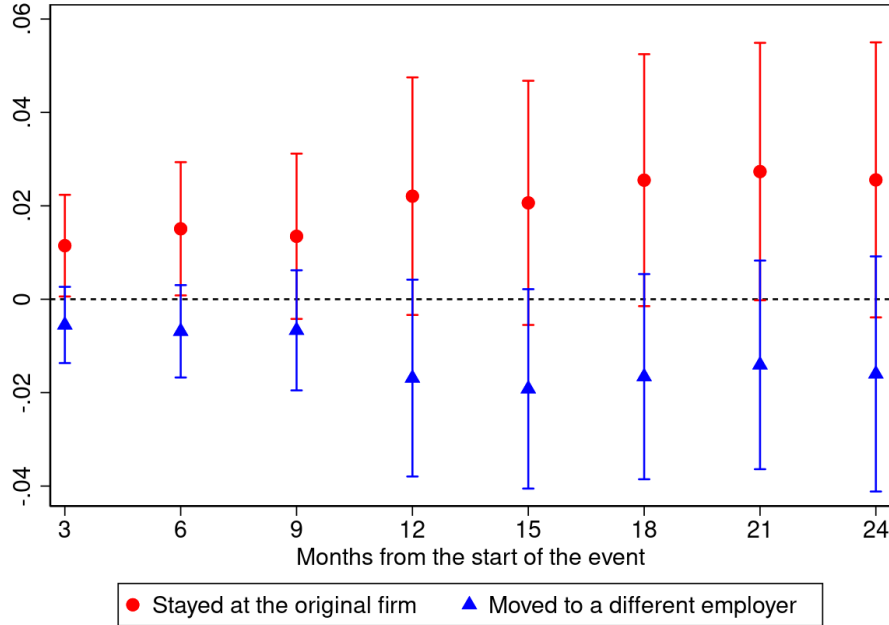
Note: Only connected sample. The sample is split according to whether the worker fixed effect is below or above the median. Standard errors are clustered at the firm level. The discontinuity is estimated using a local linear regression and the robust estimator and the optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

5.3 Additional results

5.3.1 Dynamic effects

Due to search frictions, genuine differences in worker search behaviour may take some time to materialize. Dynamic results are thus informative about the overall extent of job flows following the onset of STW at a firm. Figure 6 plots the coefficient β of equation (1) at different time horizons, and for the two main dependent variables of interest (Appendix Table A.10 reports the full set of estimates over time). The red dots identify the effect of the higher replacement rate on the probability that the worker is still with the same firm. This impact is positive and significant already in the first three months after the CIG event, but then further increases to 2.2 at $t = 12$ and eventually reaches 2.8 at $t = 24$. The impact on the probability that the worker joins a different firm is instead negative and also broadly grows over time. Finally, the effect on the probability that the worker is inactive is zero across the entire time span (see Appendix Table A.10).

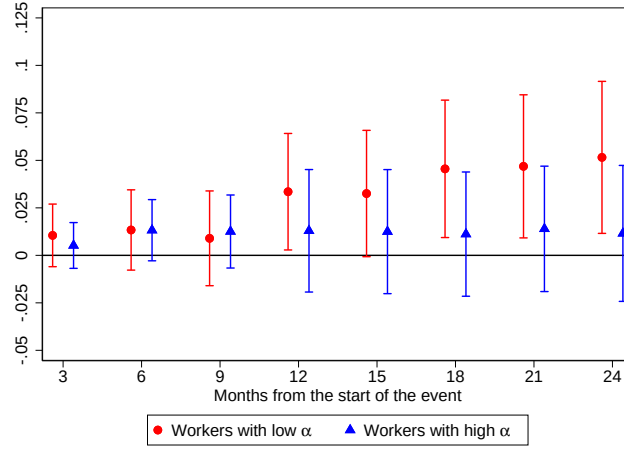
Figure 6: Discontinuity in the main outcomes at the threshold at different t



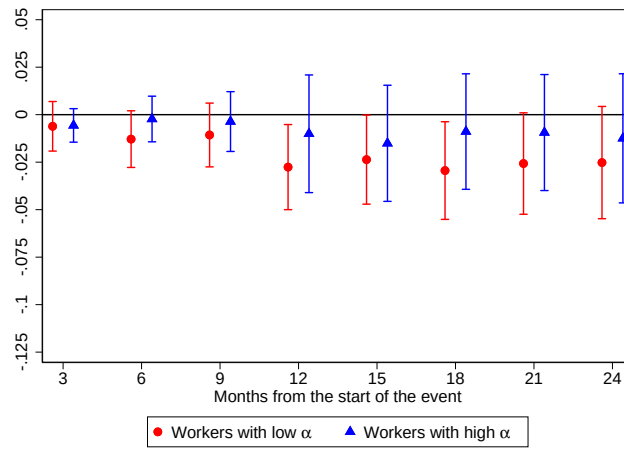
Note: Main sample. The Figure plots the estimated discontinuity in the main outcomes at the threshold at different time since the event, with a 95% confidence interval. The discontinuity is estimated using the biased-corrected robust RDD estimator of Calonico et al. (2017) with a linear local polynomial.

Figure 7: Discontinuity in the main outcomes at the threshold at different t , by worker fixed effects (α)

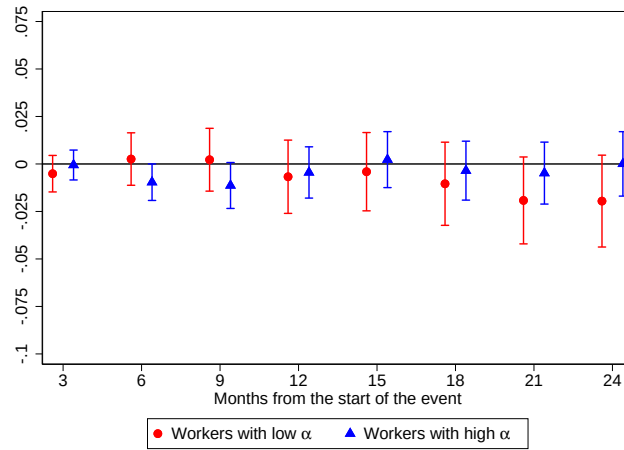
(a) Stayed at the original firm



(b) Moved to another firm



(c) Not in salaried employment anymore



Note: The Figure plots the estimated discontinuity in the main outcomes at the threshold at different time since the event, with a 95% confidence interval. See Table 2 for other details.

We also test for dynamic effects for different worker types. Figure 7 indicates that the likelihood of staying and of moving to another firm for those below the median is stronger and more distinct than the other group over time. This is not the case for the likelihood of moving to non-employment.

5.3.2 Effects by local labour market slack, age and sector

We also explore whether the results are heterogeneous across labour markets characterised by different levels of slack. Theoretically, the relationship between slack and the relevance of the replacement rate is ambiguous. On the one hand, a low meeting probability may dampen the effect of the replacement rate on the probability of moving. On the other hand, slack labour markets may feature a disproportionate incidence of low-productivity firms, which makes workers more reactive to the replacement rate as more outside options are close to their reservation productivity. In our empirical analysis, we use three different measures of labour market slack: the unemployment rate, the vacancy rate, and labour market tightness (vacancy rate over unemployment rate).²⁸ We estimate our baseline model separately for local labour markets above and below the median value of these variables. The results reported in Figure 8 show that the benefit has stronger effects when there is more labour market slack, i.e., when there is higher unemployment, a lower vacancy rate, and lower labour market tightness. In sum, workers in slack labour markets respond to the replacement rate similarly to workers with low AKM fixed effects, suggesting that the impact of the benefit is driven by local outside options around the worker’s reservation productivity.

In Figure 9 we report additional results. First, we show that the effects are primarily driven by younger workers. In general, those approaching retirement (aged 55 and over) have little incentive to move to other firms under any circumstances. As a result, the impact of the benefit on their probability of staying with the firm is close to zero.

Second, we turn to differences across sectors. The shock affecting the individual firm that decided to put its workers on STW might be correlated at the sectoral level. If most firms are hit by a similar shock and also largely utilise STW, it might be unlikely that workers can move to other employers. Consistently, Figure 9 shows that

²⁸The unemployment rate is estimated at the local labour market level (LLM) over 2006–2008, before the Great Recession kicks in to avoid potential concerns about congestion confounders; LLMs are sets of municipalities that are meant to proxy commuting areas and are defined by the Italian National Statistical Institute through an algorithm applied to 2011 Census data. The vacancy rate is the ratio between the total number of vacancies and the sum of vacancies and employed persons averaged over the period 2004–2008, again before the Great Recession, and measured at the regional level (data at the sub-regional level are not available; see Carta and Rizzica (2018)). Labour market tightness is measured as the ratio between the vacancy rate and the unemployment rate at the regional level, averaged over the same period, 2004–2008.

the effects are smaller in distressed sectors, i.e., where the average exit rate of firms exceeds the entry rate. The results are similar when we split the sample by whether the share of firms using CIG in the year of analysis is below or above the median.

5.4 Robustness

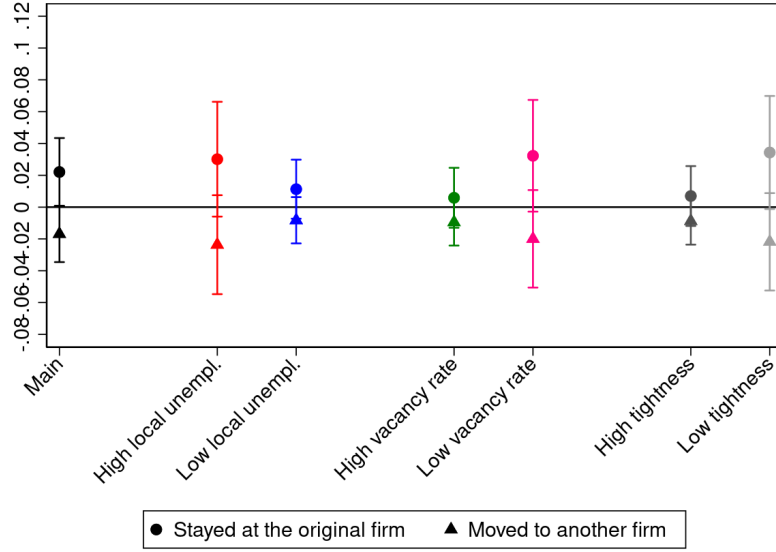
We consider different bandwidth selectors as proposed by Calonico et al. (2014), all estimated through a local linear regression: (i) the common bandwidth MSE estimator (MSERD 1), (ii) the coverage error-rate-optimal common bandwidth (CERRD 1), and (iii) the coverage error-rate-optimal bandwidth separate on either side of the threshold (CERRD 2). In all cases the results are very similar to our baseline that uses the mean squared error bias-corrected estimator and separate bandwidths on either side of the threshold. Last, we show the estimates using a simple local average (MSETWO 0) and a second-order polynomial for the local regression (MSETWO 2). Our baseline results are confirmed, although the simple average tends to underestimate the effects at the threshold. Similarly, results are unchanged if we include in the sample the very small firms that, in principle, should not have access to the CIG. These sensitivity checks can be found in Figure A.5 in the Appendix. Furthermore, we test the robustness of the estimates to selecting manually the bandwidth length, symmetric across the threshold (and for different degrees of the local polynomial). Although these latest checks are not preferable to the main estimates — because the bandwidth is chosen ad hoc and not selected optimally — the results are very similar to the main ones across different specifications (Figure 6).

To further assess the robustness of our results, we compare our main results with estimates of the regression discontinuity at different values of z_i where the benefit schedule is continuous. In these instances, we do not expect to identify any significant effect. Figures A.7 and A.8 in the Appendix report the estimated coefficients and the relative z -statistic for the probability of staying in the same firm and the likelihood of transitioning to another firm, respectively. The discontinuity at our cutoff (0) clearly stands out as the most statistically significant. While there are some other sizeable point estimates, they generally occur either at the left end of the distribution of the running variable — close to a kink in the benefit schedule — or at the right end of it, where observations are relatively sparse.

Finally, as anticipated in Section 3.2, we examined whether the results differ before 2013, when the unemployment benefit schedule closely mirrored the CIG threshold. Although the estimates remain imprecise, the point estimates in Appendix Figure A.9 are qualitatively consistent both before and after 2013. If anything, the effects appear

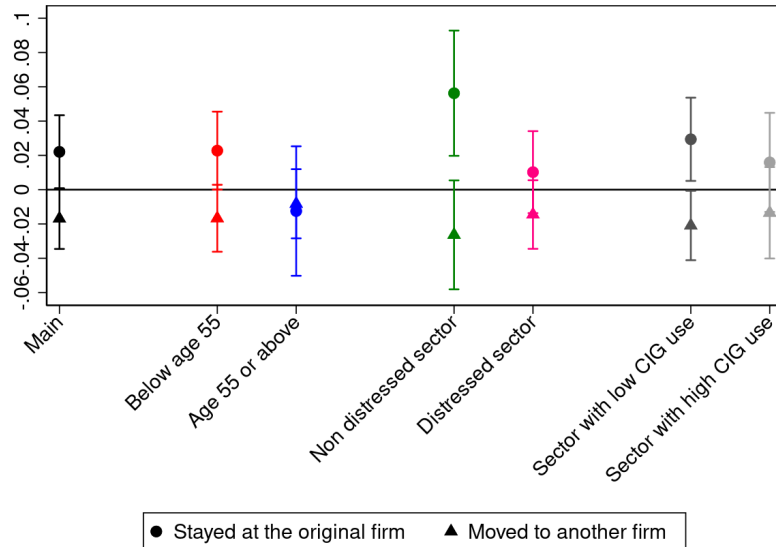
more pronounced after 2013, when the benefit schedule diverged. This suggests that this issue does not significantly affect our findings.

Figure 8: Heterogeneity in the main outcomes at $t = 12$ by labour market slack



Note: The Figure plots the estimated discontinuity in the main outcomes at the threshold at $t = 12$ for the main results (Figure 5 and Table 1) and in three sample splits: (i) local labour markets with above or below median unemployment rate; (ii) regions with above or below median vacancy rate; (iii) regions with above or below median tightness. The discontinuity is estimated using the biased-corrected robust RDD estimator of Calonico et al. (2014, 2017) with a locally linear polynomial.

Figure 9: Heterogeneity in the main outcomes at $t = 12$ by age and sector-level shocks



Note: The Figure plots the estimated discontinuity in the main outcomes at the threshold at $t = 12$ for the main results (Figure 5 and Table 1) and in three sample splits: (i) by workers' age at the time of the CIG spell onset; (ii) sectors (NACE 2 digit) where the average annual exit rate of firms is above the average annual entry rate, calculated over the period of analysis; (iii) sectors (NACE 2 digit) with above or below median share of firms in which the CIG spell starts. The discontinuity is estimated using the biased-corrected robust RDD estimator of Calonico et al. (2014, 2017) with a locally linear polynomial.

