*Outside options and worker motivation**

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Abstract

We study the relationship between outside options and workers' incentives to exert effort. We first set up a relational contracting model where effort is constrained by the future value of an employment relationship. To test the predictions from this model, we evaluate changes in outside options arising from age and experience cutoffs in the Austrian unemployment insurance (UI) system. Results indicate that a 9-week UI benefit extension increases worker absenteeism at the intensive margin by 0.5 days per half-year, on average. Consistent with our model predicting that these effort reductions are more pronounced if the perceived relationship value is small, we find that our effects are stronger for workers with higher potential cost of unemployment, for older workers, in declining rather than in growing firms, in low-wage firms, and for women as well as workers with children.

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I. Introduction

Effectively incentivizing workers is a fundamental determinant of firm performance (Lemieux et al. 2009, Prendergast 1999). If contracting frictions prevent the adoption of formal incentive contracts, agreements based on informal, "subjective," performance measures are widely used instead (Frederiksen et al. 2017, Kampkötter & Sliwka 2016). Their efficiency relies on the (future) value of an employment relationship, which is not only determined by its inherent productivity and stability, but also by workers' and firms' outside options. While a large theoretical literature has explored the link between outside options and incentives using models of efficiency wages (Shapiro & Stiglitz 1984, Yellen 1984) and relational contracts (MacLeod & Malcomson 1989, Malcomson 2013), systematic empirical evidence remains scarce.

In this paper, we show that better outside options indeed reduce worker effort. We first set up a theoretical model of an infinitely repeated firm-worker relationship, in which effort benefits the firm and is observable but not verifiable. Therefore, formal, court-enforceable contracts to motivate the worker are not feasible. Instead, self-enforcing relational contracts are used, where the future relationship value provides incentives to perform today. We demonstrate that the maximum effort level that can be implemented is higher for a larger future relationship surplus. While this surplus increases in the stability and inherent productivity of an employment relationship, it decreases in workers' and firms' outside options. Thus, we predict that better outside options reduce effort.

We test this prediction by exploiting age and experience cutoffs in the Austrian unemployment insurance (UI) system, which provide variation in workers' outside options by increasing the potential payoff when becoming unemployed. In particular, workers above the age of 40 gain an additional 9 weeks of potential UI benefits if they have worked at least 6 out of the last 10 years. To construct counterfactuals, we use same-age workers that are not eligible for the benefit extension because they do not fulfill the 6-year experience criterion, comparing eligible and ineligible workers before and after the benefit extension kicks in. As an empirical proxy for non-verifiable worker effort, we follow the literature (e.g., Bennedsen et al. 2019, Ichino & Maggi 2000, Ichino & Riphahn 2005) and use worker *absenteeism,* which can be measured consistently across occupations and industries.

In line with our theoretical predictions, we find that the 9-week benefit extension leads to increases in absenteeism by 0.5 days of sick leave per half-year, on average. Consistent with this being an effort response, we find that (1) the 9-week benefit extension does not affect healthcare utilization, (2) effects are much stronger when we only consider sick leaves due to 'easy-to-shirk' diagnoses such as common cold or low back pain, (3) effects are stronger for sick leaves on days with good weather, and (4) we find zero effects for placebo tests using sick leaves due to cancer as the outcome.

Our empirical design requires that ineligible workers are a valid counterfactual for eligible workers. We provide several results in support of this assumption. Most importantly, we show that eligible and ineligible workers follow parallel trends in absenteeism prior to the change in outside options becoming important. To address dynamic selection—in the sense that ineligible workers may become eligible when accumulating more time on the labor market—we show that: (1) despite significant baseline differences, the composition of eligible and ineligible workers does not change around the age-40 cutoff, (2) results are similar when we fix eligibility at a certain age, and (3) our results are robust to omitting workers that switch eligibility status after age 25.

Additionally, we provide evidence from an alternative identification strategy that exploits changes in early retirement age (ERA) laws to validate our baseline findings.[1] In particular, the Austrian government enacted two reforms that gradually increased the ERA from 60 to 65 for men and 55 to 60 for women based on quarter-of-birth cohorts. We argue that a higher ERA increases the future value of the relationship because the expected end of a worker's career at the margin of retiring is postponed. Therefore, we predict that a higher ERA leads to a decrease in absenteeism. Indeed, results from a fixed effects model which is identified by changes in the ERA between sick leaves of a worker indicate that a one-year increase in the ERA decreases average sick leave durations by around 0.5 days, on average.

[1] The reason we use this second source of identification solely as a validation check is because the potential complier population—that is, people old enough to be affected by the reforms and on the margin of retiring early—is much smaller than in our main model.

To gain a better understanding of the mechanisms underlying our main result, we use our theoretical model to derive additional predictions. Some of these predictions make use of the result that the negative effect of a higher outside option on worker motivation is more pronounced if the initial perceived relationship value is smaller. The reason is that, if a small relationship value effectively constrains effort, a further reduction in the relationship value causes a stronger effort response than if this constraint is less relevant to begin with. We then identify variables in our data which we argue shape the relationship surplus.

First, we find that our effect is stronger for blue-collar workers, who face a higher risk of becoming unemployed and tend to remain unemployed longer than white-collar workers. Therefore, in blue-collar jobs, the relationship stability is systematically smaller than in white-collar jobs, and unemployment benefits assume a larger role in a worker's outside option. Second, the effect becomes larger over time as workers age, which we argue is because of the smaller relationship surplus of workers that are closer to retirement. Theoretically, this can be explained by an eventually declining probability of continuing employment. Third, the negative effect of a higher outside option on effort is stronger for declining rather than growing firms, which again reflects a lower job stability among the former. Fourth, the effect is smaller for high-wage firms, which we identify using estimated Abowd, Kramarz & Margolis (1999, AKM hereafter) firm wage fixed effects. Following the literature, we suggest that firms with higher AKM fixed effects are inherently more productive, which increases the value of their employment relationships. Fifth, workers with children react stronger to higher UI benefits than those without, and this effect is particularly pronounced for women. The reason is that the replacement rate provided by UI benefits is substantially higher for workers with dependents, who therefore benefit more when being eligible for an extension and reaching the threshold; moreover, we argue that being responsible for one's children corresponds to higher opportunity costs of exerting effort and thus implies a smaller relationship surplus. Finally, the effect is smaller if we compare eligible and non-eligible workers within the same firm, which we claim can be an indicator for some firms having formed multilateral relational contracts (Levin 2002). In such a case, a firm has a relational contract with its workforce as a whole, and changes in the outside options of some workers affect all within-firm relationships alike.

We also argue that the observed link between UI benefits and worker absenteeism cannot be generated by alternative models such as a competitive labor market and search-and-matching models. If the labor market is competitive, i.e., firms can directly "purchase" the worker's productivity (for example because frictionless formal contracts can be used to motivate effort), one would rather expect a negative (if any) effect of higher UI benefits on absenteeism or a positive effect on inherent worker productivity. If employed workers conduct on-the-job search and more search raises their chances of being fired because of the reduced motivation, higher UI benefits would indeed increase absenteeism. But then we would expect the more intense search to generate more outside offers and consequently more job-to-job transitions. To the contrary, we do not find that these transitions go up for eligible workers who pass the cutoff age.

Our paper ties together several strands of the literature. Most importantly, we complement recent work by Jäger, Schoefer, Young & Zweimüller (2020), who show that changes in potential UI benefits have no effects on worker wages. This result seems at odds with the Nash bargaining model widely used in labor search theory, which predicts relatively large positive wage changes when the worker's outside option increases. Our findings suggest that changes in outside options can have real consequences on employment relationships and, in particular, affect their efficiency.[2]

While other papers have considered effort responses to changes in outside options, these papers are mostly descriptive single-firm case studies. Cappelli & Chauvin (1991) show that higher wage premia and higher local unemployment are both associated with fewer disciplinary problems in a large US manufacturing firm. Similarly, Lazear, Shaw & Stanton (2016) use data from a US services firm and observe that worker productivity was significantly higher during the 2009 recession, and that this increase was particularly strong in areas with high unemployment.[3] We are aware of only one design-based paper that studies a similar question. Lusher et al. (2022) use scanner data and

[2]We note that Jäger et al. (2020) show results on the share of months spent on sick leave in an appendix too, but they only observe sick leaves longer than 6 to 12 weeks, depending on job tenure, where social security steps in and picks up half of the worker's wage bill. We can use more granular data on individual sick leaves, which allows us to identify even small effects at the intensive margin of leave taking.

[3]Further support for a positive correlation between the unemployment rate and worker effort is provided by Scoppa & Vuri (2014) for Italy and Burda et al. (2020) for the US.

US state-level variation in UI benefit levels to estimate effects on supermarket cashier productivity. They find that transaction length increases by 2.4 seconds or 2 percent for cashiers who experienced an 18-week increase in the potential benefit duration.

We also contribute to the literature providing empirical evidence for the presence and characteristics of relational contracts (Fahn et al. 2017, Gil et al. 2022, Gil & Marion 2013, Gil & Zanarone 2018, Macchiavello 2022, Macchiavello & Morjaria 2015, 2021). While these contributions mostly rely on between-firm relationships in single industries or markets, we have access to the universe of employment relationships in Upper Austria and present evidence indicating that the respective theoretical mechanisms are also relevant in *within-firm* relationships—that is, between firms and their employees. Furthermore, we propose an additional test to distinguish a relational contracting mechanism from potential alternative explanations. Therefore, our predictions are not only based on changes in the relationship value or reneging temptations (the standard approach in the literature), but also on comparisons of effect sizes based on the ex-ante relationship value.

To conclude, we make three distinct contributions. First, we provide causal evidence for the effect of outside options on productivity for an entire workforce instead of a single firm (Cappelli & Chauvin 1991, Lazear et al. 2016) or a specific occupational group (Lusher et al. 2022). Second, we suggest an additional tool to better assess the mechanisms driving this effect by applying the theoretical result the effect of relationship value or reneging temptations on effort is more pronounced if the surplus has initially been smaller. Third, using absenteeism as a proxy for non-verifiable effort, we suggest a measure that is more readily available in register data than alternative measures of productivity in individual employment relationships and can therefore be used to analyze additional effort dimensions within firms.

The rest of the paper is structured as follows. In section II, we set up the relational contracting model and derive first predictions. In section III, we turn to the empirical analysis and first discuss the institutional setting, the design we use, and the data, before we turn to presenting our main result and several robustness checks that lend support to our findings. In section IV, we then derive additional theoretical predictions that we each test using our empirical model. Section VI concludes.

II. The relationship between outside options and worker effort

In this section, we derive a relational contracting model to formalize the relationship between outside options and worker effort. We organize this section as follows: In subsection II.1, we lay out the model environment. In subsection II.2, we formalize payoffs and the first best. In subsection II.3, we introduce relational contracts to the model. In subsection II.4, we describe the optimization problem. In subsection II.5, we derive comparative statics. Finally, in subsection II.6, we discuss the role of outside options and potential UI benefits in the model.

II.1. Environment

In every period t of an infinite time horizon, a risk-neutral principal/firm ("she") makes an employment offer to a risk-neutral agent ("he"). The offer contains an upfront wage $w_t \geq 0$ and a discretionary bonus $b_t \geq 0$. We describe the agent's acceptance decisions with $d_t \in \{0, 1\}$, where $d_t = 1$ corresponds to an acceptance and $d_t = 0$ to a rejection. Upon acceptance, the agent receives w_t and chooses an effort level $e_t \in \mathbb{R}_+$ which is associated with effort costs $c(e_t)$, where $c'(\cdot)$, $c''(\cdot) > 0$, $c'''(\cdot) \geq 0$ and $c(0) = c'(0) = 0$. Effort generates an (expected) output $e_t\theta$ (with $\theta > 0$) which is subsequently consumed by the principal.

Future payoffs are discounted with a common factor $\delta \in (0, 1]$; δ not only captures time preferences, but also reflects the probability with which the relationship is continued. Continuation probabilities can be driven by industry- or firm-wide, as well as personal characteristics.

II.2. Payoffs and first best

If the agent rejects the principal's offer in a given period, both consume their outside options which are $\bar{\pi} \in \mathbb{R}^+$ for the principal and $\bar{u}_t \in \mathbb{R}^+$ for the agent. Thus, players' discounted payoff streams in a period t are

$$U_t \equiv \sum_{\tau=t}^{\infty} \delta^{\tau-t} \left[d_\tau \left(w_\tau + b_\tau - c(e_\tau) \right) + (1-d_\tau)\, \bar{u}_\tau \right]$$

$$\Pi_t \equiv \sum_{\tau=t}^{\infty} \delta^{\tau-t} \left[d_\tau \left(e_\tau \theta - w_\tau - b_\tau \right) + (1-d_\tau)\, \bar{\pi} \right].$$

Moreover, $\bar{U}_t \equiv \sum_{\tau=t}^{\infty} \delta^{\tau-t} \bar{u}_\tau$ and $\bar{\Pi} \equiv \bar{\pi}/(1-\delta)$. We also define

$$S_t \equiv \Pi_t + U_t$$

as the period-t surplus, and $\bar{S}_t \equiv \bar{\Pi} + \bar{U}_t$. Thus, the per-period surplus if $d_t = 1$ equals $e_t\theta - c(e_t)$, and first-best effort e^{FB} is characterized by

$$\theta - c'(e^{FB}) = 0. \tag{1}$$

For the following, we assume that

$$e^{FB}\theta - c(e^{FB}) > \bar{\pi} + \bar{u}_t \tag{2}$$

holds for all t. Therefore, it is efficient for the agent to work for the principal if e^{FB} is implemented.

II.3. Contractibility, payoffs, and relational contract

We assume that effort and output are observable but not verifiable to a third party. Therefore, formal incentive contracts are not possible, and only a self-enforcing *relational contract* can (potentially) be formed. In our setting without asymmetric information, it determines a subgame perfect equilibrium of the game.

We derive a relational contract that maximizes the total surplus at the onset of the game, S_1. However, note that predictions would be the same if our objective was to maximize the principal's profits, the agent's utility, or any weighted average of those (Levin 2003).

Discussion of Assumptions. Before deriving equilibrium outcomes and empirical predictions, we briefly discuss our modelling assumptions. First, effort in our setting relates to the agent's motivation on the job, and we abstract from other, easily measurable, aspects such as working hours. Such measurable dimensions could be taken care of by not-further-modeled incentive contracts. Then, our results survive as long as, without this "extra" motivation, employing the agent would not be valuable. Second, observability of the agent's effort is not important for our results. In section D.2 of the Web Appendix, we demonstrate that if effort is the agent's private information and generates an observable output measure, all our predictions survive. Third, although we use the term "wage" when referring to the agent's compensation, it might go beyond monetary payments. In particular if salaries are constrained by collective bargaining agreements or contractual obligations, firms can be restricted in setting them. Therefore, the wage in our setup reflects everything that is costly to the principal and valued by the agent. For example, it might include good working conditions, flexibility in working times, or perks. Moreover, in section D.1 of the Web Appendix, we show that our predictions hold up even if the agent's compensation is taken as given and only firing threats are used to provide incentives. Thereby, we also link our approach to classic efficiency-wage models and demonstrate that the underlying mechanisms are closely related.

II.4. Optimization Problem

Our objective is to maximize S_1. Outcomes are restricted by a number of constraints which must be satisfied in all periods t. In the following, we display these constraints for an equilibrium in which, on the equilibrium path, the agent accepts the employment offer in every period. Later on, we make precise which conditions must hold for this to be optimal.

Since all deviations from equilibrium play are publicly observable, it is optimal to punish any deviation by a reversion to the worst possible outcome for the deviator (Abreu 1988), which in our case is consuming their outside options forever thereafter.[4] First, it must be in the agent's interest

[4] However, note that equilibrium outcomes would be the same if a deviation did not lead to a termination, but instead to a continuation of the relationship in which the deviator would only receive their outside option (Levin 2003).

to accept the firm's offer, which is captured by his participation constraint (PC),

$$U_t \geq \bar{U}_t. \tag{PC}$$

Given he has accepted the contract, equilibrium effort e_t must satisfy his incentive compatibility constraint (IC),

$$-c(e_t) + b_t + \delta U_{t+1} \geq \delta \bar{U}_{t+1}. \tag{IC}$$

This takes into account that, if the agent decides to deviate from equilibrium effort, he will choose zero effort instead. Note that, if effort were verifiable, a formal incentive contract would only need to satisfy these sets of constraints. Such a contract could induce the agent to choose e^{FB} in every period, for example by setting $w_t = \bar{u}_t$ and $b_t = c(e^{FB})$.

Since formal incentive contracts are not feasible, also the principal needs incentives to pay the b_t specified by the relational contract. This is captured by so-called dynamic enforcement (DE) constraints,

$$-b_t + \delta \Pi_{t+1} \geq \delta \bar{\Pi}. \tag{DE}$$

In words, the (DE) constraint implies that paying b_t must be profitable for the principal which it is only if the subsequent continuation profits are sufficiently high compared to her payoff after a termination. Finally, given $b_t \geq 0$, a participation constraint for the principal ($\Pi_t \geq \bar{\Pi}$) is implied by (DE) and can hence be omitted. To conclude, our objective is to maximize S_1, subject to (PC), (IC), and (DE) which must hold in every period t.

II.5. Results

II.5.1. Preliminaries

First, we simplify the optimization problem and obtain the following results.

Lemma 1. *The equilibrium is sequentially efficient, i.e., maximizing S_1 is equivalent to maximizing $e_t\theta - c(e_t)$ in every period t. Moreover, given $\Pi_t \geq \bar{\Pi}$ and $U_t \geq \bar{U}_t$, the following enforceability* *constraint* (EC) *is necessary and sufficient for implementing equilibrium effort* e_t^**:*

$$-c(e_t^*) + \delta S_{t+1} \geq \delta \bar{S}_{t+1}. \tag{EC}$$

These results follow from Levin (2003). Sequential efficiency implies that destroying surplus on the equilibrium path cannot improve incentives, hence on the equilibrium path the agent is never fired. The (EC) constraint is obtained by adding the (IC) and (DE) constraints. It states that the cost of exerting effort today must be covered by the net future value of continuing the relationship and captures the fundamental mechanism of relational contracts. Sufficiency follows because of substitutability between current and future incentives (if this condition holds, a payment scheme exists that satisfies the individual constraints stated above).

Lemma 1 implies that, if e^{FB} satisfies (EC) in period t, $e_t^* = e^{FB}$. Otherwise, $e_t^* < e^{FB}$ and is determined by the binding (EC) constraint. Finally, for the employment relationship never to be terminated on the equilibrium path, $e_t^*\theta - c(e_t^*) \geq \bar{\pi} + \bar{u}_t$ must hold for all t.

II.5.2. Comparative Statics

Next, we demonstrate how the level of the inherent marginal productivity of effort, θ, and the discount factor determine equilibrium outcomes.

Lemma 2. *S_t strictly increases in θ and δ. It weakly decreases in $\bar{S}_{t+1}$.*

The proof of Lemma 2 can be found in Appendix B. Higher θ and δ have a direct positive effect on the equilibrium surplus, which is further amplified by a relaxed (EC) constraint. An increase in outside options tightens (EC) and thus potentially reduces the equilibrium surplus. Next, we present a general result that is the foundation of the predictions we derive later on.

Proposition 1. *Assume* (EC) *binds in a t. Then, equilibrium effort e_t^* decreases in $\bar{S}_{t+1}$. This effect is more pronounced if S_{t+1} has initially been smaller or if $\bar{S}_{t+1}$ has been larger. If* (EC) *in t is slack, equilibrium effort e_t^* is unaffected by a marginal change in $\bar{S}_{t+1}$.*

The proof of Proposition 1 can be found in Appendix B.

II.6. Outside options

Now, we put more structure on the development of the agent's outside option to better relate to the empirical environment we study. There, we analyze the consequences of an anticipated increase of employees' unemployment benefits at the age of 40 which we argue corresponds to a permanent increase in the agent's outside option. Hence, we assume that there is a $T > 1$ such that

$$\bar{u}_t = \begin{cases} \bar{u} & \text{for } t < T \\ \bar{u}^H & \text{for } t \geq T, \end{cases} \tag{3}$$

with $\bar{u}^H > \bar{u}$.

Proposition 2. *For all periods $t \geq T-1$, equilibrium effort $\tilde{e}$ is constant; moreover, there is a $\tilde{\delta} < 1$ such that $\tilde{e} = e^{FB}$ for $\delta \geq \tilde{\delta}$ and determined by $-c(\tilde{e}) + \delta\left[\tilde{e}\theta - \left(\bar{\pi} + \bar{u}^H\right)\right] = 0$ otherwise.*

If $\delta \geq \tilde{\delta}$, $e = e^{FB}$ also in all previous periods $t < T-1$. If $\delta < \tilde{\delta}$, $e_t > \tilde{e}$ for all $t < T-1$. Then, there exist $\tilde{\delta}_t < \tilde{\delta}$ such that $e_t = e^{FB}$ for $\delta \geq \tilde{\delta}_t$ and determined by the binding (EC) otherwise, with $\tilde{\delta}_t$ increasing. Finally, $e_t \geq e_{t+1}$, with a strict inequality if $\delta < \tilde{\delta}_{t+1}$.

The proof of Proposition 2 can be found in Appendix B. Proposition 2 states that equilibrium effort is smaller in later than in earlier periods. This effort reduction is caused by the change in the agent's outside option which permanently increases in period t. Importantly, the effort reduction already unfolds in period $T-1$ or earlier because today's effort is constrained by the *future* relationship value. Moreover, moving to earlier periods diminishes the weight of the higher outside option which steadily increases the (future) value of the relationship. Therefore, equilibrium effort goes up as we move backwards, either until the very first period or until e^{FB} can be implemented.

II.6.1. First Prediction

As an agent's outside option also contains his payoff when being unemployed, Proposition 2 directly yields the first prediction.

Prediction 1. *A permanent increase in UI benefits at a given age permanently reduces an affected worker's effort. This effort reduction already materializes earlier, before the increase in UI benefits is realized.*

III. Empirical analysis

In this section we establish our main empirical result. It is organized as follows. We first discuss the institutional setting in subsection III.1, covering details about the social security system, the labor market, and sick leaves in Austria. In subsection III.2, we discuss our empirical design. In subsection III.3, we summarize our data. In subsection III.4, we show the main result for the effect of outside options on worker incentives. In subsections III.5 and III.6, we provide several robustness checks and discuss whether our proxy for worker motivation is viable. Finally, in subsection III.7 we show results from an alternative identification strategy as a validation exercise.

III.1. Institutional setting

III.1.1. Social security and the labor market

Austria has a *Bismarckian* social security system with universal access to public healthcare, pension, disability, and unemployment benefits. Workers are automatically enrolled to the system, but insurance is also extended to spouses and children, unemployed people, pensioners, and disabled people. In this paper we focus on Upper Austria, which is one of the nine Austrian federal states with around 1.5 million residents or 20 percent of the Austrian population. The labor market is characterized by broad institutional regulation with centrally bargained wages and working conditions. At the same time, the labor market is highly flexible, with particularly weak job protection (OECD 2020) and high turnover (Böheim 2017).[5] Employment contracts can generally be terminated without specifying a reason, but unilateral terminations require a notice period be observed. Note that, although our model seems to allow for more wage-setting flexibility than the

[5] In terms of the OECD employment protection legislation indicator, Austria places 33$^{\text{rd}}$ of 37 countries, with the United States ranking last. Job turnover rates are 7.9 percent for men and 8.3 percent for women, which are larger than the European Union averages of 6.7 percent for men and 7.4 percent for women.

Austrian labor market, we show in subsection D.1 that our results can also be generated in a model where compensation is given and only firing threats are used to incentivize workers.

III.1.2. Unemployment insurance

Austria's UI program is compulsory and funded through a 6 percent payroll tax that is shared equally by workers and firms. It applies to all workers who earn more than the marginal employment threshold, which was €438.05 per month in 2018. The minimum replacement rate amounts to 55 percent of daily net income, which is calculated based on pre-unemployment wages. Workers with dependents can be eligible for replacement rates of up to 80 percent. A requisite to receiving UI benefits is that claimants are willing and able to work. This implies that they must prove they frequently apply for new jobs and undergo retraining, if necessary. Benefits for laid-off workers are payable immediately upon entry into unemployment, for job quitters there is a one-month waiting period. After UI benefits are exhausted, unemployed people are eligible for means-tested income support.

The potential duration for UI benefits changes discontinuously at age and experience cutoffs. Baseline eligibility is 20 weeks for workers that have been employed for at least one year. After a total of three years of employment, the potential benefit duration is 30 weeks. At age 40, benefits are extended to 39 weeks, provided that the worker has been employed for at least 6 out of the last ten years prior to claiming UI. At age 50, benefits are extended up to a year conditional on having worked for at least 9 out of the last 15 years. In this paper, we focus on the age-40 cutoff for two main reasons. First, almost all workers in the labor market are eligible for at least 30 weeks of benefits, so the experience cutoff extending eligibility from 20 to 30 weeks is not informative.[6] Second, we know from previous literature (Ahammer & Packham 2020, Nekoei & Weber 2017) that the age-50 cutoff does not affect actual UI benefit takeup for people in unemployment, hence it is also unlikely that it represents a change in outside options for workers not currently unemployed.

[6]In 2018, 92 percent of all 30 year old non-marginally employed workers were eligible for at least 30 weeks of benefits. At age 40, this share increases to 97 percent.

III.1.3. Sick leaves

Sick leave insurance in Austria compensates workers' earnings losses due to both occupational and nonoccupational disease. Workers are entitled to full wage compensation for 6 to 12 weeks, depending on job tenure. After this period, workers receive 80 percent of their wage for another 4 weeks, but the wage bill is shared equally between firms and social security. After these 4 weeks, workers are entitled to public sickness benefits that replace 60 percent of the current wage (Halla et al. 2015).

To take sick leave, workers have to produce a sick note to the employer. These sick notes are usually issued by primary-care physicians, who also directly report the sick leave to the health insurance provider. Sick notes do not reveal a medical diagnosis to the employer, and it is forbidden for employers to ask workers to disclose a diagnosis. Importantly, workers are not obliged to produce sick notes for leaves of less than 4 days, unless the firm explicitly requires it. Firms are generally free to enforce such a rule, and there are no further contracts or agreements necessary. In our data, it is, in fact, quite common: 97.6 percent of firms have at least one short sick leave per year and 50 percent of their sick leaves are short leaves. This will cause some measurement error in our regressions, which will lead us to underestimate the effect of outside options on absenteeism.

III.2. Design

To test for the effect of a change in outside options on worker absenteeism, we estimate differences in sick leave takeup before and after a 9-week UI benefit extension kicks in at age 40 for workers with enough experience (the 'eligible group'). To avoid comparing older with younger workers, we use workers without enough experience (the 'non-eligible' group) to partial out age effects. Before we discuss the specifics of our empirical strategy, we note two important design choices. First, we focus mostly on intensive margin responses in absenteeism. This is because we care most about the marginal decision whether to stay home a day more or not, conditional on being on sick leave. Therefore, we construct our data set on the sick leave level, and our primary outcome measures the duration of a sick leave j of individual i. Second, because Prediction 1 states that we expect workers to already react to the change in outside options at age 40 *before* they actually turn 40, we need to consider changes relative to some arbitrary reference period b. For now, we fix b at 37.5 years of age, but we show in section III.5 that the reference period choice does not affect our overall conclusions.

The estimand we are interested in is

$$\beta = \left(\overline{\text{duration}}_{t>b}^{\text{eligible}} - \overline{\text{duration}}_{t\leq b}^{\text{eligible}}\right) - \left(\overline{\text{duration}}_{t>b}^{\text{ineligible}} - \overline{\text{duration}}_{t\leq b}^{\text{ineligible}}\right), \tag{4}$$

where we subtract the change in average sick leave duration before and after the UI benefit extension becomes important at age b between workers that are eligible for the benefit extension (the left term) and those not eligible (the right term). We can write this in regression form as

$$\text{duration}_{jit} = \beta \left(\text{eligible}_{jit} \times \mathbb{1}[t > b]\right) + X'\gamma + u_{jit}, \tag{5}$$

where duration_{jit} is the duration of sick leave j of individual i at age t, eligible_{jit} is one if worker i is eligible for the benefit extension at time t and zero else, $\mathbb{1}[t > b]$ indicates the post-treatment period, and X is a vector of covariates that includes flexible tenure and year fixed effects, a blue collar dummy, a female dummy, as well as the eligibility and post-treatment dummy individually. The main coefficient of interest is β, which measures the average effect of the UI benefit extension on eligible workers. In event study form, equation (5) reads

$$\text{duration}_{jit} = \sum_{k\neq b} \beta_k \left(\text{eligible}_{jit} \times \mathbb{1}[t = k]\right) + X'\gamma + u_{jit}, \tag{6}$$

where $\mathbb{1}[t = k]$ indicates age k.

A key assumption is that the difference in sick leave taking between eligible and ineligible workers would remain constant absent the UI benefit extension at age 40. In support of this assumption, we show that eligible and ineligible workers follow parallel trends in sick leave takeup before the change in outside options becomes important at age b. We also note that, in our baseline specification, workers are allowed to switch from being ineligible to being eligible. This is because we track worker outcomes over a long period of time, and we do not want to lose information of workers that switch eligibility status in our baseline model. However, we provide robustness checks which show that our results are not sensitive to (1) fixing eligibility status at age 37.5 and (2) dropping workers that switch eligibility status at some point.

III.3. Data

We combine two sources of administrative data, which allow us to track workers over time and observe their sick leave taking. First, we use data from the *Austrian Social Security Database* (ASSD, Zweimüller et al. 2009). The ASSD is structured as a matched employer-employee panel that covers the universe of Austrian workers from 1972 to today. We use the ASSD to obtain individual-level employment histories, wages, and basic demographic information. A drawback of the ASSD is that it does not contain working hours and that wages are top-coded at a social security contribution cap.