6 Some consideration about the optimal STW replacement rate

The theoretical framework outlined in Section 3.1 allows us to go one step further to understand the fiscal externality induced by the level of STW benefits. Paralleling the optimal UI literature, we study the problem of a benevolent social planner who sets the STW replacement rate to balance the social gains from redistributing consumption to poorer agents against the social costs of higher taxation, internalizing workers' optimal search behaviour described in Section 3.1. In doing so, we develop a heterogeneous-agent counterpart to the workhorse representative-agent model of Chetty (2006) and apply it to the STW replacement rate—rather than to the UI replacement rate.

Suppose that the government (i) sets a wage-specific STW replacement rate myopically to maximize workers' expected utility, and (ii) mandates how many hours h to work in STW.²⁹ The government tops up the wage of employees in STW, replacing $100\alpha(w)\%$ of their full-time hourly wage w for each hour not worked. Hence, the government spends $\tilde{\alpha}(w)w = \alpha(w)(1 - h)w$ for each employee, where $h \in (0, 1)$ denotes hours worked (paid by the firm) and w the full-time wage. The government finances STW subsidies and unemployment benefits through a payroll tax on full-time workers. Hence, differently from the baseline model introduced above, the STW replacement rate is wage specific, $\tilde{\alpha}(w)$, and full-time workers earn an after-tax wage $(1 - \tau)w$.

Let e denote the employment rate, μ the share of employed workers on STW, and $\gamma(w)$ the full-time wage density among workers on STW. Let $F(x, y, \epsilon; w)$ denote the joint distribution of worker ability, firm productivity, and the match-specific shock for workers on STW earning full-time wage w , and let $H_U(x)$ denote the ability distribution among unemployed workers. Let $R^*(x)$ denote the reservation productivity of an unemployed worker with ability x , with associated job-finding probability $f(x, R^*) \equiv \lambda [G(x) - G(R^*(x))]$.³⁰ The optimal wage-specific gross replacement rate maximizes

²⁹We implicitly assume that the mandated hours reduction induces optimal STW take-up for a given replacement rate.

³⁰The reservation productivity of an unemployed worker with ability x solves $\mathbb{E}_{\tilde{\epsilon}} [u((1 - \tau)w(x, R^*(x), \tilde{\epsilon}))] = u(b)$.

workers' expected utility ("social welfare") and solves:

$$\begin{aligned}
& \max_{\substack{\tilde{\alpha}(w) \in [0, 1-h], \\ \tau \in [0, 1]}} e(1-\delta) \left\{ \mu \int \gamma(w) \int [m(x, y^*) \mathbb{E}[u((1-\tau)w(x, \tilde{y}, \tilde{\epsilon})) \mid \tilde{y} \in (y^*, x)] \right. \\
& \quad + (1-m(x, y^*)) [\theta u((h + \tilde{\alpha}(w))w) + (1-\theta)u((1-\tau)w)] dF(x, y, \epsilon; w) dw \\
& \quad \left. + (1-\mu) \int u((1-\tau)w(x, y, \epsilon)) d\hat{F}(x, y, \epsilon) \right\} \\
& + \int [e\delta + (1-e)(1-f(x, R^*))] u(b) dH_U(x) \\
& + (1-e) \int f(x, R^*) \mathbb{E}[u((1-\tau)w(x, \tilde{y}, \tilde{\epsilon})) \mid \tilde{y} \in (R^*, x)] dH_U(x) \\
& \text{s.t. } \mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)\tau = \mathcal{B}(\mathbf{y}^*, \mathbf{R}^*, \tilde{\alpha}, b), \\
& \quad y^* = y^*(x, y, \epsilon; \tilde{\alpha}, \tau), \quad R^* = R^*(x; b, \tau),
\end{aligned}$$

where $(\mathbf{y}^*, \mathbf{R}^*)$ are vectors of reservation productivities for workers on STW and unemployed workers, and $\tilde{\alpha}$ is the vector of wage-specific gross replacement rates, respectively. Workers employed at the beginning of the period remain so with probability $1 - \delta$. A worker whose match is exogenously destroyed enters unemployment and receives unemployment benefits b . Among workers initially unemployed with ability x , a fraction $f(x, R^*)$ finds an acceptable full-time job, while the remaining fraction continues to receive unemployment benefits.

The government faces three constraints. First, the budget-balance constraint pins down the payroll tax rate τ required to finance the gross replacement rate schedule $\tilde{\alpha}(w)$ and unemployment benefits. The tax base $\mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)$ and the total outlays

$\mathcal{B}(\mathbf{y}^*, \tilde{\alpha})$ are defined as follows:

$$\begin{aligned} \mathcal{T}(\mathbf{y}^*, \mathbf{R}^*) = e(1 - \delta) & \left\{ \mu \int \gamma(w) \int \left[m(x, y^*) \mathbb{E}[w(x, \tilde{y}, \tilde{\epsilon}) \mid \tilde{y} \in (y^*, x)] + (1 - m(x, y^*)) \right. \right. \\ & \left. \left. (1 - \theta)w \right] dF(x, y, \epsilon; w) dw + (1 - \mu) \int w(x, y, \epsilon) d\hat{F}(x, y, \epsilon) \right\} \\ & + (1 - e) \int f(x, R^*) \mathbb{E}[w(x, \tilde{y}, \tilde{\epsilon}) \mid \tilde{y} \in (R^*(x), x)] dH_U(x), \end{aligned}$$

$$\begin{aligned} \mathcal{B}(\mathbf{y}^*, \mathbf{R}^*, \tilde{\alpha}, b) &= e(1 - \delta) \int \mu \gamma(w) \theta \int (1 - m(x, y^*)) dF(x, y, \epsilon; w) w \tilde{\alpha}(w) dw \\ &+ \int [e\delta + (1 - e)(1 - f(x, R^*))] b dH_U(x) \\ &= e(1 - \delta) \int B(w) \tilde{\alpha}(w) dw + \int [e\delta + (1 - e)(1 - f(x, R^*))] b dH_U(x), \end{aligned}$$

where $\hat{F}(x, y, \epsilon)$ denotes the joint distribution of worker ability, firm productivity, and the match-specific shock for employed workers *not* on STW, and $B(w) \equiv \mu \gamma(w) \theta \int (1 - m(x, y^*)) dF(x, y, \epsilon; w) w$ is the wage-specific STW subsidy base. Second, the government internalizes the static reservation productivity rule y^* followed by employed workers in their mobility decisions. Third, it internalizes the reservation productivity rule R^* followed by unemployed workers in their job-acceptance decisions.

For a given tax rate, the direct welfare effects associated with the endogenous adjustment of the reservation productivities are summarized by the following lemma.

Lemma 1

For a given tax rate, workers' endogenous adjustment of their reservation productivities has zero first-order direct effect on social welfare.

For employed workers, this result follows from the fact that y^* makes a worker indifferent between remaining at the current STW-using firm and moving to a full-time job at a firm with productivity y^* . For unemployed workers, R^* makes a worker indifferent between remaining unemployed and accepting a full-time job at a firm with productivity R^* . Consequently, changes in either threshold have no first-order direct effect on workers' expected utility choice – just like optimal search effort in the standard Baily-Chetty formula. The proof is provided in Appendix B.2. The government nevertheless internalizes the indirect fiscal effects of workers' mobility decisions through STW expenditure and the payroll-tax base.

In Appendix B.2, we show that the government optimality condition pinning down the wage-specific gross replacement rate reflects the Baily–Chetty trade-off between

the marginal value from redistribution (RV) and the moral hazard cost (MH):

$$RV(\tilde{\alpha}(w)) = 1 + MH(\tilde{\alpha}(w)). \quad (2)$$

The left-hand side of (2) represents the marginal social benefit from redistributing consumption shares to poorer workers through a higher STW subsidy. The right-hand side equals the fiscal cost of raising the replacement rate, including both the mechanical cost of the additional transfer and the fiscal externality generated by workers' behavioural response. Intuitively, by making the STW subsidy more generous, workers move less frequently away from STW-using firms, increasing STW outlays and reducing payroll-tax revenues.

Let $\mathcal{V}^B(w)$ denote the expected indirect utility of workers with full-time wage w remaining on STW, and let $\mathcal{V}^T$ denote the expected indirect utility of workers in full-time jobs and therefore paying the payroll tax. The marginal social value from redistribution is:³¹

$$RV(\tilde{\alpha}(w)) \equiv \frac{\partial \mathcal{V}^B(w)/\partial \tilde{\alpha}(w)}{\partial \mathcal{V}^T/\partial \tau} \frac{\mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)}{B(w)}.$$

Notice that the marginal expected indirect utility of each group of agents is normalized by the respective population measure.

If workers were risk neutral, the marginal social value from redistribution would equal one. Intuitively, under risk neutrality, there is no social gain from redistributing consumption shares to poorer agents. By contrast, under risk aversion, STW provides a positive social gain from redistribution as long as subsidized workers are poorer than the average taxpayer. The optimal replacement rate is attained when this social gain from redistribution exactly offsets the fiscal externality, which depends on the moral hazard cost:

$$MH(\tilde{\alpha}(w)) = \int q(x, y^*) \epsilon_{1-m(x, y^*), \tilde{\alpha}(w)} \left(1 + \frac{\mathbb{E}[w(x, y^*, \tilde{\epsilon})] - (1 - \theta)w}{\theta \tilde{\alpha}(w)w} \tau \right) dF(x, y, \epsilon; w),$$

where $q(x, y^*) \equiv \frac{1-m(x, y^*)}{\int [1-m(x, y^*)] dF(x, y, \epsilon; w)}$ is the probability of staying at the current firm relative to the average among workers on STW at the same full-time wage. As in the standard Baily–Chetty formula, the moral hazard cost depends on the elasticity of the probability of remaining in the subsidized state with respect to benefit generosity, $\epsilon_{1-m(x, y^*), \tilde{\alpha}(w)} \equiv \frac{\partial(1-m(x, y^*))/\partial \tilde{\alpha}(w)}{1-m(x, y^*)} \tilde{\alpha}(w)$. Unlike the standard Baily–Chetty formula,

³¹As is standard in the literature, we assume that the marginal change in the tax base from the reservation productivity adjustment induced by a change in the payroll tax rate is negligible.

this elasticity is weighted by an opportunity-cost multiplier reflecting the reduction in tax revenues associated with longer STW spells, $1 + \frac{\mathbb{E}[w(x, y^*, \bar{\epsilon})] - (1 - \theta)w}{\theta \bar{\alpha}(w)w} \tau$. If STW were an absorbing state for subsidized firms ($\theta \rightarrow 1$) and there were no wage dispersion ($w = \bar{w}$), the multiplier would reduce to $\frac{\bar{\alpha}(w) + \tau}{\bar{\alpha}(w)}$, as in Kolsrud et al. (2018) for UI benefits.

The behavioural response entering the moral hazard cost can also be expressed in terms of the expected duration of STW. With constant transition probabilities, a worker leaves STW either by moving to another firm, which occurs with probability $m(x, y^*)$, or by remaining at the current firm if the latter exits STW, which occurs with probability $(1 - m(x, y^*))(1 - \theta)$. Expected STW duration is therefore: $D(x, y^*) = [m(x, y^*) + (1 - m(x, y^*))(1 - \theta)]^{-1} = [1 - \theta(1 - m(x, y^*))]^{-1}$. Taking the elasticity with respect to the wage-specific gross replacement rate gives:

$$\epsilon_{D(x, y^*), \bar{\alpha}(w)} \equiv \frac{\partial D(x, y^*)}{\partial \bar{\alpha}(w)} \frac{\bar{\alpha}(w)}{D(x, y^*)} = \epsilon_{1 - m(x, y^*), \bar{\alpha}(w)} \frac{\theta(1 - m(x, y^*))}{1 - \theta(1 - m(x, y^*))}.$$

Thus, the elasticity of STW duration is proportional to the elasticity of the probability of remaining at the current firm. The proportionality factor depends on the persistence of STW and on the probability of remaining with the current employer.

Quantitative analysis. The optimality condition for the replacement rate (2) requires information on unemployment risk, UI expenditure, unemployment-to-employment transitions, and the wages of workers entering employment. Because our empirical variation identifies the mobility response of workers already on STW, we conduct the quantitative exercise in the limit as unemployment risk vanishes, by letting $e \rightarrow 1, \delta \rightarrow 0$. Under this approximation, the terms involving $R^*, f(x, R^*), H_U(x)$, and UI expenditure disappear. Note that this approximation only affects the baseline level of government outlays and the tax rate. It does not shut down any of the forces underlying the policy trade-off (2). For STW to have a nonzero value from redistribution, we need risk-averse workers. For parsimony of data requirements, we perform our calibration by assuming log utility—a conservative benchmark with respect to standard micro-estimates of the degree of relative risk aversion. Moreover, we assume that the responses of both the probability of moving and the average log wage conditional on moving to the replacement rate are approximately linear around the discontinuity in the replacement-rate schedule, which allows us to map our causal estimates into the theoretical derivatives.

Under log utility, the marginal social value from redistribution, i.e., the left-hand

side of equation (2), is:

$$RV(\tilde{\alpha}(w)) = \frac{\mu \int (1 - m(x, y^*)) \frac{\theta}{h + \tilde{\alpha}(w)} dF(x, y, \epsilon; w) \gamma(w)}{\mu \int \int \left[(1 - m(x, y^*)) \frac{1-\theta}{1-\tau} + m(x, y^*) \frac{1}{1-\tau} \right] dF(x, y, \epsilon; \tilde{w}) \gamma(\tilde{w}) d\tilde{w} + \frac{1-\mu}{1-\tau} B(w)} \frac{\mathcal{T}(\mathbf{y}^*)}{B(w)}.$$

Since the fiscal externality only features observable moments, all the objects entering the government FOC have empirical counterparts in our data. Table 14 in the Appendix provides the mapping between empirical moments and theoretical objects.

Quantitatively, we find that the marginal value from redistribution of STW at the discontinuity equals 1.33, while the fiscal cost equals 1.15. The latter is driven by an estimated elasticity of the probability of remaining in the current STW-using firm with respect to the replacement rate, $\epsilon_{1-m(x, y^*), \tilde{\alpha}(w)}$, of 0.15, mapping into an elasticity of STW duration of 0.07. Hence, even under the conservative assumption of log utility, the marginal value from redistribution locally exceeds the fiscal cost by a factor of 1.16. Since workers on STW at the discontinuity earn less than the average full-time worker, the marginal value from redistribution locally decreases with the replacement rate. This implies that the replacement rate at the threshold is below its optimal level.