Second, we have access to health records from the *Upper Austrian Health Insurance Fund* (UAHIF). The UAHIF is the main health insurance provider in Upper Austria, covering all private-sector employees apart from those working in railway and mining.[7] The UAHIF database contains information on healthcare service utilization in both the inpatient and outpatient sector, including drug prescriptions, hospital days, and physician visits. Most importantly, however, the UAHIF tracks sick leaves for all private-sector employees, which we can match to our employer-employee data set. Sick leaves are recorded as the actual number of days an employee stays away from work, not the expected duration on sick notes (these may not necessarily coincide, for example if the worker returns to work early). We also observe a primary diagnosis for each sick leave, which is coded according to the *International Classification of Diseases, 2010 revision* (ICD-10).

To construct our sample, we first draw all sick leaves between 1998 and 2018 that can be matched to employment spells in the ASSD data. If a worker has multiple jobs, we match the one where they

[7] We have no health information on public sector employees, including teachers, as well as farmers and self-employed persons.

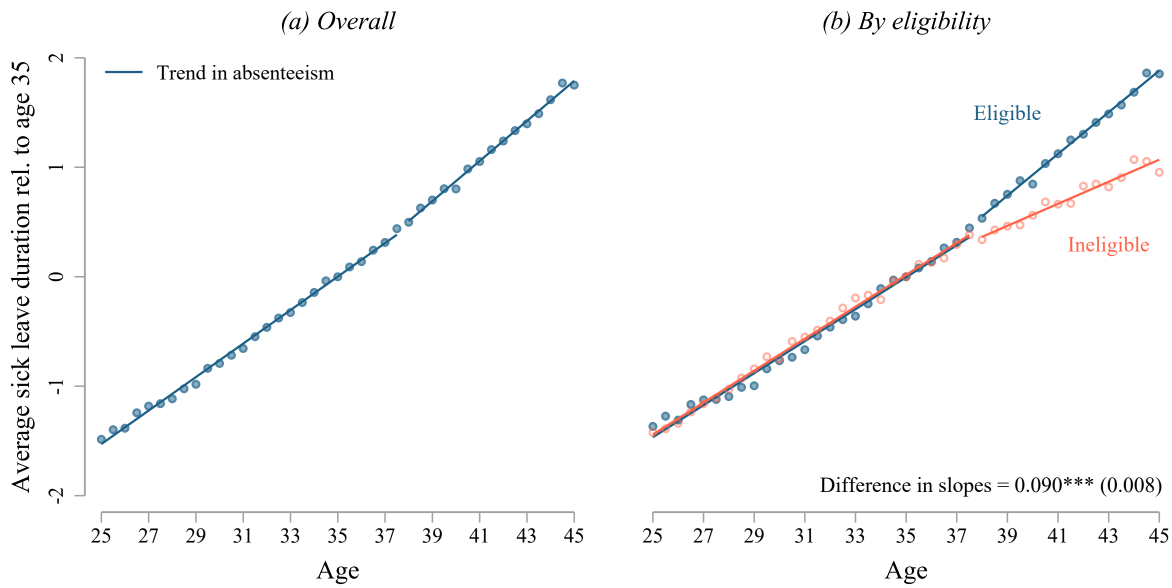
earn the highest wage in a given year. This gives us a total of 9,577,046 sick leave spells for 889,889 workers. We then drop 258,660 sick leaves that are taken by marginally employed workers, because they are not covered by UI. For our baseline analysis, we only consider workers aged 25 through 45, which leaves us with an unbalanced panel of 4,664,982 sick leaves for 558,290 workers.

Descriptive statistics are in Appendix Table A.1. The average duration of a sick leave in our data is 7.22 days and average experience is 10 years. Perhaps unsurprisingly, there are significant baseline differences between eligible and ineligible workers. Eligible workers are much less likely to be female, slightly less well educated, and much less likely to be parttime workers. Also, they have longer job tenure and higher wages, on average. Importantly, however, the composition of eligible and ineligible workers is remarkably stable over time and does not change around the age-40 cutoff. We report shares of the variables in panel (c), where we discretize tenure and wage using their respective sample medians, in Appendix Figure A.1. All variables move almost in parallel, which suggests that systematic compositional changes at age 40 are very unlikely to explain the findings below.

III.4. The effect of outside options on worker outcomes

We start by descriptively examining absenteeism patterns over worker age. In Figure 1, panel (a), we plot average sick leave durations by half-year of age for the entire workforce. We see that sick leaves generally become longer the older workers get, with leaves being, on average, 2 days longer at age 45 than at age 35. In panel (b), we divide workers by whether they experience a sharp increase in outside options because they are eligible for the UI benefit extension at age 40. Importantly, eligible and ineligible workers are on almost exactly the same trend prior to the change in outside options. Consistent with our theoretical prediction that an anticipated change in outside options affects behavior already in earlier periods, we see a gap opening around age 38, with eligible workers taking longer and longer sick leaves compared to ineligible workers. The difference in slopes amounts to 0.09 days of sick leave per half-year of age, which is significantly different from zero at the 1 percent level. We consider this as strong evidence that outside options play a role for

Figure 1 — Absenteeism over age by UI benefit extension eligibility



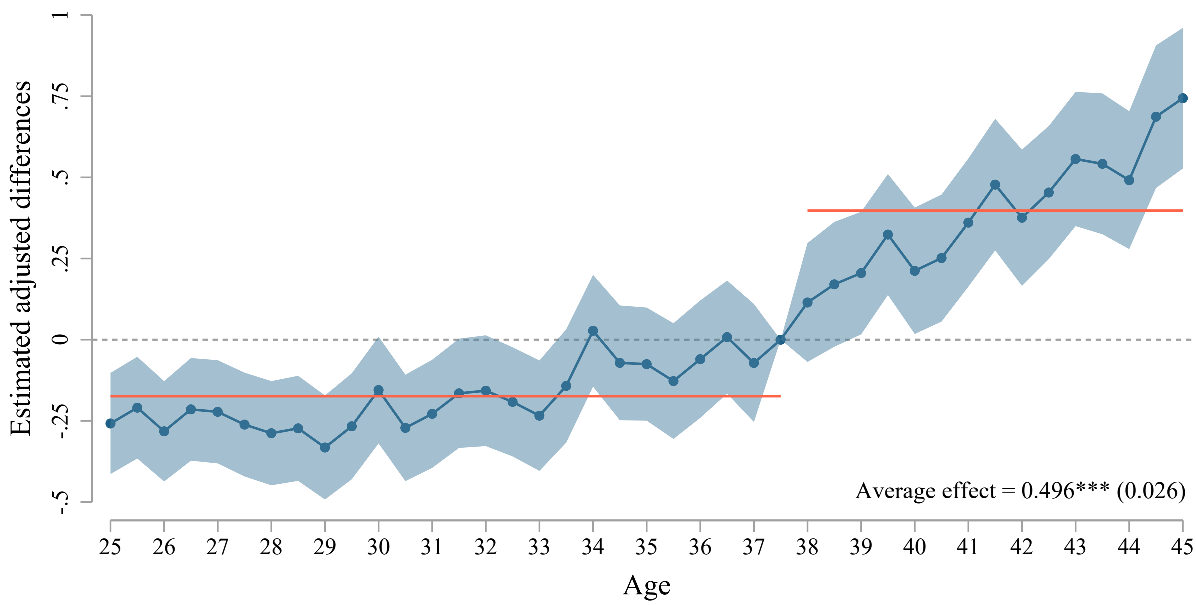
Notes: These figures show trends in average sick leave durations, conditional on taking at least one sick leave, per half-year of age over all Upper Austrian workers (panel a) and by whether workers are eligible for the 9-week UI benefit extension that, we argue, constitutes an increase in outside options (panel b). In both graphs, we center average sick leave durations around the average sick leave duration for workers at age 35 in the sample. In panel (b), we do this separately for eligible and ineligible workers.

absenteeism and that this pattern is already apparent in the raw data.

In Figure 2, we plot the differences between the trends for eligible and ineligible workers (the red and blue lines) from Figure 1 in an event study, similar to equation (6). Here we control also for tenure and year fixed effects, occupation, and gender. Our estimates suggest that the gap in sick leave taking between eligible and ineligible workers starts slowly opening up around age 34. At age 37.5, however, the difference becomes progressively larger. Comparing absenteeism before and after age 37.5 suggests that the 9-week UI benefit extension increases absenteeism, on average, by around 0.5 days per half-year of age.[8] Compared to the average sick leave duration in our sample,

[8]We also estimate our event study on wages instead of sick leaves in Appendix Figure A.2. Consistent with Jäger et al. (2020), we find no evidence that a change in the value of unemployment that comes with the UI benefit extension has an effect on log wages. While the average effect is positive, almost none of the event study coefficients is significantly different from zero. Reassuringly, we also find no evidence that eligible and ineligible workers are on different wage trajectories prior to treatment. Note that our theoretical approach does not yield wage predictions since we maximize the total surplus and allow for any distribution of it. However, if we assumed that each gets their outside option plus a fixed share of this surplus (as standard bargaining models do), our wage predictions would be ambiguous and therefore in line with an absent wage effect: Whereas a higher outside option directly increases the agent's compensation, the

FIGURE 2 — The effect of an increase in outside options on absenteeism



Notes: This figure plots event study estimates from equation (6) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on sick leave duration. Because we expect eligible workers to react already prior to age 40, we fix the reference period to $b = 37.5$, and point estimates can be interpreted as changes in average sick leave duration due to the benefit extension at a given age relative to age 37.5. The shaded area represents a 95 percent confidence band. The red horizontal lines indicate averages of age-specific estimates for both the pretreatment period $t < b$ and the posttreatment period $t \geq b$. The average effect estimate is from equation (5). All regressions control for tenure and year fixed effects, gender, and a blue collar dummy.

this is equivalent to a 6.9 percent increase. In section III.5 we discuss how the estimate changes if we vary the reference period b.

III.5. Robustness

In this section we discuss how robust the effect of outside options on absenteeism we established before is to different specification and design choices. In Appendix Table A.2, we first test how robust our average effect estimates are to using different regression specifications. In column (1), we estimate model (5) without any covariates. In column (2), we add controls for occupation, gender, and tenure as well as year fixed effects. This is our preferred specification. In column (3), we additionally control for parttime status, education, and wage. Because we do not observe parttime

indirect effect of a reduced surplus is negative.

status and education for all workers, we replace missing values with zero and add a missing indicator dummy to the regression. Because these control variables are potentially endogeneous, we omit them from our other regressions. The average effect estimate, however, increases in magnitude. In columns (4) and (5), we estimate the model using worker fixed effects, without covariates (column 4) and controlling for occupation and tenure and year fixed effects (column 5). This also leads to very similar average effect estimates compared to our baseline.

One important aspect of our design is the choice of an appropriate reference period. Since Prediction (1) tells us we should expect workers to react already before the UI extension actually kicks in, it is not immediately obvious what reference period to evaluate effects against. In Appendix Figure A.3, we therefore estimate equation (5) for different reference periods b from age 30 through age 39.5, which is the last half-year before the increase in outside options actually kicks in. This does not change our main conclusion at all, and estimates are remarkably stable across reference periods. In fact, point estimates range from $\hat{\beta} = 0.429$ with $b = 30$ and $\hat{\beta} = 0.5$ with $b = 38$, so the difference between the smallest and the largest estimate is only 0.07 days of sick leave per half-year of age.

Another important design choice we make is that workers are allowed to switch from being ineligible to being eligible and vice versa. This is because we consider sick leaves over a relatively long time span, and we want to allow workers to contribute to estimated effects both at times when they are ineligible and when they are eligible. In Appendix Table A.3 we provide evidence from two alternative constraints on eligibility status. First, in column (2), we fix eligibility at age 37.5. This means that we consider workers that did not accumulate enough experience by age 37.5 as untreated, even if they cross the experience threshold later on.[9] Second, in column (3), we drop workers that switch between being ineligible and being eligible at some point. This does not affect our main conclusions. In fact, point estimates become much larger if we use more restrictive treatment definitions.

[9]This is similar to how other papers in the literature define treatment status in settings where treatment status can change over time (e.g., Harju, Jäger & Schoefer 2021).

III.6. Do changes in absenteeism actually measure changes in motivation?

A crucial assumption we rely on is that absenteeism is indeed a good proxy for worker motivation. We are not the first to use this approach. For example, Bennedsen, Tsoutsoura & Wolfenzon (2019) estimate AKM models to separate worker and firm components in motivation, which they measure using days of sick leave, Ichino & Riphahn (2005) consider absenteeism to test for the effects of employment protection on worker effort, and Ichino & Maggi (2000) use sick leave taking to study regional shirking differentials in Italy. Absenteeism has several advantages; it can be consistently measured across occupations and industries and is readily available in administrative data sources. Also, we find that sick leaves are negatively correlated with the local unemployment rate (Figure A.4), which is consistent with earlier results (Lazear et al. 2016). A natural question is, however, if the change in outside options affects worker health and, if so, whether we can separate the health effect from the motivation effect.

Generally, we believe that health effects are not a problem for our design, and we provide several pieces of evidence in support of this assertion. Most importantly, there is no evidence that the UI extension affects healthcare takeup. In Appendix Figure A.5, we run our event study from equation (6) on total healthcare expenses, which is the sum of physician fees, drug expenses, and inpatient expenses. Our estimates suggests that there is no difference in expenses between eligible and ineligible workers, neither before nor after UI benefits are extended. We note that, even in the presence of health effects, our model is still informative. Assume a worker takes sick leave for legitimiate reasons—that is, they are actually sick. On the margin, motivation can still play a role for the decision whether to extend the sick leave by a day or not. This is one of the reasons why we focus on intensive margin effects in this paper, studying the length of individual sick leaves and not days of sick leave overall.

Additionally, in Appendix Figure A.6, we provide heterogeneity estimates by diagnosis of the sick leave and by weather conditions. We find that the change in outside options affects sick leaves with easy-to-fake diagnoses (common cold and low back pain) and those starting on days with good weather more.[10] If sick leaves are taken purely for health reasons, we should not see this effect heterogeneity. This is consistent with research showing that only little of the variation in sick leaves can be explained by differences in health status. Ahammer & Schober (2020) find that only 28 percent of the variance in sick leaves can be explained by patient observables or variation in physician prescribing behavior, the rest remains unexplained. Also in our data, a simple regression of aggregate days of sick leave in a year on cubics in physician fees, drug expenses, and inpatient expenses returns an R^2 of around 0.1. This suggests that only 10 percent of the variation in absenteeism days can be explained by observable healthcare variables.

Finally, we use sick leaves due to cancer as a placebo check. Cancer cannot possibly be affected by outside options, especially in the short-and medium-run, unless there are secular trend differentials between eligible and ineligible workers we fail to take into account. In Figure A.7, we therefore run our event study on the probability of having a sick leave due to cancer in a given half-year of age. This gives a robust zero effect, suggesting no effect of the change in outside options on getting cancer.