Discussion. Our estimated local elasticity of remaining with the current employer in STW to the replacement rate of 0.15 implies an elasticity of STW duration of 0.07. The latter is lower than standard estimates of the elasticity of unemployment duration to UI generosity. Reviewing estimates from the United States, Dube (2021) concludes that a likely range for the elasticity of unemployment duration to benefit generosity is between 0.1 and 0.9, with a central estimate of 0.5. For Western Europe, the review by Card et al. (2015) points to estimated elasticities between 0.3 and 2. However, our estimated partial effect of STW generosity on the job-to-job transition probability is remarkably similar to that estimated by Grindaker and Simmons (2024) with respect to UI benefits.³²

The condition for the optimal replacement rate we derive assumes that STW take-up is a firm-level decision that individual workers take as given, as in Niedermayer and Tilly (2016). This is a reasonable assumption in the Italian setting, as firms cannot discriminate across employees when taking up STW. By explicitly modeling the worker-level response to the STW replacement rate in terms of job-to-job transitions, our framework allows us to quantify the reallocation effects of STW *driven by workers' behaviour*. Our partial-equilibrium approach, however, holds firms' job creation

³²Grindaker and Simmons (2024) find that a 1-per cent increase in UI benefits lowers the annual probability of a job-to-job transition by approximately 0.1 percentage points. We estimate a partial effect of similar magnitude in our data.

decisions and the contact rate fixed and therefore lacks a notion of allocative efficiency and congestion externalities. Extending the framework to general equilibrium would incorporate the response of labour market tightness through changes in the meeting probability λ . Whether increasing the STW replacement rate improves or deteriorates allocative efficiency and congestion externalities would then likely depend on the direction of the baseline inefficiency (Cooper et al., 2017; Landais et al., 2018).

Taking firm-level decisions as given also means that our baseline analysis abstracts from the responses of the share of workers in STW (μ) and hours worked in STW (h) to changes in wage-specific replacement rates.³³ Focusing on these margins, Giupponi and Landais (2023) estimate a firm-side fiscal-cost multiplier of 1.38 for the Italian STW policy. To relate their estimate to ours, we now describe how our model can be extended to allow firms to choose hours worked and STW take-up.

Suppose that, for a given statutory replacement-rate schedule α , firms choose optimal hours worked according to

$$h^*(w; \alpha) = \arg \max_{h \in [0,1]} \Pi(h; w, \alpha).$$

Both the gross STW replacement rate, $\tilde{\alpha}(w) \equiv \alpha(w) [1 - h^*(w; \alpha)]$, and the share of workers on STW, $\mu^*(\alpha)$, then become endogenous. The planner continues to choose the replacement-rate schedule, but does so subject to firms' optimal hours and take-up decisions, as well as workers' optimal mobility decisions. Letting $\mathcal{V}$ denote the planner's objective, the extended problem can be written compactly as

$$\max_{\alpha, \tau} \mathcal{V}(\mathbf{y}^*, \tilde{\alpha}, \mathbf{h}^*, \mu^*; \tau) \quad \text{s.t.} \quad \tau \mathcal{T}(\mathbf{y}^*, \mathbf{h}^*, \mu^*) = \mathcal{B}(\mathbf{y}^*, \alpha, \mathbf{h}^*, \mu^*).$$

The planner therefore internalizes the responses of firms' hours and take-up choices to the replacement-rate schedule. Holding the tax rate fixed, the induced change in the government's financing requirement can be decomposed as

$$\begin{aligned} \left. \frac{d\mathcal{B}}{d\alpha(w)} \right|_{\tau} - \tau \left. \frac{d\mathcal{T}}{d\alpha(w)} \right|_{\tau} &= \frac{\partial \mathcal{B}}{\partial \alpha(w)} - \tau \frac{\partial \mathcal{T}}{\partial \alpha(w)} \\ &+ \int \left[\frac{\partial \mathcal{B}}{\partial h(\tilde{w})} - \tau \frac{\partial \mathcal{T}}{\partial h(\tilde{w})} \right] \frac{\partial h^*(\tilde{w}; \alpha)}{\partial \alpha(w)} d\tilde{w} + \left[\frac{\partial \mathcal{B}}{\partial \mu} - \tau \frac{\partial \mathcal{T}}{\partial \mu} \right] \frac{\partial \mu^*(\alpha)}{\partial \alpha(w)}. \end{aligned}$$

The first line on the right-hand side contains the mechanical effect and the worker-mobility response studied in our framework, while the second line captures the fiscal

³³These margins may be relevant if workers and firms share the additional surplus generated by the STW subsidy through bargaining, both in terms of windfall effects (Albertini et al., 2022; Cahuc et al., 2024) and savings on UI benefits (Braun and Brügemann, 2014; Stiepelmann, 2024).

effects of firms' endogenous hours and take-up choices. Under one-worker-one-firm matches and in the absence of cross-wage responses, only firms employing workers at wage w respond to a pointwise change in $\alpha(w)$. The hours component then corresponds closely to the sufficient statistic derived by Giupponi and Landais (2023). Letting $n(w)$ denote STW headcount employment at wage w and normalizing full-time hours to one, their wage-specific firm-side fiscal-cost multiplier can be written as

$$FE^F(w) = 1 + \epsilon_{n(w),\alpha(w)} \left[1 - \frac{b}{\alpha(w) [1 - h^*(w)] w} \right] - \epsilon_{h^*(w),\alpha(w)} \frac{h^*(w)}{1 - h^*(w)}.$$

The second term captures the employment response and the associated savings on UI benefits, while the last term captures the additional STW expenditure generated by firms' endogenous reduction in hours. See Appendix B.2 for the full derivation.

Giupponi and Landais (2023) use the discrete effect of access to STW on firms' hours and employment to estimate these two components in response to a common change in benefit generosity. By contrast, we identify the reduced-form worker-mobility response to a local change in $\alpha(w)$. Because the latter is estimated from observed mobility outcomes, we interpret it as a *total* worker response, i.e., already incorporating any effect of the local policy change on hours worked. Since the policy experiments are different, an exact pointwise comparison would require wage-specific estimates of firms' hours and employment responses. As a back-of-the-envelope calculation, if their average firm-side multiplier also applies locally at the wage threshold and the worker- and firm-side margins are additive when normalized by the same mechanical expenditure, the combined fiscal-cost multiplier is

$$FE^{\text{total}} \simeq 1 + \underbrace{0.15}_{\text{workers}} + \underbrace{0.38}_{\text{firms}} = 1.53.$$

Under this approximation, worker search behaviour accounts for approximately 30% of the combined net fiscal externality. The worker-mobility margin is therefore quantitatively meaningful even relative to the net firm-side response. While each endogenous margin, considered in isolation, points to limited moral hazard, together they provide a more complete measure of the fiscal externality generated by STW schemes, bringing the resulting multiplier closer to those typically estimated for UI systems.

7 Conclusions

Short-time work schemes are widely implemented across OECD countries to help firms address temporary crises without resorting to layoffs, thereby preserving ex-

isting matches. From the workers' perspective, the wage-replacement benefit provided by these schemes is crucial for minimising welfare losses and facilitating consumption smoothing. This paper uncovers a new aspect of STW benefits. As these benefits become more generous, they increase workers' incentives to remain with their current employer, thereby reducing job-to-job flows. While this facilitates the preservation of existing job matches, it also constrains labour mobility, potentially limiting the reallocation of workers to better-performing firms.

We develop this intuition within a model of on-the-job search, where workers face the possibility of remaining in STW for an extended period (irrespective of their own will), and set their reservation productivity for alternative employers to warrant a job-to-job transition. Our findings are consistent with the model's predictions: an increase in the replacement rate discourages job-to-job transitions.

On the one hand, this outcome strengthens the primary goal of STW schemes, i.e., preserving existing job matches. On the other hand, our analysis shows that the associated reduction in labour mobility mostly affects workers with low earning potential, who might otherwise have transitioned to higher-paying firms.

We derive a Baily–Chetty-type optimal benefit formula within our framework, highlighting that the key fiscal externality hinges on the elasticity of remaining in STW in response to benefit generosity. Our estimates indicate that this elasticity is around 15 per cent, which is relatively low compared to estimates for unemployment benefits. Our findings complement existing literature centred on firm decisions by documenting significant, albeit quantitatively limited, moral hazard induced by STW schemes on worker search behaviour.

Overall, our results suggest that the fiscal cost of more workers remaining in STW due to an increased replacement rate is relatively small. However, this comes with a reduction in job-to-job flows among low-AKM FE workers, leading to a deterioration in the composition of the workforce in distressed firms, and a reduction in flows towards better-paying firms. Therefore, decisions regarding the generosity of STW benefits should carefully balance the need for preserving both job matches and workers' consumption smoothing against the potential disincentives for workers to move across jobs. This latter outcome would be particularly costly if job-to-job mobility towards more dynamic and thriving firms is impaired, thus contributing to hindering labour reallocation during an economic recovery.

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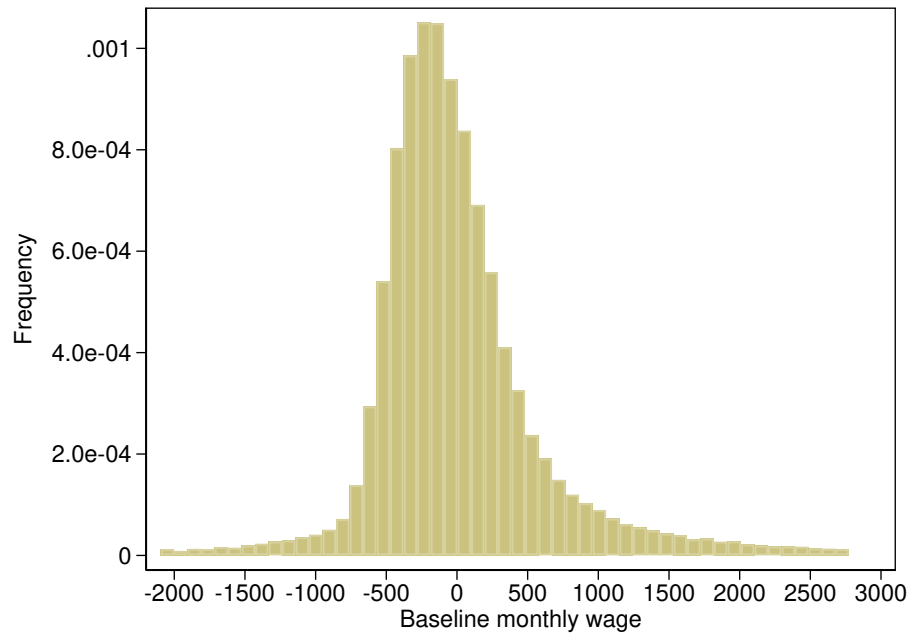
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Appendix

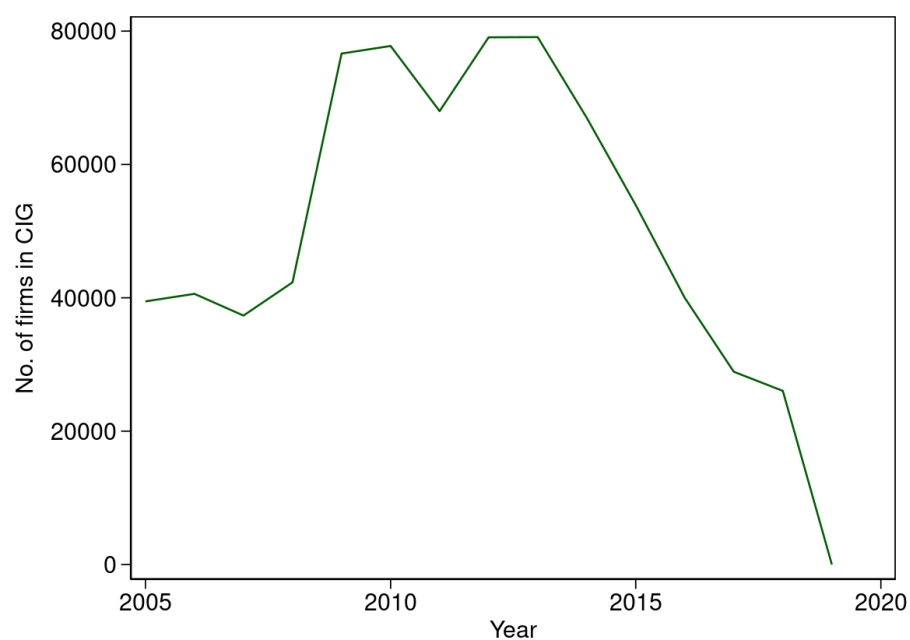
A. Additional figures and tables

A.1: Baseline monthly wage distribution for all workers in STW



Note: The histogram reports the distribution of the baseline monthly wages in the sample of workers we observe in STW in the data.

A.2: Firms in CIG between 2005 and 2019; non-farm private sector



Note: The number of firms reported here corresponds to all Italian firms using CIG for at least one hour as reported in the Italian National Social Security Institute data. Therefore, it does include also firms that are not part of our sample.

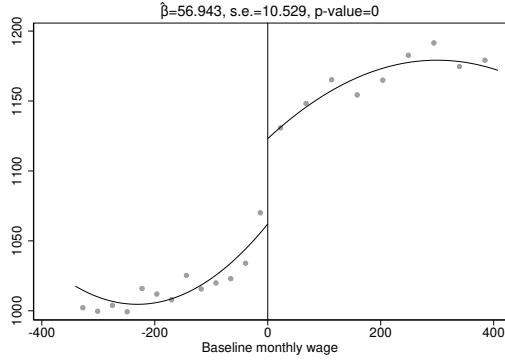
A.3: Differences in the observed and theoretical benefits paid to the workers in the direct payment sample



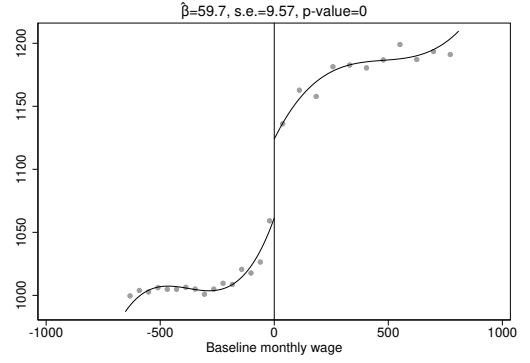
Note: The figure reports the theoretical benefit schedule along with the observed one in the direct payment sample (local polynomial fit of order zero). It highlights the errors in the implied replacement rate in this subsample of data.