III.7. Alternative identification strategy

So far we have relied solely on changes in the potential UI benefit length to identify the value of outside options. This strategy has the big advantage that it allows us to study a large population of workers that is potentially affected. As a validation exercise, we now consider an entirely different source of variation, namely exogeneous changes in the Austrian early retirement age (ERA). To this end, we exploit two reforms in Austria that gradually increased the ERA from 60 to 65 for men and from 55 to 60 for women based on birth cohorts. In Appendix C, we provide an in-depth discussion of these two reforms and the relevant institutional setting. Our model assumes an infinite time

[10] We define common cold and low back pain as potential shirking diagnoses because their symptoms are not immediately visible to a doctor and hence difficult to verify. In this definition of shirking diseases we follow previous research that finds that employers are likely to connect such sick leaves to shirking (Ahammer 2018). In our sample approximately 4 percent of sick leaves are due to a shirking diagnosis. We define good weather days as having conditions that are at least 0.5 standard deviations better than the monthly regional average. From April to September we use the average daily temperature and hours of sunshine while from October to March we use the amount of fresh snow (hinting at the fact that skiing could be possible). In our sample approximately 15 percent of sick leave spells start with a good weather day.

horizon, thus it does not capture a fixed retirement date. However, retirement could be incorporated into our model by having the discount factor decrease over time and thereby resemble a situation in which retirement becomes increasingly likely once a certain age is reached. Alternatively, we could assume that there is a predetermined last period and extend the model to still allow for positive effort. For example, as in Fahn (2023), the agent's preferences might also contain history-dependent social preferences which disappear after the principal reneged on a bonus payments. In all these cases, the total expected relationship surplus decreases as time passes. Therefore, a change in the ERA has the opposite effect of a change in UI benefits. If it becomes more difficult for workers to retire early, the expected relationship surplus with their current employer increases. Therefore, we would predict that effort increases and consequently absenteeism *decreases* with a higher ERA.

Prediction 2. *An increase in the ERA is associated with higher equilibrium effort.*

To test this prediction, we restrict our sample to workers that experienced a change in the ERA, namely those born between 1940 and 1957, and estimate a simple fixed effects model,

$$\text{duration}_{jiy} = \alpha \cdot \text{ERA}_{iy} + \theta_i + X'\delta + \varepsilon_{jiy}, \tag{7}$$

where the duration of sick leave j of individual i in year y is regressed on the statutory ERA_{iy} in the year the sick leave is taken and worker fixed effects θ_i. Additionally, we control for flexible age and quarter-year fixed effects, which are summarized in X. Importantly, because we include worker fixed effects in model (7), the parameter α is identified only through changes in the ERA between different sick leaves of a worker. The ERA is statutory and thus not a choice variable, which mitigates endogeneity concerns.

We report these estimates in Table 1 for different sets of control variables. In column (1), we estimate model (7) with OLS, only controlling for quarter-year fixed effects that account for seasonal trends in sick leave taking. This gives a small negative coefficient, suggesting that the ERA is indeed negatively related to average sick leave durations. In column (2), we add age fixed effects, which hardly changes our estimate. In column (3), we add worker fixed effects, which allows us to

Table 1 — Effects of an increase in the ERA on absenteeism

	OLS		Fixed effects	
	(1)	(2)	(3)	(4)
Statutory ERA	–0.015**	–0.020***	–0.492***	–0.205***
	(0.007)	(0.006)	(0.033)	(0.039)
Worker fixed effects	No	No	Yes	Yes
Age fixed effects	No	Yes	No	Yes
Quarter fixed effects	Yes	Yes	Yes	Yes

Notes: This table reports OLS and fixed effects estimates for the effect of a 1-year increase in the statutory ERA on sick leave duration. The sample is based on all sick leaves in Upper Austria taken by workers born between 1940 and 1957. The average sick leave duration in this sample is 12.2 days, the number of observations in all cells is 1,714,371. Stars indicate significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

exploit changes in ERA between sick leaves of a single worker. The coefficient is now much larger in magnitude. Our preferred specification is in column (4), where we control for both worker and age fixed effects. The estimate suggests that, consistent with our theoretical predictions, a one-year increase in the statutory ERA decreases the average duration of sick leaves by around half a day, which is significant at the 1 percent level. We conclude that, even when using a different source of identifying variation compared to our main empirical design, we find evidence that changes in outside options do matter for worker incentives.

IV. Mechanisms

Having established the general negative link between outside options and worker motivation, we now derive additional results that lend support to the underlying mechanism which we argue is caused by a relational contract between firms and their employees.

IV.1. Threat of unemployment

Being unemployed does not affect all workers' outside options to the same extent. Some will immediately find an alternative job, for others this is more difficult and unemployment potentially more costly. Table A.4 compares the risk of becoming unemployed and, conditional on being unemployed, the average unemployment duration for blue-collar workers with white-collar workers in our data. Blue-collar workers face more than twice the risk of becoming unemployed as white-collar workers and, if they become unemployed, the average duration until they find a job is twice as long. Both differences are statistically significant. Therefore, a potential extension of UI benefits is more relevant for blue-collar workers. This implies that we expect the resulting reduction in equilibrium effort to be more pronounced for them.

Prediction 3. *The effort reduction caused by an increase in UI benefits is more pronounced for blue-collar workers who are more likely to face extended phases of unemployment.*

Prediction 3 not only follows from a larger weight UI benefits have for $\bar{u}$, but also from the fact that a higher likelihood to actually become unemploymed is equivalent to a smaller discount factor δ. Since a smaller δ implies a lower surplus (Lemma 2), Proposition 1 contributes as well, which states that the effort reduction caused by higher (future) outside options is more pronounced if the relationship surplus has initially been small. This link is further applied in the subsequent predictions.

To test Prediction 3, we run our main regression separately for blue-collar and white-collar workers and report estimated effects in Table 2, panel (a). Both groups react significantly to the UI extension by increasing absenteeism. While white-collar workers increase sick leave takeup by 0.16 days per half-year of age, on average, the estimted effect for blue-collar workers is much larger at 0.68 days per half-year of age. The difference between the two groups is statistically significant at the 1 percent level.

IV.2. Size of Relationship Surplus

Our next predictions are based on Proposition 1, i.e., that the negative effect of higher UI benefits on effort is more pronounced if the relationship value is smaller, in combination with factors we suggest shape this relationship surplus. There, we first pick up the discussion underlying Prediction 2, where we argue that the relationship surplus decreases as time passes. This interaction also has implications for the negative consequences of higher UI benefits on effort, and indicates

Prediction 4. *The effort reduction caused by an increase in UI benefits increases over time.*

Table 2 — Mechanisms

		(a) By occupation	
	Baseline	Blue collar	White collar
	(1)	(2)	(3)
Average effect	0.496***	0.680***	0.156***
	(0.026)	(0.033)	(0.042)
t-test for effect homogeneity		−9.77 (p = 0.000)	
Avg. sick leave duration in $t < b$		7.65	6.65
Number of observations	4,648,387	2,705,903	1,942,484

		(b) By gender	
	Baseline	Women	Men
	(1)	(2)	(3)
Average effect	0.496***	0.711***	0.254***
	(0.026)	(0.037)	(0.040)
t-test for effect homogeneity		−8.38 (p = 0.000)	
Avg. sick leave duration in $t < b$		7.06	7.34
Number of observations	4,648,387	1,956,733	2,691,654

		(c) By firm growth	
	Baseline	Firm shrinks	Firm grows
	(1)	(2)	(3)
Average effect	0.486***	0.541***	0.399***
	(0.027)	(0.036)	(0.037)
t-test for effect homogeneity		−2.90 (p = 0.004)	
Avg. sick leave duration in $t < b$		7.08	7.22
Number of observations	4,308,755	2,453,082	1,855,673

		(d) By AKM fixed effect	
	Baseline	Low wage firm	High wage firm
	(1)	(2)	(3)
Average effect	0.500***	0.609***	0.454***
	(0.026)	(0.036)	(0.039)
t-test for effect homogeneity		−2.93 (p = 0.003)	
Avg. sick leave duration in $t < b$		7.31	7.13
Number of observations	4,640,660	2,221,688	2,418,972

Notes: This table reports average effect estimates from equation (5) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on sick leave duration for different subgroups. In panel (a), we split the sample by occupational collar. In panel (b), we split by worker gender. In panel (c), we split the sample by whether the worker's firm has been shrinking or growing compared to the previous year. In panel (d), we split at the sample median of the estimated AKM firm fixed effect distribution. The heterogeneity variables could not be computed for some workers, hence we report the baseline for all observations with non-missing heterogeneity variables in column (1). All regressions control for tenure and year fixed effects, gender, and a blue collar dummy. The t-test indicates whether estimates from column (2) and (3) are statistically different—it comes from a separate model where we fully interact the average effect and all covariates with the heterogeneity split variable. Standard errors are clustered on the worker level. Stars indicate significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Prediction 4 is supported by Figure 2, where we can see that the difference in absenteeism between treatment and control group increases over time.

Next, the size of δ in a specific employment relationship is affected by its future prospects. If these are worse, the likelihood that any match may be terminated is larger. The reason is that either the firm's bankruptcy risk or the chances that it has to lay off further employees are higher, and both imply a smaller continuation probability of a given match. As an indicator for the perceived stability of an employment relationship, we assess whether the firm's workforce increases or decreases in a given year.

Prediction 5. *The effort reduction caused by an increase in UI benefits is larger in firms with a shrinking than with a growing workforce.*

We test Prediction 5 in Table 2, panel (b), where we report effects by whether a worker's firm in a given year is smaller or larger than in the year prior. Indeed, we find that workers in shrinking firms react stronger, and the difference in effects between shrinking and growing firms is statistically significant at the 1 percent level.

Now, we take into account that the efficiency of an employment relationship is not only determined by its future surplus, but also by the firm's or worker's inherent productivity, which in our model corresponds to the value θ. If θ is larger, the same effort generates a higher output. Therefore, if (EC) binds and effectively constrains equilibrium effort, a higher θ increases the (future) surplus and consequently reduces the negative consequences of a higher outside option on equilibrium effort. To explore such a link empirically, we follow the literature and use AKM firm-wage fixed effects as a proxy for a firm's inherent productivity.[11] This yields

Prediction 6. *The effort reduction caused by an increase in UI benefits is more pronounced in low-wage firms.*

Note that θ also determines first-best effort, with $de^{FB}/d\theta > 0$. To generate Prediction 6 for all values of θ, we therefore must additionally demonstrate that a higher θ makes it "more likely"

[11] We estimate the AKM model on a full panel of all Austrian workers between 1998–2021 with wage information in a given year. This is different from the data we use for our other analyses. The reason is that we have information on sick leave taking only for Upper Austria, and these data end in 2018.

that de^{FB} can be implemented. Indeed, in Appendix B we show that this holds for commonly used effort cost functions.

In Table 2, panel (c), we estimate effects by whether the AKM firm fixed effect is above or below the sample median. We find that workers in low-wage firms react more strongly to the UI benefit extension, and the difference between workers in low-wage and high-wage firms is significant at the 1 percent level.

Furthermore, the unemployment rate is an important dimension of a worker's outside option. Several studies have confirmed a positive correlation between unemployment rate and worker effort (Lazear et al. 2016), a link that can also be found in our data (Figure A.4). Therefore, the (local- or sector-specific) unemployment rate may affect by how much effort goes down in response to the UI benefit extension. However, this relationship is ambiguous because the lower chances of finding a new job when facing higher unemployment induce two countervailing forces. On the one hand, the relationship value goes up, which would imply a weaker treatment effect. On the other hand, UI benefits assume a more prominent role in a worker's outside option (as we have argued for blue-collar workers), which would imply that the treatment effect is larger. Without making strong assumptions on the functional forms of the components of our model, we are not able to state that one effect dominates the other.

Prediction 7. *The effort reduction caused by an increase in UI benefits may be more or less pronounced for workers facing a higher unemployment rate.*

In Appendix Figure A.8, we provide estimates by quartiles of the sectoral unemployment rate in a region, separately for blue-collar and white-collar workers. It appears that effects are generally stronger when the unemployment rate is very low, but we can also not reject relatively large effects when the unemployment rate is very high, especially for blue-collar workers.

IV.3. Care obligations

In this section, we argue that the link between outside options and worker incentives is intensified for workers who have care obligations for their children or elderly parents. There, the most obvious channel is that the replacement rate provided by UI benefits is substantially higher for workers with dependents for whom it can assume values of up to 80 percent of one's daily net income, compared to the minimum replacement rate of 55 percent. Therefore, the increase in outside options for eligible workers with care obligations, and consequently the expected treatment effect, is larger than for eligible workers without. On top, care obligations may increase the opportunity costs of going to work versus taking/extending sick leave. To formalize this approach, assume that effort costs are $c(e, k)$, with $c_k, c_e, c_{ee}, c_{kk} > 0$ and $c_{ek} \geq 0$, and that being responsible for more care activities corresponds to a higher k. Then, we argue in Appendix B that a higher k can indeed magnify the negative effect of a higher outside option on effort. Note that we do not have evidence on individual time spent for care activities, however there is an abundance of evidence that in Austria women are responsible for most of family care work (e.g., Danzer et al. 2022). Therefore, we would argue that k is larger for women than for men, especially for those who have children. All this yields the following prediction.

Prediction 8. *The increase in absenteeism caused by higher UI benefits is more pronounced for women and workers with children.*

We test Prediction 8 in Table 2, panel (b). Women react much stronger to the change in outside options than men, and the difference in point estimates is significant at the 1 percent level. In Appendix Table A.6, we additionally split the sample by whether workers have children or not.[12] We find that effects are generally stronger for workers with children. For women, the difference in point estimates amounts to around 0.17 days, which is significant at the 5.4 percent level.

IV.4. Multilateral Relational Contracts

Our theoretical analysis is based on an individual arrangement between one firm and one worker. However, it might also be possible that the firm forms a *multilateral* relational contract with

[12]For men, our information on children is incomplete. We observe births for fathers only until 2007 and if the child was born in wedlock. We augment this information with data from childcare subsidies, which allow us to identify fathers that are not in the birth register but claim these subsidies. Nevertheless, some men we categorize as not having children, may, in fact, have children. For women, births are recorded in the ASSD employer-employee panel, hence we have complete information on children regardless of the year of birth or marital status.

its workforce (Levin 2002). There, reneging on one agent causes a breakdown of all involved relationships. If agents are heterogeneous and, for example, differ in their outside options, such multilateral relational contracts can increase the total value of a firm's employment relationships. The reason is that multilateral arrangements allow for the pooling of all future surpluses and using them most effectively by equalizing marginal effort costs. Then, however, a change in the outside option of one worker affects equilibrium effort of all involved workers, and we would not be able to detect a treatment effect if we compared eligible and non-eligible workers who are part of the same agreement. However, multilateral relational contracts are not always feasible. For example, they rely on the mutual observability of the state of workers' relationships. The following prediction therefore takes into account that some firms may use of them.

Prediction 9. *The effort reduction caused by an increase in UI benefits is less pronounced if we consider eligible and non-eligible workers from the same firm.*

To test this prediction, we reestimate our model with firm fixed effects, which uses only within-firm variation in outside options for identification. The results in Table A.5 indicate that our estimates indeed are smaller when using firm fixed effects.