A.4: Robustness on the discontinuity in the benefit at the threshold

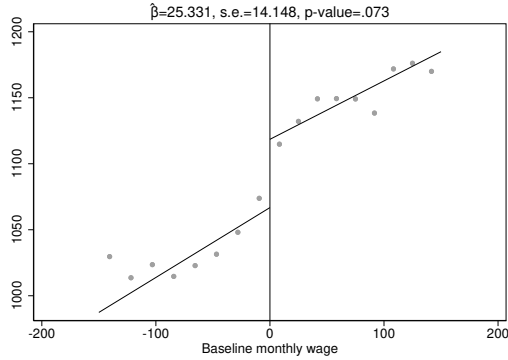
(a) Monthly benefits. Direct payment subsample; weaker outlier detection rule; $p=2$



(b) Monthly benefits. Direct payment subsample; weaker outlier detection rule; $p=3$



(c) Monthly benefits. Direct payment subsample; weaker outlier detection rule; $bw=150$

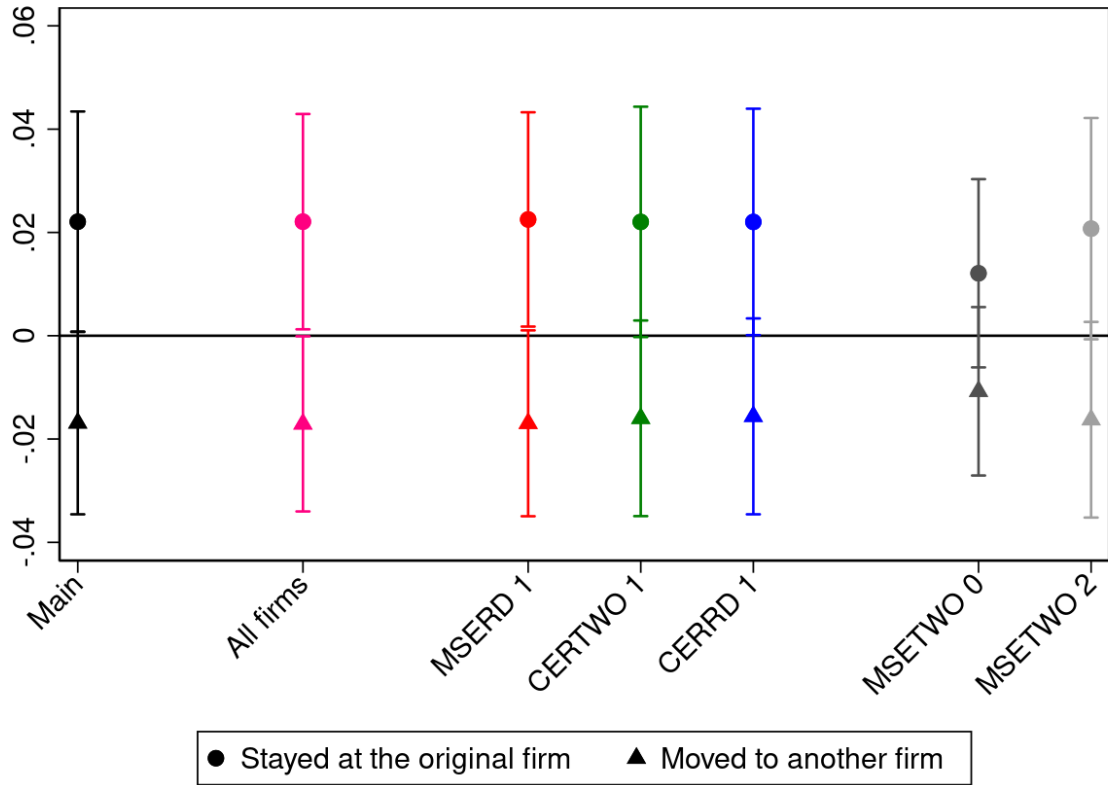


(d) Monthly benefits. Direct payment subsample; weaker outlier detection rule; $bw=500$



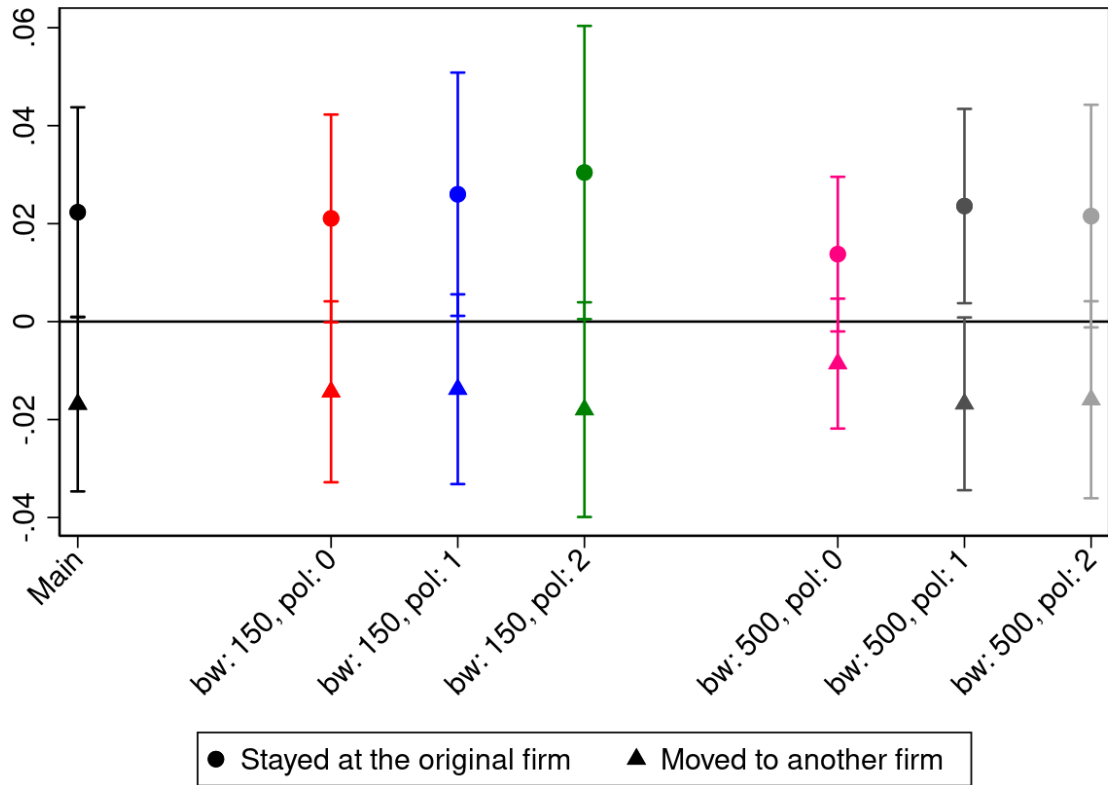
Note: sample of workers with direct payment from the Social Security Institute. The Figure presents several robustness checks on the the RD estimates at the threshold determined by the baseline wage at the time of the worker's first STW event for the sample of CIG direct payment workers, dropping only observations with a replacement rate above 80% (weaker outlier detection rule), which is theoretically not possible (panel a of Figure 2). Panels a and b plot the checks in terms of polynomial degree of the local regressions ($p=2$ and $p=3$, respectively). Panels c and d report the same RD estimates at the threshold, but in terms of different bandwidth (150 and 500 euro bandwidth symmetric around the threshold, respectively). Note that these bandwidths are chosen manually and are inferior to the method used for the main estimates (i.e., optimal mean squared error robust estimator of Calonico et al. (2014, 2017) on either side of the threshold).

A.5: Discontinuity in the main outcomes at the threshold at $t = 12$: Robustness



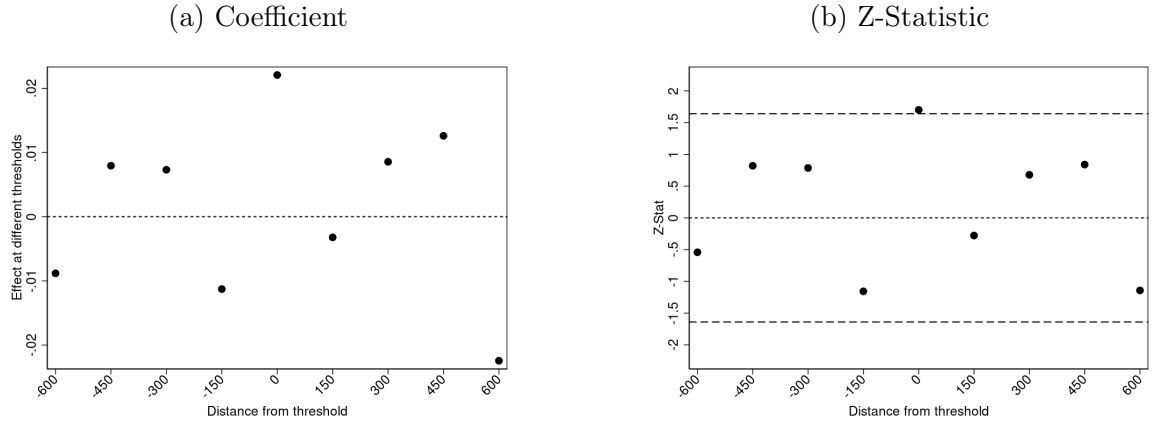
Note: The figure reports the RD estimates of the outcomes of interest at 12 month after the CIG event, i.e., being at the same firm and separate and join a different firm, for the following specifications: (i) the main specification as reported in Table 1 and Figure 5, (ii) the full sample of firms, also including very small firms, (iii) the common bandwidth MSE estimator and local linear regression (MSERD 1), (iv) the coverage error-rate-optimal common bandwidth and local linear regression (CERRD 1), (v) the coverage error-rate-optimal bandwidth separate on either side of the threshold and local linear regression (CERRD 2), (vi) a simple local average (polynomial of order 0, MSETWO 0) and (vii) a second-order polynomial for the local regression (MSETWO 2). The reported intervals correspond to the 95 confidence level.

A.6: Discontinuity in the main outcomes at the threshold at $t = 12$: Robustness on the bandwidth and degree of polynomial



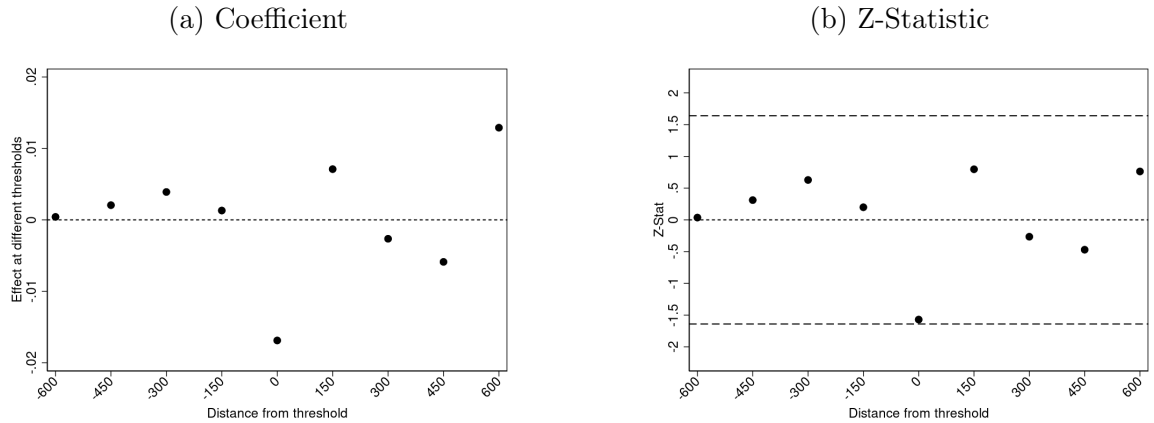
Note: The figure reports the RD estimates of the outcomes of interest at 12 month after the CIG event, i.e., being at the same firm and separate and join a different firm, for different bandwidth and degree of polynomial: (i) the main specification as reported in Table 1 and Figure 5, (ii) 150 euro hand-selected symmetric bandwidth and degree 0, (iii) 150 euro hand-selected symmetric bandwidth and degree 1, (iv) 150 euro hand-selected symmetric bandwidth and degree 2, (v) 500 euro hand-selected symmetric bandwidth and degree 0, (vi) 500 euro hand-selected symmetric bandwidth and degree 1 and (vii) 500 euro hand-selected symmetric bandwidth and degree 2. The reported intervals correspond to the 95 confidence level.

A.7: Discontinuity in the probability of staying at the original firm at $t = 12$ at different thresholds



Note: Main sample. The Figure plots the estimated discontinuities in the probability of staying at the original firm at 12 months since the beginning of the CIG event, estimated at different thresholds. The discontinuity is estimated using the biased-corrected robust RDD estimator of Calonico et al. (2017) with a linear local polynomial. The left panel plots the estimated coefficients, the right one plots the related z -statistic. The area within the dashed-lines identifies the values such that $p > 0.10$

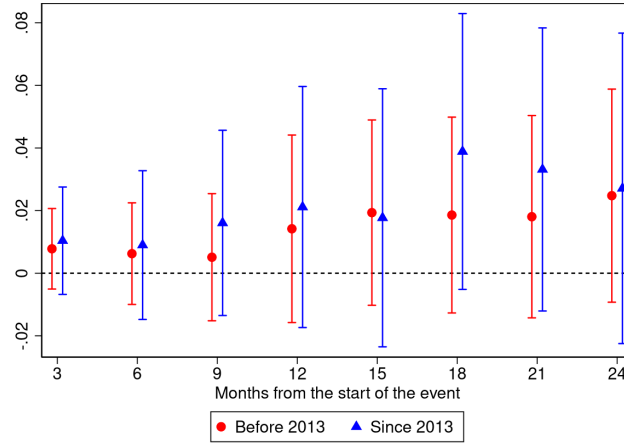
A.8: Discontinuity in the probability of moving to a different employer at $t = 12$ at different thresholds



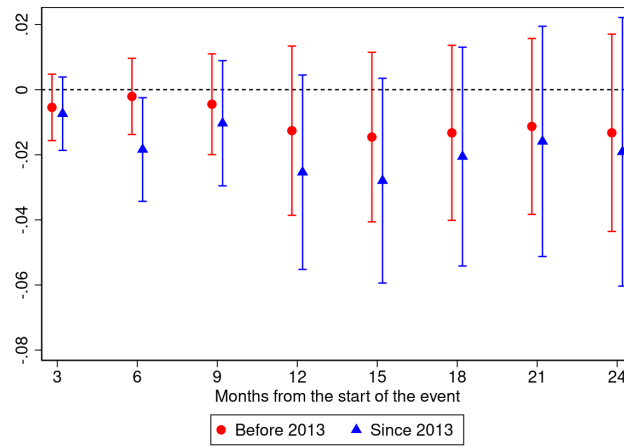
Note: Main sample. The Figure plots the estimated discontinuities in the probability of moving to a different employer in 12 months since the beginning of the CIG event, estimated at different thresholds. The discontinuity is estimated using the biased-corrected robust RDD estimator of Calonico et al. (2017) with a linear local polynomial. The left panel plots the estimated coefficients, the right one plots the related z -statistic. The area within the dashed-lines identifies the values such that $p > 0.10$

A.9: Discontinuity in the main outcomes at the threshold at different t , before and after 2013

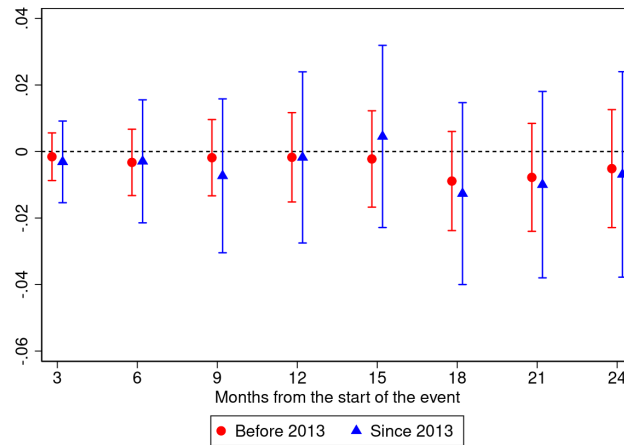
(a) Stayed at the original firm



(b) Moved to another firm



(c) Not in salaried employment anymore



Note: The Figure plots the estimated discontinuity in the main outcomes at the threshold at different time since the event, with a 95% confidence interval. See Table 2 for other details.