V. Alternative models of the labor market

Our mechanism relies on two building blocks. First, workers' productivity depends on their costly effort, thus they need to be motivated accordingly. Second, formal, court-enforceable, contracts are not feasible for that purpose and relational contracts based on the future relationship value used instead. These features are sufficient for generating our predictions, and we do not need to rely on one specific model setup (see section D of the Web Appendix). Now, we argue that this mechanism also does a better job explaining our observations compared to other models of the labor market. For this comparison, we focus on the most widely used alternatives, the competitive model as well as search search-and-matching models.

V.1. Competitive model of the labor market

We argue that a standard, competitive model of the labor market cannot generate our predictions. To do so, we first identify a potential link between UI benefits and worker absenteeism. For example, higher UI benefits might reduce stress levels on the job, in which case we would expect affected workers' sick days to go down instead of up. Alternatively, to integrate our interpretation that worker absenteeism relates to productive effort into a competitive labor market model, we could assume that firms are able to use formal contracts to motivate effort. With formal contracts (and regardless of the allocation of rents), first-best effort characterized by $\theta - c'(e^{FB}) = 0$ would always be implemented. Thus, only if a worker's outside option had an effect on his inherent productivity θ, the increase in UI benefits could influence effort. If there was such a link, we would rather expect a more generous safety net to increase inherent productivity instead of reducing it. Moreover, even if for some reason inherent productivity was reduced, it would be difficult to generate predictions 4, 5, 6, and 9.

V.2. Search models

Next, we discuss a model in which labor markets are characterized by search-and-matching frictions. There, we focus on on-the-job search as a means that can potentially affect worker absenteeism.[13] Consider an agent who spends some of his working time searching for alternative jobs and where more search increases absenteeism. Then, higher UI benefits can affect his incentives to conduct search if there is some probability that he will lose his job. If this probability is independent of the extent of on-the-job search, one reason to conduct search would be to have an offer available in case of being fired. But then, one would expect higher UI benefits to reduce on-the-job search and consequently absenteeism which would contradict our predictions. Therefore, let us suppose that more search increases the agent's chances of being fired. In this case, higher UI benefits

[13]Naturally, more generous UI benefits could also reduce incentives to search for those without a job and consequently increase the unemployment rate, thereby influencing incentives for the employed. However, we do not expect such a link to matter in our setting because the policy we utilize is not a labor market reform but instead an institutional feature that affects some employees differently than others.

would indeed cause absenteeism to go up. We argue, however, that this link generates a mechanism that is very similar to the one captured by our model: Higher search increases absenteeism and thus decreases productive effort. Thereby, the agent captures private benefits (as with an effort reduction), and the principal's payoff goes down. Furthermore, if higher absenteeism indeed was caused by more on-the-job search instead of other consequences of a reduced motivation, this would result in better job offers and consequently more job-to-job transitions (presuming that at least some of these outside offers are not matched by the current employer). In Appendix Figure A.9, we therefore test whether being eligible for more generous UI benefits increases job-to-job transitions. We find no evidence that this is the case. If anything, job-to-job transitions decrease slightly, but this effect is not significant at the 5-percent level.

VI. Conclusion

This paper has demonstrated that better outside options can decrease workers' incentives to exert effort. We have presented causal evidende for this link and on top found support for a proposed mechanism based on the interaction of outside options with the relational contracts firms have with their workers. To put the effect size of an increase in UI benefit duration into perspective, we can compare it to Lusher et al. (2022). In our case a 9-week increase in UI benefit duration leads to a 7 percent decrease in effort. Lusher et al. (2022) find a 1 percent decrease in effort (2 percent decrease for 18-week increase in UI benefit duration). We suggest that these differences arise due to the different settings that are studied. Whereas our effort measure, sick days, captures many dimensions of a worker's motivation, Lusher et al. (2022) use one specific aspect, high-frequency scanner data of carried out transactions. Moreover, there is evidence that more generous UI benefits increase unemployment (Schmieder & von Wachter 2016). Jessen et al. (2023) establish this link for Poland, Hartung et al. (2022) find that less generous UI benefits in Germany lowered unemployment. But higher unemployment may increase equilibrium effort (see our discussion in Section IV) which then would counteract the negative effect of more generous UI benefits. This might further explain an apparently smaller effect size in Lusher et al. (2022) who study state-wide extensions of UI benefits.[14] In our setting where benefit extensions happen on an individual level once a certain age cutoff is reached, feedback effects seem much less likely. In any case, Lusher et al. (2022) as well as the present study indicate that policies such as UI benefits can have consequences that go beyond the well-studied effects of improving the welfare of the unemployed or reducing their incentives to search for and take up a new job. Instead, such policies also have an indirect effect on employed workers, by changing their outside options and consequently their incentives to exert effort. This link can inform firms when responding to policies that are actually aimed at the unemployed. For example, we have demonstrated that a negative impact of higher UI benefits on equilibrium effort is less pronounced for high-productivity and stable relationships. Therefore, investments into firm-specific human capital not only increase productivity directly, they also have an indirect positive equilibrium effect by relaxing the enforceability constraint on effort and mitigating potential negative consequences of better outside options. Follow-up work may consider recent developments such as the technological process that will affect relational contracts and thereby the role of outside options. For example, the monitoring of employees' activities could improve and make more dimensions of their effort verifiable. In any case, labor market studies should not neglect workers' incentives to exert effort, especially in times in which firms are struggling with phenomena like quiet quitting, i.e. employees only do what they are contractually obliged to. Then, it is particularly important to consider the role of relational contracts that can incentivize workers beyond the levels specified by their formal employment contracts.

[14]Note that, although Lusher et al. (2022) argue that such endogeneity is not present in their setting, they focus on the opposite causal direction that higher unemployment may cause higher UI benefits. Their finding that UI benefits are not correlated with *lagged* unemployment rates therefore does not rule out that it is the other way around.

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Web appendix

This web appendix contains additional tables and figures for the paper *"Outside options and worker motivation"* by Alexander Ahammer, Matthias Fahn, and Flora Stiftinger.

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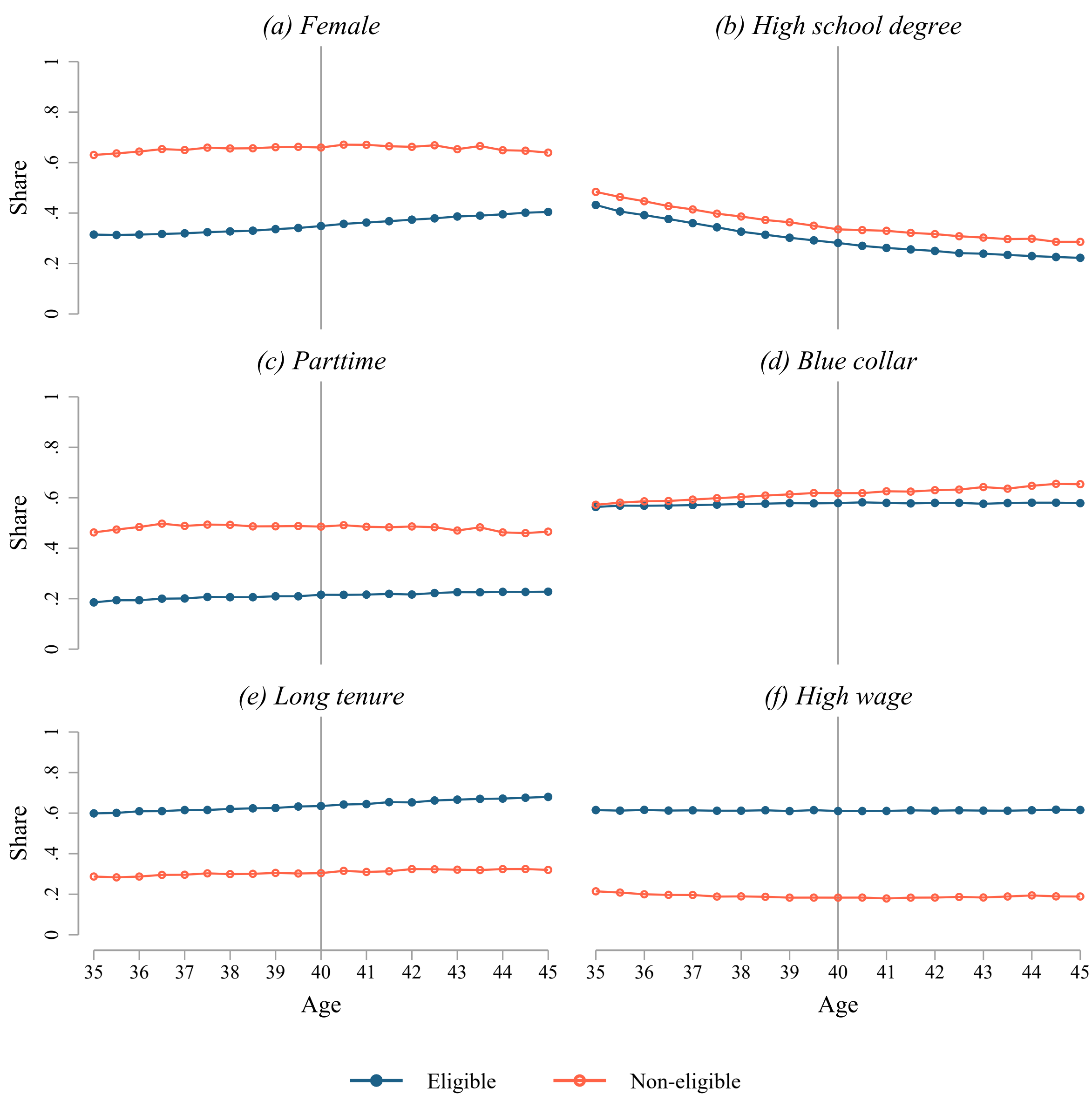
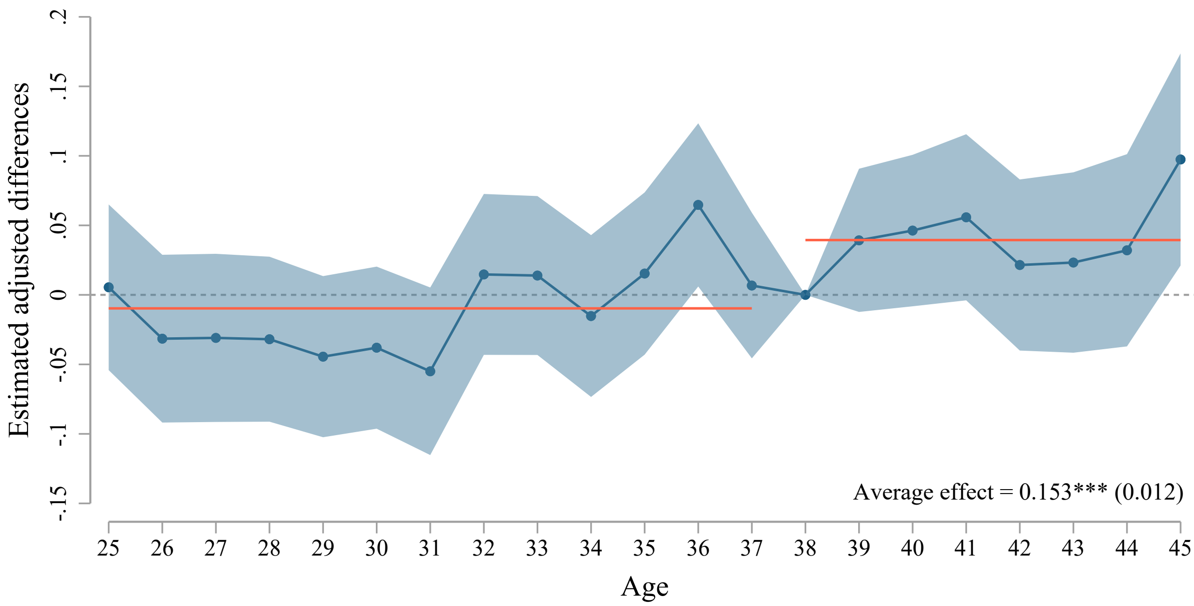
A. Additional tables and figures

Figure A.1 — Composition around the age-40 UI benefit extension cutoff



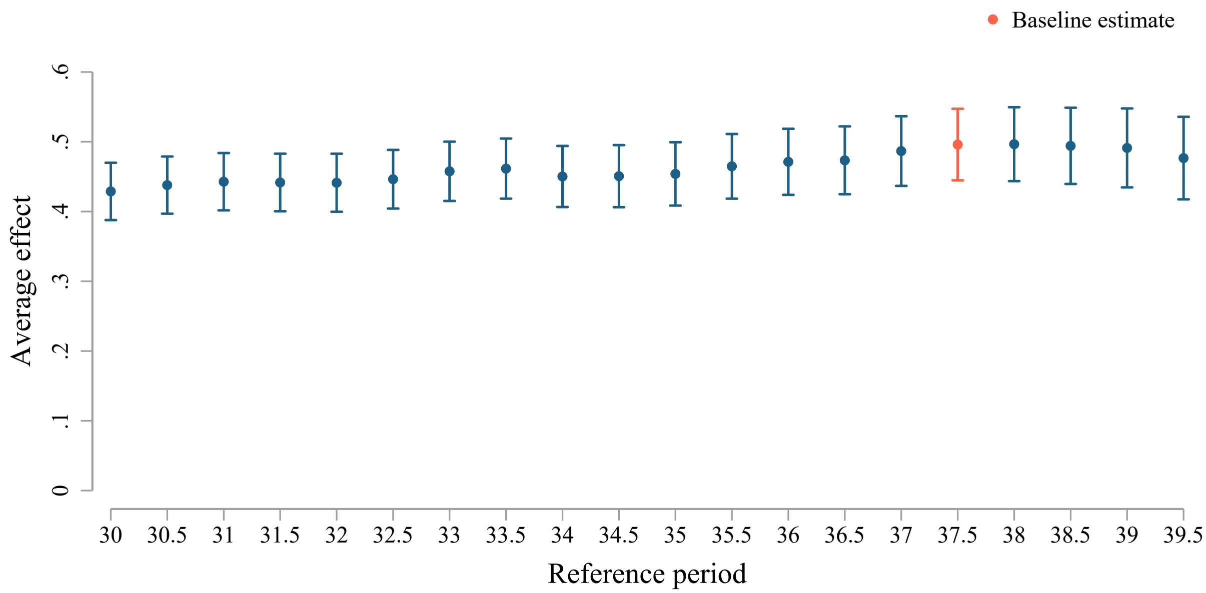
Notes: This figure reports the share of female workers (panel a), workers with a highschool degree (*'Matura,'* panel b), parttime workers (panel c), blue-collar workers (panel d), workers with above-median tenure (the median is 1.96 years, panel e), and workers with above-median wage (€ 22,902, panel f) for eligible and ineligible workers at a given age.

Figure A.2 — The effect of an increase in outside options on log wages



Notes: This figure provides event study estimates similar to those in equation (6) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on log annual wages. Because we expect eligible workers to react already prior to age 40, we fix the reference period to $b = 37.5$, and point estimates can be interpreted as log changes in wages due to the benefit extension at a given age relative to age 37.5. The shaded area represents a 95 percent confidence band. The red horizontal lines indicate averages of age-specific estimates for both the pretreatment period $t < b$ and the posttreatment period $t \geq b$. The average effect estimate is from a model similar to equation (5). All regressions control for tenure and year fixed effects, gender, and a blue collar dummy.

Figure A.3 — Robustness to different reference period choices

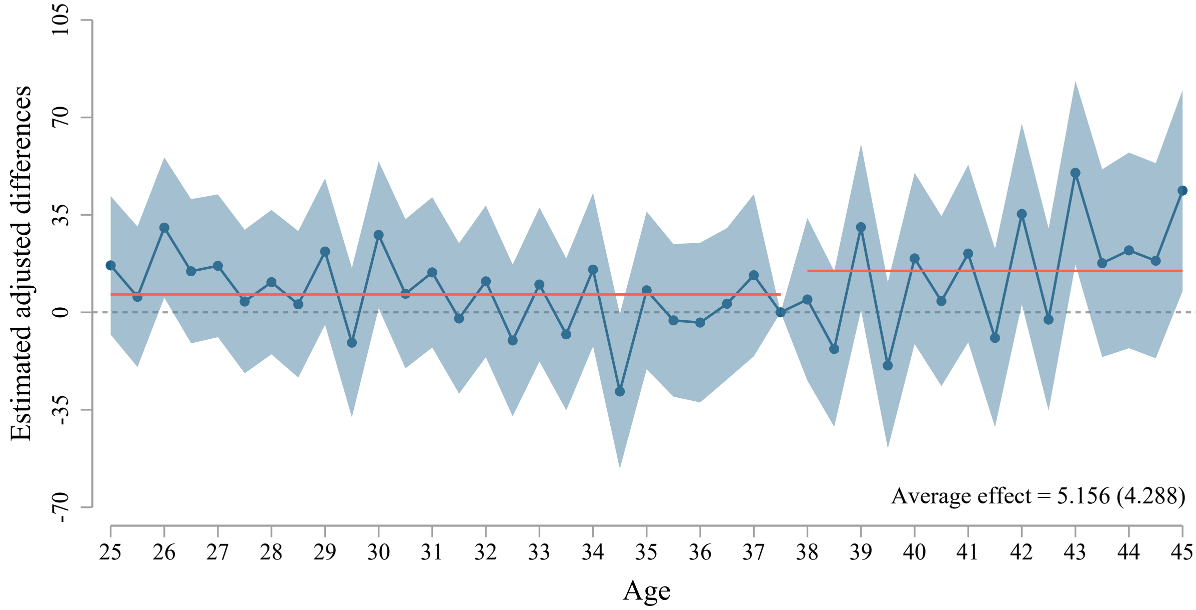


Notes: This figure shows estimates from equation (5) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers for different reference periods b. All regressions control for tenure and year fixed effects, gender, and a blue collar dummy.

FIGURE A.4 — Relationship between sick leave duration and the unemployment rate

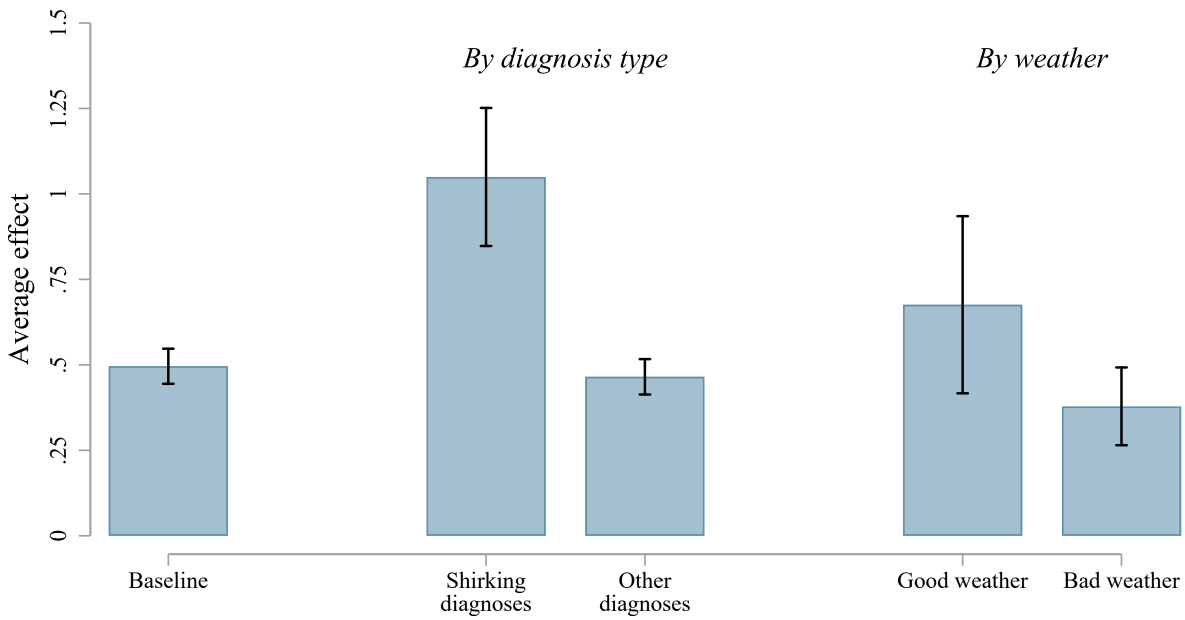
Notes: This graph depicts the relationship between sick leave duration and the sectoral unemployment rate in a region. The unemployment rate is calculated for every NACE95 2-digit sector and NUTS 3 combination. Both the scatters and the regression line are weighted by the number of workers at each point.

FIGURE A.5 — The effect of an increase in outside options on total healthcare expenses



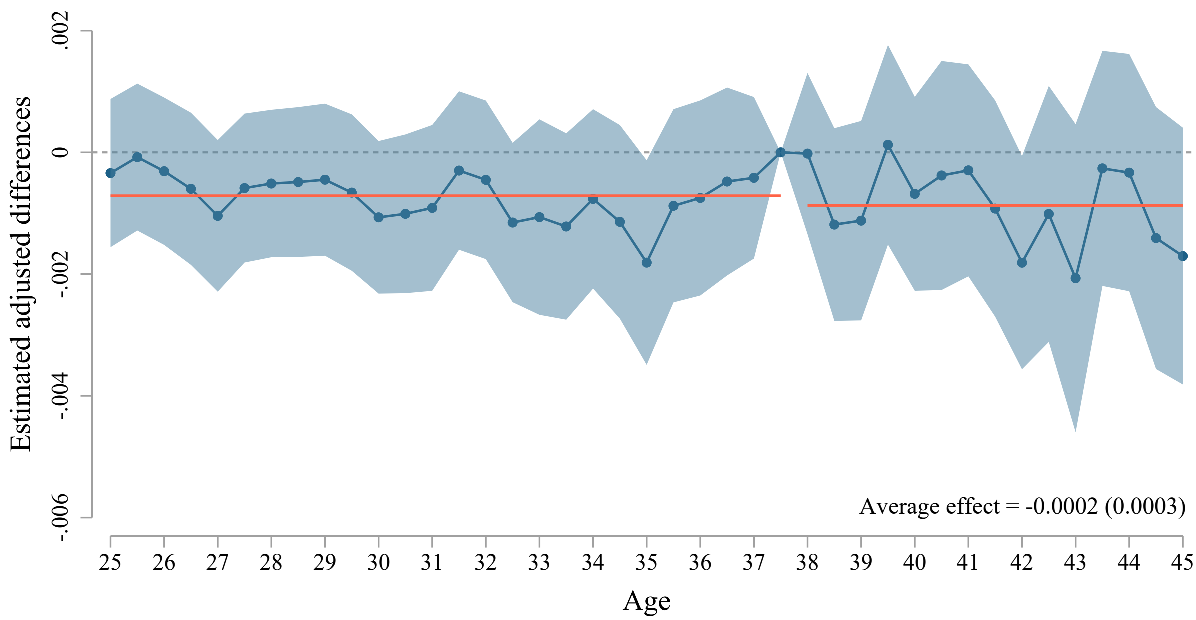
Notes: This figure provides event study estimates similar to those in equation (6) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on total healthcare expenses, which is the sum of physician fees, drug expenses, and hospital expenses. Because we expect eligible workers to react already prior to age 40, we fix the reference period to $b = 37.5$, and point estimates can be interpreted as changes in average healthcare expenses due to the benefit extension at a given age relative to age 37.5. The shaded area represents a 95 percent confidence band. The red horizontal lines indicate averages of age-specific estimates for both the pretreatment period $t < b$ and the posttreatment period $t \geq b$. The average effect estimate is from a model similar to equation (5). All regressions control for tenure and year fixed effects, gender, and a blue collar dummy.

FIGURE A.6 — Heterogeneity by diagnosis type and weather



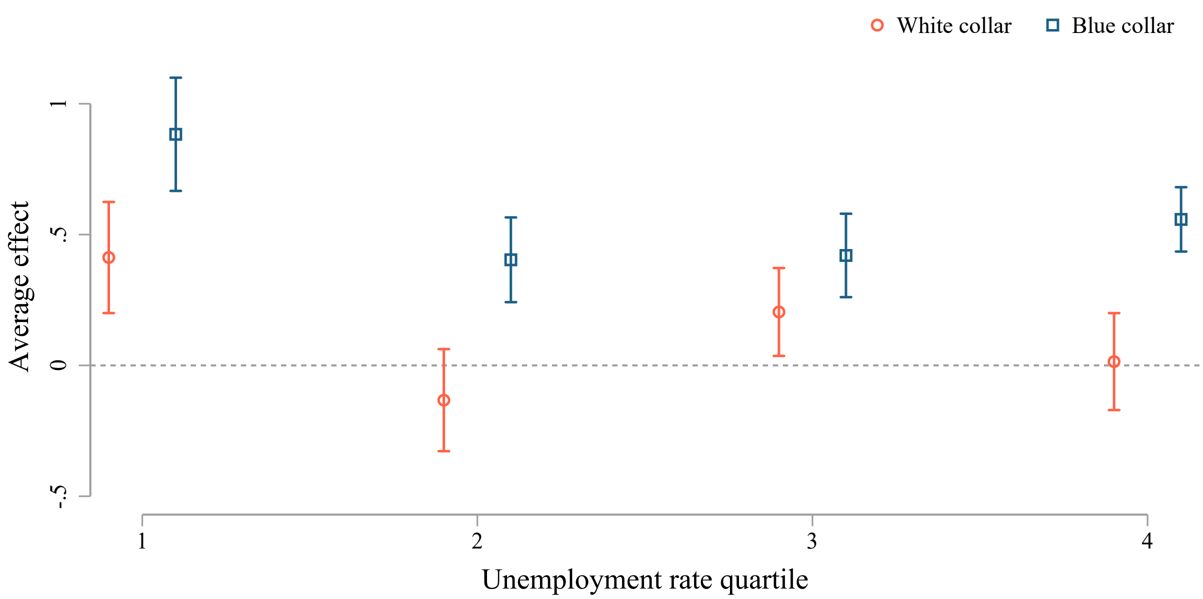
Notes: This figure plots estimates from equation (5) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on sick leave duration by diagnosis type and weather. The error bars indicate 95 percent confidence intervals. Shirking diagnoses are defined as common cold and low back pain. Good weather is defined differently for Summer and Winter. In Summer (April–September), we define good weather as the temperature on the first day of the sick leave being half of a standard deviation higher and the sunshine duration on the first day of the sick leave being half a standard deviation longer than the monthly average in a zip code. During Winter (October–March), good weather is defined as fresh snow on the first day of the sick leave being half a standard deviation higher than the monthly average in a zip code. All regressions control for tenure and year fixed effects, gender, and a blue collar dummy.

FIGURE A.7 — The effect of an increase in outside options on sick leaves due to cancer



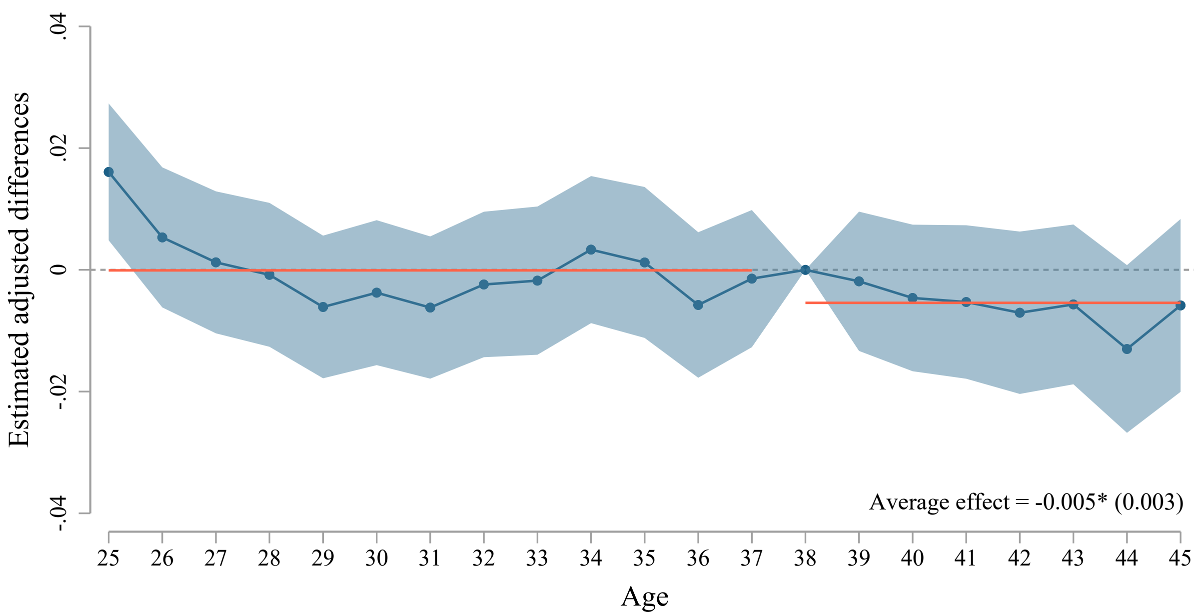
Notes: This figure plots event study estimates from equation (6) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on the probability of having a sick leave due to cancer in a certain half-year of age. We fix the reference period to $b = 37.5$, and point estimates can be interpreted as changes in average sick leave duration due to the benefit extension at a given age relative to age 37.5. The shaded area represents a 95 percent confidence band. The red horizontal lines indicate averages of age-specific estimates for both the pretreatment period $t < b$ and the posttreatment period $t \geq b$. The average effect estimate is similar to that from equation (5). All regressions control for tenure and year fixed effects, gender, and a blue collar dummy.

Figure A.8 — The effect of an increase in outside options on sick leaves by quartiles of the sectoral unemployment rate in a region, separately for blue collar and white collar workers



Notes: This figure plots estimates from equation (5) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on sick leave duration by quartiles of the sectoral unemployment rate in a region and occupation. The unemployment rate is calculated for every NACE95 2-digit sector and NUTS 3 combination. The error bars indicate 95 percent confidence intervals. All regressions control for tenure and year fixed effects, gender, and a blue collar dummy.