A.1: Descriptive statistics

	Selected sample	50-euro interval around the threshold
Share manufacturing	0.64	0.62
Share born in Italy	0.78	0.80
Average age	41.73	42.35
Share female	0.17	0.13
Average worker fixed effect (logs)	-0.02	-0.05
Average firm fixed effect (logs)	0.20	0.22
Average log firm employment size	4.44	4.57
Average log firm average wage	7.56	7.60
Average local labour market unemployment rate (percent)	5.62	5.64
Average local labour market vacancy rate	0.89	0.89
Average local labour market tightness	0.08	0.08
Share stayed at the original firm after 12 months	0.82	0.83
Share moved to a new firm after 12 months	0.09	0.09
Share not in salaried employment anymore after 12 months	0.09	0.08
Observations	145,595	12,972

A.2: Discontinuity in benefits and CIG intensity at the threshold

	Repl. rate (1)	Repl. rate No outliers (2)	Benefit (3)	Benefit No outliers (4)	Intensity (5)	Intensity No outliers (6)
Panel a. Conventional estimator						
$\beta_{threshold}$	0.028***	0.076***	60.580***	183.900***	0.013	0.020
s.e.	0.005	0.002	9.865	2.696	0.061	0.016
p-value	0.000	0.000	0.000	0.000	0.833	0.210
Panel b. Bias-corrected estimator						
$\beta_{threshold}$	0.026***	0.075***	54.743***	182.393***	0.018	0.022
s.e.	0.005	0.002	9.865	2.696	0.061	0.016
p-value	0.000	0.000	0.000	0.000	0.771	0.175
Panel c. Robust estimator						
$\beta_{threshold}$	0.026***	0.075***	54.743***	182.393***	0.018	0.022
s.e.	0.005	0.002	10.786	2.987	0.066	0.019
p-value	0.000	0.000	0.000	0.000	0.788	0.238
Actual obs. left	4886	2984	4142	3883	5206	8181
Actual obs. right	4095	3644	3952	3565	5134	4698
Left bandwidth	181.4	148.6	156.8	185.8	191.9	310.3
Right bandwidth	232.2	287.0	221.9	279.0	320.0	418.6
Avg. Outcome left	0.469	0.447	1065.4	1001.9	0.517	0.448
Average outcome	0.536	0.541	1045.4	1042.7	0.557	0.468
Observations	41107	35282	41107	35282	41107	35282

Note: Intensity is the share of working hours spent in STW in the month. Sample of workers whose benefit is paid directly by the Social Security Institute. Columns (1), (3) and (5) drop only observations with a replacement rate above 80% (weaker outlier detection rule), which is theoretically not possible; columns (2), (4) and (6) further remove all observations with a monthly benefit more than 10% different from the theoretical value. Standard errors are clustered at the firm level. The discontinuity, calculated at $t = 0$, is estimated using a local linear regression and the robust optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.3: Discontinuity in workers' characteristics around the threshold

	Manufacturing (1)	Foreign born (2)	Age (3)	Female (4)
Panel a. Conventional estimator				
$\beta_{threshold}$	-0.002	0.007	0.170	-0.014
s.e.	0.019	0.009	0.203	0.009
p-value	0.922	0.405	0.403	0.118
Panel b. Bias-corrected estimator				
$\beta_{threshold}$	0.002	0.004	0.090	-0.015*
s.e.	0.019	0.009	0.203	0.009
p-value	0.935	0.655	0.656	0.097
Panel c. Robust estimator				
$\beta_{threshold}$	0.002	0.004	0.090	-0.015
s.e.	0.020	0.010	0.226	0.010
p-value	0.940	0.681	0.690	0.126
Actual obs. left	22278	31453	30338	17443
Actual obs. right	22351	36543	34972	23132
Left bandwidth	154.0	214.6	209.4	123.9
Right bandwidth	201.2	409.8	385.3	211.9
Avg. Outcome left	0.620	0.799	42.29	0.139
Average outcome	0.642	0.784	41.73	0.169
Observations	145595	145595	144236	144236

Note: Main sample (workers whose CIG benefits are advanced by the employer). Standard errors are clustered at the firm level. The discontinuity, calculated at $t = 0$, is estimated using a local linear regression and the robust optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.4: Discontinuity in firms' characteristics around the threshold

	Log total employment (1)	Log average wage (2)	Year of entry (3)	Year of observation (4)
Panel a. Conventional estimator				
$\beta_{threshold}$	0.034	-0.001	-3.354	-0.036
s.e.	0.169	0.012	9.600	0.085
p-value	0.838	0.939	0.727	0.673
Panel b. Bias-corrected estimator				
$\beta_{threshold}$	0.086	0.001	-5.615	-0.035
s.e.	0.169	0.012	9.600	0.085
p-value	0.610	0.938	0.559	0.677
Panel c. Robust estimator				
$\beta_{threshold}$	0.086	0.001	-5.615	-0.035
s.e.	0.170	0.013	9.383	0.093
p-value	0.612	0.941	0.550	0.704
Actual obs. left	32091	27845	32653	50961
Actual obs. right	29394	29415	28127	40769
Left bandwidth	218.6	192.0	222.2	345.3
Right bandwidth	288.0	288.3	272.0	505.6
Avg. Outcome left	4.554	7.597	379.5	2011.0
Average outcome	4.436	7.558	382.2	2011.0
Observations	145595	145595	145595	145595

Note: Main sample (workers whose CIG benefits are advanced by the employer). Standard errors are clustered at the firm level. The discontinuity, calculated at $t = 0$, is estimated using a local linear regression and the robust optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.5: Discontinuity in the likelihood of being put in CIG

	In CIG within # months:				
	1	3	6	9	12
Panel a. Conventional estimator					
$\beta_{threshold}$	0.0003	0.0006	0.0009	0.0014	0.0018
s.e.	0.0004	0.0009	0.0013	0.0016	0.0018
p-value	0.528	0.540	0.496	0.395	0.335
Panel b. Bias-corrected estimator					
$\beta_{threshold}$	0.0003	0.0005	0.0008	0.0010	0.0014
s.e.	0.0004	0.0009	0.0013	0.0016	0.0018
p-value	0.563	0.579	0.564	0.515	0.436
Panel c. Robust estimator					
$\beta_{threshold}$	0.0003	0.0005	0.0008	0.0010	0.0014
s.e.	0.0005	0.0009	0.0014	0.0017	0.0019
p-value	0.567	0.588	0.579	0.530	0.452
Actual obs. left	7184285	6986345	6772745	6352084	5965841
Actual obs. right	8096506	8954305	9031903	8665251	8372395
Left bandwidth	197.4	199.8	205.0	204.4	204.5
Right bandwidth	356.9	429.1	469.9	480.2	496.3
Avg. outcome left	0.00412	0.00836	0.0119	0.0143	0.0160
Average outcome	0.00390	0.00841	0.0131	0.0172	0.0204
N	51462725	48970276	45635215	42633606	39898796

Note: Sample of all private sector workers. Each observation is a worker-firm pair at risk of CIG; observations after the first CIG episode at the worker-firm level are dropped. The dependent variable is a dummy for being in CIG # months after, hence the results can be interpreted as discontinuities in the likelihood of ending up in CIG. Standard errors are clustered at the firm level. The discontinuity is estimated using a local linear regression and the robust mean squared error optimal bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.6: Discontinuity in worker and firm fixed effects at the threshold

	Worker effect 1=Above median (1)	Firm effect 1=Above median (2)
Panel a. Conventional estimator		
$\beta_{threshold}$	-0.007	0.024
s.e.	0.011	0.020
p-value	0.539	0.225
Panel b. Bias-corrected estimator		
$\beta_{threshold}$	-0.009	0.031
s.e.	0.011	0.020
p-value	0.431	0.120
Panel c. Robust estimator		
$\beta_{threshold}$	-0.009	0.031
s.e.	0.013	0.021
p-value	0.489	0.137
Actual obs. left	37540	24387
Actual obs. right	30128	31203
Left bandwidth	267.0	179.1
Right bandwidth	322.2	347.0
Avg. Outcome left	0.479	0.549
Average outcome	0.500	0.500
Observations	138613	137529

Note: Only connected sample. The dependent variables are indicators for having a worker fixed effect or a firm fixed effect above the respective sample median. Standard errors are clustered at the firm level. The discontinuity, calculated at $t = 0$, is estimated using a local linear regression and the robust optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.7: Logit regressions for the likelihood of moving to a different firm at $t = 12$ (odds ratios)

	(1)	(2)	(3)
1[Above median worker eff.]	1.295*** (0.042)	1.299*** (0.040)	1.174*** (0.037)
log(wage)	0.727*** (0.047)	0.824 (0.296)	0.536* (0.175)
log(wage) ²		0.991 (0.031)	1.037 (0.030)
1[foreign born]			1.228*** (0.045)
1[female]			0.950 (0.068)
1[manufacturing]			0.471*** (0.035)
age			0.978*** (0.001)
Observations	138613	138613	138382
Year fixed eff.	X	X	X

Note: The reference group consists of workers with below-median AKM worker fixed effects. Only connected sample, as in Table 6. The observations are at $t = 0$. Standard errors in parentheses are clustered at the firm level. *, **, *** indicate significance at .10, .05 and .01 level.

A.8: Discontinuity in local labour market characteristics around the threshold

	Urban area	High unempl. rate	High tightness	High vacancy rate
	(1)	(2)	(3)	(4)
Panel a. Conventional estimator				
$\beta_{threshold}$	-0.003	0.007	-0.012	-0.007
s.e.	0.014	0.017	0.018	0.016
p-value	0.823	0.686	0.491	0.681
Panel b. Bias-corrected estimator				
$\beta_{threshold}$	-0.002	0.006	-0.012	-0.005
s.e.	0.014	0.017	0.018	0.016
p-value	0.872	0.746	0.493	0.744
Panel c. Robust estimator				
$\beta_{threshold}$	-0.002	0.006	-0.012	-0.005
s.e.	0.016	0.019	0.019	0.018
p-value	0.884	0.760	0.516	0.762
Actual obs. left	32501	40307	42479	38406
Actual obs. right	42480	41140	40235	38608
Left bandwidth	221.7	273.0	287.4	259.6
Right bandwidth	558.9	515.9	492.6	453.6
Avg. Outcome left	0.609	0.507	0.489	0.493
Average outcome	0.602	0.503	0.485	0.491
Observations	145311	145595	145595	145595

Note: Main sample (workers whose CIG benefits are advanced by the employer). Standard errors are clustered at the firm level. The discontinuity, calculated at $t = 0$, is estimated using a local linear regression and the robust optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.9: Discontinuity in earnings in a new firm at the threshold at $t = 12$, conditional on moving

	Earnings in the new firm	Log earnings in the new firm
Panel a. Conventional estimator		
$\beta_{threshold}$	1.853	0.023
s.e.	46.443	0.039
p-value	0.968	0.556
Panel b. Bias-corrected estimator		
$\beta_{threshold}$	6.114	0.026
s.e.	46.443	0.039
p-value	0.895	0.505
Panel c. Robust estimator		
$\beta_{threshold}$	6.114	0.026
s.e.	53.783	0.045
p-value	0.909	0.566
Actual obs. left	4141	4380
Actual obs. right	2941	3311
Left bandwidth	340.8	364.9
Right bandwidth	345.9	419.6
Avg. outcome left	1733.1	7.345
Average outcome	1841.6	7.376
Observations	13474	13474

Note: Main sample, conditional on having moved to a new firm at $t = 12$. Earnings are winsorized at the 1% level. Standard errors are clustered at the firm level. The discontinuity is estimated using a local linear regression and the robust optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.10: Discontinuity in the main outcomes at the threshold at different t

		Month t from the beginning of CIG event									
		3	6	9	12	15	18	21	24		
Panel a. Employed at the original firm at t											
$\beta_{threshold}$	0.011**		0.015**	0.013	0.022*	0.021	0.026*	0.027*		0.026*	
s.e.	0.006		0.007	0.009	0.013	0.013	0.014	0.014		0.015	
p-value	0.039		0.038	0.135	0.089	0.122	0.064	0.052		0.089	
Actual obs. left	34873	30999	30999	30829	35020	39749	44809	47247	44447		
Actual obs. right	24544	21695	25867	24930	26373	26373	26182	26060	26951		
Left bandwidth	236.5	212.2	210.9	237.4	268.6	268.6	303.3	319.9	301.1		
Right bandwidth	226.1	193.8	243.2	231.3	249.2	249.2	246.8	245.1	256.4		
Avg. Outcome left	0.948	0.911	0.870	0.821	0.870	0.781	0.750	0.722	0.691		
Average outcome	0.947	0.904	0.863	0.816	0.775	0.775	0.744	0.714	0.683		
Panel b. Employed at a different firm at t											
$\beta_{threshold}$	-0.006		-0.007	-0.007	-0.017	-0.019*	-0.017	-0.014		-0.016	
s.e.	0.004		0.005	0.007	0.011	0.011	0.011	0.011		0.013	
p-value	0.186		0.174	0.311	0.116	0.078	0.139	0.218		0.213	
Actual obs. left	44461	38522	38047	48149	47801	47801	49642	49587	47199		
Actual obs. right	26913	24551	27446	27996	30085	30085	33262	34135	37616		
Left bandwidth	301.3	260.3	256.8	326.4	323.9	323.9	336.3	335.8	319.6		
Right bandwidth	255.7	226.1	263.0	270.3	298.1	298.1	349.3	364.7	431.8		
Avg. Outcome left	0.0260	0.0419	0.0618	0.0905	0.111	0.111	0.130	0.141	0.158		
Average outcome	0.0268	0.0475	0.0681	0.0925	0.112	0.112	0.131	0.145	0.159		
Panel c. Not employed at t											
$\beta_{threshold}$	-0.004		-0.004	-0.003	-0.003	0.000	-0.009	-0.010		-0.007	
s.e.	0.003		0.005	0.005	0.006	0.007	0.007	0.007		0.008	
p-value	0.210		0.400	0.590	0.635	0.970	0.236	0.184		0.366	
Actual obs. left	34379	29267	32085	27524	35243	35243	33531	45778	42525		
Actual obs. right	33177	30736	38593	29803	30577	30577	25835	27293	27153		
Left bandwidth	233.6	201.2	218.6	189.7	238.6	238.6	228.2	309.8	287.9		
Right bandwidth	347.5	308.1	453.3	293.7	305.7	305.7	242.8	260.9	259.1		
Avg. Outcome left	0.0257	0.0463	0.0678	0.0861	0.107	0.107	0.119	0.137	0.150		
Average outcome	0.0260	0.0487	0.0694	0.0919	0.113	0.113	0.126	0.141	0.157		
N	145595	145595	145595	145595	145595	145595	145595	145595	145595		145595

Note: Main sample. Standard errors are clustered at the firm level. The discontinuity is estimated using the local linear regression and the robust mean squared error optimal bandwidth of Calonico et al. (2014, 2017), letting the bandwidth to differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.11: Discontinuity in the main outcomes at the threshold at $t = 12$, by quintile of worker fixed effects