Figure A.9 — The effect of an increase in outside options on job-to-job transitions



Notes: This figure plots event study estimates from equation (6) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on the probability of changing jobs in a certain half-year of age. We fix the reference period to $b = 37.5$, and point estimates can be interpreted as changes in average sick leave duration due to the benefit extension at a given age relative to age 37.5. The shaded area represents a 95 percent confidence band. The red horizontal lines indicate averages of age-specific estimates for both the pretreatment period $t < b$ and the posttreatment period $t \geq b$. The average effect estimate is similar to that from equation (5). All regressions control for tenure and year fixed effects, gender, and a blue collar dummy.

Table A.1 — Summary statistics

	Mean	Std. dev.	By eligibility status: Eligible	Ineligible	Difference
	(1)	(2)	(3)	(4)	(5)
(a) Outcome					
Sick leave duration (days)	7.43	9.12	7.65	6.93	–0.72***
(b) Treatment assignment information					
Experience (years)	10.17	4.87	12.29	5.17	–7.12***
(c) Socioeconomic and job information					
Female	0.41	0.49	0.34	0.58	0.24***
High school degree†	0.56	0.50	0.54	0.62	0.08***
Parttime worker†	0.22	0.42	0.16	0.37	0.21***
Blue-collar worker	0.55	0.50	0.57	0.51	–0.07***
Tenure (years)	3.54	3.24	4.21	2.00	–2.20***
Annual wage (€ 1,000)	23.51	12.41	25.90	18.08	–7.82***

Notes: This table provides summary statistics for the overall sample (means in column 1 and standard deviations in column 2) and by whether the worker is eligible for the UI benefit extension at age 40 (columns 3 and 4). Column (5) gives the difference between columns (3) and (4), with the stars indicating p-values from a two-sample t-test with significance levels * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$. The number of observations is 4,648,387.

† Parttime status is missing for 44.02 percent of observations and education is missing for 0.8 percent of observations. Means and standard deviations are calculated based on all observations with non-missing values.

Table A.2 — Robustness to different regression specifications

	OLS			Fixed effects	
	(1)	(2)	(3)	(4)	(5)
Average effect	0.437***	0.496***	0.536***	0.495***	0.421***
	(0.027)	(0.026)	(0.026)	(0.038)	(0.038)
Covariates					
Blue-collar worker		0.991***	0.702***		–0.005
		(0.013)	(0.014)		(0.030)
Female		0.271***	0.000		
		(0.013)	(0.015)		
Parttime worker			0.124***		
			(0.018)		
High school degree			–0.468***		
			(0.014)		
Annual wage (€ 1,000)			–0.023***		
			(0.001)		
Tenure and year fixed effects	No	Yes	Yes	No	Yes
Worker fixed effects	No	No	No	Yes	Yes

Notes: This table reports average effect estimates from equation (5) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on sick leave duration using different regression specifications. In column (1), we run model (5) using OLS without any covariates. In column (2), we control for a blue collar dummy, a gender dummy, as well as tenure and year fixed effects. In column (3), we additionally control for parttime work, education, and wage. In column (4), we estimate model (5) with worker fixed effects but no other covariates. In column (5), we add a blue collar dummy as well as tenure and year fixed effects to the fixed effects model from column (4). Whenever we control for education and parttime work, we replace missing values with zero and add a missing indicator dummy to the regression. The average sick leave duration in the pretreatment period $t < b$ is 7.22, the number of observations in each cell is 4,648,387. Stars indicate significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.3 — Robustness to different eligibility status constraints

	Baseline	Eligibility constraints: Fixed at age 37.5	Eligibility constraints: Drop switchers
	(1)	(2)	(3)
Average effect	0.496***	0.542***	0.791***
	(0.026)	(0.032)	(0.035)
Tenure fixed effects	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes
Gender and occupation	Yes	Yes	Yes
Average sick leave duration in $t < b$	7.22	7.22	7.33
Number of observations	4,648,387	4,228,846	3,138,720

Notes: This table reports average effect estimates from equation (5) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on sick leave duration using different eligibility constraints. In column (2), we consider all workers (in)eligible that are (in)eligible at age 37.5, regardless of whether they change eligibility status later. In column (3), we drop workers from the sample that switch eligibility status at some point. All regressions control for tenure and year fixed effects, gender, and a blue collar dummy. Stars indicate significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.4 — Probability of becoming unemployed by occupation

	Overall (1)	White collar (2)	Blue collar (3)	Difference (4)
Probability of becoming unemployed	0.18	0.10	0.24	−0.14***
	(0.39)	(0.30)	(0.42)	
Unemp. duration | being unemployed	15.39	8.94	19.60	−10.66***
	(42.27)	(33.89)	(46.44)	

Notes: This table reports the probability of becoming unemployed and the average duration of being unemployed, conditional on being unemployed, by occupational collar for workers aged 40 or older in our data. We do not consider workers that are recalled to the same firm within 9 months as unemployed. The value in column (4) is the difference between columns (2) and (3) and the stars indicate whether the difference is statistically significantly different from zero based on a two-sample t-test with significance levels * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

TABLE A.5 — Change in estimates when using firm fixed effects

	Baseline (1)	+ firm FEs (2)	+ worker FEs (3)
Average effect	0.496***	0.433***	0.347***
	(0.026)	(0.037)	(0.036)
Worker fixed effects	No	No	Yes
Firm fixed effects	No	Yes	Yes
Tenure fixed effects	Yes	Yes	Yes
Year fixed effects	Yes	Yes	Yes
Gender and occupation	Yes	No	No

Notes: This table reports average effect estimates from equation (5) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on sick leave duration when using firm fixed effects in our specification. In column (1), we report our baseline specification. In column (2) add firm fixed effects, considering only within-firm variation in sick leaves. In column (3), we add both worker and firm fixed effects. The average sick leave duration in the pretreatment period $t < b$ is 7.22, the number of observations in each column is 4,648,387. Stars indicate significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

Table A.6 — Effects by whether workers have children

	No children (1)	Children (2)
(a) Women		
Average effect	0.426***	0.597***
	(0.078)	(0.044)
t-test for effect homogeneity	1.93 (p = 0.054)	
Avg. sick leave duration in $t < b$	6.48	7.56
Number of observations	761,194	1,195,539
(b) Men		
Average effect	0.180***	0.338***
	(0.050)	(0.072)
t-test for effect homogeneity	1.80 (p = 0.072)	
Avg. sick leave duration in $t < b$	7.04	8.04
Number of observations	1,669,912	1,021,742

Notes: This table reports average effect estimates from equation (5) for the differential impact of the UI benefit extension at age 40 between eligible and ineligible workers on sick leave duration for workers with and without children at the time of the sick leave, separately for women (panel a) and men (panel b). For men, we observe children only if they had been married at the time of birth (but only for births until 2007) or if they claim childcare subsidies. For women we observe all births. All regressions control for tenure and year fixed effects, gender, and a blue collar dummy. Stars indicate significance levels: * $p < 0.10$, ** $p < 0.05$, *** $p < 0.01$.

B. Proofs

Proof of Lemma 2. Assume a profit-maximizing equilibrium with effort levels e_t^*, and that θ goes up to θ'. Holding e_t^* constant, this directly increases each per-period surplus $e_t^*\theta - c(e_t^*)$ and thus each S_t. Moreover, all (EC) are relaxed in t and higher effort levels can and will be implemented because a higher θ also increases e^{FB}. Holding equilibrium effort levels constant, a higher δ increases S_t, which further relaxes all (EC) constraints. Finally, a higher $\bar{S}_{t+1}$ has no direct effect on S_t, but tightens (EC) in t and therefore reduces S_t if the constraint binds. ■

Proof of Proposition 1. Assume $e_t^* < e^{FB}$ is characterized by the binding (EC) constraint. Then, the implicit function theorem yields

$$\frac{de_t^*}{d\delta_t(S_{t+1} - \bar{S}_{t+1})} = \frac{1}{c'(e_t^*)} > 0,$$

which implies $de_t^*/d\bar{S}_{t+1} < 0$. Moreover,

$$\frac{d^2 e_t^*}{d\left(\delta_t(S_{t+1} - \bar{S}_{t+1})\right)^2} = -\frac{c''(e_t^*)}{\left(c'(e_t^*)\right)^2}\frac{1}{c'(e_t^*)} < 0.$$

The last statement follows since $e_t^* = e^{FB}$, and e^{FB} is unaffected by outside options. ■

Proof of Proposition 2. The stationarity of effort in all periods $t \geq T-1$ follows from standard arguments (see Levin 2003). Regarding existence of $\tilde{\delta}$, note that $\tilde{e}$ is constrained by $-c(\tilde{e}) + \delta\left[\tilde{e}\theta - \left(\bar{\pi} + \bar{u}^H\right)\right] \geq 0$. For $\left[\tilde{e}\theta - \left(\bar{\pi} + \bar{u}^H\right)\right] > 0$, the left hand side is increasing in δ. For $\delta \longrightarrow 1$, the constraint holds for e^{FB}, for $\delta \longrightarrow 0$, it is violated for e^{FB}.

Now, take period $T-2$. There, effort is constrained by

$$-c(e_{T-2}) + \delta\frac{\tilde{e}\theta - c(\tilde{e})}{1-\delta} \geq \delta\left(\bar{\pi} + \bar{u} + \delta\frac{\bar{\pi} + \bar{u}^H}{1-\delta}\right), \tag{EC}$$

which can be rewritten to

$$-c(e_{T-2}) + \delta\left[\tilde{e}\theta - \left(\bar{\pi} + \bar{u}^H\right)\right] + \delta\left[c(e_{T-2}) - c(\tilde{e}) + \left(\bar{u}^H - \bar{u}\right)(1-\delta)\right] \geq 0. \tag{8}$$

If $\tilde{e} = e^{FB}$, (EC) also holds for $e_{T-2} = e^{FB}$. If $\tilde{e} < e^{FB}$ and $\delta\left[\tilde{e}\theta - \left(\bar{\pi} + \bar{u}^H\right)\right] \geq c(\tilde{e})$, (EC) becomes $\delta\left(\bar{u}^H - \bar{u}\right) \geq \left(c(e_{T-2}) - c(\tilde{e})\right)$, thus $e_{T-2} > \tilde{e}$. Existence of $\tilde{\delta}_{T-2}$ is immediate: At $\tilde{\delta}$, $\bar{u} < \bar{u}^H$ implies that (EC) is slack for $e_{T-2} = e^{FB}$. Continuity and monotonicity in δ deliver the stated properties.

The rest of the proposition follows: The earlier a period, the smaller the weight of $\bar{u}^H$ relative to $\bar{u}$ in the respective (EC) constraint. ■

Formal Discussion of 6. Assume two productivity levels, θ^l and θ^h. If the (EC) constraint binds for both values, the prediction follows from Lemma 2, which states that the surplus increases in θ, as well as Proposition 1. It thus remains to show that (EC) is "more likely to bind" for θ^l than for θ^h. To do that, we compute the critical discount factors above which first-best effort can be implemented and explore whether it is indeed larger for θ^l.

For that, we focus on period $T-1$ after which the relational contract is stationary. Then, first-best effort – characterized by $\theta - c'(e^{FB}) = 0$ – can be implemented if it satifies $-c(e^{FB}) + \delta\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right) \geq 0$, or if

$$\delta \geq \bar{\delta} \equiv \frac{c(e^{FB})}{\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right)}.$$

Moreover,

$$\begin{aligned}
\frac{d\bar{\delta}}{d\theta} =& -\frac{c(e^{FB})e^{FB}}{\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right)^2} + \frac{c'(e^{FB})\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right) - c(e^{FB})\theta}{\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right)^2}\frac{de^{FB}}{d\theta} \\
=& -\frac{c(e^{FB})e^{FB}}{\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right)^2} + \frac{c'(e^{FB})\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right) - c(e^{FB})\theta}{\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right)^2 c''(e^{FB})} \\
=& \frac{c'(e^{FB})\theta e^{FB} - c(e^{FB})\theta - c''(e^{FB})c(e^{FB})e^{FB}}{\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right)^2 c''(e^{FB})} \\
& - \frac{c'(e^{FB})\left(\bar{u}^H + \bar{\pi}\right)}{\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right)^2 c''(e^{FB})} \\
=& \frac{c'(e^{FB})c'(e^{FB})e^{FB} - c(e^{FB})c'(e^{FB}) - c''(e^{FB})c(e^{FB})e^{FB}}{\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right)^2 c''(e^{FB})} \\
& - \frac{c'(e^{FB})\left(\bar{u}^H + \bar{\pi}\right)}{\left(\theta e^{FB} - \bar{u}^H - \bar{\pi}\right)^2 c''(e^{FB})}.
\end{aligned}$$

Since $\bar{u}^H + \bar{\pi} > 0$, for $d\bar{\delta}/d\theta < 0$ to hold it is sufficient that the numerator of the first term is non-positive. For a standard effort cost function $c(e) = \frac{e^n}{n}$, with $n \geq 2$, this term becomes (for any e)

$$\begin{aligned}
& e^{2n-1} - \frac{e^{2n-1}}{n} - (n-1)\frac{e^{2n-1}}{n} \\
=& e^{2n-1}\left(\frac{n-1}{n} - \frac{n-1}{n}\right) = 0.
\end{aligned}$$

■

Formal Discussion of Prediction 8. Assume effort costs are $c(e,k)$, with $c_e, c_k, c_{ee}, c_{kk} > 0$ and $c_{ek} \geq 0$. We want to explore whether the negative effect of a higher $\bar{u}$ is more pronounced if k is larger. We again focus on period $T-1$ after which the relational contract is stationary. Then, the (EC) constraint equals $-c(e,k) + \delta(e\theta - \bar{u} - \bar{\pi}) \geq 0$. First, we assume that (EC) binds, hence

$$\begin{aligned}\frac{de}{dk} &= \frac{c_k}{-c_e + \delta\theta} < 0 \\ \frac{de}{d\bar{u}} &= \frac{\delta}{-c_e + \delta\theta} < 0 \\ \frac{d^2 e}{dk d\bar{u}} &= \frac{c_{ek}\left(-c_e + \delta\theta\right) + c_k c_{ee}}{\left(-c_e + \delta\theta\right)^2} \frac{de}{d\bar{u}}.\end{aligned}$$

This term is negative if $c_{ek}\left(-c_e + \delta\theta\right) + c_k c_{ee}$ is positive, which holds for $c_{ek} = 0$. For $c_{ek} > 0$, assume the standard effort cost function $c(e,k) = k\frac{e^n}{n}$, for which

$$\begin{aligned}&c_{ek}\left(-c_e + \delta\theta\right) + c_k c_{ee} \\ =&e^{n-1}\left(-ke^{n-1} + \delta\theta\right) + k\frac{e^n}{n}(n-1)e^{n-2} \\ =&e^{n-1}\left(\delta\theta - \frac{ke^{n-1}}{n}\right) \\ =&e^{n-1}\left(\bar{u} + \bar{\pi}\right) > 0.\end{aligned}$$