	Worker fixed effect:				
	1st quintile (1)	2nd quintile (2)	3rd quintile (3)	4th quintile (4)	5th quintile (5)
Panel a. Outcome: Stay at the original firm					
$\beta_{threshold}$	0.037	0.047**	0.027	0.013	0.006
s.e.	0.023	0.020	0.018	0.020	0.021
p-value	0.104	0.022	0.144	0.533	0.775
Panel b. Outcome: At a different firm					
$\beta_{threshold}$	-0.019	-0.028*	-0.016	-0.009	-0.004
s.e.	0.016	0.015	0.015	0.019	0.018
p-value	0.233	0.058	0.290	0.645	0.834
Panel c. Outcome: Not in salaried employment					
$\beta_{threshold}$	0.001	-0.015	-0.012	-0.004	0.001
s.e.	0.017	0.013	0.011	0.010	0.013
p-value	0.958	0.258	0.254	0.699	0.952
Observations	27722	27724	27724	27722	27722

Note: Only connected sample. The sample is split by quintile of the worker fixed effect. Standard errors are clustered at the firm level. The discontinuity is estimated using a local linear regression and the robust estimator and the optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.12: Discontinuity in the main outcomes at the threshold at $t = 12$, by initial firm fixed effects

	Firm fixed effect:	
	Below median (1)	Above median (2)
Panel a. Outcome: Stay at the original firm		
$\beta_{threshold}$	0.011	0.026*
s.e.	0.021	0.014
p-value	0.605	0.069
Panel b. Outcome: At a different firm		
$\beta_{threshold}$	-0.016	-0.017
s.e.	0.019	0.012
p-value	0.411	0.147
Panel c. Outcome: Not in salaried employment		
$\beta_{threshold}$	0.007	-0.007
s.e.	0.010	0.007
p-value	0.456	0.281
Observations	68774	68756

Note: Only connected sample. The sample is split according to whether the initial firm's fixed effect is below or above the median. Standard errors are clustered at the firm level. The discontinuity is estimated using a local linear regression and the robust estimator and the optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

A.13: Discontinuity in the probability of being with another firm at the threshold at $t = 12$

	At a different firm with fixed effects:	
	Below median (1)	Above median (2)
Panel a. Conventional estimator		
$\beta_{threshold}$	-0.002	-0.011
s.e.	0.002	0.009
p-value	0.258	0.230
Panel b. Bias-corrected estimator		
$\beta_{threshold}$	-0.003	-0.014
s.e.	0.002	0.009
p-value	0.175	0.143
Panel c. Robust estimator		
$\beta_{threshold}$	-0.003	-0.014
s.e.	0.002	0.010
p-value	0.239	0.196
Actual obs. left	41111	47871
Actual obs. right	32233	30581
Left bandwidth	279.3	326.1
Right bandwidth	334.7	308.5
Avg. outcome left	0.0170	0.0689
Average outcome	0.0195	0.0683
Observations	144832	144832

Note: Only connected sample. The dependent variables indicate whether the worker is employed at $t = 12$ by a different firm whose fixed effect is below or above the median, respectively. Standard errors are clustered at the firm level. The discontinuity, calculated at $t = 0$, is estimated using a local linear regression and the robust optimal mean squared error bandwidth of Calonico et al. (2014, 2017), letting the bandwidth differ on either side of the threshold. *, **, *** indicate significance at .10, .05 and .01 level.

B. Theoretical derivations and proofs

B.1 Theoretical Framework

The value of an employed worker in a full-time job reads:

$$W(x, y, \epsilon) = u(w(x, y, \epsilon)) + \beta(1 - \delta) \left[\lambda (G(x) - G(y)) \int_{y^F}^x \mathbb{E}_{\tilde{\epsilon}} [W(x, \tilde{y}, \tilde{\epsilon})] dG(\tilde{y} | \tilde{y} \in (y^F, x)) + (1 - \lambda (G(x) - G(y))) W(x, y, \epsilon) \right] + \beta \delta U,$$

where $y^F = y^F(x, y, \epsilon)$ denotes the reservation productivity of a full-time worker with ability x at a firm with productivity y and match-specific shock ϵ .

The reservation productivity for a worker on STW solves:

$$\begin{aligned} y^*(x, y, \epsilon; \tilde{\alpha}) : \mathbb{E}_{\tilde{\epsilon}}[W(x, y^*, \tilde{\epsilon})] &= \mathbb{E}_{\theta}[W^{STW}(x, y, \epsilon; \tilde{\alpha})], \\ \iff \mathbb{E}_{\tilde{\epsilon}}[u(w(x, y^*, \tilde{\epsilon}))] &= \mathbb{E}_{\theta}[u^{STW}(w(x, y, \epsilon); \tilde{\alpha})] - \beta(1 - \delta)(1 - \theta)\lambda \left[(G(x) - G(y^*)) \right. \\ &\quad (\mathbb{E}_{\tilde{y}, \tilde{\epsilon}}[W(x, \tilde{y}, \tilde{\epsilon}) | \tilde{y} \in (y^*, x)] - \mathbb{E}_{\tilde{\epsilon}}[W(x, y^*, \tilde{\epsilon})]) - (G(x) - G(y^F)) \\ &\quad \left. (\mathbb{E}_{\tilde{y}, \tilde{\epsilon}}[W(x, \tilde{y}, \tilde{\epsilon}) | \tilde{y} \in (y^F, x)] - W(x, y, \epsilon)) \right] - \beta(1 - \delta)(1 - \theta) (\mathbb{E}_{\tilde{\epsilon}}[W(x, y^*, \tilde{\epsilon})] - W(x, y, \epsilon)), \\ \iff \mathbb{E}_{\tilde{\epsilon}}[u(w(x, y^*(x, y, \epsilon; \tilde{\alpha}), \tilde{\epsilon}))] &\approx \mathbb{E}_{\theta}[u^{STW}(w(x, y, \epsilon); \tilde{\alpha})], \end{aligned}$$

where $\mathbb{E}_{\theta}[u^{STW}(w(x, y, \epsilon); \tilde{\alpha})] \equiv \theta u((h + \tilde{\alpha})w(x, y, \epsilon)) + (1 - \theta)u(w(x, y, \epsilon))$, and the last line approximates the intertemporal mobility decision with a static rule assuming that the differential continuation value is small enough. The propositions in the main text are comparative statics of this static reservation rule.

Proof of Proposition 1. The derivative of the reservation productivity with respect to the replacement rate is:

$$\frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} = \frac{\theta u'((h + \tilde{\alpha})w(x, y, \epsilon)) w(x, y, \epsilon)}{w_y \mathbb{E}_{\tilde{\epsilon}}[u'(w(x, y^*, \tilde{\epsilon}))]} > 0,$$

where w_y is the derivative of the wage function with respect to firm productivity evaluated at the reservation productivity y^* , $w_y(x, y^*, \tilde{\epsilon})$. Due to linearity and additive separability of the wage function, it equals the constant parameter w_y . Hence, a higher replacement rate reduces employees' probability of moving:

$$\frac{\partial m(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} = -\lambda G'(y^*) \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} < 0.$$

□

Proof of Proposition 2. The second derivative of the moving probability with respect to the replacement rate and worker ability holding the wage constant reads:

$$\left. \frac{\partial^2 m(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w = -\lambda G''(y^*) \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial x} \Big|_w \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} - \lambda G'(y^*) \frac{\partial^2 y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \Big|_w,$$

where:

$$\begin{aligned} \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial x} &= -\frac{w_x}{w_y} \left(1 - \frac{\theta(h + \tilde{\alpha})u'((h + \tilde{\alpha})w(x, y, \epsilon)) + (1 - \theta)u'(w(x, y, \epsilon))}{\mathbb{E}_{\tilde{\epsilon}}[u'(w(x, y^*, \tilde{\epsilon}))]} \right) < 0, \\ \left. \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial x} \right|_w &= -\frac{w_x}{w_y} < 0, \\ \left. \frac{\partial^2 y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w &= -\frac{\theta u'((h + \tilde{\alpha})w(x, y, \epsilon))w(x, y, \epsilon)}{w_y} \frac{\mathbb{E}_{\tilde{\epsilon}} \left[u''(w(x, y^*, \tilde{\epsilon})) w_y \left(\frac{w_x}{w_y} + \frac{\partial y^*}{\partial x} \Big|_w \right) \right]}{\mathbb{E}_{\tilde{\epsilon}}[u'(w(x, y^*, \tilde{\epsilon}))]^2} \\ &= 0. \end{aligned}$$

Notice that, for given firm productivity, taking the derivative of one variable with respect to worker ability holding the wage constant is equivalent to constraining the match-specific shock to be a linear function of ability, that is, $w_x x + w_y \epsilon = \bar{w}$.

Overall, it follows that:

$$\begin{aligned} \left. \frac{\partial^2 m(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w &= \frac{w_x}{w_y} \lambda G''(y^*(x, y, \epsilon; \tilde{\alpha})) \frac{\theta u'((h + \tilde{\alpha})w(x, y, \epsilon)) w(x, y, \epsilon)}{\mathbb{E}_{\tilde{\epsilon}}[u'(w(x, y^*, \tilde{\epsilon})) w_y]} \\ &> 0 \iff G''(y^*(x, y, \epsilon; \tilde{\alpha})) > 0. \end{aligned}$$

□

B.1.2 Alternative models of search behaviour.

We now survey how the predicted cross-sectional impact of a change in the replacement rate on workers' moving probability varies depending on the specific modeling of workers' search behaviour. For analytical transparency, we focus on two-period models. To isolate the role of search behaviour, we focus on risk-neutral workers.

Random search with ex ante heterogeneous jobs (Baseline). We first study our baseline model with risk-neutral workers, i.e, $u(w) = w$. Under risk neutrality, the

reservation productivity solves:

$$\mathbb{E}_{\tilde{\epsilon}}[w(x, y^*(x, y, \epsilon; \tilde{\alpha}), \tilde{\epsilon})] = [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, y, \epsilon).$$

It follows that:

$$\begin{aligned} \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} &= \frac{\theta w(x, y, \epsilon)}{w_y} > 0, \\ \implies \frac{\partial m(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} &= -\lambda G'(y^*) \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} < 0. \end{aligned}$$

Moreover,

$$\begin{aligned} \left. \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial x} \right|_w &= -\frac{w_x}{w_y} < 0, \\ \left. \frac{\partial^2 y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w &= 0. \\ \implies \left. \frac{\partial^2 m(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w &= -\lambda G''(y^*) \left. \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial x} \right|_w \frac{\partial y^*(x, y, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} \\ &> 0 \iff G''(y^*(x, y, \epsilon)) > 0. \end{aligned}$$

Hence, even under risk neutrality, increasing local density, $G''(y^*(x, y, \epsilon)) > 0$, is a sufficient and necessary condition for the effect of the replacement rate on the moving probability to be locally increasing in ability.

Random search with ex ante homogeneous jobs and endogenous search effort. In this alternative model, firms (jobs) are ex ante equally productive and workers of any ability are qualified for all jobs (formally, $y = 0$ in our baseline model). However, workers can choose how much search effort s to exert, subject to an increasing and convex search effort cost function ψ .³⁴

Optimal search effort solves:

$$\begin{aligned} s(x, \epsilon; \tilde{\alpha}) &= \arg \max_s s \lambda (\mathbb{E}_{\tilde{\epsilon}}[w(x, \tilde{\epsilon})] - [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, \epsilon)) - \psi(s) \\ &= (\psi')^{-1} (\lambda (\mathbb{E}_{\tilde{\epsilon}}[w(x, \tilde{\epsilon})] - [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, \epsilon))) \\ &= (\psi')^{-1} (\lambda (w_x x - [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, \epsilon))), \end{aligned}$$

where we use the linearity and additive separability of the wage function to simplify

³⁴Absent endogenous search effort, workers would simply accept all job offers (provided that the current match-specific shock is sufficiently low). As a result, the moving probability would be independent of the replacement rate.

the expected wage function. Let $\varphi \equiv (\psi')^{-1}$.

In this alternative model, the moving probability is defined as:

$$m(x, \epsilon; \tilde{\alpha}) \equiv s(x, \epsilon; \tilde{\alpha})\lambda.$$

Hence, the response of the moving probability to the replacement rate tracks the response of search effort.

$$\begin{aligned} \frac{\partial s(x, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} &= -\varphi'(\lambda(w_x x - [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, \epsilon))) \lambda \theta w(x, \epsilon) < 0, \\ \implies \frac{\partial m(x, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} &= \lambda \frac{\partial s(x, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} < 0. \end{aligned}$$

Hence, a higher replacement rate reduces job-to-job transitions, as in our baseline.

Moreover,

$$\begin{aligned} \left. \frac{\partial^2 s(x, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w &= -w_x \varphi''(\lambda(w_x x - [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, \epsilon))) \lambda^2 \theta w(x, \epsilon) \\ \implies \left. \frac{\partial^2 m(x, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w &= \lambda \left. \frac{\partial^2 s(x, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w \\ &> 0 \iff \varphi''(.) < 0 \iff (\psi^{-1})''(.) < 0. \end{aligned}$$

Hence, the response of the moving probability to the replacement rate locally increases with ability if and only if the search effort cost function is sufficiently convex. With the standard iso-elastic functional form, $\psi(s) = \frac{s^{1+\eta}}{1+\eta}$, this condition would map into a more-than-quadratic cost function, i.e., $\eta > 1$ – a larger degree of convexity than normally used.

Directed search with ex ante homogeneous jobs. Rather than passively receiving random job offers, workers may actively direct their search towards specific segments of the labour market (*sub-markets*) offering some terms of trade. Therefore, in this alternative model we allow workers to direct their search.

The targeted sub-market is chosen by trading-off higher wages with lower job find-

ing probability:

$$\begin{aligned}
\bar{w}(x, \epsilon; \tilde{\alpha}) &= \arg \max_{\bar{w}} \lambda(\vartheta(\bar{w})) (\mathbb{E}_{\tilde{\epsilon}}[\bar{w} + w_{\epsilon}\tilde{\epsilon}] - [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, \epsilon)), \\
\lambda'(\vartheta)\vartheta'(\bar{w}(x, \epsilon; \tilde{\alpha})) (\bar{w}(x, \epsilon; \tilde{\alpha}) - [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, \epsilon)) + \lambda(\vartheta) &= 0, \\
\implies \bar{w}(x, \epsilon; \tilde{\alpha}) &= [\theta(h + \tilde{\alpha}) + (1 - \theta)]w(x, \epsilon) - \frac{\lambda(\vartheta)}{\lambda'(\vartheta)\vartheta'(\bar{w}(x, \epsilon; \tilde{\alpha}))},
\end{aligned}$$

where the second line represents the first-order condition solving the maximization problem in the first row.

The worker search problem can be solved by a schedule of wages and job finding probabilities. To pin down the equilibrium combination, we impose a free-entry condition on the firm side and iso-elastic, CRS meeting functions governing matching in each sub-market, $\mathcal{M}(S, V) = V^{\chi}S^{1-\chi}$, $\chi \in (0, 1)$.

The free-entry condition reads:

$$\begin{aligned}
\kappa &= \beta \vartheta(\bar{w}(x, \epsilon; \tilde{\alpha}))^{\chi-1} [p(x) - \bar{w}(x, \epsilon; \tilde{\alpha})] \quad \forall \bar{w}, \\
\vartheta(\bar{w}(x, \epsilon; \tilde{\alpha})) &= \left(\frac{\beta}{\kappa} [p(x) - \bar{w}(x, \epsilon; \tilde{\alpha})] \right)^{\frac{1}{1-\chi}}.
\end{aligned}$$

It follows that:

$$\begin{aligned}
\vartheta'(\bar{w}(x, \epsilon; \tilde{\alpha})) &= -\frac{1}{1-\chi} \frac{\beta}{\kappa} \lambda(\vartheta(\bar{w}(x, \epsilon; \tilde{\alpha}))), \\
\lambda'(\vartheta) &= \chi \vartheta^{\chi-1}, \\
\frac{\lambda(\vartheta)}{\lambda'(\vartheta)\vartheta'(\bar{w}(x, \epsilon; \tilde{\alpha}))} &= -\frac{1-\chi}{\chi} [p(x) - \bar{w}(x, \epsilon; \tilde{\alpha})].
\end{aligned}$$

Plugging these equations into the wage function yields:

$$\begin{aligned}
\bar{w}(x, \epsilon; \tilde{\alpha}) &= [\theta(h + \tilde{\alpha}) + (1 - \theta)] w(x, \epsilon) + \frac{1-\chi}{\chi} [p(x) - \bar{w}(x, \epsilon; \tilde{\alpha})] \\
&= \chi [\theta(h + \tilde{\alpha}) + (1 - \theta)] w(x, \epsilon) + (1 - \chi)p(x).
\end{aligned}$$

In this alternative model, the moving probability is defined as:

$$m(x, \epsilon; \tilde{\alpha}) \equiv \lambda(\vartheta(\bar{w}(x, \epsilon; \tilde{\alpha}))) = \left[\frac{\beta\chi}{\kappa} (p(x) - [\theta(h + \tilde{\alpha}) + (1 - \theta)] w(x, \epsilon)) \right]^{\frac{\chi}{1-\chi}}.$$

For an interior solution with positive match surplus, the response of the moving prob-

ability to the replacement rate is:

$$\frac{\partial m(x, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha}} = -\frac{\chi}{1-\chi} \left(\frac{\beta\chi}{\kappa} \right)^{\frac{\chi}{1-\chi}} \theta w(x, \epsilon) (p(x) - [\theta(h + \tilde{\alpha}) + (1-\theta)] w(x, \epsilon))^{\frac{\chi}{1-\chi}-1} < 0.$$

Hence, a higher replacement rate reduces job-to-job transitions, as in our baseline. Moreover,

$$\begin{aligned} \left. \frac{\partial^2 m(x, \epsilon; \tilde{\alpha})}{\partial \tilde{\alpha} \partial x} \right|_w &= -\frac{\chi}{1-\chi} \left(\frac{\chi}{1-\chi} - 1 \right) \left(\frac{\beta\chi}{\kappa} \right)^{\frac{\chi}{1-\chi}} \theta w(x, \epsilon) \\ &\quad (p(x) - [\theta(h + \tilde{\alpha}) + (1-\theta)] w(x, \epsilon))^{\frac{\chi}{1-\chi}-2} p'(x), \\ &> 0 \iff \chi < \frac{1}{2}. \end{aligned}$$

Hence, the response of the moving probability to the replacement rate locally increases with ability if and only if the elasticity of the meeting function with respect to vacancies is lower than 1/2. With the standard equally weighted functional form, $\chi = 1/2$, the directed search model would predict a constant response of the moving probability to the replacement rate across the ability distribution.

B.2 Optimal replacement rate

The Lagrangian associated with the government problem is:

$$\begin{aligned} \mathcal{L}(\tilde{\alpha}, \tau) &= e(1-\delta) \left\{ \mu \int \gamma(w) \int [m(x, y^*) \mathbb{E}[u((1-\tau)w(x, \tilde{y}, \tilde{\epsilon})) \mid \tilde{y} \in (y^*, x)] \right. \\ &\quad + (1-m(x, y^*)) [\theta u((h + \tilde{\alpha}(w))w) + (1-\theta)u((1-\tau)w)] dF(x, y, \epsilon; w) dw \\ &\quad \left. + (1-\mu) \int u((1-\tau)w(x, y, \epsilon)) d\hat{F}(x, y, \epsilon) \right\} \\ &\quad + \int [e\delta + (1-e)(1-f(x, R^*))] u(b) dH_U(x) \\ &\quad + (1-e) \int f(x, R^*) \mathbb{E}[u((1-\tau)w(x, \tilde{y}, \tilde{\epsilon})) \mid \tilde{y} \in (R^*(x), x)] dH_U(x) \\ &\quad + \xi [\mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)\tau - \mathcal{B}(\mathbf{y}^*, \mathbf{R}^*, \tilde{\alpha}, b)], \end{aligned}$$

where ξ is the multiplier attached to the government budget constraint.

In an interior solution, the optimal wage-specific gross replacement rate is pinned

down implicitly by:

$$\begin{aligned}
& e(1 - \delta)\mu\gamma(w) \int (1 - m(x, y^*)) \theta u'((h + \tilde{\alpha}(w))w) w dF(x, y, \epsilon; w) \\
& - \xi \left[\frac{d\mathcal{B}(\mathbf{y}^*, \mathbf{R}^*, \tilde{\alpha}, b)}{d\tilde{\alpha}(w)} \Big|_{\tau} - \tau \frac{d\mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)}{d\tilde{\alpha}(w)} \Big|_{\tau} \right] \\
& = e(1 - \delta)\mu\gamma(w) \int \left\{ \lambda G'(y^*) (\mathbb{E}[u((1 - \tau)w(x, \tilde{y}, \tilde{\epsilon})) \mid \tilde{y} \in (y^*, x)] - \mathbb{E}_{\theta}[u^{STW}(w(x, y, \epsilon); \tilde{\alpha}(w))]) \right. \\
& \quad \left. - m(x, y^*) \frac{\partial \mathbb{E}[u((1 - \tau)w(x, \tilde{y}, \tilde{\epsilon})) \mid \tilde{y} \in (y^*, x)]}{\partial y^*} \right\} \frac{\partial y^*}{\partial \tilde{\alpha}(w)} dF(x, y, \epsilon; w). \tag{3}
\end{aligned}$$

Conditional on the tax rate, $R^*(x; b, \tau)$ does not respond directly to $\tilde{\alpha}(w)$. Hence, the unemployment component of welfare does not enter the right-hand side of equation (3). The factor $e(1 - \delta)$ multiplies both the marginal utility gain of STW beneficiaries and the effects arising through employed workers' mobility decisions.

The left-hand side of equation (3) represents the net social value of increasing the replacement rate. It equals the marginal utility gain of STW beneficiaries from higher transfers minus the social value of the government resources required to balance the budget. The right-hand side collects the direct welfare effect (with opposite sign) arising from the adjustment of the reservation productivity y^* induced by the change in the replacement rate, for a given tax rate. This adjustment has two effects. First, a higher replacement rate raises the reservation productivity and reduces job-to-job transitions, $\partial m(x, y^*)/\partial \tilde{\alpha}(w) < 0$. Second, by raising the reservation productivity, it increases expected utility conditional on moving. By Lemma 1, these two effects exactly offset each other at the margin, so that the adjustment of y^* has zero first-order direct effect on social welfare.

The first-order condition therefore simplifies to:

$$\frac{e(1 - \delta)\partial \mathcal{V}^B(w)/\partial \tilde{\alpha}(w)}{\xi} = \frac{d\mathcal{B}(\mathbf{y}^*, \mathbf{R}^*, \tilde{\alpha}, b)}{d\tilde{\alpha}(w)} \Big|_{\tau} - \frac{d\mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)}{d\tilde{\alpha}(w)} \Big|_{\tau}, \tag{4}$$

where $\frac{\partial \mathcal{V}^B(w)}{\partial \tilde{\alpha}(w)} \equiv \mu \int (1 - m(x, y^*)) \theta u'((h + \tilde{\alpha}(w))w) w \gamma(w) dF(x, y, \epsilon; w)$ is the marginal expected indirect utility of STW beneficiaries with respect to the replacement rate, conditional on being employed and on the match surviving.

Conditional on τ and R^* , the unemployment-related components of the tax base and government outlays do not depend directly on $\tilde{\alpha}(w)$. Therefore, the right-hand

side of equation (4) equals:

$$\begin{aligned}
& \left. \frac{d\mathcal{B}(\mathbf{y}^*, \mathbf{R}^*, \tilde{\alpha}, b)}{d\tilde{\alpha}(w)} \right|_{\tau} - \left. \frac{d\mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)}{d\tilde{\alpha}(w)} \right|_{\tau} \tau \\
&= e(1-\delta)\mu\gamma(w) \int \theta [1 - m(x, y^*)] w dF(x, y, \epsilon; w) \\
&\quad + e(1-\delta)\mu\gamma(w) \int \theta \frac{\partial[1 - m(x, y^*)]}{\partial\tilde{\alpha}(w)} \tilde{\alpha}(w) w dF(x, y, \epsilon; w) \\
&\quad + e(1-\delta)\mu\gamma(w) \int \frac{\partial[1 - m(x, y^*)]}{\partial\tilde{\alpha}(w)} [\mathbb{E}[w(x, y^*, \tilde{\epsilon})] - (1-\theta)w] dF(x, y, \epsilon; w) \tau \\
&= e(1-\delta)B(w) \left[1 + \int q(x, y^*) \epsilon_{1-m, \tilde{\alpha}(w)} \left(1 + \frac{\mathbb{E}[w(x, y^*, \tilde{\epsilon})] - (1-\theta)w}{\theta\tilde{\alpha}(w)w} \tau \right) dF(x, y, \epsilon; w) \right],
\end{aligned}$$

where $q(x, y^*) \equiv \frac{1-m(x, y^*)}{\int [1-m(x, y^*)] dF(x, y, \epsilon; w)}$ is the probability of staying at the current firm relative to the average among workers on STW at the same full-time wage, and $B(w) \equiv \mu\gamma(w) \int \theta [1 - m(x, y^*)] w dF(x, y, \epsilon; w)$. The common factor $e(1-\delta)$ cancels from equation (4), yielding:

$$\frac{\partial \mathcal{V}^B(w) / \partial \tilde{\alpha}(w)}{\xi} = B(w) [1 + MH(\tilde{\alpha}(w))], \quad (5)$$

where $MH(\tilde{\alpha}(w)) \equiv \int q(x, y^*) \epsilon_{1-m, \tilde{\alpha}(w)} \left(1 + \frac{\mathbb{E}[w(x, y^*, \tilde{\epsilon})] - (1-\theta)w}{\theta\tilde{\alpha}(w)w} \tau \right) dF(x, y, \epsilon; w)$ and $\epsilon_{1-m, \tilde{\alpha}(w)} \equiv \frac{\partial[1-m(x, y^*)] / \partial \tilde{\alpha}(w)}{1-m(x, y^*)} \tilde{\alpha}(w)$.

The first-order condition with respect to the payroll tax equalizes the marginal cost of public funds with its marginal utility cost for full-time taxpayers:

$$\xi \mathcal{T}_{\tau}^{\text{eff}}(\mathbf{y}^*, \mathbf{R}^*) = \frac{\partial \mathcal{V}^T}{\partial \tau}, \quad (6)$$

where the effective marginal tax base is defined as:

$$\begin{aligned}
\mathcal{T}_{\tau}^{\text{eff}}(\mathbf{y}^*, \mathbf{R}^*) &= \mathcal{T}(\mathbf{y}^*, \mathbf{R}^*) + \left[\tau \frac{\partial \mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)}{\partial \mathbf{R}^*} - \frac{\partial \mathcal{B}(\mathbf{y}^*, \mathbf{R}^*, \tilde{\alpha}, b)}{\partial \mathbf{R}^*} \right] \frac{\partial \mathbf{R}^*}{\partial \tau} \\
&\quad + \left[\tau \frac{\partial \mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)}{\partial \mathbf{y}^*} - \frac{\partial \mathcal{B}(\mathbf{y}^*, \mathbf{R}^*, \tilde{\alpha}, b)}{\partial \mathbf{y}^*} \right] \frac{\partial \mathbf{y}^*}{\partial \tau} \\
&\simeq \mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)
\end{aligned}$$

On top of the mechanical effect, the government internalizes the fiscal effect of the response of $\mathbf{R}^*$ and $\mathbf{y}^*$ to the payroll tax: a higher payroll tax raises those thresholds, reduces unemployment- and employment-to-employment transitions, contracts the tax base, and increases UI and STW expenditure. Following the literature, we assume that

those fiscal effects are small and approximate the effective marginal tax base with the tax base itself.

The marginal utility cost for full-time taxpayers equals:

$$\begin{aligned} \frac{\partial \mathcal{V}^T}{\partial \tau} \equiv & e(1 - \delta) \left\{ \mu \int \int [(1 - m(x, y^*)) (1 - \theta) u'((1 - \tau)\tilde{w}) \tilde{w} \right. \\ & + m(x, y^*) \mathbb{E}[u'((1 - \tau)w(x, \tilde{y}, \tilde{\epsilon})) w(x, \tilde{y}, \tilde{\epsilon}) \mid \tilde{y} \in (y^*, x)]] \gamma(\tilde{w}) d\tilde{w} dF(x, y, \epsilon; \tilde{w}) \\ & + (1 - \mu) \int u'((1 - \tau)w(x, y, \epsilon)) w(x, y, \epsilon) d\hat{F}(x, y, \epsilon) \Big\} \\ & + (1 - e) \int f(x, R^*) \mathbb{E}[u'((1 - \tau)w(x, \tilde{y}, \tilde{\epsilon})) w(x, \tilde{y}, \tilde{\epsilon}) \mid \tilde{y} \in (R^*(x), x)] dH_U(x). \end{aligned}$$

The first component is the marginal utility loss borne by full-time taxpayers whose matches survive. The second is the marginal utility loss borne by workers who enter full-time employment from unemployment. Workers remaining unemployed do not enter this expression because b is held fixed.

Substituting equation (6) into equation (5) yields:

$$RV(\tilde{\alpha}(w)) = 1 + MH(\tilde{\alpha}(w)),$$

where $RV(\tilde{\alpha}(w)) \equiv \frac{\partial \mathcal{V}^B(w)/\partial \tilde{\alpha}(w)}{\partial \mathcal{V}^T/\partial \tau} \frac{\mathcal{T}(\mathbf{y}^*, \mathbf{R}^*)}{B(w)}$ is the marginal value from redistribution.