To assess whether a higher k also makes it "less likely" that the (EC) binds, we again state the critical discount factor above which e^{FB} can be implemented,

$$\bar{\delta} \equiv \frac{c(e^{FB}, k)}{\theta e^{FB} - \bar{u} - \bar{\pi}},$$

with

$$\begin{aligned}\frac{d\bar{\delta}}{dk} =& \frac{c_k}{\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right)} + \frac{c_e\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right) - c(e^{FB}, k)\theta}{\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right)^2} \frac{de^{FB}}{dk} \\ =& \frac{c_k}{\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right)} - c_{ek} c_e \frac{\left(c_e e^{FB} - c(e^{FB}, k)\right)}{c_{ee}\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right)^2} \\ &+ \frac{c_{ek} c_e}{c_{ee}} \frac{\left(\bar{u} + \bar{\pi}\right)}{\left(\theta e^{FB} - \bar{u} - \bar{\pi}\right)^2},\end{aligned}$$

which clearly is positive for $c_{ek} = 0$. For $c_{ek} > 0$, we again assume the standard effort cost function

$c(e,k) = k\frac{e^n}{n}$, with $n \geq 2$, for which this term becomes (for any e)

$$
\begin{aligned}
&\frac{\frac{e^n}{n}}{(\theta e - \bar{u} - \bar{\pi})}\left(1 - \frac{ke^n}{(\theta e - \bar{u} - \bar{\pi})}\right) + \frac{c_{ek}c_e}{c_{ee}}\frac{(\bar{u} + \bar{\pi})}{(\theta e - \bar{u} - \bar{\pi})^2} \\
=&- \frac{e^n}{n}\frac{(\bar{u} + \bar{\pi})}{(\theta e - \bar{u} - \bar{\pi})^2} + \frac{e^n}{(n-1)}\frac{(\bar{u} + \bar{\pi})}{(\theta e - \bar{u} - \bar{\pi})^2} \\
=&\frac{e^n}{n(n-1)}\frac{(\bar{u} + \bar{\pi})}{(\theta e - \bar{u} - \bar{\pi})^2} > 0.
\end{aligned}
$$

■

C. Alternative identification strategy

As a validation exercise, we use changes in the Austrian early retirement age (ERA) as an alternative source of variation in outside options. In this appendix we describe the institutional setting and the two pension reforms we consider in more detail. We draw heavily from Staubli & Zweimüller (2013) and Manoli & Weber (2016), who study employment effects and actual pension takeup of the ERA reforms.

C.1. The pension system

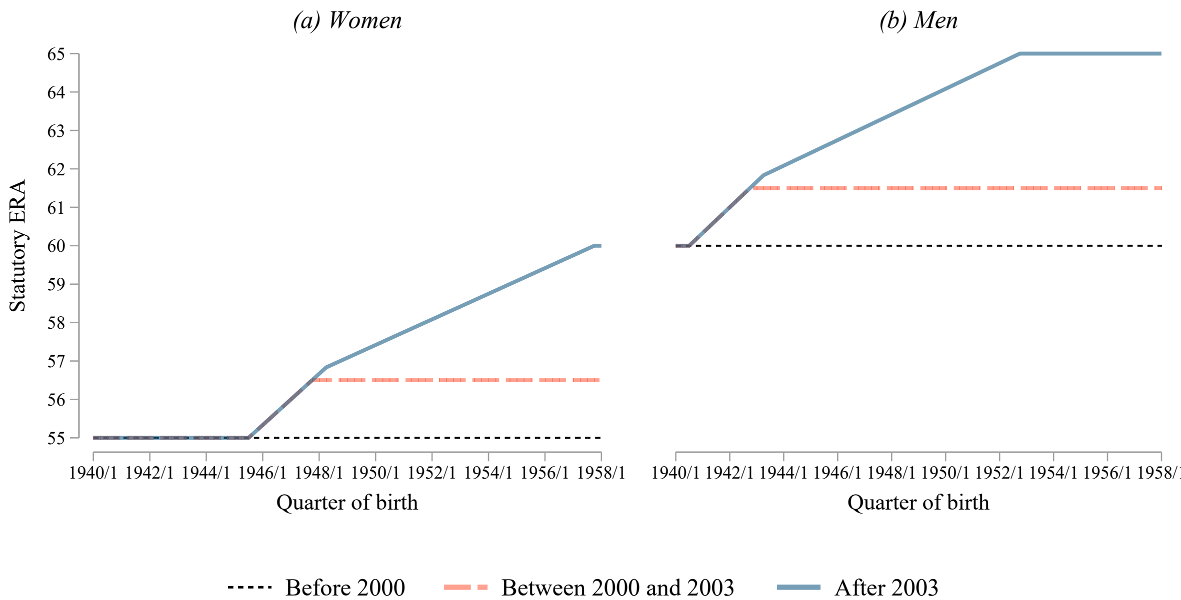
The Austrian public pension system is universal and private-sector workers are automatically enrolled. Financing is based on a pay-as-you-go scheme, with payroll taxes that are withheld from the worker's salary up to a contribution cap. The system covers three types of pension: regular old age pension, early retirement, and disability pension. Currently, the statutory retirement age is 65 years for men and 60 years for women. Until 2000, early retirement was possible from age 60 for men and age 55 for women, provided they had worked for at least 35 years or if they had been long-term unemployed but had worked for at least 15 years in total. Pension benefits are assessed based on pre-retirement wages multiplied with a 'pension coefficient,' which is a factor of up to 0.8 that increases with work experience. The penalty for retiring early amounts to around 2 percentage points per year of retiring before the regular retirement age.

C.2. The pension reforms

For the purpose of fiscal consolidation, the Austrian government enacted two pension reforms in 2000 and 2003 that increased the ERA gradually by quarter of birth cohort. Figure C.1 provides a graphical representation. The 2000 reform, which became effective on October 1, 2000, increased the ERA by 1.5 years in two-month steps for every quarter of birth for both men and women. The first affected cohort for men was 1940/q4, where the ERA was increased to 60 years and 2 months, and 1945/q4 for women, where the ERA was increased to 55 years and 2 months. The last affected cohorts were 1942/q4 (men) and 1947/q4 (women), for whom the ERA was increased to 61.5 (men) and 56.5 (women). The statutory ERAs by cohort after 2000 are represented by the red dashed line in Figure C.1.

The second reform in 2003, which became effective on January 1, 2004, effectively abolished early retirement by increasing the ERA to the regular old age pension age (65 for men and 60 for women). Again, the ERA increase was phased-in based on quarter of birth cohorts. For men, the ERA was raised in two-month steps for those born in 1943/q1 and 1943/q2 and then in one-month steps for those born between 1943/q3 and 1952/q4. Similarly, for women, the ERA was increased

Figure C.1 — Statutory ERAs over time, by quarter of birth, and by gender



Notes: This figure depicts the statutory ERAs by quarter of birth and for both genders. The black dashed line represents the initial situation where the ERA was 55 for female workers and 60 for male workers. The red dashed line represents the situation between after the first reform in 2000 and before the second reform became effective in 2004. The blue line represents the current ERAs that are in effect since January 1, 2004.

in two-month steps for those born in 1948/q1 and 1948/q2 and in one-month steps for those born between 1948/q3 and 1957/q4. This is depicted as the blue line in Figure C.1.

D. Theoretical Robustness

Our predictions are based on a model of a relational contract which we assume is optimally designed for an individual agent. Here, we argue that the specific interpretation of the model is not important for our results. In fact, it only matters that the future relationship surplus determines today's actions.

D.1. Fixed wages and a standard efficiency wage model

Although a worker's compensation contains many components besides monetary payments, one might argue that our model allows for too much flexibility in determining individual compensation when our objective is to capture the situation in Austria. As described above, the Austrian labor market is characterized by centrally bargained wages and working conditions, and on top by weak job protection. In the following, we therefore show that our results do not rely on the principal's ability to tailor compensation systems to individual workers, but can also be generalized in a more constrained setting. Suppose wages are exogenously given and incentives are solely provided by firing threats upon non-performance. We thus rule out the use of an informal performance-based bonus. Such a setting resembles classic models of efficiency wages which generally are relational contracting models with restrictions on the forms of compensation (see MacLeod & Malcomson 2023). With a given compensation, the only individualized aspect of the employment relationship is the agent's effort. Thus, assume that the agent is supposed to exert effort e_t^*. If he complies, he remains employed, otherwise he is fired at the end of the period. We focus on an equilibrium in which the agent remains employed on the path of play—which implicitly requires the wage to be high enough to satisfy the agent's participation constraint and low enough to satisfy the principal's participation constraint—hence his utility in a period t is

$$U_t = w - c(e_t^*) + \delta U_{t+1}.$$

Equilibrium effort is constrained by his (IC) constraint,

$$-c(e_t^*) + \delta U_{t+1} \geq \delta \bar{U}_{t+1}, \tag{IC}$$

where $\bar{U}_{t+1}$ is defined as in our main model. Now, equilibrium effort and comparative statics are not determined by the total future surplus, but by the agent's continuation payoff. It is immediate that $\delta\left(U_{t+1} - \bar{U}_{t+1}\right)$ increases in w and δ (given $U_\tau - \bar{U}_\tau > 0 \,\forall\, \tau$) and decreases in $\bar{U}_{t+1}$. If this continuation rent is large enough, $e_t = e^{FB}$, otherwise, e_t^* is determined by the binding (IC). Equivalently to the proof to Proposition 1, we can show that e_t^* increases in $\delta\left(U_{t+1} - \bar{U}_{t+1}\right)$ if (IC) binds and otherwise does not respond to it. Moreover, the effect of a higher continuation rent on effort is more pronounced if this rent has initially been smaller. If we also suppose that

$\delta\left(U_{t+1} - \bar{U}_{t+1}\right)$ goes down as time passes, this setting can as well generate Predictions 1–5. If a higher inherent productivity θ and consequently a higher relationship rent also increases w_t, we can also generate Prediction 6. Therefore, treating wages as exogenously given and providing incentives only via firing threats does not change our predictions. The reason is that the main mechanism, that workers are motivated by future rents from employment, still drives the results.

D.2. Effort is private information

In this subsection, we demonstrate that our results do not rely on the principal being able to observe the agent's effort. Suppose effort is the agent's private information, and the principal can observe a non-verifiable output measure $y_t = \{0, \theta\}$, where $y_t = \theta$ with probability e_t and $y_t = 0$ with probability $1 - e_t$. Moreover, we assume effort is between 0 and 1 and θ sufficiently small to always guarantee an interior solution. Now, the agent is incentivized if his payoffs are larger after a high than after a low output. As in our main model, such a reward can either take the form of a bonus paid at the end of a period or higher payments in the future. Here, we assume that only future payoffs are used for that purpose, thus no bonuses are paid but only wages. This assumption is without loss of generality because any bonus paid after a high output provides the same incentives as an equivalent increase of next period's wage (multiplied by δ). We do so to not have to deal with potentially negative bonuses which we would otherwise have to consider after a low output (and consequently a (DE) constraint for the agent). In the following we use the superscript "+" to indicate continuation payoffs after a success, and "–" for continuation payoffs after a failure. Therefore,

$$\begin{aligned} U_t =& w_t + e_t \delta U^+_{t+1} + (1 - e_t)\delta U^-_{t+1} - c(e_t) \\ \Pi_t =& e_t\left(\theta + \delta \Pi^+_{t+1}\right) + (1 - e_t)\delta \Pi^-_{t+1} - w_t. \end{aligned}$$

Now, the following constraints must be satisfied in a self-enforcing relational contract:

$$-c'(e_t) + \delta\left(U^+_{t+1} - U^-_{t+1}\right) = 0 \tag{IC}$$

$$\Pi^+_{t+1}, \Pi^-_{t+1} \geq \bar{\Pi} \tag{PCP}$$

$$U^+_{t+1}, U^-_{t+1} \geq \bar{U}_{t+1}, \tag{PCA}$$

where (PCP) and (PCA) indicate the principal's and the agent's participation constraints, respectively, and effort in (IC) is determined by the agent's first-order condition.

As Levin (2003) has demonstrated, we can again separate the provision of incentives from the allocation of the resulting surplus. Therefore, we once more focus on maximizing the relationship surplus at the beginning of the relationship. Moreover, this problem is sequentially efficient, hence maximizing the initial surplus is equivalent to maximizing the surplus in every period t. This also implies that no firing threats are used to provide incentives, and equilibrium effort and incentives in a period are independent of whether a success or failure has been previously observed. Finally, it is without loss to let the period-t wage constitute the only difference between U_t^+ and U_t^-, and Π_t^+ and Π_t^-. Thus, $U_t^+ > U_t^-$ and $\Pi_t^+ < \Pi_t^-$, as well as $U_t^+ + \Pi_t^+ = U_t^- + \Pi_t^- \equiv S_t$, and the problem becomes to maximize S_t in every period t, subject to

$$-c'(e_t) + \delta\left(U_{t+1}^+ - U_{t+1}^-\right) = 0 \tag{IC}$$

$$\Pi_{t+1}^+ \geq \bar{\Pi} \tag{PCP}$$

$$U_{t+1}^- \geq \bar{U}_{t+1}. \tag{PCA}$$

Multiplying both sides of (PCP) and (PCA) with δ and adding all constraints yields the enforceability constraint

$$-c'(e_t) + \delta\left(S_{t+1} - \bar{S}_{t+1}\right) \geq 0, \tag{EC}$$

where $\bar{S}_{t+1} = \bar{\Pi} + \bar{U}_{t+1}$.

Levin (2003) shows that, as long as each player at least gets their outside option, this constraint is necessary and sufficient for obtaining equilibrium effort e_t. Therefore, $e_t = e^{FB}$ if it satisfies the condition, otherwise e_t is determined by the binding (EC) constraint. In the latter case, e_t increases in the future net surplus $\delta\left(S_{t+1} - \bar{S}_{t+1}\right)$, thus in δ and θ, and decreases in $\bar{S}_{t+1}$. Hence, Predictions 1 and 3 would also be generated in a setting with private effort. For our other predictions, it can immediately be shown that the positive effect of the future net surplus is more pronounced for an initially smaller surplus if and only if $c''' > 0$.

The same holds if, as in Section D.1, the principal cannot tailor compensation to an individual employment relationship and is not able to use performance-based compensation. Then, firing threats are the only means to provide incentives. Here, we assume that the principal fires the agent after a low output with some probability $1 - \alpha \in [0, 1]$. For simplicity, we take α as given; if α was set to maximize profits, its level would not be stationary and potentially depend on the whole history of the game (Fong & Li 2017). Such an analysis is beyond the scope of this paper, though, and would not affect our results qualitatively. The agent's utility in such a setup is

$$U_t = w_t - c(e_t) + \delta\left[e_t U_{t+1} + (1 - e_t)\left(\alpha U_{t+1} + (1-\alpha)\bar{U}_{t+1}\right)\right], \tag{9}$$

hence effort is characterized by

$$c'(e_t) = \delta(1-\alpha)\left(U_{t+1} - \bar{U}_{t+1}\right).$$

It follows that, for $c''' > 0$, comparative statics are as in Section D.1.