Endogenous firm responses and interacted fiscal externality. A natural extension to our framework allows firms' hours and STW take-up decisions to respond to the replacement-rate schedule. In this case, the policy instrument is the replacement rate per hour not worked, $\alpha(w)$, while the gross subsidy per worker, $\tilde{\alpha}(w) = \alpha(w) [1 - h^*(w; \boldsymbol{\alpha})]$, becomes endogenous. The planner does not choose hours or STW take-up directly, but internalizes the policy functions $h^* = h^*(w; \boldsymbol{\alpha})$, $\mu^* = \mu^*(\boldsymbol{\alpha})$ induced by firms' optimal choices. We assume no cross-wage responses and no direct effect of aggregate STW take-up on an individual worker's mobility decision. Because hours affect workers' income while on STW, the total worker response is

$$\left. \frac{d[1 - m(x, y^*)]}{d\alpha(w)} \right|_{\tau} = \left. \frac{\partial[1 - m(x, y^*)]}{\partial \alpha(w)} \right|_{h, \tau} + \frac{\partial[1 - m(x, y^*)]}{\partial h} \frac{\partial h^*(w; \boldsymbol{\alpha})}{\partial \alpha(w)}. \quad (7)$$

Our estimated worker-level response is interpreted as the total derivative on the left-hand side of equation (7). It therefore already incorporates the interaction between firms' hours choices and workers' mobility choices.

Let $\mathcal{C} \equiv \mathcal{B} - \tau \mathcal{T}$ denote the government's net financing requirement. Holding μ , h ,

and y^* fixed, the mechanical cost of increasing $\alpha(w)$ is

$$\mathcal{M}(w) \equiv e(1 - \delta)\mu\gamma(w)\theta [1 - h^*(w)] w \int [1 - m(x, y^*)] dF(x, y, \epsilon; w). \quad (8)$$

At a fixed payroll tax, the complete budget response is

$$\begin{aligned} \left. \frac{d\mathcal{C}}{d\alpha(w)} \right|_{\tau} &= \mathcal{M}(w) \\ &+ e(1 - \delta)\mu\gamma(w) \int [\theta\alpha(w) (1 - h^*(w)) w + \tau\Delta w(x, w)] \left. \frac{d[1 - m(x, y^*)]}{d\alpha(w)} \right|_{\tau} dF(x, y, \epsilon; w) \\ &- e(1 - \delta)\mu\gamma(w)\theta\alpha(w)w \int [1 - m(x, y^*)] dF(x, y, \epsilon; w) \frac{\partial h^*(w; \boldsymbol{\alpha})}{\partial \alpha(w)} \\ &+ \left[\frac{\partial \mathcal{B}}{\partial \mu} - \tau \frac{\partial \mathcal{T}}{\partial \mu} \right] \frac{\partial \mu^*(\boldsymbol{\alpha})}{\partial \alpha(w)}. \end{aligned} \quad (9)$$

Define the total worker elasticity as $\epsilon_{[1-m(x, y^*)], \alpha(w)}^{total} \equiv \left. \frac{d[1-m(x, y^*)]/d\alpha(w)}{[1-m(x, y^*)]} \right|_{\tau} \alpha(w)$. Normalizing equation (9) by the mechanical cost gives

$$\begin{aligned} FE^{total}(\alpha(w)) &\equiv \frac{d\mathcal{C}/d\alpha(w)|_{\tau}}{\mathcal{M}(w)} \\ &= 1 + MH^W(\alpha(w)) + MH^h(\alpha(w)) + MH^{\mu}(\alpha(w)), \end{aligned} \quad (10)$$

where

$$\begin{aligned} MH^W(\alpha(w)) &\equiv \int q(x, y^*) \epsilon_{[1-m(x, y^*)], \alpha(w)}^{total} \left[1 + \frac{\tau\Delta w(x, w)}{\theta\alpha(w) (1 - h^*(w)) w} \right] dF(x, y, \epsilon; w), \\ MH^h(\alpha(w)) &\equiv - \epsilon_{h^*(w), \alpha(w)} \frac{h^*(w)}{1 - h^*(w)}, \\ MH^{\mu}(\alpha(w)) &\equiv \frac{\left[\frac{\partial \mathcal{B}}{\partial \mu} - \tau \frac{\partial \mathcal{T}}{\partial \mu} \right] \frac{\partial \mu^*(\boldsymbol{\alpha})}{\partial \alpha(w)}}{\mathcal{M}(w)}. \end{aligned}$$

There is no additional worker-hours interaction term in equation (10) on top of that already contained in the total worker elasticity. This decomposition nevertheless requires the maintained assumption that a change in aggregate take-up does not itself alter individual mobility probabilities.

To map the take-up term into Giupponi and Landais (2023), let $n(w)$ denote the wage-specific mass of workers retained in STW and suppose that each additional worker in $n(w)$ would otherwise receive unemployment benefits b . Ignoring payroll-tax effects

and normalizing full-time hours to one gives

$$MH^F(\alpha(w)) = \epsilon_{n(w), \alpha(w)} \left[1 - \frac{b}{\alpha(w) [1 - h^*(w)] w} \right] - \epsilon_{h^*(w), \alpha(w)} \frac{h^*(w)}{1 - h^*(w)}.$$

It is $n(w)$, rather than the conditional share μ alone, that corresponds to the employment object in Giupponi and Landais (2023). The mapping from μ to $n(w)$ additionally requires the employment rate and the wage composition of workers entering or leaving STW to remain fixed.

Finally, the worker envelope condition eliminates the direct welfare effect of the adjustment of y^* , but it does not generally eliminate the welfare effects of firms' choices. The planner's first-order condition is

$$\left. \frac{\partial \mathcal{V}}{\partial \alpha(w)} \right|_{h, \mu, \tau} + \frac{\partial \mathcal{V}}{\partial h} \frac{\partial h^*(w; \boldsymbol{\alpha})}{\partial \alpha(w)} + \frac{\partial \mathcal{V}}{\partial \mu} \frac{\partial \mu^*(\boldsymbol{\alpha})}{\partial \alpha(w)} = \xi \mathcal{M}(w) FE^{total}(\alpha(w)).$$

Thus, the left-hand side is not a pure redistribution term unless further assumptions make firms' hours and take-up choices socially optimal at the margin.

For a back-of-the-envelope comparison, suppose that the average firm-side net fiscal externality estimated by Giupponi and Landais (2023) applies locally at the wage threshold, that both components are normalized by the same mechanical expenditure, and that the difference in policy experiments can be ignored. Because our estimated worker response is a total response, no further worker-hours correction is required. Under these assumptions,

$$FE^{total} \simeq 1 + \underbrace{0.15}_{\text{workers}} + \underbrace{0.38}_{\text{firms}} = 1.53.$$

Worker search behaviour therefore accounts for approximately 28% of the combined behavioural net fiscal externality. Notice, however, that this calculation is illustrative rather than an exact aggregation because the two estimates correspond to different policy experiments.

Proof of Lemma 1. The RHS of the government FOC (3) collects, with the opposite sign, the first-order welfare effect arising through the adjustment of employed workers' reservation productivity, holding the tax rate constant. To isolate this effect, apply the chain rule pointwise in (x, y, ϵ) , holding fixed the direct argument $\tilde{\alpha}(w)$ of period utility while allowing the reservation productivity to vary in the policy-induced direction

$\partial y^*/\partial \tilde{\alpha}(w)|_\tau$:

$$\begin{aligned}
& e(1-\delta)\mu\gamma(w) \int \frac{\partial}{\partial y^*} \left\{ \lambda \int_{y^*}^x \mathbb{E}_{\tilde{\epsilon}} [u((1-\tau)w(x, \tilde{y}, \tilde{\epsilon}))] dG(\tilde{y}) \right. \\
& \quad \left. + [1 - \lambda(G(x) - G(y^*))] \mathbb{E}_{\theta} [u^{STW}(w(x, y, \epsilon); \tilde{\alpha}(w))] \right\} \frac{\partial y^*}{\partial \tilde{\alpha}(w)} \Big|_{\tau} dF(x, y, \epsilon; w) \\
&= e(1-\delta)\mu\gamma(w) \int \lambda G'(y^*) \left\{ \mathbb{E}_{\theta} [u^{STW}(w(x, y, \epsilon); \tilde{\alpha}(w))] \right. \\
& \quad \left. - \mathbb{E}_{\tilde{\epsilon}} [u((1-\tau)w(x, y^*, \tilde{\epsilon}))] \right\} \frac{\partial y^*}{\partial \tilde{\alpha}(w)} \Big|_{\tau} dF(x, y, \epsilon; w) \\
&= 0.
\end{aligned}$$

The last equality follows from the reservation-productivity condition:

$$\mathbb{E}_{\tilde{\epsilon}} [u((1-\tau)w(x, y^*, \tilde{\epsilon}))] = \mathbb{E}_{\theta} [u^{STW}(w(x, y, \epsilon); \tilde{\alpha}(w))].$$

Thus, the endogenous adjustment of y^* has zero first-order direct effect on social welfare. Equivalently, the RHS of equation (3) is zero.

Following the same steps, the component of expected utility associated with an initially unemployed worker of ability x that depends on $R^*(x)$ is

$$\begin{aligned}
\mathcal{V}^U(R^*(x)) &= (1-e)\lambda \int_{R^*(x)}^x \mathbb{E}_{\tilde{\epsilon}} [u((1-\tau)w(x, \tilde{y}, \tilde{\epsilon}))] dG(\tilde{y}) \\
& \quad + (1-e) [1 - \lambda(G(x) - G(R^*(x)))] u(b).
\end{aligned}$$

Therefore,

$$\frac{\partial \mathcal{V}^U}{\partial R^*(x)} = (1-e)\lambda G'(R^*(x)) \{u(b) - \mathbb{E}_{\tilde{\epsilon}} [u((1-\tau)w(x, R^*(x), \tilde{\epsilon}))]\} = 0,$$

where the last equality follows from the reservation-productivity condition:

$$\mathbb{E}_{\tilde{\epsilon}} [u((1-\tau)w(x, R^*(x), \tilde{\epsilon}))] = u(b).$$

Because this derivative is zero for every x , its product with any policy-induced response of $R^*(x)$ is also zero. Thus, workers' endogenous adjustment of either reservation productivity has zero first-order direct effect on social welfare. $\square$

A.14: Calibration Table

Moment	Description	Values
μ	Avg share of workers on STW	0.043
h	Avg fraction of hours worked on STW	0.47
w	FT wage at the threshold	25,224
$\mathbb{E}^{STW}[w]$	Average FT wage in STW	24,222
$\mathbb{E}[w]$	Average FT wage (excl. workers on STW)	25,794
$\gamma(w)$	Avg share of workers in STW around the threshold	0.179
$\alpha(w)$	STW replacement rate at the threshold	0.556
$\mathbb{E}[\alpha(w)]$	Avg STW replacement rate	0.518
θ	Prob. of remaining on STW in the next period	0.37
$\int \gamma(w) \int m(x, y^*) dF(x, y, \epsilon; w) dw$	Avg prob. of moving for workers on STW	0.099
$\int \gamma(w) \int m(x, y^*) \mathbb{E}[w(x, \tilde{y}, \tilde{\epsilon}) \mid \tilde{y} \in (y^*, x)] dF(x, y, \epsilon; w) dw$	Post-transition-wage-weighted prob. of moving for workers on STW	2,260
$\int \gamma(w) \int [1 - m(x, y^*)] w dF(x, y, \epsilon; w) dw$	FT-wage-weighted prob. of staying for workers on STW	22,695
$\int m(x, y^*) dF(x, y, \epsilon; w)$	Prob. of moving for workers just below the threshold	0.099
$\int \lambda G'(y^*(x, y, \epsilon)) \frac{\partial y^*}{\partial \alpha} dF(x, y, \epsilon; w)$	Partial effect of replacement rate on prob. of staying at the same firm, for workers on STW just below the threshold	0.446
$\int \lambda G'(y^*(x, y, \epsilon)) \frac{\partial y^*}{\partial \alpha} \mathbb{E}[w(x, \tilde{y}, \tilde{\epsilon}) \mid \tilde{y} \in (y^*, x)] dF(x, y, \epsilon; w)$	Partial effect of replacement rate on prob. of staying at the same firm, weighted by avg wage conditional on moving for workers on STW just below the threshold	8,294
$\int m(x, y^*) \frac{\partial \mathbb{E}[w(x, \tilde{y}, \tilde{\epsilon}) \mid \tilde{y} \in (y^*, x)]}{\partial y^*} \frac{\partial y^*}{\partial \alpha} dF(x, y, \epsilon; w)$	Partial effect of replacement rate on avg wage conditional on moving weighted by prob. of moving for workers on STW just below the threshold	12.280

Note: Annual frequency. Numeraire: EUR per year. The quantitative exercise imposes the zero-unemployment approximation, $e = 1$ and $\delta = 0$.

The empirical moments reported in Table 14 map directly into the theoretical policy trade-off. Under the zero-unemployment approximation, the wage-weighted probabilities of staying and moving determine the payroll-tax base and, together with μ , θ , and the average gross replacement rate, the implied tax rate:

$$\mathcal{T} = \mu \left[(1 - \theta) \overline{W}^S + \overline{W}^M \right] + (1 - \mu) \mathbb{E}[w], \quad \tau = \frac{\mu \theta \mathbb{E}[\alpha(w)] (1 - h) \overline{W}^S}{\mathcal{T}},$$

where $\overline{W}^S = 22,695$ and $\overline{W}^M = 2,260$ denote the wage-weighted probabilities of staying and moving reported in the table. Under log utility, the marginal value from

redistribution is

$$RV(\tilde{\alpha}(w)) = \frac{\frac{\mu\gamma(w)\theta[1-m(w)]}{h+\tilde{\alpha}(w)}}{\frac{\mu[(1-\theta)(1-\bar{m})+\bar{m}]+(1-\mu)}{1-\tau}} \frac{\mathcal{T}}{\mu\gamma(w)\theta[1-m(w)]w},$$

where $\bar{m} = 0.099$ is the average moving probability among workers on STW and $m(w) = 0.099$ is the corresponding probability at the threshold. The estimated partial effect of the gross replacement rate on the probability of staying implies

$$\epsilon_{1-m,\tilde{\alpha}(w)} = \frac{0.446}{1-0.099} [0.556(1-0.47)] \simeq 0.146.$$

Notice that $\epsilon_{1-m,\tilde{\alpha}(w)}$ denotes the elasticity of the average probability of remaining at the current firm among workers on STW earning w . Equivalently, it is the staying-probability-weighted average of the corresponding worker-level elasticities:

$$\epsilon_{1-m,\tilde{\alpha}(w)} \equiv \int q(x, y^*) \epsilon_{1-m(x,y^*),\tilde{\alpha}(w)} dF(x, y, \epsilon; w) = \frac{\partial \int [1-m(x, y^*)] dF(x, y, \epsilon; w) / \partial \tilde{\alpha}(w)}{\int [1-m(x, y^*)] dF(x, y, \epsilon; w)} \tilde{\alpha}(w).$$

Moreover, the two wage-response moments recover the response-weighted wage at the reservation productivity as $\bar{w}^* = \frac{8,294-12,280}{0.446} \simeq 18,600$. The moral-hazard cost can therefore be computed as

$$MH(\tilde{\alpha}(w)) = \epsilon_{1-m,\tilde{\alpha}(w)} \left[1 + \frac{\bar{w}^* - (1-\theta)w}{\theta\tilde{\alpha}(w)w} \tau \right].$$

Using the moments in the table gives $RV(\tilde{\alpha}(w)) = 1.33$ and $1 + MH(\tilde{\alpha}(w)) = 1.15$. The same estimated staying elasticity implies an elasticity of expected STW duration of 0.07.